Minimum Wages and Capital-Labor Substitution

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Abstract

This paper studies the capital-labor substitution effects associated with higher minimum wages, using Costa Rica's rich administrative data. I exploit this country's occupation-based setting to estimate average and sector-specific elasticities of substitution between capital and labor. In this case, the policy establishes a relevant minimum wage for both low and higher-skilled occupations. I find elasticities consistently below one, suggesting that the substitution away from labor towards capital is not large enough to reduce the labor share after a minimum wage increase. Specifically, I compute an elasticity of 0.59 for all firms, and significant heterogeneity across representative sectors, stressing differences in the production technologies across industries. The estimated value is higher in manufacturing (0.81) and tradable sectors (0.76) but smaller in non-tradable sectors (0.46).

JEL Codes: D22, D24, E24, J23, J24, J31, J38

1 Introduction

This paper examines the capital-labor substitution adjustment margin to higher minimum wages. I exploit Costa Rica's occupation-based minimum wage setting, as it binds to an extensive segment of the labor market. Specifically, the policy establishes a relevant minimum wage for low and higher-skilled occupations. Besides, minimum wage levels

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are relatively high, and they increased significantly during the period of analysis. Consequently, the policy extends to sectors for which standard minimum wage settings would not be binding or would have modest impacts. Furthermore, prior work exploiting the minimum wage inherently restricts to the dynamics between capital and low-skilled labor. The analysis conducted for this country would yield a more general interpretation of the interaction between these two production inputs.

I assemble a comprehensive dataset covering the universe of firms and workers in Costa Rica's formal sector to estimate a firm-level exposure measure. I define exposure as the increase in the total wage bill that a firm has to pay to satisfy the new minimum wage requirements. Specifically, it computes the cost of compliance for firms when minimum wages increase. A one percentage point, for instance, means that the firm has to increase its wage bill by one percent to bring all of its current employees up to the new minimum wage levels. Hence, low-paying firms would have high exposure when the minimum wages increase. This variable, however, is potentially endogenous. Therefore, I construct an instrument exploiting the fact that firms are differentially exposed to the common minimum wage adjustments based on their occupational composition. More precisely, the instrument consists of the occupation-specific minimum wage increases, weighted by the firm's occupational composition in 2007. Afterward, I estimate a sequence of regressions estimating the effect of a minimum wage change on firm outcomes at different year horizons. In other words, I examine if differential exposure to the minimum wage leads to differential changes in relevant firm outcomes.

The results of the analysis suggest that higher minimum wages induce firms to increase their capital-labor ratios. A one percent increase in the labor costs induced by the minimum wage initially leads to a rise of 0.26 percent in the capital-labor ratio. However, the magnitude of the elasticity increases as the horizon expands to stabilize around 0.5. Such dynamics reflect firms reducing their employment levels and increasing their capital stocks in response to the policy.

I utilize the estimated firm responses to compute the elasticity of substitution of labor and capital (σ_{KL}) using a simple model of labor demand with imperfect competition in the output market as in Hamermesh (1993); Aaronson and French (2007); Harasztosi and Lindner (2019). Given Costa Rica's occupation-specific minimum wage setting binding most of the labor market segment, σ_{KL} can be interpreted as an average elasticity, instead of capital-low-skilled labor or an industry-specific parameter. I compute an elasticity of 0.59 for all firms and significant heterogeneity across representative sectors. The estimated elasticity of substitution is larger in manufacturing (0.81) and tradable sectors (0.76) but smaller in non-tradable sectors (0.46). Overall, the values are precisely estimated, i.e., with small standard errors. The contrast in the estimated parameters reflects the differences in the production technologies across sectors, stressing the importance of extending the analysis to different industries when assessing the minimum wage's incidence on firms.

Despite the heterogeneity across sectors, the estimated elasticities are consistently below one. Such a pattern supports the minimum wage's positive impact on the firm-level labor shares documented in Garita (2020). In other words, the substitution away from labor to capital is not large enough to reduce the labor share.

The rest of the paper unfolds as follows. Section 2 summarizes Costa Rica's minimum wage setting. Section 3 describes the data and Section 4 discusses the empirical strategy of the paper. Section 5 presents and discusses the estimation results.

Related literature and contribution: This paper contributes to the literature analyzing the extent to which firms can substitute labor with capital. Most recently, Harasztosi and Lindner (2019); Chen (2019); Hau et al. (2020) exploit firm-level data from Hungary, U.S. manufacturing, and China to show that firms substitute capital for labor in response to higher minimum wages. This paper reaches similar conclusions but differentiates in two major ways. First, I examine adjustment dynamics to the minimum wage. Consistent with adjustment costs, I find that the capital-labor substitution effects expand as the horizon extends. Second, by exploiting Costa Rica's occupation-based setting, I do not restrain to capital and low-skilled labor. Hence, I can speak about input substitutability in a broader sense.

The paper also provides new estimates of the elasticity of substitution between capital and labor, σ_{KL}. I exploit a comprehensive increase in the labor costs induced by the minimum wage to compute such a parameter. Costa Rica's occupation-based minimum wage setting allows me to account for and identify differences across relevant sectors. As stressed by Herrendorf et al. (2015), sectoral production functions differ in capital and labor substitutability. Moreover, since tradable sectors have less space to increase their prices to absorb higher minimum wages (Aaronson, 2001; Aaronson and French, 2007), input demand adjustments are potentially more relevant. Chen (2019) estimates σ_{KL} using U.S firm-level manufacturing data within a minimum wage context. She reports a value around 0.85, remarkably similar to the one estimated in this paper for such a sector. Harasztosi and Lindner (2019) estimate an elasticity for all firms, arguing that Hungary's atypical minimum wage increase in the early 2000s binds to most economic sectors. Using a similar theoretical framework, these authors infer an average elasticity of 3.35 and values between 2.60 and 4.63 for other sectors. Such a striking difference could be reflecting that Hungary's minimum wage increase mostly affected the relative cost of low-skilled labor. I examine the difference between firms intensive in low-skilled and high-skilled occupations, finding a higher elasticity for the former group, consistent with such a story. However, in both cases the magnitude of the parameter is below one.

Other studies have estimated the capital-labor substitution using alternative frameworks and contexts. Although there is substantial heterogeneity across the empirical estimates, most studies report a value for σ_{KL} below one. Oberfield and Raval (Forthcoming) estimate firm-level elasticities of substitution, exploiting cross-sectional differences in local wages for the U.S. manufacturing sector, within a range of 0.3 and 0.5.[1] Other literature focused on aggregate estimates have provided similar ranges. Herrendorf et al. (2015) report an aggregate elasticity of 0.84, comparable to manufacturing (0.80) and slightly above services (0.75). Chirinko (2008); Raval (2019) summarize the literature using variation in the user cost and price of capital, pointing out a range of 0.40-0.60.

As explained by Elsby et al. (2013), the elasticity of substitution is central to understand the evolution of the labor share. The global decline in the labor share (Karabarbounis and Neiman, 2014) has raised concerns about expanding inequality, wage stagnation, and consumer purchasing power deterioration. Blanchard and Giavazzi (2003); Piketty (2015); Azmat et al. (2012) stress the importance of labor market regulation and minimum wage policies as factors shaping the labor share and possible remedies to revert the decline. In Garita (2020), I find that higher minimum wages have a persistent positive impact on firm-level labor shares, despite the capital-labor substitution effects. A degree of substitutability between these two inputs below one precisely provides support to such a conclusion.

2 Institutional Context

2.1 Minimum Wage Structure

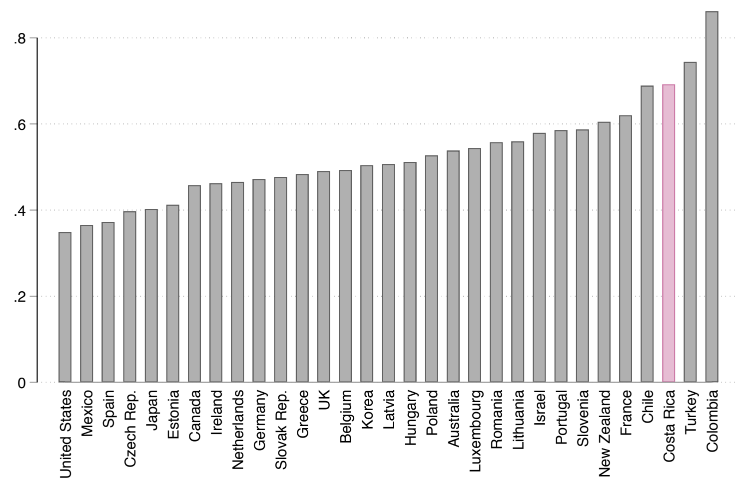
The minimum wage policy in Costa Rica is substantially more differentiated than in most of the OECD countries. This country implements a multi-tiered system of legal wage floors that vary by occupation, so minimum wage rates are essentially set by skill level. Adjustments are made twice a year, with new levels becoming effective in January and July, and decisions are carried out by the National Council of Salaries (NCS), a national-level tripartite commission formed of three representatives from labor unions, three from the Chamber of Commerce (private-sector companies) and three from the Central Gov-

[1] These authors exploit the firm-level elasticities to build up an aggregate value. They find a similar range for the aggregated elasticity, between 0.5 and 0.7.

ernment. The negotiating process is widely publicized, and the central purpose of the policy is to protect low-wage workers by establishing a wage floor that ensures basic living conditions to these individuals.

Overall, Costa Rica has a highly binding minimum wage. Figure 1 offers an international comparison, placing Costa Rica as one of the economies with the highest minimum wage.

Figure 1: Kaitz Index Across OECD countries
(Percentage of median wage. 2015)



Notes: Minimum relative to median wages of full-time workers
Source: OECD LFS

Workers are organized into three broad categories. The first group is of occupations associated with the production process (blue-collar workers). The second one, generic, applies to white-collar or administrative occupations. The third one covers specific occupations such as domestic workers and reporters. The first two groups are further divided into four skill categories: unskilled, semi-skilled, skilled, and specialized. Finally, there is an additional legal wage floor for workers with a bachelor's degree (undergraduate diploma) and university graduates (5-year university degree or *Licenciatura*). Table 1 summarizes the most important categories.

Table 1: Costa Rica: Minimum Wages by Skill Groups

	Minimum Wage (Low Skilled=100)	Kaitz Index	Percentage Increase 2006-2017	
			Nominal	CPI-Deflated
Low Skilled	100	0.82	122.8	27.7
Semi Skilled	122	0.77	118.0	25.0
Skilled	127	0.74	113.1	22.1
Technical Low-Skilled	143	0.61	107.3	18.8
Specialized	146	0.67	108.2	19.3
Technical High-Skilled	194	0.68	107.3	18.8
Bachelors University	216	0.55	107.3	18.8
University Graduate	290	0.45	107.3	18.8

Notes: The Kaitz Index is defined as the ratio of minimum wage to median wage. The monthly minimum wage for a low skilled worker in 2020 is 316,965 CRC, approximately US$560.
Source: Ministry of Labor and Social Security (MTSS)

Starting 2009, Costa Rica experienced a rapid decline in the inflation rate, a direct result of the adoption of an inflation-targeting regime, and the abrupt decrease in the international price of commodities due to the great recession (See Figure 3). These elements lead to an automatic and significant increase in the minimum wage between 2009 and 2016, as the 1998 agreement opened the room for negotiation only in cases of atypically high inflation rates and given the fact that inflation expectations slowly adjusted to the new inflationary steady state. In late 2011, the NCS and the Central Government agreed upon a new formula that takes into account recent but now expected inflation and GDP per capita during the past five years. Such a transition explains why the minimum wage behavior stabilizes in real terms after 2016.

Figure 2: Minimum Wage Minimorum (CPI-2015 Deflated. January 2000=100)

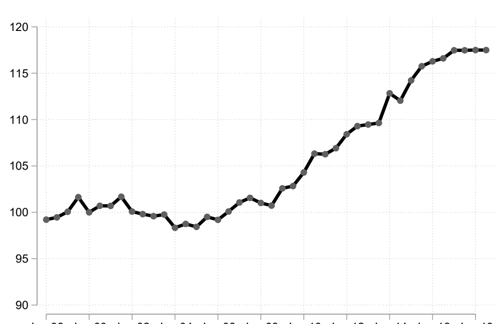
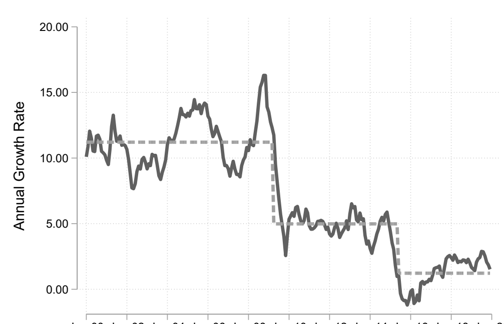


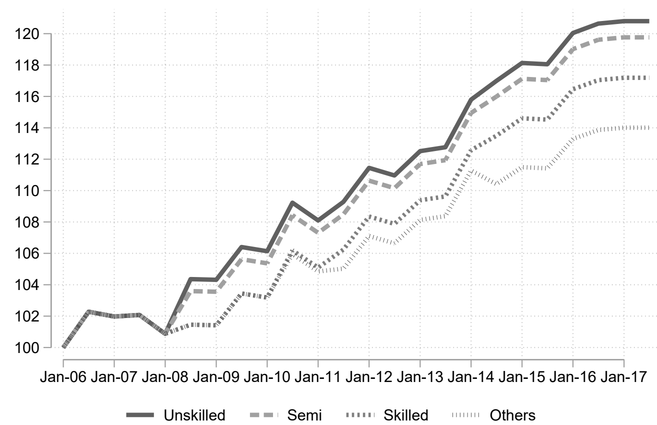
Figure 3: Annual Inflation Rate (CPI-2015)



Notes: The Minimum wage minimorum is the lowest level of the multi-tier system, corresponding to low-skilled occupations
Source: Ministry of Labor and Social Security (MTSS) and Central Bank of Costa Rica (BCCR)

The NCS decided to increase the minimum wage of the lower-skilled categories relatively more on three occasions (2008, 2012 and 2014). Hence, by 2017, low-skilled occupations experienced a sharper increase in the legal wage floor (see Figure 4). As it can be read from the NCS minutes that contain the discussion around each minimum wage adjustment decision (MTSS, 2008, 2012, 2014), the resolution of increasing low-skilled legal wage floors relatively more was mostly because under the new inflation rates, the indexation would lead to a small increase that would break a long period of two-digit growth rates, causing some social and political discontent. In other words, inflationary inertia was the main factor behind the decision-making process and the upward trend observed between 2008 and 2016.

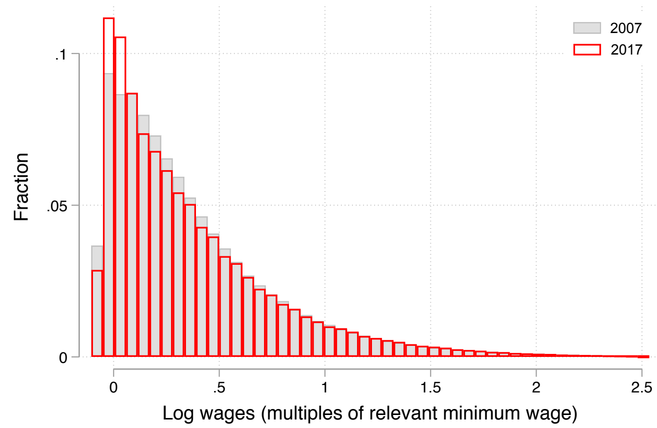
Figure 4: Costa Rica: Minimum Wage by Skill Groups
(CPI-2015 deflated. January 2006=100)



Notes: "Others" include specialized occupations, and university graduates.
Source: Ministry of Labor and Social Security (MTSS)

The steady increase in the real minimum wage translated into a higher bite of the minimum wage into the wage distribution. As shown in Figure 5, the mass of earnings around the relevant minimum wage significantly increased in 2017 relative to 2007.

Figure 5: Wage Distribution



Notes: Figure shows the frequency distribution of monthly log earnings in 2008 (last year before the steady increase in the real minimum wage), and in 2017 (when the adjustments stabilized in real terms). The red outlined bars show the earning distribution in 2017, and the grey solid bars show 2008. Labor earnings are CPI-2015 deflated. Sample selection restricts to full-time workers aged 18-60 employed by the private-sector.

3 Data and Descriptive Statistics

3.1 Main Dataset

I combine different administrative datasets that collectively comprise the universe of workers and firms in Costa Rica's formal sector. The first source of information is a monthly linked employer-employee data (CR-LEED) that I construct using raw firm-level records reported to the Costa Rican Social Security Fund and secured by the Central Bank of Costa Rica (BCCR). This data matches workers and employers from 2006 to 2017 and identifies each person with the legal person identifier and each employer with a legal tax identifier that facilitates the merging with other related information. By nature, these reports exclude part of the informal sector since they only include individuals contributing to social security. For each worker, I observe sociodemographics such as age, nationality, sex, and residence. In terms of the job match, I observe monthly labor earnings, full-time status, and if the employee is on paid-leave (maternity or sick-leaves, for example). Jobs are likewise organized into occupations according to the tasks and duties that are undertaken in the job, consistent with the International Standard Classification of Occupations (ISCO) at a 4-digit level.

The second dataset comes from the universe of corporate tax returns presented by firms from 2005 to 2018 (REVEC), which consists of annual balance sheets and income statements. I construct firm-level measures of performance and productivity from these records. Since both workers and firms are identified using the same legal identifiers, it is straightforward to combine both data sources. The outcome is a clean and comprehensive picture of the labor market, representing a significant advantage concerning existent literature, as most of the related studies lack at least one dimension of information. For instance, the administrative structure of it allows tracking with high precision firm entry and exit and, additionally, identifying and labeling employment flows and job-to-job transitions. Furthermore, I can observe the workforce and wage bill composition of each firm at a high detail to compute accurate and granular measures of exposure to the minimum wage.

One limitation, however, is that employers do not report the number of hours the employee worked. I overcome this shortcoming by restricting to full-time workers and exploiting the longitudinal history and panel structure to identify atypical wage reports. In Garita (2020), I provide more details about the data cleaning process.

For the remaining of the analysis, I restrict the sample to full-time workers aged 18 to 60 employed by a private-sector firm. Hence, I exclude self-employed individuals, households, non-profit firms, and state-owned enterprises, representing around 30 percent of total firm-year observations in the dataset.

3.2 Descriptive Statistics

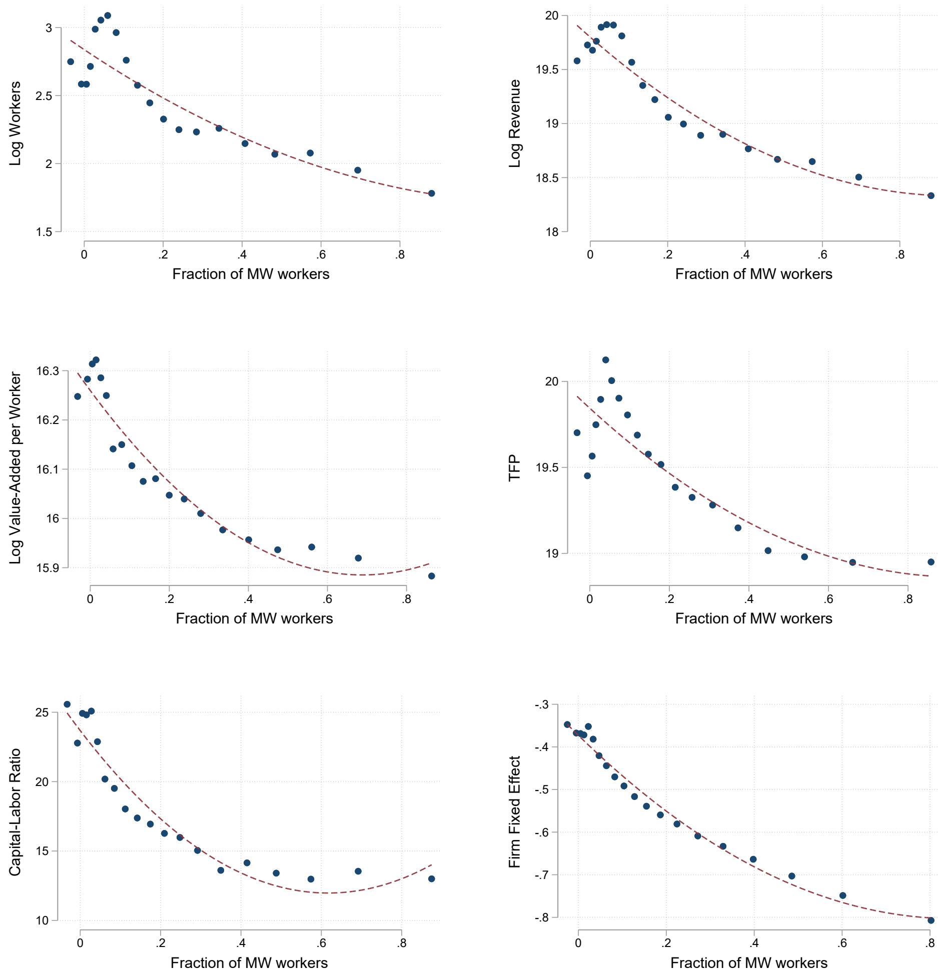
Table 2 summarizes the primary descriptive statistics for firms in the sample in 2007. Similarly, Figure 6 display binned scatterplots illustrating the non-parametric relationship between key firm characteristics and exposure to the policy. Overall, exposure is negatively correlated with firm size (both total employment and total revenues). Besides, highly exposed firms are less productive and more labor-intensive. Additionally, low paying firms are more exposed to the policy: the firm pay premium and the average wage are negatively associated with exposure. I include these firm characteristics as controls in the regression analysis to account for this pattern of correlation.

Table 2: Summary Statistics by Exposure Intensity. 2007

	All	Fraction of Minimum Wage Workers			
		0-25	25-50	50-75	75-100
Wage Bill	16.54	17.09	16.31	16.04	15.57
Average Wage	13.99	14.24	13.92	13.76	13.50
Revenue	18.70	19.14	18.47	18.27	17.98
Workers	41.06	64.65	21.35	16.39	11.39
Labor Share	0.17	0.18	0.16	0.16	0.14
Export Share	0.03	0.04	0.03	0.02	0.01
Capital-Labor Ratio	7.35	8.41	7.24	6.66	4.99
Exit Rate	0.11	0.09	0.10	0.11	0.17
Firms	18,646	9,835	3,452	2,287	3,072
(Fraction of total)	100	52.7	18.5	12.3	16.5

Notes: Export and labor share as proportion of revenues, profitability defined as profits per revenue. Export share include firms with zero exports. Capital-Labor ratio (fixed assets divided by number of workers) in millions of 2012 CRC.
Source: CR-LEED

Figure 6: Firm Characteristics and Minimum Wage Exposure



Notes: Figures show the binned scatterplot relating the firm-level fraction of workers exposed to the the minimum wage and firm outcomes, with the red line representing the best quadratic fit. Regressions include 2-digit industry and year fixed effects. Nominal variables deflated using 2012 GDP deflator. Total factor productivity index (TFP) estimated using a control function approach a la Ackerberg et al. (2015). Capital-Labor Ratio in millions of 2012 CRC. Firm fixed-effects estimated using a time-varying AKM model

4 Empirical Strategy

4.1 Minimum Wage Exposure

The key to identify effects of minimum wages on firms-level outcomes is to define a firm-specific minimum wage exposure measure. I define minimum wage exposure as the percentage increase in firm j's wage bill required to bring all of its current employees up to the new minimum wage[2]:

[2]Both wages and minimum wages are deflated using the CPI

$$Exposure_{j,t} = \frac{\sum_{i,o} \max\left(w_{o,t}^{min} - w_{i,j,o,t-1}, 0\right)}{\sum_{i,o} w_{i,j,o,t-1}}$$

This variable can also be interpreted as a firm-level compliance cost or a firm-specific minimum wage increase. It measures the distance between each worker's wage and the next year's minimum wage level.[3] By definition, this exposure measure requires complete worker-level detail for an accurate estimation.[4] The granular detail in the Costa Rican data represents a pivotal advantage to overcome these limitations, as I can construct accurate exposure measures for each firm in the labor market, regardless of its size or industry. If $Exposure_{j,t}$ increases by one percentage point, then the minimum wage policy is forcing the firm to increase its wage bill by one percent. Draca et al. (2011) also used a similar metric to measure minimum wage exposure, calling it the wage gap.

$Exposure_{jt}$ is measured based on the labor composition the period before the minimum wage change. In other words, it measures the firm-level increase in the wage bill induced by the minimum wage if the employer does not change its employment structure. Using current minimum wage changes and individual wages could be misleading as it would capture adjustments that the firm already implemented to comply with the policy. However, $Exposure_{jt}$ is still potentially endogenous, as it could be correlated to unobservables affecting firm outcomes. For example, an unobserved productivity shock can lead to changes in the employment composition and levels, simultaneously affecting minimum wage exposure and changes in outcome variables. Additionally, exposure could be correlated to unobservables that simultaneously put the firm closer to the minimum wage and the exit margin. To address this issue, define $z_{j,o,t}$ as the occupational share: the number of workers employed in occupation o relative to the total employment within the firm. Then, the exposure measure can be decomposed as the weighted average of exposure in each occupation category:

$$Exposure_{j,t} = \sum_{o} z_{j,o,t} Exposure_{j,o,t}$$

This structure precisely emphasizes that firms are going to be differentially exposed to the common minimum wage adjustments based on their occupational composition.

[3]Between 2006 and 2015, minimum wages were adjusted in January and June of each year. I use the January level for constructing the exposure measure.

[4]The existing literature has proven that such data requirement is difficult to meet, as there are not many information sources with such detail. Most of prior work measures of treatment intensity based on firm average wages, due to lack of worker-level data. As explain by Draca et al. (2011); Mayneris et al. (2018), any continuous measure of treatment intensity based on firm average wage is potentially noisy, especially when defining groups based on treatment.

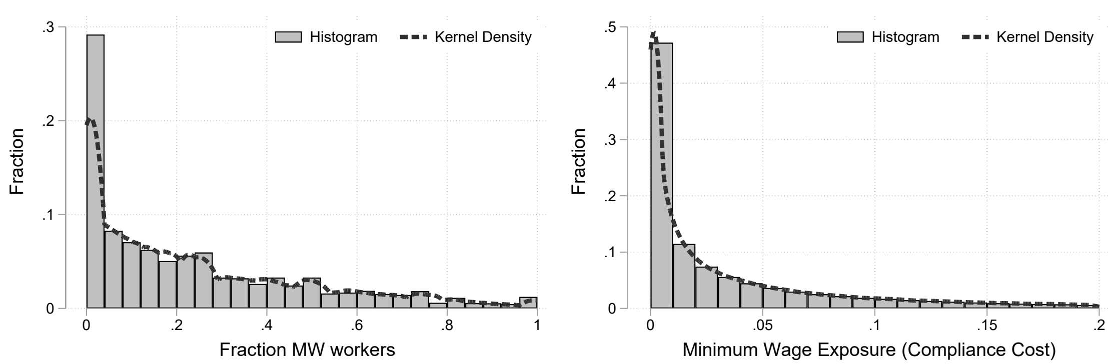
Hence, I consider $Exposure^{IV}_{j,t}$, an instrument for $Exposure_{j,t}$, defined as follows:

$$Exposure^{IV}_{j,t} = \sum_o z_{j,o,2007} mw_{o,t+1}$$

Where $mw_{o,t+1}$ is the percent change in the real minimum wage for occupation o relative to 2007 levels and $z_{j,o,2007}$ is the respective occupational share in firm j, estimated in 2007. By fixing the occupational shares to the 2007 levels, I analyze if firms with a particular occupational composition experience differential changes in outcomes following the minimum wage increases. The $Exposure^{IV}_{j,t}$ variable can be interpreted as a firm-level minimum wage increase, using the initial occupational shares as weights. The instrument is, by nature, a shift-share instrument and, as shown by Goldsmith-Pinkham et al. (2020), the empirical strategy is numerically equivalent to a generalized method of moments (GMM) estimator with the occupation shares as instruments and a weight matrix composed by the occupation-specific minimum wage increases.

Figure 7 summarizes the distribution of minimum wage exposure across firms in 2007. Consistent with the nature of the policy, there is a considerable concentration of firms with zero exposure. However, there is substantial variation in the degree of exposure among the rest of the firms. Besides, Table 3 confirms Costa Rica's setting extends to a large proportion of the labor market segment. Low binding minimum wage policies are characterized by an unpromising variation of minimum wage exposure, forcing researchers to restrict the analysis on specific sectors and demographics. Moreover, such a lack of variation has been a point of debate. Part of the literature argues that a low binding minimum wage policy is insufficient to detect the policy's true impact on firms (e.g. Sorkin (2015); Meer and West (2016); Neumark (2019); Clemens and Wither (2019)).

Figure 7: Histogram of Minimum Wage Exposure Measures (2007)



Notes: Figures show the histrogram of the firm-year estimated exposure to the minimum wage. Figure on the left shows the fraction of minimum wage workers employed by each firm. Figure on the right displays the wage gap (benchmark minimum wage exposure measure).
Source: CR-LEED

Table 3: Minimum Wage Exposure by Industry

Industry (2-digit ISIC Rev. 4)	Mean	Median
Agriculture, fishing and mines	0.25	0.19
Manufacturing	0.22	0.15
—Food products	*0.21*	*0.14*
—Wearing apparel	*0.30*	*0.27*
—Wood and of products of wood and cork	*0.31*	*0.30*
—Rubber and plastics products	*0.15*	*0.06*
—Computer, electronic and optical products	*0.12*	*0.07*
—Manufacture of machinery and equipment	*0.19*	*0.11*
Electricity , gas and water	0.17	0.07
Construction	0.20	0.14
Wholesale and retail trade	0.17	0.11
Accommodation and food service activities	0.20	0.14
Transportation and storage	0.19	0.11
Information and communication	0.15	0.08
Financial and insurance activities, real estate	0.12	0.04
Professional, scientific and technical activities	0.18	0.13
—Management consultancy activities	*0.11*	*0.03*
—Advertising and market research	*0.18*	*0.13*
—Security and investigation activities	*0.21*	*0.16*
Education	0.24	0.18
Human health and social work activities	0.22	0.18
Arts, entertainment and recreation	0.18	0.11
Other service activities	0.21	0.17

Notes: Table shows the fraction of minimum wage workers by industry in 2006-2007 (average.

4.1.1 Identification Assumptions

In this case, the implied IV strategy is that the initial occupation shares measure the differential exposure to the minimum wage increases. As I show in the next subsections, I focus on differential *changes* in the outcome variables (e.g., log cumulative *changes* in employment). The occupational shares could be correlated with outcome levels without representing an identification threat. The central identification assumption states that the initial shares are exogenous to the error term conditional on observables. In other words, there is an issue if the initial occupation composition predicts *changes* in outcomes through channels other than minimum wage exposure.

First, fixing the occupational composition to 2007 levels seeks to support the identification assumptions. As described in Section 2, such a year concludes almost a decade in which all minimum wages were stable in both real and relative terms (See Figure 2 and 4). Minimum wage adjustments were explicitly made to compensate for past inflation, so the minimum wage variation before 2008 was confounded with the wage adjustments employers make to compensate workers for inflation. Hence, before 2008, differential exposure to the minimum wage should not have driven differential changes in the studied outcomes if the identification assumption holds.

Still, the identification assumption previously stated is not directly testable. I discuss its plausibility based on the recommended strategy by Goldsmith-Pinkham et al. (2020). First, I explore how much the 2007 occupational shares are correlated with other potential confounders, also measured in 2007. For such purposes, I focus on five main occupation groups: low skilled, semi-skilled, skilled, specialized, and college graduate.[5] I estimate the corresponding occupational share for each group, and I compute the correlation between the occupational shares and key firm characteristics: labor share, capital share, export share, import share, and profitability. This set of variables cover the relative importance of labor and capital, firm size, and international trade exposure, as well as how close the firm is to the margin of exit. The idea is to explore if the instruments (occupational shares) are correlated with initial firm characteristics that could lead to confounding channels other than the minimum wage. For example, firms with high low-skilled occupational shares could also operate closer to the exit margin, so these firms may be prone to reductions in their employment levels. Similarly, firms with high skilled occupations could be significantly exposed to international trade, so productivity improvements could be linked to this channel. Reassuringly, Table 4 shows no apparent systematic pattern of correlation.

[5] Goldsmith-Pinkham et al. (2020) suggest estimating Rotemberg weights to identify the most representative shares driving the identification power. Instead, I consider the main occupational groups, although it should lead to similar conclusions. Since I am constructing firm-level instruments, computing Rotemberg weights become computationally intensive.

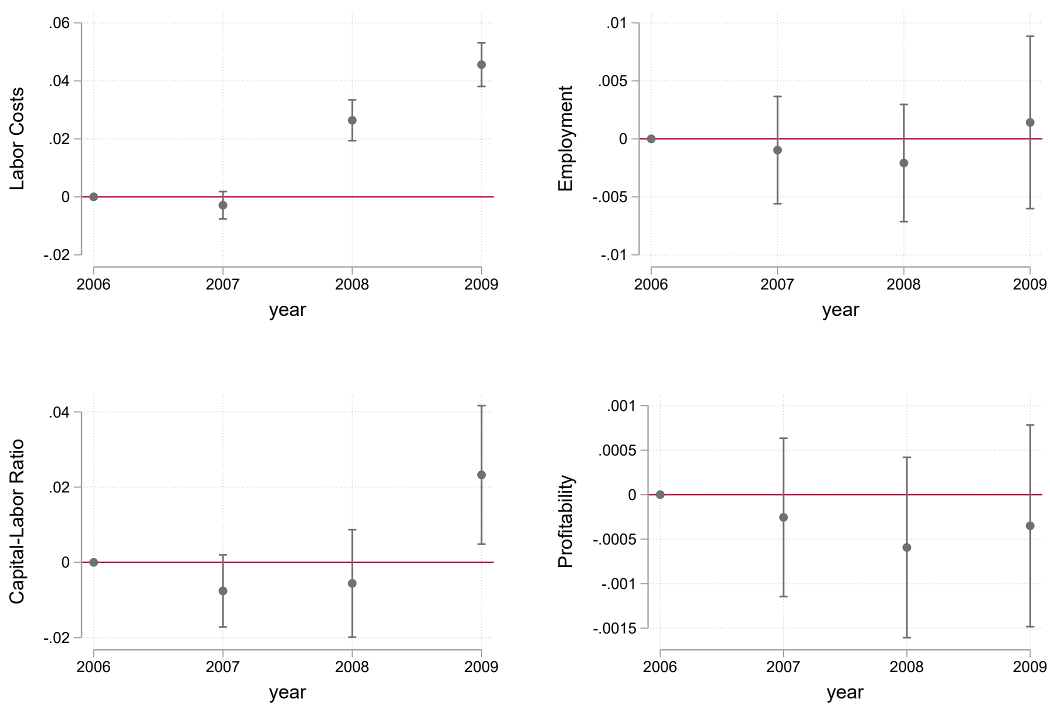
Table 4: Correlation between occupational shares and firm characteristics in 2007

	Low Skilled	Semi-Skilled	Skilled	Specialized	College Graduate	Aggregate Instrument
Labor Share	-0.405	-0.070	-0.188	0.104	0.146	-0.040
	(0.316)	(0.091)	(0.153)	(0.080)	(0.160)	(0.025)
Capital Share	-0.002	-0.004	-0.004	-0.000	0.004	-0.001
	(0.006)	(0.006)	(0.003)	(0.001)	(0.004)	(0.001)
Export Share	-0.209	-1.980	0.080	0.532	0.179	-0.177
	(0.633)	(1.511)	(0.388)	(0.589)	(0.402)	(0.125)
Import Share	0.109	-1.620	0.070	0.166	0.369	0.015
	(0.092)	(1.047)	(0.046)	(0.429)	(0.417)	(0.092)
Profitability	0.006	-0.010	-0.006	-0.002	0.013	-0.000
	(0.037)	(0.022)	(0.017)	(0.004)	(0.013)	(0.001)
Obs.	13,586	13,586	13,586	13,586	13,586	13,586
R^2	0.366	0.316	0.202	0.309	0.200	0.311

Notes: Table shows the result of a regression of the occupational share on firm characteristics, both estimated in 2007. The final column corresponds to the aggregate instrument, i.e., the interaction of all occupational shares with the respective minimum wage increase. Standard errors in parenthesis. For legibility, coefficients and standard errors are scaled by 10,000,000. * p<0.05

Second, I explore for parallel pre-trends. As mentioned previously, the idea is to explore if firms showed a different behavior depending on minimum wage exposure before 2008. Unfortunately, there is no available information before 2006 to extend the analysis beyond this point, but the available information discards the presence of pre-trends. Figure 8 displays the results of a set of regression of the log changes in firm outcomes (relative to 2007) on the minimum wage exposure instrument in 2007, i.e., the inner product of the occupational shares and the occupation-specific minimum wage increases, both in 2007. The plots indicate that before 2007, and in some cases 2008, there is no significant difference in the evolution of the firm characteristics between exposed and non-exposed firms.

Figure 8: Pre-Trends on Selected Firm Outcomes



Notes: Figures report results of regressing the log outcome changes relative to 2006 on the minimum wage exposure instrument in 2007, alongside 95% confidence intervals estimated using robust standard errors.

Finally, results suggest a strong first-stage and no evidence of misspecification. The F-statistics are sufficiently large to support the relevance assumptions behind the strategy and the overidentification tests fail to reject the null of misspecification (See Table 5). I will discuss these statistics in more detail in Section 5.

4.2 Firm-Level Analysis

The main goal of the paper is to estimate if differential exposure to the minimum wage leads to differential changes in relevant firm outcomes. To account for dynamics in the response, I estimate a sequence of regressions based on the local projection framework proposed by Jordà (2005). For firm j at year t and horizon $h = 1, ..., 5$, I consider the following specification[6]:

$$\Delta_h \ln Y_{j,t+h} = \alpha_h + \beta^h Exposure_{j,t} + \sum_{i=0}^{h-1} b_i Exposure_{j,t+1+i} + \gamma_h X_j + \nu_{s,t+h} + u_{j,t+h} \quad (1)$$

[6]For $h = 0$, consider $\Delta_0 \ln Y_{j,t} = \alpha_0 + \beta_0 Exposure_{j,t} + \gamma_0 X_j + \nu_{s,t} + u_{j,t}$

With $Y_{j,t}$ denoting firm's j outcome (e.g., capital stock, employment), $\Delta_h \ln Y_{j,t+h} = \ln Y_{j,t+h} - \ln Y_{j,t-1}$ the cumulative difference at horizon h.[7] $\nu_{s,t+h}$ denotes a set of industry (2-digit)-year controls and X_j a battery of firm-level characteristics in 2006-2007.[8] As discussed above, one p.p. increase in $Exposure_{j,t}$ means that the minimum wage policy is pushing firms to increase their wage bills by one percent to comply with the new requirements.

In case of a single and permanent minimum wage increase, a local projection of $\Delta_h \ln Y_{j,t+h}$ on $Exposure_{j,t}$ would be enough to capture short and longer-term responses to a single period minimum wage change at t. However, minimum wages also vary between $t+1$ and $t+h$ following the initial change captured in $Exposure_{j,t}$. Therefore, the h-period cumulative change in outcome Y combines the impact of the initial and subsequent minimum wage changes. To account for this staggered nature, equation (1) controls for those minimum wage changes between $t+1$ and $t+h$ through the $\sum_{i=0}^{h-1} b_i Exposure_{j,t+1+i}$ term. Hence, β^h would be the coefficient of interest: the firm-level response to a minimum wage change in t at different year horizons, controlling for subsequent minimum wage changes.

As mentioned previously, one issue is that $Exposure_{j,t}$ is likely to be endogenous. Then, for each relevant year horizon h, I instrument the exposure term using the instrument discussed previously, $Exposure^{IV}_{j,t+h} = \sum_o z_{j,o,2007} mw_{o,t+h}$.

To have a more comparable and intuitive estimate, I translate the impulse responses to minimum wage elasticities ε^y_{t+h}, defined as the percent change in the outcome variable y due to a one percent increase in the labor costs *induced* by the minimum wage:

$$\varepsilon^y_{t+h} = \frac{\Delta_h \ln Y_{j,t+h}}{\Delta_h \ln W_{j,t+h}} = \frac{\beta_{y,t+h}}{\beta_{\text{Wage Bill},t+h}}$$

Capturing long-term responses to minimum wage increases is particularly challenging. As argued by Sorkin (2015), in lack of indexation and if adjustments are phased over several years, nominal minimum wage increases are temporary and even not binding. Inflation and real wages in the relevant labor market erode the initial pressure from the policy rapidly. Costa Rica's setting is pivotal for exploring longer-term responses, as the strong indexation leads to more permanent minimum wage increases.

[7] For firm exit, define $\Delta_h Exit_{j,t+h} = Exit_{j,t+h} - Exit_{j,t-1}$, with $Exit_{j,t}$ an indicator taking the value of one if the firm exits after $t+1$, 0 otherwise.

[8] I measure and fix these characteristics in the 2006-2007 as these two years represent the ending of a long period of real minimum wage stability, as previously discussed. Variables include export share, import share, profitability, labor share, capital share, average industry-level exposure. These covariates control for the relative importance of capital and labor within the firm, international trade exposure, firm size, and how close the firm is to the exit margin. I additionally include the square of these variables, and the average industry-level exposure in 2006-2007

5 Estimation Results

Table 5 summarizes the estimated dynamic responses following equation (1), jointly with the first-stage F-statistics and the overidentification tests. Overall, these statistics confirm a robust first-stage and no evidence of misspecification. More precisely, for all horizons, the overidentification tests do not reject the null that the instruments are valid, i.e., uncorrelated with the error term, and that the excluded instruments are correctly excluded from the estimated equation.

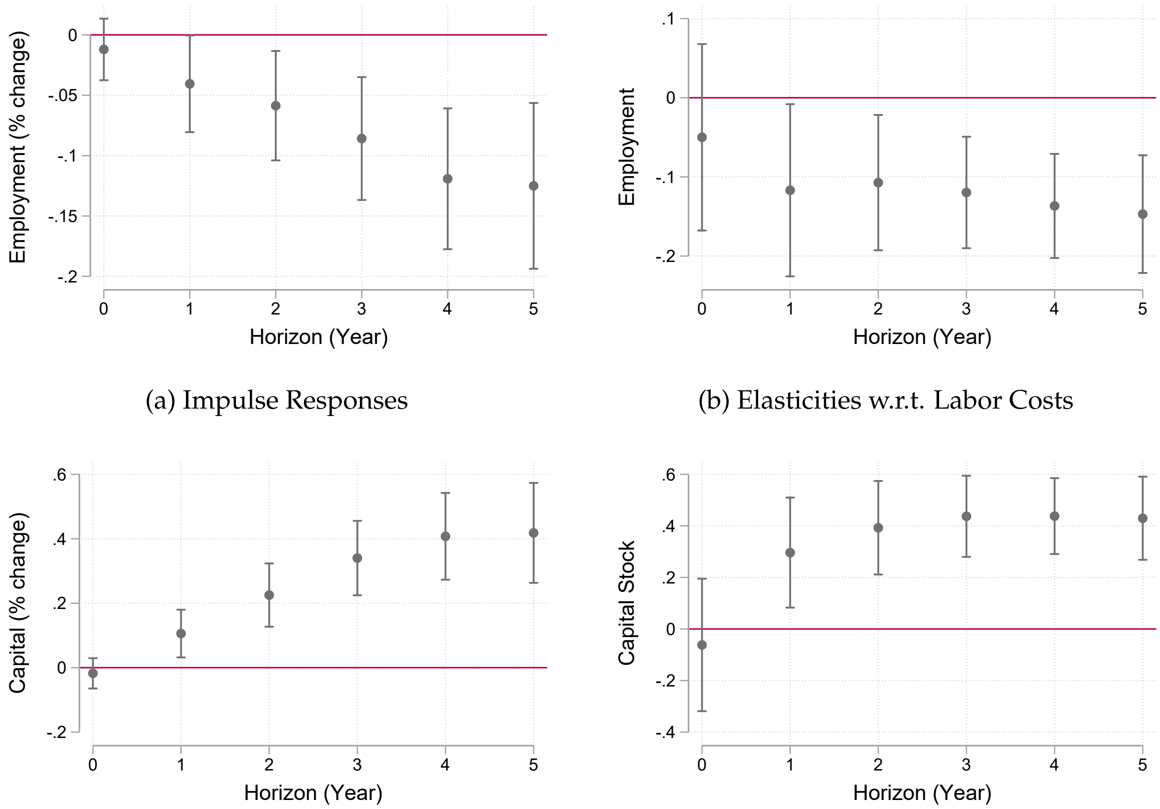
Table 5: Firm Outcome Responses to Minimum Wage Exposure

	Horizon (Year)					
	0	1	2	3	4	5
Wage Bill	0.135*	0.358*	0.577*	0.783*	0.944*	0.998*
	(0.012)	(0.021)	(0.026)	(0.031)	(0.036)	(0.040)
Employment	-0.012	-0.041*	-0.059*	-0.086*	-0.119*	-0.125*
	(0.013)	(0.020)	(0.023)	(0.026)	(0.030)	(0.035)
Capital	-0.018	0.106*	0.225*	0.340*	0.408*	0.418*
	(0.024)	(0.038)	(0.050)	(0.059)	(0.069)	(0.079)
Capital-Labor Ratio	0.034	0.130*	0.257*	0.369*	0.437*	0.519*
	(0.021)	(0.038)	(0.052)	(0.061)	(0.068)	(0.079)
Observations	142,360	120,310	101,791	85,657	71,258	57,805
F-Statistic	4,815.2	3,318.4	383.2	288.9	205.2	145.7
Overidentification Test	[0.219]	[0.127]	[0.128]	[0.505]	[0.942]	[0.571]

Notes: Table shows the log changes in the outcome variable to a one percent point increase the compliance cost to the minimum wage, following equation (1). Robust standard errors in parenthesis. Overidentification test reports the p-value (in brackets) for the null hypothesis that the instruments are valid (no misspecification). * p<0.05

Figure 9 shows that higher minimum wages induce firms to reduce their employment levels and increase their capital stock. Nevertheless, the employment adjustment is relatively small and takes time to materialize. The associated elasticity stabilizes around -0.14. In other words, a one percent increase in the labor costs induced by the minimum wage leads to a 0.14 percent decline in the employment level. For the capital stock, the changes also require a year to become significant. Capital stock elasticities are smaller during the first years after exposure (0.3) and then converge to approximately 0.43.

Figure 9: Minimum Wages, Employment and Capital Stock



(a) Impulse Responses

(b) Elasticities w.r.t. Labor Costs

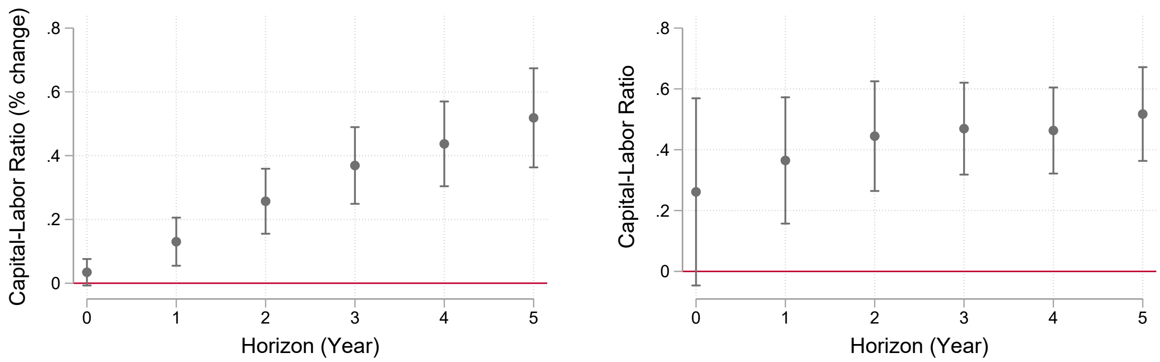
(c) Impulse Responses

(d) Elasticities w.r.t. Labor Costs

Notes: Figures on the left shows the response to a one percent point increase the compliance cost to the minimum wage, computed using equation (1), alongside 95 percent confidence intervals estimated using robust standard errors. Figures on the right shows the respective elasticity (percent change in the outcome variable due to one percent increase in the labor costs induced by the minimum wage) and 95 percent confidence intervals estimated using boostrapped standard errors.

As a first approximation of the elasticity of substitution between capital and labor, I estimate the capital-labor ratio response to minimum wage changes. Figure 10 summarizes the main results. Consistent with greater capital adoption and employment reduction, capital-labor ratios increase after a minimum wage increase. During the first years after the minimum wage change, a one percent increase in the labor costs induced by the minimum wage increases the firm-level capital-labor ratio by 0.26-0.36 percent. In the longer-term horizons, the magnitude of the increase escalates to 0.46-0.51. Such findings emphasize the capital-labor adjustment dynamics after a minimum wage change and a degree of substitutability between these two inputs below one.

Figure 10: Minimum Wages and Capital-Labor Ratios



(a) Impulse Responses (b) Elasticities w.r.t. Labor Costs

Notes: Figures on the left shows the response to a one percent point increase the compliance cost to the minimum wage, computed using equation (1), alongside 95 percent confidence intervals estimated using robust standard errors. Figures on the right shows the respective elasticity (percent change in the outcome variable due to one percent increase in the labor costs induced by the minimum wage) and 95 percent confidence intervals estimated using boostrapped standard errors.

5.1 Labor Demand Model

In Garita (2020), I document that higher minimum wages induce firms to increase their labor shares. The main purpose of this paper is to provide empirical support to such a finding by discussing the capital-labor substitution effects in more detail. Specifically, the firm-level labor share response to higher minimum wages depends on two main forces: the increase in the wage rate induced by the policy and the capital-labor substitutability within the production process. As explained by Elsby et al. (2013); Oberfield and Raval (Forthcoming), the relationship between the labor share and the capital-labor ratio is determined by the magnitude elasticity of substitution between capital and labor $\left(\sigma_{KL} = \frac{d\ln(K/L)}{d\ln(w/r)}\right)$. Under a Cobb-Douglas production function ($\sigma_{KL} = 1$), minimum wage increases would not impact the labor share. In contrast, if labor is less substitutable than in the Cobb-Douglas case ($\sigma_{KL} < 1$), then the labor share would be increasing in the capital-labor ratio (and vice versa).

To provide a theoretical foundation to the discussed results and provide a structural estimation of the labor-capital elasticity of substitution, I estimate a simple input demand model with imperfect competition in the output market based on Hamermesh (1993); Harasztosi and Lindner (2019). In this framework, firms operate in monopolistic competition, and they use capital and labor[9] to produce a differentiated good. Hence, firms can

[9] I abstract from other inputs such as materials to think in terms of value-added. Harasztosi and Lindner (2019) include materials in the same framework, which imposes an additional simplified structure that could be problematic, as it ignores simultaneous interactions between the three inputs. In fact, these authors estimate intermediate-input elasticities of substitution of zero and capital-labor elasticities between 2.3 and

set a price above their marginal cost. Two key structural parameters govern the model. The first one is the elasticity of substitution between capital and labor, measuring the ease of substituting labor for capital within the production process. The second one is the absolute value of the output demand elasticity η, determined by the firm's constant markup. The elasticity of labor demand with respect to the minimum wage is:

$$\varepsilon^{L}_{MW} = -s_L\eta - (1 - s_L)\sigma_{KL} \tag{2}$$

Where s_L is labor's share of value-added. The first term $(-s_L\eta)$ is the scale effect, underscoring the output demand as a factor shaping the disemployment effects associated with the policy. Higher labor costs induce firms to reduce their employment levels, determined by the relative importance of labor (s_L). However, firms also increase their product prices. If the demand is inelastic enough $(\eta < 1)$, revenues increase, reducing the employment demand adjustments. The second term $((1 - s_L)\sigma_{KL})$ is the substitution effect. As minimum wages increase, capital becomes relatively cheaper, triggering input substitution.

Similarly, the capital elasticity with respect to the minimum wage combines the scale effect and the degree of substitutability between capital and labor:

$$\varepsilon^{K}_{MW} = (1 - s_L)(\sigma_{KL} - \eta) \tag{3}$$

Finally, the revenue response will combine both the change in prices and the change in quantities associated with higher minimum wages:

$$\varepsilon^{R}_{MW} = s_L(1 - \eta) \tag{4}$$

The output elasticity of demand determines the sign of the revenue response. An inelastic demand $(\eta < 1)$ indicates that revenues do not fall in response to higher prices.

I use the labor share of value added s_L observed in the data and the estimated minimum wage elasticities $(\varepsilon^{L}_{MW}, \varepsilon^{K}_{MW}, \varepsilon^{R}_{MW})$[10][11] to estimate the structural parameters $\Theta = (\eta, \sigma_{KL})$ by matching the elasticities predicted by the model and the empirical elasticities:

$$\hat{\Theta} = \arg\min_{\Theta} \left(m(\Theta) - \hat{m}\right)' W \left(m(\Theta) - \hat{m}\right)$$

3.4, outside the range found in the related literature. Future work, however, will be oriented in a structural model including these three inputs and more realistic features.

[10] The revenue elasticity is estimated and discussed in Garita (2020).

[11] We can also use $(\varepsilon^{L}_{MW}, \varepsilon^{K}_{MW})$ to estimate η and σ_{KL} using equations (2) and (3). A new draft will present these results.

With $m(\Theta)$ the vector of minimum wage elasticities as function of parameters Θ, $\hat{m}$ the vector of the empirical elasticities and W a weigthing matrix (variance-covariance matrix).

Table 6: Labor Demand Model Estimation

	All	**Manufacturing**	**Tradable**	**Non-Tradable**	**Low Skilled**	**High Skilled**
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
			Outcome elasticities			
Employment Elasticity	-0.110	-0.115	-0.116	-0.088	-0.148	-0.082
Revenue Elasticity	0.35	0.31	0.28	0.40	0.32	0.36
Capital Elasticity	0.47	0.54	0.51	0.40	0.63	0.43
Labor Share s_L	0.54	0.44	0.46	0.56	0.52	0.59
			Estimated Parameters			
Capital-Labor Substitution σ_{KL}	0.585	0.813	0.757	0.461	0.783	0.501
	(0.05)	(0.08)	(0.07)	(0.05)	(0.06)	(0.04)
Output Demand Elasticity η	0.095	0.116	0.154	0.053	0.071	0.130
	(0.13)	(0.14)	(0.13)	(0.13)	(0.13)	(0.13)
SSE	0.207	0.213	0.185	0.220	0.237	0.220

Notes: Table shows the estimated parameters of the labor demand model based on Hamermesh (1993), using a minimum-distance estimator. Column (1) includes all firms in the sample. Column (5) restricts to firms with . Standard errors in parenthesis and SSE denotes the weighted sum of squared errors.

Table 6 summarizes the main results and reports significant heterogeneity in the estimated elasticities and parameters across six main groups: all firms, manufacturing, tradable, non-tradable, firms intensive in low-skilled occupations, and firms intensive in high-skilled occupations. The estimated capital-labor elasticity of substitution is consistently below 1. The computed value for all firms is 0.59. Manufacturing (0.81) and tradable sectors (0.76) exhibit a larger elasticity, while non-tradable sectors a smaller value (0.46). This pattern is consistent with manufacturer workers more likely to perform routine tasks that are more substitutable for capital. But moreover, manufacturing firms produce more tradable goods, so the price adjustment channel is more limited than in services and non-tradable industries, stressing the input demand adjustment channel. The estimated elasticity of demand η supports this argument: the demand seems less inelastic in manufacturing and tradable sectors.

Minimum wages are often exploited to compute the elasticity of substitution between capital and labor. Nevertheless, nearly all the analyzed settings have a single minimum wage affecting low-paid workers and low-skilled occupations. To expand the paper's external validity, I restrict the estimation to firms mainly employing low-skilled and high-skilled occupations. Occupations in the dataset are classified based on the International Standard Classification of Occupations (ISCO-08). Hence, I can easily group occupations based on their skill level. I define a low-skilled intensive firm as the one in which 45 percent of its workers are in occupations involving the performance of simple and routine physical or manual tasks. Similarly, I define a high-skilled intensive firm as the one employing more than 45 percent of its labor force in occupations that involve the performance of tasks requiring complex problem-solving, decision-making, and creativity based on an extensive body of theoretical and factual knowledge in a specialized field.[12] 6 Columns 4 and 5 of Table report the results, showing that the elasticity of substitution is higher in low-skilled intensive firms (0.78) and lower in high-skilled intensive firms (0.50), relative to all firms. Such a difference is consistent with previous studies suggesting that low-skilled labor is less substitutable than high-skilled labor (e.g., Lordan and Neumark (2018); Aaronson and Phelan (2019); Baqaee and Farhi (2019)). Nevertheless, for both groups, the elasticity is below one.

6 Conclusions

This paper offers new empirical evidence on the degree of substitutability between capital and labor. I document that increases in the cost of labor, induced by the minimum wage, leads to a decline in employment and an increase in capital stocks. I exploit Costa Rica's minimum wage setting, which has several advantages. First, the policy binds to an extensive segment of the labor market. Hence, I can account for and report heterogeneity across relevant sectors. Second, the results speak about the substitutability between capital and labor in a broader scope. Prior work exploiting minimum wage hikes restricts to the adjustment dynamics between capital and low-skilled labor.

I find elasticities of substitution consistently below one, suggesting that the substitution away from labor towards capital is not big enough to reduce the labor share after a minimum wage increase. The estimated elasticity of substitution is larger in manufacturing (0.81) and tradable sectors (0.76) but smaller in non-tradable sectors (0.46). Additionally, firms intensive in low-skilled labor exhibit a higher degree of substitutability than firms intensive in high-skilled labor. The contrast in the estimated parameters reflects the differences in the production technologies across sectors, stressing the importance of extending the analysis to different industries when analyzing minimum wage policies.

[12] These two occupational groups represent skill levels 1 and 4 as defined by ISCO-08.

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