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Minimum Wage and Firm Variety

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Minimum Wage and Firm Variety

Abstract

Exploiting minimum-wage variation within multi-state commuting zones, we document a negative relationship between minimum wages and firm variety in the U.S. restaurant and retail-trade industries. To explain this finding, we construct a heterogeneous-firm model with a monopsonistic labor market and endogenous firm variety. The decentralized equilibrium underprovides the mass of firms compared to the outcome achieved by a welfare-maximizing planner. A binding minimum wage further reduces the mass of firms, exacerbating the distortion. Workers value employer variety, and thus, by reducing firm variety the minimum wage reduces workers' welfare even if the average wage increases.

JEL-Codes: J380, J420.

Keywords: minimum wage, number for firms, love of employer variety.

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1 Introduction

There is a large body of empirical research on the labor market implications of minimum wages, with the main focus being on the impact of the minimum wage on employment and average wages (see Neumark and Shirley, 2021 and Manning, 2021 for recent surveys). Although the evidence is mixed, studies that do not find a negative employment response to minimum wages (see, *e.g.*, Card and Krueger, 1994) often mention that a potential explanation is that labor markets are monopsonistic: facing an upward sloping labor supply, firms offer a wage that is below their marginal labor cost, and thus a minimum wage could lead to an increase in employment. This analysis, however, is incomplete because it assumes a fixed number of firms, and this number can be affected by a minimum wage. Simply put, a binding minimum wage increases firms' costs, which may lead to the exit of the least productive firms and reduce entry incentives, causing a decline in the mass of firms. Surprisingly, in spite of its first order importance, the vast minimum-wage literature has ignored the relationship between minimum wages and the number of firms. The objective of this paper is to fill this void by empirically analyzing this relationship, and by studying the welfare implications of minimum wages when employer variety matters.

In section 2, we document a negative and statistically significant relationship between minimum wages and the number of firms in the U.S. restaurant and retail-trade industries. Using yearly U.S. data at the commuting zone–state level for the 1990-2011 period, we estimate specifications that build on the empirical literature of minimum wages and employment, starting with a simple specification where the dependent variable is the log of the number of establishments in the industry of interest and the main explanatory variable is the log of the prevalent minimum wage in the state. Inspired by Dube, Lester, and Reich (2010), we then estimate a specification that only uses multi-state commuting zones. This is our preferred specification, as it allows us to exploit minimum wage variation within commuting zones while controlling for time-varying spatial heterogeneity at the local level. For both industries, the inverse relationship between minimum wages and the number of firms is stronger in the stricter specification.

To understand the mechanisms at play, in section 3 we expand the monopsonistic labor market framework to endogenize the number of firms in a setting with heterogeneous firms. In our model, changes in the number of firms affect welfare because workers love variety of employers: the larger the number of firms a worker could choose to work for, the higher her welfare is. Whereas the role of the number of firms is well understood for consumers' welfare—more firms imply a larger variety of goods—and the Dixit-Stiglitz model with love-of-variety preferences is the workhorse framework

in industrial organization and international trade, the role of the number of firms in workers' welfare has received less attention. Card, Cardoso, Heining, and Kline (2018) discuss a model where workers have idiosyncratic preferences for employers due to factors such as location, work hours, and work culture. As first pointed out by Thisse and Toulemonde (2010), in such a setting the larger the mass of employers the greater the maximized expected utility of workers.[1]

Using a Melitz-type structure, we find that the decentralized equilibrium and the social planner's problem yield similar firm sizes, so that in spite of a monopsonistic labor market, there is no misallocation of resources across firms.[2] However, the mass of firms and total employment are less in the decentralized equilibrium compared to the planner's problem solution, as a result of firms not taking into account the positive effect that the creation of an extra firm has on labor supply and welfare. In this setting, a binding minimum wage wipes out the least productive firms, affecting the allocation of labor across firms, and reducing the total mass of employers. Since the mass of firms was suboptimal in the decentralized equilibrium, a binding minimum exacerbates the existing distortion and reduces total employment and welfare, even if the average wage increases.

In related research, Berger, Herkenhoff, and Mongey (2019) show that a binding minimum wage in their oligopsony model can alleviate the distortion induced by firms' markdowns on wages, but in contrast to our model, they abstract from the welfare effects of employer variety because they consider a fixed number of firms. Close to the endogenous-labor-supply feature of our model, Bhaskar and To (1999) study the implications of minimum wages in a setting where firms' monopsony power emerges due to workers' commuting costs, so that firms must pay higher wages to attract workers from farther away. They find that a binding minimum wage has an ambiguous effect on labor force participation and consequently on employment.

Dustmann, Lindner, Schönberg, Umkehrer, and Vom Berge (2020) study the reallocation effects of minimum wages following the introduction of a nationwide minimum wage in Germany that affected 15 percent of all employees. They find that the minimum wage raised wages and did not lower employment, but induced low-wage workers to move from small, low-paying firms to larger, higher-paying firms at the expense of increased commuting time. The latter finding highlights the importance of idiosyncratic, non-pecuniary factors in workers' preferences of for employers. More related to our focus on employer variety, Luca and Luca (2019) use Yelp ratings data from San Francisco restaurants and find that the minimum wage increases exit of one-star restaurants, whereas

[1] In related work, in Jha and Rodriguez-Lopez (2021) we study the effects of international trade on inequality and welfare when the labor market is monopsonistic and employer variety matters.

[2] This parallels the result in Dhingra and Morrow (2019) where with CES preferences, monopolistic competition is constrained efficient.

it does not affect exit in five-star restaurants. Hence, consistent with our model, there is evidence of exit among low productivity restaurants. As well, Chava, Oettl, and Singh (2019) find evidence of increased exit probability in states where the federal minimum wage binds compared to states where the federal minimum wage does not bind.

2 Minimum Wages and the Number of Establishments: Empirical Assessment

This section documents a negative and statistically significant relationship between minimum wages and the number of establishments in the U.S. restaurant and retail-trade industries, which are the two U.S. industries where minimum wages are more likely to bind.

2.1 Data

From the Census's County Business Patterns (CBP) we obtain yearly establishment counts, employment counts, and annual pay from 1990 to 2011. We follow Acemoglu, Autor, Dorn, Hanson, and Price (2016) (AADHP hereafter)—and make extensive use of their detailed programs—to process the CBP data into 479 industries and 722 commuting zones, with the difference that in our data commuting zones are also split by state. At the commuting zone–state level there are 585 single-state commuting zones, 129 two-state commuting zones, and 8 three-state commuting zones.[3] We also follow AADHP to obtain yearly working-age population at the commuting zone–state level from the Census of Population Estimates. Lastly, yearly minimum-wage data at the state level—defined as the largest of the federal minimum wage and the state minimum wage—is obtained from Vaghul and Zipperer (2016).[4]

From these sources we construct two datasets. The first dataset includes all available 866 commuting zone–state entities, while the second dataset includes only the 137 multi-state commuting zones (corresponding to 281 commuting zone–state entities).[5] Following Dube, Lester, and Reich (2010), DLR hereafter, the purpose of the second dataset is to exploit local differences in minimum wages to be able to control for spatial heterogeneity, which may bias national estimates. As shown in Figure A-1 in the Appendix, the 137 multi-state commuting zones are fairly distributed within the continental United States.[6]

[3] As in AADHP, we exclude Alaska and Hawaii from our analysis. We also exclude the District of Columbia, which appears in the CBP starting in 2004.

[4] The minimum wage data is available in this link.

[5] The 866 commuting zone–state entities come from 585 single-state commuting zones, 129×2 two-state commuting zones, and $8 \times 3 - 1$ three-state commuting zones—the District of Columbia, which is not included in our data, is part of a three-state commuting zone.

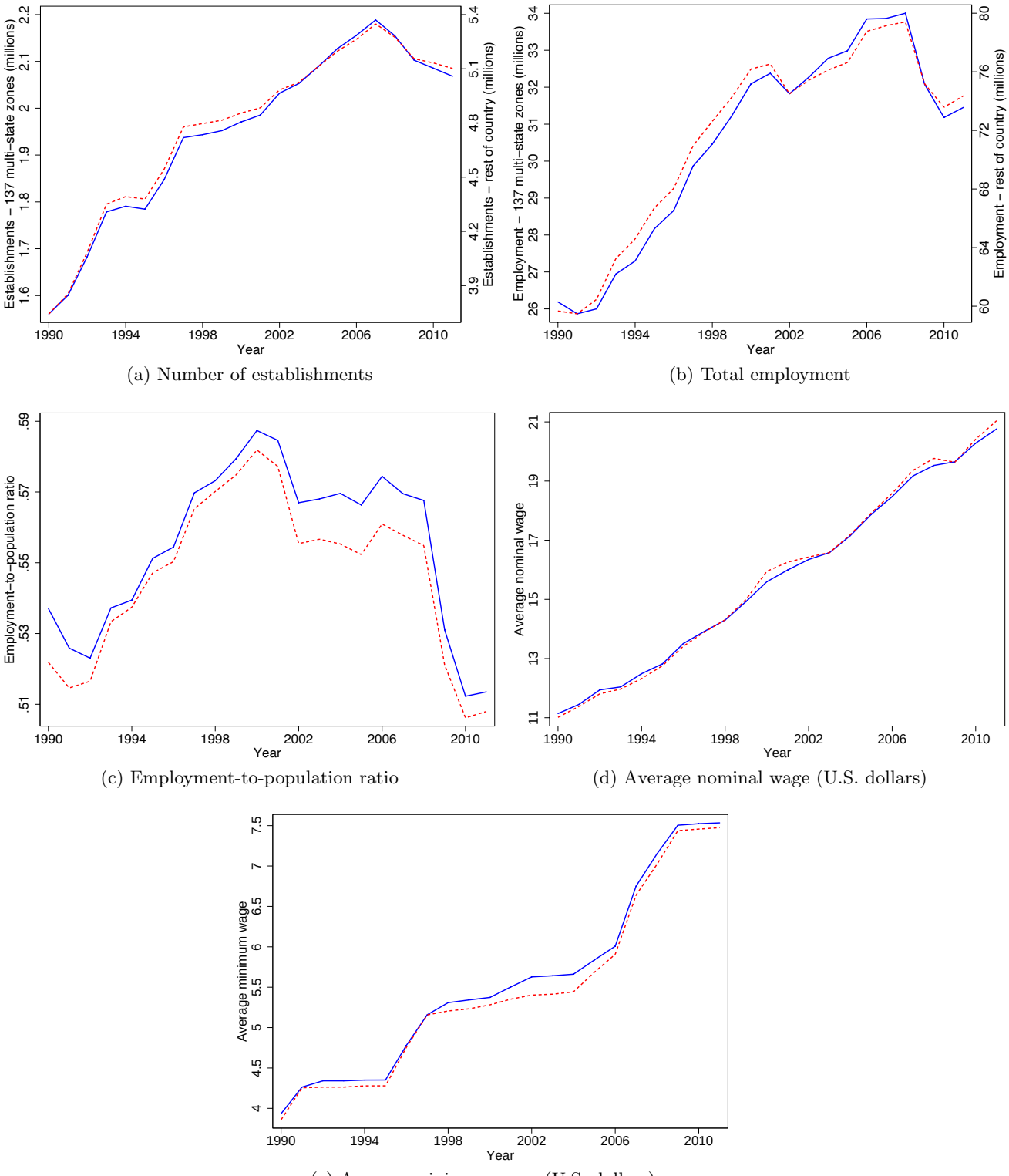
[6] The Appendix is available at http://www.socsci.uci.edu/~jantonio/Papers/minwage_firmvariety_app.pdf.

Throughout the 1990-2001 period, the multi-state commuting zone group accounts on average for 29.88 percent of U.S. employment, 29 percent of U.S. establishments, and 29.57 percent of the U.S. working-age population.[7] Moreover, all the variables of interest in this exercise are very similar in the multi-state group (137 commuting zones) and in the rest-of-the-country group (585 single-state commuting zones). To show this, Figure 1 presents a comparison between the groups along establishment and employment counts, employment-to-populations ratios, average nominal wages, and average minimum wages.

Figures 1a–1c show that both groups follow the same patterns for establishment counts, employment counts, and employment-to-population ratios. The only noticeable difference is that employment-to-population ratios are slightly higher in the multi-state group—the average throughout the period is 55.46 percent for the multi-state group and 54.63 percent for the single-state group. Figure 1d shows similar values and patterns for the average nominal *hourly* wage, calculated for each group as the total annual pay divided by total annual working hours ($2,087 \times$ employment). Finally, Figure 1e shows that the average minimum wage—weighted by commuting zone–state working-age population—has a similar evolution in both groups, increasing from about \$3.9 in 1990 to about \$7.5 in 2011.

As in DLR, we focus on the restaurant and retail-trade industries—industries 5812 (Eating and drinking places) and 5210 (Retail trade) in the Standard Industrial Classification used by AADHP—which are the industries where the minimum wage is likely to have a larger impact. After aggregating the remaining 477 AADHP industries into 18 industries, Table A-1 in the Appendix shows industry-level average nominal wages and wage rankings for 1990 and 2011. Note that in both years the restaurant and retail-trade industries have the lowest average nominal wages. However, there is a large wage gap between both industries, with the retail-trade industry paying an average wage that is 75 percent higher in 1990 and 77 percent higher in 2011. Table A-1 also presents industry employment shares, showing that our two industries of interest jointly account for 21.15 percent of U.S. employment in 1990 (7.21 percent in restaurants and 13.94 percent in retail trade) and for 21.81 percent in 2011 (8.72 percent in restaurants and 13.08 percent in retail trade). For both the restaurant and retail-trade industries, Figures A-2 and A-3 in the Appendix show that establishment counts, employment, and wages follow the same patterns in the multi-state commuting zone group and the rest of the country.

[7] These shares are very stable over time. They range between 29.55 and 30.5 percent for employment, between 28.83 and 29.43 percent for establishments, and between 29.43 and 29.9 percent for working-age population.



(a) Number of establishments

(b) Total employment

(c) Employment-to-population ratio

(d) Average nominal wage (U.S. dollars)

(e) Average minimum wage (U.S. dollars)

Figure 1: Comparison between commuting-zone groups: 137 multi-state commuting zones (solid blue) and rest of the country (dashed red)

2.2 Econometric Specifications

Our specifications for the relationship between the number of establishments and minimum wages follow on the employment specifications of DLR, who expand on the work of Neumark and Wascher (1992) by rigorously controlling for time-varying spatial heterogeneity. Using the full-country dataset, we estimate equation

$$\ln e_{it} = \alpha + \beta \ln MW_{it} + \gamma \ln E_{it}^{-} + \delta \ln P_{it} + \eta_i + \tau_{ct} + \kappa_S \mathbb{1}_S \cdot T + \varepsilon_{it}, \tag{1}$$

where for commuting zone–state i in year t, e_{it} is the number of restaurant or retail establishments, MW_{it} is the minimum wage, E_{it}^{-} is the total number of establishments in commuting zone–state i minus e_{it}, P_{it} is the working-age population, η_i is a commuting zone–state i fixed effect, τ_{ct} accounts for time fixed effects for each of the nine Census regional divisions, and $\mathbb{1}_S \cdot T$ represents state-level trends ($\mathbb{1}_S$ is a dummy variable taking the value of 1 if entity i belongs to state S and T denotes a time trend).

Although specification (1) attempts to account for spatial heterogeneity by including Census-division time fixed effects and state-level trends, DLR argue that this is not enough to account for local economic conditions and introduce a multi-state county-pair approach to control for these. Along these lines, we exploit minimum-wage variation within commuting-zone pairs using the 137 multi-state commuting zone dataset and the econometric model

$$\ln e_{ipt} = \alpha + \beta \ln MW_{it} + \gamma \ln E_{it}^{-} + \delta \ln P_{it} + \eta_i + \tau_{pt} + \nu_{it}, \tag{2}$$

where subscript p identifies a commuting-zone pair for entity i, and τ_{pt} denotes pair–time fixed effects, which control for spatial heterogeneity at the local level. Among our 137 multi-state commuting zones there are 151 pairs: one pair for each of the 129 two-state commuting zones, three pairs for each of 7 of the three-state commuting zones, and one more pair corresponding to Virginia and Maryland in the DC-VA-MD commuting zone (recall that DC is excluded from our data).

As in DLR, we estimate (2) using stacked-pair differences: if entities i and j belong to pair p, we subtract $\ln e_{jpt}$ from (2) to obtain

$$\ln e_{ipt} - \ln e_{jpt} = \beta \left(\ln MW_{it} - \ln MW_{jt}\right) + \gamma \left(\ln E_{it}^{-} - \ln E_{jt}^{-}\right) + \delta \left(\ln P_{it} - \ln P_{jt}\right) + \eta_{ij} + \nu_{ijt}, \tag{3}$$

where $\eta_{ij} \equiv \eta_i - \eta_j$ is the pair (i, j) fixed effect, and $\nu_{ijt} = \nu_{it} - \nu_{jt}$ is the error term. We can then estimate (3) as a panel with 151 pairs and 22 years.

Table 1: Estimation of Number-of-Establishments Responses to Minimum Wages

	Full sample			Multi-state zones		
	(1)	(2)	(3)	(4)	(5)	(6)
A. Restaurant industry						
ln(minimum wage)	-0.120***	-0.093**	-0.044*	-0.122*	-0.186**	-0.169**
	(0.032)	(0.038)	(0.023)	(0.069)	(0.079)	(0.081)
ln(establishments^{-})	0.160**	0.319***	0.355***	0.174	0.294**	0.271*
	(0.067)	(0.067)	(0.066)	(0.152)	(0.121)	(0.144)
ln(population)	0.938***	0.687***	0.604***	0.985***	0.838***	0.755***
	(0.085)	(0.079)	(0.079)	(0.206)	(0.176)	(0.198)
B. Retail-trade industry						
ln(minimum wage)	-0.023	-0.061**	-0.029	-0.007	-0.113**	-0.127**
	(0.021)	(0.025)	(0.019)	(0.043)	(0.053)	(0.055)
ln(establishments^{-})	0.427***	0.418***	0.402***	0.394***	0.389***	0.376***
	(0.039)	(0.041)	(0.042)	(0.077)	(0.076)	(0.083)
ln(population)	0.401***	0.428***	0.467***	0.409***	0.308**	0.247*
	(0.046)	(0.049)	(0.057)	(0.084)	(0.125)	(0.131)
Year effects	Y			Y		
Region-year effects		Y	Y			
State trends			Y			
Pair-year effects					Y	Y
Number of pairs					151	129
Observations	19,052	19,052	19,052	6,182	3,322	2,838

Notes: This table reports $\hat{\beta}$, $\hat{\gamma}$, and , $\hat{\delta}$ from the estimation of equation (1) (in columns 1-4) and equation (3) (in columns 5 and 6) for the restaurant and retail industries. The dependent variable is the log number of establishments in columns 1-4, and the difference in log number of establishments for each pair within a commuting zone in columns 5-6. Columns 1-3 use the full sample with 722 commuting zones (866 commuting zone-state entities), columns 4 and 5 use the 137 multi-state commuting zones, and column 6 uses only the 129 two-state commuting zones. All regressions include entity fixed effects. Standard errors (in parentheses) are clustered at commuting-zone level. The coefficients are statistically significant at the *10%, **5%, or ***1% level.

2.3 Results

From the previous econometric specifications, our coefficient of interest is β, which denotes the elasticity of the number of establishments to the minimum wage. Table 1 presents our main empirical results from the estimation of (1) and (3) for the restaurant and retail-trade industries for the 1990-2011 period. All regressions include entity fixed effects and standard errors are clustered at the commuting-zone level.

Using the full sample, column 1 in Table 1 starts with a version of (1) that only includes year fixed effects, so that it abstracts from controlling for time-varying spatial heterogeneity. For restaurants, the estimated minimum-wage elasticity of the number of establishment is -0.12, so that a 1

percent increase in the minimum wage is associated with a 0.12 percent decline in the number of establishments. Both controls yield positive and significant coefficients, indicating that the number of restaurant establishments moves together with the number of establishments in other industries, and that larger populations are related to more establishments. Columns 2 and 3 show that as we include time-varying spatial heterogeneity controls, first with Census region-year effects in column 2 and then adding state trends in column 3, the minimum-wage elasticity declines in magnitude and significance, though it remains statistically different from zero at a 10 percent level.

Using only the multi-state commuting-zone dataset, column 4 re-estimates the specification in column 1 and shows that the minimum-wage elasticity coefficient barely changes. However, the main advantage of the multi-state zones dataset is that it allows us to estimate equation (3), which controls for time-varying spatial heterogeneity at the local level. The estimation of (3) is presented in columns 5 and 6, with column 5 including the 151 pairs spanning from the 137 multi-state commuting zones, whereas column 6 includes only the 129 pairs from the 129 two-state commuting zones.[8] For the restaurant industry, columns 5 and 6 show that after strictly controlling for spatial heterogeneity, the minimum-wage elasticity is larger in magnitude (either -0.186 or -0.169) and statistically significant at a 5 percent level. Hence, our results in Table 1 show strong support for a negative and significant relationship between the minimum wage and the number of establishments in the restaurant industry.

For the retail industry, the estimates of the minimum-wage elasticity are also negative, but they are smaller in magnitude than those of the restaurant industry (this is expected, as the average wage in the retail industry is around 75 percent higher), and three out of six are not statistically significant. Importantly, the estimated elasticity is the largest in magnitude and is significant at a 5 percent level when using the stricter local specification from equation (3) (see columns 5 and 6).

Although we are mostly interested in the relationship between minimum wages and the number of establishments, we can also delve into the most popular and controversial topic in the minimum-wage literature: the existence (or not) of a negative and significant minimum-wage elasticity of employment. The main point of DLR is that the negative relationship between employment and minimum wages in the restaurant industry weakens as more strict controls of spatial heterogeneity are included, and disappears—flipping to a positive, but not significant, coefficient—when using the local specification of contiguous county-pairs sharing a state border. Table 2 presents the results from the estimation of equations (1) and (3), but using employment instead of the number of establishments (on the right-hand side of these equations, we replace the total number of establishments in the other

[8]The purpose of showing also the results with only the two-state zones is to make sure that the over representation of pairs from three-states zones (there are three pairs in each three-state zone, whereas there is only one pair in two-state zones) is not an important driver of the results.

Table 2: Estimation of Employment Responses to Minimum Wages

	Full sample			Multi-state zones		
	(1)	(2)	(3)	(4)	(5)	(6)
A. Restaurant industry						
ln(minimum wage)	-0.298***	-0.149***	-0.060*	-0.294***	-0.325***	-0.326***
	(0.040)	(0.045)	(0.032)	(0.071)	(0.094)	(0.096)
ln(employment$^-$)	-0.027	0.074*	0.095**	-0.122	0.137	-0.018
	(0.045)	(0.044)	(0.047)	(0.087)	(0.108)	(0.111)
ln(population)	0.960***	0.794***	0.801***	1.136***	0.926***	0.929***
	(0.066)	(0.071)	(0.075)	(0.118)	(0.202)	(0.236)
B. Retail-trade industry						
ln(minimum wage)	-0.030	-0.025	0.037	-0.018	-0.110	-0.132*
	(0.025)	(0.035)	(0.027)	(0.051)	(0.072)	(0.075)
ln(employment$^-$)	0.151***	0.118***	0.114***	0.020	0.070	-0.003
	(0.034)	(0.037)	(0.037)	(0.065)	(0.088)	(0.084)
ln(population)	0.665***	0.684***	0.712***	0.885***	0.637***	0.861***
	(0.046)	(0.052)	(0.056)	(0.089)	(0.173)	(0.148)
Year effects	Y			Y		
Region-year effects		Y	Y			
State trends			Y			
Pair-year effects					Y	Y
Number of pairs					151	129
Observations	19,052	19,052	19,052	6,182	3,322	2,838

Notes: Using employment instead of the number of establishments, this table reports $\hat{\beta}$, $\hat{\gamma}$, and , $\hat{\delta}$ from the estimation of equation (1) (in columns 1-4) and equation (3) (in columns 5 and 6) for the restaurant and retail industries. The dependent variable is the log employment in columns 1-4, and the difference in log employment for each pair within a commuting zone in columns 5-6. Columns 1-3 use the full sample with 722 commuting zones (866 commuting zone-state entities), columns 4 and 5 use the 137 multi-state commuting zones, and column 6 uses only the 129 two-state commuting zones. All regressions include entity fixed effects. Standard errors (in parentheses) are clustered at commuting-zone level. The coefficients are statistically significant at the *10%, **5%, or ***1% level.

industries with total employment in the other industries).

As in DLR, the full-sample results for the restaurant industry in columns 1-3 from Table 2 show that the minimum-wage elasticity of employment declines in magnitude as we include more time-varying spatial heterogeneity controls—the elasticity changes from -0.298 in column 1 to -0.06 in column 3. However, and in stark contrast to DLR, the elasticity is larger in magnitude (and highly significant) when estimating the local specification exploiting minimum-wage differentials within commuting zones: columns 5 and 6 show similar elasticities of about -0.325. Thus, there is evidence of a negative and statistically significant relationship between minimum wages and restaurant-industry employment in the United States, even after controlling for time-varying local economic conditions.

Table 3: Estimation of Average Wage Responses to Minimum Wages

	Full sample			Multi-state zones		
	(1)	(2)	(3)	(4)	(5)	(6)
A. Restaurant industry						
ln(minimum wage)	0.214***	0.182***	0.167***	0.171***	0.239***	0.265***
	(0.032)	(0.042)	(0.041)	(0.054)	(0.070)	(0.076)
ln(average wage^{-})	0.198***	0.177***	0.160***	0.112	-0.053	0.060
	(0.043)	(0.047)	(0.047)	(0.070)	(0.125)	(0.095)
ln(population)	0.084**	0.108***	0.113**	0.005	0.081	0.077
	(0.043)	(0.041)	(0.046)	(0.072)	(0.088)	(0.074)
B. Retail-trade industry						
ln(minimum wage)	0.039*	0.041*	0.044**	0.029	0.052	0.065*
	(0.020)	(0.025)	(0.019)	(0.035)	(0.041)	(0.039)
ln(average wage^{-})	0.168***	0.173***	0.157***	0.156**	0.159	0.083
	(0.032)	(0.035)	(0.034)	(0.066)	(0.110)	(0.112)
ln(population)	-0.002	0.005	0.014	-0.036	-0.173	-0.028
	(0.025)	(0.029)	(0.032)	(0.050)	(0.127)	(0.074)
Year effects	Y			Y		
Region-year effects		Y	Y			
State trends			Y			
Pair-year effects					Y	Y
Number of pairs					151	129
Observations	19,052	19,052	19,052	6,182	3,322	2,838

Notes: Using the average hourly wage instead of the number of establishments, this table reports $\hat{\beta}$, $\hat{\gamma}$, and , $\hat{\delta}$ from the estimation of equation (1) (in columns 1-4) and equation (3) (in columns 5 and 6) for the restaurant and retail industries. The dependent variable is the log average wage in columns 1-4, and the difference in log average wage for each pair within a commuting zone in columns 5-6. Columns 1-3 use the full sample with 722 commuting zones (866 commuting zone-state entities), columns 4 and 5 use the 137 multi-state commuting zones, and column 6 uses only the 129 two-state commuting zones. All regressions include entity fixed effects. Standard errors (in parentheses) are clustered at commuting-zone level. The coefficients are statistically significant at the *10%, **5%, or ***1% level.

For the retail industry the relationship is much weaker, being only mildly significant in the local specification that uses the 129 two-state commuting zones (see column 6).

Finally, Table 3 shows estimates of the minimum-wage elasticity of the average wage. Instead of the number of establishments, we estimate (1) and (3) using the average hourly wage in either the restaurant or the retail-trade industry (and the average hourly wage in the rest of the industries on the right-hand side of the specifications). For the restaurant industry, all columns show a positive and highly significant (at a 1 percent level) elasticity, ranging between 0.167 and 0.265. Taking our preferred specification in column 5, a 10 percent increase in the minimum wage increases the average wage in restaurants by 2.39 percent. For the retail industry the elasticity is also positive, but much

smaller and not significant in two of the six columns, ranging from a non-significant 0.029 to a mildly significant 0.065.

2.4 Discussion

The main finding of DLR is that the negative and significant relationship between minimum wages and employment disappears after estimating a specification similar to equation (3), and conclude that previous studies obtained a negative and significant elasticity because they failed to control for time-varying local economic conditions. We obtain the opposite result: Table 2 shows that the estimated negative minimum-wage elasticity of employment increases in magnitude and is statistically significant at a 1 percent level when estimating the local specification.

There are at least two reasons that may explain why we find a large, negative, and significant elasticity in the local specification whereas DLR do not: *(i)* differences in the frequency of the data (we use yearly data from the Census's CBP, while DLR use quarterly data from the BLS's QCEW), and *(ii)* differences in the geographic level of aggregation (we focus on commuting zones, while DLR focus on counties). Regarding *(i)*, we think that quarterly data yield lower elasticities because they only capture the very short-term effects of minimum wages. Regarding *(ii)*, comparing our map in Figure A-1 against DLR's map in their Figure 2, we can see that although they are relatively similar in coverage, DLR include many contiguous counties across the country that may have very little economic activity between them. By using instead commuting zones, which by definition attempt to capture areas of joint economic activity, it is possible that we are better able to capture employment variation driven by cross-state minimum-wage differentials. Fortunately, our CBP data allows us quantify the importance of the geographic level of aggregation in our results. Table A-2 and section A.2 in the Appendix present and discuss the estimation of (3) for the number of establishments, employment, and average wages using the county pairs of DLR, showing that indeed, the level of geographic aggregation matters.

3 The Model

This section presents a model with a monopsonistic labor market, heterogeneous firms, and endogenous firm variety. The model shows that a binding minimum wage reduces the mass of firms, which then reduces welfare because workers love employer variety. Here we present the relevant parts of the model, and leave the most technical details for section B in the Appendix.

3.1 Household Preferences and Production

The economy produces a single final good that is chosen as the numéraire. The final good is traded in a perfectly competitive market, and is produced by a finite mass, M, of heterogeneous firms. Firms produce the final good using labor, which is procured from a unit measure of households participating in a monopsonistic labor market.

As in Berger, Herkenhoff, and Mongey (2019), the utility function of the representative household is

$$\mathbb{U} \equiv C - \frac{N^{1+\frac{1}{\psi}}}{1+\frac{1}{\psi}}, \tag{4}$$

where C is the consumption of the final good, N is the labor-supply index, and ψ is the Frisch elasticity of labor supply. The labor-supply index is defined as

$$N = \left(\int_{\omega \in \Omega} l(\omega)^{\frac{1+\theta}{\theta}} d\omega \right)^{\frac{\theta}{1+\theta}}, \tag{5}$$

where $l(\omega)$ is the amount of labor supplied to firm ω, Ω is a set of measure M of firms/employers offering jobs, and $\theta > \psi$. The parameter $\theta > 0$ is the elasticity of substitution across jobs from different firms, and accounts for workers' cost of mobility: the lower the value of θ, the higher the mobility costs, and thus the greater the monopsony power of firms.

The above specification captures the love-of-variety for employers. To see this clearly, suppose there are M identical firms, so that $l(\omega) = l$ is the same across firms and $N = M^{\frac{\theta}{1+\theta}} l$. For a constant amount of supplied labor, $L = Ml$, note that the labor-supply index can be written as $N = \frac{L}{M^{1/(1+\theta)}}$. Hence, the disutility from supplying the same amount of labor is lower the larger M is.

Given the utility function in (4) and wages $w(\omega)$, for $\omega \in \Omega$, the representative household maximizes its utility by choosing, C, N, and allocating labor, $l(\omega)$, across firms. As shown in the Appendix, this maximization exercise yields the firm-level labor supply function

$$l(\omega) = \frac{w(\omega)^{\theta}}{W^{\theta-\psi}}, \tag{6}$$

for $\omega \in \Omega$, where W is an index of the wages available to the representative household, which is defined as

$$W = \left(\int_{\omega \in \Omega} w(\omega)^{1+\theta} \right)^{\frac{1}{1+\theta}}. \tag{7}$$

Firms take W as given, and hence, the function in (6) features a constant wage-elasticity of labor supply, θ.[9] The firm-level labor supply is increasing in $w(\omega)$, and thus, the firm has monopsony power in the labor market.

[9] Jha and Rodriguez-Lopez (2021) and Egger, Kreickemeier, Moser, and Wrona (2019) also derive a constant elasticity labor supply function by using a random utility framework where workers have idiosyncratic preferences for employers.

The maximization problem also yields $N = W^{\psi}$ and hence, our assumption that $\theta > \psi$ simply means that the elasticity of the labor-supply index (also known as the Frisch elasticity) is smaller than the elasticity of firm-level labor supply. The wage bill of firms, which is given by $\int_{\omega\in\Omega} w(\omega)l(\omega)d\omega = WN = W^{1+\psi}$, equals the consumption of the household, C. Therefore, the welfare of the representative household in (4) can be rewritten as

$$\mathbb{U} = WN - \frac{N^{1+\frac{1}{\psi}}}{1+\frac{1}{\psi}} = \frac{W^{1+\psi}}{1+\psi}, \tag{8}$$

Thus, welfare is increasing in the wage index, W.

As in Melitz (2003), after incurring an entry cost of $f_{_E}$ in terms of the final good, each firm draws its productivity φ from a distribution $G(\varphi)$. The firm produces for the market if and only if it can cover a fixed cost of operation, f, also in terms of the final good.[10] After meeting the fixed cost, the production function for a firm with productivity φ is $y(\varphi) = \varphi l$. Whereas the goods market is perfectly competitive, the labor market is monopsonistically competitive. Given (6) and using φ instead of ω to identify each firm, the profit maximization problem of a firm with productivity φ yields as solution

$$l(\varphi) = \left(\frac{\theta}{1+\theta}\right)^{\theta} \frac{\varphi^{\theta}}{W^{\theta-\psi}} \qquad \text{and} \qquad w(\varphi) = \left(\frac{\theta}{1+\theta}\right)\varphi. \tag{9}$$

The solution shows that employment and wages are increasing in φ: higher productivity firms employ more workers and pay higher wages. Note also that the solution for $w(\varphi)$ shows that a worker is paid a fraction $\frac{\theta}{1+\theta}$ of the value of her marginal product, φ (*i.e.*, the proportional markdown on wages, $\frac{1}{1+\theta}$, is the same for every firm). It follows that the gross profit function of a firm with productivity φ, $\pi(\varphi) = \varphi l(\varphi) - w(\varphi)l(\varphi)$, is given by

$$\pi(\varphi) = \left[\frac{\theta^{\theta}}{(1+\theta)^{1+\theta}}\right] \frac{\varphi^{1+\theta}}{W^{\theta-\psi}}. \tag{10}$$

3.2 Decentralized Equilibrium and the Social Planner's Problem

As in Melitz (2003), there is a cutoff level of productivity, $\hat{\varphi}$, such that firms with productivity below $\hat{\varphi}$ cannot cover their fixed cost of operation and hence do not survive. The zero-cutoff-profit condition is $\pi(\hat{\varphi}) = f$, which from (10) implies that $W^{\theta-\psi} = \left[\frac{\theta^{\theta}}{(1+\theta)^{1+\theta}}\right] \frac{\hat{\varphi}^{1+\theta}}{f}$, so that we can rewrite firm-level employment in (9) as $l(\varphi) = \frac{(1+\theta)f\varphi^{\theta}}{\hat{\varphi}^{1+\theta}}$ and the gross profit function in (10) as $\pi(\varphi) = \left(\frac{\varphi}{\hat{\varphi}}\right)^{1+\theta} f$. Firms enter up to the point that the expected value of entry, $\int_{\hat{\varphi}}^{\infty} [\pi(\varphi) - f]\, g(\varphi)d\varphi$, is equal to the

[10] We model entry and fixed costs in terms of the final good so that it is the same for all firms. Modeling it in terms of labor will make it different for firms because each firm pays a different wage in equilibrium.

entry cost, f_E; therefore, the free entry condition is given by

$$\int_{\hat{\varphi}}^{\infty} \left[\left(\frac{\varphi}{\hat{\varphi}} \right)^{1+\theta} - 1 \right] f g(\varphi) d\varphi = f_E. \tag{11}$$

Once we solve for $\hat{\varphi}$ from (11), we obtain the equilibrium W from the zero-cutoff-profit condition, and thus, using (8) we obtain that welfare is $\mathbb{U} = \frac{1}{1+\psi} \left\{ \left[\frac{\theta^\theta}{(1+\theta)^{1+\theta}} \right] \frac{\hat{\varphi}^{1+\theta}}{f} \right\}^{\frac{1+\psi}{\theta-\psi}}$.

Letting M_E be the mass of entrants, the mass of producing firms, M, is simply the fraction of entrants that survive: $M = [1 - G(\hat{\varphi})] M_E$. The equilibrium M is determined from the definition of W in (7), which can be rewritten as

$$W = \left[M \int_{\hat{\varphi}}^{\infty} w(\varphi)^{1+\theta} g(\varphi | \varphi \geq \hat{\varphi}) d\varphi \right]^{\frac{1}{1+\theta}}. \tag{12}$$

Finally, total employment is obtained by aggregating firm-level employment across all producing firms, $L = M \int_{\hat{\varphi}}^{\infty} l(\varphi) g(\varphi | \varphi \geq \hat{\varphi}) d\varphi$, which can be rewritten as

$$L = \frac{(1+\theta) f M}{\hat{\varphi}^{1+\theta}} \int_{\hat{\varphi}}^{\infty} \varphi^\theta g(\varphi | \varphi \geq \hat{\varphi}) d\varphi. \tag{13}$$

In the following, we refer to the decentralized equilibrium values as $\hat{\varphi}_D$, $l_D(\varphi)$, $\mathbb{U}_D$, M_D, and L_D, where subscript D denotes 'dencentralized'.

A social planner chooses $\hat{\varphi}$, $l(\varphi)$, and the mass of entrants, M_E, so as to maximize (4) subject to (5) and final-good consumption (C) being equal to total final-good production minus final-good requirements to cover firms' entry and fixed costs. Section B.3 in the Appendix shows the details of this standard maximization problem. Letting $\hat{\varphi}_P$, $l_P(\varphi)$, $\mathbb{U}_P$, M_P, and L_P denote the solution values from the planner's problem, the Appendix shows that $\hat{\varphi}_P = \hat{\varphi}_D$, $l_P(\varphi) = l_D(\varphi)$ for every φ,

$$\frac{M_P}{M_D} = \frac{L_P}{L_D} = \left(\frac{1+\theta}{\theta} \right)^{\frac{(1+\theta)\psi}{\theta-\psi}} > 1 \quad \text{and} \quad \frac{\mathbb{U}_P}{\mathbb{U}_D} = \frac{\theta-\psi}{(1+\theta)} \left(\frac{1+\theta}{\theta} \right)^{\frac{\theta(1+\psi)}{\theta-\psi}} > 1.$$

Thus, the decentralized equilibrium has the same productivity cutoff and the same firm level employment as in the planner's problem, but the planner chooses a higher level of M, which results in a higher L and higher welfare. In other words, the decentralized equilibrium provides a sub-optimal mass of firms.

Intuitively, for a given amount of supplied labor, the representative household's disutility from labor supply is smaller the larger the mass of firms (due to the household's love-of-variety for employers). The planner recognizes this and therefore, the perceived cost of providing an additional firm is lower in the planner's problem than in the decentralized case, so the planner ends up choosing a higher mass of firms. Importantly, the choice of labor across firms is not distorted in the decentralized case—the planner chooses the same allocation of resources across firms as in the decentralized

outcome. This is a consequence of the constant elasticity of firm-level labor supply, which implies that all firms mark down wages below the value of the marginal product by the same proportion, resulting in an efficient allocation of resources across firms.

A natural question is whether there is a policy that can correct the distortion and restore optimality in the decentralized equilibrium. In section B.4 in the Appendix we show that this is indeed the case: a wage subsidy of $\frac{1}{1+\theta}$ financed by a lump-sum tax on households restores optimality, so that the mass of firms, labor supply, and welfare in the decentralized case correspond to the planner's problem solution.[11]

3.3 Impact of a Minimum Wage Regulation

Denote a binding minimum wage by $\underline{w}$. Since the lowest productivity firm that survives in a decentralized equilibrium without any policy intervention offers a wage of $w(\hat{\varphi}_D)$, for the minimum wage to be binding it must be the case that $\underline{w} > w(\hat{\varphi}_D)$. Let $\underline{\varphi}$ denote lowest productivity level for which the desired wage offered by the firm, $\left(\frac{\theta}{1+\theta}\right)\underline{\varphi}$, equals $\underline{w}$. Therefore, $\underline{\varphi} = \frac{(1+\theta)\underline{w}}{\theta}$, and firms with $\varphi \geq \underline{\varphi}$ are unconstrained by the minimum wage, with their gross profits given by $\underline{\pi}(\varphi) = \left[\frac{\theta^\theta}{(1+\theta)^{1+\theta}}\right]\frac{\varphi^{1+\theta}}{\underline{W}^{\theta-\psi}}$, where $\underline{W}$ is the wage index.

Let $\underline{\hat{\varphi}}$ denote the cutoff level of productivity in an equilibrium with a binding minimum wage, so that firms with $\varphi \in [\underline{\hat{\varphi}}, \underline{\varphi})$ are minimum-wage constrained. These firms' gross profits are directly proportional to their employed labor, $\underline{\pi}(\varphi) = (\varphi - \underline{w})l(\varphi)$, and therefore, each of them hires $\frac{\underline{w}^\theta}{\underline{W}^{\theta-\psi}}$, which from (6) we know is the amount of labor that workers supply to a firm paying wage $\underline{w}$. The zero-cutoff-profit condition, $\underline{\pi}(\underline{\hat{\varphi}}) = f$, is then $(\underline{\hat{\varphi}} - \underline{w})\frac{\underline{w}^\theta}{\underline{W}^{\theta-\psi}} = f$. Lastly, the free entry condition is $\int_{\underline{\hat{\varphi}}}^{\infty}[\underline{\pi}(\varphi) - f]g(\varphi)d\varphi = f_E$, which can be rewritten as

$$\int_{\underline{\hat{\varphi}}}^{\underline{\varphi}}\left[\frac{(\varphi - \underline{w})\underline{w}^\theta}{\underline{W}^{\theta-\psi}} - f\right]g(\varphi)d\varphi + \int_{\underline{\varphi}}^{\infty}\left\{\left[\frac{\theta^\theta}{(1+\theta)^{1+\theta}}\right]\frac{\varphi^{1+\theta}}{\underline{W}^{\theta-\psi}} - f\right\}g(\varphi)d\varphi = f_E. \tag{14}$$

Using the definition of $\underline{\varphi}$, we solve for the equilibrium values of $\underline{\hat{\varphi}}$ and $\underline{W}$ from the zero-cutoff-profit condition and the free entry condition. We then solve for the equilibrium mass of firms ($\underline{M}$), total employment ($\underline{L}$), and welfare ($\underline{\mathbb{U}}$) using similar expressions to those in section 3.2. As well, the average wage, $\underline{\bar{w}}$, can be calculated as the ratio of the total wage bill ($\underline{W}^{1+\psi}$) and total employment. The following proposition presents our model's main results.

[11] Cahuc and Laroque (2014) study optimal policies in a monopsony model where workers are heterogeneous in productivity and in working opportunity costs—heterogeneous opportunity costs generate an upward-sloping labor supply curve, giving rise to monopsony power (as in Bhaskar and To, 1999). Similar to our findings, the monopsony distortion leads to suboptimal employment, which is corrected by a wage subsidy.

Proposition 1. ***(The effects of a binding minimum wage)***

A binding minimum wage wipes out the least productive firms, and reduces the total mass of firms, the wage index, and workers' welfare. If productivity follows a Pareto distribution, then a binding minimum wage also reduces total employment and increases the average wage.

The proof of this proposition is in section B.5 in the Appendix. Given that the minimum-wage equilibrium approaches the decentralized equilibrium when $\underline{w} \to w(\hat{\varphi}_D)$, the proof simply shows that *(i)* for every productivity distribution, it holds that $\frac{d\underline{\hat{\varphi}}}{d\underline{w}} > 0$, $\frac{d\underline{M}}{d\underline{w}} < 0$, $\frac{d\underline{W}}{d\underline{w}} < 0$, $\frac{d\underline{\mathbb{U}}}{d\underline{w}} < 0$, and thus $\underline{\hat{\varphi}} > \hat{\varphi}_D$, $\underline{M} < M_D$, $\underline{W} < W_D$, and $\underline{\mathbb{U}} < \mathbb{U}_D$; and that *(ii)* if the productivity distribution is Pareto, then it also unambiguously holds that $\frac{d\underline{L}}{d\underline{w}} < 0$ and $\frac{d\underline{\bar{w}}}{d\underline{w}} > 0$, so that $\underline{L} < L_D$ and $\underline{\bar{w}} > \bar{w}_D$, where $\bar{w}_D$ is the average wage in the decentralized equilibrium.[12]

Supporting Proposition 1, our main empirical finding from section 2 is that an increase in the minimum wage is associated with reductions in the masses of firms in the U.S. restaurant and retail-trade industries. In the model, a binding minimum wage makes the survival of low-productivity firms harder. This is captured by an increase in the cutoff productivity level, which then leads to a reduction in the mass of firms. The reduction in $\underline{M}$ causes a decline in the wage index, $\underline{W}$, which then translates into a reduction in welfare, $\underline{\mathbb{U}} = \frac{\underline{W}^{1+\psi}}{1+\psi}$. Welfare declines due to workers' love-of-variety for employers. To see this, it is useful to refer to the definition of the wage index in (12), which depends on the mass of firms and firm-level wages. Whereas firm-level wages increase to $\underline{w}$ for constrained firms, which pushes for an increase in $\underline{W}$, the reduction in $\underline{M}$ works in the opposite direction and is the dominant force.[13]

Notice that in spite of Melitz-type reallocation of resources from less productive firms to more productive firms, the minimum wage does not increase welfare. To understand why this is the case, recall from section 3.2 that the allocation of resources in the decentralized equilibrium is efficient ($\hat{\varphi}_P = \hat{\varphi}_D$ and $l_P(\varphi) = l_D(\varphi)$ for every φ). Therefore, any reallocation induced by a binding minimum wage cannot be a source of welfare gain. On the other hand, we also know from section 3.2 that the

[12] Since Chaney (2008), the Pareto distribution has been used extensively in applications of the Melitz model, as it brings substantial tractability gains. This is also true for our case, where it allows us to obtain closed-form solutions for $d\underline{L}/d\underline{w}$ and $d\underline{\bar{w}}/d\underline{w}$. Although the second part of Proposition 1 cannot be proved for a general productivity distribution, we verified numerically that it holds for a lognormal distribution, $g(\varphi) = \frac{1}{\varphi\sqrt{2\pi\rho}} \exp\left(-(\ln\varphi - \mu)^2/2\rho\right)$, with parameter values $\mu = -0.02$ and $\rho = 0.35$. These values are from Combes, Duranton, Gobillon, Puga, and Roux (2012), who find that firm-level productivity of French firms is better approximated by a mix 95% lognormal and 5% Pareto, and that restricting the distribution to be 100% lognormal yields our assumed values.

[13] Note that firm-level wages for unconstrained firms do not change ($w(\varphi) = \theta\varphi/(1+\theta)$ for $\varphi \geq \underline{\varphi}$). If we assume instead a firm-level production function with decreasing returns to scale, $y(\varphi) = \varphi l^{\lambda}$, for $\lambda < 1$, the equilibrium firm-level wage for unconstrained firms is given by $w(\varphi) = \left\{[\lambda\theta/(1+\theta)]\,\underline{W}^{(\theta-\phi)(1-\lambda)}\varphi\right\}^{1/[1+\theta(1-\lambda)]}$, which is increasing in $\underline{W}$. The result $d\underline{W}/d\underline{w} < 0$ still holds with $\lambda < 1$, and therefore, a binding minimum wage reduces unconstrained firms' wages.

mass of firms is suboptimal in the decentralized equilibrium. By further reducing the mass of firms, a binding minimum exacerbates the existing distortion.

In accordance with Proposition 1, our empirical results show that an increase in the minimum wage is associated with lower employment and a higher average wage in the U.S. restaurant industry. These results are useful to quantify welfare changes. Given that the wage bill, $\underline{L}\,\underline{w}$, equals $\underline{W}^{1+\psi}$, we can rewrite the expression for welfare as $\underline{\mathbb{U}} = \frac{\underline{L}\,\underline{w}}{1+\psi}$. Therefore, the elasticity of welfare with respect to the minimum wage equals the sum of the elasticities of total employment and the average wage. Based on our preferred elasticity estimates in column 5 of Tables 2 and 3, we obtain that the minimum-wage elasticity of welfare is $-.0325 + 0.239 = -0.086$ for the restaurant industry, so that—as predicted by Proposition 1—a 10% increase in the minimum wage is associated with 0.86% loss in welfare.[14]

4 Conclusion

Whereas we presented a simple model with competitive firms and monopsonistic labor markets, its key insights regarding the impact of a binding minimum wage on the mass of firms and employment will go through even if the product market were imperfectly competitive, as long as the number of firms is endogenous. The welfare effects would be different, however, depending on the precise specification of imperfect competition. For example, with standard CES preferences and monopolistic competition, a minimum-wage-induced reduction in the mass of firms will reduce welfare through the consumption channel, in addition to the welfare loss discussed in the paper through the employer variety channel.

[14] Tables 2 and 3 show that the signs of the elasticities of employment and the average wage for the retail-trade industry are also as predicted by Proposition 1, but they are smaller in magnitude and non-significant for our preferred specification in column 5. Based on those estimates, the elasticity of welfare for the retail-trade industry is $-0.11 + 0.052 = -0.058$.

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