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## What's Across the Border? Re-Evaluating the Cross-Border Evidence on Minimum Wage Effects

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# What's Across the Border? Re-Evaluating the Cross-Border Evidence on Minimum Wage Effects

## Abstract

**Dube, Lester, and Reich (2010)** argue that state-level minimum wage variation can be correlated with economic shocks, generating spurious evidence that higher minimum wages reduce employment. Using minimum wage variation within contiguous county pairs that share a state border, they find no relationship between minimum wages and employment in the U.S. restaurant industry. We show that this finding hinges critically on using cross-border counties to define local economic areas with which to control for economic shocks that are potentially correlated with minimum wage changes. We use, instead, multi-state commuting zones, which provide superior definitions of local economic areas. Using the same within-local area research design—but within cross-border commuting zones—we find a robust negative relationship between minimum wages and employment.

JEL-Codes: J230, J380.

Keywords: minimum wage, employment, commuting zones.

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# 1 Introduction

In an influential paper in the research literature on the employment effects of the minimum wage in the United States, [Dube, Lester, and Reich \(2010\)](#)—DLR hereafter—find a small, non-negative, and insignificant relationship between minimum wages and employment in the U.S. restaurant industry. The core evidence in DLR is that the negative relationship between employment and minimum wages in that industry disappears—flipping to a positive, but not significant, coefficient—when using a specification that is intended to control for time-varying local economic conditions by using minimum wage variation within cross-border county pairs.

DLR’s study is cited frequently (1,391 Google scholar cites as of April 25, 2022). More significantly, it is often cited in both policy debate and research summaries as one of the key papers overturning the prediction from the competitive labor market model—and from a great deal of evidence (see, for example, [Neumark and Wascher, 2008](#))—that a higher minimum wage reduces employment among lower-skilled workers.<sup>1</sup>

In short, the key contention of DLR is that one needs local controls to credibly identify the employment effects of minimum wages. Their general strategy is to study minimum wage variation *within* local economic areas to control for economics shocks that may be correlated with minimum wages. DLR define these local economic areas as contiguous county-pairs sharing a state border. They estimate minimum wage effects using variation within cross-border county pairs, and doing this changes the answer from finding of a negative employment effect near  $-0.2$  (an elasticity) to close to zero. And this result has been widely touted—including by the authors—as overturning the conclusion from a large prior literature, and in arguing for large minimum wage increases.

In this paper, however, we show DLR’s conclusion relies critically on defining the local economic area used to capture spatial economic shocks as pairs of contiguous counties across state lines. If local economic areas are defined instead as commuting zones, which is far more natural—but we otherwise follow the approach of DLR, using as identifying information minimum wage variation within multi-state commuting zones—we find a negative and significant employment elasticity.

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<sup>1</sup>For example, Michael Reich, one of the authors of this study, testified before the U.S. House of Representatives that: “Economists have conducted literally hundreds of studies based on over 160 minimum wage changes in the past thirty-five years. The best of these studies do provide a credible guide to the likely employment effects of a \$15 floor. They indicate that the Act will have minimal to no adverse effects on employment and that they will have substantial positive dynamic effects on the lowest-wage areas of the U.S.” ([IRLE link](#)). And Dube, in dismissing the results of more conventional state-by-year panel data estimates, and citing DLR, writes: “a ‘fourth generation’ of ... papers ... have shown that the variation over the past two decades in minimum wages has been highly selective spatially, and employment trends for low-wage workers vary substantially across states...” and that “approaches such as comparing contiguous counties across policy boundaries ... produce employment effects close to zero” ([Dube, 2011](#), p. 764).

There is compelling reason to believe that commuting zones are better able to net out other economic shocks and hence to capture employment variation driven by cross-state minimum wage differentials than are contiguous county-pairs sharing a state border. As described by [Tolbert and Sizer \(1996\)](#), commuting zones are defined as a “groups of counties with strong commuting ties” based on Census’s *journey-to-work* data, and specifically, “commuting zones are intended for use as spatial measures of local labor markets.” This is not necessarily the case for county pairs: even if they are contiguous, two isolated U.S. counties may share little or no commuting and economic activity.<sup>2</sup> Moreover, we show that cross-border areas of commuting zones in fact have more common economics shocks than do cross-border county pairs. Hence, DLR’s main finding follows from their contiguous-county-pair approach not meeting a definition of “local economic area” that effectively captures common local shocks.

Indeed, in another (unpublished) paper written at about the same time as DLR, two of the three authors of DLR have made the same argument. There, they write that using commuting zones “is appealing because these areas are not only contiguous; they are also demonstrably linked with each other by an economically meaningful criterion” ([Allegretto, Dube, and Reich, 2009](#)).<sup>3</sup>

Our empirical approach proceeds as follows. First, in section 2 we use DLR’s replication package (1990-2006 quarterly data and programs) and re-estimate their contiguous-pair specifications, but compare results using, instead, pairs from multi-state commuting zones. We find that the estimated minimum wage elasticity changes from a non-significant value of 0.016 in DLR, to a significant value of  $-0.141$  when using pairs from multi-state commuting zones, which is very similar to the estimate without the spatial heterogeneity controls.

We also show that this same result holds using more complete data. In particular, in section 3 we estimate similar specifications to DLR but using yearly U.S. County Business Patterns (CBP) data at both the county and commuting zone–state levels for the 1990-2016 period. This dataset, which builds on the CBP commuting-zone datasets of [Autor, Dorn, and Hanson \(2013\)](#) and [Acemoglu, Autor, Dorn, Hanson, and Price \(2016\)](#), has better geographical coverage than DLR’s dataset and covers a longer period. Our CBP data yield a non-significant elasticity of  $-0.081$  when using contiguous county-pairs that share a state border, whereas they yield a significant elasticity of

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<sup>2</sup>For example, there are ten county pairs between Colorado and Kansas in DLR’s list, even though there is not any defined commuting zone between them. By definition, commuting ties between two counties in different commuting zones are weak ([Tolbert and Sizer, 1996](#)).

<sup>3</sup>This paper analyzes teen employment in the U.S. by exploiting minimum wage variation within multi-state commuting zones. For much the same reason, [Liu, Hyclak, and Regmi \(2016\)](#) estimate minimum wage effects using Bureau of Economic Analysis Economic Areas—which are micropolitan or metropolitan statistical areas and the surrounding counties that are economically related—to control for spatial heterogeneity.

−0.242 when using pairs from multi-state commuting zones.

Following DLR, we also explore the long-term effects of minimum wages and the possibility of pre-existing trends affecting our specifications. While the results using county-pairs show a non-significant long-term effect of minimum wages, echoing the findings in DLR, using pairs from multi-state commuting zones we find large and significant negative effects of minimum wages in the medium and long terms.

Finally, we show that using commuting zones as local controls performs as well or better than DLR’s cross-border counties in accounting for sources of changes in employment prior to minimum wage increases. And we present other evidence indicating that commuting zones provide better controls for local shocks.

## 2 Re-Analysis of the **Dube, Lester, and Reich (2010)** Approach

Using DLR’s data and programs (**Dube, Lester, and Reich, 2011**), this section focuses on the replication and re-analysis of DLR’s preferred **specification (6)**, which is given by

$$\ln e_{ipt} = \alpha + \beta \ln MW_{it} + \rho Z_{it} + \eta_i + \tau_{pt} + \nu_{it}, \quad (1)$$

where for entity  $i$  from pair  $p$  in period  $t$ ,  $e_{ipt}$  denotes employment,  $MW_{it}$  is the minimum wage,  $Z_{it}$  is a vector of time-variant entity level controls (which include population and total private sector employment),  $\eta_i$  is an entity  $i$  fixed effect, and  $\tau_{pt}$  denotes pair–time fixed effects, which are intended to control for spatial heterogeneity at the local level.<sup>4</sup> We also show results for DLR’s **specification (5)**, which is a restricted version of (1) with  $\tau_{pt} = \tau_t$  for every pair  $p$ —the more standard two-way fixed effects model that DLR claim gives spurious evidence of negative employment effects. In DLR an entity is a county that shares a state border, whereas we look at multi-state commuting zones, with an entity defined as a commuting zone-state.

DLR use quarterly data from the Quarterly Census of Employment and Wages (QCEW) from 1990 to 2006. They identify 1,139 counties sharing a state border and 1,181 border-county pairs.<sup>5</sup> Of these, DLR only use counties that report information for the restaurant industry for the 66 quarters in their 1990-2006 sample period (due to confidentiality practices, the QCEW suppresses information for many counties), which restricts the sample to 504 border counties. These 504 border counties appear in 754 county pairs, but only 316 pairs are “complete” in the sense that there is

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<sup>4</sup>DLR’s replication package is available at this [link](#).

<sup>5</sup>These include three pairs from California that involve the County of San Francisco, which imposed a different minimum wage than neighboring California counties.

information for both counties in the pair (*i.e.*, 438 of the 754 pairs are “incomplete” because they have information for only one county). From these 754 pairs, DLR build a sample—which they refer to as the contiguous border county-pair (CBCP) sample—with 1,070 observations each period ( $1,070 \times 66 = 70,620$  observations in total), composed of 438 observations from incomplete pairs, and  $316 \times 2 = 632$  observations from complete pairs. Given that a county can appear in multiple contiguous pairs, there are many repeated observations in the CBCP sample, with some counties appearing up to seven times each period.<sup>6</sup>

We create a similar dataset with multi-state commuting zones. From [Tolbert and Sizer \(1996\)](#), there are 137 multi-state commuting zones in the U.S. (out of 741 total commuting zones): 129 two-state zones and 8 three-state zones. These 137 zones yield 282 commuting zone-state entities ( $129 \times 2 + 8 \times 3$ ) and 153 pairs, corresponding to 129 pairs from the two-state zones, and 24 pairs from the three-state zones. To aggregate DLR’s county-level QCEW data to the commuting zone-state level, we use the county-to-commuting zone crosswalk of [Acemoglu, Autor, Dorn, Hanson, and Price \(2016\)](#). We only use counties that report restaurant-industry information in all 66 quarters, which restricts the sample to 184 commuting zone-state entities. These 184 entities appear in 128 pairs, of which 73 are complete and 55 are incomplete. Similar to DLR, from these 128 pairs we build a sample—which we refer to as the multi-state commuting zone-pair (MCZP) sample—with 201 observations each period ( $201 \times 66 = 13,266$  observations in total), composed of 55 observations from incomplete pairs, and  $73 \times 2 = 146$  observations from complete pairs. Given that a commuting zone-state can appear in at most two pairs (if it belongs to a three-state zone), the fraction of repeated observations is smaller in our MCZP sample than in DLR’s CBCP sample (each period, 167 commuting zone-state entities appear once, and 17 appear twice).

Table 1 shows the estimation results for [specifications \(5\) and \(6\)](#) in DLR’s Table 2. As in DLR, in addition to estimates of the minimum wage elasticity of employment, Table 1 also shows estimates of the elasticity of average earnings. Panel A shows DLR’s exact estimated coefficients, which use the CBCP sample, and panel B re-estimates those specifications but using instead the MCZP sample. Following [Cameron, Gelbach, and Miller \(2011\)](#), DLR report two-way clustered standard errors at the state and border segment levels, where a border segment is defined as a pair of states sharing contiguous counties.<sup>7</sup>

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<sup>6</sup>For example, county 32031 (Washoe County, NV) appears seven times each period: two times from complete pairs with California counties 6057 and 6061, five times from three incomplete pairs with California counties (6035, 6049, and 6091) and two times from incomplete pairs with Oregon counties (41025 and 41037).

<sup>7</sup>DLR’s standard errors for their [specification \(6\)](#) presented in their Table 2 are not accurate because they use data from incomplete pairs in their calculation. Specifically, controlling for pair-period effects,  $\tau_{pt}$ , in equation (1) requires complete pairs, and thus, only complete pairs are effectively used in its estimation. Therefore, columns

Table 1: Replication and re-analysis of [Dube, Lester, and Reich \(2010\)](#) using 1990-2006 QCEW data

	<b>Specification (5)</b>		<b>Specification (6)</b>	
	(1)	(2)	(3)	(4)
<b><i>A. DLR's Contiguous Border County-Pair Sample</i></b>				
<b><i>(a) ln(employment)</i></b>				
ln(minimum wage)	-0.137*	-0.112	0.057	0.016
	(0.072)	(0.079)	(0.088)	(0.076)
<b><i>(b) ln(earnings)</i></b>				
ln(minimum wage)	0.232***	0.221***	0.200***	0.189***
	(0.033)	(0.034)	(0.050)	(0.047)
Observations	70,620	70,582	41,712	41,676
<b><i>B. Multi-state Commuting Zone-Pair Sample</i></b>				
<b><i>(a) ln(employment)</i></b>				
ln(minimum wage)	-0.212***	-0.186**	-0.128*	-0.141**
	(0.069)	(0.072)	(0.070)	(0.070)
<b><i>(b) ln(earnings)</i></b>				
ln(minimum wage)	0.232***	0.226***	0.222***	0.208***
	(0.041)	(0.042)	(0.071)	(0.064)
Observations	13,266	13,264	9,636	9,634
Pair-period effects			Y	Y
Total private sector		Y		Y

Notes: This table replicates and re-analyzes the estimation results of specifications (5) and (6) of Table 2 in DLR. It uses DLR's replication package to obtain their exact reported coefficient estimates in panel A. After aggregating DLR's data at the commuting zone-state level, panel B shows the results from the re-estimation of DLR's specifications using instead cross-border pairs from multi-state commuting zones. 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 1 shows that the coefficients from the earnings regressions are very stable across specifications and across samples. But that is not the case for the employment regressions. DLR find that the estimate of the minimum wage elasticity of employment changes from a significant  $-0.137$  in the specification without pair-period effects and no total private sector employment, to

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3 and 4 in Table 1 estimate [specification \(6\)](#) using only complete pairs (the size of the CBCP sample declines to  $316 \times 2 \times 66 = 41,712$  observations, and the size of the MCZP sample declines to  $73 \times 2 \times 66 = 9,636$  observations), and report the estimated coefficients of DLR (in panel A) but corrected standard errors. This standard error correction makes a small but not material difference in the results.

a non-significant 0.016 when using the specification that controls for pair-period effects and total private sector employment. On the other hand, when using instead pairs from multi-state commuting zones, panel B shows that the estimate remains negative, sizable, and significant in DLR’s preferred specification, only declining from  $-0.212$  to  $-0.141$ .

Therefore, the definition of local economic area—whether a pair of contiguous counties or a commuting zone—is crucial for the sign, size, and significance of the minimum wage elasticity of employment. In particular, when we control for spatial heterogeneity by using minimum wage variation within commuting zones—which actually are defined to capture common economic shocks to the labor market—we continue to find evidence that minimum wages reduce employment, with elasticity estimates in the traditional or earlier “consensus” range of  $-0.1$  to  $-0.2$ .

### 3 Estimation Using the County Business Patterns Database

To examine the evidence for more complete data that is thus more generalizable, in this section we analyze the relationship between minimum wages and employment in the restaurant industry using yearly Census CBP data from 1990 to 2016. The CBP data not only include ten more years of data than the contiguous border county-pair (CBCP) and multi-state commuting zone-pair (MCZP) samples, but also has much better geographic coverage. For example, whereas the CBCP and MCZP only include 316 and 73 complete pairs, respectively, the CBP data yields 1,165 complete pairs (out of 1,181) in the county approach and 151 complete pairs (out of 153) in the commuting-zone approach: our pair coverage goes from 26.75 percent to 98.65 percent in the county approach, and from 47.71 percent to 98.69 percent in the commuting-zone approach.

DLR also use CBP data for a robustness check and obtain a non-significant minimum wage elasticity of employment of  $-0.034$ . However, DLR are skeptical about their CBP data because of changes in industry classification (from SIC to NAICS) and the fact that, due to confidentiality reasons, the CBP reports many county-industry cells as an employment range. These problems are minimal in our CBP data, which was processed using the sophisticated fixed-point imputation and industry-classification method of Autor, Dorn, and Hanson (2013) and Acemoglu, Autor, Dorn, Hanson, and Price (2016)—AADHP hereafter.<sup>8</sup>

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<sup>8</sup>AADHP’s employment-imputation algorithm makes use of three pieces of information in the CBP: (i) through 2016, when not exactly reported in a county-industry cell, employment is given as one of 12 ranges, starting from “0-19 employees” and up to “more than 100,000 employees”; (ii) through 2016, the CBP always reports the exact number of establishments in each county-industry cell and splits the total into 12 establishment-size categories, starting from the number of establishments with 0-4 employees up to the number of establishments with more than 5000 employees; and (iii) even if exact employment is not reported for a county-industry cell, the CBP reports employment at higher levels of industry aggregation.

### 3.1 Data Description

From the CBP we obtain yearly employment counts and annual pay from 1990 to 2016.<sup>9</sup> We follow AADHP—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.<sup>10</sup> We also follow AADHP to obtain yearly working-age population at the commuting zone–state level from the Census Bureau’s Population Estimates Program. Last, yearly minimum wage data at the state level—defined as the largest of the federal minimum wage and the state minimum wage—are obtained from [Vaghul and Zipperer \(2016\)](#).<sup>11</sup>

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). The 866 commuting zone–state entities come from 585 single-state commuting zones,  $129 \times 2$  two-state commuting zones, and  $8 \times 3 - 1$  three-state commuting zones.<sup>12</sup> The second dataset—which we label as the CBP-MCZP sample—is analogous to the MCZP sample from section 2, and its purpose is to exploit local differences in minimum wages to control for spatial heterogeneity at the local level. As shown in Figure 1, the 137 multi-state commuting zones are widely distributed across the continental United States.

Throughout the 1990-2016 period, the multi-state commuting zone group accounts on average for 29.8 percent of U.S. employment, 29 percent of U.S. establishments, and 29.4 percent of the U.S. working-age population.<sup>13</sup> 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 2 presents a comparison between the groups of establishment and employment counts, employment-to-populations ratios, earnings per worker, and average minimum wages.

Figures 2a–2c show that both groups follow the same patterns for establishment counts, em-

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<sup>9</sup>Although CBP data can be obtained through 2019, we do not use the 2017-2019 period because starting with the 2017 release, Census’s changes in confidentiality practices no longer allow the implementation of AADHP’s employment-imputation algorithm. In particular, since 2017 the CBP fully omits county-industry cells with less than three establishments, and thus, the CBP no longer shows full establishment counts nor employment ranges for these cells (see previous footnote).

<sup>10</sup>As in AADHP, we exclude Alaska and Hawaii from our analysis.

<sup>11</sup>The minimum wage data are available at this [link](#).

<sup>12</sup>The District of Columbia, which appears in the CBP starting in 2004 and is not included in our analysis, is part of a three-state commuting zone.

<sup>13</sup>These shares are very stable over time. They range between 29.3 and 30.5 percent for employment, 28.8 and 29.4 percent for establishments, and 29.1 and 29.7 percent for the working-age population.

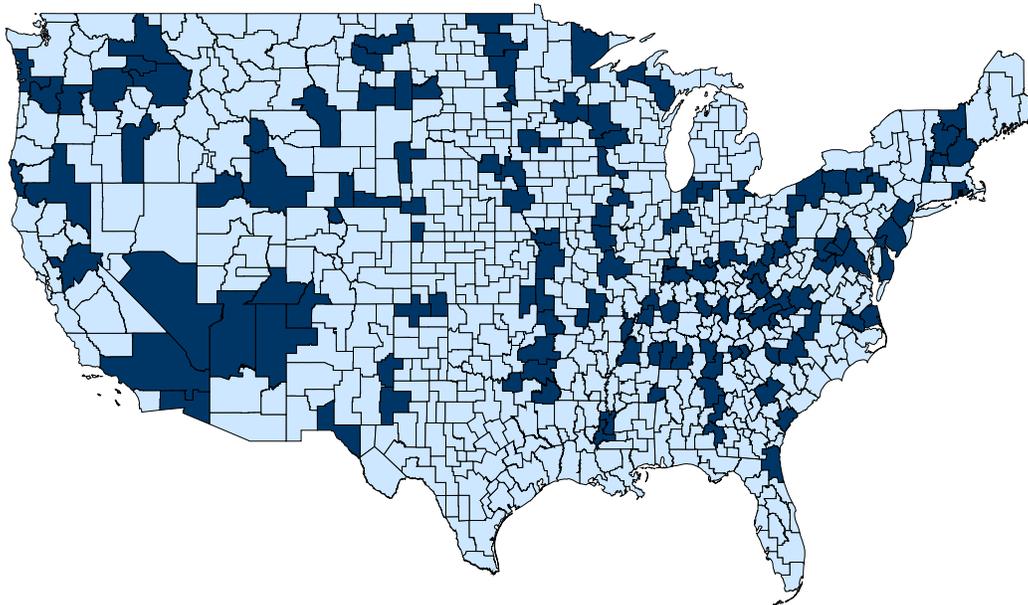


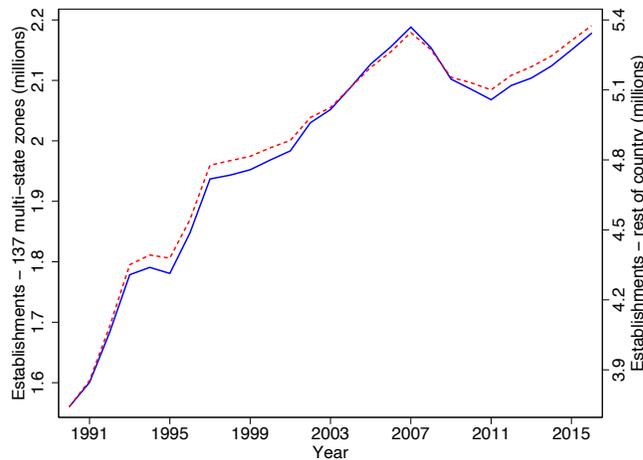
Figure 1: The 137 multi-state commuting zones (dark blue)

ployment 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 55.5 percent for the multi-state group and 54.4 percent for the single-state group. Figure 2d shows similar values and patterns for earnings per worker, calculated for each group as the total annual pay divided by total employment. Finally, Figure 2e shows that the average minimum wage—weighted by commuting zone–state working-age population—has a similar evolution in both groups, increasing from about \$3.86–\$3.93 in 1990 to about \$8.12–\$8.45 in 2016.<sup>14</sup>

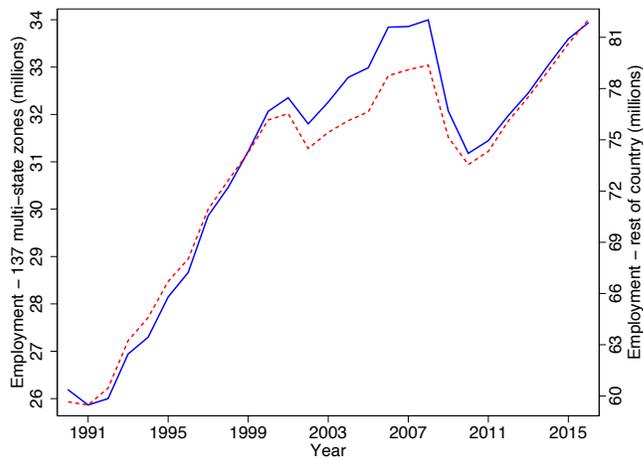
### 3.2 Econometric Specifications

Although in section 2 we focused on DLR’s preferred [specification \(6\)](#) to replicate and re-analyze their findings, here we take advantage of our CBP data. Because the application of these new data to this issue is new, we provide a more complete analysis that covers a larger set of the specifications that DLR report—building up to the key specifications using within-commuting zone minimum wage variation. As in DLR, we start with a simple specification à la [Neumark and Wascher \(1992\)](#)

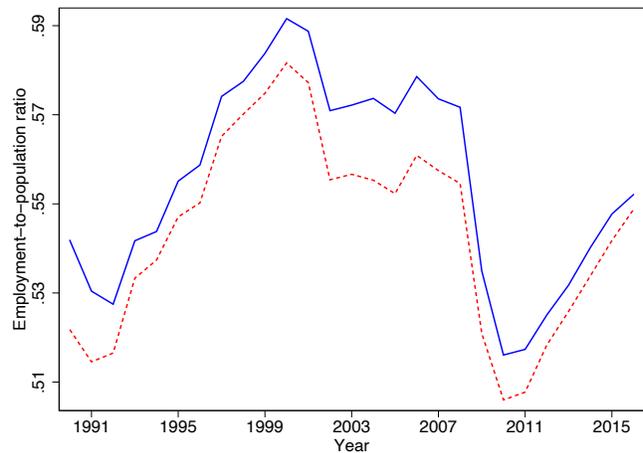
<sup>14</sup>The restaurant industry has code 5812 (Eating and drinking places) in the Standard Industrial Classification used by AADHP. After aggregating the remaining 478 AADHP industries into 19 industries, Table A-1 in the Appendix shows industry-level earnings per worker and earnings rankings in 1990 and 2016. In both years the restaurant industry has the lowest earnings per worker. Moreover, the earnings gap with the retail-trade industry (the second lowest-earnings industry) is large, with the retail-trade industry paying 76 percent more in 1990 and 56 percent more in 2016. Table A-1 also presents industry employment shares, showing that restaurants accounted for 7.21 percent of U.S. employment in 1990 and for 9.37 percent in 2016. Finally, Figure A-1 in the Appendix shows that establishment counts, employment, and earnings in the restaurant industry follow the same patterns in the multi-state commuting zone group and in the rest of the country.



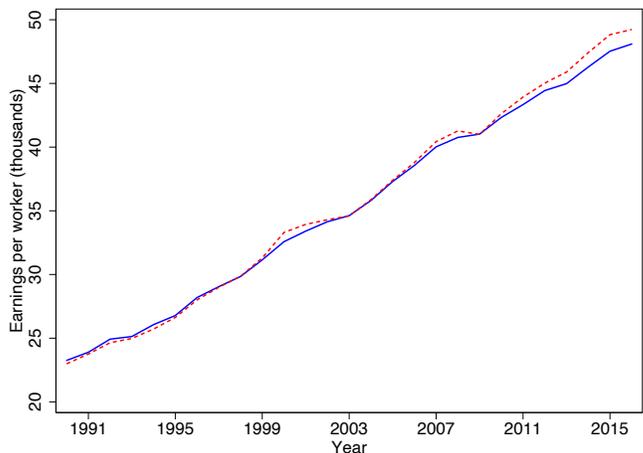
(a) Number of establishments



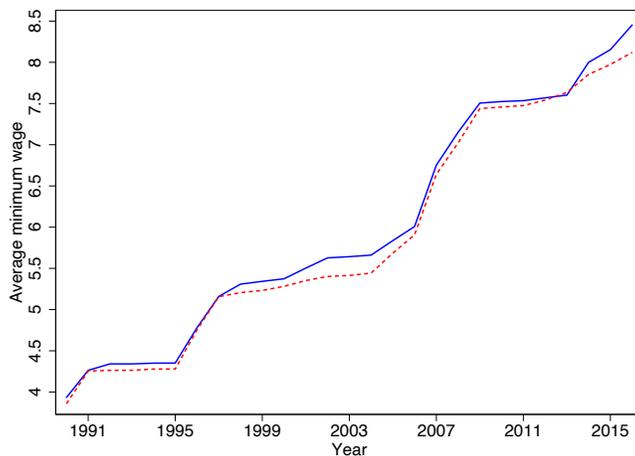
(b) Total employment



(c) Employment-to-population ratio



(d) Earnings per worker (U.S. dollars)



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

Figure 2: Comparison between commuting-zone groups: 137 multi-state commuting zones (solid blue) and rest of the country (dashed red)

that uses the full-country dataset and controls for some levels of time-varying spatial heterogeneity. We then move to our multi-state commuting zone sample and estimate a specification similar to DLR’s [specification \(6\)](#), which controls for time-varying local economic conditions.

Using the full-country dataset, we estimate equation

$$\ln e_{it} = \alpha + \beta \ln MW_{it} + \gamma \ln E_{it}^- + \delta \ln P_{it} + \eta_i + \tau_{ct} + \kappa_s \mathbb{1}_s \cdot T + \varepsilon_{it}, \quad (2)$$

where for commuting zone–state  $i$  in year  $t$ ,  $e_{it}$  is total employment in the restaurant industry,  $MW_{it}$  is the minimum wage,  $E_{it}^-$  is total employment in commuting zone–state  $i$  in all other industries,  $P_{it}$  is the working-age population,  $\eta_i$  is a commuting zone–state  $i$  fixed effect,  $\tau_{ct}$  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 \(2\)](#) 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 their novel cross-border county-pair approach in their [specification \(6\)](#). Along these lines, we exploit minimum wage variation within commuting zones by using the 137 multi-state commuting zone dataset and the econometric model

$$\ln e_{ipt} = \alpha + \beta \ln MW_{it} + \gamma \ln E_{it}^- + \delta \ln P_{it} + \eta_i + \tau_{pt} + \nu_{it}, \quad (3)$$

where subscript  $p$  identifies a pair for commuting zone-state  $i$ , and  $\tau_{pt}$  denotes pair–period fixed effects, which control for spatial heterogeneity at the local level. The only difference between [\(3\)](#) and [\(1\)](#) is that DLR use total private sector employment as a control (which we assume includes employment in restaurants), whereas we use instead total employment in the rest of the industries.<sup>15</sup>

Among our 137 multi-state commuting zones there are 151 complete pairs: one pair for each of the 129 two-state commuting zones, three pairs for each of the 7 three-state commuting zones, and

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<sup>15</sup>Using total employment as a control amounts to including our dependent variable on the right-hand side of equation [\(3\)](#), which would introduce a bias in the estimation of  $\beta$ . Recall that  $\beta$  is the effect of our variable of interest on the dependent variable holding the controls constant, and the latter condition cannot be satisfied if one of the controls includes the dependent variable. In particular, if the minimum wage reduces restaurant employment, and restaurant employment is in the total employment control (restaurant employment accounts for between 7.21% and 9.37% of total U.S. employment during the 1990–2016 period), there is attenuation of the minimum wage effect towards zero (because we effectively have a “bad control”, or we “overcontrol”). If we use instead total employment as the control in our commuting-zone specifications below, the estimated minimum wage elasticity of employment is between 75% and 94% the size of the estimated elasticity when using as control total employment in the rest of the industries. For example, it declines in size from  $-0.242$  to  $-0.188$  and from  $-0.255$  to  $-0.227$  in our results from [Table 3](#), and from  $-0.689$  to  $-0.58$  in the four-year elasticity (the  $p$ -value increases to 0.13 in the first case, but remains below 0.01 in the other two cases). In principle, one could construct a control that leaves out other low-wage industries as well, so that we are more likely picking up a cyclical control and not effects of minimum wages on employment in other industries—removing more of this overcontrol would only strengthen the minimum wage effect.

one more pair corresponding to Virginia and Maryland in the DC-VA-MD commuting zone (recall that DC is excluded from our data). Notice that the estimation of (3) requires complete pairs: if entities  $i$  and  $j$  belong to pair  $p$ , we control for  $\tau_{pt}$  by subtracting the equation for  $\ln e_{jpt}$  from (3) to obtain

$$\ln e_{ipt} - \ln e_{jpt} = \beta (\ln MW_{it} - \ln MW_{jt}) + \gamma (\ln E_{it}^- - \ln E_{jt}^-) + \delta (\ln P_{it} - \ln P_{jt}) + \eta_{ij} + \nu_{ijt}, \quad (4)$$

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 (4) as a panel with 151 pairs and 27 years, or we can directly estimate (3) by using the multi-way fixed effects estimator of [Correia \(2016\)](#) in our CBP-MCZP sample.<sup>16</sup> We follow the latter approach, as [Correia \(2016\)](#) provides a Stata package (`reghdfe`) that allows for multi-way clustering of standard errors as in [Cameron, Gelbach, and Miller \(2011\)](#).<sup>17</sup>

### 3.3 Main Results

Using the full sample, which is a panel of 866 commuting zone-state entities and 27 years, [Table 2](#) presents our results from the estimation of different versions of (2). [Panel A](#) presents the employment results, whereas [panel B](#) shows the estimation results of similar specifications for earnings per worker. All regressions include commuting zone–state effects and report standard errors that are clustered at the state level.

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<sup>16</sup>Our CBP-MCZP sample includes  $151 \times 2$  observations each year, for a maximum of 8,154 observations in the 1990-2016 period. Of these, our main regression below uses 8,134 observations, as we have 20 observations with missing data.

<sup>17</sup>Note that DLR estimate specifications to test for spillover effects within county pairs, across state borders. In the competitive model, border counties may be more likely to lose jobs because they can move across the border, compared to interior counties, leading to “amplification.” In a search model the opposite can happen (“attenuation”), because the higher minimum wage may induce more search in border counties from workers across the border, leading to more job openings. DLR’s spillover analysis rests on comparing estimates in border and interior counties. But DLR’s key contention is that there is spatial heterogeneity such that shocks at the border that might be correlated with minimum wage changes differ from shocks in the interior. If so, then their tests are invalid, since we cannot distinguish between different minimum wage effects in border and interior counties and different shocks correlated with minimum wages. If there is this kind of spatial heterogeneity, then an alternative approach to testing for spillovers is needed. If there is not—and the shocks are the same in border and interior areas—then state-level panel data estimates are valid and we do not need DLR’s border-county design. (Equivalently—as we find below using commuting zones—the estimates should be quite similar with or without the cross-border pair fixed effects.)

Spillovers would not invalidate the negative employment effect we find. If there is “amplification,” we may overstate the effect in the affected area alone, but it still must be negative (and imply that a higher minimum wage reduces employment in the affected area). And if there is “attenuation,” then our negative estimate is understated. That said, we think spillovers are unlikely to be quantitatively important in the restaurant sector, because restaurants serve very local customers ([Liu, Han, and Cohen, 2015](#)), in contrast, say to business-to-business services that could much more easily relocate across a border in the same commuting zone. Relatedly, there is evidence that restaurants are more likely to increase prices rather than relocate after a minimum wage hike (see [Romich, Allard, Obara, Althausser, and Buszkiewicz, 2020](#) and [Allegretto and Reich, 2018](#)), and whereas [Dharmasankar and Yoo \(2022\)](#) find evidence of spillovers after a minimum wage increase in Seattle, they find that the effects are concentrated in retail, not hospitality, consistent with view that people do not travel much to eat, but do to shop.

Table 2: Conventional estimation of minimum wage responses with CBP 1990-2016 data

	Full sample			Multi-state zones		
	(1)	(2)	(3)	(4)	(5)	(6)
<i><b>A. ln(employment)</b></i>						
ln(minimum wage)	-0.338*** (0.089)	-0.126** (0.051)	-0.049 (0.038)	-0.299*** (0.101)	-0.161* (0.091)	-0.128** (0.062)
ln(employment <sup>-</sup> )	0.000 (0.072)	0.115** (0.056)	0.122** (0.057)	-0.095 (0.070)	0.006 (0.066)	0.039 (0.080)
ln(population)	1.029*** (0.088)	0.866*** (0.071)	0.876*** (0.074)	1.155*** (0.125)	1.077*** (0.124)	1.025*** (0.116)
<i><b>B. ln(earnings)</b></i>						
ln(minimum wage)	0.215*** (0.037)	0.170*** (0.033)	0.168*** (0.035)	0.178*** (0.050)	0.169*** (0.037)	0.154*** (0.050)
ln(earnings <sup>-</sup> )	0.255*** (0.069)	0.221*** (0.067)	0.185*** (0.060)	0.287* (0.143)	0.248* (0.142)	0.203 (0.148)
ln(population)	0.091** (0.039)	0.111*** (0.036)	0.120** (0.045)	0.017 (0.079)	0.004 (0.076)	-0.014 (0.095)
Zone-state effects	Y	Y	Y	Y	Y	Y
Year effects	Y			Y		
Region-year effects		Y	Y		Y	Y
State trends			Y			Y
Observations	23,361	23,361	23,361	7,577	7,577	7,577

Notes: This table reports  $\hat{\beta}$ ,  $\hat{\gamma}$ , and  $\hat{\delta}$  from the estimation of equation (2) for the restaurant industry using yearly data from 1990 to 2016. In panel A, the dependent variable is log employment. Panel B uses instead log earnings per worker. Columns 1-3 use the full sample with 866 commuting zone-state entities, and columns 4-6 use the 281 multi-state commuting zone-state entities. Standard errors (in parentheses) are clustered at the state level. The coefficients are statistically significant at the \*10%, \*\*5%, or \*\*\*1% level.

Column 1 in Table 2 starts with a version of (2) that only includes commuting zone-state fixed effects and year fixed effects, so that it abstracts from controlling for time-varying spatial heterogeneity. The estimated minimum wage elasticity of employment is  $-0.338$ . Columns 2 and 3 show that as we add time-varying spatial heterogeneity controls, first with Census region-year effects in column 2 and then state trends in column 3, the minimum wage elasticity declines in magnitude, but keeps its negative sign. However, the estimated elasticity loses its statistical significance when including state trends. Although in that case the elasticity is small ( $-0.049$ ), [Meer and West \(2016\)](#) point out that the model with state trends can lead to attenuation, especially when the treatment effect grows over time. Indeed, our evidence below in section 4.1 points to minimum wage effects that increase over time.

In columns 4-6 we re-estimate the specifications from columns 1-3, but restricting the sample to

the 281 commuting zone-state entities of the 137 multi-state commuting zones. The story is similar as with the full sample, though the elasticity in the specification with state trends in column 6 ( $-0.128$ ) is more than twice the size of the elasticity in column 3 and is statistically significant.

Panel B in Table 2 shows estimates of the minimum wage elasticity of earnings per worker. Instead of employment, we estimate (2) using earnings per worker in the restaurant industry (and earnings per worker in the rest of the industries on the right-hand side of the specifications). All columns show a positive and highly significant elasticity that ranges between 0.154 and 0.215.

Using our multi-state zones dataset (the CBP-MCZP sample), we now turn to the estimation of equation (3), which controls for time-varying spatial heterogeneity at the local level, as DLR advocate (and which they claim eliminates evidence of adverse employment effects of the minimum wage). Table 3 presents the estimates of (3), with column 1 using the 151 pairs available in our CBP data, and column 2 restricting the CBP-MCZP sample to 71 of the 73 complete pairs available in the MCZP sample from DLR’s QCEW data. As in DLR, standard errors are two-way clustered at the state and border segment levels. The elasticity of employment is negative, large, and highly significant when estimating the local specification exploiting minimum wage differentials within commuting zones: columns 1 and 2 in panel A show elasticities of  $-0.242$  and  $-0.255$ . 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.

Having established that the pair-approach estimation with multi-state commuting zones yields negative and significant minimum wage elasticities of employment with both the CBP-MCZP sample (in columns 1-2 of Table 3) and the MCZP sample from DLR’s QCEW data (in Table 1), we now re-estimate equation (3) but using contiguous county-pairs that share a state border. Out of the 1,181 possible county pairs, our CBP data contains 1,165 for the restaurant industry. Columns 3 and 4 in Table 3 present the county-pair estimation results.

Column 3 estimates (3) using the 1,165 complete county pairs in the CBP data, and shows a non-significant elasticity of  $-0.081$ , which is about one third of the  $-0.242$  elasticity in column 1. The large reduction in the magnitude of the elasticity indicates, again, that the level of geographical aggregation used to construct local economic areas matters. For a comparison with DLR’s county-pair coverage, column 4 presents the estimation of (3) using 309 of the 316 complete county pairs of DLR, and reports a non-significant minimum wage elasticity of employment of  $-0.023$  (about one ninth of the elasticity in column 2). Hence, columns 3 and 4 show that for the within-county-pair approach using the CBP data, the sample that uses 309 of the 316 DLR’s complete pairs (26.16

Table 3: Pair-approach estimation of minimum wage responses with CBP 1990-2016 data

	<b>Multi-state zones</b>		<b>Contiguous counties</b>	
	(1)	(2)	(3)	(4)
<b><i>A. ln(employment)</i></b>				
ln(minimum wage)	-0.242** (0.120)	-0.255*** (0.082)	-0.081 (0.063)	-0.023 (0.056)
ln(employment <sup>-</sup> )	0.159 (0.098)	0.073 (0.089)	0.193*** (0.053)	0.154*** (0.051)
ln(population)	0.934*** (0.179)	1.116*** (0.184)	0.979*** (0.100)	1.000*** (0.081)
<b><i>B. ln(earnings)</i></b>				
ln(minimum wage)	0.163*** (0.055)	0.198*** (0.044)	0.156*** (0.044)	0.211*** (0.029)
ln(earnings <sup>-</sup> )	0.113 (0.138)	-0.047 (0.034)	0.044 (0.056)	0.017 (0.022)
ln(population)	0.085 (0.084)	0.068 (0.049)	0.027 (0.042)	0.040 (0.031)
Zone-state effects	Y	Y		
County effects			Y	Y
Pair-period effects	Y	Y	Y	Y
DLR data pairs		Y		Y
Number of pairs	151	71	1,165	309
Observations	8,134	3,830	62,228	16,670

Notes: This table reports  $\hat{\beta}$ ,  $\hat{\gamma}$ , and  $\hat{\delta}$  from the estimation of equation (3) for the restaurant industry from CBP 1990-2016 yearly data. Columns 1-2 use pairs within multi-state commuting zones, and columns 3-4 use contiguous county pairs. The dependent variable in panel A is log employment, whereas in panel B it is log earnings per worker. Column 1 uses the 151 multi-state commuting zone pairs, and column 2 uses 71 of the 73 DLR complete pairs used in columns 3-4 of panel B in Table 1 (we lose the two DC pairs). Column 3 uses all county pairs available in our dataset, and column 4 uses the complete pairs of DLR that are available in the CBP data (309 out of 316—we lose four pairs involving DC and three pairs involving the County of San Francisco). Panel B presents similar estimation results for the average wage. 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.

percent of all possible pairs) underestimates the minimum wage elasticity of employment when compared to a sample that uses 98.65 percent of contiguous county pairs that share a state border, although, as we have seen, the much more substantive impact comes from the definition of local economic areas.

Finally, all columns in panel B of Table 3 show statistically significant point estimates for the minimum wage elasticity of earnings in the restaurant industry, with values ranging between 0.156

and 0.211. Taking our preferred specifications in columns 1 and 2, a 10 percent increase in the minimum wage is associated with an average earnings increase in restaurants between 1.63 and 1.98 percent. The fact that the elasticities for earnings in the restaurant industry are not much affected when using the county-pair approach simply reflects that no matter the location, there is always a sizable fraction of restaurant employees paid the minimum wage. Hence, an increase in the minimum wage in one county will be met by an increase in the average earnings gap between this county and the control counties (irrespective of whether a county pair has any joint commuting or economic activity).

### 3.4 Evolution of the Minimum Wage Elasticity of Employment

Our analysis of the CBP data includes ten more years of available data than DLR’s QCEW data analysis. In this section we study the evolution of the minimum wage elasticity of employment between 2006 (the last year in DLR) and 2016 (the last year of our CBP data). We do this by estimating with our CBP data the four specifications of equation (3) in Table 3 for different periods, starting with DLR’s 1990–2006 period and ending with 1990–2016. Table 4 shows the estimated elasticities.

Our preferred specification in column 1, which uses all available pairs in our multi-state commuting zone data, consistently shows a large and significant elasticity. (It follows a bit of a U pattern, decreasing from  $-0.291$  in 1990–2006 to  $-0.341$  in 1990–2012, and then rising to our last value of  $-0.242$ .) We observe similar evidence in column 2 when we restrict the sample to the DLR pairs, with the elasticity ranging between  $-0.214$  and  $-0.285$ . Importantly, the 1990–2006 value of  $-0.214$  is not statistically different from the  $-0.141$  value obtained in Table 1 with DLR’s QCEW data.<sup>18</sup>

Columns 3 and 4, which use contiguous-county pairs, show three important results. First, the estimated non-significant elasticity of  $-0.059$  from using DLR pairs with CBP data for the period 1990–2006 is not statistically different from DLR’s estimated elasticity of 0.016. Second, the DLR-pairs specification in column 4 consistently underestimates the size of the elasticity when compared to the sample using all available pairs in column 3, being on average less than half the size.<sup>19</sup> And third, column 3 shows that when using all available county pairs, the minimum wage

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<sup>18</sup>As a potential explanation of the observed U pattern, [Clemens and Wither \(2019\)](#) document particularly strong minimum wage effects on employment as a consequence of the large hike in the federal minimum wage (from \$5.15 to \$7.25) between July 2007 and July 2009, whereas the effects are negative but weaker for minimum wage changes implemented in post-Great Recession years ([Clemens and Strain, 2018, 2021](#)).

<sup>19</sup>On the other hand, when using multi-state commuting zones, the coefficients in column 2 are on average about 87 percent of the size of the column 1 coefficients, and are larger in the last three periods.

Table 4: Evolution of the minimum wage elasticity of employment with pair approach and CBP data

Period	Multi-state zones		Contiguous counties	
	(1)	(2)	(3)	(4)
1990–2006	-0.291*** (0.095)	-0.214** (0.088)	-0.166** (0.065)	-0.059 (0.058)
1990–2007	-0.301*** (0.092)	-0.239*** (0.084)	-0.169*** (0.061)	-0.076 (0.055)
1990–2008	-0.327*** (0.093)	-0.262*** (0.084)	-0.175*** (0.061)	-0.086 (0.055)
1990–2009	-0.337*** (0.092)	-0.270*** (0.084)	-0.170*** (0.058)	-0.086 (0.055)
1990–2010	-0.338*** (0.094)	-0.275*** (0.086)	-0.159*** (0.055)	-0.083 (0.054)
1990–2011	-0.333*** (0.094)	-0.275*** (0.087)	-0.141** (0.054)	-0.075 (0.054)
1990–2012	-0.341*** (0.095)	-0.284*** (0.087)	-0.127** (0.053)	-0.071 (0.054)
1990–2013	-0.323*** (0.099)	-0.285*** (0.087)	-0.106* (0.053)	-0.064 (0.056)
1990–2014	-0.275** (0.112)	-0.278*** (0.089)	-0.081 (0.056)	-0.048 (0.059)
1990–2015	-0.258** (0.117)	-0.260*** (0.086)	-0.084 (0.057)	-0.033 (0.059)
1990–2016	-0.242** (0.120)	-0.255*** (0.082)	-0.081 (0.063)	-0.023 (0.056)
Zone-state effects	Y	Y		
County effects			Y	Y
Pair-period effects	Y	Y	Y	Y
DLR data pairs		Y		Y
Number of pairs	151	71	1,165	309

Notes: This table reports  $\hat{\beta}$  from the estimation of equation (3) for the restaurant industry using CBP yearly data for different periods. Columns 1-2 use pairs within multi-state commuting zones, and columns 3-4 use contiguous county pairs. The dependent variable is log employment, the main regression is the log minimum wages, and the controls (not reported) are employment in the rest of the industries and working age population. Column 1 uses the 151 multi-state commuting zone pairs, and column 2 uses 71 of the 73 DLR complete pairs used in columns 3-4 of panel B in Table 1. Column 3 uses all county pairs available in our dataset, and column 4 uses the complete pairs of DLR that are available in the CBP data (309 out of 316). 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.

elasticity of employment is statistically significant in all periods between 1990–2006 and 1990–2013, reaching values up to  $-0.175$ . Therefore, not only does the county-pair approach underestimate

the employment elasticity when compared to using pairs within multi-state commuting zones, but the last two findings indicate that DLR’s original county-pair results depend on both the selection of counties and the sample period.

## 4 Time Paths, Pre-Trends, and Common Shocks

DLR explore both the long-term effects of minimum wages and the possibility of pre-existing trends that may affect their specifications. For their preferred [specification \(6\)](#) for contiguous county pairs sharing a state border, DLR find *(i)* a minimum wage elasticity of employment that is stable around zero with no delayed effects after four years of the minimum wage change, and *(ii)* no evidence of existing pre-trends. This section revisits these findings using instead pairs from multi-state commuting zones. Moreover, we also look at differences between the two types of geographical aggregation in capturing common shocks.

### 4.1 Long-Term Employment Effects of Minimum Wages

The core motivation for DLR’s border-county design is to control for spatial shocks that might be correlated with minimum wage variation. They argue that the more conventional panel data estimator using state variation (with fixed state and year effects) is biased from such shocks. Although there is good reason to believe that commuting zones capture local economic shocks, and do this better than cross-border counties, it is at least in principle possible that commuting zones fail to do this. We thus replicate the analysis DLR present to try to argue that their cross-border county design controls for spatial heterogeneity. We find that, on the criterion that DLR use, commuting zones capture local economic shocks, and appear to do so better than do cross-border county pairs.

DLR estimate a distributed-lag model with eight quarters of leads and sixteen quarters of lags. In particular, the distributed-lag version of DLR’s preferred [specification \(6\)](#) is

$$\ln e_{ipt} = \alpha + \sum_{k=-4}^7 \beta_{-2k} \Delta_2 \ln MW_{i,t-2k} + \beta_{-16} \ln MW_{i,t-16} + \rho Z_{it} + \eta_i + \tau_{pt} + \nu_{it}, \quad (5)$$

where  $\Delta_2$  is a two-quarter difference operator, and the rest of the variables are defined as in equation (1). The model estimates thirteen  $\beta$  parameters that indicate the cumulative effect of minimum wages, with  $\beta_8$  denoting the lead effect eight quarters before the minimum wage change, and up to  $\beta_{-16}$ , which denotes the cumulative effect sixteen quarters after the minimum wage change.

Using DLR’s QCEW data and programs, [Figure 3](#) shows the estimates of the  $\beta$  parameters from the estimation of (5)—along with 90 percent confidence intervals—using either contiguous county

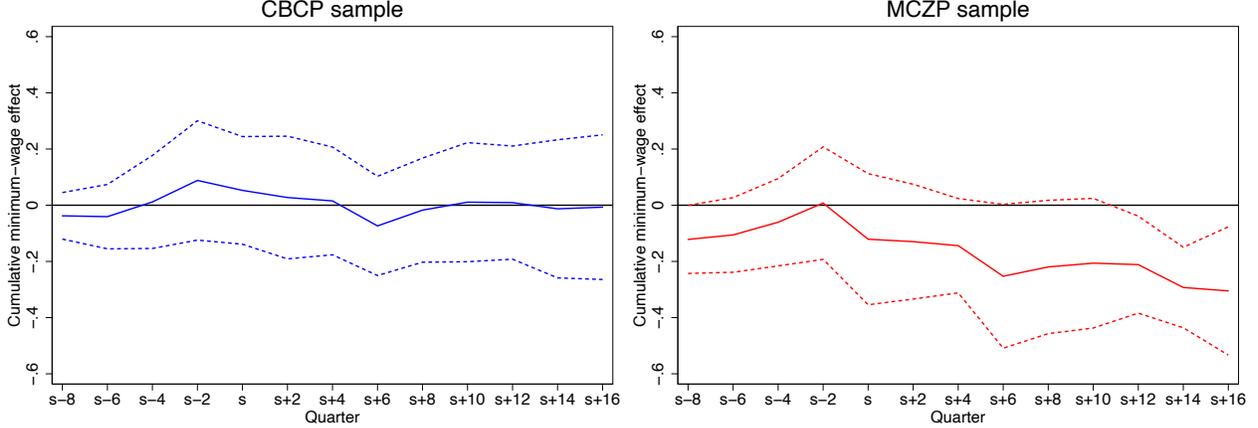


Figure 3: Time paths of minimum wage effects on employment with 90% confidence intervals using DLR’s QCEW data: Contiguous county pairs (left–blue) and pairs from multi-state commuting zones (right–red)

pairs (CBCP sample) or pairs from multi-state commuting zones (MCZP sample). Table A-2 in the Appendix shows the estimation results for both cases. We assume that the minimum wage change occurs at time  $s$ , and hence, each plot in Figure 3 starts at  $s - 8$  (with  $\hat{\beta}_8$ ) and ends at  $s + 16$  (with  $\hat{\beta}_{-16}$ ). The plot that uses the CBCP sample is reported by DLR, with the only difference being that the 90 percent interval was calculated using corrected standard errors, calculated using only the complete-pair observations used in the estimation.

Notice that although the cumulative effect of a minimum wage change on restaurant employment is stable around zero when using county pairs, the story is very different when using pairs from multi-state commuting zones. In the latter case, the cumulative effect is near zero two quarters before the minimum wage change, but then it gets into a solid negative trend, reaching a significant elasticity of  $-0.305$  after four years—this cumulative effect is more than twice as large as the  $-0.141$  contemporaneous elasticity reported in Table 1. Therefore, a simple switch of the definition of local economic area—from pairs of contiguous counties sharing a state border, to multi-state commuting zones that are actually defined as local economic areas—dramatically overturns DLR’s finding of a near zero long-term effect of minimum wages on employment in the restaurant industry.

To assess the robustness of this result, we also look at time paths for the estimated minimum wage elasticity of employment using the more comprehensive CBP yearly data. The distributed-lags version of specification (3) is

$$\ln e_{ipt} = \alpha + \sum_{k=-2}^3 \beta_{-k} \Delta \ln MW_{i,t-k} + \beta_{-4} \ln MW_{i,t-4} + \gamma \ln E_{it}^- + \delta \ln P_{it} + \eta_i + \tau_{pt} + \nu_{it}, \quad (6)$$

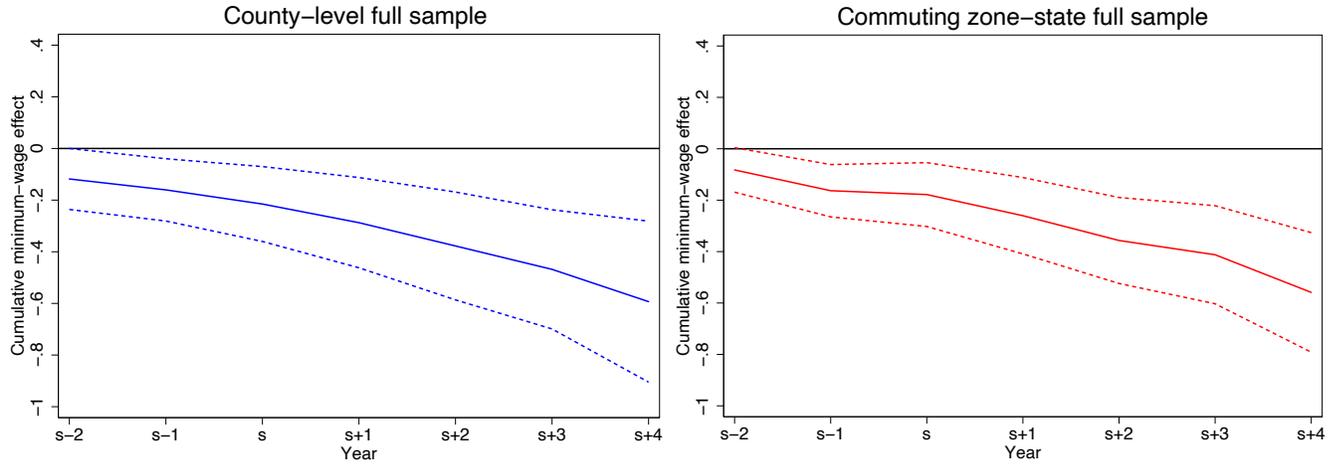
where  $\Delta$  is a one-year difference operator, and the other variables are as described in section 3.2.

Thus, similar to the specification in (5), the model in (6) includes two years of leads and four years of lags, with the difference that we use one-year rather than two-quarter changes. In total, from (6) we estimate seven  $\beta$  parameters, starting with  $\beta_2$  for the lead effect two years before the minimum wage change, and up to  $\beta_{-4}$ , which captures the cumulative effect up to four years after the minimum wage change.

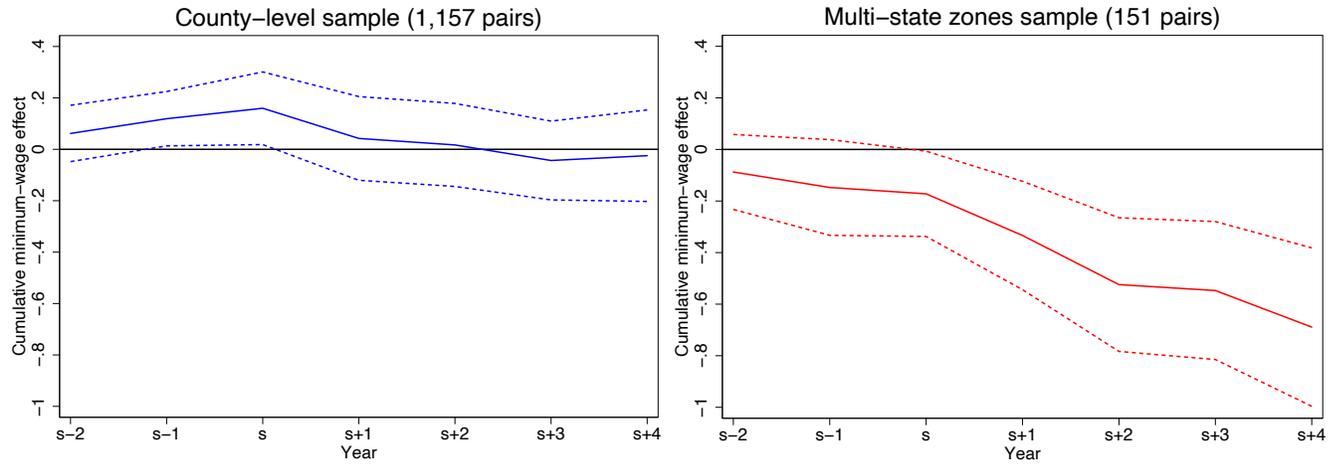
Using both contiguous county pairs and pairs from multi-state commuting zones, we start with a leads-lags version of the conventional two-way fixed effects estimation using the full-country dataset (as in column 1 of Table 2), and then we estimate equation (6) using either all complete pairs or only DLR’s complete pairs. For all these cases, Figure 4 shows plots of the estimated  $\beta$  coefficients along with 90 percent confidence intervals. (Table A-3 in the Appendix shows the point estimates and their standard errors.) Figure 4a shows the cumulative effects for the two-way fixed effects estimation, Figure 4b presents the cumulative effects from the estimation of (6) when using all available complete pairs, and Figure 4c shows the effects when we restrict the CBP sample to DLR’s complete pairs.

Whether we use the county-level sample or the commuting zone-state level sample, Figure 4a shows that the conventional two-way fixed effects estimation yields similar cumulative minimum wage responses. The cumulative minimum wage elasticity after four years is  $-0.593$  when using the county-level sample, and is  $-0.559$  when using the commuting zone-state sample. Even though there are negative leading effects, implying that minimum wages were rising more where employment had declined, there is still a noticeable change in the slope of the cumulative effect after the treatment; this is very clear in the right-hand panel of Figure 4a, but even in the left-hand panel, the relationship is not linear. Hence, even if there are some leading correlations in the two-way fixed effects estimation that, unaccounted for, would lead to negative bias, the plots in Figure 4a show that there is an employment decline attributable to the minimum wage increase.

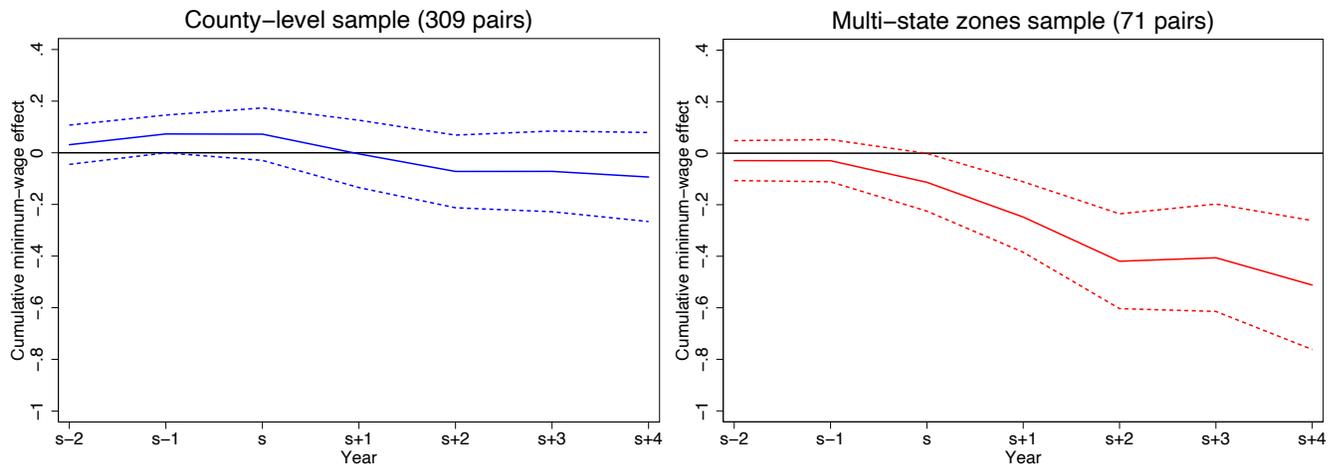
From the estimation of (6), Figures 4b and 4c reinforce our findings from Figure 3. Whereas the county-level plots show a non-significant long-term effect of minimum wages, the plots using pairs from multi-state commuting zones show large and significant negative effects of minimum wages in the medium and long terms, with stark changes in the slope of the cumulative effect occurring around time  $s$ . The cumulative minimum wage elasticity after four years is  $-0.689$  in the estimation that uses the 151 complete pairs, and  $-0.512$  in the estimation that uses the 71 complete pairs in DLR’s QCEW data. In the first case, the cumulative four-year elasticity is 2.8 times larger than the contemporaneous elasticity in column 1 of Table 3 ( $-0.242$ ), and in the second case, the four-year elasticity is two times larger than the elasticity in column 2 ( $-0.255$ ).



(a) Conventional two-way fixed effects leads-lags estimation using full sample



(b) Estimation of equation (6) using all complete pairs



(c) Estimation of equation (6) using DLR's complete pairs

Figure 4: Time paths of minimum wage effects on employment with 90% confidence intervals using CBP data: County-level samples (left-blue) and commuting zone-state samples (right-red)

Whereas when using commuting-zone pairs we find that the better-identified local control approach yields a larger estimated cumulative effect than the two-way fixed effects approach, the opposite happens when using the county-level pairs. These results point toward an upward bias in the estimation of the minimum wage elasticity of employment when using county pairs as local economic areas.

## 4.2 Pre-existing Trends

To formally test for pre-existing trends, DLR estimate a leads-only version of their [specification \(6\)](#) given by

$$\ln e_{ipt} = \alpha + \beta_{12}(\ln MW_{i,t+12} - \ln MW_{i,t+4}) + \beta_4(\ln MW_{i,t+4} - \ln MW_{it}) + \beta_0 \ln MW_{it} + \delta P_{it} + \eta_i + \tau_{pt} + \nu_{it}, \quad (7)$$

where  $\beta_{12}$  is the lead effect twelve quarters before the minimum wage change and  $\beta_4$  is the cumulative effect four quarters before the change, so that if the minimum wage change occurs in quarter  $s$ ,  $\beta_4 - \beta_{12}$  captures the pre-existing trend between  $s - 12$  and  $s - 4$ . Equation (7) includes the population control, but excludes total private sector employment. Instead, DLR estimate a version of (7) where the latter also serves as dependent variable.

Panel A in Table 5 shows  $\hat{\beta}_{12}$ ,  $\hat{\beta}_4$ ,  $\hat{\beta}_0$  and  $\hat{\beta}_4 - \hat{\beta}_{12}$  from the estimation of (7) when using complete pairs from the CBCP and MCZP samples. The county-pair results are the same as those reported in DLR's Table 3, but we fixed the standard errors so that only complete-pair observations are used in their calculation. The results show no evidence of pre-trends for restaurant employment or total private sector employment for both the contiguous-county and multi-state commuting zones samples.

To test for pre-trends in the CBP data, we estimate a yearly version of equation (7) with  $\beta_3$  denoting the lead effect three years before the change in the minimum wage,  $\beta_1$  denoting the cumulative effect up to one year before the change, and  $\beta_1 - \beta_3$  capturing the trend between year  $s - 3$  and year  $s - 1$ . Besides restaurant employment, we also test for pre-trends in total employment in the rest of the industries. Panel B in Table 5 shows  $\hat{\beta}_3$ ,  $\hat{\beta}_1$ ,  $\hat{\beta}_0$  and  $\hat{\beta}_1 - \hat{\beta}_3$  when using all complete pairs in the county-level and multi-state zones samples. Whereas the county-level sample yields a significant positive trend for restaurant employment, the multi-state zones sample shows a negative and significant trend driven by  $\hat{\beta}_1 = -0.177$ . The large negative  $\hat{\beta}_1$  in the multi-state zones sample, however, is not surprising, as minimum wage changes are typically announced months before implementation, which allows firms to prepare and adjust employment

Table 5: Testing for pre-trends in the pair-approach estimation

	County-level sample		Multi-state zones sample	
	Restaurant Employment	Aggregate Employment	Restaurant Employment	Aggregate Employment
<i>A. DLR's Quarterly Data (1990-2006)</i>				
$\hat{\beta}_{12}$	0.009 (0.052)	0.025 (0.053)	-0.090 (0.062)	0.025 (0.054)
$\hat{\beta}_4$	0.050 (0.133)	0.084 (0.112)	-0.054 (0.117)	0.056 (0.100)
$\hat{\beta}_0$	0.065 (0.123)	0.172 (0.174)	-0.173 (0.120)	0.101 (0.121)
Trend ( $\hat{\beta}_4 - \hat{\beta}_{12}$ )	0.041 (0.103)	0.058 (0.073)	0.035 (0.080)	0.031 (0.063)
Number of pairs	316	316	73	73
Observations	37,920	37,896	8,760	8,758
<i>B. CBP Yearly Data (1990-2016) — All Complete Pairs</i>				
$\hat{\beta}_3$	0.017 (0.064)	-0.034 (0.037)	-0.084 (0.070)	-0.065 (0.058)
$\hat{\beta}_1$	0.100 (0.085)	-0.019 (0.059)	-0.177** (0.076)	-0.075 (0.086)
$\hat{\beta}_0$	-0.058 (0.100)	-0.082 (0.089)	-0.451*** (0.126)	-0.215 (0.146)
Trend ( $\hat{\beta}_1 - \hat{\beta}_3$ )	0.083* (0.048)	0.015 (0.030)	-0.094* (0.048)	-0.009 (0.041)
Number of pairs	1,163	1,163	151	151
Observations	54,766	54,766	7,204	7,204
<i>C. CBP Yearly Data (1990-2016) — DLR's Complete Pairs</i>				
$\hat{\beta}_3$	0.038 (0.052)	-0.018 (0.040)	-0.057 (0.052)	-0.065 (0.063)
$\hat{\beta}_1$	0.070 (0.066)	0.014 (0.057)	-0.110** (0.053)	-0.076 (0.090)
$\hat{\beta}_0$	-0.028 (0.089)	-0.006 (0.079)	-0.365*** (0.108)	-0.117 (0.140)
Trend ( $\hat{\beta}_1 - \hat{\beta}_3$ )	0.032 (0.030)	0.033 (0.029)	-0.052 (0.032)	-0.011 (0.036)
Number of pairs	309	309	71	71
Observations	14,784	14,784	3,396	3,396

Notes: Panel A in this table reports  $\hat{\beta}_{12}$ ,  $\hat{\beta}_4$ ,  $\hat{\beta}_0$ , and  $\hat{\beta}_4 - \hat{\beta}_{12}$  from the estimation of (7) using DLR's QCEW data for contiguous county pairs and pairs from multi-state commuting zones. Panels B and C report similar coefficients but using yearly CBP data, with panel C restricting the sample to DLR's complete pairs. Aggregate employment is total private sector employment in panel A, and employment in all other industries in panels B and C. 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.

before the change goes into effect.<sup>20</sup> Moreover, the size of the trend ( $-0.094$ ) is about one third of

<sup>20</sup>The positive pre-trend in the county-level sample is harder to explain as an anticipatory effect of minimum

the change that occurs between  $s - 1$  and  $s$  ( $-0.274$ ), with the cumulative effect increasing in size from  $-0.177$  to  $-0.451$ . Thus, there is a clear change in the relationship between employment and minimum wages around time  $s$ . For employment in the rest of the industries, Panel **B** shows no significant evidence of pre-trends.

Panel **C** in Table 5 restricts the CBP sample to complete pairs used in DLR’s estimation, showing no evidence of pre-trends for both employment variables in both samples. Similar to panel **B**,  $\hat{\beta}_1$  is statistically significant in the multi-state commuting zone sample, indicating that restaurants may start adjusting employment before the minimum wage change goes into effect. However, all of the evidence indicates that using commuting zones as local economic areas points to large negative effects of minimum wages on restaurant employment.

### 4.3 Common Shocks in Local Economic Areas

DLR’s main argument is that their [specification \(6\)](#), with county-border pairs, controls for spatial heterogeneity by accounting for local economic shocks. This paper shows that the definition of local economic area as two counties sharing a state border is crucial for DLR’s result of a near-zero effect of minimum wages on employment in the U.S. restaurant industry. By instead using multi-state commuting zones to define local economic areas—consistent with the actual definition of commuting zones—we find statistically significant, negative, and persistent effects of minimum wages on restaurant employment. Hence, to help interpret the evidence, it is important to look at how each definition of local economic area captures common shocks.

Using complete pairs from both DLR’s QCEW data and the CBP data, Table 6 presents within-pair correlations for one-, two-, three-, and four-year log changes in the controls used by DLR (total private sector employment and population) and in the controls used in our CBP-data estimations (total employment in the rest of the industries and population), as well as for Bartik shocks that capture predicted changes in labor demand in each commuting zone-state—as a consequence of national industry-level employment changes—while accounting for local specialization patterns.<sup>21</sup> In Table 6, we use  $\Delta_\lambda$  to denote log changes of  $\lambda$  years. Panel **A** shows correlations using DLR’s QCEW complete pairs, panel **B** uses all complete pairs in the CBP data, and panel **C** restricts the

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wages, unless one believes that higher minimum wages increase employment, and may instead point to a problematic positive pre-trend for the counties in cross-border pairs where the minimum wage increased.

<sup>21</sup>We define Bartik shocks at the commuting zone-state level using AADHP’s 479 industries. Similar to [Autor, Dorn, Hanson, and Majlesi \(2016\)](#), the Bartik shock for commuting zone-state  $i$  from  $t_{\text{start}}$  to  $t_{\text{end}}$  is given by  $\sum_j \left( \frac{e_{ijt_{\text{start}}}}{e_{it_{\text{start}}}} \right) \left[ \ln e_{jt_{\text{end}}}^{-i} - \ln e_{jt_{\text{start}}}^{-i} \right]$ , where  $e_{ijs}$  is the commuting zone-state  $i$ ’s employment in industry  $j$  at time  $s$ ,  $e_{it_{\text{start}}}$  is total employment in commuting zone-state  $i$ ’s at time  $s$ , and  $e_{js}^{-i}$  is industry  $j$ ’s employment across all U.S. commuting zone-state entities with the exception of commuting zone-state  $i$  at time  $s$ .

Table 6: Within-pair correlations of controls and Bartik shocks

	Contiguous county pairs				Pairs within multi-state zones			
	$\Delta_1$	$\Delta_2$	$\Delta_3$	$\Delta_4$	$\Delta_1$	$\Delta_2$	$\Delta_3$	$\Delta_4$
<i>A. DLR's Quarterly Data (1990-2006)</i>								
Priv. employment	0.26	0.34	0.35	0.31	0.28	0.43	0.45	0.44
Population	0.38	0.39	0.39	0.40	0.50	0.52	0.54	0.54
<i>B. CBP Yearly Data (1990-2016) — All Complete Pairs</i>								
Employment <sup>-</sup>	0.10	0.15	0.17	0.21	0.19	0.30	0.31	0.34
Population	0.35	0.42	0.45	0.45	0.52	0.58	0.61	0.63
Bartik shock	0.72	0.76	0.72	0.74	0.83	0.84	0.82	0.83
<i>C. CBP Yearly Data (1990-2016) — DLR's Complete Pairs</i>								
Employment <sup>-</sup>	0.21	0.33	0.35	0.41	0.35	0.48	0.53	0.58
Population	0.46	0.49	0.48	0.50	0.62	0.66	0.66	0.67
Bartik shock	0.87	0.88	0.87	0.88	0.91	0.92	0.90	0.91

CBP sample to DLR's complete pairs. For all samples and all types of shocks, Table 6 shows that within-pair correlations are always larger for pairs within multi-state commuting zones than for contiguous county pairs, by an average (across all 32 comparisons) of 36%.

The previous correlations are about the pattern of aggregate shocks. They are not about shocks to restaurant employment, nor about the possibility of contemporaneous correlations between shocks and minimum wage changes, which our identification strategy rules out by assumption. There is still an open question of what causes the apparent positive bias in the cross-border county pair estimates. A possible explanation is that, because shocks are not common (at least compared to commuting zones), the minimum wage tends to be increased when there is a positive unobserved shock to restaurant/low-skilled employment (Neumark, 2019, pp. 308-309).

## 5 Conclusion

This paper shows that a simple change in the local level of aggregation—from pairs of contiguous border counties to pairs within multi-state commuting zones—overturns DLR's finding of no relationship between minimum wages and employment in the U.S. restaurant industry.

Given that multi-state commuting zones are defined to capture common economic influences, our evidence implies that accounting for spatial heterogeneity in estimating the effects of state minimum wage variation does not eliminate—nor does it even weaken—the evidence that minimum

wage increases reduce restaurant employment. Rather, the potentially more-rigorous approach of isolating state minimum wage variation from correlated economic shocks by looking within local economic areas generates evidence that reinforces the conclusion that a higher minimum wage reduces restaurant employment.

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# Appendix

## A Supporting Figures and Tables

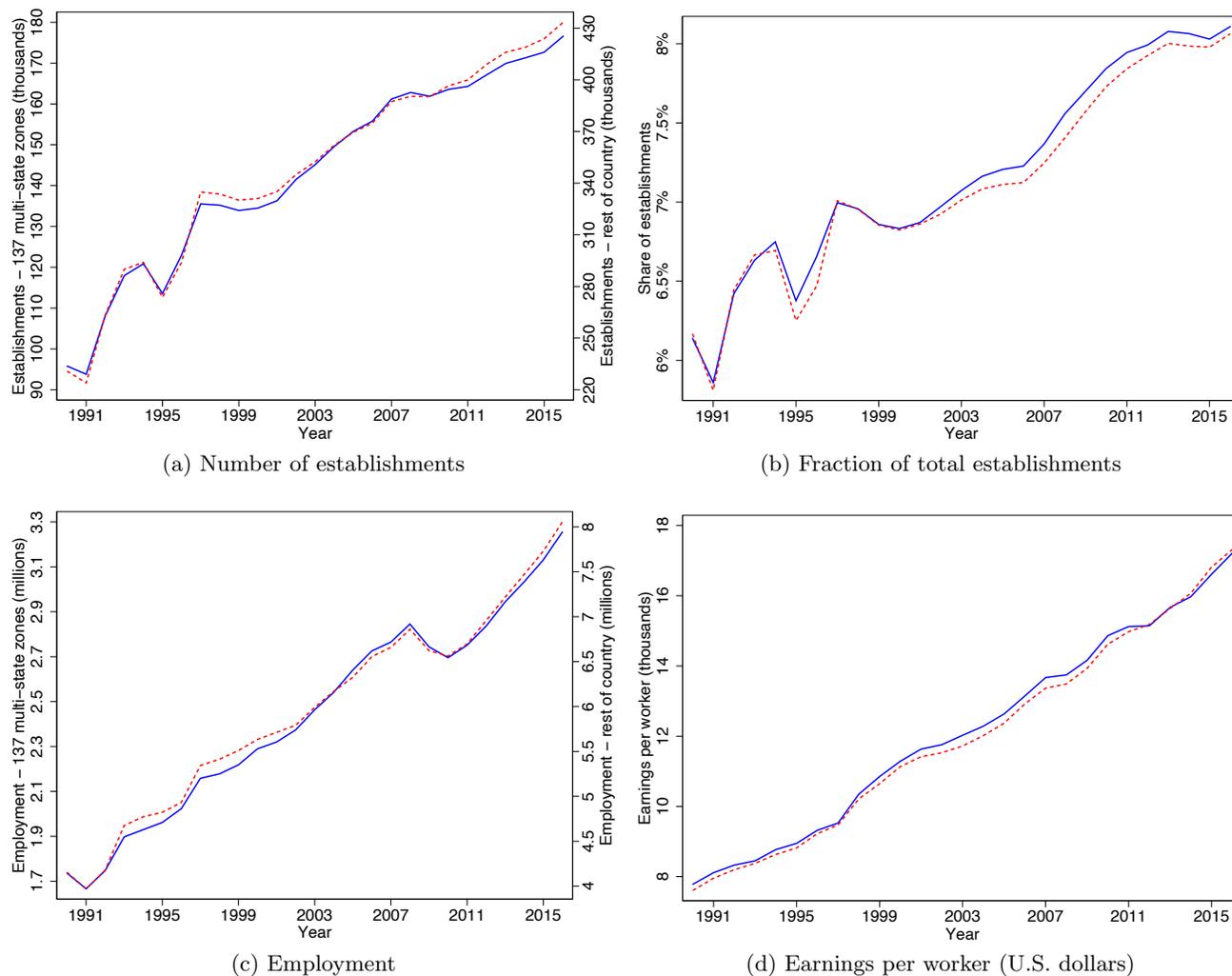


Figure A-1: Comparison between commuting-zone groups for restaurant industry: 137 multi-state commuting zones (solid blue) and rest of the country (dashed red)

Notes: This figure shows that in both groups, the restaurant industry accounted for 6.1 percent of all establishments in 1990 and that this share increased to about 8.1 percent by 2016. From Figure A-1d, note that nominal earnings per worker increased from \$7.6-\$7.8 thousand in 1990 to \$17.2-\$17.4 thousand in 2016. The annual payroll variable (from the CBP database) that we use to calculate earnings per worker includes reported tips.

Table A-1: Employment shares and earnings ranking of industries, 1990 and 2016

<b>Industry</b>	<b>1990</b>			<b>2016</b>		
	<i>Employment share</i>	<i>Worker earnings (thousands US\$)</i>	<i>Earnings ranking</i>	<i>Employment share</i>	<i>Worker earnings (thousands US\$)</i>	<i>Earnings ranking</i>
Eating and drinking places	7.21%	7.68	1	9.37%	17.36	1
Retail trade	13.94%	13.47	2	12.81%	27.08	2
Services	16.40%	14.06	3	22.92%	33.02	3
Textiles/apparel	1.99%	15.92	4	0.31%	36.20	4
Wood/furniture	1.36%	19.73	5	0.68%	41.76	5
Other manufacturing	0.44%	21.36	6	0.19%	47.40	7
Food/tobacco	1.62%	23.98	7	1.31%	45.28	6
Health services	9.83%	24.27	8	12.78%	53.87	10
Plastics, clay, stone	1.57%	24.72	9	0.91%	49.80	8
Agriculture, forestry, fishing, and mining	1.29%	25.10	10	1.52%	52.35	9
Construction	5.94%	25.22	11	5.13%	58.69	12
Paper/printing	2.44%	26.94	12	1.04%	63.76	14
Wholesale trade	6.69%	27.97	13	5.50%	66.68	16
Finance, insurance, and real estate	7.60%	28.00	14	7.05%	86.09	19
Metals	2.46%	28.31	15	1.36%	54.19	11
Transp., comm., elec., gas, and sanitary	5.88%	28.98	16	5.50%	60.16	13
Equipment	4.94%	30.33	17	2.21%	66.18	15
Legal, consulting, and computing services	5.27%	34.43	18	7.67%	91.09	20
Transportation manufacturing	2.03%	35.02	19	1.04%	69.44	17
Chemicals/petroleum	1.11%	35.94	20	0.70%	84.35	18

Table A-2: Long-term minimum wage responses with DLR's QCEW 1990-2006 data

	CBCP sample	MCZP sample
$\hat{\beta}_8$	-0.038 (0.050)	-0.122* (0.073)
$\hat{\beta}_6$	-0.041 (0.069)	-0.106 (0.080)
$\hat{\beta}_4$	0.012 (0.100)	-0.061 (0.093)
$\hat{\beta}_2$	0.088 (0.128)	0.008 (0.121)
$\hat{\beta}_0$	0.053 (0.116)	-0.121 (0.140)
$\hat{\beta}_{-2}$	0.027 (0.132)	-0.130 (0.123)
$\hat{\beta}_{-4}$	0.015 (0.116)	-0.144 (0.101)
$\hat{\beta}_{-6}$	-0.074 (0.107)	-0.253 (0.154)
$\hat{\beta}_{-8}$	-0.017 (0.112)	-0.220 (0.143)
$\hat{\beta}_{-10}$	0.011 (0.128)	-0.206 (0.139)
$\hat{\beta}_{-12}$	0.009 (0.122)	-0.211** (0.104)
$\hat{\beta}_{-14}$	-0.013 (0.149)	-0.293*** (0.086)
$\hat{\beta}_{-16}$	-0.007 (0.156)	-0.305** (0.137)
ln(priv. employment)	0.384*** (0.091)	0.354*** (0.125)
ln(population)	0.727*** (0.199)	0.717 (0.478)
Pair-period effects	Y	Y
Total private sector	Y	Y
Number of pairs	316	73
Observations	40,416	9,342

Notes: This table reports  $\hat{\beta}_k$ , for  $k \in \{8, 6, 4, 2, 0, -2, -4, -6, -8, -12, -16\}$ , from the estimation of equation (5) using either the CBCP sample or the MCZP sample. 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 A-3: Long-term minimum wage responses with CBP 1990-2016 data

	County-level sample			Multi-state zones sample		
	(1)	(2)	(3)	(4)	(5)	(6)
$\hat{\beta}_2$	-0.118 (0.070)	0.062 (0.065)	0.031 (0.045)	-0.082 (0.052)	-0.087 (0.087)	-0.029 (0.046)
$\hat{\beta}_1$	-0.160** (0.072)	0.119* (0.063)	0.073 (0.044)	-0.163*** (0.061)	-0.148 (0.111)	-0.029 (0.049)
$\hat{\beta}_0$	-0.215** (0.086)	0.160* (0.084)	0.072 (0.061)	-0.178** (0.074)	-0.172* (0.099)	-0.113* (0.066)
$\hat{\beta}_{-1}$	-0.287*** (0.104)	0.042 (0.097)	-0.004 (0.078)	-0.260*** (0.089)	-0.334** (0.125)	-0.248*** (0.081)
$\hat{\beta}_{-2}$	-0.378*** (0.124)	0.017 (0.096)	-0.072 (0.084)	-0.357*** (0.100)	-0.524*** (0.154)	-0.419*** (0.109)
$\hat{\beta}_{-3}$	-0.468*** (0.138)	-0.044 (0.091)	-0.072 (0.093)	-0.412*** (0.114)	-0.547*** (0.159)	-0.406*** (0.124)
$\hat{\beta}_{-4}$	-0.593*** (0.186)	-0.025 (0.106)	-0.094 (0.103)	-0.559*** (0.139)	-0.689*** (0.183)	-0.512*** (0.149)
ln(employment <sup>-</sup> )	0.163*** (0.039)	0.204*** (0.055)	0.115** (0.044)	0.019 (0.073)	0.079 (0.090)	0.067 (0.088)
ln(population)	1.007*** (0.072)	0.929*** (0.121)	1.015*** (0.079)	1.065*** (0.091)	0.806*** (0.182)	1.128*** (0.212)
Zone-state effects				Y	Y	Y
County effects	Y	Y	Y			
Year effects	Y			Y		
Pair-period effects		Y	Y		Y	Y
DLR data pairs			Y			Y
Number of pairs	–	1,157	309	–	151	71
Observations	64,064	47,268	12,866	18,109	6,262	2,954

Notes: This table reports  $\hat{\beta}_k$ , for  $k \in \{2, 1, 0, -1, -2, -3, -4\}$ ,  $\hat{\gamma}$ , and  $\hat{\delta}$  from the estimation of equation (6) using either the CBP county-level sample or the CBP multi-state commuting zones sample. Columns 3 and 6 restrict the sample to complete pairs in DLR's data. Although column 1 in Table 3 uses 1,165 complete pairs, the leads and lags in (6) make us lose eight pairs. We do not lose any pairs in the estimation with multi-state commuting zones. For the sample that uses DLR's complete pairs, recall that we have 309 out of 316 of DLR's complete pairs, and 71 out of 73 for the multi-state zones estimation. Standard errors (in parentheses) are clustered at the state level in columns 1 and 4, and are two-way clustered at the state and border segment levels in columns 2-3 and 5-6. The coefficients are statistically significant at the \*10%, \*\*5%, or \*\*\*1% level.