City Size, Monopsony, and the Employment Effects of Minimum Wages

Appendix — For Online Publication

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A Supporting Figures and Tables

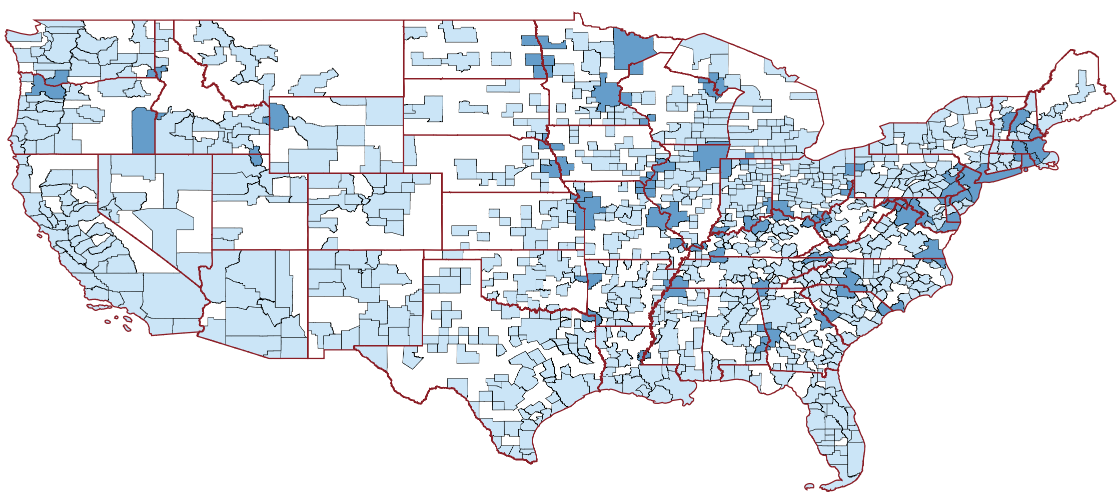
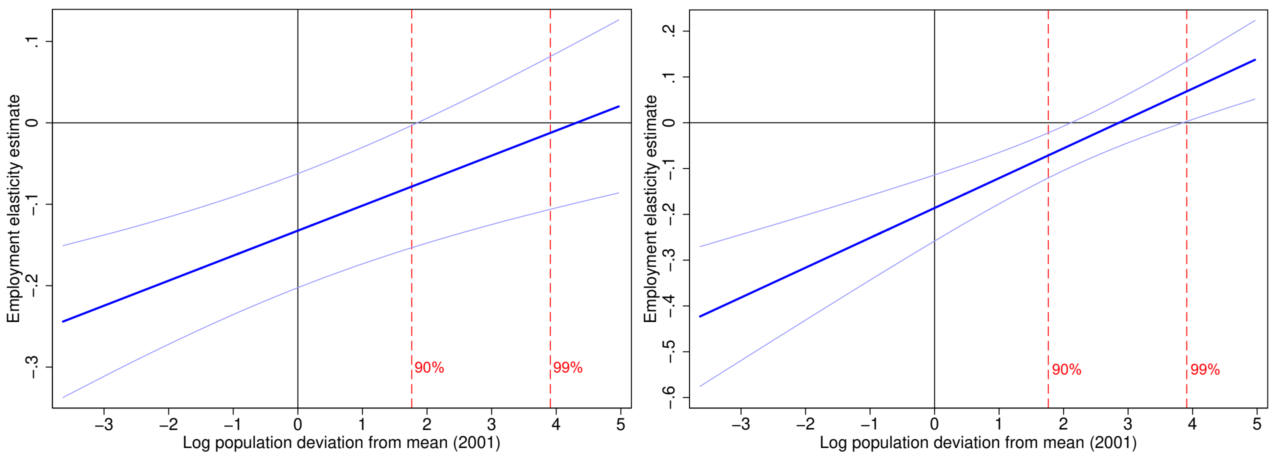
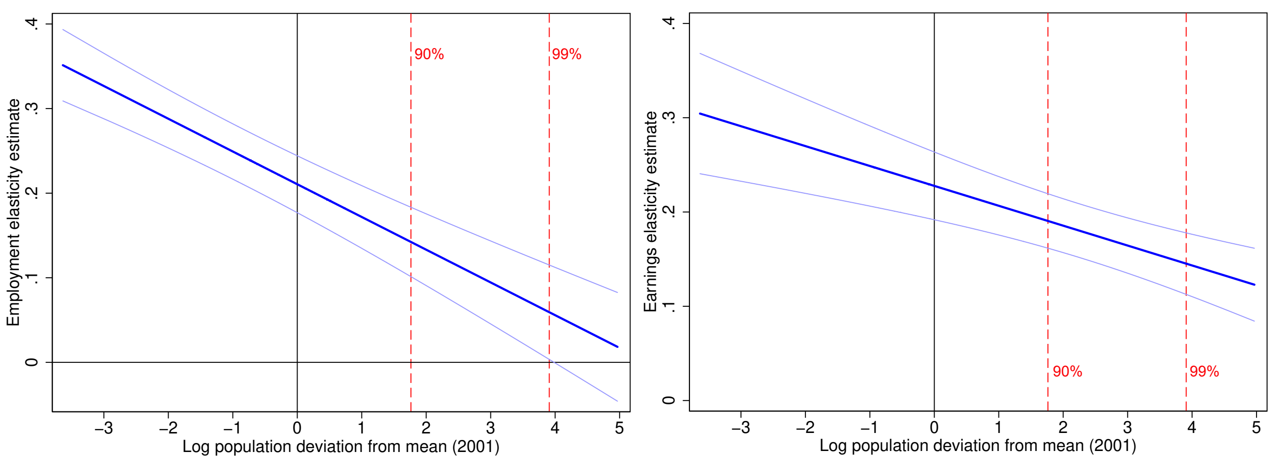


Figure A-1: The 919 Core-Based Statistical Areas (61 multi-state CBSAs in darker shade)

Notes: This map was created using Census TIGER files, which do not clip water bodies, as is particularly evident from the irregular shape of states with substantial coastlines or bodies of water, such as Michigan.

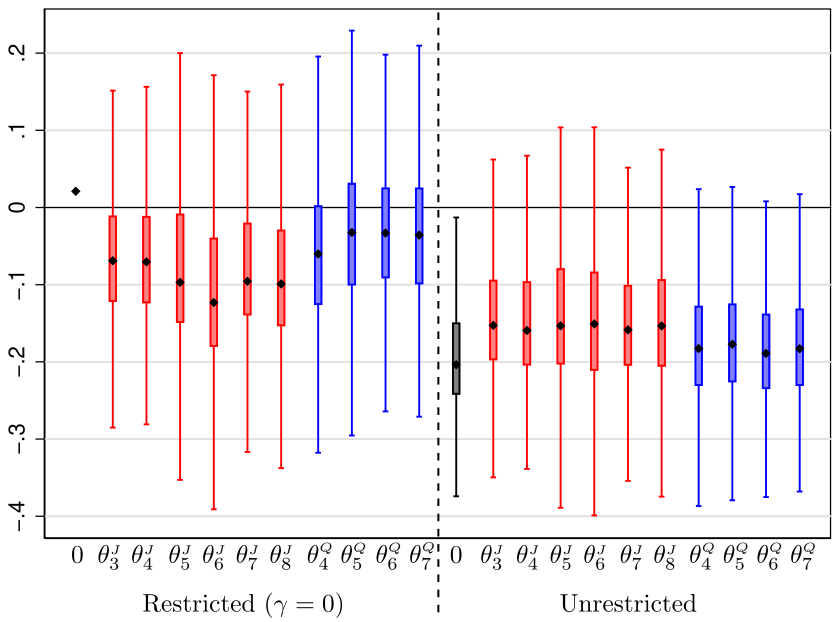


(a) Elasticity estimates for employment – unweighted (left) and population weights (right)

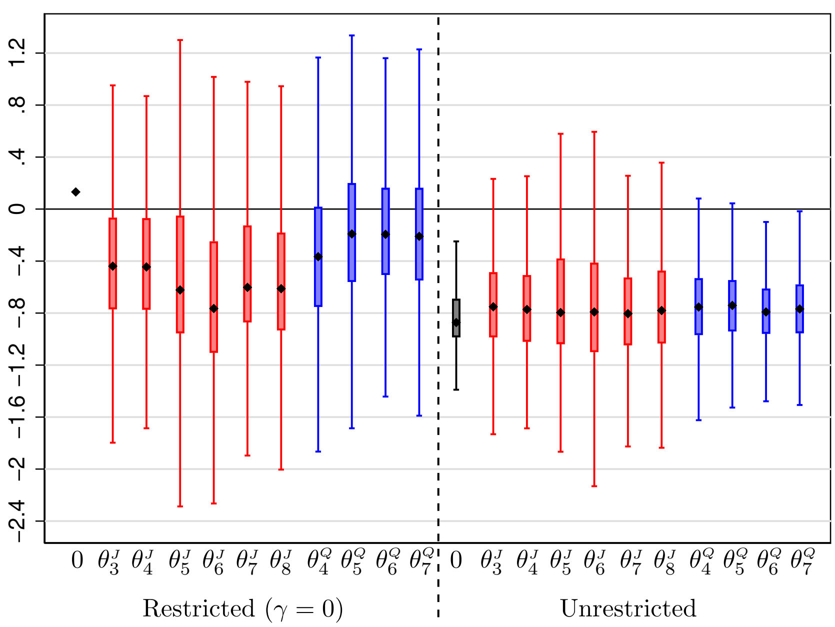


(b) Elasticity estimates for earnings – unweighted (left) and population weights (right)

Figure A-2: Estimates of the minimum wage elasticity of employment and earnings by population size and weighting method (with 90 percent confidence bands, and 90th and 99th percentiles for geography size)

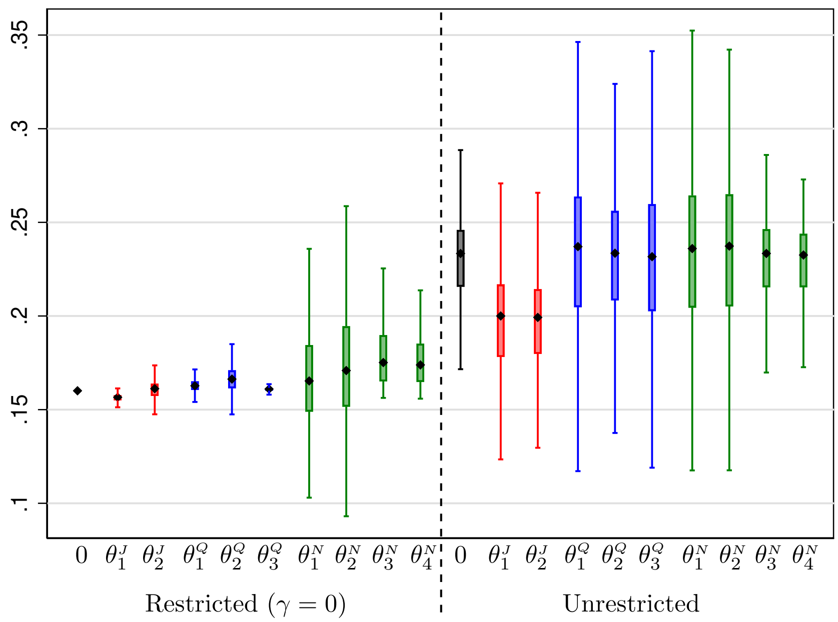


(a) Predicted minimum wage elasticities of employment

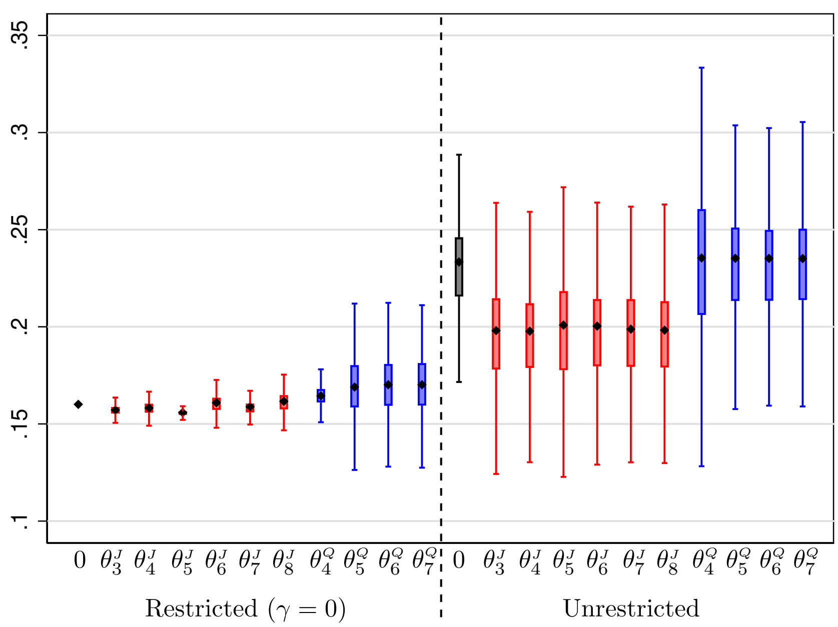


(b) Predicted own-wage elasticities (OWEs)

Figure A-3: Boxplots of predicted minimum wage elasticities of employment and OWEs for CBSA-by-state entities by fluidity monopsony-power proxy



(a) Benchmark fluidity and concentration measures



(b) Remaining fluidity measures

Figure A-4: Boxplots of predicted minimum wage elasticities of earnings for CBSA-by-state entities by monopsony-power proxy

Table A-1: Correlations between monopsony power proxies

	θ^J_1	θ^J_2	θ^J_3	θ^J_4	θ^J_5	θ^J_6	θ^J_7	θ^J_8	θ^Q_1	θ^Q_2	θ^Q_3	θ^Q_4	θ^Q_5	θ^Q_6	θ^Q_7	θ^N_1	θ^N_2	θ^N_3	θ^N_4	$\ln P$
θ^J_1	1.00																			
θ^J_2	0.94	1.00																		
θ^J_3	0.83	0.85	1.00																	
θ^J_4	0.80	0.86	0.95	1.00																
θ^J_5	0.99	0.92	0.79	0.76	1.00															
θ^J_6	0.93	0.99	0.80	0.80	0.93	1.00														
θ^J_7	0.96	0.91	0.84	0.81	0.89	0.88	1.00													
θ^J_8	0.89	0.96	0.89	0.91	0.85	0.90	0.90	1.00												
θ^Q_1	0.63	0.63	0.77	0.74	0.61	0.58	0.62	0.69	1.00											
θ^Q_2	0.64	0.68	0.77	0.80	0.62	0.63	0.62	0.73	0.92	1.00										
θ^Q_3	0.65	0.70	0.64	0.66	0.61	0.67	0.67	0.71	0.80	0.78	1.00									
θ^Q_4	0.65	0.67	0.78	0.78	0.63	0.61	0.64	0.73	0.98	0.98	0.81	1.00								
θ^Q_5	0.61	0.61	0.72	0.68	0.59	0.56	0.62	0.66	0.90	0.84	0.75	0.89	1.00							
θ^Q_6	0.59	0.64	0.70	0.72	0.57	0.59	0.59	0.70	0.82	0.89	0.71	0.88	0.86	1.00						
θ^Q_7	0.62	0.65	0.74	0.73	0.60	0.60	0.63	0.70	0.89	0.90	0.75	0.91	0.96	0.97	1.00					
θ^N_1	-0.27	-0.21	-0.14	-0.14	-0.28	-0.21	-0.25	-0.18	-0.29	-0.24	-0.44	-0.27	-0.25	-0.16	-0.21	1.00				
θ^N_2	-0.28	-0.21	-0.14	-0.14	-0.28	-0.22	-0.25	-0.19	-0.29	-0.24	-0.47	-0.27	-0.24	-0.16	-0.20	0.98	1.00			
θ^N_3	-0.04	-0.11	-0.05	-0.07	-0.03	-0.11	-0.06	-0.09	0.06	0.00	0.10	0.03	0.02	-0.02	0.00	-0.14	-0.20	1.00		
θ^N_4	-0.05	-0.12	-0.07	-0.08	-0.04	-0.12	-0.08	-0.11	0.05	-0.01	0.07	0.02	0.01	-0.03	0.00	-0.09	-0.14	0.98	1.00	
$\ln P$	-0.22	-0.30	-0.28	-0.28	-0.20	-0.29	-0.24	-0.30	-0.13	-0.18	-0.02	-0.16	-0.13	-0.20	-0.17	-0.33	-0.43	0.47	0.46	1.00

Notes: See Table 1 for definitions.

Table A-2: TWFE estimation of minimum wage effects on restaurant employment for the remaining fluidity monopsony-power measures

	Restricted ($\gamma = 0$)		**Unrestricted**		
	$\hat{\beta}$	$\hat{\delta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$
A. Job-to-Job fluidity monopsony proxies (32,643 obs.)					
Overall hir. rate (θ_3^J)	-0.066**	-2.882***	-0.189***	0.047***	-1.781***
	(0.031)	(0.569)	(0.049)	(0.016)	(0.317)
Overal sep. rate (θ_4^J)	-0.067**	-2.960***	-0.195***	0.049***	-1.797***
	(0.031)	(0.616)	(0.052)	(0.016)	(0.342)
J2J cont. hir. rate (θ_5^J)	-0.082**	-9.140***	-0.174***	0.035***	-6.984***
	(0.033)	(1.755)	(0.043)	(0.009)	(1.349)
J2J cont. sep. rate (θ_6^J)	-0.108***	-8.855***	-0.167***	0.024***	-7.386***
	(0.035)	(1.489)	(0.045)	(0.009)	(1.153)
J2J b.n.e. hir. rate (θ_7^J)	-0.077**	-14.152***	-0.188***	0.043***	-9.706***
	(0.032)	(2.947)	(0.048)	(0.013)	(1.953)
J2J b.n.e. sep. rate (θ_8^J)	-0.089***	-15.655***	-0.181***	0.037***	-11.203***
	(0.031)	(2.968)	(0.047)	(0.013)	(1.868)
B. QWI fluidity monopsony proxies (74,139 obs.)					
Turnover rate (θ_4^Q)	-0.064*	-3.229***	-0.177***	0.049***	-1.646***
	(0.033)	(0.813)	(0.046)	(0.014)	(0.457)
Stable hir. rate (θ_5^Q)	-0.038	-4.518***	-0.172***	0.053***	-2.030***
	(0.036)	(1.059)	(0.046)	(0.016)	(0.558)
Stable sep. rate (θ_6^Q)	-0.036	-4.063***	-0.180***	0.057***	-1.314
	(0.036)	(1.106)	(0.046)	(0.017)	(0.804)
Stable turn. rate (θ_7^Q)	-0.040	-4.482***	-0.176***	0.054***	-1.786**
	(0.036)	(1.132)	(0.046)	(0.017)	(0.684)

Notes: This table reports $\hat{\beta}$ and $\hat{\delta}$ from the estimation of specification (2) under $\gamma = 0$ (restricted model), and $\hat{\beta}$, $\hat{\gamma}$, and $\hat{\delta}$ from the unrestricted estimation of specification (2) for the restaurant industry using 2001-2019 QWI data and different monopsony power proxies. Regressions are weighted by initial population. Standard errors (in parentheses) are clustered at the state level. The coefficients are statistically significant at the *10%, **5%, or ***1% level.

Table A-3: TWFE estimation of minimum wage effects on restaurant average earnings for different monopsony power measures

	Restricted ($\gamma = 0$)		**Unrestricted**		
	$\hat{\beta}$	$\hat{\delta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$
A. Job-to-Job fluidity monopsony proxies (32,643 obs.)					
J2J Hiring rate (θ_1^J)	0.156***	0.111	0.221***	-0.025***	-0.901
	(0.018)	(0.526)	(0.029)	(0.007)	(0.661)
J2J Sep. rate (θ_2^J)	0.161***	0.274	0.221***	-0.025***	-0.761
	(0.018)	(0.499)	(0.029)	(0.007)	(0.671)
B. QWI fluidity monopsony proxies (74,139 obs.)					
Hiring rate (θ_1^Q)	0.163***	0.107	0.234***	-0.030***	-0.830**
	(0.019)	(0.328)	(0.023)	(0.005)	(0.404)
Separation rate (θ_2^Q)	0.166***	0.234	0.231***	-0.028***	-0.675
	(0.019)	(0.348)	(0.023)	(0.005)	(0.424)
Replacement rate (θ_3^Q)	0.161***	0.034	0.230***	-0.029***	-0.896*
	(0.020)	(0.385)	(0.023)	(0.005)	(0.457)
C. NETS concentration monopsony proxies (75,151 obs.)					
Est. p/worker (θ_1^N)	0.169***	1.657	0.236***	-0.021***	1.607
	(0.025)	(1.119)	(0.023)	(0.006)	(1.014)
Firms p/worker (θ_2^N)	0.176***	1.747*	0.236***	-0.020***	1.452
	(0.028)	(0.991)	(0.023)	(0.006)	(0.914)
$1 - HHI$ *(est.)* (θ_3^N)	0.183***	-1.016***	0.228***	-0.021***	-0.039
	(0.018)	(0.315)	(0.021)	(0.006)	(0.195)
$1 - HHI$ *(firm)* (θ_4^N)	0.180***	-0.849***	0.226***	-0.022***	0.197
	(0.017)	(0.276)	(0.021)	(0.006)	(0.230)

Notes: This table reports $\hat{\beta}$ and $\hat{\delta}$ from the estimation of specification (2) under $\gamma = 0$ (restricted model), and $\hat{\beta}$, $\hat{\gamma}$, and $\hat{\delta}$ from the unrestricted estimation of specification (2) for the restaurant industry using 2001-2019 QWI data and different monopsony power proxies. Regressions are weighted by initial population. Standard errors (in parentheses) are clustered at the state level. The coefficients are statistically significant at the *10%, **5%, or ***1% level.

Table A-4: TWFE estimation of minimum wage effects on restaurant average earnings for the remaining fluidity monopsony-power measures

	Restricted ($\gamma = 0$)		**Unrestricted**		
	$\hat{\beta}$	$\hat{\delta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$
A. Job-to-Job fluidity monopsony proxies (32,643 obs.)					
Overall hir. rate (θ_3^J)	0.157***	0.086	0.220***	-0.024***	-0.487
	(0.018)	(0.284)	(0.029)	(0.007)	(0.388)
Overal sep. rate (θ_4^J)	0.158***	0.119	0.218***	-0.023***	-0.439
	(0.019)	(0.291)	(0.028)	(0.007)	(0.384)
J2J cont. hir. rate (θ_5^J)	0.156***	0.117	0.222***	-0.025***	-1.471
	(0.018)	(0.815)	(0.029)	(0.007)	(1.029)
J2J cont. sep. rate (θ_6^J)	0.160***	0.387	0.221***	-0.025***	-1.138
	(0.018)	(0.732)	(0.029)	(0.008)	(1.010)
J2J b.n.e. hir. rate (θ_7^J)	0.158***	0.525	0.219***	-0.023***	-1.959
	(0.018)	(1.416)	(0.028)	(0.007)	(1.737)
J2J b.n.e. sep. rate (θ_8^J)	0.161***	0.903	0.219***	-0.024***	-1.974
	(0.018)	(1.537)	(0.028)	(0.007)	(1.905)
B. QWI fluidity monopsony proxies (74,139 obs.)					
Turnover rate (θ_4^Q)	0.165***	0.171	0.233***	-0.029***	-0.785*
	(0.019)	(0.345)	(0.023)	(0.005)	(0.429)
Stable hir. rate (θ_5^Q)	0.170***	0.738	0.230***	-0.024***	-0.390
	(0.018)	(0.520)	(0.022)	(0.006)	(0.620)
Stable sep. rate (θ_6^Q)	0.171***	0.741	0.230***	-0.024***	-0.398
	(0.018)	(0.503)	(0.022)	(0.005)	(0.580)
Stable turn. rate (θ_7^Q)	0.171***	0.780	0.230***	-0.024***	-0.401
	(0.018)	(0.531)	(0.022)	(0.005)	(0.630)

Notes: This table reports $\hat{\beta}$ and $\hat{\delta}$ from the estimation of specification (2) under $\gamma = 0$ (restricted model), and $\hat{\beta}$, $\hat{\gamma}$, and $\hat{\delta}$ from the unrestricted estimation of specification (2) for the restaurant industry using 2001-2019 QWI data and different monopsony power proxies. Regressions are weighted by initial population. Standard errors (in parentheses) are clustered at the state level. The coefficients are statistically significant at the *10%, **5%, or ***1% level.

B Border Discontinuity Designs

For robustness, this section presents two different border discontinuity approaches: panel estimation with pair-period fixed effects as in Dube, Lester, and Reich (2010), and an event study estimation with pair-period effects as in Dube, Reich, Bhatt, and Sosinskiy (2025b).

B.1 Panel Pair-Period Fixed Effects

In an influential minimum wage paper, Dube, Lester, and Reich (2010) (DLR hereafter) argue that conventional TWFE regressions similar to specification (1), but with $\gamma = 0$, may yield biased results because they do not control for time-varying local economic shocks that are correlated with minimum wage changes. In a generalization of Card and Krueger (1994), they define local economic areas as pairs of U.S. counties straddling state borders. Using 316 such pairs with quarterly data from 1990 to 2006, and including pair-period fixed effects to control for time-varying spatial heterogeneity, they find null effects of minimum wages on restaurant employment.

In a recent re-evaluation of DLR, Jha, Neumark, and Rodriguez-Lopez (2024) show that how we define local economic areas matters: just because two counties are adjacent does not mean one is a good control for the other, as they may not be subject to the same local shocks. They show that when local areas are instead defined using multi-state commuting zones, a DLR-style specification again yields significant negative employment effects of minimum wages.[22] Importantly, each commuting zone—like each CBSA—is defined as a collection of counties linked by strong commuting ties, based on Journey-to-Work data.[23]

Following this idea, we verify the robustness of our results from Table 3 by estimating a DLR-style specification using the 61 multi-state CBSAs. Formed by 133 CBSA-by-state geographies (see Section 2.1), these MS-CBSAs generate 85 cross-state pairs: 52 from two-state CBSAs, 21 from three-state CBSAs, and 12 pairs from four-state CBSAs. Our pair-period specification is given by

$$\ln e_{ipt} = \alpha + \left[\beta + \gamma\left(\ln P_i - \overline{\ln P}\right) + \delta\left(\theta_i - \overline{\theta}\right)\right]\ln MW_{it} + \rho \ln E_{it}^{-} + \zeta \ln P_{it} + \eta_i + \tau_{pt} + \nu_{ipt}, \quad \text{(B-1)}$$

where subscript p denotes a pair—a geography i will belong to more than one pair in MS-CBSAs that span three or more states. The term τ_{pt} captures pair-period effects, which control for time-varying spatial heterogeneity, and ν_{ipt} is the error term.

[22] Using more complete data from 1990 to 2016, Jha, Neumark, and Rodriguez-Lopez (2024) also find that restricting the county pair sample to pairs within multi-state commuting zones similarly yields negative and significant minimum wage effects on employment.

[23] Writing around the same time as DLR, Allegretto, Dube, and Reich (2009) were the first to propose using multi-state commuting zones as local areas instead of county pairs.

With complete data for all 76 quarters, a pair-period specification can have at most 12,920 observations: two observations for each of the 85 pairs in each quarter ($2 \times 85 \times 76$). However, as mentioned in Section 2, some states enter the QWI only after 2001, and we have J2J monopsony proxy measures for only 436 (out of 991) CBSA-by-state entities. In the end, our pair-approach regressions include 9,882 observations from 68 pairs (drawn from 46 MS-CBSAs with 102 CBSA-by-state entities) when using J2J fluidity proxies; 12,266 observations from 84 pairs (from 60 MS-CBSAs with 131 entities) when using QWI fluidity proxies; and 12,290 observations from 85 pairs (from 61 MS-CBSAs with 133 entities) when using NETS concentration proxies. Table B-1 presents the results from estimating (B-1), both with and without the restriction that $\gamma = 0$, using our 9 selected proxies for monopsony power.

As before, the specifications in Table B-1 show an upward bias in $\hat{\beta}$ in the restricted regressions, which even leads to positive and significant estimates when using NETS proxies. Comparing Table B-1 with Table 3, one of the main contrasts is that the estimated coefficients for the size interaction, $\hat{\gamma}$, are statistically insignificant in 6 out of the 9 unrestricted regressions in Table B-1. This is largely a consequence of sample selection when using MS-CBSAs. While the average working-age population in 2001 was 178,093 across the 991 CBSA-by-state entities, it was 464,639 across the 102 entities used in the J2J pair regressions, 366,859 across the 131 entities used in the QWI pair regressions, and 361,498 across the 133 entities in the NETS pair regressions. Because coefficients are identified from within MS-CBSA variation, and the sample consists mostly of large entities on both sides of the border, this is likely why we do not obtain the significant size effect documented in Sections 3.2 and 4.

Although the estimates of β in the unrestricted specifications are less precise—also a consequence of the pair samples consisting mostly of large entities, where minimum wages are less likely to bind—the 9 estimated coefficients in Table B-1 are all negative, ranging from -0.186 to -0.049. Hence, the pair-approach regressions also point to negative effects of minimum wages on employment for an average-sized location with an average level of monopsony power.

Note that the estimates of δ for our J2J and QWI fluidity proxies are negative in both the restricted and unrestricted models, consistent with the results in Table 3. Including the additional fluidity measures from Table B-2, these estimates are generally less precise than those in Table 3, but the pair-approach regressions using fluidity measures still convey the same message: more fluid labor markets are associated with stronger negative employment effects of minimum wages, whereas less fluid (more monopsonistic) labor markets exhibit weaker effects. In addition, the estimates of δ are biased downward when the city size interaction is excluded.

Table B-1: Pair-approach estimation of minimum wage effects on restaurant employment for different monopsony power measures

	Restricted ($\gamma = 0$)		**Unrestricted**		
	$\hat{\beta}$	$\hat{\delta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$
A. Job-to-Job fluidity monopsony proxies (9,882 obs.)					
J2J Hiring rate (θ_1^J)	-0.091	-6.466***	-0.144*	0.032	-4.455
	(0.058)	(2.318)	(0.083)	(0.030)	(3.015)
J2J Sep. rate (θ_2^J)	-0.105*	-6.141***	-0.151*	0.031	-4.214
	(0.055)	(2.162)	(0.078)	(0.031)	(2.949)
B. QWI fluidity monopsony proxies (12,266 obs.)					
Hiring rate (θ_1^Q)	-0.026	-2.518**	-0.101	0.029	-2.103*
	(0.061)	(1.090)	(0.107)	(0.028)	(1.201)
Separation rate (θ_2^Q)	0.003	-1.640	-0.091	0.037	-1.113
	(0.067)	(1.574)	(0.091)	(0.030)	(1.945)
Replacement rate (θ_3^Q)	-0.103	-4.611***	-0.186	0.031	-4.222***
	(0.066)	(1.431)	(0.112)	(0.026)	(1.520)
C. NETS concentration monopsony proxies (12,290 obs.)					
Est. p/worker (θ_1^N)	0.074**	1.301	-0.075	0.053	3.557
	(0.030)	(1.645)	(0.124)	(0.032)	(2.173)
Firms p/worker (θ_2^N)	0.081**	1.846	-0.084	0.061*	4.615**
	(0.032)	(1.534)	(0.129)	(0.031)	(1.872)
$1 - $ *HHI (est.)* (θ_3^N)	0.076*	-0.518	-0.049	0.045*	-1.212
	(0.045)	(1.827)	(0.106)	(0.025)	(1.884)
$1 - $ *HHI (firm)* (θ_4^N)	0.069	-0.153	-0.051	0.043*	-0.786
	(0.048)	(1.889)	(0.108)	(0.025)	(1.889)

Notes: This table reports $\hat{\beta}$ and $\hat{\delta}$ from the estimation of specification (B-1) under $\gamma = 0$ (restricted model), and $\hat{\beta}$, $\hat{\gamma}$, and $\hat{\delta}$ from the unrestricted estimation of specification (B-1) for the restaurant industry using 2001-2019 QWI data and different monopsony power proxies. Regressions are weighted by initial population. Standard errors (in parentheses) are two-way clustered at the state and border segment levels. The coefficients are statistically significant at the *10%, **5%, or ***1% level.

Table B-2: Pair-approach estimation of minimum wage effects on restaurant employment for the remaining fluidity monopsony-power measures

	Restricted ($\gamma = 0$)		**Unrestricted**		
	$\hat{\beta}$	$\hat{\delta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$
A. Job-to-Job fluidity monopsony proxies (9,882 obs.)					
Overall hir. rate (θ_3^J)	-0.094	-3.748*	-0.146	0.035	-2.376
	(0.077)	(2.026)	(0.092)	(0.030)	(2.165)
Overal sep. rate (θ_4^J)	-0.073	-3.345	-0.135	0.035	-2.129
	(0.081)	(2.320)	(0.102)	(0.029)	(2.256)
J2J cont. hir. rate (θ_5^J)	-0.091*	-10.718***	-0.143*	0.030	-7.716
	(0.051)	(3.307)	(0.082)	(0.031)	(4.779)
J2J cont. sep. rate (θ_6^J)	-0.106**	-9.187***	-0.150*	0.030	-6.389
	(0.051)	(3.002)	(0.077)	(0.031)	(4.448)
J2J b.n.e. hir. rate (θ_7^J)	-0.063	-13.147*	-0.126	0.035	-8.124
	(0.072)	(7.740)	(0.093)	(0.029)	(8.185)
J2J b.n.e. sep. rate (θ_8^J)	-0.087	-16.434*	-0.140	0.033	-10.776
	(0.069)	(8.206)	(0.089)	(0.030)	(9.036)
B. QWI fluidity monopsony proxies (12,266 obs.)					
Turnover rate (θ_4^Q)	-0.022	-2.348*	-0.104	0.032	-1.851
	(0.060)	(1.251)	(0.101)	(0.029)	(1.513)
Stable hir. rate (θ_5^Q)	0.025	-1.729	-0.088	0.038	-1.455
	(0.058)	(2.039)	(0.087)	(0.029)	(2.320)
Stable sep. rate (θ_6^Q)	0.049	-0.734	-0.070	0.040	-0.513
	(0.075)	(3.061)	(0.083)	(0.028)	(3.228)
Stable turn. rate (θ_7^Q)	0.035	-1.347	-0.081	0.039	-1.069
	(0.067)	(2.608)	(0.083)	(0.029)	(2.862)

Notes: This table reports $\hat{\beta}$ and $\hat{\delta}$ from the estimation of specification (B-1) under $\gamma = 0$ (restricted model), and $\hat{\beta}$, $\hat{\gamma}$, and $\hat{\delta}$ from the unrestricted estimation of specification (B-1) for the restaurant industry using 2001-2019 QWI data and different monopsony power proxies. Regressions are weighted by initial population. Standard errors (in parentheses) are two-way clustered at the state and border segment levels. The coefficients are statistically significant at the *10%, **5%, or ***1% level.

Regarding our NETS concentration-based monopsony proxies, the four estimates for δ in Table B-1 using either establishments or firms per worker—θ_1^N and θ_2^N—remain positive, with one of them statistically significant. Thus, lower concentration (i.e., higher values of θ_1^N and θ_2^N) is not associated with more adverse employment effects of minimum wages. For the $1-\text{HHI}$ measures, all estimates for δ are negative, but they have large standard errors.

B.2 Event Study Estimation

Following Dube, Reich, Bhatt, and Sosinskiy (2025b), this section introduces a border-approach implementation of the local-projections DiD estimator of Dube, Girardi, Jordà, and Taylor (2025a). A border discontinuity design permits a transparent definition of events at the border-segment level, which mitigates the window truncation caused by federal minimum wage increases. In defining events in terms of sustained changes in the cross-border minimum-wage differential—rather than level changes in a single state's minimum wage—we identify the high-minimum-wage side of the border as the treated unit. The cross-border design allows us to extend the post-treatment window for a state minimum wage increase even if the minimum wage increases on the other side of the border (whether because of a federal or state increase), as long as the cross-border gap remains above a chosen threshold.[24]

We first describe our event-study specifications, which capture heterogeneous minimum wage effects by city size and monopsony power, then we describe our events and controls, and finally we present our results.

B.2.1 Event Study Specifications

Consider a time window $\{t-R,\ldots,t-1,t,t+1,\ldots,t+T\}$ around event time t, where $k \in \{-R,...,-1,0,1,...,T\}$ indexes time relative to t. The event-study specification analogous to the pair-approach specification in (B-1) is

$$\ln e_{ipt+k} = \alpha + \left[\beta + \gamma\left(\ln P_i - \overline{\ln P}\right) + \delta\left(\theta_i - \overline{\theta}\right)\right] D_{it+k} + \rho \ln E^-_{it+k} + \zeta \ln P_{it+k} + \eta_i + \tau_{pt+k} + \nu_{ipt+k}, \tag{B-2}$$

where $D_{it+k}=1$ for $k\in\{0,\ldots,T\}$ if CBSA-by-state geography i receives the treatment and is zero otherwise. All remaining variables and terms are defined as before. The LP–DiD pooled specification is then

$$\overline{\Delta \ln e_{ipt}} = \left[\beta + \gamma\left(\ln P_i - \overline{\ln P}\right) + \delta\left(\theta_i - \overline{\theta}\right)\right] D_{it} + \rho\overline{\Delta \ln E^-_{it}} + \zeta\overline{\Delta \ln P_{it}} + \lambda_{pt} + \varepsilon_{ipt}, \tag{B-3}$$

[24] In addition, in this framework common federal changes that affect both sides equally are absorbed as pair-level shocks and difference out mechanically.

where $\overline{\Delta y_{it}}$ denotes the average difference of variable y over the post-treatment window relative to $t-1$, i.e., $\overline{\Delta y_{it}} = \frac{1}{T+1}\sum_{k=0}^{T}(y_{it+k} - y_{it-1})$, $\lambda_{pt} \equiv \overline{\Delta\tau_{pt}}$ is a pair–event-window fixed effect, and $\varepsilon_{ipt} \equiv \overline{\Delta\nu_{ipt}}$ is the error term. For a treated geography with $P_i = \overline{P}$ and $\theta_i = \overline{\theta}$, the coefficient $\hat{\beta}$ is the estimated average treatment effect over the post-treatment period.

To obtain event-study plots, we estimate a separate regression for each horizon k:

$$\ln e_{ipt+k} - \ln e_{ipt-1} = \left[\beta_k + \gamma_k\left(\ln P_i - \overline{\ln P}\right) + \delta_k\left(\theta_i - \overline{\theta}\right)\right] D_{it} + \rho_k\left(\ln E^{-}_{it+k} - \ln E^{-}_{it-1}\right) + \zeta_k\left(\ln P_{it+k} - \ln P_{it-1}\right) + \left(\tau_{pt+k} - \tau_{pt-1}\right) + \left(\nu_{ipt+k} - \nu_{ipt-1}\right), \tag{B-4}$$

with the property that, when all events share the same post-treatment window length T, the pooled estimate satisfies $\hat{\beta} = \frac{1}{T+1}\sum_{k=0}^{T}\hat{\beta}_k$, and the same aggregation holds for all other coefficients.

We restrict attention to events with complete post-treatment windows.[25] The preferred specifications in Dube, Reich, Bhatt, and Sosinskiy (2025b) omit time-varying covariates, as including them can bias the LP–DiD estimates when they are correlated with treatment. In line with this concern, most of our specifications impose $\rho = \zeta = 0$, and we treat regressions that include $\overline{\Delta \ln E^{-}_{it}}$ and $\overline{\Delta \ln P_{it}}$ as robustness checks. By contrast, our main specification allows for systematic heterogeneity in treatment effects through interactions of the treatment indicator with pre-treatment characteristics, population size (P_i) and monopsony power (θ_i), so that β captures the effect for a treated geography with $P_i = \overline{P}$ and $\theta_i = \overline{\theta}$, while γ and δ parameterize how the effect varies with these pre-treatment characteristics.

Our treatment interactions with P_i and θ_i rely on parallel trends conditional on these pre-treatment characteristics. This assumption is not demanding when pair-period effects are included, since a central advantage of the border-discontinuity design is that properly defined border counterparts provide strong controls. Once pair-period effects are introduced, which become pair–event-window effects when moving from (B-2) to (B-3), identification of the treatment-effect coefficients in a specification without time-varying controls comes exclusively from treated pairs: the only relevant comparison counterfactuals are the border counterparts to the treated sides. Hence, untreated pairs only contribute to the estimation of standard errors in our main specification.

B.2.2 Events and Controls

Given our pair approach, we define events at the border-segment level. An event satisfies two conditions: (1) both sides of the border have the same minimum wage in every quarter during the

[25] Dube, Reich, Bhatt, and Sosinskiy (2025b) allow post-treatment window lengths to vary by event. As shown by Jha, Neumark, and Rodriguez-Lopez (2025), this causes overweighting of the earlier periods—especially the time of impact—in the LP-DiD pooled estimate.

three years before the event quarter; and (2) in the event quarter the minimum wage may change on one or both sides, but the resulting minimum wage on the treated side is at least 5% higher than on the control side, and in every quarter over the subsequent five years this gap remains at or above 5%, so that the treated side's minimum wage never falls below the control side's.

Under these conditions, we obtain 16 border-segment events from 9 state events: Illinois (2004-q1), New York (2013-q4), New Jersey (2014-q1), Delaware (2014-q2), Minnesota (2014-q3), Maryland (2015-q1), Nebraska (2015-q1), South Dakota (2015-q1), and West Virginia (2015-q1). Even though the conditions allow both sides of a border segment to increase their minimum wage at the time of the event, in all 16 border-segment events only the treated side increased the minimum wage. These border-segment events span 30 multi-state CBSA pairs.[26] Although all our previous regressions use data through the last quarter of 2019, our event-study specifications extend the sample through the first quarter of 2020 in order to obtain full post-treatment windows for the 2015-q1 events.

Full control pairs stem from border segments where both sides have the same minimum wage throughout a 33-quarter event window (12 pre-treatment quarters and 21 post-treatment quarters). There are 14 border segments that span 18 multi-state CBSA full control pairs. These 18 distinct pairs appear as full control pairs for several of our six event windows (around event quarters 2004-q1, 2013-q4, 2014-q1, 2014-q2, 2014-q3, 2015-q1), for a total of 85 full control pairs. Overall, each of our event-study regressions includes at most 230 observations: 60 observations are from the 30 treated pairs (one treated side and one control side), and 170 observations are from the 85 control pairs. As mentioned above, only the 60 observations from the treated pairs identify the treatment-effect estimates in the regressions without time-varying controls, although the remaining observations also contribute to the estimation of standard errors.[27]

B.2.3 Event Study Results

To focus first on the effects of city size, we abstract from monopsony power and report in Table B-3 the results from estimating specification (B-3) under $\delta = 0$, with and without the city-size interaction. Panel A shows a negative and significant pooled estimate in the unweighted regression

[26] In chronological order, the border-segment events and the number of multi-state CBSA pairs they span are: IL–IN (1), IL–IA (3), IL–KY (1), IL–MO (4), NY–NJ (1), NY–PA (1), NJ–PA (3), DE–PA (1), MN–ND (3), MN–WI (3), MD–PA (1), MD–VA (1), NE–IA (2), SD–IA (1), WV–KY (1), and WV–VA (3). Note that there is an NY–NJ border-segment event from a New York increase in 2013-q4, and then an NJ–PA border-segment event from a New Jersey increase in 2014-q1. This does not violate our conditions, as the NY–NJ minimum wage gap remained above the 5% threshold in favor of New York throughout the entire post-period for the 2013-q4 event.

[27] For the estimation of the pooled LP–DiD specification (B-3), the 230 observations are constructed from the $230 \times 21 = 4{,}830$ quarterly observations in the post-treatment period.

Table B-3: Event study estimation of minimum wage effects on restaurant employment without and with city size interaction

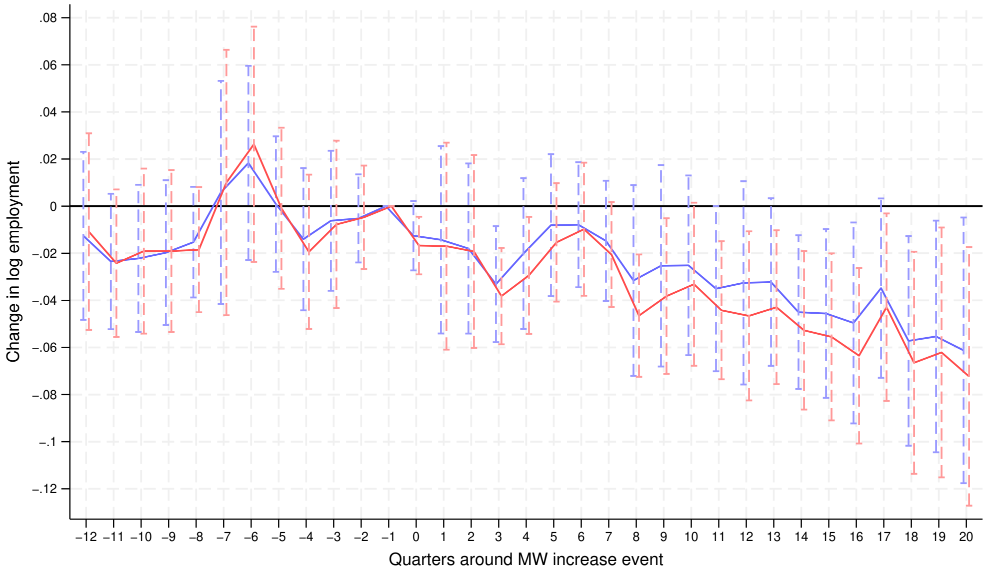
	(1)	(2)	(3)
A. Without city size interaction			
Event indicator	-0.031*	-0.029	0.007
	(0.018)	(0.018)	(0.015)
B. With city size interaction			
Event indicator	-0.040**	-0.038**	-0.042***
	(0.015)	(0.016)	(0.013)
Event $\times$ $(\ln P - \overline{\ln P})$	0.016**	0.015**	0.013***
	(0.006)	(0.007)	(0.004)
Pair-window effects	Y	Y	Y
Controls		Y	
Population weights			Y

Notes: This table reports $\hat{\beta}$ from the estimation of specification (B-3) under $\gamma = 0$ and $\delta = 0$ in Panel A, and $\hat{\beta}$ and $\hat{\gamma}$ from the estimation of specification (B-3) under $\delta = 0$ in Panel B. All regressions include 230 observations: 60 observations from the 30 treated pairs (one treated side and one control side) and 170 observations from the 85 control pairs. Column (1) presents the unweighted estimation without controls, column (2) includes controls, and column (3) presents the weighted estimation (using population at the beginning of the window as weights) without controls. Standard errors (in parentheses) are clustered at the border-segment level. Coefficients are statistically significant at the *10%, **5%, or ***1% level.

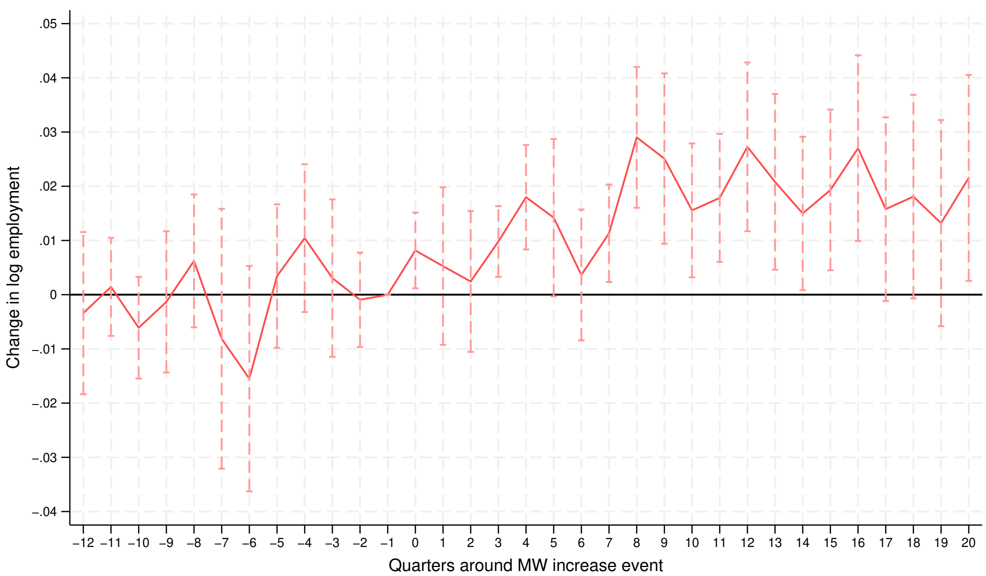
without controls (column (1)), which becomes insignificant when including controls (column (2)) and switches to a near-zero positive coefficient when we use population weights (column (3)).[28] Panel B shows that once we include the city-size interaction, the estimated coefficient for β is negative and significant (and similar in size) across all specifications, while the coefficient on the city-size interaction, γ, is positive and significant in all cases. Despite the sample selection toward larger entities when using MS-CBSAs (as documented in Section B.1), which would tend to underestimate γ, the results in Table B-3 strongly confirm our main findings from the TWFE estimation in Table 2.

Figure B-1 shows the event-study plots from estimating specification (B-4) corresponding to the unweighted specification in column (1) of Table B-3—as mentioned above, the average of the estimated coefficients over the post-treatment period equals the pooled estimate. Figure B-1a not

[28] Our TWFE regressions in the main body of the paper use initial population weights. For the weighted event-study regressions we instead use as weights the population at the beginning of the window (at $k = -R$).



(a) Event effect ($\hat{\beta}$) – without (blue) and with (red) city-size interaction



(b) Event×city-size effect ($\hat{\gamma}$)

Figure B-1: Event study estimates of employment effects of minimum wages with 90 percent confidence bands

only shows that excluding the city-size interaction understates the magnitude of β, but also that the estimated effect becomes more negative over time, reaching -0.072 by the end of the post-treatment period in the estimation with unrestricted γ. Figure B-1b shows a city-size effect with an increasing trend over the first eight quarters, reaching 0.029, and then fluctuating between 0.013 and 0.027 for the rest of the post-window. The figures do not show any evidence of pre-trends for either β or γ.

Table B-4 presents the results from estimating (B-3) using different measures of monopsony power. As in Table B-3, we report unweighted estimates with and without controls, and weighted estimates without controls. Across all specifications, the estimate for β is negative and significant. The estimates for γ are all positive, but they are insignificant in 11 of the 27 regressions—and, in particular, in the 6 regressions that use the J2J fluidity proxies. This reflects the smaller J2J sample and its bias toward larger CBSA-by-state entities (see Section B.1), which in this case uses only 23 out of 30 treated pairs and 190 observations in total. By contrast, the estimates for δ are all negative and mostly significant when using fluidity measures (Panels A and B), and positive and significant when using NETS establishments or firms per worker, while they are generally insignificant with conflicting signs when using the NETS HHI measures. These results confirm our findings from the TWFE estimation in Table 3.[29]

[29] Table B-5 shows the event-study estimation results using the remaining fluidity measures, which are similar to those in Panels A–B of Table B-4.

Table B-4: Event study estimation of minimum wage effects on restaurant employment for different monopsony-power measures

	Unweighted – No controls			**Unweighted – Controls**			**Weighted – No controls**		
	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$
A. Job-to-Job fluidity monopsony proxies (190 obs.)									
J2J Hiring rate (θ_1^J)	-0.039**	0.006	-1.456*	-0.036*	0.003	-1.587**	-0.036**	0.002	-1.343*
	(0.017)	(0.007)	(0.748)	(0.017)	(0.008)	(0.705)	(0.014)	(0.008)	(0.742)
J2J Sep. rate (θ_2^J)	-0.038**	0.006	-1.252*	-0.034*	0.002	-1.358**	-0.035***	0.001	-1.195*
	(0.017)	(0.007)	(0.665)	(0.017)	(0.008)	(0.626)	(0.013)	(0.007)	(0.631)
B. QWI fluidity monopsony proxies (230 obs.)									
Hiring rate (θ_1^Q)	-0.066***	0.018***	-0.721*	-0.065***	0.017**	-0.764*	-0.040***	0.009***	-0.415***
	(0.019)	(0.006)	(0.394)	(0.020)	(0.007)	(0.394)	(0.011)	(0.003)	(0.084)
Separation rate (θ_2^Q)	-0.047**	0.016**	-0.255	-0.045*	0.015**	-0.256	-0.043***	0.009***	-0.488**
	(0.022)	(0.006)	(0.458)	(0.023)	(0.007)	(0.462)	(0.012)	(0.003)	(0.212)
Replacement rate (θ_3^Q)	-0.067***	0.018***	-0.773**	-0.066***	0.017**	-0.822**	-0.039***	0.006	-0.697***
	(0.022)	(0.006)	(0.334)	(0.023)	(0.007)	(0.339)	(0.011)	(0.004)	(0.179)
C. NETS concentration monopsony proxies (230 obs.)									
Est. p/worker (θ_1^N)	-0.059***	0.024***	1.096**	-0.058***	0.024***	1.204**	-0.037***	0.012***	1.152**
	(0.019)	(0.007)	(0.463)	(0.020)	(0.007)	(0.473)	(0.013)	(0.003)	(0.420)
Firms p/worker (θ_2^N)	-0.058***	0.025***	1.026**	-0.057***	0.024***	1.129**	-0.035**	0.012***	1.061***
	(0.019)	(0.007)	(0.456)	(0.020)	(0.007)	(0.464)	(0.013)	(0.003)	(0.364)
$1 - $ *HHI (est.)* (θ_3^N)	-0.036**	0.012	0.332	-0.035*	0.012	0.263	-0.036***	0.015***	-0.567
	(0.017)	(0.010)	(0.555)	(0.018)	(0.011)	(0.621)	(0.012)	(0.005)	(0.512)
$1 - $ *HHI (firm)* (θ_4^N)	-0.037**	0.012	0.338	-0.035*	0.012	0.273	-0.040***	0.014**	-0.232
	(0.017)	(0.010)	(0.605)	(0.019)	(0.011)	(0.670)	(0.012)	(0.006)	(0.599)

Notes: This table reports $\hat{\beta}$, $\hat{\gamma}$, and $\hat{\delta}$ from estimating specification (B-3) using different monopsony power proxies. The column headings indicate whether the estimation is unweighted or weighted, and whether controls are included. Standard errors (in parentheses) are clustered at the border-segment level. Coefficients are statistically significant at the *10%, **5%, or ***1% level.

Table B-5: Event study estimation of minimum wage effects on restaurant employment for the remaining fluidity monopsony-power measures

	Unweighted − no controls			**Unweighted − controls**			**Weighted − no controls**		
	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$
A. Job-to-Job fluidity monopsony proxies (190 obs.)									
Overall hir. rate (θ_3^J)	-0.055***	0.007	-1.185*	-0.051**	0.004	-1.162*	-0.047***	0.009	-0.454
	(0.018)	(0.008)	(0.617)	(0.019)	(0.008)	(0.596)	(0.013)	(0.007)	(0.505)
Overal sep. rate (θ_4^J)	-0.052**	0.008	-1.069*	-0.048**	0.005	-1.087*	-0.045***	0.011	-0.250
	(0.019)	(0.007)	(0.571)	(0.019)	(0.007)	(0.565)	(0.014)	(0.007)	(0.502)
J2J cont. hir. rate (θ_5^J)	-0.037**	0.006	-2.294**	-0.033*	0.002	-2.424**	-0.030**	-0.001	-2.330**
	(0.017)	(0.007)	(1.058)	(0.017)	(0.007)	(0.984)	(0.012)	(0.007)	(1.057)
J2J cont. sep. rate (θ_6^J)	-0.035**	0.006	-1.603	-0.031*	0.003	-1.759*	-0.033***	-0.001	-1.942*
	(0.017)	(0.007)	(0.989)	(0.017)	(0.008)	(0.923)	(0.011)	(0.007)	(0.955)
J2J b.n.e. hir. rate (θ_7^J)	-0.042**	0.009	-3.346	-0.038**	0.005	-3.885*	-0.043***	0.009	-1.657
	(0.019)	(0.007)	(2.259)	(0.018)	(0.008)	(2.163)	(0.015)	(0.008)	(2.344)
J2J b.n.e. sep. rate (θ_8^J)	-0.046**	0.006	-4.917**	-0.042**	0.002	-5.149**	-0.040**	0.006	-2.577
	(0.016)	(0.007)	(1.974)	(0.017)	(0.008)	(1.918)	(0.015)	(0.007)	(1.619)
B. QWI fluidity monopsony proxies (230 obs.)									
Turnover rate (θ_4^Q)	-0.060***	0.017***	-0.622	-0.059***	0.016**	-0.651	-0.042***	0.009***	-0.471***
	(0.020)	(0.006)	(0.417)	(0.021)	(0.007)	(0.416)	(0.011)	(0.003)	(0.118)
Stable hir. rate (θ_5^Q)	-0.068***	0.017**	-1.297*	-0.065***	0.016**	-1.291*	-0.044***	0.011***	-0.498**
	(0.021)	(0.006)	(0.665)	(0.022)	(0.007)	(0.670)	(0.011)	(0.003)	(0.226)
Stable sep. rate (θ_6^Q)	-0.050**	0.015**	-0.742	-0.047**	0.014**	-0.663	-0.043***	0.011***	-0.496
	(0.022)	(0.006)	(0.799)	(0.022)	(0.006)	(0.776)	(0.012)	(0.003)	(0.469)
Stable turn. rate (θ_7^Q)	-0.062**	0.015**	-1.300*	-0.059**	0.014**	-1.256*	-0.043***	0.011***	-0.537*
	(0.023)	(0.006)	(0.740)	(0.023)	(0.007)	(0.719)	(0.011)	(0.003)	(0.314)

Notes: This table reports $\hat{\beta}$, $\hat{\gamma}$, and $\hat{\delta}$ from estimating specification (B-3) using different monopsony power proxies. The column headings indicate whether the estimation is unweighted or weighted, and whether controls are included. Standard errors (in parentheses) are clustered at the border-segment level. Coefficients are statistically significant at the *10%, **5%, or ***1% level.

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