Big Bad Banks?

The Impact of U.S. Branch Deregulation on Income Distribution

Thorsten Beck, Ross Levine, and Alexey Levkov

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Abstract: Policymakers and economists disagree about the impact of bank regulations on the distribution of income. Exploiting cross-state and cross-time variation, we test whether liberalizing restrictions on intra-state branching in the United States intensified, ameliorated, or had no effect on income distribution. We find that branch deregulation lowered income inequality. Deregulation lowered income inequality by affecting labor market conditions, not by boosting the business income of the poor, nor by enhancing educational attainment. Reductions in the earnings gap between men and women and between skilled and unskilled workers account for the bulk of the explained drop in income inequality.

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Beck: World Bank, tbeck@worldbank.org; Levine: Brown University and the NBER, ross_levine@brown.edu; Levkov: Brown University, alexey_levkov@brown.edu. Martin Goetz and Carlos Espina provided exceptional research assistance. We especially thank Phil Strahan for providing extensive comments and sharing his data and Yona Rubinstein, whose suggestions substantively improved the quality of our research. We received very helpful comments at the World Bank conference, "Access to Finance," the Brown University conference, "The Causes and Consequences of Income Inequality," and seminars at the International Monetary Fund, European Central Bank, Georgia State University and Mannheim University. The many remaining weaknesses are our responsibility. This paper's findings, interpretations, and conclusions are entirely those of the authors and do not necessarily represent the views of the World Bank, its Executive Directors, or the countries they represent.

Throughout the history of the United States, income distributional considerations have shaped policies toward banks (Hammond, 1957). Besides constitutional concerns, Thomas Jefferson's fears that concentrated banking power would help only the wealthy spurred him to fight against the Bank of the United States. Similar anxieties fueled the termination of this bank in 1811 and Andrew Jackson's veto of the re-chartering of the Second Bank of the Unites States in 1832.

The view that the unconstrained expansion of powerful banks hurts the poor finds support in modern economic theory and continues to motivate bank regulations around the world. If banking is a natural monopoly, then unregulated, monopolistic banks may earn rents through high fixed fees that disproportionately curtail the economic opportunities of the poor. Based on this argument, politicians in many U.S. states implemented and maintained regulatory restrictions on bank branching for much of the 20^{th} century (Southworth, 1928; White, 1982). Furthermore, most countries regulate bank mergers and acquisitions, with the advertised goals of enhancing competition, reducing borrowing costs, and expanding access to bank loans (Barth et al, 2006).

Countervailing arguments, however, challenge the view that regulations on bank expansion help the poor. These regulations could hinder competition and thereby raise lending fees and hurt the poor. Indeed, Flannery (1984) shows that U.S. branching restrictions prevented banks from competing in distant areas. This created and protected local banking monopolies, allowing them to maintain high lending rates and fees (Jayaratne and Strahan, 1998). Furthermore, from a political economy perspective, Acemoglu (2007), Haber (2007), Kroszner and Strahan (1999), and White (1982) argue that governments frequently enact entry barriers to protect favored groups, not to promote aggregate efficiency and economic growth. From this perspective, income distributional effects frequently play a leading role in determining bank regulations.

In this paper, we provide the first assessment of the impact of bank branching regulations on the distribution of income. From the 1970s through the 1990s, most states in the United States removed restrictions on intra-state branching. We test whether liberalizing these restrictions intensified, ameliorated, or had no effect on income inequality. We also evaluate different theories of the particular channels linking deregulation and income distribution.

Our examination builds directly on past work showing that liberalizing restrictions on intra-state branching (i) increased the average size of banks through consolidation (Calem, 1994; Savage, 1993) and (ii) lowered the borrowing costs paid by firms (Jayaratne and Strahan, 1998). We do not re-examine the empirical chain running from deregulation to intensified bank competition and lower borrowing costs. Rather, we assess the impact of intra-state branch deregulation on the distribution of income, which has been a central -- if not the central -- front in the battle over bank regulations.

Methodologically, the deregulation of intra-state branching provides a natural setting for identifying and assessing the impact of regulatory reform on the distribution of income. Kroszner and Strahan (1999) show that national technological innovations triggered deregulation, which was exogenous to income distributional changes within individual states. Specifically, (1) the invention of automatic teller machines (ATMs), in conjunction with court rulings that ATMs are not bank branches, weakened the geographical bond between customers and banks; (2) checkable money market mutual funds facilitated banking by mail and telephone, which weakened local bank monopolies; and, (3) improvements in communications technology lowered the costs of using distant banks. These innovations reduced the monopoly power of local banks, and therefore weakened their ability and desire to fight deregulation. Kroszner and Strahan (1999) further show that cross-state variation in the timing of deregulation reflects the interactions of these technological innovations with preexisting conditions. For example, deregulation occurred later in states where politically powerful groups viewed large, multiple-branch banks as potential competitors. Thus, the driving forces behind deregulation and its timing were largely independent of state-level changes in income distribution. Consequently, we exploit cross-state, cross-year variation in income distribution and deregulation to assess the impact of a single policy change on different state economies.

The paper's major finding is that deregulation of branching restrictions reduced income inequality. This finding is robust to using different measures of income inequality, examining different components of income, controlling for time-varying state characteristics, and conditioning on state and year fixed effects. Furthermore, the impact of deregulation on income distribution varies across states in ways that are fully consistent with Kroszner and Strahan's (1999) political economy assessment of branch deregulation. Moreover, we find no evidence that reverse causality affects the results. While income inequality widened in the U.S. during the sample period, we show that branch deregulation lowered income inequality relative to this national trend. The magnitude is consequential. Deregulation explains about 60% of the de-trended variation of income inequality relative to state and year averages. Moreover, deregulation reduced income inequality by exerting a disproportionately positive impact on the poor, not by hurting the rich.

The two major theoretical explanations for this finding stress the enhanced ability of the poor to access banking services following deregulation. The first explanation focuses on the funding of human capital accumulation. In Galor and Zeira (1993), for example, the high cost of acquiring human capital together with capital market imperfections prevent the poor from borrowing to fund education, which intensifies the inter-generational persistence of relative income differences. Deregulation that reduces credit market imperfections, therefore, permits more poor individuals to finance their education, reducing income inequality. A second explanation focuses on the ability of the poor to become entrepreneurs (Banerjee and Newman, 1993). Financial imperfections are particularly binding on the poor because they lack collateral and because their incomes are relatively low compared to the fixed costs of obtaining a bank loan. Thus, branch deregulation that improves bank efficiency by lowering collateral requirements and borrowing costs will disproportionately benefit the poor by expanding their access to bank credit.

We find, however, that neither of these two theories directly accounts for much of the reduction in income inequality explained by bank deregulation. First, branch deregulation has no effect on high school or college educational attainment. Second, the impact of branch deregulation levels-off quickly. Relative to state and year fixed effects, the Gini coefficient of income inequality falls for six years after deregulation and then stops falling. Deregulation has a level effect that fully materializes in six years, not a trend effect. The time pattern seems inconsistent with the education explanation, which implies a growing effect on income distribution as the poor accumulate education. Third, the impact of deregulation on proprietor income accounts for only 7% of the explained reduction in income inequality. Although Black and Strahan (2002) show that branch deregulation spurred the entry of new firms, changes in entrepreneurial income account for exceptionally little of branch deregulation's impact on income distribution.

Rather, we show that a tightening of the distribution of wage earnings accounts for almost 70% of the explained reduction in total income inequality, suggesting that branch deregulation reduces inequality primarily by affecting labor market conditions. In particular, deregulation has differential effects on labor by skill and gender. First, deregulation increases the wage income of unskilled workers relative to skilled workers. Since unskilled workers tend to earn less than skilled workers, this narrowing of the earnings gap reduces overall income inequality. Second, deregulation reduces the wage income gap between men and women, pushing up women's wage income relative to men's income. Since, on average, women earn less than men, deregulation helped reduce total income inequality by reducing the gender income gap.

This paper relates to an enormous literature on the determinants and consequences of the distribution of income. First, the international policy community increasingly emphasizes the benefits of providing the poor with greater access to financial services as a vehicle for fighting poverty and reducing income inequality. While the experience of branch deregulation within U.S. states may not generalize perfectly to developing economies, the results do raise the possibility that financial development helps the poor primarily by boosting wage income, not by increasing the business income of the poor. This warrants further research. Second, many authors examine income redistribution as a mechanism for reducing the inefficient propagation of relative incomes across generations (Aghion, Caroli, and Garcia-Penalosa, 1999; Galor and Moav, 2006). Unlike financial liberalization, however, redistribution has adverse incentive effects. Thus, our work contributes to the policy debate on how to address socially inefficient income inequality. Third, we contribute to recent cross-country analyses of finance and the distribution of income. Beck, Demirguc-Kunt and Levine (2007) find that banking sector development reduces income inequality and poverty. In contrast, we analyze the impact of a specific, exogenous policy change on the distribution of income, rather than examining an overall index of financial sector development. By using the differences-in-differences approach across U.S. states, a consistent source of income inequality data, and a specific policy event, we increase the power of the econometric tests, reduce potential biases due to measurement error, and reduce concerns about omitted variables and endogeneity.

This paper also relates to a substantive body of work on the effects of branch deregulation. Researchers examine the impact of intra-state branch deregulation on economic growth (Jayaratne and Strahan, 1996; Huang, 2007), the entry of non-financial companies (Black and Strahan, 2002), and income volatility (Demyanyk, Ostergaard, and Sorensen, 2007; Acharya, Imbs and Sturgess, 2007). Focusing on rent-sharing and discrimination, Black and Strahan (2001) show that deregulation reduced the wage rate gap between men and women bank executives. We look beyond the banking industry and ask whether bank deregulation affected the overall distribution of income distribution, as well as the gender and skill gap throughout the entire economy. Demyanyk (2007) finds a positive impact of branch deregulation on income growth of proprietors, with the impact stronger for women and minorities than for men, while Jerzmanowski and Nabar (2007) argue that bank deregulation increases the skill premium. In this paper, we show that (1) deregulation reduces overall income inequality, inequality among proprietors, and inequality among wage earners, (2) deregulation reduces overall income inequality primarily by tightening the distribution of wage income, while the reduction in income inequality among proprietors accounts for very little of the reduction in overall income inequality and (3) deregulation reduces income inequality primarily by reducing the income gap between men and women and between skilled and unskilled workers.

The remainder of the paper proceeds as follows. Section 1 describes the data and econometric methodology. Section 2 provides the core results on the impact of deregulation on the distribution of income, while Section 3 provides further evidence on how deregulation influences labor market conditions. Section 4 concludes.

1. Data and Methodology

To assess the effect of branching deregulation on income distribution, we gather data on the timing of deregulation, income distribution, and other banking sector and state-level characteristics. This section presents the data and describes the econometric methodologies.

1.1. Branch deregulation

Historically, most U.S. states had restrictions on branching within and across state borders. Beginning in the early 1970s, however, many states started relaxing these restrictions, allowing bank holding companies to consolidate subsidiaries into branches and permitting de novo branching throughout the state. This deregulation led to significant entry into local banking markets (Amel and Liang, 1992), consolidation of smaller banks into large bank holding companies (Calem, 1994) and conversion of existing bank subsidiaries into branches (McLaughin, 1995). This relaxation, however, came gradually, with the last state lifting restrictions following the 1994 passage of the Riegle-Neal Interstate Banking and Branching Efficiency Act. An extensive literature has assessed the impact of this gradual branch deregulation on economic growth, entrepreneurial activity and other banking sector and real economy outcomes. The intra-state branching deregulation was often accompanied by inter-state branching deregulation, which allowed banking holding companies to expand across state borders. The literature, however, has found little effect of inter-state branching deregulation on banking market structure or real outcomes. Similarly, we find no relation between inter-state branch deregulation and the distribution of income. Hence, we focus on intra-state branch deregulation.

Consistent with Amel (1993), Jayaratne and Strahan (1996), and others, we choose the date of deregulation as the date on which a state permitted branching via mergers and acquisitions (M&As) through the holding company structure. This was typically the first step in the deregulation process, followed by de novo branching. Table A1 presents the deregulation dates.[1] Twelve states deregulated before the start of our sample period in 1976. Arkansas, Iowa and Minnesota were the last states to deregulate, only after the passage of the Riegle-Neal Act in 1994.

1.2. Income distribution data

Information on the distribution of income is from the March Supplement of the Current Population Survey (CPS), which is a survey of about 60,000 households across the states of the U.S. The CPS provides information on total personal income, income from wages, proprietor income, as well as detailed demographic information. The CPS is not a true panel; it is cross-sectional survey that is repeated each year. The CPS does not trace individuals through time, but rather replaces 50% of the sample each year, so that each household is included for two consecutive years. Each individual is assigned a sample weight corresponding to his or her representativeness in the sample, which we use in our analyses.

We measure the distribution of income for each state and year over the period 1976-2005 in four ways. First, the natural logarithm of the Gini coefficient of income distribution is derived from the Lorenz curve (Gini), where larger values imply greater income inequality. Our second measure of income distribution is the coefficient of variation of income, i.e. the variance of income divided by the mean (Variance). Third, we use the difference between the natural logarithm of incomes of those at the 90th percentile and those at the 10th percentile (90/10). Finally, we use the difference between the natural logarithm of incomes of those at the 75th percentile and those at the 25th percentile (75/25).

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[1] We have data for 50 states and the District of Columbia. Consistent with the literature on branch deregulation, we drop Delaware and South Dakota because the structure of their banking systems were heavily affected by laws that made them centers for the credit card industry, and because South Dakota lacks some financial and banking data.

We start our analysis using income data for 1976 because this is when the CPS begins identifying individuals by state. Consistent with studies of the U.S. labor market, our main sample (a) includes prime-age (25-54) black and white civilians, (b) excludes people living in group quarters, those with missing observations on demographic characteristics, and a few individuals for which the CPS assigns a zero sampling weight, and (c) corrects for top-coding and translates income to constant 1990 dollars.[2] Table 1, Panel A provides details on the construction of the sample, while Panel B presents summary statistics on the individuals in the sample.

Table 1 Panel C presents descriptive statistics on the logarithm of the Gini coefficient of income inequality. These data are calculated on a state-year basis, so there are 1,433 observations. For total individual income, the average Gini across states and over time is 0.48, corresponding to an average of -0.74 for the log of Gini. The standard deviation of the log of Gini within states over years is 0.56.

1.3. Control variables

To control for time-varying changes in a state's economy, we use the U.S. Department of Commerce data to calculate the growth rate of per capita Gross State Product (GSP). We also control for the unemployment rate, obtained from the Bureau of Labor. We also condition on government taxes to personal income and government expenditures to personal income in order to account for governmental redistribution policies. These data were obtained from the U.S. Census Bureau. In further robustness tests, we test whether the impact of deregulation on income inequality varies in a predictable way with different state characteristics at the time of

[2] Following the usual practice in analyses of income distribution using the CPS data, we inflate top-coded incomes by a factor of 1.5. For purposes of confidentiality the CPS does not report exact incomes above a certain threshold. Instead, these incomes are grouped into a single category called the "top-code". The top-code changes across years and types of income. Furthermore, consistent with usual practices, we drop all allocated incomes, i.e., incomes that were originally missing but were assigned a non-missing value by the CPS based on demographic characteristics of the respondents.

deregulation. As we discuss below, we control for the interaction of branch deregulation with a unit banking indicator, the small bank share, the small firm share and population dispersion, each of which we measure in the year before deregulation. Data on the share of small firms and banks are from Kroszner and Strahan (1999). Data on population dispersion are from the U.S. Census Bureau.

1.4. Methodology

We use a differences-in-differences specification to assess the relationship between branch deregulation and income distribution, based on the following regression set-up:

$$Y_{s,t} = \alpha_s + \beta_t + \gamma D_{s,t} + \delta X_{s,t} + \varepsilon_{s,t}, \qquad s=1,\ldots,49; \quad t=1976,\ldots,2005 \qquad (1)$$

where $Y_{s,t}$ is a measure of income distribution in state s in year t, α and β are vectors of state and year fixed-effects, $X_{s,t}$ is a set of time-varying, state-level variables and ε is the error term. The variable of interest is $D_{s,t}$, a dummy variable that takes on the value one the year after state s deregulates. The coefficient, γ, therefore indicates the impact of branch deregulation on income distribution. A positive and significant γ suggests that deregulation exerts a positive effect on the degree of income inequality, while a negative and significant γ indicates that deregulation pushed income inequality lower. Consistent with past work on branch deregulation, we drop the year of deregulation. In total, we have data for 48 states plus the District of Columbia, over 30 years, minus 37 state-years in which deregulation occurred. Thus, 1,433 state-year observations serve as the basis for our analysis.

The differences-in differences estimation technique allows us to control for omitted variables. We include time-specific dummy variables to control for economy-wide shocks that might drive income distribution over time, such as business cycles, long-term trends in income distribution, and changes in female labor force participation across the country. We include state-specific dummy variables to control for time-invariant, unobserved state characteristics that shape income distribution across states. The coefficient γ therefore measures the effect of branching deregulation relative to the average income distribution over time in state i and relative to average income distribution across all states in year t. We estimate regression (1) allowing for state-level clustering, i.e. allowing for correlations in the error terms over time within states.[3]

2. Branch deregulation and income distribution

2.1. Main results

The Table 2 results show a strong negative and significant relationship between branching deregulation and the Gini coefficient. The branching deregulation dummy enters negatively and significantly at the 5% level in all seven regressions. The column 1 result suggests that relative to the state- and year-specific averages, branching deregulation results in a 2.6 percentage point drop in the Gini coefficient. To gauge the economic effect of this result, we compare the coefficient estimate to the de-trended standard deviation of Gini, which is computed after accounting for state and year effects. This standard deviation is 0.041 (Table 1), suggesting that branching deregulation explains about 60% of the variation of log Gini relative to state and year averages. In contrast to the influential view that regulatory restrictions on bank branching protected the poor from the adverse effects of large banks, we find that deregulation reduced income inequality.

[3]Consistent with Bertrand, Duflo, and Mullainathan (2004), we also confirm our results when reducing the time-series dimension of the data as follows: (i) estimate a version of equation (1) that includes only state and year dummy variables, (ii) compute the residuals, $e_{s,t,}$, (iii) average over the pre-deregulation years to obtain $e_{s,pre}$ and average over the post-deregulation years to obtain $e_{s,post}$, and (iv) for only those states that experienced deregulation during the sample period, run a pooled regression of $e_{s,post}$ and $e_{s,pre}$ on a constant and a dummy variable, d_s, that equals one for $e_{s,post}$ and zero for $e_{s,pre}$. We find that d_s enters negatively and significantly, indicating that deregulation reduces income inequality.

Regressions 2 through 7 show the robustness of our findings to controlling for other potential time-varying state-level factors associated with income inequality. The column 2 regressions shows that there is no significant relationship between income inequality and the growth rate of per capita Gross State Product (GSP), while the column 3 results indicate that a higher unemployment rate is associated with higher income inequality. This does not imply that per capita economic growth is unrelated to income inequality; rather, it suggests that state-level per capita growth is not significantly associated with deviations of state-level income inequality from the state and year-average Gini coefficient. The column 4 regression shows that higher tax burden on income is associated with higher income inequality, while the column 5 regression suggests that higher government spending is associated with lower income inequality. Furthermore, jointly controlling for per capita income growth, the unemployment rate, government taxation and spending as reported in column 6 does not change the finding that deregulation reduces income inequality. The deregulation dummy continues to enter negatively and significantly and with a similar coefficient. In column 7, we add the lagged growth rate of Gini. The deregulation dummy variable continues to enter negatively and significantly. In unreported regressions, we further show that these results also hold when controlling for the size of each state's aggregate economy, the growth rate of each state's aggregate economy, the level of real per capita income in each state, or lagged values of each state's Gini coefficient. Further, this paper's results hold when examining family income, rather than focusing on the income of individuals.

As a robustness test, and to clarify the timing of the impact of deregulation on the distribution of income, we examine the dynamics of the relationship between deregulation and the distribution of income. We do this by including a series of dummy variables in the standard regression. In particular, D_x_j equals one for all observations in state j that are x years before deregulation, while Dx_j equals one for the year x year after deregulation in state j. Figure 1 plots the results and the 95% confidence intervals, centering the estimates around year 0, the year of deregulation.

The results indicate that innovations in the distribution of income did not precede deregulation and the impact of deregulation on inequality materializes over the six years after deregulation. As shown, D_x_j is insignificantly different from zero for all years before deregulation. This suggests that changes in the distribution of income did not precede deregulation. Next, note that inequality falls after deregulation, as D1 is negative. The impact of deregulation on inequality grows and becomes significantly negative three years after deregulation. After three years, the impact continues to grow slightly. The full effect of deregulation is observed within six years. Deregulation has a level effect on the distribution of income, but no trend effect on the Gini coefficient of income inequality.

2.2. Alternative categories of income and different measures of income distribution

The negative impact of deregulation on income inequality holds when using different categories of income. While Table 2 focused on the Gini coefficient computed using total income, Table 3 provides regression results using the Gini coefficient based on wage and salary income and on proprietor income.[4] As shown in the first column of Table 3, deregulation induces a significant drop in inequality using each of these three categories of income.

The negative relation between removing restrictions on intra-state bank branching and income inequality is quite robust to using different measures of income distribution. Beside Gini, we present results with (i) the log ratio of the 90th and the 10th percentiles of the income

[4] Besides proprietor income and wage and salary income, total income also includes government transfer payments as well as income from interest, dividends, and rent.

distribution (90/10), (ii) the log ratio of the 75th and the 25th percentiles of the income distribution (75/25), and (iii) the log coefficient of variation (Variance). Since we report the results for each category of income, Table 3 presents the findings from twelve separate regressions while controlling for the unemployment rate, as well as state and year fixed effects. While the negative impact of deregulation on income inequality is slightly weaker for the two measures of income inequality measures that put greater weight on the tails of the distribution of income (90/10 and Variance), the results are very robust when using the Gini and 75/25 measures of inequality.

The negative relation between deregulation and inequality is significant at the five percent level for all of the categories of income when using the Gini and 75/25 measures of inequality. Moreover, the 75/25 measure of inequality drops because the income the poorest quarter of the population increases, not because the richest 25% become poorer. In particular, the incomes of those at the 25th percentile of the distribution jump 13% in real terms after deregulation (significant at the 5% level), while the incomes of those at the 75th percentile do not change significantly. In sum, when not putting too much weight on outliers, deregulation is associated with a tightening of the distribution of income when using different categories of income and alternative indexes of inequality.

2.3. The impact of deregulation as a function of initial conditions

We next assess whether the impact of deregulation on the distribution of income varies in predictable ways across states with different initial conditions. If the impact of deregulation on income distribution varies in a theoretically predictable manner, this provides greater confidence in the conclusions, sheds empirical light on the channels through which deregulation influences income distribution, and also reduces concerns about reverse causality.

Following Kroszner and Strahan (1999), we expect the impact of deregulation to be stronger in states where branch regulation had a more distorting effect on bank behavior prior to deregulation. We consider four initial conditions that reflect the distorting effects of branch regulation on the economy. First, unit banking was the most extreme form of branching restriction. Thus, we expect that deregulation exerted a particularly large, negative impact on income inequality in states that had unit banks before they deregulated. Second, one characteristic that signals the distorting effect of branching is the degree to which small banks compose a relatively large proportion of total state banking assets prior to deregulation. Thus, we expect that deregulation had an especially large impact on inequality in states with a comparatively high ratio of small banks at the time of deregulation. Third, small firms tend to have a harder time than large firms obtaining credit from banks that are farther away (Berger et al., 1998). This suggests that local banking monopolies protected by branch regulations were particularly strong in states dominated by small firms. Thus, we expect that branching deregulation had a bigger impact in states with a large proportion of small firms prior to deregulation. Finally, we examine the dispersion of the population. Local banking monopolies will be particularly well-protected if the population is diffuse, so that other banks are more likely to be far away. This suggests that deregulation would have a bigger negative effect on inequality in states with a high initial geographic dispersion of people.

The results support each of these four predictions, as shown in Table 4. Specifically, branch deregulation reduces income inequality more in states that had (a) unit banking (column 2), (b) a higher share of small banks (column 3), (c) a larger share of small firms (column 4), and (d) a more dispersed population (column 5).[5] In the unit banking and population dispersion regressions, the interactions of unit banking with deregulation and population dispersion with deregulation enter negatively and significantly at the 5% level (columns 2 and 5), and the deregulation dummy variable enters negatively in both of these specifications. These findings suggest that the effect of deregulation on income distribution is stronger both for states that initially had unit banking and those with more dispersed populations. In the small banks and small firms regressions (columns 3 and 4), the interaction of the share of small banks and the deregulation dummy enters significantly and negatively at the five percent level. In these two regressions, the deregulation dummy variable by itself enters positively. Thus, for these two regressions, we also evaluate the overall effect of deregulation at different levels of the small bank share and small firm share, respectively. The results indicate that there was a negative and significant effect of deregulation on income inequality only in states with a small bank share or small firm share at or above the median, while the impact is insignificant in the other states. In sum, branch deregulation reduced income inequality more in states where regulatory restrictions on bank branching distorted banking behavior comparatively more before deregulation. These results are consistent with the view that deregulation weakened banking monopolies, intensified bank competition, and tightened the distribution of income.

[5] Following Kroszner and Strahan, we classify the following states as having unit banking before deregulation: Arizona, Colorado, Florida, Illinois, Iowa, Kansas, Minnesota, Missouri, Montana, Nebraska, North Dakota, Oklahoma, Texas, Wisconsin, West Virginia, and Wyoming,

2.4. Testing for reverse causality

To test directly for reverse causality, we use a hazard model that assesses whether changes in the distribution of income spurred branch deregulation. Following the procedure in Kroszner and Strahan (1999), Table 5 reports tests of whether income inequality influences the likelihood that a state deregulates in a specific year given that it has not already deregulated. We only include the 37 states that deregulated after 1976.[6] As in Kroszner and Strahan (1999), we find that states with a larger share of small banks and better capitalized small banks deregulate later, while states with a higher share of small firms deregulate earlier. States where banks are allowed to sell insurance products deregulate later if the insurance sector is relatively larger. Finally, states where the Democratic Party holds a larger share of political power deregulate later.

Consistent with our interpretation, the Gini coefficient does not enter significantly in any of the Table 5 regressions. The likelihood that a state deregulates does not change with higher or lower income inequality. This also confirms that our finding of a negative association between deregulation and income inequality is not driven by reverse causation.

[6] While the Kroszner and Strahan sample period starts in 1970, we do not have Gini data available before 1976. Also, since we use the original Kroszner and Strahan dataset, our sample period ends in 1994, when there were three states that had not deregulated yet – Arkansas, Iowa and Minnesota.

3. How Does Deregulation Affect the Distribution of Income?

We now empirically assess the two major theoretical explanations of how branch deregulation reduces income inequality. These theories are not mutually exclusive. They both rely on branch deregulation improving the ability of the poor to access banking services directly. One stresses that the poor use this improved access to purchase more education. The second emphasizes that the poor use the improved access to become entrepreneurs. Since we do not find much empirical support for either explanation, we end this section by providing a detailed decomposition of how deregulation affects the distribution of wage and salary income.

3.1. Assessing education theories

Galor and Zeira (1993) do not directly examine branch deregulation, but their theory of the co-evolution of economic growth and the distribution of income provides a unique framework for deriving predictions of how an improvement in the financial sector affects both aggregate growth and the distribution of income. In their model, the high cost of accumulating human capital together with financial market imperfections combine to prevent the poor from borrowing to fund education. This produces a socially inefficient allocation of schooling that both slows aggregate economic growth and perpetuates inefficiently high levels of income inequality. In the context of their theoretical framework, financial reforms that ease financial market imperfections will both accelerate growth and reduce inefficient income inequality by allowing talented, but poor, individuals to borrow and purchase education.

The empirical evidence suggests that human capital accumulation was not the channel through which branch deregulation helped reduce income inequality across states in the U.S. First, Figure 1 shows that the full effects of deregulation on the distribution of income are realized quickly. These dynamics seem inconsistent with the view that branch deregulation eases credit constraints, permitting the poor to accumulate more human capital. The education explanation suggests a growing effect of deregulation on income distribution as the poor accumulate skills. There is no evidence of a trend effect of deregulation on inequality.

More directly, Table 6 shows that branch deregulation did not result in higher educational attainment as predicted by theory. Specifically, we regress the share of the population with at least a high school degree and the share of population with at least a college degree on a set of state and time dummy variables and the deregulation dummy (columns 1 and 3). Though it would be useful to examine enrollment rates, such information is unavailable on a state-year basis. To assess the possible long-run effects of deregulation on education, we also use a specification in which we employ a set of deregulation dummies that take on the value one for the period (i) five to nine years after deregulation, (ii) 10 to 14 years after deregulation, (iii) 15-19 years after deregulation, (iv) 20 to 24 years after deregulation, (v) 25 to 29 years after deregulation, and (vi) more than 30 years after deregulation (column 2 and 4). Using a set of dummies for different time periods after deregulation allows us to test whether there was a medium- to long-term effect of deregulation on education. None of the dummies enters significantly. In sum, the analysis provides little support for the view that deregulation boosts education.

3.2. Assessing entrepreneurship theories

In Banerjee and Newman (1993), financial imperfections represent a particularly severe impediment to poor individuals seeking to become entrepreneurs. This might arise because the poor lack collateral or the fixed costs of borrowing are prohibitively high. In this context, financial imperfections (a) retard economic growth by hindering the efficient allocation of capital to poor, but talented, individuals and (b) increase inequality by limiting the opportunities of the poor. By improving credit markets, deregulation can spur growth and reduce inequality.

Our empirical findings suggest that enhanced entrepreneurship directly accounts for very little of the impact of deregulation on income distribution. In particular, we decompose the impact of deregulation on the distribution of income into that part accounted for by proprietor income and that part accounted for by non-proprietor income, which is wage income, income from interest and dividends, and transfer payments. We do this decomposition by using the variance of the distribution of income, rather than the Gini coefficient, so that we can employ standard variance decomposition techniques described in the Appendix. As shown in Figure 2, the impact of deregulation on proprietor income accounts for only 7% of the reduction in total income variance explained by branch deregulation. This does not necessarily suggest that entrepreneurship is unimportant for explaining the drop in income inequality. Rather, this findings implies that direct changes in entrepreneurial income account for very little of the reduction in income inequality.

3.3. Decomposing the impact of deregulation on wage income

In light of these results on education and proprietor income, we now provide more information on wage income. First, we simply decompose the drop in the variance of total income explained by branch deregulation into that part accounted for by the impact of deregulation on wage income and that part accounted for by the impact of deregulation on all other forms of income. Thus, we follow the same procedure outlined above for assessing the role of proprietor income. When focusing only on wage income, Figure 2 shows that the explained drop in wage income accounts for almost 70% of the explained reduction in the variance of total income.

Since changes in wage and salary income account for the bulk of branch deregulation's impact on the distribution of total income, we explore (1) the difference between skilled and unskilled workers and (2) the difference between male and female workers. An extensive literature documents the expanding skill premium in the U.S. labor market, where wage rates of skilled workers are rising faster than those of unskilled workers (e.g., Juhn, Murphy, and Pierce, 1993; Katz and Autor, 1999). Similarly, an active literature examines the gender gap (e.g., Mincer and Polacheck, 1974; Mulligan and Rubinstein, 2007a, b, c).

We decompose the explained tightening in the distribution of wage income into (i) the component associated with a tightening in the distribution of wage income *among* skilled workers, (ii) the component associated with a tightening in the distribution of wage income *among* unskilled workers, and (iii) the component associated with a tightening in the distribution wage income *between* skilled and unskilled workers. We define skilled workers as those who have graduated from a four-year college. As shown in Figure 3, deregulation induced a tightening in all three components.

Critically, 75% of the explained reduction in the variance of wage income is accounted for by the explained reduction in the income gap between skilled and unskilled workers. In terms of wage income, deregulation helps unskilled workers relative to skilled ones. Since, unskilled workers, on average, earn less than skilled workers, this is consistent with the findings that deregulation reduces overall income inequality.

Next, we conduct the same type of decomposition but focus on gender gap differences. We decompose the explained reduction in wage income inequality into that part accounted for by a reduction in inequality among men, that part accounted for by a reduction in inequality among women, and that part accounted for by a reduction in inequality between men and women.

As shown in Figure 3, about half of the explained reduction in the variance of wage income is accounted for by the explained reduction in the gender income gap. Deregulation boosts the incomes of female workers relative to male workers. Since, on average, women tend to earn less than men, this results helps account for the paper's major finding: deregulation reduces overall income inequality.

4. Conclusions

Policymakers and economists disagree sharply about the impact of bank regulations on the distribution of income. While some argue that the unregulated expansion of large banks will increase banking fees and reduce the economic opportunities of the poor, others hold that regulations restrict competition, protect monopolistic banks, and widen the distribution of income. More generally, an influential political economy literature stresses that income distribution considerations, rather than efficiency and growth considerations, frequently exert the dominant influence on bank regulations (see Claessens and Perotti, 2007, for an overview).

We find that liberalizing restrictions on intra-state branching across the states of the U.S. tightened the distribution of income. This finding is robust to an array of sensitivity analyses. We find no evidence that reverse causality drives the results. Moreover, the impact of deregulation on income distribution varies in a theoretically predictable manner across states with distinct economic, financial, and demographic characteristics at the time of deregulation. Deregulation tightened the distribution of income by disproportionately helping the poor, not by hurting the rich. These findings support the view that branch regulation in the United States restricted competition, protected local banking monopolies, and impeded the economic opportunities of the relatively poor.

Even among theories predicting that deregulation will reduce inequality, there are disagreements about the mechanisms. The two most influential views, both in terms of economic theory and in terms of the policy recommendations of international financial institutions, stress direct access. These two views hold that greater direct access to financial services will disproportionately expand the economic opportunities of the poor by enhancing their educational and entrepreneurial opportunities. We find no evidence that branch deregulation enhances educational attainment. Furthermore, while branch deregulation stimulates entrepreneurship (Black and Strahan, 2002), changes in entrepreneurial income account for very little of the reduction in income inequality generated by branch deregulation in the United States.

Rather, deregulation lowered income inequality by affecting labor market conditions. Most of the explained drop in income inequality is due to (1) a reduction in the wage income differential between skilled and unskilled workers and (2) a reduction in the wage income gap between men and women.

Appendix: Variance Decomposition

To provide additional information on the impact of deregulation on income distribution, we decompose the explained reduction in income inequality into different components. We use the variance of the distribution of income, rather than the Gini coefficient, so that we can perform a standard, variance decomposition.

In particular, let $Z \equiv X+Y$, where, for example, Z is total income, X is income from labor, and Y is income from non-labor sources, such as proprietor income and income from dividends and interest payments. By definition, $Var(Z) = Var(X) + Var(Y) + 2Cov(X,Y)$.

Next, define the change in the variation of Z explained by bank deregulation as

$$
\begin{aligned}
\Delta Var(Z) &\equiv [Var(Z)]_{after} - [Var(Z)]_{before} \\
&= [Var(X)+Var(Y)+2Cov(X,Y)]_{after} - [Var(X)+Var(Y)+2Cov(X,Y)]_{before} \\
&= \Delta Var(X) + \Delta Var(Y) + 2\Delta Cov(X,Y).
\end{aligned}
$$

In this context, "before" and "after" refer to the period before and after bank deregulation, respectively. To link this directly with the survey data, we must complicate this formula by accounting for the different weights on the sources of income, so that $Z \equiv \alpha X + (1-\alpha)Y$, where α is the share of income from labor and $1-\alpha$ is the share of income from non-labor sources. Thus,

$$\Delta Var(Z) = \alpha^2\Delta Var(X) + (1-\alpha)^2\Delta Var(Y) + 2\alpha(1-\alpha)\Delta Cov(X,Y).$$

$$1 \quad = [\alpha^2\Delta Var(X)/\Delta Var(Z)[+ [(1-\alpha)^2\Delta Var(Y)/\Delta Var(Z)] + [2\alpha(1-\alpha)\Delta Cov(X,Y)/ \Delta Var(Z)].$$

The first term, $[\alpha^2\Delta Var(X)/\Delta Var(Z)]$, represents the fraction of the change in the variance of Z accounted for by the change in the variance <u>within</u> X; the second term, $[(1-\alpha)^2\Delta Var(Y)/\Delta Var(Z)]$, represents the fraction of the change in the variance of Z accounted for by the change in the variance <u>within</u> Y, while the final term, $[2\alpha(1-\alpha)\Delta Cov(X,Y)/ \Delta Var(Z)]$, represents the fraction of the change in the variance of Z accounted for by the change between X and Y.

We compute ΔVar(Z), ΔVar(X), ΔVar(Y), and Cov(X,Y) from four regressions. Specifically, controlling for state and year fixed effects, we regress Var(Z), Var(X), Var(Y), and Cov(X,Y) on the state-specific bank deregulation indictor, $D_{i,t}$, that equals one after the state deregulated branching restrictions, and zero before the deregulation. Let β, γ, δ, and ψ be the estimated coefficients on $D_{i,t}$ in the regressions where Var(Z), Var(X), Var(Y), and Cov(X,Y) are the dependent variables respectively.

Thus, the explained changes in the variances and covariances are ΔVar(Z) = β, ΔVar(X) = γ, ΔVar(Y) = δ, and ΔCov(X,Y) = ψ. Since we also know α we can decompose the total explained change in the variation of Z into the change in variation within X, the change in variation within Y, and the change in variation between X and Y.

In some cases, we further decompose the covariance term and allocate the between effect to either X or Y. For example, in the case of decomposing total income into wage- and non-wage income, we want to assess whether changes in the variance of wage income inequality dominate the explained reductions in total income inequality. For the purposes of this paper, we are not interested in assessing the fraction of the explained reduction in income inequality accounted for by changes in the variance of income between wage and non-wage income. However, when we decompose income by gender, we are very interested in assessing the fraction of the explained reduction in income inequality that is accounted for by the explained reduction in incomes between men and women.

Thus, for some decompositions, we allocate the covariance term as follows:

$$\text{(A2)} \quad \Delta Var(Z) = [\alpha^2\Delta Var(X)+2\alpha(1-\alpha)w\Delta Cov(X,Y)] + [(1-\alpha)^2\Delta Var(Y)+2\alpha(1-\alpha)(1-w)\Delta Cov(X,Y)]$$

where w is the change in average wage income (x) relative to the change in average total income z, so that $w=\Delta x/\Delta z$. Intuitively, the component of the explained change in the variance of total income accounted for by a reduction in the variance between groups X and Y, $2\alpha(1-\alpha)\Delta Cov(X,Y)$, can be further decomposed into that part associated with a change in average wage income and that part associated with a change in average non-wage income. Therefore, the change in total variance attributable to the change in variance of wage income is the change of variance within wage income plus the weighted contribution of wage income to the covariance term.

The first term in the above expression, $[\alpha^2\Delta Var(X)+2\alpha(1-\alpha)w\Delta Cov(X,Y)]$, is the explained change in total variance attributable to the explained change in X (wage income), and the second term, $[(1-\alpha)^2\Delta Var(Y)+2\alpha(1-\alpha)(1-w)\Delta Cov(X,Y)]$, is the explained change in total income attributable to the explained change in Y (non-wage income). In the paper, we conduct this decomposition for various definitions of Z, X, and Y.

REFERENCES

Acemoglu, Daron (2007). Oligarchic and Democratic Societies. *Journal of the European Economic Association*, forthcoming.

Acharya, Viral, Jean Imbs and Jason Sturgess (2007). Finance and Efficiency: Do Bank Branching Regulations Matter?, London Business School mimeo.

Aghion, Philippe, Caroli, Eve, and Cecilia Garcia-Penalosa (1999). Inequality and Economic Growth: The Perspective of the New Growth Theories. *Journal of Economic Literature* 37, 1615-1660.

Amel, Dean and Nellie Liang (1992). The Relationship between Entry into Banking Markets and Changes in Legal Restrictions on Entry. *The Antitrust Bulletin* 37, 631-49.

Ashenfelter, Orley and Timothy Hannan (1986). "Sex Discrimination and Product Market Competition: The Case of the Banking Industry." *Quarterly Journal of Economics* 101, 149-73.

Banerjee, Abhijit and Andrew F. Newman (1993). Occupational Choice and the Process of Development, *Journal of Political Economy* 101, 274-98.

Beck, Thorsten, Asli Demirgüç-Kunt, and Ross Levine (2007). Finance, Inequality and the Poor: Cross-Country Evidence. *Journal of Economic Growth*, forthcoming.

Becker, Gary S. (1957). *The Economics of Discrimination*. Chicago: Chicago University Press.

Berger, Alan; Anthony Saunders; Scalise, Joseph and Udell, Gregory (1997): The Effects of Bank Mergers and Acquisitions on Small Business Lending, Board of Governors of the Federal Reserve System, Finance and Economics Discussion Series 1997-28

Bertrand, Marianne, Esther Duflo and Sendhil Mullainathan (2004): How Much Should We Trust Differences-in-Differences Estimates, Quarterly Journal of Economics 119, 249-275.

Black, Sandra E. and Elizabeth Brainerd (2004). Importing Equality? The Effects of Globalization on Gender Discrimination, *Industrial and Labor Relations Review*, 57(4), 540-559.

Black, Sandra E. and Philip Strahan (2001). The Division of Spoils: Rent-Sharing and Discrimination in a Regulated Industry. *American Economic Review*, 91(4), 814-831.

Black, Sandra E. and Philip Strahan (2002). Entrepreneurship and Bank Credit Availability, *Journal of Finance* 57, 2807-2833.

Calem, Paul (1994). The Impact of Geographic Deregulation on Small Banks, *Business Review*, Federal Research Bank of Philadelphia, 17-31.

Claessens, Stijn and Enrico Perotti (2007). Finance and Inequality: Channels and Evidence, University of Amsterdam mimeo.

Demirguc-Kunt, Asli and Ross Levine (2007). Finance and Opportunity: Financial Systems and Intergenerational Persistence of Relative Incomes, World Bank mimeo.

Demyanyk, Yuliya (2007). U.S. Banking Deregulation and Self-Employment: A Differential Impact on Those in Need, Federal Reserve Bank of St. Louis Working Paper.

Demyanyk, Yuliya, Charlotte Ostergaard, and Bent Sorensen (2007). U.S. Banking Deregulation, Small Businesses, and Interstate Insurance of Personal Income, *Journal of Finance*, forthcoming.

Flannery, Mark (1984). The Social Costs of Unit Banking Restrictions, *Journal of Monetary Economics* 13, 237-49.

Galor, Oded and Joseph Zeira. (1993). Income Distribution and Macroeconomics. *Review of Economic Studies* 60, 35-52.

Galor, Oded and Omer Moav (2004). From Physical to Human Capital Accumulation: Inequality and the Process of Development. *Review of Economic Studies* 71, 1001-26.

Galor, Oded and Omer Moav. 2006. Das Human-Kapital: A Theory of the Demise of the Class Structure, *Review of Economics Studies* 73, 85-117.

Greenwood, Jeremy and Boyan Jovanovic. (1990). Financial Development, Growth, and the Distribution of Income, *Journal of Political Economy* 98, 1076-1107.

Haber, Stephen. (2007). Political Institutions and Financial Development: Evidence from the Political Economy of Banking Regulation in the New World Economies, mimeo, Stanford University.

Hammond Bray. (1957). *Bank and Politics in America: From the Revolution to the Civil War*. Princeton University Press.

Huang, Rocco (2007). Did Branching Deregulation Accelerate Growth? *Journal of Financial Economics*, forthcoming.

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

Jayaratne, Jith and Phil Strahan (1998). Entry Restrictions, Industry Evolution and Dynamic Efficiency: Evidence from Commercial Banking, *Journal of Law and Economics* 41, 239-74.

Jerzmanowski, Michal and Malhar Nabar (2007). Financial Development and Wage Inequality: Theory and Evidence. Mimeo.

Juhn, Chinhui; Kevin M. Murphy, and Brooks Pierce (1993). Wage Inequality and the Rise in Returns to Skill, *Journal of Political Economy* 101(3), 410-42.

Katz, Lawrence F. and David H. Autor (1999). Changes in the Wage Structure and Earnings Inequality, in Orley Ashenfelter and David Card, eds., *Handbook of Labor Economics* Vol. 3A. Amsterdam: Elsevier Science, North-Holland, 1463-1555.

Kroszner, Randall and Philip Strahan (1999). What Drives Deregulation? Economics and Politics of the Relaxation of Bank Branching Deregulation, *Quarterly Journal of Economics* 114, 1437-67.

McLaughin, Susan (1995). The Impact of Interstate Banking and Branching Reform: Evidence from the States, *Current Issues in Economics and Finance* 1, Federal Reserve Bank of New York.

Mincer, Jacob and Solomon Polachek (1974). Family Investments in Human Capital; Earnings of Women, *Journal of Political Economy* 82(2), S76-S108.

Mulligan, Casey and Yona Rubinstein (2007a). Selection, Investment, and Women's Relative Wages Since 1975, Brown University mimeo.

Mulligan, Casey and Yona Rubinstein (2007b). The Closing Gender Gap as a Roy Model Illusion, Brown University Mimeo.

Mulligan, Casey and Yona Rubinstein (2007c). Specialization, Inequality, and the Labor Market for Married Women, Brown University Mimeo.

Savage, Donald T. (1993). Interstate Banking: A Status Report, *Federal Reserve Bulletin* 79, 1075-89.

Southworth, Shirley D. 1928. *Branch Banking in the United States*. New York, NY: McGraw-Hill.

White, Eugene N. (1982). The Political Economy of Banking Regulation, 1864-1933. *Journal of Economic History* 42, 33-40.

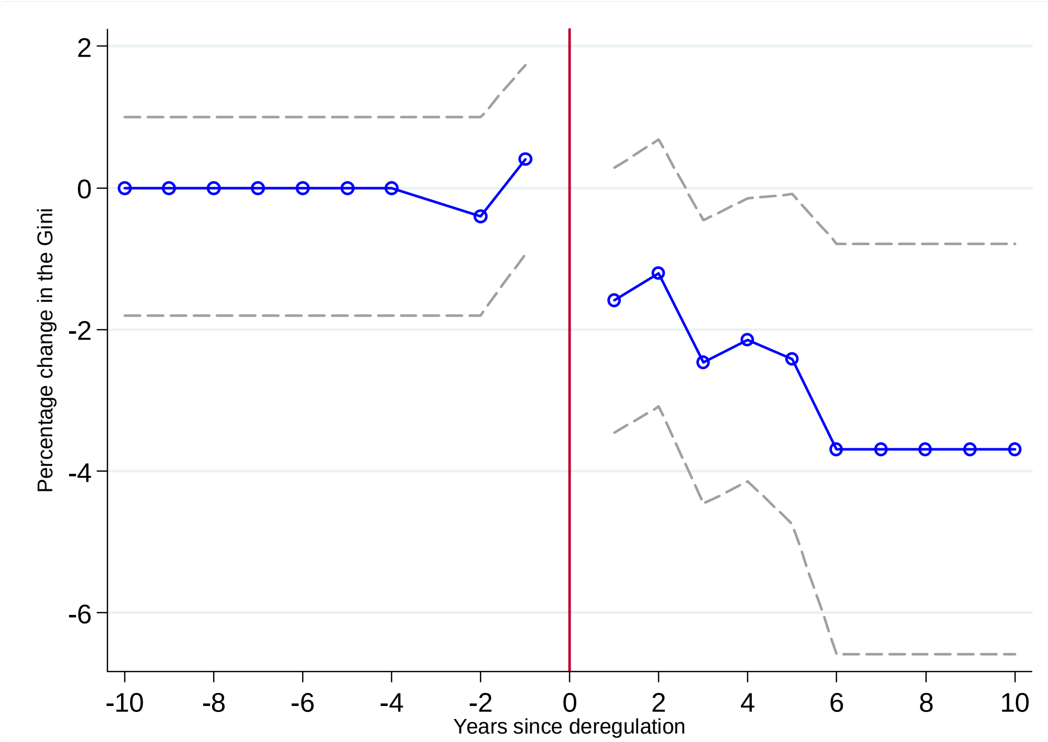

Figure 1 – The Effect of Deregulation on the Gini Index

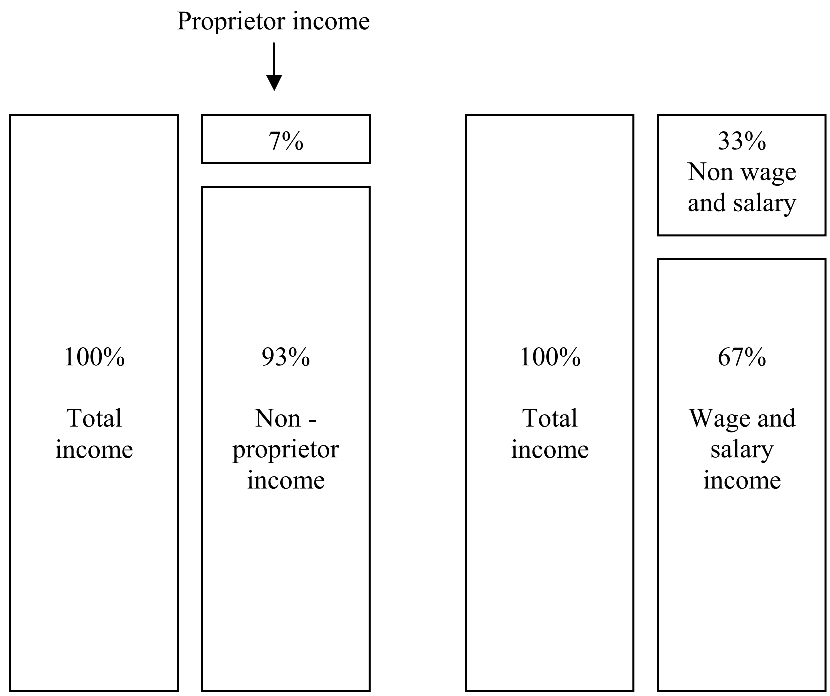

Figure 2 – Decomposing the Reduction in Inequality by Type of Income

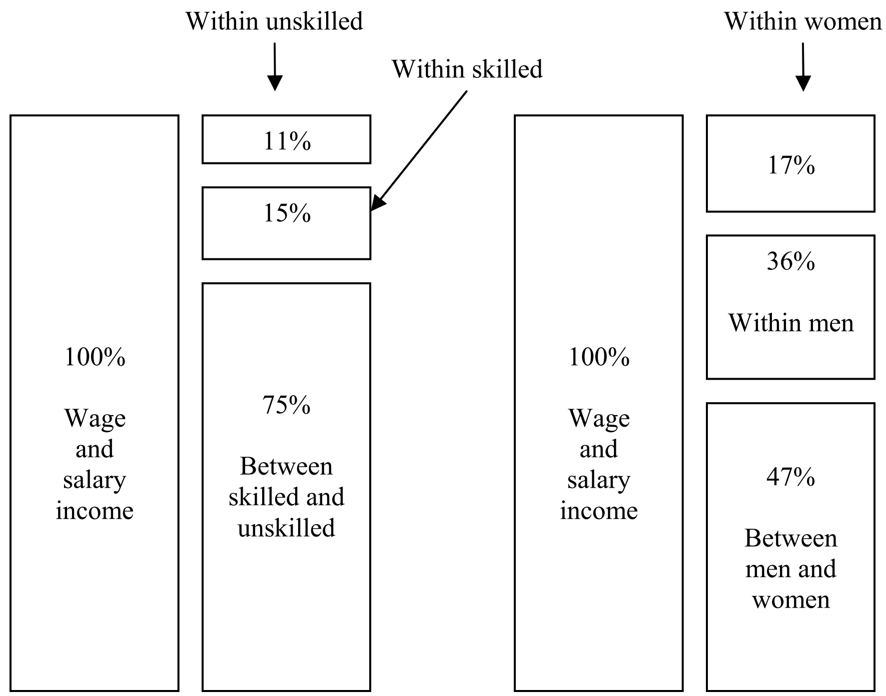

Figure 3 – Decomposing the Reduction in Inequality by Skills and Gender

Table 1: Descriptive Statistics

A. Data Construction

Restriction	Persons deleted	% deleted	Persons left
CPS sample years 1977-2006			4,878,496
Adult civilians, aged 25-54, in households	3,044,675	62.4%	1,833,821
Whites and blacks only	300,726	16.4%	1,533,095
Non-allocated incomes with positive sampling weights	98,773	6.4%	1,434,322

B. Sample Statistics	N	Mean	Min	Max
White	1,434,322	0.87	0	1
Education: High school or more	1,434,322	0.89	0	1
Education: College or more	1,434,322	0.27	0	1
Self-employed	1,434,322	0.09	0	1
Work for wages	1,434,322	0.73	0	1
Not in the labor force	1,434,322	0.18	0	1

C. Statistics on the Gini Coefficient of Income Distribution

					Standard deviation		
	N	Mean	Min.	Max.	Cross states	Within states	Within state-years
Log(Gini):							
Total income	1,433	-0.745	-1.013	-0.531	0.056	0.056	0.041
Wage and salary income	1,433	-0.947	-1.174	-0.715	0.056	0.055	0.049
Proprietor (self-employed) income	1,433	-0.368	-1.199	0.127	0.11	0.116	0.093

Note: A) The table is based on the March Demographic Supplements to the Current Population Surveys (CPS) corresponding to income received in years 1976-2005 for prime-age (25-54) adult civilians. The sample is restricted to whites and blacks who do not reside in institutions and excludes individuals with allocated incomes. Finally, we eliminate 21 persons that are assigned a zero sampling weight by the CPS. B) For panel B, we use the sampling weights provided by the CPS. C) The Gini indexes were calculated for each state and year, for the 30 years between 1976 and 2005 and for 48 states and the District of Columbia, excluding South Dakota and Delaware. We exclude the year of deregulation, which drops 37 observations (37 states deregulated between 1976 and 2005). Thus, there are 1,433 state-year observations in the sample. All incomes are corrected for top-coding.

Table 2 - The Effect of Deregulation on the Distribution of Income

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Deregulation	-0.026***	-0.025***	-0.024**	-0.026***	-0.024**	-0.023**	-0.021***
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.008)
GSP per capita growth		-0.055				-0.045	-0.030
		(0.045)				(0.046)	(0.040)
Unemployment rate			0.005***			0.006***	0.007***
			(0.002)			(0.002)	(0.002)
Gov taxes / income				0.295**		0.436**	0.342***
				(0.114)		(0.170)	(0.099)
Gov expenditures / income					-0.301*	-0.479**	-0.441**
					(0.152)	(0.179)	(0.210)
Gini growth, t-1							0.114***
							(0.033)
Observations	1433	1433	1433	1335	1335	1335	1239
R-squared	0.464	0.465	0.473	0.480	0.481	0.500	0.470

Note: Standard errors are adjusted for state-level clustering and appear in parentheses. All models control for state and year fixed effects. The dependent variable is ln(Gini) which is based on total personal income. The sample is described in the note to Table 1. Data on government taxes and expenditures are only available until 2003. A * indicates significance at 10%; ** significance at 5%; and *** significance at 1%.

Table 3 - The Effect of Deregulation on Income Distribution by Type of Income

	Gini	90/10	75/25	Variance
A. Total income				
Deregulation	-0.024**	-0.250	-0.155***	-0.065*
	(0.009)	(0.199)	(0.055)	(0.037)
Unemployment rate	0.005***	0.150***	0.017**	-0.005
	(0.002)	(0.040)	(0.008)	(0.007)
Observations	1433	1433	1433	1433
R-squared	0.473	0.816	0.599	0.592
B. Wage and salary income				
Deregulation	-0.021***	-0.102***	-0.050***	-0.058
	(0.007)	(0.030)	(0.013)	(0.036)
Unemployment rate	0.010***	0.091***	0.024***	0.001
	(0.002)	(0.010)	(0.002)	(0.008)
Observations	1433	1433	1433	1433
R-squared	0.247	0.470	0.359	0.595
C. Proprietor income				
Deregulation	-0.047**	-0.318*	-1.471***	-0.124**
	(0.020)	(0.169)	(0.459)	(0.059)
Unemployment rate	-0.004	-0.076**	-0.064	-0.013
	(0.003)	(0.034)	(0.068)	(0.011)
Observations	1433	1394	1432	1433
R-squared	0.369	0.039	0.335	0.165

Note: Standard errors are adjusted for state-level clustering and appear in parentheses. All models control for state and year fixed effects. The sample is described in the notes to Table 1. The dependent variables are as follows: The natural logarithm of the Gini coefficient of income distribution (Gini) in column (1), the difference between natural logarithm of the 90th and 10th percentiles of the distribution of income (90/10) in column (2); the difference between natural logarithm of the 75th and 25th percentiles of the distribution of income (75/25) in column (3), and the natural logarithm of the coefficient of variation of income (Variance), which is the variance divided by the mean, in column (4). For some state-years the 10th, and in one case the 25th, percentiles of the total income distribution are zero. In order to take logs we therefore inflated zero income by 1 dollar for the 10th percentile and by 2 dollars for the 25th percentile. We followed a similar procedure for proprietor income. A * indicates significance at 10%; ** significance at 5%; and *** significance at 1%.

Table 4 - The Effect of Deregulation on Gini as a Function of Initial State Characteristics

	(1)	(2)	(3)	(4)	(5)
Deregulation	-0.026***	-0.009	0.024	0.577*	-0.017*
	(0.009)	(0.010)	(0.015)	(0.294)	(0.009)
(Deregulation x unit banking)		-0.032***			
		(0.010)			
(Deregulation x initial share of small banks)			-0.351***		
			(0.129)		
(Deregulation x initial share of small firms)				-0.672*	
				(0.334)	
(Deregulation x initial population dispersion)					-0.210***
					(0.071)
Linear combinations					
Deregulation + (Deregulation x initial share of small banks):					
evaluated at the 25th percentile of the small banks distribution			-0.000		
			(0.008)		
evaluated at the median of the small banks distribution			-0.013*		
			(0.007)		
evaluated at the 75th percentile of the small banks distribution			-0.026***		
			(0.009)		
Deregulation + (Deregulation x initial share of small firms):					
evaluated at the 25th percentile of the small firms distribution				-0.008	
				(0.008)	
evaluated at the median of the small firms distribution				-0.015**	
				(0.007)	
evaluated at the 75th percentile of the small firms distribution				-0.025***	
				(0.009)	
Observations	1433	1433	1133	1133	1433
R-squared	0.464	0.476	0.541	0.531	0.473

Note: Standard errors are adjusted for state-level clustering and appear in parentheses. All models control for state and year fixed effects. The sample is described in the notes to Table 1. The dependent variable is the natural logarithm of the Gini coefficient of total personal income (Gini). The first column replicates the results reported in the second column in Table 2. Since we control for state fixed effects and since unit banking, initial share of small banks, initial share of small firms, and initial population dispersion are time-invariant, the direct effect of unit banking, initial share of small banks, initial share of small firms, and initial population dispersion are dropped from the regressions. The data on the share of small banks and small firms is obtained from Kroszner and Strahan (1999). These data exclude 10 states that deregulated in 1960 and therefore have 300 fewer state-year observations. A * indicates significance at 10%; ** significance at 5%; and *** significance at 1%.

Table 5 - Reverse Causality

	(1)	(2)	(3)	(4)	(5)	(6)
Small bank share	4.562***	4.554***	3.767***	3.776***	3.710**	3.498***
	(1.199)	(1.242)	(1.167)	(1.183)	(1.467)	(0.938)
Relative small bank capital-asset ratio	7.470***	7.469***	5.978**	5.997**	6.058**	3.955
	(2.811)	(2.802)	(2.557)	(2.803)	(2.792)	(2.706)
Relative Insurance Value added	-0.956***	-0.946***	-1.338***	-1.016***	-1.027***	-0.442
	(0.299)	(0.354)	(0.316)	(0.265)	(0.338)	(0.316)
Banks allowed to sell insurance	-1.168**	-1.164**	-1.155**	-1.148**	-1.143**	-0.439
	(0.561)	(0.568)	(0.484)	(0.559)	(0.563)	(0.424)
Relative Insurance Value added*Banks	3.198**	3.185**	3.147***	3.089**	3.069**	1.044
allowed to sell insurance	(1.299)	(1.333)	(1.113)	(1.297)	(1.315)	(1.014)
Share of firms that are small	-5.505***	-5.523***	-4.570**	-5.116***	-5.074***	-10.176***
	(2.031)	(1.935)	(1.857)	(1.836)	(1.943)	(2.119)
Share of govt. controlled by Democrats	0.241**	0.240**	0.218**	0.194**	0.192**	0.135*
	(0.095)	(0.095)	(0.085)	(0.084)	(0.088)	(0.072)
Single party controls whole state govt.	-0.025	-0.025	-0.016	-0.002	-0.002	0.108
	(0.061)	(0.063)	(0.057)	(0.058)	(0.058)	(0.077)
Gini coefficient	-0.013	-0.013	-0.019	-0.007	-0.007	-0.015
	(0.013)	(0.013)	(0.012)	(0.013)	(0.014)	(0.010)
Interest rate spread		0.164				
		(3.323)				
Bank failure rate			1.999***			
			(0.774)			
Unit banking				0.122**	0.124**	0.214***
				(0.055)	(0.061)	(0.077)
State changes bank insurance powers					-0.011	-0.127
					(0.109)	(0.105)
Includes regional dummies	No	No	No	No	No	Yes
Constant	8.005***	8.011***	7.734***	7.473***	7.464***	12.423***
	(1.778)	(1.736)	(1.503)	(1.573)	(1.582)	(1.719)
LnP	1.799***	1.802***	1.952***	1.830***	1.830***	2.058***
	(0.148)	(0.193)	(0.175)	(0.149)	(0.148)	(0.156)
RC	0.000	0.000	0.000	0.000	0.000	0.000
Log-Likelihood	9.46	9.46	10.79	10.81	10.82	21.72
Number of Observations	408	408	408	408	408	408
p-value of Chi-Squared	0.000	0.000	0.000	0.000	0.000	0.000

Note: This table replicates Table III from Kroszner and Strahan (1999). The model is a Weibull hazard model where the dependent variable is the log expected time to deregulation. All variables, including the Gini index, are included in levels. Sample period Is 1976 to 1994 and the sample comprises 37 states that deregulated after 1977. States drop from the sample once they deregulate. Robust standard errors are in parentheses. A * indicates significance at 10%; ** significance at 5%; and *** indicates significance at 1%.

Table 6 - The Effect of Deregulation on Educational Attainment

	% High school +		% College +	
Deregulation	0.006		-0.003	
	(0.007)		(0.005)	
5-9 yrs after deregulation		-0.000		0.002
		(0.004)		(0.002)
10-14 yrs after deregulation		0.001		0.001
		(0.007)		(0.004)
15-19 yrs after deregulation		0.005		0.005
		(0.011)		(0.006)
20-24 yrs after deregulation		0.006		0.012*
		(0.013)		(0.007)
25-29 yrs after deregulation		0.002		0.011
		(0.013)		(0.008)
30+ yrs after deregulation		-0.015		0.009
		(0.018)		(0.009)
Observations	1385	1421	1385	1421
R-squared	0.901	0.904	0.806	0.808

Notes: Standard errors are adjusted for state-level clustering and appear in parentheses. All models control for state and year fixed effects. The sample is similar to the one described in the notes to Table 1, except that in this table the period is 1977-2005 rather than 1976-2005. In columns (1) and (3) the year of deregulation is dropped. In columns (2) and (4), each right-hand side variable equals one for the indicated period. A * indicates significance at 10%.

Table A1 -- Branching Deregulation Events

State	State_code	Year of deregulation
Alabama	AL	1981
Alaska	AK	1960
Arizona	AZ	1960
Arkansas	AR	1994
California	CA	1960
Colorado	CO	1991
Connecticut	CT	1980
District of Columbia	DC	1960
Florida	FL	1988
Georgia	GA	1983
Hawaii	HI	1986
Idaho	ID	1960
Illinois	IL	1988
Indiana	IN	1989
Iowa	IA	1999
Kansas	KS	1987
Kentucky	KY	1990
Louisiana	LA	1988
Maine	ME	1975
Maryland	MD	1960
Massachusetts	MA	1984
Michigan	MI	1987
Minnesota	MN	1993
Mississippi	MS	1986
Missouri	MO	1990
Montana	MT	1990
Nebraska	NE	1985
Nevada	NV	1960
New Hampshire	NH	1987
New Jersey	NJ	1977
New Mexico	NM	1991
New York	NY	1976
North Carolina	NC	1960
North Dakota	ND	1987
Ohio	OH	1979
Oklahoma	OK	1988
Oregon	OR	1985
Pennsylvania	PA	1982
Rhode Island	RI	1960
South Carolina	SC	1960
Tennessee	TN	1985
Texas	TX	1988
Utah	UT	1981
Vermont	VT	1970
Virginia	VA	1978
Washington	WA	1985
West Virginia	WV	1987
Wisconsin	WI	1990
Wyoming	WY	1988