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MANAGEMENT SCIENCE
Articles in Advance, pp. 1–23
ISSN 0025-1909 (print), ISSN 1526-5501 (online)

Does a One-Size-Fits-All Minimum Wage Cause Financial Stress for Small Businesses?

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Received: July 8, 2020
Revised: April 22, 2021; October 15, 2021; January 17, 2022; March 22, 2022
Accepted: April 17, 2022
Published Online in Articles in Advance: January 6, 2023

https://doi.org/10.1287/mnsc.2022.4620

Abstract. Using intertemporal variation in the bounding of a state's minimum wage by the federal rate and business credit-score data for 15.2 million establishments, we find that the increase in labor costs caused by a higher federal minimum wage leads to lower business credit scores and worsens the financial health of small businesses in the affected states. In particular, small, young, labor-intensive, and minimum-wage-sensitive establishments located in affected states and those located in competitive and low-income areas experience higher financial stress. Increases in the minimum wage are associated with employment reductions and a higher exit rate for small businesses. Our results document some potential costs of a one-size-fits-all nationwide minimum wage for some small businesses.

History: Accepted by Gustavo Manso, finance.
Funding: This work was supported by the Ewing Marion Kauffman Foundation Junior Faculty Fellowship.
Supplemental Material: The data and internet appendix are available at https://doi.org/10.1287/mnsc.2022.4620.

Keywords: small business • financial distress • minimum wage • credit score

1. Introduction

Substantial debate exists among academics and policymakers on the costs and benefits of mandated minimum wages (Belman and Wolfson 2014, 2019). However, evidence on the impact of an increase in the minimum wage on the financial health of small businesses in the United States is limited (Drucker et al. 2021). In this paper, we examine the effects of one-size-fits-all federal minimum wage increases on the financial health of small businesses. We use a new measure of a small business's financial health based on its payment performance to suppliers and vendors, namely, its establishment-level credit score (PAYDEX score), and document that higher minimum wages negatively impact the financial health of small businesses. Our paper contributes to the ongoing policy debate in the United States on the impact of one-size-fits-all federal minimum wage increases on small businesses.

We focus on the financial health of small businesses, as they are a vital component of the U.S. economy and account for almost 50% of the nonfarm gross domestic product. The opening and closing of small businesses—that is, those with fewer than 10 employees—accounted for more than 70% of job gains and losses in 2018 (Bureau of Labor Statistics). For credit-rationed small businesses, trade credit is an essential source of financing (Petersen and Rajan 1997). Delaying payments to vendors is one method by which small businesses absorb financial shocks.[1] However, delayed payments to vendors can adversely affect the credit scores of small businesses, their future payment terms, and their credit availability, thus increasing their probability of default (Petersen and Rajan 1994, Kallberg and Udell 2003, Barrot and Nanda 2020).

Wages comprise a significant fraction of the operational costs of many small businesses. An increase in labor costs due to an increase in the minimum wage may not cause financial stress to a small business if it has the flexibility to immediately adjust its capital-to-labor ratio or pass on the increased costs to its customers. Alternately, the firm can maintain profit margins by reducing other costs or by increasing productivity. However, a business's inability to take these steps may financially stress the firm by adversely affecting profit margins, lowering free cash flows, delaying payments to vendors, and lowering credit scores.[2]

In this paper, we study the impact of one-size-fits-all federal minimum wage increases on the financial health of small establishments located in states where the effective minimum wage is equal to the federal rate (the *bounded* states) relative to those in states where wage rates are higher than the federal rate (the *unbounded* states). Further, we study how firm-level, industry-level,

and local economic conditions that may impact the ability of these businesses to pass on the increased labor costs to consumers moderate or amplify these wage increases.

We focus on federal minimum wage changes for two reasons. First, although labor market conditions (e.g., availability, productivity, the bargaining power of workers) and economic conditions (e.g., local product market competition) vary substantially with geography, 14 states in the United States have a minimum wage rate equal to the federal rate.[3] Second, federal minimum wage changes are less likely to be affected by the state-level or county-level local economic conditions and are plausibly exogenous to the financial health of small businesses. Finally, we also use matching methods and fuzzy triple differences to further isolate potentially confounding factors and show robustness to the main effect of minimum wage changes on the financial health of small businesses that we identify.

We use intertemporal variation in whether a state's minimum wage is bound by the federal minimum wage and the Dun & Bradstreet (D&B) PAYDEX credit score data for approximately 15.2 million establishments for 1989–2013. The PAYDEX score is a dollar-weighted numerical indicator of how a firm paid its bills based on trade experiences reported to D&B through its 4,000 trade exchange participants in the United States. Lenders, insurance providers, and lessors use credit scores for setting terms on credit, loans, lease payments, and insurance premiums. We find that a one-dollar increase in the federal minimum wage corresponds to an almost 1.0-point reduction in an establishment's PAYDEX credit score in a *bounded* state compared with a similar establishment in an *unbounded* state. In terms of days payable outstanding, the 1.0-point reduction implies a delay of 1.4 days beyond the due date (i.e., a 10% increase compared with the sample mean of 14.4 days). A back-of-the-envelope calculation suggests that, for the average firm in our sample, a one-dollar increase in the minimum wage (without a price pass-through) can reduce its profit margin by up to 21.9%.[4]

We find that even a 1.0-point decline in credit score can significantly affect a business's likelihood of survival. We find significant threshold effects wherein an establishment's 1.0-point decline in credit score from 80 points (i.e., payment within terms, perfect score) to 79 points (i.e., payment two days beyond terms) arising from an increase in the federal minimum wage corresponds to an increase in the exit probability by 2.2 percentage points (pp) or a 25% increase from the 8.5% unconditional annual exit probability. We find similar effects for other thresholds. Further, independent of the credit score, we also find an increase in exit probability for the average small business by 10% after an increase in the federal minimum wage.[5]

Since two of the three federal minimum wage increases were enacted during recession years (i.e., 1990–1991 and 2007–2009), one potential concern is whether the results can instead be attributed to business cycles at the national or state level.

More broadly, we may not be able to identify our effect if the federal government's decision to adjust minimum wages is affected by, or correlated with, some other observable and unobservable differences in the economies of bounded versus unbounded states.

However, mitigating the above concerns, we find that the bounded and unbounded states followed similar business cycles before and after the federal minimum wage increases. We control for various state-, county-, and ZIP-code-level observable characteristics in the regressions and also match on these dimensions to identify the right control group. Further, we follow Dube et al. (2010) and control for local economic conditions by analyzing establishments located in contiguous state-border counties. In our estimation, we include county-pair × year fixed effects, and sometimes county-pair × NAICS4 × year fixed effects, to control for time-varying industry and local-area-specific unobservables. We find consistent results for establishments in the bordering counties of the bounded states.

Under FLSA enterprise coverage, employer businesses with annual sales of $500,000 or less are exempt from mandated federal minimum wages. We implement fuzzy triple differences to identify the control group around the $500,000 threshold within the bounded states and include state × year fixed effects to control for time-varying state-specific unobservables. We find that the average PAYDEX score declines by almost 2.20 points in the year of minimum wage increase for the nonexempted firms located within the bounded states, compared with those located in states not bounded by federal minimum wage.

Note that the affected establishments may not experience any financial stress if they can immediately pass on these increased wage costs entirely to their customers. On the other hand, any constraints faced by small businesses in passing on their wage costs may impact their financial health. We find that establishments within the same industry, those located in more competitive counties, and those located in low-income ZIP codes experience a more significant decrease in their credit score. We also find that small and young establishments, which are more likely to have financial constraints, experience a more significant decrease in their credit scores. Similarly, we find that this negative impact is more pronounced in industries that employ a higher number of minimum wage workers (e.g., restaurants, retail), but it is not limited to these industries.[6]

Finally, our county-industry-level analysis shows an increase in the exit rate and a reduction in the entry rate for counties in the bounded states after a federally mandated minimum wage increase. We find that aggregate employment declines significantly more for restaurants

(9.5%) and for retail businesses (8.2%) in bounded states, with the negative effect more prominent in counties with lower personal income.

To the best of our knowledge, our paper is the first to document the impact of federal minimum wage changes on trade credit payments of small businesses using a large sample of small businesses in the United States. Further, we are the first to document the possible spillover effect of an increase in federal minimum wage on exempted very small businesses (i.e., those with less than $500,000 in sales). Our evidence complements results in Drucker et al. (2021) that studies the income redistribution effect and the impact of higher minimum wage on Israeli small businesses. Other studies that document the impact of minimum wage changes on firm profitability (Draca et al. 2011, Bell and Machin 2018, Mayneris et al. 2018, Harasztosi and Lindner 2019, Hau et al. 2020), utilize minimum wage law changes in the UK, China, Hungary, or other non-U.S. countries. These economies have very different labor laws and economic policies to support small businesses compared with the United States. Our study is also related to that of Luca and Luca (2019), which documents an increase in price and exit rate for restaurants after a minimum wage increase. We contribute to the voluminous literature on the effect of minimum wages on employment.[7] Finally, we also contribute to the literature on trade credit (Petersen and Rajan 1997, Breza and Liberman 2017, Barrot and Nanda 2020) and document the implications of minimum wage changes on supply chain relationships. More broadly, our paper also adds to the literature that analyzes the interactions between labor costs and firm outcomes.[8] The rest of the paper proceeds as follows. We discuss our empirical methodology and identification concerns in Section 2. Section 3 describes our data and summary statistics. Our main empirical results are presented in Section 4, and we conclude in Section 5.

2. Identification Challenges and Empirical Strategy

In this section, we discuss threats to identifying the impact of minimum wage changes on the health of small businesses and our empirical strategy to ameliorate these concerns.

Although the FLSA mandates broad minimum wage coverage, states are able to set separate minimum wage rates that differ from those mandated by the federal government. Under the provisions of the FLSA, employers must pay workers the highest minimum wage as prescribed by either federal, state, or local law. Adjusting state minimum wage rates is typically done in one of two ways: (a) through legislatively scheduled rate increases that may include one or more increments, or (b) by using a measure of inflation to index the value of the minimum wage to the general change in prices.

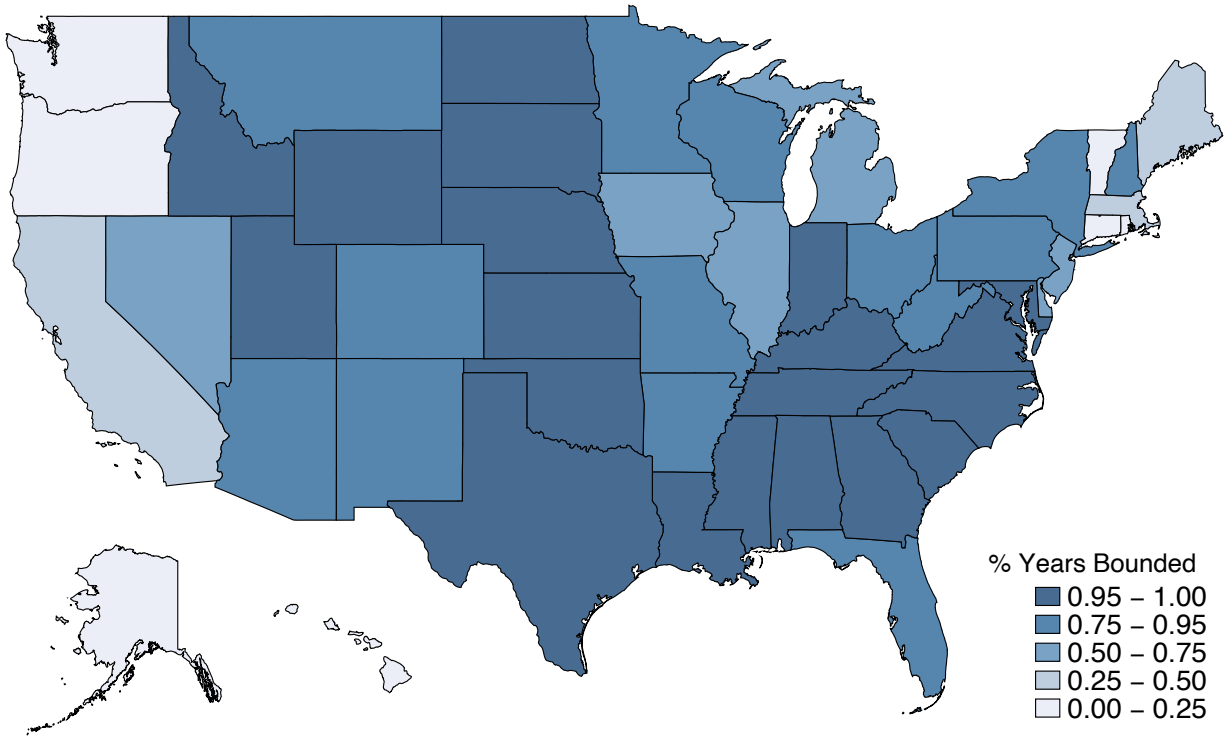
A natural starting point for empirical examination of the impact of minimum wages is to exploit the staggered state-level minimum wage changes in a difference-in-differences setup. However, estimates in this framework are likely to be biased, as the introduction of state-level minimum wage increases are likely to occur at nonrandom times and may be correlated with local economic conditions. For example, Allegretto et al. (2017) show that states that increase minimum wages have different business cycle severity, increased income inequality, and differing composition of their labor force.

However, not all states voluntarily increase their minimum wages. After the introduction of higher federal minimum wage requirements,[9] states with effective minimum wages below the federal minimum are bound and must immediately match the federal minimum wage. For this study, we refer to states with minimum wage rates that are higher than the federal rate as *unbounded* (i.e., not bounded to the federal minimum wage rate), and we refer to states with effective minimum wage rates equal to the federal rate as *bounded* (i.e., bounded to the federal rate).

We utilize this bounding feature to examine the differential effect of federal minimum wage increases on the financial health of establishments located in bounded states versus unbounded states. During the past three decades, federal minimum wage increased in 1990–1991, 1996–1997, and 2007–2009.[10] During those same periods, there have been numerous changes in state minimum wage policies. In 1990, at the beginning of our sample, the federal minimum wage was $3.80 per hour. In Figure 1, we graphically show the percentage of years that each state in our sample was bounded by the federal minimum wage. Notice that the federal minimum wage is always bounded in states such as Alabama, Georgia, Texas, and many others. This means that employers in these states have always had their minimum wage rates defined by federal laws rather than state laws. Our strategy exploits the fact that an increase in the federal minimum wage rate affects states with minimum wage rates equal to or less than the federal minimum wage (i.e., bound states) more directly than states with higher minimum wages. As specified in the FLSA, certain small business within the bounded states are exempt from the minimum wage requirements. As such, even though these exempt businesses may not be required by the FLSA to increase wages, the law may still force them to do so through local labor market competition from firms affected by the law (Berger et al. 2022). Thus, although we are unsure about the extent to which these exempted businesses will respond to local labor market forces, their inclusion in our treatment group provides lower-bound estimates for the credit-oriented minimum wage effects that we explore in this paper.[11]

In any given year, the exact number of states with a minimum wage rate above the federal rate may vary

Figure 1. (Color online) Bounded States by Year



Notes. The map plots the percentage of years during 1989–2013 that a given state has an average minimum wage bounded by the federal minimum wage. The dark shade reflects states that are mostly bounded by the federal minimum wage. Source: Bureau of Labor Statistics.

depending on the interaction between the federal rate and the mechanisms in place to adjust the state minimum wage rate. Before 1987, Alaska and the District of Columbia were the only two states that consistently had minimum wage rates that exceeded the federal rate. Since 1987, many states have adopted higher minimum wage rates, resulting in a divergence between the average state minimum wage and the federal rate. Because the federal and state minimum wage rates change at different times and in different increments, the share of the labor force for which the federal rate is the binding wage floor has changed over time, with many states alternating between being bound and unbound over time.

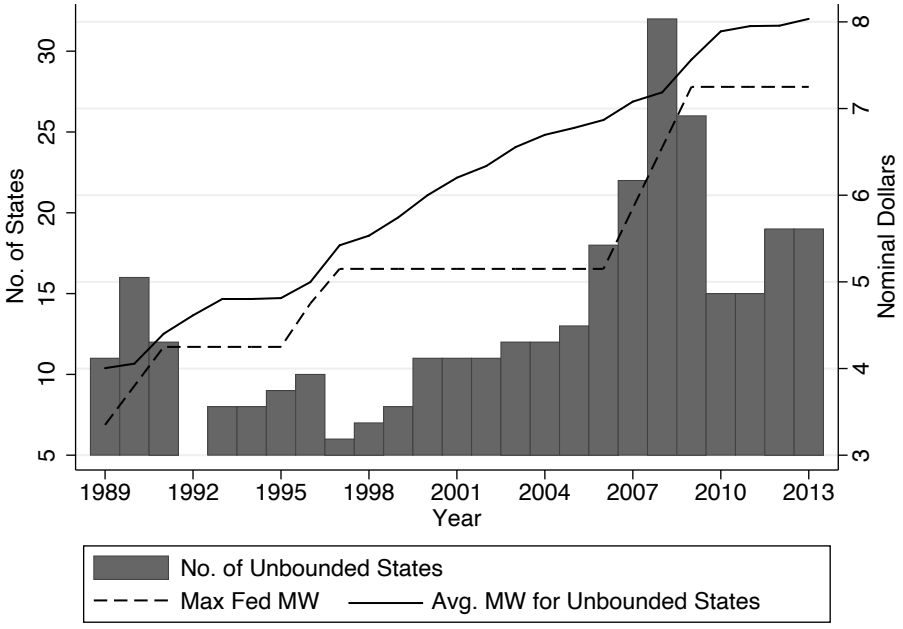
Figure 2 demonstrates this variation over time: The bars show for a given year the number of states with an average minimum wage above the average federal minimum wage. The dashed line plots the average federal minimum wage (in nominal dollars), and the solid line plots the average minimum wage in unbounded states. We estimate the following difference-in-differences equation to quantify the differential impact of the federal minimum wage changes on the financial health of establishments located in bounded states versus unbounded states:

$$Y_{it} = \alpha_1 Bound_{s,t-1} \times \Delta MW(F)_t + \alpha_2 Bound_{s,t-1} + \kappa X_{i,t-1} + \nu_i + \omega_t + \epsilon_{ist}, \quad (1)$$

where subscripts i, s, and t index establishments, states, and years, respectively. Our dependent variable, Y_{it}, is the average PAYDEX score, which is our measure for an establishment's financial health. The PAYDEX score is a business credit score generated by Dun & Bradstreet (D&B) that captures an establishment's payment performance (i.e., whether it pays its bills on time). Each establishment is assigned a numerical score from 1 to 100, with 100 signifying a perfect payment history. We explain this variable in more detail in our data section, Section 3. The expression $\Delta MW(F)_t$ measures the nominal dollar increase in the federal minimum wage in year t; it equals zero in years with no increases. The indicator $Bound_{s,t-1}$ is a dummy variable equal to 1 if, at the beginning of fiscal year t, the establishment's state s has a minimum wage less than or equal to the federal minimum wage. We include establishment fixed effects, ν_i, to control for time-stable unobserved heterogeneity at the establishment level, and we include year fixed effects, ω_t, to control for time-specific macro-level shocks. In addition, we include a full set of establishment-level control variables ($X_{i,t-1}$) in our regressions: size (measured as log(*Sales*)), age (log(*Age*)), the number of employees (log(*Employees*)), and sales growth. These variables are winsorized at their 1st and 99th percentiles.[12]

Since all of our identifying variation is within-establishment due to the inclusion of ν_i, we can interpret our main coefficient of interest, α_1, as the differential effect of a federally mandated minimum wage increase for bounded firms above and beyond the effects of

Figure 2. Minimum Wage and Unbounded States



Notes. The bar (left axis) shows by year the number of states with a minimum wage above the federal minimum wage (unbounded states) in each year between 1989 and 2013. The dashed and solid lines (right axis) plot, respectively, the average federal minimum wage per hour and the average minimum wage in unbounded states. Source: Bureau of Labor Statistics.

bounding on an establishment's performance (as captured by α_2). Our standard errors are clustered at the state level.

Our baseline specification can identify the effect of minimum wage increases to the extent that the PAYDEX scores of establishments in bounded and unbounded states are similar in level and trends. We conduct various tests to verify this assumption.[13] The decision by state governments to set their minimum wages at (or above) the federal level may not be random. States that increase their minimum wages tend to differ in their business-cycle severity, their economic inequality, and the composition of their labor force (Allegretto et al. 2017). We test for state-level variables that may affect a state's decision to keep minimum wages at the federal level. We find that states with large populations and states with Democratically controlled senates are more likely to keep state minimum wages above the federal minimum wage (see Table IA2 in the internet appendix).

Hence, in our regressions, we explicitly control for these state-level variables. In addition, counties and ZIP codes in unbounded and bounded states may differ in other economic conditions such as unemployment rate, per capita income, house prices, and aggregate demand. We present results that directly control for a battery of local economic conditions (see Section IA1.4 in the internet appendix). Further, we test the robustness of results if some bounded and unbounded states increase their state-level minimum wages in response to a federal minimum wage increase (Section 4.1.3). We use geographic border discontinuity design (Section 4.1.4) to control for time-varying unobservables characteristics. We also test if states adopt more generous policies to help businesses around the minimum wage changes (Figure IA4 in the internet appendix). Further, we implement fuzzy triple differences using the FLSA enterprise coverage rule to identify counterfactual within the bounded states and include state × year fixed effects (Section 4.1.5). These specifications utilize the variations in credit scores of establishments located within the same state and/or county and should mitigate some of these concerns.

In summary, conditional on a variety of approaches, we show that after an increase in the federal minimum wage, firms in unbounded states serve as an appropriate counterfactual to firms in bounded states, ameliorating concerns about our identification strategy.

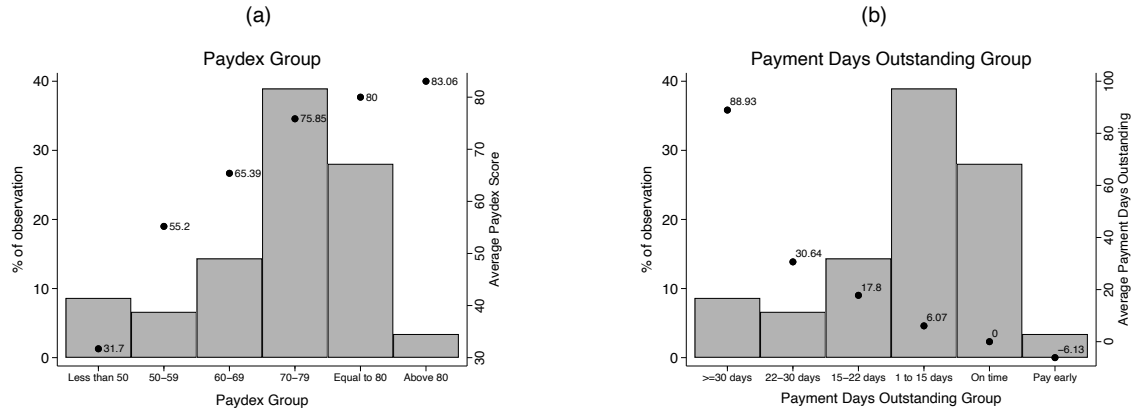
3. Data and Descriptive Statistics

3.1. Data Sources and Sample Selection

We use establishment-level data for all the establishments in the United States from the National Establishment Time-Series (NETS) Database.[14] This database provides an annual record for a large part of the U.S. economy that includes establishment-level employment counts, sales figures, establishment failures, market segment, corporate affiliations, and historical D&B credit and payment ratings.

This database covers almost 50 million U.S. businesses. Among these firms, 15 million firms have data on their PAYDEX scores over 25 years (1989–2013). We exclude establishments with only one employee (almost three

Figure 3. *PAYDEX Score* and *Days Payable Outstanding* Group Summary



Notes. Panel (a) plots the summary based on PAYDEX score. Panel (b) utilizes the one-to-one mapping of *PAYDEX Score* to *Days Payable Outstanding* provided by Dun & Bradstreet. The bars in the figure plot the percentage of observations (left axis) in each group, and the dots represent the mean *PAYDEX Score* or *Days Payable Outstanding* (right axis) in each group.

million businesses). From the remaining 12.79 million establishments, we further remove non-stand-alone businesses (i.e., we omit 900,000 establishments affiliated with large firms). In addition, we exclude 3.8 million establishments in finance, real estate, utilities, and professional services that are less likely to employ minimum wage workers.[15] Finally, to construct our measure of lagged sales growth (which is one of our control variables in our baseline specification), we need at least three observations, and, therefore, we lose an additional three million establishments. However, our results are robust if we include these businesses in our sample (see Section 4.1.3). As such, our final sample consists of 4.4 million small businesses that survived for three years or more.[16]

3.2. Summary Statistics

We next provide the summary statistics of our data set. We first describe our primary variable of interest (i.e., PAYDEX score) and how it relates to various firm characteristics. Then, we provide summary statistics on state and federal minimum wage changes.

3.2.1. PAYDEX Score. The PAYDEX score is a business credit score calculated as a dollar-weighted numerical indicator of how a firm paid its bills based on trade experiences reported to D&B through its trade exchange program. D&B acquires its trade data from over 12,000 trade exchange participants globally in 35 markets, including 4,200 in the United States. The PAYDEX score compares payments to terms of sale. It is dollar-weighted and calculated based on the overall manner of payments reported to D&B. The score rates the likelihood that a business will make payments to suppliers or vendors on time. Like a personal credit score, it is primarily used to measure the financial risk to lenders, and it can affect the premiums and interest rates that companies pay for bank loans or credit cards.

In addition to lenders, the PAYDEX score is used by vendors, who often deliver goods and services and invoice a business for payment afterward. As a result of this process, vendors have some financial risk of not being paid. The PAYDEX score is one metric that such suppliers can use to determine whether a new client or business partner might present possible risks going forward. Poor scores may make suppliers reluctant to do business or may limit the size and scope of the services to which they are willing to agree.

Many businesses try to make payments on time or early to maintain future access to credit. D&B provides a one-to-one mapping of credit score with average payment days and categorizes these data into six groups based on the payment behavior. Panel (a) of Figure 3 plots the summary based on *PAYDEX Score*, and panel (b) plots the summary based on *Days Payable Outstanding* compared with the payment terms. The bars in the figure plot the percentage of observations (left axis) in each group, and the dots represent the mean *PAYDEX Score* or *Days Payable Outstanding* (right axis) in each group.[17]

A PAYDEX score of 80 means that a business makes its payments on time, above 80 means in advance, and a perfect score of 100 implies that a business makes payments one month in advance of when they are due. The NETS data are available at a yearly frequency. Whereas the PAYDEX score is calculated on a rolling 12-month basis, the NETS data only report the maximum and minimum PAYDEX score during the reported year. We compute the mean of the minimum and maximum PAYDEX score over the year as the *Average PAYDEX Score* and use the one-to-one mapping to create *Average Days Payable Outstanding*. In our sample, only 3.4% of establishments

make early payments, averaging 6.13 days before the due date. Moreover, 28% of businesses make payments on time, and about 39% of businesses make payments within six days of the due date. In our sample, the median of *Average PAYDEX Score* is about 76.5, which implies that the median business makes payments five days after the typical 30-day term.

3.2.2. Establishment Characteristics. Table 1, panel A, provides the summary statistics of our establishment sample. From our 4.4 million small businesses with a PAYDEX score, we obtain just over 31 million establishment-year observations. Whereas the PAYDEX score is available for only 42.9% of our total observations, we report information on the approximately 41 million establishment-year observations of firms that do not have PAYDEX scores. We do not utilize these data, but we present them here for comparison. Based on observable characteristics, establishments with PAYDEX scores have lower exit rates, higher sales, and more employees. In addition, these establishments are older and more labor intensive (with more employees per million in sales), and they compete in more concentrated industries as measured by a higher Herfindahl-Hirschman Index (HHI) (defined at the five-digit NAICS level over establishment sales).

Table 1. Summary Statistics

Panel A: Establishment sample

	All			With PAYDEX score			Without PAYDEX score		
	72,375,466			31,083,694			41,291,772		
	Median	Mean	Standard deviation	Median	Mean	Standard deviation	Median	Mean	Standard deviation
PAYDEX Score									
Minimum	74	67.35	17.13	74	67.35	17.13	—	—	—
Maximum	80	73.92	12.74	80	73.92	12.74	—	—	—
Average	76.5	70.63	13.12	76.5	70.63	13.12	—	—	—
Days Payable Outstanding									
Minimum	9	19.3	32.71	9	19.3	32.71	—	—	—
Maximum	0	9.46	24.4	0	9.46	24.4	—	—	—
Average	5.5	14.4	27.3	5.5	14.4	27.3	—	—	—
Exit	0	0.05	0.21	0	0.03	0.16	0	0.06	0.24
Sales ($ millions)	0.25	0.71	1.33	0.49	1.41	3.12	0.17	0.34	0.68
Number of employees	3	6.8	9.7	5	10.90	17.68	2	4.29	5.82
Age (in years)	12	17.7	17.0	17	22	18.77	9	14.5	14.7
Employee-to-sales	14.28	18.28	15.39	12.0	15.34	12.99	16.1	20.48	16.30
HHI Index	0.07	0.17	0.24	0.09	0.19	0.25	0.06	0.16	0.23

Panel B: Minimum wage

	N	Median	Mean	Standard deviation
All				
Average federal minimum wage ($ per hour)	1,275	5.15	5.25	1.13
Average state minimum wage ($ per hour)	1,275	5.15	5.50	1.29
$Bound_{t-1}$	1,275	1	0.74	0.44
$\Delta MWDummy(F)$	1,275	0.00	0.44	0.50
For $\Delta MWDummy(F) = 1$				
$\Delta MW(F)$ ($ per hour)	561	0.34	0.35	0.22
$\%\Delta MW(F)$	561	0.06	0.07	0.04
Bounded states				
Average federal minimum wage ($ per hour)	939	5.15	5.14	1.09
Average state minimum wage ($ per hour)	939	5.15	5.17	1.10
For $\Delta MWDummy(F) = 1$				
$\Delta MW(S)$($ per hour)	399	0.34	0.34	0.26
$\%\Delta MW(S)$	399	0.06	0.07	0.05
Unbounded states				
Average federal minimum wage ($ per hour)	336	5.15	5.55	1.18
Average state minimum wage ($ per hour)	336	6.75	6.42	1.35
For $\Delta MWDummy(F) = 1$				
$\Delta MW(S)$ ($ per hour)	162	0.15	0.24	0.26
$\%\Delta MW(S)$	162	0.03	0.04	0.05

Notes. This table reports summary statistics for our sample. Panel A reports the summary statistics of establishment data. Panel B reports summary statistics for federal and state minimum wages during 1989–2013.

3.2.3. Minimum Wage. Table 1, panel B, reports the summary statistics on federal and state minimum wages and their growth rates. We find that the average annual state minimum wage is about \$5.50 per hour, which is above the federal minimum wage (i.e., \$5.25 per hour). This is especially true for unbounded states. Note that, whenever the federal government decides to change the minimum wage, the average level of change or growth is much higher for bounded states than for unbounded states. For example, the median $\%\Delta MW(S)$ is about 6.0% for bounded states but only 3.0% for unbounded states.

4. Results

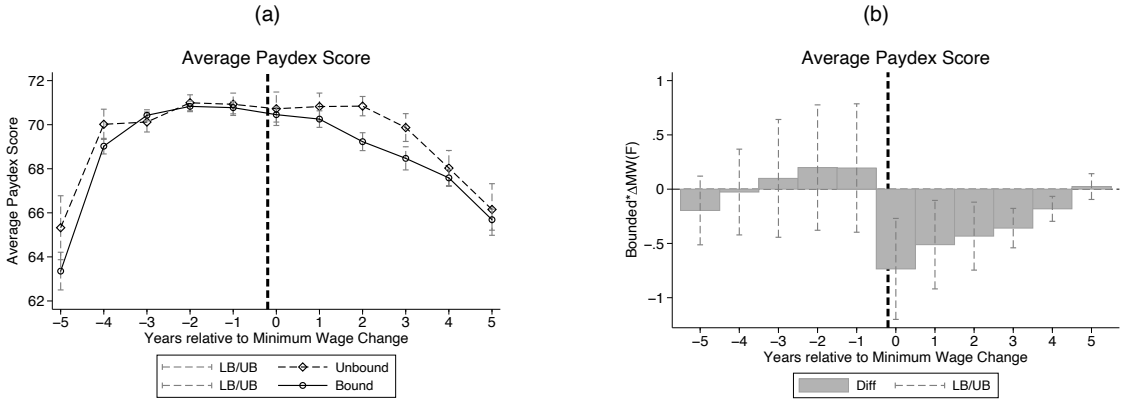
4.1. Do Increases in the Federal Minimum Wage Affect Small Business PAYDEX Scores?

In this subsection, we discuss our baseline PAYDEX results (Section 4.1.1) for Equation (1). We conduct tests for pre– and post–minimum wage change dynamics (Section 4.1.2). We show that our results are robust for variations in the baseline model (Section 4.1.3). We utilize bordering-county discontinuity tests (Section 4.1.4) and the FLSA exemption rule (Section 4.1.5) to further address endogeneity concerns.

4.1.1. Baseline Results. We begin our analysis by plotting the average PAYDEX score for establishments in bounded states and unbounded states around the years before and after federal minimum wage increases. Figure 4(a) plots the average score with a 95% confidence interval. The solid line with circled data points plots the average PAYDEX score for establishments located in bounded states, and the dashed line with diamond data points plots the average of the PAYDEX score for unbounded states. The bold dashed line indicates the period immediately before the federal minimum wage change. As can be seen, the average PAYDEX score for bounded and unbounded states followed parallel trends before the minimum wage enactment. Second, within two years of a federal minimum wage increase, there is a sharp decline in the PAYDEX score for establishments in bounded states. Finally, we observe that the difference between the PAYDEX score for establishments in the bounded and unbounded states converges after three to five years. One possible reason for this may be that stronger establishments survive and may be able to pass on the increased labor costs to their customers. We explore this issue in Section 4.1.2.

It should be noted that these results do not take into account firm-specific and time-specific unobserved heterogeneity that may lead to lower credit scores for establishments located in bounded states. To account for this potential unobserved heterogeneity, we estimate our difference-in-differences Equation (1). Note that the interaction term, α_1, as captured by $Bound_{s,t-1} \times \Delta MW(F)_t$, identifies the differential effect of federally mandated minimum wage increases over and above the effect of state-level variation caused by a change in the state-determined minimum wage and the changing status of the focal state from bound to unbound (or vice versa). As previously discussed in Section 2, the number of states that are bound by the federal minimum wage changes over time. In addition, we control for establishment fixed effects and year fixed effects to ensure that identification arises only from within-establishment variation after controlling for macroeconomic trends. We report these results in Table 2.

Figure 4. PAYDEX Score Dynamics



Notes. We test the dynamics of the differential effect of the federal minimum wage on the PAYDEX scores for establishments located in bounded vs. unbounded states. Panel (a) plots the average PAYDEX score with a 95% confidence interval. The solid line with circle plots the average PAYDEX score for establishments located in bounded states, and the dashed line with diamonds plots the average PAYDEX score for unbounded states. The bold dashed line indicates the period immediately before the federal minimum wage change. Panel (b) plots the regression coefficients of Equation (2) with a 95% confidence interval. The bars plot the regression coefficients of the interaction term identifying bounded states for five years before and after the federal minimum wage increase, while the dashed lines plot the 95% confidence intervals. Standard errors are clustered at the state level.

Table 2. Effect of One-Size-Fits-All Minimum Wage on Business Credit Scores and Days Payable Outstanding

Dependent variables	Panel A: *PAYDEX Score*					
	(1) Minimum	(2) Maximum	(3) Average	(4) Minimum	(5) Maximum	(6) Average
$Bound_{s,t-1} \times \Delta MW(F)_t$	−0.90***	−0.60**	−0.75***	−0.87***	−0.58***	−0.73***
	[0.30]	[0.23]	[0.24]	[0.30]	[0.21]	[0.23]
$Bound_{s,t-1}$	−0.04	−0.00	−0.02	−0.07	−0.03	−0.05
	[0.16]	[0.13]	[0.14]	[0.16]	[0.12]	[0.14]
Adjusted R^2	0.59	0.56	0.61	0.59	0.56	0.62
Number of establishments		4,447,312			4,447,312	
Number of observations		31,031,426			31,031,426	
Dependent variables	Panel B: *Days Payable Outstanding*					
$Bound_{s,t-1} \times \Delta MW(F)_t$	1.77***	1.16***	1.47***	1.72***	1.12***	1.42***
	[0.51]	[0.42]	[0.41]	[0.52]	[0.39]	[0.40]
$Bound_{s,t-1}$	−0.05	−0.06	−0.06	0.00	−0.01	−0.01
	[0.27]	[0.22]	[0.24]	[0.27]	[0.20]	[0.23]
Adjusted R^2	0.59	0.56	0.61	0.59	0.56	0.62
Number of establishments		4,447,312			4,447,312	
Number of observations		31,031,426			31,031,426	
Establishment fixed effects	✓	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓	✓
Establishment controls				✓	✓	✓

Notes. Panel A of this table reports results from our baseline regression Equation (1) estimating the differential effect of the federally mandated minimum wage on an establishment's credit score using the *PAYDEX Score* as a dependent variable. In columns (1)–(3), we estimate the regression equation without establishment controls, and in columns (4)–(6), we report results with a full set of establishment-level control variables ($X_{i,t-1}$) in our regressions: size (measured as log(*Sales*)), age (log(*Age*)), number of employees (log(*Employees*)), and sales growth. These variables are winsorized at their 1st and 99th percentiles. Columns (1) and (4) report results for a minimum *PAYDEX Score* during the year, and columns (2) and (5) report results for a maximum *PAYDEX Score* during the year. In columns (3) and (6), we report results for an average score during the year measured as the mean of the minimum and maximum score during the year. In panel B, we utilize the one-to-one mapping of credit score to average payment days and rerun the similar regressions with *Days Payable Outstanding* as the dependent variable. Note that $Bound_{s,t-1}$ is a dummy variable equal to 1 if, at the beginning of fiscal year t, the establishment's state s has a state minimum wage less than or equal to the federal minimum wage. The variable $\Delta MW(F)_t$ measures the nominal dollar increase in the federal minimum wage in year t; otherwise, it equals zero. Therefore, the interaction term, $Bound_{s,t-1} \times \Delta MW(F)_t$, identifies the differential effect of the federally mandated minimum wage over and above the effect of state-level variation caused by a change in the state-determined minimum wage and change in status from bounded to unbounded or vice versa. Standard errors are in brackets and are clustered at the state level.

*$p < 0.10$; **$p < 0.05$; ***$p < 0.01$.

In columns (1)–(3), we estimate the regression equation without establishment controls, and in columns (4)–(6), we report results with a full set of establishment-level control variables ($X_{i,t-1}$) in our regressions: size (measured as log(*Sales*)), age (log(*Age*), number of employees (log(*Employees*)), and sales growth, all of which are winsorized at their 1st and 99th percentiles. Columns (1) and (4) report results for a minimum PAYDEX score during the year, and columns (2) and (5) report results for a maximum PAYDEX score during the year. In columns (3) and (6), we report results for an average score during the year, measured as the mean of the minimum and maximum score during the year.

Our preferred specification is presented in column (6). It shows a point estimate of −0.73, implying that for a one-dollar increase in the federal minimum wage, establishments in bounded states experience a reduction in their average PAYDEX score by 0.73 points relative to changes in the PAYDEX score of establishments in unbounded states. To understand the point estimate, we utilize the one-to-one mapping of credit score to average payment days and rerun the similar regressions with *Days Payable Outstanding* as the dependent variable. In panel B, we find that the point estimate is 1.42 days. The median establishment in our sample pays bills an average of five and a half days beyond the term, with a mean of 14.4 days. The results imply a 9.8% increase in payment days outstanding compared with the mean (26% increase compared with the median) after the federal minimum wage increases by a dollar. For the analysis going forward, we use *Average PAYDEX Score* as our primary measure of interest. In Section 4.3, we discuss how the decline in PAYDEX scores can impact business survival probabilities.

4.1.2. Dynamics in the PAYDEX Scores of Affected Establishments Before and After Minimum Wage Increases.

As previously discussed in Section 2, our above results can only identify the effect of minimum wage increases to the extent that the PAYDEX score of establishments in bounded and unbounded states are following similar trends around the time that the federal

government adjusts minimum wages. We test this assumption in this subsection. We estimate the following equation:

$$Y_{it} = \sum_{j=-5}^{5} \alpha_j BD_{s,t}(j) + \sum_{j=-5}^{5} \alpha_j Bound_{s,t}(j) + \kappa X_{i,t-1} + \nu_i + \omega_t + \epsilon_{ist}. \quad (2)$$

In the above equation, $BD_{s,t}$ is defined as $Bound_{s,t-1} \times \Delta MWDummy(F)_t$, and all the controls are similar to those included in Equation (1). The inclusion of the dummy $Bound_{s,t-1}$ for both pre- and post-window controls for the changing status of bounded to unbounded or vice versa. Here, we estimate these interaction terms for the five years before and the five years after the minimum wage increase.

We present our regression results graphically in Figure 4(b). The bar plots the regression coefficients of the interaction term, identifying bounded states for five years before and after the federal minimum wage increase, and the dashed lines correspond to 95% confidence intervals. The bold dashed line indicates the period immediately before the federal minimum wage change.

Similar to Figure 4(a), we observe that establishments in bounded states did not experience differential trends prior to the introduction of federal minimum wage changes. Second, we note that in the year of the federal minimum wage increase, there is a sharp decline in the PAYDEX score for establishments in bounded states. This is consistent with our baseline results reported in Section 4.1.1. Finally, the difference between the PAYDEX score for establishments in the bounded and unbounded states converges over three to five years. One possible reason for this could be that establishments that managed to survive, which is potentially an indicator of the stronger establishments, may be able to pass on some of these extra labor costs to customers over an extended period.[18]

Next, we track the before and after changes in PAYDEX score using alternative long-term estimation. This alternative long-term estimation allows us to include case-firm and case-time fixed effects. We discuss these results in Section IA1.2 of the internet appendix. Overall, we find similar dynamic patterns if we limit our sample to firms that exist one year before a minimum wage increase and drop firms that exit after the minimum wage increase. The results suggest that our results are not primarily driven by some businesses getting affected by changes in minimum wage and exiting. The evidence suggests that the decline was temporary for the surviving businesses, and businesses can adjust to the increased labor costs over time.

4.1.3. Robustness Tests. In this section, we test the robustness of our main result reported in column (6) of Table 2 to various potential confounding factors. We present the results of these robustness checks in Table 3.

One potential concern with the interpretation of the results presented so far is that they may be driven by the entry of numerous small, unhealthy firms into bounded states. To account for this, we interact all establishment controls with the bounded dummy, and we report results in column (1). We find that the negative effect declines from −0.73 to −0.70 but remains statistically significant. Another potential concern may be that industry-specific, time-varying, unobserved heterogeneity in an establishment's PAYDEX score is driving our results. We address this concern by including NAICS4 × year fixed effects in column (2) instead of year fixed effects in our baseline specification. Our results remain robust. In order to address any concerns about the construction of our sample, we include all the establishments that we omit from our baseline in column (3). We make use of 90 million observations for 15 million establishments. Although the magnitude falls by 0.10 points, it remains significant. In column (4), we include all the industries that we omitted from our baseline sample, and we find similar results.

In our baseline specification, we excluded multiestablishment firms since they are typically larger, they are different from single-establishment firms, and they are less likely to be affected by minimum wage increases. In column (5), we include multiestablishment businesses, and we find that our negative effect reduces the magnitude but remains statistically significant at the 1% level. In column (6), we find an almost insignificant effect on the credit scores for businesses connected with multiple-establishment firms.

In columns (7) and (8), we test the robustness of our baseline results to our definition of minimum wage changes. We define $\Delta MWDummy(F)_t$ as an indicator variable equal to 1 if there is an increase in the federal minimum wage in year t and 0 otherwise. Moreover, $\%\Delta MW(F)_t$ captures the percentage change in the minimum wage by the federal government in year t, and equals 0 if no change occurred in year t. For example, in the year 2007, the federal minimum wage increased from \$5.15 to \$7.25, which implies an increase of almost 40%. We replace $\Delta MW(F)_t$ with $\%\Delta MW(F)_t$ and $\Delta MWDummy(F)_t$, respectively, and we report the regression results. The regression coefficient suggests a decline in PAYDEX scores by (0.40 × 3.85=) 1.85 points. This reduction in score implies a delay in payment by nearly three days. We find consistent results when we use a dummy instead of a change measure.

In column (8), we report the dynamics of the three main federal minimum wage increases under examination: $\Delta MWDummy(F)_t$ Yr1 is an indicator variable equal to 1 that identifies the years 1990, 1996, and 2007; $\Delta MWDummy(F)_t$ Yr2 identifies the years 1991, 1997, and 2008; and $\Delta MWDummy(F)_t$ Yr3 identifies the year

Table 3. Robustness Tests

Dependent variables	*Average PAYDEX Score*								
	(1) Bound × controls	(2) NAICS4 × year fixed effects	(3) No drop	(4) All industries	(5) Include multiestablishment	(6) Only multiestablishment	(7) Percentage change	(8) Dummy	(9) Series
$Bound_{s,t-1}$ × $\Delta MW(F)_t$	−0.70***	−0.74***	−0.63***	−0.68***	−0.65***	−0.24*			
	[0.24]	[0.23]	[0.21]	[0.23]	[0.22]	[0.13]			
$Bound_{s,t-1}$ × $\%\Delta MW(F)_t$							−3.85***		
							[1.33]		
$Bound_{s,t-1}$ × $\Delta MWDummy(F)_t$								−0.43***	
								[0.15]	
$Bound_{s,t-1}$ × $\Delta MWDummy(F)_t$ Yr1									−0.48*
									[0.25]
$Bound_{s,t-1}$ × $\Delta MWDummy(F)_t$ Yr2									−0.40**
									[0.17]
$Bound_{s,t-1}$ × $\Delta MWDummy(F)_t$ Yr3									−0.42**
									[0.17]
$Bound_{s,t-1}$	−0.57	−0.03	−0.04	−0.07	−0.07	−0.20	−0.06	−0.07	−0.07
	[0.55]	[0.13]	[0.12]	[0.13]	[0.14]	[0.16]	[0.13]	[0.14]	[0.13]
Adjusted R^2	0.62	0.62	0.59	0.62	0.60	0.44	0.62	0.62	0.62
Number of establishments	4,447,312	4,447,312	15,046,496	6,632,327	4,921,821	474,509	4,447,312	4,447,312	4,447,312
Number of observations	31,031,426	31,031,419	90,291,929	45,172,568	37,235,728	6,204,302	31,031,426	31,031,426	31,031,426

Notes. In this table, we report results for various robustness tests of our baseline specification (i.e., column (6) of Table 2). In column (1), we report results in which we interact the establishment controls with the bound dummy. Column (2) reports regression results in which we include NAICS4 industry × year fixed effects. In column (3), we do not omit any data, and we report regression results on the full sample. In column (4), we include all the industries that we omit from our baseline specification. In column (5), we also include multiestablishment businesses to our baseline specification. In column (6), we report results for multiestablishment businesses. We replace $\Delta MW(F)_t$ with $\%\Delta MW(F)_t$ and $\Delta MWDummy(F)_t$ and report results in columns (7) and (8), respectively. Note that $\Delta MWDummy(F)_t$ is an indicator variable equal to 1 if there is an increase in federal minimum wage in year t, and 0 otherwise; $\%\Delta MW(F)_t$ is a change measure indicating the percentage increase in the minimum wage by the federal government in year t, and 0 otherwise (in column (9), we report the dynamics); $\Delta MWDummy(F)_t$ Yr1 is an indicator variable equal to 1 that identifies the years 1990, 1996, and 2007, and 0 in all other years; $\Delta MWDummy(F)_t$ Yr2 identifies the years 1991, 1997, and 2008, and is 0 in all other years; and $\Delta MWDummy(F)_t$ Yr3 identifies 2009. All regressions are with establishment fixed effects and year fixed effects, and also include establishment controls. Standard errors are in brackets and are clustered at the state level.

*$p < 0.10$; **$p < 0.05$; ***$p < 0.01$.

2009. We find that the effect is negative and reduces over time.

Additional robustness tests of our baseline specification (Equation (1)) are presented in Section IA1.3 in the internet appendix. As shown in Figure 2, many states switch from being bounded to unbounded before a federal minimum wage increase. Also, some bounded and unbounded states may increase their state minimum wages in response to federal minimum wage increases.[19] In our robustness tests, we omit all observations in the state-years in which bounded and unbounded states respond to federal minimum wage increases. It is possible that not all unbound states may serve as a good control group because their state minimum wage may be much higher than the federal minimum wage, and these states may have different policies in place to support small businesses. Results in Section IA1.3.1 of the internet appendix suggest that all unbound states can serve collectively as appropriate control states. We also test if establishments located in states that are newly bounded by the federal minimum wage are impacted differently (see Section IA1.3.2 in the internet appendix). We further consider weighted regression and find consistent results (see Section IA1.3.3 in the internet appendix). As a further robustness test, we explicitly control for these state-level, county-level, and ZIP code–level observables in our regressions (see Section IA1.4 in the internet appendix). We find similar results. Further, we match establishments in bounded states (treatment group) with those in unbounded states (control group) based first on the pre-shock level of the establishments' credit scores, followed by trends in the treated establishments' credit scores, as well as establishment-, state-, county- and ZIP code–level observable characteristics. We find that for establishments in the bounded states, after a one-dollar increase in the federal minimum wage, the PAYDEX score declines by an additional 0.75 points, compared with similar establishments located in the unbounded states. It also appears that selection on the levels and trends in the observable characteristics are not driving our main results (see Section IA1.5 in the internet appendix for details).

In addition to the matching estimates presented in Section IA1.5 in the internet appendix, the bordering-county estimates presented in Section 4.1.4 and fuzzy triple differences method using the FLSA Exemption Rule in Section 4.1.5 further strengthen our counterfactual analysis through a more stringent construction of the control group.

4.1.4. Controlling for Local Economic Conditions: Bordering Counties. We attempt to further control for local economic conditions by analyzing the establishments located in the contiguous counties next to state borders. The underlying assumption for this identification strategy requires that the adjacent counties at state borders have similar economic conditions along all dimensions other than the minimum wage.

We begin our bordering-county analysis by utilizing the federal minimum wage increase of 2007 and test how establishments located in the contiguous counties of bounded and unbounded states are affected. We follow Dube et al. (2010) and identify 1,135 unique contiguous counties at state borders. For 2006, we found 497 unique contiguous counties and 982 county-pairs in which, across the state border, one state was bounded by the federally mandated minimum wage and the other was not. We start our analysis with the establishments in 2006 that are located in county-pairs in which *bounded* states bordered *unbounded* states. Then, we track their PAYDEX score from 2004 to 2009. We estimate Equation (1) with establishment-level fixed effects and include county-pair × year fixed effects to control for any time-varying county-pair-level unobservables. We also control for time-varying establishment-, state-, county-, and ZIP code–level observable characteristics. The inclusion of county-pair × year fixed effects and time-varying observable characteristics should satisfy the identification assumption, namely, the assumption that adjacent counties at state borders have similar economic conditions except for their respective minimum wage bounded status.[20]

In column (1) of Table 4, we keep establishments located in counties where the distance between the centroids of the bordering counties is less than 25 miles. We find that for establishments in the bordering county of the bounded state, a one-dollar increase in the federal minimum wage reduces the PAYDEX score by 0.52 points more than similar establishments located in the bordering county of the unbounded state after controlling for time-varying establishment-, state-, county-, and ZIP code–level observable characteristics and time-varying county-pair-specific unobservables.

According to the 2015 Current Population Survey, restaurants (NAICS 72) and retail trade (NAICS 44, 45) are the only industries in which over 10% of employees earn the minimum wage. We test whether the magnitude of the impact is higher for such industries, by interacting Equation (1) for the bordering-county sample with three dummies identifying restaurants, retail establishments, and establishments in other sectors. The results in column (2) suggest that the negative effect is present in all sectors but is more substantial for restaurants. Restaurants experience a statistically significant 0.89-point additional decline in credit score as compared with other establishments. In column (3), we include establishments located in counties where the distance between the centroids of the bordering counties is less than 50 miles and test how these effects vary based on the distance between the centroid of the bordering counties. We find that the decline in PAYDEX score is more considerable when the distance between

Table 4. Controlling for Local Economic Conditions: Bordering Counties

Dependent variables	*Average PAYDEX Score*					
	(1)	Industry (2)	Distance (3)	All years (4)	(5)	(6)
$Bound_{s,t-1} \times \Delta MW(F)_t$	−0.52***			−0.77***	−0.72***	−0.61***
	[0.09]			[0.12]	0.12]	[0.10]
$Bound_{s,t-1}$	−0.03			0.19	0.19	0.12
	[0.13]			[0.13]	[0.13]	[0.14]
$Bound_{s,t-1} \times \Delta MW(F)_t$ × *Restaurants*		−1.38***				
		[0.39]				
$Bound_{s,t-1} \times \Delta MW(F)_t$ × *Retail*		−0.43***				
		[0.15]				
$Bound_{s,t-1} \times \Delta MW(F)_t$ × *Others*		−0.48***				
		[0.12]				
$Bound_{s,t-1} \times \Delta MW(F)_t \times \mathbb{1}_{Distance \leq 25}$			−0.58***			
			[0.10]			
$Bound_{s,t-1} \times \Delta MW(F)_t \times \mathbb{1}_{25<Distance \leq 50}$			−0.33**			
			[0.13]			
Establishment fixed effects	✓	✓	✓	✓	✓	✓
County-pair × year fixed effects	✓	✓	✓	✓	✓	
Interaction terms		✓	✓			
NAICS4 × year fixed effects					✓	
County-pair × NAICS4 × year fixed effects						✓
Adjusted R^2	0.74	0.74	0.75	0.71	0.71	0.72
Number of observations	928,765	928,765	1,964,260	1,415,354	1,415,308	1,376,962
Restaurants–others		−0.89				
p-value		0.05				
$\mathbb{1}_{Distance \leq 25}$ - $\mathbb{1}_{25<Distance \leq 50}$			−0.25			
p-value			0.18			

Notes. This table reports our baseline regression Equation (1), where we control for local economic conditions by analyzing the establishments located in the contiguous counties at the state borders. See Section 4.1.4 for more details. In columns (1)–(3), we keep data for establishments that exist in 2006 and utilize their data for 2004 to 2009, and in columns (4)–(6), we redo our analysis for all years for establishments located in the contiguous counties. In column (1), we keep establishments located in counties where the distance between the centroid of the bordering counties is less than 25 miles. We interact Equation (1) for the bordering-county sample with three dummies identifying restaurants (NAICS 72), retail establishments (NAICS 44, 45), and establishments in other sectors. Column (2) reports the regression results. In column (3), we include establishments located in counties where the distance between the centroids of the bordering counties is less than 50 miles and interact Equation (1) with the dummy variables $\mathbb{1}_{Distance \leq 25}$ and $\mathbb{1}_{25<Distance \leq 50}$. The dummy variable $\mathbb{1}_{Distance \leq 25}$ identifies the county-pairs where the distance between the centroid of the bordering counties is less than 25 miles. Other dummies are defined similarly. We reestimate the specification of column (1) for all federal minimum wage changes in our sample and report results in column (4). We include NAICS4 × year and county-pair × NAICS4 × year fixed effects and report results in columns (5) and (6). In all regressions, we control for time-varying establishment-, state-, county-, and ZIP code–level observable characteristics, as in Section IA1.4 in the internet appendix. Standard errors are in brackets and are clustered at the state level.

*$p < 0.10$; **$p < 0.05$; ***$p < 0.01$.

the centroid of the bordering counties is less than 25 miles.

We reestimate the specification of column (1) for all federal minimum wage changes in our sample and report results in column (4). For all federal minimum wage changes, in the bordering-county sample, we find a decline in the PAYDEX score of 0.77 points. Next, we include NAICS4 × year and county-pair × NAICS4 × year fixed effects and report results in columns (5) and (6). We find that our results are robust and qualitatively similar after controlling for within county-pair time-varying industry unobservables.

4.1.5. FLSA Exemption Rule. In the previous section, we control for local economic conditions by analyzing the establishments located in the contiguous counties next to state borders. However, it is possible that the contiguous counties method may not sufficiently control for potential differences in policies across bounded and unbounded states in the baseline estimates. For example, states with more generous minimum wage rates (i.e., unbounded states) may also have other policies in place that allow businesses to absorb negative shocks like increased labor costs more easily. They may also adopt more generous policies to help businesses during a crisis period, which coincides with minimum wage increases.

First, we test if states adopt more generous policies to help businesses around the minimum wage changes and if these policies vary across bounded and unbounded states. We utilize the Upjohn Institute's Panel Database on Incentives and Taxes. We estimate regression Equation (1) at the state level for the years 1985 to 2012 and business incentives as the dependent variable

(see Figure IA4 in the internet appendix for regression results). We find that the bounded and unbounded states provide similar business incentives before and after federal minimum wage changes.

Further, as specified in the FLSA enterprise coverage, businesses with at least two employees and annual sales of less than $500,000—within the bounded states—are exempted from minimum wage requirements. Although these exempted businesses may not be mandated to increase wages, the labor market impact of the law may still force them to do so as a result of increased local labor market competition (Berger et al. 2022). To control these effects in our baseline regression, we implement the fuzzy triple differences method.[21] We keep establishments in both bounded and unbounded states. We begin our analysis on the 2007 to 2009 wave, keeping data between 2006 to 2011. The year 2006 serves as the base year. We identify nonexempted establishments with a $Non\text{-}Exempt_{i,t-1}$ dummy, which is equal to 1 if the establishment is not exempt from FLSA based on the $500,000 threshold for sales as defined by enterprise coverage under FLSA. To capture the $2.10 increase to the federal minimum wage that was phased-in over the three years, we define $\Delta MW(F)_0$ as $0.7 for the year 2007, and 0 otherwise; $\Delta MW(F)_1$ is defined as $1.4 for the year 2008, and 0 otherwise. Similarly, $\Delta MW(F)_2$, $\Delta MW(F)_3$, and $\Delta MW(F)_4$ are defined at $2.1 for the year 2009, 2010, and 2011, respectively, and 0 otherwise.

Table 5 reports the regression results. Our coefficients of interest are $Non\text{-}Exempt_{i,t-1} \times Bound_{s,t-1} \times \Delta MW(F)_I$, where I ranges from 1 to 4. In column (1), we keep sample establishments with ex ante sales (one year before) around (±$5,000) the $500,000 threshold. Thus, the treatment group is establishments with sales between $500,000 and $505,000. The control group includes establishments with sales less than $500,000 and more than $495,000 (see Table IA22 in the internet appendix for information on sample selection). We include county × NAICS4 × year fixed effects across all specifications. These fixed effects ensure that we are comparing firms within the same county-industry-year. We find a strong negative impact on average PAYDEX score by almost 2.5 points in the year of a minimum wage increase for nonexempted firms located within bounded states, but this negative effect diminishes with time. We observe similar negative and significant effects when we increase

Table 5. Evidence from Enterprise Coverage of FLSA

Dependent variables	*Average PAYDEX Score*					
	Sales threshold around $500,000 exemption limit					
	(±5,000) (1)	(±10,000) (2)	(±15,000) (3)	(±5,000) (4)	(±10,000) (5)	(±15,000) (6)
$Non\text{-}Exempt_{i,t-1} \times Bound_{s,t-1}$						
$\times \Delta MW(F)_0$	−2.55*	−2.04***	−2.12***	−2.49*	−2.00***	−2.05***
	[1.29]	[0.75]	[0.68]	[1.27]	[0.73]	[0.67]
$\times \Delta MW(F)_1$	−1.98***	−0.92**	−0.99***	−1.93***	−0.94**	−0.99***
	[0.68]	[0.43]	[0.35]	[0.65]	[0.43]	[0.36]
$\times \Delta MW(F)_2$	−0.12	−0.05	0.04	−0.12	−0.06	0.05
	[0.45]	[0.32]	[0.25]	[0.44]	[0.32]	[0.25]
$\times \Delta MW(F)_3$	−0.61*	−0.14	−0.04	−0.64**	−0.16	−0.06
	[0.33]	[0.24]	[0.29]	[0.32]	[0.24]	[0.29]
$\times \Delta MW(F)_4$	−1.50	0.02	0.09	−1.50	0.01	0.08
	[0.95]	[0.60]	[0.50]	[0.95]	[0.60]	[0.51]
Establishment fixed effects	✓	✓	✓	✓	✓	✓
$Non\text{-}Exempt_{i,t-1} \times \Delta MW(F)_i$	✓	✓	✓	✓	✓	✓
Other interaction terms	✓	✓	✓	✓	✓	✓
County × NAICS4 × year fixed effects	✓	✓	✓	✓	✓	✓
Controls				✓	✓	✓
Adjusted R^2	0.68	0.67	0.68	0.68	0.68	0.68
Number of observations	149,497	168,342	185,842	149,497	168,342	185,842

Notes. This table reports the regression results where we utilize the within-state variation for bounded states using the $500,000 threshold for sales as defined by enterprise coverage under the FLSA. We focus on the 2007–2009 wave and keep samples for 2006–2011. The year 2006 serves as the base year. The $Non\text{-}Exempt_{i,t-1}$ dummy is equal to 1 if the establishment is not exempt from FLSA based on the $500,000 threshold for sales as defined by enterprise coverage under FLSA; $\Delta MW(F)_0$ is defined as $0.7 for the year 2007, and is 0 otherwise; $\Delta MW(F)_1$ is defined as $1.4 for the year 2008, and is 0 otherwise. Similarly, $\Delta MW(F)_2$, $\Delta MW(F)_3$, and $\Delta MW(F)_4$ are defined as $2.1 for the years 2009, 2010, and 2011, respectively, and 0 otherwise. In column (1), we keep sample establishments with ex ante sales (one year before) around (±$5,000) the $500,000 threshold. Thus, the treatment group includes establishments with sales between $500,000 and $505,000. The control group includes establishments with sales less than $500,000 and more than $495,000. We keep establishments in both bounded and unbounded states. Similarly, other thresholds are defined in columns (2) and (3). County × NAICS4 × year fixed effects and establishment fixed effects are included in all specifications. In columns (4)–(6) we also include establishment-level controls. Standard errors are in brackets and are clustered at the state level.

*$p < 0.10$; **$p < 0.05$; ***$p < 0.01$.

the sales intervals around the \$500,000 threshold to ±\$10,000 and ±\$15,000 (columns (2) and (3), respectively). We find similar results in columns (4)–(6) when we include establishment-level controls. Finally, we also estimate the fuzzy triple design tests for the 1996–1997 wave. We find an immediate decline in score by almost 4.7 points (see Section IA1.6, Table IA11, in the internet appendix).[22]

Overall, the results from Sections 4.1.3, 4.1.4 and 4.1.5 indicate that after an increase in the federal minimum wage, establishments in bounded states experienced a reduction in their average PAYDEX score relative to changes in the PAYDEX score of establishments in unbounded states. These results are robust after controlling for various time-invariant and time-varying observable and unobservable characteristics that may be correlated with the timing of the federal minimum wage increase. In the next section, we test heterogeneity of the above results based on affected firms and geographies.

4.2. How Does the Impact of One-Size-Fits-All Federal Minimum Wage Increases Vary Across Firms and Geography?

As discussed before, there is significant heterogeneity across firms, industries, and labor market conditions across the United States. In this section, we analyze how one-size-fits-all federal minimum wage increases impact the cross-section of firm and geography in the United States. In particular, we examine how the minimum wage–induced PAYDEX score effects vary with the establishment's labor intensity, size, age, local competition, and local personal income.

4.2.1. Labor Intensity. We next test the differential effect of a federal minimum wage increase on the establishment's financial health based on its labor utilization. In our data, the median establishment employs 12 employees per \$1 million in sales. We hypothesize that the negative effect of a federal minimum wage increase should be more severe for labor-intensive businesses. First, we partition our sample into quintiles based on labor intensity one year before the federal minimum wage change. Then, we reestimate Equation (1), where we interact the equation by each quintile group. In Figure 5, we plot the regression coefficient on triple interaction terms with a 95% confidence interval. We find that, with a minimum wage increase, more labor-intensive establishments are more adversely affected than less labor-intensive establishments. To ensure that local economic conditions in the bounded versus unbounded states are not driving our results, we also control for state-year fixed effects in our cross-sectional results. Table 6 reports the triple interaction results.

We partition our sample into two groups using the median establishment labor-intensity one year before the federal minimum wage change. We define *MoreLabor* as 1 if the establishment's labor-intensity measure is above median labor-intensity, and 0 otherwise. In column (1), we include establishment controls, establishment fixed effects, and state-year fixed effects. We are able to hold all state-year-specific heterogeneity constant through the inclusion of these state-year fixed effects, and we identify our triple interaction effect through within state-year across firm-size variation by interacting *MoreLabor* with our main coefficient ($Bound_{s,t-1} \times \Delta MWDummy(F)_t$) from Equation (1). We also include NAICS4 × year fixed effects to absorb any industry-year-specific heterogeneity that may exist. We find that the effect is stronger for more labor-intensive establishments relative to less labor-intensive establishments within the bounded states.

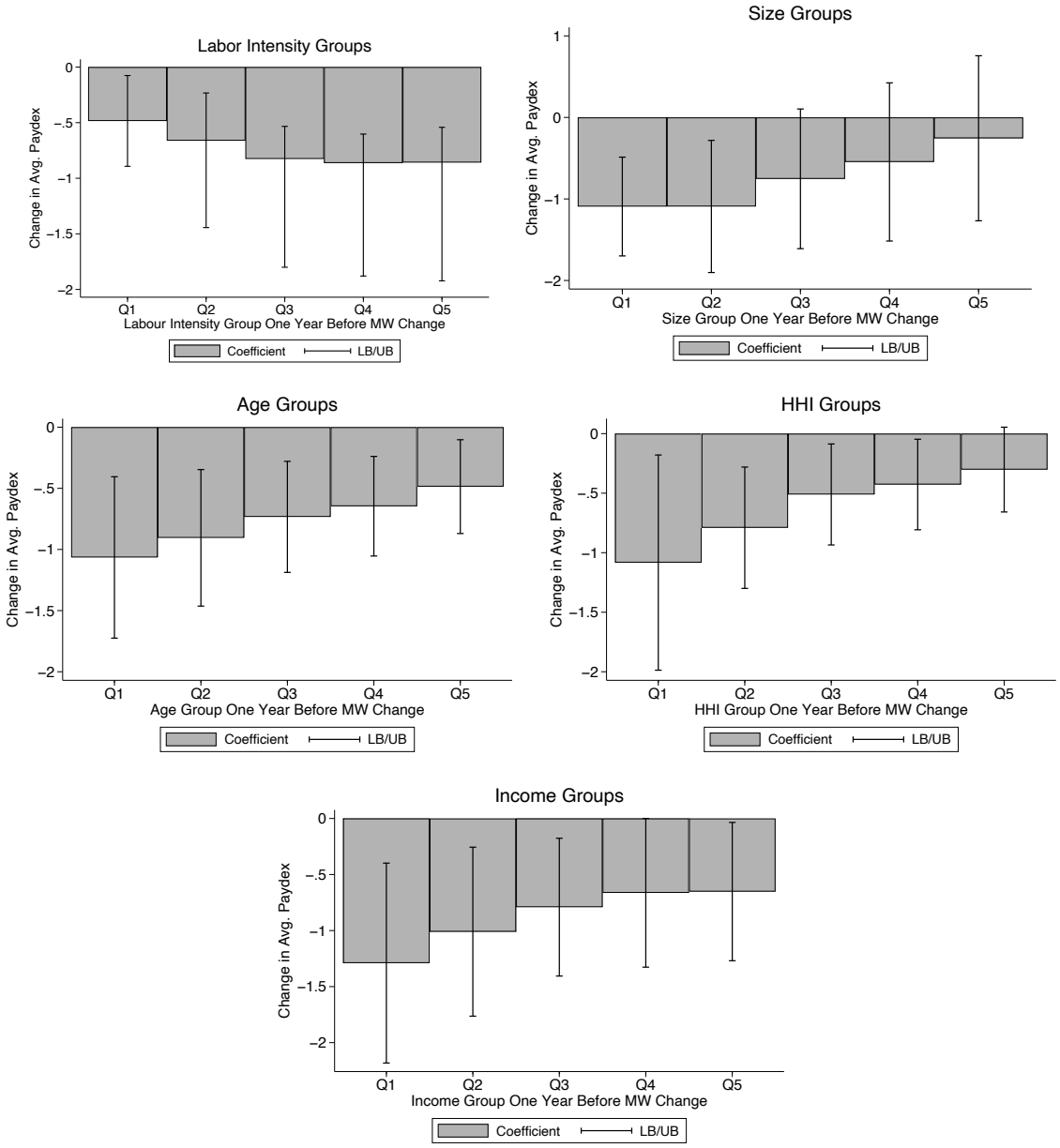
We extend our analysis by calculating labor cost instead of labor intensity. We measure the establishment's labor cost as (number of employees × average salary)/sales. We use Quarterly Census of Employment and Wages (QCEW) data to estimate the average compensation at the NAICS4 industry-County level. Table 6, column (2), reports the regression results. The results are similar to the labor-intensity results. We find consistent negative results for labor-intensive businesses using dynamic regressions (see Figure IA5 in the internet appendix).[23] Overall, we find that, with the minimum wage increase, establishments with more employees are more adversely affected than establishments with fewer employees.

4.2.2. Establishment Size and Age. In this subsection, we test the differential effect of a federal minimum wage increase on an establishment's financial health based on its size and age. These measures may serve as a proxy for the ability of the businesses to absorb the financial shock caused by an increase in labor cost.

Similar to our treatment of labor intensity, we partition our sample into quintiles based on size (measured by sales) and age one year before the federal minimum wage change (see Figure 5). We find that small and young establishments are more adversely affected by the minimum wage increase.[24] We find similar results using triple interactions (see Table 6, column (3)). We conduct the same analysis for establishment age and find similar results (see Table 6, column (4)): younger firms experience larger decreases in their PAYDEX score than older firms.

4.2.3. Local Product Market Competition. With an increase in labor costs, the cost of goods sold increases for businesses. If establishments can completely and immediately pass on these increased costs to their customers, then they may not feel any additional financial stress as a result. In this subsection, we test this possibility by examining the relative local competitiveness in a given firm's industry. The establishments in our sample are relatively small businesses, and local competition

Figure 5. PAYDEX Score: Heterogeneity



Notes. We test the differential effect of the federal minimum wage on PAYDEX scores for establishments located in bounded vs. unbounded states based on the establishment's (a) labor intensity, (b) size, (c) age, (d) local competition, and (e) local personal income. These panels plot the regression coefficients of Equation (1) with a 95% confidence interval, where we interact the equation with each quintile group based on the above measures one year before the federal minimum wage change.

determines their cash flows. We expect that an establishment within the same industry and located in a less competitive neighborhood may find it easier to pass on the increased labor costs compared with other establishments, and they may experience a smaller reduction in PAYDEX scores.

To test the effect of local competition on a firm's ability to pass on these costs, we measure local product market competition using the HHI index, measured at the NAICS5 Industry-County-year level. To create the HHI index, we use the full set of 50 million establishments in the NETS data set. Similar to the previous subsection, we first partition our sample into quintiles based on the HHI index one year before a federal minimum wage change. We find that, with a minimum wage increase, establishments in more competitive locations are adversely affected, whereas establishments in less competitive locations are not negatively affected at all (see Figure 5). We also partition our sample into two groups and split the HHI index at its

Table 6. Heterogeneity

Dependent variables	*Average PAYDEX Score*					
	Labor		Establishment		Local	
Group	Intensity *MoreLabor* (1)	Cost *MoreLabor* (2)	Size *Small* (3)	Age *Young* (4)	Competition *HighCompetition* (5)	Personal Income *LessIncome* (6)
$Group \times Bound_{s,t-1} \times \Delta MW(F)_t$	−0.24***	−0.14**	−0.50***	−0.25**	−0.21**	−0.23***
	[0.08]	[0.08]	[0.08]	[0.12]	[0.10]	[0.06]
$Group \times Bound_{s,t-1}$	−0.01	0.05	0.09**	−0.35***	0.03	0.11***
	[0.03]	[0.03]	[0.04]	[0.04]	[0.03]	[0.03]
$Group \times \Delta MW(F)_t$	0.27***	0.23***	0.62***	0.58***	0.29***	0.29***
	[0.04]	[0.04]	[0.05]	[0.09]	[0.08]	[0.04]
Establishment fixed effects	✓	✓	✓	✓	✓	✓
Establishment controls	✓	✓	✓	✓	✓	✓
NAICS4 × year fixed effects	✓	✓	✓	✓	✓	✓
State × year fixed effects	✓	✓	✓	✓	✓	
County × year fixed effects						✓
Adjusted R^2	0.62	0.62	0.62	0.62	0.62	0.67
Number of establishments	3,603,144	3,095,275	4,423,967	4,420,503	4,420,503	3,885,212
Number of observations	25,739,939	21,005,381	30,902,244	30,902,244	30,871,111	21,149,326

Notes. This table reports heterogeneity for our baseline regression Equation (1) based on the establishment's labor intensity, size, age, local competition, and local personal income. We measure an establishment's labor intensity as the number of employees per $1 million in sales, labor cost as (number of employees × average salary)/sales (where we use QCEW data to estimate the average compensation at the NAICS4 industry-County level), size as sales one year before the federal minimum wage change, and age using the year of establishment. Finally, we measure local product market competition using the HHI index (based on sales) measured at the NAICS5 Industry-County-year level, and we measure local personal income using IRS data at the ZIP code level. We define group dummies *MoreLabor*, *Small*, *Young*, *HighCompetition*, and *LessIncome* based on the medians of the above measures. We do analysis using triple interaction. We include establishment controls, establishment fixed effects, NAICS4 × year fixed effects, and state × year fixed effects or county × year fixed effects Standard errors are in brackets and are clustered at the state level.

*$p < 0.10$; **$p < 0.05$; ***$p < 0.01$.

median one year before the federal minimum wage change. In Table 6, column (5), we include a triple interaction to identify our effect of interest. As such, our tests effectively compare two establishments in the same industry and the same bounded state, and we exploit only variation in competition across industry-states. We find a strong negative effect for establishments located in counties with more competition. We find that the effect is very strong and dominant for establishments in more competitive areas in subsample analysis (see columns (1) and (2) of Table IA18 in the internet appendix).

Overall, our results suggest that some small businesses located in bounded states, especially those located in more competitive counties, are more affected by one-size-fits-all federal minimum wage increases. These establishments may not be able to completely pass on the increased costs to their customers, and hence they experience some financial stress.

4.2.4. Local Personal Income. Similar to local competitiveness, the personal income of a firm's local customers may determine the firm's ability to immediately increase prices. The increase in the minimum wage, on one hand, increases labor costs for businesses, but, at the same time, it increases the per capita local income. If businesses can pass on these costs to customers in low-income ZIP codes, then we should not find a decline in their PAYDEX scores. Otherwise, we should expect a more negative effect in low-income neighborhoods.

Similar to the previous subsection, we first partition our sample into quintiles based on ZIP code–level IRS data on personal income one year before a federal minimum wage change (Figure 5). We find that, with the minimum wage increase, establishments in the lowest-income neighborhoods are the most adversely affected. We find similar results when we partition our sample into two groups and include establishment controls, establishment fixed effects, and county-year fixed effects (see Table 6, column (6), and Table IA18, columns (5)–(8), in the internet appendix). Here, our tests essentially compare two establishments in the same industry and in the same bounded state; we find a strong negative impact on establishments located in ZIP codes with low income.[25]

Overall, the results in this section suggest that small and young establishments, which are more likely to have financial constraints, experience a more significant decrease in their credit scores after minimum wage changes. Further, labor-intensive businesses (i.e., those with high labor costs), those located in counties with more competition, and those located in the low-income

ZIP codes appear to have difficulty in passing-through their increased labor costs to consumers; consequently, they experience a more significant decrease in their credit score.

4.3. Threshold Effects and Exit of Establishments After Minimum Wage Increases

In the previous sections, we find that, with an increase in the minimum wage by the federal government, there is a differential effect on the PAYDEX scores of establishments located in bounded versus unbounded states. This effect is stronger for labor-intensive, small, and young businesses and those businesses located in low-income and competitive neighborhoods. In this section, we document how decline in credit scores with minimum wage increases affects the probability of exit and also overall impact of minimum wage increases on the probability of exit, independent of credit scores.

Note that the PAYDEX score is one of the metrics that suppliers can use to determine whether a new client or business partner might present risks going forward. Low PAYDEX scores may make suppliers reluctant to do business with a firm or may limit the size and scope of the services to which they are willing to agree. Businesses with healthier PAYDEX scores are more likely to be approved for financing, and for better terms (i.e., higher loan amounts, longer repayment terms, and lower interest rates). Good PAYDEX scores can also put businesses in a position to negotiate better terms with lender, as they show that businesses are reliable with their debt obligations. Figure 3 suggests that about 70% of the establishments in our database have a score of more than 70, whereas the average of 80.35 for the group "80 and above" suggests a lumping of data at 80. Next, we test the importance of making payments "on time" or having a "perfect score" (i.e., 80 points) and how this score affects the probability of exit.

Our dependent variable is $Exit_{t+1}$, a dummy variable equal to 1 if the establishment exits in year $t+1$. We interact our baseline Equation (1) with $SameScore_{it}$ (80), a dummy variable identifying establishment-years in which the establishment does not observe any change in its score (i.e., 80) from year $t-1$ to year t. Table 7 reports the regression results. Column (1) reports the coefficient from this triple interaction term, which is negative and significant, indicating that establishments located in bounded states that do not observe a decline in their credit score after federal minimum wage changes have a lower exit probability in the following year. Next, in column (2), we test the differential effect on establishments that

Table 7. Establishment Exit and *PAYDEX Score*

Dependent variables	$Exit(=1)_{t+1}$					
	(1)	(2)	(3)	(4)	(5)	(6)
$Bound_{s,t-1} \times \Delta MW(F)_t$	−0.016***					−0.016***
$\times SameScore_{it}$ (80)	[0.005]					[0.005]
$Bound_{s,t-1} \times \Delta MW(F)_t$		0.022***				0.018***
$\times Downgrade_{it}$ (80 to 79)		[0.003]				[0.003]
$Bound_{s,t-1} \times \Delta MW(F)_t$			−0.004			−0.006
$\times Downgrade_{it}$ (81 to 80)			[0.005]			[0.004]
$Bound_{s,t-1} \times \Delta MW(F)_t$				0.021**		0.019*
$\times Downgrade_{it}$ (70 to 69)				[0.010]		[0.010]
$Bound_{s,t-1} \times \Delta MW(F)_t$					0.016*	0.014
$\times Downgrade_{it}$ (71 to 70)					[0.009]	[0.009]
$Bound_{s,t-1} \times \Delta MW(F)_t$	−0.007	−0.009**	−0.008*	−0.008*	−0.008*	−0.007
	[0.005]	[0.004]	[0.004]	[0.004]	[0.004]	[0.005]
$Bound_{s,t-1}$	0.008***	0.008***	0.008***	0.008***	0.008***	0.008***
	[0.003]	[0.003]	[0.003]	[0.003]	[0.003]	[0.003]
Interaction terms	✓	✓	✓	✓	✓	✓
Establishment fixed effects	✓	✓	✓	✓	✓	✓
Establishment controls	✓	✓	✓	✓	✓	✓
Year fixed effects	✓	✓	✓	✓	✓	✓
Adjusted R^2	0.18	0.18	0.18	0.18	0.18	0.18
Number of establishments	3,274,815	3,274,815	3,274,815	3,274,815	3,274,815	3,274,815
Number of observations	22,316,622	22,316,622	22,316,622	22,316,622	22,316,622	22,316,622

Notes. This table reports the results from regression Equation (1) estimating the differential effect of the federally mandated minimum wage on the probability of small business exits. We measure *Exit* as the last year of the establishment in the NETS database. Thus, $Exit(=1)_{t+1}$ is a dummy variable measuring the probability of exit in year $t+1$; $SameScore_{it}$ (80) is a dummy variable identifying establishment-years in which the establishment does not observe any change in *PAYDEX Score* (80) from year $t-1$ to year t; and $Downgrade_{it}$ (80 to 79) is a dummy variable identifying establishment-years in which the establishment observes a drop in average score from 80 in year $t-1$ to 79 in year t. All other downgrade variables are similarly defined. In addition to the reported coefficients, we include the dummy for each group and its interaction term with the bound dummy in all regressions. Standard errors are in brackets and are clustered at the state level.

*$p < 0.10$; **$p < 0.05$; ***$p < 0.01$.

observe a decline in credit score across a threshold. We find that a 1.0-point decline in their credit score from 80 to 79 implies a 2.2% increase in the probability of exit. The unconditional probability of exit is 8.5%. Therefore, a one-dollar increase in the federal minimum wage, along with a 1.0-point decline in credit score from 80 to 79, increases the exit probability by 25% more for establishments located in bounded states. We do not find such an exit effect for the decline in credit scores from 81 to 80 (column (3)). However, we find a similar effect for a 1.0-point decline in credit scores from 70 to 69 (column (4)), whereas the results are weaker for declines in credit scores from 71 to 70 (column (5)). In column (6), we include all the groups and find similar results.

We explore these threshold effects for different levels of ex ante PAYDEX scores and find that a 1.0-point decline in PAYDEX score increases the exit probability for most thresholds between 71 to 80 (see Figure IA7 in the internet appendix).[26]

Overall, we identify significant threshold effects wherein an establishment's 1.0-point decline in credit score across different thresholds (80 to 79 and 70 to 69) that arises from an increase in the federal minimum wage corresponds to an increase in the exit probability by almost two percentage points (a more than 20% increase in the hazard of exit from the unconditional exit probability of 0.085).

Next, we test the impact of minimum wage increases on the probability of exit, independent of credit scores. Similar to Section 4.1.4, we confine our sample to the establishments located in the contiguous counties next to state borders, our strictest specification, to control for local economic conditions. The average exit rate for this sample is 4.2%. The magnitude of 0.005 in

Table 8. Establishment Exit: Bordering Counties

Dependent variables	*Exit*(= 1)			
	(1)	Industry (2)	Distance (3)	All years (4)
$Bound_{s,t-1} \times \Delta MW(F)_t$	0.006***			0.003**
	[0.002]			[0.002]
$Bound_{s,t-1}$	−0.004**			−0.002
	[0.001]			[0.001]
$Bound_{s,t-1} \times \Delta MW(F)_t$		0.020***		
$\times$ *Restaurants*		[0.005]		
$Bound_{s,t-1} \times \Delta MW(F)_t$		0.007		
$\times$ *Retail*		[0.005]		
$Bound_{s,t-1} \times \Delta MW(F)_t$		0.005**		
$\times$ *Others*		[0.002]		
$Bound_{s,t-1} \times \Delta MW(F)_t$			0.007***	
$\times \mathbb{1}_{Distance \leq 25}$			[0.002]	
$Bound_{s,t-1} \times \Delta MW(F)_t$			0.004	
$\times \mathbb{1}_{25<Distance \leq 50}$			[0.004]	
Establishment fixed effects	✓	✓	✓	✓
County-pair × year fixed effects	✓	✓	✓	✓
NAICS4 × year fixed effects	✓	✓	✓	✓
Adjusted R^2	0.32	0.32	0.37	0.44
Number of observations	928,726	928,726	1,964,241	1,182,527
Restaurants–others		0.015		
p-value		0.01		
$\mathbb{1}_{Distance \leq 25}$ and $\mathbb{1}_{25<Distance \leq 50}$			0.002	
p-value			0.56	

Notes. This table reports regression results for our tests assessing the impact of minimum wage increases on the probability of exit for establishments in our sample. Similar to Section 4.1.4, we confine our sample to the establishments located in the contiguous counties next to state borders, our strictest specification, to control for local economic conditions. Our dependent variable is *Exit*, a dummy variable equal to 1 if the establishment exits during the year of minimum wage increase. In columns (1)–(3), we keep data for establishments that exist in 2006 and utilize their data for 2004 to 2009, and in column (4), we redo our analysis for all years for establishments located in the contiguous counties. In column (1), we keep establishments located in counties where the distance between the centroids of the bordering counties is less than 25 miles. We interact Equation (1) for the bordering-county sample with three dummies identifying restaurants (NAICS 72), retail (NAICS 44, 45) establishments, and establishments in other sectors. Column (2) reports the regression results. In column (3), we include establishments located in counties where the distance between the centroids of the bordering counties is less than 50 miles and interact Equation (1) with the $\mathbb{1}_{Distance \leq 25}$ and $\mathbb{1}_{25<Distance \leq 50}$ dummy variables. Note that $\mathbb{1}_{Distance \leq 25}$ is a dummy variable that identifies the county-pairs where the distance between the centroids of the bordering counties is less than 25 miles. Other dummies are defined similarly. We reestimate the specification of column (1) for all federal minimum wage changes in our sample and report results in column (4). In all regressions, we control for time-varying establishment-, state-, county-, and ZIP code–level observable characteristics. Standard errors are in brackets and are clustered at the state level.

*$p < 0.10$; **$p < 0.05$; ***$p < 0.01$.

column (1) suggests that, for a one-dollar increase in the federal minimum wage, the exit probability increases by almost 12% for establishments located in the bordering county of the unbounded state after controlling for time-varying establishment-, state-, county-, and ZIP code–level observable characteristics and time-varying county-pair-specific unobservables. Again, the results are much stronger for restaurants but are not limited to those businesses (see Table 8). We also estimate exit regressions with a fixed and constant sample (see Table IA21 in the internet appendix). We include time fixed effects, allowing us to obtain our variation only from minimum wage increase events. The average exit rate for this sample is 15%. We find that 2% more establishments exit for bounded states, whereas almost 10% more restaurants shut down three years after the minimum wage increase. We find similar results when we aggregate our data at the county-industry level (see Section IA1.9 in the internet appendix). We also test the aggregate employment effect (see Section IA1.9 in the internet appendix). We find that, with an increase in the federal minimum wage, some of the establishments located in bounded states may not be able to absorb the increase in wage costs. As a result, there is a decline in aggregate employment in the affected states, especially among industries sensitive to the minimum wage and industries located in low-income areas.

Overall, we find that establishments located in bounded states experience a decline in their credit scores around the time of federal minimum wage increases, and they are more likely to exit in the following year. We also find that, in states bounded by the federal minimum wage, there is an increase in exits and a decline in entries for all industries after an increase in the minimum wage, including in those sensitive to the minimum wage.

5. Conclusion

In this paper, we analyze the impact of one-size-fits-all federal minimum wage increases on the financial health of affected small businesses. We use a new measure of a small business's financial health based on the business's payment speed to suppliers and vendors (i.e., its establishment-level PAYDEX score). Using intertemporal variation in whether a state's minimum wage is bounded by the federal minimum wage, and using credit score data for approximately 15.2 million establishments for 1989–2013, we find that increases in labor costs caused by a higher federal minimum wage lead to delayed payments to trade creditors, lower business credit scores, and, ultimately, worse financial health of small businesses in the affected states. Small, young, labor-intensive, and minimum-wage-sensitive establishments operating in competitive and low-income areas that were located in these bounded states experienced higher financial stress, eventually leading to a higher rate of exit.

Wages comprise a significant portion of operating costs of small businesses. Our results suggest that some of the affected small businesses may not have the flexibility to immediately adjust their capital-to-labor ratio or pass on the increased costs to their customers. Overall, our results document the unintended negative effect of one-size-fits-all federal minimum wage increases on the financial health of some small establishments located in states where effective minimum wages are equal to the federal rate.

Acknowledgments

The authors thank Ajay Agrawal, Ashwini Agrawal (discussant), Tania Babina, Emek Basker, Andrew Ellul, Bart Hamilton (discussant), Shane Johnson (discussant), Sari Kerr, Filippo Mezzanotti, Ramana Nanda, Paige Ouimet, Roberto Pinto (discussant), Antoinette Schoar, Elena Simintzi (discussant), Scott Stern, Peter Thompson, Liu Yang, Miao (Ben) Zhang (discussant), conference participants at the 2019 NBER Summer Institute, 2019 Labor and Finance Conference, 2019 SFS Cavalcade, 2019 European Finance Association Conference, 2019 American Finance Association Conference, 2019 Midwest Finance Association Conference, 2018 Colorado Finance Summit, 2018 UT Dallas Finance Conference, and 2018 Workshop on Workforce Development Initiatives, and seminar participants at Deakin University, Georgia Tech, Melbourne Business School and the University of New South Wales for comments that helped to improve the paper. The contents of this publication are solely the responsibility of the authors.

Endnotes

[1] According to the Federal Reserve Bank (2019), 28% of the surveyed firms responded that they delay payment to vendors to address any financial challenge.

[2] One limitation of our study is that we cannot completely distinguish whether a higher minimum wage leads to lower free cash flows, and thus lower credit scores, or whether a higher minimum wage leads to a reduction in credit scores and thus a reduction in free cash flows.

[3] Also, two states have rates below the federal rate, and five states have no state minimum wage requirement. Since 1981, seven federal changes occurred: during 1990–1991, 1996–1997, and 2007–2009. Under the provisions of the Fair Labor Standards Act (FLSA), employers have to pay workers the highest minimum wage prescribed by federal, state, and local law.

[4] An average firm in our sample (where a Paydex score is available) employs about 11 employees (with accompanying sales revenue of $1.4 million). A one-dollar wage increase would translate to additional labor costs of $1 \times 11 \times 8 \times 261 = \$22{,}968$ (assuming eight hours worked per day and 261 working days in a year). In our data, we do not observe establishment profitability. However, using IRS tax-filing data (https://www.irs.gov/statistics/soi-tax-stats-integrated-business-data) in 2013 (the last year in our sample), an average S-corporation with annual business revenue of $1.5 million generates an income of $104,790. Therefore, the increase in labor cost by a dollar can reduce the profit margin by up to 21.9% ($=100 \times \$22{,}968/\$104{,}790$). Consistent with Lopresti and Mumford (2016) and Gopalan et al. (2021), here we assume an increase in wages for all workers.

[5] It is possible that less productive firms may be more negatively affected by the minimum wage increase. Unfortunately, in our data, we do not have any firm productivity measures. However, we find

that even the 30% of the establishments in our sample that make payments to their vendors on time or early begin to delay payments to their vendors after the minimum wage increase. These results, and the significant threshold effects discussed earlier, suggest that efficient firms are also adversely affected.

[6] One possible explanation is a spillover effect on other sectors. Barrot and Nanda (2020) find that accelerated payments by the federal government to small business contractors can have a significant positive impact on employment. However, we are unable to analyze spillover effects due to data availability.

[7] We provide a detailed summary of recent literature documenting the impact of minimum wage on firms in Table IA1 in the internet appendix. For employment, see Katz and Krueger (1992), Card and Krueger (1994), Neumark and Wascher (2000), Card and Krueger (2000), Dube et al. (2010), Giuliano (2013), Sorkin (2015), Meer and West (2016), Cengiz et al. (2019), Clemens and Wither (2019), and Gopalan et al. (2019). For wage dispersion, see Dinardo et al. (1996), Lee (1999), MaCurdy (2015), and Autor et al. (2016). For price levels, see Aaronson (2001) and Aaronson and French (2007). For personal finance, see Aaronson et al. (2012), Tonin (2011), and Agarwal et al. (2019).

[8] Firms with higher pay inequality are also larger, have higher valuations, and have stronger operating performance than other firms (Mueller et al. 2017). Reductions in labor unemployment risks allow firms to increase leverage by mitigating workers' exposure to unemployment risk (Agrawal and Matsa 2013). Government-provided employment guarantee programs can also push firms to reduce their permanent work force (Agarwal et al. 2021).

[9] Since July 24, 2009, the federal government has mandated a nationwide minimum wage of $7.25 per hour. As of January 2019, 29 states and the District of Columbia have minimum wage rates above the federal rate of $7.25 per hour, with rates ranging from $7.50 to $13.25. Two states have minimum wage rates below the federal rate, and five states have no state minimum wage requirement. The remaining 14 states have minimum wage rates equal to the federal rate.

[10] The 1990–1991 increase was enacted on November 17, 1989, with the federal minimum wage increasing in two waves: from $3.35 to $3.80 on April 1, 1990, and to $4.25 on April 1, 1991. For the 1996–1997 change, the law was enacted on August 20, 1996, and the federal minimum wage increased in two waves: from $4.25 to $4.75 on October 1, 1996, and to $5.15 on September 1, 1997. The most recent federal minimum wage change was enacted on May 25, 2007, and rates were increased from $5.15 in three waves to $5.85, $6.66, and $7.25, effective July 24, 2007, July 24, 2008, and July 24, 2009, respectively.

[11] According to the FLSA, an employee can be covered by the law using (a) enterprise coverage that includes businesses with at least two employees and an annual dollar volume of sales or business done of at least $500,000, and (b) individual coverage that can be used when there is no enterprise coverage and employees are protected by the FLSA if their work regularly involves them in commerce between states ("interstate commerce") broadly defined. For some small establishments, NETS data impute the sales and employment information. Crane and Decker (2019) recommend to be cautious while using this data. Therefore, with this caveat, we utilize this variation to identify within-state counterfactual data and report results in Section 4.1.5.

[12] We report absolute effects instead of elasticities because all treated states have the same initial minimum wage (i.e., federal minimum wage).

[13] The two waves of federal government–mandated minimum wage increases occurred during recession years (1990–1991 and 2007–2009). This overlap may confound our analysis if the economies of firms (and thus their financial health) in bounded states are more correlated with the U.S. economy as a whole. In Section IA1.1 of the internet appendix, we test if bounded and unbounded states followed similar business cycles before and after federal minimum wage changes. We find that the amplitude of business cycle swings is slightly more pronounced in unbounded states than in bounded states.

[14] Walls and Associates convert Dun & Bradstreet (D&B) archival establishment data into a time-series database of establishment information.

[15] Specifically, we omit establishments in the following industries: utilities (NAICS 22), finance and insurance (NAICS 52), real estate (NAICS 53), professional services and management of companies (NAICS 54, 55), educational services (NAICS 61), healthcare (NAICS 62), religious organizations (NAICS 813), and public administration (NAICS 92).

[16] Crane and Decker (2019) show that the NETS database imputes employment data for some small establishments. The imputation practice must vary systematically and geographically for the imputation to materially affect our results. However, Crane and Decker (2019) do not find that this imputation varies systematically with geography. They also do not explore imputation practices in the NETS database for credit information, which is our main dependent variable. Figure IA6 in the internet appendix shows that our results are robust to establishments with more than 10 employees, reducing any concerns that the systematic imputation of credit scores may be driving our results. We discuss these results further in Section 4.2.1.

[17] See https://www.dandb.com/glossary/paydex/#1.1 for more information on mapping of credit scores to payment days.

[18] Our data are left-censored for the 1990–1991 shock precluding us from properly estimating it dynamically. This left-censoring is alleviated for the 1996–1997 and 2007–2009 shocks; for example, we consider 1991–1995 as the preperiod for the 1996–1997 shock and 2002–2006 as the preperiod for the 2007–2009 shock.

[19] As an example, in response to the 2007 federal minimum wage increase (from $5.15 to $7.25 through three consecutive annual increases of $0.70), the state of Ohio increased its state minimum wage from $5.15 to $7.30, with annual increments of $1.00, $0.15, and $0.30 in the years 2007, 2008, and 2009, respectively. Similarly, the state of California increased its state minimum wage from $6.75 to $8.00, with annual increments of $0.75 and $0.50 in 2007 and 2008, respectively.

[20] Some minimum wage workers near state borders may commute across the state border to earn a higher minimum wage, potentially causing market wages in the bounded stated to converge toward those of the unbounded state. The potential for these spillover effects implies that our results provide a lower-bound estimate of the effect of minimum wages in our setting.

[21] A caveat of this approach is that sales data from NETS contain imputation errors (see Section 3.1).

[22] Further, we conduct placebo tests where we use placebo thresholds of $400,000 and $600,000, and redefine the $Non\text{-}Exempt_{i,t-1}$ dummy. As expected, the placebo threshold results are statistically insignificant (see Section IA1.6, Table IA12, in the internet appendix). These results suggest that the increase in federal minimum wage similarly impacts firms within the bounded states around the placebo thresholds. Among the bounded states, six states had their own exemption rules (see Table IA13 in the internet appendix). In Table IA14 of the internet appendix, we drop states with their own exemption rule and report the baseline regression estimates from the treatment group. We find that the magnitude for baseline results increases from −0.73 to −0.77. Finally, we utilize the state's variation for bounded states with their own exemption rules. Table

IA15 of the internet appendix reports the regression results. We include state × year fixed effects in addition to establishment fixed effects and firm controls to ensure that different unobservable local economic conditions are not driving our results. We find consistent results. We discuss these results in Section IA1.7 of the internet appendix.

[23] We also conduct subsample analysis and find consistent results (see Internet Appendix, Table IA16). As discussed in Section 3, Crane and Decker (2019) recommend that we should be cautious when using the NETS data set, especially for small firms. As a robustness test, we reestimate Equation (1), where we interact the equation with different labor groups. Figure IA6 in the internet appendix plots the regression coefficient on the triple interaction terms with a 95% confidence interval. We find statistically significant effects for each group.

[24] Further, we partition our sample into two groups divided along the median of sales and age (see columns (1)–(4) of Table IA17 in the internet appendix). Note that we find strong negative results for both small and large establishments, but the negative effect is greater for small establishments. We estimate the regression for each size quintile separately. We find that small firms are more adversely impacted. Only for the top quintile do we find negative and statistically insignificant impact. In the subsample analysis, we find negative and significant impact on firms in the third and fourth quarters, which is not the case when we implement the triple interaction.

[25] Lenders use the PAYDEX score to assess financial risk of potential borrowers. Therefore, we analyze publicly available small business loan data provided by the U.S. Small Business Administration (SBA). We find a decline in bank credit and an increase in loan defaults after a minimum wage increase. However, due to data limitations, we can't completely disentangle demand-side effects from supply-side effects. See Section IA1.8 in the internet appendix for details.

[26] In our data, 13% of the firms have a PAYDEX score between 60 to 70. The mean probability of exit is higher for businesses with PAYDEX scores between 60 and 70, but, with fewer businesses on the common support, these estimates are less precisely estimated. In Table IA20 in the internet appendix, we report results in which we replace the $SameScore_{it}$ (80) dummy with the $SameGroup_{it}$ (80+) dummy. Here, $SameGroup_{it}$ (80+) is a dummy that identifies the establishment-years in which the establishment retains an "80 and above" score both in year $t-1$ and year t. We continue to observe similar threshold effects.

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