DECLINING WORKER POWER
AND CORPORATE INVESTMENT*

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ABSTRACT

Declining worker bargaining power has been advanced as an explanation for major macroeconomic shifts in the U.S., such as declining labor share, reduced consumer purchasing power, and rising inequality. This paper explores microeconomic implications, focusing on the impact of declining worker power on firm-level investment responses to a labor cost shock arising from mandated minimum wage increases. Over four decades, we observe a shift in investment-wage sensitivities from negative to insignificant as management gains flexibility through enhanced outside options. This shift is more pronounced in firms that are more exposed to globalization, technological change, and declining unionization, reflecting broader trends in weakening worker power and its influence on corporate decision-making.

KEYWORDS: DECLINING WORKER POWER, CORPORATE INVESTMENT, LABOR SUBSTITUTION, GLOBALIZATION, U.S.-CHINA, AUTOMATION

JEL CODES: E2, J38, G31, F61

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. The literature proposes several non-mutually exclusive explanations for the decline, including 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 (Goldschmidt and Schmieder, 2017), and capital-biased technological change or automation (Acemoglu and Restrepo, 2018). Collectively, these factors indicate a decrease in workers' bargaining power vis-a-vis their employers, driven by the improved outside options available to employers (Stansbury and Summers, 2020).[1] 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.[2] In this paper, we consider microeconomic implications by examining the extent to which declining worker power has affected firm-level investment decisions—previously understudied labor market friction. Our empirical strategy focuses on the impact of worker power on firm investment responses to adverse labor cost shocks, specifically those arising from mandated changes in the minimum wage.[3] Intuitively, if capital and labor are complements, an increase in la-

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[1]Another prominent hypothesis emphasizes the emergence of "superstar firms" and heterogeneity across firms (Autor, Dorn, Katz, Patterson and Van Reenen, 2017, 2020).

[2]For exposition purposes, we use "worker power" and "worker bargaining power" interchangeably throughout the paper.

[3]This empirical setting has several features that are useful for our investigation. First, using changes in the minimum wage allows us to focus on low-skilled workers that arguably drive the decline in labor share. For example, Eisfeldt, Falato and Xiaolan (2022) find that the share of national income going to skilled labor has not declined over time. Second, mandated changes in the minimum wage provide a plausibly exogenous shock to wages that allow for clean identification and the ability to make causal statements regarding the effect of declining worker power on firm investment decisions. Finally, worker power is particularly "binding" when firms encounter negative labor cost shocks, enabling us to identify its impact on investment decisions.

bor costs typically results in a decrease in the demand for capital. As worker power declines, firms gain a better ability to substitute capital for labor (e.g., thanks to technology improvements) and/or replace U.S. workers with cheaper foreign labor (e.g., due to globalization), thereby dampening the negative impact of labor cost shocks on capital investment. We illustrate the conceptual link between worker power and the elasticity of substitution using an endogenously derived production function under the mechanization framework (Zeira, 1998; Nakamura and Nakamura, 2008).

We begin our analysis by examining changes in the sensitivity of firm-level investment to minimum wage changes over the 1984 to 2017 time period. We estimate the sensitivity of investment to minimum wage changes (hereafter "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. The results are striking: 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. In sharp contrast, for the post-2000 period, the estimated investment-wage sensitivity is 0.001 and statistically insignificant at conventional levels.[4]

Our main set of tests of the declining worker bargaining power hypothesis focuses on examining the impact on investment-wage sensitivities of the various forces that have been advanced as driving the decline in worker bargaining 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. Our empirical strategy is to compare changes in investment-wage sensitivities for firms that are more versus less exposed to the underlying drivers of declining worker power. Our expectation is that firms that

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[4]Consistent with the framework and findings in Gustafson and Kotter (2022), we find more significant results here, and throughout the rest of our analyses, when focusing on industries that are more exposed to minimum wage labor.

are more exposed to the underlying factors will experience greater changes in investment-wage sensitivities.

To start, we investigate the effect of globalization. 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 earned 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. To identify firms that are more likely to benefit from greater access to cheap Chinese labor, we use information from 10-k filings and define U.S. 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. In sharp contrast, for 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.[5]

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.[6] The Chinese economic reforms in the 1980s and 1990s

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

[6]Córdoba, Isojärvi and Li (2023) propose the constant elasticity of substitution (CES) matching function that endogenizes the matching elasticity and bargaining weights. The CES model highlights the decline in bargaining power arising from a decrease in labor demand.

resulted in rapid productivity growth and a consequent surge in Chinese exports. Its export growth was reinforced by China's entry into 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.[7] 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 for 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 industries that are more highly exposed 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 investment by firms that are more exposed to Chinese import competition is significantly less sensitive to minimum wage shocks as Chinese import competition grows. In contrast, investment-wage sensitivities remain negative and significant for firms in industries that are less exposed to Chinese import competition. 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).[8]

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[7]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.

[8]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 operations. To test this hypothesis, we identify leader (laggard) firms for each industry as those with above-median (below-median) *Tobin's q*, following

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 bargaining 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 part of our sample period. Graetz and Michaels (2017) find that routine-intensive jobs are particularly susceptible to replacement by new technologies. Hence, we measure the extent to which industries are subject to labor-saving 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 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 response to mandated minimum wage increases. To address potential endogeneity concerns, we use the passage of right-to-work (RTW) laws as a plausibly exogenous shock to union bargaining power. In states with RTW laws, mandated union membership or payment of union dues is prohibited, which limits a union's access to resources, thereby weakening its 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 bargaining 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

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Gutiérrez and Philippon (2017). We find that with increasing Chinese import competition, leaders tend to invest more relative to laggards in response to minimum wage increases. Notably, laggards still reduce their investment significantly following minimum wage increases in the later period (2001–2017). These results are robust to using sales and total assets to identify industry leaders.

to increases in the minimum wage became statistically and economically insignificant in the later time period. Importantly, 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 enriching the literature on worker power, our study contributes to the literature on corporate investment. A stream of existing research has extensively examined how capital market frictions (e.g., Fazzari et al., 1988; Almeida and Campello, 2007; Hennessy and Whited, 2007; Chava and Roberts, 2008) or labor market frictions (e.g., Bai, Fairhurst and Serfling, 2020; Hau, Huang and Wang, 2020; Cho, 2021; Gustafson and Kotter, 2022; Jeffers, 2024) affect capital investment. Our study enhances the current understanding of these dynamics by investigating the role of worker power—a previously underexplored determinant —in shaping corporate investment decisions, particularly in response to labor cost shocks. Our findings suggest that declining worker power mitigates the negative impact of labor cost shocks on capital investment, highlighting the influence of worker power on real outcomes.

Our study also contributes to the literature on the effects of mandated minimum wage increases and has policy implications for minimum wage laws. 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 offers a potential explanation grounded in worker power for the mixed results in earlier studies that examine the impact of minimum wage on corporate investment (e.g., the negative effects reported in Cho (2021) and Gustafson and Kotter (2022) and the positive effects documented in Geng, Huang, Lin and Liu (2022) and Hau et al. (2020)).[9] By focusing on this unique perspective, our study sheds light on the intricate relationship between minimum wage policies and corporate investment decisions and calls for a more comprehensive approach to labor market regulation.

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[9]Consistent with the notion of a higher degree of substitutability leading to the positive effects in China, Manu, McAdam and Willman (2022) estimate China's aggregate elasticity of substitution between capital and labor well above unity (1.2), which is roughly twice that documented for many developed countries.

II. Theoretical Motivation

This section illustrates a conceptual link between worker power and the elasticity of substitution between capital and labor using an endogenously derived production function under the mechanization framework (Zeira, 1998; Nakamura and Nakamura, 2008).[10] In this framework, technology is represented by the degree of mechanization, with its implementation difficulties—determined by worker bargaining power or outside options of employees—linked to the elasticity of substitution. Consider a representative profit-maximizing firm producing one final good by using a continuum of intermediate goods i in the interval $i \in [0, 1]$. These intermediate goods are produced by two different technologies: manual technology utilizing only labor or industrial technology using only capital (e.g., machines, robots). Specifically, the production function for intermediate goods is given by

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y(i) = \begin{cases} \eta(i)l(i), & manual\ technology\ (labor\ intermediate\ good) \\ \theta(i)k(i), & industrial\ technology\ (capital\ intermediate\ good), \end{cases}
\tag{1}
$$

where $l(i)$ and $k(i)$ are the units of labor and capital inputs in the production of intermediate good i, respectively, and $\eta(i) > 0$ and $\theta(i) > 0$ represent the factor productivity of labor and capital, respectively. Mechanization takes place when an increasing portion of intermediate goods are produced using capital instead of labor.

The intermediate goods are ordered by the relative efficiency of capital to labor, $\Psi(i) \equiv \theta(i)/\eta(i)$, in that $\Psi'(i) < 0$, where $\Psi(i) \geq 0$ and $\Psi(0) > 0$. Therefore, using industrial technology becomes increasingly difficult as i increases. Moreover, the shape of the function $\Psi(i)$ determines the difficulty of implementing mechanization. For example, the fast food services industry, with its standardized and repetitive tasks, is better suited to replacing workers with machines. As a result, it exhibits a relatively high efficiency of capital to labor (i.e., higher values of $\Psi(i)$) and a relatively

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[10]This framework can be easily generalizable to illustrate the elasticity of substitution between domestic and foreign workers.

flat shape. In contrast, the educational services industry relies heavily on human interaction and specialized expertise, making it less amenable to mechanization. This leads to lower efficiency of capital to labor (i.e., lower values of $\Psi(i)$) and a steeper curve.

The firm chooses the optimal share of intermediate goods and the amount of capital and labor for each intermediate good by solving the following problem with the resource constraints:

$$\max_{\{a,\; k(i)|i\in[0,a],\; l(j)|j\in[a,1]\}} lnY = \int_0^1 ln\; y(i)di = \int_0^a ln\; \theta(i)k(i)di + \int_a^1 ln\; \eta(j)l(j)dj, \quad (2)$$

$$s.t. \quad K = \int_0^a k(i)di \quad and \quad L = \int_a^1 l(j)dj$$

where the final good production function follows the Cobb-Douglas form and a is the fraction of capital intermediate goods. Nakamura and Nakamura (2008) derive the global production function as the envelope of local production functions. The derived global production function takes the constant-elasticity-of-substitution (CES) form, assuming $\Psi(i)$ decreasing log-linearly over i:

$$Y = e^{-\gamma}[bK^{1/(1+\gamma)} + (1-b)L^{1/(1+\gamma)}]^{1+\gamma}, \quad (3)$$

$$\Psi(i) = \left(\frac{b}{1-b}\right)^{1+\gamma}\left(\frac{1-i}{i}\right)^{\gamma}, \quad (4)$$

where $\gamma > 0$ is the shape parameter capturing the difficulty of implementing industrial technology and $0 < b < 1$. A smaller γ implies easier implementation. Then, the elasticity of substitution of the global production function equals $1 + 1/\gamma$. In our setting, if worker bargaining power becomes weaker, firms are better able to implement industrial technology when facing a labor cost shock. This is reflected in a lower value of the shape parameter γ, implying a higher elasticity of substitution and making firms less sensitive to the labor cost shock.

III. Minimum Wage and Corporate Investment

A. *Institutional Details*

The Fair Labor Standards Act (FLSA) contains the federal minimum wage standards for employees in the U.S. All employees are covered by the FLSA as long as their employers do interstate business and have annual sales of at least $500,000. The Act also covers hospitals, schools, and government agencies regardless of their dollar volume of business. Although enterprises do not meet the minimum annual sales threshold, their employees are covered during the workweek when workers are individually engaged in interstate commerce. In addition, the Act provides rules for overtime compensation, recordkeeping, and youth employment for workers in the private and public sectors. Since its enactment in 1938, it has undergone several amendments, mostly to raise the federal minimum wage. The FLSA is administered by the Department of Labor's Wage and Hour Division and provides protection to more than 143 million workers (about 93% of the civilian labor force in the United States) at more than 9.8 million establishments as of July 2009.

There are also many states with their own minimum wage regulations. Some states fix their minimum wage rates at the federal rate, link the rates to inflation, adjust the rates based on the scheduled legislative action, or combine the first three options. The federal minimum wage standards may not match the state minimum pay rates. When an employee is subject to both regulations, Section 18 of the FLSA states that the employee is entitled to the higher of the two standards.

B. *Identification Strategy*

Our identification for investigating the relation between corporate investment and minimum wage assumes that changes in the minimum wage rates are orthogonal to firm outcomes. The federal and state minimum wage rates fluctuate at different intervals and in different amounts. These changes are depicted in Figure B.1 for a sample of three geographically distant states: California, Connecticut, and Illinois. As shown in Figure B.1, the timing of minimum wage changes varies

at the federal level and across the states. In our tests, we exploit this staggered timing of changes in minimum wage rates. The potential issue of systematic disparities between the treatment and control groups would be mitigated since minimum wage increases are staggered, making it feasible for firms to be in both the treatment and control groups at various periods.

As outlined earlier, each state has its own process for adjusting minimum wage rates. One method is to index minimum wage rates to inflation. This adjustment procedure is problematic for our identification strategy because inflation might directly affect corporate investment.[11] To address this concern, we exclude all firms headquartered in states that index minimum wage rates to inflation.[12]

The second mechanism for adjusting minimum wages is to include specific future dates for specific minimum wage rates in legislation. Our identifying assumption could be violated if such legislation were driven by local economic conditions (e.g., the lawmakers' anticipation of improvement in investment opportunities). To address this, we include state-by-year fixed effects in our conditional analyses to account for the possibility of state-level minimum wage policies being endogenous to local economic conditions.[13]

The third adjustment mechanism used by states is to set their minimum wage rates based on the federal rate. This mechanism is similar to specifying rates through state law. We emphasize

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[11]In principle, there are two conflicting effects of inflation on corporate investment (Hochman and Palmon, 1983): depreciation allowances and interest deductions effects. First, since depreciation allowances are determined by historical costs (not current nominal values), the real tax advantage of depreciation declines with inflation. Second, the real tax benefit of interest rises with inflation because firms deduct interest expenses at nominal interest rates (not at real rates). Thus, the question of whether inflation influences corporate investment positively or negatively is one of the empirical studies (Feldstein, 1982).

[12]These 15 states are Alaska, Arizona, Colorado, Florida, Michigan, Minnesota, Missouri, Montana, Nevada, New Jersey, Ohio, Oregon, South Dakota, Washington, and Vermont. Our results are robust to including those states in our analysis.

[13]One approach used in the literature (e.g., Dube, Lester and Reich, 2010; Allegretto, Dube and Reich, 2011; Geng et al., 2022) to deal with endogeneity is to focus on border counties. However, we opt not to compare cross-border county pairs since neighboring counties across the state border may not be the best control group (Neumark, Salas and Wascher, 2014). Intuitively, those cross-border counties do not necessarily possess a similar business or regulatory environment that is subject to different state-level policies (e.g., state tax rates, local government spending on infrastructure, state-level labor regulations, etc.). Moreover, we do not observe corporate investment at the county level and firms do not usually operate along geographical borders.

that changes in *federal* minimum wage regulations may be seen as exogenous to the *state-level* macroeconomic conditions that could influence individual firm outcomes. Insofar as the federal minimum wage regulation is orthogonal to state-level economic circumstances, the identification enables us to isolate the impact of unobservable state-level macroeconomic shocks on capital expenditures. 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 U.S. states from the Tax Policy Center.[14] 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 hand-collect the data from the U.S. Department of Labor.[15] 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 exclude observations from financial institutions (SIC codes 6000–6999) and those with negative *Tobin's q*. We also eliminate firm-year observations with asset or sales growth exceeding 100% because these firms exhibit large jumps in business fundamentals in terms of size and sales, and these jumps are typically caused by major corporate events, such as mergers and acquisitions or reorganizations. The last step is to discard extremely tiny companies with capital of less than $10 million because linear investment models might not be suitable for such companies, as described in Gilchrist and Himmelberg (1995). Using the consumer price index for all urban consumers (CPI-U), we convert all dollar-valued variables into December 2014 constant dollars. The final sample has 59,096 firm-year observations.

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[14]https://www.taxpolicycenter.org/statistics/state-minimum-wage-rates
[15]https://www.dol.gov/whd/state/stateMinWageHis.html

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 the book value of assets.

For state-level variables, 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 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). 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.[16] These numbers indicate that there is considerable within-state variation in minimum wage rates.

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 the latest headquarters location.[17] In our dataset,

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[16]The absence of investment-wage sensitivity in the post-2000 period might be explained if within-state variation in minimum wage rates decreased with time. For the later sample period, however, the within-state standard deviation is higher: $0.72 ($1.00) for the pre-2000 (post-2000) period.

[17]We obtain robust results using the data on historical headquarters states provided by Bai et al. (2020) and the University of Notre Dame's Software Repository for Accounting and Finance (Loughran and McDonald, 2016). Another

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). The detailed definition of each variable is provided in Appendix A.

[Insert Table I here.]

E. Baseline Results

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

$$\frac{I_{i,s,t}}{K_{i,s,t-1}} = \alpha_i + \alpha_t + \beta_1 \, Tobin's \, q_{i,s,t-1} + \beta_2 \, \frac{CF_{i,s,t}}{K_{i,s,t-1}} + \beta_3 \, w_{i,s,t-1} + \beta_4 \, X_{i,s,t-1} + \epsilon_{i,s,t}, \quad (5)$$

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.[18] 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

concern is that a firm's operation may be geographically dispersed, hence its headquarters location does not necessarily reflect a relevant minimum wage rate for which most of its workforce is located. To address this concern, we count the number of times each 10-K mentions a U.S. state name and use the relative state counts for each state as a proxy for a firm's operational intensity of each state (García and Norli, 2012). We then restrict firm-year observations to have a state with an operational intensity greater than 50% and use that state to identify a firm's corresponding minimum wage rate. Since the mandatory filing via the EDGAR system was implemented after May 1996, we use the information in 1997 for the years before 1997. We confirm the declining sensitivity.

[18]The political orientation of state governors may affect the state-level minimum wage policy and the business environment that affects corporate investment decisions. Therefore, using the data collected from the National Governors Association, we include an indicator for the political parties of state governors (Democrat, Republican, or Independent) and confirm that the results are robust. Using real minimum wage also yields qualitatively similar results.

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 (5) 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, the 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×-0.017). This 270 basis point reduction corresponds to an 11.0% decrease, relative to the sample mean 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).

Columns (2) and (3) report results, respectively, from estimating equation (5) for the first- and second-half of the sample: subperiods 1984–2000 and 2001–2017. The results show a dramatic drop in the sensitivity of investment to minimum wage increases. More specifically, column (2) shows that, for the first half of the sample period, the estimated investment-wage sensitivity is −0.038 (more than twice the estimate for the full sample period) and statistically significant at the 1% 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 sharp contrast, column (3) shows that for the second half of the sample, firm-level investment is no longer sensitive to minimum wage increases: the estimated investment-wage sensitivity is 0.001 and statistically insignificant at conventional levels. The investment-wage sensitivities between the two sub-periods are statistically different at the 1% level of significance (χ^2-statistic = 12.52).[19] This trend is consistent with the rising elasticity of substitution between capital and labor

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[19]The economy of the United States has transitioned toward service and technology-based enterprises (e.g., toward white-collar jobs). Corrado and Hulten (2010) document that the share of intangible capital out of firms' total capital has increased from 25% (1973–1994) to 34% (1995–2007). This change could account for the lack of investment-wage sensitivity in the later sample period. However, we do not observe any substantial changes in industry composition in

documented in the literature (Cantore, Levine, Pearlman and Yang, 2015; Chirinko and Mallick, 2017; La Grandville, 2017). The pattern is also in line with a sharp increase in manufacturing firms' monopsony power (as measured by wedges between marginal revenue products of labor and wages) since the early 2000s (Yeh, Macaluso and Hershbein, 2022). We further examine this time-varying nature of investment-wage sensitivity and conduct a formal structural break analysis in Appendix B.A.

We acknowledge that the estimated sensitivity may identify a general-equilibrium effect that reflects the potential product or labor market spillovers. For example, workers may cross the border to pursue the minimum wage hike in the neighboring state, affecting corporate investment decisions. Although such spillovers are possible, there is little empirical evidence in the literature for their existence or quantitative importance (Aaronson, French, Sorkin and To, 2018).

F.   Minimum Wage Sensitive Industries and Time-varying State Specific Economic Conditions

To lend further credence to our empirical setting, we examine whether minimum wage sensitive industries exhibit more pronounced negative investment-wage sensitivities. We identify industries that are most subject to minimum wage changes using the percentage of workers who are paid at or below the federal minimum wage. The data are obtained from the Labor Force Statistics (2017) by the Bureau of Labor Statistics. The minimum wage sensitive industries include manufacturing, retail trade, leisure and hospitality, and services, accounting for 95% of minimum-wage workers in the private sector. We reestimate Panel A of Table II by interacting the minimum wage variable with an indicator for minimum wage sensitive industries.

[Insert Panel C of Table II here.]

The results are reported in column (1) in Panel C of Table II. The estimated coefficient on the

the two sub-periods: for instance, manufacturing industries (NAICS codes = 31, 32, or 33) account for 47% (45%) of firm-year observations in the sample from 1984 to 2000 (from 2001 to 2017). Also, we obtain similar results using the firm-year observations with zero R&D expenses.

interaction term is −0.006 and statistically significant at the 1% level, verifying our identification strategy. This result is consistent with findings documented in Gustafson and Kotter (2022). Furthermore, to the extent that labor's losses to capital are largely driven by low-wage workers who are less likely to be human capitalists (Eisfeldt et al., 2022), the declining sensitivity documented in Panel A would be more pronounced for those minimum wage sensitive industries. The results in columns (2) and (3) support this view. The estimated coefficient on the interaction term is negative in the pre-2000 period whereas it is positive in the post-2000 period. These results imply that the elimination of the investment-wage sensitivity over time is largely driven by the industries most subject to minimum wage policy. To address a concern that variations in minimum wage rates might be endogenous to state economic conditions, we include state-by-year fixed effects in columns (4)–(6). The results are similar to those in columns (1)–(3), confirming that our sensitivity estimates are less likely to be affected by unobserved time-varying state-specific economic conditions.

G. A Decrease in the Share of Minimum-Wage Workers

In estimating the sensitivity, we ideally want to consider the following two-stage system:

$$\frac{I_{i,s,t}}{K_{i,s,t-1}} = \beta \, Wage_{i,s,t} + \epsilon_{i,s,t}, \tag{6}$$

$$Wage_{i,s,t} = \delta \, w_{i,s,t-1} + \eta_{i,s,t},$$

where we omit the fixed effects and control variables for ease of exposition and $Wage$ indicates the average hourly wages of workers in firm i. If $w_{i,s,t-1}$ is orthogonal to $\epsilon_{i,s,t}$, an IV regression of investment on wages using the minimum wage variable as an instrument would recover the sensitivity, β. In contrast, a reduced-form regression of investment on the minimum wage variable would identify $\beta \times \delta$ where δ reflects a fraction of minimum-wage or low-skilled workers. Since we have limited information on $Wage$, we employ the reduced-form approach in estimating

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the sensitivity[20]. Therefore, the declining sensitivity could be driven by a declining fraction of minimum-wage workers over our sample period, even in the absence of declining worker power. To check this possibility, we reestimate the investment-wage sensitivity for industries with a non-decreasing percentage of minimum-wage workers. Comparing the fraction of minimum wage workers for the years 2003 and 2017, we identify the following industries with the non-decreasing percentage: mining, manufacturing, retail trade, and transportation and utilities.[21]

[Insert Panel D of Table II here.]

The results are documented in Panel D of Table II. The estimated sensitivities are similar to those obtained in Panel A. A formal statistical test to evaluate the null hypothesis of the equality of sensitivities between the two sub-periods is rejected at the 5% level. Thus, declining sensitivity is not a manifestation of decreasing fraction of minimum-wage workers.

H.   *Robustness of Baseline Results*

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

1.   **Placebo Test**

We perform a placebo test to check whether a *pseudo* minimum wage increase affects investment. Specifically, we repeat the estimation of equation (5) for the pre-2000 sample using a pseudo minimum wage variable. To construct the pseudo minimum wage variable ($w_{i,s,t-1}^{Pseudo}$), we randomly assign a firm i to a state s by ensuring 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 $w_{i,s,t-1}^{Pseudo}$ as the minimum wage at time $t-1$ in state s where firm i's *hypothetical* headquarters is located. Once all firms i in the sample are assigned in this manner, we estimate equation (5) using the simulated data and store the coefficient on $w_{i,s,t-1}^{Pseudo}$. This pro-

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[20]Only about 10% of firm-year observations in Compustat have non-missing values of the total labor costs.

[21]The data on the percentage of minimum-wage workers are available from 2002; however, the industry classification in 2002 was different from that in other years.

cedure constitutes one run of simulation and is repeated 1,000 times, and a distribution of $w_{i,s,t-1}^{Pseudo}$ coefficients is generated.

[Insert Panel B of Table II]

Panel B of Table II reports the empirical distribution of the coefficient on $w_{i,s,t-1}^{Pseudo}$. The mean and median of this distribution are 0.002 and 0.001, respectively, and both are close to zero. This suggests that, on average, there is no investment-wage sensitivity in our simulation. We plot the empirical distribution of the coefficient on $w_{i,s,t-1}^{Pseudo}$ in Figure B.2. The green line is the estimated nonparametric kernel density of the coefficient on $w_{i,s,t-1}^{Pseudo}$ coefficient. The red vertical line indicates the coefficient on $w_{i,s,t-1}$ (−0.038) obtained from 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), suggesting that our estimated sensitivity in the 1984-2000 period is not likely due to chance.

2. Measurement Error in Tobin's q

Tobin's q, a proxy for investment opportunities, typically contains measurement error that can generate biased estimates of the investment-wage sensitivity. To mitigate this issue, Erickson, Jiang and Whited (2014) propose minimum distance estimators for a traditional errors-in-variables model that includes variables with and without measurement error. Using these 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 Panel A of Table II for easier comparison. 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

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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)).

3. Strict Exogeneity Assumption Tests

In equation (5), 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 Panel A of Table II 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.

IV.   Worker Bargaining Power and Investment-Wage Sensitivity

As documented in the previous section, over the past four decades, corporate investment has become insensitive (statistically and economically) to increases in the minimum wage. In this section, we investigate the hypothesis that this trend reflects declining worker bargaining power, vis-a-vis their employers, due to both the decline in unionization and the increased availability of employer outside options made possible by globalization and labor-saving technological advances.

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

In this section, we use the 1999 U.S.-China Bilateral Agreement to investigate whether increased access to lower-cost Chinese labor improved the bargaining power of U.S. employers vis-a-vis their workers and thereby reduced the sensitivity of investment to mandated increases in the minimum wage.[22]   As documented by Devereaux and Lawrence (2004), among others, the 1999 U.S.-China bilateral agreement opened the economy of China to U.S. multinational firms by improving contracting institutions and by allowing U.S. firms to capture a greater share of the profits from their Chinese operations.[23]   In effect, the agreement opened up the Chinese labor market to U.S. firms through the enhanced ability it provided to maintain or establish more profitable operations in China. This, in turn, weakened the bargaining power of U.S. workers by increasing the outside options of their employers, i.e., firms faced with a minimum wage shock to their U.S. labor costs could more credibly threaten to move capital investments outside the U.S. to take advantage of cheaper labor.[24]

To identify firms that are more likely to benefit from the enhanced ability to access 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 first focusing on U.S. firms with at least one subsidiary in China as of 1997, two years prior to the bilateral agreement, and then consider those firms that newly established subsidiaries in China following the agreement. To identify U.S. firms with Chinese

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[22]Devereaux and Lawrence (2004) show that the bilateral agreement was largely unexpected due to strong opposition in the U.S. Congress making it useful as a means of identifying a causal relation between U.S. worker bargaining power and investment-wage sensitivity. Ceglowski and Golub (2012) show that relative unit labor costs in Chinese 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.

[23]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.

[24]Using a similar empirical strategy, Bena and Simintzi (2019) show that improved access to cheaper Chinese labor led U.S. firms to reduce their process innovation activities in the U.S., i.e., cheaper Chinese labor substituted for cost-reducing process innovation activities at home.

subsidiaries in our sample, we use hand-collected information from 10-k filings in 1997.

Our tests are based on the following difference-in-differences regression:

$$
\frac{I_{i,s,t}}{K_{i,s,t-1}} = \alpha_i + \alpha_t + \beta_1 \, Tobin's \, q_{i,s,t-1} + \beta_2 \, \frac{CF_{i,s,t}}{K_{i,s,t-1}} + \beta_3 \, w_{i,s,t-1} \tag{7}
$$
$$
+ \beta_4 \, Agreement_t \times w_{i,s,t-1} + \beta_5 \, China97_i \times Agreement_t \times w_{i,s,t-1}
$$
$$
+ \beta_6 \, China97_i \times w_{i,s,t-1} + \beta_7 \, China97_i \times Agreement_t + \beta_8 \, X_{i,s,t-1} + \epsilon_{i,s,t},
$$

where $China97_i$ 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; and $X_{i,s,t-1}$ 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 $Agreement_t$.[25] The outcome of interest in this generalized difference-in-differences framework is β_5 which captures the change in investment-wage sensitivity for firms operating in China as of 1997 relative to firms not operating in China. As shown earlier, our baseline estimate of the investment-wage sensitivity prior to 1999 is negative, (Panel A of Table II). If firms with greater access to cheap Chinese labor increase their bargaining power over their workers, we expect that their investment will become less sensitive to minimum wage increases, and hence, we expect β_5 to be positive and significant.

[Insert Table III here.]

Table III presents the estimation results for equation (7). In column (1), the difference-in-differences in the investment-wage sensitivity (β_5) is 0.035 and significant at the 1% level. We interpret this finding as follows: For the treated firms with Chinese subsidiaries at the time of the agreement, 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 significant, i.e., prior to the agreement, the treated firms significantly

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

adjusted their investment in response to mandated changes in the minimum wage. The investment-wage sensitivity for the treated firms 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, both economically and statistically insignificant, indicating that 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 effect of the agreement on the control group of firms that did not have Chinese subsidiaries at the time of the agreement. 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; prior to the agreement, the control firms, like the treated firms, significantly adjusted their investment in response to mandated changes in the minimum wage. The investment-wage sensitivity for the control firms 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. Thus, in contrast to the treated firms, the estimated investment-wage sensitivity for the control firms continues to be economically and statistically significant (albeit smaller in magnitude than before).

[Insert Figure II (a) here.]

Figure II (a) summarizes the results of column (1). It plots the changes in investment-wage sensitivities around the bilateral agreement for the treated and control firms. The solid dots indicate the point estimates of the 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 U.S.-China bilateral agreement significantly shifted bargaining power away from U.S. workers to U.S. firms because of the treated firms' improved access to cheap labor in China. As a result, these U.S. firms' investment decisions are not sensitive to minimum wage shocks. Together, these findings are consistent with the hypothesis that greater access to cheaper Chinese labor increases the bargaining power of employers vis-a-vis their workers causing a sig-

nificant decline in the investment response to minimum wage increases.[26]

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 U.S.-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 (7). 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.[27] Figure B.3 plots these results.

Figure B.3 is also consistent with the revealed preference theory, which can be used to analyze the China subsidiary choices of firms and to compare the influence of the U.S.-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

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[26]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.4 in Appendix B are qualitatively similar to those reported in Table III.

[27]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.4 in Appendix B.

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.

B. 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 U.S.-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 U.S. 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 U.S. 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 U.S. market and thus intensify the import competition for U.S. firms across many sectors of the economy. This increase in competition can be measured by the increase in the import penetration ratio of U.S. 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

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as U.S. gross output plus U.S. imports minus U.S. exports.[28] 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; Agarwal, Ayyagari and Kosová, 2024). 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 a capital-intensive production process, displacement of workers, automation (e.g., Dai and Qiu, 2023), 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 U.S. firms is eliminated after China's export surge (i.e., after 2001), and this effect is more pronounced for firms in 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.

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

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[28]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.

the amount of its imports and exports is at least \$10,000 per worker, or if the total amount exceeds \$500M.[29]

To test our hypothesis, we estimate the following difference-in-differences regression, which is similar to equation (7):

$$\frac{I_{i,s,t}}{K_{i,s,t-1}} = \alpha_i + \alpha_t + \beta_1 \, Tobin's \, q_{i,s,t-1} + \beta_2 \, \frac{CF_{i,s,t}}{K_{i,s,t-1}} + \beta_3 \, w_{i,s,t-1} + \beta_4 \, WTO_t \times w_{i,s,t-1} \quad (8)$$
$$+ \beta_5 \, Exposure_{UC,i} \times WTO_t \times w_{i,s,t-1} + \beta_6 \, Exposure_{UC,i} \times w_{i,s,t-1}$$
$$+ \beta_7 \, Exposure_{UC,i} \times WTO_t + \beta_8 \, X_{i,s,t-1} + \epsilon_{i,s,t}.$$

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.[30] 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 a 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 (8) 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 ($Exposure_{UC,i} = 1$) and control ($Exposure_{UC,i} = 0$) firms is 0.041 and significant at the 1% level. We interpret this

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

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

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 β_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 (8) 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.

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To address this concern and 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.[31] 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, which 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.[32] 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 (8). 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, $\beta_5 = 0.023$ is statistically significant at the 1% level and larger than the corresponding OLS estimate of 0.019 in column (2).

[Insert Figure II (b) here.]

We present the 2SLS estimates of investment-wage sensitivity for two groups of firms in Figure II (b). The first group is U.S. firms with no exposure to import competition with $Exposure_{UC,i} = 0$ (firms not vulnerable to import competition). The second group is U.S. firms with $Exposure_{UC,i}$

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

[32] 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.

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 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 into 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.[33] Motivated by their findings, we examine whether the differential changes in investment-wage sensitivities following China's accession to 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 (8) 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

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[33]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).

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.

C. *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 the U.S. 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).[34] We first define the routine-task intensity (RTI) score for each occupation as

$$RTI_k = ln(T_k^{routine}) - ln(T_k^{abstract}) - ln(T_k^{nonroutine\ manual}) \tag{9}$$

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

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

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 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 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 $Exposure_{tech,i} = 0$ are termed as firms with *Low Exposure* to technological change respectively. We then estimate a difference-in-differences regression that is similar to equation (7).

$$\frac{I_{i,s,t}}{K_{i,s,t-1}} = \alpha_i + \alpha_t + \beta_1\, Tobin's\ q_{i,s,t-1} + \beta_2\, \frac{CF_{i,s,t}}{K_{i,s,t-1}} + \beta_3\, w_{i,s,t-1} + \beta_4\, Post \times w_{i,s,t-1} \quad (10)$$
$$+ \beta_5\, Exposure_{tech,i} \times Post \times w_{i,s,t-1} + \beta_6\, Exposure_{tech,i} \times w_{i,s,t-1}$$
$$+ \beta_7\, Exposure_{tech,i} \times Post + \beta_8\, X_{i,s,t-1} + \epsilon_{i,s,t}.$$

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 (10) using $Exposure_{tech,i}$ as a continuous variable. Our inferences are qualitatively unchanged, and we discuss these results in Figure II (c).

[Insert Figure II (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.

D.   *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 I (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

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(on both 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 Figure I (a), we also plot the time series of investment-wage sensitivity estimated using 15-year rolling window regressions in Panel A of Table B.3 (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 a 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. Since union bargaining power is endogenously determined, we exploit the staggered passage of right-to-work (RTW) laws by the U.S. 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).

[Insert Table VI here.]

In column (1) 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 II (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 II (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, which indicates that the passage of RTW laws also has a direct positive impact on investment. These results are consistent with findings in Hirsch (1992), Fallick and Hassett (1999), and 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 U.S. firms over time.[35]

V. 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 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

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[35]In Appendix C, we conduct a simple counterfactual exercise to gauge the overall economic effects of a minimum wage increase on the 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 the general-equilibrium effects of the minimum wage increase on factor or output prices.

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 the 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 (5). 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 III.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 the 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*** [12.52]	
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 $w_{i,s,t-1}^{Pseudo}$. Panel C reestimates Panel A by including an indicator for minimum wage sensitive industries ($MW\ Ind$) and its interaction with the minimum wage variable. We identify industries that are most subject to minimum wage changes using the percentage of workers who are paid at or below the federal minimum wage, which is obtained from the Labor Force Statistics (2017) by the Bureau of Labor Statistics. The minimum wage sensitive industries include manufacturing, retail trade, leisure and hospitality, and services. Panel D reestimates Panel A within the subsample of industries with a non-decreasing fraction of minimum-wage workers by comparing the years 2003 and 2017. The definitions of all other 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.

Panel B. Placebo Test: Regression Coefficients from Bootstrapped Sample

	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

Dependent Variable: $Investment_{i,s,t}$ / Sample from 1984 to 2000

Panel C. Conditional Analyses on Minimum Wage Sensitive Industries

Dependent Variable: $Investment_{i,s,t}$

	Full Sample	1984 to 2000	2001 to 2017	Full Sample	1984 to 2000	2001 to 2017
	(1)	(2)	(3)	(4)	(5)	(6)
$w_{i,s,t-1}$	-0.013	-0.023**	-0.004			
	(0.009)	(0.010)	(0.005)			
$w_{i,s,t-1} \times MW\ Ind$	-0.006***	-0.020***	0.007*	-0.004**	-0.018***	0.009**
	(0.002)	(0.006)	(0.004)	(0.002)	(0.005)	(0.004)
$MW\ Ind$	0.031***	0.064***	-0.040	0.018*	0.051**	-0.049
	(0.010)	(0.024)	(0.036)	(0.010)	(0.021)	(0.036)
Controls / Firm FEs	Yes	Yes	Yes	Yes	Yes	Yes
Year FEs	Yes	Yes	Yes	No	No	No
State-by-Year FEs	No	No	No	Yes	Yes	Yes
# of Firm-Year Obs.	59,096	31,408	27,688	59,096	31,408	27,688
Adjusted R^2	0.140	0.136	0.123	0.427	0.454	0.488

Panel D. Industries with Non-decreasing Fraction of Minimum-Wage Workers

Dependent Variable: $Investment_{i,s,t}$

	Full Sample	1984 to 2000	2001 to 2017
	(1)	(2)	(3)
$w_{i,s,t-1}$	-0.018*	-0.035**	0.003
	(0.010)	(0.016)	(0.003)
H_0: (2)[$w_{i,s,t-1}$] - (3)[$w_{i,s,t-1}$] = 0		-0.038** [4.99]	
Controls / Firm and Year FEs	Yes	Yes	Yes
# of Firm-Year Obs.	40,389	22,136	18,253
Adjusted R^2	0.141	0.134	0.121

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

Column (1) presents difference-in-differences regressions of investment on minimum wages interacted with two indicators, $China97$ and $Agreement$, in equation (7). $w_{i,s,t-1}$ 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 U.S.-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 (8). $w_{i,s,t-1}$ 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 into 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 (10). 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 $Exposure_{tech,i}$ 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 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 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

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Table VI: Weakening Union Power: Passage of Right-to-Work (RTW) Laws

This table 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.

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: Investment Sensitivity to Minimum Wage Over Four Decades
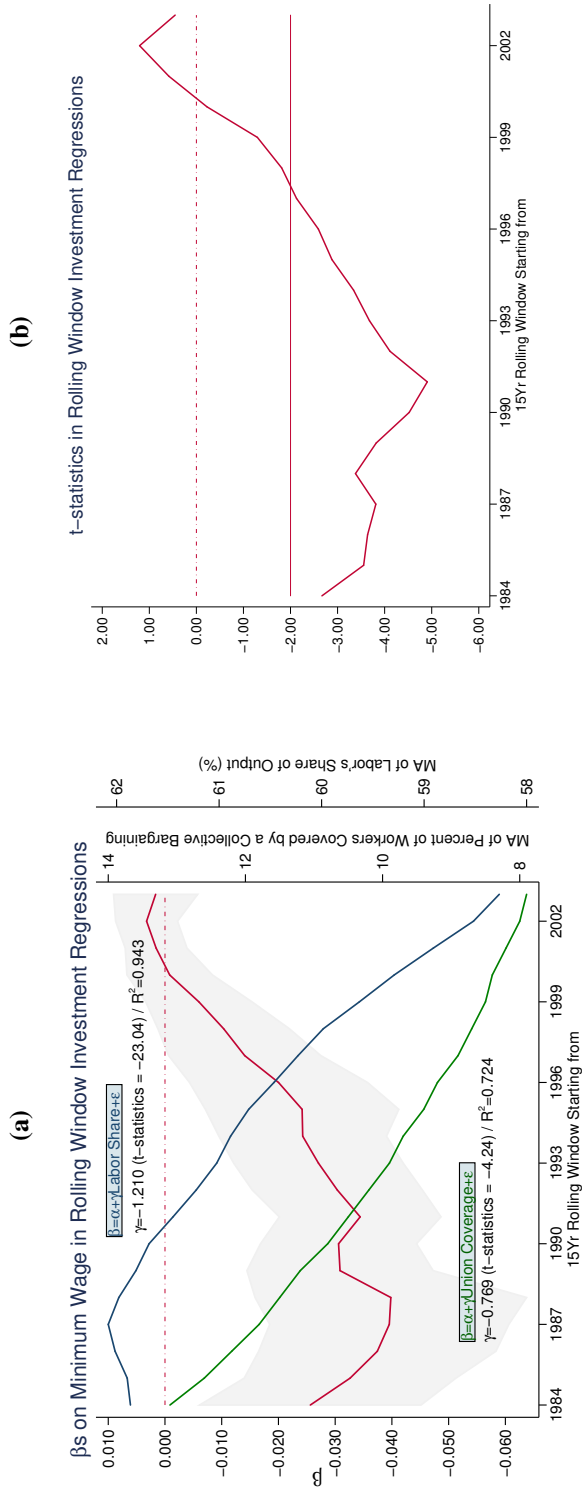
In Panel (a), the red line plots the time-series of estimated investment-wage sensitivity (β) from the 15-year rolling window investment regressions in Panel A of Table B.3. The grey shaded area indicates the 95% confidence interval. In Panel (b), the red line plots the time-series of corresponding t-statistics 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 A of Table B.3. 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 II: Declining Worker Power and Changes in Investment Sensitivity to Minimum Wage

(a) 1999 U.S.-China Bilateral Agreement

(b) Chinese Import Competition

(c) Technological Change / Automation

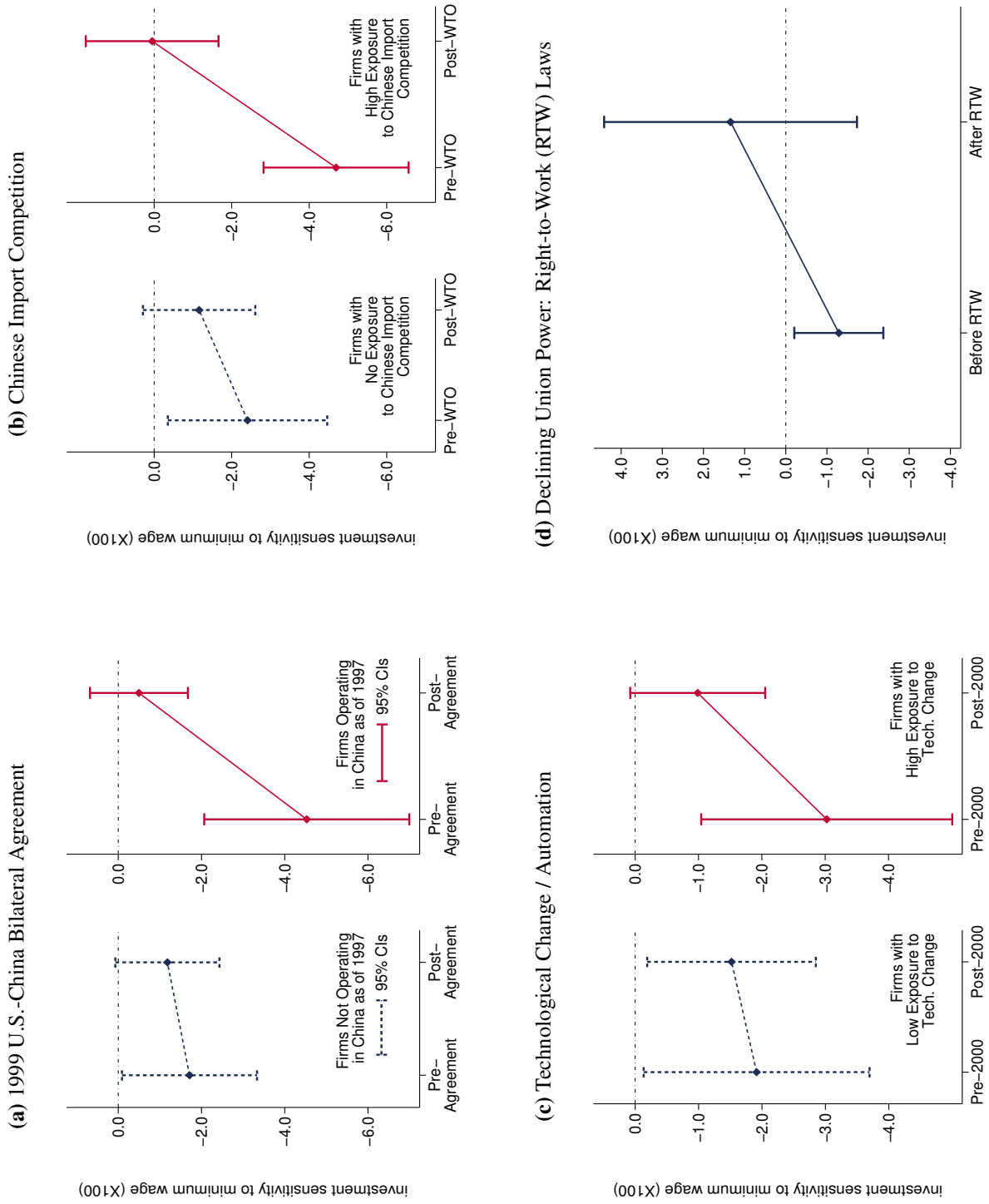**(d)** Declining Union Power: Right-to-Work (RTW) Laws

Figure II: 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 U.S.-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 U.S. 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 components 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.

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
$MW\ Ind$	An indicator variable set to one for firms in minimum wage sensitive industries. We identify those industries using the percentage of workers who are paid at or below the federal minimum wage, which is retrieved from the Labor Force Statistics from the Current Population Survey (2017) by the Bureau of Labor Statistics. The minimum wage sensitive industries include manufacturing, retail trade, leisure and hospitality, and services. They account for 95% of minimum-wage workers in the private sector.
Agreement	An indicator variable for the time period after the U.S.-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 U.S.-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 U.S.-China bilateral agreement in 1999), and zero otherwise

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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 percentage 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

A.   *Rolling Window Regressions and Structural Break Analysis*

To examine dynamic changes in investment-wage sensitivity over time, we run 15-year rolling window regressions. Specifically, we estimate equation (5) in the main text for twenty sub-sample periods. The first sample period runs from 1984 to 1998 and the last sample period runs from 2003 to 2017. We then obtain a time series of twenty estimates of investment-wage sensitivity and corresponding t-statistic (Panel A of Table B.3). 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 sensitivity (in magnitude) peaks in the samples from 1987 to 2001 and from 1988 to 2002. The coefficient then steadily decreases in magnitude (while continuing to be statistically significant) for all 15-year sub-periods with starting dates from 1988 to 1998. However, after 1998, the 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 I in the main text. In panel (a) of Figure I, the red line plots the time-series of estimated sensitivity and the grey shaded area, the 95% confidence intervals. The corresponding t-statistics are plotted in panel (b) with a horizontal line indicating t-statistics of −2.

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:

$$\frac{I_{i,s,t}}{K_{i,s,t-1}} = \alpha_i + \alpha_t + \beta_1\, w_{i,s,t-1} + \beta_2\, Z_{i,s,t-1} + d_t(k)\Big[\beta_3\, w_{i,s,t-1} + \beta_4\, Z_{i,s,t-1}\Big] + \epsilon_{i,s,t}, \quad (11)$$

where $Z_{i,s,t-1}$ is a set of firm- and state-level control variables used in equation (5), 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. We adopt two statistical methods to identify the best fit model: Akaike's information criterion (AIC)

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and Bayesian (or Schwarz's) information criterion (BIC). The lower the value of the criteria, the better the quality of the model. In Panel B of Table B.3, we plot AIC and BIC as a function of assumed break year k. Both AIC and BIC achieve their minimum value if $k = 1999$, indicating that a break date of 1999 best fits the data. These results are also consistent with the rolling window regression results.

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 the market value of assets to the book value of assets. Columns (1–2) report 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)

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

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: Rolling Window Regressions and Structural Break Analysis

Panel A repeats the estimation in Panel A of Table II in the main text 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 A. Time-Varying Effects of Minimum Wage on Investment: 15-Year Rolling Window Regressions

Sample Period		β_3	t-statistics
From	To	(Coefficient on $w_{i,s,t-1}$)	
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 B.3: Rolling Window Regressions and Structural Break Analysis (continued)

In Panel B, 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 the firm- and state-level control variables used in equation (5). $d_t(k)$ 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 a structural break.

Panel B. Analysis of Structural Breaks: Single Known Break

Table B.4: Robustness Results on 1999 U.S.-China Bilateral Agreement

Panel A repeats Table III in the main text by using 1998 (instead of 1997) as the year to check the 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 U.S.-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.4: Robustness Results on 1999 U.S.-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 U.S. 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: Minimum Hourly Wage Across some U.S. States (California, Connecticut, and Illinois), 1983-2017
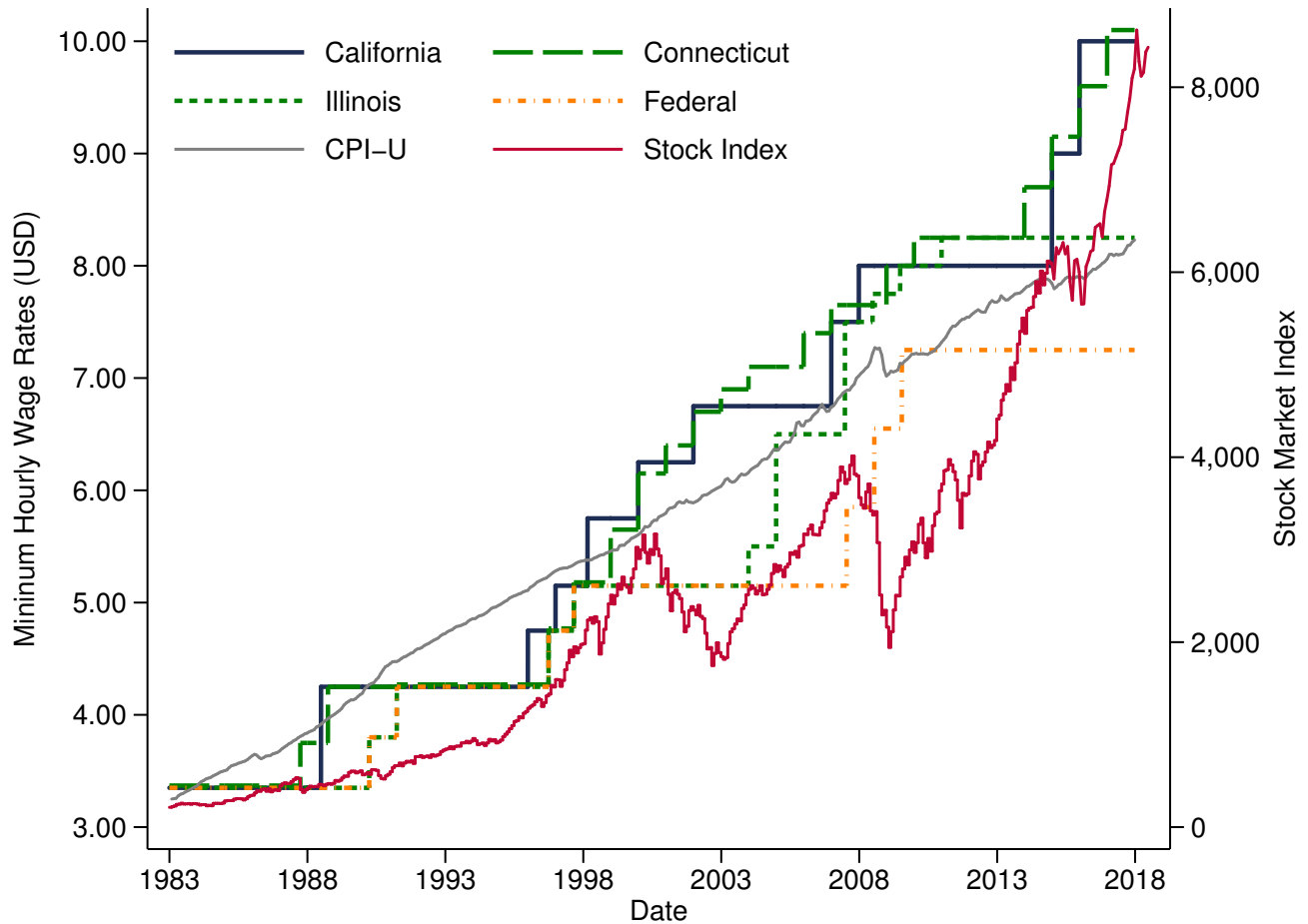
This figure shows the time-series of minimum hourly wage rates for three states and the federal government as an example for the time period 1983 to 2017. For 1983–2014, we obtain the historical minimum wage rates for non-farm private sector employment from the Tax Policy Center. These data are originally 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. We also plot the Consumer Price Index for All Urban Consumers (in a grey solid line) by setting the index value in January 1983 to a wage rate of $3.25 on the left axis and the value-weighted stock market index (NYSE/AMEX/NASDAQ) on the right axis.

## Figure B.2: Investment Sensitivity to Minimum Wage: Placebo Test
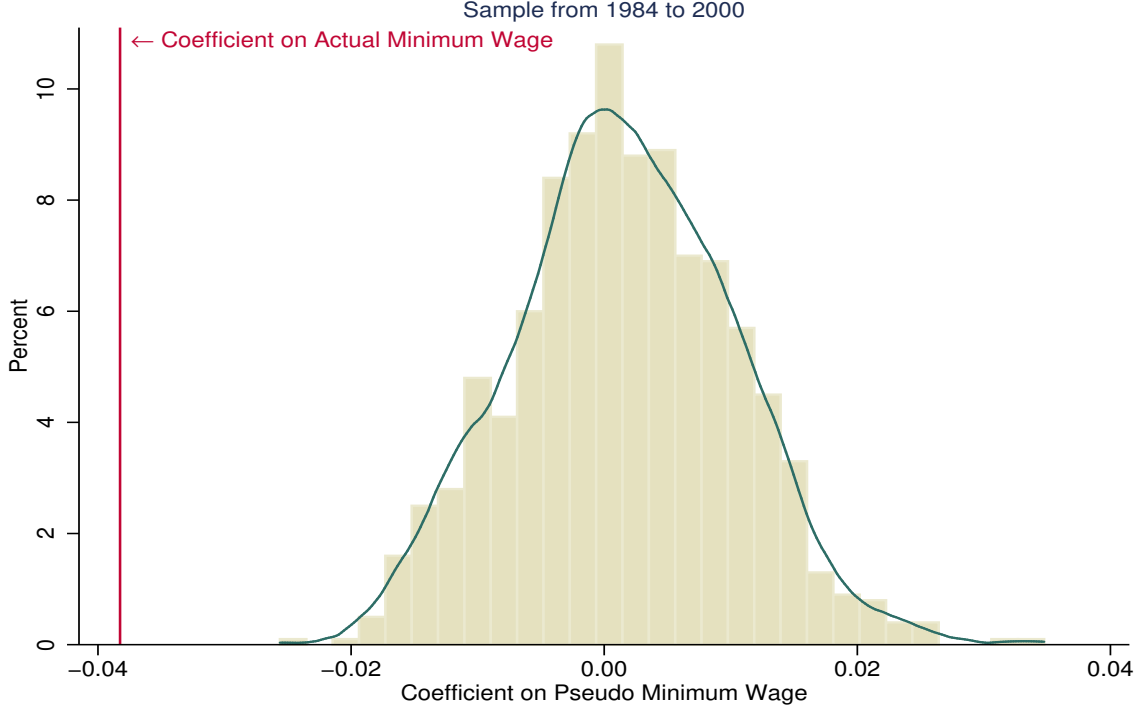
This figure is based on the following regressions:

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

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. We construct a pseudo minimum wage variable ($w_{i,s,t-1}^{Pseudo}$), by randomly assigning each firm i to state s. Using the pseudo state, we define $w_{i,s,t-1}^{Pseudo}$ as the minimum wage rate at time $t-1$ in state s where firm i's *hypothetical* headquarters is located. We then 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 displays the empirical distribution of the estimated coefficient on $w_{i,s,t-1}^{Pseudo}$. The green line is the estimated non-parametric kernel density. The red vertical line depicts 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.

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Figure B.3: 1999 U.S.-China Bilateral Agreement: Investment Sensitivity to Minimum Wage
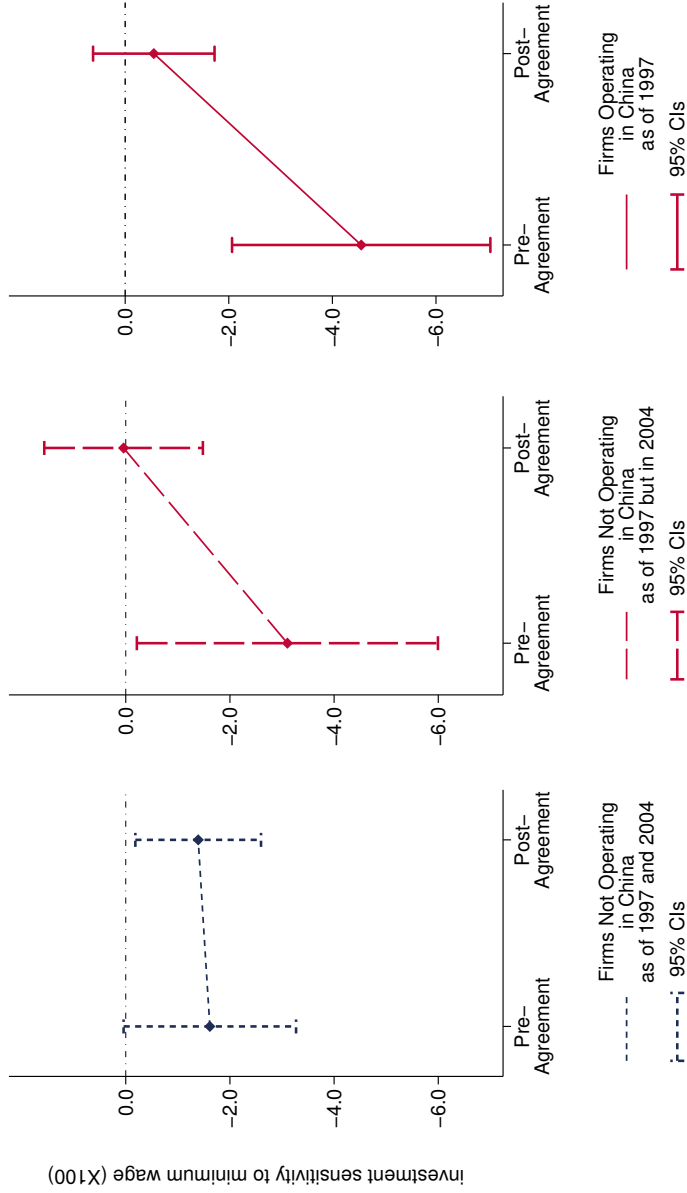
These figures plot the changes in investment-wage sensitivity for three groups of firms before and after the 1999 U.S.-China bilateral agreement. The first group (in the left figure) consists of firms that do not operate in China as of 1997, two years prior to the agreement, and remain not operating 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 U.S. 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.

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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.[36] 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.[37] 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.

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 the 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*

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[36]The Washington Post reported on Oct. 23, 2020, for example; "Biden wants to raise federal minimum wage. Trump doesn't. A look at their debate disagreement."

[37]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/`

workforce. We assume conservatively that all hourly-paid workers would fully benefit from a 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 of 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.

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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 U.S. 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.]

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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 of 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]		Parameter / Calculation	Description
Pre-2000	Post-2000		
-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$0.90	$\beta_3 \cdot \Delta w_{min} \cdot K \cdot (L/K) \cdot w$	opportunity cost of job losses through investment cuts
M$[-56.87, -11.86]	M$[-4.22, 5.93]		(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
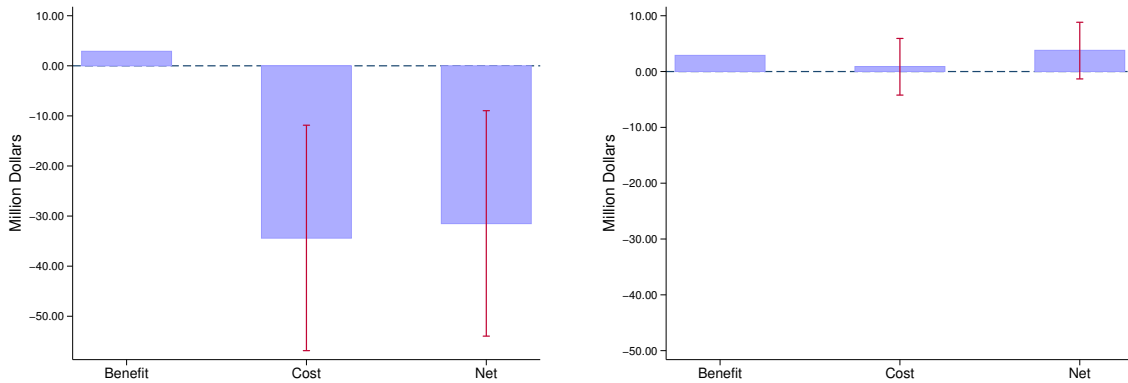
Panel C. Net Cost (Benefit) of Minimum Wage Increase to Workforce: For Firms that are Subject to a Minimum Wage Shock

(a) 1999 U.S.-China Bilateral Agreement   (b) Chinese Import Competition

(c) Technological Change / Automation   (d) Declining Union Power
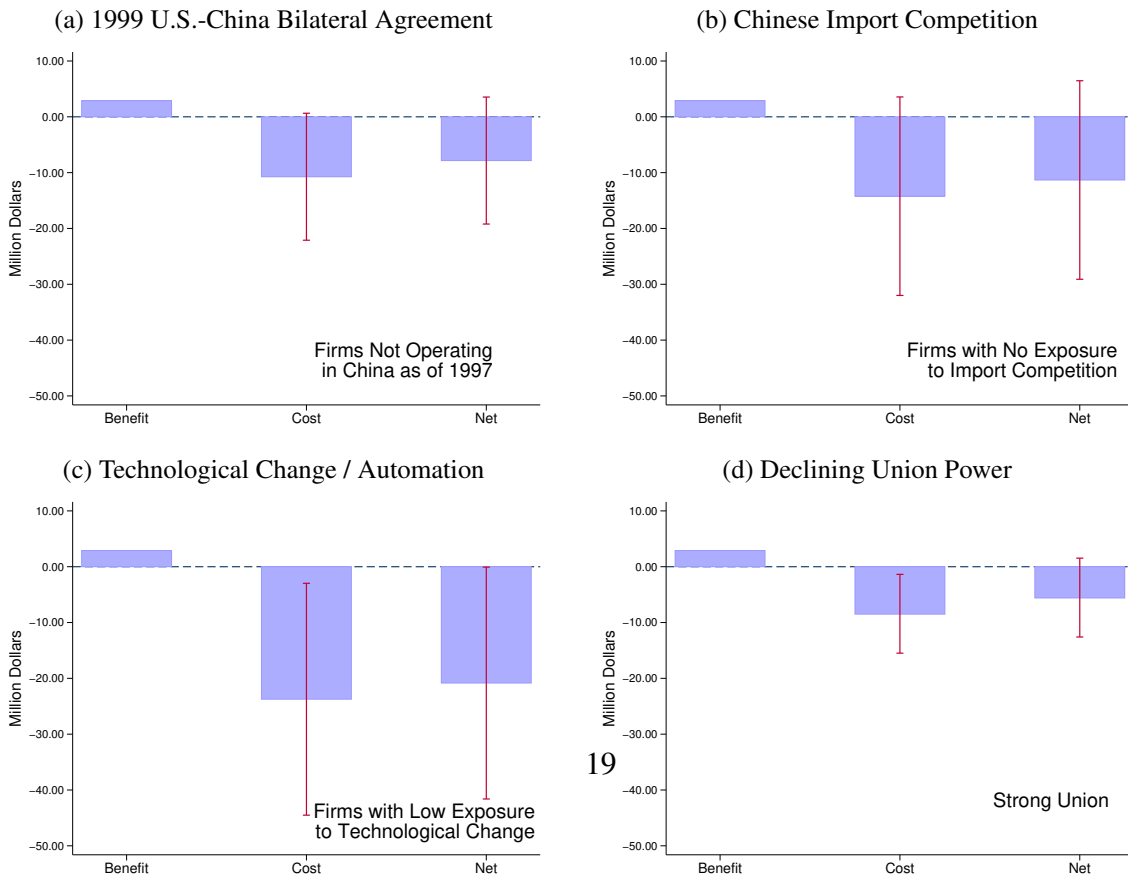

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Table C.2: Cost-Benefit Analysis of Minimum Wage Increase to Workforce: Counterfactual Analysis Using Alternative Parameter Values

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 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 capital is based on the average number of workers per unit of 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

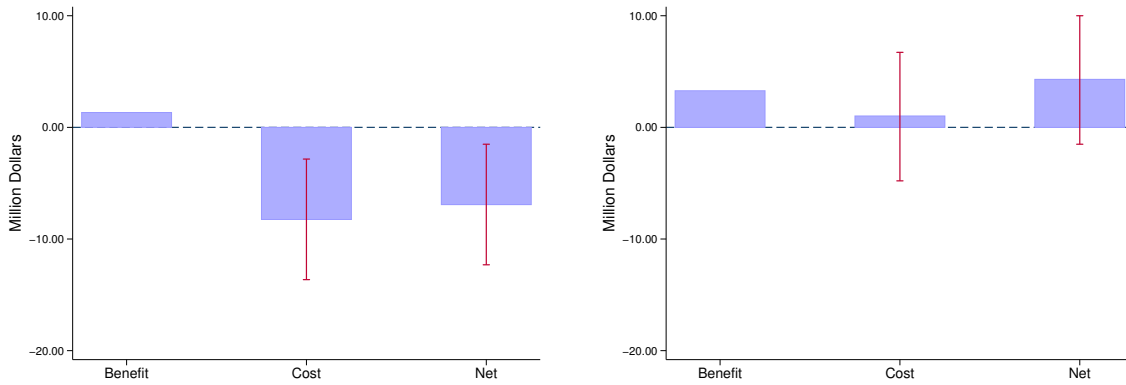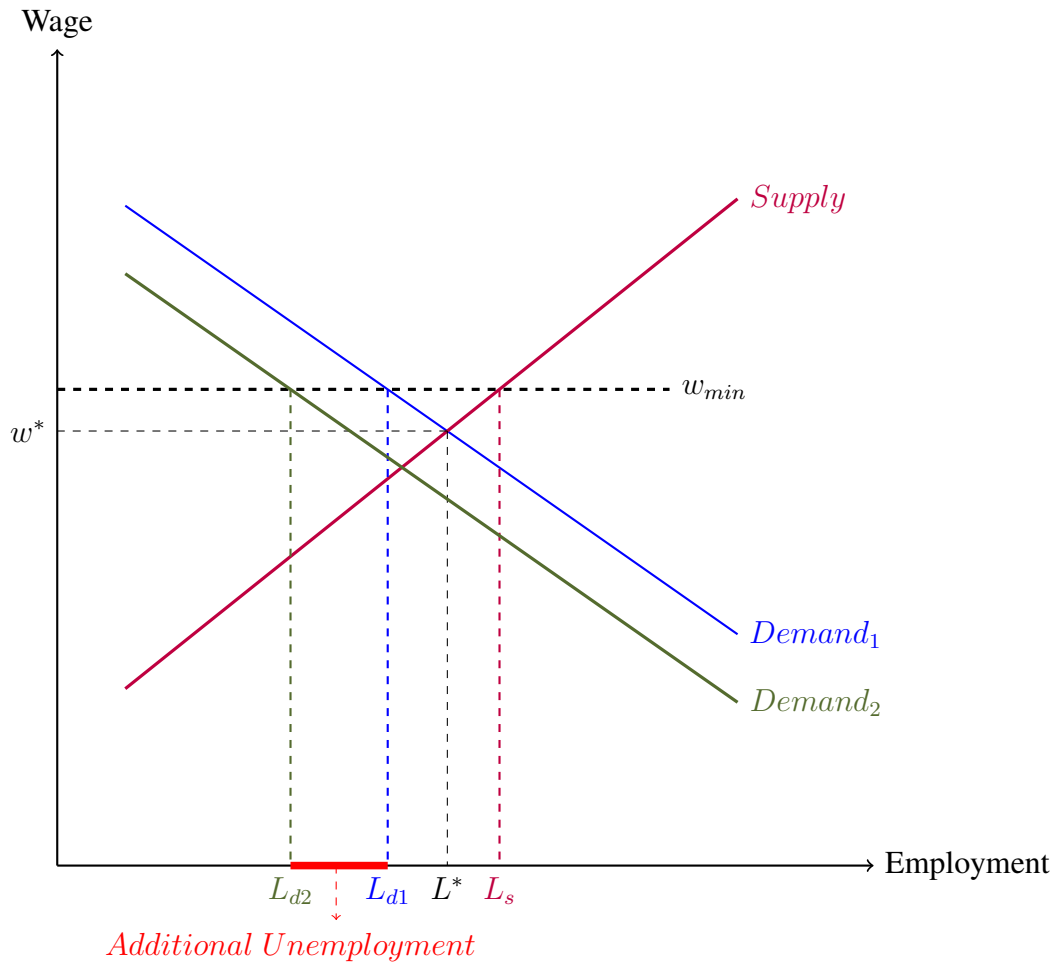(a) Pre-2000   (b) Post-2000

Figure C.1: Investment Cut and Unemployment

Wage

$Supply$

w_{min}

w^*

$Demand_1$

$Demand_2$

L_{d2} L_{d1} L^* L_s Employment

Additional 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.