EMPLOYER CONCENTRATION AND LABOR FORCE PARTICIPATION

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Abstract

This paper examines the association between employer concentration and labor outcomes (labor force participation and employment). It uses restricted data from the U.S. Census Bureau's Longitudinal Business Database to estimate, at the county level, to what extent more concentrated labor markets have lower labor force participation rates and lower employment. The analysis also examines whether unionization rates and education levels mediate these associations.

The paper found that:

- Labor force participation is strongly negatively correlated with employer concentration.
- Employment is weakly negatively associated with employer concentration.
- The relationship between concentration and labor outcomes is weaker where union coverage is higher, while education does not play a major role in moderating these associations.

The policy implications of the findings are:

- Places where labor markets are concentrated might have a greater prevalence of low eligibility for Social Security benefits due to lower participation rates.
- Antitrust policy applied to the labor market might improve future Social Security benefits for individuals in concentrated labor markets.

Introduction

Even prior to the COVID-19 pandemic, the labor-force participation (LFP) of prime-age workers had been declining steadily over the past two decades. Individuals who do not work lose access to a wide array of benefits, ranging from Medicaid (in some states) to the accrual of credits towards Old Age, Survivors, and Disability Insurance. Simultaneously, an accumulation of evidence suggests that when few employers employ increasingly larger shares of the local workforce – a rise in the concentration of employers – firms possess greater bargaining power in employment negotiations, potentially driving down wages and LFP. The evidence has begun to filter through to policy, with a recent presidential executive order instructing the Federal Trade Commission to consider labor-market concentration, in addition to product-market concentration, when evaluating mergers.[1] This paper examines whether markets with higher employer concentration are associated with lower LFP and whether the relationship is weaker for employees with more bargaining power, such as those covered by unions.

The analysis in this paper fills in a missing link between employer concentration and lower wages by directly estimating the correlations between concentration and employment, and between concentration and LFP. Using restricted data from the Census Bureau's *Longitudinal Business Database* (LBD), a Herfindahl-Hirschman Index (HHI) of employment concentration is calculated at the county level. This index is then used in an OLS regression that uses employment or LFP (from the LBD) as the independent variables. In addition, the analysis explores whether the correlation of concentration and labor-market outcomes is weaker when workers have bargaining power to counteract employer bargaining power, by estimating the interaction effect of concentration and union coverage. To also explore whether higher-education workers have a better negotiating position, the analysis interacts concentration with educational structure of the population. Both these latter variables are acquired from the *Current Population Survey* (CPS).

The rest of the paper is organized as follows. Section 2 describes the state of the literature on employer monopsony. Section 3 describes the data and methods for the analysis. Section 4 presents the results. The final section concludes that employer concentration is

[1] See Irwin (2021). See also U.S. Department of the Treasury (2022) for a review of relevant literature and policy on this issue.

negatively associated with employment and LFP, while union coverage mitigates this effect and education has no clear relationship with the effect of concentration.

Background

From its recent peak in 2000 to its trough in 2015, the LFP of prime-age workers dropped from 84 percent to 80.9 percent (Federal Reserve Bank of St. Louis 2021).[2] Prime-age LFP recovered slightly prior to the COVID-19 pandemic in 2020, but never returned to the levels of the late 1990s. A large literature has examined potential drivers of this decline, including trade with China, the emergence of robots, changes in the minimum wage, the age-structure of the population, and changing preferences for leisure activities. Thus far, no consensus has emerged with respect to the importance of these and other factors (for example, see Abraham and Kearney 2018 or Krueger 2017).[3]

At the same time, a growing body of evidence suggests that employer bargaining power has increased in U.S. labor markets, resulting in lower wages (Azar, Marinescu, and Steinbaum 2022 (forthcoming); and Benmelech, Bergman, and Kim 2022 (forthcoming)). In theory, lower wages come about in a local "monopsony" because employers hire fewer workers than in a competitive market. When the marginal cost of workers is increasing, monopsonistic employers hire fewer workers than do competitive ones: instead of paying the high marginal cost of labor for the last of many employees hired in a competitive market, monopsonistic employers can pay a wage equal to the lower marginal cost of the last employee of a smaller group hired to maximize profit. This dynamic is especially strong when workers have little negotiating power, as do those without representation by labor unions.

Lower wages would, in turn, lead to lower LFP among workers whose attachment to the labor market is weak. For example, Katz and Notowidigdo (2017) show that if wages in the bottom two quintiles of the income distribution had grown proportionally to productivity, as in a

[2] The secular drop in prime-age LFP during this period affected both men and women. While 3.1 percentage points of LFP may not seem large, consider that between the peak and trough surrounding the Great Recession, prime-age LFP fell by only 0.7 percentage points. The decline in LFP from February 2020, before the COVID-19 recession, to its nadir in April 2020 was itself only 3.1 percentage points, a decline which quickly shrank to just 1.4 percentage points by June 2020.

[3] Abraham and Kearney (2018) look at employment, a different measure than LFP, and conclude that declining demand for U.S. labor, driven by trade with China and the automation of routine tasks, is the clearest explanation for the recent reduction. Krueger (2017) presents correlational evidence that health and opioid abuse account for a large share of the decline in LFP in the prime-age population; however, he acknowledges concerns of reverse causality.

competitive market, then the LFP of prime-age men would be as high today as it was in 1980. However, the direct link between employer concentration, employment, and LFP has not been studied.

The labor market effects of employer concentration have not received much attention historically, but this topic has gained traction among economists in recent years. Early models of monopsony envisioned a small company town characterized by a single employer facing an upward-sloping supply curve for labor (for a review, see Ashenfelter, Farber, and Ransom 2010). The key insight from these static models – parallel to the standard view of monopoly in the product market – is that the employer hires fewer workers than would have been employed in a perfectly competitive labor market and pays them a wage below their marginal revenue product. The resulting "wage gap" is inversely proportional to the elasticity of labor supply, implying that the employer has more bargaining power when workers have a strong desire to work. Although company towns are increasingly rare, empirical studies have shown that certain labor markets behave as if employers face little competition, such as the markets for minimum-wage workers, nurses, teachers, and even software engineers in Silicon Valley (Belman and Wolfson 2014; Council of Economic Advisors 2016, Merrifield 1999; Ransom and Sims 2010; Staiger, Spetz, and Phibbs 2010; Quinby and Wettstein 2022 (forthcoming)).

In an attempt to move away from the company-town model of monopsony, a more recent line of thought considers how job-search costs and differentiated human capital give employers bargaining power even in seemingly competitive labor markets. Similar to the static framework, these dynamic models relate employer bargaining power to the firm-specific elasticity of labor supply (Burdett and Mortensen 1998; Bhaskar, Manning, and To 2002; Manning 2003). Supporting this hypothesis, a number of empirical studies estimate relatively low but positive elasticities of labor supply for prime-age workers across a range of settings (on the order of 0.3, see Chetty et al. 2011; and Peterman 2016 for reviews of this literature). A recent body of research also reveals that the cost of searching for jobs remains substantial even in the age of the Internet (Kuhn and Mansour 2011; Cardoso, Loviglio and Piemontese 2016).

The static and dynamic monopsony models are conceptually similar, particularly in a world where prime-age workers are increasingly unlikely to move geographically (Moretti 2011; and Marinescu and Rathelot 2018). Workers with industry-specific human capital may have many job prospects nationally but few in the commuting zone where they currently live. Until recently, however, labor economists had no direct measure of employer concentration that could be used to test the monopsony hypothesis. Consequently, they relied on calibrations of a highly stylized macroeconomic model; one such estimate places the overall employment and GDP loss from monopsony around 13 percent, and the decline in labor's share of GDP around 22 percent (Naidu, Posner, and Weyl 2012 (forthcoming)).

A breakthrough arrived when Azar, Marinescu, and Steinbaum (2022, forthcoming) and Benmelech, Bergman, and Kim (2022, forthcoming) independently developed indices of labor market concentration at the local level similar to indices of product market concentration. They find that the typical local labor market is highly concentrated according to the DOJ-FTC guidelines used to evaluate mergers and acquisitions in the product market, although the degree of concentration varies considerably across industries and localities. As predicted, the variation in labor market concentration is also strongly associated with wage levels. Similar conclusions were reached by Arnold (2020) and Prager and Schmitt (2021), who found that mergers lead to wage declines when taking a difference-in-differences approach. Consistent with slow wage growth being driven by increasing labor-market concentration, a number of studies also find that the effect of employer concentration is greater in areas where workers are less unionized (Arnold 2020; Benmelech, Bergman, and Kim 2022 (forthcoming); Prager and Schmitt 2021).

Furthermore, product market concentration in the U.S. corporate sector has been growing over time (e.g., Autor et al. 2017).[4] While this trend does not necessarily indicate increasing labor market concentration, it is consistent with it.[5] As pointed out by Naidu, Posner, and Weyl (2022, forthcoming), under the regulatory framework of anti-trust in the United States, firms that can demonstrate that their market power does not result in higher consumer prices can avoid anti-trust action. In this setting, a possible alternative for such firms to increase pure profits is to reduce labor costs instead of raising product prices.

[4] Whether labor-market concentration itself has also been growing over time is controversial. Rinz (2018), Berger et al. (2019), Grossman and Oberfield (2021), and Rossi-Hansberg et al. (2021) present evidence of declining labor-market concentration, while Benmelech, Bergman, and Kim (2022, forthcoming) show slightly increasing concentration in recent decades.

[5] Other evidence for the theory includes the slowing of real wage growth, especially at lower education levels (Goldin and Katz 2008; Acemoglu and Autor 2011), the weakening of the link between productivity and wage growth (Mishel 2012; Bivens and Mishel 2015; Ugoccioni 2016; Stansbury and Summers 2017), and the decline of the national labor share of income since the early 2000s (Karabarbounis and Neiman 2014).

Data and Methods

The established link between employer concentration and lower wages is theoretically predicated on employment being depressed (relative to a perfectly competitive benchmark) in highly concentrated labor markets. To test for this effect, the analysis proceeds in two steps. First, local employer concentration is indexed, and the values of the index are estimated. Second, the association of the index with employment levels and LFP is estimated. Finally, the analysis checks whether the relationship between concentration and labor market outcomes is attenuated when workers possess bargaining power.

Data

The typical measure of market concentration is the HHI, which varies between 0 (extremely diffuse) and 1 (a monopsony). To apply this index at a local geographic level, the analysis uses data on firms and their employment at the county level. This index must then also be linked to county characteristics, in particular employment and LFP.

The analysis relies on three datasets to link employer concentration at the county level to LFP. First, the HHI index is defined using the restricted LBD maintained by the U.S. Census Bureau. The data span the years 1995-2013, and provide the most comprehensive information on firm-level employment within counties.[6]

Data on county LFP is acquired for years 1995-2013 from the CPS. County population and demographic characteristics – such as education levels and unionization rates – are also obtained from the CPS.

The measure of concentration $M_{c,t}$ in each county, *c*, at each year, *t*, is defined in two steps, following the approach in Benmelech, Bergman, and Kim (2022, forthcoming). First, $HHI_{c,i,t}$ is calculated at the county-industry-year level.[7] This index measures concentration based on the share of each firm j's employment in county *c*, industry *i*, and year *t*, $s_{j,c,i,t}$, by calculating:

[6] In fact, the data also cover 1990-1994; however, those years are not included in the analysis because many county definitions are inconsistent between this early period and later years. The 2013 end point was chosen for comparability with past work, but can be expanded to include more recent years in future research.

[7] This measure is calculated at the industry level, as the skills of workers in one industry may not transfer well to another. Industries are defined at the 3- and 4-digit NAICS level, and results are assessed for both classifications.

$$HHI_{c,i,t} = \sum_j s_{j,c,i,t}^2.$$

In the second step, this industry-specific index is aggregated to the county-year level, weighted by industry shares of employment in the county at the start of the sample:

$$M_{c,t} = \sum_i I_{c,i,1} * HHI_{c,i,t}\,.$$

$M_{c,t}$ is thus a measure of the average employer concentration in the county-year, and is the key independent variable in the analysis.

The main dependent variables of interest are employment/population ratios and LFP. Both are acquired from the CPS. The analysis considers four different age segments of the population for these outcomes: the full population, young workers (18-29), prime-age workers (30-54), and older workers (55 and over). The employment ratios take as their denominator the population of interest. Finally, a few different county-level control variables, such as the racial and education makeup of the population, are also based on the CPS.

To further test the theory that the correlation of employer concentration and labor outcomes is due to *employer* bargaining power, the analysis also assesses whether the relationship is mediated by *employee* bargaining power by using union coverage as a proxy for workers' ability to negotiate effectively.[8] Union coverage rates at the county-year level are also estimated from the CPS, where all workers covered by collective bargaining are counted as "unionized".

Table 1 displays the summary statistics of the main variables included in the analysis. Regarding the key outcome variables, LFP is about 65 percent over the entire sample, and higher for men than for women (72 percent versus 58 percent). Employment, conditional on being in the labor force, is 94 percent for both genders.

With respect to concentration, the HHI is 0.17 when defined based on 3-digit NAICS codes, and 0.24 when using 4-digit codes. Notably, these numbers are very similar to those in Rinz (2018), but much lower than those in Benmelech, Bergman, and Kim (2022, forthcoming). Rinz uses NAICS codes, as this analysis does, while Benmelech et al. use SIC codes; the difference likely stems from this choice. Given the lower estimate of concentration using the 3-digit definition, the rest of the discussion will focus on the 3-digit definition to be conservative;

[8] This measure includes both union members and non-members covered by collective bargaining agreements.

results using the 4-digit definition are the Appendix A and are qualitatively very similar to the results in the main text.

Methods

With measures of employer concentration and labor-market outcomes in hand, the correlation of the two is estimated using the following OLS equation:

$$L_{c,t} = \alpha + \beta M_{c,t} + \gamma X_{c,t} + \varepsilon_{c,t}.$$

$L_{c,t}$ is either the labor force participation rate or the employment ratio (among those in the labor force) within a county during a given year. The hypothesis is that β would be negative, i.e., a high degree of employer concentration is associated with lower LFP and employment rates.[9] Alternative specifications consider male and female labor outcomes separately, since the two groups display divergent time trends. $X_{c,t}$ is a vector of state and year fixed effects, and demographic controls.

Conceptually, employees in a better negotiating position – such as those covered by collective bargaining through a union – should be less affected by the increase in employer concentration (Goldin and Katz 2008).[10] The analysis explores whether employee bargaining power counteracts employer concentration with two heterogeneity tests. First, it examines whether unions mitigate the effect of a dominant employer by giving workers more bargaining power. A positive coefficient on the interaction term would be consistent with a causal relationship between employer concentration and low employment and LFP.

Second, the analysis explores whether the relationship between employer concentration and labor outcomes differs across education groups by interacting $M_{c,t}$ with the percentage of workers in each county falling into broad education categories (no high school, high school, some college, college degree, graduate degree). The hypothesis is that high levels of education will attenuate the negative correlation between employer concentration and labor outcomes under the assumption that highly educated workers have more bargaining power because their skills are in greater demand. However, this assumption is less clear-cut than with regards to

[9] Observations will be weighted by the population in each county in 1995, since migration may be endogenous to labor market conditions in a county (as shown, for example, in Blanchard and Katz 1992). Standard errors in this regression are clustered at the county level to account for serial correlation within county over time.

[10] Although union coverage rates are low in the general population, they have varied over time (from 16.6 percent of workers in 1995 to 12.4 percent in 2013) and across regions in 2013 (from 39.5 percent of workers in Bronx County, NY, to almost none in Anderson County, SC). See Hirsch and Macpherson (2021).

unionization, since highly specialized workers may actually face a less competitive labor demand, as only a small number of employers make use of their skills.

Results

The analysis produces four broad categories of results, described below. The first and second are for the association between employer concentration and the two labor market outcomes – employment and LFP. A variety of specifications are presented for each outcome. The third and fourth subsections explain how these associations are modulated by union coverage and education.

Association Between Employer Concentration and Employment

Table 2 shows the results of the regressions with the employment rate as the dependent variable.[11] Column 1 has no controls besides year and state fixed effects; Column 2 adds controls for the racial composition and age structure of counties; and Column 3 adds in education controls in anticipation of the heterogeneity analysis below. Columns 4 and 5, respectively, display results for men and women separately.

Broadly, the estimates are consistent with a negative association between employment and employer concentration. However, only some of the specifications are statistically significant; in particular, once education is controlled for the association between concentration and employment becomes insignificant. A weak association between concentration and employment echoes recent causal estimates in the hospital sector (Prager and Schmitt 2021). Notably, when breaking results out by age group, only young workers display a significant association of employment and concentration (even with all the controls). For these workers, a move from perfect competition (HHI of 0) to monopsony (an HHI of 1) is associated with a 5-precentage point decline in LFP (Table 3, column 1; $p<0.01$). In contrast, prime age and older workers' employment is uncorrelated with HHI.

Given the large recent literature showing that concentration leads to wage declines, the finding that employment is only tenuously related to concentration gives some indication that the elasticity of labor supply is small. This conclusion, too, is supported by a large quantity of evidence (see review in Chetty et al. 2011).

[11] Appendix Table 1 shows these estimates using the 4-digit NAICS definition of concentration.

However, even when only a weak association between concentration and employment prevails, employer monopsony can still lead to substantial effects on LFP because the effects of concentration on employment and on LFP concern different marginal workers. In particular, the effect of concentration on LFP is one of workers marginally attached to the labor market.

Association Between Employer Concentration and LFP

Indeed, we do find large and significant negative correlations between employer concentration and LFP. Table 4 shows these estimates in various specifications, following the same convention as Table 2.[12] For the specification with all the demographic controls, in Column 3, the estimate implies that going from perfect competition (infinite fragmentation of employers) to a monopsony is associated with a 4-percentage point decline in LFP ($p<0.05$). Scaling the effect by the standard deviation of HHI, 0.0847 (see Table 1), implies that a one-standard-deviation increase in concentration is associated with a 0.37-percentage point decline in LFP.

The results are qualitatively robust across the different specifications, and indicate that places where employers are more concentrated tend to have lower LFP. This pattern holds in a variety of different cuts of the population: both for men and women (columns 4 and 5 of Table 4, respectively) and among young, prime-age, and older workers (Table 5). Unlike with respect to employment, the age pattern of the LFP association is inverted-U shaped, with the weakest association for prime-age workers, and a similarly negative relationship between HHI and LFP for the youngest and oldest workers.

How can we reconcile a large association between concentration and LFP with a weak association between concentration and employment? The effect must be driven by a decline in unemployment in highly-concentrated labor markets. One theory could be search costs: such costs might be high when there are numerous potential employers, but low when there are only a few relevant employment options. Consider a labor market where employers consolidate: the market would initially see a high unemployment rate as many workers search for jobs; such workers are not employed but are still in the labor force. After the consolidation, only few potential workers are unemployed at any given time because the initially unemployed either promptly find work at one of the few employers left, or they eschew bearing search costs,

[12] Appendix Table A2 shows the parallel results using 4-digit NAICS codes.

knowing the wages (which are relatively low in the absence of competition among employers) are not worth the effort of searching for a better job which simply does not exist.[13]

Taken together, the weak associations between employer concentration and employment and the strong association with LFP suggest that the wage effects of concentration estimated in the past serve to push marginally attached workers out of the labor force entirely, while many of those with lower elasticity of labor remain employed, but have their surplus appropriated by their employer.

Interaction with Union Coverage

Thus far, the facts are consistent with employer concentration conferring on employers a greater ability to reduce wages, leading to some decline in employment, and to a more substantial decline in LFP through the departure of marginal workers from the labor force altogether. Those who remain employed receive lower wages than they would in a more diffuse labor market and the difference is captured by employer rents.

However, other interpretations are also possible. For example, places with low LFP might have poor economic prospects, and thus few employers locate there. The interpretation of the facts as evidence of employer bargaining power would be bolstered if, when workers had more bargaining power, additional employer concentration were less impactful.

Such attenuation of the association between employer concentration and LFP is what we find, echoing the results on wages in Benmelech, Bergman, and Kim (2022, forthcoming).[14] Table 6 shows that at the mean of HHI, 0.1655 (see Table 1), for every additional percentage point of workers covered by collective bargaining, the association between concentration and LFP is smaller by (0.998*0.1655)/100=0.17 percentage points ($p<0.01$).[15]

[13] A similar hypothesis on search frictions is advanced by Prager and Schmitt (2021) to explain their null findings of the effect of monopsony on employment, even as they find a negative effect on wage growth. The findings in this analysis further flesh out the mechanism behind this theory.

[14] The results with respect to employment are in Appendix Table A3.

[15] The main effect of union coverage on LFP is, itself, negative, with a one percentage point increase in unionization associated with a decline of 0.2 percentage points in LFP at an HHI of 0. This pattern, too, is consistent with a simple model of supply and demand for labor.

Interaction with Education

At first glance, education might seem to operate much like unionization in terms of attenuating employer bargaining power. That is, one might anticipate that more educated workers would have more ability to negotiate their wages with employers. However, unlike the case of union coverage, education also involves specialization which can serve to further restrict the set of relevant employers. For example, while a Ph.D. in economics is a relatively rare qualification that employers might compete to hire, the set of employers willing to engage in such competition is relatively small. Similarly, the market for software engineers appears highly concentrated, as is the market for schoolteachers (Council of Economic Advisors 2016; and Quinby and Wettstein, 2022 (forthcoming)).

As a result of this ambiguity, it is not *a priori* obvious that any pattern should emerge with respect to the association of employer concentration and education. In fact, no such pattern is apparent (Table 6). None of the interactions of education with employer concentration are statistically significantly different from zero.[16] Thus, this analysis offers no evidence that the education level of employees affects their relative bargaining power.

Conclusion

Recently documented declines in wages when employer concentration increases have led to a presumption of declines in employment in concentrated markets, consistent with a monopsonistic model of labor demand. However, this relationship had not been explicitly documented. Neither had the possible negative relationship of concentration with LFP, as marginally attached workers leave the labor force when confronted with lower wages. Both employment and LFP are important indicators in themselves, as they impact economic policymaking at a societal level and determine eligibility for various benefits, such as disability and old-age insurance, on the individual level.

This paper directly analyzed the relationship between employer concentration and employment ratios and LFP. The main findings show a weak negative relationship between

[16] Some of the interactions are significant when the outcome is employment, rather than LFP (Appendix Table A3). However, the pattern of significant results still does not suggest that more education improves worker bargaining power, with attenuation in the association of HHI occurring for those with high school education and college, but not for those with some college or graduate degrees. This pattern suggests that, at least, the modulation of HHI's association with employment is not linear in education.

employer concentration and employment, and a more robust negative relationship with LFP. Furthermore, the analysis supported the interpretation of the results as evidence of employer bargaining power by finding that the negative association of employer concentration and LFP is attenuated when workers have their own bargaining power through unions. One possible mechanism (consistent with the divergent results on LFP and employment) is that unions drive up wages, making more non-employed workers more willing to bear job-search costs, leaving them in the pool of unemployed, versus out of the labor force.

The implications of this analysis reinforce some of the conclusions of past work on employer bargaining power and wages. The results point toward noncompetitive labor markets as a real phenomenon, providing an explanation for the prevalence of measures to correct market failures, such as minimum wages and unionization. Other policy levers that have been suggested in the past, such as application of anti-trust regulation to the labor market, are also possible responses to concentrated labor markets. The impact of employer concentration not only on wages but also on employment and LFP raises the question of whether places and industries that are more concentrated affect the eligibility of workers in those sectors for Social Security and other work-contingent benefits.

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Tables

Table 1. *Summary Statistics*

	Mean	S.D.
In labor force	0.6487	0.0434
In labor force, men	0.7189	0.0461
In labor force, women	0.5833	0.0461
Employed	0.9381	0.0238
Employed, men	0.9361	0.0271
Employed, women	0.9403	0.0226
HHI (3-digit NAICS)	0.1655	0.0847
HHI (4-digit NAICS)	0.2398	0.1043
White	0.6779	0.1863
Black	0.1175	0.0981
Hispanic	0.142	0.1468
Asian	0.0464	0.0618
Other race	0.0162	0.0239
Share young (18-29)	0.1621	0.0196
Share prime age (30-54)	0.3569	0.0236
Share older (55+)	0.2271	0.0415
Covered by collective bargaining	0.1483	0.0702
Less than high school	0.3667	0.053
High school	0.2415	0.0473
Some college	0.2044	0.0301
College	0.1247	0.032
Graduate degree	0.0627	0.0247

Sources: Authors' calculations from the 1995-2013 LBD and CPS.

Table 2. *Regressions of Employment Rates, with HHI Defined on 3-Digit NAICS Codes*

Variables	(1) No controls	(2) No education controls	(3) Base	(4) Male	(5) Female
HHI	-0.0255**	-0.0342***	-0.0084	-0.00334	-0.0152
	(0.0120)	(0.0100)	(0.00943)	(0.0104)	(0.0102)
High school			0.0479	0.0212	0.0806**
			(0.0302)	(0.0315)	(0.0318)
Some college			0.0377	0.0212	0.0564**
			(0.0253)	(0.0271)	(0.0271)
College			0.175***	0.177***	0.176***
			(0.0258)	(0.0306)	(0.0256)
Graduate degree			0.0497	0.0431	0.0577
			(0.0385)	(0.0432)	(0.0396)
Observations	4,100	4,100	4,100	4,100	4,100
R-squared	0.669	0.745	0.761	0.738	0.678
Age	No	Yes	Yes	Yes	Yes
Race	No	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
State FE	Yes	Yes	Yes	Yes	Yes

Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.05.
Sources: Authors' calculations from the 1995-2013 LBD and CPS.

Table 3. *Regressions of Employment Rates by Age, with HHI Defined on 3-Digit NAICS Codes*

Variables	(1) Young (18-29)	(2) Prime age (30-54)	(3) Older (55+)
HHI	-0.0509***	-0.00805	0.00662
	(0.0167)	(0.00872)	(0.00826)
High school	0.0286	0.0709***	0.0213*
	(0.0279)	(0.0268)	(0.0124)
Some college	0.0660***	0.0450**	0.0335**
	(0.0235)	(0.0221)	(0.0150)
College	0.163***	0.123***	0.0359**
	(0.0258)	(0.0227)	(0.0155)
Graduate degree	0.169***	0.0747***	0.0651***
	(0.0431)	(0.0263)	(0.0145)
Observations	4,100	4,100	4,100
R-squared	0.643	0.742	0.548
Age	No	No	No
Race	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
State FE	Yes	Yes	Yes

Note: *** p<0.01, ** p<0.05, * p<0.1.
Source: Authors' calculations from the 1995-2013 LBD and CPS.

Table 4. *Regressions of Labor Force Participation, with HHI Defined on 3-Digit NAICS Codes*

Variables	(1)	(2)	(3)	(4)	(5)
	No controls	No education controls	Base	Male	Female
HHI	-0.181***	-0.104***	-0.0436**	-0.0664***	-0.021
	(0.0362)	(0.0216)	(0.0175)	(0.0217)	(0.0168)
High school			0.344***	0.220***	0.416***
			(0.0476)	(0.0484)	(0.0526)
Some college			0.277***	0.114**	0.420***
			(0.0505)	(0.0524)	(0.0539)
College			0.440***	0.387***	0.460***
			(0.0414)	(0.0490)	(0.0489)
Graduate degree			0.316***	0.191***	0.428***
			(0.0669)	(0.0710)	(0.0726)
Observations	4,100	4,100	4,100	4,100	4,100
R-squared	0.573	0.796	0.826	0.796	0.779
Age	No	Yes	Yes	Yes	Yes
Race	No	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
State FE	Yes	Yes	Yes	Yes	Yes

Note: *** p<0.01, ** p<0.05.
Source: Authors' calculations from the 1995-2013 LBD and CPS.

Table 5. *Regressions of Labor Force Participation by Age, with HHI Defined on 3-Digit NAICS Codes*

Variables	(1) Young (18-29)	(2) Prime age (30-54)	(3) Older (55+)
HHI	-0.105***	-0.0514***	-0.0991***
	(0.0300)	(0.0195)	(0.0316)
High school	0.159***	0.221***	0.0933**
	(0.0342)	(0.0429)	(0.0403)
Some college	0.00414	0.193***	0.258***
	(0.0319)	(0.0458)	(0.0514)
College	0.171***	0.284***	0.353***
	(0.0374)	(0.0380)	(0.0391)
Graduate degree	0.157**	0.222***	0.409***
	(0.0716)	(0.0429)	(0.0547)
Observations	4,100	4,100	4,100
R-squared	0.638	0.665	0.695
Age	No	No	No
Race	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
State FE	Yes	Yes	Yes

Note: *** p<0.01, ** p<0.05.
Source: Authors' calculations from the 1995-2013 LBD and CPS.

Table 6. *Heterogeneity in the Association of Labor Force Participation and HHI by Unionization and Education, with HHI Defined on 3-Digit NAICS Codes*

Variables	(1) Including union interaction	(2) Including education interaction
HHI	-0.186***	-0.242
	(0.0296)	(0.148)
HS	0.344***	0.246***
	(0.0424)	(0.0701)
Some college	0.267***	0.285***
	(0.0483)	(0.0849)
College	0.393***	0.333***
	(0.0398)	(0.0876)
Graduate degree	0.300***	0.414***
	(0.0672)	(0.124)
HHI x HS		0.612
		(0.383)
HHI x Some college		0.0199
		(0.388)
HHI x College		0.792
		(0.585)
HHI x Graduate degree		-0.829
		(0.729)
Union coverage	-0.224***	
	(0.0357)	
HHI x Union coverage	0.998***	
	(0.181)	
Observations	4,100	4,100
R-squared	0.834	0.827
Age	Yes	Yes
Race	Yes	Yes
Year FE	Yes	Yes
State FE	Yes	Yes

Note: *** $p<0.01$.
Source: Authors' calculations from the 1995-2013 LBD and CPS.

Appendix A. Supplementary Tables

Table A1. *Regressions of Employment Rates, with HHI Defined on 4-Digit NAICS Codes*

Variables	(1) No controls	(2) No education controls	(3) Base	(4) Male	(5) Female
HHI	-0.0174	-0.0353***	-0.0112	-0.00486	-0.0192**
	(0.0112)	(0.00785)	(0.00842)	(0.00948)	(0.00890)
High school			0.047	0.0213	0.0786**
			(0.0299)	(0.0313)	(0.0314)
Some college			0.0362	0.02	0.0544**
			(0.0246)	(0.0267)	(0.0259)
College			0.168***	0.174***	0.165***
			(0.0264)	(0.0315)	(0.0260)
Graduate degree			0.0499	0.0433	0.0576
			(0.0382)	(0.0433)	(0.0387)
Observations	4,100	4,100	4,100	4,100	4,100
R-squared	0.668	0.747	0.761	0.738	0.679
Age	No	Yes	Yes	Yes	Yes
Race	No	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
State FE	Yes	Yes	Yes	Yes	Yes

Note: *** p<0.01, ** p<0.05.
Source: Authors' calculations from the 1995-2013 LBD and CPS.

Table A2. *Regressions of Labor Force Participation, with HHI Defined on 4-Digit NAICS Codes*

Variables	(1) No controls	(2) No education controls	(3) Base	(4) Male	(5) Female
HHI	-0.135***	-0.0901***	-0.0353**	-0.0596***	-0.0104
	(0.0341)	(0.0163)	(0.0158)	(0.0181)	(0.0166)
HS			0.347***	0.223***	0.433***
			(0.0477)	(0.0481)	(0.0534)
Some college			0.282***	0.120**	0.429***
			(0.0518)	(0.0531)	(0.0558)
College			0.437***	0.375***	0.479***
			(0.0411)	(0.0483)	(0.0513)
Graduate degree			0.319***	0.195***	0.428***
			(0.0671)	(0.0710)	(0.0737)
Observations	4,100	4,100	4,100	4,100	4,100
R-squared	0.566	0.796	0.826	0.796	0.777
Age	No	Yes	Yes	Yes	Yes
Race	No	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
State FE	Yes	Yes	Yes	Yes	Yes

Notes: Robust standard errors in parentheses. *** p<0.01, ** p<0.5.
Source: Authors' calculations from the 1995-2013 LBD and CPS.

Table A3. *Heterogeneity in the Association of Employment and HHI by Unionization and Education, with HHI Defined on 3-Digit NAICS Codes*

Variables	(1) Including union interaction	(2) Including education interaction
HHI	-0.000773	-0.295***
	(0.0174)	(0.0894)
HS	0.0548*	-0.0635
	(0.0286)	(0.0498)
Some college	0.0430*	0.00314
	(0.0245)	(0.0407)
College	0.167***	0.107**
	(0.0258)	(0.0507)
Graduate degree	0.0517	0.0814
	(0.0372)	(0.0770)
HHI x HS		0.720***
		(0.204)
HHI x Some college		0.341
		(0.215)
HHI x College		0.510**
		(0.253)
HHI x Graduate degree		-0.306
		(0.436)
Union coverage	-0.0271	
	(0.0282)	
HHI x Union coverage	-0.0608	
	(0.120)	
Observations	4,100	4,100
R-squared	0.763	0.766
Age	Yes	Yes
Race	Yes	Yes
Year FE	Yes	Yes
State FE	Yes	Yes

Note: *** p<0.01, ** p<0.05, * p<0.1.
Source: Authors' calculations from the 1995-2013 LBD and CPS.

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