NBER WORKING PAPER SERIES

STRONG EMPLOYERS AND WEAK EMPLOYEES:
HOW DOES EMPLOYER CONCENTRATION AFFECT WAGES?

Efraim Benmelech
Nittai Bergman
Hyunseob Kim

Working Paper 24307
http://www.nber.org/papers/w24307

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
February 2018

We thank Eric Kim for research assistance, Bert Grider at the TRFSRDC for helping with data and clearance requests, and Evan Buntrock at the NYFSRDC (Cornell) for administrative assistance. Any opinions and conclusions expressed herein are those of the authors and do not necessarily represent the views of the U.S. Census Bureau. All results have been reviewed to ensure that no confidential information is disclosed. The views expressed herein are those of the authors and do not necessarily reflect the views of the National Bureau of Economic Research.

Strong Employers and Weak Employees: How Does Employer Concentration Affect Wages?
Efraim Benmelech, Nittai Bergman, and Hyunseob Kim
NBER Working Paper No. 24307
February 2018
JEL No. E24,J21,J23,J31,J42

ABSTRACT

We analyze the effect of local-level labor market concentration on wages. Using Census data over the period 1977–2009, we find that: (1) local-level employer concentration exhibits substantial cross-sectional and time-series variation and increases over time; (2) consistent with labor market monopsony power, there is a negative relation between local-level employer concentration and wages that is more pronounced at high levels of concentration and increases over time; (3) the negative relation between labor market concentration and wages is stronger when unionization rates are low; (4) the link between productivity growth and wage growth is stronger when labor markets are less concentrated; and (5) exposure to greater import competition from China (the “China Shock”) is associated with more concentrated labor markets. These five results emphasize the role of local-level labor market monopsonies in influencing firm wage-setting behavior and can potentially explain some of the stagnation of wages in the United States over the past several decades.

Efraim Benmelech
Kellogg School of Management
Northwestern University
2001 Sheridan Road
Evanston, IL 60208
and NBER
e-benmelech@kellogg.northwestern.edu

Nittai Bergman
Coller School of Management
and Berglas School of Economics
Tel Aviv University
Tel Aviv-Yafo
Israel
nbergman@tauex.tau.ac.il

Hyunseob Kim
Cornell University
Johnson Graduate School of Management
321 Sage Hall
114 East Avenue
Ithaca, NY 14853
hk722@cornell.edu

I. INTRODUCTION

Wages in the United States have exhibited puzzling behavior in recent decades. Real wage growth has slowed since the early 1970s, particularly at lower values of the wage distribution (Acemoglu and Autor 2011). Although productivity growth has also diminished somewhat, the link between productivity growth and real increases in hourly compensation appears to have declined since the early 1980s (see, e.g., Mishel 2012; Bivens and Mishel 2015; and Uguccioni, 2016). Consistent with these trends, the labor share of national income in the United States has diminished, particularly since the early 2000s (e.g., Karabarbounis and Neiman 2014).

Many economists have sought to understand these phenomena. A number of potential—and nonmutually exclusive—explanations for the stagnation in wages and the decline in the labor income-share are proposed in the literature. Among these are increased globalization, with an emphasis on increased trade and competition with China (e.g., Autor, Dorn, and Hanson 2013; and Elsby, Hobijn, and Şahin 2013); changes in technology and the relative cost of capital goods, particularly when stemming from increased industrial automation (see, e.g., Karabarbounis and Neiman 2014; Brynjolfsson and McAfee 2014; and Acemoglu and Restrepo 2016); and an increase in rents flowing to landowners as opposed to labor (Rognlie 2015). Another important avenue of research in understanding the behavior of labor income focuses on secular changes in the industrial composition of firms within the United States, with an emphasis on the increase in market concentration within the corporate sector in the United States since 1997 (see, e.g., Barkai 2016; and Autor et al. 2017).

In the spirit of this literature on market concentration, we use micro-level Census data to analyze how monopsony power in local labor markets affects wage behavior. Whereas much prior work on the labor income share focuses on the decrease in competition within the *product* market—that is, on firms' ability to exploit market power to increase prices and markups, thereby extracting rents (see, e.g., Barkai 2016; Grullon, Larkin, and Michaely 2016; Autor et al. 2017; and De Locker and Eeckhout 2017)—we focus on concentration within the market for *labor*: that is, the ability of monopsonist employers in the corporate sector to exploit their market power to reduce wages. Further, because the relevant market associated with job searching is largely *local*—U.S. labor mobility has declined significantly and job switches often occur between positions in the same area (Moretti 2011; Molloy, Smith, and Wozniak 2014)—our analysis emphasizes the importance of measuring labor market concentration within relatively localized geographic areas.

Our analysis combines two main sources of data from the U.S. Census Bureau over the sample period 1977–2009. First, to measure local-level labor market concentration, we use the Longitudinal Business Database (LBD) to construct the Herfindahl-Hirschman Index (HHI) of firm employment at the *county-industry-year* level: HHI is calculated at the local industry level under the plausible assumption that employees' specific human capital and mobility costs constrain their job searches toward firms within the same industry and geographical area. These HHI measures of employer concentration are then related to measures of average wages and productivity at the establishment level constructed from the Census of Manufacturers (CMF) and the Annual Survey of Manufacturers (ASM).

We achieve five main results. First, local-level employer concentration exhibits significant cross-sectional and time series variation. During 1977–2009, the standard deviation of local-level employer HHI (defined at the 4-digit SIC-level) was 0.334. Further, the data show that local-level employer concentration has increased considerably over time, with the employment-weighted mean four-digit county-level HHI increasing by 5.8%, from 0.698 during 1977–1981 to 0.756 during 2002–2009.

Our second result establishes our baseline finding: there is a negative relation between the local-level HHI measures of employer concentration and wages. Employers operating in areas with more concentrated labor markets thus appear able to exploit monopsony power in order to reduce employee wages. One important concern regarding local-level HHI as a measure of local employer concentration is that it may be correlated with other observable and unobservable differences across plants, industries, and local areas (see, e.g., Boal and Ransom 1997). To mitigate this concern, we show that these results continue to hold after controlling for a host of observables likely to affect wages, including establishment-level labor productivity and local labor market size, as well as firm-by-year fixed effects. Identification is thus achieved using within firm-year variation in which (after controlling for such observables as productivity and market size), within a given year, two establishments belonging to the same firm but located in areas with varying levels of labor market concentration are compared. The baseline finding is also robust to the use of industry-by-year fixed effects together with firm-by-year fixed effects.[5] This battery of controls removes any cross-industry variation within firms, thereby alleviating cross-industry heterogeneity as an alternative channel that drives wage differences. Furthermore, our baseline results continue to hold in a subsample of firms that operate multiple plants in only one industry segment. Using this subsample in conjunction with firm-by-year fixed effects, we largely sidestep cross-industry heterogeneity as an alternative explanation.

The economic magnitude of the baseline negative relation between employer concentration and wages effect is significant over the sample period, a one standard deviation increase in the HHI measure of local employer concentration reduces wages by between 1% and 2%, depending on the specification. Importantly, we show that this baseline relation is highly nonlinear and convex, with labor market concentration at high levels of employer HHI exhibiting a particularly strong effect on wages—plausibly due to reduced worker bargaining power at these levels of concentration. Further, our results indicate that the negative relation between employer concentration and wages doubles in magnitude over the sample period: during 1997–2001, a one standard deviation in local-level concentration is associated with a 1.37% wage reduction, whereas the equivalent effect in 1977–1981 is only 0.63%. This strengthening of the sensitivity of wages to employer concentration is strongly consistent with a secular decline in worker bargaining power over time, as would be predicted by the reduction in labor mobility in the United States (constraining the choice set of workers as they search for employment and negotiate over compensation) or by the drop in unionization rates within the United States beginning in the 1970s (Card 1992).

Our third finding concerns the impact of unionization on the relation between employer concentration and wages. We hypothesize that by improving employee bargaining power, unionization may diminish the ability of employers to lower wages in concentrated labor markets, and indeed, the data bear this out. The negative relation between the HHI labor market concentration measure and wages is significantly weaker among plants in industries with high unionization rates. For example, when measuring local labor market concentration at the three-digit SIC code level, at one standard deviation of the unionization rate above its mean, there is no statistically significant effect of labor market concentration on wages.

[5] We obtain identification of both industry-by-year as well as firm-by-year fixed effects because multi-establishment firms may operate establishments in more than one industry.

Fourth, we investigate how local-level labor market concentration affects the transmission of productivity growth into higher wages. We hypothesize that high levels of employer concentration impede the translation of productivity growth to wage increases, as employers use their monopsony power to prevent wage increases. In contrast, when labor markets are more competitive, productivity increases should give rise to wage growth as employers compete for workers. Put another way, productivity growth should translate into a rise in wages when employee bargaining power is sufficiently high.

Measuring labor productivity at the establishment level using Census data, we first confirm an annual link between wage growth and productivity growth (see, e.g., Stansbury and Summers 2017). Importantly, consistent with our hypothesis on the role of monopsony power in labor markets, we find that the link between wage growth and productivity growth is significantly larger when local-level employer concentration is small. Decreasing the employer-based HHI measure of concentration by one standard deviation is associated with an increase in the elasticity of wages to productivity from 0.38% to 0.47%.[6]

Last, we investigate the impact of increased trade with China—the "China Shock"—on the level of local labor market concentration. As such, in contrast to our prior findings, we are interested here in the determinants, rather than the effects, of labor market concentration. We hypothesize that increased import competition from China may lead to an increase in local labor market concentration by causing employers to shutter their operations. The data indeed confirm that a rise in industry-level import competition from China is associated with increased employer concentration in the local labor market. Aside from labor displacement and an associated decline in wages due to a reduced demand for labor, then, increased import competition from China may also have the effect of reducing wages of *nondisplaced* workers due to an increase in employer concentration.

Our article relates to a growing literature dealing with monopsony in labor markets.[7] However, most prior empirical work on monopsony power in labor markets and firms' ability to control wages is indirect. Examples include anecdotal evidence on lawsuits against employer collusion using antipoaching agreements (U.S. Department of Justice Office of Public Affairs

[6] In related work, Benmelech, Bergman, and Kim (2017) show that the relation between wage growth and productivity growth is increasing in unionization rates. As with the results above, when worker bargaining power rises, increased productivity translates into wage growth.
[7] For reviews, see Boal and Ransom (1997) and Council of Economic Advisers Issue Brief (2016).

2014; Whitney 2015) and studies beginning with Card and Krueger (1994) showing that minimum wage increases are not followed by reductions in employment—consistent with nonperfectly competitive labor markets. Other studies show that the propensity of workers to leave their jobs after wage declines is smaller than would be expected in competitive markets, suggestive of employers' wage-setting ability (Dube, Lester, and Reich 2010, 2016).

In contrast to these studies, we provide a direct measure of the degree of local-level labor market competition and relate it to wage behavior on a host of margins. As such, our article is most closely related to Azar, Marinescu, and Steinbaum (2017), which uses online job postings from the website CareerBuilder.com to construct a local-level measure of employer concentration based on commuting zones and occupations. Azar et al.'s chief finding is that increased local-level labor market concentration within a given occupation and commuting zone predicts lower online wage postings. Though similar, our article is differentiated along a number of dimensions. We use actual wages (at the establishment level) as taken from the Census Bureau, whereas Azar et al. analyze wage *postings* offered by firms. Our study uses a series of data that spans more than three decades and covers the entire manufacturing sector, whereas Azar et al. examine online wage posting within CareerBuilder.com (that is, not solely focused on manufacturing) over a relatively short time (preceding the 2008–2009 crisis). Further, our study focuses on manufacturing, so we can control directly for productivity—an important covariate when analyzing wages—and indeed we do so at the establishment-year level. Last, we provide evidence on the time-series evolution of employer concentration, the impact of unionization on the relation between local employer concentration and wages, how labor market concentration affects the transmission of productivity changes into wage growth, and the effect of the "China Shock" on local employer concentration. Interestingly, we also find that the negative relation between concentration and wages is convex and more pronounced in a setting that approaches textbook-like monopsony—that is, when the local HHI employer concentration measure approaches one. In contrast, Azar et al. show that the negative relation in their setting is not driven by monopsony.

The rest of the article is organized as follows. Section II presents the data and summary statistics. Section III presents the empirical analysis, and Section IV concludes.

II. Data and Key Variables

II.A. Data Sources and Sample Construction

1. Plant-Level Data. We obtain data on manufacturing establishments ("plants") from the Census of Manufacturers (CMF) and the Annual Survey of Manufacturers (ASM) maintained by the U.S. Census Bureau to construct measures of wages and labor productivity. The CMF covers all manufacturing plants in the United States for years ending in 2 and 7 ("Census years"), resulting in roughly 300,000 plants in each Census. The ASM covers about 50,000 plants for "non-Census years." The ASM includes all plants with more than the applicable number of employees, which increases from 250 to 1,000 during our sample period;[8] those with fewer employees are sampled randomly, with the probability of inclusion increasing in size. Although the ASM is called a survey, reporting is mandatory if the plant is selected, and misreporting is subject to legal penalties and fines. Both databases provide operating information at the plant level, including the total value of shipments, wages, labor hours, and material and energy costs. A key advantage of the CMF and ASM data over standard firm-level databases of public firms such as Compustat is that they comprise both privately and publicly owned plants, covering a significant fraction of U.S. workers. We also use the Longitudinal Business Database (LBD) to construct measures of concentration of firms' employment within a county-industry. The LBD is a comprehensive data set of manufacturing and nonmanufacturing establishments in the United States, tracking more than 5 million establishments every year. The database provides variables for the number of employees, annual payroll, industry classifications, and geographical location (i.e., counties and states).

2. Union Coverage Data. We obtain data on collective bargaining coverage from the Union Membership and Coverage Database.[9] This database provides private- and public-sector labor union membership, coverage, and density estimates computed from the monthly household Current Population Survey (CPS) using methods that the Bureau of Labor Statistics (BLS) employs to calculate union statistics at the national level. Prior research on financial and labor economics uses the database to estimate industry-level collective bargaining coverage (e.g., Matsa 2010; and Elsby, Hobijn, and Şahin 2013). Hirsch and Macpherson (2003) provide

[8] The thresholds are 250 employees before 1999, 500 from 1999 to 2003, and 1,000 after 2003.
[9] The database is available at www.unionstats.com.

detailed information on the construction of the union data set. Below we give a detailed account of the data source and construction of the data that are relevant for our empirical analysis.

The Hirsch-Macpherson database provides union coverage estimates by Census detailed industry code (CIC) beginning in 1983 (aggregate-level coverage is available from 1973).[10] In order to match the union coverage data with Census plant-level data sets, we make the following adjustments. First, we use the Census Bureau concordances between the CIC and Standard Industry Classification (SIC) codes for the years 1983–2002. For 2003–2009, during which the CIC is based on a NAICS classification (as opposed to SIC), we use the Census Bureau crosswalk between the 1990 and 2000 CIC codes to match the 2000 CIC to SIC codes. Second, we impute missing industry-level collective bargaining coverages before 1983 using the information from the 1983 CPS. Third, we lag the union coverage information by one year relative to plant-year observations. For the majority of industries, the matching between the CIC and SIC codes results in linkages at the three-digit SIC level, and less frequently at the two- or four-digit levels.

We focus on manufacturing plant-year observations in the CMF and ASM from 1977 to 2009 (the first and last years of data coverage). Given that the starting year of the union coverage database by detailed industry (1983) is later than 1977, we impute the rate of collective bargaining coverage using the 1983 information for years between 1977 and 1982. We require each plant observation to have the variables necessary to estimate average wages and labor productivity, including the SIC code,[11] total value of shipments, production-worker equivalent hours, total wages, and number of employees (both production and nonproduction). Given that our identification strategy relies on within-firm variation in wages and employer concentration across plants, we require that firm-years have at least two plants under their ownership. In addition, we require that each plant-year have a one-year lagged observation, which is needed to compute changes in average wages and productivity in part of our analysis. This selection procedure yields approximately 656,000 plant-years in our sample.[12]

II.B. Measuring Employer Concentration and Wages

---

[10] They use the 1980 CIC (based on SIC codes) from 1983 to 1991, the 1990 CIC from 1992 to 2002 (based on SIC codes), and the 2000 CIC (based on NAICS codes) from 2003 and on.
[11] The ASM and CMF provide SIC codes until 2002 and NAICS codes thereafter. We impute SIC codes after 2002.
[12] The numbers of observations are rounded to the nearest thousand to follow the Census Bureau's disclosure rules.

We construct HHI of employment by firms at the county-industry-year level as our main measures of employer concentration using the LBD data. Specifically, we measure the employment share of every firm in a given county-industry-year cell as: $s_{f,j,c,t} = \frac{emp_{f,j,c,t}}{\sum_{f=1}^{N} emp_{f,j,c,t}}$, where *emp* represents total employment and *f* represents a firm that operates in industry *j* and county *c* in year *t.* We next calculate HHI at the county-industry-year level as the sum of the squared employment shares in the country-industry-year level: $HHI_{j,c,t} = \sum_{f=1}^{N} s_{f,j,c,t}^{2}$. We define two variants of HHI using either of the three- or four-digit SIC codes.

Given that we are interested in understanding the effect of employer concentration on wages, we use the log of average wages per worker at the plant-year level as our dependent variable. We compute the average wages at the plant-year level as total wage bills divided by the number of workers. We also use labor productivity (per hour) as a control variable, defined as output divided by total labor hours. Output is computed as the sum of the total value of shipments (TVS) and the net increase in inventories of finished goods and works in progress. To account for industry-level changes in output price, we divide output by the four-digit SIC-level output price deflator from the NBER-CES manufacturing industry database constructed by Bartelsman, Becker, and Gray (2000). In addition, we use "production-worker equivalent hours" as our measure of total labor hours (Lichtenberg 1992; Kovenock and Phillips 1997). Specifically, total labor hours are constructed as the total production-worker hours multiplied by total wages divided by wages for production ("blue-collar") workers. The underlying assumption in constructing this measure of labor hours is that the relative wages for production and nonproduction workers represent the ratio of their marginal products. We winsorize all potentially unbounded variables at the 1% tails.

II.C. Descriptive Statistics

Table I reports descriptive statistics for the characteristics of the plant observations over the period 1977–2009. The average plant in the sample has \$96 million in total value of shipments ("sales") and 349 employees. Focusing on the measures of employer concentration—the mean HHI defined at the three-digit SIC level is 0.545 with a standard deviation of 0.350, while the mean four-digit SIC-based HHI is 0.682 with a standard deviation of 0.334. We also define dummy variables that take the value of one in counties in which HHI equals one, and zero otherwise. These HHI-based dummy variables are designed to capture counties with an absolute monopsony power of one large employer. As Table I shows, 22.7% of plant-year observations are in counties in which there is only one employer in a given three-digit SIC industry. The concentration of employer power is even stronger when we define industries at the four-digit SIC level. For example, 37.9% of plant-year observations are in counties with only one employer in the four-digit SIC industry.

The average worker in these plants earns wages of $41,840 per year, and the average log labor productivity is 4.61. The average log employment of a given county-industry is 6.24 (5.66) when the industry is defined at the three-digit (four-digit) level. The average union coverage rate for the manufacturing plants in the sample is 22.8%, which is higher than the average coverage rate across all industries during a similar period. Segment and firm size are measured by the log number of plants in a given industry segment of a given firm and the log number of all plants of a given firm, respectively. The average plant in the sample has approximately 10 and 30 plants (including itself) within the segment-firm and firm it is associated with. Plant age is defined as the number of years since a plant's inception—identified by the flag for plant inception in the LBD—or its first appearance in the CMF or ASM database, whichever is the earliest. The starting year is censored in 1972, when the coverage of the Census databases begins. The average age of plants in the sample is 17.1 years. This set of control variables is standard among research that analyzes plant-level data using the CMF and ASM data (e.g., Schoar 2002; Giroud 2013).

Figure I plots trends in local-level employer concentration, which exhibits considerable time-series variation from 1977 to 2009. During 1977–1981, the employment-weighted mean local-level employer HHI (defined at the four-digit SIC level) is 0.698, but the mean four-digit county-level HHI increases by 5.8 percentage points to 0.756 during 2002–2009. This upward trend in local employer HHI is consistent with increasing concentration and power of firms in the labor market in the U.S. over the past few decades.

III. EMPIRICAL ANALYSIS

III.A. Relation between Employer Concentration and Wages

We begin by investigating a reduced-form baseline relation between employer concentration and wages. To this end, we run the following regression:

$$log(avg.\ wages)_{pfjt} = \boldsymbol{\beta_0} + \boldsymbol{\beta_1} HHI_{jct-1} + \boldsymbol{\beta_2} X_{pfjt} + \boldsymbol{\beta_3} \boldsymbol{Z}_{jct-1} + \boldsymbol{\delta}_{jt} + \boldsymbol{\mu}_{ft} + \boldsymbol{\varepsilon}_{pfjt}, \quad (1)$$

where $log(avg.\ wages)_{pfjt}$ is log average wage per worker. The subscripts are plant p, in firm f, within industry j, in year t. HHI_{jct-1} is one-year lagged employment-based measure of concentration at either the three- or four-digit SIC level in county j, $\boldsymbol{X}_{pijt}$ is a vector of plant-level control variables comprising the log of labor productivity, the log of the number of plants per segment within the firm, the log of the number of plants per firm, and plant age, $\boldsymbol{Z}_{jct-1}$ is the one-year lagged log number of employment at the county-industry level, $\boldsymbol{\delta}_{jt}$ is a vector of industry by year fixed effects, and $\boldsymbol{\mu}_{ft}$ is a vector of firm-by-year fixed effects. All standard errors are clustered at the county level.

Panel A of Table II provides the results from estimating regression (1) using three-digit SIC industries to compute employment concentration. As column (1) of Table II shows, the coefficient estimate for log(employment) is 0.042 and is statistically significant at the 1% level, suggesting that wages are higher in local labor markets with more workers. This finding is consistent with economic forces related to agglomeration (e.g., Ellison, Glaeser, and Kerr 2010; and Greenstone, Hornbeck, and Moretti 2010). The coefficient on log labor productivity is 0.156 and significant at the 1% level, suggesting that workers' wages are higher when the plant is more productive, plausibly due to rent sharing between the employer and workers (e.g., Nickell and Wadhwani 1990; and Benmelech, Bergman, and Kim 2017).

Turning to our main variable of interest—namely, employer concentration—column (1), which includes year fixed effects as a control in addition to log employment at the three-digit industry-county level and plant-level labor productivity, shows that log average wage is negatively associated with HHI. This relation is statistically significant at the 1% level, and the economic magnitude is sizable—a one standard deviation increase in employer concentration (0.35) is associated with a 0.88% decrease (= 0.35 × −0.025) in average wages. Similarly, the relation between log average wages and employer concentration is statistically and economically significant when other plant-level controls are added (column (2)) and industry fixed effects are included (column (3)). In particular, by including industry fixed effects, we exploit within-industry variation of employer concentration to identify the effect. Thus, the estimates in column (3) are based off of variation in employer concentration across different local labor markets, instead of variation of concentration of firms in the national market across industries.

Furthermore, we also include firm or firm-by-year fixed effects in the last two columns of the table, which enables us to identify off of variation within firms across plants that are located in counties and industries with different employer concentration. Our plant-level explanatory variables—in particular, log(labor productivity) and plant age—control for plant-level productivity, alleviating the concern that our results are driven by productivity differences across plants within a firm. With the inclusion of firm fixed effects in column (4), the semi-elasticity of wages to employer concentration is −0.028. Thus, increasing HHI by 0.35, the sample standard deviation, decreases wages by 0.98% (= 0.35 × −0.028). With the addition of firm-by-year fixed effects, this semi-elasticity becomes −0.023 (column (5)). Wages thus appear to be negatively correlated with employer concentration, consistent with the prediction that concentration of employers in the local labor market gives firms monopsony power that they use to exploit workers in the form of lower wages.

While the analyses in columns (4) and (5) mitigate the concern that lower wages in more concentrated labor markets are due to heterogeneity across firms, a remaining concern is that a given firm may pay different levels of wages to employees across industries and locations due to other omitted variables, such as product market competition and productivity differences. To address this concern, we include industry-by-year fixed effects, which control for time-varying differences across industries, in addition to firm-by-year fixed effects. In this analysis, the identifying variation comes from plants in the same industry and the same firm but that are located in counties with different employer concentration. Including these industry-by-year fixed effects, column (6) shows that the semi-elasticity of wages to employer concentration is −0.049, which is more than twice the effect without controlling for time-varying industry shocks in column (5). This result suggests that once we compare plants in the same industry and firm, those located in a more concentrated local labor market pay significantly lower wages. In conjunction with later results in Table IV, which limits the analysis to firm-years that operate in one industry segment, results in this table suggest that industry-level heterogeneity (in productivity in particular) is unlikely to drive the association between employer concentration and wages.

Panel B of Table II repeats the analysis in Panel A using four-digit SIC industries to compute HHI of employer concentration at the county level. The negative relation between concentration and employee wages is evident in this panel as well. The economic magnitude of the effect is larger than that in Panel A, with a semi-elasticity ranging from −0.038 to −0.063 relative to −0.023 to −0.049 in Panel A. In sum, the results in Table II exhibit a statistically and economically significant relation between wages and employer concentration, suggesting that employer concentration results in lower wages.

One concern about the negative relation between employer concentration and wages presented in Table II is that our local measure of HHI is correlated with labor productivity, which could mean that the results capture the effect of low productivity rather than employer concentration on wages. Even though all of our regressions in Table II also control for log(labor productivity), defined as the natural log of output scaled by labor hours, it might be measured with error and thus is likely an imperfect control for labor productivity. To address this concern, in Table III we add a control variable for labor productivity—namely, value added (VA) by labor, defined as the total value of shipments plus the net increase in inventories of finished goods and works in progress minus material and energy costs scaled by total labor hours—in the regression. All the components of the variable are deflated using four-digit SIC industry-level deflators available from the NBER-CES manufacturing database. One key difference of "labor VA" relative to "labor productivity" is that it nets out intermediate inputs (i.e., material and energy), which are presumably not affected by labor input.

Table III, Panel A, shows that the coefficient on log(labor VA) is positive and significant at the 1% level except for columns (4) and (5), which include firm and year and firm-by-year fixed effects, consistent with the notion that employees of plants who add more value per labor hour are paid more. However, the coefficients on HHI remain statistically significant and their magnitudes are almost identical, suggesting that the baseline relation between HHI and wages is unlikely to be driven by omitted controls for value added by workers or labor productivity. Panel B, which uses an industry based on the four-digit SIC level, shows similar results—that the coefficient on local employment concentration is largely unaffected by adding a control for value added by labor.

Among the specifications we employ in Table II, the inclusion of firm-by-year fixed effects in column (5) identifies the effect of local employer concentration on wages off of variation within firms in a given year. Although this specification with the detailed fixed effects likely provides a tighter identification, a remaining concern is that the negative association between local employer concentration and wages is driven by omitted differences across different industries even within firms in a given year. For example, a firm may have a more productive industry segment (e.g., machinery) in a less concentrated local labor market, and a less productive segment (e.g., chemical) in a more concentrated market, which may lead to a spurious correlation between concentration and wages due to a difference in productivity.

We address this concern in two ways. First, in addition to firm-by-year fixed effects, we include industry-by-year fixed effects in Table II, column (6), to absorb time-varying industry-level shocks in market concentration, wages, and other omitted variables (such as productivity). Second, we focus our estimation on a subsample of firms that operate multiple plants in only one industry segment (defined at either the three- or four-digit SIC level). Combined with firm-by-year fixed effects, use of this subsample removes any cross-industry variation within firms, thereby sidestepping cross-industry heterogeneity as an alternative channel that drives wage differences.

Table IV presents the results from estimating equation (1) using a subsample of firms that operate multiple plants within one industry segment, defined at the three- (Panel A) or four-digit SIC level (Panel B). As Table IV illustrates, across all columns the coefficient on HHI is negative and significant at the 1% level, consistent with the baseline result in Table II. In terms of economic magnitudes, the coefficients are approximately twice larger than those in Table II (e.g., −0.061 and −0.023 in column (5), which includes firm-by-year fixed effects), indicating that the effect is larger for the subset of firms that are unlikely to be affected by cross-industry heterogeneity in productivity and wages. Panel B, which uses four-digit SIC industries, shows a qualitatively similar result, with economic magnitudes being again larger than those in Panel B of Table II. In sum, our results are robust to the inclusion of various measures of labor productivity and hold in firms that operate multiple plants within one industry, which we believe alleviates concerns that an omitted productivity variable could be driving our HHI results.

III.B. Subsample Periods

We next investigate how the relation between wages and employer concentration evolves over the sample period. To this end, Table V divides the full sample from 1977 to 2011 into five-year subperiods and reruns the wage regressions in equation (1), which includes HHI of employers (i.e., firms) at the county-industry level as a key independent variable.[13] Our five-year

[13] The final subperiod is 2002–2009.

subperiods in Table V match the subperiods in Figure I, which illustrates that local-level employer concentration increases over time. As shown in Panel A, the coefficient on HHI—defined at the three-digit SIC level—increases over time. In the first two subsamples, 1977–1981 and 1982–1986, the coefficients on HHI are −0.018 (with a *t*-statistic of −2.75) and −0.012 (with a *t*-statistic of −1.78), respectively. This coefficient increases to −0.031 in 1992–1996 (with a *t*-statistic of 4.59) and −0.039 in 1997–2001 (with a *t*-statistic of 5.31). In the final subsample, the coefficient is −0.029 and is statistically significant at the 1% level. The lower wages in highly concentrated local labor markets are thus clustered in the later part of the sample period. In Panel B, which uses four-digit SIC industries to compute HHI, the increasing pattern is muted, although the effect of concentration on wages is the greatest between 1992 and 2001 (−0.044 to −0.047).

A potential explanation for the increasing effect over time of employer concentration is that labor mobility declines over the sample period across both economic sectors and geographical areas (see, e.g., Murphy and Topel 1987; and Molloy, Smith, and Wozniak 2014). The effect of labor market monopsony hinges on a lack of labor mobility across markets (for a review, see Boal and Ransom 1997). Thus, to the extent that workers have become less mobile in the United States over the past few decades, our definition of local labor markets using counties and industries is more likely to bind for them in recent decades relative to such earlier periods as 1977–1981 and 1982–1986.

III.C. Employer Concentration, Unions, and Wages

While their monopsony power may lead employers to pay lower wages, unionization strengthens labor's bargaining position and enables employees to diminish employers' monopsony power. We next empirically test whether unionization mitigates the ability of firms to reduce wages in concentrated markets. More specifically, we interact our local measure of employer concentration (HHI) with the degree of unionization at the industry level, as in the following regression:

$$log(avg.\ wages)_{pfjt} = \beta_0 + \beta_1 HHI_{jct-1} + \beta_2 HHI_{jct-1} \times Union_{jt-1} + \beta_3 Union_{jt-1} + \beta_4 \boldsymbol{X}_{pfjt} + \beta_5 \boldsymbol{Z}_{jct1} + \boldsymbol{\delta}_{jt} + \boldsymbol{\mu}_{ft} + \boldsymbol{\varepsilon}_{pfjt}\,, \qquad (2)$$

where $Union_{jt-1}$ is the unionization rate for industry j in which plant i operates in year $t-1$, and the other variables are defined in equation (1). The main coefficient of interest in this regression is β_2, which measures the degree to which unionization rates affect the marginal effect of employer concentration on wages.

Table VI presents estimates of regression (2) using plant-level data from 1977 to 2009.[14] Columns (1) and (2) in Panel A show that, as in Table II, controlling for year fixed effects, and plant-level and county-industry-level controls, average workers' wages at manufacturing plants are lower when the local labor market (defined at the county level) is more concentrated in a given three-digit SIC industry. More importantly, this negative effect of employer concentration on wages is mitigated for plants that operate in industries with higher unionization rates. For example, the regression estimates in column (2) suggest that a one standard deviation increase in unionization rates (12.9%) from the average (22.8%) would reduce the semi-elasticity of wages to HHI from −0.061 (= −0.153 + 0.228 × 0.402) to −0.009 (= −0.153 + (0.228 + 0.129) × 0.402). This result is consistent with unions giving workers bargaining power in wage negotiations, mitigating a negative effect of local labor market concentration on wages.

Columns (3)–(5) further include industry, firm, and firm-by-year fixed effects. The results are consistent with those in columns (1) and (2)—the negative effect of employer concentration is mitigated by high unionization rates at the industry level. In terms of economic magnitude, estimates in column (5) indicate that a one standard deviation increase in unionization rates (12.9%) from the average (22.8%) would reduce the semi-elasticity of wages to HHI from −0.035 (= −0.087 + 0.228 × 0.228) to −0.005 (= −0.087 + (0.228 + 0.129) × 0.228). Column (6) further includes industry-by-year fixed effects to control for time-varying industry shocks, and finds consistent evidence that the presence of union bargaining power mitigates a negative association between employer concentration and wages.

In Panel B of Table VI we repeat the analysis in Panel A using four-digit SIC industries to compute HHI of employer concentration at the county level.[15] The results reported in Panel B are similar to those found in Panel A—we find that labor unions mitigate the effect of employer concentration on wages. The economic magnitudes of the effects are somewhat smaller than those documented in Panel A. For example, the estimates in column (5) suggest that a one standard deviation increase in unionization rates (12.9%) from the average rates (22.8%) would

[14] As explained in Section II.A, data on industry-level unionization rates are available from 1983 only. Thus, we impute the unionization rates before 1983 (i.e., 1977–1982) using values for corresponding industries in 1983.

[15] In Panel B and throughout the article in which we use four-digit SIC industries to compute employer concentration, we use the same industry-level of unionization rates, which are based on CIC (see Section II.A), as in the case of using three-digit SIC industries.

reduce the semi-elasticity of wages to HHI from −0.046 (= −0.080 + 0.228 × 0.151) to −0.026 (= −0.080 + (0.228 + 0.129) × 0.151). Nevertheless, as our calculation above shows, a one standard deviation in unionization rate reduces the negative effect of employer concentration on wages by 42% relative to the mean.

III.D. Employer Concentration and Rent Sharing

The analysis thus far has focused on the association between employer concentration in the local labor market and wages. In addition to wage levels, concentration of employers may also affect the transmission of productivity growth into wage changes. Previous research on rent sharing suggests that workers' bargaining power increases the portion of productivity growth that they gain in the form of increased wages (e.g., Stansbury and Summers 2017). Thus, we empirically test whether high levels of employer concentration impede the translation of productivity growth to wage increases, as employers use their monopsony power to avoid wage hikes. More specifically, we examine whether sensitivities of wage growth to productivity growth are affected by our local measure of employer concentration (HHI) using the following regression:

$$\Delta log(avg.\ wages)_{pfjt} = \beta_0 + \beta_1 HHI_{jct-1} + \beta_2 \Delta log(labor\ productivity)_{pfjt} + \beta_3 HHI_{jct-1} \times \Delta log(labor\ productivity)_{pfjt} + \beta_4 \boldsymbol{X}_{pfjt} + \beta_5 \boldsymbol{Z}_{jct1} + \boldsymbol{\delta}_{jt} + \boldsymbol{\mu}_{ft} + \boldsymbol{\varepsilon}_{pfjt}, \quad (3)$$

where $\Delta log(avg.\ wages)_{pfjt}$ is the growth rate of average wages, and $\Delta log(labor\ productivity)_{pfjt}$ is the growth rate of labor productivity of plant p in industry j, county c, and year t; the other variables are defined in equation (1).

Table VII reports estimation results. Across columns in Panel A of Table VII, the coefficient on Δ*log(labor productivity)* is positive and significant at the 1% level, confirming that there is a positive association between wage growth and productivity growth at the plant level (see, e.g., Stansbury and Summers 2017). Importantly, consistent with our prediction regarding the role of monopsony power in labor markets, the coefficient on *HHI*×Δ*log(labor productivity)* is negative and significant at the 10% level. In terms of economic magnitude, estimates in column (5) suggest that a one standard deviation decrease in HHI from its mean would increase the elasticity of wages to productivity from 0.38% (= 0.005 − 0.002 × 0.544) to 0.47% (= 0.005 − 0.002 × (0.544 − 0.350)). This finding is consistent with our prediction that when labor markets are more competitive, productivity increases should give rise to wage growth, as employers compete for workers.

III.E. Robustness Tests

1. Nonlinear Effect of Employer Concentration on Wages. Regressions in equations (1) and (2) assume a linear effect of employer concentration, measured by HHI, on average wages. Is the effect concentrated among county-industries that are highly concentrated by few firms or among those that are close to competitive local labor markets with many employers? Table I shows that there is indeed a significant fraction of manufacturing plants that are located in these counties with extreme concentration—namely, 22.7% and 37.9% of plant-year observations based on three-digit and four-digit SIC industries, respectively. We examine this question by including a dummy variable that is equal to one for county-industries that have only one firm, and thus HHI is equal to one in the following regression:

$$log(avg.\ wages)_{pfjt} = \beta_0 + \beta_1 HHI_{jct-1} + \beta_2 HHI_{jct-1} \times Union_{jt-1} + \beta_3 Union_{jt-1} + \beta_4 D(HHI=1)_{jct-1} + \beta_5 1(HHI=1)_{jct-1} \times Union_{jt-1} + \beta_6 \boldsymbol{X}_{pfjt} + \beta_7 \boldsymbol{Z}_{jct1} + \boldsymbol{\delta}_{jt} + \boldsymbol{\mu}_{ft} + \boldsymbol{\varepsilon}_{pfjt}, \quad (4)$$

where $D(HHI=1)_{jct-1}$ is a dummy variable equal to one if HHI of industry j in county c in year $t-1$ is equal to one, and zero otherwise; the other variables are defined in equation (1). Given that we include firm-by-year fixed effects ($\boldsymbol{\mu}_{ft}$), interpretation of the coefficients on β_4 is the effect of absolute monopsony power of one large employer on wages within firms and across different local labor markets.

Table VIII reports estimation results. Column (1) includes the dummy variable but excludes the baseline continuous variable for HHI, as well as interaction terms with unionization rates. The coefficient on the dummy is −0.017, suggesting that relative to other plants in nonperfectly concentrated local labor markets, plants in perfectly monopsonistic labor markets pay 1.7% lower wages, controlling for plant- and county-industry-level determinants of wages. Column (2) includes the continuous variable HHI, in addition to the dummy, and shows that employer concentration has an additional negative effect on wages (−0.011) when HHI increases to one, in addition to the general effect of HHI (−0.013). These results suggest that the negative effect of employer concentration on wages is more pronounced when the local labor market is close to monopsony (i.e., one employer in a given market).

Next, in column (3), we exclude the continuous variable HHI but interact the dummy for HHI equals to one with industry-level unionization rates. Consistent with previous estimates in Table VI, the column shows that unionized labor significantly mitigates the negative effect of local market monopsony. At the average rate of unionization in the sample (22.8%), local labor markets with perfect concentration would have wages that are 2.5% lower (= −0.062 + 0.162 × 0.228) than otherwise similar markets within firms. But a one standard deviation increase in unionization rates would lead to a reduction in the semi-elasticity to a mere −0.4%, an 84% reduction (in absolute magnitude) from the average effect.

In Panel B, we compute employer concentration using four-digit SIC industries and find results that are consistent with those in Panel A—local labor markets with only one firm would pay wages that are 3.1% lower than otherwise similar plants within firms (column (1)). Moreover, estimates in column (2) suggest that the bulk of the effect of employer concentration on wages is at the margin around complete monopsony-like labor markets: the continuous variable HHI is not significant, while the dummy for HHI equals to one is −0.029 and significant at the 1% level. Column (3) in Panel B tells a similar story to that in Panel A—at the average rate of unionization in the sample (22.8%), local labor markets with perfect concentration would have wages that are 3.5% lower (= −0.058 + 0.101 × 0.228) than otherwise similar markets within firms. And a one standard deviation increase in unionization rates would reduce this semi-elasticity to −2.2%, a 37% reduction from the average effect. Overall, results in Table VIII suggest that the effect of employer concentration in local labor markets is more pronounced among those closest to the textbook notion of labor market monopsony (i.e., only one employer in a given market).

2. Controlling for National Trends in Employer Concentration. We measure labor market concentration at the local (i.e., county) level. The premise of this approach is that relevant labor markets for manufacturing workers are "local" due to the costs of moving across geographical areas (e.g., Moretti 2011; and Molloy, Smith, and Wozniak 2014). Meanwhile, Autor et al. (2017) show that product markets in the United States have become more concentrated in the past few decades, giving rise to "superstar firms" that are highly productive yet have lower labor shares in a given industry. They argue that this trend can account for declining labor shares documented by previous research (e.g., Elsby, Hobjin, and Şahin 2013; and Karabarbounis and

Neiman 2014). One concern about our results related to Autor et al.'s finding is that if product market (which is usually defined at the national level for "traded sectors" like manufacturing) and local labor market concentrations are positively correlated, then the negative association between local employment concentration and wages may merely reflect a negative relation between (national) product market concentration and labor shares.

We control for this effect by constructing HHI at the national level as follows. We first measure the employment share of every firm in a given industry-year cell as: $s_{f,j,t} = \frac{emp_{f,j,t}}{\sum_{f=1}^{N} emp_{f,j,t}}$, where *emp* represents total employment and *f* represents a firm that operates in industry *j* and year *t.* We next calculate the industry-year HHI as the sum of the squared employment shares in the industry-year level: $HHI_{j,t} = \sum_{f=1}^{N} s_{f,j,t}^2$. As with the local-level HHI, we define two variants of the national-level HHI using either of the three- or four-digit SIC codes. Then we include the national-level HHI as a control for product market concentration in the regression in equation (1).

Table IX, Panel A, shows that the coefficient on *HHI (SIC3-year)* is significantly positive for the most part (except for column (3), in which it is slightly negative but insignificant), which is inconsistent with the story that national-level product market concentration is associated with lower wages. Importantly, even with the inclusion of national HHI, the coefficient on local HHI remains positive and the magnitudes are sizable. For example, the coefficients on *HHI (SIC3-county-year)* are −0.023 and −0.043, respectively, in Tables II and IX, Panel A, column (5), which use the same set of controls. This result is inconsistent with an alternative mechanism that emphasizes concentration in the product market, as opposed to the labor market, which tends to be more local. Panel B, which uses the four-digit SIC level as the definition of industry, shows a similar result that the coefficient on HHI does not change much when national HHI is added as a control.

III.F. Wages, Market Concentration, and the China Shock

One concern about our findings is that they are driven by the well-documented "China Shock," in which import penetration from China has led to declining employment and wages in the U.S. manufacturing sector. For example, Autor, Dorn, and Hanson (2013) and Autor et al. (2014) find that local labor markets that are more affected by import competition from China exhibit higher unemployment, lower labor-force participation, and reduced wages. In this subsection, we examine the possibility that import penetration from China is driving our results in two ways. First, we control directly for a China effect by adding a control variable that captures the degree of import penetration from China at the industry-by-year level to our regression specification.

Second, we show that even though our results are not directly driven by import penetration, there is an interesting interaction among import penetration, employer concentration, and wages. More specifically, we examine whether import penetration from China has an indirect effect on wages by increasing employer concentration in U.S. local labor markets. That is, imports from China not only curtail the local demand for labor but also lead to increased market power of U.S. firms, which results in even lower wages.

For these two types of analyses, we construct a measure of import competition from China. We measure import exposures from China to the United States at the industry by year level as the industry-level dollar value of imports scaled by the total value of shipments in the industry.[16] More specifically, we define import competition from China as: $China\ exposure_{j,t} = \frac{import\ from\ China_{j,t}}{\sum_{i=1}^{N} shipments_{i,j,t}}$, in which *import from China* represents the dollar value of imports from China to the United States, *shipments* represents the total value of shipments (TVS), and i represents a plant that operates in industry j and year t. For the analysis of China import exposure only, we limit our analysis to the plant-years 1992–2008, given that the import penetration data employed are available from 1991 to 2007 and that we use lagged China exposure as an independent variable.

We address the concern that our measure of employer concentration is capturing direct exposure to Chinese imports by including *China exposure* as an additional control in equation (1). Table X, Panel A, presents the results and shows that the coefficient on *China exposure* measured at the three-digit SIC level is negative and significant at the 1% level, consistent with existing research documenting negative consequences of the China shock on labor market outcomes. The economic magnitude of the effect is sizable as well—a one standard deviation increase in China exposure is associated with a 0.95% (= −0.001 × 9.508) reduction in average wages. Importantly, however, including this direct control for the China exposure does not significantly affect our estimates for HHI—the coefficient on HHI remains negative and

[16] The data sources for the import and total value of shipment data are U.N. Comtrade and the ASM and CMF. We thank David Dorn for making the import penetration data available on his website.

significant at the 1% level. Panel B, which uses four-digit SIC codes as the definition of industry, shows a similar result—that controlling for the China shock does not alter the baseline result.

We now turn to analyzing the indirect effect of the China shock on wages through increased employer concentration. We begin by regressing HHI on our measure of competition from China. Specifically, we run the following regression:

$$HHI_{jct} = \beta_0 + \beta_1 China\ exposure_{jt} + \beta_2 \boldsymbol{X}_{pfjt} + \beta_3 \boldsymbol{Z}_{jct1} + \boldsymbol{\delta}_{jt} + \boldsymbol{\mu}_{ft} + \boldsymbol{\varepsilon}_{pfjt}\ , \qquad (5)$$

where HHI_{jct} is an employment-based measure of concentration at either the three-digit or four-digit SIC level in county c, $China\ exposure_{jt}$ is defined as the total value of imports from China to the United States scaled by the total value of shipments for industry j in year t, and the other variables are defined in equation (1).

Panel A of Table XI presents the results for estimating equation (5) using three-digit SIC industries. We find that the coefficient on China exposure is significantly positive at the 1% level, suggesting that industry-level import competition is associated with increased concentration of employers in the local labor market. In terms of economic magnitude, estimates in column (4), which includes firm-by-year fixed effects, suggest that a one standard deviation increase in Chinese import competition leads to a 0.018 (= 0.002 × 9.508) increase in HHI, which is sizable relative to the mean HHI of 0.544. We obtain similar results when we define industries based on the four-digit SIC level.[17]

Having shown that local employer concentration is higher in industries that are more exposed to imports from China, it appears natural to assume that such exposure to Chinese imports will have an indirect effect on wages through increased concentration. However, it is empirically challenging to disentangle the direct and indirect effects of *China exposure* in the same specification. Nevertheless, we use an instrumental-variables (IV) approach in which we instrument HHI using import competition from China as in the following regression:

$$log(avg.\ wages)_{pfjt} = \beta_0 + \beta_1 \widehat{HHI_{jct-1}} + \beta_2 \boldsymbol{X}_{pfjt} + \beta_3 \boldsymbol{Z}_{jct1} + \boldsymbol{\delta}_{jt} + \boldsymbol{\mu}_{ft} + \boldsymbol{\varepsilon}_{pfjt}\ , \qquad (6)$$

where $China\ exposure_{jt}$ is defined as the total value of imports from China to the United States scaled by the total value of shipments, $\widehat{HHI_{jct-1}}$ is the instrumented HHI using China exposure in industry j, county c, and year t, and the other variables are as in equation (5).

The chief caveat with the estimates in (6) is that the "instrument" *China exposure* does not meet the exclusion restriction in that it may also affect wages directly through a decline in

[17] We omit these results for brevity.

demand for labor, for example. Nevertheless, we present results from the second-stage regression (6) to illustrate the magnitude of the indirect effect of the China shock on wages. However, given the concerns about the exclusion restriction in equation (6), we are wary of drawing a causal interpretation from this analysis on the indirect effect of competition from China on wages in the United States through an employer concentration channel.

Panel B presents the results for estimating the second-stage regression in equation (6) and shows that HHI instrumented using the China import exposure is indeed significantly (at the 1% level) negatively associated with average wages. This finding is consistent with our argument that import competition from China leads to increased concentration of employers in a local labor market (presumably due to the departure or closure of existing firms), which in turn gives remaining firms in the market more power to reduce wages. The economic magnitude of the second-stage estimates is much larger than that of the baseline estimates—for example, the coefficient on HHI is −0.567 in Table XI, Panel B, column (4), while the comparable coefficient in Table II, Panel A, column (5) is −0.023. However, we interpret these relative magnitudes with caution, given that the second-stage estimates might also be affected by lower labor demand caused by import penetration from China. We also find consistent results using the four-digit SIC level as the definition of industries (unreported).

IV. CONCLUSION

We use manufacturing plant-level data from the U.S. Census Bureau from 1977 to 2009 to provide evidence that wages are significantly lower in local labor markets in which employers are more concentrated. In particular, the negative effect of employer concentration on wages appears to be concentrated in a labor market that resembles a monopsonistic market—one in which there is one significant employer of a given set of workers. We also show that this sensitivity of wages to employer concentration increases in the later years of our sample, such as 1997–2001. We argue that the results are consistent with firms exploiting workers in the form of lower wages (than a competitive market level) in monopsonistic labor markets, particularly when labor bargaining power is weak and worker mobility is limited. We suggest that the decline in U.S. unionization and labor mobility during the 1980s and 1990s is important in explaining the stagnation in wages. In addition, we show how higher employer concentration impairs the transmission of productivity growth into wage increases. Finally, we document an indirect China effect in which competition with Chinese exporters leads to a higher concentration of employers, resulting in even lower worker wages.

REFERENCES

Acemoglu, Daron, and David H. Autor, "Skills, Tasks and Technologies: Implications for Employment and Earnings," in Handbook of Labor Economics, vol. 4B, David Card and Orley Ashenfelter, eds. (Amsterdam: Elsevier, 2011), 1043–1171.

Acemoglu, Daron, and Pascual Restrepo, "The Race between Machine and Man: Implications of Technology for Growth, Factor Shares and Employment," NBER Working Paper 22252, 2016.

Autor, David H., David Dorn, and Gordon H. Hanson, "The China Syndrome: Local labor Market Effects of Import Competition in the United States," *American Economic Review,* 103 (2013), 2121–2168.

Autor, David H., David Dorn, Gordon H. Hanson, and Jae Song, "Trade Adjustment: Worker-Level Evidence," *Quarterly Journal of Economics*, 129 (2014), 1799–1860.

Autor, David, David Dorn, Lawrence F. Katz, Christina Patterson, and John Van Reenen, "The Fall of the Labor Share and the Rise of Superstar Firms," Working Paper, 2017.

Azar, Jose, Ioana Marinescu, and Marshall I. Steinbaum, "Labor Market Concentration," NBER Working Paper 24147, 2017.

Barkai, Simcha, "Declining Labor and Capital Shares," Stigler Center for the Study of the Economy and the State, New Working Paper Series, 2016.

Bartelsman, Eric, Randy Becker, and Wayne Gray, "NBER-CES Manufacturing Industry Database," National Bureau of Economic Research, 2000.

Benmelech, Efraim, Nittai Bergman, and Hyunseob Kim, "Why Did Rent-Sharing Vanish in the United States?," Working Paper, 2017.

Bivens, Josh, and Lawrence Mishel, "Understanding the Historic Divergence between Productivity and a Typical Worker's Pay," EPI Briefing Papers 406, 2015.

Boal, William M., and Michael R. Ransom, 1997, "Monopsony in the Labor Market," *Journal of Economic Literature*, 35 (1997), 86–112.

Brynjolfsson, Erik, and Andrew McAfee, The Second Machine Age: Work, Progress, and Prosperity in a Time of Brilliant Technologies (New York: W. W. Norton, 2014).

Card, David, "The Effect of Unions on the Distribution of Wages: Redistribution or Relabelling?," NBER Working Paper 4195, 1992.

Card, David, and Alan B. Krueger, "Minimum Wages and Employment: A Case Study of the

Fast-Food Industry in New Jersey and Pennsylvania," *American Economic Review*, 84 (1994), 772–793.

Council of Economic Advisers Issue Brief, "Labor Market Monopsony: Trends, Consequences, and Policy Responses," 2016.

De Locker, Jan, and Jan Eeckhout, "The Rise of Market Power and the Macroeconomic Implications," Princeton University, working paper, 2017.

Dube, Arindrajit, William T. Lester, and Michael Reich, "Minimum Wage Effects across State Borders: Estimates Using Contiguous Counties," *Review of Economics and Statistics,* 92 (2010), 945–964.

———, "Minimum Wage Shocks, Employment Flows and Labor Market Frictions," *Journal of Labor Economics,* 34 (2016), 663–704.

Ellison, Glenn, Edward L. Glaeser, and William R. Kerr, "What Causes Industry Agglomeration? Evidence from Coagglomeration Patterns," *American Economic Review*, 100 (2010), 1195–1213.

Elsby, Michael W. L., Bart Hobijn, and Ayşegül Şahin, "The Decline of the U.S. Labor Share," *Brookings Papers on Economic Activity* (2013), 1–52.

Giroud, Xavier, "Proximity and Investment: Evidence from Plant-Level Data," *Quarterly Journal of Economics,* 128 (2013), 861–915.

Greenstone, Michael, Richard Hornbeck, and Enrico Moretti, "Identifying Agglomeration Spillovers: Evidence from Winners and Losers of Large Plant Openings," *Journal of Political Economy*, 118 (2010), 536–598.

Grullon, Gustavo, Yelena Larkin, and Roni Michaely, "Are US Industries Becoming More Concentrated?," Working Paper, 2016.

Hirsch, Barry T., and David A. Macpherson, "Union Membership and Coverage Database from the Current Population Survey: Note," *Industrial and Labor Relations Review,* 56 (2003), 349–354.

Karabarbounis, Loukas, and Brent Neiman, "The Global Decline of the Labor Share," *Quarterly Journal of Economics,* 129 (2014), 61–103.

Kovenock, Dan, and Gordon Phillips, "Capital Structure and Product Market Behavior: An Examination of Plant Exit and Investment Decisions," *Review of Financial Studies,* 10 (1997), 767–803.

Lichtenberg, Frank, Corporate Takeovers and Productivity (Cambridge, MA: MIT Press, 1992).

Matsa, David, "Capital Structure as a Strategic Variable: Evidence from Collective Bargaining," *Journal of Finance,* 65 (2010), 1197–1232.

Mishel, Lawrence, "The Wedges between Productivity and Median Compensation Growth," Issue Brief, Economic Policy Institute, 2012.

Molloy, Raven, Christopher L. Smith, and Abigail Wozniak, "Declining Migration within the U.S.: The Role of the Labor Market," NBER Working Paper 20065, 2014.

Moretti, Enrico, 2011, "Local Labor Markets," in Handbook of Labor Economics, vol. 4B, David Card and Orley Ashenfelter, eds. (Amsterdam: Elsevier, 2011), 1237–1313.

Murphy, Kevin M., and Robert H. Topel, "The Evolution of Unemployment in the United States: 1968–1985," *NBER Macroeconomic Annual* (1987), 11–57.

Nickell, Stephen J., and Sushil Wadhwani, "Insider Forces and Wage Determination," *Economic Journal*, 100 (1990), 496–509.

Rognlie, Matthew, "Deciphering the Fall and Rise in the Net Capital Share: Accumulation or Scarcity?," Brookings Papers on Economic Activity, Spring 2015.

Schoar, Antoinette, "Effects of Corporate Diversification on Productivity," *Journal of Finance,* 57 (2002), 2379–2403.

Stansbury, Anna, and Lawrence Summers, "Productivity and Pay: Is the Link Broken?," Working Paper, 2017.

Uguccioni, James, "Decomposing the Productivity-Wage Nexus in Selected OECD Countries, 1986–2013," CSLS Research Report, 2016.

U.S. Department of Justice Office of Public Affairs, "Justice Department Requires eBay to End Anticompetitive 'No Poach'" Hiring Agreements," 2014.

Whitney, Lance, "Apple, Google, Others Settle Antipoaching Lawsuit for $415 Million," CNET.com, September 3, 2015.

Figure I: Trends in Average Local-Level Employment Concentration, 1977–2009

This figure plots trends in the employment-weighted average of the Herfindahl-Hirschman Index (HHI) of employment by firms computed at the county-three-digit industry-year level. The computed HHI is averaged across county-three-digit industry-year cells within each of the five-year periods (the last period includes eight years, 2002–2009) using the number of employees in each cell as the weight. Thus, the average HHI represents the degree of employer concentration the average worker faces in the labor market.

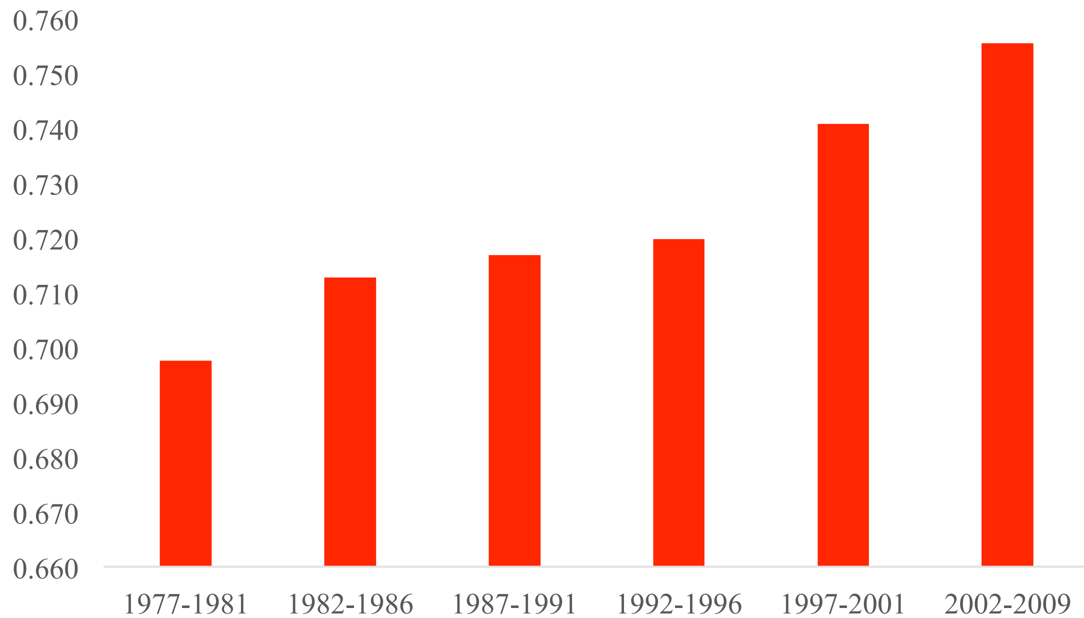

Table I: Summary Statistics on Plant Observations from the CMF and ASM Sample

This table presents descriptive statistics on the manufacturing plant-year observations used in the analysis from the CMF and ASM databases for the period 1977–2009. We require each observation in the sample to have all variables necessary to compute labor productivity (and lagged value) and value added per worker (and lagged value). "Total value of shipments" is TVS in the CMF and ASM databases and a measure of sales from plants in million dollars; "Total wage" is the sum of wages for production and nonproduction workers in million dollars; "Total employees" is the number of total employees; "Total labor hours" is the production worker equivalent man hours in thousands; "HHI (SIC3 or 4-county-year)" is the Herfindahl-Hirschman Index (HHI) of employment by firms at the county-industry (three- or four-digit)-year level; "Dummy HHI (SIC3 or 4-county-year) = 1" is a dummy variable equal to one if HHI = 1, and zero otherwise; "HHI (SIC3 or 4- year)" is the Herfindahl-Hirschman Index (HHI) of employment by firms at the industry (three- or four-digit)-year level; "log(employment, SIC3 or 4-county-year)" is the log number of employees at the county-industry (three- or four-digit)-year level; "Labor productivity" is defined as output divided by total labor hours (a quantity-based measure of labor productivity); "Average wage" is computed as total wage divided by total employees (in thousand dollars); "log(employment, SIC3 or 4-county-year)" is the log number of employment at the county-industry (three- or four-digit)-year level; "Plants per segment" is the number of plants in a given three-digit SIC industry segment of a given firm; "Plants per firm" is the total number of plants of a given firm; "Plant age" is the number of years since a plant's birth, which is proxied either by the flag for plant birth in the Longitudinal Business Database (LBD) or by its first appearance in the CMF or ASM database, whichever is earliest; and "Unionization rate" is the percentage of the workforce covered by collective bargaining collected by Hirsch and Macpherson (2003). The number of observations is rounded to the nearest thousand to follow the Census Bureau's disclosure rules.

	Mean	STD
Total value of shipment ($m)	95.95	379.45
Total wage ($m)	16.51	50.63
Total employees	348.56	816.68
Total labor hours (000)	817.25	3955.94
HHI (SIC3-county-year)	0.545	0.350
HHI (SIC4-county-year)	0.682	0.334
Dummy HHI (SIC3-county-year) = 1	0.227	0.419
Dummy HHI (SIC4-county-year) = 1	0.379	0.485
HHI (SIC3-year)	0.022	0.029
HHI (SIC4-year)	0.044	0.048
Log labor productivity	4.61	0.94
Average wage ($000)	41.84	14.24
Log average wage ($000)	3.67	0.35
log(employment, SIC3-county-year)	6.24	1.68
log(employment, SIC4-county-year)	5.66	1.76
Plants per segment (SIC3)	9.85	15.20
Plants per firm	31.89	39.14
Plant age	17.09	8.99
Unionization rate	0.228	0.129
Observations (plant-years)	656,000	—

Table II: Local Employer Concentration and Wages

This table examines the basic effects of employer concentration in a local labor market on the wages of plants. The dependent variable is the log of average wages per worker as defined in Table I. "HHI" and "log(employment)" are lagged by one year. Panels A and B present estimates using three- and four-digit SIC industries to compute HHI. The *t*-statistics based on standard errors adjusted for sample clustering at the county level are reported below the coefficient estimates. The numbers of observations are rounded to the nearest thousand to follow the Census Bureau's disclosure rules.

Panel A: 3-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Var.	Log avg. wages					
HHI (SIC3-county-year)	−0.025	−0.044	−0.037	−0.028	−0.023	−0.049
	−2.87	**−5.41**	**−5.29**	**−5.64**	**−4.31**	**−9.41**
log(employment, SIC3-county-year)	0.042	0.039	0.036	0.026	0.027	0.026
	19.08	**17.16**	**21.62**	**26.99**	**26.74**	**27.28**
log(labor productivity)	0.156	0.155	0.102	0.095	0.094	0.065
	47.19	**46.31**	**55.77**	**49.09**	**44.94**	**36.78**
log(plant per segment)	—	−0.042	−0.015	−0.017	−0.020	−0.008
	—	**−19.20**	**−10.68**	**−11.26**	**−11.42**	**−5.08**
log(plant per firm)	—	0.041	0.021	−0.016	—	0.000
	—	**26.63**	**19.28**	**−8.96**	—	**0.00**
Plant age (/100)	—	0.422	0.425	0.405	0.412	0.358
	—	**17.61**	**24.27**	**23.23**	**23.31**	**21.94**
Year fixed effects	Y	Y	Y	Y		
Industry fixed effects			Y			
Industry-year fixed effects						Y
Firm fixed effects				Y		
Firm-year fixed effects					Y	Y
Observations	656,000	656,000	656,000	656,000	656,000	656,000
R^2	20.45%	22.56%	43.90%	54.99%	62.55%	67.18%

Panel B: 4-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Var.			Log avg. wages			
HHI (SIC4-county-year)	−0.042	−0.063	−0.054	−0.043	−0.038	−0.055
	−4.21	**−6.51**	**−8.24**	**−8.34**	**−6.90**	**−10.73**
log(employment, SIC4-county-year)	0.034	0.031	0.026	0.020	0.021	0.019
	18.93	**17.36**	**27.84**	**25.25**	**24.82**	**28.54**
log(labor productivity)	0.152	0.151	0.090	0.094	0.093	0.060
	43.24	**43.51**	**53.47**	**47.27**	**43.40**	**35.08**
log(plant per segment)	—	−0.036	−0.013	−0.015	−0.017	−0.007
	—	**−18.51**	**−11.16**	**−11.01**	**−11.13**	**−5.01**
log(plant per firm)	—	0.034	0.018	−0.019	0.000	0.000
	—	**24.81**	**18.03**	**−11.21**	**0.00**	**0.00**
Plant age (/100)	—	0.423	0.430	0.414	0.421	0.374
	—	**17.30**	**26.31**	**23.67**	**23.60**	**23.32**
Year fixed effects	Y	Y	Y	Y		
Industry fixed effects			Y			Y
Firm fixed effects				Y		
Firm-year fixed effects					Y	Y
Observations	656,000	656000	656,000	656000	656,000	656,000
R^2	19.44%	21.28%	48.37%	54.67%	62.23%	69.22%

Table III: Local Employer Concentration and Wages Controlling for Labor Value-Added

This table examines the effects of employer concentration in a local labor market on the wages of plants including an additional control for labor productivity—valued added (total value of shipments + net increase in inventories of finished goods and works in progress—material and energy costs) scaled by labor hours. The dependent variable is the log of average wages per worker as defined in Table I. "HHI" and "log(employment)" are lagged by one year. Panels A and B present estimates using three- and four-digit SIC industries to compute HHI. The *t*-statistics based on standard errors adjusted for sample clustering at the county level are reported below the coefficient estimates. The numbers of observations are rounded to the nearest thousand to follow the Census Bureau's disclosure rules.

Panel A: 3-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Var.			Log avg. wages			
HHI (SIC3-county-year)	−0.025	−0.044	−0.038	−0.028	−0.023	−0.049
	−2.91	**−5.45**	**−5.37**	**−5.61**	**−4.29**	**−9.47**
log(employment, SIC3-county-year)	0.041	0.038	0.036	0.027	0.027	0.026
	17.95	**16.16**	**21.30**	**26.99**	**26.74**	**27.10**
log(labor productivity)	0.144	0.144	0.098	0.097	0.096	0.061
	43.59	**43.70**	**48.87**	**45.73**	**41.63**	**30.01**
log(labor VA)	0.016	0.015	0.005	−0.002	−0.001	0.004
	9.86	**9.14**	**5.26**	**−2.65**	**−1.49**	**5.35**
log(plant per segment)	—	−0.041	−0.014	−0.017	−0.020	−0.007
	—	**−19.01**	**−10.57**	**−11.32**	**−11.46**	**−5.03**
log(plant per firm)	—	0.041	0.021	−0.016	—	—
	—	**26.65**	**19.27**	**−8.91**	—	—
Plant age (/100)	—	0.424	0.424	0.405	0.413	0.357
	—	**17.64**	**24.23**	**23.25**	**23.34**	**21.93**
Year fixed effects	Y	Y	Y	Y		
Industry fixed effects			Y			
Industry-year fixed effects						Y
Firm fixed effects				Y		
Firm-year fixed effects					Y	Y
Observations	656,000	656,000	656,000	656,000	656,000	656,000

R^2	20.67%	22.74%	43.92%	55.00%	62.55%	67.19%

Panel B: 4-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Var.			Log avg. wages			
HHI (SIC4-county-year)	−0.042	−0.063	−0.055	−0.043	−0.038	−0.055
	−4.28	**−6.58**	**−8.36**	**−8.33**	**−6.90**	**−10.82**
log(employment, SIC4-county-year)	0.033	0.030	0.026	0.020	0.021	0.019
	17.77	**16.24**	**27.42**	**25.21**	**24.77**	**28.22**
log(labor productivity)	0.139	0.138	0.082	0.095	0.094	0.054
	38.94	**39.93**	**39.99**	**43.71**	**39.83**	**26.51**
log(labor VA)	0.019	0.018	0.009	−0.001	−0.001	0.006
	11.08	**10.57**	**10.04**	**−1.59**	**−0.60**	**7.93**
log(plant per segment)	—	−0.035	−0.013	−0.015	−0.017	−0.007
	—	**−18.14**	**−11.14**	**−11.07**	**−11.16**	**−5.01**
log(plant per firm)	—	0.034	0.018	−0.019	—	—
	—	**24.89**	**18.09**	**−11.17**	—	—
Plant age (/100)	—	0.425	0.428	0.414	0.421	0.372
	—	**17.36**	**26.19**	**23.69**	**23.62**	**23.30**
Year fixed effects	Y	Y	Y	Y		
Industry fixed effects			Y			
Industry-year fixed effects						Y
Firm fixed effects				Y		
Firm-year fixed effects					Y	Y
Observations	656,000	656,000	656,000	656,000	656,000	656,000
R^2	19.73%	21.53%	48.41%	54.67%	62.23%	69.24%

Table IV: Subsample of Firms with One Industry Segment across Multiple Plants

This table examines the effects of employer concentration in a local labor market on the wages of plants using a subsample of plants owned by firms that have multiple plants in only one industry segment. The dependent variable is the log of average wages per worker as defined in Table I. "HHI" and "log(employment)" are lagged by one year. Panels A and B present estimates using three- and four-digit SIC industries to compute HHI. The *t*-statistics based on standard errors adjusted for sample clustering at the county level are reported below the coefficient estimates. The numbers of observations are rounded to the nearest thousand to follow the Census Bureau's disclosure rules.

Panel A: 3-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Var.			Log avg. wages			
HHI (SIC3-county-year)	−0.065	−0.080	−0.071	−0.058	−0.061	−0.070
	−5.29	**−6.58**	**−6.90**	**−7.14**	**−5.48**	**−6.59**
log(employment, SIC3-county-year)	0.034	0.032	0.025	0.019	0.019	0.018
	14.81	**13.81**	**12.79**	**13.88**	**10.88**	**10.70**
log(labor productivity)	0.147	0.142	0.110	0.094	0.097	0.076
	37.06	**35.74**	**32.87**	**30.28**	**22.31**	**16.83**
log(plant per firm)	—	0.031	0.022	−0.014	—	—
	—	**14.13**	**12.49**	**−3.91**	**—**	**—**
Plant age (/100)	—	0.288	0.327	0.333	0.353	0.335
	—	**7.19**	**10.18**	**9.26**	**6.98**	**7.04**
Year fixed effects	Y	Y	Y	Y		
Industry fixed effects			Y			
Industry-year fixed effects						Y
Firm fixed effects				Y		
Firm-year fixed effects					Y	Y
Observations	110,000	110,000	110,000	110,000	110,000	110,000
R^2	18.77%	20.13%	38.09%	56.32%	72.42%	77.58%

Panel B: 4-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Var.			Log avg. wages			
HHI (SIC4-county-year)	−0.080	−0.096	−0.078	−0.069	−0.071	−0.078
	−5.70	**−6.82**	**−8.39**	**−8.03**	**−6.74**	**−8.07**
log(employment, SIC4-county-year)	0.026	0.024	0.019	0.014	0.014	0.013
	15.30	**14.02**	**17.17**	**13.44**	**10.75**	**11.27**
log(labor productivity)	0.140	0.135	0.103	0.092	0.093	0.074
	33.28	**32.16**	**34.56**	**28.96**	**23.57**	**19.04**
log(plant per firm)	—	0.030	0.019	−0.015	—	—
	—	**15.55**	**13.14**	**−4.63**	**—**	**—**
Plant age (/100)	—	0.323	0.366	0.346	0.351	0.319
	—	**8.63**	**13.12**	**10.53**	**8.36**	**8.16**
Year fixed effects	Y	Y	Y	Y		
Industry fixed effects			Y			
Industry-year fixed effects						Y
Firm fixed effects				Y		
Firm-year fixed effects					Y	Y
Observations	154,000	154,000	154,000	154,000	154,000	154,000
R^2	17.12%	18.61%	43.24%	55.03%	68.79%	77.54%

Table V: Employer Concentration and Wages by Five-Year Time Period

This table examines the basic effects of employer concentration in a local labor market on the wages of plants by five-year period from 1977 to 2009. The dependent variable is the log of average wages per worker as defined in Table I. "HHI" and "log(employment)" are lagged by one year. Panels A and B present estimates using three- and four-digit SIC industries to compute HHI. The *t*-statistics based on standard errors adjusted for sample clustering at the county level are reported below the coefficient estimates. The numbers of observations are rounded to the nearest thousand to follow the Census Bureau's disclosure rules.

Panel A: 3-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Subsample period:	1977–1981	1982–1986	1987–1991	1992–1996	1997–2001	2002–2009
Dep. Var.:			Log avg. wages			
HHI (SIC3-county-year)	–0.018	–0.012	–0.015	–0.031	–0.039	–0.029
	–2.75	**–1.78**	**–1.95**	**–4.59**	**–5.31**	**–4.18**
log(employment, SIC3-county-year)	0.029	0.031	0.030	0.025	0.021	0.023
	24.35	**23.09**	**20.38**	**16.45**	**10.98**	**11.75**
log(labor productivity)	0.122	0.109	0.095	0.089	0.072	0.080
	40.09	**36.56**	**30.13**	**29.30**	**23.02**	**29.07**
log(plant per segment)	–0.009	–0.014	–0.024	–0.028	–0.020	–0.025
	–3.87	**–5.32**	**–10.36**	**–11.70**	**–8.37**	**–10.21**
Plant age (/100)	1.681	0.971	0.591	0.514	0.473	0.267
	16.24	**16.31**	**12.85**	**18.37**	**18.71**	**14.48**
Firm-year fixed effects	Y	Y	Y	Y	Y	Y
Observations	114,000	87,000	101,000	110,000	102,000	143,000
R^2	66.02%	65.65%	62.84%	62.30%	56.36%	60.87%

Panel B: 4-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Subsample period:	1977–1981	1982–1986	1987–1991	1992–1996	1997–2001	2002–2009
Dep. Var.:			Log avg. wages			
HHI (SIC4-county-year)	–0.044	–0.038	–0.031	–0.044	–0.047	–0.033
	–6.87	**–5.51**	**–3.88**	**–6.08**	**–6.15**	**–4.67**
log(employment, SIC4-county-year)	0.022	0.022	0.024	0.018	0.017	0.021
	21.97	**18.87**	**18.10**	**13.78**	**10.05**	**11.81**
log(labor productivity)	0.121	0.107	0.093	0.088	0.072	0.080
	40.01	**35.11**	**28.54**	**27.87**	**22.61**	**28.92**
log(plant per segment)	–0.008	–0.012	–0.023	–0.024	–0.017	–0.021
	–3.63	**–5.34**	**–10.42**	**–10.67**	**–7.42**	**–8.40**
Plant age (/100)	1.728	1.001	0.603	0.538	0.476	0.264
	16.59	**16.44**	**12.78**	**19.09**	**18.71**	**14.32**
Firm-year fixed effects	Y	Y	Y	Y	Y	Y
Observations	114,000	87,000	101,000	110,000	102,000	143,000
R^2	65.70%	65.24%	62.51%	61.91%	56.12%	60.70%

Table VI: Local Employer Concentration, Unions, and Wages

This table examines the interactive effects of employer concentration in a local labor market and industry unionization rates on the wages of plants. The dependent variable is the log of average wages per worker as defined in Table I. "HHI," "log(employment)," and "Union" are lagged by one year. Panels A and B present estimates using three- and four-digit SIC industries to compute HHI. The *t*-statistics based on standard errors adjusted for sample clustering at the county level are reported below the coefficient estimates. The numbers of observations are rounded to the nearest thousand to follow the Census Bureau's disclosure rules.

Panel A: 3-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Var.	Log avg. wages					
HHI (SIC3-county-year)	−0.135	−0.153	−0.070	−0.094	−0.087	−0.079
	−8.05	**−9.40**	**−6.41**	**−11.62**	**−10.39**	**−9.86**
log(employment, SIC3-county-year)	0.039	0.036	0.036	0.025	0.025	0.026
	17.68	**15.92**	**21.86**	**25.31**	**25.09**	**27.63**
log(labor productivity)	0.148	0.148	0.102	0.092	0.092	0.065
	48.73	**48.21**	**55.77**	**49.09**	**44.90**	**36.81**
Union	0.108	0.109	0.136	0.097	0.109	0.169
	2.14	**2.15**	**4.96**	**4.47**	**4.78**	**1.45**
HHI x Union	0.407	0.402	0.145	0.237	0.228	0.131
	7.29	**7.20**	**4.09**	**8.99**	**8.59**	**5.59**
log(plant per segment)	—	−0.044	−0.015	−0.018	−0.020	−0.008
	—	**−19.97**	**−10.82**	**−11.98**	**−11.98**	**−5.14**
log(plant per firm)	—	0.040	0.021	−0.016	—	—
	—	**26.85**	**19.28**	**−8.79**	—	—
Plant age (/100)	—	0.408	0.427	0.397	0.405	0.358
	—	**17.45**	**24.58**	**22.88**	**22.86**	**21.88**
Year fixed effects	Y	Y	Y	Y		
Industry fixed effects			Y			
Industry-year fixed effects						Y
Firm fixed effects				Y		
Firm-year fixed effects					Y	Y
Observations	656,000	656,000	656,000	656,000	656,000	656,000

R^2	21.89%	23.98%	43.98%	55.37%	62.88%	67.20%

Panel B: 4-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Var.			Log avg. wages			
HHI (SIC4-county-year)	−0.111	−0.131	−0.072	−0.086	−0.080	−0.074
	−5.40	**−6.52**	**−6.50**	**−9.07**	**−8.41**	**−8.21**
log(employment, SIC4-county-year)	0.031	0.029	0.026	0.019	0.020	0.019
	17.23	**15.75**	**28.03**	**23.29**	**22.86**	**28.41**
log(labor productivity)	0.145	0.144	0.090	0.091	0.091	0.060
	44.23	**44.66**	**53.28**	**46.90**	**43.04**	**34.98**
Union	0.160	0.162	0.145	0.116	0.127	—
	2.52	**2.58**	**4.67**	**4.36**	**4.71**	**—**
HHI × Union	0.253	0.251	0.078	0.156	0.151	0.084
	3.84	**3.84**	**2.10**	**5.25**	**5.18**	**3.16**
log(plant per segment)	—	−0.038	−0.013	−0.017	−0.019	−0.007
	—	**−19.74**	**−11.16**	**−12.08**	**−12.02**	**−5.01**
log(plant per firm)	—	0.033	0.018	−0.018	—	—
	—	**25.00**	**17.99**	**−10.75**	**—**	**—**
Plant age (/100)	—	0.405	0.431	0.406	0.414	0.373
	—	**16.87**	**26.44**	**23.32**	**23.18**	**23.24**
Year fixed effects	Y	Y	Y	Y		
Industry fixed effects			Y			
Industry-year fixed effects						Y
Firm fixed effects				Y		
Firm-year fixed effects					Y	Y
Observations	656,000	656,000	656,000	656,000	656,000	656,000
R^2	20.60%	23.98%	43.98%	55.37%	62.88%	62.51%

Table VII: Employer Concentration and Sensitivities of Wage Changes to Productivity Changes

This table examines how employer concentration shapes sensitivities of changes in wages to changes in labor productivity. The dependent variable is the log change in average wages per worker as defined in Table I. "HHI" is lagged by one year. Panels A and B present estimates using three- and four-digit SIC industries to compute HHI. The *t*-statistics based on standard errors adjusted for sample clustering at the county level are reported below the coefficient estimates. The numbers of observations are rounded to the nearest thousand to follow the Census Bureau's disclosure rules.

Panel A: 3-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Var.			ΔLog avg. wages			
HHI (SIC3-county-year)	0.006	0.006	0.006	0.005	0.005	0.005
	10.40	**10.99**	**10.99**	**10.59**	**9.56**	**9.81**
Δlog(labor productivity)	0.002	0.002	0.002	0.002	0.002	0.003
	2.50	**2.63**	**3.22**	**2.47**	**2.64**	**3.42**
HHI × Δlog(labor productivity)	−0.002	−0.002	−0.002	−0.002	−0.002	−0.002
	−1.70	**−1.82**	**−1.73**	**−1.72**	**−1.96**	**−1.90**
log(plant per segment)	—	−0.001	−0.001	0.000	0.000	0.000
	—	**−3.13**	**−2.62**	**−0.88**	**−1.34**	**−0.84**
log(plant per firm)	—	0.000	0.000	0.000	—	—
	—	**1.47**	**1.95**	**−0.61**	—	—
Plant age (/100)	—	−0.033	−0.032	−0.030	−0.029	−0.028
	—	**−10.98**	**−10.56**	**−9.81**	**−8.62**	**−7.87**
Year fixed effects	Y	Y	Y	Y		
Industry fixed effects			Y			
Industry-year fixed effects						Y
Firm fixed effects				Y		
Firm-year fixed effects					Y	Y
Observations	656,000	656,000	656,000	656,000	656,000	656,000
R^2	0.98%	1.00%	1.05%	3.80%	24.36%	25.18%

Panel B: 4-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)	(6)
Dep. Var.			ΔLog avg. wages			
HHI (SIC4-county-year)	0.007	0.007	0.007	0.006	0.005	0.005
	10.98	**11.16**	**9.74**	**10.08**	**8.46**	**8.52**
Δlog(labor productivity)	0.003	0.003	0.003	0.003	0.003	0.004
	2.63	**2.75**	**3.38**	**2.52**	**2.65**	**3.43**
HHI × Δlog(labor productivity)	−0.003	−0.003	−0.003	−0.003	−0.003	−0.003
	−2.04	**−2.14**	**−2.05**	**−1.94**	**−2.13**	**−2.11**
log(plant per segment)	—	−0.001	0.000	0.000	0.000	0.000
	—	**−3.56**	**−0.55**	**−1.32**	**−1.47**	**−0.41**
log(plant per firm)	—	0.000	0.000	0.000	—	—
	—	**1.43**	**1.38**	**−0.54**	—	—
Plant age (/100)	—	−0.033	−0.032	−0.030	−0.029	−0.027
	—	**−10.95**	**−10.73**	**−9.81**	**−8.62**	**−7.72**
Year fixed effects	Y	Y	Y	Y		
Industry fixed effects			Y			
Industry-year fixed effects						Y
Firm fixed effects				Y		
Firm-year fixed effects					Y	Y
Observations	656,000	656,000	656,000	656,000	656,000	656,000
R^2	0.99%	1.00%	1.10%	3.79%	24.36%	26.74%

Table VIII: Nonlinear Effect of Local Employer Concentration on Wages

This table examines nonlinear effects of employer concentration in a local labor market on the wages of plants. The dependent variable is the log of average wages per worker as defined in Table I. "HHI," "log(employment)," and "Union" are lagged by one year. "Dummy HHI (SIC3 or 4-county-year) = 1" is equal to one if HHI based on three- or four-digit SIC industries is equal to one, and zero otherwise. Panels A and B present estimates using three- and four-digit SIC industries to compute HHI. The *t*-statistics based on standard errors adjusted for sample clustering at the county level are reported below the coefficient estimates. The numbers of observations are rounded to the nearest thousand to follow the Census Bureau's disclosure rules.

Panel A: 3-digit SIC Industries

	(1)	(2)	(3)
Dep. Var.		Log avg. wages	
Dummy HHI (SIC3-county-year) = 1	−0.017	−0.011	−0.062
	−5.19	**−2.79**	**−11.59**
HHI (SIC3-county-year)	—	−0.013	—
	—	**−1.99**	**—**
log(employment, SIC3-county-year)	0.028	0.027	0.027
	26.12	**26.47**	**24.63**
log(labor productivity)	0.094	0.094	0.092
	44.86	**44.88**	**44.64**
Union	—	—	0.196
	—	**—**	**11.81**
HHI × Union	—	—	0.162
	—	**—**	**9.23**
log(plant per segment)	−0.020	−0.020	−0.020
	−11.59	**−11.55**	**−12.00**
Plant age (/100)	0.409	0.412	0.399
	23.29	**23.25**	**22.73**
Firm-year fixed effects	Y	Y	Y
Observations	656,000	656,000	656,000
R^2	62.55%	62.55%	62.87%

Panel B: 4-digit SIC Industries

	(1)	(2)	(3)
Dep. Var.		Log avg. wages	
Dummy HHI (SIC4-county-year) = 1	−0.031	−0.029	−0.058
	−9.74	**−7.43**	**−10.94**
HHI (SIC4-county-year)	—	−0.006	—
	—	**−0.78**	**—**
log(employment, SIC4-county-year)	0.021	0.021	0.020
	22.34	**23.15**	**20.24**
log(labor productivity)	0.093	0.093	0.091
	43.27	**43.23**	**42.85**
Union	—	—	0.186
	—	**—**	**10.08**
HHI × Union	—	—	0.101
	—	**—**	**6.08**
log(plant per segment)	−0.018	−0.018	−0.020
	−11.35	**−11.34**	**−12.19**
Plant age (/100)	0.419	0.419	0.408
	23.39	**23.5**	**22.87**
Firm-year fixed effects	Y	Y	Y
Observations	656,000	656,000	656,000
R^2	62.28%	62.28%	62.56%

Table IX: Local Employer Concentration and Wages Controlling for National Concentration

This table examines the effects of employer concentration in a local labor market on the wages of plants, including an additional control for employer concentration at the national level. The dependent variable is the log of average wages per worker as defined in Table I. "HHI" and "log(employment)" are lagged by one year. Panels A and B present estimates using three- and four-digit SIC industries to compute HHI. The *t*-statistics based on standard errors adjusted for sample clustering at the county level are reported below the coefficient estimates. The numbers of observations are rounded to the nearest thousand to follow the Census Bureau's disclosure rules.

Panel A: 3-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)
Dep. Var.			Log avg. wages		
HHI (SIC3-county-year)	−0.069	−0.088	−0.037	−0.047	−0.043
	−7.88	**−10.40**	**−5.27**	**−9.39**	**−7.90**
HHI (SIC3-year)	1.556	1.527	−0.055	0.940	0.995
	24.68	**24.44**	**−0.91**	**19.58**	**17.57**
log(employment, SIC3-county-year)	0.036	0.033	0.036	0.024	0.025
	16.55	**14.82**	**21.68**	**24.21**	**23.88**
log(labor productivity)	0.151	0.150	0.102	0.093	0.092
	46.58	**45.89**	**55.75**	**48.31**	**43.93**
log(plant per segment)	—	−0.043	−0.015	−0.019	−0.021
	—	**−19.43**	**−10.61**	**−11.89**	**−11.94**
log(plant per firm)	—	0.040	0.021	−0.016	0.000
	—	**26.20**	**19.32**	**−8.69**	**0.00**
Plant age (/100)	—	0.413	0.425	0.400	0.408
	—	**17.03**	**24.32**	**22.72**	**22.83**
Year fixed effects	Y	Y	Y	Y	
Industry fixed effects			Y		
Industry-year fixed effects					
Firm fixed effects				Y	
Firm-year fixed effects					Y
Observations	656,000	656,000	656,000	656,000	656,000
R^2	21.88%	23.94%	43.90%	55.34%	62.88%

Panel B: 4-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)
Dep. Var.			Log avg. wages		
HHI (SIC4-county-year)	−0.092	−0.110	−0.054	−0.060	−0.055
	−8.78	**−10.75**	**−8.20**	**−11.28**	**−9.67**
HHI (SIC4-year)	0.994	0.963	−0.157	0.442	0.464
	30.62	**29.74**	**−4.65**	**16.06**	**14.88**
log(employment, SIC4-county-year)	0.028	0.026	0.026	0.019	0.020
	16.87	**15.47**	**27.82**	**23.86**	**23.33**
log(labor productivity)	0.146	0.147	0.090	0.092	0.092
	42.29	**43.25**	**53.21**	**46.63**	**42.43**
log(plant per segment)	—	−0.038	−0.013	−0.016	−0.019
	—	**−19.59**	**−11.00**	**−11.74**	**−11.79**
log(plant per firm)	—	0.031	0.018	−0.018	0.000
	—	**23.83**	**18.13**	**−11.18**	**0.00**
Plant age (/100)	—	0.393	0.431	0.405	0.412
	—	**16.06**	**26.43**	**22.96**	**22.91**
Year fixed effects	Y	Y	Y	Y	
Industry fixed effects			Y		
Industry-year fixed effects					
Firm fixed effects				Y	
Firm-year fixed effects					Y
Observations	656,000	656,000	656,000	656,000	656,000
R^2	21.07%	22.78%	48.38%	54.88%	62.44%

Table X: China Import Penetration, Local Employer Concentration, and Wages

This table examines the effects of employer concentration in a local labor market on wages controlling for the effect of imports from China on wages. The dependent variable is the log of average wages per worker as defined in Table I. "HHI" and "log(employment)" are lagged by one year. "China exposure" is defined as total value of import from China to the U.S. scaled by total value of shipments at the industry by year level. Panels A and B present estimates using three- and four-digit SIC industries to compute the HHI. The *t*-statistics based on standard errors adjusted for sample clustering at the county level are reported below the coefficient estimates. The numbers of observations are rounded to the nearest thousands to follow the Census Bureau's disclosure rules.

Panel A: 3-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)
Dep. Var.			Log avg. wages		
HHI (SIC3-county-year)	−0.041	−0.065	−0.035	−0.032	−0.043
	−4.36	**−7.25**	**−6.39**	**−5.31**	**−4.93**
China exposure (SIC3-year)	−0.001	−0.001	−0.001	−0.001	−0.001
	−4.01	**−5.96**	**−5.01**	**−5.01**	**−7.22**
log(employment, SIC3-county-year)	0.039	0.035	0.023	0.024	0.033
	14.29	**12.14**	**17.81**	**17.03**	**15.63**
log(labor productivity)	0.138	0.137	0.080	0.082	0.134
	36.94	**37.33**	**36.75**	**34.18**	**41.29**
log(plant per segment)	—	−0.051	−0.023	−0.025	−0.048
	—	**−24.25**	**−13.63**	**−13.28**	**−23.96**
log(plant per firm)	—	0.043	−0.011	0.000	0.042
	—	**24.47**	**−4.75**	**0.00**	**24.61**
Plant age (/100)	—	0.432	0.384	0.388	0.423
	—	**17.79**	**21.92**	**21.18**	**20.04**
Year fixed effects	Y	Y	Y		Y
Industry fixed effects					
Firm fixed effects			Y		
Firm-year fixed effects				Y	
County fixed effects					Y
Observations	346,000	346,000	346,000	346,000	346,000
R^2	15.60%	18.35%	54.05%	60.36%	21.18%

Panel B: 4-digit SIC Industries

	(1)	(2)	(3)	(4)	(5)
Dep. Var.			Log avg. wages		
HHI (SIC4-county-year)	−0.051	−0.076	−0.046	−0.042	−0.046
	−4.53	**−7.13**	**−7.69**	**−6.57**	**−4.71**
China exposure (SIC4-year)	−0.001	−0.001	−0.001	−0.001	−0.002
	−5.76	**−7.75**	**−4.75**	**−4.56**	**−9.09**
log(employment, SIC4-county-year)	0.031	0.027	0.018	0.019	0.027
	13.88	**11.78**	**16.45**	**15.59**	**15.26**
log(labor productivity)	0.134	0.135	0.079	0.081	0.131
	34.20	**35.04**	**35.57**	**33.13**	**39.55**
log(plant per segment)	—	−0.044	−0.020	−0.022	−0.041
	—	**−22.12**	**−12.28**	**−11.96**	**−21.86**
log(plant per firm)	—	0.034	−0.015	0.000	0.033
	—	**21.24**	**−6.38**	**0.00**	**21.66**
Plant age (/100)	—	0.434	0.392	0.395	0.421
	—	**17.66**	**22.33**	**21.47**	**19.72**
Year fixed effects	Y	Y	Y		Y
Industry fixed effects					
Firm fixed effects			Y		
Firm-year fixed effects				Y	
County fixed effects					Y
Observations	346,000	346,000	346,000	346000	346,000
R^2	14.65%	17.09%	53.73%	60.04%	20.31%

Table XI: Indirect of Effect of China Import Penetration on Wages through Employer Concentration

This table examines how exposure to China imports indirectly affects wages through employer concentration using an instrumental variables (IV) approach. This table uses three-digit SIC industries to compute HHI and China exposure. Panels A and B present estimates for the first and second stages. In Panels A and B, the dependent variables are HHI and the log of average wages per worker as defined in Table I. "HHI" and "log(employment)" are lagged by one year. The *t*-statistics based on standard errors adjusted for sample clustering at the county level are reported below the coefficient estimates. The numbers of observations are rounded to the nearest thousand to follow the Census Bureau's disclosure rules.

Panel A: China Exposure and Local Employer Concentration (First Stage)

	(1)	(2)	(3)	(4)	(5)
Dep. Var.			HHI (SIC3-county-year)		
China exposure (SIC3-year)	0.003	0.003	0.001	0.002	0.003
	11.32	**11.41**	**5.42**	**6.49**	**12.39**
log(employment, SIC3-county-year)	−0.119	−0.120	−0.130	−0.132	−0.104
	−24.59	**−25.15**	**−37.02**	**−35.97**	**−48.25**
log(labor productivity)	−0.009	−0.011	−0.003	−0.002	−0.004
	−2.55	**−3.20**	**−1.17**	**−0.76**	**−1.80**
log(plant per segment)	—	−0.019	−0.001	0.000	−0.020
	—	**−7.45**	**−0.57**	**−0.04**	**−9.10**
log(plant per firm)	—	0.021	−0.022	—	0.020
	—	**11.30**	**−8.57**	**—**	**12.17**
Plant age (/100)	—	0.397	0.270	0.273	0.386
	—	**13.82**	**12.01**	**11.23**	**15.19**
Year fixed effects	Y	Y	Y		Y
Firm fixed effects			Y		
Firm-year fixed effects				Y	
County fixed effects					Y
Observations	346,000	346,000	346,000	346,000	346,000
R^2	29.94%	31.10%	51.40%	54.65%	36.23%

Panel B: Second Stage

	(1)	(2)	(3)	(4)	(5)
Dep. Var.			Log avg. wages		
Inst'd HHI (SIC3-county-year)	−0.327	−0.493	−0.660	−0.567	−0.545
	−4.27	**−6.07**	**−3.22**	**−3.56**	**−6.97**
log(employment, SIC3-county-year)	0.005	−0.017	−0.058	−0.047	−0.015
	0.46	**−1.47**	**−2.20**	**−2.25**	**−3.08**
log(labor productivity)	0.135	0.133	0.078	0.080	0.116
	33.99	**31.48**	**22.78**	**23.47**	**35.58**
log(plant per segment)	—	−0.059	−0.024	−0.025	−0.050
	—	**−20.85**	**−5.01**	**−5.51**	**−18.64**
log(plant per firm)	—	0.051	−0.025	—	0.049
	—	**19.11**	**−3.31**	**—**	**19.22**
Plant age (/100)	—	0.602	0.553	0.534	0.564
	—	**13.62**	**9.25**	**10.92**	**17.34**
Year fixed effects	Y	Y	Y		Y
Firm fixed effects			Y		
Firm-year fixed effects				Y	
County fixed effects					Y
Observations	346,000	346,000	346,000	346,000	346,000
R^2	9.96%	5.86%	3.59%	5.26%	10.29%