Gender Disparities in the Welfare Effect of the Minimum Wage*

Decio Coviello, Erika Deserranno, Nicola Persico

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Abstract

Using administrative data from a major US retailer, we study how women and men respond to, and benefit from, a minimum wage increase. We find that, in terms of overall welfare, the average female worker benefits more from the minimum wage than the average male worker, primarily because women hold lower-paying positions. However, this result flips when comparing workers within the same position: *ceteris paribus*, the welfare of women increases less with the minimum wage hike than that of men, even though both receive comparable pay raises. We show that this occurs because women exert more effort compared to men as a response to the minimum wage hike, a behavior driven by their greater need for job retention due to less favorable external employment options. This evidence points to a generalizable mechanism whereby disparities *outside* the firm beget welfare disparities in the impact of an important gender-neutral policy (i.e., the minimum wage) *inside* the firm.

Keywords: gender inequality, minimum wage, welfare, outside option.

JEL Classification: M52, J30, J24, J60.

*Decio Coviello (HEC Montreal). Erika Deserranno (Northwestern University and Bocconi University). Nicola Persico (Northwestern University). This paper was previously circulated under the title "Gender Differences in Worker Response to the Minimum Wage." We thank Paul Oyer, Maddalena Ronchi, Alan Manning, Till von Wachter for insightful comments on the paper. We also thank participants at the NBER Law and Economics Group, CEPR Labour Workshop, B.L.O.W. Workshop, and seminar participants at the Tilburg University, Tinbergen Institute, Bocconi University, IIES, University Laval, European University Institute, Bicocca University, CEMFI. We are grateful to Elizabeth Jaramillo, Paulina Koenig and Virginia Manzo for their (more than) excellent research assistance. The authors declare they have no relevant or material financial interests that relate to the research described in this paper. This work has been screened to ensure no confidential information is revealed. Data and institutional background are provided in such a way as to not disclose information that may allow the firm to be identified. This paper is undertaken, in part, thanks to funding from the Canada Research Chairs program. E-mails: decio.coviello@hec.ca; erika.deserranno@kellogg.northwestern.edu; nicola@nicolapersico.com.

1 Introduction

This paper addresses an important fairness question: when a job's working conditions improve, do women and men benefit equally? We ask this question in the context of the minimum wage, using data from salespeople at a major US retailer.

We provide a new formula for quantifying the welfare effect of the minimum wage on workers in minimum-wage supported jobs. We show that, in our empirical setting, the *average* female worker benefits more from the minimum wage than the *average* male worker – primarily because women are disproportionately employed in lower-paying positions and so, mechanically, their pay is more exposed to minimum wage adjustments. However, the result flips when comparing workers within the same position: *ceteris-paribus*, women benefit less from a minimum wage increase than men, despite receiving an equal pay raise. This disparity arises because women's pay gains are offset by endogenous effort. Women, we show, put forth more effort than men in response to the minimum wage because they care more about retaining their job due to worse employment opportunities outside our firm. Overall, our findings show that while minimum wage policies can foster gender equalization, this primarily happens because female workers often occupy lower-paying positions than men. However, within the same job position in the firm, the minimum wage disproportionately benefits men, largely due to their more favorable employment options elsewhere. These insights emphasize the need to assess policies by considering their effects on the gender welfare gap, rather than focusing solely on the gender pay gap, as these may not align.

Our evidence comes from salespeople who work at a large US retailer employing more than 10% of department store employees nationwide, and operating more than 2,000 stores across all fifty states. The sample population is broadly representative of US hourly-paid workers, which represent almost 60% of all US workers. Our workers' pay is, in part, based on productivity (sales per hour) which is recorded administratively. When the worker's average hourly pay falls below the minimum wage, the employer is required to pay a "top-up" to make up the difference.

Our data cover 70 minimum wage increases at the state and local levels. Using a border-discontinuity research design, we study the differential gender effects of the minimum wage on pay and welfare, by comparing gender differences in stores where the minimum wage has increased ("treated" stores) with those in stores where it has not ("control" stores) across the *same* county border. Because our stores are composed of two departments, and women disproportionately work in the lower-paying one, our *ceteris-paribus* specification includes department×store fixed effects, and thus effectively compares women and men in the same working conditions within the firm (although the workers' outside options are not held fixed). We also include worker fixed effects to account for potential differences in innate characteristics by gender, such as ability. In the *non-ceteris-paribus* specification, we remove the department×store fixed effects, enabling the estimates to capture within-store differences in department allocation across genders. We first discuss the *ceteris paribus* results, followed by the more "mechnical" *non-ceteris-paribus* results.

Our findings reveal that, *ceteris paribus*, women receive a pay raise comparable to that of men when the minimum wage increases. However, women respond by putting forth additional productivity (larger increase in sales per hour), and are rewarded with extra job stability (larger increase in retention). We argue that this stronger productivity response from women is due to them exerting more effort when the minimum wage increases. This conclusion is reached after ruling out other potential explanations such as gender differences in worker selection, firm adjustments, demand shocks, and pre-trends.

We show that the primary reason women exert more effort following a minimum wage increase is that they have worse outside options compared to men, thereby creating a stronger incentive to retain a job that becomes more attractive with the minimum wage. This is supported by the observation that women's productivity response to the minimum wage is stronger than the men's when, and only when, their outside options (average market wages in the county) are substantially lower than those of men. The disparity in productivity response between genders instead disappears when women and men face similar outside options. This implies that the differences in productivity response do not inherently reflect gender itself, but rather, the differences in external employment opportunities based on gender.[1]

[1] We show that the results are robust to accounting for hours, unemployment duration and hedonic factors when estimating the gender gap in outside option.

To be clear, the gender differential response to the minimum wage in our setting is not driven by a gender differential impact of the minimum wage *on the outside option*. Indeed, we show that the impact of the minimum wage on the outside option is second-order empirically, both for men and for women, because minimum wage jobs are a small fraction of all the jobs that represent the outside option. This is not to say that the minimum wage has zero effect on the outside option, i.e., on economy-wide wages. However, for workers who are *currently* paid the minimum wage, these effects are second-order compared to the first-order effect coming from the protection in their current job.

Having documented gender differentials in the response to the minimum wage, we turn to welfare. Empirically, the effect of the minimum wage on our workers' welfare reflects several countervailing forces. On the one hand, women benefit less than men because they work extra hard after a minimum wage increase (effort cost) and, also, because their pay is topped up less often (because they work harder). On the other hand, women benefit more than men because their variable pay increases by more and, also, they are retained more. To boil down these countervailing effects to a single number, we turn to theory.

We derive a novel (to our knowledge) theory-based formula for the impact of the minimum wage on the welfare of minimum-wage supported workers. The formula says that a worker's welfare gain from a minimum wage increase is the product of two terms neither of which, conveniently, requires information about the (unobservable) effort cost. The first term is akin to a discount factor that captures a worker's expected tenure in her current job, and is higher for women. The second term, the flow benefit of an increased minimum wage, tends to be larger for men because *ceteris paribus* – i.e., comparing women and men in the same department – the men's pay is more frequently topped up. In our calibrations, the second effect dominates, leading us to conclude that the welfare benefits of the minimum wage are larger for male than female workers, *ceteris paribus*.

The welfare estimates flip in the *non-ceteris-paribus* analysis: now, women benefit more than men from the minimum wage. This finding reflects the fact that in our firm (as in the whole economy) women are disproportionately employed in lower-paying positions (in our case, lower-paying departments) and so, mechanically, their pay is topped up more often by the minimum wage. This result supports the idea that the minimum wage is a force for gender equalization because female workers are dissimilarly situated than male workers – even though, among similarly situated workers, a higher minimum wage disproportionately benefits men.

The data in this paper are substantively the same as in Coviello et al. (2022). That paper reveals that workers respond to the minimum wage as predicted by the efficiency wage model.[2] Relative to that paper – and to the literature at large – this paper makes two key contributions. First, it is the first paper, to our knowledge, that quantifies the impact of a gender-neutral policy (here, the minimum wage) on the gender gap *in welfare*, as opposed to pay. This is important because, when costly effort is endogenous, and workers care about retention, pay is not welfare. Second, our case study illustrates empirically that disparities *outside* the firm (gender differences in the outside option) beget disparities *inside* the firm – in our case, disparate welfare impact of a gender-neutral policy. Therefore, a "systemic" gender disparity outside the firm determine optimal regulation inside the firm and may require affirmative correction even within a scrupulously gender-neutral firm.

Our paper contributes to several literatures. First, it contributes to the literature on the disparate impact of the minimum wage by gender. Caliendo & Wittbrodt (2022); Blau et al. (2023); Paul-Delvaux (2023) study the differential gender effect of the minimum wage on wages. In line with our *non-ceteris-paribus* results, they find that a higher minimum wage reduces the gender pay gap because women tend to be overrepresented in lower-paying positions.[3] Whereas these papers focus on wages, we also document the disparate effects of the minimum wage on a rich set of outcomes including retention and, most notably, welfare, in addition to wages. Furthermore, the existing estimates in the literature are not *ceteris paribus*, i.e., they do not compare women and men in the same role. However, when evaluating the "fairness" of a policy, we show that it is important to also make comparisons among workers in similar positions, as these may differ (or even reverse) from comparisons between women and men in different positions.

[2] In Coviello et al. (2022), workers are not separated by gender. Hence, Coviello et al. (2022) does not contribute to understanding the effect of the minimum wage on the gender pay and welfare gap.

[3] Similar results have been documented in the literature on the effect of the minimum wage on the distribution of wages – e.g., DiNardo et al. (1996); Lee (1999); Autor et al. (2016); Cengiz et al. (2019); Fortin et al. (2021); Dustmann et al. (2022). Other papers have studied the non-ceteris-paribus effects of the minimum wage by other "low-paid" categories such as teenagers (Card, 1992; Katz & Krueger, 1992) and minorities (Derenoncourt & Montialoux, 2021; Derenoncourt et al., 2021).

Unrelated to the minimum wage, a number of papers have studied the role of the outside option on the workers' incentives to exert effort and on their productivity. Lazear et al. (2016) show that workers employed in a large US firm are less productive in times of low unemployment, when their outside option is better, and attribute this effect to lower individual effort. However, their analysis focuses broadly on individual worker productivity without differentiating by gender. Separately, improvements in workers' outside options, measured with an unemployment insurance benefit extension, have been shown to increase worker absenteeism in Austria (Ahammer et al., 2023) and to reduce productivity among cashiers in the US (Lusher et al., 2022). We contribute to this literature by shifting the focus to worker welfare. Specifically, we show that policies that improve the *pay* of lower-paid workers (e.g., women) do not necessarily reduce the *welfare gap*.

Finally, we contribute to the literature on gender disparities caused by ostensibly gender-neutral policies. Carry (2022) shows that introducing a legal minimum on working time increases the men's welfare more than the women's, mainly due to the replacement of female part-time workers with male full-time workers. Unlike our paper, the analysis compares men and women in different working conditions. Biasi & Sarsons (2022) show that an increase in wage flexibility benefits men more than similarly situated women, largely because men tend to have better bargaining abilities than women. Antecol et al. (2018) show that, in high-skilled professions, the adoption of gender-neutral tenure clock stopping policies increase gender gaps in tenure. We complement these papers by studying another policy (i.e., the minimum wage) and by focusing on welfare, which has received less attention than pay in the gender gap literature.

The paper proceeds as follows. Section 2 describes the institutional setting and identification strategy. Sections 3 to 5 present the *ceteris-paribus* impact of the minimum wage by gender, and shed light on the important role of the outside option. Section 6 quantifies the *non-ceteris-paribus* impact of the minimum wage by gender, and discusses the external validity of the results. Section 7 concludes.

2 Data and Identification Strategy

2.1 Institutional setting and worker-level data

Our data cover more than 40,000 consultative sales associates working in more than 2,000 stores at a nationwide US retailer from February 2012 to June 2015. Restricting the sample to border stores as per our research design (described in Section 2.2), our analysis covers a sub-sample of more than 200 stores with over 10,000 consultative sales associates, about 7,000 of which are administratively classified as men. Henceforth, all the information we report refers to the restricted "border store sample."

Consultative sales associates assist walk-in customers by answering their queries and demonstrating product features. These tasks, collectively referred to as "exerting effort", involve warmly greeting the customers, patiently explaining and persuading, up-selling higher-margin products, and cross-selling items such as warranties, loans, and credit cards. All sales associates nationwide are subject to the same compensation scheme composed of a fixed and a variable portion; the latter is based on customer purchases which each associate claims as her own sales.

Each store has, on average, 16 consultative sales associates, a manager and, sometimes, one or more assistant managers. In what follows, we describe the summary statistics of male and female consultative sales associates. These statistics are reported in Table A.1. The data vary at the monthly level, and one observation is a worker$\times$month.

Age, tenure, and termination The average worker is 36 years old, and the median age is 27. These numbers are comparable for women and men, and indicate that both populations are relatively young. Measured from the hiring dates listed in the HR records, the average tenure is 58 months for women and 44 months for men (median tenures of 27 and 22 months, respectively). Women exhibit lower termination rates, with 4.1% of women terminated each month compared to 4.8% of men. For both genders, approximately one third of these terminations are classified as "involuntary" by the company. Due to the recognized subjectivity of this classification, we primarily report results that combine voluntary and involuntary terminations. However, we will also present the results for involuntary terminations alone.

As expected, termination (both voluntary and involuntary) is positively correlated with low productivity (low sales per hour); moreover, the function that maps productivity into termination – the termination rule – is the same for both women and men (Table A.2, columns 1 and 2). No consultative sales associates are promoted to managerial positions in our setting, and as a result, career advancement opportunities are limited within the company for both male and female workers.

Pay and department allocation Our consultative sales associates are paid by the hour. Their *regular pay* includes a fixed component (*base hourly pay*) and a variable component (*commissions* based on customer purchases which each associate claims as her own sales). On average, they earn $12 per hour as regular pay, with $6 stemming from the fixed component and another $6 from commissions. In addition, if the weekly average of a worker's regular pay per hour falls below the minimum wage, the employer is required to make up the difference as prescribed by the Fair Labor Standards Act.[4] We create a variable called "minimum wage *top-up*," which equals the average hourly amount paid by the employer to comply with the minimum wage. Approximately 42% of our workers receive some top-up in at least one week of a month and, among these workers, the average top-up amount is $0.50 per hour. However, only 3.2% of our workers receive a top-up in every week of the month (and so are paid exactly minimum wage in that month). Later in our analysis, we will refer to *total pay* as regular pay plus any top-up.

Within a store, employees work in different units that sell different product types. Following an internal company classification, we group units into two "departments," denoted A and B for confidentiality. Employees in department A earn significantly more than their counterparts in department B: refer to Figure A.1, panel A.[5] Since the compensation schemes (base hourly rates and commission rates) and the tasks performed are essentially the same between the two departments,[6] we attribute the difference in pay to the fact that

[4]Under this law, commissioned workers can occasionally be deemed "exempt" and thus not receive a top-up. Based on administrative records, however, all of the workers in our sample are non-exempt.

[5]Median pay is $12.6 per hour in department A and $9.9 in department B. The share of workers who receive a "top-up" over the course of a month is 34% in department A vs. 69% in department B.

[6]See Table A.3, panel A and B. In both departments, salespeople are responsible for making customers happy, providing them with information, increasing sales, helping to maintain the sales floor appearance,

the items sold in department B are more expensive and more popular, resulting in higher sales per hour.

The gender composition differs across departments, with men making up 75% of workers in department A and only 9% in department B. As shown in Figure A.1 (panel B), men earn more than women in our firm and are also substantially less likely to be situated at the lower end of the pay distribution.[7] This is in line with most of the literature on gender disparities in the US economy, which finds that women earn less than men (Goldin, 2014; Olivetti & Petrongolo, 2016; Blau & Kahn, 2017; Petrongolo & Ronchi, 2020) and, therefore, are more likely to benefit from the minimum wage (Blau et al., 2023).

In our setting, these gender pay disparities are entirely explained by the disproportionate allocation of women to the lower-paying department B.[8] In fact, within a department, the gender pay gap disappears. If anything, women appear to earn slightly *more* than men despite facing a similar compensation scheme as men.[9] Furthermore, Figure A.1 (panels C and D) reveals that, within both departments, women are also substantially *less likely* to be at the lower end of the pay distribution, which suggests that, mechanically, they are less likely to benefit from the minimum wage relative to the men in their *same* department. We show next that this is because a substantial portion of worker compensation is linked to "sales per hour," which is less often at the bottom of the distribution for women.[10]

Sales per hour/productivity Both female and male sales associates work an average of 28 hours per week.[11] We compute "sales per hour" as the value of sales divided by the number of hours worked. We refer to sales per hour interchangeably as "productivity."

facilitating customer transactions as needed, and generally cooperating with other employees.

[7]The gender pay gap is 4.5% in our firm: median pay is $11.3 for men and $10.8 for women. Men are more likely to be at the bottom of the distribution: 36% of them receive a "top-up" during the month, compared to 53% of women.

[8]The reasons why women are overrepresented in the lower-paying department B is an important question that extends beyond the scope of this study.

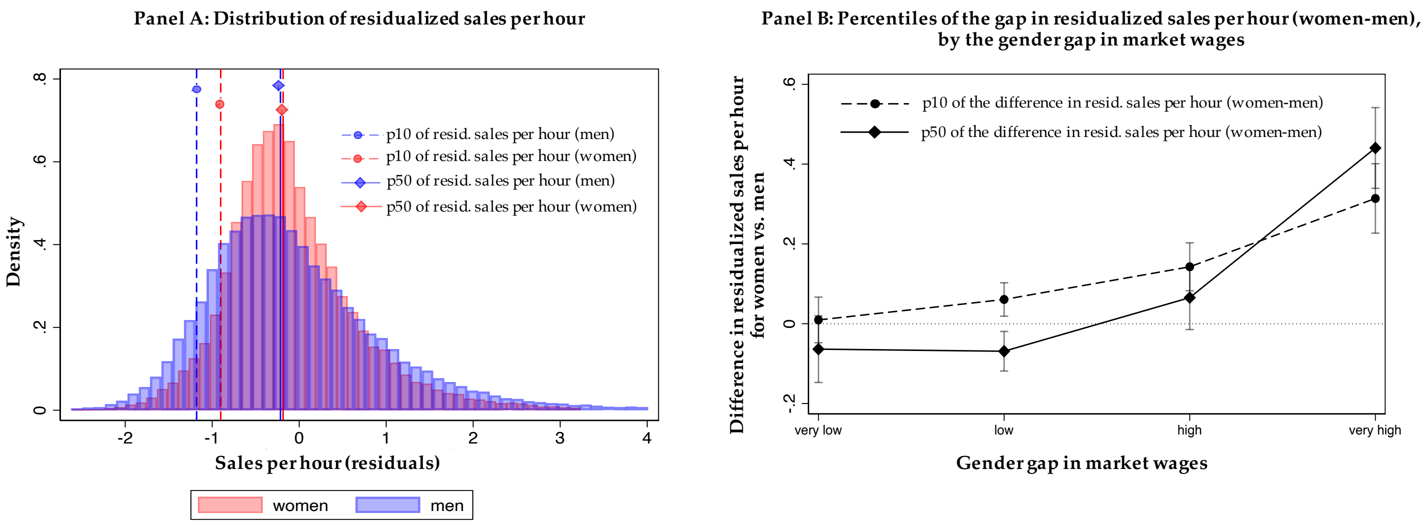
[9]Panels A and B in Table A.3 show that, within each department, women and men have a similar base hourly rate and commission rate. Within department A, the median pay per hour of women is $12.2 vs. $11.3 for men. Within department B, the median pay per hour of women is $9.6 vs. $9.2 for men.

[10]Note that comparisons between men's and women's pay in the *same* pay-for-performance position within the same team are rare in the literature as they require very granular data.

[11]74% of the employees work 35 hours per week (part-time), 35% work 35 hours per week and 7% work 40 hours per week. The distribution of hours is the same for both genders.

Because women disproportionately work in department B, they sell less than men on average;[12] however, this is no longer the case when holding the department fixed. Notably, within a department, women's sales per hour are less likely to be at the lower end of the distribution compared to men's, as shown in Figure 1, panel A. This suggests that, within a department, women are less likely to benefit from the minimum wage top-up.

Figure 1: Sales per Hour (Residuals) by Gender, and by Gender Gap in Market Wages



Notes: Panel A plots the residuals from a regression of sales per hour on worker's department, by gender. One observation is a worker-month. For visual reasons, we remove the top 1% of the residuals. Panel B plots the gender difference in the median (p50) and bottom decile (p10) of the women's vs. men's residuals distribution of panel A when the difference in average market wages in the county (men-women) is very low (below the 10th percentile), low (between the 10th and 50th percentile), high (between the 50th and 90th percentile), very high (above the 90th percentile). The higher the value of the x-axis, the larger is the difference between men's average market wage relative to women's. Bars are 95% confidence intervals.

Figure 1 panel B plots the gender difference (women − men) in the median and bottom decile of the sales per hour distribution in panel A, as a function of the gender gap in labor market wages around the store.[13] We see that the women's productivity premium is limited to periods when their labor market wages are worse than men's. Multiple mechanisms could account for this pattern including women being positively selected when their outside options are worse. In Section 4, we will show that these patterns appear to

[12] Average sales per hour stand at 1.7 for women and 2.3 for men, with units shrouded for confidentiality reasons. The value is rescaled by a factor between $1/50$ and $1/150$ relative to its value in dollars.

[13] We provide details of our measure of labor market wages in Section 4.1. We concentrate on the lowest decile since it encompasses workers who are most likely impacted by changes in the minimum wage.

be mostly explained by women working harder than men (i.e., putting more effort) when their outside options are worse.

Minimum wage variation From February 2012 to June 2015, stores in our sample were affected by 70 minimum wage increases: 49 at the state level and 21 at the county or city level. The prevailing minimum wage in a locality is the highest between the state, county or city level. It has a mean of \$7.84 per hour in our sample, a median of \$7.70 per hour and a standard deviation of half a dollar. The mean minimum wage increase is \$0.54. Appendix B presents a map and a full list of the minimum wage changes, and discusses the data sources.[14]

2.2 Identification strategy

Our empirical specification implements a border discontinuity design in the spirit of Card & Krueger (2000), and closely follows Dube et al. (2010) and Allegretto et al. (2011). Specifically, workers on the side of the border where the minimum wage increased (treatment group) are compared to workers on the other side, where the minimum wage did not increase (control group). This research design has the advantage of ensuring that, apart from the minimum wage change, treated and control groups are similarly situated in terms of local economic conditions and demand shocks. The main disadvantage of this approach is the risk of cross-border worker movements from control to treated stores (Neumark et al., 2014). We will show that this risk is minimal in our setting (see Section 3.3).

Following Card & Krueger (2000), Dube et al. (2010, 2016) and Allegretto et al. (2017), we restrict our sample to stores (and their respective workers) located in adjacent counties that share a border and whose centroids are less than 75 km apart. This subset comprises over 200 stores and more than 10,000 salespeople, half of which experiences variations in the minimum wage during our study period.[15]

[14]In the Appendix, we document the limited overlap of the minimum wage adjustments and other labor market policies that could elicit similar worker responses to those we identify, such as the extension of unemployment benefits.

[15]See Appendix B.2 for details on data construction. We use the 75km threshold because it has been used in the literature. In Section 3.3, we will show that our main results are similar if we use different thresholds, e.g., stores in bordering counties with centroids less than 37.5km or 18.75km apart. This is

***Ceteris-paribus* impact of the minimum wage by gender** In Section 3, we will assess the causal effect of the minimum wage for women vs. men *under the same working conditions*. We estimate the following specification:

$$Y_{idjpt} = \alpha + \beta M_{jt} + \gamma M_{jt} * Woman_i + \eta X_{idjpt} + \delta_i + \zeta_{dj} + \phi_{pt} + \varepsilon_{idjpt}. \quad (1)$$

Y_{idjpt} is the outcome variable of interest (pay, retention, productivity, and, later, welfare) for worker i in department d of store j of county-pair p in month t. $Woman_i$ is an indicator for whether worker i is a woman. M_{jt} is the prevailing minimum wage in store j's jurisdiction in month t, expressed in dollars. The coefficients β and $\beta + \gamma$ capture the effect of increasing the minimum wage by \$1 on men and women, respectively.[16] The coefficient γ captures the differential effect of the minimum wage by gender, which is the focus of this paper. To ensure that this differential effect does *not* capture different working conditions across gender, equation (1) includes department×store fixed effects ζ_{dj}, thus effectively comparing women and men in the same department within the same store. We also include worker fixed effects δ_i to account for time-invariant worker characteristics such as ability.[17]

We implement the border discontinuity design by including county-pair × month fixed effects in equation (1), thus effectively restricting the comparison to "treated" and "control" stores/workers on either side of the *same* border.[18] We estimate this equation by "stacking" our data, meaning that stores/workers located in a county sharing a border with n other counties appear n times in the final sample. The standard errors are two-way clustered at the state level and at the border-segment level.[19]

In our main specification, X includes $M_{jt} * Department_d$. This control accounts, for

reassuring because by narrowing down the definition of bordering counties, we increase the comparability between treated and control stores, with the caveat that it reduces the sample size.

[16]A \$1 increase corresponds to an increase of two standard deviations in the minimum wage, or a 13% rise relative to the average minimum wage level.

[17]We can identify worker and department×store fixed effects because we observe nearly 20% of our workers switching department or store. In Section 3.2, we will show that the minimum wage does not affect the likelihood that female and male workers switch department or store.

[18]These fixed effects interact 113 unique county-pair identifiers with 41 month dummies.

[19]We cluster standard errors this way because the presence of a single county in multiple pairs along a border segment induces a mechanical correlation across county-pairs, and potentially along an entire border segment (Dube et al., 2010). Refer to Appendix B.2 for more details.

example, for the fact that a higher minimum wage may increase demand for products in one department more than another. As a robustness check, we will show that the results are unchanged if we include the following correlates of gender, interacted with the minimum wage, in the vector X: worker tenure, age, work-home distance, and child-bearing age. The economic logic for introducing these controls is explained in Section 3.3. In a separate robustness check, we will show that the results are unchanged if we include department×store×month fixed effects (ζ_{djt}) and thus control for any time-varying department characteristics (e.g., quality of the manager).[20] We will also show that the results are comparable if we estimate equation (1) separately in department A and department B with store instead of store×department fixed effects.

***Non-ceteris-paribus* impact of the minimum wage by gender** Replacing the department×store and worker fixed effects in equation (1) with store fixed effects yields estimates of the differential effect of the minimum wage by gender (γ) which reflect the fact that women are disproportionately represented in the low-paying department relative to men. These *non-ceteris-paribus* results are presented in Section 6.

Summary of Section 2

Every store has two departments. While, within a store, men earn more than women, these gender pay disparities are entirely explained by the disproportionate allocation of women to the lower-paying department. In fact, within each department, women have higher sales per hour, and earn *more* than men. Therefore, the estimates that include department×store and worker fixed effects (which we call the *ceteris-paribus* specification) will be different from those that only control for store fixed effects (*non-ceteris-paribus*).

[20] We do not use these fixed effects in our main specification because their addition does not allow us to estimate the coefficient β.

3 Ceteris-Paribus Impact of the Minimum Wage by Gender

This section documents the impact of the minimum wage on pay, retention probability, and individual productivity (sales per hour) – by gender. Section 3.1 presents the main results: *ceteris paribus*, the minimum wage only modestly reduces the gender pay gap. However, it significantly increases retention for women. We show that this is because women's productivity responds more strongly to the minimum wage than the men's. Section 3.2 argues that the women's stronger productivity response to the minimum wage is, in fact, a stronger *effort* response and rules out other potential mechanisms. Section 3.3 assesses the robustness of our findings and explores two potential threats to identification: violation of the common trends assumption, and cross-border movements.

3.1 Results

The results are presented in Table 1, which reports the estimates of β and γ from equation (1). The p-value for $H_0 = \beta + \gamma = 0$ is presented at the bottom of the table.

The effect of the minimum wage on total pay per hour by gender is documented in column 1. A \$1 increase in the minimum wage increases total pay by \$0.638 per hour (+5.3%) for women and \$0.556 per hour (+4.5%) for men. The effect is not statistically different by gender, though it is higher for women.

Despite the fact that total pay increases similarly for both genders, it is instructive that the *pay composition* (regular pay vs. top-up) does not. Indeed, the increase in regular pay is more than two times higher for women (\$0.553 for women and \$0.215 for men), while the increase in top-up is four times larger for men (\$0.341 for men and \$0.085 for women); see columns 2 and 3.[21] There are two potential explanations for why regular pay – which is composed of both fixed and variable pay – increases more for women than men. First, the

[21] The minimum wage increases the share of men and women receiving some top-up over the month by 18.9 and 12.6 percentage points respectively. While these effects are both large and significant, the effect is significantly larger for men. Similar results are obtained for the number of weeks per month in which the worker receives a top-up. See Table A.4, columns 1 and 2.

Table 1: Impact of the Minimum Wage on Pay, Productivity and Retention by Gender (Ceteris-Paribus Analysis)

	(1)	(2)	(3)	(4)	(5)
		Pay		**Productivity**	**Retention**
Dep.Var.	Total pay per hour = col.(2)+(3)	Regular pay per hour (fixed + variable)	MinW top-up per hour	Sales per hour	Retained
MinW	0.556***	0.215	0.341***	0.059	-0.004
	(0.127)	(0.162)	(0.061)	(0.040)	(0.005)
MinW * Woman	0.082	0.338**	-0.256*	0.055***	0.020***
	(0.163)	(0.124)	(0.126)	(0.016)	(0.003)
Observations	215,558	215,558	215,558	217,746	217,746
Mean Dep.Var.	12.271	12.046	0.225	2.085	0.954
p-value for H_0: MinW+MinW*Woman=0	0.027	0.037	0.330	0.024	0.020
Effect MinW for Men (%)	4.5%	1.8%	194.9%	2.5%	-0.4%
Effect MinW for Women (%)	5.3%	4.7%	26.6%	6.8%	1.7%

Notes: All regressions include store*department fixed effects, worker fixed effects, pair-month fixed effects and control for MinW*department. Standard errors are two-way clustered at the state and border-segment level. *** p<0.01, **p<0.05, * p<0.1. "Total pay per hour" is the monthly total pay (in $ per hour). "Regular pay per hour" is the total amount earned from the base hourly rate and variable pay (commission rate*sales per hour), without the top-up. "MinW top-up per hour" is the monthly total minimum wage adjustment paid by the company to the worker (in $ per hour). The sample size is smaller for the pay variables because we trim the top 1% of the observations due to presence of outliers. "Sales per hour" are the sales per hour rescaled by a factor between 1/50 and 1/150 relative to its $ value. "Retained" is a dummy variable that equals one if the worker is retained that month (i.e., not terminated). "MinW" is the predominant minimum wage in deviation from its sample mean (in $). "Effect MinW for Men (%)" [resp., "Effect MinW for Women (%)"] is the percent effect of a $1 increase in MinW relative to the mean of the outcome variable for men [resp., women].

firm may have changed the compensation *scheme* for women more than men as a response to the minimum wage (i.e., higher base or commission rate). Second, women may have put forth extra productivity, which causes the variable component of their pay to go up and "overtake" the top-up. Our data support the second (productivity) explanation. Table 1, column 4 shows that the productivity of women increases by 6.8% with the minimum wage (significant at the 5% level), whereas the productivity of men increases only by 2.5% (not statistically significant). The gender difference in productivity response is significant at the 1% level. The first explanation, by contrast, is not supported by the data: we find that the minimum wage does not affect the compensation *scheme* by gender.[22] Therefore,

[22]See Table A.4, columns 3-4. This makes sense because we are here comparing women and men in the same department, and our firm has a compensation scheme that is uniform within department.

the gender-balanced increase in total pay hides an extra productivity put forth by women, relative to men.

But, if women become more productive after a minimum wage increase compared to men, how are they rewarded for their extra productivity? Table 1, column 5 shows that they are rewarded in the form of greater retention: female retention goes up by 1.6 percentage points (1.7%) with the minimum wage, with no corresponding effect on the retention of male workers (coefficient of -0.004, not statistically significant).[23] Similar results are obtained when terminations are limited exclusively to "involuntary terminations," as opposed to also considering voluntary termination: see Table A.4, column 5. In Table A.5, we show that the minimum wage does not affect the termination rule – i.e., the function that maps high productivity into lower termination – by gender. Thus, the increase in retention among women is consistent with them working extra hard relative to men.

Overall, this section demonstrates that, *ceteris paribus*, the minimum wage only marginally narrows the gender pay gap. Its effects manifest themselves in two less-studied dimensions: productivity, and retention. The minimum wage increases the productivity of women more than the men's, and women are rewarded for this with higher retention.

3.2 Interpreting the differential productivity response by gender as a differential effort response

Here, we argue that the differential productivity response to the minimum wage which has been documented above is, in fact, an *effort* response. The case is made by ruling out other potential mechanisms that could have produced this gender disparity without a variation in effort response.

Worker selection It could be that, after a minimum wage increase, women's productivity (sales per hour) increases more than the men's because of selection into and out of our worker pool by ability, in a way that differs by gender. For example, stores may have

[23]We will show in the next section that women are not rewarded by being transferred to the highest-paying department or by being assigned more/different hours.

retained more-able women and shed less able ones after a minimum wage increase. Also, hiring into our worker pool may differ by innate ability in a way that is correlated with gender. The presence of worker fixed effects in our specification mitigates these concerns because we effectively compare the "same" worker at two minimum wage levels. Moreover, we replicate our findings in the "non-selected" subsample of workers who were present on the first and the last day of our sample period.[24] When we do this, the sample size drops but the results on productivity are similar to the main sample: see Table A.6, column 1. The results in Section 4.4 will further rule out this selection story.

Firm adjustments It could be that, after a minimum wage increase, women's productivity (sales per hour) increased because the firm: disproportionately reduced the working hours of women; moved women to better shifts or higher-paying departments; or increased the final product prices for women only. Next, we discuss (and reject) each of these mechanisms in turn. In Section 4.4, we show that the firm also did not adjust the monitoring protocols differentially by gender.

The firm might have reduced the working hours for women more than for men in response to a minimum wage increase. This could have resulted in higher productivity per hour for women, especially if fewer working hours reduce "fatigue per hour." Table A.7, column 1, refutes this hypothesis by showing that the minimum wage has no effect on the hours worked by women and men.

Alternatively, the firm might have disproportionately moved women either to better shifts (busier shopping hours) within a department, from department B to department A within a store, or to a more "popular" store. All of these stories are unlikely in our setting. Better shifts are allocated to full-time workers in our firms, and the minimum wage does not disproportionately increase the share of women who become full-time worker relative to men (Table A.7, column 2). Moreover, minimum wage has no differential effect on the likelihood that women vs. men switch department or store (columns 3-4).[25]

[24] A caveat: not all the workers in this subsample are employed *continuously* throughout our sample period. Restricting to continuously employed workers leaves us with few observations for our analysis.

[25] Unlike Dustmann et al. (2022), our workers are not upgraded to "better" stores after the introduction of the minimum wage. They are also not upgraded to "better" departments.

The minimum wage could also have increased women's sales per hour more than men's if, within the same department, the firm increased the prices of feminine more than masculine SKUs. This is unlikely because both genders sell the same products within a department and the company employs nationwide pricing.

Demand shift It could be that, after a minimum wage increase, the demand for female sale representatives jump disproportionately. For this mechanism to confound our estimates, such a shock must affect women more than men *within a given department*.[26] While this possibility seems unlikely, we are able to show directly that the minimum wage causes no demand shifts – gender-specific or not – in our setting. We use the stores' parking lot occupancy as a proxy for store-level demand: see Figure A.2 for an example of the satellite picture from which data on parking lot areas are digitized.[27] Table A.8 shows that the parking lot occupancy rate does not vary with the minimum wage.

Business stealing It could be that, after a minimum wage increase, part of the higher effort put forth by women goes into stealing business from their male colleagues. In this case, the productivity of male workers will be artificially depressed. If business stealing is an issue, its effect should vary with the department's gender composition. However, we find that the estimated impact of the minimum wage on men's productivity is unaffected by controlling for the department's proportion of female employees interacted with the minimum wage (Table A.9, column 6), or when the analysis is limited to male workers who are in departments composed of only men (Table A.6, column 2).

[26] Recall that our estimates include department×store fixed effects (ζ_{dj}) and control for possible differential effects of the minimum wage across departments: $M_{jt} * Department_d$.

[27] We leverage 51,000 satellite images covering 93% of our stores (an average of 2.6 images per store per month). The images are digitized using a machine learning and computer vision algorithm which identifies parking lot areas around each store, counts the number of parking spaces in the parking lot, and counts the number of cars parked. Parking lot occupancy has been used by financial traders to forecast revenues for nationwide retailers, and they are suitable for our purposes because it captures customer volume, which is exogenous to worker effort, as opposed to quantity purchased which is not.

3.3 Robustness checks and threats to identification

This section shows that our main results (Table 1) are robust to adding more fixed effects and controls. It then explores two potential threats to identification: violation of the common trends assumption, and cross-border worker movements.

Robustness We perform several robustness checks. First, we show that the results on pay, productivity and retention are robust to controlling more flexibly for time: Table A.9, column 1, presents the results in a specification that includes department×store×month fixed effects. Second, we show that the results are robust to controlling for potential correlates of gender – worker tenure, age, childbearing age, work-home distance – and their interaction with the minimum wage (Table A.9, columns 2-5). This suggests that the heterogeneous effects by gender are not explained by disparities in these potential correlates of gender. We will provide further evidence consistent with this finding in Section 4.4. Third, we show that the findings are robust to using a log-log specification (Table A.10). Fourth, we show in Table A.6 (columns 3-4) that the results are comparable when restricting the sample to workers in department A or B only.

Pre-trends One might be worried that the differential gender effects we have identified are explained by the outcome variables being on a different trend, within a department, for women and men (even absent the minimum wage increase). We test for differential pre-trends by gender in the twelve months preceding the minimum wage change using a similar autoregressive distributed lag model as in Dube et al. (2010). (This model has the advantage of taking into account the sequential occurrence of changes in the minimum wage level.)

$$\begin{aligned} Y_{idjpt} \quad = \quad & \alpha + \eta_{12-1}(M_{j,t+12} - M_{j,t+1}) + \eta_{1-0}(M_{j,t+1} - M_{j,t}) + \theta_{12-1}(M_{j,t+12} - M_{j,t+1}) * Woman_i \\ & + \theta_{1-0}(M_{j,t+1} - M_{j,t}) * Woman_i + \rho M_{j,t} + X_{idjt}\eta + \delta_i + \zeta_{dj} + \phi_{pt} + \varepsilon_{ijpt}. \end{aligned} \tag{2}$$

Here, $M_{j,t+m}$ is the minimum wage m months after month t, and all other variables are defined as in equation (1). η_{12-1} (η_{1-0}) is a leading coefficient that captures variations in sales per hour during the months -12 to -1 (resp., -1 to 0) from each change in the minimum wage for men. θ_{12-1} (θ_{1-0}) are the corresponding differences across gender. We assess whether men and women are on different trends before the minimum wage increase by estimating whether $\theta_{1-0} - \theta_{12-1}$ is statistically different than zero. Table A.11 shows no gender differential pre-trends preceding changes in the minimum wage.[28]

Cross-border movements Border-discontinuity research designs are vulnerable to the concern that workers may move across borders from control to treated counties (Neumark et al., 2014). This becomes an issue for our identification strategy if women are less inclined to cross borders compared to men, and if men who cross borders are of particularly low/high ability, leading to a change in the ability composition of female vs. male workforce in both treated and control counties following the minimum wage increase. In our specification, this "worker selection" confounder is mitigated by the presence of worker fixed effects, and the fact that very few of our workers transfer to a different store on the opposite side of the same county.[29] Moreover, the results are similar if we restrict our analysis to bordering counties with centroids less than 37.5km, or 18.75km apart, instead of using the 75km threshold (Table A.6, columns 5-6). If cross-border movements were an issue, we should observe changes in the results as we narrow down the definition of bordering counties.

Summary of Section 3

After a minimum wage increase, the total pay of female workers increases comparably to that of males in their same department. However, the productivity of female workers increases more strongly with the minimum wage, for which the women are rewarded with increased job stability. We argue that the women's productivity boost is mediated by effort; we do this by ruling out other mechanisms (differential worker selection, firm adjustments, demand shocks) that could have produced this boost without variation in effort.

[28]The specification is estimated for the sample of 110 thousand workers-months who are continuously employed for 12 months before the minimum wage event in panel A. In panels B-C, we show that there are also no differential pre-trends in the 6 and 3 months preceding the minimum wage change, using the larger sample of 150 and 180 thousand workers-months who are continuously employed for 6 and 3 months before the minimum wage event. The estimated coefficient for $\theta_{1-0} - \theta_{12-1}$ decreases from panel A to C, but it is never statistically significant and a joint test never rejects the lack of pre-trends.

[29]Only 1.2% of male and female workers transfer to a store on the opposite side of the same county.

4 Outside Option as a Mechanism for the Differential Response to the Minimum Wage by Gender

The previous section has shown that the women put forth more effort than men in response to the minimum wage, and that this has implications on their pay composition and their retention. Here, we show that women's stronger productivity/effort response arises only when they have worse outside options than men (Section 4.2). We provide a theoretical argument for why this is the case (Section 4.3), and we rule out alternative mechanisms, other than the outside option, that could account for why women work extra hard after a minimum wage increase (Section 4.4).[30] Before doing all this, we first discuss our proxy of the gender gap in outside option (Section 4.1).

4.1 Proxy for the workers' outside option

The outside option captures the worker's expected future welfare (in net present value terms) right after separating from the firm. Although our data provide detailed information while workers are employed at our firm, we do not observe them after separation. Consequently, we are unable to utilize recent structural methods for calculating the outside option (Caldwell & Danieli, 2023; Schubert et al., 2022). Instead, we will estimate a worker's outside option using gender-specific average market wages in department stores at the county-quarter level, as supplied by the Quarterly Workforce Indicators (QWI).

Using market wages as a proxy for the worker's outside option is an admittedly minimalist approach. However, as we will show in Section 4.2, our findings are robust to considering other plausible determinants of the outside option, beyond market wages:

1. *Hedonic factors:* One could allow the utility differential between being employed by the firm and separating to depend on hard-to-measure hedonic factors which are thought to differ systematically by gender, including preference for commuting or

[30] The important role of the outside option, it is worth remarking, has already surfaced earlier in the paper. Panel B of Figure 1 shows that women are more productive than men only when their outside option is worse.

the presence of children at home.[31] We will show that our results are robust to accounting for proxies of these hedonic factors.

2. *Hours worked:* One could consider that women typically work fewer hours than men on average in the labor market, which means that the gender gap in hourly wages is actually smaller than the gap in total wages. Controlling for gender-specific hours is possible, but only at a lower level of granularity given the available data.[32] We will show that the results are robust to accounting for hours.

3. *Unemployment duration:* Unemployment duration is another contributor to a worker's post-separation welfare. Controlling for gender-specific unemployment duration is also possible, but again only at a lower level of granularity.[33] We will show that the results are robust to accounting for unemployment duration.

4. *Wages across all industries:* We use "department store wages" instead of "wages across all industries" as our proxy for the outside option in our main specification because, empirically, our workers' productivity is much more responsive to the former than the latter (suggesting that workers view department store wages as the relevant outside option).[34] That said, we will show that our results are robust to using the average market wage across all industries as the proxy of outside option (thus, implicitly, assuming that the worker's potential new job could be selected from any industry).

The robustness of our findings to considering other plausible determinants of the outside option suggests that, despite its minimalism, our approach "works:" i.e., the gender

[31] Dislike for commute has been argued to affect the outside option in a way that differs by gender (Le Barbanchon et al., 2021). Note however that, in our sample, the distance between work and home is essentially the same for men and women (mean: 8.9 vs. 8.4 km).

[32] Data on gender-specific hours are available from the Current Population Survey (CPS) at varying levels of granularity depending on the store's location: at the county×year level for some of our stores, at the state×year level for others, and at the year level for the rest. We utilize the most detailed level of data available for each location. Note that the CPS data also contain information on hourly wages. We do not use these data as our main measure of outside option due to the lack of granularity.

[33] During our sample period, gender-specific unemployment duration is only available at the year level, averaged across all US counties and all industries. Unemployment duration in the US was very similar for women and men: the average and median duration was 40 and 20 weeks for men, and 39 and 19 weeks for women (Labor Force Statistics, 2012).

[34] See Table A.2, columns 3 and 4.

gap in local market wages captures enough of the variation in the true unobservable gender gap that folding some or all of the above variables into some (necessarily arbitrary) index of the gap in outside option is unlikely to overturn our results.

Finally, Table A.3 (panels C and D) shows that there is no correlation between the variation in local market wages and the compensation scheme within our firm, which is a measure of the workers' "inside option;" this makes sense because, as explained earlier, our nationwide firm sets the compensation scheme uniformly across stores.

4.2 Gender difference in productivity response co-varies with the gender gap in the outside option

We test whether the gender differential productivity response to the minimum wage varies with the gender gap in outside options by estimating the following interacted version of equation (1):

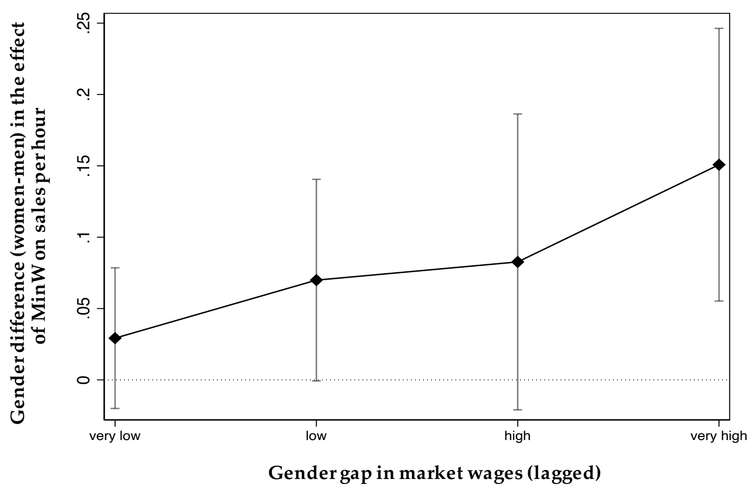
$$\begin{aligned} Y_{idjpt} \quad = \quad & \alpha + \beta M_{jt} + \sum_{k=1}^{4} \lambda_k \mathbb{1}(OutsideGap)^k_{j,t-1} + \gamma M_{jt} * Woman_i + \sum_{k=1}^{4} \rho_k \mathbb{1}(OutsideGap)^k_{j,t-1} * M_{jt} \\ & + \sum_{k=1}^{4} \mu_k Woman_i * \mathbb{1}(OutsideGap)^k_{j,t-1} + \sum_{k=1}^{4} \theta_k M_{jt} * Woman_i * \mathbb{1}(OutsideGap)^k_{j,t-1} \\ & + X_{idjt}\eta + \delta_i + \zeta_{dj} + \phi_{pt} + \varepsilon_{idjpt}, \end{aligned} \tag{3}$$

where $\mathbb{1}(OutsideGap)^k_{j,t-1}$ are four indicators for *within-county* percentile bins. These indicators assess the disparity (men − women) in outside options (market wages) within store j's county at time $t-1$, categorizing it as very low (bottom decile), low (between the 10^{th} and 50^{th} percentile), high (between the 50^{th} and 90^{th} percentile), and very high (top decile), within that county's gender gap distribution.[35] We use these categories – particularly the bottom and top deciles – because we are interested in assessing whether the gender differential productivity response documented in Table 1 (column 4) vanishes when the gender gap in outside options is nearly non-existent (i.e., bottom decile; θ_1), and

[35] As illustrated in Figure A.3, the outside option (measured with department store wages) is higher for male workers than for female workers in 98% of the county-quarters. Therefore, even in the bottom decile, women tend to earn less than men. The bottom (top) decile is \$200 (\$800) per month in favor of men. The median "male premium" is \$392 per month.

compare this to times when the gap is significantly larger (top decile; θ_4). We leverage *within-county* variation in the gender gap to rule out the possibility that our results reflect variation in gender norms, as these presumably remain fixed within a county during our sample period (three years).[36,37] Finally, we lag the gender gap to ensure that it is pre-determined and exogenous to subsequent minimum wage changes.

Figure 2: Gender Differential Impact of the Minimum Wage on Productivity by Gender Gap in Market Wages



Notes: The figure plots the effect of the minimum wage on sales per hour for women relative to men, as a function of the gender gap in average market wages. The higher the value of the x-axis, the larger is the difference between men's average market wage relative to women's. The estimates are obtained from an empirical specification that interacts the minimum wage with being a woman and with four indicators: whether the difference in average market wages (men-women) in department stores in county c and quarter q-1 is very low (below the 10th percentile of the county's distribution), low (between the 10th and 50th percentile), high (between the 50th and 90th percentile), very high (above the 90th percentile). Bars are 95% confidence intervals. The difference between the first and fourth coefficient is significant at the 5% level.

The estimates of the θ's coefficients are presented in Figure 2. We find that female workers respond more strongly than male workers to the minimum wage only at times in which their outside options are substantially lower than the men's, in particular when the gender gap is "very high" (top decile). When the outside options of women and men

[36]Table A.12, panel A, reveals that the within-county variation in the gender pay gap is substantial, accounting for more than half of the between-county variation. In panel B, we show that, within a county, men earn more than women in those times when the share of establishments in male (vs. female) dominated industries is high, and when the labor market is less tight for men than for women (i.e., the number of hires as a function of employment is higher for men than women).

[37]The results are very similar when we leverage between-county variation in the gender gap, indicating that gender norms might in fact not play a major role here.

are similar (bottom decile), the productivity reaction of women is comparable to that of men.[38]

The estimates of the θ's coefficients are robust to extending specification (3) to control for proxies of hedonic factors that may differ systematically by gender – i.e., preference for commuting and the presence of children at home – and their interaction with M_{jt}, $Woman_i$, and $M_{jt} * Woman_i$: see Figure A.4 (panel A).[39] The results are also robust to controlling for local gender-specific hours worked (panel B) and gender-specific unemployment duration (panel C). These results are reassuring as they indicate that hedonic factors, hours and unemployment duration are unlikely to confound the heterogeneous effect by labor market wages. Finally, the results are robust to measuring the outside option using "wages across all industries" rather than "wages in department stores" (panel D).

Overall, the results indicate that the gender gap in outside option plays an important role in understanding whether women and men react differently to changes in the minimum wage. Figure A.5 further highlights the outside option as a key determinant of worker's response to the minimum wage: it shows that workers of both genders respond less strongly to the minimum wage when the level of their own outside option goes up.

4.3 Why the outside option matters

This section makes an intuitive theoretical argument: as the gap between current working conditions and outside option widens due to a minimum wage increase, workers will respond by working harder to reduce the probability of being terminated (efficiency wage theory à la Shapiro & Stiglitz (1984) or Rebitzer & Taylor (1995)). Importantly, though, this effect should be stronger for workers with a worse outside option, as they should care most about keeping their job. The theoretical argument, then, is that the same improvement in working conditions is processed differently by workers with different levels of outside option. Since women have a worse outside option than men, this theoretical

[38] The fourth coefficient in Figure 2 is more than three times larger than the first, with the difference being significant at the 5% level. The second and third coefficients are twice as large as the first. While large, these last two differences are not statistically significant.

[39] To proxy for "preferences for commuting," we use distance between work and home. To proxy for "children at home," we use an indicator for the worker being in childbearing age.

argument neatly organizes the findings in Table 1, Figure 2, and even Figure A.5.

While neat, this theoretical argument hinges on two key empirical assumptions: that, for workers who currently benefit from the minimum wage, increasing the minimum wage actually improves current working conditions more than it improves the outside option; and that this effect is not smaller for women than for men. The rest of this section is devoted to validating these assumptions. The arguments that follow acknowledge that the outside option depends on unemployment duration in addition to labor market wages.

It is reasonable to expect, speaking generally and without specific reference to our setting, that, for workers who *currently* benefit from the minimum wage, a higher minimum wage should widen the gap between inside and outside option. We expect the outside option (in net present value, NPV) to be less sensitive to the minimum wage than the inside option because: (a) unemployment duration is long (average of 39 weeks and median of 20 weeks), and neither its flow value to the worker nor its duration depend on the minimum wage.[40] Therefore, mechanically, a big chunk of the NPV of becoming unemployed is independent of the minimum wage. Furthermore, (b) it is unlikely that our terminated workers will find a new job that is as supported by the minimum wage as her current job is.[41] This is not to say that the minimum wage necessarily has zero effect on the outside option. However, for workers who are *currently* protected by the minimum wage, these effects are likely second-order compared to the first-order effect coming from stronger protection in their current job.

In our empirical setting, we can dig somewhat deeper. Using a similar border discontinuity design as in equation (1) with QWI data, we find that the minimum wage does

[40] Pedace & Rohn (2011); Dube et al. (2016); Gittings & Schmutte (2016) show that unemployment duration is essentially unaffected the minimum wage or, if anything, it increases a bit, which strengthens our argument.

[41] We peg this probability at between 3% and 25%. The US Bureau of Labor Statistics reports that only 3% of hourly workers earned the minimum wage in 2015. Therefore, if an unemployed worker randomly drew her next job from the nationwide job supply, the probability of landing a minimum wage job would be very low. Data from the Atlanta Fed's Wage Growth Tracker, which is constructed using data from the Current Population Survey (CPS), indicate that, in our sample period, about 25% of workers with similar pay as ours transition to an equally- or lower-paying job. This estimate probably overstates our workers' likelihood of transitioning to a minimum wage job, considering that young workers like ours are more likely to be on an upward pay trajectory, and that not all lower-paying jobs are supported by the minimum wage.

not materially impact labor market wages, nor unemployment duration, for the average department store worker of either gender.[42] Therefore, quantitatively, the impact of the minimum wage on the outside option appears small for both genders. Since the minimum wage improves employment conditions equally by gender, it follows that, for currently employed workers, a higher minimum wage indeed widens the gap between the working conditions and outside option equally for women and men. But women respond more sharply because, on average, they have a lower outside option (even after the minimum wage increase).

4.4 Ruling out mechanisms other than the outside option

This section rules out alternative mechanisms, other than the outside option, that could account for why females work extra hard after a minimum wage increase; we discuss some prominent ones next.

Difference in innate characteristics Men and women in our sample could differ in some innate characteristics: ability, risk aversion, propensity to reciprocate (as in the gift-exchange model), or, potentially, cognitive ability to deal with the complexity of employment contracts. For example, Figure 1 panel A could be read as showing a gender difference in innate ability. Such difference could lead men and women to respond differently to the minimum wage: e.g., men may have lower incentive to exert extra effort if their total projected pay (without the minimum wage adjustment) is more likely to fall below the new minimum wage. Our evidence suggests that this is not the case because, when their outside options are the same, men and women respond equally to the minimum wage, suggesting that any such innate differences are unlikely driving our findings.

Difference in job fit The results are also robust to controlling for factors that could affect job fit differentially by gender – specifically, childbearing age, and home-to-work

[42] We estimate $Y_{pt} = \alpha + \beta M_{pt} + \eta X_{pt} + \phi_{pt} + \varepsilon_{pt}$, where Y_{pt} is the outcome in county-pair p in quarter t. Columns 1-3 of Table A.13 reveal that a \$1 increase in the minimum wage raises men's and women's average monthly market wages (in department stores) by 0.4 and 0.8% respectively (neither are statistically significant), and does not significantly affect the wage gender gap. Columns 4-6, similarly, show null effects for unemployment duration.

distance – and their interaction with the minimum wage: see Table A.9, columns 4-5. This indicates that plausible job fit dimensions that are potentially correlated with gender – e.g., women valuing their current job more when they are in childbearing age or when they commute less – do not explain away the gender productivity effect.

Changes in within-firm incentives and monitoring As explained in Section 2, the function that maps sales per hour into lower termination – the termination rule – is the same for both women and men. In principle, it could be that, after a minimum wage increase, the termination rule shifts in a way that accounts for the gender difference in effort response: for instance, monitoring protocols and/or managerial pressure might have changed more for women than men, leading them to exert more effort to avoid being terminated. But, as shown earlier, the termination rules for either gender are *unaffected* by the minimum wage (Table A.5). Furthermore, as depicted in Table A.6 (column 7), the gender-specific effects are similar (though slightly less precise) when restricting the sample to departments overseen by female supervisors, who may be less prone to ratcheting up monitoring differentially on female subordinates.

Alternatively, the firm may have modified the compensation scheme (base and commission rates) exclusively for women, and this may have triggered an increase in effort on their part only. As shown in Table A.4, columns 3-4, the minimum wage does not affect the compensation scheme for either gender. Thus, within-firm incentives did not change differentially by gender with the minimum wage increase.

Summary of Section 4

We show that women's productivity/effort response to the minimum wage is stronger than the men's because they have worse outside options, and maintaining a job that becomes more desirable is more important to them. Interestingly, when men and women have comparable outside options, they have a similar response to the minimum wage. We rule out alternative mechanisms, other than worse outside options, that could account for why females work extra hard after a minimum wage increase, including gender differences in innate characteristics, job fit, and within-firm incentives.

5 Welfare Effect of the Minimum Wage by Gender

This section quantifies the effect of the minimum wage on the welfare of our female vs. male workers. Section 5.1 provides a model, within which Section 5.2 derives the formula for the welfare effect of increasing the minimum wage. Section 5.3 describes how we take the formula to data separately by gender, and Section 5.4 provides the calibration results: *ceteris paribus*, women benefit less than men from the minimum wage increase. Section 6 will show that, *non-ceteris paribus*, the results flip and explains why.

5.1 Model

The model that follows is in the spirit of Rebitzer & Taylor (1995)'s efficiency wage model.[43] A worker (in our empirical setting, a salesperson whose job is to interact with a customer) chooses effort under two incentives: the probability of being terminated, and the wage. The probability of termination is decreasing in worker effort. The expected wage is based on individual performance (in our setting, sales per hour) and is increasing in effort. By law, the wage cannot fall below the minimum wage. The fine details about the model are provided in Appendix C.1.

Primitives Worker effort is denoted by e and has cost $c(e)$. Worker performance (in our case, sales per hour) is a random variable $Y(e)$ that enjoys the strict monotone likelihood ratio property (MLRP) in e. Intuitively, the MLRP means that greater effort produces stochastically higher output.[44]

Consider any continuous nondecreasing compensation scheme $\overline{w}(\cdot)$ that transforms performance into pay. For example, $\overline{w}(Y) = b + cY$, where b represents the base salary and c the commission rate. Since in our firm all workers nationwide are subject to the same compensation scheme, in our model we cannot assume that the compensation scheme $\overline{w}(\cdot)$ is optimally adapted to the local parameters, including the minimum wage M. We assume, instead, that when a locality increases M, $\overline{w}$ does not change.[45] Thus, in a store

[43] Section 4 has shown that efficiency wage theory is "the right model" to describe how workers of both genders respond to the minimum wage.

[44] The MLRP implies first-order stochastic dominance.

[45] This assumption is validated empirically in Table A.4, columns 3 and 4, where we show that when a

that is subject to a local minimum wage M, the expected wage is:

$$w\left(e;M\right)=\mathbb{E}\left(\max\left[M,\overline{w}\left(Y\left(e\right)\right)\right]\right). \tag{4}$$

The function $w\left(e;M\right)$ is bounded below by M and is nondecreasing in all its arguments.[46]

The worker's effort choice problem The worker's effort choice problem is:

$$V^{E}(M)=\max_{e} w(e;M)-c\left(e\right)+\frac{1}{(1+r)}\left[\pi(e)V^{E}(M)+(1-\pi(e))V^{U}(M)\right]. \tag{5}$$

Here, $V^E(M)$ represents the lifetime welfare of a worker who is currently employed by our firm. The numbers $r>0$ and $V^U(M)$ represent, respectively, the discount rate and the lifetime value of becoming unemployed. The function $\pi\left(e\right)$ represents the probability of continued employment, which is assumed to be strictly increasing and continuously differentiable over $[0,1]$.

To simplify the worker's problem, subtract the equation $\left[r/\left(1+r\right)\right]V^U(M)=u^U(M)$ from (5). We get:

$$V(M)=\max_{e} u(e;M)+\frac{1}{(1+r)}\pi(e)V(M), \tag{6}$$

where $V(M)=V^E(M)-V^U(M)$ represents the additional lifetime welfare of a worker who is currently employed by our firm relative to being unemployed, and

$$u(e;M)=w(e;M)-c\left(e\right)-u^{U}(M)$$

represents the flow value of employment, net of flow opportunity cost $u^U(M)$, of a worker who is currently employed and exerts effort e.

To ensure that the maximization problem in (6) is strictly concave in e, we assume $u_{ee}<0$ and $\pi_{ee}\leq 0$. Concavity of u in e may be imparted to u by either of its components, w and c. For example, $u_{ee}<0$ if the wage w is identically equal to the minimum wage, provided that the cost function is strictly convex in e. These assumptions guarantee that

locality increases M, base pay and commission rates in the store do not change. This is not surprising as our firm is nationwide and sets the compensation scheme uniformly across stores and departments.

[46]It is obviously nondecreasing in M. It is nondecreasing in e by stochastic dominance, because the function $\max\left[M,w\left(Y\right)\right]$ is nondecreasing in Y.

the worker's optimal effort $e^*(M)$ is the unique solution to the first-order conditions for problem (6).

The next lemma establishes that, for fixed M, $e^*(M)$ is strictly decreasing in the worker's outside option $V^U(M)$.

Proposition 1. *For any given minimum wage level, the worker's optimal effort $e^*(M)$ is strictly decreasing in the worker's outside option.*

Proof. See Appendix C.1. ■

This property is intuitive: in an efficiency wage model, the worker is motivated to exert effort by the fear of being terminated. When the consequences of being terminated improve, this fear factor attenuates, and effort decreases. Conversely, workers exert more effort when their outside option worsens.

5.2 Formula for the effect of the minimum wage on worker welfare

Next, we compute a formula for how the welfare of a generic worker changes as a function of M. In the following sections, the formula will be taken to the data separately for men and women, meaning that we do not need to *assume* that women and men have the same pay or retention schedules, or the same cost of effort. In fact, the worker's cost of effort happens to drop out of the formula, which is helpful because this function is unobservable.

First (and most insightful) step to compute the welfare formula Rewrite problem (6) as follows:

$$V(M) = u(e^*; M) + \frac{1}{1+r}\pi(e^*)V(M), \tag{7}$$

and rearrange (7) to get:

$$V(M) = \underbrace{\frac{1+r}{1+r-\pi(e^*)}}_{\text{dynamic factor}} \cdot \underbrace{u(e^*; M)}_{\text{static factor}}. \tag{8}$$

Intuitively, the lifetime welfare $V(M)$ of an employed relative to an unemployed worker is the product of two factors. The second factor is the flow difference between the employed and unemployed state; we call this a static factor. The first factor is a discount factor that converts flows into stocks, and depends on the probability π that the worker remains employed; we call this a dynamic factor. Both terms depend on the effort level e^* chosen by the worker.

Differentiate with respect to M and use the envelope (or first-order) condition for problem (6) to get:

$$\frac{dV(M)}{dM} = \left[\frac{1+r}{1+r-\pi(e^*)}\right] u_M(e^*; M). \tag{9}$$

The calculation is presented in Appendix C.2. This expression represents the formula for the change in $V(M)$ due to a change in the minimum wage M. A significant empirical advantage is that formula (9) does not depend on $c(e)$, the worker's cost function in her current employment: conveniently, this term dropped out due to an envelope condition whose economic content is discussed next.

Intuition for formula (9) To get an intuition for expression (9), observe that this formula is simply the partial derivative of $V(M)$ with respect to M, without accounting for the changes in the worker's optimal effort $e^*(M)$. Technically, the reason why these changes drop out of the algebra is the envelope (or first-order) condition for problem (6). Intuitively, the reason why the change in effort does not affect the worker's welfare is that by definition the effort level e^* maximizes the worker's lifetime welfare $V(M)$, so any small change in effort around the baseline level e^* only has second-order effects on $V(M)$.

Even more intuitively, the worker sets e^* to optimally balance two countervailing effects: increasing e increases the probability of retention, and hence the first (dynamic) factor in (8); and it decreases the second (static) factor $u(e; M)$ because, as is apparent from inspecting equation (6), the function $\pi(\cdot)$ being strictly increasing motivates the worker to exert *excessive effort* relative to what is justified solely by static incentives. At the optimal effort choice e^*, changing the worker's effort causes these two effects to move in opposite directions in a way that *exactly offsets* each other. As a result, the only effect on welfare is the one directly caused by the variable M which is not under the worker's control. In other words, when M increases, the envelope condition implies that any welfare change that is mediated by a shift in effort (e.g., increased effort cost, increased wage due to more effort, and increased probability of retention due to a change in effort) has no welfare implications. Only the direct effect of the minimum wage on pay (and, potentially, on the outside option) matters for the welfare calculation.

Expression (9) is deceptively elegant, but its empirical implementation is delicate because it requires computing a counterfactual. Indeed, e^* represents the counterfactual effort that the worker would have exerted *in the absence of a change in the minimum wage*. We will deal with this empirical challenge in the next subsection.

Second step: the complete welfare formula We are interested in the change in lifetime welfare, *inclusive of post-separation future*, of a current employee at our firm. Hence, using the notation of equation (5), we are interested in:

$$\frac{dV^E(M)}{dM} = \frac{dV(M)}{dM} + \frac{dV^U(M)}{dM}.$$

After some algebra presented in Appendix C.3, we get:

$$\frac{dV^E(M)}{dM} = \left[\frac{1+r}{1+r-\pi(e^*)}\right]\left[w_M(e^*;M) + \frac{[1-\pi(e^*)]}{r}u_M^U(M)\right]. \tag{10}$$

Naturally, this formula reduces to (9) when $u_M^U(M) = 0$, i.e., when the post-separation future does not depend on the minimum wage.

5.3 Calibrating the welfare formula, by gender

We estimate the following version of the welfare formula (10) *separately by gender*:

$$\frac{dV^E(M)}{dM} = \left[\frac{1+r}{1+r-\pi\left(e_{t-1}^*\right)}\right]\left[w_M\left(e_{t-1}^*;M_t\right) + \frac{\left[1-\pi\left(e_{t-1}^*\right)\right]}{r}u_M^U(M_t)\right]. \tag{11}$$

Compared with formula (10), the terms involving e^* are lagged relative to M. This is because, in formula (10), e^* represents the counterfactual effort that the worker would have exerted in the absence of a change in the minimum wage. Empirically, using e_t^* would be incorrect whenever the minimum wage changes at t, because contemporaneous effort is endogenous to the prevailing minimum wage M_t. Therefore, we use e_{t-1}^* to proxy for the *counterfactual effort* that the worker would have exerted had the minimum wage remained at level M_{t-1}.

Formula (11) involves calibrated parameters and estimates.

1. For $\pi\left(e^*_{t-1}\right)$ we plug in the average retention rate in month $t-1$, by gender. Hence, within either gender, $\pi\left(e^*_{t-1}\right)$ is the fraction of workers who were retained at t among those who were employed at time $t-1$. We demonstrate the robustness of our results to alternative ways of calibrating π in Appendix D.

2. We set the monthly discount rate r to 2.5%. This level of discounting, which is consistent with Coviello et al. (2022), is larger than is normally assumed in welfare analyses, but is in line with field-experimental evidence on the personal discount factor.[47] We demonstrate the robustness of our results to alternative calibrations of r in Appendix D.

3. We set $u^U_M(M_t)$ to zero. This choice makes sense because we have shown in the previous section that increasing the minimum wage has negligible effect on the worker's future welfare. We relax this assumption in Appendix D and show that the results remain mostly unchanged.

Using the above calibrations, the right-hand side of welfare formula (11) becomes:

$$\frac{1+r}{1+r-\pi\left(e^*_{t-1}\right)} \cdot w_M\left(e^*_{t-1}; M_t\right). \tag{12}$$

To estimate this expression, we proceed in three steps. As a first step, we create the following variable for each worker i:

$$w\left(e^*_{i,t-1}; M_{jt}\right) = \max\left[M_{jt}, \overline{w}\left(Y\left(e^*_{i,t-1}\right)\right)\right], \tag{13}$$

which we refer to as i's "synthetic pay per hour." This variable is the empirical counterpart to expression (4) in the theory. This variable involves a counterfactual: it is the hourly pay that the company would have paid worker i in store j in a month t when the minimum

[47] Yearly personal discount rates are estimated at 28% in a representative sample of the Danish population (Harrison et al., 2002, p. 1612) and as large as 35% for enlisted military personnel (Warner & Pleeter, 2001, p. 49).

wage increases, had the worker made the same sales as in the pre-increase regime.[48] As a second step, we use (13) to create the following variable:

$$\frac{1+r}{1+r-\pi\left(e_{t-1}^{*}\right)}\cdot w\left(e_{i,t-1}^{*};M_{jt}\right), \tag{14}$$

which we refer to as i's "discounted synthetic pay per hour." The third step is to estimate (12) by regressing (14) on M_{jt} (note that (12) has the subscript M but (14) does not) using specification (1).

5.4 Results: Welfare effect of the minimum wage, by gender

A priori, it is ambiguous whether expression (12) should be larger for men or women. Indeed, the first term in (12) is larger for women because, in our data, women have a higher probability of retention π; but the second term is larger for men because, *ceteris paribus*, men are more likely to be at the bottom of the productivity distribution (Figure 1, panel A) and, therefore, to benefit from the top-up (Figure A.1, panels C and D).

The estimates of the welfare effect of the minimum wage (expression 12) by gender are presented in Table 2. We find that a \$1 increase in the minimum wage increases men's lifetime welfare by 10.4% and women's lifetime welfare by 5.8%, where the percentage is expressed relative to the mean of the dependent variable (synthetic pay per hour). This difference, which reveals that men benefit twice as much as women from the minimum wage, is statistically significant at the 5% level. This is the paper's headline finding.[49]

Table A.14 shows that these welfare results are robust to using different values for r and different calibrations of π. They are also robust to relaxing the assumption that the minimum wage does not impact the outside option, i.e., to allowing $u_M^U(M_{jt})$ to be positive in equation (11). These robustness results are discussed in Appendix D.[50]

[48]Specifically, $w\left(e_{i,t-1}^{*};M_{jt}\right)$ is computed as the total pay in the pre-increase regime. plus any top-up if that amount is below the new minimum wage.

[49]It may be worth emphasizing that these coefficients do not express the increase in welfare *as a percentage of baseline welfare*. Indeed, the latter is unobservable: the theory does not provide a method for recovering welfare *levels* due to the presence of the unobservable term $c(e)$.

[50]Table A.14, column 1, reports the *wrong* estimates of the welfare effects where, *incorrectly*, contemporaneous pay replaces synthetic pay in formula (12): the estimates are very different from Table 2. This difference shows that using synthetic pay is essential to estimate the welfare effects correctly.

Table 2: Impact of the Minimum Wage on Welfare by Gender (Ceteris-Paribus Analysis)

	Dep.Var: Discounted synthetic pay per hour
MinW	20.307***
	(3.913)
MinW * Woman	-9.007**
	(3.792)
Observations	197,333
Mean Dep.Var.	195.876
p-value for H_0: MinW+MinW*Woman=0	0.028
Effect MinW for Men (%)	10.4%
Effect MinW for Women (%)	5.8%

Notes: The regression reports estimates from a regression of expression (14) on the prevailing minimum wage interacted with gender, using specification (1). These estimates are interpreted as the effect of the minimum wage on welfare by gender. The regression includes store*department fixed effects, worker fixed effects, pair-month fixed effects and controls for MinW*department. Standard errors are two-way clustered at the state and border-segment level. *** p<0.01, **p<0.05, * p<0.1. "Discounted synthetic pay per hour" is the synthetic pay per hour multiplied by the discount factor $[(1+r)/(1+r-\pi)]$ where r is the monthly discount rate and π is the average monthly retention rate, by gender (lagged by one period). The synthetic pay per hour is the hourly pay that the company would have paid the worker had the worker made the same sales as in the month before the minimum wage increase. This is calculated as the maximum value between the total pay per hour in time t-1 and the minimum wage in time t. "MinW" is the predominant minimum wage in deviation from its sample mean (in $). "Effect MinW for Men (%)" [resp., "Effect MinW for Women (%)"] is the percent effect of a $1 increase in MinW relative to the mean of the synthetic pay per hour for men [resp., women].

Expression (12) reveals that, when effort is endogenous, the contemporaneous increase in pay (i.e., $w_M(e_t^*; M_t)$ where, as opposed to (12), effort is not lagged) is not necessarily a good measure of welfare. So, the fact that the minimum wage happens to increase women's and men's pay by nearly the same amount (Table 1, column 1) is uninformative about welfare. This discrepancy between pay and welfare is due to the fact that women earn their pay boost through a greater effort response. In sum, pay is not welfare.

Summary of Section 5

Several competing forces create gender differences in the welfare effect of the minimum wage. On the one hand, women benefit less than men because they work extra hard after a minimum wage increase (effort cost) but receive a similar pay increase. On the other hand, women benefit more than men because they are retained more. We provide a new formula for the welfare effect of the minimum wage on a worker, which boils down these countervailing effects to a single number. After calibrating this formula for our male and female workers separately, we find that, *ceteris paribus*, the minimum wage increases the welfare of women less than that of men.

6 Non-Ceteris-Paribus Impact of the Minimum Wage by Gender, and External Validity

So far, we have compared female and male workers *ceteris paribus*, i.e., in the same working conditions within the firm (although women's outside option could be, and sometimes was, less favorable). Section 6.1 presents *non-ceteris-paribus* effects of the minimum wage that take into account that, within our firm, women are disproportionately represented in the low-paying department relative to men. Section 6.2 supports the external validity of our *ceteris-paribus* and *non-ceteris-paribus* findings.

6.1 Non-ceteris-paribus effects

Replacing the department×store and worker fixed effects in equation (1) with store fixed effects, makes the gender comparison more similar to papers that study the gender gap in the US economy using establishment fixed effects – e.g., Goldin et al. (2017); Barth & Olivetti (2021). In this *non-ceteris-paribus* specification, the estimates of the effects of the minimum wage on pay and welfare are expected to be larger for women relative to men, at least when compared to the *ceteris-paribus* specification. This is because women are overrepresented in the low-paying department relative to men.[51]

The *non-ceteris-paribus* results on pay are presented in Table 3, column 1. As expected, total pay per hour increases more sharply for women compared to men when the minimum wage increases. This is in contrast to the equal pay increase observed in Table 1, column 1. The results regarding productivity and retention – presented in Table A.15 – remain consistent with our previous findings: women's productivity response is more pronounced than that of men, and their retention increases more.[52] In this *non-ceteris-*

[51]The literature has shown that a significant fraction of the overall gender pay gap is explained by women sorting into lower-paying establishments and, within an establishment, into lower-paying occupations – e.g., Bayard et al. (2003); Goldin et al. (2017); Barth & Olivetti (2021). In our context, the gap is also explained by a more granular type of sorting: that of women into lower-paying departments *within* an occupation×establishment. Indeed, as noted in Section 2, men make up 75% of workers in high-pay department A and only 9% in low-pay department B.

[52]Women's regular pay still increases more than that of men, but women are now equally (rather than less) likely to be "topped-up" after the minimum wage increase.

paribus specification, therefore, women benefit more than men from the minimum wage not only in terms of retention but, also, in terms of pay.

Not surprisingly then, the minimum wage increases welfare (expression 12) for women more than men: welfare increases by 7.4% for women vs. 4.0% for men (Table 3, column 2). Thus, once we move beyond *ceteris paribus* and allow comparison between men and women working in different departments, the welfare gap flips: now, women benefit more than men from the minimum wage. This is because, in our welfare estimates, the fact that women work in the lower-paying department outweighs the fact that they benefit less from the minimum wage in terms of welfare *ceteris paribus*.

Table 3: Impact of the Minimum Wage on Pay and Welfare by Gender (Non-Ceteris-Paribus Analysis)

	(1)	(2)
Dep.Var.	Total pay per hour	Discounted synthetic pay per hour
MinW	0.197	7.656***
	(0.149)	(1.382)
MinW * Woman	0.357***	7.473***
	(0.095)	(2.202)
Observations	215,565	198,033
Mean Dep.Var.	12.27	195.9
p-value for H_0: MinW+MinW*Woman=0	<0.001	<0.001
Effect MinW for Men (%)	1.6%	4.0%
Effect MinW for Women (%)	4.6%	7.4%

Notes: All regressions include store fixed effects, pair-month fixed effects and control for the uninteracted woman dummy. Standard errors are two-way clustered at the state and border-segment level. *** p<0.01, **p<0.05, * p<0.1. Column (2) reports the estimates from a regression of expression (14) on the prevailing minimum wage interacted with gender, using store fixed effects and pair-month fixed effects. "Discounted synthetic pay per hour" is the synthetic pay per hour multiplied by the discount factor $[(1+r)/(1+r-\pi)]$ where r is the monthly discount rate and π is the average monthly retention rate, by gender (lagged by one period). The synthetic pay per hour is the hourly pay that the company would have paid the worker had the worker made the same sales as in the month before the minimum wage increase. This is calculated as the maximum value between the total pay per hour in time t-1 and the minimum wage in time t. "MinW" is the predominant minimum wage in deviation from its sample mean (in $). "Effect MinW for Men (%)" [resp., "Effect MinW for Women (%)"] is the percent effect of a $1 increase in MinW relative to the mean of the outcome variable for men [resp., women].

6.2 External validity

This paper has analyzed a single firm, albeit a very large one. In this section, we discuss the external validity of our findings: Can our findings about the differential impact of the minimum wage by gender generalize to the broader economy? To answer affirmatively, a necessary (though, admittedly, not sufficient) condition is to show that our setting reasonably approximates the conditions of male and female workers *in the economy at large*, in several dimensions: worker pay, retention, and other factors.

Given the absence of comprehensive economy-wide data that compare female and male workers at the same level of granularity as we do,[53] we compute summary statistics at our firm's level (i.e., without controlling for department or store) and juxtapose them to statistics for US workers who are "paid by the hour." These workers are the majority (58%) of all US workers and, likely, a vast majority of minimum wage recipients.[54]

Our workers are paid by the hour and are reasonably representative of the US workers who were paid by the hour in 2015 (henceforth, hourly workers for short). Among hourly workers, 4.1% of women and 2.5% of men were paid "at or below" the minimum wage (US Bureau of Labor Statistics, 2015a). These percentages are similar to the fraction of our workers who receive minimum wage top-ups for four weeks in a month (4.9% of women, 2% of men). The bottom decile of hourly earnings is also very similar to our setting: $9 per hour for male hourly workers and $8.3 for female hourly workers (compared to $8.9 and $8.4 in our setting). Our workers are somewhat younger: the average and median ages are 36 and 27 years old in our sample, compared to a median and average age of 40 years old among all hourly workers.[55] For hourly workers, median hourly earnings was

[53] Indeed, outside of lab experiments, we are unaware of any economy-wide study in the US that controls for job characteristics finer than establishment×occupation, and Figure A.1 panel A shows that such fine distinctions make a big difference for our purposes.

[54] Hourly workers are defined by the US Bureau of Labor Statistics as "the employed wage and salary workers who report that they are paid at an hourly rate on their job." Hourly workers earn substantially less than other workers and are thus presumably much more affected by the minimum wage. In January 2015, for example, the bottom decile (median) of weekly earnings was $187 ($535) for hourly workers vs. $400 ($1,057) for non-hourly workers in the CPS data.

[55] The statistics on the bottom decile of pay and on age are taken from the CPS data, January 2015. In our specific study context, 70% of the workforce comprises men, compared to 44% in the wider retail sector (US Census Bureau, 2020). This higher proportion of men in our sample is attributed to our focus on retail sales roles rather than other lower-wage retail positions, such as cashiers.

\$12.6 for women and \$14.6 for men; for our workers, these figures are slightly lower at \$10.8 and \$11.3, respectively, but the gender pay gap is comparable. The gender disparity in monthly termination rates is also comparable: 2.9% for women and 3.6% for men,[56] as opposed to 4.1% and 4.8% in our setting.

Considering the resemblance between our workers and other hourly-paid US workers, we believe that our results can be generalized to this crucial subpopulation. True, not all workers have a variable component to their pay: but formula (12) continues to hold even for fixed pay workers (the term w_M will simply be replaced by 1), so the welfare methodology provided in this paper applies even though, obviously, the precise empirical estimates might vary.

Summary of Section 6

Once we move beyond *ceteris paribus*, that is, once we compare men and women who are not necessarily in the same department, our setting reproduces several key features of the US labor market for hourly-paid workers: in particular, women earn less and are more likely to benefit from the minimum wage than men.
In this analysis, i.e., without store×department and worker fixed effects, our welfare calibration indicates that women benefit more than men from the minimum wage due to their disadvantaged positions within the firm.

7 Conclusion

This paper has examined an important fairness question: when the minimum wage increases, do both workers benefit equally? To address this question empirically, we study the differential effect of the minimum wage on pay and welfare by gender among more than 10,000 hourly paid salespeople whose pay is partly based on performance, and who are employed by a large US retailer that operates more than 2,000 stores. The sample population is broadly representative – in terms of pay, termination rates, and share of

[56] These rates are calculated as the number of female (male) workers who lost their job and have been searching for a new one in the last month, as a share of the number of employed female (male) workers (US Bureau of Labor Statistics, 2015b).

women – of US hourly-paid workers, which represent almost 60% of all US workers.

We have shown that women benefit less from the minimum wage in welfare terms than men in *ceteris-paribus* working conditions, despite experiencing a similar pay increase. Hence, the paper's first contribution: to demonstrate empirically that the impact of a "facially neutral" improvement (minimum wage increase) in *ceteris-paribus* working conditions can lead to differential worker response by gender, and to a disparate welfare impact. This disparity, we argue, is due to baseline "systemic disparities" that are not under the employer's control – in our case, gender differences in the outside option. Extrapolating from our specific setting, this paper points out that an improvement in working conditions can have disparate welfare impacts on two identically-situated co-workers who differ only in their outside options.

The paper's second contribution is a cautionary point: when effort is endogenous, differences in pay do not necessarily track welfare differences. Indeed, empirically, we find that, *ceteris paribus*, the minimum wage benefits men strictly more than women in welfare terms – but not in pay terms. The cautionary point that "pay is not welfare" speaks to the growing literature on the gender pay gap.

Once we move beyond *ceteris paribus* and allow comparison between men and women working in different departments, the welfare gap flips: now, women benefit more than men from the minimum wage. This finding reflects the fact that in our firm (as in the whole economy) women are disproportionately employed in lower-paying positions and so, mechanically, their pay is more exposed to minimum wage adjustments. These *non-ceteris-paribus* welfare estimates support the idea that the minimum wage is a force for gender equalization because female workers are in disadvantaged positions compared to male workers. However, among similarly situated workers, a higher minimum wage disproportionately benefits men.

References

Ahammer, A., Fahn, M., & Stiftinger, F. (2023). Outside Options and Worker Motivation. *Working Paper*.

Allegretto, S. A., Dube, A., & Reich, M. (2011). Do Minimum Wages Really Reduce Teen Employment? Accounting for Heterogeneity and Selectivity in State Panel Data. *Industrial Relations: A Journal of Economy and Society*, *50*(2), 205–240.

Allegretto, S. A., Dube, A., Reich, M., & Zipperer, B. (2017). Credible Research Resigns for Minimum Wage Studies: A Response to Neumark, Salas, and Wascher. *Industrial and Labor Relations Review*, *70*(3), 559–592.

Antecol, H., Bedard, K., & Stearns, J. (2018). Equal but Inequitable: Who Benefits from Gender-Neutral Tenure Clock Stopping Policies? *American Economic Review*, *108*(9), 2420–2441.

Autor, D. H., Manning, A., & Smith, C. L. (2016). The Contribution of the Minimum Wage to US Wage Inequality over Three Decades: a Reassessment. *American Economic Journal: Applied Economics*, *8*(1), 58–99.

Barth, K. S. P., E., & Olivetti, C. (2021). The Dynamics of Gender Earnings Differentials: Evidence from Establishment Data. *European Economic Review*, *134*(103713).

Bayard, K., Hellerstein, J., Neumark, D., & Troske, K. (2003). New Evidence on Sex Segregation and Sex Differences in Wages from Matched Employee-Employer Data. *Journal of Labor Economics*, *21*(4), 887–922.

Biasi, B., & Sarsons, H. (2022). Flexible Wages, Bargaining, and the Gender Gap. *Quarterly Journal of Economics*, *137*(1), 215–266.

Blau, F. D., Cohen, I., Comey, M. L., Kahn, L., & Boboshko, N. (2023). The Minimum Wage and Inequality Between Groups. *Working Paper*.

Blau, F. D., & Kahn, L. M. (2017). The Gender Wage Gap: Extent, Trends, and Explanations. *Journal of Economic Literature*, *55*(3), 789–865.

Caldwell, S., & Danieli, O. (2023). Outside Options in the Labor Market. *Working Paper*.

Caliendo, M., & Wittbrodt, L. (2022). Did the Minimum Wage Reduce the Gender Wage Gap in Germany? *Labour Economics*, *78*, 102228.

Card, D. (1992). Using Regional Variation in Wages to Measure the Effects of the Federal Minimum Wage. *Industrial and Labor Relations Review*, *46*(1), 22–37.

Card, D., & Krueger, A. B. (2000). Minimum Wages and Employment: A Case Study of the Fast-Food Industry in New Jersey and Pennsylvania: Reply. *American Economic Review*, *90*(5), 1397–1420.

Carry, P. (2022). The Effects of the Legal Minimum Working Time on Workers, Firms and the Labor Market. *Working Paper*.

Cengiz, D., Dube, A., Lindner, A., & Zipperer, B. (2019). The Effect of Minimum Wages on Low-wage Jobs. *Quarterly Journal of Economics*, *134*(3), 1405–1454.

Coviello, D., Deserranno, E., & Persico, N. (2022). Minimum Wage and Individual Worker Productivity: Evidence from a Large US Retailer. *Journal of Political Economy*, *130*(9), 2315–2360.

Derenoncourt, E., Gérard, F., Lagos, L., & Montialoux, C. (2021). Racial Inequality, Minimum Wage Spillovers, and the Informal Sector. *Working Paper*.

Derenoncourt, E., & Montialoux, C. (2021). Minimum Wages and Racial Inequality. *Quarterly Journal of Economics*, *136*(1), 169–228.

DiNardo, J., Fortin, N., & Lemieux, T. (1996). Labor Market Institutions and The Distribution of Wages, 1973-1992: A Semiparametric Approach. *Econometrica*, *64*(5), 1001–1044.

Dube, A., Lester, T. W., & Reich, M. (2010). Minimum Wage Effects Across State Borders: Estimates Using Contiguous Counties. *The Review of Economics and Statistics*, *92*(4), 945–964.

Dube, A., Lester, T. W., & Reich, M. (2016). Minimum Wage Shocks, Employment Flows, and Labor Market Frictions. *Journal of Labor Economics*, *34*(3), 663–704.

Dustmann, C., Lindner, A., Schönberg, U., Umkehrer, M., & Vom Berge, P. (2022). Reallocation Effects of the Minimum Wage. *Quarterly Journal of Economics*, *137*(1), 267–328.

Fortin, N. M., Lemieux, T., & Lloyd, N. (2021). Labor Market Institutions and the Distribution of Wages: The Role of Spillover Effects. *Journal of Labor Economics*, *39*(S2), S369–S412.

Gittings, R. K., & Schmutte, I. M. (2016). Getting Handcuffs on an Octopus: Minimum Wages, Employment, and Turnover. *Industrial and Labor Relations Review*, *69*(5), 1133–1170.

Goldin, C. (2014). A Grand Gender Convergence: Its Last Chapter. *American Economic Review*, *104*(4), 1091–1119.

Goldin, C., Kerr, S. P., Olivetti, C., & Barth, E. (2017). The Expanding Gender Earnings Gap: Evidence from the LEHD-2000 Census. *American Economic Review*, *107*(5), 110–114.

Harrison, G. W., Lau, M. I., & Williams, M. B. (2002). Estimating Individual Discount Rates in Denmark: A Field Experiment. *American Economic Review*, *92*(5), 1606–1617.

Katz, L. F., & Krueger, A. B. (1992). The Effect of the Minimum Wage on the Fast-Food Industry. *Industrial and Labor Relations Review*, *46*(1), 6–21.

Lazear, E. P., Shaw, K. L., & Stanton, C. (2016). Making Do with Less: Working Harder During Recessions. *Journal of Labor Economics*, *34*(S1), S333–S360.

Le Barbanchon, T., Rathelot, R., & Roulet, A. (2021). Gender Differences in Job Search: Trading off Commute against Wage. *Quarterly Journal of Economics*, *136*(1), 381–426.

Lee, D. S. (1999). Wage Inequality in the United States during the 1980s: Rising Dispersion or Falling Minimum Wage? *Quarterly Journal of Economics*, *114*(3), 977–1023.

Lusher, L., Schnorr, G. C., & Taylor, R. L. (2022). Unemployment insurance as a worker indiscipline device? evidence from scanner data. *American Economic Journal: Applied Economics*, *14*(2), 285–319.

Neumark, D., Salas, J. I., & Wascher, W. (2014). Revisiting the Minimum Wage—Employment Debate: Throwing Out the Baby with the Bathwater? *Industrial and Labor Relations Review*, *67*(3_suppl), 608–648.

Olivetti, C., & Petrongolo, B. (2016). The Evolution of Gender Gaps in Industrialized Countries. *Annual Review of Economics*, *8*, 405–434.

Paul-Delvaux, L. (2023). Minimum Wage and Gender Gaps: Evidence from Morocco. *Working Paper*.

Pedace, R., & Rohn, S. (2011). The Impact of Minimum Wages on Unemployment Duration: Estimating the Effects Using the Displaced Worker Survey. *Industrial Relations: A Journal of Economy and Society*, *50*(1), 57–75.

Petrongolo, B., & Ronchi, M. (2020). Gender Gaps and The Structure of Local Labor Markets. *Labour Economics*, *64*, 101819.

Rebitzer, J. B., & Taylor, L. J. (1995). The Consequences of Minimum Wage Laws Some New Theoretical Ideas. *Journal of Public Economics*, *56*(2), 245–255.

Schubert, G., Stansbury, A., & Taska, B. (2022). Employer Concentration and Outside Options. *Working Paper*.

Shapiro, C., & Stiglitz, J. E. (1984). Equilibrium Unemployment as a Worker Discipline Device. *American Economic Review*, *74*(3), 433–444.

US Bureau of Labor Statistics. (2015a). *Characteristics of Minimum Wage Workers.* https://www.bls.gov/opub/reports/minimum-wage/2015/home.htm.

US Bureau of Labor Statistics. (2015b). *Labor Force Characteristics by Race and Ethnicity.* https://www.bls.gov/opub/reports/race-and-ethnicity/2015/.

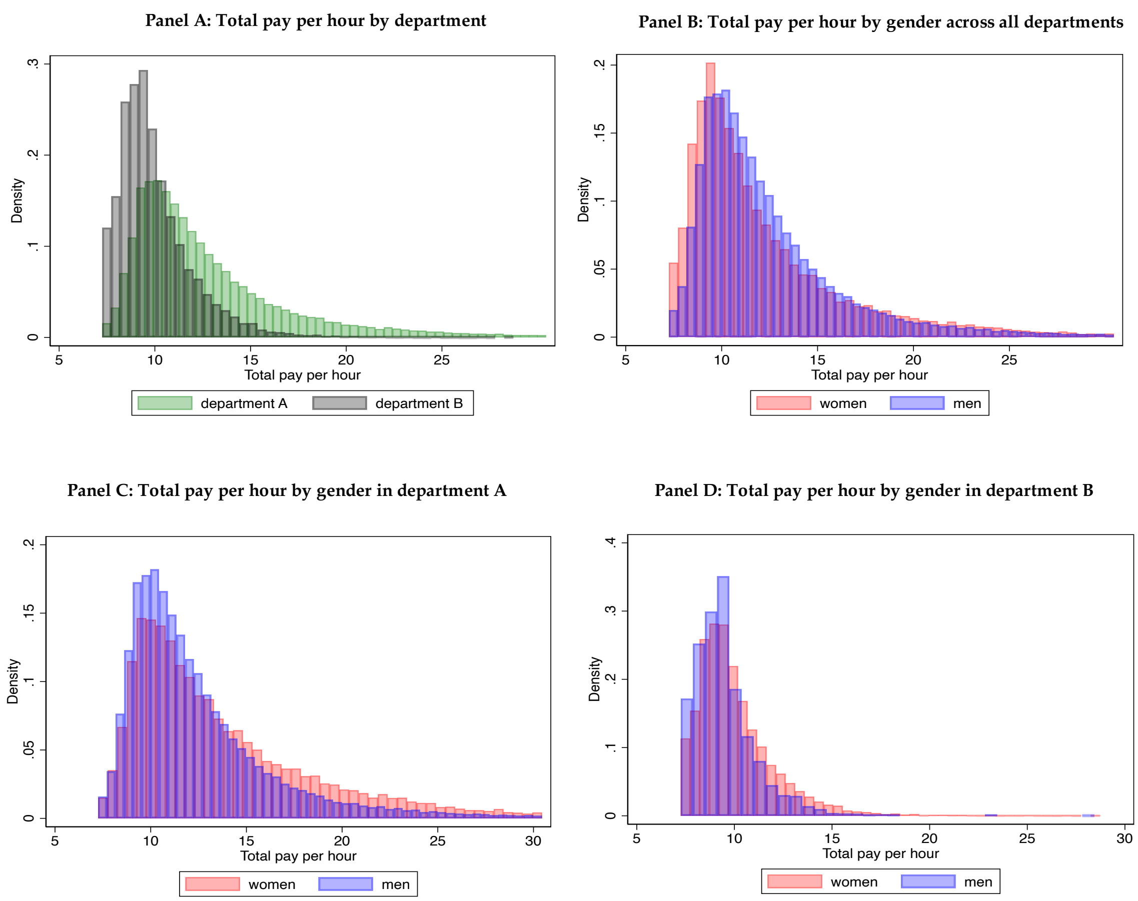
US Census Bureau. (2020). *Retail Jobs Among the Most Common Occupations.* https://www.census.gov/library/stories/2020/09/profile-of-the-retail-workforce.html.

Warner, J. T., & Pleeter, S. (2001). The Personal Discount Rate: Evidence from Military Downsizing Programs. *American Economic Review*, *91*(1), 33–53.

ONLINE APPENDIX

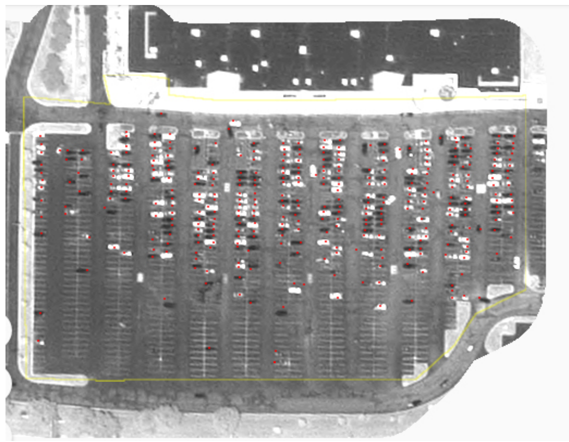
A Appendix Figures and Tables

Figure A.1: Distribution of Total Pay per Hour by Gender Across and Within Department



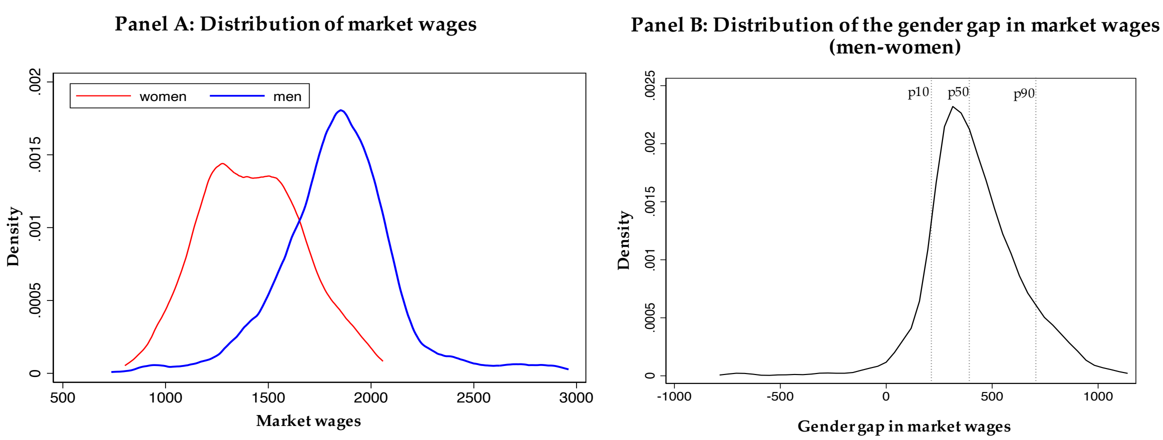
Notes: This figure presents the distribution of total pay per hour by department (panel A), by gender (panel B), by gender within department A (panel C), by gender within department B (panel D). One observation is a worker-month. For visual reasons, we remove observations in which total pay per hour is below the minimum wage (0.6% of the sample).

Figure A.2: Satellite Image of a Parking Lot



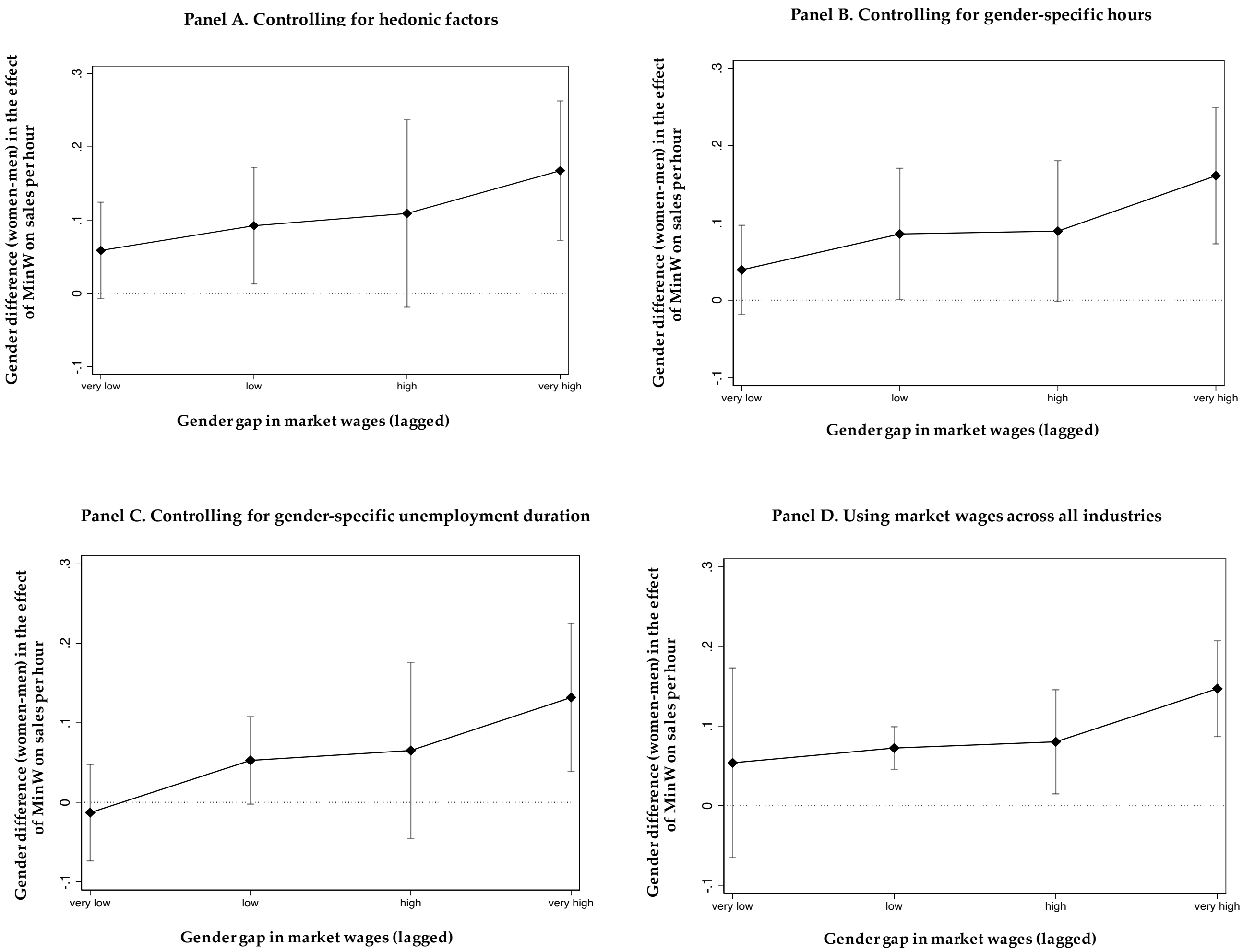
Notes: Satellite imagine of one store's parking lot area with car counts highlighted.
©2018 RS Metrics; Imagery © (CNES) 2018; Distribution Airbus DS Imagery © 2018 DigitalGlobe.

Figure A.3: Gender Gap in Market Wages



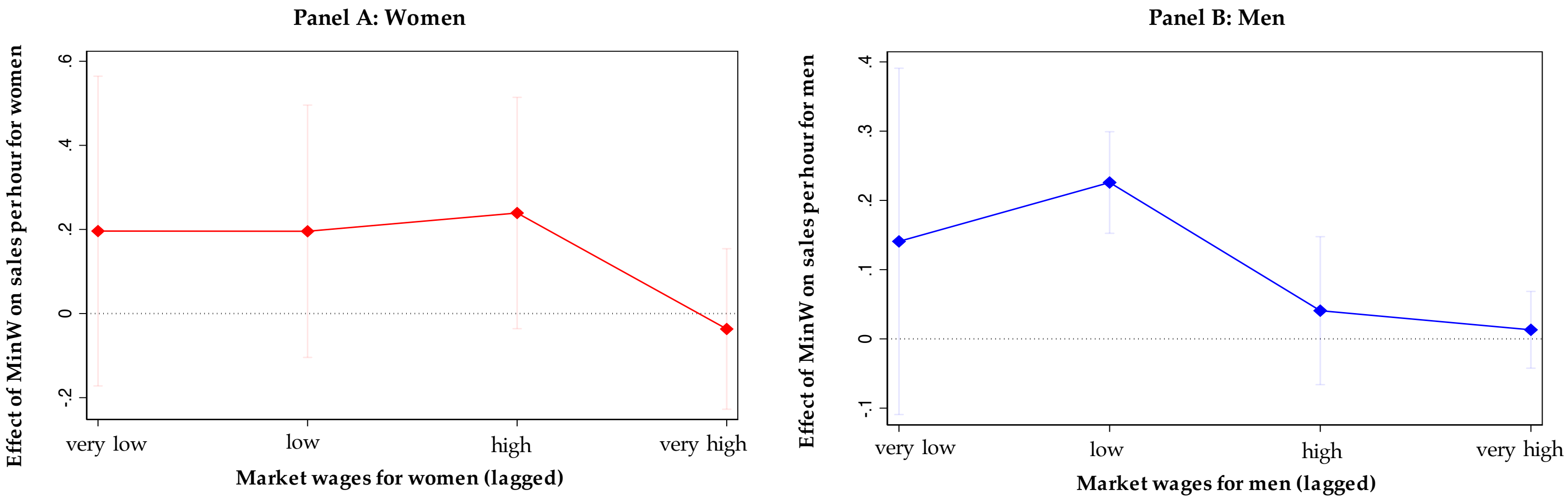
Notes: Panel A plots the county-level distribution of the average market wage in department stores for women and men, separately. Panel B plots the county-level distribution of the gender gap in average market wage (men-women). The average market wage is taken from the QWI data and is defined for employees who were on the payroll on the last day of the reference quarter in a given county. One observation is a county*quarter. The sample is restricted to the counties in our border sample from 2012 to 2015. For visual reasons, we drop the top 1%.

Figure A.4: Gender Differential Impact of the Minimum Wage on Productivity by Gender Gap in Market Wages – Robustness



Notes: The figure plots the effect of the minimum wage on sales per hour for women relative to men, as a function of the gender gap in average wages in department stores (Panels A, B, C) and average wages across all industries (Panel D). The higher the value of the x-axis, the larger is the difference between men's average market wage relative to women's. The estimates are obtained from an empirical specification that interacts the minimum wage with being a woman and with four indicators: whether the difference in average market wages (men-women) in county c and quarter q-1 is very low (below the 10th percentile of the county's distribution), low (between the 10th and 50th percentile), high (between the 50th and 90th percentile), very high (above the 90th percentile). In Panel A, we control for a workers' childbearing age and work-home distance, interacted with the minimum wage and with minimum wage*gender. In Panel B, we control for the gender-specific average number of hours worked per week. Data on gender-specific hours are available from the CPS ("typical hours per week from all jobs") at varying levels of granularity depending on the store's location: at the county*year level for some of our stores, at the state*year level for others, and at the year level for the rest. We utilize the most detailed level of data available for each location. In Panel C, we control for gender-specific unemployment duration. Unemployment duration is available at the year-level by gender (averaged across all US counties). Bars are 95% confidence intervals.

Figure A.5: Impact of the Minimum Wage on Women's and Men's Productivity by Market Wages



Notes: The figure plots the effect of the minimum wage on sales per hour for women (panel A) and men (panel B), as a function of their own respective average market wages. The coefficients are estimated from two separate regressions (one for women and one for men) of sales per hour on the minimum wage interacted with whether the average market wage in department stores in quarter q-1 is very low (below the 10th percentile of the county's distribution), low (between the 10th and 50th percentile), high (between the 50th and 90th percentile), very high (above the 90th percentile). Bars are 95% confidence intervals.

Table A.1: Summary Statistics by Gender

	(1)	(2)	(3)	(4)	(5)	(6)
Sample		Women			Men	
	Mean	S.D.	Median	Mean	S.D.	Median
# observations [#workers-months]		76,336			141,410	
Worker characteristics, tenure and termination						
Age (in years)	36.37	16.74	28.20	35.54	17.15	27
Work-to-home distance (in km)	8.407	8.588	6.485	8.869	9.189	7.008
Tenure (in months)	57.74	73.09	27	44.16	59.65	22
Terminated = {0, 1}	0.041	0.199	0	0.048	0.214	0
Department allocation						
Department A (vs. department B)	0.601	0.490	1	0.977	0.148	1
Compensation structure						
Base hourly rate (in $)	6.097	1.235	6	6.144	1.112	6
Commission rate (in %)	2.871	1.763	2.435	2.441	1.447	2.065
Pay: total, regular and top-up						
Total pay per hour (in $)	12.14	4.177	10.78	12.34	3.786	11.27
Regular (fixed+variable) pay per hour (in $)	11.82	4.727	10.58	12.17	4.126	11.17
MinW top-up per hour (in $)	0.319	2.069	0.048	0.175	1.540	0
Minimum wage top-up frequency						
MinW top-up at least one week of the month = {0, 1}	0.534	0.499	1	0.359	0.480	0
MinW top-up all weeks of the month = {0, 1}	0.049	0.215	0	0.020	0.142	0
Number of weeks with minW top-up (1 to 4)	1.019	1.197	1	0.595	0.954	0
Hours worked						
Number of hours per week	27.62	4.847	25	27.570	4.802	25
Part-time = {0, 1}	0.620	0.485	1	0.593	0.491	1
Productivity						
Sales per hour (shrouded units)	1.665	1.353	1.411	2.311	1.477	2.094

Notes: This table reports summary statistics for women and men separately, across all departments. "Terminated" is a dummy variable that equals one if the worker is terminated that month (i.e., not retained). "Base hourly rate" is the monthly base rate per hour worked (in $ per hour). "Commission rate" is the earnings from commissions divided by sales (in %). "Total pay per hour" is the monthly total pay (in $ per hour). "Regular pay per hour" is the total amount earned from the base hourly rate and variable pay (commission rate* sales per hour), without the top-up. "MinW top-up per hour" is the monthly total minimum wage adjustment paid by the company to the worker (in $ per hour). "Number of weeks with minW top-up" is the number of weeks over the months in which the worker is paid a positive minimum wage adjustement by the firm (1 to 4). "Sales per hour" are the sales per hour rescaled by a factor between 1/50 and 1/150 relative to its $ value.

Table A.2: Predictors of Termination and Sales per Hour

	(1)	(2)	(3)	(4)
Dep.Var.	Terminated (voluntary or involuntary)	Involuntary terminated	Sales per hour	
Woman	-0.014***	-0.009***		
	(0.002)	(0.002)		
Sales per hour	-0.004***	-0.005***		
	(0.001)	(0.001)		
Sales per hour * Woman	0.002	0.001		
	(0.001)	(0.001)		
Average market wage (department stores)			-0.105**	
			(0.041)	
Average market wage (all industries)				-0.041**
				(0.020)
Tenure (in years)			-0.002	-0.001
			(0.006)	(0.005)
Tenure^2			0.000	0.000
			(0.000)	(0.000)
Observations	217,746	217,746	217,746	217,746
Mean Dep.Var.	0.046	0.018	2.085	2.085

Notes: All the regressions include store*department fixed effects, and pair-month fixed effects. The regressions in columns (1) and (2) also control for sales per hour interacted with the department. Standard errors are two-way clustered at the state level and at the border-segment level. *** p<0.01, **p<0.05, * p<0.1. "Terminated" is a dummy variable that equals one if the worker is terminated that month (i.e., not retained). "Involuntary termination" is a dummy variable that equals one if the worker is teminated that month and the termination is categorized as "non-voluntary." "Sales per hour" are the sales per hour rescaled by a factor between 1/50 and 1/150 relative to its $ value. "Average market wage (department stores) [resp., (all industries)]" is the gender-specific average monthly wage in department stores [resp., across all industries] in the county in which the employee works. It is measured for women and men separately using the QWI data, and is expressed in thousands of $. "Tenure" is expressed in years.

Table A.3: Summary Statistics by Gender for Each Department and by Gender Gap in Market Wages

	(1)	(2)	(3)	(4)	(5)	(6)
Sample		**Women**			**Men**	
	Mean	S.D.	Median	Mean	S.D.	Median
Panel A: Department A						
# observations [#workers-months]		45,878			138,226	
Base hourly rate (in $)	5.809	1.323	6	6.146	1.104	6
Commission rate (in %)	2.898	1.766	2.375	2.437	1.443	2.060
Total pay per hour (in $)	13.57	4.615	12.17	12.41	3.796	11.33
Regular (fixed+variable) pay per hour (in $)	13.30	5.337	11.97	12.24	4.133	11.23
MinW top-up per hour (in $)	0.167	1.554	0	0.270	2.625	0
Panel B: Department B						
# observations [#workers-months]		30,458			3,184	
Base hourly rate (in $)	6.523	0.944	6.500	6.041	1.386	6.500
Commission rate (in %)	2.836	1.758	2.528	2.629	1.592	2.326
Total pay per hour (in $)	10.03	2.074	9.549	9.477	1.651	9.209
Regular (fixed+variable) pay per hour (in $)	9.641	2.302	9.223	8.951	1.970	8.800
MinW top-up per hour (in $)	0.391	0.642	0.184	0.526	0.673	0.254
Panel C: Gender gap (men-women) in market wages below the median						
# observations [#workers-months]		35,636			59,334	
Base hourly rate (in $)	6.160	1.276	6	6.175	1.136	6
Commission rate (in %)	2.838	1.764	2.382	2.46	1.475	2.082
Total pay per hour (in $)	12.28	4.173	10.89	12.34	3.746	11.25
Regular (fixed+variable) pay per hour (in $)	11.94	5.094	10.66	12.13	3.835	11.13
MinW top-up per hour (in $)	0.341	2.811	0.066	0.213	0.591	0
Department A (vs. department B)	0.601	0.490	1	0.972	0.165	1
Panel D: Gender gap (men-women) in market wages above the median						
# observations [#workers-months]		40,700			82,076	
Base hourly rate (in $)	6.028	1.185	6	6.117	1.088	6
Commission rate (in %)	2.908	1.76	2.497	2.424	1.42	2.049
Total pay per hour (in $)	11.99	4.176	10.66	12.34	3.822	11.29
Regular (fixed+variable) pay per hour (in $)	11.7	4.281	10.48	12.2	4.371	11.20
MinW top-up per hour (in $)	0.295	0.516	0.032	0.141	2.048	0
Department A (vs. department B)	0.601	0.490	1	0.982	0.131	1

Notes: This table reports summary statistics for women and men separately in different samples: workers in department A (panel A), workers in department B (panel B), workers in counties*months with below median gender gap in market wages (panel C), workers in counties*months with above median gender gap in market wages (panel D). The gender gap in market wages is the difference in average market wages (men-women) in department stores in county c and quarter q-1. "Base hourly rate" is the monthly base rate per hour worked (in $ per hour). "Commission rate" is the earnings from commissions divided by sales (in %). "Regular pay per hour" is the total amount earned from the base hourly rate and variable pay (commission rate* sales per hour), without the top-up. "MinW top-up per hour" is the monthly total minimum wage adjustment paid by the company to the worker (in $ per hour).

Table A.4: Impact of the Minimum Wage on Top-up, Compensation Scheme and Involuntary Terminations by Gender (Ceteris-Paribus Analysis)

	(1)	(2)	(3)	(4)	(5)
	Top-up		**Compensation scheme**		**Termination**
Dep.Var.	MinW top-up at least one week of the month	Number of weeks with minW top-up	Base hourly rate	Commission rate	Involuntary terminated
MinW	0.189***	0.539***	-0.093	0.044	0.003
	(0.013)	(0.030)	(0.061)	(0.031)	(0.005)
MinW * Woman	-0.063***	-0.125***	0.032	0.017	-0.014***
	(0.015)	(0.029)	(0.108)	(0.030)	(0.003)
Observations	215,558	215,558	215,558	192	217.746
Mean Dep.Var.	0.423	0.743	6.128	2.583	0.018
p-value for H_0: MinW+MinW*Woman=0	<0.001	<0.001	0.511	0.209	0.046
Effect MinW for Men (%)	52.4%	90.6%	-1.5%	1.8%	7.3%
Effect MinW for Women (%)	23.4%	40.6%	-1.0%	2.1%	-35.1%

Notes: All regressions include pair-month fixed effects, worker fixed effects and control for MinW*department. All the regressions. Standard errors are two-way clustered at the state level and at the border-segment level. *** p<0.01, **p<0.05, * p<0.1. "Number of weeks with minW top-up" is the number of weeks over the months in which the worker is paid a positive minimum wage adjustement by the firm. "Base hourly rate" is the monthly base rate per hour worked (in $ per hour). "Commission rate" is the earnings from commissions divided by sales (in %). The value is missing for workers with zero sales per hour (hence, the smaller sample size). "Involuntary termination" is a dummy variable that equals one if the worker is teminated that month and the termination is categorized as "non-voluntary." "MinW" is the predominant minimum wage in deviation from its sample mean (in $). "Effect MinW for Men (%)" [resp., "Effect MinW for Women (%)"] is the percent effect of a $1 increase in MinW relative to the mean of the outcome variable for men [resp., women].

Table A.5: Impact of the Minimum Wage on the Termination Rule by Gender

	(1)
Dep.Var.	Retained
MinW	-0.004
	(0.006)
MinW * Woman	0.021***
	(0.005)
Sales per hour	0.011***
	(0.001)
Sales per hour * Woman	-0.002***
	(0.001)
MinW * Sales per hour	-0.000
	(0.001)
MinW * Sales per hour * Woman	-0.001
	(0.002)
Observations	217,746
Mean Dep.Var.	0.954

Notes: The regression includes store*department fixed effects, pair-month fixed effects, worker fixed effects and controls for MinW*department and sales per hour*woman. Standard errors are two-way clustered at the state and border-segment level. *** p<0.01, **p<0.05, * p<0.1. "MinW" is the predominant minimum wage in deviation from its sample mean (in $). "Sales per hour" are the sales per hour rescaled by a factor between 1/50 and 1/150 relative to its $ value.

Table A.6: Impact of the Minimum Wage on Pay, Productivity, and Retention by Gender in Different Subsamples (Ceteris-Paribus Analysis)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Sample restricted to …	Workers with 40 months of tenure or more	Males workers in departments with only men	Department A	Department B	Stores located in counties with centroids <37.5 km apart	Stores located in counties with centroids <18.75 km apart	Workers with a female supervisor
Panel A: Dep.Var. = Total Pay per Hour							
MinW	0.404***	0.626	0.593***	0.432	0.507**	0.781	0.662*
	(0.115)	(0.411)	(0.140)	(0.405)	(0.199)	(0.420)	(0.363)
MinW * Woman	0.097		0.099	-0.136	0.042	0.166	0.237
	(0.085)		(0.176)	(0.163)	(0.192)	(0.218)	(0.287)
Observations	78,412	13,550	181,919	33,474	150,750	98.931	58,020
Mean Dep.Var.	12.803	11.857	12.694	9.980	12,387	12,585	12.179
p-value for H_0: MinW+MinW*Woman=0	0.009		0.029	0.351	0.146	0.146	0.176
Panel B: Dep.Var. = Sales per Hour							
MinW	0.078	0.340	0.086**	0.215*	0.025	0.086	0.022
	(0.083)	(0.215)	(0.036)	(0.111)	(0.045)	(0.063)	(0.058)
MinW * Woman	0.069***		0.047**	0.155**	0.054**	0.081**	0.043
	(0.016)		(0.022)	(0.065)	(0.020)	(0.025)	(0.042)
Observations	79,714	13,595	184,099	33,482	152,523	100,212	58,514
Mean Dep.Var.	2.063	2.408	2.265	1.095	2.025	1.973	2.059
p-value for H_0: MinW+MinW*Woman=0	0.911		0.011	0.390	0.155	0.090	0.385
Panel C: Dep.Var. = Retained							
MinW	-0.001	0.004	-0.007	0.008	0.005	-0.002	-0.016
	(0.010)	-0.039	(0.004)	(0.028)	(0.006)	(0.001)	(0.011)
MinW * Woman	0.009***		0.020***	0.034	0.016***	0.015***	0.022***
	(0.003)		(0.003)	(0.021)	(0.003)	(0.004)	(0.003)
Observations	79,714	13,595	184,099	33,482	152,523	100,212	58,514
Mean Dep.Var.	0.979	0.950	0.954	0.954	0.956	0.957	0.957
p-value for H_0: MinW+MinW*Woman=0	0.521		0.026	0.032	0.011	0.002	0.634

Notes: Regressions are estimated in a subsample of observation, described in the panel headings. Column 3 onward include store*department fixed effects, worker fixed effects, pair-month fixed effects and control for MinW*department. Columns 1 and 2 include store fixed effects, worker fixed effects, and pair-month fixed effects. Standard errors are two-way clustered at the state and border-segment level. *** p<0.01, **p<0.05, * p<0.1. "Total pay per hour" is the monthly total pay (in $ per hour). "Sales per hour" are the sales per hour rescaled by a factor between 1/50 and 1/150 relative to its $ value. "Retained" is a dummy variable that equals one if the worker is retained that month (i.e., not terminated). "MinW" is the predominant minimum wage in deviation from its sample mean (in $).

Table A.7: Impact of the Minimum Wage on Hours and Moves by Gender (Ceteris-Paribus Analysis)

	(1)	(2)	(3)	(4)
	Hours		**Moves**	
Dep.Var.	Hours per week	Part-time worker	Move to high-pay department within same store	Move to another store
MinW	0.273	-0.021	0.001	0.001
	(0.255)	(0.017)	(0.001)	(0.008)
MinW * Woman	0.065	-0.024	0.001	-0.000
	(0.056)	(0.020)	(0.001)	(0.003)
Observations	217,746	217,746	217,746	217,746
Mean Dep.Var.	27.590	0.603	0.084	0.086
p-value for H_0: MinW+MinW*Woman=0	0.179	0.010	0.300	0.946
Effect MinW for Men (%)	1.0%	-3.5%	1.5%	1.1%
Effect MinW for Women (%)	1.2%	-7.3%	2.2%	0.6%

Notes: All regressions include pair-month fixed effects, worker fixed effects and control for MinW*department. Regressions in columns (1) and (2) also include store*department fixed effects. Regression in column (3) includes store fixed effects. Standard errors are two-way clustered at the state level and at the border-segment level. *** p<0.01, **p<0.05, * p<0.1. "Hours per week" is the average number of hours worked in a week. "Move to high-pay department within same store" is a dummy variable for whether a worker moved from the low- to high-pay department (B to A) within the same store. "Move to another store" is a dummy variable for whether a worker moved to another store, regardless of whether she/he moved to the same or a different department. "MinW" is the predominant minimum wage in deviation from its sample mean (in $). "Effect MinW for Men (%)" [resp., "Effect MinW for Women (%)"] is the percent effect of a $1 increase in MinW relative to the mean of the outcome variable for men [resp., women], with the exception column (3) in which it is the percent effect relative to the fraction of workers who switched departments.

Table A.8: Impact of the Minimum Wage on Demand (Store-Level Analysis)

	(1)
Dep.Var.	Parking occupancy (in %)
MinW	0.301
	(0.892)
Observations	17,529
Mean Dep.Var.	22.45%

Notes: One observation is a store*month. The regression includes pair-month fixed effects and store fixed effects. Standard errors are two-way clustered at the state level and at the border-segment level. *** p<0.01, ** p<0.05, *p<0.1. "MinW" is the predominant monthly minimum wage (in $). "Parking occupancy" is the average occupancy rate of the store's parking lot in a given month (in %).

Table A.9: Impact of the Minimum Wage on Pay, Productivity, and Retention by Gender, with Additional Controls (Ceteris-Paribus Analysis)

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Dep.Var. = Total Pay per Hour						
MinW		0.601***	0.572***	0.553***	0.547***	0.547***
		(0.124)	(0.130)	(0.124)	(0.127)	(0.125)
MinW * Woman	0.124	0.117	0.108	0.087	0.091	0.103
	(0.138)	(0.155)	(0.164)	(0.164)	(0.173)	(0.147)
Observations	215,312	215,558	215,558	215,558	208,308	215,558
Mean Dep.Var.	12.271	12.271	12.271	12.271	12.271	12.271
p-value for H_0: MinW+MinW*Woman=0		0.010	0.021	0.024	0.031	0.018
Panel B: Dep.Var. = Sales per Hour						
MinW		0.046	0.048	0.060	0.067	0.055
		(0.041)	(0.041)	(0.040)	(0.042)	(0.041)
MinW * Woman	0.058**	0.047**	0.040**	0.052***	0.049***	0.068***
	(0.024)	(0.019)	(0.018)	(0.015)	(0.015)	(0.015)
Observations	217,500	217,746	217,746	217,746	210,302	217,746
Mean Dep.Var.	2.085	2.085	2.085	2.085	2.085	2.085
p-value for H_0: MinW+MinW*Woman=0		0.086	0.097	0.024	0.016	0.007
Panel C: Dep.Var. = Retained						
MinW		-0.011*	-0.007	-0.003	-0.002	-0.004
		(0.005)	(0.006)	(0.005)	(0.006)	(0.005)
MinW * Woman	0.020***	0.016***	0.016***	0.019***	0.019***	0.018***
	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.004)
Observations	217,500	217,746	217,746	217,746	210,302	217,746
Mean Dep.Var.	0.954	0.954	0.954	0.954	0.954	0.954
p-value for H_0: MinW+MinW*Woman=0		0.361	0.088	0.018	0.019	0.070
Extra controls in the regression:						
Department*store*month fixed effects	✓					
Tenure (above median) & MinW*Tenure		✓				
Age (above median) & MinW*Age			✓			
Childbearing age & MinW*Childbearing age				✓		
Work-home distance & MinW*Work-home distance					✓	
Share of women & MinW*Share of women						✓

Notes: All regressions include store*department fixed effects, worker fixed effects, pair-month fixed effects, MinW*department and the extra controls indicated at the bottom of the table. Standard errors are two-way clustered at the state and border-segment level. *** p<0.01, **p<0.05, * p<0.1. The dependent variables are in the panel headings. "Total pay per hour" is the monthly total pay (in $ per hour). "Sales per hour" are the sales per hour rescaled by a factor between 1/50 and 1/150 relative to its $ value. "Retained" is a dummy variable that equals one if the worker is retained that month (i.e., not terminated). "MinW" is the predominant minimum wage in deviation from its sample mean (in $). "Tenure" and "age" are dummy variables that indicate whether tenure/age are above the median. "Childbearing age" is proxied with a dummy variable that takes value one if the worker is between 25 and 55 years old. "Work-home distance" is the distance in km between the house of the worker and the store. The variable is missing values for 7k worker-months. "Share of women" is the share of workers who are women in the department*store in which the employee works in month t. All the variables are de-meaned such that the coefficient for "MinW" picks up the effect of the minimum wage for men when the variables are equal to the sample mean, and the results are comparable across columns.

Table A.10: Impact of the Minimum Wage on Pay and Productivity by Gender with a Log-Log Specification (Ceteris-Paribus Analysis)

	(1)	(2)
Dep.Var.	Log total pay per hour	Log sales per hour
Log MinW	0.327***	0.292
	(0.118)	(0.304)
Log MinW * Woman	0.057	0.188**
	(0.096)	(0.078)
Observations	215,558	217,746
Mean Dep.Var. (not log)	12.271	2.085
p-value for H_0: MinW+MinW*Woman=0	0.004	0.078

Notes: Log-log specification. All regressions include store*department fixed effects, worker fixed effects, pair-month fixed effects, and control for MinW*department. Standard errors are two-way clustered at the state and border-segment level. *** p<0.01, **p<0.05, * p<0.1. "Log total pay per hour" is the log of the monthly total pay (in $ per hour). "Log of sales per hour" are the log of the sales per hour rescaled by a factor between 1/50 and 1/150 relative to its $ value. "Log MinW" is the log of the predominant minimum wage (in $).

Table A.11: Test of Pre-trends by Gender

	(1)	(2)	(3)	(4)
		Pay		Productivity
Dep.Var.	Total pay per hour = col.(2)+(3)	Regular pay per hour (fixed + variable)	MinW top-up per hour	Sales per hour
Panel A: 12-Months Pre-Trend				
Pre-trend (12 months)	0.094	0.116	-0.022	-0.081
	(0.303)	(0.080)	(0.357)	(0.147)
Pre-trend (12 months) * Woman	0.051	-0.120	0.172	0.159
	(0.216)	(0.114)	(0.317)	(0.120)
Observations	111,933	111,933	111,933	113,648
Mean Dep.Var.	12.271	12.046	0.225	2.085
p-value for H_0: Pre-trend+Pre-trend*Woman=0	0.604	0.915	0.612	0.428
Panel B: 6-Months Pre-Trend				
Pre-trend (6 months)	-0.011	0.049	-0.060	-0.029
	(0.188)	(0.049)	(0.224)	(0.142)
Pre-trend (6 months) * Woman	-0.294	-0.051	-0.243	0.023
	(0.245)	(0.051)	(0.289)	(0.105)
Observations	149,010	149,010	149,010	150,924
Mean Dep.Var.	12.271	12.046	0.225	2.085
p-value for H_0: Pre-trend+Pre-trend*Woman=0	0.236	0.958	0.271	0.955
Panel C: 3-Months Pre-Trend				
Pre-trend (3 months)	0.163	0.063	0.100	0.104
	(0.215)	(0.058)	(0.252)	(0.144)
Pre-trend (3 months) * Woman	-0.444*	-0.010	-0.435	-0.067
	(0.232)	(0.054)	(0.271)	(0.139)
Observations	178,394	178,394	178,394	180,466
Mean Dep.Var.	12.271	12.046	0.225	2.085
p-value for H_0: Pre-trend+Pre-trend*Woman=0	0.234	0.252	0.200	0.671

Notes: "Pre-trend (j months)" corresponds to the estimate of $\eta(1{-}0) - \eta(j{-}1)$ in the specification reported in the paper and "Pre-trend (j months)*Woman" to the estimate of $\theta(1{-}0) - \theta(j{-}1)$, where j is equal to 12 in panel A, 6 in panel B and 3 in panel C. All regressions include store*department fixed effects, worker fixed effects, pair-month fixed effects and control for MinW*department. Standard errors are two-way clustered at the state and border-segment level. *** p<0.01, **p<0.05, * p<0.1. "Total pay per hour" is the monthly total pay (in $ per hour). "Regular pay per hour" is the total amount earned from the base hourly rate and variable pay (commission rate*sales per hour), without the top-up. "MinW top-up per hour" is the monthly total minimum wage adjustment paid by the company to the worker (in $ per hour). "Sales per hour" are the sales per hour rescaled by a factor between 1/50 and 1/150 relative to its $ value. The sample size is smaller for the pay variables because we trim the top 1% of the observations due to presence of outliers.

Table A.12: Within-County Variation in the Gender Pay Gap (County-Level Analysis)

	(1)	(2)	(3)	(4)	(5)	(6)
Dep.Var.	Gender gap (men-women) in market wages					
Panel A: Variance decomposition						
Overall	273.72					
Between	237.08					
Within	134.20					
Panel B: Correlates						
Difference in the number of establishments in male vs.		23.637*				19.993
female dominated industries		(12.658)				(13.128)
Gender gap (men-women) in labor market slackness			19.828**			18.482**
			(8.032)			(8.507)
Average income per capita				-11.688		31.907
				(72.235)		(52.754)
Unemployment rate					25.992	28.374**
					(17.566)	(11.372)
Observations		916	843	888	916	815
R-squared		0.762	0.772	0.762	0.762	0.781

Notes: The data are the quarter QWI data. One observation is a county-quarter. In panel B, all regressions include county and quarter*year fixed effects. Standard errors are clustered at the county level. *** p<0.01, **p<0.05, * p<0.1. "Gender gap in market wages" is the difference between men's average market wage in department stores relative to women's. "Difference in the number of establishments in male vs. female dominated industries" is the difference between the number of establishments in construction vs. education/services. "Gender gap (men-women) in labor market slackness" is the difference (men-women) between the number of workers hired in a quarter as a function of employed workers in quarter t-1. The coefficients represent the effect of a one standard deviation increase in the row variable on the gender gap in market wages.

Table A.13: Impact of the Minimum Wage on Market Wages and Unemployment Duration (County-Level Analysis)

	(1)	(2)	(3)	(4)	(5)	(6)
	Market wages			Unemployment duration		
	Men (in $)	Women (in $)	Gap: Men vs. Women (in %)	Men (in quarters)	Women (in quarters)	Gap: Men vs. Women (in %)
MinW	6.618	11.527	-0.031	0.005	0.009	-0.002
	(86.271)	(48.079)	(0.049)	(0.014)	(0.011)	(0.004)
Observations	11,024	11,149	10,958	78,036	78,124	78,123
Mean Dep.Var.	1,780	1,377	0.1471	1.971	1.808	0.076

Notes: The regressions use QWI data. One observation is a county-quarter. All regressions include pair-quarter fixed effects, and control for the log of county population and total private sector employment (as in Dube et al. 2016). Standard errors are two-way clustered at the state level and the border-segment level. *** p<0.01, **p<0.05, * p<0.1. For each dependent variable, we present the effect for men, women and the percentage difference between the two (women-men/women). "Market wages" (in $) measures the county-level monthly average market wage in the years 2012-2015 in department stores, by gender. "Unemployment duration" measures the county-level unemployment duration (expressed in quarters) for terminated workers, in all the years before 2012 (which explains the larger sample size), and across all industries. Data on unemployment duration lack industry specificity and are not accessible for our sample years.

Table A.14: Impact of the Minimum Wage on Welfare by Gender – Robustness (Ceteris-Paribus Analysis)

Dep.Var: Discounted synthetic pay per hour

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
	Estimating the wrong welfare results	**Sensitivity of welfare results to aggregation of π**			**Sensitivity of welfare results to r**		**Sensitivity of welfare results to assumptions on the outside option**		
Assumptions	Use *contemporaneous* pay per hour rather than *synthetic* pay per hour	π = average retention rate *across all months*, by gender	π = average retention rate, measured *each month*, *by gender* department*	π = average retention rate, measured *each month*, *by gender* department*store*	r = 0.41%	r = 1.5%	outside option u^U = average market wage across all industries, by gender	outside option u^U = average market wage in department stores, by gender	outside option u^U = average market wage in department stores weighted for unemployment duration, by gender)
	$\frac{1+r}{1+r-\pi(e_t^*)} \cdot w(e_{it}^*; M_{jt})$	$\frac{1+r}{1+r-\pi(e^*)} \cdot w(e_{i,t-1}^*; M_{jt})$	$\frac{1+r}{1+r-\pi_d(e_{t-1}^*)} \cdot w(e_{i,t-1}^*; M_{jt})$	$\frac{1+r}{1+r-\pi_{dj}(e_{t-1}^*)} \cdot w(e_{i,t-1}^*; M_{jt})$	$\frac{1+r}{1+r-\pi(e^*)} \cdot w(e_{i,t-1}^*; M_{jt})$		$\frac{1+r}{1+r-\pi(e^*)} \cdot \left(w(e_{i,t-1}^*; M_{jt}) - \frac{1-\pi(e^*)}{r} u^u(M_{jt})\right)$		
MinW	9.714***	20.790***	23.541***	100.716***	29.139***	24.049***	22.306***	23.832***	23.383***
	(2.841)	(3.991)	(4.224)	(19.533)	(5.526)	(4.595)	(3.364)	(5.278)	(5.470)
MinW * Woman	1.224	-12.418***	-20.515***	-58.012***	-17.670***	-14.451***	-13.873***	-17.158***	-15.663***
	(5.753)	(1.682)	(5.191)	(16.585)	(2.487)	(1.985)	(3.034)	(0.746)	(0.662)
Observations	215,558	197,333	197,333	197,333	197,333	197,333	197,333	151,745	151,745
Mean Dep.Var.	191.856	179.451	197.566	387.359	250.200	207.161	285.555	230.147	193.910
p-val. MinW+MinW*Woman=0	0.066	0.052	0.602	0.080	0.062	0.056	0.077	0.269	0.188
Effect MinW for Men (%)	5.2%	11.9%	12.2%	28.1%	12.2%	12.0%	7.5%	10.4%	12.4%
Effect MinW for Women (%)	5.4%	4.4%	1.5%	9.7%	4.3%	4.4%	3.2%	2.9%	3.8%

Notes: All regressions include store*department fixed effects, worker fixed effects, pair-month fixed effects and control for MinW*department. Standard errors are two-way clustered at the state and border-segment level. *** p<0.01, **p<0.05, * p<0.1. In column (1), "discounted pay per hour (contemporaneous; not synthetic)" computes the welfare formula with the contemporaneous pay [w(e_i,t)] rather than the synthetic pay [w(e_i,t-1)]. In columns (2)-(9), the dependent variable is the discounted synthetic pay per hour using the synthetic pay [w(e_i,t-1)]. How we measure the discount factor, [(1+r)/(1+r-π)], is explained in the column headings. In columns (6)-(8), we allow the outside option to vary with the minimum wage. In column (7) [resp., (8)], the outside option is measured with the mean pay for women (men) in department stores [resp., across all industries] in the county in which the female (male) employee works. In column (9), it is measured with the average monthly pay expected within 5 years of an hypothetical termination from our firm, assuming that the worker earns an income of zero while unemployed and the average monthly pay in the county once employed. Due to the lack of data, unemployment duration is measured using yearly average unemployment duration by gender. "MinW" is the predominant minimum wage in deviation from its sample mean (in $). "Effect MinW for Men (%)" [resp., "Effect MinW for Women (%)"] is the percent effect of a $1 increase in MinW relative to the mean of the outcome variable for men [resp., women].

Table A.15: Impact of the Minimum Wage on Pay, Productivity, and Retention by Gender (Non-Ceteris-Paribus Analysis)

	(1)	(2)	(3)	(4)
	Pay		**Productivity**	**Retention**
Dep.Var.	Regular pay per hour (fixed + variable)	MinW top-up per hour	Sales per hour	Retained
MinW	-0.081	0.278***	-0.061	0.000
	(0.152)	(0.050)	(0.044)	(0.006)
MinW * Woman	0.352**	0.005	0.129**	0.005**
	(0.132)	(0.061)	(0.050)	(0.002)
Observations	215,565	215,565	217,746	217,746
Mean Dep.Var.	12.05	0.225	2.085	0.954
p-value for H_0: MinW+MinW*Woman=0	0.048	<0.001	0.209	0.399
Effect MinW for Men (%)	-0.7%	158.9%	-2.6%	0.0%
Effect MinW for Women (%)	2.3%	88.8%	4.1%	0.5%

Notes: All regressions include store fixed effects, pair-month fixed effects and control for the uninteracted woman dummy. Standard errors are two-way clustered at the state and border-segment level. *** p<0.01, **p<0.05, * p<0.1."Regular pay per hour" is the total amount earned from the base hourly rate and variable pay (commission rate*sales per hour), without the top-up. "MinW top-up per hour" is the monthly total minimum wage adjustment paid by the company to the worker (in $ per hour). "Sales per hour" are the sales per hour rescaled by a factor between 1/50 and 1/150 relative to its $ value. "Retained" is a dummy variable that equals one if the worker is retained that month (i.e., not terminated). "MinW" is the predominant minimum wage in deviation from its sample mean (in $). "Effect MinW for Men (%)" [resp., "Effect MinW for Women (%)"] is the percent effect of a $1 increase in MinW relative to the mean of the outcome variable for men [resp., women].

B Data and Identification Appendix

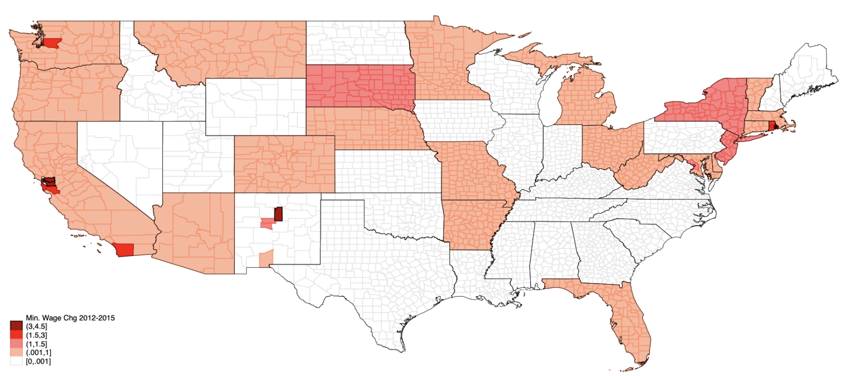
B.1 Minimum wage data

Our data contain information on the geographical location of stores (latitude and longitude), which we match with the monthly statutory minimum wage level in that store, extracted from the public dataset maintained by the Washington Center for Equitable Growth. Variations in minimum wage take place at state, county, and city levels; with city and county minimum wages always set to be higher than the state minimum wage.

From February 2012 to June 2015, our sample of stores is affected by 70 variations in minimum wage: 49 variations are at the state level, and 21 are at the county or city level. The exact timing of each minimum wage change is reported in Table B.1 and presented visually in Figure B.1.

Of all the variations in minimum wage present in our sample, the three changes in Florida coincide with state-level variations in unemployment insurance potential benefits duration (see Lusher et al. (2022) Online Appendix 2, page 6). State-level changes in unemployment insurance potential benefits duration in Arkansas, Illinois, Michigan, Georgia, North and South Carolina, and Missouri occurred either before our sample period or sufficiently distant in time from the minimum wage changes employed in our research design.

Figure B.1: Variations in Minimum Wage from February 2012 to June 2015



Notes: Store locations are withheld for confidentiality reasons.

Table B.1: Changes in Minimum Wages from February 2012 and June 2015

State	State	Date C.1	W_{t-1}	W_t	Date C.2	W_{t-1}	W_t	Date C.3	W_{t-1}	W_t	Date C.4	W_{t-1}	W_t
Alaska	AK	2015m2	7.75	8.75									
Arkansas	AR	2015m1	7.25	7.5									
Arizona	AZ	2013m1	7.65	7.8	2014m1	7.8	7.9	2015m1	7.9	8.05			
California	CA	2014m7	8	9									
Colorado	CO	2013m1	7.64	7.78	2014m1	7.78	8	2015m1	8	8.23			
Connecticut	CT	2014m1	8.25	8.7	2015m1	8.7	9.15						
DC	DC	2014m7	8.25	9.5									
Delaware	DE	2014m6	7.25	7.75	2015m6	7.75	8.25						
Florida	FL	2013m1	7.67	7.79	2014m1	7.79	7.93	2015m1	7.93	8.05			
Hawaii	HI	2015m1	7.25	7.75									
Massachusetts	MA	2015m1	8	9									
Maryland	MD	2015m1	7.25	8									
Michigan	MI	2014m9	7.4	8.15									
Minnesota	MN	2014m8	7.25	8									
Missouri	MO	2013m1	7.25	7.35	2014m1	7.35	7.5	2015m1	7.5	7.65			
Montana	MT	2013m1	7.65	7.8	2014m1	7.8	7.9	2015m1	7.9	8.05			
Nebraska	NE	2015m1	7.25	8									
New Jersey	NJ	2014m1	7.25	8.25	2015m1	8.25	8.38						
New York	NY	2013m12	7.25	8	2014m12	8	8.75						
Ohio	OH	2013m1	7.7	7.85	2014m1	7.85	7.95	2015m1	7.95	8.1			
Oregon	OR	2013m1	8.8	8.95	2014m1	8.95	9.1	2015m1	9.1	9.25			
Rhode Island	RI	2013m1	7.4	7.75	2014m1	7.75	8	2015m1	8	9			
South Dakota	SD	2015m1	7.25	8.5									
Vermont	VT	2014m1	8.6	8.73	2015m1	8.73	9.15						
Washington	WA	2013m1	9.04	9.19	2014m1	9.19	9.32	2015m1	9.32	9.47			
West Virginia	WV	2015m1	7.25	8									
County	State	Date C.1	W_{t-1}	W_t	Date C.2	W_{t-1}	W_t	Date C.3	W_{t-1}	W_t			
Bernalillo	NM	2013m7	7.5	8	2014m1	8	8.5	2015m1	8.5	8.65			
Johnson	IA	2015m11	7.25	8.2									
Montgomery	MD	2014m10	7.25	8.4									
Prince George's	MD	2014m10	7.25	8.4									
Santa Fe	NM	2014m4	7.5	10.66	2015m3	10.66	10.84						
City	State	Date C.1	W_{t-1}	W_t	Date C.2	W_{t-1}	W_t	Date C.3	W_{t-1}	W_t	Date C.4	W_{t-1}	W_t
Alburquerque	NM	2013m1	7.5	8.5	2014m1	8.5	8.6	2015m1	8.6	8.75			
Berkeley	CA	2014m10	9	10									
Las Cruces	NM	2015m1	7.5	8.4									
Oakland	CA	2015m3	9	12.25	2016m1	12.25	12.55						
Richmond	CA	2015m1	9	9.6	2016m1	9.6	11.52						
San Diego	CA	2015m1	9	9.75									
San Francisco	CA	2013m1	10.24	10.55	2014m1	10.55	10.74	2015m1	10.74	11.05	2015m5	11.05	12.25
San Jose	CA	2013m3	8	10	2014m1	10	10.15	2015m1	10.15	10.3			
Santa Fe	NM	2012m3	9.5	10.29	2013m3	10.29	10.51	2014m3	10.51	10.66	2015m3	10.66	10.84
SeaTac	WA	2013m1	9.04	9.19	2014m1	9.19	15						
Seattle	WA	2013m1	9.04	9.19	2014m1	9.19	9.32	2015m1	9.32	9.47	2015m4	9.47	11
Sunnyvale	CA	2015m1	9	10.3									
Tacoma	WA	2013m1	9.04	9.19	2014m1	9.19	9.32	2015m1	9.32	9.47			
Washington	DC	2014m7	8.25	9.5									

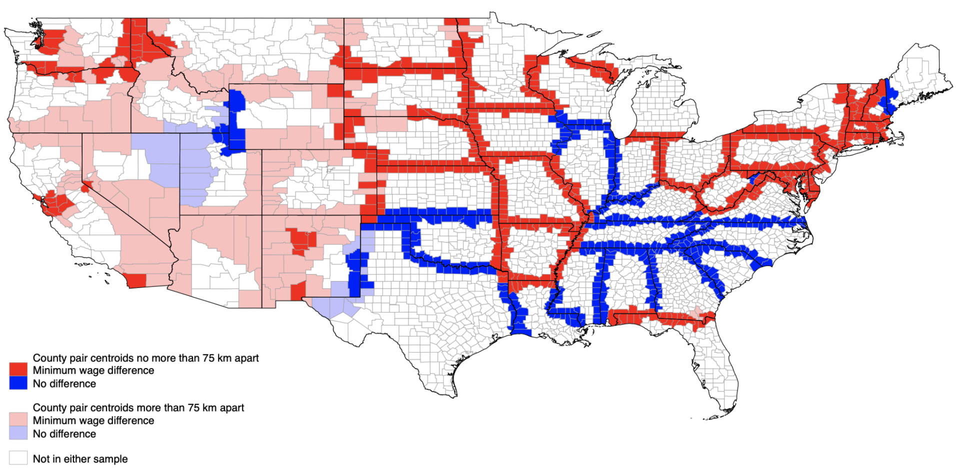
Notes: This table reports all state/county/city variations in statutory minimum wage from 2/1/2012 to 6/30/2015, irrespective of whether there is a store located in that state/county/city. The data are extracted from the public dataset maintained by the Washington Center for Equitable Growth. Our identification strategy effectively leverages only a sub-sample of these changes (70 out of 89), i.e., those that affect at least one store in our sample. We do not report which ones are the 70 variations we leveraged in the paper for confidentiality reasons. W_t (W_{t-1}) refers to the minimum wage level after (before) the change. The states with no change in minimum wage from February 2012 and June 2015 are: AL, GA, IA, ID, IL, IN, KS, KY, LA, ME, MS, NC, ND, NH, NM, NV, OK, PA, SC, TN, TX, UT, VA, WI, WY.

B.2 Border discontinuity design

We use a border discontinuity design, as implemented in Card & Krueger (2000), Dube et al. (2010, 2016) and Allegretto et al. (2011, 2017). This approach exploits minimum wage policy discontinuities at the state- or county-border by comparing workers on one side of the border where the minimum wage increased (treatment group) to workers on the other side where the minimum wage did not increase (control group). As shown in Dube et al. (2010), this research design has desirable properties for identifying minimum wage effects since workers on either side of the border are more likely to face similar economic conditions and are likely to experience similar shocks at the same time. The main disadvantage of this design is the risk of cross-border worker movements from control to treated stores (Neumark et al., 2014). We alleviate this concern in Section 3.3.

Following Card & Krueger (2000), Dube et al. (2010, 2016) and Allegretto et al. (2017), we restrict our sample to stores (and their respective workers) located in adjacent counties that share a border. For *state*-level minimum wage variations, we keep stores located in county pairs that: share a *state* border, and whose centroids are within 75 km of each other: see Figure B.2.

Figure B.2: Variations in the Minimum Wage in Bordering Counties



Notes: Store locations are withheld for confidentiality reasons.

For *county*-level minimum wage variations, we "seed" the sample with stores located in those counties that increased their minimum wage, and then add as controls all adjacent counties whose centroids are within 75 km of the seed county. Minimum wage changes at the *city* level are attributed only to stores within the city limits, but not to stores in the county containing that city. Such stores are included as controls, as are stores in all neighboring counties. (In our sample there are no municipalities that lie in more than one county). For instance, for the city of San Francisco (which increased its minimum wage) we include all the counties that share a county-border with San Francisco County and whose centroids are within 75 km of its centroid (i.e., the counties of Marin, Alameda, and San Mateo).

C Theory Appendix

The material in the model is a streamlined version of the model in Coviello et al. (2022), except for Proposition 1 and formulas (9) and (10) which are new. The main simplification here is that we abstract from heterogeneity in worker ability. This simplification is appropriate because, in this paper, we are interested in the average effect of the minimum wage across workers of all abilities.

C.1 Modeling details and proof of Proposition 1

The function $c(e)$ is strictly increasing in e. We assume that the marginal cost of effort vanishes at zero and is infinite at 1; these assumptions help ensure that optimal effort is interior to $[0, 1]$. Worker performance (in our case, sales per hour) is a non-negative random variable $Y(e)$ that is uniformly bounded from above across all e. Its density $f_Y(y; e)$ has interval support, is twice continuously differentiable in both its arguments, and enjoys the strict monotone likelihood ratio property (MLRP) in e.[57]

Assumption 1 (concavity). $u_{ee} < 0$ *and* $\pi_{ee} \leq 0$.

Under Assumption 1, the worker's optimal effort $e^*(M)$ is the unique solution to the first-order condition:

$$u_e(e; M) + \frac{1}{(1+r)}\pi'(e)V(M) = 0. \tag{15}$$

[57] This means that the ratio $f_Y(y; e)/F_Y(y; e)$ is strictly increasing in e whenever $f > 0$.

Proof of Proposition 1

Proof. Fix M. The function $u(e;M)$ shifts down if $u^U(M)$ increases, and thus also if $V^U(M)$ increases. Coviello et al (2022, Lemma 3 part 2 in Online Appendix B) shows that if the function $u(e;M)$ shifts down, the worker's net value $V(M)$ decreases. Therefore, as $u^U(M)$ increases, both functions of e in (15) shift down, hence $e^*(M)$ decreases. ■

C.2 Computing expression (9)

Differentiating (8) with respect to M yields:

$$\frac{dV(M)}{dM} = (1+r)\left[\frac{u_e(e^*;M)[1+r-\pi(e^*)] + u(e^*;M)\cdot\pi'(e^*)}{[1+r-\pi(e^*)]^2}\cdot\frac{de^*}{dM} + \frac{u_M(e^*;M)}{1+r-\pi(e^*)}\right].$$

This formula simplifies because the numerator of the first fraction inside the brackets is zero. Indeed, substituting (8) into the first-order conditions yields:

$$u_e(e^*;M) + \frac{\pi'(e^*)}{1+r-\pi(e^*)}\cdot u(e^*;M) = 0.$$

Therefore equation (9) holds.

C.3 Computing expression (10)

We have:

$$\begin{aligned}
\frac{dV^E(M)}{dM} &= \frac{dV(M)}{dM} + \frac{dV^U(M)}{dM} \\
&= \left[\frac{(1+r)}{1+r-\pi(e^*)}\right]u_M(e^*;M) + \left[\frac{(1+r)}{r}\right]u_M^U(M) \\
&= \left[\frac{(1+r)}{1+r-\pi(e^*)}\right]\left[w_M(e^*;M) - u_M^U(M)\right] + \left[\frac{(1+r)}{r}\right]u_M^U(M) \\
&= \left[\frac{(1+r)}{1+r-\pi(e^*)}\right]w_M(e^*;M) + \left[\frac{(1+r)}{r} - \frac{(1+r)}{1+r-\pi(e^*)}\right]u_M^U(M),
\end{aligned}$$

where the second line used the definition

$$u^{U}(M)=[r/(1+r)]V^{U}(M).$$

Substituting:

$$\begin{aligned}
& \frac{(1+r)}{r}-\frac{(1+r)}{1+r-\pi\left(e^{*}\right)} \\
= & \frac{(1+r)\left[1+r-\pi\left(e^{*}\right)\right]-r(1+r)}{r\left[1+r-\pi\left(e^{*}\right)\right]} \\
= & \frac{(1+r)\left[1-\pi\left(e^{*}\right)\right]}{r\left[1+r-\pi\left(e^{*}\right)\right]},
\end{aligned}$$

we get:

$$\begin{aligned}
\frac{dV^{E}(M)}{dM} & = \left[\frac{(1+r)}{1+r-\pi\left(e^{*}\right)}\right] w_{M}\left(e^{*};M\right)+\frac{(1+r)\left[1-\pi\left(e^{*}\right)\right]}{r\left[1+r-\pi\left(e^{*}\right)\right]} u_{M}^{U}(M) \\
& = \left[\frac{(1+r)}{1+r-\pi\left(e^{*}\right)}\right]\left[w_{M}\left(e^{*};M\right)+\frac{\left[1-\pi\left(e^{*}\right)\right]}{r} u_{M}^{U}(M)\right].
\end{aligned}$$

D Results Appendix: Robustness of the Welfare Effects of the Minimum wage by Gender

This section assesses the sensitivity of our welfare results to using different calculations of π, values for r, and assumptions on the outside option. The results are presented in Table A.14.

Columns 2-4 show that the results are not sensitive to how π is calculated. The results are similar if we calculate $\pi\left(e^{*}\right)$ across all time periods (rather than month by month), by gender (column 2). They are also similar if we calculate $\pi_{d}\left(e_{t-1}^{*}\right)$ each month, and aggregated at the gender$\times$department level (column 3) or at the gender$\times$department$\times$store level (column 4).

Columns 5-6 show that the results are very similar if we use a monthly discount rate

r of 0.41% or 1.5%. The former corresponds to an annual rate of 5% and the latter to a quarterly $\beta = 1/1 + r$ of 0.96, commonly used in macroeconomics calibrations.

Columns 7-9 present the sensitivity of our results to different assumptions on the outside option $u_M^U(M_{jt})$. Recall that in our main welfare calculations we set $u_M^U(M_{jt})$ to zero. Results remain very similar if we remain agnostic about $u_M^U(M_{jt})$ and run specification (1) with the following outcome variable:

$$\left[\frac{(1+r)}{1+r-\pi\left(e^{*}\right)}\right]\left[w_M\left(e_{i,t-1}^{*}; M_{jt}\right)+\frac{\left[1-\pi\left(e^{*}\right)\right]}{r}u^U\left(M_{jt}\right)\right], \tag{16}$$

where u^U is measured with the county-level average monthly wages in department stores (or alternatively, across all industries) by gender, or with a similar value that takes into account the fact that terminated workers go through a period of unemployment in which we assume their pay to be zero.