The Effect of the Minimum Wage on Employment Growth in the Longer Horizon: An Event-Study Framework

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Abstract

This paper estimates the effect of the minimum wage on employment growth in the short run and in the longer horizon using an event-study framework. We construct a general equilibrium model and show that the minimum wage's employment effect depends on capital-labor substitution and the labor market monopsony power. A measure of exposure to the minimum wage at the county level is constructed by ranking the industries by their fraction of binding minimum wage workers. The event-study regression shows that a higher exposure to the medium-binding industries decreases the employment growth rates, with a short-run elasticity of -0.04 which evolves to -0.1 in the longer horizon. On the other hand, the effect is insignificant for the high-binding industries which are often thought to be more affected by the minimum wage. The results suggest that industries often ignored in the minimum wage literature are in fact quite responsive in employment adjustment following a minimum wage increase.

Keywords: minimum wage, employment growth.

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1 Introduction

Recent minimum wage hikes spurred ongoing debate on the employment effect of the minimum wage (see e.g. Cengiz et al. (2019), Neumark (2018)). The literature is separated into two strands, with one finding zero or positive employment effect and the other small negative effects.[1] As Baker et al. (1999) and Neumark and Wascher (1992) note, much of the evidence of a zero or positive employment effect is based on first-difference estimates of the elasticity, while the negative estimates are more likely the result of single panel data fixed-effects estimation. In both cases, the estimates are the short-run average employment effects of the minimum wage. They are not informative about the employment effect in the longer horizon, which might be larger as firms substitute alternative factors easier in longer horizons, a point emphasized by Lordan and Neumark (2018),Sorkin (2015), and Baker et al. (1999).

We take a different approach in the paper. In particular, we focus on two waves of the US minimum wage increases and use an event-study framework to estimate the employment effects of the minimum wage. Doing so allows us to examine the immediate short-run effect of the minimum wage on employment growth as well as its effects several years later. Being able to examine pre-trends also helps validate the empirical approach.

We first construct a general equilibrium model with the minimum wage. In the model, the production function is industry specific. Each industry produces an industry-specific intermediate good by combining a continuum of tasks as in Acemoglu and Restrepo (2018) and Acemoglu and Restrepo (2020). The industries differ in the range of tasks that are automatable, which is determined by industry-specific innovation. The firms have monopoly power in the goods market and monopsony power in the labor market. The firms can hire low-skill labors at the minimum wage or rent capital to complete automatable tasks.

The model produces an equation specifying the relation between the employment growth rates and the minimum wage which we use as the basis for the regression. The equation suggests that an increase in the minimum wage affects the low-skill employment growth rate by two channels. First, it changes the relative price between the low-skill labor and capital and hence the demand for the low-skill labor. We refer to the it as the labor demand effect. Consistent with theories and recent empirical evidence (see e.g. Card and Krueger (1994), Azar et al. (2019), Corella (2020)), larger monopsony power implies more positive employment response to the minimum wage increase.

Second, the minimum wage affects the industry's incentive to innovate, which affects the automatibility of tasks and hence capital labor substitution. The channel is consistent

[1] For an overview of the literature, see e.g. Card and Krueger (2015), Neumark and Wascher (2008).

with recent evidence in e.g. Lordan and Neumark (2018) and Aaronson and Phelan (2017). We refer to the second channel as the substitution effect. The second channel's effect on employment growth depends on the industry's willingness to innovate and the fraction of low-skill workers in the industry. The combination of the willingness to innovate and the fraction of low-skill workers gives a measure of exposure to the minimum wage.

We conduct the empirical estimation by calculating a county-level exposure to the minimum wage measure. Specifically, we first calculate the fraction of workers with binding minimum wages for each industry using Current Population Survey (CPS) from 2008 to 2017. We rank the industries by the fraction from high to low, where the industries ranked highest are most dependent on minimum wage workers. We refer to the industries in the top 25 percentiles as the high-binding industries, and the industries ranked in the next 25 percentiles (25th to 50th percentile) as the medium-binding industries. The county-level exposure to the minimum wage is the fraction of employment in the high- or medium-binding industries respectively, where the employment data uses the Quarterly Census of Employment and Wages (QCEW).

We leverage two waves of the federal minimum wage increases in 1996 and 2007, which were accompanied by various state-level minimum wage increases. We choose 1995 and 2006 as the base years and identify the states with clear pre-periods without further changes in the minimum wage. In particular, we require that the binding minimum wage needs to remain the same four years prior to the minimum wage increase, and we estimate the effect for the first four years after the minimum wage increase. We calculate the employment growth rates for the high(medium)-binding industries relative to the base years at the county level. Using within-state, cross-county variation in industry composition, we estimate the effects of a higher initial exposure to the high(medium)-binding industries on the employment growth rates of high(medium)-binding industries in the event study framework.[2]

The results show that minimum wage increases has little impact on the high-binding industries' employment growth rates. In the 1996 wave, the point estimates for the post minimum wage periods are insignificant and there is significant pre-trend. In the 2007 wave, while the point estimates for the post minimum wage periods are significant, it captures both the effects of the minimum wage and the Great Recession. After controlling for the Great Recession shocks that also correlates with the exposure to the minimum wage measure, the estimates become close to zero and insignificant.

[2]We do not expect the minimum wage to decrease the overall employment growth rates. As a robustness check, we examine the effect of a higher exposure to the high- and medium-binding industries on the employment growth rates of the high- and medium-binding industries and find significant negative effects. The majority of the decline comes from the medium-binding industries.

In comparison, higher initial exposure to the medium-binding industries implies a much larger decline in the corresponding employment growth rates, especially in the longer horizon. The first-year response to a minimum wage increase in counties with a 10% higher exposure to medium-binding industries is a 0.4% decrease in the employment growth rate, which extends to a 1% decrease four years later. The elasticity in the short run is about -0.04 which magnifies to about -0.1 in the longer horizon. The estimates are significant at the 5% level for the post periods and there is no pre-trend in the estimates. The results from the medium-binding industries suggest that the minimum wage's effect on the employment growth rates can last in the longer horizon, possibly due to the slow adjustment in the production process.

We control for potential confounding factors by estimating the effect of exposure to the low-binding industries on the employment growth rates of the low-binding industries, where the low-binding industries rank in the bottom 25 percentile in the ranking of the fraction of minimum wage workers. By subtracting the results, we control for factors such as mean reversion and interest rate increases.[3] After the subtraction, the effect for the high-binding industries remains small and insignificant, while there are still significant negative effects for the medium-binding industries.

Because the results are qualitatively and quantitatively similar in both waves, we argue that the minimum wage decreases the medium-binding industries' employment growth both in the short run and in the longer horizon. The results call for further investigation into industries that often left out in the analysis of the minimum wage's effects.

This paper makes several contributions to the literature. First, to the best of our knowledge, our paper is the first to estimate the effects of the minimum wage on employment growth in the event study framework, which adds to the literature on the employment effect of the minimum wage (see e.g. Card and Krueger (1994), Neumark (2001), Meer and West (2016), Dube et al. (2010), Neumark (2018), Cengiz et al. (2019), Lindner and Harasztosi (2019), Clemens and Wither (2019)). As Baker et al. (1999) mentioned, the first difference or the panel fixed effect regression obtain different results likely because they weigh the minimum wage changes depending on their frequency. We circumvent the frequency problem by using the event study framework and focusing on the recent two

[3]The federal funds rate increased from 3% in 1994 to 5.5% in 1995. It increased again from 1% in 2004 to 5.25% in 2006. The interest rate increase could be a potential confounding factor because it is a negative aggregate demand shock which could lead to lower employment. However, we think that it is unlikely that the interest rate increase would affect minimum-wage-worker-dependent industries more. Moreover, the interest rate dropped dramatically in 2007, but the employment effect continues to decline in our estimation. The interest rate increase could disincentivize capital investment, which can lead to smaller employment effect. Nonetheless, we argue that the aggregate demand affects industries similarly so that subtracting the coefficients of the low-binding industries controls for the confounding factor.

waves of minimum wage increases. The event study framework also allows us to validate the identification assumption by examining the pre-trends and estimate the employment growth rate response in the short run and in the longer horizon, i.e. four years after the initial minimum wage increase. Cengiz et al. (2019) estimate the long-run effect of the minimum wage on employment levels, while we focus on employment growth, similar to Meer and West (2016).

We show that higher initial exposure to the medium-binding industries implies a large decline in their employment growth rates. The result is important because the negative employment growth response could lead to a larger welfare decline for the low-skill workers. Even if the employment level does not decrease in the short run, the minimum wage can decrease employment growth in the longer horizon, which makes it difficult for the low-skill workers to find employment, as noted earlier in Kremer (1993) and recently in Lordan and Neumark (2018) and Aaronson and Phelan (2017).

Second, we show that the negative employment growth rate response in the longer horizon implies that the estimates in the two-way fixed effect regression and the panel DiD regression are biased upward. Meer and West (2016) notes a similar point, and Goodman-Bacon (2018) formalizes the econometric analysis. From 1992 to 2012, most of the variations in minimum wages center on small time windows around the two waves of minimum wage increases. If there are negative employment effects in the longer horizon, the states that increased the minimum wages less than four years ago are no longer valid controls because the pre-trend assumption is violated. We sidestep the issue by synchronizing the minimum wage increase and utilizing intra-state variations in the exposure to the minimum wage instead of inter-state variation in the timing of minimum wage increase.

Third, the results suggest that the minimum wage's effect on employment growth is not limited only to the high-binding or low-wage industries. In particular, the effect is most significant on the medium-binding industries in the longer horizon. The findings show that even if the minimum wage increase is not permanent as in Sorkin (2015), the effect can still be large on the employment growth rate if we focus on the longer horizon. The result is consistent with Lordan and Neumark (2018) who find that minimum wage increases cause low-skilled workers in automatable jobs to become nonemployed or employed in worse jobs. Baker et al. (1999) also emphasizes the point based on Canada context. We reach similar conclusions using the US administrative data.

Outline The organization of the rest of the paper is: section 2 lays out the model and derives the estimation equation. We show that the construction of the exposure to the minimum wage measure in section 3. We specify the regression framework in section 4 and present the results in section 5. We discuss the implications of the results in section 6. Section 7 concludes. All proofs are in the appendix section B.

2 Model

We layout a framework to study the minimum wage's effect on employment growth. The model is closest to Acemoglu and Restrepo (2018) and Acemoglu and Restrepo (2020). The economy consists of $|\mathcal{I}|$ industries, each produces one industry-specific good $Y_t(i)$. At time t, the final good production function is

$$Y_t = \left(\int_0^1 v_i^{\frac{1}{\sigma}} Y_t(i)^{\frac{\sigma-1}{\sigma}} di \right)^{\frac{\sigma}{\sigma-1}}, \quad \int_0^1 v_i di = 1, \quad \sigma > 0 \tag{1}$$

Y_t is the numeraire. In each industry i, the industry-specific final output $Y_t(i)$ is equal to the industry-specific output $Y_t^g(i)$ net of the cost of technology production $q_t(i)$. Industries use intermediate goods $X_t(i)$ and services $S_t(i)$ produce the industry-specific output $Y_t^g(i)$ with a Cobb-Douglas production function, augmented by the technology production $q_t(i)$. The production function for the industry-specific output $Y_t^g(i)$ is

$$Y_t^g(i) = \frac{\eta^{-\eta}}{1-\eta} \left[X_t(i)^{\alpha(i)} S_t(i)^{1-\alpha(i)} \right] q_t^{1-\eta} \tag{2}$$

High skill labor $h_t(i)$ produces service $S_t(i)$ in a one-to-one fashion. The production of the industry-specific intermediate goods $X_t(i)$ combines a continuum of industry-specific tasks $x_t(i,z)$, similar to Acemoglu and Restrepo (2018):

$$X_t(i) = \left(\int_0^1 x_t(i,z)^{\frac{\zeta-1}{\zeta}} dz \right)^{\frac{\zeta}{\zeta-1}} \tag{3}$$

Some of the tasks are automatable and use capital k_t, while the others need low-skill labor input l_t. The production technology of each task is

$$x(i,z) = \begin{cases} k_t(i,z) & \text{for } z \in [0, \theta_t(i)] \\ A_L l_t^{\gamma}(i,z) & \text{for } z \in (\theta_t(i), 1] \end{cases} \tag{4}$$

The formulation suggests that the tasks in the range $[0, \theta_t(i)]$ are automatable. Firm use capital to perform tasks in the range. The tasks in the range $(\theta_t(i), 1]$ can only use low-skill labor input. $\theta_t(i)$ measures industry-specific labor substitutability. If $\theta_t(i) = 0$, the tasks in industry i are labor-intensive, and capital cannot substitute labor in the production process. If $\theta_t(i) = 1$, the tasks are perfectly automatable, and firms do not hire low-skill labors. We normalize capital to map one-to-one to output. A_L denotes the productivity of labor relative to capital.

There are both monopoly power in the goods market and monopsony power in the labor market. ζ determines the monopoly power in the goods market, where a smaller ζ implies that the intermediate inputs $x_t(i,z)$ are complements. A smaller ζ hence means a larger monopoly power. In the labor market, γ measures the monopsony power of the firms. A smaller γ means a higher marginal cost of an extra worker, which is the result of a larger monopsony power.

Each period, firms produce capital by investment:

$$K_t = D(1+\xi) I_t^{\frac{1}{1+\xi}} \tag{5}$$

Capital fully depreciates after one period. Equation (5) implies that the rental rate of capital R_t is

$$R_t = D^{-1-\xi}(1+\xi)^{-\xi} K_t^{\xi} \tag{6}$$

Firms hire low-skill workers at the minimum wage M_t. The price of the industry intermediate good $X_t(i)$ is then

$$P_{X,t}(i) = \left(\theta_t(i) R_t^{1-\zeta} + (1-\theta_t(i)) \left(\frac{M_t^{1/\gamma}}{A_L} \right)^{1-\zeta} \right)^{\frac{1}{1-\zeta}} \tag{7}$$

The share of low-skill workers in the production of $X_t(i)$ is the payment to low-skill workers over the total revenue of producing $X_t(i)$. Denote it by $s_L(i)$. It depends on the substitutability of labor, the minimum wage, monopsony power and monopoly power:

$$s_L(i) = (1-\theta_t(i)) \left(\frac{M_t^{1/\gamma}}{A_L P_X(i)} \right)^{1-\zeta} \tag{8}$$

A technology monopolist determines the technology output $q_t(i)$. The technology output $q_t(i)$ affects both the output $Y_t^g(i)$ and the costs of technology production. We normalize the marginal cost of the technology output to be $(1-\eta)$ units of the industry-specific output $Y_t^g(i)$. The normalization implies that the optimal technology output is

$$q_t(i) = \frac{1}{\eta} X_t(i)^{\alpha(i)} S_t(i)^{1-\alpha(i)} \tag{9}$$

The technology monopolist also determines the technology innovation rate $\theta_t(i)$ in equation (4). The technology innovation rate $\theta_t(i)$ affects the monopolist's profit in two ways. First, it affects the price of the intermediate output $X_t(i)$ and hence the revenue of the monopolist. Second, as a higher innovation rate $\theta_t(i)$ determines the extent to which capital can substitute labor, it is costly to set a higher innovation rate. We assume that the marginal cost of increasing the technology innovation rate is $C_i(\theta_t(i))$. The marginal cost is industry specific, indicating that automation potential varies across industries. The monopolist would like to maximize her profits given by $\frac{1-\eta}{2-\eta}P_{Y,t}(i)^{1-\sigma}Y_t(1-C_i(\theta_t(i)))$. She takes the output Y_t as given. A log transformation of the profit equation means that her objective is

$$\max_{\theta_t(i)} (1-\sigma)\ln P_{Y,t}(i) + \ln(1-C_i(\theta_t(i))) \tag{10}$$

Let the wage paid to high-skill workers be W_t. We derive in the appendix that the price of the industry-specific final output $P_{Y,t}(i)$ is

$$P_{Y,t}(i) = \lambda(i)P_{X,t}(i)^{\alpha(i)}W_t^{1-\alpha(i)} \tag{11}$$

in which $\lambda(i) = (1-\eta)\alpha(i)^{-\alpha(i)}(1-\alpha(i))^{-(1-\alpha(i))}$. By equation (11) and equation (7), the profit maximizing equation of the monopolist is

$$\max_{\theta_t(i)} (1-\sigma)\alpha(i)\ln\left(\theta_t(i)R_t^{1-\zeta} + (1-\theta_t(i))\left(\frac{M_t^{1/\gamma}}{A_L}\right)^{1-\zeta}\right) + \ln(1-C_i(\theta_t(i))) \tag{12}$$

Equation (12) shows that the technology monopolist's incentive to innovate depends on the minimum wage M_t. The technology innovation in turn affects the employment growth rate of the low-skill workers. The following proposition provides a mathematical characterization.

Proposition 1. *Let $e_t(i)$ denote the low-skill employment share of industry i, defined as the fraction of low-skill workers out of the total low-skill workers employed in industry i. Taken price and technology innovation as given, the percentage change in low-skill employment satisfies the following equation*

$$\begin{aligned} d\ln L_t = &-\sum_i e_t(i)\frac{d\theta_t(i)}{1-\theta_t(i)} + \frac{1-(\zeta+\gamma)}{\gamma}d\ln M_t - \sum_i(1-\zeta)e(i)\chi(i)d\ln P_{X,t}(i) \\ &+ (1-\sigma)\sum_i e_t(i)(1-\alpha(i))d\ln W_t + d\ln Y_t \end{aligned} \tag{13}$$

in which $\chi(i) = 1-\alpha(i)(1-\sigma)/(1-\zeta)$.

Equation (13) suggests that the effect of minimum wages on employment growth has two parts. First, there is a labor demand effect captured by $(1-\zeta-\gamma)/\gamma d\ln M$. The sign of the effect depends on both the monopoly power ζ and the monopsony power γ. The second effect is capital-labor substitution given by $\sum e(i)d\theta(i)/(1-\theta(i))$. The capital-labor substitution might take several periods to realize and it depends on industry-specific substitutability of labor $\theta(i)$ and the industry's employment as a fraction of the total employment $e(i)$.

Let us first consider the labor demand effect. The sign of the labor demand effect depends on the monopoly power in the goods market ζ and the monopsony power in the labor market γ. Both are necessary to imply an increase in labor demand after the minimum wage increase. In particular, if $(\zeta+\gamma)<1$, an increase in the minimum wage leads to an increase in low-skill labor demand. The model emphasizes the role of monopoly power which is often ignored in the minimum wage literature: if the firms possess no monopoly power, any minimum wage that pushes the marginal cost above a certain level results in firm closure.

For a fixed level of monopoly and monopsony power, i.e. when $(\zeta+\gamma)$ is constant, the magnitude of the labor demand effect depends on the monopsony power γ. If $(\zeta+\gamma)<1$, the labor demand effect is positive. More labor market monopsony power (a smaller γ) implies that the labor demand effect is more positive. The result is consistent with Azar et al. (2019) and Corella (2020) who show empirically that minimum wage's employment effect is more positive in high concentration markets where firms have more monopsony power.

Next we consider the minimum wage's effect on the technology innovation rate, which determines capital-labor substitution. We make some assumptions about the marginal cost of technology innovation. Specifically, we assume it has the functional form

$$C_i(\theta_t(i)) = 1-(1-G(\theta_t(i)))^{\frac{1}{\rho(i)}} \tag{14}$$

in which G is an increasing and convex function such that $G'(0)=0$, $\lim_{x\to 1} G(1)=1$, and $g(x)\geqslant 1/(1-x)$, in which $g(x)=G'(x)/(1-G(x))$. The exponent $\rho(i)$ captures industry-specific technological possibilities for automation. Under a higher $\rho(i)$, it is easier for the technology monopolist to increase innovation rate.

We first state the proposition that qualitative characterizes the effect of the minimum wage on technology innovation.

Proposition 2. *Under the assumptions about the marginal cost of technology innovation, the technology innovation rate $\theta^*(M_t)$ is non-decreasing in the minimum wage M_t. A higher minimum* *wage causes the technology monopolist to choose a higher rate of innovation, leading to a larger measure of automatable tasks θ^*.*

The assumptions about the marginal cost function ensures the supermodularity between the innovation rate $\theta_t(i)$ and the minimum wage M_t. The supermodularity implies that a higher minimum wage leads to a higher technology innovation rate. As a result, a higher minimum wage induces capital-labor substitution. The increase in minimum wage makes low-skill workers more expensive relative to the capital, and the technology monopolist would set a higher innovation rate so that more tasks are automatable.

The higher innovation rate decreases low-skill employment. The magnitude of the decrease depends on the fraction of low-skill workers in a industry $e_t(i)$. If a industry employs a large fraction of the low-skill workers and it increases the technology innovation rate, the negative effect on employment growth is large. It also depends on the industry-specific marginal cost of technology innovation, which determines labor substitutability for an industry and hence the magnitude of $d\theta_t(i)$. If capital cannot substitute labor in an industry, the effect of the minimum wage would be small.

Equation (13) suggests that the employment effect of the minimum wage in a two-way fixed effect regression is not robust to the inclusion of state-specific time trend. In the partial equilibrium in which the firms take prices and the technology innovation as given, the minimum wage M_t could change the technology innovation $\theta_t(i)$. Let us consider the following two-way fixed effect regression:

$$\text{Employment}_{st} = \alpha + \beta \ln M_{st} + \text{State Fixed Effect} + \text{Time Fixed Effect} + \epsilon_{st} \tag{15}$$

The state and time fixed effect partially absorbs the price of industry-specific intermediate output $P_{X,t}(i)$, wage of high-skill workers W_t, and the aggregate output Y_t. β captures the effects of both the capital-labor substitution $\sum e_t(i)d\theta(i)/(1-\theta(i))$ and the labor demand effect $(1-\zeta-\gamma)/\gamma d \ln M_t$. Now let us consider including state-specific time trend:

$$\text{Employment}_{st} = \alpha + \text{State} \times t + \beta \ln M_{st} + \text{State Fixed Effect} + \text{Time Fixed Effect} + \epsilon_{st} \tag{16}$$

If the technology innovation or capital labor substitution takes place over several periods, the state-specific time trend will partly capture the capital-labor substitution effect $\sum e_t(i)d\theta_t(i)/(1-\theta_t(i))$, which changes the estimate β.

We can analyze the direction of the bias in the two-way fixed effect regression when we make assumptions about the capital-labor substitution effect. If the substitution effect is negative, in which case the minimum wage induces capital to substitute labor in the longer horizon, the coefficient β in the two-way fixed effect regression will be biased upward.

Consider the following example: there are two states A and B. There are three periods. In the first period, state A increases its minimum wage while state B does not. In the second period, state B increases but state A does not. Let the employment growth rate of A be $\{a_1, a_2, a_3\}$ and B be $\{b_1, b_2, b_3\}$. If there is no substitution effect so that the minimum wage only affects the employment in the period of the increase, state A can serve as the control state when state B increase its minimum wage in the third period. In a fixed effect or a panel DiD regression, the effect of the minimum wage on employment growth is then a weighted average of the two minimum wage increases, shown in figure 2 (a).[4] In particular, for the minimum wage increase in period 2, the estimated effect is $(b_3 - b_2) - (a_3 - a_2)$ in a standard DiD framework.

Assume there are negative substitution effects of the minimum wage on employment growth rates, as illustrated in figure 2 (b). The decrease in employment growth can last more than one period. In this case, state A is no longer a valid control state for the minimum wage increase in state B in period 3. In other words, state A would not be on a parallel trend to state B in the absence of the minimum wage change in state B. The estimated effect in period 3 is now biased upward: in (a), $(a_3 - a_2) = 0$ while in (b) $(a_3 - a_2) < 0$, which makes $-(a_3 - a_2) > 0$. Depending on the strength of the substitution effect, the bias could cancel the treatment effect $(b_3 - b_2)$ and even make the estimated effect to be positive.

The partial equilibrium equation (13) inspires the event-study regression in later analysis. For completeness, we solve the general equilibrium in the appendix. In the general equilibrium, only the technology innovation and the minimum wage are relevant for determining employment growth. We state the result in proposition 3 and leave the details in the appendix.

Proposition 3. *In the general equilibrium, the effect of the minimum wage on the low-skill employment growth rate is*

$$d \ln L_t = A \sum e_t(i) \frac{d\theta_t(i)}{1 - \theta_t(i)} + B d \ln M_t \tag{17}$$

The coefficient A and B depends on the model parameters.

While equation (13) gives the event study regression equation, we cannot directly use employment shares per period as an empirical measure of $\sum e_t(i) d\theta_t(i)/(1 - \theta_t(i))$, if we expect any employment effect of the minimum wage. We construct a measure of exposure to minimum wage in section 3 to estimate the effect of the minimum wage on employment

[4] Goodman-Bacon (2018) formalizes the simple example here.

growth.

3 Measuring Local Exposure to Minimum Wages

The effect of the minimum wage on employment growth in the longer horizon depends on $\sum e_t(i)d\theta_t(i)/(1-\theta_t(i))$. Since $d\theta_t(i)$ correlates with minimum wage changes and industry-specific capital-labor substitutability, we can instead use $\sum e_t(i) \times$labor substitutability$(i) \times \mathbb{I}_{dM_t>0}$ as the dependent variable. The question is then to measure empirically $e_t(i) \times$ labor substitutability(i).

We construct a measure of local exposure to the minimum wage at the county level. We choose 1990 to 2014 as our sample period, during which the minimum wage changes took place at the state level.[5] County level variables allow us to difference out the labor demand effect $(1-\zeta-\gamma)/\gamma d\ln M_t$ in equation (13) by including state-by-year fixed effects, if the counties within the state have the same distribution of monopoly and monopsony power. In reality, the assumption is not likely to hold. Even if the average monopoly or monopsony power might be similar across counties in the same state, if different industries across counties have different monopoly or monopsony power, we cannot completely cancel the labor demand effect if we analyze specific industries. Thus in the short run, the estimate is likely to capture both the labor demand effect and the substitution effect.

Regardless of the assumption on the distribution of monopoly and monopsony power, the substitution effects $\sum e_t(i)d\theta_t(i)/(1-\theta_t(i))$ vary at the county level. San Francisco county is unlikely to be affected by California minimum wage law as it mainly has high-wage industries. From the point of view in section 2, its "exposure" to minimum wages, defined as $\sum e_t(i)d\theta_t^{SF}(i)/(1-\theta_t^{SF}(i))$ is small because the share of employment $e_t(i)$ that is affected by the minimum wage is small. We are likely to see a larger effect in Fresno county as its industries ranked lower in average wages as a comparison.

To construct such a local exposure measure, we calculate the degree to which minimum wages bind for the industries. Specifically, we use CPS data from 2008 to 2017 to calculate the fraction of workers with binding minimum wages for each industry at Census 2002 4-digit code level. Excluding agriculture, military, food industry, and public administration industries, there are 257 industries. We rank the industries by the fraction of minimum wage workers from high to low. We map the industries to NAICS 4-digit code in QCEW.[6] A higher ranked industry has a larger fraction of employment with binding

[5]The exceptions are a few counties in New Mexico and District of Columbia which had county-level minimum wages. We exclude them because they do not satisfy our sample criteria below.

[6]Because we use the NAICS code in QCEW, our sample can only start at 1990 when the NAICS code is available. The data availability restricts us from analyzing other waves of minimum wage increases.

minimum wages. We define the high-binding industries as the ones that rank top 25 percentile in the fraction of binding minimum workers and medium-binding industries as those that rank top 25 percentile to 50 percentile. We define the local exposure measure as the fraction of total county-level employment in high- or medium-binding industries in the year before the minimum wage increase. With the measure, Fresno county would have a higher exposure to the high-binding industries than San Francisco county, because a larger fraction of the employment in Fresno is in the low wage industries.

Concretely, we define the local exposure measure as

$$\text{Exposure}_c = \frac{\sum_j \text{Emp}_{jc,\text{base year}} \mathbb{1}(\text{treated}_j = 1)}{\sum_j \text{Emp}_{jc,\text{base year}}} \tag{18}$$

That is, the exposure measure is defined as the county-level fraction of total employment working in industries that are considered treated (either high or medium-binding) in the base year. The measure is time-invariant and captures the initial exposure to subgroups of industries.

We define the base year as one year before a minimum wage increase. The base year needs to satisfy another condition: there are at least four years prior to the base year with no minimum wage increases.[7] Figure 1 gives two examples. Figure 1 (a) represents the minimum wage policy in California from 2002 to 2013. We can choose 2006 to be the base year and construct equation (18) accordingly. On the other hand, we cannot define a base year for figure 1 (b) which represents the minimum wage policy in Washington from 2002 to 2013.[8]

We utilize two waves of federal minimum wage increases: in 1996, the federal minimum wage was increased from \$4.25 to \$4.38, followed by a \$4.88 minimum wage in 1997, and eventually reached \$5.15 in 1998 and plateaued until 2007. The second wave of federal minimum wage changes started in 2007, with an increase from \$5.15 to \$5.50, followed by \$6.20 in 2008 and \$6.90 in 2009, and stopped at \$7.25 in 2010. During these two waves, many states concurrently adjust the state-level minimum wages upward, even if they exceed the federal minimum wage. We pool all the states that satisfy the base year definition in each wave. The restriction leaves us with 26 eligible states for the analysis of the 1996 wave and 27 states for the 2007 wave. Table 2 lists the states.

In the 1996 wave of federal minimum wage changes, we set 1995 as the base year, and

[7] We refer to minimum wage increases as increases in the binding minimum wage at the state level. For example, if a state's state-level minimum wage is lower than the federal minimum wage, and it increases its state-level minimum wage to the federal level, we do not consider the event as a minimum wage increase.

[8] The minimum wage increases twice in figure 1 (a). Ideally we would like a one-time minimum wage increase. Such occurrences are very rare, so we consider figure 1 (a) as a valid instance.

study the effect of minimum wages from 1992 to 1999. In the 2007 wave, the base year is 2006, and periods of analysis is from 2003 to 2010. Both waves allow four years before the base year without minimum wage changes. However, the federal minimum wage experienced a 27% increase from 1990 to 1992. There is likely a pre-trend in the outcome variable during the 1996 wave.

There is considerable variation in the exposure measure in each states. For example, in Glacier, Montana, less than 9% of the employment is in the high-binding industries while Gallatin County in the same state has 46% of its employment in the high-binding industries. We summarize the statistics of the exposure measure in the appendix.

4 Estimating Equation

Let b denote the base year. Based on section 2 and section 3, we estimate the effect of the exposure to the minimum wage on the employment growth rates using the following event study framework:

$$\Delta \text{Emp}_{ct} = \alpha + \sum_{y=b-4}^{b+4} \beta_y \left[\log\left(\text{Exposure}_c\right) \times \mathbb{1}(t=y)\right] + X_c' \gamma_t + \mu_{st} + \epsilon_{ct} \tag{19}$$

The outcome variable is the county-level employment growth rate of high- or medium-binding industries with respect to the base year:

$$\Delta \text{Emp}_{ct} = \frac{\sum_i \text{Emp}_{ict} - \sum_i \text{Emp}_{ic,\text{base year}}}{\sum_i \text{Emp}_{ic,\text{base year}}} \mathbb{I}_{(i=\text{treated})} \tag{20}$$

Note that the indicator function applies to both the numerator and the denominator. The specification means that the outcome variable is the employment growth rate for the subgroup of industries.

We choose the subgroups because we do not expect the minimum wage to affect the aggregate employment growth rates. We weight the regression by the ratio between the county-level high or medium-binding employment to the total high- or medium-binding employment in the base year:

$$\omega_c = \frac{\sum_i \text{Emp}_{ic,\text{base year}}}{\sum_c \sum_i \text{Emp}_{ic,\text{base year}}} \mathbb{I}_{(i=\text{treated})} \tag{21}$$

We restrict our sample to counties with employment growth between -80% and 80%. The employment data in QCEW covers establishments which report to the Unemployment

Insurance program of the United States. Report errors could result in large fluctuations in the employment growth rates. For example, the largest growth rate observation in our sample is 4400%. There are also observations in which the employment growth rate is -100%. Such occurrences more likely present misreporting than what actually happened. We run robustness checks in the appendix section A.3 by varying the cutoffs as well as including the full sample. The results are consistent both qualitatively and quantitatively across most of the specifications.

The identification assumption of equation (19) is that ϵ_{ct} is not correlated with the exposure measure in a way that is unrelated to minimum wages. By differencing, the outcome variable cancels out permanent level differences in county employment that can be correlated with the exposure measure.

We include in equation (19) the state-year fixed effect. The state-year fixed effect account for the effect of state policies that vary each year. In particular, the labor demand effect of the minimum wage is captured by the state-year fixed effect if the counties have the same distribution of monopoly and monopsony power.

We include additional sets of county-level controls X_c that could affect the employment growth rate. The first set of controls includes the log of county population, unemployment rate, and Herfindahl index in the base year. The variables reflect the county fundamentals that could have the predictive power of its future employment growth. The second set captures the effect of trade liberalization and technology changes concurrent with the minimum wage increase during the sample periods. In particular, we include county-level NAFTA index, exposure to China imports, and the 10-year change in the exposure to China imports. The details of the construction of the index is in the appendix section A.2.

There are concerns with the two waves of minimum wage increases. During the 1996 wave, the economy was in an expansion, so our specification is unlikely to capture county-specific trends that decrease employment but are unrelated to the minimum wage increase. On the other hand, there was a federal minimum wage increase in 1992. In particular, the percentage change in the federal minimum wage from 1989 to 1992 is 41%. The large increase makes it difficult to guarantee that there is no pre-trend.

In the 2007 wave, the drawback is that the minimum wage increase coincided with the Great Recession, making it challenging to interpret the estimates. For example, let us suppose the exposure measure is constructed using the high-binding industries, and the outcome variable is the employment growth rate of high-binding industries. If the Great Recession significantly reduces employment of high binding industries, the coefficients for the exposure measure should be negative in the absence of the minimum wage increase. In other words, the coefficients are capturing the employment effect of both minimum wages and the Great Recession on high-binding industries. We address this by using low-binding industries, defined as the bottom 25 percentile of the industries in the ranking of the fraction of binding minimum wage workers, to construct the exposure measure and estimating the effect on low-binding industries as a comparison.

5 Results

5.1 The 1996 Wave

We first examine the effect of exposure to the high-binding industries on the employment growth rates of high-binding industries, as well as exposure to the medium-binding industries on their employment growth. We include the main results for both waves in table 1.

For the high-binding industries, there are significant increasing pre-trend before the minimum wage increase. The results suggest no evidence of the employment effect after the minimum wage increase. Figure 3 (a) shows that a 10% more initial exposure to high-binding industries implies a 0.2% decrease in the employment growth rate of high-binding industries in the first year of the minimum wage increase, i.e. 1996 or β_1. The estimate for β_1 is significant at the 5% level. Four years later, the effect evolves to a decline of 0.4%, although the estimate is insignificant.[9]

For the medium-binding industries, there is no significant pre-trend in the medium-binding industries' employment growth rates. After the minimum wage increase, the employment growth rate significantly declines, with an effect larger than the effects for the high-binding industries. In particular, a 10% more initial exposure to the medium-binding industries would decrease its employment growth rate by 0.4% in the first year of the minimum wage increase, which magnifies to 1.4% four years after the initial increase. In the longer horizon, the -0.14 elasticity is much larger than the -0.04 short run elasticity. The result is consistent with Lordan and Neumark (2018) who suggest that the minimum wage affects employment in medium wage industries by inducing capital-labor substitution. Overtime, the decline in the employment growth rates is almost linear. The linearity is consistent with the firms gradually substituting capital for labor to smooth factor adjustment. The magnitude of the estimate is consistent with Meer and West (2016) who find

[9] Although the outcome variable is the employment growth rate instead of the log of the employment growth rate, the numerical values are small so that the log transformation produces similar results. In other words, we can interpret the results as elasticity instead of semi-elasticity. Meer and West (2016) use the same interpretation. We use the log of the employment growth rate as the outcome variable and the results are almost identical, which are available upon request.

a short-run elasticity of -0.05 when the outcome variable is overall employment growth rates in a two-way fixed effect regression. We note in section 2 that their result is likely a weighted average of the effects over time.[10]

Note that we include state-year fixed effect, which captures other state-wide shocks not correlated with the exposure measure that change every year. For example, if the counties within the states have the same distribution of monopoly and monopsony power, section 2 shows that the labor demand effect of the minimum wage is the same for every county in a state, which will be captured by the fixed effects. A issue rises when the shocks are correlated with industry compositions but uncorrelated with the set of controls in section 4. In this case, the shocks could confound the effect of minimum wages. One such example is mean reversion. In the presence of mean reversion, counties with larger share of baseline employment in the high-binding industries may experience slower high-binding employment growth over time.

We use low-binding industries to capture the effect of the shocks correlated with the exposure measure other than the minimum wage. The low-binding industries are industries ranked in the bottom 25 percentiles in the fraction of binding minimum wage workers. The minimum wages are unlikely to affect the employment growth of these industries. The key assumption that we make is that the shocks do not affect industries differently based on the industry's dependency on the minimum wage workers. For example, the effect of mean reversion satisfies the assumption. Under the assumption, the estimates would capture the effect of shocks that affect the employment growth rates net of the minimum wage effects. If the results are similar to figure 3, it means that these shocks are driving the result. Otherwise, the minimum wage is the more likely driver of the results.

The results in figure 4 shows that there is no evidence that the low-binding industries experience employment growth slow down in the 1996 wave. In particular, there is no significant pre trend, and the estimates for the post-periods are all insignificant. If we subtract the point estimates in figure 3 (a) by the point estimate in figure 4, the coefficients for the post periods become -0.008(0.013), -0.014(0.022), -0.028(0.033), and -0.006(0.041). The results suggest that after controlling for other shocks that could correlate with the exposure measure, the effect of minimum wages on employment growth is small and insignificant for the high-binding industries.[11]

On the other hand, the effects remain large for medium-binding industries. After subtracting the estimates in figure 4, the post-base year estimates are -0.032(0.014), -0.065(0.021),

[10]See Goodman-Bacon (2018) for an econometric formalization. We replicate their results in section A.1.

[11]The standard errors are calculated using the delta method assuming independence of the estimates. There is no correlation between the exposure measures for the high-binding and the low-binding industries.

-0.093(0.030), and -0.097(0.036). Namely, under the previous assumption, a 10% more medium-binding exposed county experiences a 0.3% decrease in employment growth rate the first year after the minimum wage increase. It aggravates to 1% four years after the initial minimum wage increase.

The assumption that the shocks other than the minimum wage affect each group of industries equally is strong. However, all of the coefficients for the low-binding industries are insignificant, suggesting that during the 1996 wave, the existence of other shocks lacks evidence. Moreover, the effect of the minimum wage on the medium-binding industries' employment remains large and significant at the 5% level even after subtracting the effects of the other shocks.

Another potential problem is that the exposure measure might be correlated with time-varying county specific characteristics that also affect the employment growth rate. For example, it is unlikely that the counties have the same distribution of monopoly and monopsony power, so that time-varying county distribution of monopoly and monopsony power will imply dispersion in the labor demand effect. The issue is less severe because it simply means the estimates, especially the short-run estimates (β_1 and β_2) likely capture both the labor demand effect and the substitution effect. In other cases, county-specific characteristics could decrease employment growth in the absence of the minimum wage increase.

We address the issue by using county-level characteristics to identify commonality among counties experiencing a decrease in the employment growth rate in the high- or medium-binding industries before the minimum wage increase. We test whether these common characteristics changed after the minimum wage increase. In particular, we use the variables that are unlikely to be severely affected by the minimum wage but are important for employment growth, such as county-level population density. We leave the details in the appendix.[12]

Lastly, we combine the high- and medium-binding industries and examine the effect of exposure to these industries on their combined employment growth rates. The results in figure 6 are almost the un-weighted averages of the separate results in figure 3. This is consistent with the lack of a significant correlation between the high- and medium-binding exposure measures.

Overall, we conclude that the 1996 minimum wage increases induce a sustained decline in the employment growth rates of the medium-binding industries, even when we subtract the coefficients for the low-binding industries. There is little evidence that the minimum

[12] Azar et al. (2019) use population density to rule out confounding factors. They also use the average wage, which is not suitable for our case because the average wage is arguably affected by the minimum wage.

wage decreases the employment growth rates of the high-binding industries.

5.2 The 2007 Wave

The results in the 2007 have similarities and differences compared to the results in the 1996 wave. Overall, the point estimates are much larger in magnitude. For the high-binding industries, the coefficients are more precisely estimated. In particular, the minimum wage significantly decreases the employment growth rates of the high-binding industries. It suggests that a 10% more high-binding exposed county would face a 0.5% decrease in employment growth rate in 2007. The decline reaches 1% in 2010. Similar to the 1996 wave, the increasing pre-trend is reversed after 2007. The effect is likely a combination of the minimum wage and the Great Recession.

For the medium-binding industries, the 2007 wave reserve the qualitative results from the 1996 wave. Figure 7 (b) shows that the minimum wage significantly decreases the employment growth rates of the medium-binding industries. We do not find significant pre-trend of the effect for the medium-binding industries at the 5% level, although the point estimates are quite large compared to those in the 1996 wave. The results show that a county with 10% more initial exposure to the medium-binding industries suffers a 1% decrease in the employment growth the first year after the minimum wage increase. In four years, the effect damps to a 2.1% decrease.

While the overall qualitative results are consistent across the two waves, the point estimates see a large decrease in the 2007 wave. The larger magnitude could come from the larger percentage increase in minimum wages in the 2007 wave: the federal minimum wage increases by 41% from 2006 to 2009, compared to 21% from 1996 to 1998. It could also be that the Great Recession affects the industries in a way that is correlated with the exposure to the minimum wage measure.

To subtract the effect of the Great Recession, we regress the low-binding industries' employment growth on the exposure measure to the low-binding industries similar to section 5.1. The results in figure 8 show some evidence of the Great Recession's impact on the employment growth of the low-binding industries: β_1 is significant at the 5% level and β_3 significant at the 10% level. There is also clear evidence of pre-trend in the employment growth rates, unlike in the 1996 wave.

Regardless, we subtract the estimates in figure 8 from the ones in figure 7. The point estimates for the high-binding industries adjust to -0.021(0.014), -0.058(0.021), -0.055(0.026), and -0.079(0.029). For medium-binding industries, the numbers are -0.069(0.028), -0.148(0.034), -0.158(0.035), and -0.203(0.041). Even after subtraction, most of the adjusted results are larger in magnitude compared to the 1996 wave. Besides the larger increases in the minimum wage in the 2007 wave, it could be that the Great Recession affects industries differently, so that the estimates using the low-binding industries do not capture the effects of the Great Recession on the high- and medium-binding industries.

To account for the confounding factors, we assume that the minimum wage has no impact on the high-binding industries' employment growth during the 2007 wave. Under the assumption, the coefficients for the high-binding industries are only capturing the effect of the Great Recession. We also assume that the Great Recession has similar impact on the high- and medium-binding industries. Subtracting the estimates for the medium-binding industries by the ones for the high-binding industries, the estimates after the minimum wage increase are -0.048(0.029), -0.091(0.037), -0.102(0.040), and -0.124(0.046). The numbers are similar to the ones in the 1996 wave and the estimates are significant at the 5% level except for β_1.

Taking together the results from the 1996 and the 2007 wave, we can see that despite the different aggregate economic conditions, the results are similar both qualitatively and quantitatively for the high- and medium-binding industries. In particular, there is little evidence that the minimum wage decreases the employment growth rates of the high-binding industries, while it significantly decreases the employment growth rates of the medium-binding industries. The point estimates suggest that the first year after the minimum wage has an elasticity of -0.04, which evolves to a elasticity of -0.1 four years after the initial minimum wage increase.

To study the overall effects, we run the regression by pooling the high- and medium-binding industries as in section 5.1. The results again are close to the unweighted average of the separate estimates. Figure 10 shows that a county with 10% more initial exposure to the high- and medium-binding industries has a -0.6% decline in the corresponding employment growth rate in the first year after the minimum wage increase. The effect becomes -1.4% in four years. The estimates for the post-minimum-wage-increase periods are significant at the 5% level. There are some evidence of pre-trends. Overall, the result is consistent with those in the 1996 wave that the minimum wage decreases employment growth for the combined high- and medium-binding industries.

6 Discussion

We employ the event study framework to circumvent the use of panel data methods which are known to produce disagreeing results in the minimum wage literature. The framework is well motivated by the structural model in section 2.

Whenever the minimum wage's effect last for more than one period, estimates in the panel DiD regression and the two-way fixed effect regression are biased.[13] In particular, if the effect is negative in the longer horizon as shown section 5, the estimates are biased upwards. One evidence of the upward bias is a positive pre-trend in the panel DiD regression. Using CPS data, we confirm that there are significant positive trend in the number of jobs below the new minimum wage.[14] Another evidence is that the estimates in the two-way fixed effect regression decrease after we include the state-specific time fixed effect. The state-specific time fixed effects capture the graduate employment adjustment induced by the minimum wage and partly undo the upward bias.

The results in section 5 are similar in both waves of minimum wage increases. The consistent findings are that the minimum wage has little effect on the employment growth rates of the high-binding industries while it significantly decreases the employment growth rates of the medium-binding industries. The short-run elasticity is -0.04, which evolves to about -0.1 four years after the initial minimum wage increase. Several recent papers emphasize the capital-labor substitution channel in the paper (Lordan and Neumark (2018), Baker et al. (1999), Sorkin (2015), Clemens and Wither (2019), Aaronson and Phelan (2017)). In particular, the small short-run employment effect of the minimum wage does not imply the same in the longer horizon. While the employment effect and the displacement effect is small, the minimum wage can induce the firms to substitute capital for labor and create fewer vacancies. The decrease in vacancy would lower the employment growth rates in the longer horizon, just as we see in section 5.

For the medium-binding industries, the decline in the employment growth rates is almost linear. The result is consistent with firms smoothing capital investment over time. Several recent papers start to explore the minimum wage's implication on industries or workers often ignored in the minimum wage literature. The cost of displacing workers can be large. For example, the firms might face law suit when displacing workers, and there could be productivity drop associated with displacement. The productivity drop can happen not only to the displaced workers, but to the coworkers as well.

When the cost of displacing workers is large, the firms might not layoff workers when the minimum wage increases. More and more recent evidence suggests that the firms are likely to adjust the production factors. For example, Aaronson et al. (2018) show that restaurants with lower ratings tend to exit the market following a minimum wage increase, ensued by an increase in entry of higher-rated restaurants. It could well be the case that the new entrants depend less on labor, which would lead to a decrease in the employment

[13] In our current empirical specification, one period is a year.
[14] The details are in the appendix.

growth rates. In essence, our model captures all capital-labor substitution of the sort induced by the minimum wage increase. It could be that within the same establishment, as fast-food restaurants replace workers with self-order machines. It could also happen that the old establishment is replaced with another one that is less labor intensive.

Even if the total employment does not change, there can be reallocation of employment opportunities, as capital replaces low-skill worker and increases demand for high-skill worker. Kremer (1993) note the reallocation theoretically, and recent papers find empirical support (e.g. Lordan and Neumark (2018), Aaronson and Phelan (2017)). The medium-binding industries' employment could decline while the other industries experience employment increase.

Another related evidence is the polarization of the US labor market, which sees drops in employment and wages in routine cognitive jobs in particular. Autor and Dorn (2013) and Autor et al. (2015) note that trade shock might induce automation, which leads to the hollowing of medium income jobs. The trade shock in effect increases the relative cost of low-skill labor. In the case of the minimum wage, it also lead to an increase in the cost of low-skill labor. The paper find evidence that indeed the medium-binding industries shift away from labor in the recent two waves of the minimum wage increase.

7 Conclusion

We use an event study framework to estimate the long-run effects of minimum wages on employment growth. We exploit cross-industry variation in minimum wage binding degree interacted with county-level variation in industry concentration. The results show that minimum wages have persistent long-run negative effects on employment growth, and the effect is most significant on the medium-binding industries rather than the high-binding industries. The estimates suggest that a 1% increase in the exposure to medium minimum wage binding industries reduces the employment growth rate by 0.1% to 0.4% percent in the first of the minimum wage increase and 1.6% to 3.4% percent in the fourth year of the initial minimum wage increase.

We interpret the results as evidence that minimum wage induces industries to substitute labor with capital. The effect is concentrated on the medium-binding industries because their production technology is more flexible or because automation already starts to replace workers, as is the case for the aging manufacturing industries documented in Ericksson et al. (2019). In the latter case, the minimum wage accelerates the substitution of labor with capital.

The results have important welfare implications for low-skill workers: even if one assumes no employment effect of minimum wages, it causes long-run employment growth deceleration, making it more difficult for low-skill workers to find jobs. Compared to the literature that focus on average short-run employment effect of minimum wages in food/retail sale industries, we emphasize the need to examine the effect of minimum wages on the long-run employment growth of the medium-binding industries that are capable of substituting labor with capital. The negative long-run trends in the employment growth rates also imply that estimates in the two-way fixed effect model and the panel DiD model are biased upward. Namely, if the true effects are negative, estimates in these regressions are biased towards being zero and even positive.

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Tables and Figures

Table 1: The Effect of Minimum Wages on Employment Growth Rate

	1996 Wave				2007 Wave			
	(1) High	(2) Medium	(3) Low	(4) High & Medium	(1) High	(2) Medium	(3) Low	(4) High & Medium
β_{-4}	-0.079***	-0.022	0.005	-0.035	-0.144***	-0.030	-0.006	-0.090***
	(0.023)	(0.034)	(0.021)	(0.024)	(0.025)	(0.037)	(0.028)	(0.035)
β_{-3}	-0.061***	-0.010	0.008	-0.023	-0.125***	-0.041	-0.031*	-0.068**
	(0.021)	(0.035)	(0.017)	(0.022)	(0.021)	(0.034)	(0.017)	(0.028)
β_{-2}	-0.038**	-0.012	-0.001	-0.021	-0.063***	-0.034	-0.039***	-0.034
	(0.017)	(0.037)	(0.013)	(0.019)	(0.017)	(0.027)	(0.012)	(0.022)
β_{-1}	-0.025**	0.014	-0.003	-0.010	-0.037***	-0.029*	-0.027***	-0.017
	(0.012)	(0.009)	(0.007)	(0.013)	(0.011)	(0.015)	(0.009)	(0.012)
β_1	-0.019**	-0.044***	-0.012	-0.021**	-0.047***	-0.096***	-0.026***	-0.057***
	(0.009)	(0.011)	(0.010)	(0.010)	(0.012)	(0.026)	(0.008)	(0.020)
β_2	-0.027*	-0.077***	-0.012	-0.049***	-0.069***	-0.160***	-0.011	-0.097***
	(0.016)	(0.013)	(0.016)	(0.016)	(0.018)	(0.032)	(0.011)	(0.026)
β_3	-0.035	-0.100***	-0.007	-0.068***	-0.078***	-0.180***	-0.023*	-0.112***
	(0.022)	(0.018)	(0.024)	(0.023)	(0.023)	(0.033)	(0.013)	(0.030)
β_4	-0.044	-0.135***	-0.038	-0.075**	-0.090***	-0.214***	-0.012	-0.135***
	(0.030)	(0.024)	(0.027)	(0.030)	(0.026)	(0.038)	(0.013)	(0.034)
Observations	15226	10762	12042	15341	16474	14537	15306	16614
State-Year FE	Y	Y	Y	Y	Y	Y	Y	Y

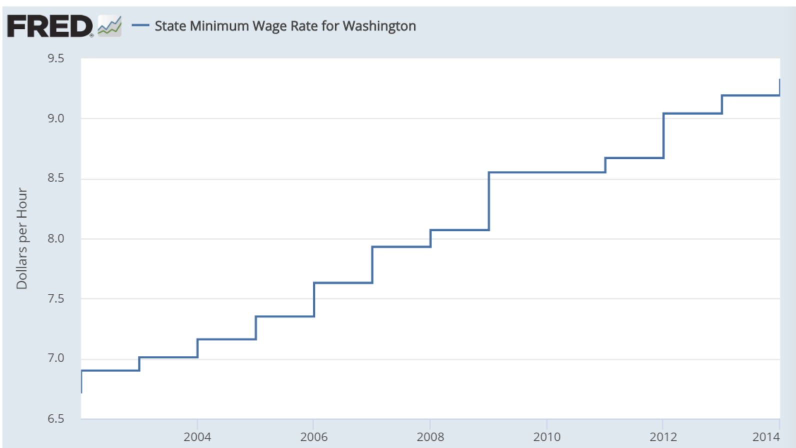
Notes. Table 1 shows the estimates of the effect of the minimum wage on employment growth. The regression is equation (19). The standard error is clustered at the county level. * means significant at the 10% level, ** 5% level, and *** 1% level. The set of controls that are common in both waves are the county-level log of population, unemployment rate, exposure to China imports, and the 10-year change in the exposure to China Imports, county-level Herfindahl index, and tariff index. The construction of the set of index is in section A.2.

Table 2: Feasible State-level Minimum Wage Changes

1996	1996	2007	2007
Alabama	Montana	Alabama	Mississippi
Colorado	Nevada	Arkansas	Nevada
Georgia	New Hampshire	California	New Hampshire
Idaho	North Dakota	Georgia	New Jersey
Illinois	Oklahoma	Idaho	North Dakota
Indiana	Pennsylvania	Indiana	Oklahoma
Iowa	South Carolina	Iowa	Pennsylvania
Kansas	South Dakota	Kansas	South Carolina
Kentucky	Tennessee	Kentucky	South Dakota
Louisiana	Texas	Louisiana	Tennessee
Michigan	Utah	Massachusetts	Texas
Minnesota	Virginia	Michigan	Utah
Mississippi	Wyoming	Minnesota	Virginia
			Wyoming

Notes. Table 2 lists the states that have valid minimum wage increases during the 1996 and 2007 wave. The validity criterion is that the state needs to have fours prior to the minimum wage increase with any other minimum wage increase.

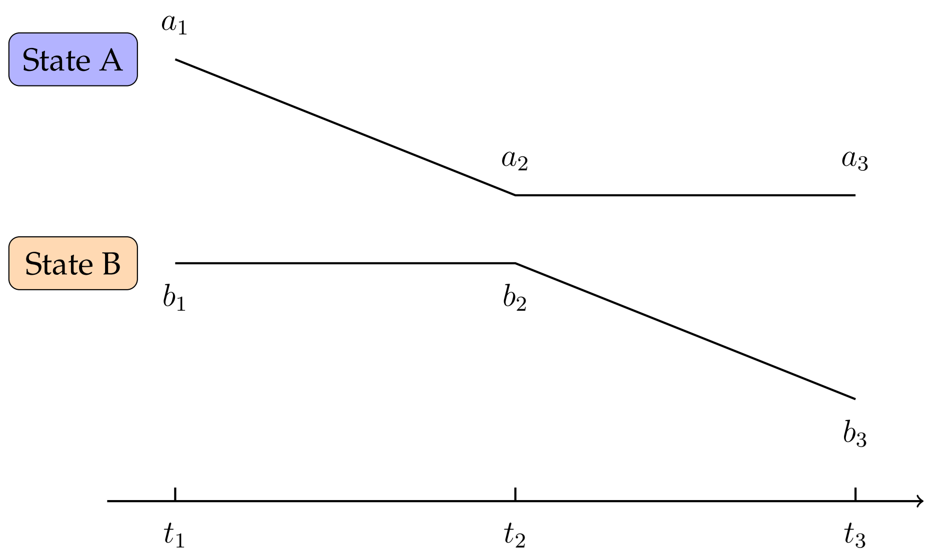
(a) California Minimum Wage



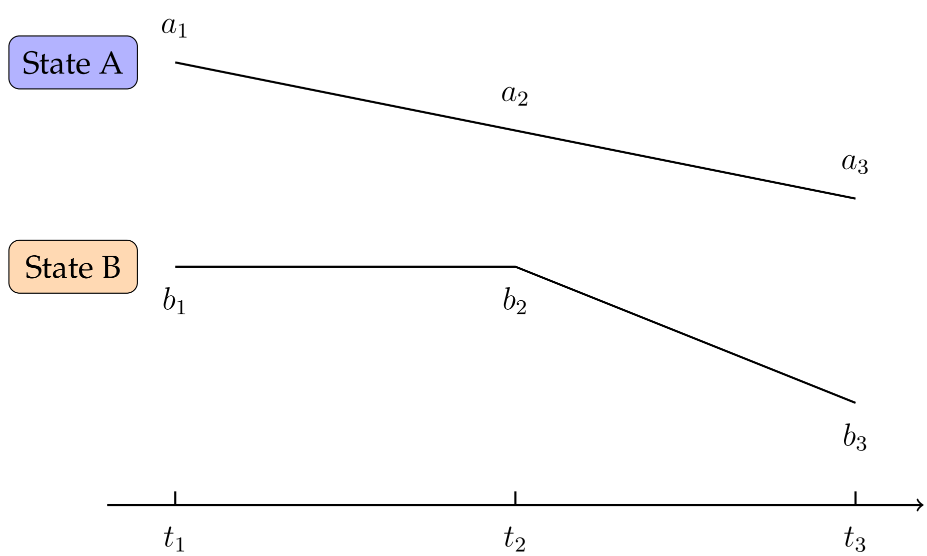
(b) Washington Minimum Wage

Figure 1: Base Year Example

Notes. Figure 1 shows an example of a valid minimum wage increase and an invalid one. California had no minimum wage increase from 2002 to 2006, increases its minimum wage in 2007 and 2008, and kept at the 2008 level until 2014. Washington increases its minimum wage every year from 2002 to 2014, so that it does not have valid minimum wage increases during the period.



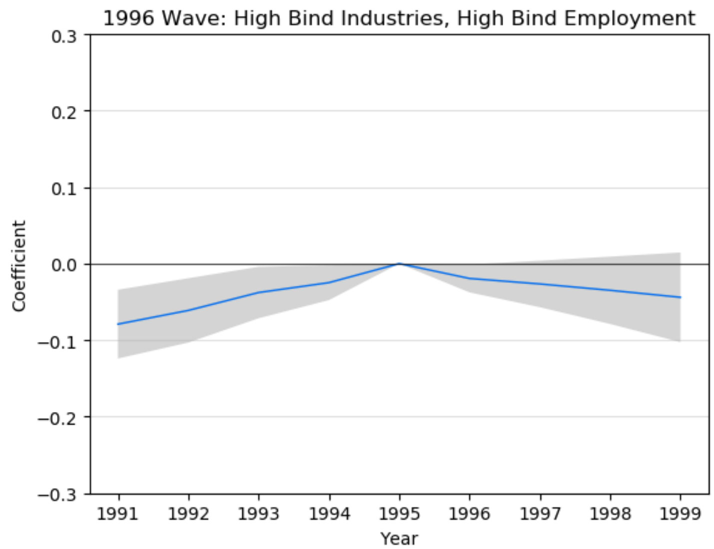
(a) No Effect in Longer Horizon



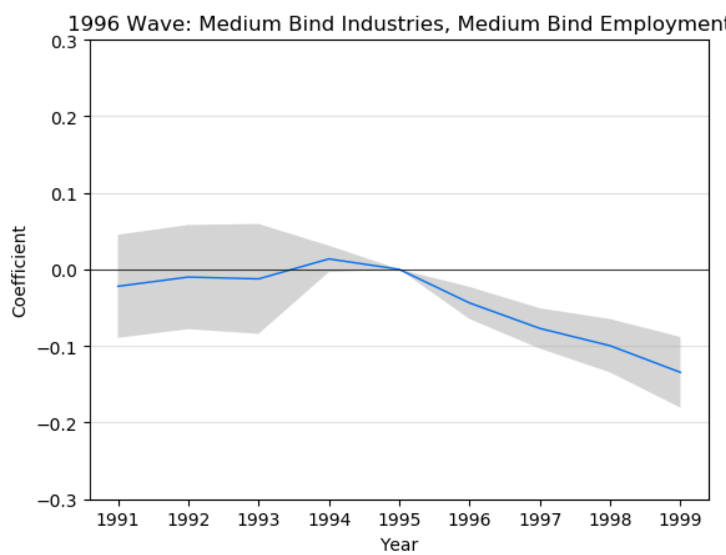
(b) Negative Effects in Longer Horizon

Figure 2: Short-Run DiD Estimates of the Effect on Employment Growth

Notes. Figure 2 shows how the negative effect of the minimum wage in the longer horizon biases the estimates upward. State A increases its minimum wage in t_1 and state B in t_2. Consider t_2 when state B increases its minimum wage. The employment growth rate decreases from b_2 to b_3. If the minimum wage has only short-run effect (the effect lasts for one period), state A is a valid control in the sense that the parallel trend assumption is satisfied, as in (a). The treatment effect is $(b_3 - b_2) - (a_3 - a_2) = b_3 - b_2$. If the minimum wage has negative effects in the longer horizon (the effects last for more than one periods), the parallel trend assumption at t_2 is violated, the estimated treatment effect is $(b_3 - b_2) - (a_3 - a_2) > b_3 - b_2$ since $a_3 - a_2 < 0$. It shows that the estimate is biased upward.



(a) H.B. Industries, H.B. Employment



(b) M.B. Industries, M.B. Employment

Figure 3: Minimum Wage's Employment Effect: The 1996 Wave

Notes. Figure 3 shows the estimates of equation (19). In (a), the outcome variable is the employment growth rates of the high-binding industries, and the covariate is the exposure measure to the high-binding industries. In (b), the outcome variable is the employment growth rates of the medium-binding industries, and the covariate is the exposure measure to the medium-binding industries. The standard error is clustered at the state level.

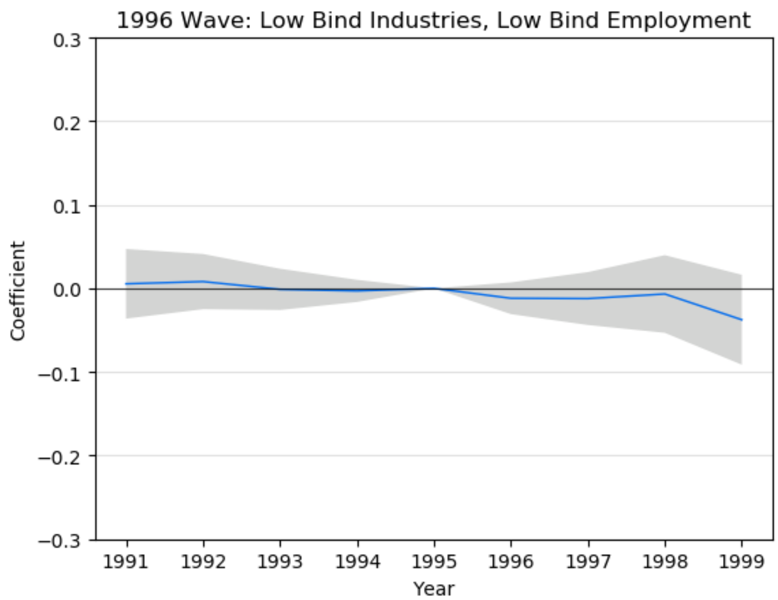
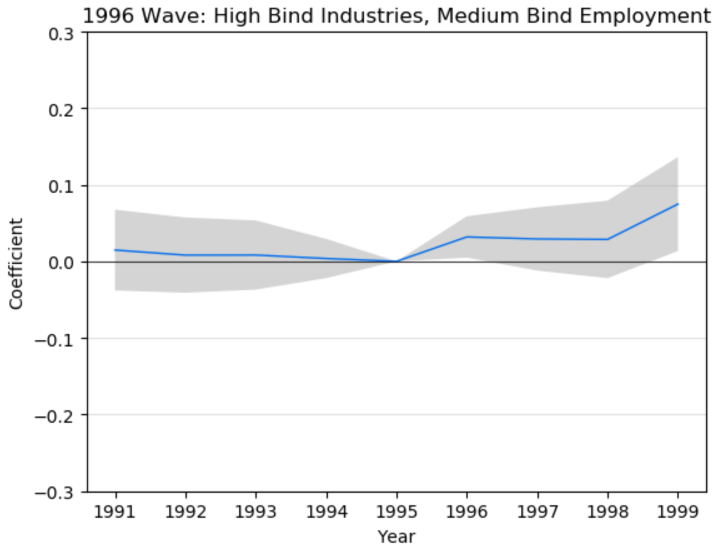
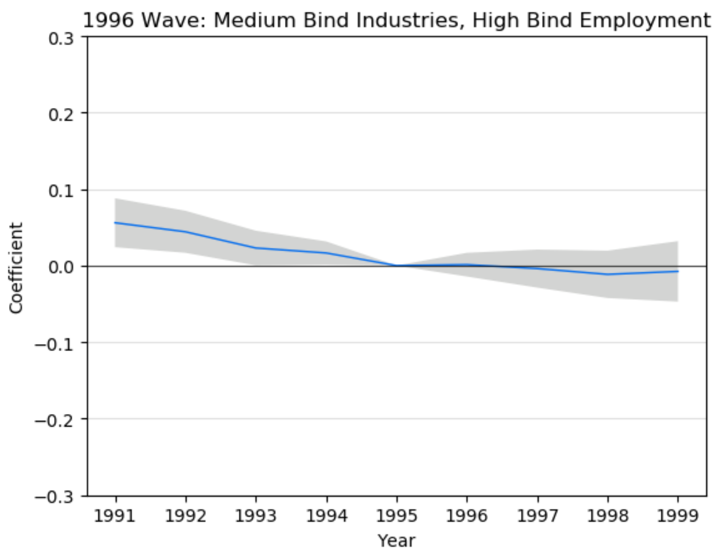


Figure 4: 1996 Wave: Low-Binding Industries

Notes. Figure 4 shows the estimates of equation (19) using the employment growth rates of the low-binding industries as the outcome variable and the exposure measure to the low-binding industries as the explanatory variable. Two out of the nine coefficients are significant, indicating that there is little evidence that there are other shocks correlated with the exposure measure that affect employment growth rates.



(a) H.B. Industries, M.B. Employment



(b) M.B. Industries, H.B. Employment

Figure 5: Minimum Wage's Employment Effect: The 1996 Wave

Notes. Figure 5 shows the effect of exposure to high(medium)-binding industries on the employment growth of medium(high)-binding industries. In particular, (a) regresses the employment growth rates of the medium-binding industries on the exposure to the high-binding industries. (b) does the reverse. There is little evidence that there are factors correlated with the exposure measure that drives the decline in the employment growth rates, in which case we would expect the estimates to be negative after the minimum wage increase.

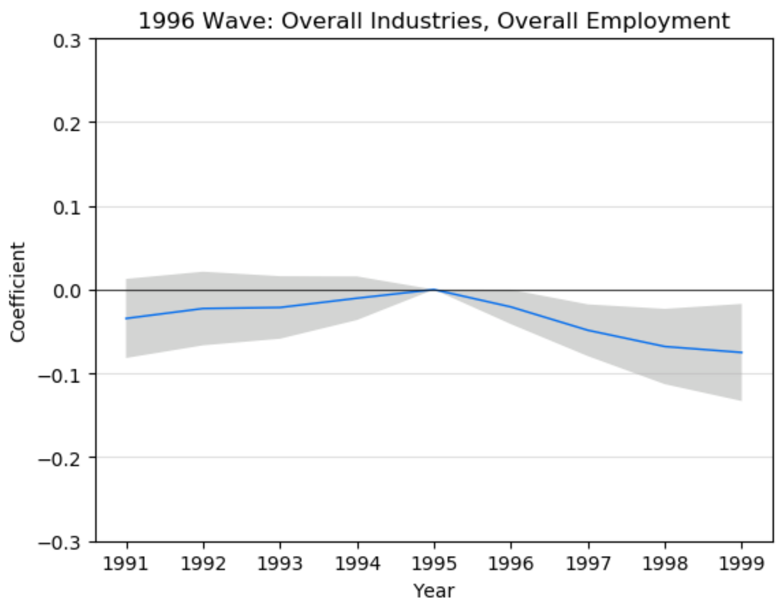
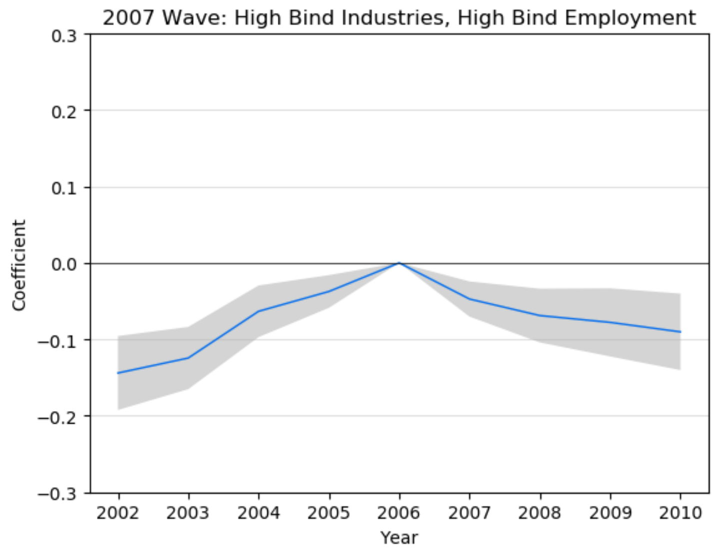
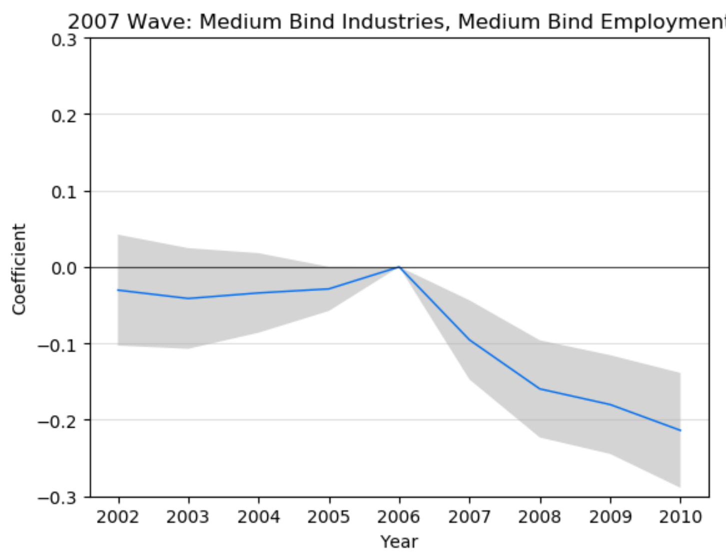


Figure 6: 1996 Wave: High- and Medium Binding Industries

Notes. Figure 6 plots the estimates of the exposure measure to the high- and medium-binding industries on the employment growth rates of the high- and medium-binding industries. The results are almost the un-weighted averages of the separate results in figure 3. This is expected since there is no significant correlation between the high- and medium-binding exposure measure.



(a) H.B. Industries, H.B. Employment



(b) M.B. Industries, M.B. Employment

Figure 7: Minimum Wage's Employment Effect: The 2007 Wave

Notes. Figure 7 plots the estimates using equation (19) and the 2007 wave of minimum wage increases. The results are similar to the ones in figure 3.

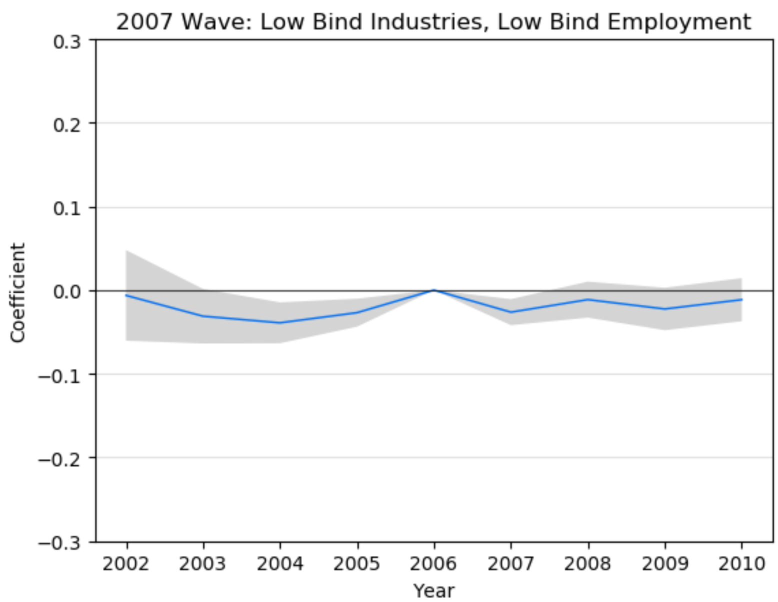
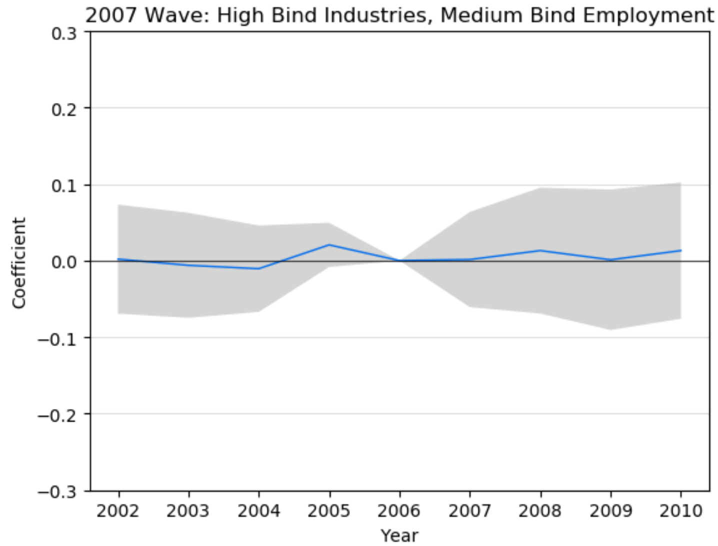
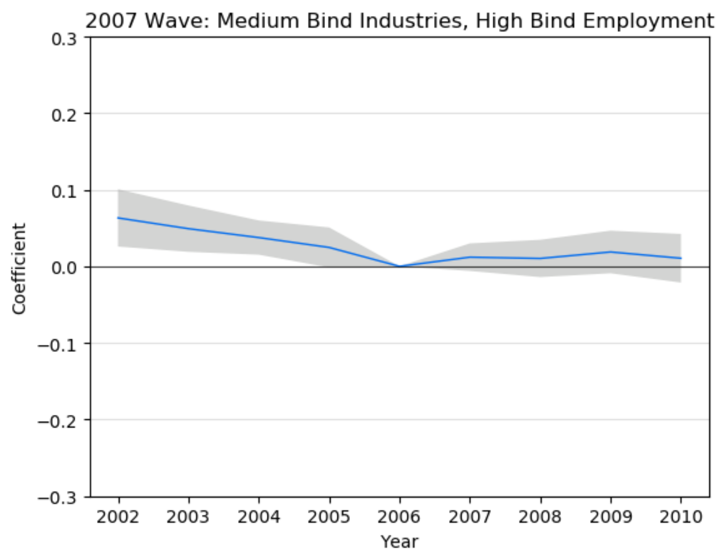


Figure 8: 2007 Wave: Low-Binding Industries

Notes. Figure 8 shows the similar results of figure 4 with the 2007 wave of minimum wage increases. There are more evidence of the existence of other shocks that correlate with the exposure measure.



(a) H.B. Industries, M.B. Employment



(b) M.B. Industries, H.B. Employment

Figure 9: Minimum Wage's Employment Effect: The 2007 Wave

Notes. Figure 9 shows the effect of exposure to high(medium)-binding industries on the employment growth of medium(high)-binding industries. In particular, (a) regresses the employment growth rates of the medium-binding industries on the exposure to the high-binding industries. (b) does the reverse. Overall, there is little evidence that there are factors correlated with the exposure measure that drives the decline in the employment growth rates, similar to figure 5.

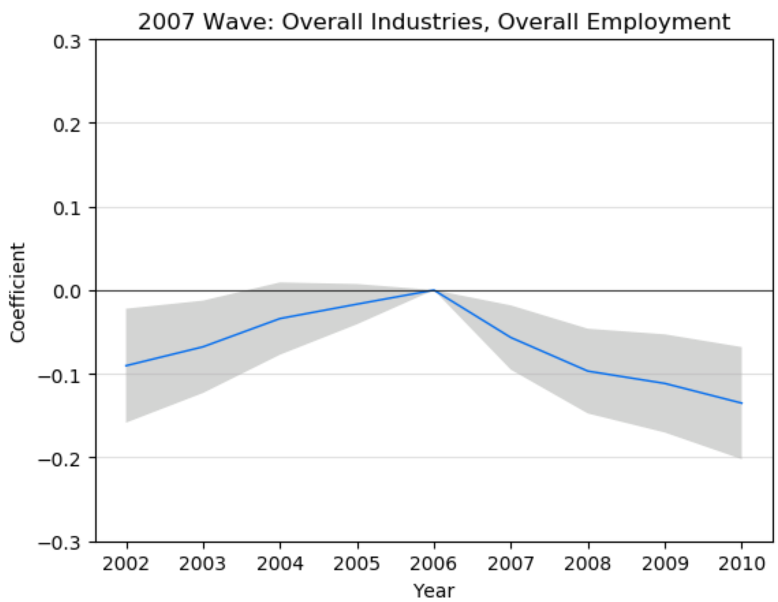


Figure 10: 2007 Wave: High- and Medium Binding Industries

Notes. Figure 10 plots the estimates of the exposure measure to the high- and medium-binding industries on the employment growth rates of the high- and medium-binding industries.

Appendices

A Data

A.1 Two-Way Fixed Effect Regression on the Employment Growth Rates

We use the two-way fixed effect regression to study the minimum wage's effect on the employment growth rates. Section 5 shows that the minimum wage has heterogeneous effects on the employment growth rates, depending on whether we study the high or the medium-binding industries. We follow the analysis in section 5 and construct three outcome variables: the high-binding industries' employment growth rates, the medium-binding industries' employment growth rates, and their combined employment growth rates. To compare with Meer and West (2016) who use the same data, we also study the effect on the employment growth rates of all industries. To be consistent with section 5, we include only employment growth rates within the $[-80\%, 80\%]$ range. We include all the industries without restrictions on the employment growth rates and the result is shown in column (5) in table A.1.

The regression specification is the following two-way fixed effect model with county and quarter-year fixed effects.

$$\Delta\text{Emp}_{ct} = \alpha + \beta \ln MW_{ct} + \lambda_c + \delta_t + \epsilon_{ct} \tag{A.1}$$

We also show the results where we add the state-specific time trend $\gamma_s \times t$.

Table A.1 shows that the minimum wage decreases the overall employment growth rates. A 10% minimum wage increase leads to a 0.1% decrease in the employment growth rates of all industries, which implies an elasticity of -0.01. After including the full sample, namely lifting the restriction on the employment growth rates, the result -0.05 is the same as in Meer and West (2016) who use a longer sample period and run the regression at the state level. Recall that section 5 shows that the minimum wage decreases the medium-binding industries' employment growth rates. In table A.1 we do not see a significant effect on the medium-binding industries' employment growth. One reason for the difference is that the event-study regression is mostly capturing the substitution effect while the two-way fixed effect is capturing both the labor demand effect and the substitution effect. It could be that the labor demand effect is small or positive, shown extensively in the literature (see e.g. Card and Krueger (1994); Cengiz et al. (2019); Lindner and Harasztosi (2019)).

Another reason for the difference could be that the estimates table A.1 are biased upwards, which was explained in section 2. Note that the results in table A.1 include data from all the states while section 5 only uses a subset of states. The states that do not satisfy our criterion are the states that change their minimum wage more often during these two waves. According to section 2, they potentially poses more issue when used as control states. Including the states that increase the minimum wage often in table A.1 could exacerbate the upward bias.

Table A.1 also shows the importance for restricting the sample. Comparing column (4) and (5), the number of observations only increases by 2231 or 0.7%, while the estimate increases by 300% in magnitude. The small fraction of extreme values are unlikely realistic representation of the economy, but impose large bias when we include them in the regression. Nonetheless, we run the same regression in the main section but varying the restrictions on the sample in section A.3. The results are consistent with the main results in section 5.

The comparison between the results in table A.1 and section 5 show the importance of the potential upward bias. The literature emphasizes the potential bias because the controls might not satisfy the parallel trend assumption (see e.g. Neumark (2018), Meer and West (2016)). We formalize the idea by constructing a structural model and use the event study frame to study the employment effect in the longer horizon. The framework allows us to show that for the medium-binding industries there is no significant pre-trend, and the employment growth rates decrease after the minimum wage increase. The result is both qualitatively and quantitatively consistent in the two waves of minimum wage increases.

A.2 Control Construction

We discuss the construction of the control variables in equation (19). Two important changes occurred in the US trade policy during our periods of interest. The first is the implementation of North American Free Trade Agreement (NAFTA). Introduced in 1994, NAFTA immediately eliminated tariff on nearly half of all manufactured products imported from Mexico and gradually lowered tariffs on the rest of the products between 1994 and 2004. The existing evidence suggests that NAFTA considerably lowered the wage growth in the US manufacturing sector (Hakobyan and McLaren (2016)).

The second is the US legislative action in 2001 to permanent designate China as a most favored nation (MFN). The permanent MFN status allowed Chinese manufacturing firms to specialize in exports to the US, which led to substantial increase in US import of Chinese

Table A.1: The Effect of Minimum Wages on Employment Growth Rate

	(1) High Binding	(2) Medium Binding	(3) Combined	(4) All Industries	(5) Full Sample
			No State-Specific Time Trend		
$\ln MW_{ct}$	-0.010	0.001	-0.004	-0.011**	-0.045*
	(0.008)	(0.005)	(0.006)	(0.006)	(0.026)
			With State-Specific Time Trend		
$\ln MW_{ct}$	-0.007	0.001	-0.002	-0.009	-0.047*
	(0.009)	(0.007)	(0.007)	(0.006)	(0.027)
N	281900	233662	283782	291229	293460
County FE	Y	Y	Y	Y	Y
Quarter-Year FE	Y	Y	Y	Y	Y

Notes. Table A.1 shows the estimates of the short-run effect of the minimum wage on employment growth, i.e. equation (A.1). The standard error is clustered at the state level. We use the real minimum wage deflated by regional price index. The first row shows the results without state-specific time trend.

goods. The additional trade volume led to employment loss (Autor et al. (2013); Pierce and Schott (2016)) and reallocation of jobs from manufacturing to service (Bloom et al. (2020)) in the US.

To control for the trade-related employment loss in our analysis, we include two additional variables that capture the degree of trade liberalization with Mexico and China, respectively. To control for trade liberalization with Mexico, we use NAFTA index, a shift-share measure of local exposure to competition against Mexican imports formulated by Hakobyan and McLaren (2016). The index is designed to use variation in local industry composition, similar to how our exposure measure is constructed, to discern which counties experienced a more significant negative shock due to import competition.

NAFTA index is defined as

$$l\tau_{it} = \frac{\sum_j^J w_{ij,1990} \times RCA_{j,1990} \times \tau_{jt}}{\sum_j^J w_{ij,1990} \times RCA_{j,1990}}$$

where $w_{ij,1990}$ is the employment count of industry j in county i in 1990, τ_{jt} is the weighted average of tariff rates on all products produced by the industry j in year t, and $RCA_{j,1990}$ is the relative competitiveness of Mexico in industry j. The employment count is constructed using County Business Patterns (CBP). The censored cells in CBP are imputed using David Dorn's program described in Autor et al. (2013). The tariff rates come from the Harmonized Schedule of Tariffs and we aggregate the product level tariffs by industry (Feenstra et al. (2002)). Lastly, the competitiveness of Mexican exports in industry j is measured as the ratio between Mexico's export volume and the worldwide trade volume. The inclusion of this variable ensures that tariff levied on an industry is weighted more heavily in the construction of NAFTA index ($l\tau_{it}$) if Mexico's export volume in the industry is greater. The trade data used to construct RCA comes from the UN Comtrade database.

We capture the local employment effect of Chinese imports to the US by constructing the following variable

$$ci_{it} = \frac{\sum j^J w_{ij} \times cv_{jt}}{\sum j^J w_{ij}}$$

where w_{ij} is the county-by-industry employment count, and cv_{jt} is the total value of goods imported from China to the US in industry j on year t. The data on employment and trade come from CBP and the UN Comtrade database, respectively. The variable, ci_{it}, therefore captures the county-level variation in the degree of import competition that US firms experience.

A.3 Robustness

A.3.1 Confounding Factors

Our local exposure measure is constructed so that it relates to the minimum wage. It could be correlated with other county time-varying characteristics that potentially confound the results. To address the potential violation of the identification assumption, we use time-varying county-level factors to see if there is other underlying changes outside of the minimum wage increase. If some time-varying unobservable causes a decrease in the employment growth rates, observable characteristics unrelated to the minimum wage are likely to change when the minimum wage increases. We choose a set of time-varying observables to discern evidence that there are reasons outside the minimum wage that could drive the results. In particular, we replace the employment growth rate with these variables in the event-study framework to find any evidence of significant changes that not caused by but coincide with the minimum wage increase.

A.3.2 Changing Sample Restriction

In our main results we restrict the employment growth rates to be between -80% and 80%. In the section, we choose various cutoffs and show that the main conclusions are robust to the changes. Table A.2 and table A.3 show the results using equation (19) and different sample selection based on the employment growth rates.

We first look at the 1996 wave. For the high-binding industries, the first column changes the cutoff so that the sample includes employment growth rates between $[-100\%, 100\%]$. The estimates for the first two years after the minimum wage are now significantly negative at the 5% level, compared to table 1. β_1 remains the same in both regressions, while β_2 decreases by 29%. The large decrease is driven by merely 90 observations (15326-15226). In table A.2, β_3 and β_4 are significant at the 10% level. Overall, including more observations lead to a decrease in the estimates. The decrease can be quite large: for β_{-2}, it decreases by 49% compared to the one in table 1. Even so, qualitatively and quantitatively, there are still significant negative pre-trend for the high-binding industries, and the decrease in the employment growth rates after the minimum wage increase is small compared to the medium-binding industries.

Changing the cutoff to $[-100\%, 120\%]$ further decreases the estimates substantially, which is only driven by 50 more observations. When we run the regression using the full sample, all estimates are significantly negative at the 5% level. The large decrease in the estimates is again the result of a small fraction of the total observation. However, compared to the baseline estimates for the medium-binding industries, the employment growth decline is moderate. In particular, the estimates after the minimum wage increase for the high-binding industries is consistently about 50% that of the estimates for the medium-binding industries.

As a comparison, the estimates for the medium-binding industries are robust to changes in cutoffs. Qualitatively, the results exhibit no pre-trend in table A.2 and the estimates are significantly negative after the minimum wage increase. Quantitatively, the change in estimates from the baseline to $[-100\%, 100\%]$ and to $[-100\%, 120\%]$ is moderate. Only until we include the full sample do we see relatively large decreases in the estimates, which is the consequence of a few hundreds observations.

For the low-binding industries, the results in table A.2 are similar to the baseline results. The estimates decrease after we include more outliers, but the decrease is on par with the decrease for the high-binding industries. The similar decrease implies that after subtracting the estimates for the low-binding industries from the estimates for the high-binding industries, the results would be close to zero, which is the same as in section 5. To that extent, the results in section 5 for the high-binding industries are robust to the

Table A.2: The Effect of Minimum Wages on Employment Growth Rate: The 1996 Wave

	1996 Wave								
	High-Binding			**Medium-Binding**			**Low-Binding**		
	(1) -100% to 100%	(2) -100% to 120%	(3) Full Sample	(1) -100% to 100%	(2) -100% to 120%	(3) Full Sample	(1) -100% to 100%	(2) -100% to 120%	(3) Full Sample
β_{-4}	-0.099***	-0.099***	-0.102***	-0.022	-0.022	-0.029	0.002	0.001	0.002
	(0.023)	(0.023)	(0.023)	(0.034)	(0.034)	(0.035)	(0.022)	(0.022)	(0.022)
β_{-3}	-0.077***	-0.078***	-0.082***	-0.010	-0.011	-0.015	0.008	0.006	0.009
	(0.022)	(0.022)	(0.022)	(0.035)	(0.035)	(0.035)	(0.017)	(0.017)	(0.017)
β_{-2}	-0.054***	-0.056***	-0.057***	-0.013	-0.013	-0.016	-0.001	-0.004	-0.003
	(0.020)	(0.020)	(0.020)	(0.037)	(0.037)	(0.036)	(0.013)	(0.013)	(0.013)
β_{-1}	-0.025**	-0.026**	-0.030**	0.014	0.014	0.012	-0.004	-0.006	-0.007
	(0.012)	(0.012)	(0.012)	(0.009)	(0.009)	(0.009)	(0.007)	(0.008)	(0.008)
β_{1}	-0.021**	-0.021**	-0.024**	-0.045***	-0.047***	-0.052***	-0.012	-0.013	-0.017*
	(0.010)	(0.009)	(0.009)	(0.011)	(0.011)	(0.011)	(0.010)	(0.010)	(0.010)
β_{2}	-0.035**	-0.037**	-0.045***	-0.083***	-0.088***	-0.098***	-0.013	-0.014	-0.023
	(0.016)	(0.016)	(0.016)	(0.014)	(0.014)	(0.014)	(0.016)	(0.016)	(0.016)
β_{3}	-0.042*	-0.048**	-0.065***	-0.107***	-0.111***	-0.137***	-0.008	-0.011	-0.025
	(0.023)	(0.023)	(0.023)	(0.018)	(0.018)	(0.019)	(0.024)	(0.024)	(0.024)
β_{4}	-0.050*	-0.061**	-0.083***	-0.141***	-0.153***	-0.195***	-0.041	-0.043	-0.068**
	(0.030)	(0.030)	(0.030)	(0.023)	(0.024)	(0.025)	(0.027)	(0.027)	(0.028)
Observations	15316	15367	15613	10845	10899	11209	12129	12168	12469

Notes. Table A.2 shows the estimates of the effect of the minimum wage on employment growth with different cutoffs. The regression is equation (19). The standard error is clustered at the county level. * means significant at the 10% level, ** 5% level, and *** 1% level.

Table A.3: The Effect of Minimum Wages on Employment Growth Rate: The 2007 Wave

	2007 Wave								
	High-Binding			**Medium-Binding**			**Low-Binding**		
	(1) -100% to 100%	(2) -100% to 120%	(3) Full Sample	(1) -100% to 100%	(2) -100% to 120%	(3) Full Sample	(1) -100% to 100%	(2) -100% to 120%	(3) Full Sample
β_{-4}	-0.153***	-0.157***	-0.180***	-0.104**	-0.110**	-0.141**	-0.011	-0.013	-0.032
	(0.025)	(0.025)	(0.028)	(0.049)	(0.049)	(0.051)	(0.028)	(0.027)	(0.028)
β_{-3}	-0.133***	-0.136***	-0.151***	-0.044	-0.048	-0.071**	-0.035**	-0.039**	-0.051***
	(0.021)	(0.021)	(0.024)	(0.034)	(0.034)	(0.033)	(0.017)	(0.017)	(0.018)
β_{-2}	-0.070***	-0.073***	-0.089***	-0.039	-0.041	-0.055**	-0.044***	-0.045***	-0.055***
	(0.018)	(0.018)	(0.021)	(0.027)	(0.027)	(0.026)	(0.013)	(0.013)	(0.013)
β_{-1}	-0.039***	-0.040***	-0.048***	-0.031**	-0.033**	-0.038**	-0.029***	-0.029***	-0.035***
	(0.011)	(0.011)	(0.012)	(0.015)	(0.015)	(0.015)	(0.009)	(0.008)	(0.009)
β_1	-0.049***	-0.051***	-0.052***	-0.099***	-0.101***	-0.108***	-0.027***	-0.027***	-0.034***
	(0.012)	(0.012)	(0.012)	(0.026)	(0.026)	(0.026)	(0.008)	(0.008)	(0.008)
β_2	-0.070***	-0.074***	-0.080***	-0.165***	-0.167***	-0.182***	-0.011	-0.013	-0.022*
	(0.018)	(0.018)	(0.019)	(0.033)	(0.033)	(0.032)	(0.012)	(0.012)	(0.012)
β_3	-0.081***	-0.083***	-0.095***	-0.189***	-0.191***	-0.210***	-0.021	-0.024*	-0.035**
	(0.023)	(0.023)	(0.024)	(0.032)	(0.032)	(0.032)	(0.013)	(0.013)	(0.014)
β_4	-0.098***	-0.104***	-0.115***	-0.221***	-0.225***	-0.245***	-0.014	-0.016	-0.028*
	(0.026)	(0.026)	(0.027)	(0.037)	(0.037)	(0.037)	(0.013)	(0.014)	(0.015)
Observations	16646	16725	16974	14787	14883	15606	15701	15835	16605

Notes. Table A.3 shows the estimates of the effect of the minimum wage on employment growth with different cutoffs. The regression is equation (19). The standard error is clustered at the county level. * means significant at the 10% level, ** 5% level, and *** 1% level.

changes in the cutoffs.

Turning to the 2007 wave, we focus on the medium-binding industries. In the baseline regression, the only significant pre-trend is the year before the base year for the medium binding industries. However, when we include employment growth rates between [-100%, 100%], β_{-4} becomes significant at the 10% level. In particular, β_4 decreases by about 320%, driven by only 250 or 1.7% of the observations. The comparison highlights the necessity for restricting the sample employment growth rates. The extreme outliers contribute to the large decrease in the estimates. While we do not know what causes the extreme values in the employment growth rates, we think it is unlikely an accurate reflection of the economy during the 2007 wave. To the extent that we would like to estimate the average treatment effect, excluding the outliers might make the estimates more informative.

For the high- and low-binding industries, the qualitative results remain the same as in the baseline regression. Expanding the sample decreases the estimates for both the high- and low-binding industries. As a result, their differences remain relatively unchanged.

B Proofs

Proof of proposition 1.

Proof. Let the employment share of low-skill workers in the production of $X(i)$ be given by

$$s_L(i) = (1-\theta^c(i))\left(\frac{M}{\gamma_L(i)P_X(i)}\right)1-\zeta \tag{B.1}$$

In the equilibrium, the wage bill equation gives

$$\begin{aligned} ML(i) &= \eta\alpha(i)s_L(i)P_Y(i)Y^g(i) \\ &= \eta\alpha(i)s_L(i)P_Y(i)Y(i)\frac{Y^g(i)}{Y(i)} \\ &= \frac{1}{2-\eta}\alpha(i)s_L(i)v_iP_Y(i)^{1-\sigma}Y \\ &= \frac{1}{2-\eta}\alpha(i)(1-\theta^c(i))M^{1-\zeta}\gamma_L(i)^{-(1-\zeta)}P_X(i)^{-(1-\zeta)}v_iP_Y(i)^{1-\sigma}Y \end{aligned} \tag{B.2}$$

Dividing both sides by the minimum wage M and defining $\epsilon(i) = 1-\alpha(i)(1-\sigma)/(1-\zeta)$, the labor demand for low-skilled workers is

$$L(i) = \frac{1}{2-\eta}\alpha(i)\gamma_L(i)^{-(1-\zeta)}\lambda(i)^{1-\sigma}(1-\theta^c(i))M^{-\zeta}P_X(i)^{-(1-\zeta)\epsilon(i)}W^{(1-\sigma)(1-\alpha(i))}v_iY \tag{B.3}$$

Take the log and differentiate equation (B.3), we have, for a single industry i, the change in labor demand is

$$d\ln L(i) = -\frac{d\theta^c(i)}{1-\theta^c(i)} - \zeta\ln M - (1-\zeta)\epsilon(i)d\ln P_X(i) + (1-\sigma)(1-\alpha(i))d\ln W + d\ln Y \tag{B.4}$$

Aggregating equation (B.4) and denoting the employment share of low-skill workers in industry i by $l(i)$, we arrive at equation (13). □

Proof of proposition 2.

Proof. To ensure that $r^*(M)$ is increasing in M, we need to show that equation (10) displays supermodularity. Taking the derivative with respect to M and r, it is easy to see that it is positive. Hence $\partial r^*(M)/\partial M \geqslant 0$. □

Proof of proposition 3.

Proof. Let χ_i denote the value-added share of industry i. Let s_L denote the average labor share. It relates to the industry specific labor share $s_L(i)$ by

$$s_L l(i) = s_L(i)\chi_i \tag{B.5}$$

Taking the log of equation (7) and differentiating, the change in the price of X is

$$d\ln P_X(i) = (1 - s_L(i))d\ln R + s_L(i)d\ln M - s_L(i)\frac{\pi(i)l(i)}{\chi_i}\frac{d\theta^c(i)}{1-\theta^c(i)} \tag{B.6}$$

$\pi(i)$ is the cost savings from capital given by

$$\pi(i) = \frac{1}{1-\zeta}\left[1 - \left(\frac{A_L R_t}{M_t^{1/\gamma}}\right)^{1-\zeta}\right] \tag{B.7}$$

The ideal price index implies that

$$0 = \sum \chi_i \alpha(i) d\ln P_X(i) + \sum (1-\alpha(i))d\ln W \tag{B.8}$$

Plugging equation (B.8) into equation (B.6), we arrive at the following equation

$$(1-s_L)d\ln R + s_L d\ln M - \sum \chi_i(1-\alpha(i))d\ln W = \pi s_L \sum l(i)\frac{d\theta^c(i)}{1-\theta^c(i)} \tag{B.9}$$

By equation (6), we have

$$d\ln R = \xi d\ln K \tag{B.10}$$

Equation (2) implies that the aggregate demand for high-skill labor is

$$SW = (1-\alpha)Y$$

Differentiating this equation and noting that S is fixed, we have

$$d\ln W = d\ln Y \tag{B.11}$$

Lastly, the capital share satisfies

$$RK = \frac{1}{2-\eta}\sum \alpha(i)(1-s_L(i))v_i P_Y(i)^{1-\sigma}Y$$

Taking the log, differentiating, and plugging in equations (7) and (11), it gives us

$$
\begin{aligned}
d\ln K =& \left(\frac{s_L}{1-s_L} + [(1-\zeta)-\alpha(1-\sigma)]\pi s_L\right)\sum l(i)\frac{d\theta^c(i)}{1-\theta^c(i)} \\
& - (\zeta + (1-s_L)[(1-\zeta)-\alpha(1-\sigma)])d\ln R \\
& - [(1-\zeta)-\alpha(1-\sigma)]s_L d\ln M + (1-\sigma)(1-\alpha)d\ln W + d\ln Y
\end{aligned} \tag{B.12}
$$

Together there are five equations equations (B.9) to (13) and five unknowns $\{d\ln L, d\ln R, d\ln W, d\ln K, d\ln Y\}$. Solving for $d\ln L$, we arrive at equation (17). The coefficients A and B are given

$$
\begin{aligned}
A &= -(1+\pi s_L\Lambda_0) + \Lambda_0(1-s_L)\frac{\xi}{\Lambda_1}\Lambda_2 + \left(\frac{\Lambda_0(1-s_L)\xi\Lambda_4}{\Lambda_1} + \Lambda_4\right)\frac{\Lambda_6}{\Lambda_5} \\
B &= \Lambda_0 s_L - \xi + \Lambda_0(1-s_L)\xi\frac{\Lambda_3}{\Lambda_1} + \left(\frac{\Lambda(1-s_L)\xi\Lambda_4}{\Lambda_1} + \Lambda_4\right)\frac{\Lambda_7}{\Lambda_5}
\end{aligned} \tag{B.13}
$$

in which $\{\Lambda_i\}_{i=0,1,\ldots,7}$ are given by

$$
\begin{aligned}
\Lambda_0 &= (1-\sigma)\alpha - (1-\zeta) & \Lambda_1 &= 1-(1-s_L)\Lambda_0 + \xi \\
\Lambda_2 &= \frac{s_L}{1-s_L} - \Lambda_0\pi s_L & \Lambda_3 &= s_L\Lambda_0 \\
\Lambda_4 &= (1-\sigma)(1-\alpha) + 1 & \Lambda_5 &= (1-\alpha) - \frac{1-s_L}{\Lambda_1}\xi\Lambda_4 \\
\Lambda_6 &= -\pi s_L + \frac{1-s_L}{\Lambda_1}\xi\Lambda_2 & \Lambda_7 &= \frac{1-s_L}{\Lambda_1}\xi\Lambda_3 + s_L
\end{aligned} \tag{B.14}
$$

□