# Bank Regulation and Income Distribution:

## Evidence from Branch Deregulation

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**Abstract:** Policymakers and economist disagree about the impact of bank regulations on the distribution of income. We test whether liberalizing restrictions on intra-state branching in the states of the United States intensified, ameliorated, or had no effect on income distribution. We find that branch deregulation lowered income inequality, but the results contradict the major theoretical explanations for this effect. Deregulation lowered income inequality by affecting labor market conditions, not by providing the poor with greater access to financial services. About half of the explained drop in income inequality is due to a reduction in the income gap between men and women.

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Throughout the history of the United States, political leaders have argued that large banks benefit only the rich (Hammond, 1957). Besides constitutional considerations, Thomas Jefferson's fears of concentrated banking power spurred him to fight against Alexander Hamilton's proposal for the founding of the Bank of the United States. Similar anxieties fueled the termination of this bank in 1811, Andrew Jackson's veto of the re-chartering of the Second Bank of the Unites States in 1832, and regulatory prohibitions on banks opening branches both across and within state boundaries.

The view that large banks primarily help the wealthy finds support in modern theory and continues to shape bank regulations around the world. In Greenwood and Jovanovic's (1993) model, banks assist those wealthy enough to pay the fee associated with accessing banking services. In practice, most countries regulate bank mergers, acquisitions, and bank activities with the stated goal of constraining large banks (Barth et al, 2006). In the U.S., many intra- and inter-state branching regulations lasted through the mid-1990s, where proponents explicitly argued that these regulations would limit the concentration of economic power (Southworth, 1928; White, 1982).

Countervailing arguments, however, challenge the view that impeding the formation of large banks helps the poor. In theories by Galor and Zeira (1993) and Galor and Moav (2004), regulatory restrictions that impede competition and raise the price of financial services will disproportionately hurt the poor by reducing their access to financial services. In the United States, for example, Flannery (1984) shows that state branching restrictions prevented banks from competing in distant areas, which created and protected small banks with local monopolies. Indeed, White (1982), Haber (2007), and others argue that governments enact banking regulations precisely to assist favored banks and firms. From this perspective, branching restrictions protected firms with close connections to local banks and therefore limited the economic opportunities of many poor individuals without those connections.

In this paper, we provide the first assessment of whether liberalizing restrictions on intra-state branching affected the distribution of income within states. More specifically, from the mid-1970s through the mid-1990s, most states in the United States removed regulatory restrictions on branching. We test whether removing these restrictions intensified, ameliorated, or had no effect on the skewness of income distribution. Past work shows that liberalizing restrictions on intra-state branching (i) increased the average size of banks through consolidation (Calem, 1994; Savage, 1993), (ii) improved bank efficiency (Jayaratne and Strahan, 1998), (iii) boosted the entry of new, non-financial companies (Black and Strahan, 2002), and (iv) accelerated average per capita income growth (Jayaratne and Strahan, 1996). We examine the impact of bank 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 powerful insurance companies 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 distribution, examining different components of income, controlling for time-varying state characteristics, and conditioning on state and year fixed effects. While income inequality widened in the U.S. during this period, we show that branch deregulation lowered income inequality relative to this national trend by using year fixed effects. The magnitude is consequential: Deregulation explains 45% of the variation of income inequality during the sample period relative to state and year averages; or put differently, deregulation's estimated effect eliminates about 15% of the large trend increase in income inequality that the U.S. experienced during the last decades. Furthermore, deregulation reduces income inequality by exerting a disproportionately positive impact on the poor, not by hurting the rich.

We next evaluate the two major theoretical explanations for this finding. The first explanation focuses on the degree to which poor individuals can directly access financial services to fund their accumulation of human capital. In the model by Galor and Zeira (1993), the high cost of acquiring human capital together with capital market imperfections prevent the poor from borrowing to fund education. This implies that comparatively talented, but poor, individuals obtain too little schooling. This socially inefficient allocation of schooling slows economic growth and intensifies the persistence of relative income differences across generations. From this perspective, deregulation that reduces credit market imperfections permits more individuals to finance their education, which accelerates growth and reduces inefficient income inequality.

A second explanation for the finding that deregulation reduces income inequality focuses on the increased ability of the poor to access financial services and become entrepreneurs (Banerjee and Newman, 1993). Financial imperfections are particularly binding on the poor because they lack collateral and access fees are high relative to their incomes. The severity of these imperfections determines the extent to which poor individuals can raise external funds to initiate projects. Again, financial imperfections (i) reduce aggregate economic growth by hindering the efficient allocation of capital and (ii) maintain a skewed distribution of income by preventing the poor from accessing financial services. Thus, branch deregulation that improves bank efficiency may exert a particularly beneficial impact on the poor by lowering borrowing costs and reducing collateral requirements.

Our empirical estimates, however, do not support either of these two leading theories. First, the impact of branch deregulation levels-off quickly. Relative to state and year fixed effects, the Gini coefficient of income inequality falls for four years after deregulation and then stops falling. Deregulation has a level effect that fully materializes in four years, not a trend effect. The time pattern seems inconsistent with the education explanation, which implies a growing effect on income distribution as poor individuals accumulate education. Second, we assess how much of the impact of deregulation on the distribution of income is accounted for by entrepreneurship. The impact of deregulation on proprietor income accounts for only 4% of the reduction in total income variance explained by branch deregulation. Education and entrepreneurship explain exceptionally little of branch deregulation's impact on income distribution.

In light of these findings, we advance and evaluate a labor market explanation for why branch deregulation tightens the distribution of income. Our explanation does not rely on the poor directly accessing financial services to purchase education or become entrepreneurs. As previously noted branch deregulation exposed local banking monopolies to greater competition, boosted banking efficiency, and enhanced the efficiency of capital allocation among non-financial firms. This scenario suggests at least three interrelated channels linking deregulation to changes in labor market conditions. First, improved capital allocation will increase the marginal product of labor, boosting wages. If this effect falls primarily on low-skilled workers, branch deregulation will reduce income inequality among wage earners. This first channel does not necessarily differentiate by gender or race. The second channel also starts with improved capital allocation boosting wages, but focuses on female labor force participation. Holding reservation wages constant, deregulation that boosts wages will pull new women into the work force with higher reservation wages than women who were already working. This will reduce income inequality. Third, as deregulation spurs bank competition and hence breaks calcified bonds between banks and inefficient firms, it will also spur competition among nonfinancial firms. According to Becker (1957), this competition makes gender and racial discrimination more expensive;[1] thus, branch deregulation lowers income inequality by reducing the income gap between genders and races.

We find considerable support for the general hypothesis that deregulation reduces income inequality by affecting labor market conditions, not by providing the poor with direct access to financial services. In particular, we provide four additional results that provide a richer picture of the impact of deregulation on the economy, though we do not yet fully discriminate among the specific mechanisms potentially linking deregulation with labor market conditions. First, most of the explained reduction in the variance of income is due to a reduction of the income gap between men and women. While deregulation reduces income inequality among men, and among women for some categories of income, the largest proportion of the explained drop in total income inequality is accounted for by a reduction in the gender income gap. Second, deregulation increased women's labor force participation by about one percent above the positive trend in women's labor participation

[1] Also, see Ashenfelter and Hannan (1986), Black and Strahan (2001), and Black and Brainerd (2004).

during this period. Deregulation works on the extensive margin by pulling women into the labor force and reducing income inequality. Third, deregulation boosts wages. Specification, the wage rates of both unskilled and skilled women rise more than the corresponding wage rates of men, and the wages of skilled workers rise more than those of unskilled workers for both genders. Finally, none of this paper's results are explained by minorities, nor do the results hold for minorities. While all of the results hold when including or excluding minorities from the analyses, deregulation does not reduce the income gap between whites and blacks, or between whites and non-whites, nor does deregulation reduce inequality within minority groups.

While these additional findings suggest avenues for future research, they emphasize three simple conclusions within the more limited confines of this paper. First, deregulation lowered income inequality primarily by affecting labor market conditions, not by providing the poor with more direct access to financial services. Second, a big part of this effect operates by reducing the income gap between men and women. Finally, deregulation both increased female labor force participation, and increased female wages relative to men's wages, for both skilled and unskilled workers.

This paper relates to an enormous literature on the determinants and consequences of the distribution of income. From a policy perspective, the literature primarily examines 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). Redistribution, however, has adverse incentive effects that are not a shortcoming of financial reforms. Thus, our work contributes to the policy debate on how to address socially inefficient income inequality. Furthermore, the international policy community is increasingly emphasizing the benefits of providing the poor with greater access to financial services as a vehicle for fighting poverty. While branch deregulation within U.S. states may not provide clear lessons for developing economies, the results do raise the possibility that financial development may help the poor primarily by intensifying competition and boosting wages, not by expanding direct access to financial services. This warrants further research. Finally, 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.[2] In contrast, we analyze the impact of a specific, exogenous policy change on the distribution of income, rather than examining an overall index of banking sector development. By using the difference-in-differences approach across U.S. states we increase the power of the econometric tests and reduce concerns about omitted variables and endogeneity.

Before continuing, we highlight two analytical limitations that we are in the process of rectifying. First, we do not trace individuals through time. We use the Current Population Survey, which is a regular, cross-sectional survey, but it is not a longitudinal study of particular people. With these data, therefore, we cannot examine exactly who was helped and hurt by deregulation. By using the Panel Study of Income Dynamics, we plan to provide more direct evidence on how deregulation influenced individuals in the next version of this paper. Second, the findings do not indicate that bank deregulation is everywhere and always beneficial. As emphasized, ample evidence suggests that branch deregulation in the U.S. from the 1970s through the 1990s intensified competition. We interpret the results as suggesting that the intensification of competition both increased economic growth and reduced income inequality. Thus, it is not deregulation per se that drives the results; it is deregulation that intensifies competition. We plan to provide more direct evidence on this channel in the next version of this paper.

[2] Beck, Demirguc-Kunt, and Levine (2007) find no effect of stock market development on income distribution, but also see Das and Mohapatra (2003).

## 1. Data and Methodology

To assess the effect of branching deregulation on income distribution, we gather data on the timing of deregulation, income distribution across states and time, 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. While the intra-state branching deregulation was in many cases accompanied by inter-state branching deregulation, allowing banking holding companies to expand across state borders, the literature has found little effect of inter-state branching deregulation on banking market structure or real outcomes. We therefore focus on intra-state branching 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.[3] Fifteen states deregulated before the start of our sample period in 1977. 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 United States.[4] The CPS contains information at the individual and household level on the sources of income. CPS provides information on total household income, total personal income, income from wages, proprietor's income, etc. Furthermore, we also categorize these groups by gender and race. 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.[5]

We measure the distribution of income for each state and year over the period 1977-2003 in three ways.[6] First, the Gini coefficient of income inequality is derived from the Lorenz curve, where larger values imply greater income inequality. Second, the variance of the distribution provides an alternative measure of the dispersion of incomes within a state during a given year. Third, we use the difference between the incomes of those at the 90$^{th}$ percentile and those at the 10$^{th}$ percentile as a final measure of income distribution. We start our analysis in 1977 since the 1978 March survey is the first survey in which we can distinguish between households and other institutions. We restrict

[3] 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.
[4] Jayaratne and Strahan (1996) rely on Department of Commerce data to examine the impact of deregulation on income growth. The Commerce data represent the aggregate income of the entire population, not simply information from a survey. Unfortunately, since these are aggregate data, they cannot be used to compute income distribution measures.
[5] Each household and each of its members are assigned a sample weight which corresponds to their representativeness in the population, which we use in our analysis.
[6] Income years 1977-2003 correspond to surveys conducted in 1978-2004.

our analysis to households with at least one member aged 15+ with non-missing income and drop households with negative total household income.[7]

Table 1 presents descriptive statistics. The average Gini coefficient across states and over time is 0.42, corresponding to an average of -0.86 for the log of Gini, with a range from 0.33 in New Hampshire in the year 1978 to 0.56 in District of Colombia in the year 1998. The standard deviation of the log of Gini is 0.71, with very similar standard deviations across states (0.49) and within states over years (0.52). Table 1 also shows that the Gini for individual income is on average higher than the household Gini, but shows a smaller standard deviation. The Gini coefficient for wage and salary incomes is of similar magnitudes as the household income Gini, but has a lower standard deviation. Computing separate Gini coefficients for male and female wage and salary earners, however, shows a large standard deviation in the Gini of male wage and salary incomes across states and over time. Finally, the Gini of proprietors' income is on average substantially higher than the average household income Gini and also shows a much higher standard deviation.

### 1.3. Control variables

We use the U.S. Department of Commerce data to calculate the growth rate of Gross State Product (GSP) and per capita personal income for the years 1977-2003. We use current and lagged growth rates of GSP as time-varying controls for the state of the economy. We also control for 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. Finally, we control for educational attainment across states and over time by including the ratio of college graduates to total population. In the era of skill-biased technological change, higher

[7] 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.

ratio of college graduates to total population results in higher income inequality. These ratios were calculated using the March Demographic Supplement described earlier.

### 1.4. Methodology

We use the differences-in-differences estimation technique to assess the relationship between branch deregulation and income distribution. Specifically, we utilize the following regression set-up:

$$Y_{i,t} = \alpha_i + \beta_t + \gamma D_{i,t} + \delta X_{i,t} + \varepsilon_{i,t}, \quad i=1,\ldots,49; \quad t=1977,\ldots,2003 \qquad (1)$$

where Y is a measure of income distribution, $\alpha$ and $\beta$ are vectors of state and year fixed-effects, X is a set of time-varying state-level variables and $\varepsilon$ is the error term. The variable of interest is D, a dummy variable that takes on the value one after a state deregulates. The coefficient, $\gamma$, therefore indicates the impact of branch deregulation on income distribution. A positive and significant $\gamma$ suggests that deregulation exerts a positive effect on the degree of income inequality, while a negative and significant $\gamma$ indicates the deregulation pushed income inequality lower.

The difference-in difference 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 $\gamma$ 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. Following recommendations by Bertrand, Duflo, and Mullainathan (2004), we estimate regression (1) allowing for state-level clustering, i.e. allowing for correlations in the error terms over time within states.

When running regression (1), we drop the year in which deregulation took place. In total, we have data for 48 states plus the District of Columbia, over 27 years, minus 36 state-years in which deregulation took place. This leaves us with 1,287 state-year observations which serve as the base for our analysis.

## 2. Results

### 2.1. Branch deregulation and income distribution

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 least at the 5% level in all five regressions. The column 1 result suggests that relative to the state-specific average and relative to the year-specific average, branching deregulation results in a 1.5 percentage point drop in the Gini coefficient. To gauge the economic effect of this result, we compare the coefficient estimate to the standard deviation of log Gini from state and year averages. This standard deviation is 0.034 (Table 1), suggesting that branching deregulation explains almost 45% of the variation of log Gini relate 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

Regressions 2 through 5 show the robustness of our findings to controlling for other potential time-varying state-level factors associated with income inequality. The column 2 regression shows that higher tax burden on income is associated with higher income inequality, while higher government spending is associated with lower income inequality. We do not, however, find any significant association between Gini and of economic growth, as measured by growth in Gross State Product (GSP). This does not imply that economic growth is unimportant in explaining income inequality; rather, it suggests that state-level growth is not associated with deviations of state-level income inequality from the state and year-average Gini coefficient. The column 3 through 5 regression suggest that there might be a negative relationship between lagged GSP growth and income inequality, however, once we control for the lagged Gini coefficient, the coefficient on lagged GSP growth turns insignificant. GSP growth lagged by two years never enters significantly. The share of college graduates in the total population is not significantly associated with deviations of the Gini coefficient from state- and year-average either. Finally, the column 5 regression suggests a strong persistence of the Gini coefficient over time. When controlling for the lag of the Gini coefficient, the size of the coefficient of the branching deregulation dummy drops to -0.01, but continues to enter significantly at the 5% level.

As a robustness test, and to clarify 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, $D\_4_j$ equals one for all observations in state j that are four or more years before deregulation, $D\_3_j$ equals one for the year three years before deregulation in state j, etc. $D1_j$ equals one for the year one year after deregulation in state j, $D2_j$ equals one for the year two years after deregulation in state j, etc. Figure 1 plots the results and the 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 four years after deregulation. As shown, $D\_x_j$ is insignificantly different from zero for all years before to deregulation. This suggests that changes in the distribution of income did not induce 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. The impact continues to grow slightly for one additional year. Thus, the full effect of deregulation is observed within four years, and is maintained thereafter. Deregulation has a level effect on the distribution of income that is realized within four years, but no trend effect on the Gini coefficient of income inequality.

In unreported regressions, we show that the relationship between branching deregulation and income inequality is weaker in states where there was higher political resistance against deregulation from the insurance industry and where deregulation therefore took place later. We interact the branching deregulation dummy with the size of the insurance sector (total assets in that state and year) relative to the sum of banking plus insurance sector assets. This interaction term of branching deregulation enters significantly and positively, partly off-setting the negative effect of branching deregulation. This is consistent with findings by Kroszner and Strahan (1999) that states where banks were allowed to distribute insurance products and insurance companies had a larger market share relative to banks deregulated later. Given that technology was undermining the effectiveness of branching restrictions, we expect the effect of branching deregulation therefore to be smaller for states that deregulated later, as part of the positive impact on income equity has most likely already been captured. This is what we find. Comparing the coefficient sizes on the deregulation dummy and its interaction with the relative economic power of insurance companies indicates that at values above one standard deviation of the average ratio of insurance to insurance and banking assets, the relationship between branching deregulation and income distribution falls, turning insignificant.

### 2.2. Deregulation and alternative measures of income distribution

The negative impact of deregulation on income inequality holds when using different measures of income inequality. Table 3 reports the analyses based on total household income, total individual income, and individual wage and salary income, both when computing the Gini coefficient and when using the variance of income. Thus, we report the results from six regressions. As shown, we confirm that deregulation reduces income inequality for these different categories of income and when measuring inequality either with the Gini coefficient or with the variance of income across households or individuals.

### 2.3. Assessing education and entrepreneurship theories

Theory advances two candidate explanations of why deregulation reduces income inequality. Galor and Zeira (1993) do not directly examine branch deregulation, but their theory of the co-evolution of economic growth and the distribution 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.

Banerjee and Newman (1993) instead focus on entrepreneurship. In their model, financial imperfections are particularly harmful to the poor. This might arise because the poor lack collateral or the fixed costs of accessing financial services are prohibitively high. In this context, financial imperfections both retard aggregate economic growth by hindering the efficient allocation of capital to poor, but entrepreneurially talented individuals and maintain an inefficiently skewed income distribution by preventing the poor from accessing financial services. By reducing credit market imperfections, branch deregulation can spur growth and reduce inequality.

Although our findings are consistent with the predictions from these theories that deregulation lowers inequality, we now provide a more precise evaluation of the particular channels proposed by these theories. We test whether deregulation reduces inequality by promoting education and entrepreneurship. First, there are two results that suggest that education is not driving the results. In Table 2, controlling for college education does not affect the coefficient estimates or the standard errors. These findings suggest that deregulation is not providing the poor with easier access to credit to fund a college education that reduces income inequality. Furthermore, Figure 1 shows that the full effects of deregulation on the distribution of income are realized in four years. These dynamics seem inconsistent with the view that branch deregulation eases credit constraints, which permits the poor to more easily accumulate 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.

Second, the findings suggest that entrepreneurship 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 also decompose the impact of deregulation on the distribution of income into that part accounted for by wage income and that part accounted for by non-wage income, including proprietor income. 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. The Appendix describes the decomposition in greater detail. As shown in Figure 2, the impact of deregulation on proprietor income accounts for only 4% of the reduction in total income variance explained by branch deregulation. When focusing only on wage income, Figure 2 shows that all sources of non-wage income account for only 15% of the reduction in the reduction in the variance of income distribution explained by deregulation. While entrepreneurship may be part of the story of how deregulation affects the distribution of income, it plays a small role.

## 3. Deregulation, Labor Market Conditions, and the Distribution of Income

In this section, we explore alternative, labor market explanation for why branch deregulation tightened income distribution. As discussed in the Introduction, there are at least three interconnected mechanisms linking deregulation, labor market conditions, and income distribution. First, deregulation improves capital allocation, which boosts the marginal product of labor and wages. If the improvement in capital allocation is particularly beneficial to low-skilled workers, branch deregulation will reduce income inequality. Second, deregulation improves capital allocation and boosts wages, which draw more women into the labor force. If reservation wages do not change, deregulation will increase female labor force participation and reduce income inequality by pulling new women into the work force with higher reservation wages than previously working women. Third, deregulation improves capital allocation and intensifies competition, within banking and across non-financial firms. Greater competition will make discrimination more expensive and hence reduce the wage gap between discriminated and non-discriminated groups. These labor market explanations do not rely on the poor directly accessing financial services to purchase education or become entrepreneurs.

Consequently, we provide a richer, empirical picture of the labor market effects of branch deregulation. First, we decompose the impact of branch deregulation on the variance of income distribution into (i) the part accounted for by changes in the income gap between men and women, (ii) the part accounted for by changes in the variance among men, and (iii) the part accounted for changes in the variance among women. This is illustrated in Figure 3. As shown, 51% of the explained reduction in the variance of total income is due to a reduction of the income gap between men and women, 15% is due to a reduction of the variance of income across women, and 34% is accounted for by a drop in the variance of income across men. The same basic decomposition holds when excluding all forms of income except for wage income. Impressively, when focusing only on the wage income of full-time, full-year works of prime working age, the reduction in the gender income gap accounts for 100% of the drop in income inequality! Thus, while income inequality falls among men, and among women for some measures of income, a large proportion of the drop in income inequality from deregulation was due to a reduction in the gender income gap.

Second, deregulation boosts wages, especially women's wages. In Table 4, we present regressions based on individuals, rather than on the distribution of income within a state during a particular year. The dependent variable is a wage rate, computed as annual income divided by annual hours worked. The regressions controls for standard traits, education, education, marital status, etc. As shown, women's wage rates rise by 7% after deregulation, while branch deregulation boosts men's wages by less than 5%. Furthermore, the wage rates of both skilled (having at least some college education) and unskilled (having no college education) women rise by more than men's rates, and wage rates of skilled women rise by more than the wage rates of unskilled workers. Thus, deregulation reduced the wage rate gap between men and women

Third, deregulation boosted female labor force participation. We test whether deregulation increased the number of full-time, full-year FTFY females. Again, we control for state and year fixed effects. Thus, we are assessing whether deregulation pulled more women into the labor market relative to trend increases in female labor force participation. As reported in Table 5, deregulation increased FTFY women by about one percent above trend and state-fixed effects. Thus, the combined results reported in Figure 3, Table 4, and Table 5 suggest that deregulation pulls women into the labor force, reduces the gender income gap, and also closes the gender wage gap.

These results do not fully identify the mechanisms linking deregulation and the reduction in income inequality, but they do provide some guidance. As detailed above, deregulation did not reduce inequality primarily by increasing access to financial services. Rather, deregulation lowered income inequality primarily by affecting labor market conditions. Deregulation reduced the income gap between men and women by boosting the earnings of women and pulled women into the labor force.

## 4. Conclusions

Policymakers and economists disagree sharply about the impact of bank regulations on the distribution of income. In this paper, we tested whether liberalizing restrictions on intra-state branching in the states of the United States from the mid-1970s to the mid-1990s intensified, ameliorated, or had no effect on the Gini coefficient of income distribution. We find that bank deregulation reduced income inequality; it did not increase inequality as predicted by some policymakers and theories.

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 recommendation 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. Our works suggests that the education and entrepreneurship explanations account for very little of the reduction in income inequality generated by deregulation.

In contrast to key theories and policy initiatives by pivotal institutions, we show that deregulation lowered income inequality by affecting labor market conditions, not by providing the poor with greater access to financial services. Deregulation boosted the wages of women and increased women's labor force participation. Furthermore, most of the explained drop in income inequality is due to a reduction in the 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. Thus, for example, $Z \equiv \alpha X + (1-\alpha)Y$, where $\alpha$ is the share of income from salary and $1-\alpha$ is the share of income from non-salary sources, so that

(A1) $\Delta Var(Z) = \alpha^2\Delta Var(X) + (1-\alpha)^2\Delta Var(Y) + 2\alpha(1-\alpha)\Delta Cov(X,Y).$

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

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 within 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 $\Delta Var(Z)$, $\Delta Var(X)$, $\Delta Var(Y)$, and $Cov(X,Y)$ from four regressions. Specifically, controlling for state and year fixed effects, we regression 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 $\beta$, $\gamma$, $\delta$, and $\psi$ 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 $\Delta Var(Z) = \beta$, $\Delta Var(X) = \gamma$, $\Delta Var(Y) = \delta$, and $\Delta Cov(X,Y) = \psi$. Since we also know $\alpha$ 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 in 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:

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

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.

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Table1 -- Descriptive Statistics

| | Mean | Median | St. Dev. | Min. | Max. | Standard deviation: Cross states | Standard deviation: Within states | Standard deviation: Within state-years |
|---|---|---|---|---|---|---|---|---|
| Log(Gini): | | | | | | | | |
| Total household income | -.864 | -.866 | .071 | -1.123 | -.578 | .049 | .052 | .034 |
| Total individual income | -.618 | -.616 | .053 | -.775 | -.474 | .037 | .039 | .027 |
| Wage and salary income | -.804 | -.802 | .050 | -.963 | -.634 | .027 | .042 | .038 |
| Female | -.840 | -.839 | .060 | -1.058 | -.660 | .034 | .050 | .047 |
| Male | -.870 | -.867 | .079 | -1.146 | -.605 | .041 | .068 | .050 |
| Gross State Product growth | .023 | .024 | .039 | -.348 | .260 | .010 | .038 | .032 |
| Gov. taxes / Personal income | .068 | .064 | .021 | .026 | .334 | .018 | .012 | .011 |
| Gov. expenditure / Personal income | .140 | .131 | .049 | .068 | .457 | .046 | .016 | .012 |
| Proportion of college grads. | .386 | .387 | .093 | .153 | .608 | .053 | .077 | .019 |

Note: The sample is at the state-year level, for the years 1977-2003 and 49 states. South Dakota and Delaware are dropped. The year of deregulation is dropped as well. The resulting number of state-year observations is 1,287.

Table 2 -- The Effect of Deregulation on Household Income Inequality

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| Deregulation indicator | -.015 (.006)*** | -.016 (.005)*** | -.015 (.005)*** | -.014 (.005)*** | -.010 (.004)** |
| GSP growth | | -.056 (.043) | -.034 (.037) | -.036 (.037) | -.045 (.039) |
| Gov. taxes / Personal income | | .442 (.068)*** | .566 (.071)*** | .572 (.070)*** | .406 (.058)*** |
| Gov. expenditure / Personal income | | -.163 (.092)* | -.249 (.104)** | -.241 (.105)** | -.234 (.096)** |
| GSP growth 1-year lag | | | -.076 (.030)** | -.074 (.030)** | -.056 (.028)** |
| GSP growth 2-year lag | | | -.075 (.034)** | -.075 (.035)** | -.047 (.037) |
| College grads. | | | | -.072 (.065) | -.034 (.055) |
| Log(Gini) 1-year lag | | | | | .299 (.038)*** |
| R-squared | .77 | .77 | .77 | .77 | .79 |
| Num. obs | 1,287 | 1,287 | 1,287 | 1,287 | 1,239 |

Note: Standard errors (in parentheses) are adjusted for state-level clustering. The sample is at the state-year level, for the years 1977-2003 and 49 states (in the last column the sample is constrained to the years 1978-2003). South Dakota and Delaware are dropped. The year of deregulation is dropped as well. The dependent variable is log(Gini), which is based on total household income (in 1990 dollars) and calculated separately for each state and year. Households with non-positive total household income are dropped. Household sample weights are used in calculation. All models control for state- and year fixed effects. "*" indicates 10% significance; "**" - 5% significance; and "***" - 1% significance.

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

| | Dependent variable is log(Gini) of: | | | Dependent variable is log(variance) of: | | |
|---|---|---|---|---|---|---|
| | Total household income | Total individual income | Wage and salary income | Total household income | Total individual income | Wage and salary income |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Deregulation indicator | -.015 | -.018 | -.016 | -.092 | -.100 | -.084 |
| | (.005)*** | (.006)*** | (.005)*** | (.039)** | (.037)*** | (.036)** |
| GSP growth | -.034 | -.091 | -.083 | -.064 | -.255 | -.248 |
| | (.037) | (.033)*** | (.030)*** | (.200) | (.197) | (.183) |
| Gov. taxes / Personal income | .566 | .426 | .232 | 2.374 | 2.298 | 2.854 |
| | (.071)*** | (.089)*** | (.098)** | (.359)*** | (.399)*** | (.456)*** |
| Gov. expenditure / Personal income | -.249 | -.414 | -.256 | -.928 | -1.535 | -2.120 |
| | (.104)** | (.109)*** | (.112)** | (.539)* | (.644)** | (.708)*** |
| GSP growth 1-year lag | -.076 | -.092 | -.104 | .148 | .150 | .090 |
| | (.030)** | (.021)*** | (.018)*** | (.162) | (.173) | (.246) |
| GSP growth 2-year lag | -.075 | -.072 | -.136 | .491 | .557 | .515 |
| | (.034)** | (.022)*** | (.028)*** | (.317) | (.298)* | (.238)** |
| R-squared | .77 | .69 | .69 | .83 | .82 | .85 |
| Num. obs | 1,287 | 1,287 | 1,287 | 1,287 | 1,287 | 1,287 |

Note: Standard errors (in parentheses) are adjusted for state-level clustering. The first column is a replication of the third column in Table 2. The sample is at the state-year level, for the years 1977-2003 and 49 states. South Dakota and Delaware are dropped. The year of deregulation is dropped as well. The dependent variable in columns (1)-(3) is log(Gini) of the relevant type of income; the dependent variable in columns (4)-(6) is log(variance) of the relevant type of income. The Gini index and the variance of income (in 1990 dollars) are calculated separately for each state and year. Households with non-positive total household income are dropped. Household and personal sample weights are used in calculation. All models control for state- and year fixed effects. "*" indicates 10% significance; "**" - 5% significance; and "***" - 1% significance.

Table 4 -- The Effect of Deregulation on Log Hourly Wages
(Entries multiplied by 100)

| | Women | | | Men | | |
|---|---|---|---|---|---|---|
| | All | Skilled | Unskilled | All | Skilled | Unskilled |
| Deregulation indicator | 7.04 | 8.42 | 6.05 | 4.78 | 7.19 | 2.64 |
| | (2.23)*** | (2.59)*** | (2.20)*** | (2.03)** | (2.32)** | (2.21) |
| High school graduate | 28.66 | | 29.29 | 31.34 | | 32.01 |
| | (1.87)*** | | (1.75)*** | (2.38)*** | | (2.18)*** |
| Some college | 45.88 | | | 45.67 | | |
| | (2.29)*** | | | (3.00)*** | | |
| College graduate | 70.98 | 24.63 | | 71.10 | 25.03 | |
| | (2.20)*** | (0.48)*** | | (3.12)*** | (0.66)*** | |
| College + | 84.33 | 38.61 | | 80.07 | 34.86 | |
| | (2.05)*** | (1.05)*** | | (3.30)*** | (0.74)*** | |
| Experience | 1.83 | 2.19 | 1.40 | 3.38 | 3.99 | 2.74 |
| | (0.07)*** | (0.08)*** | (0.12)*** | (0.60)*** | (0.08)*** | (0.09)*** |
| Experience-squared | -.032 | -.044 | -.021 | -.053 | -.072 | -.038 |
| | (.002)*** | (.002)*** | (.002)*** | (.001)*** | (.002)*** | (.002)*** |
| Married | .052 | .612 | -.963 | 17.12 | 17.07 | 16.66 |
| | (.361) | (.491) | (.366) | (0.63)*** | (0.70)*** | (0.60)*** |
| R-squared | .22 | .13 | .07 | .24 | .16 | .16 |
| Num. obs | 314,663 | 174,265 | 140,398 | 475,610 | 258,852 | 216,757 |

Note: Standard errors (in parentheses) are adjusted for state-level clustering. The sample is at the individual level, for the years 1977-2003. "Skilled" is defined as having at least some college education; "unskilled" is defined as having high school diploma or being a high school drop-out. The sample is restricted to white prime-age (25-54) wage earners, not in institutions, not self-employed and not in armed forces, who work full-time-full-year, and have hourly wages (in 1990 dollars) of at least $1.90 (one half of the minimum wage in 1990), and with potential experience less than 39 years. Potential experience is defined as "age - years of education - 6". We drop households with non-positive total household income. Sample weights are used in all estimations. All models control for time trend. "*" indicates 10% significance; "**" - 5% significance; and "***" - 1% significance.

Table 5 -- The Effect of Deregulation on Participation* in the Labor Market
(Entries multiplied by 100)

| | Women | | Men | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| Deregulation | 1.27 | 1.05 | -0.03 | -0.24 |
| indicator | (0.43)*** | (0.43)** | (0.65) | (0.57) |
| GSP growth | | 5.06 | | 9.75 |
| | | (2.77)* | | (3.08)*** |
| Gov. taxes / | | 8.47 | | 11.71 |
| Personal income | | (7.39) | | (12.93) |
| Gov. expenditure / | | -9.17 | | -54.75 |
| Personal income | | (10.66) | | (12.17)*** |
| GSP growth | | 9.84 | | 14.47 |
| 1-year lag | | (2.62)*** | | (2.36)*** |
| GSP growth | | 7.73 | | 14.36 |
| 2-year lag | | (3.00)** | | (3.02)*** |
| R-squared | .88 | .89 | .66 | .69 |
| Num. obs | 1,287 | 1,287 | 1,287 | 1,287 |

Note: * Participation is defined as being employed at least 40 weeks per year and at least 35 hours per week. Standard errors (in parentheses) are adjusted for state-level clustering. The sample is at the state-year level, for the years 1977-2003 and 49 states and includes only whites. South Dakota and Delaware are dropped. The year of deregulation is dropped as well. The number of observation in all models is 1,287. All models control for state- and year fixed effects. "*" indicates 10% significance; "**" - 5% significance; and "***" - 1% significance.

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 |

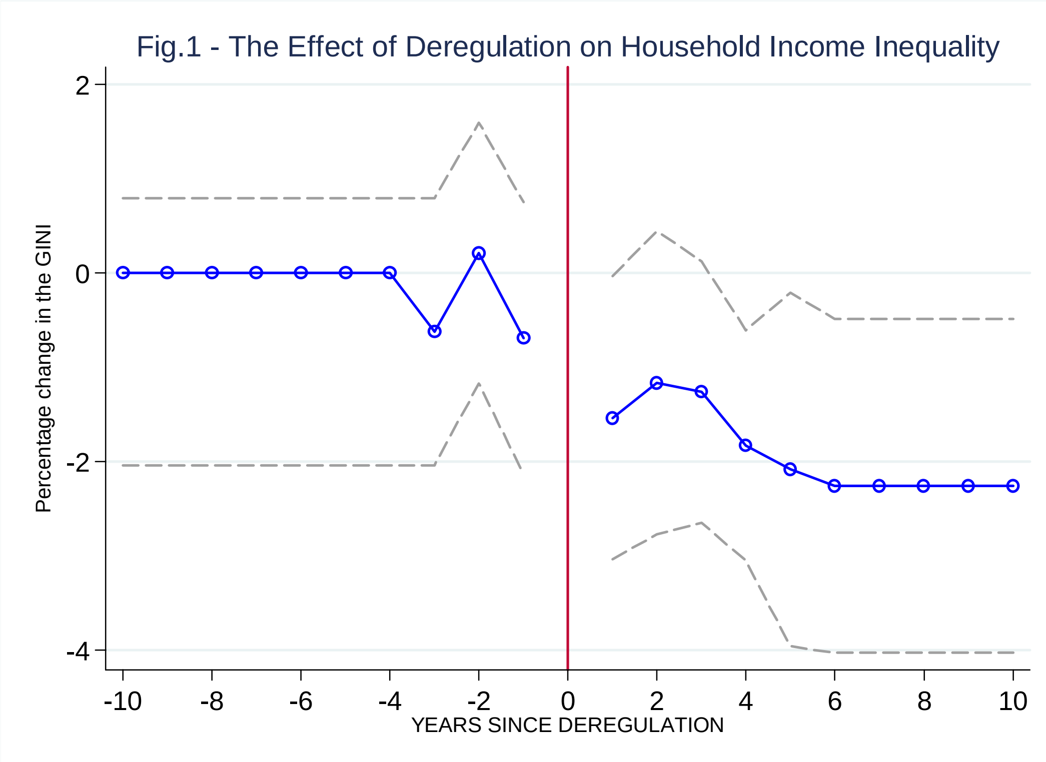

Fig.1 - The Effect of Deregulation on Household Income Inequality
Percentage change in the GINI
2
0
-2
-4
-10
-8
-6
-4
-2
0
2
4
6
8
10
YEARS SINCE DEREGULATION

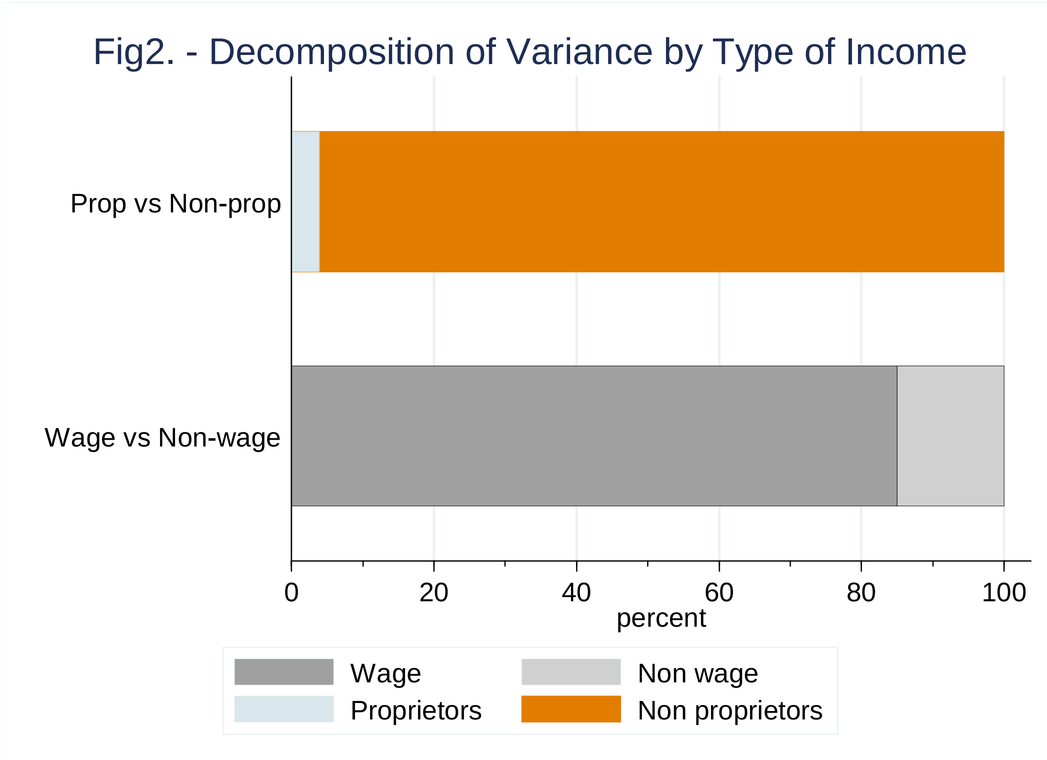
Fig2. - Decomposition of Variance by Type of Income
Prop vs Non-prop
Wage vs Non-wage
0
20
40
60
80
100
percent
Wage
Non wage
Proprietors
Non proprietors

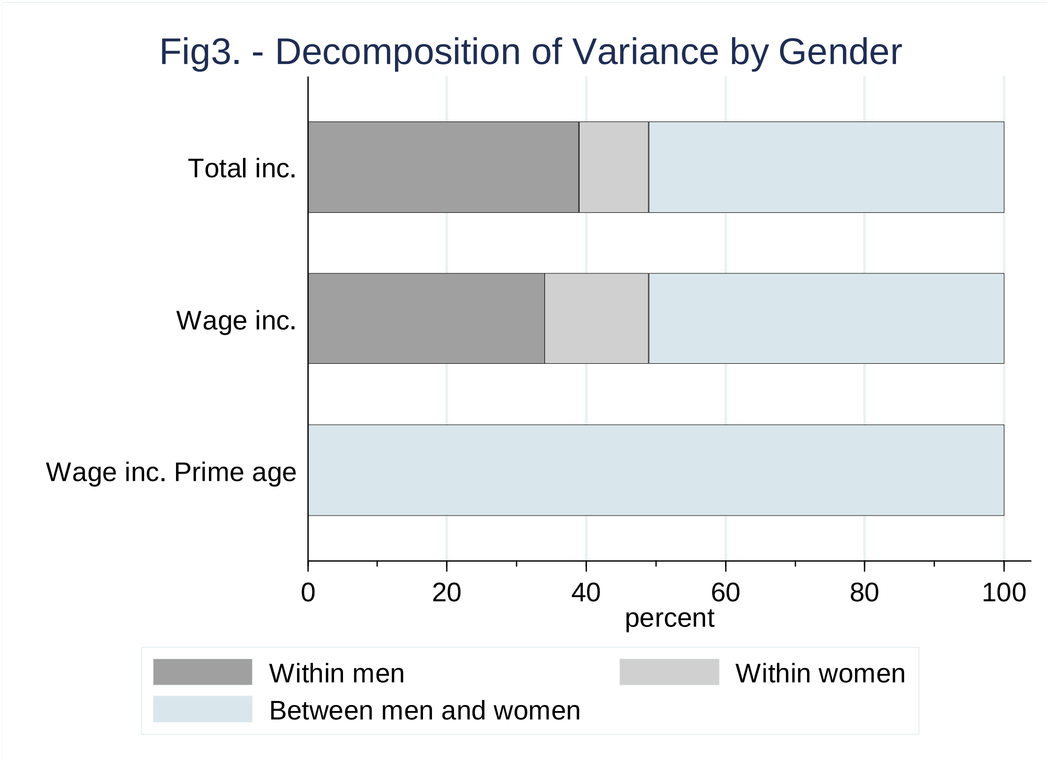
Fig3. - Decomposition of Variance by Gender
Total inc.
Wage inc.
Wage inc. Prime age
0
20
40
60
80
100
percent
Within men
Within women
Between men and women