HOW DOES DECLINING WORKER POWER AFFECT INVESTMENT SENSITIVITY TO MINIMUM WAGE?*

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ABSTRACT

Declining worker power has been advanced as an explanation for dramatic generational changes in the U.S. macroeconomic environment such as the substantial decline in labor's share of the national income, the loss of consumer purchasing power, and growing income and wealth inequality. In this paper, we consider microeconomic implications by examining the extent to which declining working power affects firm-level investment decisions as reflected in firm responses to mandated increases in the minimum wage. Over our sample period, we find that investment-wage sensitivities go from negative (when worker power constrains management) to insignificant (when management becomes less constrained and can pursue outside options). We also provide evidence on the channel through which declining worker power affects firm investment responses, by showing that changes in investment-wage sensitivities are more significant for firms that are more exposed to globalization, technological change, and declining unionization.

KEYWORDS: DECLINING WORKER POWER, CORPORATE INVESTMENT, MINIMUM WAGE, GLOBALIZATION, US-CHINA

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I. Introduction

Labor's share of national income in the U.S. has been declining since the 1980s (Karabarbounis and Neiman, 2014; Gutiérrez and Piton, 2020). The decline has been the focus of much public and academic scrutiny due to its importance in understanding income and wealth inequality, slowing income growth, and the loss of consumer purchasing power - which is an important driver of the economy. Explanations for the decline have largely focused on decreasing worker power vis-a-vis their employers.[1] As discussed in Stansbury and Summers (2020), the decline in worker power reflects not only the significant decline in unionization over this period, but also from increased employer bargaining power due to improved outside options made possible by technological advances and by the ability to substitute labor from low wage countries arising from globalization. In addition, increases in shareholder power and shareholder activism have led to pressures on companies to cut labor costs.

Much of the academic literature has focused on the implications of declining worker power for the macroeconomy. In this paper, we consider microeconomic implications by examining the extent to which declining worker power has affected firm-level investment decisions. Our empirical strategy focuses on the effect of worker power on firm investment responses to mandated changes in the minimum wage. When worker power is high and firms are constrained in making labor force adjustments, traditional neoclassical, cost-of-adjustment, and q-theory models of investment predict that minimum wage increases will result in less investment because increased labor costs lower the future cash flows of new investment projects, thereby reducing the optimal level of firm investment. As worker power declines and firms are less constrained in making workforce adjustments, the ability to substitute capital for labor (enhanced by technology improvements) and/or replace U.S. workers with cheaper foreign labor (due to globalization) dampens the negative effect of minimum wage increases on investment (i.e., the isoquant of production function becomes less convex).

We begin our analysis by examining changes in the strength of investment sensitivity to mini-

[1] The literature has proposed several non-mutually exclusive explanations that broadly fall into the weakening worker power: for instance, rising international trade and associated substitution of labor from low-wage countries (Elsby, Hobijn and Sahin, 2013), a decline in the relative price of capital (Karabarbounis and Neiman, 2014), domestic-outsourcing (contracting out) workers (Goldschmidt and Schmieder, 2017), and capital-biased technological change (automation) (Acemoglu and Restrepo, 2018). Another prominent hypothesis emphasizes the emergence of "superstar firms" and heterogeneity across firms (Autor, Dorn, Katz, Patterson and Van Reenen, 2017, 2020).

mum wage changes over the 1984 to 2017 time period. We estimate investment-wage sensitivity by augmenting standard investment regressions (e.g., Fazzari, Hubbard and Petersen, 1988) with a minimum wage variable. Using staggered changes in minimum wage rates across U.S. states, we start with a simple comparison of pre-2000 and post-2000 investment-wage sensitivities. Consistent with declining worker power, we find minimum wage changes had a significant negative effect on capital expenditures in the pre-2000 sample period but had no effect on capital expenditures in the post-2000 sample period. For the pre-2000 period, the estimated investment-wage sensitivity is –0.038 and statistically significant at the 1% level, which corresponds to a 24.6% decrease relative to the sample mean. However, for the post-2000 period, the estimated investment-wage sensitivity is 0.001 and statistically not significant at conventional levels.

To provide a more granular picture, we next investigate changes in investment-wage sensitivities using 15-year rolling window regressions. The negative impact of minimum wage on investment peaks in magnitude for the 15-year regressions using the 1987-2001 and 1988-2002 windows. The investment-wage sensitivity becomes statistically insignificant in the regressions following the 1999 to 2013 window. Consistent with these findings, a formal test for regime shifts in the structural relation between corporate investment and minimum wage suggests that the model with a structural break at 1999 fits the data best.

Our main sets of tests focus on examining the impact on investment-wage sensitivity of the various forces that have been advanced as driving the decline in worker power over the past four decades: globalization (which allowed easier access to cheap foreign labor as well as increased import penetration), technological change and the associated automation of the workplace, and weakening union power. To investigate the impact of globalization on worker power, as manifested by changes in investment sensitivity to minimum wage increases, we begin by examining the impact of the 1999 U.S.-China bilateral agreement that enabled U.S. firms to secure a greater fraction of the profits from their Chinese operations. In effect, the agreement provides U.S. firms with better access to the Chinese labor market, which weakens the bargaining power of U.S. workers by increasing the outside options of their employers. As a result, firms are less restricted in how they respond to a shock to their labor costs caused by mandated changes in the minimum wage.

Our empirical strategy is to compare changes in investment-wage sensitivities around the 1999 bilateral agreement for firms that are more versus less likely to benefit from greater access to cheaper Chinese labor. Using information from 10-k filings about U.S. firms' subsidiaries and their location, we define firms with at least one subsidiary in China as of 1997 (two years prior to the agreement) as most likely to benefit from access to cheap Chinese labor (i.e., treated firms) and treat all other firms as control firms. Using a generalized difference-in-differences framework, we find that firms operating in China as of 1997 experienced a dramatic decline in investment-wage sensitivities, moving from –0.045 (significant at the 1% level) before the agreement to –0.005 (statistically insignificant) after the agreement. For the control firms without subsidiaries in China prior to the agreement, investment-wage sensitivities are negative and highly significant both before and after the agreement. These findings are consistent with the hypothesis that firms with greater outside options afforded by globalization are less constrained in responding to minimum wage shocks.[2]

The shock to the U.S. labor market that was triggered by the 1999 U.S.-China bilateral agreement was a supply-driven shock. For additional evidence on the impact of globalization, we next examine a demand-side change in the U.S. labor market that was induced by a dramatic increase in the Chinese share of U.S. imports. The Chinese economic reforms in the 1980s and 1990s resulted in rapid productivity growth and a consequent surge in Chinese exports during this period. Its export growth was reinforced by China's entry to the World Trade Organization (WTO) in 2001. In particular, the Chinese share of U.S. imports increased from 4.0% in 1991 to 9.0% in 2001, before surging to 18.4% in 2011 (21.9% in 2017), which imposes stronger competition on U.S. firms.[3] Since firms in a more competitive environment are less able to shift rising labor costs to their consumers (Harasztosi and Lindner, 2019), they have stronger incentives to displace their workers and/or replace them with machines when they are hit by a shock to labor costs. In particular, minimum wage increases will lead to a competitive disadvantage of U.S. firms (especially those in the tradable sector) relative to Chinese rivals that are not subject to such a minimum wage shock. Therefore, we hypothesize that the investment-wage sensitivity decreases after China's export surge, and this change is more pronounced for firms in the industries that are highly exposed

[2]We provide additional evidence by dividing our control firms (without Chinese subsidiaries prior to the agreement) into those that do form Chinese subsidiaries after the agreement (treated firms) and those that do not (control firms). Consistent with an improvement in outside options made possible by globalization, we find a significant decline in investment sensitivities for the treated firms and no change for the control firms.

[3]A similar observation is noted in Autor, Dorn and Hanson (2013): the import penetration ratio for U.S. imports from China rose sharply over 1991–2007 with an inflection point in 2001.

to Chinese import competition.

We measure Chinese import exposure as of 1999, two years prior to China's accession to the WTO in 2001. Specifically, we define a U.S. industry's exposure to imports from China as the log of the Chinese import penetration ratio (Bernard, Jensen and Schott, 2006) if firms are classified as in the tradable sector (Mian and Sufi, 2014), and zero otherwise. Using a similar difference-in-differences approach, we find that rising exposure to Chinese import competition indeed expedites changes in investment-wage sensitivity moving toward zero. Our results are robust to using a non-U.S. trade exposure to Chinese imports as an instrument, following Autor et al. (2013) and Acemoglu, Autor, Dorn, Hanson and Price (2016).[4]

We next investigate the effects of technological innovations over the last several decades on investment-wage sensitivities. Technological advances in workplace automation serve to weaken worker power by providing employers with greater opportunities to substitute capital for labor. Coupled with the empirical observation that labor has indeed become more substitutable in the post-2000 period due to technological advances (Acemoglu and Restrepo, 2019a), this substitution hypothesis may explain a decrease in investment-wage sensitivity in the later sample period. Graetz and Michaels (2017) find that routine-intensive jobs are particularly susceptible to replacement by new technologies. Hence, we measures the extent to which industries are subject to technological change using an industry-level share of routine-task labor (Zhang, 2019). Our difference-in-differences estimates indicate that firms that are more exposed to automation experience a larger decrease in investment-wage sensitivity in the post-2000 period.

Finally, motivated by the continuing decline in union coverage rates over our sample period (Açıkgöz and Kaymak, 2014), we examine whether weakening union power contributes to decreased investment-wage sensitivity in the post-2000 period. Weakening labor union power enables firms to adjust their workforce (at both the extensive and intensive margins) more flexibly in

[4]To further support our findings, we examine whether the differential changes in investment-wage sensitivity following increased competition are mainly caused by industry leaders. Motivated by findings in Khanna and Tice (2000) and Gutiérrez and Philippon (2017), we conjecture that greater exposure of U.S firms to Chinese import competition motivates industry leaders to invest more to compete with their Chinese rivals when they face a minimum wage shock; however, such a large exposure forces laggards to exit or to downsize their operation. To test this hypothesis, we identify leader (laggard) firms for each industry as those with above-median (below-median) *Tobin's q*, following Gutiérrez and Philippon (2017). We find that rising Chinese import competition leads leaders to invest more relative to laggards in response to minimum wage increases. Notably, laggards still reduce their investment significantly following the minimum wage increase in the later period (2001–2017). These results are robust to using sales and total assets to identify industry leaders.

response to mandated minimum wage increases. We first measure weakening labor union as the decline in the union coverage rate at the state- or industry-level for the entire sample period. Using this measure, we find that firms that experience a larger drop in union coverage are less sensitive to minimum wage shocks. To address potential endogeneity concerns, we also use the passage of right-to-work (RTW) laws as a plausibly exogenous shock to union bargaining power. In states with RTW legislation, mandated union membership or payment of union dues is prohibited, which limits a union's access to resources thereby weakening union power. We find that, after the passage of RTW laws, corporate investment responds less negatively to minimum wage increases.

In sum, consistent with declining worker power, we find micro-level evidence that employers have become less constrained in their response to exogenous shocks to labor costs associated with mandated changes in the minimum wage. Whereas firms reduced investment following minimum wage increases in the early (pre-2000) part of our sample, the sensitivity of investment to increases in the minimum wage became statistically and economically insignificant in the later time period. Furthermore, we show that declines in investment-wage sensitivities are tied to forces that have been advanced to explain declining worker power: globalization, technological advances, and declining union power.

In addition to contributing to the literature on worker power, our study also contributes to the literature on the real effects of mandated minimum wage increases. Since much of that literature focuses on the effects of minimum wages on employment, our focus on the investment side, and how it interacts with worker power, adds additional perspective. In this regard, for example, our paper provides a potential explanation for the mixed results in earlier studies that examine the effect of minimum wage on corporate investment (e.g., the negative effects reported in Cho (2021) and Gustafson and Kotter (2021) and the positive effects documented in Geng, Huang, Lin and Liu (forthcoming) and Hau, Huang and Wang (2020)).

II. Minimum Wage and Corporate Investment

A. Institutional Details

The federal minimum wage provisions for employees in the U.S. are contained in the Fair Labor Standards Act (FLSA). Employers whose annual sales are at least $500,000 and who engage in interstate commerce are subject to the FLSA. All of their employees are covered by the Act. The Act also covers employees engaged in interstate commerce even if their employers' revenue is less than $500,000. The Act establishes overtime pay, recordkeeping, and youth employment standards for workers in the private sector as well as in federal, state, and local governments. It was enacted in 1938 and has been amended many times since, mainly to increase the federal minimum wage. As of July 2009, more than 143 million workers (about 93% of the U.S. civilian labor force) in more than 9.8 million workplaces are protected by the FLSA, which is enforced by the Wage and Hour Division of the Department of Labor.

In addition, many states also have their own minimum wage laws. Some states index their minimum wage rates to inflation, increase the rates in legislatively scheduled increments, set the rates at the federal rate, or a mix of these three methods. The state minimum wage rates may differ from those set by the federal statutes. Under Section 18 of the FLSA, when an employee is subject to both the federal and state minimum wage laws, the employee is entitled to the higher of the two standards.

B. Identification Strategy

Our identification strategy for investigating the relation between corporate investment and minimum wage assumes that changes in the federal and state level minimum wage rates are exogenous to individual firm outcomes. The federal and state minimum wage rates change at various times and in various increments. These changes are depicted in Figure I at the federal level and for a sample of three geographically distant states: California, Connecticut, and Illinois.

[Insert Figure I here.]

As shown in Figure I, the timing of minimum wage changes varies at the federal level and across the states. For example, wage rate changes in Connecticut and California seem to lead the federal wage increases over the period 1983–2017. Illinois' wage rate changes moved in lockstep with the federal wage rate changes till 2003; however, after 2003, wage rate changes in Illinois led to the federal wage rate increases. In our econometric tests, we exploit this staggered timing of changes in minimum wage rates. Since minimum wage increases are staggered, it is possible for firms to be in both the treatment (i.e., minimum wage increases) and control groups at different times, which alleviates the potential problem of systematic differences between treatment and control groups.

As outlined earlier, each state uses its own adjustment mechanisms for minimum wage rates. One of the common methods is to index minimum wage rates to inflation. In Figure I, we plot the CPI-U (Consumer Price Index for All Urban Consumers) index over the same period by setting the index value in January 1983 to a wage rate of $3.25 per hour (on the left axis). The figure also shows that California, Connecticut, and Illinois wage rates lagged the CPI index till the year 2000 but led the CPI index after 2000. The federal wage rates lagged the CPI index over the entire sample period. For the states shown in this figure, it appears that inflation does not seem to trigger wage increases. This is desirable for our identification strategy because inflation might directly affect corporate investment. However, there are some states that maintain the real value of the minimum wage rates over time by indexing the rates to inflation. Thus, if inflation triggers a minimum wage increase, our identifying assumption for estimating the investment sensitivity to minimum wage (hereafter investment-wage sensitivity) may be violated. In general, inflation has two direct conflicting effects on corporate investment (Hochman and Palmon, 1983): depreciation and interest effects. On the one hand, the real tax benefit of depreciation decreases with inflation because depreciation allowances are based on historical costs, rather than on current nominal values. On the other hand, the real tax benefit of interest deductions increases with inflation because firms deduct interest expenses at nominal interest rates, rather than at real rates. Therefore it remains an empirical question whether inflation increases or decreases corporate investment. Feldstein (1982) empirically finds that inflation is negatively associated with firm investment under the structure of U.S. tax rules. Therefore, to make our identification strategy more convincing, we exclude all the firms headquartered in the 15 U.S. states that have indexed their minimum wage rates to some measures of inflation.[5]

[5]These 15 states are Alaska, Arizona, Colorado, Florida, Michigan, Minnesota, Missouri, Montana, Nevada, New

The second adjustment mechanism that some states use is to write in specific future dates for specific minimum wage rates in legislation. If such legislation were motivated by an anticipated but unobservable improvement in investment opportunities by the legislators, our identifying assumption (the exogeneity of wage increases to firm investment) may be violated. To examine this possibility, we plot the time-series of the value weighted stock market (NYSE/AMEX/NASDAQ) index (as a proxy for investment opportunity) over the sample period on the right axis in Figure I. The figure shows that minimum wage increases occur in both up and down stock markets either at the federal or state level. If the unobservable component of expected investment opportunity is positively associated with future changes in minimum wage rates, our tests are likely to find a positive relation between minimum wage increases and capital expenditure. In results tabulated later in the paper, we find instead a negative relation between minimum wage and corporate investment. Therefore, we conclude that this legislative issue is unlikely to affect our results.

The third and last adjustment mechanism used by states is to set their minimum wage rates based on the federal rate. Therefore, from the identification perspective, these states are similar to the case where states specify their rates in state legislation. Moreover, a change in *federal* minimum wage laws can largely be regarded as exogenous to the *state-level* macroeconomic conditions that may affect individual firm outcomes. This enables us to isolate the effect of unobservable state-level macroeconomic shocks on corporate investment to the extent that federal minimum wage policy is orthogonal to the state-level economic conditions. To directly control for the macroeconomic conditions of the U.S. economy that may facilitate a change in federal minimum wage laws, we include year fixed effects in our investment regressions.

C. Sample Construction

We obtain the historical changes in minimum wages for non-farm private sector employment under state laws for all US states from the Tax Policy Center.[6] These data are sourced from the Wage and Hour Division of the U.S. Department of Labor and from the *Monthly Labor Review* by the Bureau of Labor Statistics, and span the 1983 to 2014 time period. For 2015–2017, we

Jersey, Ohio, Oregon, South Dakota, Washington, and Vermont. We obtain qualitatively similar results when including those 15 states in our analysis.

[6] https://www.taxpolicycenter.org/statistics/state-minimum-wage-rates

hand-collect the data from the U.S. Department of Labor.[7] To match this time period, we construct a sample of firms in Compustat following the sample selection criteria of Almeida, Campello and Galvao (2010). We eliminate observations from financial institutions (SIC codes 6000–6999). In addition, we discard firm-years that display asset or sales growth exceeding 100% to eliminate firms that exhibit large jumps in business fundamentals in terms of size and sales, because these jumps are usually associated with major corporate events, such as mergers and acquisitions or reorganizations. We also remove very small firms for which capital is less than \$10 million, because linear investment models may not be appropriate for those firms, as discussed by Gilchrist and Himmelberg (1995). Finally, we eliminate firm-years that have negative *Tobin's q*. All dollar-valued variables are converted into December 2014 constant dollars using the consumer price index for all urban consumers (CPI-U). The final sample has 59,096 firm-year observations.

D. Descriptive Statistics

Table I provides summary statistics of the main variables used in this study. The firm-year-level data consists of 59,096 firm-year observations from 1984 to 2017, consisting of 6,376 firms. We define investment rates for firm i in state s in year t as capital expenditures (I) normalized by the beginning-of-year capital stock (K) in which capital stock is measured as property, plant, and equipment. This variable is named as $Investment$ $(= \frac{I_{i,s,t}}{K_{i,s,t-1}})$. The sample mean and median of investment rates are 24.5% and 18.5% respectively, which implies that the empirical distribution of *Investment* is right-skewed. Cash flow is calculated as earnings before extraordinary items plus depreciation (CF), normalized by the beginning-of-the-year capital stock: $Cash\ Flow = \frac{CF_{i,s,t}}{K_{i,s,t-1}}$. $Cash\ Flow$ is also right-skewed and has a very high standard deviation (89.5%). $Tobin's\ q$ is a proxy for investment opportunities, which is measured as the ratio of the market value of assets to book value of assets where the market value of assets is defined as total assets plus market equity minus book equity. In line with the overall Compustat dataset, our sample also has an average $Tobin's\ q$ of 1.641.

For state-level variables ($w_{s,t-1}$, *GDP Growth*, *Population*, and *Unemployment*), we report their descriptive statistics based on 1,190 state-year observations from 1983 to 2016. We define $w_{s,t-1}$ as the minimum hourly wage rate for year t–1 in state s. The average of $w_{s,t-1}$ in nominal dollars

[7] `https://www.dol.gov/whd/state/stateMinWageHis.html`

is \$5.31 and its standard deviation is \$1.54. $w_{s,t-1}$ displays both cross-sectional variation (across states) and time-series variation (within-state variation over time). The across-state standard deviation (the cross-sectional standard deviation of state-level time-series averages) of $w_{s,t-1}$ is \$0.25 and within-state standard deviation (the average of time-series standard deviations for all states) is \$1.52. These numbers indicate that there is considerable within-state across-time variation in minimum wage rates. The staggered nature of these changes across time and states helps us to identify the investment-wage sensitivity.

We define $w_{i,s,t-1}$ as the minimum hourly wage rate for year $t-1$ in state s where firm i's headquarters is located, which is used in our regressions. We obtain information about firms' headquarters from the Compustat data that provide only the latest headquarters location.[8] In our dataset, there are many state-years that have more than one minimum wage rate in effect during the year. In these cases, we compute a weighted average minimum wage rate where the weights are given by the number of days the minimum wage rate is in effect. The average of $w_{i,s,t-1}$ is \$5.28 and its (overall) standard deviation is \$1.59 (not reported in the table). Summary statistics of other state-level control variables (*GDP Growth*, *Population*, and *Unemployment*) used in our analyses are also reported in this table. The detailed definition of each variable is provided in Appendix A.

[Insert Table I here.]

E. Estimation

To estimate the investment-wage sensitivity, we augment standard investment regressions (e.g., Fazzari et al., 1988) with a minimum wage variable as follows:

$$\frac{I_{i,s,t}}{K_{i,s,t-1}} = \alpha_i + \alpha_t + \beta_1 \, Tobin's \; q_{i,s,t-1} + \beta_2 \, \frac{CF_{i,s,t}}{K_{i,s,t-1}} + \beta_3 \, w_{i,s,t-1} + \beta_4 \, X_{i,s,t-1} + \epsilon_{i,s,t}, \qquad (1)$$

[8] Since locations rarely change, and even when they do, the new and old locations are usually not far apart and are within the same state, we follow an extensive line of research (e.g., Serfling, 2016) and proxy for each firm's location of headquarters using the Compustat data. This measurement error may bias our findings. If a firm is coded as being located in a state with minimum wage increases but in fact was not, the regression results will show that the firm's investment is not responsive to changes in the minimum wage. Hence our estimated investment-wage sensitivity will be a fraction of the true value. Likewise, if a firm is coded as not being located in a state with minimum wage increases but in fact was located in such a state, the regression tests will show that the firm's investment changed despite the absence of minimum wage changes in our data. Therefore, this type of measurement error will bias our tests to find no effect of minimum wage on investment. In sum, the measurement error in headquarters location will produce a conservative estimate of investment-wage sensitivity.

where i, s, and t index firms, states, and years; α_t is a set of year fixed effects, which absorb time-varying macroeconomic shocks faced by all firms; and α_i is a set of firm fixed effects, which absorb time-invariant unobservable firm characteristics. We also control for state-level macro variables, $X_{i,s,t-1}$, including real GDP growth rates, log of population, and unemployment rates. The definitions of all variables are provided in Appendix A. We cluster standard errors of regression coefficients at the state-level, instead of the firm-level. Given that the minimum wage laws vary by state, potential correlations in unobserved factors that affect different firms in the same state may lead to inconsistent estimates of standard errors. Hence this method accounts for cross-firm correlations of error terms within a state, which is more general than firm-level clustering.

[Insert Panel A of Table II here.]

Column (1) in Panel A of Table II reports the estimated coefficients in equation (1) for the entire sample from 1984 to 2017. The coefficient on the minimum hourly wage rate ($w_{i,s,t-1}$, hereafter minimum wage) is negative and statistically significant at the 10% level. The estimate of −0.017 (investment-wage sensitivity) indicates that when a state's minimum wage increases, firms headquartered in that state reduce their investment rates. The magnitude of the regression coefficient implies that the effect is economically large. Following a one standard deviation increase in the minimum wage ($1.588), firms reduce their investment rates by 270 basis points ($= 1.588 \times -0.017$). This 270 basis points reduction corresponds to an 11.0% decrease, relative to the sample mean of investment rate (24.5%). In spite of the sizable economic impact of minimum wage on investment, the statistical significance of the result is quite weak (at the 10% level). According to the report by McKinsey Global Institute, the labor's share of income in the nonfarm US business sector has undergone tremendous changes during our sample period. The share has fallen from about 62.7% to 56.7%, a 10% decline, over the period between 1980 and 2016.[9] This tremendous change motivates an examination of changes in the strength of investment-wage sensitivity over our sample period.

In columns (2) and (3), we reestimate the same investment regression by evenly splitting the full sample into two sub periods: from 1984 to 2000 and from 2001 to 2017. The results from

[9]These numbers are excerpted from Exhibit 1 of the report by McKinsey Global Institute. (`https://www.mckinsey.com/featured-insights/employment-and-growth/a-new-look-at-the-declining-labor-share-of-income-in-the-united-states`).

two sub periods are starkly different. In the first sub period (column(2)), the investment-wage sensitivity is estimated to be –0.038 which is 2.2 times the estimate for the full sample period. The statistical significance of the coefficient in column (2) is very strong (at the 1% level) compared to that in column (1) (at the 10% level). A one standard deviation increase in the minimum wage leads firms to reduce their investment rates by 603 basis points, which corresponds to a 24.6% decrease, relative to the sample mean, in the 1984 to 2000 period. In the second sub period (column(3)), the investment-wage sensitivity is estimated to be 0.001 which is practically a null effect compared to the full sample period. Further, the estimate is not statistically significant. A formal statistical test to evaluate the null hypothesis of the equality of investment-wage sensitivity coefficients between the two sub periods is rejected at the 1% level of significance (χ^2-statistic = 12.54). These findings suggest that the investment-wage sensitivity is likely to vary over time. We examine this time-varying nature of investment-wage sensitivity and conduct a formal structural break analysis in Section III.A.

F. Robustness

In this section, we present several robustness checks of the estimated investment-wage sensitivity in Section II.E.

1. Placebo Test

We first perform a placebo test to check whether a *pseudo* minimum wage increase affects investment. Specifically, we repeat the estimation of equation (1) for the pre-2000 sample using a pseudo minimum wage variable. To construct this pseudo minimum wage variable ($w^{Pseudo}_{i,s,t-1}$), we first randomly assign a firm i to a particular state s. In the process of this random assignment, we take care to ensure that the distribution of the number of firms in each state is identical to our main sample. The timing of the state level minimum wage changes is also identical to our main sample. We define $w^{Pseudo}_{i,s,t-1}$ as the minimum wage at time t–1 in state s where firm i's *hypothetical* headquarters is located. All other control variables are based on the firm's characteristics as well as the firm i's assigned state s' macro variables for this simulation run. Once all firms i in the sample are assigned in this manner, we perform the investment regression (1) on the simulated panel of data and store the coefficient on $w^{Pseudo}_{i,s,t-1}$. This procedure constitutes one run of the simulation and is repeated 1,000 times and a distribution of $w^{Pseudo}_{i,s,t-1}$ coefficients is generated.

[Insert Panel B of Table II and Figure B.1 here.]

Panel B of Table II reports the empirical distribution of the coefficient on $w^{Pseudo}_{i,s,t-1}$, based on this simulation. The mean and median of this distribution are 0.002 and 0.001 respectively, and both close to zero. This suggests that on average there is no investment-wage sensitivity in our simulation. We note that the coefficient on $w^{Pseudo}_{i,s,t-1}$ is –0.017 at the first percentile of this distribution. We plot the empirical distribution (histogram) of the coefficient on $w^{Pseudo}_{i,s,t-1}$ in Figure B.1. The green line in the figure is the estimated nonparametric kernel density (probability density function) of the coefficient on $w^{Pseudo}_{i,s,t-1}$ coefficient. The red vertical line in the figure indicates the coefficient on $w_{i,s,t-1}$ (–0.038) obtained from the investment regression using the actual data (column (2) of Panel A of Table II). The actual value of –0.038 is far below –0.017, the first percentile of the simulated distribution. This result suggests that our estimated significant investment-wage sensitivity in the 1984-2000 period is not likely due to chance.

2. Measurement Error in Tobin's q

The empirical proxy (*Tobin's q*) for marginal q or investment opportunities is likely to contain measurement errors which in turn produces biased coefficients in investment regressions. To overcome this problem, Erickson, Jiang and Whited (2014) develop minimum distance estimators for a classical errors-in-variables model with multiple mismeasured and multiple perfectly measured regressors on panel data. These, in turn, produce unbiased coefficients. The underlying estimating equations are linear in the third- and higher-order polynomial functions of moments (cumulants) of the joint distribution of the observable variables. Using Erickson et al. (2014)'s high-order cumulant estimators, we assess the robustness of the investment-wage sensitivity when the proxy for investment opportunities is subject to measurement errors.

In columns (1–2) of Table B.1 in Appendix B, we report the baseline fixed effect OLS estimates from Table II, Panel A, columns (2–3) for easier comparison with the cumulant estimates. Columns (3–4) display the regression coefficients estimated using the higher-order cumulant estimators. Consistent with Erickson et al. (2014), the coefficients on *Tobin's q* (*Cash Flow*) based on the cumulant estimation are larger (smaller) than those from fixed effect OLS estimation. The investment-wage sensitivity continues to remain significant for the pre-2000 period, and economic significance increases, as the magnitude of the coefficient becomes larger (−0.041 compared to −0.038). The investment-wage sensitivity for the post-2000 period is virtually zero and insignificant, similar to what was obtained from the fixed effect OLS estimation (comparing the coefficient on $w_{i,s,t-1}$ across columns (2) and (4)). Overall, these results show that the relation between minimum wage and investment is robust to measurement error in *Tobin's q*.

3. Strict Exogeneity Assumption Tests

In equation (1), the consistency of the fixed effects estimator crucially depends upon the strict exogeneity assumption, as noted by Wooldridge (2011). The strict exogeneity assumption asserts that $E(\epsilon_{i,s,t}|w_{i,s,\tau}, \alpha_i) = 0$ for all t and τ. Therefore, we conduct strict exogeneity assumption tests for all our estimations in Table II, Panel A. As suggested by Wooldridge (2011) and Grieser and Hadlock (2019), we include the one period lead value of the key variable of interest ($w_{i,s,t}$) in the investment regressions. Wooldridge (2011) notes that the coefficient on this lead variable is zero under the null hypothesis of strict exogeneity. We report the results of this exercise in Table B.2 of Appendix B. In columns (1–3) of Table B.2, we report the baseline fixed effect OLS estimates from Table II, Panel A, columns (1–3) for easier comparison. Columns (4–6) display the regression coefficients estimated using the procedure suggested by Wooldridge (2011). The estimated coefficient on $w_{i,s,t}$ is close to zero and statistically insignificant in all specifications. We, therefore, conclude that the strict exogeneity assumption is satisfied in our empirical setting.

III. Investment Sensitivity to Minimum Wage Over Time

As documented in Panel A of Table II, the investment-wage sensitivity is very different in the first and second half of our sample period. Corporate investment becomes insensitive (statistically and economically insignificant) to an increase in the minimum wage in the 21st century. In this section, we document the changes in investment-wage sensitivity over time and then attempt to provide potential explanations for these changes.

A. *Rolling Window Regressions and Structural Break Analysis*

To examine changes in investment-wage sensitivity over time, we run 15-year rolling window regressions over the entire sample period. Specifically, we repeat the estimation of equation (1) for twenty sub sample periods. The first sample period runs from 1984 to 1998. For every subsequent sub-sample, we roll forward the starting date by one year and then estimate a 15-year regression estimate of (1). The last sample period runs from 2003 to 2017. We obtain a time series of twenty estimates of β_3 (investment-wage sensitivity) and corresponding t-statistic.

[Insert Panel C of Table II and Figure II here.]

Panel C of Table II shows the time-series of estimated investment-wage sensitivity and corresponding t-statistics for these twenty sub samples. The pattern is clear: investment-wage sensitivity is negative and strongly significant for all 15-year sub samples with starting dates from 1984 to 1998. The negative investment-wage sensitivity (in magnitude) peaks in the samples from 1987 to 2001 and from 1988 to 2002 (–0.040 and −0.040 with t-statistics of −3.813 and −3.380, respectively). The coefficient then steadily decreases in economic magnitude (while continuing to be statistically significant) for all 15-year sub periods with starting dates from 1988 to 1998. However, after 1998, investment-wage sensitivity is statistically and economically zero in all the recent 15-year sub periods which start from 1999 to 2003. This pattern is displayed in Figure II. In panel (a) of Figure II, the red line plots the time-series of estimated investment-wage sensitivity and the grey shaded area, the 95% confidence intervals. The figure indicates that the investment-wage sensitivity of firms has disappeared over time. The corresponding t-statistics are plotted in panel (b) with a horizontal line indicating t-statistics of −2. Panel (b) shows that the statistical significance of estimated investment-wage sensitivity has also disappeared over the sample period.

We now formally test for a regime shift in the relation between corporate investment and minimum wage. We assume a single, known structural break and allow all the coefficients to change after the structural break date:

$$\frac{I_{i,s,t}}{K_{i,s,t-1}} = \alpha_i + \alpha_t + \beta_1\, w_{i,s,t-1} + \beta_2\, Z_{i,s,t-1} + d_t(k)\Big[\beta_3\, w_{i,s,t-1} + \beta_4\, Z_{i,s,t-1}\Big] + \epsilon_{i,s,t}, \qquad (2)$$

where $Z_{i,s,t-1}$ is a set of firm- and state-level control variables used in equation (1), α_t is a set of year fixed effects, and α_i is a set of firm fixed effects. $d_t(k)$ equals one if t is greater than or equal to the assumed year of structural break k, and zero otherwise. We require at least five years of data for both the pre- and post-break periods in our estimation and hence k runs from 1989 to 2013. The next step is to pick k that produces the best fit model according to some statistical criteria.

We adopt two statistical methods to identify the "best" model: Akaike's information criterion (AIC) and Bayesian (or Schwarz's) information criterion (BIC). Both criteria are an estimator of prediction error and thereby help compare the relative quality of statistical models for a given set of data. The lower the value of the criteria, the better is the quality of the model. We calculate the AIC as $-2\ ln(L[k]) + 2p$ where $ln(L[k])$ is the maximized log-likelihood of the model in which the assumed structural break is year k and p is the number of parameters estimated. We calculate the BIC as $-2\ ln(L[k]) + p\ ln(N)$ where N is the sample size. In Panel D of Table II, we plot both AIC and BIC as a function of assumed break year k. As shown in this figure, both AIC and BIC achieve their minimum value if $k = 1999$. This means that an estimated structural break date of 1999 best fits the data. These results are also consistent with the rolling window regressions in Panel C of Table II: the investment-wage sensitivity becomes statistically insignificant the first time for the 15-year period from 1999–2013.

[Insert Panel D of Table II here.]

B. Explanations for the Changes in Investment-Wage Sensitivity

The results in Section III.A indicate that corporate investment decisions became insensitive (statistically and economically) to changes in minimum wage starting from 1999 and confirmed by a structural break analysis. Motivated by the *Declining Worker Power Hypothesis* proposed by Stansbury and Summers (2020), we attempt to provide potential explanations consistent with these changes in investment-wage sensitivity over time. Stansbury and Summers (2020) propose the reduction in worker power vis-a-vis corporate management from 1980 as a unified explanation for many of the broad macroeconomic trends in the U.S. such as rising corporate valuation and markups, sluggish wage growth, declining labor's share of national income, and low unemployment and inflation rates. Our starting point is the hypothesis that factors weakening worker power as identified by Stansbury and Summers (2020) also lead to insignificant investment-wage sensitivity over time. In this section, we examine changes in the U.S. labor markets along four dimensions that may weaken the power of U.S. workers over the past four decades and study its impact on investment-wage sensitivity: (a) easier access to cheap labor, (b) increased Chinese import competition, (c) technological change and automation, and, (d) weakening union power. We note that these explanations are not mutually exclusive or exhaustive.

1. U.S. Firms' Access to Cheap Labor: 1999 US-China Bilateral Agreement

Devereaux and Lawrence (2004) document that the bilateral agreement signed between the U.S. and China opened the economy of China to U.S. multinational firms by improving contracting institutions: for instance, relaxing foreign ownership restrictions, eliminating foreign exchange balancing requirements, removing local content requirements, lifting requirements of any kind including offsets, transfer of technology, or requirements to conduct research and development in China, etc. Before the agreement, Chinese partners accrued a large fraction of the profits earned by U.S. firms' Chinese subsidiaries. The agreement enabled U.S. firms to capture a greater share of the profits from their Chinese operations. Thus, the elimination of the investment restrictions in China was widely expected to increase U.S. multinational firms' investment in China. At the same time, Ceglowski and Golub (2012) show that relative unit labor costs in manufacturing vis-a-vis the U.S. (which accounts for relative productivity, relative wages, and real exchange movements) was about 22% in 1998, based on the World Bank estimates. Therefore, the bilateral agreement opened up the Chinese labor market to U.S. firms through their potential capital investments. This, in turn, weakens the bargaining power of U.S. workers by increasing the outside options of the firms (employers). Thus, firms faced with a shock to their U.S. labor costs due to mandated changes in minimum wage could credibly move capital investments outside the US. Devereaux and Lawrence (2004) also show that the bilateral agreement was largely unexpected due to strong opposition in the U.S. Congress and we use this largely exogenous shock to the U.S. worker power to identify its causal effect on changes in investment-wage sensitivity. Consistent with this view, Bena and Simintzi (2019) document that U.S. firms operating in China reduced their process innovation activities in the U.S. that would have lowered their production costs. It is posited that after the agreement, improved access to cheaper Chinese labor would be a substitute for process innovation.

We, therefore, conjecture that U.S. firms' better access to cheap labor in China, which was triggered by the US-China bilateral agreement signed in November 1999 (which incidentally coincides with a structural break in investment-wage sensitivity identified in Section III.A), makes firms' investment less sensitive to minimum wage shocks.

Specifically, we examine whether the heightened ability to source cheap foreign labor eliminates the investment-wage sensitivity, after the bilateral agreement. To identify firms that can benefit from the agreement, by exploiting cheap Chinese labor, we focus on U.S. firms with subsidiaries in China, following the strategy employed in Bena and Simintzi (2019). Since firms may endogenously choose to operate in China after the agreement, we begin by focusing on U.S. firms with at least one subsidiary in China as of 1997, two years prior to the bilateral agreement.[10] We use hand collected information from 10-k filings in 1997 to identify all U.S. firms with Chinese subsidiaries in our dataset. We employ the following difference-in-differences regression by extending our baseline specification in equation (1):

$$\begin{aligned}\frac{I_{i,s,t}}{K_{i,s,t-1}} = {} & \alpha_i + \alpha_t + \beta_1\, Tobin's\ q_{i,s,t-1} + \beta_2\, \frac{CF_{i,s,t}}{K_{i,s,t-1}} + \beta_3\, w_{i,s,t-1} \\ & + \beta_4\, Agreement_t \times w_{i,s,t-1} + \beta_5\, China97_i \times Agreement_t \times w_{i,s,t-1} \\ & + \beta_6\, China97_i \times w_{i,s,t-1} + \beta_7\, China97_i \times Agreement_t + \beta_8\, X_{i,s,t-1} + \epsilon_{i,s,t},\end{aligned} \tag{3}$$

$China97_i$ is an indicator variable set to one if firm i has at least one subsidiary in China as of 1997, and zero otherwise; $Agreement_t$ is an indicator variable set to one for the time period after the agreement (including 1999), and zero otherwise; $X_{i,s,t-1}$ includes the same set of control variables used in Panel A of Table II. We also include interaction terms of all control variables with $China97_i$ and $Agreement_t$.[11] We note that, in this generalized difference-in-differences framework, the outcome of interest is the *change* in the slope coefficient on minimum wage (i.e., investment-wage sensitivity) captured by β_5. A similar difference-in-differences framework is employed in a number of papers, e.g., Gormley, Kim and Martin (2012). The coefficient β_5 captures the change in investment-wage sensitivity for firms operating in China as of 1997, following the bilateral agreement as compared to years before the agreement, relative to firms not operating in

[10]We also check the robustness of our results for firms with at least one subsidiary in China as of 1998, one year prior to the bilateral agreement. The results documented in Panel A of Table B.3 in Appendix B are qualitatively similar to those reported in Table III.

[11]The $China$ and $Agreement$ indicators are absorbed by the firm and year fixed effects, respectively.

China. Our baseline estimate of the investment-wage sensitivity before 1999 is negative, as documented in Panel A of Table II. Firms with access to cheap Chinese labor will be able to undo this sensitivity as discussed above and hence, we expect β_5 to be positive and significant.

[Insert Table III here.]

Table III presents the regression results for equation (3). In column (1), the difference-in-differences in the investment-wage sensitivity (β_5) after the agreement between treated ($China97$ = 1) and control ($China97$ = 0) firms is 0.035 and significant at the 1% level. We interpret this effect as follows:

For the treated firms, the investment-wage sensitivity before the agreement is calculated as the sum of regression coefficients: $\beta_3 + \beta_6 = -0.017 - 0.028 = -0.045$ with a t-statistic of –3.73, both economically and statistically very significant. This means, these treated firms significantly adjusted their investment in response to mandated changes in the minimum wage before the agreement. For the very same group of firms, the investment-wage sensitivity after the agreement is calculated as the sum of four regression coefficients: $\beta_3 + \beta_4 + \beta_5 + \beta_6 = -0.017 + 0.005 + 0.035 - 0.028 = -0.005$ with a t-statistic of –0.86. The estimated investment-wage sensitivity is economically and statistically insignificant and the treated firms do not adjust their investment in response to mandated changes in the minimum wage after the agreement.

We now turn to the control group of firms who were not operating in China at the time of the bilateral agreement ($China97$ = 0; the omitted group). For these firms, the investment-wage sensitivity before the agreement is the regression coefficient $\beta_3 = -0.017$ with a t-statistic of –2.14, both economically and statistically significant. This means, these control firms significantly adjusted their investment in response to mandated changes in the minimum wage before the agreement. For the very same group of firms, the investment-wage sensitivity after the agreement is calculated as the sum of two regression coefficients: $\beta_3 + \beta_4 = -0.017 + 0.005 = -0.012$ with a t-statistic of –1.92. The estimated investment-wage sensitivity continues to be economically and statistically significant (but smaller in magnitude than before) and the control firms do indeed adjust their investment in response to mandated changes in the minimum wage even after the agreement.

In conclusion, the bilateral agreement causally eliminates the investment-wage sensitivity of the treated firms (firms with operations in China) while the control firms continue to have a negative investment-wage sensitivity even after the agreement.

[Insert Figure III (a) here.]

Figure III (a) summarizes the results of column(1): It plots the changes in investment-wage sensitivities around the bilateral agreement for the two groups of firms: firms not operating in China as of 1997 and firms operating in China as of 1997. The solid dots in the figure indicate the point estimates of the investment-wage sensitivity from the regression and the vertical lines around these point estimates denote the corresponding 95% confidence intervals for the point estimates. The figure shows that the 1999 US-China bilateral agreement significantly shifts bargaining power away from the U.S. workers to U.S. firms because of these firms' improved access to cheap labor in China. As a result, these U.S. firms' investment decisions are not sensitive to minimum wage shocks.

In Table III column (2), we also examine the changes in investment-wage sensitivity of those select firms that changed their operational status in China, following the US-China bilateral agreement. These firms newly established their subsidiaries in China after the agreement. Since setting up operations in a new country is not instantaneous, we reasonably choose 2004 (five years after the agreement) as the year to check the operational status of these firms in China. Specifically, we define $China04_i$ as an indicator variable set to one for those firms without any subsidiary in China as of 1997, but have at least one subsidiary in China as of 2004, and zero otherwise. We then introduce this indicator variable with its interactions in equation (3). In column (2), the omitted group consists of firms that have no operations in China; that is, $China04$ = $China97$ = 0. For firms that changed their operational status in China after the agreement ($China04$ = 1), the results are as follows: The investment-wage sensitivity changes from –0.031 ($= -0.016 - 0.015$, a t-statistic of –2.18) before the agreement to 0.000 ($= -0.031 + 0.002 + 0.029$, a t-statistic of 0.06) after the agreement. The result is consistent with the view that firms move their operations to China after the agreement to source cheap labor for their operations and eliminate their investment-wage sensitivity. Figure B.2 plots these results.

[Insert Figure B.2 here.]

Figure B.2 is also consistent with the revealed preference theory which can be used to analyze the China subsidiary choices of firms and for comparing the influence of the US-China bilateral agreement on firm behavior. Firms with operations in China in 1997 before the agreement presumably had the most to gain by reducing the negative impact of minimum wage shocks on investment and were willing to incur costs of doing business in China, even before the agreement. Consistent with this view, these firms had the most negative investment-wage sensitivity of all firms in our sample before the agreement and also gained the most by eliminating this investment-wage sensitivity after the agreement. Firms with operations in China in 2004 but not in 1997 also gained presumably because the agreement lowered their costs of doing business in China enough to overturn their earlier decision of not having a Chinese subsidiary. Indeed these firms had a negative impact of minimum wage shocks on investment (but not as much as the firms that were operating in China as of 1997) and were willing to incur costs of doing business in China, only after the agreement lowered their costs to do business in China. This group of firms was successful in eliminating their investment-wage sensitivity after the agreement. Finally, the firms with no operations in China before and after the agreement had their negative investment-wage sensitivity virtually unchanged after the agreement.

Robustness Check: To track U.S. firms' entry into China after the bilateral agreement, we also construct a time-varying indicator, $China_{i,t}$, that takes a value of one if firm i has at least one subsidiary in China in year t, and zero otherwise. Since comprehensive reporting of subsidiary information in the EDGAR (Electronic Data Gathering, Analysis, and Retrieval) database from the U.S. Securities and Exchange Commission (SEC) is only available starting from 1997, we use information as of 1997 for all years prior to 1997. Using this time-varying indicator instead of $China97$, we continue to obtain similar results. These robustness results are reported in Panel B of Table B.3 in Appendix B.

2. U.S. Firms' Increased Exposure to Chinese Import Competition: 2001 China's Accession to WTO

Devereaux and Lawrence (2004) note that after the passage of the US-China bilateral agreement in November 1999, the Clinton Presidential administration announced its strong support for Permanent Normal Trade Relations (the PNTR bill) with China. After a ten-month-long effort, the US House and Senate passed the PNTR bill into law in September 2000, which would be in force once China's accession to the World Trade Organization (WTO) was completed. In December 2001, China became the 143rd member of the WTO, and the US extended the PNTR status to China as of January 2002. The net effect of these laws was to grant China's producers access to the US market and thus intensify the import competition for US firms across many sectors of the economy. This increase in competition can be measured by the increase in the import penetration ratio of US imports from China for each sector and for the entire economy. Import penetration, measured as of 1999 (two years prior to China's accession into the WTO) is defined as the ratio of U.S. imports from China to total U.S. expenditure on goods in which U.S. expenditure is measured as U.S. gross output plus U.S. imports minus U.S. exports.[12] Consistent with this view, Autor et al. (2013) document that the import penetration ratio for Chinese goods rose from 0.6 percent in 1991 to 4.6 percent in 2007, with an inflection point in 2001.

Product market competition affects a firm's ability to raise their product prices in response to minimum wage shocks (e.g., Harasztosi and Lindner, 2019). In a more competitive industry (higher import penetration ratio), firms are less able to shift rising labor costs to their consumers through an increase in prices without losing their competitive advantage (e.g., without causing a large drop in output). Therefore, these exposed firms have a stronger incentive to find a way out of rising labor costs when they face an increase in labor costs, for instance, transition to the capital-intensive production process, displacement of workers, automation, etc. In addition, Gutiérrez and Philippon (2017) show that greater exposure of U.S. firms to Chinese import competition motivates industry leaders to invest more to compete with Chinese rivals when they face a labor cost shock; however, it forces industry laggards to exit or to downsize their operation. Therefore, we hypothesize that the investment-wage sensitivity of US firms is eliminated after China's export surge (i.e., after 2001) and this effect is more pronounced for firms in the industries that are highly exposed to Chinese import competition and for industry leaders. We use this demand-driven shock to U.S. labor markets to identify the causal effect of China's export surge on changes in investment-wage sensitivity.

[12] Specifically, we construct the import penetration ratio as

$$IP_{j,1999} = \frac{M_{j,1999}^{UC}}{Y_{j,1999} + M_{j,1999} - E_{j,1999}}$$

where for each industry j, $M_{j,1999}^{UC}$ is U.S. imports from China in 1999; $Y_{j,1999}$ is industry shipments; $M_{j,1999}$ refers to industry imports; and $E_{j,1999}$ is industry exports. We thank Peter Schott for making the trade flows data used in his paper (Bernard et al., 2006) available on his website.

We define a U.S. industry's exposure to imports from China, $Exposure_{UC}$, as the log of the import penetration ratio if an industry is classified to be in the tradable sector, and zero otherwise. We use the log transformation of the import penetration ratio due to its highly right-skewed distribution. Following Mian and Sufi (2014), we classify a four-digit NAICS industry as tradable if the amount of its imports and exports is at least \$10,000 per worker, or if the total amount exceeds \$500M.[13]

To test our hypothesis, we estimate the following difference-in-differences regression which is similar to equation (3):

$$
\begin{aligned}
\frac{I_{i,s,t}}{K_{i,s,t-1}} = {} & \alpha_i + \alpha_t + \beta_1\, Tobin's\; q_{i,s,t-1} + \beta_2\, \frac{CF_{i,s,t}}{K_{i,s,t-1}} + \beta_3\, w_{i,s,t-1} + \beta_4\, WTO_t \times w_{i,s,t-1} \qquad (4)\\
& + \beta_5\, Exposure_{UC,i} \times WTO_t \times w_{i,s,t-1} + \beta_6\, Exposure_{UC,i} \times w_{i,s,t-1}\\
& + \beta_7\, Exposure_{UC,i} \times WTO_t + \beta_8\, X_{i,s,t-1} + \epsilon_{i,s,t}.
\end{aligned}
$$

We define $Exposure_{UC,i}$ as firm i's exposure to imports from China, as of 1999 in two ways: (i) As a continuous variable, we assume it to be equal to $Exposure_{UC}$ for the four-digit NAICS industry to which firm i belongs. (ii) As an indicator variable, we set $Exposure_{UC,i} = 1$ for all firms in the industries with above-median $Exposure_{UC}$, and zero otherwise. WTO_t indicates the time period after China's WTO accession (including the year 2001). $X_{i,s,t-1}$ includes the same set of control variables used in Panel A of Table II. $X_{i,s,t-1}$ also includes all interaction terms of these control variables with $Exposure_{UC,i}$ and WTO_t.[14] The coefficient of the triple interaction term ($Exposure_{UC,i} \times WTO_t \times w_{i,s,t-1}$) captures the difference-in-differences in investment-wage sensitivity before and after 2001 across firms that are subject to different degree of import competition. Our hypothesis predicts β_5 to be positive.

[Insert Panel A of Table IV here.]

The first column of Panel A, Table IV estimates equation (4) using $Exposure_{UC,i}$ as an indicator variable. In column (1), the difference-in-differences in the investment-wage sensitivity (β_5) after China's accession to the WTO (WTO accession, hereafter) between treated ($Exposure_{UC,i}$ =

[13]We thank the authors for making their full list of industry classification available in the Supplemental Material (Mian and Sufi, 2014).

[14]The $Exposure_{UC,i}$ and WTO variables are absorbed by the firm and year fixed effects, respectively.

1) and control ($Exposure_{UC,i}$ = 0) firms is 0.041 and significant at the 1% level. We interpret this effect as follows:

For the treated firms, the investment-wage sensitivity before the WTO accession is calculated as the sum of regression coefficients: $\beta_3 + \beta_6 = -0.021 - 0.027 = -0.048$ with a t-statistic of −5.69, both economically and statistically very significant. This means, these treated firms significantly adjusted their investment in response to mandated changes in the minimum wage before the WTO accession. For the very same group of firms, the investment-wage sensitivity after the agreement is calculated as the sum of four regression coefficients: $\beta_3 + \beta_4 + \beta_5 + \beta_6 = -0.021 + 0.009 + 0.041 - 0.027 = 0.002$ with a t-statistic of 0.34. The estimated investment-wage sensitivity is economically and statistically insignificant and the treated firms do not adjust their investment in response to mandated changes in the minimum wage after the WTO accession.

We now turn to the control group of firms ($Exposure_{UC,i}$ = 0; the omitted group). For these firms, the investment-wage sensitivity before the WTO accession is the regression coefficient $\beta_3 = -0.021$ with a t-statistic of −2.05, both economically and statistically significant. This means, these control firms significantly adjusted their investment in response to mandated changes in the minimum wage before the WTO accession. For the very same group of firms, the investment-wage sensitivity after the WTO accession is calculated as the sum of two regression coefficients: $\beta_3 + \beta_4 = -0.021 + 0.009 = -0.012$ with a t-statistic of −1.65. The estimated investment-wage sensitivity continues to be economically significant but barely statistically significant with a p-value of 0.109 (and smaller in magnitude than before). The control firms do indeed adjust their investment in response to mandated changes in the minimum wage even after the WTO accession. In Column (2), we estimate equation (4) using $Exposure_{UC,i}$ as a continuous variable. Our inferences are qualitatively unchanged.

One important concern about $Exposure_{UC,i}$ is that it may also be correlated with domestic shocks to U.S. industries that in turn affect U.S. import demand especially if China were the least-cost producer of the demanded products. Therefore, to the extent that corporate investment may be correlated with unobserved shocks to U.S. product demand, the OLS estimates using $Exposure_{UC,i}$ as an RHS variable could be biased and inconsistent.

To identify the causal effect of an increase in import competition on investment-wage sensitivity, we employ an instrumental-variables strategy used in Autor et al. (2013). We instrument our exposure measure ($Exposure_{UC,i}$) with a non-U.S. trade exposure to Chinese imports ($Exposure_{OC,i}$) that is constructed using data on imports from China in eight other high-income countries excluding the United States.[15] These countries are Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland. This instrument is motivated by the fact that other high-income countries are similarly exposed to China's export surge that is mostly driven by supply shocks in China. The identifying assumption of this strategy is that (unobserved) import demand shocks are uncorrelated across high-income countries.[16] We estimate the predictive power of (non-U.S.) high-income-country instrument ($Exposure_{OC}$) for U.S. trade exposure to Chinese imports ($Exposure_{UC}$) by regressing $Exposure_{UC}$ by industry ($n = 386$) on the corresponding $Exposure_{OC}$ for the year 1999. The estimated coefficient on the instrument is 1.12 (The first stage $F - statistic = 49.68$) and the adjusted R-squared is 0.87. These results confirm the strong predictive power (in the first-stage regression) of other high-income countries' trade exposure on the U.S. trade exposure to Chinese imports.

Column (3) in Table IV reports the second stage two-stage least squares (2SLS) estimates of equation (4). The estimates are qualitatively similar to the OLS estimates in column (2). The 2SLS estimate of the coefficient on triple interaction term, the difference-in-differences in the investment-wage sensitivity, β_5 = 0.023 is statistically significant at the 1% level and larger than the corresponding OLS estimate of 0.019 in column (2).

[Insert Figure III (b) here.]

We present the 2SLS estimates of investment-wage sensitivity for two groups of firms in Figure III (b). The first group is US firms with no exposure to import competition with $Exposure_{UC,i} = 0$ (firms not vulnerable to import competition). The second group is US firms with $Exposure_{UC,i}$ value that is one standard deviation above the sample mean (firms more vulnerable to import competition). Figure 3(b) shows that the magnitude of the investment-wage sensitivity of the not vulnerable firms is lower than the investment-wage sensitivity of the more vulnerable firms in the pre-WTO accession period. In the post-WTO accession period, both groups of firms move towards eliminating their investment-wage sensitivity. The not vulnerable firms now have a lower magnitude of investment-wage sensitivity than before (but statistically significant). However, the

[15] We thank David Dorn for making the data available on his website (`http://www.ddorn.net/data.htm`).

[16] Since the U.S. total expenditure may be endogenous to import shocks, we check the robustness of our results by measuring industry total expenditure one year prior to 1999 and obtain similar results.

more vulnerable firms have no investment-wage sensitivity (The point estimate is about zero and is statistically insignificant) after the WTO accession.

Gutiérrez and Philippon (2017) document theoretical and empirical evidence that industry leaders invest more (compared to industry laggards) in response to a sharp increase in import competition after China's entry to the WTO. They argue that it is optimal for the leaders to invest more either because of the increased elasticity of substitution between different firms in the same industry or their desire to re-establish their leadership. In contrast, following a massive influx of Chinese products, the laggards are likely to exit or to downsize their investment.[17] Motivated by their findings, we examine whether the differential changes in investment-wage sensitivities following China's accession in WTO are driven by industry leaders.

[Insert Panel B of Table IV here.]

Following Gutiérrez and Philippon (2017), we identify leader (laggard) firms for each SIC industry as those firms with above-median (below-median) *Tobin's q* as of 1999, two years prior to China's accession to the WTO. We also identify leaders using different criteria: firms with above-median sales or total assets as of 1999. We estimate a regression model similar to equation (4) with an indicator variable $Leader$, which equals one if the firm is an industry leader (based on *Tobin's q*, sales or total assets) and zero otherwise. We also include the interaction terms of $Leader$ with WTO_t and $w_{i,s,t-1}$ in the regression, to examine the differential effects between leaders and laggards.

Columns (1), (2), and (3) in Panel B, Table IV present the results using *Tobin's q*, sales, total assets respectively to identify industry leaders. In all three specifications, we find that industry leaders reduce the magnitude of their investment-wage sensitivity compared to industry laggards after China's accession to WTO. These results are consistent with Gutiérrez and Philippon (2017)'s evidence that industry leaders invest more (compared to industry laggards) in response to a sharp increase in import competition after China's entry into the WTO.

[17]Using the data on new entry by Wal-Mart into the local markets, Khanna and Tice (2000) find similar results: larger and more profitable incumbents invest more (i.e., expansion in the number of stores) in response to Wal-Mart's entry, while highly levered incumbents shrink (i.e., retrenchment in the number of stores).

3. Technological Change and Automation

In response to minimum wage shocks, some firms might attempt to substitute labor with capital if the productivity per unit of capital cost exceeds the productivity per unit of labor cost, other things equal. Acemoglu and Restrepo (2019b) document that robots competing against humans reduce employment and wages for workers in US local labor markets between 1990 and 2007. In order to examine how technology affects investment-wage sensitivity, we use the observation that routine-intensive jobs are particularly susceptible to replacement by new robot technologies (Graetz and Michaels, 2017). Specifically, we measure the extent to which industries are subject to technological change (automation displacing labor) using an industry-level share of routine-task labor, $Exposure_{tech}$.

To construct $Exposure_{tech}$, we closely follow the procedure employed in Zhang (2019).[18] We first define the routine-task intensity (RTI) score for each occupation as

$$RTI_k = ln(T_k^{routine}) - ln(T_k^{abstract}) - ln(T_k^{nonroutine\ manual}) \tag{5}$$

where $T_k^{routine}$, $T_k^{abstract}$, and $T_k^{nonroutine\ manual}$ are the routine, abstract, and nonroutine manual task skill levels (scaled from 1 to 10) required by occupation k obtained from the revised fourth edition of the *Dictionary of Occupational Titles* by the U.S. Department of Labor. We classify workers as routine-task labor if their occupations' RTI score falls in the top quintile of the RTI distribution. Next, we obtain data on the number of employees and their wages for each occupation-industry pair in the year 1999, from the *Occupational Employment Statistics* by the Bureau of Labor Statistics. $Exposure_{tech}$ is the proportion of routine-task labor costs to the total industry labor costs in the year 1999. Higher the $Exposure_{tech}$ variable, the greater is the likelihood that automation would replace labor in that industry.

Using this industry-level measure of exposure to technological change ($Exposure_{tech}$), we define $Exposure_{tech,i}$ as firm i's exposure to technological change, as of 1999 in two ways: (i) As a continuous variable, we set $Exposure_{tech,i}$ to be equal to $Exposure_{tech}$ for the three-digit SIC industry to which firm i belongs; (ii) As an indicator variable, we set $Exposure_{tech,i} = 1$ for

[18]We thank Miao Ben Zhang for making the data available on his website (`https://www.miaobenzhang.com`).

firms if their continuous $Exposure_{tech,i}$ measure is above the median value of the $Exposure_{tech,i}$ distribution, and zero otherwise. Firms with $Exposure_{tech,i}$ = 1, are termed as firms with ***High Exposure*** to technological change and the firms with $Exposure_{tech,i}$ = 0, termed as firms with *Low Exposure* to technological change respectively. We then estimate a difference-in-differences regression that is similar to equation (3).

$$\begin{aligned}\frac{I_{i,s,t}}{K_{i,s,t-1}} &= \alpha_i + \alpha_t + \beta_1\, Tobin's\ q_{i,s,t-1} + \beta_2\, \frac{CF_{i,s,t}}{K_{i,s,t-1}} + \beta_3\, w_{i,s,t-1} + \beta_4\, Post \times w_{i,s,t-1} \\ &+ \beta_5\, Exposure_{tech,i} \times Post \times w_{i,s,t-1} + \beta_6\, Exposure_{tech,i} \times w_{i,s,t-1} \\ &+ \beta_7\, Exposure_{tech,i} \times Post + \beta_8\, X_{i,s,t-1} + \epsilon_{i,s,t}.\end{aligned} \tag{6}$$

We interact $w_{i,s,t-1}$, the minimum wage, with two variables, $Exposure_{tech,i}$ and $Post$ (indicating the time period after 2000). Following Acemoglu and Restrepo (2019b), we assume that the post-2000 period is more technologically advanced than the pre-2000 period. The coefficient of the triple interaction term ($Exposure_{tech,i} \times Post \times w_{i,s,t-1}$) captures difference-in-differences in investment-wage sensitivity after the year 2000 across firms that are differentially susceptible to technological change in the 21st century. We include the same set of control variables used in Panel A of Table II. We also include all interaction terms of these control variables with $Exposure_{tech,i}$ and $Post$ to capture differential effects of control variables on investment after the year 2000 across firms with different degrees of exposure.

[Insert Table V here.]

Table V presents the estimates of the relation between technological change exposure and changes in investment-wage sensitivity. We first discuss the regression results using $Exposure_{tech,i}$ as an indicator variable. Column (1), Panel A of Table V estimates that the difference-in-differences in the investment-wage sensitivity after the year 2000 between the *High Exposure* and *Low Exposure* firms (the triple interaction term, $Exposure_{tech,i} \times Post \times w_{i,s,t-1}$) is 0.020. This estimate is economically large and significant at the 1% level. We interpret this effect as follows.

For the *High Exposure* firms, the pre-2000 investment-wage sensitivity is calculated as the sum of regression coefficients: $\beta_3\ +\ \beta_6 = -0.018 - 0.012 = -0.030$ with a t-statistic of –2.91, both economically and statistically very significant. This means, these firms significantly reduced their investment in response to mandated changes in the minimum wage before the year 2001. For the same group of firms, the post-2000 investment-wage sensitivity is calculated as the sum of four regression coefficients: $\beta_3 + \beta_4 + \beta_5 + \beta_6 = -0.018 + 0.001 + 0.020 - 0.012 = -0.009$ with a t-statistic of -1.89. The estimated investment-wage sensitivity is economically close to zero and statistically significant marginally at the 10% level. This means, the *High Exposure* firms virtually do not adjust their investment in response to mandated changes in the minimum wage after the year 2000.

We now turn to the *Low Exposure* firms (the omitted group). For these firms, the investment-wage sensitivity in the pre-2000 period is the regression coefficient $\beta_3 = -0.018$ with a t-statistic of -2.17, both economically and statistically significant. *Low Exposure* firms significantly adjusted their investment in response to mandated changes in the minimum wage before the year 2000. For the same firms, the investment-wage sensitivity in the post-2000 period, is calculated as the sum of two regression coefficients: $\beta_3 + \beta_4 = -0.018 + 0.001 = -0.017$ with a t-statistic of -2.66. The estimated investment-wage sensitivity continues to be economically significant with almost the same magnitude as before, and strongly statistically significant. The estimation results in column (1) suggest that technological change-induced automation replacing labor is significantly associated with the change in investment-wage sensitivity observed over our sample period 1984–2017. In Column (2), we estimate equation (6) using $Exposure_{tech,i}$ as a continuous variable. Our inferences are qualitatively unchanged and we discuss these results in Figure III (c).

[Insert Figure III (c) here.]

In this figure, the firms with *Low* (*High*) *Exposure* to technological change are defined using $Exposure_{tech,i}$ as a continuous variable, as follows: firms with $Exposure_{tech,i}$ value one-standard deviation below (above) the mean value of the variable are *Low* (*High*) *Exposure* firms respectively. The figure shows that the *Low Exposure* firms do not change their investment-wage sensitivity over the entire sample period. The *High Exposure* firms eliminate the investment-wage sensitivity in the second half of the sample period because of technological change-induced substitution of labor with capital.

4. Weakening Union Power

It has been well documented that labor union membership in the United States has been declining over the last 50 years (Açıkgöz and Kaymak, 2014). We define annual union coverage as the percent of private-sector workers that are covered by a collective bargaining agreement each year from 1984–2017. We obtain the annual union coverage data from Hirsch and Macpherson (2003) and calculate the 15-year moving average of the annual union coverage variable. In Figure II (a), we plot the time-series of this variable (green solid line). Consistent with the well-documented decline in unionization in the United States, the 15-year moving average of union coverage drops from 13% to 8% between 1984 and 2017 (the annual union coverage decreases from 17% to 7.3% over the same period).

This declining trend in union coverage reduces the bargaining power of unions vis-a-vis firm management. Because of the weakened labor union power, firms are able to adjust their workforce (on both the extensive and intensive margins) more flexibly when faced with mandated minimum wage increases. This response is likely to make these firms less constrained by a minimum wage shock and may contribute to the elimination of the investment-wage sensitivity (i.e., negative sensitivity moving toward zero) in the post-2000 period.

First, we offer some preliminary evidence consistent with this hypothesis. In II (a), we also plot the time series of investment-wage sensitivity estimated using 15-year rolling window regressions in Panel C of Table II (red solid line). The figure shows that as union coverage declines over time, investment-wage sensitivity increases towards zero from a negative value. We estimate a univariate time-series regression of investment-wage sensitivity at time t on the 15-year moving average of the annual *Union Coverage* variable. The estimated coefficient on *Union Coverage* is −0.769 and is statistically significant with t-statistic of −4.24 (We use the Newey-West standard error, which is robust to heteroskedasticity and autocorrelation up to 14 lags, to take into account the fact that the dependent variable is estimated using the overlapping samples). The R^2 of this regression is 0.724, which suggests that changes in union coverage density can explain almost three-quarters of the variation in changes in investment-wage sensitivity over our sample period. Although this result is not causal, it is sensible and consistent with our hypothesis.

We also develop firm-level panel data evidence on whether the declining union bargaining power plays a role in explaining the changes in investment-wage sensitivity. To this end, we conduct two tests. In the first test, we measure the weakening power of labor unions as a decline in union coverage at the state or industry level. For each firm, we calculate the annualized change in union coverage of the state in which a firm is headquartered (or the annualized change in the union coverage of the industry in which a firm operates) between the first and the last year, when each firm appears in the panel data. Then we define $Large\ Decline$ as an indicator variable set to one if a firm has a below-median annualized change in the union coverage (using the firm-level distribution of this variable), and zero otherwise. Therefore, firms with $Large\ Decline = 1$ experience a larger drop in union coverage than those with $Large\ Decline = 0$ at the state or industry level. Our empirical model is as follows:

$$\begin{aligned}\frac{I_{i,s,t}}{K_{i,s,t-1}} &= \alpha_i + \alpha_t + \beta_1\ Tobin's\ q_{i,s,t-1} + \beta_2\ \frac{CF_{i,s,t}}{K_{i,s,t-1}} + \beta_3\ w_{i,s,t-1} \\ &\quad + \beta_4\ Large\ Decline_i \times w_{i,s,t-1} + \beta_5\ X_{i,s,t-1} + \epsilon_{i,s,t}.\end{aligned} \tag{7}$$

In our investment regression equation (7), we interact minimum wage $w_{i,s,t-1}$ with ***Large Decline*** to estimate the differential investment-wage sensitivities of firms with large declines in (i.e., weak union) and small declines in (i.e., strong union) union coverage. All other control variables are identical to equation (1). Although this method is subject to the typical endogeneity concerns, it provides an intuitive interpretation of the results and also captures a slow-moving trend in the union coverage rate.

[Insert Table VI here.]

Panel A of Table VI presents the results of this estimation. The $Large\ Decline$ variable is omitted estimation because it is absorbed by the firm fixed effects included in the estimation. In column (1), we use changes in state-level union coverage as a measure of union power. The estimated coefficient on the minimum wage variable ($w_{i,s,t-1}$) is –0.025 and is statistically significant at the 5% level. The interaction term ($Large\ Decline \times w_{i,s,t-1}$) is 0.013 and significant at the 5% level. These results imply that investments of firms that experience a large decline in state-level union coverage over the time period (i.e., a large decline in union power) are less sensitive to a minimum wage shock. The investment-wage sensitivity of the $Large\ Decline = 1$ group (–0.012 = –0.025 + 0.013) is about a half of the investment-wage sensitivity of $Large\ Decline = 0$ group

(–0.025). We obtain similar results in column (2) when we use industry-level union coverage to measure the bargaining power of labor unions. Overall, these results are consistent with the notion that, due to weakened union power, firms become more flexible in making investment decisions when there is a minimum wage shock.

Since union bargaining power is endogenously determined, we develop a second test. In this test, we exploit the staggered passage of right-to-work (RTW) laws by US states as an exogenous source of variation in union strength (e.g., Matsa, 2010; Chava, Danis and Hsu, 2020). In states with RTW legislation, a union and an employer cannot compel firm employees to join the union or pay membership dues as a condition of employment. As a result, unions under these laws have limited access to financial resources and manpower, which weakens the union's bargaining power with the firm management. Employees who do not join the union are still protected by the collective bargaining agreement negotiated by the union and thus the passage of RTW laws exacerbates the free-rider problem within unionized firms. We define RTW as an indicator variable that assumes the value one if the state where a firm is headquartered has passed RTW legislation as of year t and zero otherwise. We remove states that introduced an RTW law before 1984, which is the beginning of our sample period, following Chava et al. (2020).

In column (1), Panel B of Table VI, we interact the minimum wage variable $w_{i,s,t-1}$ in the investment regression with RTW to examine whether weakened union power causes corporate investment to be less sensitive to a minimum wage shock. We expect the interaction term to be positive and significant, consistent with our weakening labor unions hypothesis. In Column (1), the estimated coefficient of $w_{i,s,t-1}$ is –0.013 and statistically significant at the 5% level. The estimate on the interaction term ($RTW \times w_{i,s,t-1}$) is 0.026 and statistically significant at the 10% level. This suggests that after the passage of RTW laws, corporate investment is less sensitive to minimum wage increases. The investment-wage sensitivity before and after the passage of the law is illustrated in Figure III (d). The figure shows that the investment-wage sensitivity becomes positive after the passage of RTW law (0.013 = –0.013 + 0.026) although it is not statistically significant at the conventional level. We conclude that the passage of RTW laws eliminated the investment-wage sensitivity for firms affected by these laws in our sample.

[Insert Figure III (d) here.]

In column (2), we introduce an additional indicator variable, $Large\ Decline_{RTW}$, to isolate the effects of states with a large decline in union coverage around the RTW adoption year. We compute the three-year average of state-level union coverage before and after the RTW adoption year. The difference between these values measures the change in union coverage around the passage of RTW laws which is expected to be negative. We then define $Large\ Decline_{RTW}$ as an indicator variable set to one if a state (where a firm is headquartered) experienced a below-median change around the adoption year, and zero otherwise. The estimate on the interaction term ($RTW \times Large\ Decline_{RTW} \times w_{i,s,t-1}$) is 0.043 and statistically significant at the 1% level. This result is consistent with our hypothesis that weakening union power causes corporate investment decisions to be less sensitive to minimum wage increases. The effect is also mainly driven by states with a larger decline in union strength following the passage of RTW laws. We also note that the coefficients on RTW are positive and significant in both columns. This indicates that the passage of RTW laws also has a direct positive impact on investment, which is consistent with findings in Chava et al. (2020).

Overall, the results in this section suggest that weakening union power during 1984–2017 is an important causal mechanism that explains the elimination of investment-wage sensitivity for our sample of US firms over time.[19]

IV. Conclusion

The continuing decline in labor's share of national income in the U.S. since the 1980s has generated substantial interest and contention among academics, the press, and the public. Much of the academic literature has proposed explanations for the decline that rely on decreasing worker power vis-a-vis their employers. Whereas the literature has mostly focused on the macroeconomic implications of weakening worker power, in this article, we study microeconomic impacts by examining the extent to which declining worker power has affected firm investment responses to mandated changes in the minimum wage. In doing so, we focus on the various forces that have

[19] In Appendix C, we conduct a simple counterfactual exercise to gauge the overall economic effects of a minimum wage increase to workforce taking into account job losses due to investment cuts. The results indicate that the opportunity cost of job losses resulting from the investment cut is much larger than the benefit of wage increase for the pre-2000 period in which the investment-wage sensitivity is negative. We note that this analysis is simplistic in that it does not take into account general-equilibrium effects of the minimum wage increase on factor or output prices.

been advanced in the literature as driving the decline in worker power: globalization, technological change and the associated automation of the workplace, and weakening union power.

Our evidence on the effect of globalization on worker power comes from the ascension of China in world markets. We show that firms operating in China as of 1997 experienced a larger decrease in investment-wage sensitivities after the 1999 U.S.-China bilateral agreement, which allowed easier access to cheap Chinese labor. We also show that firms more exposed to Chinese import competition exhibited a more considerable decrease in investment-wage sensitivities after China's accession to the WTO in 2001. Regarding the effect of technological change on worker power, we provide evidence that firms more exposed to automation experienced a larger decrease in investment-wage sensitivities in the post-2000 period, during which labor has become more substitutable. Finally, corporate investment responds less negatively to minimum wage increases after the passage of the right-to-work laws that weaken union power. Collectively, these findings show that declines in investment-wage sensitivities are tied to forces that arguably have been driving the decline in worker power. In addition to adding to the academic debate, our findings on the microeconomic consequences of weakening worker power should be informative for workers, corporate managers, and policymakers.

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Table I: Descriptive Statistics

This table provides descriptive statistics of the main variables used in this study. The firm-year sample consists of 59,096 firm-year observations from 1984 to 2017. $Investment$ is measured as capital expenditures normalized by the beginning-of-year capital stock (property, plant, and equipment). $Cash\ Flow$ is calculated as earnings before extraordinary items plus depreciation, normalized by the beginning-of-the-year capital stock. $Tobin's\ q$ is a proxy for investment opportunities, which is measured as the ratio of the market value of assets to book value of assets where the market value of assets is defined as total assets plus market equity minus book equity. For state-level variables, we report their descriptive statistics based on 1,190 state-year observations. $w_{s,t-1}$ is the minimum wage at time t−1 in state s. For 1983–2014, we obtain the historical changes in minimum wages for non-farm private sector employment under state and federal laws from the Tax Policy Center. These data are sourced from the Wage and Hour Division of the U.S. Department of Labor and from the *Monthly Labor Review* by the Bureau of Labor Statistics. For 2015–2017, we hand-collect the data from the U.S. Department of Labor. Under Section 18 of the Fair Labor Standard Act, when an employee is subject to both the federal and state minimum wage laws, the employee is entitled to the higher of the two standards. The across-state variation is the cross-sectional standard deviation of state-level time-series averages of $w_{s,t-1}$ whereas the within-state variation is the average of time-series standard deviations for all states. $GDP\ growth$ is state-level annual growth rate (in percentage) of real GDP from the Bureau of Economic Analysis; $Population$ is intercensal estimates of the resident population (in thousands) for each state from the U.S. Census Bureau; $Unemployment$ is state-level unemployment rate (in percentage) from the Bureau of Labor Statistics. The definitions of all variables are provided in Appendix A.

Variables	Mean	Median	Std.Dev.	# of Obs.
Firm-Year-Level Data				
$Investment$	0.245	0.185	0.217	59,096
$Cash\ Flow$	0.379	0.269	0.895	59,096
$Tobin's\ q$	1.641	1.331	0.978	59,096
State-Year-Level Data				
$w_{s,t-1}$ ($)	5.307	5.150	1.535	1,190
$w_{s,t-1}$ ($) (across-state variation)			0.248	
$w_{s,t-1}$ ($) (within-state variation)			1.516	
$GDP\ Growth$ (%)	2.576	2.400	2.801	1,190
$Population$ (thousands)	5,687	3,506	6,736	1,190
$Unemployment$ (%)	5.792	5.400	2.002	1,190

Table II: Investment Sensitivity to Minimum Wage

Panel A presents fixed effect OLS regressions of corporate investment on minimum wages in equation (1). The dependent variable is *Investment*, measured as capital expenditures normalized by the beginning-of-the-year capital stock (property, plant, and equipment). $w_{i,s,t-1}$ is the minimum wage at time $t-1$ in state s where firm i's headquarters is located. For 1983–2014, we obtain the historical changes in minimum wages for non-farm private sector employment under state and federal laws from the Tax Policy Center. These data are sourced from the Wage and Hour Division of the U.S. Department of Labor and from the *Monthly Labor Review* by the Bureau of Labor Statistics. For 2015–2017, we hand-collect the data from the U.S. Department of Labor. Under Section 18 of the Fair Labor Standard Act, when an employee is subject to both the federal and state minimum wage laws, the employee is entitled to the higher of the two standards. We exclude 15 states that have indexed their minimum wage rates to inflation for the identification reason discussed in Section II.B. We measure *Cash Flow* as earnings before extraordinary items plus depreciation normalized by the beginning-of-the-year capital stock and *Tobin's q* as a ratio of the market value of assets to book value of assets. We also control for state-level macro-variables: real GDP growth rates, log of population, and unemployment rates. The definitions of all variables are provided in Appendix A. Standard errors in parentheses are robust to heteroskedasticity and clustered by state. The full sample period runs from 1984 to 2017. H_0: (2)[$w_{i,s,t-1}$] - (3)[$w_{i,s,t-1}$] = 0 is based on a two-tailed test with χ^2-statistics in squared bracket.

Panel A. Minimum Wage and Corporate Investment: Full Sample, Pre-, and Post-2000

	Dependent Variable: $Investment_{i,s,t}$		
	Full Sample	1984 to 2000	2001 to 2017
	(1)	(2)	(3)
$w_{i,s,t-1}$	-0.017*	-0.038***	0.001
	(0.009)	(0.012)	(0.003)
$Cash\ Flow$	0.043***	0.098***	0.029***
	(0.002)	(0.005)	(0.002)
$Tobin's\ q$	0.063***	0.066***	0.053***
	(0.003)	(0.003)	(0.003)
$GDP\ growth$	0.002**	0.003**	0.002**
	(0.001)	(0.001)	(0.001)
ln($Population$)	-0.108	-0.145*	-0.192***
	(0.065)	(0.076)	(0.049)
$Unemployment$	0.0002	0.001	0.002
	(0.001)	(0.001)	(0.002)
H_0: (2)[$w_{i,s,t-1}$] - (3)[$w_{i,s,t-1}$] = 0		-0.039*** [12.54]	
Firm and Year FEs	Yes	Yes	Yes
# of Firm-Year Obs.	59,096	31,408	27,688
Adjusted R^2	0.140	0.135	0.122

Table II: Investment Sensitivity to Minimum Wage (continued)

Panel B repeats the estimation of column (2) of Panel A, using 1,000 simulated samples where we randomly assign each firm to a particular state. The panel shows the empirical distribution of the coefficient on $w^{Pseudo}_{i,s,t-1}$. The sample period runs from 1984 to 2000. Panel C repeats the estimation of column (1) of Panel A using a 15-year rolling window sample that starts from 1984. The total number of estimated regressions is twenty. Standard errors in parentheses are robust to heteroskedasticity and clustered by state.

Panel B. Placebo Test: Regression Coefficients from Bootstrapped Sample

	Dependent Variable: $Investment_{i,s,t}$ / Sample from 1984 to 2000										
	Col. (2) Panel A	Mean	p1	p5	p10	p25	p50	p75	p90	p95	p99
$w_{i,s,t-1}$	-0.038	0.002	-0.017	-0.013	-0.010	-0.004	0.001	0.007	0.012	0.015	0.022

Panel C. Time-Varying Effects of Minimum Wage on Investment: 15-Year Rolling Window Regressions

Sample Period From	Sample Period To	β_3 (Coefficient on $w_{i,s,t-1}$)	t-statistics
1984	1998	-0.026**	-2.667
1985	1999	-0.033***	-3.554
1986	2000	-0.037***	-3.638
1987	2001	-0.040***	-3.813
1988	2002	-0.040***	-3.380
1989	2003	-0.031***	-3.820
1990	2004	-0.031***	-4.521
1991	2005	-0.034***	-4.906
1992	2006	-0.030***	-4.115
1993	2007	-0.027***	-3.677
1994	2008	-0.024***	-3.340
1995	2009	-0.024***	-2.881
1996	2010	-0.020**	-2.592
1997	2011	-0.014**	-2.128
1998	2012	-0.010*	-1.813
1999	2013	-0.006	-1.296
2000	2014	-0.001	-0.224
2001	2015	0.002	0.585
2002	2016	0.003	1.206
2003	2017	0.002	0.451

Table II: Investment Sensitivity to Minimum Wage (continued)

In Panel D, we test for a regime shift in the relation between corporate investment and minimum wage. We assume a single, known structural break and allow all the coefficients to change after the structural break year:

$$\frac{I_{i,s,t}}{K_{i,s,t-1}} = \alpha_i + \alpha_t + \beta_1 w_{i,s,t-1} + \beta_2 Z_{i,s,t-1} + d_t(k)\Big[\beta_3 w_{i,s,t-1} + \beta_4 Z_{i,s,t-1}\Big] + \epsilon_{i,s,t},$$

where i, s, and t index firms, states, and years; α_i and α_t is a set of firm and year fixed effects, respectively; $Investment$ $(= \frac{I_{i,s,t}}{K_{i,s,t-1}})$ is investment rates; $w_{i,s,t-1}$ is minimum wage at time t −1 in state s where firm i's headquarters is located; $Z_{i,s,t-1}$ is a set of firm- and state-level control variables used in equation (1). $d_t(k)$ equals one if t is greater than or equal to the assumed year of structural break k, and zero otherwise. We require at least five years of data for both periods (pre- and post-break) and hence k runs from 1989 to 2013. We calculate the Akaike's information criterion (AIC) as $-2ln(L[k]) + 2p$ where $ln(L[k])$ is the maximized log-likelihood of the model in which the assumed structural break is year k and p is the number of parameters estimated. The Bayesian (or Schwarz's) information criterion is defined as $-2ln(L[k]) + pln(N)$ where N is the sample size. The figure plots the Akaike's and Bayesian information criteria for each assumed year of structural break.

Panel D. Analysis of Structural Breaks: Single Known Break

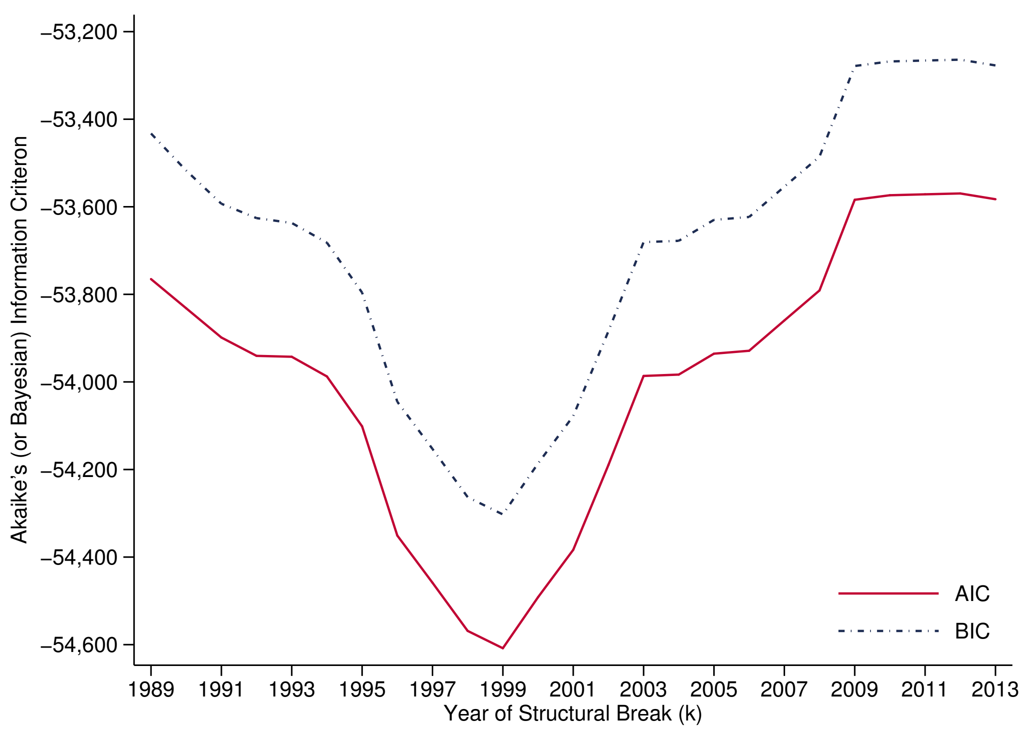


Table III: U.S. Firms' Access to Cheap Labor: 1999 US-China Bilateral Agreement

Column (1) presents difference-in-differences regressions of investment on minimum wages interacted with two indicators, $China97$ and $Agreement$, in equation (3). $w_{i,s,t-1}$ is the minimum wage at time $t-1$ in state s where firm i's headquarters is located. $China97$ is an indicator variable set to one if a firm has at least one subsidiary in China as of 1997, two years prior to the US-China bilateral agreement in 1999, and zero otherwise; $Agreement$ indicates the time period after the agreement (including 1999). We use hand collected information from 10-k filings to identify U.S. firms' Chinese subsidiaries. The coefficient of the triple interaction term ($China97 \times Agreement \times w_{i,s,t-1}$) captures difference-in-differences in investment-wage sensitivity after the agreement between treated ($China97$ = 1) and control ($China97$ = 0) firms. In column (2), we introduce another group by defining $China04$ as an indicator variable set to one for firms without any subsidiary in China as of 1997 but having at least one subsidiary as of 2004 (five years after the agreement), and zero otherwise. The omitted group consists of firms that have no operations in China, that is, $China97 = China04 = 0$. In all columns, we include the same set of control variables used in Panel A of Table II and all interaction terms of these control variables with $China97$ (or $China04$) and $Agreement$ indicators. $China97$ ($China04$) and $Agreement$ indicators are absorbed by firm and year fixed effects, respectively. The definitions of all variables are provided in Appendix A. Standard errors in parentheses are robust to heteroskedasticity and clustered by state. The sample period runs from 1984 to 2017.

Dependent Variable →	$Investment_{i,s,t}$	
	Two Groups	Three Groups
	(1)	(2)
$w_{i,s,t-1}$	-0.017**	-0.016*
	(0.008)	(0.008)
$Agreement \times w_{i,s,t-1}$	0.005	0.002
	(0.005)	(0.006)
$China97 \times Agreement \times w_{i,s,t-1}$	0.035***	0.038***
	(0.010)	(0.010)
$China04 \times Agreement \times w_{i,s,t-1}$		0.029**
		(0.014)
$China97 \times w_{i,s,t-1}$	-0.028***	-0.029***
	(0.008)	(0.009)
$China04 \times w_{i,s,t-1}$		-0.015
		(0.012)
$China97 \times Agreement$	-0.093	-0.051
	(0.179)	(0.192)
$China04 \times Agreement$		0.408*
		(0.239)
Investment Sensitivity to Minimum Wage [t-stat]		
Before (omitted group)	-0.017** [-2.14]	-0.016* [-1.98]
After (omitted group)	-0.012* [-1.92]	-0.014** [-2.34]
Before ($China04$ = 1)		-0.031** [-2.18]
After ($China04$ = 1)		0.000 [0.06]
Before ($China97$ = 1)	-0.045*** [-3.73]	-0.045*** [-3.71]
After ($China97$ = 1)	-0.005 [-0.86]	-0.005 [-0.96]
Controls / Interaction of Controls / Firm and Year FEs	Yes	Yes
# of Firm-Year Obs.	59,096	59,096
Adjusted R^2	0.157	0.158

Table IV: U.S. Firms' Increased Exposure to Chinese Import Competition: 2001 China's Accession to WTO

This table presents difference-in-differences regressions of investment on minimum wages interacted with $Exposure_{UC,i}$ and WTO in equation (4). $w_{i,s,t-1}$ is the minimum wage at time $t-1$ in state s where firm i's headquarters is located. WTO indicates the time period after China's entry to the World Trade Organization in 2001. For each U.S. industry, we first define its exposure to imports from China, $Exposure_{UC}$, as the log of the import penetration ratio (Bernard et al., 2006) if the industry is classified to be in the tradable sector(Mian and Sufi, 2014), and zero otherwise. We then measure firm i's exposure as of 1999 ($Exposure_{UC,i}$), two years prior to China's accession to the WTO, in two ways: (i) As a continuous variable, we set $Exposure_{UC,i} = Exposure_{UC}$ for the four-digit NAICS industry to which firm i belongs; (ii) As an indicator variable, we set $Exposure_{UC,i} = 1$ for firms in the industries with above-median $Exposure_{UC}$, and zero otherwise. The coefficient of the triple interaction term ($Exposure_{UC,i} \times WTO \times w_{i,s,t-1}$) captures difference-in-differences in investment-wage sensitivity before and after the year 2001 across firms that are differentially exposed to import competition. In column (3), we instrument for $Exposure_{UC,i}$ with $Exposure_{OC,i}$, Chinese import exposure for eight other high-income countries. In all columns, we include the same set of control variables used in Panel A of Table II and all interaction terms of these control variables with $Exposure_{UC,i}$ and WTO variables. In column (1), we define *No* (*High*) *Exposure* firms as firms with $Exposure_{UC,i} = 0$ (1) as an indicator variable. In columns (2) and (3), we define *No* (*High*) *Exposure* firms as firms with zero (one-standard deviation above the mean) value of $Exposure_{UC,i}$ as a continuous variable. $Exposure_{UC,i}$ and WTO variables are absorbed by firm and year fixed effects, respectively. The definitions of all variables are provided in Appendix A. Standard errors in parentheses are robust to heteroskedasticity and clustered by state. The sample period runs from 1984 to 2017.

Panel A. Full Sample

Dependent Variable →	$Investment_{i,s,t}$		
	OLS		2SLS
Exposure to Import Competition →	Dummy	Continuous	Continuous
	(1)	(2)	(3)
$w_{i,s,t-1}$	-0.021**	-0.025**	-0.024**
	(0.010)	(0.011)	(0.010)
$WTO \times w_{i,s,t-1}$	0.009	0.013	0.012
	(0.008)	(0.008)	(0.008)
$Exposure_{UC,i} \times WTO \times w_{i,s,t-1}$	0.041***	0.019***	0.023***
	(0.004)	(0.006)	(0.006)
$Exposure_{UC,i} \times w_{i,s,t-1}$	-0.027***	-0.008**	-0.015***
	(0.004)	(0.003)	(0.004)
$Exposure_{UC,i} \times WTO$	-0.094	-0.048	-0.013
	(0.090)	(0.050)	(0.074)
Investment Sensitivity to Minimum Wage [t-stat]			
Before (*No Exposure*)	-0.021** [-2.05]	-0.025** [-2.35]	-0.024** [-2.30]
After (*No Exposure*)	-0.012 [-1.65]	-0.013 [-1.59]	-0.012 [-1.57]
Before (*High Exposure*)	-0.048*** [-5.69]	-0.040*** [-4.45]	-0.047*** [-4.92]
After (*High Exposure*)	0.002 [0.34]	0.006 [0.66]	0.001 [0.06]
Controls / Interaction of Controls	Yes	Yes	Yes
Firm and Year FEs	Yes	Yes	Yes
# of Firm-Year Obs.	46,168	46,168	46,104
Adjusted R^2	0.152	0.151	0.152

Table IV: U.S. Firms' Increased Exposure to Chinese Import Competition: 2001 China's Accession to WTO (continued)

Panel B presents difference-in-difference regressions of investment on minimum wages interacted with two indicators, $Leaders$ and WTO. $Leaders$ is an indicator variable set to one for firms with above-median *Tobin's q*, sales, or total assets for each two-digit SIC industry as of 1999, two years prior to China's entry to the WTO, and zero otherwise. $Leaders$ and WTO indicators are absorbed by firm and year fixed effects, respectively. *Laggards* indicates firms that are not industry leaders (i.e., $Leaders = 0$). The definitions of all variables are provided in Appendix A. Standard errors in parentheses are robust to heteroskedasticity and clustered by state. The sample period runs from 1984 to 2017.

Panel B. Industry Leaders vs. Laggards

Dependent Variable →	$Investment_{i,s,t}$		
Leaders vs. Laggards Based on →	Tobin's q	Sales	Total Assets
	(1)	(2)	(3)
$w_{i,s,t-1}$	-0.016**	-0.009	-0.013
	(0.007)	(0.010)	(0.010)
$WTO \times w_{i,s,t-1}$	-0.002	-0.009	-0.006
	(0.008)	(0.009)	(0.009)
$Leader \times WTO \times w_{i,s,t-1}$	0.021***	0.036***	0.031***
	(0.005)	(0.005)	(0.007)
$Leader \times w_{i,s,t-1}$	-0.015**	-0.028***	-0.022***
	(0.006)	(0.005)	(0.005)
$Leader \times WTO$	0.057	-0.013	-0.064
	(0.121)	(0.092)	(0.126)
Investment Sensitivity to Minimum Wage [t-stat]			
Before (*Laggards*)	-0.016** [-2.27]	-0.009 [-0.95]	-0.013 [-1.29]
After (*Laggards*)	-0.017*** [-2.92]	-0.018** [-2.56]	-0.019** [-2.62]
Before (*Leaders*)	-0.031*** [-3.16]	-0.037*** [-3.58]	-0.035*** [-3.73]
After (*Leaders*)	-0.012** [-2.11]	-0.009* [-1.73]	-0.010* [-1.77]
Controls / Interaction of Controls	Yes	Yes	Yes
Firm and Year FEs	Yes	Yes	Yes
# of Firm-Year Obs.	37,484	38,844	38,829
Adjusted R^2	0.179	0.178	0.178

Table V: Technological Change and Automation

This table presents difference-in-differences regressions of investment on minimum wages interacted with two variables, $Exposure_{tech,i}$ and $Post$, in equation (6). We measure the extent to which industries are subject to technological change using an industry-level share of routine-task labor ($Exposure_{tech}$). To construct this variable, we follow the procedure employed in Zhang (2019). We first define the routine-task intensity (RTI) score for each occupation as $RTI_k = ln(T_k^{routine}) - ln(T_k^{abstract}) - ln(T_k^{nonroutine\ manual})$ where $T_k^{routine}$, $T_k^{abstract}$, and $T_k^{nonroutine\ manual}$ are the routine, abstract, and nonroutine manual task skill levels (scaled from 1 to 10) required by occupation k. Each occupation's required skill level data are obtained from revised fourth edition of the *Dictionary of Occupational Titles* by the U.S. Department of Labor. We classify workers as routine-task labor if their occupations' RTI score falls in the top quintile of the RTI distribution. We then construct $Exposure_{tech}$ as the proportion of routine-task labor costs to the total industry labor cost. We obtain data on the number of employees and their wages for each occupation-industry from the *Occupational Employment Statistics* by the Bureau of Labor Statistics. With the industry-level share of routine-task labor, we define $Exposure_{tech,i}$ as firm i's exposure to technological change, as of 1999 in two ways: (i) As a continuous variable, we set $Exposure_{tech,i}$ to be equal to $Exposure_{tech}$ for the three-digit SIC industry to which firm i belongs; (ii) As an indicator variable, we set $Exposure_{tech,i} = 1$ for firms if their continuous $Exposure_{tech,i}$ measure is above the median value of the $Exposure_{tech,i}$ distribution, and zero otherwise. $w_{i,s,t-1}$ is the minimum wage at time $t-1$ in state s where firm i's headquarters is located; $Post$ indicates period after 2001. The coefficient of triple interaction term ($Exposure_{tech,i} \times Post \times w_{i,s,t-1}$) captures difference-in-differences in investment-wage sensitivity after the year 2001 across firms that are differentially susceptible to technological change in the 21st century. In all columns, we include the same set of control variables used in Panel A of Table II and their interaction terms with $Exposure_{tech,i}$ and $Post$. In column (1), we define *Low* (*High*) *Exposure* group as firms with $Exposure_{tech,i} = 0$ (1) as an indicator variable. In column (2), we define *Low* (*High*) *Exposure* group as firms with $Exposure_{tech,i}$ (as a continuous variable) value one-standard deviation below (above) the mean value of $Exposure_{tech,i}$. $Exposure_{tech,i}$ and $Post$ variables are absorbed by firm and year fixed effects, respectively. The definitions of all variables are provided in Appendix A. Standard errors in parentheses are robust to heteroskedasticity and clustered by state. The sample period runs from 1984 to 2017.

Dependent Variable →	$Investment_{i,s,t}$	
Exposure to Technological Change →	Dummy	Continuous
	(1)	(2)
$w_{i,s,t-1}$	-0.018**	-0.016*
	(0.008)	(0.009)
$Post \times w_{i,s,t-1}$	0.001	-0.001
	(0.008)	(0.010)
$Exposure_{tech,i} \times Post \times w_{i,s,t-1}$	0.020***	0.073**
	(0.006)	(0.031)
$Exposure_{tech,i} \times w_{i,s,t-1}$	-0.012**	-0.049*
	(0.005)	(0.025)
$Exposure_{tech,i} \times Post$	0.083	0.405
	(0.114)	(0.507)
Investment Sensitivity to Minimum Wage [t-stat]		
Before (*Low Exposure*)	-0.018** [-2.17]	-0.019** [-2.19]
After (*Low Exposure*)	-0.017** [-2.66]	-0.015** [-2.32]
Before (*High Exposure*)	-0.030*** [-2.91]	-0.030*** [-3.10]
After (*High Exposure*)	-0.009* [-1.89]	-0.010* [-1.88]
Controls / Interaction of Controls	Yes	Yes
Firm and Year FEs	Yes	Yes
# of Firm-Year Obs.	36,213	36,213
Adjusted R^2	0.176	0.176

Table VI: Weakening Union Power

Panel A reports fixed effect OLS regression estimates on relation between labor union power and investment-wage sensitivity in equation (7). We measure the weakening power of labor unions as a decline in union coverage at the state- or industry-level. Union coverage is defined as the percentage of private-sector workers that are covered by a collective bargaining agreement. We first calculate annualized changes in union coverage of the state in which a firm is headquartered (or those of industry in which a firm operates) over the entire sample period. Then we define $Large\ Decline$ as an indicator variable set to one if a firm has a below-median annualized change in union coverage (using the firm-level distribution), and zero otherwise. We obtain the annual state- and industry-level union coverage data from Hirsch and Macpherson (2003). $w_{i,s,t-1}$ is the minimum wage at time $t-1$ in state s where firm i's headquarters is located. In all columns, we include the same set of control variables used in Panel A of Table II and all interaction terms of these control variables with $Large\ Decline$. The definitions of all variables are provided in Appendix A. Standard errors in parentheses are robust to heteroskedasticity and clustered by state. The sample period runs from 1984 to 2017.

Panel A. Large Decline in Union Coverage

Dependent Variable →	$Investment_{i,s,t}$	
Union Power is Measured as Change in →	State-level Union Coverage	Industry-level Union Coverage
	(1)	(2)
$w_{i,s,t-1}$	-0.025**	-0.026***
	(0.012)	(0.009)
$Large\ Decline \times w_{i,s,t-1}$	0.013**	0.015***
	(0.006)	(0.004)
$Large\ Decline$	(omitted)	(omitted)
Controls / Interaction of Controls	Yes	Yes
Firm and Year FEs	Yes	Yes
# of Firm-Year Obs.	59,096	55,974
Adjusted R^2	0.143	0.138

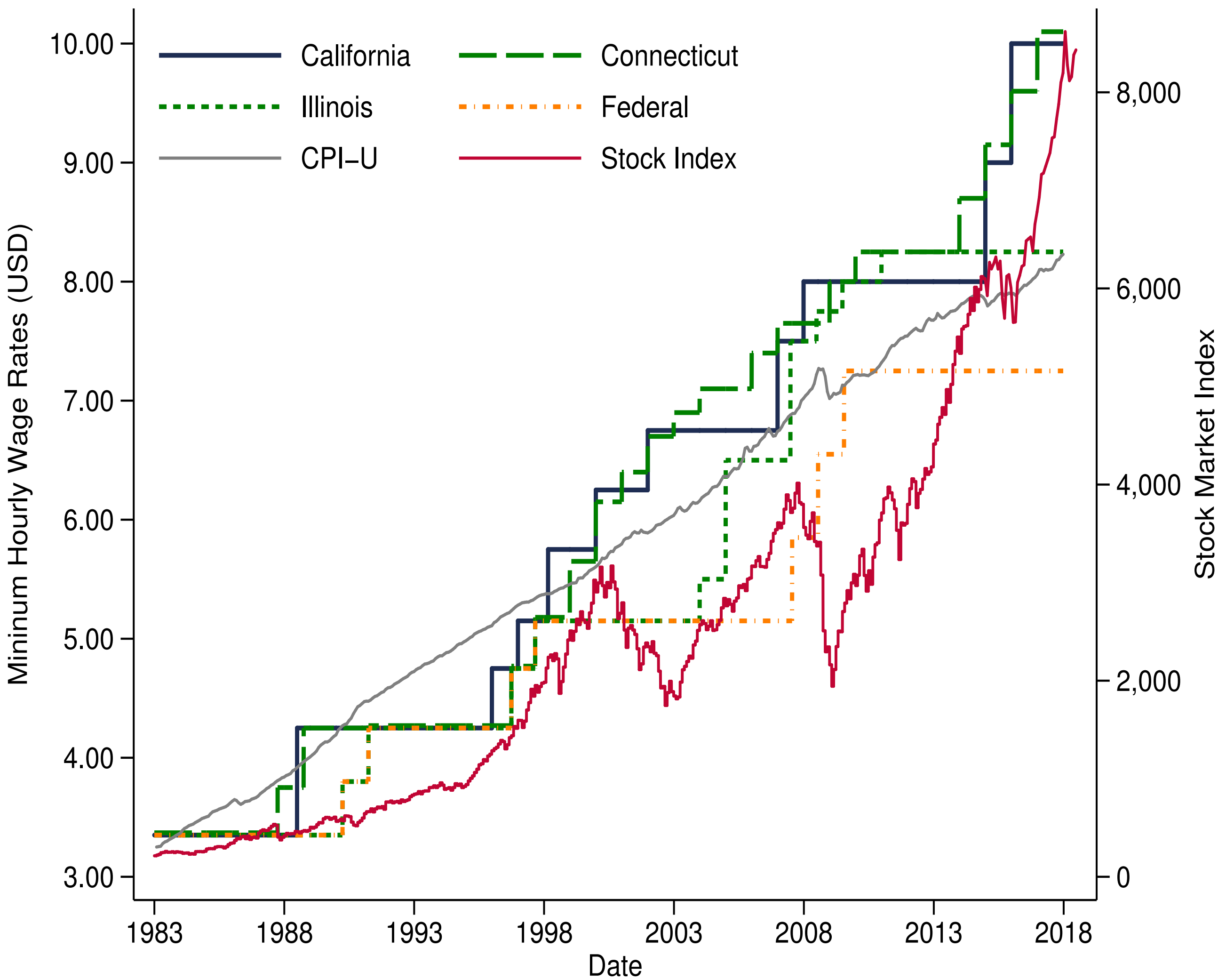
Table VI: Weakening Union Power (continued)

Panel B presents difference-in-differences regressions of investment on minimum wages interacted with an RTW indicator variable. We use the passage of right-to-work (RTW) laws to measure the weakening power of labor unions. We define RTW as an indicator variable that assumes the value one, if the state where a firm is headquartered has passed RTW legislation as of year t, and zero otherwise. $w_{i,s,t-1}$ is the minimum wage at time $t-1$ in state s where firm i's headquarters is located. We remove states that introduced an RTW law before 1984, which is the beginning of our sample period. In column (2), we introduce an additional indicator, $Large\ Decline_{RTW}$, to isolate the effects of states with a large decline in union coverage around the RTW adoption year. We compare the three-year average of state-level union coverage before and after the RTW adoption year. The difference between these values measures the change in union coverage around the passage of RTW laws which is expected to be negative. We then define $Large\ Decline_{RTW}$ as an indicator variable set to one if a state (where a firm is headquartered) experienced a below-median change in union coverage rate around the adoption year, and zero otherwise. In all columns, we include the same set of control variables used in Panel A of Table II and all interaction terms of these control variables with RTW. $Large\ Decline_{RTW}$ is absorbed by firm fixed effects. The definitions of all variables are provided in Appendix A. Standard errors in parentheses are robust to heteroskedasticity and clustered by state. The sample period runs from 1984 to 2017.

Panel B. Passage of Right-to-Work (RTW) Laws

Dependent Variable $\rightarrow$	$Investment_{i,s,t}$	
		States With Large Decline in Union Coverage Around the Adoption Year
	(1)	(2)
$w_{i,s,t-1}$	-0.013**	-0.012**
	(0.005)	(0.005)
$RTW \times w_{i,s,t-1}$	0.026*	0.011
	(0.015)	(0.009)
$RTW \times Large\ Decline_{RTW} \times w_{i,s,t-1}$		0.043***
		(0.009)
$Large\ Decline_{RTW} \times w_{i,s,t-1}$		0.002
		(0.007)
RTW	0.751*	1.208***
	(0.385)	(0.419)
$RTW \times Large\ Decline_{RTW}$		-0.306***
		(0.055)
Controls / Interaction of Controls	Yes	Yes
Firm and Year FEs	Yes	Yes
# of Firm-Year Obs.	37,111	37,111
Adjusted R^2	0.144	0.144

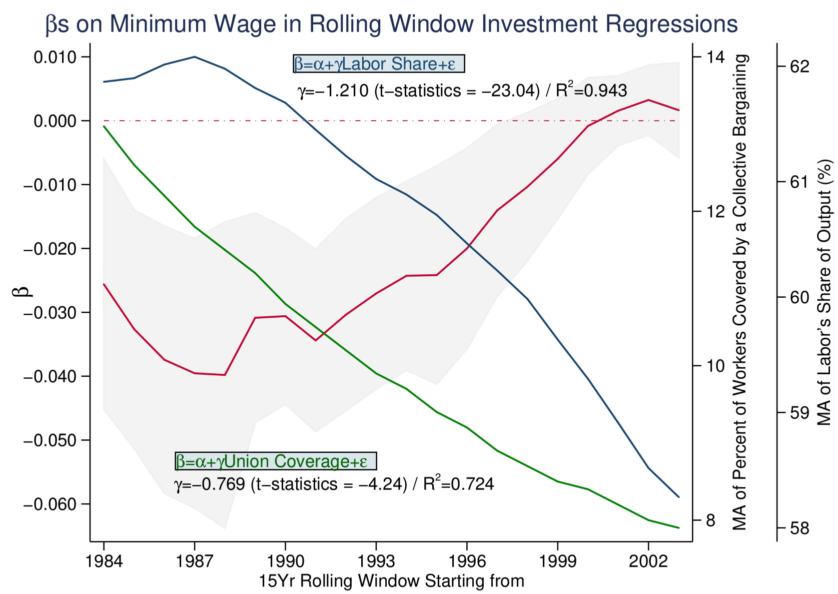
Figure I: Minimum Hourly Wage Across some U.S. States (California, Connecticut, and Illinois), 1983-2017



This figure shows the time-series of minimum hourly wage rates for California, Connecticut, Illinois, and the federal government as an example for the time period 1983 to 2017. For 1983–2014, we obtain the historical changes in minimum wages for non-farm private sector employment under state and federal laws from the Tax Policy Center. These data are sourced from the Wage and Hour Division of the U.S. Department of Labor and from the *Monthly Labor Review* by the Bureau of Labor Statistics. For 2015–2017, we hand-collect the data from the U.S. Department of Labor. Under Section 18 of the Fair Labor Standard Act, when an employee is subject to both the federal and state minimum wage laws, the employee is entitled to the higher of the two standards. We also plot the time-series of the Consumer Price Index for All Urban Consumers (CPI-U) by setting the index value in January 1983 to a wage rate of $3.25 per hour on the left axis and the time-series of the value-weighted stock market (NYSE/AMEX/NASDAQ) index on the right axis.

Figure II: Investment Sensitivity to Minimum Wage Over Four Decades

(a)

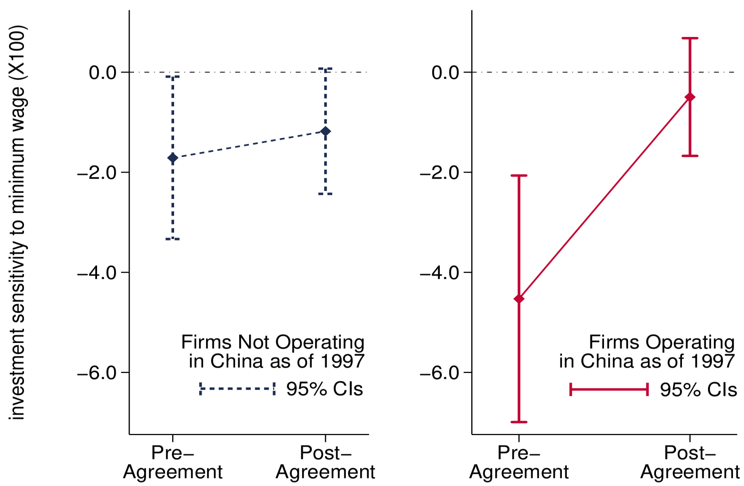


(b)

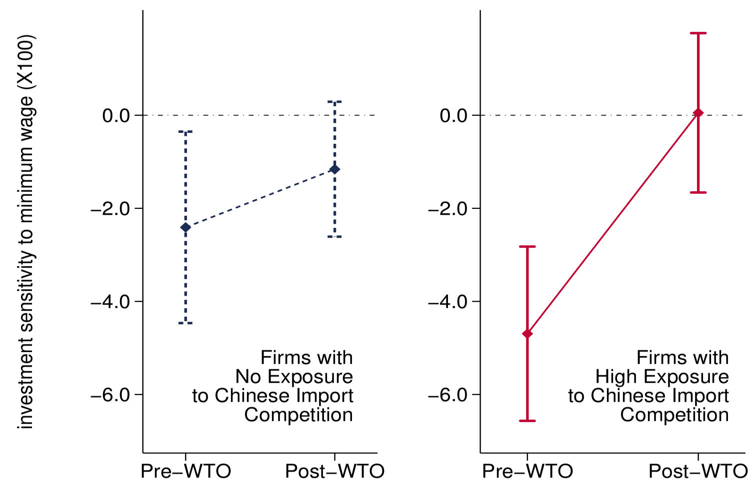
In Panel (a), the red line plots the time-series of estimated investment-wage sensitivity (β) from the 15-year rolling window investment regressions in Panel C of Table II. The grey shaded area indicates the 95% confidence intervals. In Panel (b), the red line plots the time-series of corresponding t-statistics with a horizontal line indicating t-statistics of −2. The x-axis in Panels (a) and (b) refers to the starting years of the 15-year rolling windows. The estimated coefficients and t-statistics are tabulated in Panel C of Table II. In Panel (a), we also plot two additional time-series: the green line plots the union coverage (defined as the percentage of private-sector workers that are covered by a collective bargaining agreement), and the blue line plots the labor's share of output in the nonfarm business sector. We plot the 15-year moving average of these variables using the same sample windows as used in rolling window investment regressions. The union coverage data come from Hirsch and Macpherson (2003) and the labor share data come from the U.S. Bureau of Labor Statistics. We estimate a univariate time-series regression of investment-wage sensitivity at time t on the 15-year moving average of the annual union coverage (labor share). The estimated coefficient on union coverage (labor share) is −0.769 (−1.210) with t-statistics of −4.24 (−23.04) based on the Newey-West standard error that is robust to heteroskedasticity and autocorrelation up to 14 lags. The R^2 of these regressions are 0.724 and 0.943, respectively.

Figure III: Declining Worker Power and Changes in Investment Sensitivity to Minimum Wage

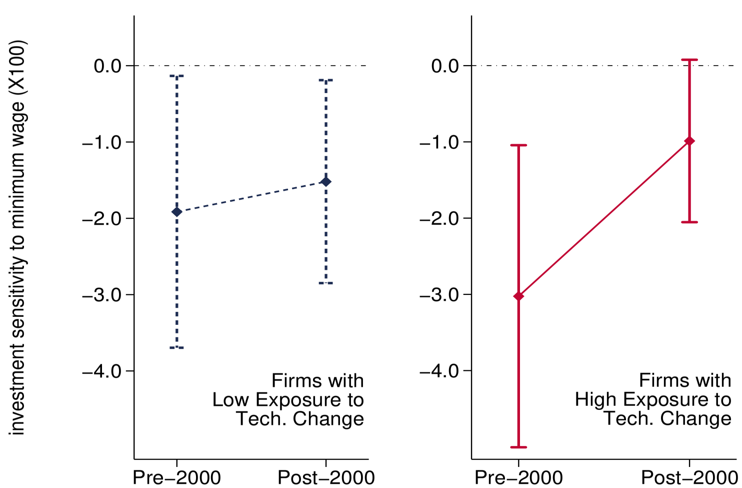
(a) 1999 US-China Bilateral Agreement



(b) Chinese Import Competition



(c) Technological Change / Automation



(d) Declining Union Power: Right-to-Work (RTW) Laws

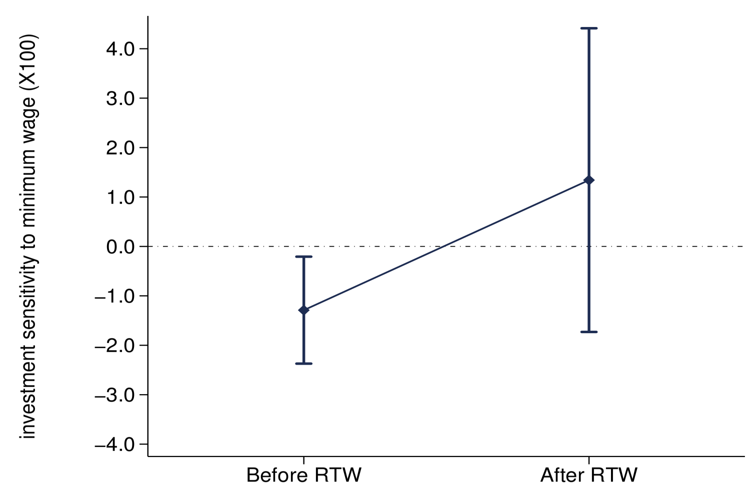


Figure III: Declining Worker Power and Changes in Investment Sensitivity to Minimum Wage (continued)

These figures plot changes in investment-wage sensitivity for two groups of firms that are identified in Tables III to VI. Figure (a) is based on the estimates in column (1) of Table III. The first group consists of firms that do not operate in China as of 1997 ($China97 = 0$), two years prior to the 1999 US-China bilateral agreement; the second group consists of firms operating in China as of 1997 ($China97 = 1$). We identify firms operating in China (i.e., having at least one subsidiary in China) using hand collected information from 10-K filings on US firms' Chinese subsidiaries. Dots indicate the estimated investment-wage sensitivity and the vertical lines around these point estimates are 95% confidence intervals. Figure (b) is based on the estimates in column (3) of Table IV. We define a U.S. industry's exposure to imports from China, $Exposure_{UC}$, as the log of the import penetration ratio if an industry is classified to be in the tradable sector, and zero otherwise. We define firm i's exposure ($Exposure_{UC,i}$) to imports from China as of 1999, two years prior to China's accession to the World Trade Organization, as $Exposure_{UC}$ for the four-digit NAICS industry to which firm i belongs. We define the first group as firms with no exposure to import competition ($Exposure_{UC,i} = 0$) and the second group as firms with one-standard-deviation above the average exposure ($Exposure_{UC,i} = \mu + \sigma$). To capture supply-driven component in U.S. imports from China, we instrument for $Exposure_{UC,i}$ with $Exposure_{OC,i}$ in which $Exposure_{OC,i}$ is Chinese import exposure for eight other high-income countries. Figure (c) is based on the estimates in column (2) of Table V. We measure the extent to which industries are subject to technological change using an industry-level share of routine-task labor ($Exposure_{tech}$). To construct this variable, we follow the procedure employed in Zhang (2019). We first define the routine-task intensity (RTI) score for each occupation as $RTI_k = ln(T_k^{routine}) - ln(T_k^{abstract}) - ln(T_k^{nonroutine\ manual})$ where $T_k^{routine}$, $T_k^{abstract}$, and $T_k^{nonroutine\ manual}$ are the routine, abstract, and nonroutine manual task skill levels (scaled from 1 to 10) required by occupation k. Each occupation's required skill level data are obtained from the revised fourth edition of the *Dictionary of Occupational Titles* by the U.S. Department of Labor. We classify workers as routine-task labor if their occupations' RTI score falls in the top quintile of the RTI distribution. We then construct $Exposure_{tech}$ as the proportion of routine-task labor costs to the total industry labor cost. We obtain data on the number of employees and their wages for each occupation-industry from the *Occupational Employment Statistics* by the Bureau of Labor Statistics. With the industry-level share of routine-task labor, we define firm i's exposure ($Exposure_{tech,i}$) to technological change as of 1999 as $Exposure_{tech}$ for the three-digit SIC industry to which firm i belongs. We define firms with low (high) exposure to technological change as firms with $Exposure_{tech,i}$ value one-standard deviation below (above) the mean value of $Exposure_{tech,i}$. Figure (d) is based on the estimates in column (1) of Panel B in Table VI. We use the passage of right-to-work (RTW) laws to measure the weakening power of labor unions. Specifically, we identify firms that are headquartered in a state as having weak (strong) union power if the state has (not) passed the RTW legislation. The detailed definitions of all variables are provided in Appendix A.

For Online Publication

Appendices

Appendix A. Variable Definitions

Variables	Definition [Compustat designations where appropriate]
Investment	Capital expenditures [CAPX] normalized by the beginning-of-the-year capital stock (property, plant, and equipment) [PPENT]
$w_{i,s,t-1}$	Minimum wage at time $t-1$ in state s where firm i's headquarters is located; We use the historical changes in minimum wages under state laws reported by the Tax Policy Center which uses data from the Wage and Hour Division of the U.S. Department of Labor and from the *Monthly Labor Review* by the Bureau of Labor Statistics. In cases where an employee is subject to both the state and federal minimum wage laws, the employee is entitled to the higher of the two under Section 18 of the Fair Labor Standard Act.
Cash Flow	Earnings before extraordinary items [IB] plus depreciation [DP] normalized by the beginning-of-the-year capital stock [PPENT]
Tobin's q	The ratio of the market value of assets to book value of assets [AT] where the market value of assets is defined as total assets [AT] plus market equity minus book equity in which market equity is defined as common shares outstanding [CSHO] times fiscal-year closing price [PRCC_F]; book equity is calculated as stockholders' equity [SEQ] minus preferred stock liquidating value [PSTKL] plus balance sheet deferred taxes and investment tax credit [TXDITC] when available minus post-retirement assets [PPROR] when available
GDP growth	State-level annual growth rate of real GDP from the Bureau of Economic Analysis
ln(Population)	Log of intercensal estimates of the resident population for each state from the U.S. Census Bureau
Unemployment	State-level unemployment rate from the Bureau of Labor Statistics
Agreement	An indicator variable for the time period after the US-China bilateral agreement in 1999 (including 1999)
China97	An indicator variable set to one for firms with at least one subsidiary in China two years prior to the US-China bilateral agreement in 1999, and zero otherwise. We use hand-collected information from 10-k filings to identify U.S. firms' Chinese subsidiaries.
China04	An indicator variable set to one for firms without any subsidiary in China as of 1997 but having at least one subsidiary in China as of 2004 (five years after the US-China bilateral agreement in 1999), and zero otherwise
WTO	An indicator for the time period after China's entry to the World Trade Organization (WTO) in 2001
$Exposure_{UC,i}$	We first define a U.S. industry's exposure to imports from China ($Exposure_{UC}$) as Chinese import penetration ratio (Bernard et al., 2006) if firms are classified as tradable sector (Mian and Sufi, 2014), and zero otherwise. We then define firm i's exposure to imports from China as of 1999 ($Exposure_{UC,i}$) in two ways: (i) As a continuous variable, we set it to be equal to $Exposure_{UC}$ for the four-digit NAICS industry to which firm i belongs; (ii) As an indicator variable, we set $Exposure_{UC,i} = 1$ for firms in the industries with above-median $Exposure_{UC}$, and zero otherwise.

$Exposure_{OC,i}$	Chinese import exposure for eight other high-income countries excluding the United States (including Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland)
Leader	An indicator variable set to one for firms with above-median *Tobin's q*, sales [SALE], or total assets [AT] for each industry (SIC two-digit) as of 1999, two years prior to China's entry to the WTO, and zero otherwise
Post	An indicator for the time period after 2001
$Exposure_{tech,i}$	We first measure the extent to which industries are subject to technological change using an industry-level share of routine-task labor ($Exposure_{tech}$). To construct this variable, we follow the procedure employed in Zhang (2019). We first define the routine-task intensity (RTI) score for each occupation as $RTI_k = ln(T_k^{routine}) - ln(T_k^{abstract}) - ln(T_k^{nonroutine\ manual})$ where $T_k^{routine}$, $T_k^{abstract}$, and $T_k^{nonroutine\ manual}$ are the routine, abstract, and nonroutine manual task skill levels (scaled from 1 to 10) required by occupation k. Each occupation's required skill level data are obtained from the revised fourth edition of the *Dictionary of Occupational Titles* by the U.S. Department of Labor. We classify workers as routine-task labor if their occupations' RTI score falls in the top quintile of the RTI distribution. We then construct $Exposure_{tech}$ as the proportion of routine-task labor costs to the total industry labor cost. We obtain data on the number of employees and their wages for each occupation-industry from the *Occupational Employment Statistics* by the Bureau of Labor Statistics. Using this industry-level measure of exposure to technological change ($Exposure_{tech}$), we define $Exposure_{tech,i}$ as firm i's exposure to technological change, as of 1999 in two ways: (i) As a continuous variable, we set $Exposure_{tech,i}$ to be equal to $Exposure_{tech}$ for the three-digit SIC industry to which firm i belongs; (ii) As an indicator variable, we set $Exposure_{tech,i} = 1$ for firms if their continuous $Exposure_{tech,i}$ measure is above the median value of the $Exposure_{tech,i}$ distribution, and zero otherwise.
Large Decline	For each firm, we first calculate the annualized change in union coverage of the state in which a firm is headquartered (or the annualized change in the union coverage of the industry in which a firm operates) between the first and the last year, when each firm appears in the panel data. We then define *Large Decline* as an indicator variable set to one if a firm has a below-median annualized change in the union coverage (using the firm-level distribution of this variable), and zero otherwise. Union coverage is a percent of private-sector workers that are covered by a collective bargaining agreement. We use the annual state- and industry-level union coverage data from Hirsch and Macpherson (2003).
RTW	An indicator variable set to one if the state where a firm is headquartered has passed the right-to-work (RTW) legislation as of year t, and zero otherwise
$LargeDecline_{RTW}$	We first compute the three-year average of state-level union coverage before and after the RTW adoption year. The difference between these values measures the change in union coverage around the passage of RTW laws. We then define $LargeDecline_{RTW}$ as an indicator variable set to one if a state (where a firm is headquartered) experienced a below-median change around the adoption year and zero otherwise

Appendix B. Additional Tables and Figures

Table B.1: Measurement Error in *Tobin's q*: High-Order Cumulant Equations

This table presents the results of regressing corporate investment on minimum wage using the linear high-order cumulant equations (Erickson et al., 2014) to address measurement error in *Tobin's q*. The dependent variables in all columns are *Investment*, measured as capital expenditures normalized by the beginning-of-the-year capital stock (property, plant, and equipment). We measure *Cash Flow* as earnings before extraordinary items plus depreciation normalized by the beginning-of-the-year capital stock and *Tobin's q* as a ratio of market value of assets to book value of assets. Columns (1–2) reports the fixed effect OLS regression results in columns (2–3) of Panel A, Table II in the main text. In columns (1–2), standard errors in parentheses are robust to heteroskedasticity and clustered by state. In columns (3–4), bootstrapped standard errors that are robust to within state correlation are reported in parentheses. ρ^2 is an estimate of the R^2 of the regression and τ_Q^2 is an index of measurement quality, which ranges from 0 to 1, for the proxy variable with standard errors in parentheses. We set the highest order of cumulants to be five.

	Dependent Variable: $Corporate\ Investment_{i,s,t}$			
	OLS-FE		EJW High-order Cumulant Estimator	
	Pre-2000	Post-2000	Pre-2000	Post-2000
	(1)	(2)	(3)	(4)
$w_{i,s,t-1}$	-.038***	.001	-.041**	.005
	(.012)	(.003)	(.018)	(.005)
$Cash\ Flow$	.098***	.029***	.026**	.003
	(.005)	(.002)	(.011)	(.004)
$Tobin's\ q$	.066***	.053***	.268***	.220***
	(.003)	(.003)	(.021)	(.040)
$GDP\ Growth$	.003**	.002**	-.0003	.001
	(.001)	(.0008)	(.002)	(.0009)
$ln(Population)$	-.145*	-.192***	-.029	-.177
	(.076)	(.049)	(.096)	(.115)
$Unemployment$	.001	.002	-.001	.009
	(.001)	(.002)	(.003)	(.006)
Firm and Year FE	Yes	Yes	Yes	Yes
# of Firm-Year Obs.	31,408	27,688	31,408	27,688
Adjusted R^2	0.135	0.122		
ρ^2			0.257	0.268
τ_Q^2			0.317***	0.322***
			(0.023)	(0.023)

Table B.2: Investment Sensitivity to Minimum Wage: Strict Exogeneity Assumption Tests

In this table, we conduct strict exogeneity assumption tests for all estimations in Panel A of Table II in the main text. As suggested by Wooldridge (2011) and Grieser and Hadlock (2019), we include the one period lead value of the key variable of interest (i.e., $w_{i,s,t}$) in the investment regressions. In columns (1–3) of this table, we report the baseline fixed effect OLS estimates from Panel A, Table II of the main text for easier comparison. Columns (4–6) present the estimated coefficients on $w_{i,s,t}$. The dependent variable is *Investment*, measured as capital expenditures normalized by the beginning-of-the-year capital stock (property, plant, and equipment). $w_{i,s,t-1}$ is the minimum wage at time $t-1$ in state s where firm i's headquarters is located. For 1983–2014, we obtain the historical changes in minimum wages for non-farm private sector employment under state and federal laws from the Tax Policy Center. These data are sourced from the Wage and Hour Division of the U.S. Department of Labor and from the *Monthly Labor Review* by the Bureau of Labor Statistics. For 2015–2017, we hand-collect the data from the U.S. Department of Labor. Under Section 18 of the Fair Labor Standard Act, when an employee is subject to both the federal and state minimum wage laws, the employee is entitled to the higher of the two standards. We measure *Cash Flow* as earnings before extraordinary items plus depreciation normalized by the beginning-of-the-year capital stock and *Tobin's q* as a ratio of market value of assets to book value of assets. We also control for state-level macro-variables. Standard errors in parentheses are robust to heteroskedasticity and clustered by state. The full sample period runs from 1984 to 2017. Note that the difference in number of observations between columns (3) and (6) (consequently columns (1) and (4)) arises from some firms in the fiscal year 2017 of which fiscal year ended 2018.

	Dependent Variable: $Corporate\ Investment_{i,s,t}$					
	Minimum Wage and Corporate Investment			Strict Exogeneity Assumption Tests		
	Full Sample	1984 to 2000	2001 to 2017	Full Sample	1984 to 2000	2001 to 2017
	(1)	(2)	(3)	(4)	(5)	(6)
$w_{i,s,t-1}$	-.017*	-.038***	.001	-.014*	-.038***	.004
	(.009)	(.012)	(.003)	(.007)	(.009)	(.003)
$w_{i,s,t}$				-.004	-.001	-.003
				(.006)	(.010)	(.005)
$Cash\ Flow$	.043***	.098***	.029***	.043***	.098***	.029***
	(.002)	(.005)	(.002)	(.002)	(.005)	(.002)
$Tobin's\ q$	.063***	.066***	.053***	.063***	.066***	.053***
	(.003)	(.003)	(.003)	(.003)	(.003)	(.003)
$GDP\ growth$	.002**	.003**	.002**	.002**	.003**	.002**
	(.001)	(.001)	(.001)	(.001)	(.001)	(.001)
$ln(Population)$	-.108	-.145*	-.192***	-.109	-.145*	-.195***
	(.065)	(.076)	(.049)	(.065)	(.075)	(.050)
$Unemployment$	.00002	.001	.002	-.00003	.001	.002
	(.001)	(.001)	(.002)	(.001)	(.001)	(.002)
Firm and Year FEs	Yes	Yes	Yes	Yes	Yes	Yes
# of Firm-Year Obs.	59,096	31,408	27,688	59,043	31,408	27,635
Adjusted R^2	.140	.135	.122	.140	.135	.122

Table B.3: Robustness Results on 1999 US-China Bilateral Agreement

Panel A repeats Table III in the main text by using 1998 (instead of 1997) as the year to check operational status in China. Column (1) presents a difference-in-differences regression of investment on minimum wages interacted with two indicators, $China98$ and $Agreement$. $w_{i,s,t-1}$ is the minimum wage at time $t-1$ in state s where firm i's headquarters is located. $China98$ is an indicator variable set to one if a firm has at least one subsidiary in China one year prior to the US-China bilateral agreement in 1999, and zero otherwise; $Agreement$ indicates the time period after the agreement (including 1999). The coefficient on the triple interaction term ($China98 \times Agreement \times w_{i,s,t-1}$) captures difference-in-differences in investment-wage sensitivity after the agreement between treated ($China98$ = 1) and control ($China98$ = 0) firms. In column (2), we introduce another group by defining $China04$ as an indicator variable set to one for firms without any subsidiary in China as of 1998 but having at least one subsidiary as of 2004 (five years after the agreement), and zero otherwise. The omitted group consists of firms that have no operations in China, that is, $China98 = China04 = 0$. In all columns, we include the same set of control variables used in Table III of the main text. $China98$ (or $China04$) and $Agreement$ indicators are absorbed by firm and year fixed effects, respectively. The definitions of all variables are provided in Appendix A. Standard errors in parentheses are robust to heteroskedasticity and clustered by state. The sample period runs from 1984 to 2017.

Panel A. Identifying Treated Firms as of 1998

Dependent Variable:	$Investment_{i,s,t}$	
	Two Groups	Three Groups
	(1)	(2)
$w_{i,s,t-1}$	-.018**	-.017**
	(.008)	(.008)
$Agreement \times w_{i,s,t-1}$	.006	.003
	(.006)	(.006)
$China98 \times Agreement \times w_{i,s,t-1}$	.034***	.036***
	(.007)	(.007)
$China04 \times Agreement \times w_{i,s,t-1}$		.033*
		(.016)
$China98 \times w_{i,s,t-1}$	-.019***	-.020***
	(.006)	(.006)
$China04 \times w_{i,s,t-1}$		-.019
		(.015)
$China98 \times Agreement$	0.172	0.203
	(0.165)	(0.172)
$China04 \times Agreement$		0.318
		(0.227)
Investment Sensitivity to Minimum Wage [t-stat]		
Before (baseline group)	-0.018** [-2.20]	-0.017** [-2.05]
After (baseline group)	-0.012* [-2.01]	-0.014** [-2.40]
Before ($China04$ = 1)		-0.036** [-2.19]
After ($China04$ = 1)		-0.000 [-0.04]
Before ($China98$ = 1)	-0.037*** [-3.98]	-0.037*** [-3.92]
After ($China98$ = 1)	0.003 [0.75]	0.002 [0.62]
Controls / Interaction of Controls / Firm and Year FEs	Yes	Yes
# of Firm-Year Obs.	59,096	59,096
Adjusted R^2	.157	.158

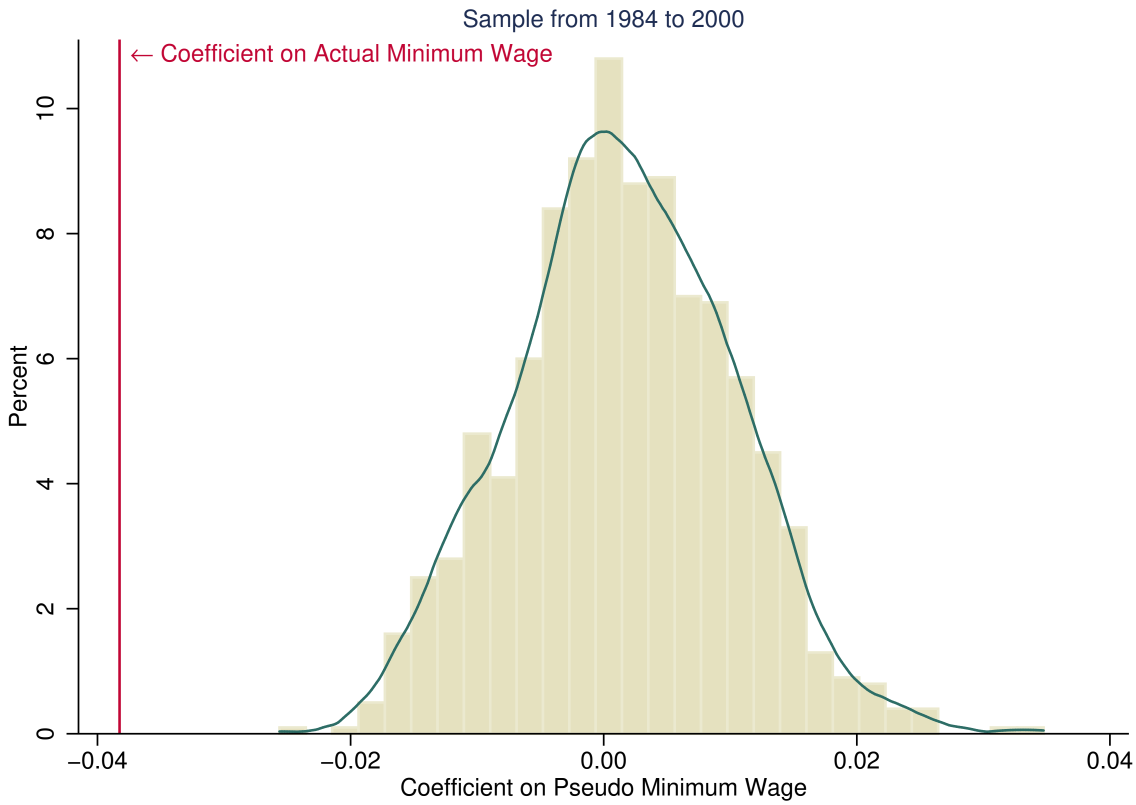
Table B.3: Robustness Results on 1999 US-China Bilateral Agreement (continued)

Panel B repeats Table III in the main text by constructing a time-varying indicator, $China_{i,t}$. We define $China_{i,t}$ as an indicator variable set to one if firm i has at least one subsidiary in China in year t, and zero otherwise. $w_{i,s,t-1}$ is the minimum wage at time $t-1$ in state s where firm i's headquarters is located; $Agreement$ indicates the time period after the agreement (including 1999). We use hand collected information from 10-k filings to identify US firms' Chinese subsidiaries in every year. Since the year 1997 is the first year of comprehensive reporting of subsidiary information, we use information as of 1997 for all years prior to 1997. The coefficient on the triple interaction term ($China \times Agreement \times w_{i,s,t-1}$) captures difference-in-differences in investment-wage sensitivity after the agreement between treated ($China$ = 1) and control ($China$ = 0) samples. We include the same set of control variables used in Table III in the main text. $Agreement$ indicator is absorbed by year fixed effects. The definitions of all variables are provided in Appendix A. Standard errors in parentheses are robust to heteroskedasticity and clustered by state. The sample period runs from 1984 to 2017.

Panel B. Fully Allowing for Entry into China after the Agreement

Dependent Variable:	$Investment_{i,s,t}$
	(1)
$w_{i,s,t-1}$	-.018**
	(.008)
$Agreement \times w_{i,s,t-1}$	.006
	(.006)
$China \times Agreement \times w_{i,s,t-1}$	.026***
	(.009)
$China$	0.196
	(0.160)
$China \times w_{i,s,t-1}$	-.020***
	(.007)
$China \times Agreement$	0.064
	(0.222)
Investment Sensitivity to Minimum Wage [t-stat]	
Before (baseline group)	-0.018** [-2.19]
After (baseline group)	-0.012* [-2.02]
Before ($China$ = 1)	-0.038*** [-3.61]
After ($China$ = 1)	-0.006 [-1.02]
Controls / Interaction of Controls	Yes
Firm and Year FEs	Yes
# of Firm-Year Obs.	59,096
Adjusted R^2	.158

Figure B.1: Investment Sensitivity to Minimum Wage: Placebo Test

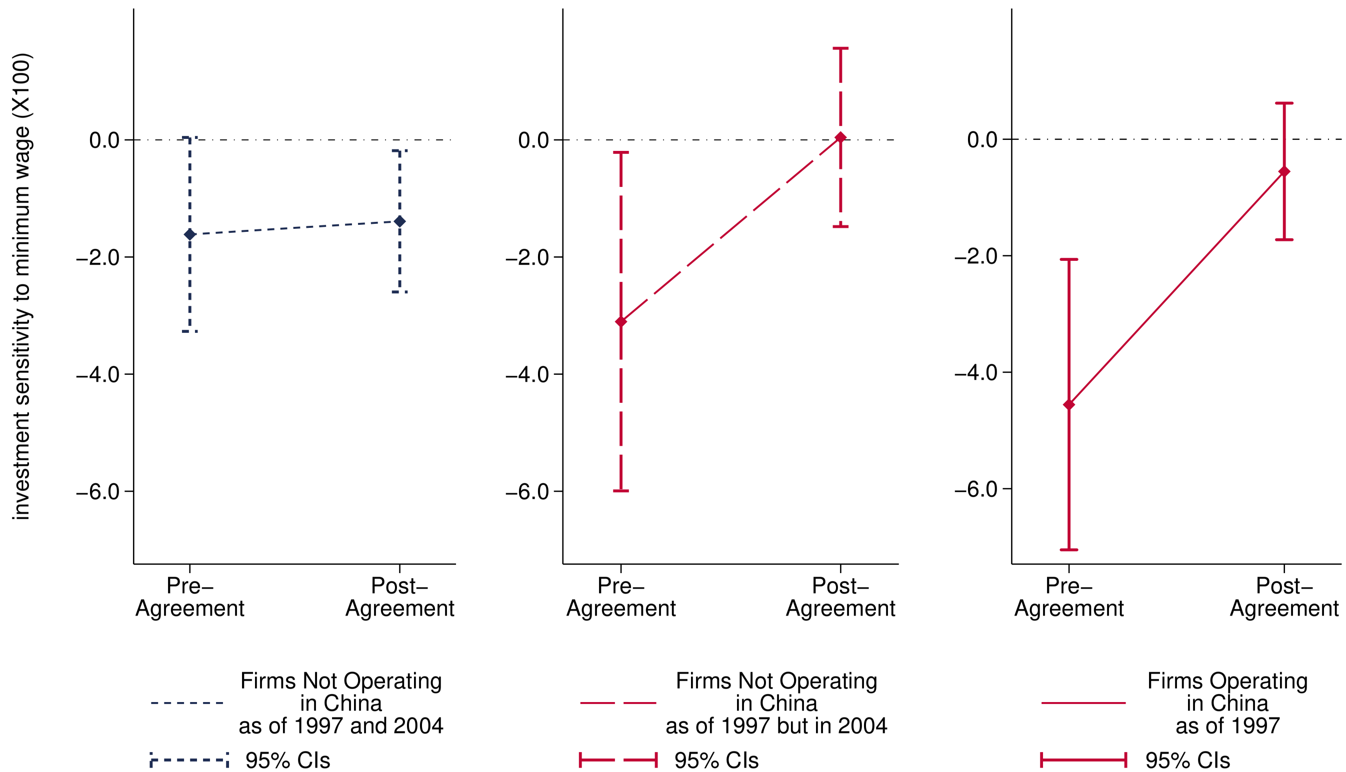


This figure is based on the following investment regressions:

$$\frac{I_{i,s,t}}{K_{i,s,t-1}} = \alpha_i + \alpha_t + \beta_1 \, Tobin's \; q_{i,s,t-1} + \beta_2 \, \frac{CF_{i,s,t}}{K_{i,s,t-1}} + \beta_3^{Pseudo} \, w_{i,s,t-1}^{Pseudo} + \beta_4 \, X_{i,s,t-1}^{Pseudo} + \epsilon_{i,s,t},$$

where i, s, and t index firms, states, and years; α_i and α_t is a set of firm and year fixed effects, respectively; $Investment \; (= \frac{I_{i,s,t}}{K_{i,s,t-1}})$ is investment rates; $Cash \;\; Flow = \frac{CF_{i,s,t}}{K_{i,s,t-1}}$ refers to cash flow; $Tobin's \; q_{i,s,t-1}$ is a proxy for investment opportunities; $X_{i,s,t-1}^{Pseudo}$ is a set of state-level macro-variables that are based on the pseudo state: real GDP growth rates, log of population, and unemployment rates. To construct a pseudo minimum wage variable ($w_{i,s,t-1}^{Pseudo}$), we randomly assign each firm i to a particular state s. With this pseudo state, we define $w_{i,s,t-1}^{Pseudo}$ as the minimum wage at time $t-1$ in state s where firm i's *hypothetical* headquarters is located. Once all firms in the sample are assigned in this manner, we estimate the investment regression and store the coefficient on $w_{i,s,t-1}^{Pseudo}$. This procedure is repeated 1,000 times and a distribution of $w_{i,s,t-1}^{Pseudo}$ coefficients is generated. The figure plots the empirical distribution of the coefficient on $w_{i,s,t-1}^{Pseudo}$. The green line is the estimated non parametric kernel density. The red vertical line indicates the investment-wage sensitivity obtained from the actual data (column (2) of Panel A in Table II). The sample period runs from 1984 to 2000. Standard errors are clustered by state.

Figure B.2: 1999 US-China Bilateral Agreement: Investment Sensitivity to Minimum Wage



These figures plot the changes in investment-wage sensitivity for three groups of firms before and after the 1999 US-China bilateral agreement. The first group (in the left figure) consists of firms that do not operate in China as of 1997, two years prior to the agreement, and remain not operating in China as of 2004, five years after the agreement. The second group (in the middle figure) indicates firms not operating in China as of 1997 but operating in 2004. The last group (in the right figure) consists of firms operating in China as of 1997. We identify firms operating in China if firms have at least one subsidiary in China, using hand collected information from 10-K filings on US firms' Chinese subsidiaries. The dots indicate the estimated investment-wage sensitivity and the vertical lines around these point estimates are the 95% confidence intervals. These estimates are based on column (2) of Table III. The detailed definitions of all variables are provided in Appendix A.

Appendix C. Cost-Benefit Analysis of Minimum Wage Increase to Workforce: Counterfactual Analysis

What are the costs and benefits of raising the minimum wage? Especially during the presidential election years, this question draws substantial interest among policymakers, the press, and the public.[20] The growing interest in recent years does not only reflect the 2016 and 2020 presidential elections, but also indicates a heated debate over the recent legislative movement towards a $15 an hour minimum wage in large cities, for example, New York, Los Angeles, San Francisco, and Seattle. Many legislators and some expert economists such as the Economic Policy Institute mainly focus on the potential positive impact of minimum wage increases on alleviating income inequality, the ability of below poverty line workers to meet their basic needs, or the unemployment of low-skill workers. Experts from the Economic Policy Institute (2021) also claim that underpaid workers will spend much of their extra earnings and this injection of wages will help stimulate the economy and spur greater business activity and job growth.[21] However, our findings in Panel A of Table II point out an important overlooked aspect of the minimum wage effect on the workforce through the investment cuts made by the firm. These investment cuts would lead to less new labor hired by the firm and lead to job losses among the workforce compared to a scenario where the minimum wage cuts were absent and hence the firm made investments and hired new labor from the workforce.[22]

To gauge the overall economic effects of a minimum wage increase taking into account job losses due to investment cuts, we conduct a simple counterfactual exercise for an average firm. We compare the actual scenario where the firm faces minimum wage increases and responds by investment cuts with a hypothetical counterfactual. The counterfactual benchmark is a situation where a firm does not experience a minimum wage increase and hence would continue to make investments. As a caveat, we note that this cost-benefit analysis is a rough back-of-the-envelope calculation which is simplistic in that it does not take into account general-equilibrium effects of the minimum wage increase on factor or output prices.

[Insert Table C.1 here.]

In Table C.1 Panel A1, we calculate the benefit of a minimum wage increase to the *existing* workforce. We assume conservatively that all hourly-paid workers would fully benefit from a

[20] CNBC News reported on Nov. 18, 2015, for example; "One of the clearest distinctions to come out of the presidential debates so far has been around the minimum wage ... Democratic candidates' support for, and the Republican candidates' opposition to, raising the federal minimum wage."

[21] The report, entitled 'Why the U.S. needs a $15 minimum wage', issued by the Economic Policy Institute on Jan. 26, 2021. `https://www.epi.org/publication/why-america-needs-a-15-minimum-wage/`

[22] "New Jersey Governor Chris Christie vetoed a bill backed by Democratic lawmakers that would have increased the state's minimum hourly wage to $15 by 2012. ... The proposed increase, he said, 'would trigger an escalation of wages that will make doing business in New Jersey unaffordable.'" (N.J.'s Christie Vetoes Minimum-Wage Bill, *Wall Street Journal*, Aug. 30, 2016)

minimum wage increase. For an average firm, we estimate the additional wages that the workforce earns from a minimum wage increase to be about \$2.90 million. This benefit is calculated as $\Delta w_{min} \cdot h \cdot L \cdot \rho_{hour}$ where Δw_{min} is the average annual change in minimum wage rates in our sample, h is the average annual hours actually worked per U.S. worker (obtained from *OECD Statistics* as of 2017), L is the average number of employees per firm (based on our sample as of the year 2017), and ρ_{hour} is the percent of hourly-paid workers out of total workers (obtained from the *Labor Force Statistics from the Current Population Survey* as of 2017).

In Table C.1 Panel A2, we calculate the cost of a minimum wage increase to the *future* workforce for the pre-2000 and post-2000 periods. We assume that the employment adjustment due to a change in the capital is based on the average number of workers per unit capital stock. This cost is calculated as $\beta_3 \cdot \Delta w_{min} \cdot K \cdot (L/K) \cdot w$ where β_3 is the estimated investment-wage sensitivity in columns (2) and (3) of Panel A, Table II, K is the average lagged capital stock in million \$ (based on our sample as of the year 2017), L/K is the average number of workers per million \$ capital stock (based on our sample as of the year 2017), and w is the average annual income per U.S. worker (obtained from *OECD Statistics* as of 2017). In the earlier period (pre-2000) in which a minimum wage increase has a strong negative impact on investment, the opportunity cost of job losses resulting from the investment cut amounts to \$34.41 million. Since β_3 in the calculation is estimated with error, the 95% confidence interval of this point estimate is (–\$56.87, –\$11.86) million. Thus, for the entire workforce, it appears that the cost is much larger than the benefit for the pre-2000 period: the net cost to the workforce of a minimum wage increase at the average firm is \$31.51 million (the 95% confidence interval of (–\$53.97, –\$8.96) million). For the post-2000 period, this cost on average is negligible because there is no negative impact on investment for the average firm (the point estimate of β_3 is close to zero): the cost estimate is \$0.90 million with the 95% confidence interval of (–\$4.22, \$5.93) million. For the entire workforce, the net effect of a minimum wage increase is \$2.00 million which is a noisy estimate as the 95% confidence interval of this estimate includes zero (the 95% confidence interval of (–\$1.32, \$8.83) million).

In Panel B of Table C.1 we summarize the results of the above counterfactual analysis by plotting the benefit, cost, and net cost (benefit) (along with the 95% confidence interval of the estimates) of a minimum wage increase to the total workforce based on calculations in Panels A1 and A2 in Table C.1. Figure (a) in Panel B is based on the estimated investment-wage sensitivity for the pre-2000 period whereas figure (b) in Panel B is based on the estimated investment-wage sensitivity for the post-2000 period. These figures emphasize an overlooked but important negative effect of minimum wages on total employment through forgone corporate investment.

As a robustness check, we repeat our counterfactual analysis using different but reasonable parameter values (changes in the minimum wage, annual hours worked, the average number of employees, percent of hourly-paid workers, average capital stock, average labor to capital ratio, and average annual income per U.S workers) for the pre-2000 period. We obtain qualitatively similar results. The details of these calculations are reported in Table C.2.

[Insert Table C.2 here.]

The results in Table C.1 Panel A2 suggest that the net benefit/cost for the average firm due to a minimum wage increase in the post-2000 period is statistically indistinguishable from zero. However, as our results in Tables III to VI show, this average result masks important heterogeneity among firms. Firms that responded to the various economic shocks (and thus have no investment-wage sensitivity) do not impose any costs on the workforce due to minimum wage increases. They have moved their operations offshore; replaced labor with automation and had weak unions to negotiate with. However, our analysis also identified a sizeable group of firms that did not respond to these economic shocks effectively. For the latter group of firms, the investment-wage sensitivity was still significantly negative. In Panel C of Table C.1 we summarize the counterfactual analysis for these groups of firms by plotting the benefit, cost, and net cost (benefit) (along with the 95% confidence interval of the estimates) of a minimum wage increase to the total workforce based on calculations similar to Panels A1 and A2 in Table C.1 for each of the economic shocks analyzed in Tables III to VI. We conclude from these results that there exist a sizeable group of US firms as of today that produces significant negative effects of minimum wages on total employment through forgone corporate investment.

In Figure C.1 we graphically illustrate this additional source of employment reduction through the investment cut triggered by a minimum wage increase. $Demand_1$ represents the labor demand curve in the absence of the minimum wage increase policy. Equilibrium occurs when supply equals demand, which generates the competitive employment L^* and wage w^*. Once the government imposes a minimum wage (w_{min}), which is greater than w^*, firms demand less labor due to the increased cost of labor. L_{d1} will be the new level of employment that is lower than L^*. Our findings suggest that this might not be the whole story of the effect of a minimum wage increase. The investment cut resulting from the minimum wage increase will shift the demand curve to the left ($Demand_2$), which amplifies the employment reduction on top of imposing the minimum wage itself. L_{d2} will be the new level of employment that is lower than L_{d1}. Thus, $L_{d1} - L_{d2}$ is the additional unemployment due to the investment cut.

[Insert Figure C.1 here.]

In conclusion, our simple counterfactual exercise suggests that the proponents of minimum wage laws must consider the unintended negative effect of minimum wages on the workforce through corporate investment. In practice, increased minimum wages will have a number of other potential benefits over and above the increase in wages itself, such as reduced income inequality or satisfying the basic needs of low-skill workers, which are usually difficult to measure. These benefits also need to be considered in the cost-benefit analysis to obtain a more complete and accurate picture of welfare implications. Nevertheless, this article provides suggestive evidence that a minimum wage increase could ultimately dampen employment growth by stifling corporate investment.

Table C.1: Cost-Benefit Analysis of Minimum Wage Increase to Workforce: Counterfactual Analysis

These tables present a cost-benefit analysis of a minimum wage increase to the workforce using a simple counterfactual analysis for an average firm. We compare the actual scenario where the firm faces minimum wage increases and responds by cutting investment with a hypothetical counterfactual. The counterfactual benchmark is a situation where a firm does not experience a minimum wage increase and hence would continue to make investments. We note that this cost-benefit analysis is a rough back-of-the-envelope calculation that does not take into account the general-equilibrium effects of a minimum wage increase on factor or output prices. Panel A1 shows the benefit of a minimum wage increase to the existing workforce. Detailed information and data source of each parameter value is shown in *Description* column. We assume conservatively that all hourly-paid workers fully benefit from a minimum wage increase. Panel A2 calculates the opportunity cost (or benefit) of a minimum wage increase to the future workforce for the pre-2000 and post-2000 periods. We assume that the employment adjustment due to a change in the capital is based on the average number of workers per unit capital stock.

Panel A1. Benefit of Minimum Wage Increase to Workforce (Pre- and Post-2000)

Value	Parameter / Calculation	Description
\$0.15	Δw_{min}	average annual changes in minimum wage rates (full sample)
1,780	h	average annual hours actually worked per U.S. worker (*OECD.Stat* as of 2017)
18,614	L	average number of employees per firm (sample as of 2017)
58.3%	ρ_{hour}	percent of hourly-paid workers out of total workers (*Labor Force Statistics from the Current Population Survey* as of 2017)
M\$2.90	$\Delta w_{min} \cdot h \cdot L \cdot \rho_{hour}$	additional wages that workforce earns from minimum wage increase

Panel A2. Cost (Benefit) of Minimum Wage Increase Due to Investment Cuts (Increases) (Pre- vs. Post-2000)

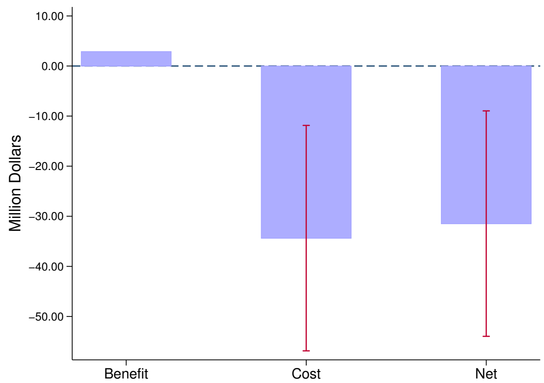
Value [95% Confidence Interval] Pre-2000	Value [95% Confidence Interval] Post-2000	Parameter / Calculation	Description
-3.83% [-6.33, -1.32]%	0.01% [-0.47, 0.66]%	β_3	investment-wage sensitivity (columns (2) and (3) of Panel A, Table II)
M\$3,409.21		K	average lagged capital stock in million \$ (sample as of 2017)
30.44		L/K	average # of workers per million \$ capital stock (sample as of 2017)
\$57,715		w	average annual income (\$) per U.S. worker (*OECD.Stat* as of 2017)
M\$-34.41 M\$[-56.87, -11.86]	M\$0.90 M\$[-4.22, 5.93]	$\beta_3 \cdot \Delta w_{min} \cdot K \cdot (L/K) \cdot w$	opportunity cost of job losses through investment cuts (or benefit of potential job openings via increased investment)

Table C.1: Cost-Benefit Analysis of Minimum Wage Increase to Workforce: Counterfactual Analysis (continued)

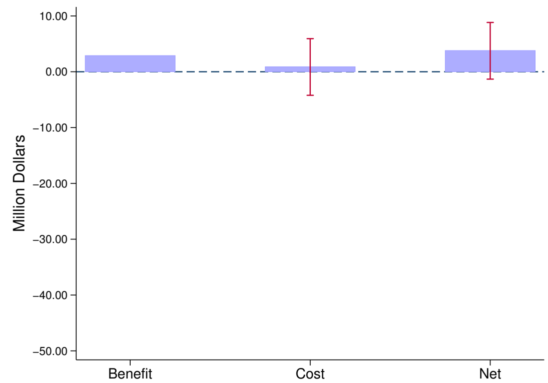
Panel B summarizes the benefit, cost, and net cost (benefit) of a minimum wage increase to the workforce based on the calculations in Panels A1 and A2. In Panel C, we repeat the same exercise to calculate the benefit, cost, and net cost (benefit) of a minimum wage increase to the workforce for the firms that are subject to a minimum wage shock based on Tables III to VI. The blue bars indicate each amount in million dollars and the red vertical lines depict the 95% confidence intervals.

Panel B. Net Cost (Benefit) of Minimum Wage Increase to Workforce: Pre-2000 and Post-2000

(a) Pre-2000

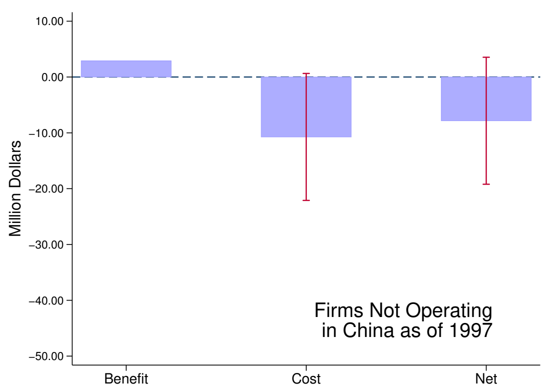


(b) Post-2000

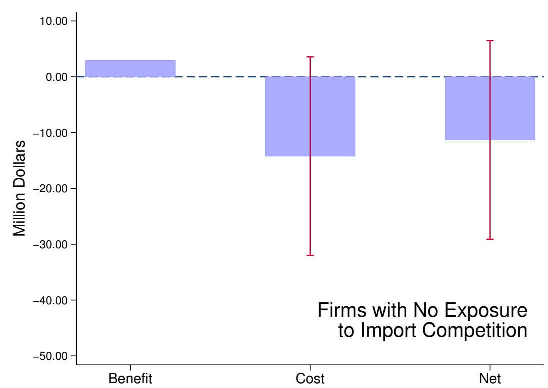


Panel C. Net Cost (Benefit) of Minimum Wage Increase to Workforce: For Firms that are Subject to a Minimum Wage Shock

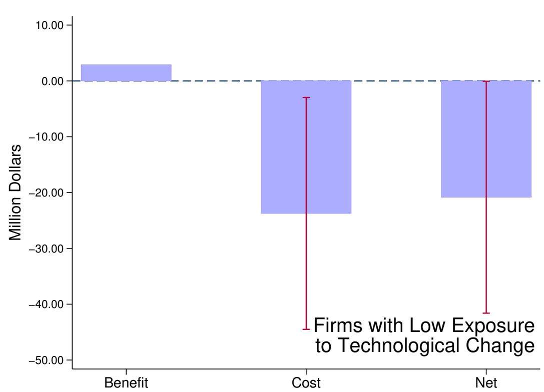
(a) 1999 US-China Bilateral Agreement



(b) Chinese Import Competition



(c) Technological Change / Automation



(d) Declining Union Power

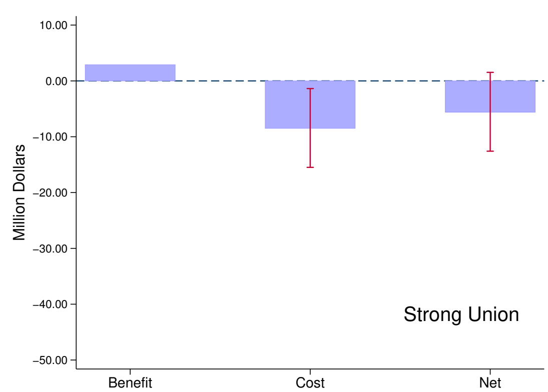


Table C.2: Cost-Benefit Analysis of Minimum Wage Increase to Workforce: Counterfactual Analysis Using Alternative Parameter Values

These tables repeat Panels A1 and A2, Table C.1 in the main text by using alternative parameter values that are used to conduct a cost-benefit analysis of a minimum wage increase to the workforce for an average firm. We compare the actual scenario where the firm faces minimum wage increases and responds by cutting investment with a hypothetical counterfactual. The counterfactual benchmark is a situation where a firm does not experience a minimum wage increase and hence would continue to make investments. We note that this cost-benefit analysis is a rough back-of-the-envelope calculation that does not take into account the general-equilibrium effects of a minimum wage increase on factor or output prices. Panel A1 shows the benefit of a minimum wage increase to the existing workforce. Detailed information and data source of each parameter value is shown in *Description* column. We assume conservatively that all hourly-paid workers fully benefit from a minimum wage increase. Panel A2 calculates the opportunity cost (or benefit) of a minimum wage increase to the future workforce for the pre-2000 and post-2000 periods. We assume that the employment adjustment due to a change in the capital is based on the average number of workers per unit capital stock.

Panel A1. Benefit of Minimum Wage Increase to Workforce (Pre- vs. Post-2000)

Value		Parameter / Calculation	Description
Pre-2000	Post-2000		
$0.12	$0.17	Δw_{min}	average annual changes in minimum wage rates (pre-2000 and post-2000 sample)
1,832	1,780	h	average annual hours actually worked per U.S. worker (*OECD.Stat* as of 2000 and 2017)
10,145	18,614	L	average number of employees per firm (sample as of 2000 and 2017)
59.6%	58.3%	ρ_{hour}	percent of hourly-paid workers out of total workers (*Labor Force Statistics from the Current Population Survey* as of 2002 and 2017)
M$1.33	M$3.28	$\Delta w_{min} \cdot h \cdot L \cdot \rho_{hour}$	additional wages that workforce earns from minimum wage increase

Panel A2. Cost (Benefit) of Minimum Wage Increase Due to Investment Cuts (Increases) (Pre- vs. Post-2000)

Value [95% Confidence Interval]		Parameter / Calculation	Description
Pre-2000	Post-2000		
-3.83% [-6.33, -1.32]%	0.01% [-0.47, 0.66]%	β_3	effect of unit minimum wage increase on investment rate (Columns (2) and (3) of Panel A, Table II)
M$1,208.43	M$3,409.21	K	average lagged capital stock in million $ (sample as of 2000 and 2017)
28.18	30.44	L/K	average number of workers per million $ capital stock (sample as of 2000 and 2017)
$52,725	$57,715	w	average annual income ($) per U.S. worker (*OECD.Stat* as of 2000 and 2017)
M$-8.25 M$[-13.64, -2.84]	M$1.02 M$[-4.79, 6.72]	$\beta_3 \cdot \Delta w_{min} \cdot K \cdot (L/K) \cdot w$	opportunity cost of job losses through investment cuts (or benefit of potential job openings via increased investment)

Table C.2: Cost-Benefit Analysis of Minimum Wage Increase to Workforce: Counterfactual Analysis Using Alternative Parameter Values (continued)

Panel B summarizes the benefit, cost, and net benefit of the minimum wage increase to the workforce based on the calculations in Panels A1 and A2. Figure (a) plots the benefit, cost, and net benefit amounts that are based on the estimation from the pre-2000 sample period. Figure (b) plots the benefit, cost, and net benefit amounts that are based on the estimation from the post-2000 period. The blue bars indicate each amount in million dollars and the red vertical lines depict the 95% confidence intervals.

Panel B. Net Cost (Benefit) of Minimum Wage Increase to Workforce: Using Different Parameter Values

(a) Pre-2000 (b) Post-2000

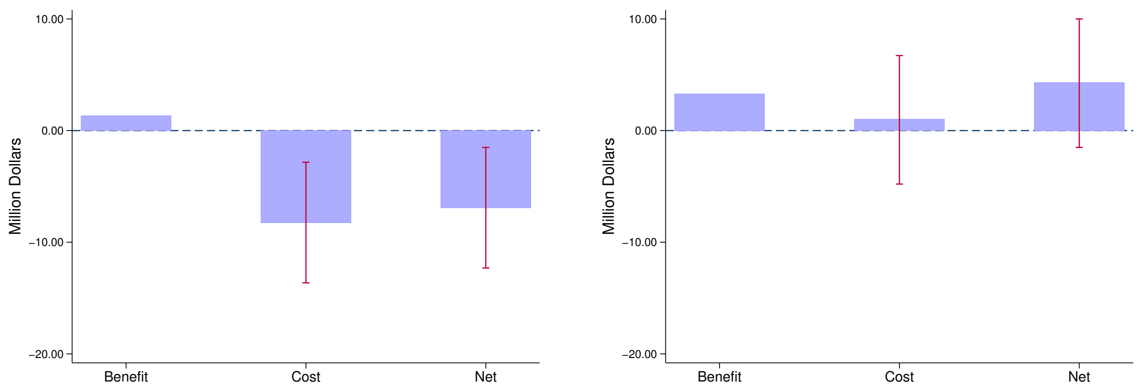
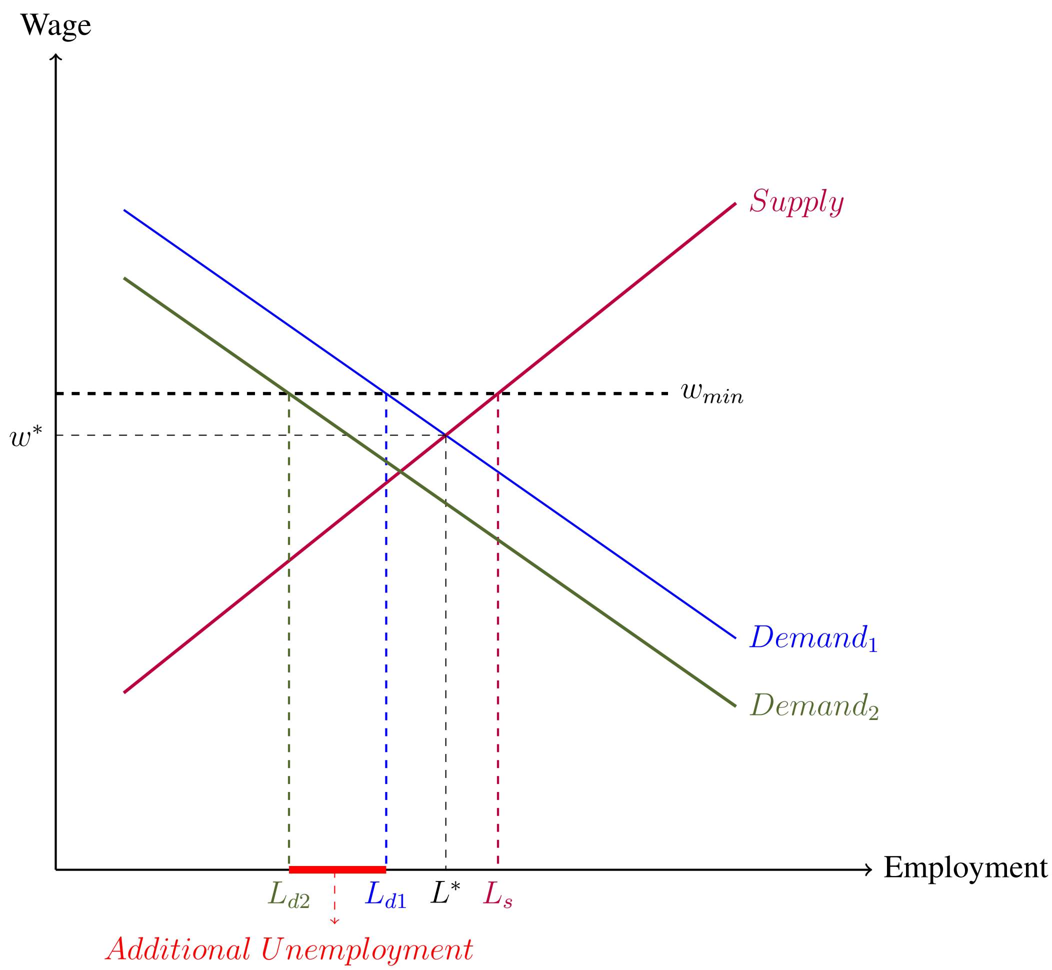


Figure C.1: Investment Cut and Unemployment



This figure illustrates an additional source of employment reduction through forgone corporate investment triggered by a minimum wage increase. $Demand_1$ represents the labor demand curve in the absence of the minimum wage increase policy. Equilibrium occurs when supply equals demand, which generates the competitive employment L^* and wage w^*. Once the government imposes a minimum wage (w_{min}), which is greater than w^*, firms demand less labor due to the increased cost of labor. L_{d1} will be the new level of employment that is lower than L^*. Our findings suggest that the investment cut resulting from the minimum wage increase will shift the demand curve to the left ($Demand_2$), which amplifies the employment reduction on top of imposing the minimum wage itself. L_{d2} will be the new level of employment that is lower than L_{d1}. Hence, $L_{d1} - L_{d2}$ is the additional unemployment due to the investment cut. As a caveat, we note that this illustration is simplistic in that it does not take into account the general-equilibrium effects of a minimum wage increase on factor or output prices.