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Taeyoung Doh, Kyoo il Kim, Sungil Kim, Hwanoong Lee, and Kyungho Song

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

Taeyoung Doh† Kyoo il Kim‡ Sungil Kim§ Hwanoong Lee¶

Kyungho Song ‖

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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. 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). 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.
JEL Classifications: J23, J31, J42.

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†Federal Reserve Bank of Kansas City;taeyoung.doh@kc.frb.org
‡Michigan State University;kyookim@msu.edu
§Duke University;sungil.kim@duke.edu
¶Korea Institute of Public Finance;holee@kipf.re.kr
‖Korea Institute of Public Finance;kyungho@kipf.re.kr

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 is in contrast to 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. Manning (2021) suggests that researchers should focus more on the turning point level of the minimum wage in which a significant negative employment effect is expected to emerge. Empirically identifying this nonlinear turning point effect is challenging because the estimated employment effect can be subject to significant uncertainty when a modest increase in the minimum wage only affects a small number of firms. Hence, to empirically test the nonlinear employment effect of the minimum wage hypothesis, it might be useful to have comprehensive micro 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 and 2019 provides this ideal ground to test a nonlinear employment effect of the minimum wage. 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 was in stark contrast to the previous two-year period (2015 and 2016) when the increase was modest. In addition, the minimum wage outpaced the increase in the typical wage as the Kaitz index that computes the level of the minimum wage relative to the median wage shows in Figure 2. This sudden shift in the pace of the minimum wage increase was mainly driven by an unexpected change in the political environment which with some confidence we can take largely exogenous to economic factors.[1]

In this paper, we use the firm level data on employment and wages to empirically test if this rapid increase in the minimum wage brought about a nonlinear employment effect. 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)

[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.

than other sectors. However, 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 minimum wage hike, we follow Harasztosi and Lindner (2019)'s 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 onto this exposure measure. Specifically, we compare the two year change from 2017 to 2019 in employment with the previous two year change from 2015 to 2017 to isolate the firm-level employment effect due to the minimum wage hike in 2018 and 2019 where the previous period plays a role of placebo test.

Our main empirical findings are summarized as follows. First, we find evidence for a significant and negative employment effect from the minimum wage hike. 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 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 disemployment was driven by firms who were previously little exposed to the minimum wage as they started to adjust employment following the minimum wage hike. On the other hand, for firms whose exposure to the minimum wage was similar between pre- and post-2017 period, the magnitude of employment adjustment did not change much since for them the shock was already absorbed. This finding suggests that the absence of the negative employment effect from a modest increase in the minimum wage could be mainly driven by a higher estimation uncertainty because not many firms were newly exposed to the shock when the minimum wage increased modestly and hence the treatment was binding only for a small number of firms. This result is not confined to the manufacturing sector as we also find a significant and negative employment effect in the service sector due to the minimum wage hike when we restrict our sample to those for whom we have information on similar control variables we use in the estimation with firms in the manufacturing sector.

Second, when we decompose the employment adjustment into the intensive margin (layoffs) and the extensive margin (plant closings), the extensive margin is estimated to contribute about 38.1% to the overall adjustment. The adjustment through the extensive margin increases with firm size for those whose employment level declines below the MMS sample cutoff value (10 people). 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. Finally, 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 although the firm level adjustment of offshoring tasks might be non-linear due to the presence of adjustment cost, the economy-wide effect aggregated over firms with heterogeneous adjustment costs can still generate a smooth linear effect from the minimum wage hike. 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. It also highlights the importance of controlling micro level heterogeneity in addressing questions related to challenging policy evaluation questions like the nonlinear employment effect 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 (2021), and Haanwinckel (2020)). In terms of methodology, we closely follow Harasztosi and Lindner (2019) and construct the firm-level minimum wage exposure measure based on the fraction of workers subject to the minimum wage. Like South Korea, the Hungarian minimum wage substantially increased both in terms of the level and the relative share to the median wage between 2000 and 2002. Compared to the 1998-2000 period when the minimum wage hike was modest, Harasztosi and Lindner (2019) find a significant and negative employment effect during the 2000-2002, when the minimum wage increased substantially, in line with our findings.[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 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. Table 6 in Harasztosi and Lindner (2019) shows that firms in the tradable manufacturing sector have about 10 times bigger disemployment effect than firms in the non-tradable service sector. We do not see such

[2] Harasztosi and Lindner (2019) 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. We also find a similar pattern from the South Korean dataset. However, if a rapid increase in the minimum wage increases the number of exposed firms more than linearly, it is possible that the employment effect of the minimum wage increase can be depicted as nonlinear.

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]

The substitution channel works for both the manufacturing and the service sector in our analysis but the difference we find across sectors is that the substitution of foreign low skill labor for domestic low skill labor contributed significantly for the employment adjustment in the manufacturing sector while automation of tasks done by low skill labor seemed to be accelerated in the service sector after the minimum wage hike in 2018 and 2019.

Engbom and Moser (2021) 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 (2020)). 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 disemployment 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 for such a reallocation hypothesis and argue that the introduction of the minimum wage raised wages but did not lower 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 for reallocation from smaller to larger firms as Dustmann et al. (2022) find from the German data. 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 (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

[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).

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, which was 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 by wage spillovers at the bottom of the wage distribution. They do not find a big role for labor-labor substitution but acknowledge that the finding might be attributed to the somewhat limited variation in the minimum wage-between 37% and 59% of the median wage. In contrast, we find empirical support for the labor-labor (or labor-capital) substitution channel as the minimum wage surpasses 60% of the median wage in South Korea.[4]

Recently, Derenoncourt et al. (2021) examine the spillover effect of a company-level minimum wage hike in large U.S. retailers such as Amazon, Costco, Target, and Walmart on the local labor market. In the case of Amazon, which introduced a $15 minimum wage in October 2018, there was a significant and positive wage spillover effect in the local labor market as well as a small but statistically significant negative employment effect unlike Cengiz et al. (2019). This finding is consistent with the significant and negative employment effect of the minimum wage hike we find in the South Korean data.

Our paper is closely related to Kreiner, Reck and Skov (2020) who use the discontinuous jump in the minimum wage at age 18 to estimate the employment effect of minimum wage on young workers. They find a substantial jump in the hourly wage (by 40%) and a significant decline (by 33%) in employment at the discontinuity. They reconcile empirical findings in the context of an equilibrium labor market search model in which individual workers differ in match-specific productivity. While our empirical findings are largely in line with theirs, we rationalize our findings through the lens of a task-based production model with firm level heterogeneity and abstract from labor market search.

Ahn, Choi and Chung (2023) study the employment adjustment of multi-national enterprises in response to 2017-2019 minimum wage hikes in South Korea. They find that multi-national enterprises increased employment in low-wage countries like Vietnam after the minimum wage in South Korea rapidly increased. The finding is consistent with the negative employment effect that we find. However, their sample of multi-national enterprises include only firms with

[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 variation 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.

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 multi-national enterprises and non-multi-national enterprises and do not study the employment effect on non-multi-national enterprises, which will be important for assessing the overall employment effect. Our analysis complements their study by using a more comprehensive firm-level dataset.

The empirical magnitude of the employment effect of the minimum wage is often tied up with the magnitude of competition in the labor market and our paper is related to the literature on the frictional labor market that limits competition. In particular, a small positive employment effect found by Card and Krueger (2015) was regarded as evidence for the monopsony model of the labor market. When the employer has monopsony power, it pays wage below the marginal value of the product by a worker. Raising the minimum wage in this model can have a positive employment effect when the employer tries to offset lower profit per output by increasing production. As Harasztosi and Lindner (2019) point out, such a model typically implies that a firm may increase employment when the minimum wage exposure is small but reduce it when the minimum wage exposure is substantial. We do not find that the relationship between the firm-level minimum wage exposure and employment effect flips the sign as the magnitude of the minimum wage exposure changes. Hence, our findings can be more easily interpreted through the lens of a competitive labor market model that highlights the substitution between different production factors. However, we do not directly test the plausibility of the monopsonistic labor market model by estimating shadow wage markdowns as in Chen Yeh and Hershbein (forthcoming) or Berger, Herkenhoff and Mongey (2022) and our findings should not be interpreted as conclusive evidence against the monopsonistic labor market model in South Korea.

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. Using this extended model we examine the employment effect of the minimum wage hike. Firstly, we introduce the difference in the comparative advantage of domestic low skill labor over domestic high skill labor to generate difference in the minimum wage exposure across firms. Secondly, we allow the firm-level heterogeneity in the adjustment cost of reorganizing the task-difficulty threshold between domestic and foreign low skill workers, who are perfect substitutes. In our model, firms will reduce the demand for domestic low skill workers by reorganizing the task-difficulty threshold once the minimum wage hike becomes more substantial and the effect will be more notable for firms who are more exposed to the minimum wage hike. Consistent with the empirical findings, our model emphasizes the substitution between low skill labor and other production factors in determining the employment effect of the minimum wage hike.

Haanwinckel (2020) 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 military dictatorship. 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.

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 represents 28.8% of total employment in South Korea as of 2017. This annual survey provides data required for understanding the structure and distribution of the mining and manufacturing industries. Specifically, the data contains not only basic characteristics such as the type of legal organization, regional

[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.

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. We use the Employment Insurance Database (EID) to decompose sample attrition from the MMS sample into plant closing and layoff. Within-firm wage distribution information is supplemented by the Employment Positions Statistics (EPS). We use NICE (National Information & Credit Evalulation, a company specialized for credit rating) Information Service Corporate Information Data Service (CIDS) for datasets related to employment in service-providing sectors. We apply the same size threshold value (10 employees) and count firms that existed for whole sample period (2015-2019), so that we can compare the results with those using MMS data. Although average salary and employee size data are included, additional control variables used in MMS are missing in CIDS. This might cause overestimation of the true employment effect if omitted control variables like the per-worker capital are correlated with the minimum wage exposure of a firm.[6] For an alternative analysis using individual characteristics of workers who are susceptible to the minimum wage hike, we use the Local Area Labor Fore 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 minimum wage context, if strong economic growth and labor shortage may alleviate the typical objection of the side that represents the employer on the minimum wage commission against the rapid increase in the minimum wage, we may not observe a substantial disemployment effect. The plot of the time series of a few key macroeconomic variables in Figure 3 shows that there is no clear evidence that economic growth was strong in 2015-2017 relative to 2017-2019, so rules out this possible confounding effect.[7]

In fact, the main reason for the rapid increase in the minimum wage was a sudden shift in

[6]For robustness, we do an empirical analysis for the sub-sample where we can obtain the same control variables as in the MMS.

[7]Average real GDP growth per year during 2015-2017 was 2.9, the same as in the 2018-2019 period. The average unemployment rate during 2015-2017 was 3.7 percent, slightly lower than 3.8 percent in the 2018-2019 period. The average inflation rate was 1.2 percent during 2015-2017 while it was 1 percent in the 2018-2019 period.

the balance of political power. The conservative president, Park Geun-hye, was impeached over a personal scandal and removed from the office on December 2016.[8] 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%) 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 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 significantly slowed down during the second half of President Moon's tenure.[9] The minimum wage rose by about 9% 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.[10]

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. However, we augment the analysis by including the different sample data of the service sector for the employment effect.

As in Harasztosi and Lindner (2019), we use a difference-in-difference method where we compare the difference in the 2017-2019 change in employment between firms more exposed to the minimum wage hike and firms less exposed to it with the difference in the 2015-2017 change in employment across same firms. We adopt the definition of the firm-level exposure to the 2019 minimum wage in Harasztosi and Lindner (2019), which is the fraction of workers earning at or below the minimum wage (FA_i for the i-th establishment).

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

[8] See `https://www.britannica.com/biography/Park-Geun-Hye`

[9] 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`.

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

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 provides the average wage information within a firm but not the distribution of wage at the firm level. The EPS data provides 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.

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.[11]

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

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 milllion Korean won (KRW).[12] 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 and define the minimum wage exposure ratio as the fraction of workers who are subject to the 2019 minimum wage. 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 calculate regression coefficients weighted by firm size. The baseline regression in our study is as follows:

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

where $y_{i,t}$ denotes the outcome variable of interest such as employment, wage, and labor productivity. $X_{i,2017}$ denotes the aforementioned 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 nonlinear effect of the minimum wage hike

[11]The functional form closely follows the one in Harasztosi and Lindner (2019) but we add an inflection point θ to better fit our dataset.

[12]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).

during the 2018-2019 period. Any significant change in the regression coefficient on the minimum wage exposure between the 2017-2019 period and the placebo sample 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 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). It is clear that 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 have to decide how to treat firms that drop out of the 2017 MMS sample in order to disentangle intensive margin 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.[13] By comparing the 2017 MMS sample with the replicated MMS sample from the EID in 2018 and 2019, we can

[13]Since industrial classification across firms are 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.

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.[14]

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 disemployment effect from the rapid increase in the minimum wage can be attributed to the adjustment through plant closing.

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 disemployment effect. Indeed, we observe a more severe disemployment 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 the MMS sample during the five years from 2015 to 2019. Alt 2 is likely to understate the true disemeployment effect of the minimum wage hike by ruling out some firms who made drastic reduction in the employment level. As expected, the magnitude of the disemployment 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 fired some workers. Despite this difference, the

[14] 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.

estimated magnitude of the disemployment effect of the minimum wage exposure in Alt 3 is similar to the baseline case.

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 the common trend but not for diverging trends across firms with different minimum wage exposure levels. 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 5. 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.[15] However, the magnitude is relatively small compared to the yearly effect in 2018.[16] This finding suggests that the violation of the common trend assumption is unlikely to be a decisive factor in explaining the significant disemployment effect of the minimum wage hike after 2017. An 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 who are exposed to the minimum wage hike may be indirectly affected. However, we do not find evidence that the minimum wage hike substantially moved the upper spectrum of wage distribution as shown in Figure 7. This mitigates a concern that the SUTVA was violated in our sample.

Finally, 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.[17] We construct the sample in the service sector comparable to the MMS sample by using the CIDS data from the NICE information service. While the CIDS data incorporates information for firms with employees fewer than 10, it does not include any firm who exited out of business in any of the five years from 2015 to 2019. Hence, we can only compare the adjustment through the intensive margin, likely underestimating negative employment effect through the extensive margin. Based on the mean annual wage information, we can similarly construct the firm level minimum wage exposure for the service sector as we did for the MMS sample. The CIDS data does not necessarily provide information for the additional control variables ($X_{i,t}$ in equation (2)) that we use for the MMS sample and we restrict our analysis for firms that have financial information in Table 5.[18] We

[15]Since we regress the difference in employment level between 2016 and 2017, the positive coefficient means a disemployment effect with the level of employment lower in 2017 than in 2016.

[16]The coefficient is 0.04 in 2016 compared to -0.1 in 2018.

[17]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.

[18]When we restrict our sample to those who have more than 10 employees but may not show financial

still find a statistically significant negative employment effect from the minimum wage exposure in the service sector but the magnitude of the coefficient is closer the value in the manufacturing sector regression when we consider only the intensive margin.[19]

3.3.3 Wage and Labor Cost

As Manning (2021) notes, the effect of a minimum wage hike on wage and labor cost 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 provide information on the average wage for firms. 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.[20]

Engbom and Moser (2021) 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 the detailed information on firm-level wage distribution from the EPS data, we estimate a nonparametric kernel density for wage distribution in the mining and manufacturing sector in Figure 7. While we observe a large shift in the lower tail of the wage distribution as well as some upward drift in the middle, the upper tail is largely unaffected. Also, the magnitude of the shift is larger between 2017 and 2019 when the minimum wage increased rapidly compared to the previous period when it increased at a modest pace. This finding on the shift in the wage distribution during the period of the rapid increase in the

information, the statistically significant negative employment effect still exists. But in this case, we cannot rule out the pre-existing trend.

[19]Figure 6 shows that it is more difficult to rule out the existence of the pre-existing trend in the service sector. But the statistical significance of the minimum wage exposure's effect on the employment change between 2015 and 2017 is much weaker than the one between 2017 and 2019. In addition, the lack of the extensive margin adjustment in the sample might have understated the negative employment effect between 2017 and 2019.

[20]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 between the 2015 year labor cost and the 2017 labor cost as the dependent variable.

minimum wage is consistent with what Harasztosi and Lindner (2019) find from the 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 before the minimum wage hike (columns (4)-(6)) show that there existed a pre-trend of steadily increasing labor productivity until 2017. Because the minimum wage had been increasing modestly before the hike, it is likely that the minimum wage increase had positive effects on LP even before the hike. One concern is that this violates the parallel trend assumption. We, however, stress that our treatment is not the minimum wage introduction nor 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 disemployment 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 disemployment 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 $\times$ 7 age groups $\times$ 2 gender groups $\times$ 3 education groups $\times$ 2 marital status groups). 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 estimate of the negative employment effect from the rapid increase in the minimum wage is more than twice as strong as what Harasztosi and Lindner (2019) find in their paper despite using the essentially same methodology and a minimum wage hike of a similar pace. The regression coefficient of the change in employment on minimum wage exposure is -0.2 in our case compared to -0.078 in Harasztosi and Lindner (2019). The difference is much smaller if we restrict our comparison to the manufacturing sector only from Harasztosi and Lindner (2019). The large difference is mainly attributed to a much stronger disemployment effect in the service sector that we find. Our analysis indicates that firm owners largely took the burden of the minimum wage increase with limited pass-through to prices. In the appendix, we provide evidence that the degree of price pass-through was indeed limited using the industry-level producer price index data.

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 exposure, the relationship between minimum wage exposure and the change in the employment is roughly linear at each level as found by Harasztosi and Lindner (2019).[21] 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. This is clearly not the case as we see in Table 3. Even if every firm adjusts along

[21] In fact, we obtain a similar negative regression coefficient of the change in employment on the minimum wage exposure if we use the 2015 MMS sample and define $\widehat{FA}_i$ based on the 2017 minimum wage level. We provide the estimation results in the appendix. However, this does not imply that the rapid increase in the minimum wage has a linear employment effect because the rapid increase can alter the minimum wage exposure more than linearly.

the extensive margin based on its own cutoff value of the minimum wage exposure (e.g., keep a plant if the minimum wage exposure is below the cutoff value and close it otherwise), the average relationship that we obtain from the regression approach like equation (2) appears to be linear if firms have different cutoff values. Our theoretical model provides the prediction consistent with this pattern.

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] 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 new administration that came into power in May 2022 is currently considering the separate minimum wage depending on the location and the industry of firms. To the extent that the negative employment effect is driven by the degree of the minimum wage exposure as we find using the firm level data, some measures mitigating it can also alleviate the negative employment effect of the minimum wage hike.[23]

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 (2021) introduce an equilibrium model of labor market search highlighting 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 (2020) sets up an equilibrium task-based production model with labor market monopsony to explain the same pattern in the Brazilian data. 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. Neumark et al. (2008)

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

[23] Berger, Herkenhoff and Mongey (2022) study the optimal minimum wage using a rich model with worker and firm heterogeneity as well as oligopsony in the local labor market. The higher minimum wage is justified mostly based on the redistribution channel which reduces the markdown of wage below the marginal revenue of product rather than the efficiency channel that moves workers from low-productivity firms to high productivity firms. If available, the magnitude of the shadow wage markdown can provide information on the optimal minimum wage across location and industries.

emphasize the substitution channel for the disemployment effect of the minimum wage hike based on a competitive model of the labor market.

Harasztosi and Lindner (2019) argue that the linear relationship between the minimum wage exposure measure and the change in employment is not compatible with models with labor market monopsony. In such a model, a modest increase in the minimum wage should mostly affect the markdown in the wage and positively affect 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. The criticism would be likely valid for the model without the heterogeneity in the degree of wage markdown while not entirely rebutting against 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.[24] 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 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 indicates that a part of adjustment is in fact offshoring. While the detailed firm level data for offshoring comparable to our MMS sample is not available, the aggregate data available from the Export-Import Bank of Korea shows that the two-year growth rate in South Korean foreign direct investment in Vietnam rose significantly during 2017-2019 compared to the 2015-2017 period as Figure 8 shows. The offshoring of low-skill tasks in a task-based production model of Grossman and Rossi-Hansberg (2008) is consistent with this fact. Besides offshoring, if firms reorganize production to reduce the threshold of the task difficulty below which they hire low-skill workers subject to the minimum wage in spite of the adjustment cost associated with the reorganization, they can generate a significant negative disemployment effect for the rapid increase in the minimum wage together with a muted employment response for the modest increase for a given firm. However, the economy-wide effect aggregated over firms with heterogeneous adjustment costs may not exhibit such a stark discontinuity as we see below in the model, consistent with our empirical findings. Now, we propose a simple task-based production model that can generate this observed pattern from the South Korean data.

[24]However, this is only indirect evidence against labor market monopsony. A more direct comprehensive test of labor market monopsony can be done by computing the firm-level shadow markdown as in Chen Yeh and Hershbein (forthcoming). For future works, we plan to estimate the firm-level production function to back out the shadow markdown in the South Korean manufacturing sector though accounting for measurement errors in capital stock at the firm level is challenging.

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.[25] 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 (2020) 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 and 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

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:

$$
\begin{aligned}
Y_j &= e^{\int_0^1 \ln y_j(i) di}, \\
y_j(i) &= A_l(i) l_j(i) + A_h(i) h_j(i),
\end{aligned}
\tag{3}
$$

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

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

workers have comparative advantages over more difficult tasks with $?A_h(i)A_l(i)$ increasing with i. Notice that all the 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 skill (high skill) workers for task $i < i^\star$ ($i > i^\star$). The firm is indifferent between a low skill worker and a high skill worker for task $i^\star$ but we assume that it hires a low skill worker to break the tie. Firm j's demand for each type of worker at task i is determined as follows:

$$l_j(i) = \{ \begin{array}{ll} \frac{y_j(i)}{A_l(i)} & \text{if } i <= i^\star, \\ 0 & \text{if } i > i^\star, \end{array} , \; h_j(i) = \{ \begin{array}{ll} 0 & \text{if } i <= i^\star, \\ \frac{y_j(i)}{A_h(i)} & \text{if } i > i^\star. \end{array}$$

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 $(\overline{L}, \overline{H})$:

$$\begin{aligned} L(w_l, w_h) &= \sum_{j=1}^{J} \int_0^1 l_j(i)di = \overline{L}, \\ H(w_l, w_h) &= \sum_{j=1}^{J} \int_0^1 h_j(i)di = \overline{H}. \end{aligned} \tag{5}$$

Let's 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:[26]

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

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

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

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 $i^\star$. That gives us the new equilibrium value of w_h. The following proposition shows that the task threshold $i^\star$ 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:

$$\begin{aligned} \int_0^{i^\star} \ln(\frac{\underline{w}_l}{A_l(i)})di + \int_{i^\star}^1 \ln(\frac{A_h(i^\star)\underline{w}_l}{A_h(i)A_l(i^\star)})di = 0\,, \\ \ln(\underline{w}_l) - \int_0^{i^\star} \ln A_l(i)di - \int_{i^\star}^1 \ln A_h(i)di + \ln(\frac{A_h(i^\star)}{A_l(i^\star)})(1-i^\star) = 0\,, \\ \ln \underline{w}_l - g(i^\star) = 0. \end{aligned}$$

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

$$\frac{di^\star}{d\underline{w}_l} = g'(i^\star) = \frac{d(\ln(\frac{A_h(i^\star)}{A_l(i^\star)}))}{di^\star}(i^\star - 1) \leq 0.$$

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 same for all the firms. Combining these two observations, we obtain the total demand for low skill labor is $\frac{i^\star\overline{Y}}{\underline{w}_l}$ where $\overline{Y}$ is $\frac{w_h\overline{H}}{1-i^\star}$. 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 $\overline{H}$ 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 $\overline{H}$ because w_h adjusts to balance demand with supply. $\square$

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}.[27] Since demand for low skill workers declines but the supply of low skill workers does not change, the binding minimum wage creates the disemployment by $\overline{L} - L(\underline{w}_l, w_h)$. 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 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, 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. On top of that, 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.[28]

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

Firms in the manufacturing sector can use foreign low skill labor through off-shoring 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:

[27] 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).

[28] A similar specification is used in Haanwinckel (2020) to introduce the worker-level heterogeneity.

$$Y^l_j(i) = [\alpha_l^{\rho_l}(A^l_k(i)K^l_j(i))^{1-\rho_l} + (1-\alpha_l)^{\rho_l} L^l_j(i)^{1-\rho_l}]^{\frac{1}{1-\rho_l}} \,,\; l \in \{m, s\}, \tag{8}$$

$$L^m_j(i) = [(A^m_{L,D}(i)L^m_{D,j}(i) + A^m_{L,F}(i)L^m_{F,j}(i))^{1-\epsilon^m} + (A^{m,g}_H(i)L^m_{H,j}(i))^{1-\epsilon^m}]^{\frac{1}{1-\epsilon^m}} \,, \tag{9}$$

$$L^s_j(i) = [(A^s_{L,D}(i)L^s_{D,j}(i))^{1-\epsilon^s} + (A^{s,g}_H(i)L^s_{H,j}(i))^{1-\epsilon^s}]^{\frac{1}{1-\epsilon^s}}. \tag{10}$$

The minimum wage ($\underline{w}_l$) is binding for domestic low skill workers and the wage of the foreign low skill worker in the manufacturing sector in case of offshoring is given by $\underline{w}_f$. The wage for a high skill 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 $\ln(\frac{A^m_{L,D}(i)}{A^m_{L,F}(i)})$ 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 $i^\star_m$ from the manufacturing firm's cost minimization condition that separates tasks performed by domestic low skill labor ($i > i^\star_m$) and those by foreign low skill labor ($i \leq i^\star_m$). At $i^\star_m$, unit cost of hiring a domestic low skill worker to perform the task is same as the one of hiring a foreign low skill worker.

$$\frac{\underline{w}_l}{A^m_{L,D}(i^\star_m)} = \frac{\underline{w}_f}{A^m_{L,F}(i^\star_m)}. \tag{11}$$

Figure 9 shows how $i^\star_m$ 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, $i^\star_m$ increases as the minimum wage substitutes away from the demand for domestic low skill labor, creating the following demand function.

$$L^m_{F,j}(i) = \{ \begin{array}{ll} \frac{Y^m_j(i)}{A^m_{L,F}(i)} & \text{if } i <= i^\star_m, \\ 0 & \text{if } i > i^\star_m, \end{array} \,,\; L^m_{D,j}(i) = \{ \begin{array}{ll} 0 & \text{if } i <= i^\star_m, \\ \frac{Y^m_j(i)}{A^m_{L,D}(i)} & \text{if } i > i^\star_m. \end{array}$$

.

Unlike two types of low skill labor, high skill labor and capital are imperfectly substitutable for low skill labor ($\frac{1}{\epsilon^m} < \infty$, $\frac{1}{\rho_l} < \infty$).[29] 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 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.

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

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

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

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.

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

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

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:

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

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

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 ($\frac{L_{D,j,t+1}^l + L_{H,j,t+1}^l}{L_{D,j,t}^l + L_{H,j,t}^l} - 1$) is lower for a firm with a higher $FA_j^{l,model}$ 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 $FA^{l,model}_{j,t}$ 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:

$$\begin{aligned}
\frac{L^l_{D,j,t+1} + L^l_{H,j,t+1}}{L^l_{D,j,t} + L^l_{H,j,t}} - 1 &= \frac{L^l_{D,j,t}}{L^l_{D,j,t} + L^l_{H,j,t}}\left(\frac{L^l_{D,j,t+1}}{L_{D,j,t}} - 1\right) = FA^{l,\text{model}}_{j,t}\left(\frac{L^l_{D,j,t+1}}{L^l_{D,j,t}} - 1\right), \\
&= FA^{l,\text{model}}_{j,t}\left(\left(\frac{w_{h,t+1}/w_{h,t}}{\underline{w}_{l,t+1}/w_{l,t}}\right)^{\frac{1}{\epsilon^l}} - 1\right).
\end{aligned}$$

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).[30] With the minimum wage increasing above the market clearing wage, the relative wage of low skill labor increases from t to $t+1$. Therefore, the coefficient on $FA^{l,\text{model}}_{j,t}$ is negative with $(\frac{w_{h,t+1}/w_{h,t}}{\underline{w}_{l,t+1}/w_{l,t}})^{\frac{1}{\epsilon^l}} < 1$.[31] $\square$

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 $i^\star_m$ and allow the heterogeneity across

[30] 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.

[31] One caveat in our model is that $FA^{l,model}_{j,t}$ is not affected by $\underline{w}_{l,t+1}$ because only low skill workers are subject to the minimum wage in any case. In reality, worker skills are differentiated by more than two types and the magnitude of the minimum wage increase can also affect how many workers are potentially exposed to the minimum wage hike with the rapid increase in the minimum wage resulting in the rise in the average exposure. In our model, that effect is not distinguishable from the relative wage effect captured by $\frac{w_{h,t+1}/w_{h,t}}{\underline{w}_{l,t+1}/w_{l,t}}$.

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:

$$\Phi_j(i^\star_{m,t}, i^\star_{m,t+1}) = f_j. \tag{17}$$

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 10 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

$$\frac{A^m_{L,D}(i^\star_m)}{A^m_{L,F}(i^\star_m)} = \frac{\underline{w}_l}{\underline{w}_f}. \tag{18}$$

Hence, if $\underline{w}_l$ increases with no change in $\underline{w}_f$, $i^\star_m$ must increase with the comparative advantage assumption that we made with $i^\star_{m,2} > i^\star_{m,1}$.[32] 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 $i^\star_{m,1}$ and $i^\star_{m,2}$ 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 $i^\star_{m,2}$ compared to $i^\star_{m,1}$ 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:

$$B(i^\star_{m,1}, i^\star_{m,2}) = \int_{i^\star_{m,1}}^{i^\star_{m,2}} [\frac{\underline{w}_{l,2}}{A_{L,D}(i)} - \frac{\underline{w}_f}{A_{L,F}(i)}]di > f_j. \tag{19}$$

Notice that by the definition of threshold,

[32]We are implicitly assuming that foreign low skill labor is abundant enough that $\underline{w}_f$ would not change by increased offshoring.

$$\frac{\underline{w}_{l,2}}{A_{L,D}(i^\star_{m,1})} > \frac{\underline{w}_{l,1}}{A_{L,D}(i^\star_{m,1})} = \frac{\underline{w}_f}{A_{L,F}(i^\star_{m,1})}, \tag{20}$$

$$\frac{\underline{w}_{l,2}}{A_{L,D}(i^\star_{m,2})} = \frac{\underline{w}_f}{A_{L,F}(i^\star_{m,2})}. \tag{21}$$

Given the comparative advantage assumption that we made ($\ln \frac{A^m_{L,D}(i)}{A_{L,F}(i)}$ is increasing in i), the term inside the bracket in equation (19) is always positive for $i \in (i^\star_{m,1}, i^\star_{m,2})$. Combine this with the fact that $i^\star_{m,2} - i^\star_{m,1}$ increases with $\frac{\underline{w}_{l,2}}{\underline{w}_{l,1}}$, we obtain the result that $B(i^\star_{m,1}, i^\star_{m,2})$ 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 ($P(f_j < B(i^\star_{m,1}, i^\star_{m,2}))$ increases with $\frac{\underline{w}_{l,2}}{\underline{w}_{l,1}}$. □

High-skill labor and capital are imperfect substitutes for low skill labor. Hence, 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 increase in the minimum wage can deepen the relative intensity of these factors ($\frac{L^m_H(i)}{L^m_D(i)}$ or $\frac{K^m(i)}{L^m_D(i)}$) under the usual assumptions on the comparative advantage that $\ln \frac{A^{m,g}_H(i)}{A^m_{L,D}(i)}$ and $\ln \frac{A^m_K(i)}{A^m_{L,D}(i)}$ 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 even after 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. In particular, firms adjusted through the extensive (plant closing) margin as well as the intensive (layoff) margin and accounting for both margins makes the estimated employment effect more negative. Consistent with the existing literature, the minimum wage increase 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 of 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 provide caution against extrapolating experiences from small and localized minimum wage hikes to large nation-wide minimum wage changes.

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

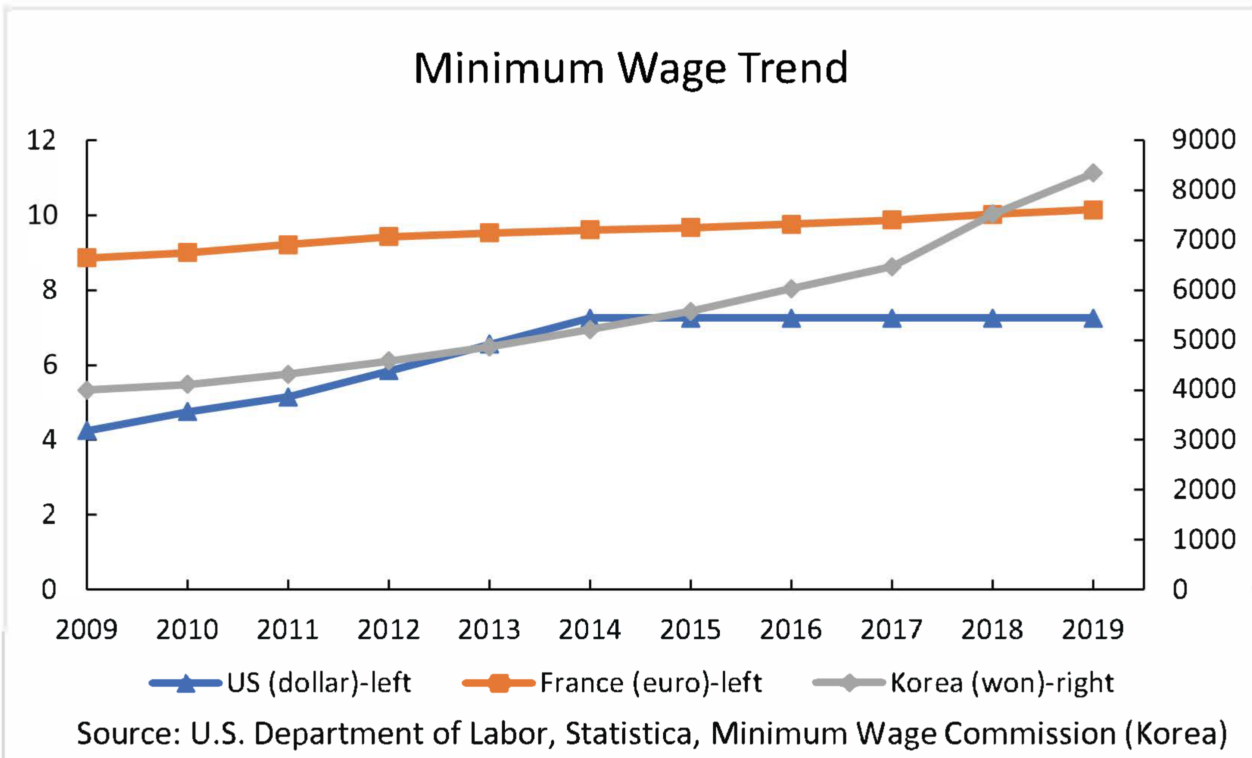


Figure 2: Time Series of the Kaitz Index in France, South Korea, and U.S.

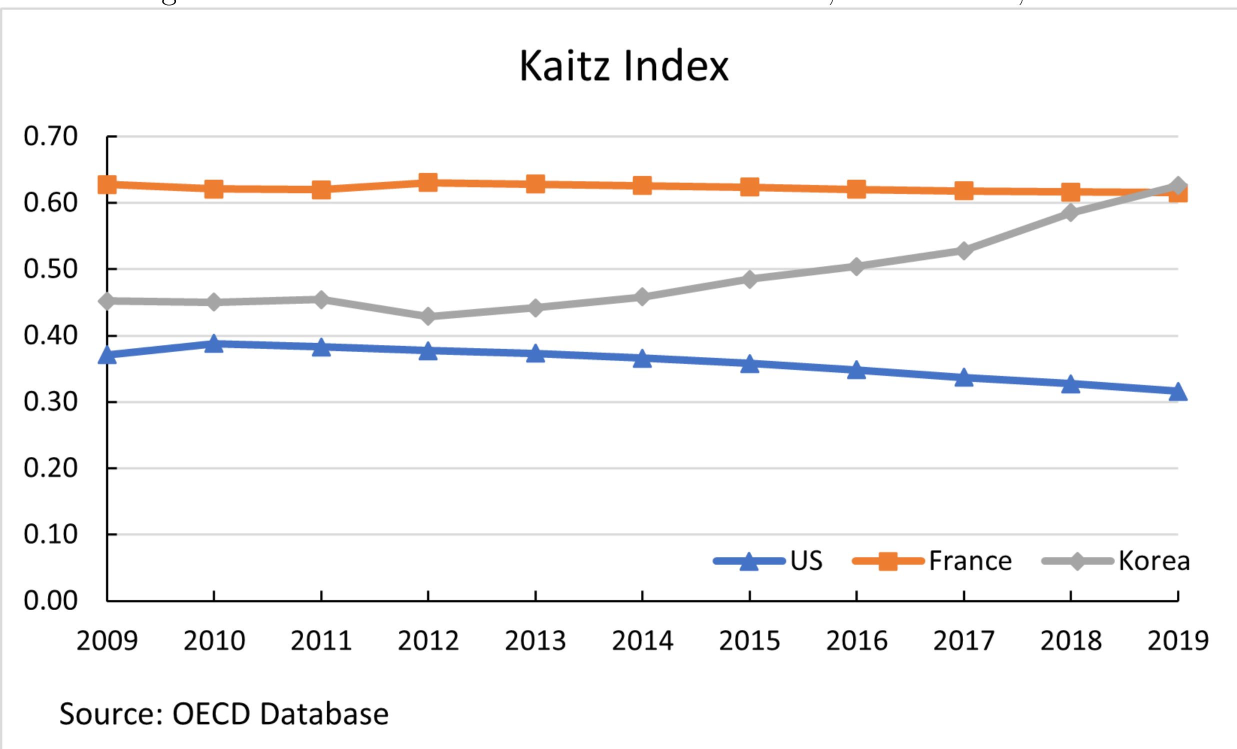


Figure 3: Macroeconomic Environment in South Korea: 2015-2019

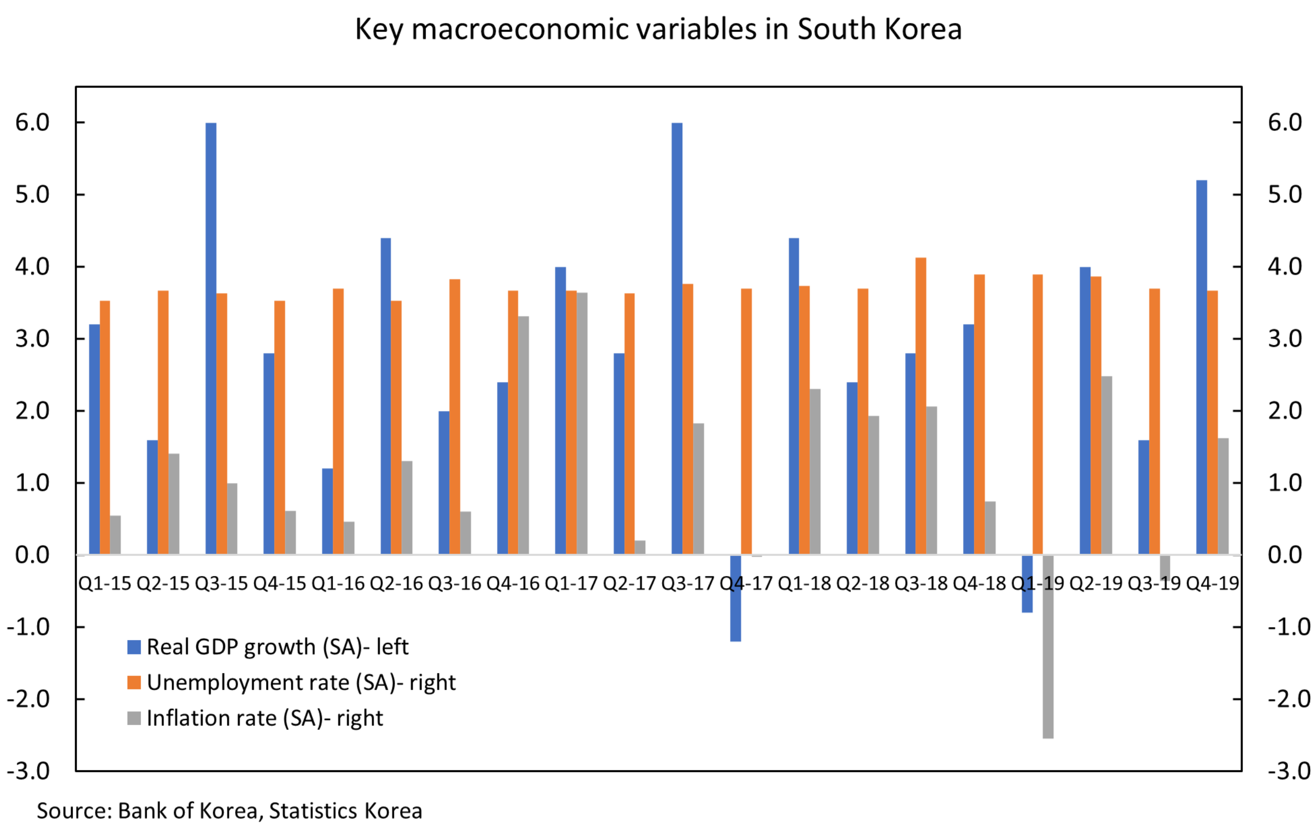
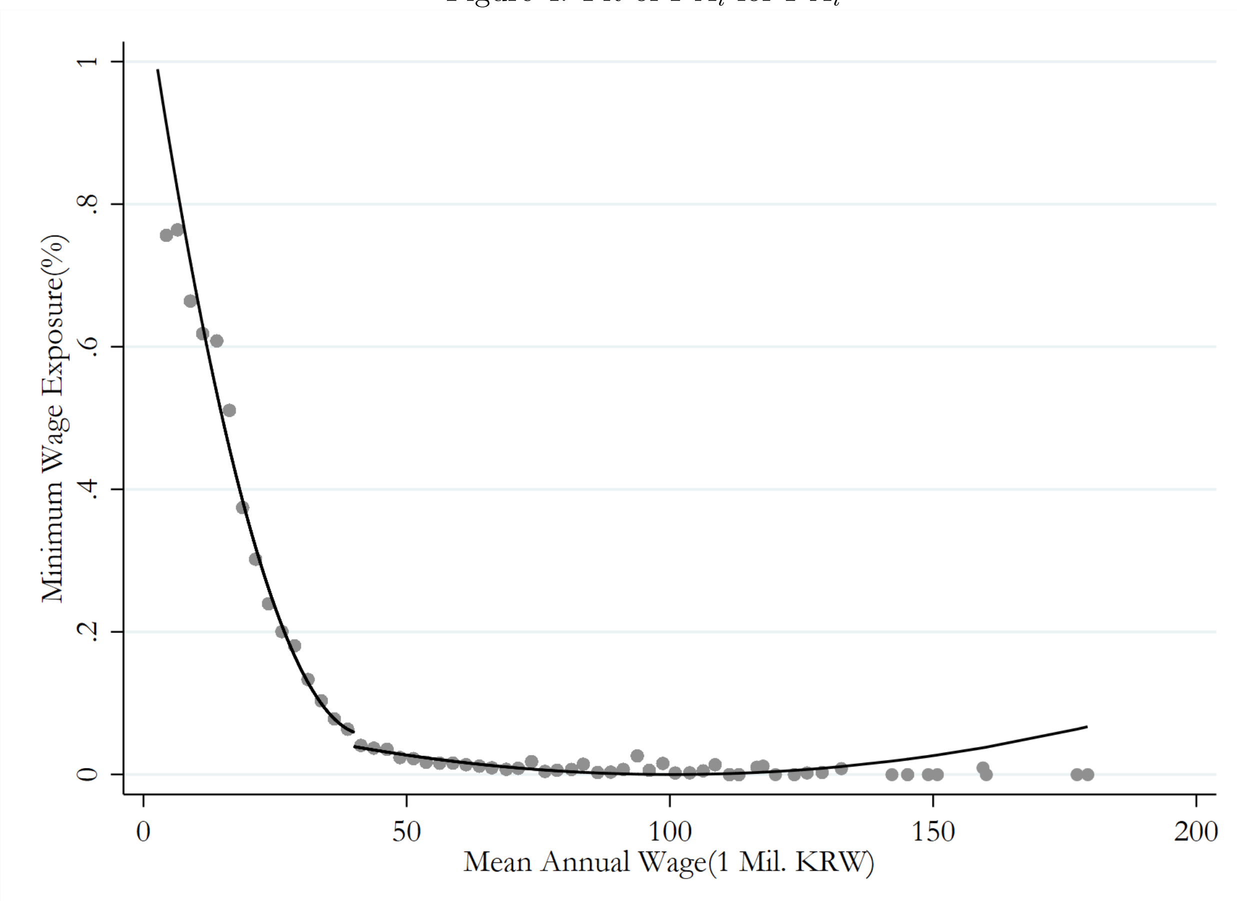
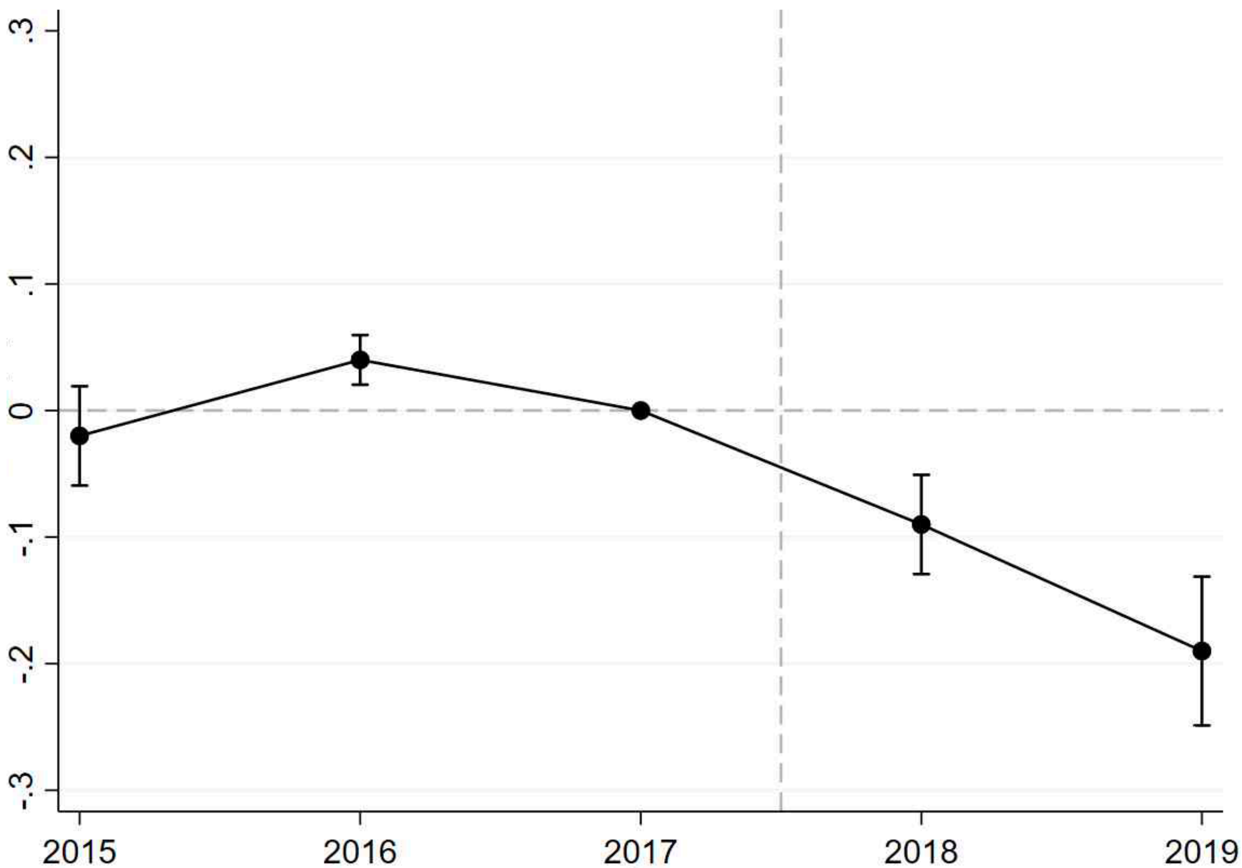


Figure 4: Fit of $\widehat{FA}_i$ for FA_i



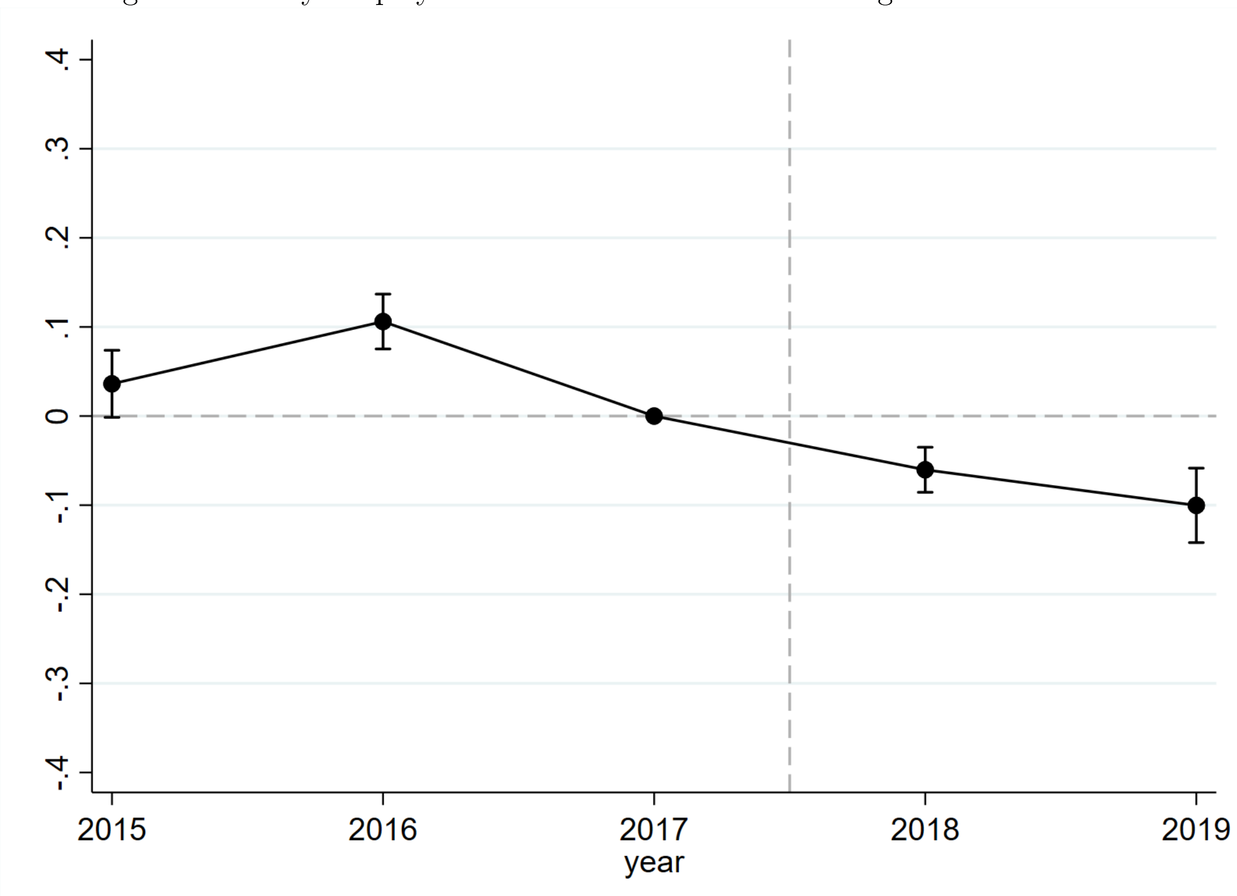
Source: Authors' calculation based on the Mining and Manufacturing Survey (MMS) and Employment Positions Statistics (EPS) Sample.

Figure 5: Yearly Employment Effect of the Minimum Wage Hike: Mining and Manufacturing Sector



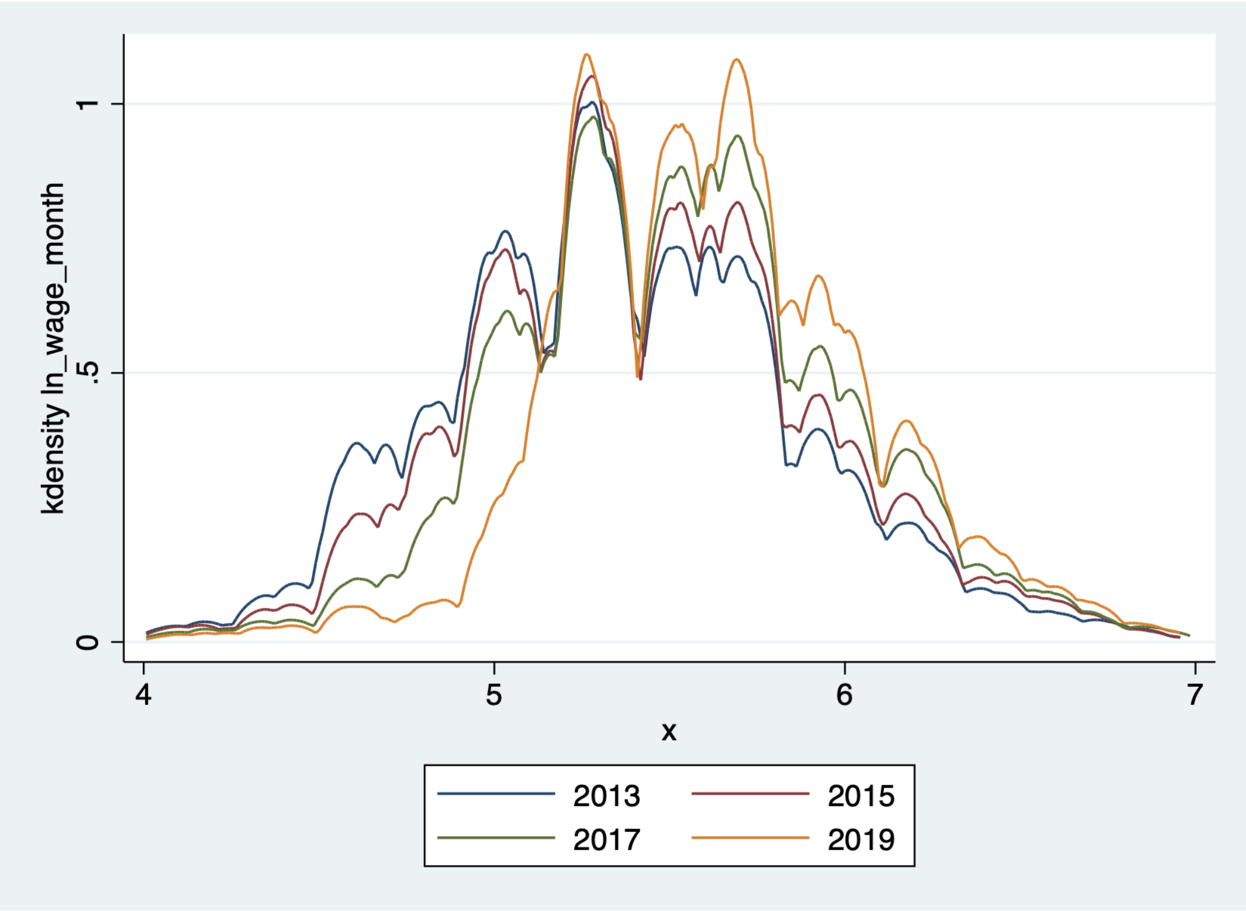
Source: Authors' calculation based on the Mining and Manufacturing Survey (MMS) Sample.

Figure 6: Yearly Employment Effect of the Minimum Wage Hike: Service Sector



Notes: Black dots represents the point estimates of yearly effects while gray dots denote the pre-existing trend.

Figure 7: Effect of the Minimum Wage Hike on Wage Dispersion in the Mining and Manufacturing Sector

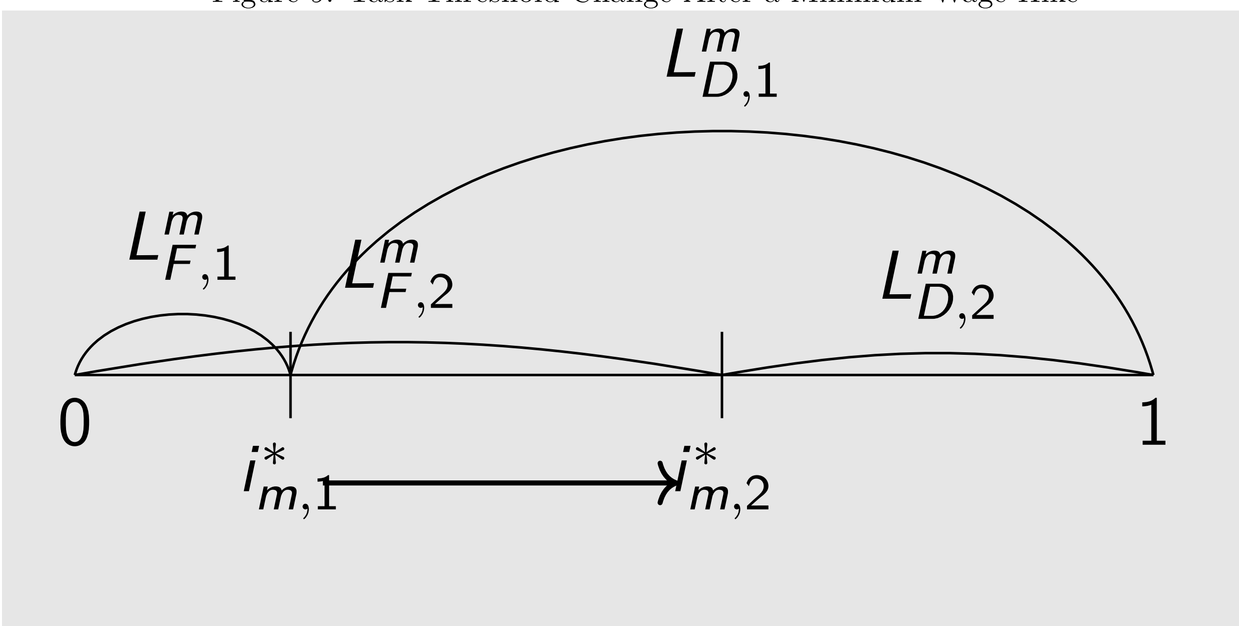


Source: Authors' calculation based on the Employment Positions Statistics (EPS) Sample. The horizontal axis represents the log of wage.

Figure 8: South Korea's Foreign Direct Investment in Vietnam (Manufacturing Sector)

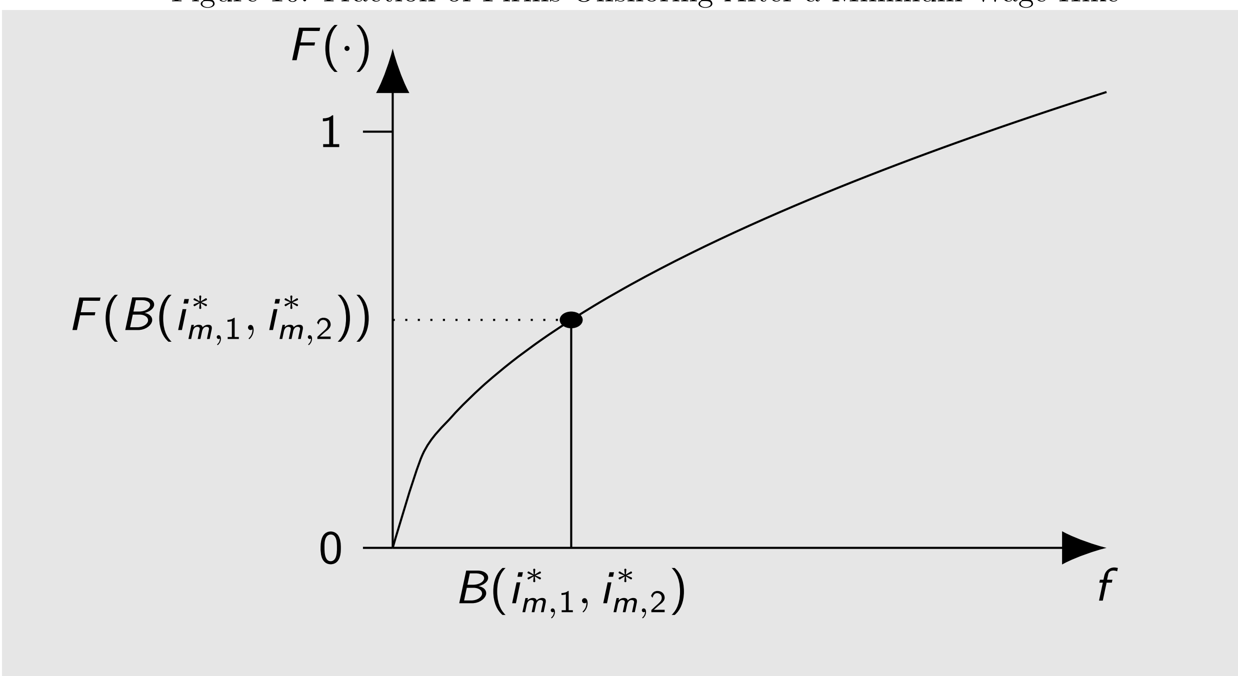
Source: Export-Import Bank of Korea.

Figure 9: Task Threshold Change After a Minimum Wage Hike



The figure 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 10: Fraction of Firms Offshoring After a Minimum Wage Hike



The figure 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 is the one-digit Korea Standard Industry Code)

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: Authors' calculation based on the 2017 Employment Positions Statistics (EPS).

Table 2: Change in Employment

	2017-2019			2015-2017		
	A. Extensive + Intensive Margin					
	(1)	(2)	(3)	(4)	(5)	(6)
$\widehat{FA}_i$	-0.20^{***}	-0.21^{***}	-0.21^{***}	0.00	-0.00	-0.02
standard error	0.02	0.02	0.03	0.02	0.02	0.02
observations	50,854	50,854	50,854	50,854	50,854	50,854
	B. Intensive Margin					
$\widehat{FA}_i$	-0.11^{***}	-0.11^{***}	-0.13^{***}	0.00	0.00	-0.02
standard error	0.02	0.02	0.02	0.02	0.02	0.02
observations	47,973	47,973	47,973	48,869	48,869	48,869
size weighted		Y	Y		Y	Y
fixed effect			Y			Y

Notes: Authors' calculation based on the 2017 Mining and Manufacturing Survey (MMS) sample. Clustered standard errors at the district (city, county) level are reported. We flip the sign of the coefficient for the 2015-2017 period because in equation (2) the negative employment effect is associated with the positive coefficient for $\widehat{FA}_i$ $***$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%

Notes: Authors' calculation based on the Employment Insurance Database (EID).

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

	baseline	Alt 1	Alt 2	Alt 3
$\widehat{FA}_i$	-0.20^{***}	-0.29^{***}	-0.09^{***}	-0.21^{***}
standard error	0.03	0.02	0.01	0.02
observations	50,854	50,854	45,122	50,854

Notes: Authors' calculation based on the Mining and Manufacturing Survey (MMS) sample. $***$ p<0.01, $**$ p<0.05, $*$ p<0.1

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

	Mining and Manufacturing	Service
$\widehat{FA}_i$	-0.19^{***}	-0.10^{***}
standard error	0.02	0.02
observations	48,869	40472
size weighted	Y	Y
fixed effect	Y	Y

Notes: Authors' calculation based on the Mining and Manufacturing Survey (MMS) sample and the Corporate Information Data Service (CIDS) sample. $***$ p<0.01, $**$ p<0.05, $*$ p<0.1. $\widehat{FA}_i$ computed at the division level (two-digit) from the Korea Standard Industry Code.

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

	2017-2019			2015-2017		
	(1)	(2)	(3)	(4)	(5)	(6)
			A. Average Wage Growth Rate			
$\widehat{FA}_i$	0.69***	0.70***	0.75***	0.16***	0.17***	0.19***
standard error	0.02	0.02	0.02	0.02	0.02	0.02
			B. Average Labor Cost Growth Rate			
$\widehat{FA}_i$	0.67***	0.69***	0.73***	0.11***	0.13***	0.15***
standard error	0.02	0.02	0.02	0.02	0.02	0.02
observations	45,122	45,122	45,122	45,122	45,122	45,122
size weighted		Y	Y		Y	Y
fixed effect			Y			Y

Notes: Authors' calculation based on the 2017 Mining and Manufacturing Survey (MMS) sample. We only use firms who never dropped out of the MMS sample for the three years (Alt 2 in Table 4). We flip the sign of the coefficient for the 2015-2017 period because in equation (2) the positive effect on wage is associated with a negative coefficient for $\widehat{FA}_i$. $***$ p<0.01, $**$ p<0.05, $*$ p<0.1.

Table 7: Change in Labor Productivity

	2017-2019			2015-2017		
	(1)	(2)	(3)	(4)	(5)	(6)
$\widehat{FA}_i$	0.40***	0.40***	0.42***	0.21***	0.22***	0.23***
standard error	0.03	0.03	0.03	0.03	0.03	0.03
observations	45,122	45,122	45,122	45,122	45,122	45,122
size weighted		Y	Y		Y	Y
fixed effect			Y			Y

Notes: Authors' calculation based on the 2017 Mining and Manufacturing Survey (MMS) sample. We only use firms who never dropped out of the MMS sample for the three years (Alt 2 in Table 4). LP is measured by the log of output/labor input. Standard errors are clustered at the district (city, county) level. We flip the sign of the coefficient for the 2015-2017 period because in equation (2) the positive productivity effect is associated with a negative coefficient for $\widehat{FA}_i$. $***$ p<0.01, $**$ p<0.05, $*$ p<0.1.

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

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

Notes: Authors' calculation based on the Local Area Labor Force Survey (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^{***}
standard error	0.0358	0.0289.
Lagged Average Probability of Being Exposed	-0.102^{***}	-0.341^{***}
standard error	0.0322	0.0296.
size weighted	N	Y

Notes: Authors' calculation based on the Local Area Labor Force Survey (LALFS) sample. $***$ p<0.01, $**$ p<0.05, $*$ p<0.1

Table 10: Average Minimum Wage Exposure in 2017 by Division (the two-digit Korea Standard Industry Code) and Location for the Manufacturing Sector

Divisio.	Seoul	Busan	Daegu	Incheon	Gwangju	Deajeon
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 Apprarel, 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 Vehiclees, 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: Authors' calculation based on the Local Area Labor Force Survey (LALFS). Location represents the six largest metropolitan cities in South Korea.

A 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 10 shows the average minimum wage exposure in 2017 by the two-digit Korea Standard Industry Code and location for the manufacturing sector.

A.2 2015 MMS Sample versus 2017 MMS Sample

[33]

Our analysis obtains the placebo sample results during 2015-2017 for firms included in the 2017 MMS sample. While it is useful to identify the pre-treatment effect for firms treated by

[33]Empirical analysis based on the 2015 MMS sample is drawn from Lee (2022).

the minimum wage hike in 2018 and 2019, we can not use the placebo sample results for the treatment effect of the minimum wage increase in 2016 and 2017 because some of affected firms already dropped out of the 2017 MMS sample. For this reason, we use the 2015 MMS sample to calculate the post-treatment effect for the 2015 sample. Table 11 shows the comparison of the post-treatment effect for the 2017 MMS sample and the 2015 MMS sample. Given a more modest increase of the minimum wage in 2016 and 2017, variations in the minimum wage exposure defined by the fraction of workers earning below the 2017 minimum wage for the 2015 MMS sample are smaller than those from the 2017 MMS sample based on the minimum wage hike in 2018 and 2019. As a result, although the point estimate of the coefficient on the employment effect is somewhat more negative, the uncertainty is much greater in the 2015 MMS sample than in the 2017 MMS sample. We interpret these findings as suggesting the following two points.

1. Conditional on the minimum wage exposure, the negative employment effect is not smaller for the modest increase in the minimum wage as the point estimate on the minimum wage exposure is actually more negative in the 2015 MMS sample.

2. However, the estimation uncertainty is much higher for the sample with the modest increase in the minimum wage.

The estimated employment effect for the manufacturing and mining sector for the 2017 MMS sample with the minimum wage exposure based on the 2019 minimum wage level is -0.21 (coefficient) $\times$ 0.148 (average exposure) $\approx$ -3% while the corresponding one for the 2015 MMS sample with the minimum wage exposure based on the 2017 minimum wage level is -0.28 (coefficient) $\times$ 0.057 (average exposure) $\approx$ -1.6%. Hence, the overall disemployment effect is bigger in the 2017 MMS sample although the point estimate of the effect of the minimum wage exposure on employment is smaller than in the 2015 MMS smaple. As mentioned by Harasztosi and Lindner (2019), the first point can be interpreted as evidence against a monopsony model of the labor market because even a modest increase is not leading to a positive employment effect. At the same time, a significant uncertainty of the point estimate suggests that little or small positive effect found in the empirical literature based on the modest increase in the minimum wage can be subject to a large degree of uncertainty.

A.3 Alternative Samples for the Service Sector

The baseline sample for the service sector is not quite comparable to the MMS sample because it contains firms with employees fewer than 10. We re-estimate the employment effect of the minimum wage hike using alternative samples in the service sector which can be more comparable to the MMS sample in terms of the number of employees or the additional control

Table 11: Change in Employment: 2017 MMS Sample versus 2015 MMS Sample

	2017 MMS: 2017-2019			2015 MMS: 2015-2017		
	A. Extensive + Intensive Margin					
	(1)	(2)	(3)	(4)	(5)	(6)
$\widehat{FA}_i$	-0.20^{***}	-0.21^{***}	-0.21^{***}	-0.28^{***}	-0.30^{***}	-0.28^{***}
standard error	0.02	0.02	0.03	0.06	0.06	0.06
observations	50,854	50,854	50,854	48,878	48,878	48,878
	B. Intensive Margin					
	(1)	(2)	(3)	(4)	(5)	(6)
$\widehat{FA}_i$	-0.11^{***}	-0.11^{***}	-0.13^{***}	-0.17^{***}	-0.19^{***}	-0.18^{***}
standard error	0.02	0.02	0.02	0.06	0.05	0.06
observations	47,973	47,973	47,973	44,377	44,377	44,377
size weighted		Y	Y		Y	Y
fixed effect			Y			Y

Notes: Authors' calculation based on the 2017 Mining and Manufacturing Survey (MMS) sample and the 2015 Mining and Manufacturing Survey (MMS) sample. Clustered standard errors at the district (city, county) level are reported. We flip the sign of the coefficient for the 2015-2017 period because in equation (2) the negative employment effect is associated with a positive coefficient for $\widehat{FA}_i$. $***$ p<0.01, $**$ p<0.05, $*$ p<0.1.

variables. Table 12 shows coefficients on the minimum wage exposure for alternative samples as well as the baseline sample. For this alternative sample, we do not require the availability of additional control variables like per-capita asset because doing so would severely reduce the sample size, distorting the representation of industries since there are very few firms in some industries. The sample size decreases by about 20% even in the case that we do not restrict the sample to firms with information on additional control variables. While the quantitative magnitude changes across samples, the qualitative pattern (the amplification of the negative employment effect in the 2017-2019 period relative to the 2015-2017 period) holds.

A.4 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

Table 12: Change in Employment: Alternative Samples for the Service Sector

	2017-2019		2015-2017	
	(1)	(2)	(3)	(4)
$\widehat{FA}_i$	-0.10^{***}	-0.26^{***}	-0.04^{**}	-0.17^{***}
standard error	0.02	0.02	0.02	0.03
observations	40,472	32,243	40,472	32,243
size weighted	Y	Y	Y	Y
fixed effect	Y	Y	Y	Y

Notes: Authors' calculation based on the alternative samples for the service sector. (1) and (3) correspond to the baseline sample while (2) and (4) restrict the baseline sample to those who have more than 10 employees comparable to the Mining and Manufacturing Survey (MMS) sample but may not have additional control variables. Clustered standard errors at the district (city, county) level are reported. We flip the sign of the coefficient for the 2015-2017 period because in equation (2) the negative employment effect is associated with a positive coefficient for $\widehat{FA}_i$. $***$ p<0.01, $**$ p<0.05, $*$ p<0.1.

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 13 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 subsample 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 disemployment effect in the service sector during 2017-2019 compared to the pre-2017 period as shown by (3) and (6) in Table 12. The price pass-through channel suggested by Table 13 should actually further amplify the disemployment effect in 2017-2019 because the sensitivity of prices to minimum wages is lower.

Table 13: Price Pass-through from Minimum Wage Hikes

Dependent Variable: $\ln PPI_{k,t}$			
Period	2011-19	2017-19	2011-17
A. Manufacturing Sector			
$\widehat{FA}_{k,t}$	0.1342	-0.0976	-0.3178
standard error	0.5668	0.2955	0.7137
$\widehat{FA}_{k,t-1}$	0.1212	-0.1704	0.2004
standard error	0.2630	0.2371	0.3595
B. Services Sector			
$\widehat{FA}_{k,t}$	0.4502**	0.2585	0.4696*
standard error	0.1846	0.1701	0.2805
$\widehat{FA}_{k,t-1}$	0.4115**	-0.1317	0.6835**
standard error	0.1998	0.2376	0.3296
observations	384	144	288

Notes: Authors' calculation based on the PPI data created by the Bank of Korea and the individual minimum wage exposure (the binary indicators for the wage below 1.2 $\times$ minimum wage) aggregated at the two-digit industry level (k is the index for the two-digit industry) from the Local Area Labor Force Survey (LALFS). $***$ p<0.01, $**$ p<0.05, $*$ p<0.1.