ISSN 1936-5330

Research Working Papers

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November 2022; updated May 2025
RWP 22-13
http://doi.org/10.18651/RWP2022-13
This paper supersedes the old version:
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FEDERAL RESERVE BANK OF KANSAS CITY

The economic effects of a rapid increase in the minimum wage: evidence from South Korea experiments*

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May 15, 2025

Abstract

South Korea's minimum wage rapidly rose from 53 percent of the median wage to 63 percent between 2017 and 2019. While the minimum wage has been increasing steadily over decades, the rapid pace in 2018-19 was largely unexpected and driven by a sudden shift in the political environment. We study the economic effects of this minimum wage hike on employment, wages, and labor productivity using South Korean manufacturing firm data, which comprehensively covers all manufacturing firms with 10 or more employees. We find a significant negative employment (3% decline in domestic employment) effect of the minimum wage hike for 2018-19 compared with its modest increase in 2015-17, as the fraction of firms exposed to the minimum wage shock substantially increased and these firms adjusted to the shock through both intensive margins (layoffs) and extensive margins (plant closings). Adjustments through extensive margins jumped up at firms whose exposure to the minimum wage hike is in the right tail. The negative employment effect is also confirmed in our supplementary analysis using population-level administrative data on firms in the service sector. At the same time, labor productivity and wages increased more for manufacturing firms with greater exposure to the minimum wage. Our empirical findings are consistent with a task-based production model with firm-level heterogeneity emphasizing the substitution between low skill labor and other production factors.

Keywords: Minimum Wage; Employment Effect; Substitution of Production Factors
JEL Classifications: J23, J31, J42

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* The opinions expressed herein are those of the authors and do not reflect the views of the Federal Reserve Bank of Kansas City, or Federal Reserve System as well as the Korea Institute of Public Finance. The empirical analysis in this paper is based on a Korea Institute Public Finance technical report (Lee and Song (2021)). The authors thank the Korea Institute of Public Finance for obtaining the dataset. We are grateful to participants in seminars at the Federal Reserve Bank of Kansas City and Federal Reserve Board as well as conference participants in the 2022 Midwest Econometrics Meeting, the 2023 KAEA-KIEP Workshop, the 2023 KLEA Summer Conference, the 2023 KER International Conference for helpful comments. All the remaining errors are ours.

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1. Introduction

The employment effect of the minimum wage increase remains an active area of debate among economists and is often associated with different views on the degree of competition in the labor market (e.g., see Card and Krueger (2015) and Neumark et al. (2008) for conflicting views). A definite conclusion on the magnitude of the employment effect remains elusive as summarized by Manning (2021), with some studies finding even a small positive employment effect and others finding a negative effect of varying degrees. This lack of consensus contrasts with the positive wage effect, which is agreeable in most studies. However, the employment effect of the minimum wage increase can shift substantially depending on the magnitude and timing of the increase, especially when it is not expected. To empirically test this turning point hypothesis of the minimum wage on employment, it would be useful to have comprehensive micro-level data which covers both the period of a modest increase in the minimum wage and the period of a rapid (and substantial) increase in the minimum wage.

Recent data on the South Korean labor market from 2015 to 2019 provide this ideal ground to test the hypothesis. As Figure 1 shows, the national minimum wage in South Korea increased rapidly and substantially during 2018 and 2019 by almost 30% over the two-year period. The pace of the increase stands in stark contrast to the previous two-year period (2015 and 2016) when the increase was modest. In addition, as shown in Figure 2, the minimum wage outpaced the increase in the typical wage as indicated by the Kaitz index, which measures the level of the minimum wage relative to the median wage. This sudden shift in the pace of the minimum wage increase was mainly driven by an unexpected change in the political environment, which can be regarded, with some confidence, as largely exogenous to economic factors.[1]

In this paper, we use the firm level data on employment and wages to empirically test whether this rapid increase in the minimum wage brought about a significant negative employment effect, which was not evident during the period of a modest increase. We focus on the mining and manufacturing sector, for which we have more comprehensive coverage on firm-level employment and wages from the annual Mining and Manufacturing Survey (MMS) than other sectors. However,

[1] In 2017, the conservative leaning president was impeached over a personal scandal and a center-left candidate won the presidential election that happened one year earlier than usual. A substantial hike in the minimum wage was a part of the incoming administration's "income-driven growth" plan. However, a conservative candidate won the presidential election in the March of 2022, partly by criticizing the negative economic effect of the minimum wage hike. We will provide more background information on the political environment in the subsequent empirical analysis.

to ensure that sector-specific factors do not drive our main results, we supplement our analysis using other data sources covering the service sector. For our main empirical strategy to identify the effects of the minimum wage hike, we use a difference-in-differences regression-type approach by constructing the firm level exposure to the minimum wage increase from wage data and regressing the employment change before and after the rapid increase in the minimum wage on this exposure measure. Specifically, we compare the two-year change in employment from 2017 to 2019 to the preceding two-year change from 2015 to 2017, using the earlier period as a placebo test to help isolate the treatment effect of the 2018–2019 minimum wage hikes.

Our main empirical findings are summarized as follows. First, we find evidence for a significant and negative employment effect from the rapid increase in the minimum wage. Our baseline estimates find a one standard deviation increase in the minimum wage exposure for a firm in the manufacturing sector is expected to decrease employment by about 4%. With the average exposure in the MMS sample rising by about three-quarters to of the standard deviation during the 2017-2019 period, our findings imply that total (domestic) employment declined by 3% due to the minimum wage hike. This substantial dis-employment was driven by two factors. On the one hand, the significant increase in the minimum wage made more firms exposed to this policy shift. On the other hand, among firms whose exposure to the minimum wage was similar between pre- and post-2017 period, the highly exposure group made more visible employment adjustments such as large layoffs and plant closings because they faced more intensive wage pressures in the post-2017 period than in the pre-2017 period. This finding suggests that the absence of the negative employment effect from a modest increase in the minimum wage could not be extrapolated to a large increase in the minimum wage, supporting the turning point hypothesis on the employment effect.

Second, when we decompose the employment adjustment into intensive (layoffs) and extensive (plant closings) margins, the extensive margin is estimated to contribute about 38.1% to the overall adjustment. The adjustment through the extensive margin increases is more pronounced at firms whose minimum wage exposure is at the right tail. Also, among those firms who dropped out of the MMS sample after the minimum wage hike because the employment was reduced below 10 people, larger firms were more likely to adjust through extensive margins.

Third, we observe a significant increase in the average wage and labor cost for firms highly exposed to the minimum wage hikes. We also find that the wage dispersion declined more with the rapid increase of the minimum wage. In addition, labor productivity also showed a significant sensitivity to minimum wage exposure during 2017-2019. Interestingly, labor productivity is higher in firms that are more exposed to the minimum wage increase, suggesting that the substitution of low-skill labor for high-skill labor or other inputs like capital might have played a role.

We introduce a task-based production model in which domestic low skill labor subject to the minimum wage can be substituted by other production factors such as domestic high skill labor, foreign low skill labor, and capital. With the two-dimensional firm level heterogeneity in 1) the comparative advantage of domestic low skill labor over domestic high skill labor in performing a variety of tasks and 2) the adjustment cost of offshoring tasks, the model can generate predictions consistent with our empirical findings. We highlight the fact that when the adjustment cost is not uniformly distributed, a rapid increase in the minimum wage can generate a sudden jump in the number of firms who make adjustments through extensive margins. Overall, our analysis supports the significant role of the substitution channel among domestic low skill labor subject to the minimum wage and other production factors in determining the economic effect of the minimum wage increase. In particular, our findings suggest that considering various margins of employment adjustment and firm-level heterogeneity is important for understanding full implications of the minimum wage increase.

Related Literature

Our paper is closely related to the empirical analysis of employment effect of the nationwide minimum wage using firm-level data (e.g., Harasztosi and Lindner (2019), Dustmann et al. (2022), Engbom and Moser (2022), and Haanwinckel (2023)). Like South Korea, the Hungarian minimum wage studied in Harasztosi and Linder (2019) rose substantially both in terms of the level and the relative share to the median wage between 2000 and 2002. In line with our results, they document a significant negative impact on employment during this period of sharp increases, in contrast to the more modest impact observed from the earlier 1998–2000 phase.[2] They show that the degree of the competition in the product market is a main factor determining the negative employment effect of the minimum wage hike. Manufacturing firms exposed to a higher degree of competition in the

[2] They argue that the relationship between the firm-level minimum wage exposure and the change in employment is roughly linear and does not depend on the size of the minimum wage increase. In contrast, our analysis suggests that while the relationship appears to be linear for most firms, among those most highly exposed to the minimum wage, the size of employment adjustments exceeds what would be predicted by the linear relationship. The finding supports the view that the employment effect may depend nonlinear on the size of the minimum wage hike at least for highly exposed firms.

product market chose to reduce employment much more than firms in the non-tradable service sector who could pass through the minimum wage shock to consumers using their market power. In their Table 6, firms in the tradable manufacturing sector show 10 times bigger dis-employment effect than firms in the non-tradable service sector. In contrast, we do not observe such a stark difference between the manufacturing sector and the service sector in the magnitude of the employment effect. Instead, we emphasize the substitution channel in the factor market. When we look at the firm-level labor productivity data, it goes up more in firms with higher exposure to the minimum wage hike, suggesting the substitution of low skill labor by other production factors might be an important factor in determining the employment effect of the minimum wage hike.[3]

Engbom and Moser (2022) examine the economic effect of the federal minimum wage hike in Brazil using firm-level data during 1996-2012. They find that the increase in the minimum wage reduces wage dispersion but has little effect on employment as the point estimate is not significantly different from zero due to a large standard error. The Kaitz index in Brazil has increased rapidly between 2004 and 2007 by more than 10 percentage points (see Saltiel and Urzúa (2022)). For other periods, the pace of the minimum wage hike was rather gradual. Not disentangling the period of the rapid increase from the period of the modest increase may add significant noises to the estimate of dis-employment effect of the rapid increase in the minimum wage hike.

Dustmann et al. (2022) study German micro data after the introduction of the nationwide minimum wage in 2015. Unlike other studies including our paper, they use an employee-employer matched dataset and can directly test the reallocation hypothesis that the minimum wage hike pushes low wage workers to more productive firms. They find evidence supporting the reallocation hypothesis and argue that the introduction of the minimum wage raised wages without reducing employment. Since we do not have the comparable data that can match employees with employers, we are not able to test the reallocation hypothesis for South Korea in the same way as Dustmann et al. (2022) do. However, we do not find evidence of reallocation from smaller to larger firms, as Dustmann et al. (2022) find in the German context. Our data show plant closing in larger firms exposed to the minimum wage shock happened at least as frequently as smaller firms. If the reallocation hypothesis were right, we would have seen less frequent plant closings in larger firms. Some reasons behind this finding are the job subsidy program limited to smaller sized firms

[3] In a frictional labor market, a higher minimum wage can push up the threshold for productivity in a successful match as shown in Kudlyak et al. (2022). But it is difficult to test this hypothesis directly without using an employee-employer matched dataset as used in Dustmann et al. (2022).

(employees of 30 or fewer) and larger firms' capability of offshoring as we detail later. Another reason that we find different results from Dustmann et al. (2022) is that the German minimum wage in 2015 was not substantially higher than the prevailing wage level of low skill workers at that time. The Kaitz index based on the 2015 German minimum wage was 0.47, which is even lower than 0.53, the Kaitz index of South Korea in 2017 before the rapid increase in the minimum wage (Doh and Van der Meer (2023)).

While the federal minimum wage in the U.S. has been flat for a long time, there have been minimum wage increases above the federal minimum wage in many states. Using 138 prominent state-level minimum wage changes between 1979 and 2016, Cengiz et al. (2019) find that the overall employment effect is small while average earnings of workers increase due to wage spillovers at the bottom of the wage distribution. They do not find a big role for skilled labor-unskilled labor substitution. In contrast, we find empirical support for the labor-labor (or labor-capital) substitution channel as the national minimum wage surpasses 60% of the median wage in South Korea.[4] Leveraging local minimum wage increases in the Seattle area, Jardim et al. (2022) investigate the effects of minimum wage hikes on low-wage employment using longitudinal data on individual workers. They find that hours worked increased for more experienced workers, whereas less experienced workers saw a reduction in hours. This finding is consistent with the skilled labor–unskilled labor substitution channel emphasized in our analysis.

Ahn, Choi and Chung (2023) study the employment adjustment of multi-national enterprises in response to the 2017-2019 minimum wage hikes in South Korea. They find that multi-national enterprises increased employment in low-wage countries such as Vietnam after the minimum wage in South Korea rapidly increased. This finding is consistent with the negative employment effect observed in our analysis. However, their sample of multi-national enterprises include only firms with more than 50 employees and is significantly narrower than the data covered in our sample. In addition, they mostly focus on the difference in the employment adjustment between multinational and non-multi-national enterprises, without directly examining the employment effects on non-multinational firms, which is an important consideration for assessing the overall impact of the policy. Our analysis complements their study by using a more comprehensive firm-level dataset.

[4] Clemens and Strain (2021) argue that large minimum wage changes had significant negative employment effects on workers most likely to be exposed to the minimum wage hike (young adults with less than a high school degree) using state-level variations in the U.S. during 2011-2019. Their results provide caution against extrapolating the employment effect from a small change in the minimum wage to a large change. Our findings align with their findings.

To understand our empirical findings in a consistent model framework, our paper extends the model in Eckardt (2022) who considers a task-based production function with a binding minimum wage for low skill workers but no heterogeneity in production technology at the firm level by introducing firm-level heterogeneity. Our task-based model is related to Haanwinckel (2023), who proposes a task-based production model with heterogeneities in both firms and workers. While this model is using a more generalized production function form and has a richer pattern of heterogeneity, it does not incorporate the adjustment cost term which can be useful for generating the significant and negative employment effect from the rapid increase in the minimum wage through the offshoring channel.[5]

The remaining part of the paper is organized as follows. Section 2 describes institutional background for the South Korean minimum wage commission and data sources. Section 3 explains the empirical framework adopted in this paper and provides results for the effects of the minimum wage hike with respect to employment, wage, and labor productivity. Section 4 discusses implications for policy and alternative models of the labor market. Section 5 concludes.

2 Institutional Background and Data

2.1 Institutional Background

In South Korea, the Minimum Wage Act was enacted in December 1986, and the Minimum Wage System was first introduced in January 1988 after the transition to democracy from an authoritarian regime. The hourly minimum wage rate is set every year by the Minimum Wage Commission which consists of 27 members with three groups of nine, each representing employees, employers, and public interest. The minimum wage law covers all businesses or workplaces with one employee or more regardless of employment status and nationality. Employers who violate the minimum wage law are subject to civil money penalty of up to about 20 million Korean Won (about $15400 at the exchange rate of June 2022) and criminal penalty of up to three years in prison. Minimum wage exceptions apply to businesses employing only family members, domestic service users, or seafarers.

[5] In our model, firms can always reorganize production in response to a minimum wage shock after paying adjustment costs unlike Hurst et al. (2022). In their model, the adjustment cost in production technology for the installed capital is high enough that it prevents the substitution of other production factors for low skill labor in the short run. However, they note that in the long run, firms may invest in new capital and substitute low skill labor for other production factors in response to an increase in the minimum wage.

2.2 Data

The main data for our analysis comes from the Mining and Manufacturing Survey of Statistics Korea (MMS) from 2015 to 2019. Mining and manufacturing represent 28.8% of total employment in South Korea as of 2017. This annual survey provides information on all mining and manufacturing firms with 10 or more employees, offering comprehensive data for analyzing the structure and distribution of these industries. Specifically, the data contain not only basic characteristics such as the type of legal organization, regional information, and the industrial classification of a business, but also items such as the number of workers, annual wages, the annual value of shipment, welfare expenses, severance pay, operating expenses, values of inventories, and tangible assets. This granularity allows us to examine the behavior of companies at fine grids after the minimum wage increases in 2018 and 2019. Our sample consists of establishments with at least 10 employees that fall under category B (Mining) and C (Manufacturing) according to the Korean Standard Industrial Classification, covering 71.2% of employment in these sectors. Within-firm wage distribution information for the MMS sample is supplemented by the Employment Positions Statistics (EPS).

We use the Employment Insurance Database (EID) to decompose sample attrition from the MMS sample into plant closures and layoffs and to analyze the employment effects of the minimum wage in the service sector. The EID is a population level administrative dataset covering all workplaces enrolled in employment insurance and all insured employees. It provides wage data for 2015–2017, allowing the precise calculation of wage distribution and minimum wage exposure within establishments. However, as it lacks wage data beyond 2017, it is limited in assessing the impact on wages and labor productivity. Thus, we use the EID primarily for supplementary analysis of the employment effects of the minimum wage.

Because EID lacks some control variables used in the MMS, omitting factors such as per-worker capital—if correlated with a firm's minimum wage exposure—could lead to a bias in the estimation of the employment effect. However, since additional control variables have only a limited impact on the estimates within the MMS sample in our analysis, we infer that any resulting bias is likely to be minimal. Additionally, for an alternative analysis of individual worker characteristics affected by the minimum wage hike and its employment impact at the individual level, we use the Local Area Labor Force Survey (LALFS) data.

3 Empirical Analysis

3.1 2017-2019 Minimum Wage Hike

The identification of the causal impact of the change in economic policy is challenging because most policy changes themselves are likely to be results of endogenously determined outcomes. For example, in the context of the minimum wage, if strong economic growth and labor shortage reduce opposition from employer representatives on the minimum wage commission, a sharp increase in the minimum wage may occur without producing a substantial dis-employment effect. However, the time-series plots of key macroeconomic variables in Figure 3 shows that economic growth during 2015–2017 was not markedly stronger than in 2017–2019, thereby helping to rule out this possible confounding effect. [6]

In fact, the main reason for the rapid increase in the minimum wage was a sudden shift in the balance of political power. The conservative president, Park Geun-hye, was impeached over a personal scandal and removed from the office in December 2016.[7] As a result, the presidential election happened 7 months earlier than normally scheduled and the liberal candidate, Moon Jae-in, won the election. The new liberal president, partly due to his long-term involvement with the labor rights and his career as a human rights lawyer, promised to raise the minimum wage from \$5.31 in 2017 by \$3 to \$8.31 during his five-year tenure, which was about a 55% increase from the 2017 level. More than a half of that goal (30% increase) was achieved during the first half of his term in power (2018 and 2019). The rapid increase in the minimum wage was a part of the broad economic strategy known as "income-led growth", which aimed to raise the labor share of national income through wage hikes and increased social expenditures. For both proponents and opponents, the real economic effect of the substantial increase in the minimum wage was highly controversial especially as the unemployment rate ticked up and the pace of the minimum wage increase significantly slowed down during the second half of President Moon's tenure.[8] The minimum wage rose by about 9%

[6] Average annual GDP growth was 3.2% in 2015–2017 and 3.0% in 2018–2019. Unemployment rose slightly from 3.7% to 3.8%, while inflation declined from 1.2% to 1.0%. Quarterly real GDP growth is reported as the compounded annualized rate of change.
[7] See https://www.britannica.com/biography/Park-Geun-Hye
[8] For the media coverage of the negative economic effect of the minimum wage hike, see https://www.wsj.com/articles/asias-most-radical-left-wing-economic-program-faces-a-harsh-reality-11550652596.

between 2019 and 2022. Yoon Suk-yeol, the conservative candidate, who became President Moon's successor in 2022, was highly critical of the steep rise in the minimum wage during Moon's tenure and promised to overhaul the minimum wage policy to mitigate the negative effect on small businesses.[9]

3.2 Empirical Specification

To tease out the economic effect of the rapid increase in the minimum wage as the policy treatment, we compare the two-year change during the 2017-2019 period (post-treatment) with the two-year change during the 2015-2017 period (pre-treatment) at the establishment level. Our main analysis is confined to the mining and manufacturing sector for which we can obtain a rather comprehensive dataset. We also extend the analysis by incorporating an alternative sample of service sector data to examine the employment effect.

We employ a difference-in-differences method with time-varying coefficients and compare the difference in the 2017-2019 change in employment between firms more exposed to the minimum wage hike and those firms less exposed to it with the difference in the 2015-2017 change in employment across the same firms. We define the firm-level exposure to the 2019 minimum wage (FA_i for the i-th establishment) as the fraction of workers earning at or below the minimum wage.

Table 1 shows the average minimum wage exposure in the mining and manufacturing sector as well as the economy-wide average. The manufacturing sector's exposure of 0.128 is somewhat lower than the economy-wide average of 0.141. The sector most exposed to the minimum wage is the accommodation and food service sector that hires many low-skilled workers. The average minimum wage exposure in this sector is 0.404, more than three times of the exposure in the manufacturing sector.

The MMS data provide the average wage information within a firm but not the distribution of wage at the firm level. The EPS data provide the firm level wage distribution data necessary for calculating FA_i but the coverage sample is more limited than the MMS for the mining and manufacturing sector.

[9] For then-candidate Yoon's position, see https://koreajoongangdaily.joins.com/2022/05/10/ business/economy/Korea-Yoon-Sukyeol-labor-policy/20220510150629155.html.

We use the following nonlinear regression to back out FA_i from the mean annual wage data for firms in the EPS who overlap with the MMS.[10]

$$\widehat{FA}_i = \max\left[0, \min\left[1, \alpha + \beta_1 \overline{w}_i + \beta_2 \overline{w}_i^2 + \beta_3 I(\overline{w}_i < \theta) + \beta_4 I(\overline{w}_i < \theta)\overline{w}_i + \beta_5 I(\overline{w}_i < \theta)\overline{w}_i^2\right]\right] \quad (1)$$

where $\overline{w}_i$ denotes the average wage at the establishment level.

We estimate all the parameters except for θ, which we fix at 40 million Korean won (KRW).[11] Figure 4 shows that the estimated minimum wage exposure ($\widehat{FA}_i$) fits the actual minimum wage exposure (FA_i) quite well especially for firms in the EPS sample whose mean annual wages are below 40 million KRW.

We set 2017 as the benchmark year and compute the change in the dependent variable (e.g., employment, wage, and labor productivity) relative to the 2017 value. We consider firms who existed in the three consecutive years from 2015 to 2017 in the MMS sample. Then, we add other control variables such as per worker fixed assets, non-wage benefit ratio, total labor cost per output, value-added per output and depreciation per output. Depending on the specification, we consider the location fixed effect and estimate regression coefficients weighted by firm size. The baseline regression in our study is as follows:

$$\frac{y_{i,t} - y_{i,2017}}{y_{i,2017}} = \alpha_t + \beta_t \widehat{FA}_i + \gamma_t X_{i,2017} + \epsilon_{i,t} \qquad (2)$$

where $y_{i,t}$ denotes the outcome variable of interest such as employment, wage, and labor productivity. $X_{i,2017}$ denotes the control variables from 2017. For this difference-in-differences type regression we allow all coefficients to be time-varying.

Using the 2015-2017 period with the modest increase in the minimum wage as the placebo sample, we can empirically identify the possible turning-point effect of the minimum wage hike during the 2018-2019 period, which changes β_t materially. Any significant change in the regression coefficient on the minimum wage exposure between the 2017-2019 period and the placebo sample

[10] The functional form closely follows the one in Harasztosi and Lindner (2019), but we add an inflection point θ to better fit our dataset.
[11] We use this inflection point because the minimum wage exposure ratio is flattened out above this level. This finding is similar to Draca, Machin and Van Reenen (2011).

period can be regarded as evidence for the effect caused by the rapid increase in the minimum wage relative to the modest increase during the previous period.

3.3 Economic Effect of the Rapid Increase in the Minimum Wage

We examine the effect of the minimum wage hike on firm-level employment, wage and labor cost, and labor productivity subsequently.

3.3.1 Employment

Panel A of Table 2 shows the two-year change in employment from the minimum wage hike during the baseline sample period (2017-2019) and the placebo sample period (2015-2017). While the minimum wage exposure has no statistically significant effect on employment when the increase was modest as in 2015-2017, it has a statistically significant and negative employment effect in 2017-2019 when the increase was rapid. For the baseline sample period, a one-standard deviation increase in the minimum wage exposure for a firm in the manufacturing sector (0.211 from Table 1) is expected to decrease employment by about 4% when we consider adjustments along both extensive margins (plant closings) and intensive margins (layoffs). Plugging in the increase in the average exposure between 2017 and 2019 of about 0.15, our estimates suggest roughly 3% decline in total domestic employment in the manufacturing sector caused by the minimum wage hike.

Since we included all the firms that were active in the MMS sample during the three-year period from 2015 to 2017 in our study, we must decide how to treat firms that drop out of the 2017 MMS sample to disentangle intensive and extensive margin adjustments. There are two reasons that firms drop out from the MMS sample. First, firms may layoff many workers, reducing the size of the employee below 10, which is the cutoff value for being in the MMS sample. Second, firms may exit the market, reducing the size of the employee count to 0. Unfortunately, the MMS does not provide explanation as to why firms in the previous year sample drop out. To address this issue, we replicate the MMS sample from the EID sample. Since the EID is the administrative dataset that covers any employer, we can classify firms who drop out from both the MMS sample and the EID sample as closing plants.[12] By comparing the 2017 MMS sample with the replicated MMS sample from the

[12] Since industrial classification across firms is not perfectly consistent with the MMS and the EID sample especially when firms have multiple workplace locations, it is difficult to perfectly replicate the MMS sample from the EID sample. The replicated MMS sample that we construct from the EID sample contains 52,810 firms while the original MMS sample has 50,856 firms. We verify that summary statistics across two samples are almost identical.

EID in 2018 and 2019, we can calculate the ratio of firms that closed plants out of those who dropped out of the MMS sample in 2018 and 2019. Table 3 provides the decomposition of sample attrition by employment size. Notably, plant closing is more likely with larger firms. For firms with employees of 10-15 that drop out the MMS sample, only 20% were found to have closed plants. In contrast, for firms with employees over 50, 92% closed plants.[13]

Using this information, we estimate the employment effect from the intensive margin only by ruling out the fraction of firms who were likely to drop out due to plant closing. Panel B of Table 2 shows that the employment effect solely from the intensive margin ranges from 55% to 61% of the total effect depending on the specification. This finding suggests that a significant part of the dis-employment effect from the rapid increase in the minimum wage can be attributed to the adjustment through plant closing.

Figure 5 shows the employment adjust patterns by the minimum wage exposure rate. While the linear relationship holds for most firms, there are some jumps at the right tail of the minimum wage exposure. The finding suggests that those most vulnerable firms made disproportionally large adjustments, in particular through extensive margins.

3.3.2 Robustness Checks for the Employment Effect

Our calculation of the total employment effect of the minimum wage hike depends on the decomposition of firms who dropped out of the 2017 MMS sample. To check how robust our baseline calculation is to alternative assumptions on the sample attrition, we re-estimate the employment effect of the minimum wage hike based on the three different scenarios as shown in Table 4. The first alternative (Alt 1) is to assume that all the firms who dropped out of the 2017 sample in the subsequent two years closed plants. Since the employment level of firms who closed plants is zero, Alt 1 is likely to overstate the true dis-employment effect. Indeed, we observe a more severe dis-employment effect with the coefficient on the minimum wage hike changing from -0.2 in the baseline case to -0.29 in Alt 1. In contrast, the second alternative (Alt 2) rules out all the firms that dropped out of the 2017 MMS sample and runs the regression only for those who had been in

[13] This calculation is conditional on the sample attrition. So, for firms who dropped out of the sample, the plant closing is more likely for large firms than the reduction in employment size. On average, we do not see much difference in terms of the magnitude of extensive margin adjustment relative to the intensive margin adjustment across different firm size because the sample attrition probability is much lower for larger firms.

the MMS sample during the five years from 2015 to 2019. Alt 2 is likely to understate the true dis-employment effect of the minimum wage hike by ruling out some firms who made a drastic reduction in the employment level. As expected, the magnitude of the dis-employment effect is reduced compared to the baseline case with the regression coefficient on the minimum wage exposure changing from -0.2 to -0.09. Finally, we consider the third alternative (Alt 3) which imputes the average employment level for firms who dropped out of the 2017 sample by taking a weighted average of plant closing and layoff cases in the EID sample. The difference between Alt 3 and the baseline is that Alt 3 imputes the same value for the employment level of all the firms who dropped out of the 2017 sample whereas in the baseline case, we distinguish the employment level of firms who closed plants from the employment level of firms who laid off some workers. Despite this difference, the estimated magnitude of the dis-employment effect of the minimum wage exposure in Alt 3 is like the baseline case.

We look at the employment effect of the minimum wage hike beyond the mining and manufacturing sector. The service sector hires a lot of low wage workers in the accommodation and food service sector, which has the highest minimum wage exposure as shown in Table 1. However, collecting the dataset for the service sector comparable to the MMS sample is challenging because most businesses in the service sector are quite small.[14] To analyze the annual changes in employment within the service sector following the minimum wage increase, we use the EID, a matched employer-employee dataset covering all firms enrolled in employment insurance and all insured employees. Additionally, the 2015–2017 EID data provide annual wage information for insured workers within firms, allowing for precise calculations of wage distribution and minimum wage exposure at the firm level.

Table 5 presents the estimated dis-employment effects of the minimum wage hike using Equation (2) based on the EID sample. For comparison with the MMS sample, we restrict the analysis to firms with at least 10 employees when estimating the employment effects. We compare the employment effects across different sectors during the 2017–2019 period: all establishments (Column 1), the manufacturing sector (Column 2), and the service sector (Column 3). Although the EID data does not necessarily provide information on the additional control variables (X_{it} in equation (2)) used in the MMS sample, the estimated results for the manufacturing sector (Column

[14] For example, firms with employees fewer than 10 account for more than 60% in the accommodation and food service sector while they account for 27.9% in the manufacturing sector.

2) based on the EID sample closely align with those from the MMS sample in Table 2, centered around -0.2. Table 5 also shows that the regression coefficient on minimum wage exposure is quite comparable across both the manufacturing and service sectors, although the estimate for the service sector is slightly smaller (-0.205 vs -0.134).[15]

One caveat in interpreting our results is that there might be separate trends for employment changes across firms with different levels of the minimum wage exposure. In equation (2), we control for a common trend but not for diverging trends across firms with different minimum wage exposure. To ensure that this does not confound our results, we estimate the yearly effect of the minimum wage hike on employment as shown in Figure 6 for the manufacturing and mining sector (Figure 7 for the service sector). While the two year change in employment from 2015 to 2017 is little affected by the minimum wage exposure, we find a slight negative employment effect of the minimum wage hike in 2016.[16] However, the magnitude is relatively small compared to the yearly effect in 2018.[17] This finding suggests that the violation of the common trend assumption is unlikely to be a factor in explaining the significant dis-employment effect of the minimum wage hike after 2017. Another important identification assumption in our study is the Stable Unit Treatment Value Assumption (SUTVA) which prohibits cross-firm spillover effects from firms with high exposure to the minimum wage to those with low exposure. For instance, if the minimum wage hike has ripple effects on the upper spectrum of wage distribution, firms not directly exposed to the minimum wage hike might still be indirectly affected. However, we do not find evidence that the minimum wage hike substantially shifted the upper spectrum of wage distribution rightward as shown in Figure 8. This finding alleviates concerns that the SUTVA was violated in our sample.

3.3.3 Wage and Labor Cost

As Manning (2021) notes, the effect of a minimum wage hike on wage and labor cost[18] is less controversial with most studies finding a significant positive effect. Our study is not an exception. One challenge in examining the effect on wage and labor cost is that the EID sample does not

[15] Figure 6 presents the event study estimates of the yearly effects, suggesting that there is no pre-existing trend in the dis-employment effect of the minimum wage hike in the service sector before 2017.
[16] Since we regress the difference in employment level between 2016 and 2017, the positive coefficient means a dis-employment effect with the level of employment lower in 2017 than in 2016.
[17] The coefficient is 0.04 in 2016 compared to -0.1 in 2018.
[18] Labor cost is defined as the sum of wages and non-wage benefits at the firm level

provide information on the average wage for each firm. Hence, we cannot back out the changes in wage and labor cost for firms who dropped out of the 2017 MMS sample and use the balanced panel data for firms between 2015 and 2019. However, the attrition bias is only valid for firms who had layoffs instead of closing plants. Because these firms are most likely to be small as we see in Table 3, we do not expect a significant attrition bias for the size-weighted coefficient.

The regression coefficient on the minimum wage exposure in average wage growth rate and average labor cost growth rate in Table 6 confirms that the minimum wage hike increased wage growth and labor cost growth. Although even the modest minimum wage hike during 2015-2017 boosted growth rates of wage and labor cost, the magnitude of the effect substantially increased during the period of the rapid increase in the minimum wage of 2017-2019.[19] This pattern is also clearly visualized in Figure 9.

Firms with high exposure to the minimum wage hikes during 2017–2019 may have experienced meaningful wage increases even during the earlier phase of gradual hikes. To gauge the wage and labor cost effects of the minimum wage policy under the pre-2017 trajectory, Figure 9 overlays the actual estimates with a counterfactual trend predicted by fitting a linear trend to the estimated β_t coefficients from the 2015–2017 period. The results show that both wages and labor costs deviated substantially from the projected trend following the sharp policy change, suggesting that the 2018–2019 increases cannot be explained by previous patterns alone.

Engbom and Moser (2022) highlight the effect of the minimum wage hike on reducing wage inequality based on the Brazilian data from 1996 to 2012. Because the minimum wage hike shifts up the lower tail of the wage distribution but is likely to have little effect on the upper tail, the wage dispersion may reduce more substantially when the pace of the minimum wage hike accelerates. Using detailed firm-level wage distribution data from the EPS, we estimate a nonparametric kernel density for wage distribution in the mining and manufacturing sector, as shown in Figure 8. We observe a substantial rightward shift in the lower tail of the wage distribution, along with a modest upward drift in the middle, while the upper tail remains largely unaffected. Moreover, the magnitude of the shift is greater between 2017 and 2019—when the minimum wage rose rapidly—compared

[19] The negative sign in the 2015-2017 regression should be interpreted as the higher minimum wage exposure increasing wage growth rate and labor cost growth rate because we define the difference in the labor cost between 2015 and 2017 as the dependent variable.

to the earlier period of more modest increases. This shift in the wage distribution during the period of sharp hikes aligns with findings from Harasztosi and Lindner (2019) based on Hungarian data.

3.3.4 Labor Productivity

The increase in the minimum wage can also change labor productivity. This is because firms with higher exposure to a minimum wage shock may respond to the shock by changing composition of workers with different skill levels or by substituting labor with other inputs. We examine the impact of the 2017-2019 minimum wage hike on labor productivity (LP) using the same difference-in-differences type regression approach in the analysis of employment effect. We measure labor productivity as the log of output per worker at the firm level and estimate the equation (2) using this measure as the dependent variable. The results are reported in Table 7.

The placebo regression results prior to the minimum wage hike (columns (4)-(6)) show that there existed a pre-trend of steadily increasing labor productivity through 2017. Because the minimum wage had been increasing modestly before the hike, it is likely that these gradual increases had positive effects on LP even before the sharp hike. One concern is that this violates the parallel trend assumption. We, however, stress that our treatment is not the minimum wage introduction nor a modest gradual increase but rather its sudden rapid hike, so this pre-trend is not necessarily threatening our identification of the effect on labor productivity from the minimum wage hike. What matters in our context is to see whether the minimum wage hike had accelerated this positive effect. Indeed, our main finding is that the effect soared and was almost doubled during the 2017-2019 minimum wage hike (columns (1)-(3)). Therefore, we may attribute this jump in labor productivity effect to the 2017-2019 minimum wage hike. The finding that labor productivity was boosted in firms more exposed to minimum wage shock is consistent with the substitution of low-skill labor for high-skill labor or other inputs like capital.

3.3.5 Individual Characteristics of Workers Exposed to the Minimum Wage Hike

We provide additional evidence for the substitution of low-skill labor as the main channel for the dis-employment effect of the minimum wage hike by examining individual characteristics of workers who are most affected by the minimum wage hike. LALFS data provides detailed information on individual characteristics of workers as well as wage information. For this analysis, we define being exposed to the minimum wage hike as the current wage below 120% of the minimum wage. Table 8 shows dominant characteristics of individual workers who are most affected by the minimum wage hike by computing the marginal R^2 from each characteristic. As in Cengiz et al. (2022), age and education are the main determinants of being exposed to the minimum wage hike. Young workers with the low level of education are typical ones who are at the highest risk of being exposed to the minimum wage hike.

We examine the dis-employment effect of being exposed to the minimum wage hike by using variations in the average probability of being exposed to the minimum wage across 1,358 subgroups (17 metro areas × 7 age groups × 2 gender groups × 3 education groups × 2 marital status groups).[20] Table 9 shows that groups who are more exposed to the minimum wage hike are less likely to be employed based on the 2011-2020 period. The rapid increase in the minimum wage tends to raise the variation in the average probability of being exposed across subgroups and will reduce the probability of being employed for the group of people who are more vulnerable to the minimum wage hike.

4 Discussion

4.1 Policy Implications

Our estimated negative employment effect from the rapid minimum wage increase is more than twice as large as that found by Harasztosi and Lindner (2019), despite the comparable pace of the minimum wage hike. The regression coefficient of the change in employment on minimum wage exposure is -0.2 in our case compared to their estimate -0.078. The difference is much smaller if we restrict our comparison to the manufacturing sector only. The large difference is mainly attributed to a much stronger dis-employment effect in the service sector that we find.[21]

One important point to note in our analysis is that we are detecting a significant and negative employment effect because the rapid increase in the minimum wage substantially increases the degree of firm-level minimum wage exposure. Conditional on different levels of minimum wage

[20] The employment effects of minimum wage increase at the group level in Table 9 were obtained using the following model. $ln(y_t^g) = \alpha + \beta_1 \widehat{FA}_t^g + \beta_2 \widehat{FA}_{t-1}^g + \gamma_g + \delta_{p,t} + \mu_{g,t}$, where g represent 1,358 subgroups, $ln(y_t^g)$ indicates the log of employment rate for group g at time t, and $\widehat{FA}_t^g$ is group g's minimum wage exposure. $\delta_{p,t}$ shows the interaction term between metro area fixed effects and year fixed effects. Although we constructed a pseudo-panel at the group level since the LALFS dataset consists of repeated cross-sectional data, the number of observations within each group varies across years, making it difficult to track employment changes relative to a specific year, as in Equation (2).

[21] Unlike their study, we do not find a significant price pass-through of the minimum wage hike in the service sector as discussed in the appendix.

exposure, the relationship between minimum wage exposure and the change in the employment is roughly linear at each level. However, the linear relationship between minimum wage exposure and the change in employment on average does not imply that firms adjust employment continuously with the change in the minimum wage exposure. In particular, as Figure 5 shows, firms at the right tail of the minimum wage exposure rate disproportionately large employment adjustment, deviating from what is implied by the linear relationship.

One reason we find that larger firms are more likely to close plants is that the job subsidy program that mitigated the negative employment effect of the minimum wage hike is limited to firms with employees of 30 or fewer.[22] [23]However, the job subsidy program was inefficient in achieving the desired effect because it was not designed to take into account a large degree of heterogeneity in the minimum wage exposure across different sectors as we see in Table 1. The administration that came into power in May 2022 initiated discussions on introducing differentiated minimum wages by region and industry. While the proposal drew attention as a potential measure to mitigate the adverse employment effects of uniform hikes, it remains at the legislative and public debate stage due to strong labor opposition, difficulty in setting objective criteria, and concerns over regional stigmatization. To the extent that the negative employment effect is driven by the degree of minimum wage exposure—as we find using firm-level data—such differentiated schemes, if properly designed with data-driven standards and broad social consensus, could help alleviate the adverse impact of minimum wage hikes.

4.2 Implications for Models of the Labor Market

The employment effect of the labor market can be used to discriminate alternative models of the labor market. Dube (2019) lists the following four factors determining the employment effect of the minimum wage hike: 1) the degree of substitutability between low-skill workers and other production inputs, 2) the degree of price pass-through of wage increases (e.g., imperfect competition in the product market), 3) imperfect competition in the labor market (e.g., monopsony), and 4) endogenous job search efforts which increase labor force participation as the minimum wage rises. Engbom and Moser (2022) introduce an equilibrium model of labor market search highlighting

[22] The program started in 2018 but was discontinued in June 2022.

[23] Hong and Moon (2024) show that firms with fewer than 30 employees—eligible for the job subsidy program—did not experience employment declines following the 2018 minimum wage hike, in contrast to larger firms not covered by the subsidy.

endogenous job search efforts to explain a small negative employment effect from the increase in the Brazilian minimum wage between 1996 and 2012. Similarly, Haanwinckel (2023) sets up an equilibrium task-based production model with labor market monopsony to explain the same pattern in the Brazilian data. In contrast, Neumark et al. (2008) emphasize the substitution channel for the dis-employment effect of the minimum wage hike based on a competitive model of the labor market. Meanwhile, Harasztosi and Lindner (2019) emphasize the heterogeneity in the degree of competition in the product market and show that the employment effect can be small if the employer can pass through the increased cost from the minimum wage hike to consumers using market power.

They argue that the linear relationship between the minimum wage exposure measure and the change in employment is not compatible with a standard model of labor market monopsony. In such models, a modest increase in the minimum wage is expected to reduce the wage markdown and lead to an increase in employment, even if more firms are exposed to the minimum wage after the increase. However, the relation may change if the pace of the increase becomes rapid and substantial. While, the criticism would be likely valid for the model without the heterogeneity in the degree of wage markdown, it does not necessarily apply to a richer model with heterogeneity. Our findings can be similarly interpreted as evidence against a simple model of labor market monopsony in which the heterogeneity in the wage markdown is limited. In addition, the fact we observe the negative employment for low wage workers goes against models emphasizing endogenous search efforts in which the search intensity and employment should increase for workers more affected by the minimum wage. Since we do not find much difference between the manufacturing sector and the service sector, the difference in the degree of product market competition across tradable (manufacturing) versus non-tradable (services) sectors does not seem to be a big factor in understanding the economic effects of the minimum wage hike in South Korea.

Our finding that larger firms are more likely to close plants suggests that a part of the adjustment occurs through offshoring. While detailed firm-level offshoring data comparable to our MMS sample is unavailable, aggregate data from the Export-Import Bank of Korea shows a significant increase in South Korean foreign direct investment in Vietnam during 2017–2019 compared to 2015–2017, as illustrated in Figure 10. This aligns with the offshoring of low-skill tasks in the task-based production model of Grossman and Rossi-Hansberg (2008).

Beyond offshoring, firms may also reorganize production to lower the task difficulty threshold at which they hire low-skill workers subject to the minimum wage. This could lead to a significant negative employment effect following a sharp minimum wage increase, while yielding a muted response to more gradual increases at the firm level due to the presence of firm-specific fixed cost of adjustment related to offshoring. However, when aggregated across firms with heterogeneous adjustment costs, the economy-wide effect may not exhibit such a stark discontinuity—consistent with both our model and empirical findings. We now propose a simple task-based production model that captures this observed pattern in the South Korean data.

4.3 An Illustrative Model: A Task-based Production Model with Skill Substitution

A task-based production model has been widely used to study the implications of technological changes on wage inequality and labor market polarization.[24] Eckardt (2022) studies the implications of a binding minimum wage in a task-based production model where low-skill workers can be replaced by machines or high-skill workers when the minimum wage increases. However, this model does not have any of the firm-level heterogeneity that we need to explain the variation in the adjustment along the extensive margin across firm sizes. Haanwinckel (2023) sets up a task-based production model with an imperfect labor market. His model has a rich pattern of heterogeneities in both workers and firms. In it, each worker is attracted to a particular firm through a non-pecuniary amenity shock, which generates monopsony power of the firm on the worker. The model is estimated to match moments of the Brazilian wage distribution in 1998 and 2012. The estimated model generates a small employment elasticity with respect to the Kaitz index since it is not focusing on what happened during the short period of time when the Kaitz index rose substantially (2004-2007).

We first introduce a simple static task-based production model in which firms are subject to a binding minimum wage for low-skill workers to examine the skill substitution effect from a minimum wage hike. Then we extend the model with firm-level heterogeneity in the comparative advantage of high skill labor over low skill labor and the cost of reorganizing the task difficulty threshold below which they hire domestic low skill workers as opposed to foreign low skill workers.

4.3.1 A Simple Model without Firm-Level Heterogeneity

[24] For example, see Acemoglu and Autor (2011) for a comprehensive review of a task-based production model.

Suppose that there are two types of workers in which L type (H type) represents the low-skill (high-skill) worker. There are J firms, each producing the final output by aggregating a variety of tasks whose difficulty level is indexed by $i \in [0,1]$. We assume the firm j's ($j \in \{1,\cdots,J\}$) final output Y_j is determined as follows:

$$Y_j = e^{\int_0^1 \ln y_j(i) di},$$

$$y_j(i) = A_l(i) l_j(i) + A_h(i) h_j(i) \tag{3}$$

where $l_j(i)(h_j(i))$ denotes the low-skill (high-skill) worker hired by firm j to perform task i.

There are $\overline{L}$ low skill workers and $\overline{H}$ high skill workers in the economy. For simplicity, we assume that low skill workers and high skill workers are perfect substitutes for each task but high skill workers have comparative advantages over more difficult tasks with $\frac{A_h(i)}{A_l(i)}$ increasing with i. Notice that all firms have the same production technology with each production factor's productivity not depending on the firm index j.

As shown in Acemoglu and Autor (2011), with this technology, a firm's labor demand can be characterized by the task difficulty threshold $i^\star$ in which the firm hires only low skilled (high skilled) workers for task $i < i^\star (i > i^\star)$. The firm is indifferent between a low skilled worker and a high skilled worker for task $i^\star$ but we assume that it hires a low skilled worker to break the tie. Firm j's demand for each type of worker at task i is determined as follows:

$$l_j(i) = \begin{cases} \frac{y_j(i)}{A_l(i)} & if \quad i \le i^\star, \\ 0 \quad if & i > i^\star, \end{cases}, \quad h_j(i) = \begin{cases} 0 & if \quad i \le i^\star, \\ \frac{y_j(i)}{A_h(i)} \quad if & i > i^\star. \end{cases}$$

We can pin down $i^\star$ based on the equalization of the unit cost for using each type of worker as

$$\frac{w_l}{A_l(i^\star)} = \frac{w_h}{A_h(i^\star)}. \tag{4}$$

Given technology parameters determining $A_l(\cdot)$ and $A_h(\cdot)$, $i^\star$ is given by a function of (w_l, w_h), which we denote by $i^\star(w_l, w_h)$, The following labor market equilibrium conditions pin down the equilibrium wage (w_l, w_h), for each type of worker as a function of technology $(A_l(\cdot), A_h(\cdot))$ and labor endowment $(\overline{L}, \overline{H})$:

$$L(w_l, w_h) = \sum_{j=1}^{J} \int_0^1 l_j(i) di = \overline{L}, \quad (5)$$

$$H(w_l, w_h) = \sum_{j=1}^{J} \int_0^1 h_j(i) di = \overline{H}.$$

Let us denote the equilibrium wage determined from the above condition by the competitive equilibrium wage (w_l^{CE},w_h^{CE}). In this case, wages are adjusting to clear markets for two types of labor, $i^{\star}$ is ultimately determined by technology ($A_l(i)$,$A_h(i)$) and aggregate labor supply from two types of workers ($\overline{L}, \overline{H}$).

Assuming the final goods as numeraire (setting its price to 1), the price of task i, denoted by $p(i)$ is given by the following condition as shown by Acemoglu and Autor (2011) as follows:[25]

$$e^{[\int_0^1 \ln p(i) di]} = e^{\int_0^{i^{\star}} \ln\left(\frac{w_l}{A_l(i)}\right) di + \int_{i^{\star}}^1 \ln\left(\frac{w_h}{A_h(i)}\right) di} = 1. \quad (6)$$

Now, we introduce the minimum wage $\underline{w_l}$ which can be binding for the low-skill worker. Suppose that $\underline{w_l}$ is greater than w_l^{CE}. Since $\underline{w_l}$ is exogenously given instead of being determined to clear the labor market for low-skill workers, only the labor market for high-skill workers clear and $\underline{w_l}$, instead of $\overline{L}$, determines the task threshold i^*. That gives us the new equilibrium value of w_h. The following proposition shows that the task threshold i^* declines monotonically as the minimum wage binding for low skill labor increases.

Proposition 1. *As the minimum wage binding for low skill labor $\underline{w_l}$ increases, the task threshold $i^{\star}$ and the demand for low skill labor decrease monotonically. The demand for high skill labor is unchanged and remains equal to the supply.*

Proof. From equation (6), we have the following condition.

$$\int_0^{i^{\star}} ln(\frac{\underline{w_l}}{A_l(i)}) di + \int_{i^{\star}}^1 ln(\frac{w_h}{A_h(i)}) di = 0.$$

We know $w_h = \frac{A_h(i^{\star})}{A_l(i^{\star})} \underline{w_l}$ from equation (4). Plugging these equations together, we can derive the following condition:

[25] The price of task i is the same as the unit cost of performing task i under the optimal allocation of labor with different skill types.

$$\int_0^{i^\star} ln(\frac{\underline{w_l}}{A_l(i)})di + \int_{i^\star}^1 ln(\frac{A_h(i^\star)\underline{w_l}}{A_h(i)A_l(i^\star)})di \ = 0,$$

$$ln(\underline{w_l}) - \int_0^{i^\star} lnA_l(i)\,di - \int_{i^\star}^1 lnA_h(i)\,di + ln\left(\frac{A_h(i^\star)}{A_l(i^\star)}\right)(1 - i^\star) = 0,$$

$$ln\,(\underline{w_l}) \ - \ g(i^\star) \ = \ 0.$$

Since the final product market is competitive, the minimum wage increase does not affect the final product price and equation (6) should hold even if $\underline{w_l}$ changes. By applying the implicit function theorem, we can show

$$\frac{di^\star}{d\underline{w_l}} = \left(\frac{1}{\underline{w_l}\,g'(i^\star)}\right) = \frac{1}{\left(\frac{d}{di^*}ln\frac{A_h(i^\star)}{A_l(i^\star)}\right)(i^\star - 1)} \leq 0.$$

In the last part, we use the comparative advantage assumption that the relative productivity of high skill labor increases as task difficulty increases while $i^\star$ is bounded above at 1. Since the log of the final output is the sum of the log of each task service output, it is optimal to produce each task service by equal amount. Also, the production technology is identical across firms. Combining these two observations, we obtain the total demand for low skill labor is $\frac{i^\star\bar{Y}}{\underline{w_l}}$ where $\bar{Y}$ is $\frac{w_h\bar{H}}{1-i^\star}$. From equation (4), the relative wage of high skill labor declines when $i^\star$ decreases given our assumption on the comparative advantage. Also, $\frac{i^\star}{1-i^\star}$. decreases as $i^\star$ decreases while $\bar{H}$ is unchanged. Putting these together, we note the total demand for low skill labor declines as $i^\star$ decreases due to the increase in $\underline{w_l}$. The demand for high skill labor is equal to $\bar{H}$ because w_h adjusts to balance demand with supply.

Proposition 1 implies that firms will use high skill labor for tasks previously performed by low skill labor as the minimum wage increases, highlighting the labor substitution channel that can create negative employment effects for low skill workers due to the minimum wage hike.

More specifically, from equation (4), $\frac{A_h(i^\star)}{A_l(i^\star)} = \frac{w_h}{\underline{w_l}} < \frac{w_h^{CE}}{w_l^{CE}}$ where w_l^{CE} and w_h^{CE} denote the wage level of low skill labor and high skill labor at a competitive equilibrium, respectively. Under the competitive equilibrium, wages are adjusted to balance demand with supply for the labor market of each skill type. $i^\star$ declines below $i^{\star,CE}$ if $\underline{w_l}$ rises above w_l^{CE}.[26] Since demand for low skill workers declines but the supply of low skill workers does not change, the binding minimum wage creates the dis-employment by $\bar{L} - L(w_l, w_h)$. The binding minimum wage distorts the allocation of workers with different skill types across tasks and reduces the aggregate output by doing so. Thus, $Y = \sum_{j=1}^{J} Y_j < \sum_{j=1}^{J} Y_j^{CE} = Y^{CE}$.

4.3.2 An Extended Model with Firm-Level Heterogeneity

The simple model previously described does not consider the firm-level heterogeneity which plays a critical role in our empirical study. In addition, while firms in the manufacturing sector can move factories to foreign countries when the domestic minimum wage increases, that kind of adjustment is not feasible for the non-tradable service sector. Instead, firms in the non-tradable service sector may automate tasks previously performed by low skill workers when the minimum wage increases. In this extended version of the baseline model, we consider two sectors $l \in \{m,s\}$ and distinguish the domestic low skill labor L_D and the foreign low skill labor L_F. Additionally, we add capital (K) to incorporate the margin for automation. The difficulty of the task level is still indexed by $i \in [0,1]$ as in the simple model. Each sector consists of a continuum of firm who are indexed by $j \in [0,J]$. We sort firms into G groups depending on the comparative advantage of domestic high skill labor over domestic low skill labor. Firms in each group $g \in \{1,\cdots,G\}$ share the same production technology. We index firms in the j-th group by $j \in [j_{g-1}, j_g]$ where $j_0 = 0$ and $j_G = J$. For any given task i, the relative productivity of domestic high skill labor over domestic low skill labor increases with a higher value for g.[27]

$$g' > g \rightarrow \ln\left(\frac{A_H^{g',l}(i)}{A_{L,D}^{l}(i)}\right) > \ln\left(\frac{A_H^{g,l}(i)}{A_{L,D}^{l}(i)}\right) \forall i \in [0,1], l \in \{m,s\}, g, g' \in \{1,\cdots,G\}. \tag{7}$$

Firms in the manufacturing sector can use foreign low skill labor through offshoring but firms in the service sector cannot. We specify the following production function of task i for firm j in sector l when j belongs to the g-th group as follows:

$$Y_j^l(i) = \left[\alpha_l^{\rho_l} A_k^l(i) K_j^l(i)^{1-\rho_l} + (1-\alpha_l)^{\rho_l} L_j^l(i)^{1-\rho_l}\right]^{\frac{1}{1-\rho_l}}, l \in \{m, s\}, \tag{8}$$

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[26] The comparative advantage assumption generates the monotonicity of $i^\star$ with respect to the relative wage of a high-skill worker $\frac{w_h}{w_l}$. If the relative wage of a high-skill worker decreases (increases), $i^\star$ also decreases (increases).

[27] A similar specification is used in Haanwinckel (2023) to introduce the worker-level heterogeneity.

$$L_j^m(i) = \left[\left(A_{L,D}^m(i)L_{D,j}^m(i) + A_{L,F}^m(i)L_{F,j}^m(i)\right)^{1-\epsilon^m} + \left(A_H^{m,g}(i)L_{H,j}^m(i)\right)^{1-\epsilon^m}\right]^{\frac{1}{1-\epsilon^m}}, \tag{9}$$

$$L_j^s(i) = \left[\left(A_{L,D}^s(i)L_{D,j}^s(i)\right)^{1-\epsilon^s} + \left(A_H^{s,g}(i)L_{H,j}^s(i)\right)^{1-\epsilon^s}\right]^{\frac{1}{1-\epsilon^s}}. \tag{10}$$

The minimum wage ($\underline{w_l}$) is binding for domestic low skill workers and the wage of the foreign low skilled worker in the manufacturing sector in case of offshoring is given by $\underline{w_f}$. The wage for a high skilled worker w_h is endogenously determined to balance demand and supply for high skill workers. Similarly, the rental rate for the capital stock is determined to balance demand and supply of capital.

Suppose that $ln\frac{A_{L,D}^m(i)}{A_{L,F}^m(i)}$ is monotonically increasing with respect to i. Since domestic low skill labor and foreign low skill labor are perfect substitutes, we can find the task threshold value $i_m^\star$ from the manufacturing firm's cost minimization condition that separates tasks performed by domestic low skill labor ($i > i_m^\star$) and those by foreign low skill labor ($i > i_m^\star$). At $i_m^\star$, the unit cost of hiring a domestic low skilled worker to perform the task is same as the one of hiring a foreign low skilled worker.

$$\frac{\underline{\mathrm{w_l}}}{\mathrm{A_{L,D}^m(i_m^\star)}} = \frac{\underline{\mathrm{w_f}}}{A_{L,F}^m(i_m^\star)}. \tag{11}$$

Figure 11 shows how $i_m^\star$ changes in response to a minimum wage increase, resulting in the substitution of foreign low skill labor for domestic low skill labor. Since domestic low skill labor has comparative advantage over foreign low skill labor in performing more complex tasks, $i_m^\star$ increases as the minimum wage substitutes away from the demand for domestic low skill labor, creating the following demand function.

$$L_{F,j}^m(i) = \begin{cases} \frac{Y_j^m(i)}{A_{L,F}^m(i)} & if\ i \le i_m^\star, \\ 0 & if\ i > i_m^\star, \end{cases} \quad ,L_{D,j}^m(i) = \begin{cases} 0 & if\ i \le i_m^\star, \\ \frac{Y_j^m(i)}{A_{L,D}^m(i)} & if\ i > i_m^\star. \end{cases}$$

Unlike two types of low skill labor, high skill labor and capital are imperfectly substitutable for low skill labor $\left(\frac{1}{\epsilon^m} < \infty, \frac{1}{\rho_l} < \infty\right)$.[28] Hence, all the tasks use these factors but with varying degrees due to the comparative advantage conditions. The firm's cost minimization leads to the following

[28] $\frac{1}{\epsilon^l}$ and $\frac{1}{\rho_l}$ are elasticities of substitution for $l \in \{m, s\}$.

first-order conditions for the allocation of high-skill labor and capital based on relative prices, where r is the rental rate for capital service when firm j belongs to the g-th group.

$$\frac{w_h}{\underline{w}_l} = \left(\frac{A_H^{m,g}(i)L_{H,j}^m(i)}{A_{L,D}^m(i)L_{D,j}^m(i) + A_{L,F}^m(i)L_{F,j}^m(i)}\right)^{-\epsilon^m} \frac{A_H^{m,g}(i)}{A_{L,D}^m(i)}, \tag{12}$$

$$\frac{\underline{w}_l}{r} = \left(\frac{1-\alpha_m}{\alpha_m}\right)^{\rho_m} \left(\frac{A_k^m(i)K_j^m(i)}{L_j^m(i)}\right)^{\rho_m} \left(\frac{L_j^m(i)}{A_{L,D}^m(i)L_{D,j}^m(i) + A_{L,F}^m(i)L_{F,j}^m(i)}\right)^{\epsilon^m} \frac{A_{L,D}^m(i)}{A_k^m(i)}. \tag{13}$$

We can derive similar cost minimization conditions for firms in the service sector. Since they cannot hire foreign low skill labor, they can either substitute high skill labor and/or capital for low skill labor when the minimum wage increases. The resulting factor demand function for firm j in the g-th group describes the optimal behavior.

$$\frac{w_h}{\underline{w}_l} = \left(\frac{A_H^{s,g}(i)}{A_{L,D}^s(i)}\right)^{1-\epsilon^s} \left(\frac{L_{H,j}^s(i)}{L_{D,j}^s(i)}\right)^{-\epsilon^s} \tag{14}$$

$$\frac{\underline{w}_l}{r} = \left(\frac{1-\alpha_s}{\alpha_s}\right)^{\rho_s} \left(\frac{A_k^s(i)K_j^s(i)}{L_j^s(i)}\right)^{\rho_s} \left(\frac{L_j^s(i)}{A_L^s(i)L_{D,j}^s(i)}\right)^{\epsilon^s} \frac{A_L^s(i)}{A_k^s(i)}. \tag{15}$$

In our empirical analysis, we define the firm-level exposure to the minimum wage as the fraction of workers subject to the minimum wage. In the model, the equivalent expression is the number of domestic low skill workers at each firm divided by the number of total domestic workers hired by the firms. We can express the model-implied measure of the firm-level exposure as follows:

$$FA_j^{l,\text{model}} = \frac{L_{D,j}^l}{L_{D,j}^l + L_{H,j}^l} = \frac{1}{1 + \left(\frac{L_{H,j}^l}{L_{D,j}^l}\right)}, l \in \{m, s\}, \tag{16}$$

where $L_{D,j}^l = \int_0^1 L_{D,j}^l(i)di$ and $L_{H,j}^l = \int_0^1 L_{H,j}^l(i)$.

Notice that $FA_j^{l,\text{model}}$ is a decreasing function of $\frac{L_{H,j}^l}{L_{D,j}}$ and given our assumption on the comparative advantage of high skill labor sorted by groups, this implies that the firm-level minimum wage exposure declines as g increases. Now we prove that total domestic employment declines more in percent terms for firms with higher exposure to the minimum wage (those that belong to the lower g group). For **Proposition 2**, we abstract from the substitution between domestic low skill labor and foreign low skill labor in the manufacturing sector that is affected by the firm-level heterogeneity in the adjustment cost, which we will introduce later (see **Proposition 3**).

Proposition 2. *Suppose that the market clearing wage for low skill labor is above the minimum wage at t. As the minimum wage binding for low skill labor* $\underline{w}_{l,t+1}$ *increases above the market clearing wage, the growth rate in the total domestic employment for firm j* $\left(\frac{L^{l}_{D,j,t+1}+L^{l}_{H,j,t+1}}{L^{l}_{D,j,t}+L^{l}_{H,j,t}}-1\right)$ *is lower for a firm with a higher* $FA^{l,\text{model}}_{j}$ *for* $l \in \{m,s\}$ *when capital per worker is controlled as in our empirical analysis (*$\beta_t < 0$ *in equation (2)).*

Proof. When the minimum wage is below the market clearing wage for low skill labor at t, the minimum wage exposure for every firm is 0. Now, $\underline{w}_{l,t+1}$ increases above the market clearing wage level. Then, firm j's exposure to the minimum wage hike is $FA^{l,\text{model}}_{j}$ when firm j belongs to the sector l ($l \in \{m,s\}$).

Since wage for high skill labor (w_h) and the rental rate for capital service (r) are perfectly flexible and the total supply for these production factors are unchanged, $w_{h,t+1}$ and r_{t+1} adjust to balance demand and supply. Hence, $L_{H,j,t+1} = L_{H,j,t}$ and $K_{j,t+1} = K_{j,t}$, implying that all the demand adjustment happens at $L_{D,j,t+1}$ because wage for low skill labor ($\underline{w}_{l,t+1}$) is set by the rule exogenous to labor market conditions and does not change to equate demand to supply at $t + 1$.

The firm-level total domestic employment change can be decomposed as follows:

$$\frac{L^{l}_{D,j,t+1}+L^{l}_{H,j,t+1}}{L^{l}_{D,j,t}+L^{l}_{H,j,t}}-1=\frac{L^{l}_{D,j,t}}{L^{l}_{D,j,t}+L^{l}_{H,j,t}}\left(\frac{L^{l}_{D,j,t+1}}{L_{D,j,t}}-1\right)=FA^{l,\text{model}}_{j,t}\left(\frac{L^{l}_{D,j,t+1}}{L^{l}_{D,j,t}}-1\right),$$

$$=FA^{l,\text{model}}_{j,t}\left(\left(\frac{w_{h,t+1}/w_{h,t}}{\underline{w}_{l,t+1}/w_{l,t}}\right)^{\frac{1}{\epsilon^{l}}}-1\right).$$

In driving the above result, we used the fact that the relative demand for labor adjust only through changes in low skill labor not high skill labor as well as relative demand shifts in equation (12) and (14).[29] With the minimum wage increasing above the market clearing wage, the relative wage of low

[29] The domestic minimum wage can also affect the margin between domestic low skill labor and foreign low skill labor but we leave $\frac{L^{l}_{F,j}}{L^{l}_{D,j}}$ unchanged to separate the role of the firm-level heterogeneity in the comparative advantage of high skill labor from that of the firm-level heterogeneity in the adjustment cost.

skill labor increases from t to t+1. Therefore, the coefficient on $FA_{j,t}^{l,\text{model}}$ is negative with $\left(\frac{w_{h,t+1}/w_{h,t}}{\underline{w}_{l,t+1}/w_{l,t}}\right)^{\frac{1}{\epsilon^l}} < 1$.[30]

While **Proposition 2** establishes the negative employment effect of the minimum wage hike based on the variation in the firm-level exposure, it does not consider the substitution of foreign low skill labor for domestic low skill labor via offshoring tasks. In our empirical study, we find that more manufacturing firms adjust along the extensive margin when the minimum wage increases rapidly. For multinational firms, offshoring factories turned out to be a significant channel of adjustment as noted by Ahn, Choi and Chung (2023). We add adjustment costs to changing the task threshold for foreign low skill labor $i_m^\star$ and allow the heterogeneity across firms in terms of the adjustment cost specification. Suppose that there is an adjustment cost per output of changing the task threshold for firm j as follows:

$$\Phi_j\left(i_{m,t}^\star, i_{m,t+1}^\star\right) = f_j. \tag{17}$$

While firms share the common comparative advantage of domestic low skill labor over foreign low skill labor, the difference in adjustment costs changes the unit cost for each firm depending on its decision on altering the threshold.

Now, we consider the two-period model with this heterogeneity in adjustment cost. We suppose the binding minimum wage rises from $\underline{w}_{l,1}$ to $\underline{w}_{l,2}$. Firms are heterogeneous in terms of adjustment costs drawn from a random distribution $F(\cdot)$. **Proposition 3** below shows that under the maintained comparative advantage assumption our model implies the rapid increase in the minimum wage increases the fraction of firms offshoring. Figure 12 illustrates how the fraction of firms offshoring tasks is determined as the minimum wage increases.

Proposition 3. *As the domestic minimum wage binding for low skill workers increases more rapidly, the fraction of firms offshoring increases.*

Proof. From equation (11), we have

30 $\frac{w_{h,t+1}/w_{h,t}}{\underline{w}_{l,t+1}/w_{l,t}}$ is likely to be less than 1 because the minimum wage pass-through to the high-skill worker wage is not perfect due to the imperfect substitutability.

$$\frac{A_{L,D}^{m}(i_m^\star)}{A_{L,F}^{m}(i_m^\star)} = \frac{\underline{w_l}}{\underline{w_f}} \tag{18}$$

Hence, if $\underline{w_l}$ increases with no change in $\underline{w_f}$, $i_m^\star$ must increase with the comparative advantage assumption that we made with $i_{m,2}^\star > i_{m,1}^\star$.[31] Consider $\underline{w_l}$ increases from $\underline{w_{l,1}}$ to $\underline{w_{l,2}}$ ($\underline{w_{l,2}} > \underline{w_{l,1}}$). In the second period, firms have two options. The first option is to pay the adjustment cost and re-optimize the task threshold based on the new minimum wage. The second option is not to change the task threshold and stick to the task threshold chosen in the first period. Let $i_{m,1}^\star$ and $i_{m,2}^\star$ be values of the task threshold with the minimum wage $\underline{w_{l,1}}$ and $\underline{w_{l,2}}$, respectively under the static optimization. Hence, the unit production cost will be lower at $i_{m,2}^\star$ compared to $i_{m,1}^\star$ if the minimum wage is $\underline{w_{l,2}}$. Firm *j* will change the task threshold in the second period if the benefit from offshoring outweighs the fixed adjustment cost as follows:

$$B\left(i_{m,1}^\star, i_{m,2}^\star\right) = \int_{i_{m,1}^\star}^{i_{m,2}^\star} \left[\frac{\underline{w_{l,2}}}{A_{L,D}(i)} - \frac{\underline{w_f}}{A_{L,F}(i)}\right] di > f_j. \tag{19}$$

Notice that by the definition of threshold,

$$\frac{\underline{w_{l,2}}}{A_{L,D}\left(i_{m,1}^\star\right)} > \frac{\underline{w_{l,1}}}{A_{L,D}\left(i_{m,1}^\star\right)} = \frac{\underline{w_f}}{A_{L,F}\left(i_{m,1}^\star\right)}, \tag{20}$$

$$\frac{\underline{w_{l,2}}}{A_{L,D}\left(i_{m,2}^\star\right)} = \frac{\underline{w_f}}{A_{L,F}\left(i_{m,2}^\star\right)}. \tag{21}$$

Given the comparative advantage assumption that we made ($ln\frac{A_{L,D}^{m}(i)}{A_{L,F}^{m}(i)}$ is increasing in *i*), the term inside the bracket in equation (19) is always positive for $i \in \left(i_{m,1}^\star, i_{m,2}^\star\right)$. Combine this with the fact that $i_{m,2}^\star - i_{m,1}^\star$ increases with $\frac{\underline{w_{l,2}}}{\underline{w_{l,1}}}$, we obtain the result that $B\left(i_{m,1}^\star, i_{m,2}^\star\right)$ also increases with $\frac{\underline{w_{l,2}}}{\underline{w_{l,1}}}$,. Suppose that f_j is distributed according to the distribution function $F(\cdot)$. Then the fraction of firms of changing the task threshold ($P(f_j < B(i_{m,1}^\star, i_{m,2}^\star)$) increases with $\frac{\underline{w_{l,2}}}{\underline{w_{l,1}}}$.

High-skill labor and capital imperfectly substitute low-skill labor, so the use of high skill labor and capital relative to low skill labor cannot be characterized by the threshold rule across task spectrum. Nonetheless, the increased relative price of low skill labor due to the rise in the minimum

[31] We assume foreign low-skill labor is abundant enough that $\underline{w_f}$ remains unchanged despite increased offshoring.

wage can deepen the relative intensity of these factors ($\frac{L_H^m(i)}{L_D^m(i)}$ or $\frac{K^m(i)}{L_D^m(i)}$) under the usual assumptions on the comparative advantage that $ln\frac{A_H^{m,g}(i)}{A_{L,D}^m(i)}$ and $ln\frac{A_K^m(i)}{A_{L,D}^m(i)}$ are increasing with respect to i.

For the substitution margin between capital and low skill labor, Yoon and Lee (2020) show that the rapid increase in the minimum wage in South Korea increased automation especially for the service sector by controlling the routine task intensity (RTI) constructed as shown in Autor and Dorn (2013). The finding suggests that the service sector may have a higher elasticity of substitution between labor and capital ($\frac{1}{\rho_s}$) than the elasticity of substitution between labor of different types ($\frac{1}{\epsilon^s}$) because a higher elasticity of substitution implies a larger adjustment in the relative intensity of input use in response to a change in the relative input price. While this can be also present for the manufacturing sector, the effect may be less noticeable in the manufacturing sector given the additional option to substitute foreign low skill labor for domestic low skill labor.

5. Conclusion

While a competitive labor market model implies the negative employment effect of a minimum wage hike, some recent empirical studies challenged the notion and found little or small positive employment effects. These studies often examine the change in the nation-wide minimum wage over the long period or the change in the local minimum change over the short period. However, both the time span and the scale of the labor market can limit overgeneralizing the conclusion found in these studies. In this paper, we leverage the unique experience in South Korea who raised the nation-wide minimum wage substantially over 2018 and 2019 after a sudden shift in the political environment. Using firm-level data, we find that the large increase in the minimum wage from 53 percent of the median wage to 63 percent had a significant negative employment effect. Firms adjusted to the minimum wage hike through the extensive (plant closing) margin as well as the intensive (layoff) margin. Accounting for both margins makes the estimated employment effect more negative. Consistent with the existing literature, the minimum wage hike unambiguously increased the average wage and depressed wage inequality. We also find that labor productivity in the manufacturing sector increased more for firms with higher exposure to the minimum wage.

Our findings have implications for both policy and the model of the labor market. Given a significant negative employment effect from minimum wage exposure and a large degree of heterogeneity in minimum wage exposure across industries and location, the minimum wage tied to these factors might be more effective in reducing the negative employment effect. For the model of the labor market, we find evidence for the significant amount substitution of low-skill workers subject to the minimum wage by other means of production even in the relatively short period when the minimum wage increased rapidly. We set up a task-based production model emphasizing the substitution of labor of different skill levels (and capital) that can be consistent with our empirical findings. Overall, our findings caution against extrapolating the effects of small and localized minimum wage hikes to large, nationwide minimum wage changes.

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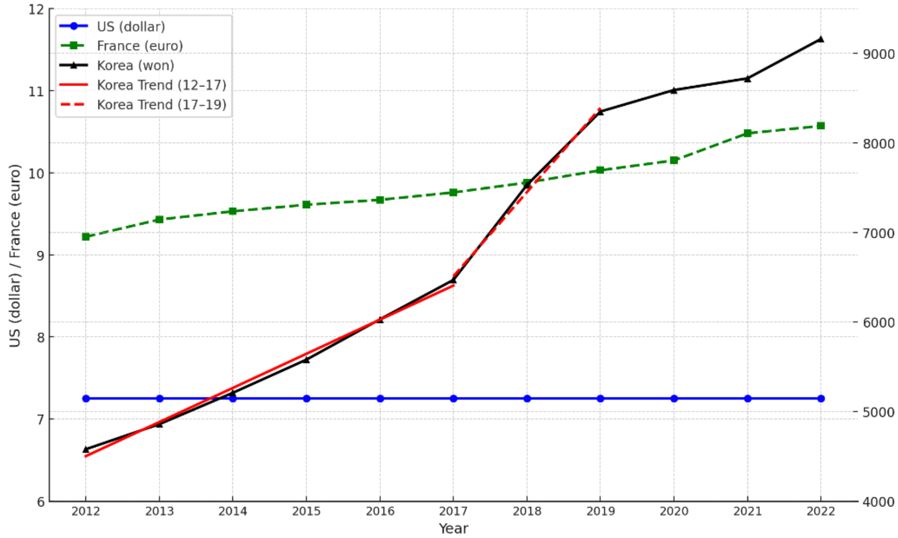
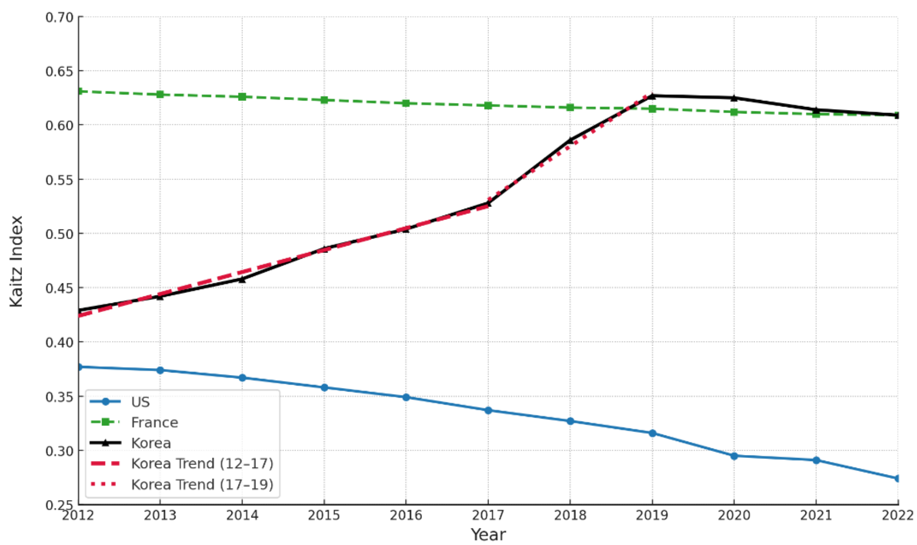
Figure 1: Time Series of the National Minimum Wage: France, South Korea, and U.S.



Sources: U.S. Department of Labor, Statista, Minimum Wage Commission (Korea)

Note: Figure 1 shows the annual changes in the minimum wage across different countries. Blue represents the United States, green represents France, and black represents South Korea. The two red lines depict the linear trends of South Korea's minimum wage for the periods 2012–2017 and 2017–2019, respectively.

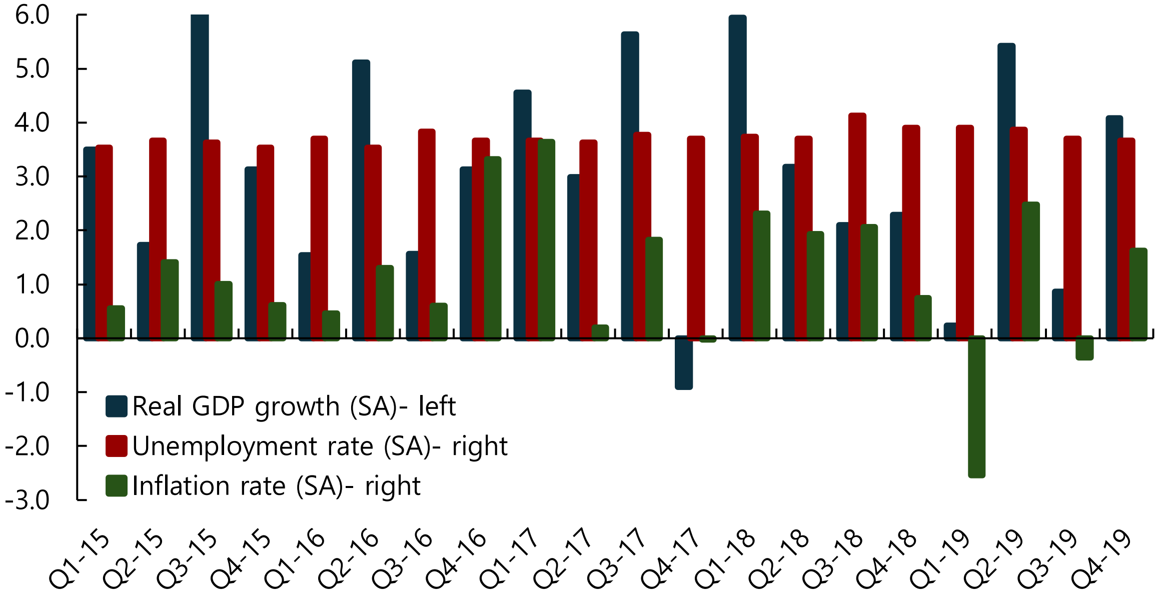
Figure 2: Time Series of the Kaitz Index in France, South Korea, and U.S.



Source: OECD Database

Note: Figure 2 shows the Kaitz index for each country. Blue represents the United States, green represents France, and black represents South Korea. The two red lines depict the linear trends of South Korea's Kaitz index for the periods 2012–2017 and 2017–2019, respectively.

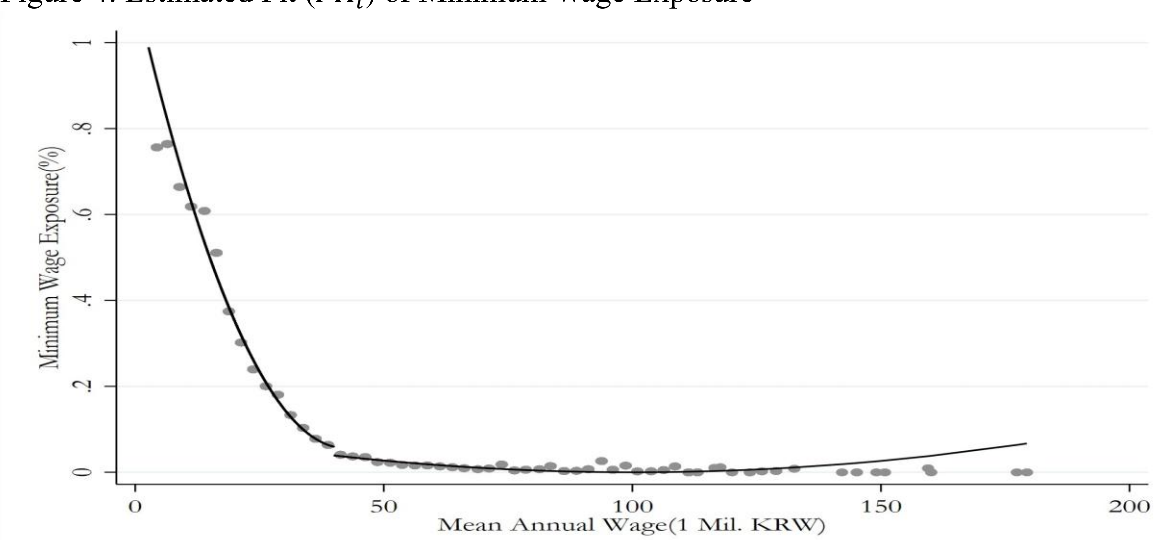
Figure 3: Macroeconomic Environment in South Korea: 2015-2019



Source: Bank of Korea, Statistics Korea

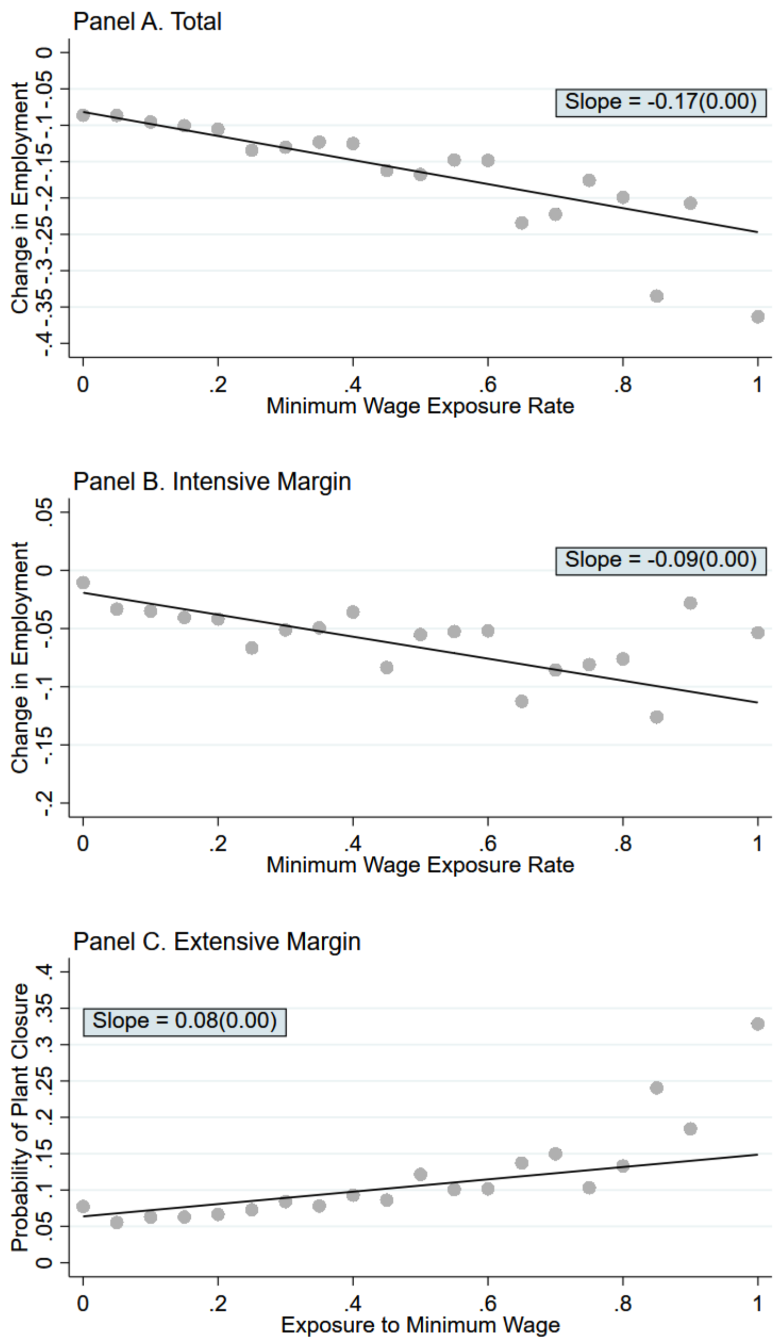
Notes: Figure 3 shows South Korea's quarterly real GDP growth (blue), unemployment rate (red), and inflation rate (green) from 2015 to 2019

Figure 4: Estimated Fit ($\widehat{FA}_i$) of Minimum Wage Exposure



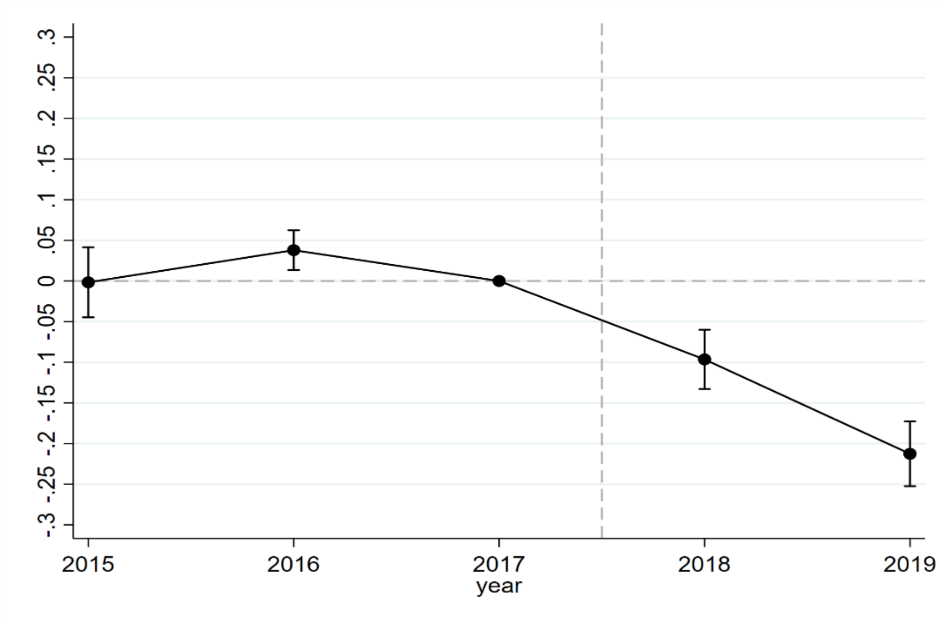
Notes: Figure 4 illustrates the nonparametric relationship between firms' annual average wages and minimum wage exposure. Each gray circle represents the average minimum wage exposure for firms grouped into 2.5 million KRW wage intervals. The black solid line depicts the relationship between the estimated minimum wage exposure, derived using equation (1), and annual average wages.

Figure 5: Employment Adjustment Patterns by Minimum Wage Exposure



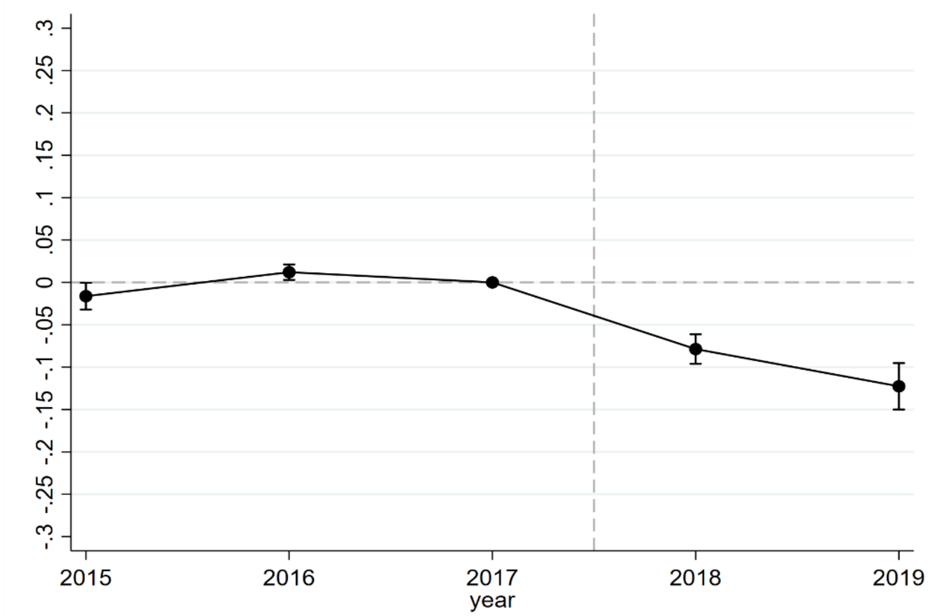
Notes: Figure 5 presents the nonparametric relationship between the minimum wage exposure rate and changes in employment outcomes between 2017–2019 in the mining and manufacturing sector. The x-axis measures the average minimum wage exposure. Firms are grouped into 5-percentage-point bins, and the y-axis shows the average change in employment outcomes (2017–2019) for each bin. Panel A reports the total change in employment at the establishment level. Panel B shows the change in employment among surviving establishments (i.e., intensive margin), while Panel C plots the probability of plant closure (i.e., extensive margin). Each panel includes a linear regression line , with slopes estimated using firm-size-weighted least squares.

Figure 6: Yearly Employment Effect of the Minimum Wage Hike: Mining and Manufacturing Sector



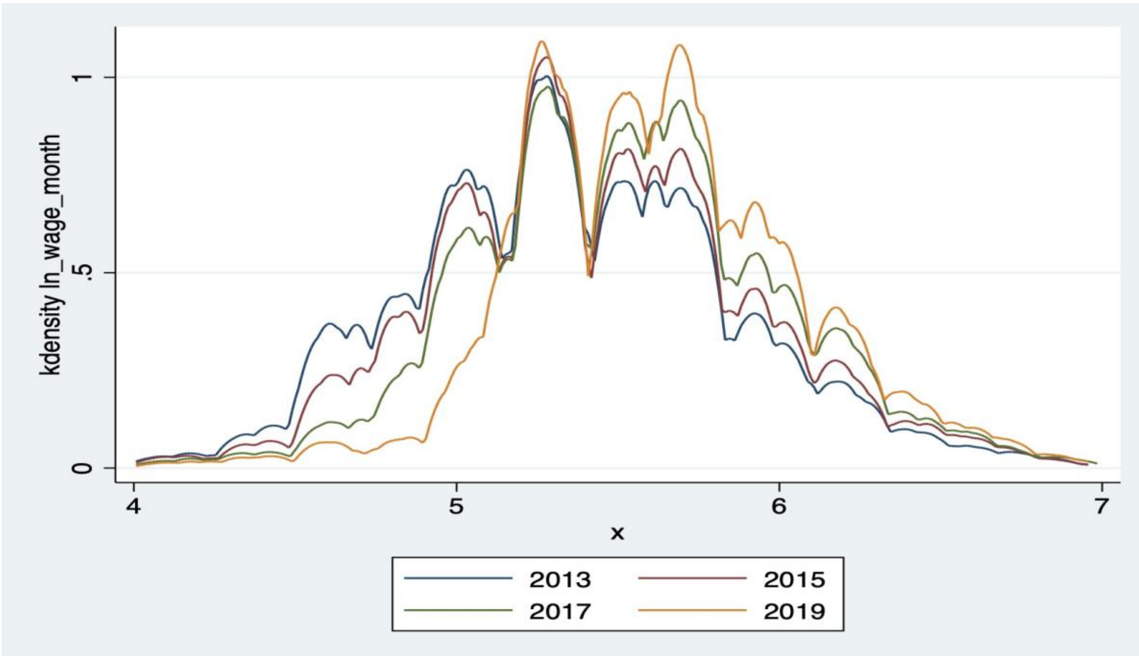
Notes: Figure 6 presents the yearly employment effects of the minimum wage hike in the MMS sample based on Equation (2). The black dots represent point estimates, with 95% confidence intervals provided.

Figure 7: Yearly Employment Effect of the Minimum Wage Hike: Service Sector



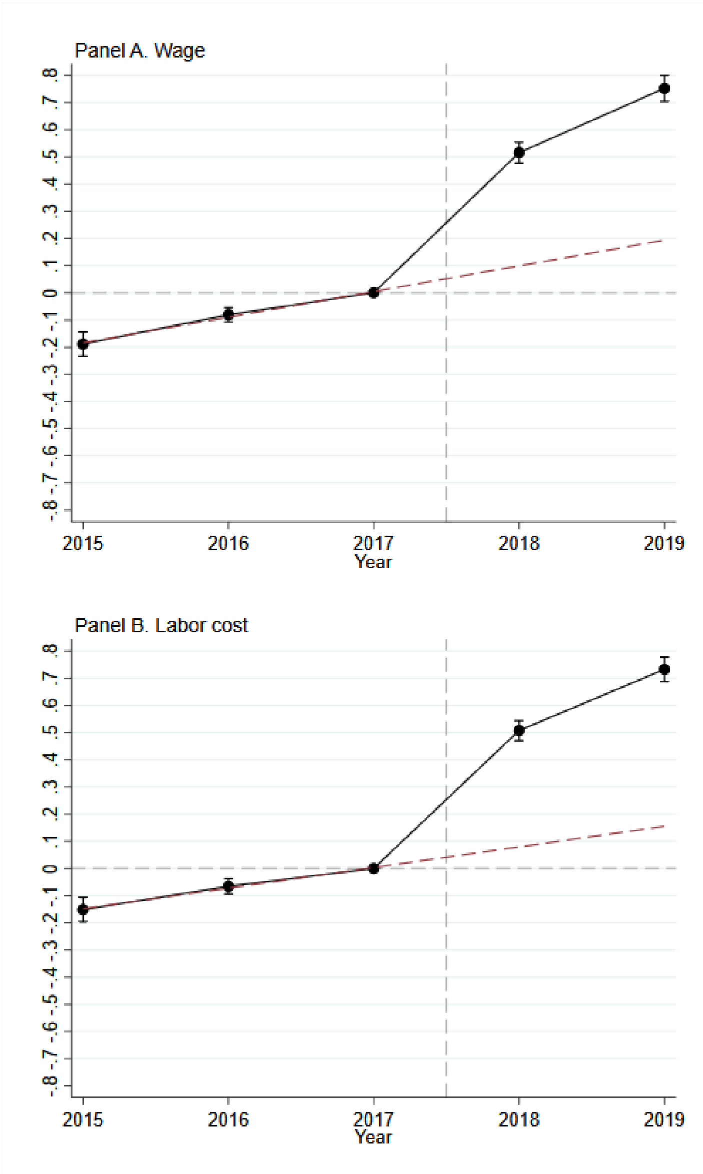
Notes: Figure 7 presents the yearly employment effects of the minimum wage hike in the EID sample, focusing on the service sector only, based on Equation (2). The black dots represent point estimates, with 95% confidence intervals provided.

Figure 8 Effect of the Minimum Wage Hike on Wage Dispersion in the Mining and Manufacturing Sector



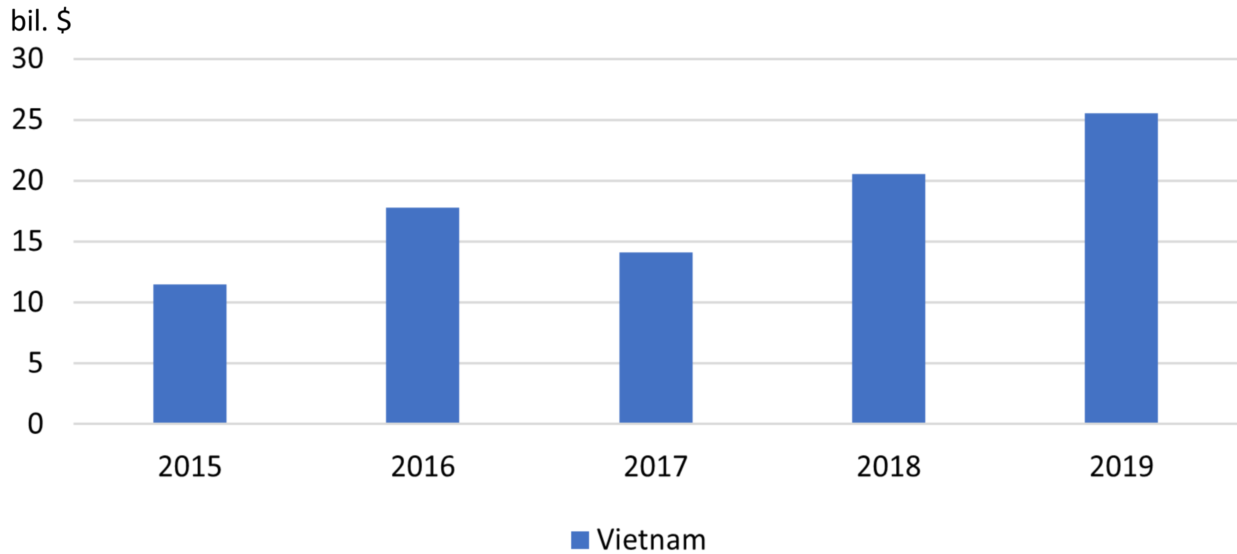
Notes: Figure 8 shows the log wage distribution for 2013, 2015, 2017, and 2019 using the Employment Positions Statistics (EPS) sample.

Figure 9. Yearly Effects of Minimum Wage Hikes on Wages and Labor Costs: Mining and Manufacturing Sector



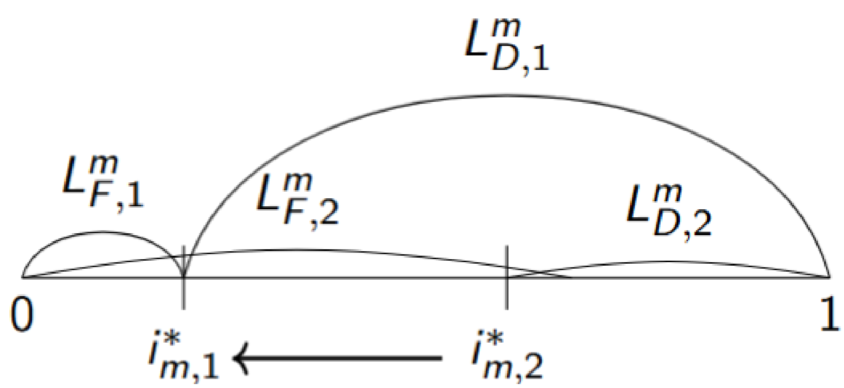
Notes: Panel A of Figure 9 presents the yearly wage effects of the minimum wage hike, while Panel B shows the corresponding labor cost effects, using the MMS sample and based on Equation (2). Black dots indicate point estimates, with 95% confidence intervals. Labor cost is defined as the sum of wages and non-wage benefits at the firm level. The red dashed line represents values predicted by fitting a linear trend to the estimated β_t coefficients from the 2015–2017 period.

Figure 10: South Korea's Foreign Direct Investment in Vietnam (Manufacturing Sector)



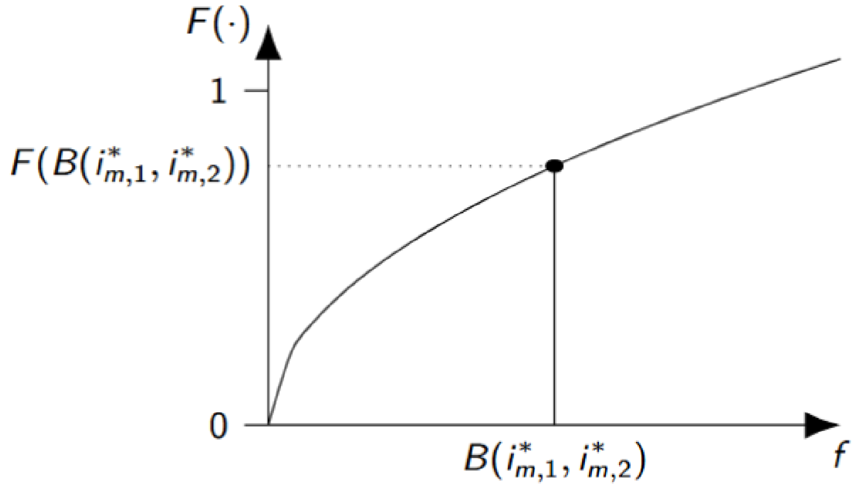
Notes: Figure 10 shows South Korea's foreign direct investment in Vietnam's manufacturing sector from 2015 to 2019, based on aggregate data from the Export-Import Bank of Korea

Figure 11: Task Threshold Change After a Minimum Wage Hike



Notes: Figure 11 describes the task threshold change when the minimum wage increases from $\underline{w}_1$ to $\underline{w}_2$. $L^m_{D,1}$ and $L^m_{F,1}$ describe domestic and foreign low skill labor in the manufacturing sector at $\underline{w}_1$, respectively, while $L^m_{D,2}$ and $L^m_{F,2}$ describe domestic and foreign low skill labor at $\underline{w}_2$

Figure 12: Fraction of Firms Offshoring After a Minimum Wage Hike



Notes: Figure 12 shows the cumulative distribution function of the adjustment cost for firms who decide to offshore factories. $B(i^*_{m,1}, i^*_{m,2})$ is the benefit of offshoring when the minimum wage increases from $\underline{w}_1$ to $\underline{w}_2$. The vertical part describes $F(B(i^*_{m,1}, i^*_{m,2}))$ which is the fraction of firms whose adjustment cost is lower than the benefit of offshoring.

Table 1: Average Minimum Wage Exposure in 2017 by Sector

Sector	Mean	Std. Dev.
Mining	0.066	0.130
Manufacturing	0.128	0.211
Electricity, gas, steam and air conditioning supply	0.016	0.042
Accommodation and food service	0.404	0.316
Economy Average	0.141	0.240

Notes: Table 1 shows the average minimum wage exposure by sector in 2017. These figures are calculated based on the 2017 EPS, with sectors classified according to the one-digit Korea Standard Industry Code.

Table 2: Employment Effect of a Minimum Wage Hike

	2019-2017			2015-2017		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. Total Effect (Extensive + Intensive Margin)						
$\widehat{FA}_i$	-0.187**	-0.193**	-0.206**	0.001	-0.002	-0.016
s.e.	0.022	0.022	0.023	0.023	0.023	0.022
Obs.	50,854	50,854	50,854	50,854	50,854	50,854
Panel B. Intensive Margin						
$\widehat{FA}_i$	-0.111***	-0.116***	-0.132***	0.001	-0.002	-0.021
s.e.	0.021	0.020	0.022	0.022	0.023	0.022
Obs.	47,965	47,965	47,965	47,965	47,965	47,965
Size weighted		Y	Y		Y	Y
Fixed effects			Y			Y

Notes: Table 2 presents the two-year change in employment due to the minimum wage hike during the baseline period (2017–2019) and the placebo period (2015–2017). Clustered standard errors at the district level are reported. ∗∗∗p<0.01, ∗∗ p<0.05, ∗ p<0.1.

Table 3: Decomposition of Sample Attrition by Employment Size

	2017-2018		2018-2019	
Employment Size	closing	job reduction	closing	job reduction
10 -15	19.79%	80.21%	33.5%	66.5%
16 - 25	57.99%	42.01%	55.61%	44.39%
26 - 50	79.62%	20.38%	73.12%	26.88%
over 50	92.49%	7.51%	83.64%	16.36%

Note: Table 3 presents the decomposition of sample attrition by employment size. The ratio of firms that closed plants among those that dropped out of the MMS sample in 2018 and 2019 is calculated using data from EID.

Table 4: Employment Effect of a Minimum Wage Hike: Robustness Check to Sample Attrition Assumptions

	baseline	Alt 1	Alt 2	Alt 3
$\widehat{FA}_i$	-0.201^{***}	-0.292^{***}	-0.090^{***}	-0.225^{***}
s.e.	0.025	0.021	0.015	0.019
Obs.	50,854	50,854	45,122	50,854

Note: Table 4 presents the employment effect of a minimum wage hike based on alternative sample attrition assumptions. Clustered standard errors at the district level are reported. ∗∗∗p<0.01, ∗∗p<0.05, ∗p<0.1.

Table 5: Employment Effect of a Minimum Wage Hike: Manufacturing versus Service

	2019-2017			2015-2017		
	(1)	(2)	(3)	(4)	(5)	(6)
	All	Manu.	Service	All	Manu.	Service
Panel A. Total Effect (Extensive + Intensive Margin)						
$\widehat{FA}_i$	-0.149^{**}	-0.205^{**}	-0.134^{**}	0.009	-0.024^{**}	0.021
s.e.	0.020	0.012	0.025	0.011	0.010	0.013
Obs.	153,967	50,070	101,176	153,967	50,070	101,176
Panel B. Intensive Margin						
$\widehat{FA}_i$	-0.096^{**}	-0.133^{**}	-0.084^{**}	-0.003	-0.045^{**}	0.011
s.e.	0.018	0.011	0.023	0.012	0.011	0.015
Obs.	141,367	46,330	92,548	141,367	46,330	92,548

Note: Table 5 shows the employment effect of a minimum wage hike based on the EID sample. $\widehat{FA}_i$ is computed at the firm level using wage information of employees affiliated with the firm from 2015 to 2017. Clustered standard errors at the district level are reported. ∗∗∗p<0.01, ∗∗p<0.05, ∗p<0.1.

Table 6: Effect of a Minimum Wage Hike on Wage and Labor Cost: Exclude Drop-out Firms

	2019-2017			2015-2017		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. Average Wage Growth Rate						
$\widehat{FA}_i$	0.687^{***}	0.703^{***}	0.753^{***}	0.157^{***}	0.173^{***}	0.189^{***}
s.e.	0.024	0.025	0.024	0.024	0.022	0.023
Panel B. Average Labor Cost Growth Rate						
$\widehat{FA}_i$	0.674^{***}	0.687^{***}	0.732^{***}	0.112^{***}	0.131^{***}	0.152^{***}
s.e.	0.022	0.023	0.023	0.024	0.023	0.023
Obs.	45,122	45,122	45,122	45,122	45,122	45,122
Size weighted		Y	Y		Y	Y
Fixed effects			Y			Y

Notes: Table 6 shows the effect of a minimum wage hike on wages and labor costs. Only firms that remained in the MMS sample for all three years (Alt 2 in Table 4) are included. Clustered standard errors at the district level are reported. ∗∗∗p<0.01, ∗∗p<0.05, ∗p<0.1.

Table 7: Effect of a Minimum Wage Hike on Labor Productivity

	2019-2017			2015-2017		
	(1)	(2)	(3)	(4)	(5)	(6)
$\widehat{FA_i}$	0.424***	0.391***	0.426**	0.214***	0.210***	0.225***
s.e.	0.027	0.027	0.027	0.024	0.026	0.024
Obs.	45,122	45,122	45,122	45,122	45,122	45,122
Size weighted		Y	Y		Y	Y
Fixed effects			Y			Y

Notes: Table 7 shows the effect of a minimum wage hike on labor productivity. Only firms that remained in the MMS sample for all three years (Alt 2 in Table 4) are included. Labor productivity (LP) is measured as the log of output per labor input. Standard errors are clustered at the district level. ∗∗∗p<0.01, ∗∗p<0.05, ∗p<0.1.

Table 8: Determinant of Minimum Wage Exposure: Marginal R^2 of Each Characteristic

	Residence	Age	Education	Marriage	Sex
Marginal R^2	0.0126	0.2879	0.0738	0.0135	0.0427

Notes: Table 8 presents the dominant characteristics of workers most affected by the minimum wage hike, based on the marginal R^2 from each characteristic following Cengiz et al. (2022), using the LALFS sample.

Table 9: Groupwise Probability of Being Employed and the Minimum Wage Exposure: 2011-2020

Coefficient	(1)	(2)
Average Probability of Being Exposed	-0.0754**	-0.268**
s.e.	0.0358	0.0289
Lagged Average Probability of Being Exposed	-0.102***	-0.341***
s.e.	0.0322	0.0296
Size Weighted	N	Y

Notes: Table 9 presents the dis-employment effect of minimum wage exposure, using variations in the average probability of exposure across 1,358 subgroups based on the LALFS sample. Clustered standard errors at the group level are reported ∗∗∗ p<0.01, ∗∗ p<0.05, ∗ p<0.1.

Appendix

A.1 Heterogeneity in the Minimum Wage Exposure by Group (two-digit KSIC)

There is a large degree of heterogeneity in the average minimum wage exposure by the detailed industry code and location. Table A1 shows the average minimum wage exposure in 2017 by the two-digit Korea Standard Industry Code and location for the manufacturing sector.

Table A1. Average Minimum Wage Exposure in 2017 by Division (the two-digit Korea Standard Industry Code) and Location for the Manufacturing Sector

Division.	Seoul	Busan	Daegu	Incheon	Gwangju	Daejeon
Food Products	0.2	0.33	0.33	0.24	0.37	0.34
Beverages	0.06	0.00	0.18	0.00	0.12	0.00
Textiles, Except Apparel	0.24	0.26	0.24	0.47	0.09	0.56
Wearing Apparel, Clothing Accessories and Fur Articles	0.39	0.59	0.64	0.42	0.55	0.74
Products of Wood and Cork Except Furniture	0.00	0.05	0.11	0.09	0.76	0.45
Pulp, Paper, and Paper Products	0.09	0.27	0.15	0.28	0.24	0.05
Printing and Reproduction of Recorded Media	0.18	0.24	0.22	0.21	0.44	0.07
Chemicals and Chemical Products. Except Pharmaceuticals, Medicinal Chemicals	0.04	0.04	0.14	0.16	0.20	0.08
Pharmaceuticals, Medicinal Chemicals, Botanical Products	0.03	0.05	0.10	0.03	0.22	0.09
Rubber and Plastic Products	0.14	0.20	0.25	0.34	0.12	0.00
Other Non-metallic Mineral Products	0.04	0.09	0.01	0.10	0.03	0.21
Basic Metal Products	0.00	0.07	0.11	0.04	0.06	0.19
Fabricated Metal Products Except Machinery and Furniture	0.07	0.11	0.17	0.15	0.18	0.04
Electronic Components, Computer, Radio, Television, Communication Equipment	0.07	0.16	0.12	0.22	0.26	0.12
Medical, Precision and Optical Instruments, Watches and Clocks	0.11	0.11	0.26	0.12	0.09	0.04
Electrical Equipment	0.09	0.14	0.27	0.21	0.11	0.16
Other Machinery and Equipment	0.08	0.09	0.09	0.09	0.16	0.08
Motor Vehicles, Trailers, Semitrailers	0.00	0.17	0.20	0.10	0.08	0.12
Furniture	0.07	0.12	0.38	0.08	0.37	0.00
Other	0.24	0.27	0.24	0.29	0.36	0.36

Notes: Table A1 presents the average minimum wage exposure in 2017 by division (two-digit Korea Standard Industry Code) and location for the manufacturing sector using LALFS. Location refers to the six largest metropolitan cities in South Korea.

A.2 Price Pass-through Difference across Sectors

Harasztosi and Lindner (2019) suggest that firms in the less competitive sector (e.g., non-tradable services) are more likely to pass-through minimum wage increase to prices while those in the more competitive sector (e.g., tradable goods) are less likely to do so. We do not have granular firm-level product price data to directly test this channel, but we have disaggregated produce price inflation (PPI) data at the two-digit industry classification codes. We regress PPI data onto the current and the lagged firm-level minimum wage exposure aggregated at the two-digit industry classification codes and examine if there was a significant difference in the degree of price pass-through across the manufacturing sector and the service sector.

Table A2 shows the sectoral difference in the degree of price pass-through from minimum wage hikes. For the manufacturing sector, we do not find any significant price pass-through for the whole sample period (2011-2019) and two sub-sample periods (2011-2017, 2017-2019). The statistically significant price pass-through is observed for the service sector but this holds only for the period including the pre-2017 sample. Hence, the price pass-through channel is unlikely to explain the relatively muted difference in the dis-employment effect in the service sector during 2017–2019 compared to the manufacturing sector, as shown in column (2) and (3) of Table 5. The absence of the price pass-through channel suggested by Table A2 in 2017-2019 should amplify the dis-employment effect from the rapid increase in the minimum wage, as firms could not offset minimum wage increases through higher prices. With lower price sensitivity, firms likely absorbed more costs, leading to greater job losses.

Table A2. Price Pass-through from Minimum Wage Hikes

Dependent Variable: $lnPPI_{k,t}$	Manufacturing Sector			Service Sector		
	2011-19	2017-19	2011-17	2011-19	2017-19	2011-17
	(1)	(2)	(3)	(4)	(5)	(6)
$\widehat{FA}_{k,t}$	0.1342	-0.0976	-0.3178	0.4502^{**}	0.2585	0.4696^{*}
s.e.	0.5668	0.2955	0.7137	0.1846	0.1701	0.2805
$\widehat{FA}_{k,t-1}$	0.1212	-0.1704	0.2004	0.4115^{**}	-0.1317	0.6835^{**}
s.e.	0.2630	0.2371	0.3595	0.1998	0.2376	0.3296
Obs.	384	144	288	384	144	288

Note: Table A2 presents the effect of minimum wage hikes on price pass-through, using PPI data from the Bank of Korea and individual minimum wage exposure (binary indicators for wages below 1.2×the minimum wage) aggregated at the two-digit industry level (k represents the industry index) from the LALFS. Clustered standard errors at the two-digit industry level are reported ∗∗∗p<0.01, ∗∗p<0.05, ∗p<0.1.