Minimum Wage Increases and Eviction Risk*

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Abstract

We extend the debate on the benefits to increasing the minimum wage by examining the impact on expenses associated with shelter, a previously unexplored area. Our analysis uses a unique data set that tracks household rental payments. Increases in state minimum wages significantly reduce the incidence of renters defaulting on their lease contracts by 1.7 percentage points over three months, relative to similar renters who did not experience an increase in the minimum wage. This represents 10.6% fewer monthly defaults. However, this effect slowly decreases over time as landlords react to wage increases by increasing rents.

JEL Classifications: J3, G13, G18, G28, R3, R31, R38

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I. Introduction

The ability to effectively identify responses to household shifts in spending and debt utilization is central to understanding and evaluating key government programs aimed at helping low-income households. One such program, first introduced in Australia and New Zealand in the 1890s, is the minimum wage, which is often the subject of contentious debate.[1] We extend this debate by examining the impact of changes to the minimum wage on expenses associated with housing, a previously unexplored area and a first-order expense for most households. Specifically, we examine how increases to state minimum wages alter the propensity for renters to default on their lease payments. Our analysis recognizes that income instability, particularly among low-income households, is often responsible for incidences of homelessness or dependency on government housing assistance (Desmond, 2016).

To contextualize our analysis, we note that households often face a variety of shocks to income (Campbell and Cocco, 2015). Whether an income shock arising from a change in the minimum wage is positive or negative, permanent or temporary is an open question. Since income volatility is directly associated with the ability of households to pay their rent, changing the minimum wage could have two possible impacts on lease default rates.

First, if an increase in the minimum wage is a realized positive, permanent shock that does not induce a corresponding increase in unemployment or reduction in hours worked (McKinnish, 2017; Draca et al., 2011), then it would reduce the probability of lease default (assuming no contemporaneous increase in housing costs). Cengiz et al. (2019)

[1]See Waltman (2000) for a concise history of the minimum wage. More recently in the U.S., the Raise the Wage Act of 2017 introduced in the U.S. Senate by Senators Sanders (I-Vt.) and Murray (D-Wa.), which would increase the federal minimum wage to $15 per hour by 2024, generated considerable discussion and debate about the effects of increasing the minimum wage and helped insert the issue of the federal minimum wage into the 2020 US presidential election (See S.1242 - Raise the Wage Act, 115th Congress (2017-2018), https://www.congress.gov/bill/115th-congress/senate-bill/1242). The $15 per hour minimum wage proposal was included in president-elect Biden's COVID-19 relief package that was announced on January 14, 2021 (https://buildbackbetter.gov/speeches/remarks-as-prepared-for-delivery-by-president-elect-joe-biden-on-the-american-rescue-plan-and-build-back-better-recovery-plan-in-wilmington-delaware/).

study job changes throughout the wage distribution and conclude that the minimum wage increases in their study (which were between 37% and 57% of the median wage) did not have a significant impact on employment outcomes for individuals in the lowest wage bins. In addition, Dettling and Hsu (2021) provide supporting evidence for the assumption that a change in the minimum wage is a positive income shock by noting that credit card delinquency rates declined following an increase in the minimum wage. In the rental market context, such an outcome may be temporary as the long run equilibrium, which hinges on local market supply and demand elasticities, requires that income shocks are capitalized into rents.[2]

Second, and in contrast to the positive view, a minimum wage hike may increase the probability of unemployment (Neumark et al., 2004; Galan and Puente, 2015), which would represent a negative transitory income shock. As a result, it could increase the probability of a rental default. Yet, evidence for a positive relation between minimum wage changes and unemployment is controversial as Dube et al. (2010) and Hoffman (2014) do not find a causal connection.

Even so, the impact of increasing the minimum wage remains the subject of debate. For example, in two recent studies of a localized change in the minimum wage, Jardim et al. (2018, 2020) find a link in Seattle, Washington between the minimum wage and hours worked. Their analysis shows that the increase in the Seattle minimum wage resulted in a reduction in hours worked among the lowest paid employment sector with an attendant decline in total income. Furthermore, Jardim et al. (2018) demonstrate that any gains associated with the increase in minimum wage accrued to employees with more experience and education. Together, these studies imply possible heterogeneous effects for changes in the minimum wage on lease payments.[3] Thus, the effect of changes

[2]See Glaeser (2008) for a comprehensive discussion of the theoretical models in urban economics that outline the connection between wages and rents.

[3]For example, the effect uncovered in Jardim et al. (2018) suggests that it is possible to have a disemployment effect associated with a minimum wage increase in the youngest and least experienced segment of the workforce (who presumably are not the primary renter) while having little impact on

to the minimum wage is an open empirical question with interpretation of the results often depending on whether it is viewed through the lens of a first or second order expense (housing versus discretionary consumption).

By focusing on the impact of changes in the minimum wage on housing, we bring to bear a policy analysis on a first-order expense that has not been studied. As a result, we contribute to the understanding of how changes in the minimum wage can affect household consumption decisions and our unique data allow us to study this question using a clean identification strategy. We employ a difference-in-differences (DID) estimation strategy where identification rests on the ability to control for property location, renter, and lease-year fixed effects, allowing us to compare the payment pattern for renters before and after an increase in the minimum wage with similar renters in states that did not experience a change in the minimum wage. Thus, we are able to establish a causal connection between the change in a state minimum wage law and the risk associated with rental housing by showing that rental defaults decline while landlords do not raise rent in the short term following an increase in the minimum wage. However, consistent with theoretical models in urban economics, we also show that this effect weakens in the long run as wage increases are capitalized into rents.

We use a unique panel data set comprising the payment performance records on individual renters to identify the change in household rental payment risk (the probability of a late or defaulted payment) following various state-level changes in minimum wage laws. Our data come from RentBureau, a national credit repository that tracked the payment patterns of individual leases in multifamily properties from January 2000 to November 2009. The advantage of this data source is that it is a national database of rental performance covering over 1.8 million individual leases from approximately 2,600 multifamily properties in 41 states. We merge the RentBureau data with Census data

more experienced, older workers who are primary renters. Agarwal et al. (2020) find supporting evidence consistent with a disemployment effect in a study of minimum wage change effects on hospitality industry revenues and occupancy.

based on the property's zip code to obtain neighborhood demographic information. Furthermore, we take advantage of the actual rents recorded from the individual leases to segment the data into high- and low-rent properties, which allows us to test the effect of changes in the minimum wage laws on the households most likely to be affected – those paying the lowest rent. Furthermore, the monthly reporting of rental payments allows us to more precisely isolate the impact of a state minimum wage change.

Our analysis shows the following empirical findings: First, property owners in states that increased the minimum wage experienced on average a 1.74 percentage-point reduction in the three-month renter default rate following the wage increase relative to the average default rate in states that did not increase the minimum wage, which corresponds to 10.6% fewer monthly defaults in relative terms. To put this into context, the results suggest, on average, a 1.1% decrease in the monthly lease default rate following a 1% increase in the minimum wage. Second, we show that renter responses to minimum wage hikes rise over time, which is consistent with the increase in wages having an immediate impact on relaxing renter budget constraints. Third, when segmenting the sample by rent level, which should correlate with renter income, we show that the intensive effect is greatest for households having the lowest rent level, which are the households most likely to be impacted by the change in the minimum wage. Fourth, we show that landlords react by increasing rents beginning approximately three months after the change in minimum wage levels. We note that, on average, rents increase about 7.6% following a minimum wage change, which implies that the increase in rents accounts for 54.6% of the average increase in income. This result is consistent with changes in the minimum wage operating through the demand channel allowing landlords to capitalize the wage increase. We conduct a variety of tests to confirm that our results are robust to concerns about multiple treatments and common treatment trends over time and across states, parallel pre-treatment trends, different measures of renter risk, location differences, housing market cycles, and potential endogeneity.

Our study brings novel insights from a first-order household expense to the literature that examines how consumption and credit use responds to income shocks. Recent studies have looked at how individual consumption decisions respond to changes in adjustable-rate mortgage payments (Di Maggio et al., 2017), sales tax holidays (Agarwal et al., 2017), increases in minimum payments on credit cards (d'Astous and Shore, 2017), tax rebates (Cui, 2017), and unanticipated fiscal policies (Agarwal and Qian, 2014). Since our results indicate that landlords partially capitalize the increase in the minimum wage through higher rents, our analysis provides an upper bound on the ability of low-income households to increase discretionary spending following an increase in the minimum wage.

We also add new evidence to the growing literature examining the economic impact of changes in policies and regulations. For example, Holmes (1998) demonstrates that state level right-to-work laws can impact business formations and locations. Hsu et al. (2018) provide evidence indicating that unemployment insurance helps reduce mortgage defaults and thus stabilizes the housing markets. On the credit supply channel, Melzer (2011) shows how state-level regulations of payday lending can impact the risk of low-income households, while Pence (2006) and Wheelock (2008) demonstrate how state-level laws governing borrower rights can affect mortgage credit availability. Furthermore, Pennington-Cross and Ho (2008) provide evidence that state laws designed to protect borrowers from predatory lending practices lead to a modest increase in credit costs. Since the results show that renter risk declines following an increase in the minimum wage, our study suggests that policies designed to stabilize lower-income households do in fact reduce the riskiness of the target households, which is consistent with the results reported in Dettling and Hsu (2021) regarding the effect of increases in the minimum wage on credit utilization among lower-income adults.[4]

[4]Furthermore, in working papers completed subsequent to ours, Dunphy (2019) and Zhang (2020) confirm our findings based on analysis of eviction fillings, which are the start of legal proceedings by landlords to remove tenants from their leased unit.

Last, our study adds novel insights to the growing literature on decisions regarding shelter. For example, Ambrose and Diop (2021) note how expansion of credit supply can alter the risk of the rental market. Contributing to the understanding of the interactions of macroeconomic policies and rental markets, our study suggests that rising incomes could offset the impact of household movement from renting to ownership. This is consistent with Abdallah and Lastrapes (2013) who provide evidence showing that state-level spending on consumption is sensitive to housing demand shocks.

The rest of our paper proceeds as follows. Section II provides a description of our administrative data set and the state-level minimum wage changes. Section III outlines the empirical method, and Section IV presents the results. Section V discusses rental market effects. Section VI describes the various robustness checks that confirm the primary findings. Section VII concludes.

II. Main Data

Because of limited financial resources, minimum wage earners are more likely to be renters than homeowners. For this group of households, making rent payments on time represents one of their most important obligations although they may face greater challenges making these payments than average households because of tighter budget constraints (Desmond, 2016). Consequently, renters represent an ideal study group when examining the effects of minimum wage increases at the household level. For this reason, we base our empirical analysis on multifamily lease performance data compiled by Experian RentBureau from 2000 to 2009.[5]

The WRDS RentBureau data represents a national snapshot of residential leases collected from property management companies. The data records lease characteristics

[5]We obtained the data from the Wharton Research Data Service (WRDS). Ambrose and Diop (2014, 2021) use the same data to examine the impact of mortgage credit expansion on the rental market and the equilibrium effects of landlord regulations on rental market outcomes, respectively. In addition, Ambrose et al. (2015) use the data to construct a time-series of monthly rents paid by a succession of tenants for each apartment unit in order to create a weighted repeat rent index.

(lease start date, lease termination date, tenant move-in date, tenant move-out date, last transaction date), property locations (city, state, and zip code), and rent payment patterns. To maintain tenant and property owner privacy, location information is limited to the property zip code level and the data contains no personally identifiable information about tenants. However, RentBureau does report the monthly rent amount and the monthly payment history, which denotes whether the tenant paid rent on time. For tenants who did not pay on time, the data reports the type of delinquency, the accrued number of late payments, and any write-off on rent and non-rental expenses due. The initial data set contains roughly 1.84 million leases on 2,648 properties located in 208 MSAs across 41 continental U.S. states. The data represent a record of tenant lease performance during the 2000s as updates are no longer available through WRDS due to changes in data collection methods.

RentBureau reports monthly lease payments in 24-digit vectors, recording historical payments over the last 24 months ending with the month of reporting or the lease maturity month. The reported payment vectors are therefore left censored since records older than 24 months are missing. However, as most residential leases are short term in nature (a year or less), issues associated with the left censoring of tenant payment records are minimized since problem tenants' leases are generally not renewed. Consistent with the commercial purpose of the database serving as a credit repository on tenants, the data contains a wealth of information about tenant rental payment history. The monthly rent payments are coded as P (on-time payment), L (late payment), N (insufficient funds or a bounced check), O (outstanding balance at lease termination), W (write-off of rent at lease termination), or U (write-off of non-rent amount owed at lease termination). We use these lease payment records to construct several lease performance measures.

Information on minimum wage increases came from Aaronson et al. (2012), who compiled the data from January issues of the *Monthly Labor Review* of the Bureau of

Labor Statistics.[6] Because of rental data availability, we restrict our study to state minimum wage increases enacted from 2000 to 2008. Table 1 lists the 25 states that passed minimum wage increases during this period. As Table 1 shows, some of these states, for example, California, experienced multiple treatments (wage increases), generally 12-months apart. In aggregate, these states enacted 76 minimum wage increases at 24 separate dates. Thus, on average, there are 3.04 wage hikes per state. The average wage increase was $0.57, representing roughly 10% of the then-prevailing wage (Table 2). Therefore, over the nine years covered by our study, the average minimum wage earner in treated states earns $1.73 ($0.57*3.04) more per hour, representing a 30.4% wage increase.

For each treatment state/date (i.e. a minimum wage change), we compile the lease payment performance for tenants in that state and in the control group over a fixed time window (three and six months) pre and post the minimum wage increase date.[7] Thus, our empirical design recognizes that we have leases from 24 treatment and control groups corresponding to the 24 separate minimum wage increase dates over the sample period. As a robustness check on this empirical design, we also construct control groups for each treated state using a synthetic controls method and report the results in section IV.C. After excluding leases with missing three-month performance data pre and post the minimum wage event dates, those with missing rent data, and winsorizing the data by eliminating extreme rent values, the final sample consists of 991,000 individual leases executed between 2000 and 2009.[8] The sample highly reflects the geographic distribution of the initial RentBureau data and contains 2,248 properties located in 173 MSAs across

[6]The original data set is listed in Table A2 of the appendix of Aaronson et al. (2012).

[7]The control group comprises 14 states (Alabama, Georgia, Idaho, Indiana, Kansas, Louisiana, Mississippi, Nebraska, Oklahoma, South Carolina, Tennessee, Texas, Utah, and Virginia) that did not change their minimum wage rate during the sample period. By including all states without a minimum wage change in the control group, we avoid arbitrarily selecting specific states to match with each treated state.

[8]We drop leases below 1% ($384) and above 99% ($2,226) of the rent distribution, which allows us to focus on tenants paying market rent and that are more likely to fall in the low to medium portion of the income distribution. However, our findings are the same when we do not winsorize the data.

39 states, 25 of which enacted minimum wage increases.

The summary statistics reported in Table 3 show an increase in lease defaults over time. For example, in Panel A we see that the average three-month (six-month) default rate was 1.19 (1.52) percentage points higher during the three-month (six-month) period following a minimum wage increase.[9] Panels B and C compare the pre- and post-treatment average default rates in control and treated states, respectively. Given the similarity of pre-treatment default rates in the treated and control states, this suggests that minimum wage increases enacted from 2000 to 2008 may have led to fewer lease defaults compared to states that did not increase their minimum wage.

III. Methodology

We analyze the effect of state minimum-wage increases on renters' payment performance using a difference-in-differences regression methodology in a manner similar to Cengiz et al. (2019). More specifically, we estimate the following DID model of renters' likelihood of lease default pre and post the minimum-wage increase dates:

$$\begin{aligned} Pr(Default_{i,t}) &= \beta_1\, MWI_t + \beta_2\, Post_t + \beta_3 (Post_t \times MWI_t) + \mathbf{X}_i'\, \Lambda \\ &\quad + \mathbf{Z}_{i,t}'\, \Theta + Y_t + \xi_{i,t}. \end{aligned} \tag{1}$$

The dependent variable, $Default_{i,t}$, is a binary variable indicating the default status of lease i during a specified observation window (either three or six months) pre and/or post the month of a minimum wage increase at date t.[10] Our default variable indicates whether a lease was ever in default during the specified time period. In our baseline

[9] The number of default observations is larger than the number of individual leases because most states passed multiple minimum wage increases. The trend in six-month defaults is larger in magnitude because of the longer observation window.

[10] Table 1 shows that states with multiple minimum wage increases generally implement them at least twelve months apart. Consequently, the risk of overlap between successive observations within a state is minimal.

model, we consider a lease to be in default in a given month if its status is not coded as on-time (P) or late (L) in the RentBureau data – we check the sensitivity of our results to this lease default definition by considering a more restrictive case in which default is defined as any lease status other than P. For leases in the treated states, we compile their performance pre and post the states' minimum wage increase dates. For the control states, we compile the tenants' performance pre and post the 24 minimum wage increase dates of the treated states.

The indicator variable, MWI_t, identifies states that passed minimum wage increases at date t with β_1 representing the difference in average lease default rates between treated and control states pre-treatment, that is three or six months before minimum wage increases. $Post_t$ is an indicator variable identifying the post-treatment period with the coefficient, β_2, indicating the average change in default in control states post treatment (that is, in the three or six months following the increase in the minimum wage). The coefficient for the interaction $Post_t \times MWI_t$ captures the average difference in default between treated and control states post treatment. Conditional on DID assumptions being met, a negative β_3 implies that increases in state minimum wages lead, on average, to lower lease defaults and vice versa, ceteris paribus. We present both unconditional estimates and estimates conditioned on lease characteristics ($\mathbf{X}_i$), housing market and macroeconomic variables ($\mathbf{Z}_{i,t}$), time fixed effects (Y_i), and state-clustered standard errors (Bertrand et al., 2004). The last element of Equation (1) represents the error term.

Table 2 summarizes the elements of $\mathbf{X}_i$ and $\mathbf{Z}_{i,t}$. First, we control for contract rent from RentBureau since lease default increases with rent, everything else the same. In order to account for heterogeneity across locations, we collect a variety of information on local housing, labor and demographic characteristics by matching the lease records to metropolitan (MSA) or state variables.[11] More specifically, our model includes local

[11]Table A.1 in the appendix documents the data sources.

(MSA) yearly fair market rent (Market Rent) from Department of Housing and Urban Development (HUD), yearly per capital income from Bureau of Economic Analysis (BEA), and annual unemployment from Bureau of Labor and Statistics (BLS). We control for state rental demand using changes in state renter population derived from Census Bureau data, state rental supply using annual rental building permits issued from Census Bureau, and state affordable housing supply using annual low-income housing tax credit (LIHTC) units from HUD.[12] Our model also includes state annual rental vacancy rates from the US Census Bureau and regional annual inflation rates (CPI) from BLS. Finally, we account for changes in local house prices at the 3-digit Zip code level using the Federal Housing Finance Agency (FHFA) quarterly house price index (HPI).

IV. Empirical Results

In this section, we discuss the baseline DID estimation and results from alternative econometric specifications, investigate the parallel trend assumption required for DID validity, and check the intensive effect of minimum wage increases over time and across renter groups. Table 4 presents the baseline DID estimation results.[13] The specification in column (1) does not control for time fixed effects while column (2) adds lease-year fixed effects.[14] Focusing on our preferred specification that incorporates lease year fixed effects (column 2), the coefficient for *Post* indicates that the average default rate in control states is significantly (1.17 percentage points) higher than treated states following a minimum wage increase.[15] The coefficient for treated states (MWI) is not statisti-

[12]Following Ambrose and Diop (2021), we define renter population as the percentage of population in the 20-34-year age group relative to the state's population. The LIHTC program is run by HUD to provide resources for the supply of affordable housing to low-income households in the U.S. (https://www.huduser.gov/portal/datasets/lihtc.html).

[13]In appendix section A and Table A.2, we explore the differences in average lease defaults using a DID model similar to Equation (1) omitting lease characteristics and macroeconomic factors but including lease-year fixed effects with standard errors clustered at the state level. Consistent with the multivariate analysis, the coefficients of the interaction term $Post \times MWI$ from the various model specifications are negative and significantly different from zero.

[14]Overall, the results confirm the unconditional findings reported in the appendix.

[15]We cluster standard errors at the state level.

cally significant, indicating that prior to treatment (an increase in the minimum wage) the treated states experienced lease defaults at a rate comparable to the control states. Finally, the estimated coefficient for the interaction $Post \times MWI$ (the DID parameter) is negative and statistically significant at the 0.1% level suggesting that treatment group states experienced on average 1.75 percentage points fewer defaults following the minimum wage increase than states in the control group. This amounts to 10.6% fewer average monthly defaults over the first three months after treatment– an economically significant effect.[16] To put these results into perspective, we note that the average minimum wage increase in our sample is $0.57/hour or 10%, and thus a 1% increase in the minimum wage corresponds to 1.1% decrease in the monthly probability of default over the first three months after minimum wage change.

To summarize, Table 4 shows that state-level minimum wage increases are strongly associated with a significant reduction in lease defaults, ceteris paribus. Furthermore, the magnitude of this effect is stronger than that derived from unconditional estimations reported in Table A.2 and the effects of the other explanatory variables included in Equation (1) are plausible, reinforcing the validity of the DID default outcome.[17] For example, contract and market rents are negatively related to default, likely because they proxy for tenant quality and rental market risk, respectively. Per capita income and inflation are positive and significant, consistent with income lagging inflation. As expected, an increase in demand (renter population) leads to fewer defaults. Affordable

[16] 1.75/(3.73+1.75)/3=10.6%, where 3.73% is the average three-month default rate in treated states post treatment reported in Panel C of Table 3.

[17] The sample used in the baseline estimations in Table 4 pools three types of lease performance data: leases for which we observe the pre-treatment period only, those for which we observe the post-treatment period only, and those for which both periods are available. To confirm that the results are not biased by the sample construction, we restrict our sample to tenants observed pre and post treatment. This restriction reduces the sample from 984,376 to 726,332 leases. These estimation results (tabulated in the appendix Table A.3) are in all respects similar to the results derived from the larger sample in Table 4, thus confirming the appropriateness of our original sample. We also check that the results persist in six-month lease default rates (see Table A.4 in the appendix) and again find that minimum wage increases are associated with lower lease defaults in treated states compared with states with no increase in minimum wage.

housing supply also has a similar effect on default as the remaining renter pool becomes less risky. On the other hand, greater rental housing supply appears to have a small positive effect on default. Finally, increases in house prices are negatively related to default, but the effect is small.

We note that, as discussed in footnote 3, the interpretation of our results hinges on the underlying implicit assumption that increases in state minimum wages do not produce a corresponding reduction in hours worked or greater unemployment for rent paying workers. Unfortunately, given the nature of our data, we are unable to test this assumption directly as we do not observe the employment status of the individual renters. However, we view this assumption as plausible since Dube et al. (2010) and Hoffman (2014) do not find a causal connection between unemployment and increases in the minimum wage while Doucouliagos and Stanley (2009) conclude that little to no evidence exists in the literature linking minimum wage hikes with unemployment.

A. Average Treatment Effect

One concern with the difference-in-differences empirical set-up in equation (1) is the use of an indicator variable (MWI_t) to infer the average impact of a change in the minimum wage. While this specification has the advantage of addressing the problem of states having multiple treatments, defining the 'pre' and 'post' periods for the control states that never raised their minimum wage is arbitrary. A traditional solution is to estimate a two-way fixed-effects model (year and state fixed effects) with the real dollar amount of the minimum wage under the common trends assumption that the treatment is homogeneous over time or across groups (e.g. Aaronson et al., 2012).[18] However, De Chaisemartin and d'Haultfoeuille (2020) raise significant concerns about the validity of the two-way fixed-effects model when this assumption is violated. Since our study contains multiple minimum wage changes in some of the 'treated' states and thus raises

[18]We thank the anonymous referee for pointing out this concern and for suggesting the DID_M estimator.

the question of the validity of the common trends assumption, we confirm the insights from our difference-in-differences specification by estimating the effect of dollar changes in the minimum wage using the new DID_M estimator proposed by De Chaisemartin and d'Haultfoeuille (2020).[19]

Table 5 reports the average intensive treatment effect of one dollar increase in minimum wage on the monthly lease defaults aggregated at the state level using the DID_M estimation.[20] Consistent with the results reported in Table 4 (the negative coefficient for the interaction of $Post \times MWI$), we note that the average treatment effect is also negative and statistically significant with a $1 wage increase corresponding to a 0.47 percentage points decrease in default. Figure 1 graphically shows the average intensive treatment effect from the multiple minimum wage increases in dollar amounts on the average monthly lease default rate. In the period prior to a change in the minimum wage, the average effect is not statistically significant, which is consistent with the common trends assumption. However, the average treatment effect post treatment (time=0) is negative and statistically significant, supporting the findings reported in Table 4. Thus, we take comfort in noting that our primary finding that minimum wage increases are strongly associated with lower lease defaults is robust across alternative econometric specifications.

B. Pre and Post Treatment Time Effects

Next, we explore the intensity of the treatment effect over time. For this, we estimate the following variant of Equation (1) comparing monthly lease defaults during the six months pre and post treatment - the variables have the same meaning as in Equation (1) with the binary variables Pre_{t-k} and $Post_{t+k}$ identifying the pre-treatment and

[19]De Chaisemartin and d'Haultfoeuille (2020) demonstrate that the DID_M estimator is robust with the data contain minimum wage changes that vary across time or states.

[20]We use the *did_multiplegt* Stata command developed by De Chaisemartin and d'Haultfoeuille (2020). This estimation uses state monthly panel data from 2000 to 2009, which explains the much smaller number of observations compared to our baseline estimation.

post-treatment months, respectively.

$$\begin{aligned} Pr(Default_{i,t,k}) &= \sum_{k=1}^{6} \beta_0^k \, Pre_{t-k} + \sum_{k=1}^{6} \beta_1^k (Pre_{t-k} \times MWI_t) + \sum_{k=1}^{6} \beta_2^k \, Post_{t+k} \\ &\quad + \sum_{k=1}^{6} \beta_3^k \, (Post_{t+k} \times MWI_t) + \mathbf{X}_i' \, \Lambda + \, \mathbf{Z_{i,t}} \Theta + Y_i + \zeta_{i,t}. \end{aligned} \quad (2)$$

Figure 2 shows the DID coefficient estimates for the fully specified model (Table A.5 in the appendix reports the coefficient estimates). The estimated coefficients of the post-treatment interaction terms (β_3^k) are negative and significant while the pre-treatment coefficients (β_1^k) are statistically insignificant, which confirms the parallel trend assumption. The clear decline in lease default rates following the minimum wage increase is consistent with minimum wage increases having an immediate positive impact on renters' ability to meet their rent payments. Unfortunately, concerns about possible contagion from other wage increases and potential confounding factors limit our ability to extend the post-treatment time analysis beyond six months.

C. Parallel Trends Test

A key condition for validity of the DID framework is the parallel trend requirement. The DID methodology implicitly assumes that the outcome of interest (i.e., lease defaults) trends similarly for the treated and control groups during the pre-treatment period and would have followed the same trend post treatment in the absence of treatment. We feel that this assumption is met based on the following three reasons.

First, our DID model controls for most factors likely to affect household risk and rental default at the MSA or state level. These factors include local rent, income, unemployment, house price, rental supply and demand factors, and regional inflation (see Table 4), which capture most causes of heterogeneity in lease defaults across states, thus allowing us to extract the true value of the treatment effect. Most importantly and consistent with the assumption of parallel trends pre-treatment, the results from the fully specified model in Table 4 indicate no statistically significant difference in average three and six month lease default rates between treated and control states in the period prior to an increase in the minimum wage. Furthermore, while not a statistical test, we note that Figure 2 clearly shows that the individual pre-treatment coefficients (β_1^k) from equation (2) are not statistically different from zero, which is consistent with the parallel trends assumption.

Second, we use a non-parametric framework to examine the behavior of average lease defaults in treated and control states pre and post treatment. To implement this, we compare each treated state with a neighboring untreated state.[21] A potential challenge is the multiplicity of treatment dates in some states. For simplicity, we focus on one treatment date for each treated state and compare its average monthly lease default rate pre and post that treatment date with its matched control state's corresponding default rate. We select the treatment dates listed in Table A.6 to minimize potential contamination from other treatments. Figure 3 shows the outcomes from these parallel trends tests for the monthly average default rates during the 12 months pre and post treatment. Although the trends in average default rates are similar across treated and control states in the pre-treatment period, they diverge in the post treatment period as average defaults in treated states edged lower. Again, this implies satisfaction of the parallel trends assumption critical for validity of the DID framework.

Third, we confirm our results using a synthetic control approach (Abadie and Gardeazabal, 2003; Abadie et al., 2010). The synthetic control method creates a counterfactual state control group for each treated state such that the average monthly lease default rate in the synthetic control state trends similarly to that of the treated state during the pre-treatment period. We then observe differences in the average monthly default rates in the post-treatment period between the treated states and the control synthetic

[21] Table A.6 in the appendix lists treated states and matched control states. When there is no neighboring untreated state, we select a non-neighboring, but relatively similar untreated state, or in rare cases a neighboring treated state with no overlapping treatment period.

states. Unlike our main lease-level analysis, the synthetic control estimation is at the state level using average monthly default rates. This results in 2,384 state-month observations compared to the 984,376 lease level observations in our main sample used in Table 4. Table 6 reports the results and shows that the default rate is significantly lower in treated states after treatment in a random effect estimation framework, which is confirmed when adding fixed effects.[22]

Thus, the results using the synthetic control method, the non-parametric framework, the inclusion of control variables in our primary specification, and the lack of statistical significance for the estimated coefficients on the pre-treatment variables assuage concerns regarding violation of the parallel trends assumption.

D. Effect by Rent Groups

MaCurdy (2015), Dettling and Hsu (2021), and Cengiz et al. (2019) show that increases in minimum wages are more consequential for low-income earners. Similarly, we expect minimum wage increases to differentially affect rental payments for various income groups, with lower-income earners likely seeing the largest effect. Normally, the intensive effect of minimum wage increases on lease default should be negatively related to household income. Unfortunately, we do not directly observe income in the RentBureau data.[23] But since the contract rent should be strongly correlated with household income, we use it as a proxy for income.

To test the income effect, we classify the leases by state and year into rent level

[22]The coefficient estimates in Table 6 are not directly comparable to the estimated coefficients in Table 4 because the dependent variable in the synthetic model is the average monthly default rate (bounded between 0 and 1), whereas our main model is a linear probability model at the lease level where default is a 0/1 variable. As expected, the difference in default pre-treatment (the coefficient for MWI) is insignificant since this method creates a synthetic group matching the treated group pre-treatment. The estimated coefficient of $Post \times MWI$ in column (2) indicates that minimum wage increases from 2000 to 2008 decreased monthly lease default rates by 11.4% (0.002 divided the average monthly default rate of 1.76% for our sample), which is similar to the estimated average effect of 10.6% from Table 4. These results persist after controlling for heteroskedasticity by using White standard errors, although the statistical significance drops due to the decline in sample size.

[23]Because of privacy concerns, RentBureau only reported limited information on renters beyond lease characteristics.

quintiles.[24] We then estimate equation (1) separately for each quintile group. This specification is similar in spirit to the analysis of Cengiz et al. (2019) who estimated the effect of minimum wage increases across various wage bins. Table 7 reports the results. Given the sample restrictions required for this analysis, we report the full-sample estimation in column (1) for the purposes of comparison and the quintile results in columns (2) to (6). We observe that the estimated coefficients for the $Post \times MWI$ interaction are negative and statistically significant across rent quintiles. Furthermore, we note that the magnitude of the estimates is largest for the lowest rent quintile and lowest for the top quintile – the difference in default post treatment between quintiles 1 and 5 is statistically significant. This confirms that renters paying the lowest rent experienced a significantly greater effect following an increase in the minimum wage. In other words, the intensive effect of a minimum wage increase is lowest for the high-rent group, as expected. Assuming that rent level is correlated with income, the results are consistent with lower-income households benefiting more from minimum wage increases than higher income households.[25]

V. Rental Market Effects

A. Landlord Response

Changes in the minimum wage do not happen in isolation. It is possible that a minimum wage increase may alter the prevailing rental market equilibrium. Everything else equal, higher wages may lead to higher rents due to the resulting increase in demand. Furthermore, depending on the structure of the rental market, landlords may try to capitalize on this opportunity by raising rents as Yamagishi (2021) documents for the

[24] We first winsorize the data to remove outliers such that the contract rent ranges from \$384 to \$2,226 (Table 2) and limit our analysis to states with at least 300 leases per year over the period from 2004 to 2009 to have sufficient observations in each quintile.

[25] Although the magnitude of the effect of a minimum wage increase in the lowest rent quintile is consistent with expectations, one concern is that we also observe a negative, albeit smaller, effect in the highest rent quintile. In Appendix B we explore the implications of this concern.

Japanese rental market. As long as any resulting increase in rents does not overwhelm the direct effect of the wage increase on lease performance, the net effect on rent default should be negative.

In this section, we test the effect of minimum wage increases on rents and thereby pin down the net benefit households derive from wage increases. To do so, we estimate the following model of rents:

$$\begin{aligned} R_{i,t,k} &= \beta_1(Pre_t \times MWI_t) + \sum_{k=1}^{6} \beta_2^k\, Post_{t+k} + \sum_{k=1}^{6} \beta_3^k (Post_{t+k} \times MWI_t) \\ &\quad + \mathbf{X}_i' \Lambda + \mathbf{Z_{i,t}}' \Theta + \zeta_i, \end{aligned} \tag{3}$$

where $R_{i,t}$ are rents on new leases in the month before and the six months after a minimum wage increase at date t. The superscript k indicates the months prior to and following the minimum wage change at date t, and the other variables have the same meaning as in Equation (1).[26]

Figure 4 plots the estimated coefficients for the DID parameters from the regression in column (2) of Table A.7.[27] Again, focusing on the interaction terms ($Post_{t+k} \times MWI_t$), we note that the estimated coefficients are positive and statistically significant starting three months after the minimum wage increase. Thus, the results suggest that landlords capitalize the permanent nature of the minimum wage shocks into rents. Based on the regression in column (2') of Table A.7, the increase in rent at the end of month post treatment is approximately \$53.9 higher than the average rent in the six months prior to the change in the minimum wage, which represents an increase of 7.3%.[28] To put this in perspective, the average of the 76 state-level minimum wage changes was \$0.57/hour

[26] We estimate equation (3) using two specifications of Pre_t to equal the average default rate one-month before the treatment or the average default rate over the six-month period before treatment.

[27] Table A.7 in the appendix reports post treatment rent adjustments relative to the first month before treatment in columns (1) and (2), and relative to the average rent in the six months before treatment in columns (1') and (2').

[28] From column (2') of Table A.7 we obtain 0.539*1000=\$53.9, which divided by average real rent of \$743 in Table 2 gives 7.3%.

or 10% and ranged from 1.6% to 35% with the average increase in rents taking up 54.6% of the average income increase.[29]

B. Tenants' Housing Decision

We have established that minimum wage increases lead to higher rents, as compared with prevailing rents in states with no wage increase. Next, we consider tenants' housing consumption decisions after wage increases. More specifically, we explore whether wage increases alter household mobility. To that effect, we estimate the likelihood of a household moving to a different rental unit as a result of a minimum wage increase using the following DID linear probability model:

$$\begin{aligned} Pr(Move_{i,t}) &= \beta_1\, Post_t + \beta_2\, MWI_t + \beta_3\,(Post_t \times MWI_t) + \mathbf{X}_i'\Lambda \,+\, \mathbf{Z_{i,t}}'\Theta \\ &\quad + Y_t + \omega_i, \end{aligned} \tag{4}$$

where $Move_{i,t}$ is a dummy variable indicating whether tenant i moved following a minimum wage increase at date t, and $\mathbf{X}_i$ and $\mathbf{Z_{i,t}}$ have the same meanings as in Equation (1). Y_t are minimum wage change year fixed effects.

We estimate this model for tenants whose leases expired within three months following a minimum wage increase and who entered into new leases between three to nine months after the wage change. Therefore, we only examine tenants with repeat leases, which results in our sample dropping to 15,046 leases. The DID coefficient estimate in column (1) of Table 8 shows that tenants in states that enacted a minimum wage increase are 8.8 percentage points more likely to move to a different unit after the wage change.

Next, we explore the likelihood of moving for different rent groups. For this analysis, we divide the tenants into two groups and identify the low-rent group as those whose

[29]The average monthly income increase is \$0.57*40*52/12=\$98.8. Thus, the ratio of rent to income is \$53.9/\$98.8=54.6%.

rents are less than or equal to their MSA's respective fair market rent. Column (2) shows that tenants in the low rent group are less likely to move following a minimum wage increase. This result makes sense because low-income renters are more likely to miss rent payments, thus more exposed to involuntary moves due to evictions. Unfortunately, our data is not rich enough to allow us to further elaborate on tenant housing decisions.

VI. Robustness Checks

A. Alternative Default Measure

So far, we have assumed a lease to be in default if its status in the RentBureau data is not coded as P (i.e., rent paid on time) or L (i.e., late rent payment). To confirm that this default measure is not driving the results, we also use an alternative, less restrictive default measure considering just non-timely rent payments as the default event. As expected, this alternative default measure leads to more defaults (see resulting average default rates in appendix Table A.8). The average three-month default rate pre (post) minimum wage increase based on this new default measure is 15.2% (16.8%) in Panel A of Table A.8, compared with 3.2% (4.4%) in Panel A of Table 3.[30] Furthermore, unconditional and multivariate estimation results in Tables A.9 and A.10 in the appendix confirm the previous DID default results. However, post-treatment effects based on this alternative default measure are larger, due to the higher incidence of default resulting from this measure.

B. Employment Location

A key assumption of our analysis is that residency and employment location are the same. Although this assumption is realistic since our analysis is at the state level, it would be problematic in MSAs that span multiple states, such as Charlotte-Concord-Gastonia MSA (NC and SC) and Cincinnati MSA (OH, KY, and IN), which include

[30] We find the same result when we use six-month defaults.

treated and untreated areas.[31] In these MSAs, it is possible that some people commute to a neighboring state for work. Furthermore, a minimum wage increase in one state may cause neighboring state residents to seek work in that state, which would muddle the DID identification strategy since some people may choose treatment. We note that 20 of the 173 MSAs making our sample span across several states. Thus, we re-estimate the model excluding leases from properties located in these 20 cross-state border MSAs. We summarize these results in the appendix in column (1) of Table A.11. Despite the smaller sample size, the previous results hold. In treated states, three-month lease defaults were 1.8 percentage points lower following minimum wage increases, compared with a mean pre-treatment default rate of 3.55% in those states.

C. Impact of the 2007-08 Crisis

Our study period spanning 2000 to 2008 almost coincides with the recent housing market boom that saw a substantial surge in homeownership that adversely affected the riskiness of the rental market as documented by Ambrose and Diop (2014). Table 1 shows that the wage increases are not uniformly distributed over that period. The distribution is negatively skewed with 2007 and 2008, probably the most critical years of that period, accounting for 50% of wage increases (38 out of 76). Even though it is unclear how the increase in rental market risk during that period differentially affected treated and control areas, prudence dictates that we check if our findings hold when we exclude the later years.[32] Even though the sample size drops considerably after excluding 2007 and 2008 minimum wage increases, the estimation results reported in the column (2) of the appendix Table A.11 confirm that the previous findings are not confined to those years. The DID three-month lease default rate estimate is -0.94 percentage points,

[31] Of the 382 MSAs listed on Bureau of Economic Analysis website, 47 span over two or more states (https://www.bea.gov/regional/docs/msalist.cfm).

[32] Although the migration of lower risk tenants to homeownership described by Ambrose and Diop (2014) should lead to increase default in the rental market, ceteris paribus. This transition of some renters to homeownership should also lower rents, hence leading to fewer defaults. Thus, the net effect is unclear. Also, it is unclear why this would affect treated and control states differently.

compared with a mean pre-treatment default rate of 3.55% in treated states.

D. Local Regulations

Local rent control and other municipal regulations, such as city-level minimum or living wage requirements, are also likely to affect local rental markets. We acknowledge that there may be heterogeneity in local responses to state minimum wage increases, and municipal wage regulations are likely to exist and may even be significant. However, our focus is on cross-state, rather than within-state, variations in lease defaults in response to state minimum wage changes. State minimum wage requirements are generally less aggressive than most cities' living wages, but tend to be more binding on employers.[33]

As far as rent control regulations are concerned, areas with these regulations should normally experience lower payment defaults as they make rents more affordable for generally riskier tenants. Consequently, these regulations should normally bias against finding significant differences in lease defaults.[34] Nonetheless, we formally control for rent control regulations by excluding from our sample states with rent control cities, namely, California, Maryland, and New York, and find that our main results are unchanged (appendix Table A.11, column 3). We also formally test for the effects of minimum wage increases on lease defaults and rent levels in the presence of rent control regulations (appendix Table A.12).[35] Due to data limitations, we limit this analysis to the 2007 and 2008 wage increases in California, the largest rent control state in our sample and, using a difference-in-difference-in-differences specification, compare leases in San Francisco and Los Angeles (both rent controlled cities) to leases from other areas in California that are

[33]For example, the City of Madison, WI, steadily increased mandated living wages from $9.01 in 2001 to $13.01 in 2018, while the state minimum wage remained at the federal level of $7.25 since 2009. However, Madison's living wage only applies to persons directly employed by the city or employed city contractors or recipient of city financial assistance. As a side note, Madison's living wage was recently nullified by 2017 Wis. Act 327, which became effective April 18, 2018 (Source: https://www.cityofmadison.com/finance/wage/factsheet.cfm). See Adams and Neumark (2005) for a discussion and evidence on the employment effects of living wage laws.

[34]Low-Income Tax Credit (LIHTC) developments should also result in lowering lease defaults. We try to limit LIHTC leases from our sample by excluding leases with rent below $384 per month.

[35]We thank an anonymous referee for this suggestion.

not subject to rent control.[36] In column (2) in Table A.12 in the appendix, we see that the estimated coefficient for the triple interaction term is not statistically significant, which indicates that the presence of rent control regulations does not alter the probability of tenant default.[37] Finally, we test the ability of landlords to capitalize wage increases into rents in the presence of rent regulations. Since rent control regulations constrain landlord ability to raise rents, the insignificant coefficient on the difference-in-differences variable in the rent model (column (4) of Table A.12) is expected and indicates that landlords in rent controlled areas do not increase rents, on average, following an increase in the minimum wage.[38]

VII. Conclusions

We estimate the impact of changes in state-level minimum wage laws on renter lease payment performance. Our analysis is based on a difference-in-differences regression method that employs property, renter, and lease-year fixed effects allowing us to compare the payment pattern for renters before and after an increase in the state minimum wage relative with similar renters in states that did not change the minimum wage.

We find four key results. First, we find a significant decline in defaults following an increase in the minimum wage relative to states that did not increase the minimum wage. Second, we report that renter responsiveness to changes in the minimum wage increases over time. This is consistent with the theory that increases in the minimum wage have an immediate impact by relaxing renter budget constraints. Third, our analysis indicates

[36] To provide a baseline for comparison, we report in Table A.12, column (1) the estimated coefficients for the lease default DID model using only leases from California. Consistent with the results reported in section IV, we note the negative and statistically significant parameter for the DID interaction term indicating a decline in average default rates post minimum wage increase.

[37] For comparison, we also report the estimation of the default model for just the leases in San Francisco and Los Angeles (column 3) as the treated group and again find a negative impact of a wage increase on lease defaults.

[38] Due to the smaller sample size, we are unable to estimate the full specification of equation 3 with post minimum wage-month interaction terms. Thus, we simply test the average rent increase over the six-months following the minimum wage increase.

that renters most likely in the lower-income segment of the population (those in the lower rent levels) experience the greatest reduction in rental default rates following an increase in the minimum wage. Finally, we find that landlords partially capitalize minimum wage increases into rents starting approximately three months following the law change.

Our study provides evidence on how households respond to income shocks with respect to shelter, a first-order expense. Furthermore, since landlords partially capitalize the income shock into future rents, our analysis suggests an upper bound on the ability of lower-income households to increase discretionary spending following an increase in the minimum wage. Finally, the results provide evidence that efforts to stabilize lower-income households via increases in the minimum wage do in fact reduce the riskiness of low-income households.

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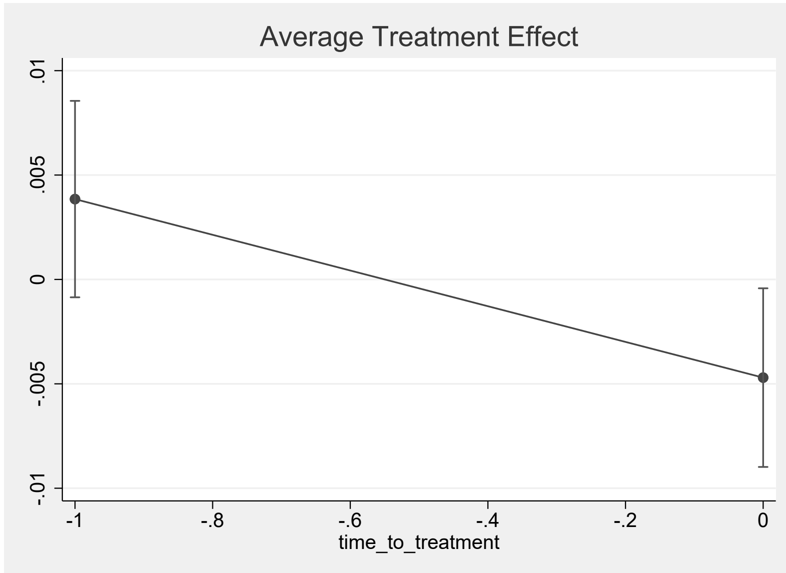

Figure 1: Average Treatment Effect of State-level Minimum Wage Changes
Note: This graph shows the average intensive treatment effect from the multiple minimum wage increases in dollar on average monthly lease default. It is estimated following the methodology developed by De Chaisemartin and d'Haultfoeuille (2020). Table 5 shows the regression result.

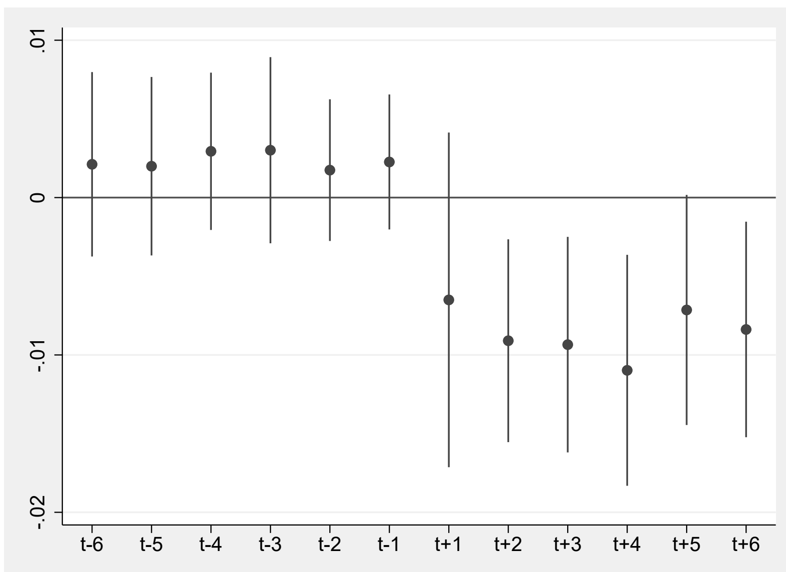

Figure 2: Effect of Minimum Wage on Lease Default Over Time
Note: This graph shows the DID regression estimates and 95% confidence intervals of monthly default in treated (minimum wage increase) states over the six months pre and post minimum wage increases as modeled in Equation (2). The corresponding regression results are reported in Table A.5, column (2).

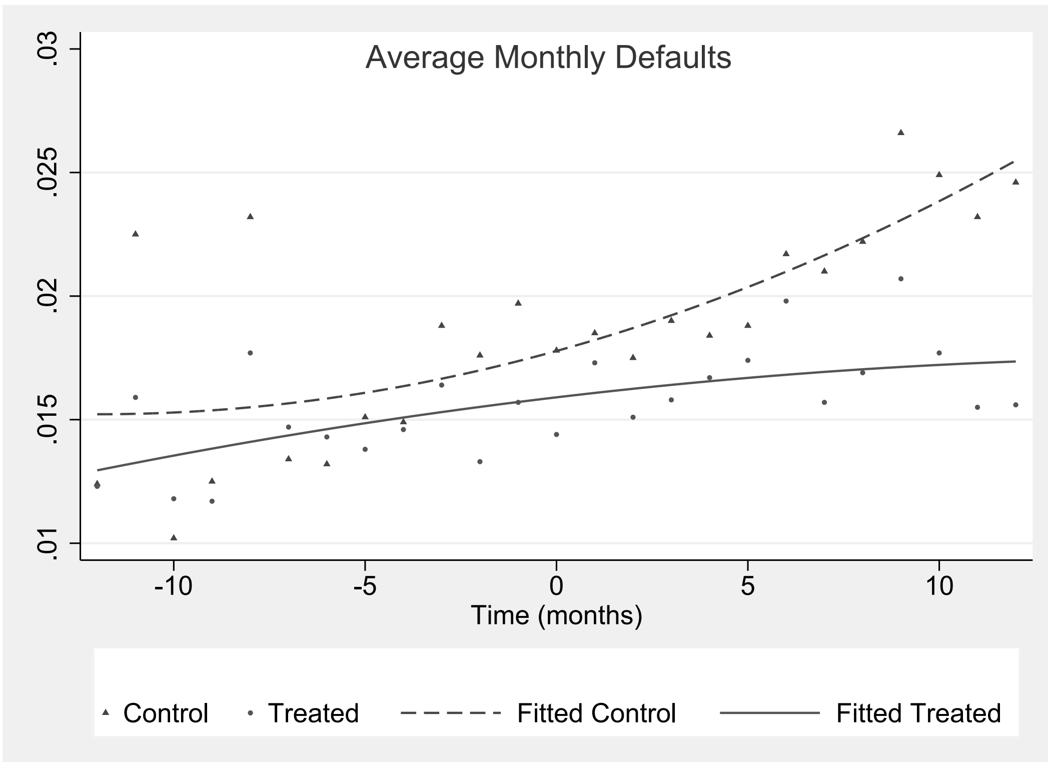

Figure 3: Average Monthly Defaults

Note: This graph shows the polynomial trend lines of average monthly default rates pre- and post treatment based on the nearest neighbor state as the control group. Each treated state is compared with a neighboring untreated state. Table A.6 in the on-line appendix lists treated states and matched control states. When there is no neighboring untreated state, we select a non-neighboring, but relatively similar untreated state, or in rare cases a neighboring treated state with no overlapping treatment period. For simplicity, we focus on one treatment date for each treated state and compare its average monthly lease default rate pre and post that treatment date with its matched control state's corresponding default rate. We select the treatment dates listed in Table A.6 to minimize potential contamination from other treatments.

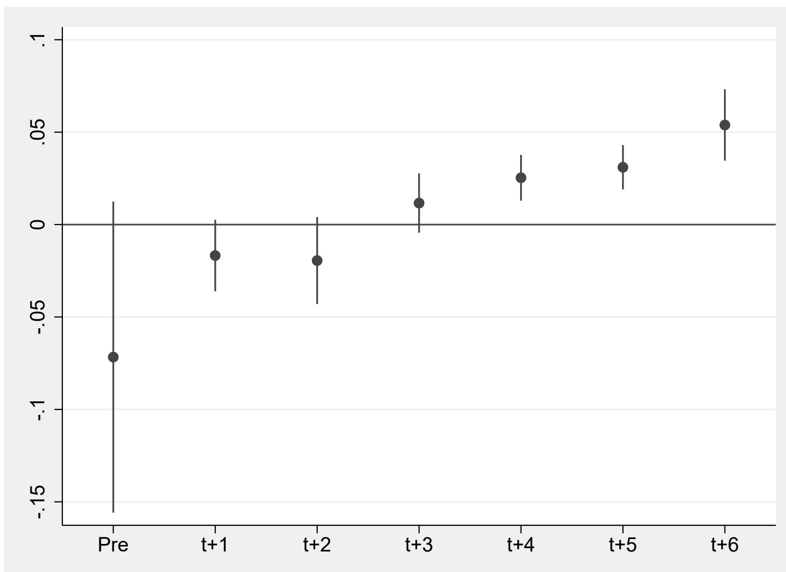

Figure 4: Effect of Minimum Wage Change on Rent Over Time
Note: This graph shows the DID regression estimates and 95% confidence intervals of rent (in '000s) in treated states over the six months following minimum wage increases as modeled in Equation (3). The corresponding regression results are reported in Table A.7, column (2).

Table 1: State Minimum Wage Increases from 2000 to 2008

State	*Date*	*Increase* ($)	*New Minimum Wage* ($)	*State*	*Date*	*Increase* ($)	*New Minimum Wage* ($)
Arizona	Jan-07	1.60	6.75	Massachusetts	Jan-01	0.75	6.75
Arizona	Jan-08	0.15	6.90	Massachusetts	Jan-07	0.75	7.50
Arkansas	Oct-06	1.10	6.25	Massachusetts	Jan-08	0.50	8.00
California	Jan-01	0.50	6.25	Michigan	Oct-06	1.80	6.95
California	Jan-02	0.50	6.75	Michigan	Jul-07	0.20	7.15
California	Jan-07	0.75	7.25	Michigan	Jul-08	0.25	7.40
California	Jan-08	0.50	8.00	Minnesota	Aug-05	1.00	6.15
Colorado	Jan-07	1.70	6.85	Missouri	Jan-07	1.35	6.50
Colorado	Jan-08	0.17	7.02	Missouri	Jan-08	0.15	6.65
Connecticut	Jan-00	0.50	6.15	Nevada	Nov-06	1.00	6.15
Connecticut	Jan-01	0.25	6.40	Nevada	Jan-07	0.18	6.33
Connecticut	Jan-02	0.30	6.70	New Hampshire	Sep-07	1.35	6.50
Connecticut	Jan-03	0.20	6.90	New Hampshire	Sep-08	0.75	7.25
Connecticut	Jan-04	0.20	7.10	New York	Jan-05	0.85	6.00
Connecticut	Jan-06	0.30	7.40	New York	Jan-06	0.75	6.75
Connecticut	Jan-07	0.25	7.65	New York	Jan-07	0.40	7.15
Delaware	Oct-00	0.50	6.15	North Carolina	Jan-07	1.00	6.15
Delaware	Jan-07	0.50	6.65	Ohio	Jan-07	1.70	6.85
Delaware	Jan-08	0.50	7.15	Ohio	Jan-08	0.15	7.00
Florida	Jan-06	1.25	6.40	Oregon	Jan-03	0.40	6.90
Florida	Jan-07	0.27	6.67	Oregon	Jan-04	0.15	7.05
Florida	Jan-08	0.12	6.79	Oregon	Jan-05	0.20	7.25
Illinois	Jan-04	0.35	5.50	Oregon	Jan-06	0.25	7.50
Illinois	Jan-05	1.00	6.50	Oregon	Jan-07	0.30	7.80
Illinois	Jan-07	1.00	7.50	Oregon	Jan-08	0.15	7.95
Illinois	Jan-08	0.25	7.75	Pennslyvania	Jan-07	1.10	6.25
Iowa	Apr-07	1.05	6.20	Pennslyvania	Jul-07	0.90	7.15
Iowa	Jan-08	1.05	7.25	Washington	Jan-00	0.80	6.50
Kentucky	Jun-07	0.70	5.85	Washington	Jan-01	0.22	6.72
Maine	Jan-02	0.60	5.75	Washington	Jan-02	0.18	6.90
Maine	Jan-03	0.50	6.25	Washington	Jan-03	0.11	7.01
Maine	Jan-05	0.10	6.35	Washington	Jan-04	0.15	7.16
Maine	Jan-06	0.15	6.50	Washington	Jan-05	0.19	7.35
Maine	Oct-06	0.25	6.75	Washington	Jan-06	0.28	7.63
Maine	Oct-07	0.25	7.00	Washington	Jan-07	0.30	7.93
Maine	Oct-08	0.25	7.25	Washington	Jan-08	0.14	8.07
Maryland	Jan-07	1.00	6.15	Wisconsin	Jun-05	0.55	5.70
Massachusetts	Jan-00	0.75	6.00	Wisconsin	Jun-06	0.80	6.50

This table documents the minimum wage increases from 2000 to 2008 passed by 25 of the 39 states represented in the RentBureau lease performance data. These 25 states constitute our initial treatment group, with our initial control group consisting of the remaining 14 states, namely Alabama, Georgia, Idaho, Indiana, Kansas, Louisiana, Mississippi, Nebraska, Oklahoma, South Carolina, Tennessee, Texas, Utah, and Virginia.

Table 2: Variable Summary Statistics

Variable Name	*Description*	*# Obs*	*Mean*	*SD*	*Min*	*Max*
Minimum Wage Increase ($)	Dollar hourly wage increase	76	0.57	0.44	0.10	1.80
Minimum Wage Increase (%)	Percents hourly wage increase	76	10.00	8.70	1.60	35.00
Rent ($000s)	Contract (lease) rent (2000 dollars)	984,376	0.743	0.289	0.303	2.119
Market Rent ($000s)	Lagged MSA fair market rent, yearly (2000 dollars)	839	0.551	0.141	0.372	1.355
PC Income	Log lagged MSA per capita annual income (2000 dollars)	839	10.23	0.18	9.53	10.99
Inflation	Lagged annual CPI region	33	189.08	15.14	163.60	215.20
Unemployment	Lagged MSA annual unemployment rate (%)	839	4.89	1.33	2.22	10.59
Change Renter Population	Lagged annual change state renter population – 20 - 30 years old (%)	179	0.67	1.04	-6.38	3.38
HPI	Lagged FHFA quarterly house price index (3-digit zip code)	2,604	184.16	50.00	110.89	374.19
Rental Supply	Log lagged number of building permits issued in state annually	179	8.91	0.97	6.68	11.27
Affordable Housing Supply	Log lagged number of LIHTC units built in state annually	179	7.54	1.01	2.94	9.80
Vacancy	Lagged annual state vacancy rate	179	0.110	0.029	0.042	0.181

This table presents summary descriptive statistics of the variables used in this study. *SD* stands for standard deviation. The data sources are in Table A.1.

Table 3: Summary Statistics of Three-Month and Six-Month Lease Defaults

	3-Month Defaults				*6-Month Defaults*			
	# Leases	*# Obs*	*Mean*	*SD*	*# Leases*	*# Obs*	*Mean*	*SD*
Panel A: Full Sample								
Pre MWI	830,319	2,678,930	0.0324	0.1770	671,117	2,017,577	0.0532	0.2244
Post MWI	934,065	2,944,237	0.0443	0.2058	819,884	2,396,450	0.0684	0.2525
Difference & t-statistic			*-0.0119*				*-0.0152*	
			(73.33)				*(66.40)*	
Panel B: Control Group								
Pre MWI	564,632	2,365,570	0.0320	0.1759	471,749	1,791,783	0.0523	0.2227
Post MWI	598,237	2,554,093	0.0454	0.2081	534,900	2,076,022	0.0692	0.2538
Difference & t-statistic			*-0.0134*				*-0.0169*	
			(76.92)				*(68.93)*	
Panel C: Treated Group								
Pre MWI	265,687	313,360	0.0355	0.1851	199,368	225,794	0.0601	0.2377
Post MWI	335,828	390,144	0.0373	0.1894	284,984	320,428	0.0635	0.2438
Difference & t-statistic			*-0.0017*				*-0.0034*	
			(3.87)				*(5.12)*	

This table reports summary statistics of 3-month and 6-month lease defaults pre and post state minimum wage increases passed from 2000 to 2008. The treated group consists of the 25 states that enacted minimum wage increases during that period and the control group consists of the 14 other states in the RentBureau data that did not pass a minimum wage increase (See Table 1). Our sample consist of 990,785 individual leases signed between 2000 and 2008. We observe the performance of each lease for up to 24 months. We consider a lease current (not in default) in a given month if RentBureau records its status as either P (on-time payment) or L (late payment). Three-month (six-month) defaults indicate whether leases defaulted in any one month during the three (six) months pre and post a minimum wage increase.

Table 4: Multivariate DID Estimation of the Effect of Minimum Wage Increases on Three-Month Lease Defaults

	(1) *3-mo. Default*	*(2)* *3-mo. Default*
Post	0.0128***	0.0117***
	(0.0011)	(0.0010)
MWI	0.0106**	0.0080
	(0.0036)	(0.0046)
Post x MWI	-0.0136***	-0.0175***
	(0.0024)	(0.0026)
Rent (Lease)	-0.0097**	-0.0064**
	(0.0031)	(0.0023)
Market Rent	-0.0053	-0.0208*
	(0.0109)	(0.0094)
PC Income	0.0229*	0.0262*
	(0.0103)	(0.0097)
Inflation	0.0001*	0.0014***
	(0.0001)	(0.0002)
Unemployment	0.0010	-0.0001
	(0.0012)	(0.0013)
Change Renter Population	-0.0007	-0.0014
	(0.0008)	(0.0008)
HPI	-0.0001**	-0.0001**
	(0.0000)	(0.0000)
Rental Supply	0.0023	0.0046
	(0.0018)	(0.0028)
Affordable Housing Supply	-0.0032	-0.0051
	(0.0022)	(0.0029)
Vacancy	0.0236	0.0604
	(0.0486)	(0.0667)
Constant	Yes	Yes
Lease-Year FE	No	Yes
State-Clustered SE	Yes	Yes
# Leases	*984,376*	*984,376*
Adjusted R-squared	*0.002*	*0.005*

This table reports multivariate DID results of lease defaults in treated and control states pre and post minimum wage increases. The dependent variable is a 3-month lease default indicator tracking whether tenants have missed a payment during the 3-month period pre and post a minimum increase. A lease is current (not in default) in a given month if RentBureau records its status as either P (on-time payment) or L (late payment). For leases in treated states, we track their performance pre and post each minimum wage increase in that state. For leases in the control group, we track lease performance pre and post the minimum wage increase dates in each treated state. *Post* stands for the post treatment period and *MWI* indicates treated states. The figures in parentheses are state-clustered standard errors. One, two, or three stars indicate statistical significance at 5%, 1%, or 0.1%, respectively.

Table 5: DID Estimation of the Average Intensive Treatment Effect of Minimum Wage Increases on Average Monthly Default Estimated Following De Chaisemartin and d'Haultfoeuille (2020)

	(1) *Monthly Default*	*(2)* *Monthly Default*
Average Treatment Effect	-0.0047*	-0.0047*
	(0.0023)	(0.0022)
Control Variables	No	Yes
Clustered SE	Yes	Yes

This table reports the average intensive treatment effect of minimum wage increases in dollar on average monthly lease default estimation using the Stata command *"did_multiplegt"* developed by De Chaisemartin and d'Haultfoeuille (2020). The list of control variables in specification (2) includes average rent, median income (in 2000 dollars), inflation, population, vacancy, and unemployment. The estimation uses state panel data with standard errors, in parentheses, clustered by state. Figure 1 shows this average effect by comparing average monthly default pre and post treatment. One star indicates statistical significance at 5%.

Table 6: DID Analysis of the Effect of Minimum Wage Increases on Default Using the *Synthetic Control Methodology*

	(1) *Monthly Default*	*(2)* *Monthly Default*	*(3)* *Monthly Default*
Post	0.0013**	0.0020***	0.0020**
	(0.0006)	(0.0007)	(0.0009)
MWI	-0.0042***	0.0450	-0.0189
	(0.0011)	(0.0850)	(0.0250)
Post x MWI	-0.0018**	-0.0020***	-0.0020*
	(0.0007)	(0.0007)	(0.0011)
Mediane Income	-0.0160***	0.0052	0.0052
	(0.0041)	(0.0104)	(0.0134)
Inflation	0.0001	-0.0001	-0.0001
	(0.0000)	(0.0001)	(0.0001)
Renter Population	-0.0028*	0.0295	0.0295
	(0.0015)	(0.0385)	(0.0530)
Rental Supply	0.0017*	0.0003	0.0003
	(0.0010)	(0.0016)	(0.0021)
Affordable Housing Supply	0.0020***	0.0004	0.0004
	(0.0006)	(0.0007)	(0.0008)
Vacancy	0.0462***	-0.0330	-0.0330
	(0.0174)	(0.0235)	(0.0261)
HPI	-0.1110***	-0.0824***	-0.0824***
	(0.0158)	(0.0170)	(0.0173)
Constant	Yes	Yes	Yes
MWI Panel FE	No	Yes	Yes
# Observations	*2,384*	*2,384*	*2,384*
R-squared	*0.249*	*0.425*	*0.425*

This table reports multivariate DID results of average lease defaults in treated and a synthetic control group pre and post minimum wage increases. The synthetic control methodology creates a counterfactual control group such that the event of interest (lease default) trends similarly during the pre-treatment period following Abadie et al. (2010). *Post* stands for the post treatment period and *MWI* indicates treated states. Columns (1) and (2) report normal standard errors whereas column (3) reports White robust standard errors. One, two, or three stars on top of coefficient estimates indicate statistical significance of coefficient estimates at 5%, 1%, or 0.1%, respectively.

Table 7: DID Estimation of the Effect of Minimum Wage Increase on Lease Performance (Three-Month Defaults) by Rent Quintiles

	(1) *Full Sample*	(2) *Quintile 1*	(3) *Quintile 2*	(4) *Quintile 3*	(5) *Quintile 4*	(6) *Quintile 5*
Post	0.0112***	0.0147***	0.0131***	0.0107***	0.0094***	0.0082***
	(0.0010)	(0.0018)	(0.0015)	(0.0007)	(0.0010)	(0.0007)
MWI	0.0087	0.0032	0.0086	0.0089	0.0071	0.0120*
	(0.0062)	(0.0084)	(0.0077)	(0.0061)	(0.0047)	(0.0044)
Post x MWI	-0.0207***	-0.0226***	-0.0212***	-0.0217***	-0.0206***	-0.0166***
	(0.0029)	(0.0048)	(0.0037)	(0.0026)	(0.0029)	(0.0026)
Control Variables	Yes	Yes	Yes	Yes	Yes	Yes
Constant	Yes	Yes	Yes	Yes	Yes	Yes
Lease-Year FE	Yes	Yes	Yes	Yes	Yes	Yes
State-Clustered SE	Yes	Yes	Yes	Yes	Yes	Yes
# Leases	*901,063*	*183,410*	*179,674*	*179,941*	*179,454*	*178,584*
Adjusted R-squared	*0.006*	*0.010*	*0.007*	*0.006*	*0.005*	*0.004*

This table reports DID regression results of 3-month lease defaults in treated and control states pre and post minimum wage increases by rent/income group. For each state, we classify annually renters into rent quintile groups according to lease contract rents and estimate our DID model for each rent quintile group separately. To ensure that we have enough observations in each quintile for our analysis, we restrict our sample to 2004 to 2009 and every year to states with at least 300 leases. We report in column (1) our DID results for the full sample and quintile results in columns (2) to (6). *Post* stands for the post treatment period and *MWI* indicates treated states. Our model includes the same control variables as in Table 4. The figures in parentheses are state-clustered standard errors clustered. One, two, or three stars indicate statistical significance at 5%, 1%, or 0.1%, respectively.

Table 8: DID Estimation of Likelihood of Renter Moving after a Minimum Wage Increase

	(1) *Pr(Move)*	*(2)* *Pr(Move)*
Post	0.2443***	0.2447***
	(0.0493)	(0.0111)
MWI	-0.0491	-0.0478*
	(0.0388)	(0.0189)
Post x MWI	0.0877*	0.1109***
	(0.0334)	(0.0249)
Post x MWI x Low Rent Group		-0.0860*
		(0.0337)
Rent (Lease)	0.0627*	0.0547**
	(0.0282)	(0.0171)
Market Rent	-0.1900*	-0.1857**
	(0.0733)	(0.0595)
Inflation	0.0011	0.0010
	(0.0028)	(0.0012)
PC Income	0.3259**	0.3315***
	(0.0952)	(0.0440)
Unemployment	0.0086	0.0084
	(0.0120)	(0.0050)
HPI	-0.0008*	-0.0008***
	(0.0004)	(0.0001)
Vacancy	-1.3100**	-1.3130***
	(0.4570)	(0.1874)
MWI Year FE	Yes	Yes
State-Clustered SE	Yes	Yes
# Leases	*15,046*	*15,046*
Adjusted R-squared	*0.095*	*0.095*

This table reports DID estimation of tenants' probability of moving after minimum wage increase. Our sample consists of tenants with leases expiring within three months after a minimum wage increase who then entered into a new lease in the following three to nine months. *Low Rent Group* is a dummy variable equal to 1 if contract rent is less or equal to MSA fair market rent. The figures in parentheses are state-clustered standard errors. One, two, or three stars indicate statistical significance at 5%, 1%, or 0.1%, respectively.

Appendix

A. Unconditional Results

We explore the differences in average lease defaults using a DID model similar to Equation (1) omitting lease characteristics and macroeconomic factors but including lease-year fixed effects with standard errors clustered at the state level. Table A.2 shows that the coefficients of the interaction term $Post \times MWI$ from the various model specifications are negative and significantly different from zero. In states that enacted minimum wage increases, renters experienced on average 0.98 to 1.17 percentage points fewer defaults, depending on the model specification, during the three months following minimum wage increases compared with the three preceding months. These figures represent 6.94% to 7.96% fewer average monthly defaults over the first three months following treatment.[39] The six-month default estimations in Table A.2 lead to the same conclusion. Renters residing in states that raised the minimum wage experienced 1.06 to 1.35 percentage points fewer defaults over the six months following the increase than renters in states with no change in minimum wage.

B. Additional Analysis on Rent Quintiles

Although the magnitude of the effect of a minimum wage increase in the lowest rent quintile is consistent with expectations, one concern is that we also observe a negative, albeit smaller, effect in the highest rent quintile. To the extent that income is correlated with rent, one would normally not expect to find an effect in the highest rent quintile as individuals in this segment of the market most likely have incomes well above the minimum wage threshold and thus are the least likely to be impacted by a change in the minimum wage. However, we note that rent is a crude approximation for income and to the extent that lower income individuals do occupy high rent units, we would expect to also find an effect in the upper rent quintiles. As a result, we recognize the possibility that our rent bin estimates are biased toward finding an effect in the upper rent categories.

To confirm whether this bias is indeed present in the data, we matched the upper limit of the first rent quintile ($565) and the lower limit of the fifth rent quintile ($945) from our RentBureau data to the distribution of renter household incomes across various housing costs buckets from the 2005 American Housing Survey (AHS).[40] Figure A.1 reports the distribution of renters segmented into ten income brackets while figures A.2 and A.3 report the renter household distributions based on incomes less than and greater than $40,000, respectively. Consistent with our assumption that income and rent are correlated, Figure A.1 shows that a higher proportion of renters in the lower income brackets have monthly housing costs below $565 (the upper limit of the first rent quintile), and a much higher proportion of high income renters (those with incomes above $100,000) have rents above $945 (the lower limit of the upper quintile). However,

[39] The average three-month default rate in treated states post treatment is 3.73% (Panel C of Table 3), which gives the following average monthly effects: 0.98/(3.73+0.98)/3=6.94% and 1.17/(3.73+1.17)/3=7.96%.

[40] We combine the AHS family income into ten groups and keep the 15 AHS housing cost bins.

it is clear that some low income renters have monthly housing costs significantly greater than the lower limit of the fifth rent quintile. We can see this clearly in Figure A.2, which shows the distribution of renter households with incomes less than $40,000. For households at this income level, which corresponds to the income level most likely to be impacted by a change in the minimum wage, we note that a non-trivial share have housing costs that fall into the upper rent quintile. As a result, the estimated coefficient for the interaction term in Table 7 indicating an impact of the minimum wage in the highest rental group is to be expected. Furthermore, Figure A.3, which reports the distribution of households with incomes above $40,000, suggests that a non-trivial number of higher income households fall into the very low rental quintile. Taken together, figures A.2 and A.3 suggest a downward bias in the estimated impact of the minimum wage change for the lower rental quintile and an upward bias in the top rental quintile reported in Table 7.

C. Supplementary Tables

Table A.1: Data Sources

Variable Name	*Source*
Minimum Wage Increase	Monthly Labor Review of the U.S. Bureau of Labor Statistics
Minimum Wage Increase	Monthly Labor Review of the U.S. Bureau of Labor Statistics
Rent	Experian RentBureau
Market Rent	U.S. Department of Housing and Urban Development (HUD)
PC Income	U.S. Bureau of Economic Analysis
Inflation	U.S. Department of Labor: Bureau of Labor Statistics
Unemployment	U.S. Bureau of Labor and Statistics
Change Renter Population	Population Estimates Program, Population Division, U.S. Census Bureau
HPI	Federal Housing Finance Agency (FHFA) Three-Digit ZIP Codes (All-Transactions Index) Housing Price Index (HPI)
Rental Supply	U.S. Census Bureau (https://www.census.gov/construction/bps/stateannual.html)
Affordable Housing Supply	U.S. Department of Housing and Urban Development (https://www.huduser.gov/portal/datasets/lihtc.html).
Vacancy	U.S. Census Bureau (https://www.census.gov/housing/hvs/index.html)

Table A.2: Unconditional DID Estimation of the Effect of Minimum Wage Increases on Three and Six Month Lease Defaults

	(1) *3-Month Default*	*(2)* *3-Month Default*	*(1')* *6-Month Default*	*(2')* *6-Month Default*
Post	0.0134***	0.0146***	0.0169***	0.0206***
	(0.0006)	(0.0006)	(0.0006)	(0.0007)
MWI	0.0036**	0.0027*	0.0078***	0.0074***
	(0.0011)	(0.0012)	(0.0019)	(0.0019)
Post x MWI	-0.0117***	-0.0098***	-0.0135***	-0.0106***
	(0.0010)	(0.0010)	(0.0014)	(0.0014)
Constant	Yes	Yes	Yes	Yes
Lease-Year FE	No	Yes	No	Yes
State-Clustered SE	Yes	Yes	Yes	Yes
# Leases	*990,785*	*990,785*	*907,035*	*907,035*
Adjusted R-squared	*0.001*	*0.003*	0.001	0.004

This table reports unconditional DID results of variations in lease defaults in treated and control states pre and post minimum wage increases. The dependent variable is a 3-month (6-month) lease default indicator tracking whether tenants have missed a payment during the 3-month (6-month) period pre and the 3-month (6-month) period post a minimum increase. The 3-month and 6-month default samples consists of 990,785 and 907,035 individual leases, respectively. We define a lease as current (not in default) in a given month if RentBureau records its status as either P (on-time payment) or L (late payment). For leases in treated states, we track their performance pre and post each minimum wage increase in that state. For leases in control states, we track lease performance pre and post each minimum wage increase date in the treated group. *Post* stands for the post treatment period and *MWI* indicates treated states. The figures in parentheses are state-clustered standard errors. One, two, or three stars indicate statistical significance at 5%, 1%, or 0.1%, respectively.

Table A.3: Multivariate DID Estimation of the Effect of Minimum Wage Increases on Three-Month Default using *Restricted Sample*

	(1) *3-mo. Default*	*(2)* *3-mo. Default*
Post	0.0200***	0.0168***
	(0.0013)	(0.0011)
MWI	0.0109**	0.0083
	(0.0037)	(0.0048)
Post x MWI	-0.0121***	-0.0170***
	(0.0033)	(0.0033)
Rent (Lease)	-0.0063	-0.0033
	(0.0032)	(0.0024)
Market Rent	-0.0093	-0.0237**
	(0.0100)	(0.0085)
PC Income	0.0215*	0.0250**
	(0.0094)	(0.0090)
Inflation	0.0001	0.0012***
	(0.0001)	(0.0002)
Unemployment	0.0010	0.0000
	(0.0011)	(0.0012)
Change Renter Population	-0.0010	-0.0017
	(0.0008)	(0.0009)
HPI	-0.0000*	-0.0001*
	(0.0000)	(0.0000)
Rental Supply	0.0024	0.0044
	(0.0018)	(0.0028)
Affordable Housing Supply	-0.0034	-0.0052
	(0.0021)	(0.0028)
Vacancy	0.0001	0.0304
	(0.0476)	(0.0638)
Constant	Yes	Yes
Lease-Year FE	No	Yes
State-Clustered SE	Yes	Yes
# Leases	*726,332*	*726,332*
Adjusted R-squared	*0.003*	*0.006*

The dependent variable is a 3-month lease default indicator showing whether a tenant has missed a payment during that period. For each lease, we measure its default status over 3 months pre and post state minimum wage increases. The above DID estimations are restricted to leases with no missing values of the dependent variable before and after minimum wage increases. *Post* stands for the post treatment period and *MWI* indicates treated states. The figures in parentheses are state-clustered standard errors. One, two, or three stars on top of coefficient estimates indicate statistical significance of coefficient estimates at 5%, 1%, or 0.1%, respectively.

Table A.4: Multivariate DID Estimation of the Effect of Minimum Wage Increases on *Six-Month Defaults*

	(1) *6-mo. Default*	*(2)* *6-mo. Default*
Post	0.0170***	0.0136***
	(0.0011)	(0.0014)
MWI	0.0184**	0.0139
	(0.0062)	(0.0077)
Post x MWI	-0.0150***	-0.0172***
	(0.0028)	(0.0032)
Rent (Lease)	-0.0080	-0.0036
	(0.0049)	(0.0040)
Market Rent	-0.0279	-0.0522**
	(0.0190)	(0.0177)
PC Income	0.0378*	0.0435**
	(0.0167)	(0.0155)
Inflation	-0.0002	0.0019***
	(0.0001)	(0.0003)
Unemployment	0.0021	0.0022
	(0.0019)	(0.0021)
Change Renter Population	-0.0023	-0.0031
	(0.0014)	(0.0017)
HPI	-0.0001**	-0.0001*
	(0.0000)	(0.0000)
Rental Supply	0.0028	0.0081
	(0.0026)	(0.0043)
Affordable Housing Supply	-0.0032	-0.0088
	(0.0023)	(0.0047)
Vacancy	-0.0933	-0.0452
	(0.0831)	(0.0970)
Lease-Year FE	No	Yes
State-Clustered SE	Yes	Yes
# Leases	*894,831*	*894,831*
Adjusted R-squared	*0.002*	*0.007*

This table reports multivariate DID results of lease defaults in treated and control states pre and post minimum wage increases. The dependent variable is a 6-month lease default indicator tracking whether tenants have missed a payment during the 6-month periods pre and post a minimum increase – a lease is current (not in default) in a given month if RentBureau records its status as either P (on-time payment) or L (late payment). *Post* stands for the post treatment period and *MWI* indicates treated states. The figures in parentheses are state-clustered standard errors. One, two, or three stars on top of coefficient estimates indicate statistical significance of coefficient estimates at 5%, 1%, or 0.1%, respectively.

Table A.5: DID Estimation of Effect of Minimum Wage Increase on Lease Performance (Default) over Time

	(1) *Monthly Default*	*(2)* *Monthly Default*
Pre x MWI(Month 6)	0.0053*	0.0021
	(0.0021)	(0.0029)
Pre x MWI(Month 5)	0.0052*	0.0020
	(0.0022)	(0.0028)
Pre x MWI(Month 4)	0.0062**	0.0029
	(0.0021)	(0.0025)
Pre x MWI(Month 3)	0.0063**	0.0030
	(0.0022)	(0.0029)
Pre x MWI(Month 2)	0.0049**	0.0017
	(0.0018)	(0.0022)
Pre x MWI(Month 1)	0.0044*	0.0022
	(0.0019)	(0.0021)
Post x MWI(Month 1)	-0.0004	-0.0065
	(0.0042)	(0.0053)
Post x MWI (Month 2)	-0.0030	-0.0091**
	(0.0024)	(0.0032)
Post x MWI (Month 3)	-0.0032	-0.0094**
	(0.0026)	(0.0034)
Post x MWI(Month 4)	-0.0048	-0.0110**
	(0.0028)	(0.0036)
Post x MWI (Month 5)	-0.0010	-0.0072
	(0.0028)	(0.0036)
Post x MWI (Month 6)	-0.0023	-0.0084*
	(0.0027)	(0.0034)
Pre Control States (Monthly)	Yes	Yes
Post Control States (Monthly)	Yes	Yes
Control Variables	Yes	Yes
Lease-Year FE	No	Yes
State-Clustered SE	Yes	Yes
# Leases	*1,010,950*	*1,010,950*
Adjusted R-squared	*0.004*	*0.007*

This table reports DID results of monthly lease defaults in treated and control states pre and post minimum wage increases during six months pre and post wage increases. For the same of conciseness, we only report results in treated (minimum wage change state). *Pre Control States (Monthly)* and *Post Control States (Monthly)* refer to pre and post coefficient estimates in control states, respectively. Our model includes the same control variables as in Table 4. The figures in parentheses are state-clustered standard errors. One, two, or three stars indicate statistical significance at 5%, 1%, or 0.1%, respectively.

Table A.6: Treated and Control State Used in Parallel Trend Analysis

Treated State	*Treatment Date*	*Control State*
Arkansas	Oct-06	Tennessee
Arizona	Jan-07	Utah
California	Jan-07	Texas
Colorado	Jan-07	Utah
Connecticut*	Jan-06	Massachussetts
Delaware	Jan-07	Virginia
Florida	Jan-06	Georgia
Illinois	Jan-05	Indiana
Iowa	Apr-07	Nebraska
Kentucky	Jun-07	Tennessee
Maine*	Oct-06	New Hampshire
Maryland	Jan-07	Virginia
Massachusetts	Jan-07	Virginia
Michigan	Oct-06	Indiana
Minnesota	Aug-05	Nebraska
Missouri	Jan-07	Kansas
Nevada	Nov-06	Utah
New Hampshire	Sep-07	Virginia
New York*	Jan-06	Pennsylvania
North Carolina	Jan-07	South Carolina
Ohio	Jan-07	Indiana
Oregon	Jan-07	Idaho
Pennslyvania	Jan-07	Virginia
Washington	Jan-07	Idaho
Wisconsin	Jun-06	Nebraska

The 25 treated states and paired control states for selected treatment dates in month and two-digit year format. Treated states with * subscript, located in the Northeast region of country (where most states are treated), are paired with neighboring states whose treatment period do not overlap.

Table A.7: DID Estimation of Effect of Minimum Wage Increases on Rent over Time

	1-Month Pre MWI Window		*6-Month Pre MWI Window*	
	(1)	*(2)*	*(1')*	*(2')*
	Rent	*Rent*	*Rent*	*Rent*
Pre x MWI	0.1627	-0.0816	0.1874	-0.0710
	(0.0974)	(0.0419)	(0.1095)	(0.0419)
Post x MWI (Month 1)	-0.0169*	-0.0009	-0.0417*	-0.0166
	(0.0082)	(0.0075)	(0.0206)	(0.0095)
Post x MWI (Month 2)	-0.0048	-0.0027	-0.0295	-0.0193
	(0.0066)	(0.0100)	(0.0152)	(0.0116)
Post x MWI (Month 3)	0.0096	0.0283**	-0.0151	0.0118
	(0.0069)	(0.0080)	(0.0166)	(0.0079)
Post x MWI (Month 4)	0.0179*	0.0418***	-0.0068	0.0254***
	(0.0080)	(0.0067)	(0.0150)	(0.0060)
Post x MWI (Month 5)	0.0251*	0.0469***	0.0004	0.0310***
	(0.0119)	(0.0080)	(0.0103)	(0.0058)
Post x MWI (Month 6)	0.0334*	0.0694***	0.0087	0.0539***
	(0.0148)	(0.0116)	(0.0114)	(0.0094)
Pre Control States (Monthly)	Yes	Yes	Yes	Yes
Post Control States (Monthly)	Yes	Yes	Yes	Yes
Control Variables	No	Yes	No	Yes
Constant	Yes	Yes	Yes	Yes
State-Clustered SE	Yes	Yes	Yes	Yes
Adjusted R-squared	*0.48*	*0.299*	*0.052*	*0.304*
# Leases	*791,974*	*789,196*	*949,046*	*949,046*

This table reports DID results of rent (in '000s) in treated states pre and post minimum wage increases as compared to rents pre and post minimum wage increase to rents. *MWI* identifies treated states. *Pre* and *Post* stand for the pre and post treatment, respectively. The figures in parentheses are state-clustered standard errors. One, two, or three stars indicate statistical significance at 5%, 1%, or 0.1%, respectively.

Table A.8: Summary Statistics of Three-Month and Six-Month Defaults Characterized as *Not On-Time Rent Payments*

	3-Month Defaults				*6-Month Defaults*			
	# Leases	*# Obs*	*Mean*	*SD*	*# Leases*	*# Obs*	*Mean*	*SD*
Panel A: Full Sample								
Pre MWI	830,319	2,678,930	0.1517	0.3587	671,117	2,017,577	0.2144	0.4104
Post MWI	934,065	2,944,237	0.1681	0.3740	819,884	2,396,450	0.2364	0.4248
Difference & t-statistc			*-0.0164*				*-0.0219*	
			(53.09)				*(54.83)*	
Panel B: Control Group								
Pre MWI	564,632	2,365,570	0.1505	0.3575	471,749	1,791,783	0.2118	0.4086
Post MWI	598,237	2,554,093	0.1692	0.3749	534,900	2,076,022	0.2370	0.4252
Difference & t-statistic			*-0.0187*				*-0.0251*	
			(56.65)				*(59.05)*	
Panel C: Treated Group								
Pre MWI	265,687	313,360	0.1608	0.3674	199,368	225,794	0.2354	0.4242
Post MWI	335,828	390,144	0.1610	0.3675	284,984	320,428	0.2325	0.4224
Difference & t-statistic			*-0.0002*				*0.0028*	
			(0.17)				*(2.44)*	

This table reports summary statistics of 3-month and 6-month lease defaults pre and post state minimum wage increases passed from 2000 to 2008. Our treated group regroups the 25 states that enacted minimum wage increases during that period, whereas the control group consists of the 14 other states in the RentBureau data that did not pass any minimum wage increase (See Table 1). Our sample consist of 990,785 unique leases signed between 2000 and 2008. We observe the performance of each lease for up to 24 months. We consider a lease current (not in default) in a given month if RentBureau records its status as P (on-time payment). Three-month (six-month) defaults indicate whether leases have been in default during any one month during the three (six) months pre and post a minimum wage increase.

Table A.9: Unconditional DID Estimation of Three and Six Month Defaults Characterized as *Not On-Time Rent Payments*

	(1) *3-mo. Default*	*(2)* *3-mo. Default*	*(1')* *6-mo. Default*	*(2')* *6-mo. Default*
Post	0.0187***	0.0194***	0.0251***	0.0269***
	(0.0014)	(0.0014)	(0.0020)	(0.0019)
MWI	0.0105	0.0039	0.0235	0.0155
	(0.0099)	(0.0104)	(0.0136)	(0.0142)
Post x MWI	-0.0186*	-0.0139	-0.0280**	-0.0209*
	(0.0077)	(0.0080)	(0.0093)	(0.0097)
Constant	Yes	Yes	Yes	Yes
Lease-Year FE	No	Yes	No	Yes
State-Clustered SE	Yes	Yes	Yes	Yes
# Leases	*990,785*	*990,785*	*907,035*	*907,035*
Adjusted R-squared	*0.001*	*0.003*	*0.001*	*0.004*

This table reports unconditional DID OLS estimation results of variations in lease defaults in treated and control states pre and post minimum wage increases. The dependent variable is a 3-month (6-month) lease default indicator tracking whether tenants have missed a payment during the 3-month (6-month) period pre and the 3-month (6-month) period post a minimum increase – a lease is current (not in default) in a given month if RentBureau records its status as P (on-time payment). For leases in treated states, we track their performance pre and post each minimum wage increase in that state. For leases in control states (i.e. states that did not pass any minimum wage increase during the study period), we track the leases' performance pre and post the minimum wage increase dates in the treated states. *Post* stands for the post treatment period and *MWI* indicates treated states. The figures in parentheses are state-clustered standard errors. One, two, or three stars on top of coefficient estimates indicate statistical significance of coefficient estimates at 5%, 1%, or 0.1%, respectively.

Table A.10: Multivariate DID Estimation of Three- and Six-Month Lease Defaults Characterized as *Not On-Time Rent Payments*

	(1) *3-mo. Default*	*(2)* *6-mo. Default*
Post	0.0138***	0.0165***
	(0.0013)	(0.0027)
MWI	0.0380*	0.0557**
	(0.0145)	(0.0191)
Post x MWI	-0.0307***	-0.0358***
	(0.0075)	(0.0084)
Rent (Lease)	-0.0921***	-0.0895***
	(0.0142)	(0.0174)
Market Rent	-0.0564	-0.0888
	(0.0390)	(0.0647)
PC Income	0.0681	0.0718
	(0.0342)	(0.0447)
Inflation	0.0029***	0.0032***
	(0.0004)	(0.0005)
Unemployment	0.0003	0.0033
	(0.0069)	(0.0077)
Change Renter Population	-0.0039	-0.0049
	(0.0035)	(0.0044)
HPI	-0.0003*	-0.0004**
	(0.0001)	(0.0001)
Rental Supply	-0.0016	0.0034
	(0.0077)	(0.0097)
Affordable Housing Supply	0.0005	-0.0028
	(0.0073)	(0.0081)
Vacancy	0.0072	-0.1904
	(0.1973)	(0.2350)
Lease-Year FE	Yes	Yes
State-Clustered SE	Yes	Yes
# Leases	*984,376*	*894,831*
Adjusted R-squared	*0.011*	*0.012*

This table reports multivariate difference-in-differences OLS estimation results of lease defaults in treated and control states pre and post minimum wage increases. The dependent variables are 3- and 6-month lease default indicators tracking whether tenants have missed a payment during the 3-month periods pre and post a minimum increase – a lease is current (not in default) in a given month if RentBureau records its status as P (on-time payment). For leases in treated states, we track their performance pre and post each minimum wage increase in that state. For leases in control states (i.e. states that did not pass any minimum wage increase during the study period), we track the leases' performance pre and post the minimum wage increase dates in the treated states. *Post* stands for the post treatment period and *MWI* indicates treated states. The figures in parentheses are state-clustered standard errors. One, two, or three stars on top of coefficient estimates indicate statistical significance of coefficient estimates at 5%, 1%, or 0.1%, respectively.

Table A.11: DID Estimation of the Effect of Minimum Wage Increases on Three-Month Lease Defaults (*Robustness Checks*)

	(1) *3-mo. Default* *Excl. Cross-State MSAs*	*(2)* *3-mo. Default* *Excl. 2007-2008*	*(3)* *3-mo. Default* *Excl. Rent Control*
Post	0.0124***	0.0106***	0.0117***
	(0.0011)	(0.0012)	(0.0010)
MWI	0.0075	0.0063*	0.0075
	(0.0048)	(0.0029)	(0.0053)
Post x MWI	-0.0180***	-0.0094*	-0.0161***
	(0.0029)	(0.0046)	(0.0028)
Control Variables	Yes	Yes	Yes
Lease-Year FE	Yes	Yes	Yes
State-Clustered SE	Yes	Yes	Yes
# Leases	*884,406*	*378,554*	*894,473*
Adjusted R-squared	*0.005*	*0.006*	*0.005*

This table reports multivariate DID results of lease defaults in treated and control states pre and post minimum wage increases. Column (1) excludes MSAs spreading over several states. Column (2) excludes minimum wage changes enacted in 2007 and 2008. Column (3) excludes rent-control states (CA, MD, and NY). The dependent variable is a 3-month lease default indicator tracking whether tenants have missed a payment during the 3-month periods pre and post a minimum increase – a lease is current (not in default) in a given month if RentBureau records its status as either P (on-time payment) or L (late payment). For leases in treated states, we track their performance pre and post each minimum wage increase in that state. For leases in control states (i.e. states that did not pass any minimum wage increase during the study period), we track the leases' performance pre and post the minimum wage increase dates in the treated states. *Post* stands for the post treatment period and *MWI* indicates treated states. The control variables are the same as in Table 4. The figures in parentheses are state-clustered standard errors. One, two, or three stars on top of coefficient estimates indicate statistical significance of coefficient estimates at 5%, 1%, or 0.1%, respectively.

Table A.12: DID Estimation of the Effects of Minimum Wage Increases on Default and Rent in Rent Control Cities in California (Los Angeles and San Francisco) Relative to the Rest of the State (*Robustness Checks*)

	(1) *All California* *3-mo Default*	*(2)* *All California* *3-mo Default*	*(3)* *Rent Control Cities* *3-mo Default*	*(4)* *Rent Control Cities* *Rent*
Post	-0.0078**	-0.0078**	-0.0097***	-0.0061
	(0.0025)	(0.0025)	(0.0018)	(0.0121)
MWI	-0.0068	-0.0067	-0.0013	0.0894
	(0.0088)	(0.0087)	(0.0090)	(0.1082)
Post x MWI	-0.0132***	-0.0134***	-0.0114***	0.0021
	(0.0021)	(0.0024)	(0.0020)	(0.0130)
Post x MWI x Rent Control		0.0010		
		(0.0023)		
Control Variables	Yes	Yes	Yes	Yes
Constant	Yes	Yes	Yes	Yes
Lease-Year FE	Yes	Yes	Yes	No
Clustered SE (State)	Yes	Yes	Yes	Yes
# Leases	*387,072*	*387,072*	*312,939*	*316,755*
Adjusted R-squared	0.006	0.006	0.008	0.563

This table reports multivariate DID default and rent results in California. Due to data limitations, we restrict our analysis to the January 2007 and 2008 minimum wage increases in that state. We compare the effects of minimum wage increases in rent-controlled cities (Los Angeles and San Francisco) using a dummy variable (Rent Control) identifying leases in those areas. The dependent variable columns 1 to 3 is lease default over the three months pre and post the wage increases – a lease is current (not in default) in a given month if RentBureau records its status as either P (on-time payment) or L (late payment). The rent regression in column 4 is based on monthly rent (at the lease level) over the six months pre and post wage increase. *Post* stands for the post treatment period and *MWI* indicates treatment. The control variables are the same as in Table 4 for the default regressions and Table A.7 for the rent regression. The figures in parentheses are state-clustered standard errors. One, two, or three stars on top of coefficient estimates indicate statistical significance of coefficient estimates at 5%, 1%, or 0.1%, respectively.

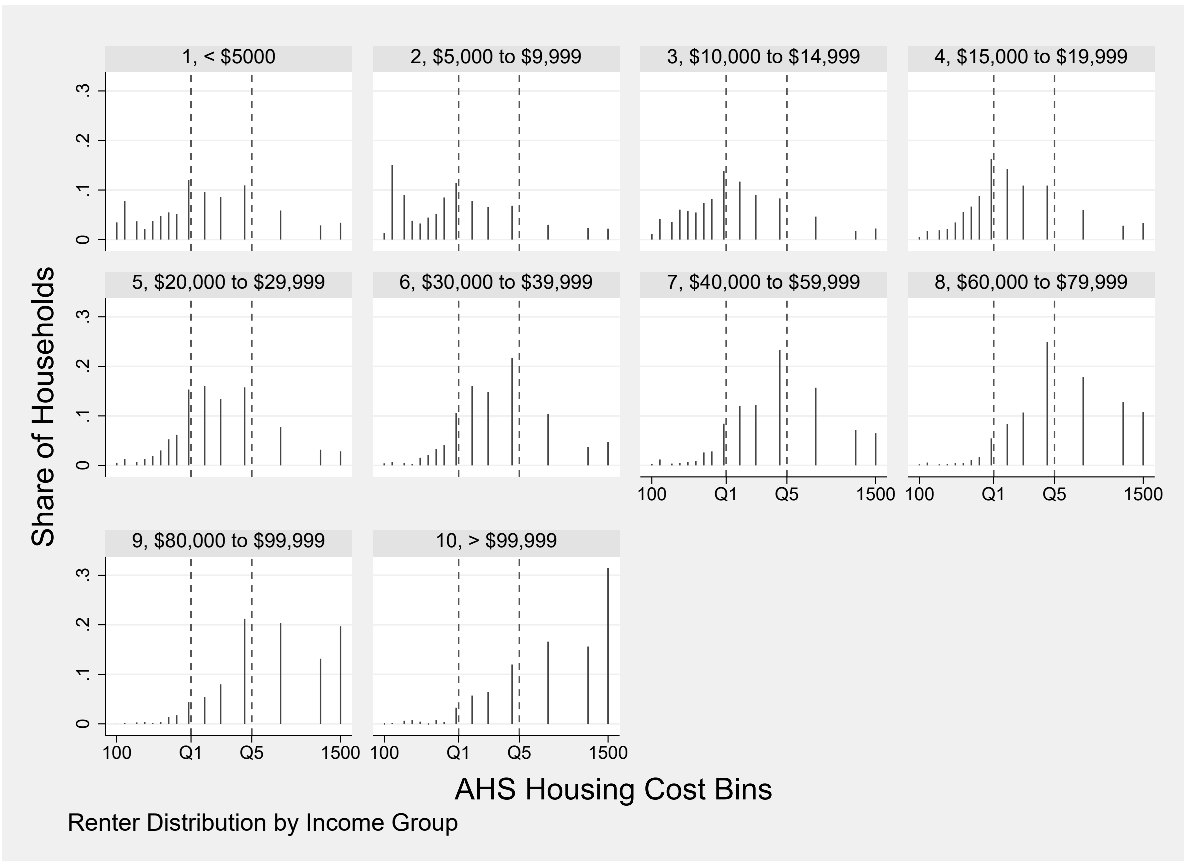

Figure A.1: Renter Distribution by Income Group
Note: These are distributions of renter households by income group across various housing cost buckets from the 2005 American Housing Survey (AHS) national data – no-cash rents are not included. We combine the AHS family income groups into ten bins, but keep the fifteen housing cost bins used by AHS. The dashed vertical lines at Q1 ($565) and Q5 ($945) are the upper limit of the first rent quintile and the lower limit of the fifth rent quintile, respectively, from the 2005 leases in the RentBureau data.

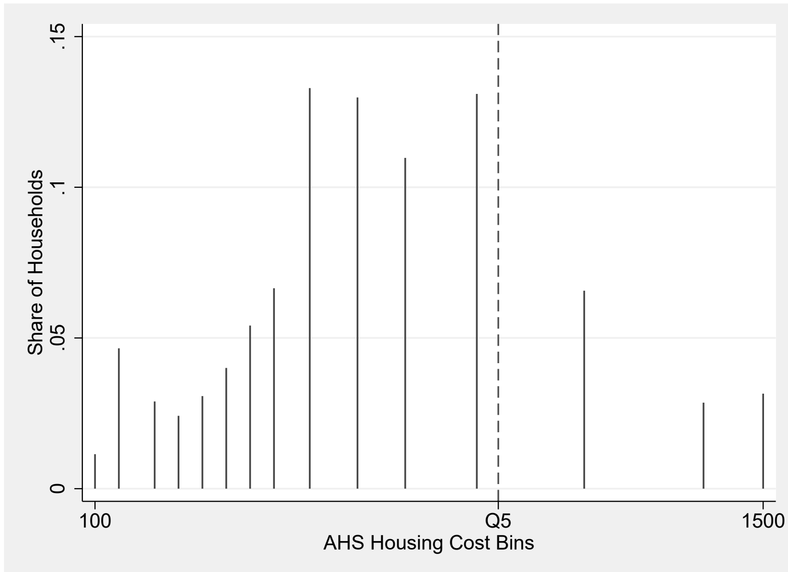

Figure A.2: Distribution of Low-Income Household Housing Cost
This graph shows the housing cost distribution of renter households earning less than $40,000 (*Low-Income Renter Households*) and the share of households in each of the fifteen housing cost bins from the 2005 American Housing Survey (AHS) national data. The vertical dashed line at Q5 ($945) is the lower limit of the fifth rent quintile from the 2005 leases in the RentBureau data.

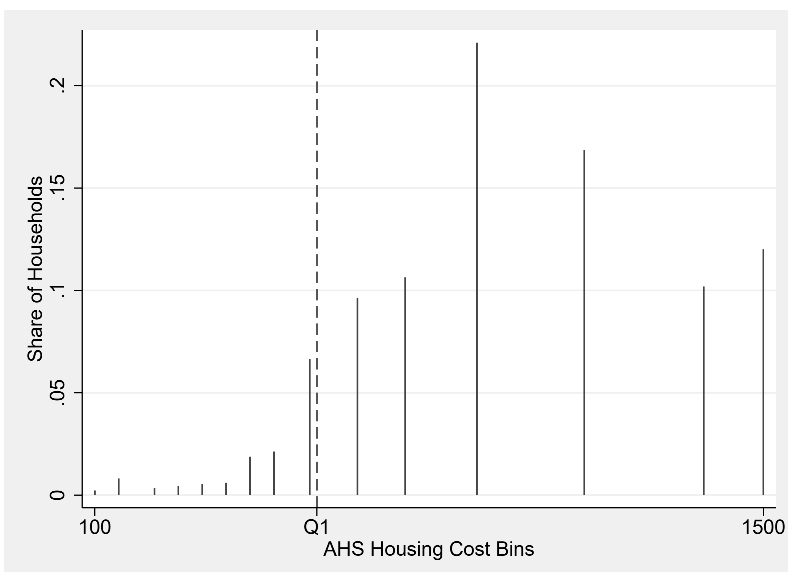

Figure A.3: Distribution of High-Income Household Housing Cost
This graph shows the distribution of renter households earning at least $40,000 (*High-Income Renter Households*) and the share of households in each of the fifteen housing cost bins from the 2005 American Housing Survey (AHS) national data. The vertical dashed line at Q1 ($565) is the upper limit of the first rent quintile from the 2005 leases in the RentBureau data.