Tipped Wage Effects on Earnings and Employment in Full-Service Restaurants*

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We exploit more than 20 years of changes in state-level tipped wage policy and estimate earnings and employment effects of the tipped wage using county-level panel data on full-service restaurants (FSR). We extend earlier work by Dube, Lester, and Reich (2010) and compare outcomes between contiguous counties that straddle a state border. We find a 10-percent increase in the tipped wage increases earnings in FSRs about 0.4 percent. Employment elasticities are sensitive to the inclusion of controls for unobserved spatial heterogeneity. In our preferred models, we find small, insignificant effects of the tipped wage on FSR employment.

Introduction

The minimum wage is one of the most researched areas in labor economics, with a vast body of literature that dates back nearly 70 years (Brown 1999). Over the last several decades, research on the minimum wage has further proliferated as economists have exploited the growing variation in state minimum wage policies. However, research, public debate, and policy have largely ignored the lesser known tipped wage received by tipped workers (sometimes referred to as the subminimum or cash wage), even as it too has ample state variation that facilitates empirical estimation. Indeed, the existence of two federal wage floors,[1] with the federal tipped wage at $2.13 since 1991, is relatively unknown.

The 1966 Fair Labor Standards Act (FLSA) amendments expanded wage protections to restaurant, hotel, and other service workers but also allowed for a "tip credit" whereby employers could use tips, provided by customers, as

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The authors are grateful for generous support from the Ford Foundation and the Russell Sage Foundation. They thank Megan Collins, Luke Reidenbach, and Rachel West for excellent research assistance and David Card, Arindrajit Dube, Bill Lester, Michael Reich, Jesse Rothstein, Ben Zipperer, and two anonymous reviewers for their valuable comments and suggestions.

[1] There is also a youth wage that allows employers to pay employees under 20 years of age a lower wage ($4.25) for a limited period (90 calendar days, not work days) after they are first employed.

Industrial Relations, Vol. 54, No. 4 (October 2015).
Published by Wiley Periodicals, Inc., 350 Main Street, Malden, MA 02148, USA, and 9600 Garsington Road, Oxford, OX4 2DQ, UK.

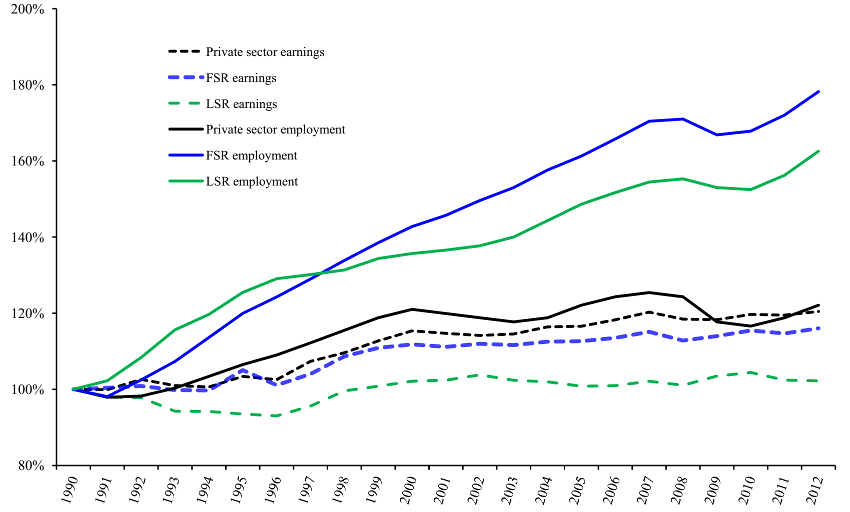
"credit" toward a worker's regular minimum wage. Today, at the federal level, the regular minimum wage is $7.25 and the allowable tipped wage is $2.13—the $5.12 difference is the maximum allowable tip credit. The $5.12 tip credit may be thought of as a customer-subsidized portion of the employer wage bill.

The paucity of research inquiry into the tipped wage and its tip credit counterpart means the policy and its effects are not well understood. Moreover, employment in the restaurant industry—a heavy user of the low-wage workforce—has been growing. Figure 1 shows that private-sector employment grew by approximately 22 percent from 1990 through 2012, while employment in the full-service (FSR) and limited-service (LSR) restaurant sectors grew by 78 percent and 62 percent, respectively. At the same time, private-sector earnings grew by 20 percent while earnings increased by 14 percent and 2 percent, respectively, for workers in the FSR and LSR sectors.

Employment of tipped workers is common in the FSR sector but not in the LSR sector. We use a panel (1990Q1–2013Q1) of data from the Quarterly Census of Employment and Wages (QCEW) that allows us to separate FSR from LSR. This separation will allow for a more nuanced analysis of the two wage floors. We expect to find earnings effects on the tipped wage for the FSR sector but not for the LSR sector—an important falsification test. We will

FIGURE 1

Employment and Earnings Growth: Private Sector, Full-Service (FSR) and Limited-Service (LSR) Restaurants, 1990–2012



Source: Authors' analysis of Quarterly Census of Employment and Wages data.

also be able to estimate minimum wage effects on earnings and employment separately by sector as the effects need not be the same.

Although the federal tipped wage has not changed since 1991, there is ample variation in state policies. We use state variation in the tipped and the regular minimum wages to identify earnings and employment effects. Although there exists little literature regarding the tipped wage, many of the empirical issues, such as the nonrandomness of wage-floor policies, parallel those found in the literature on the regular minimum wage.

Research using panel data often starts and sometimes ends with the two-way fixed-effects model. The two-way estimation strategy limits time and spatial controls to a single national time trend and state-specific fixed effects. Allegretto et al. (2013) show that observable confounds vary considerably across high and low minimum wage states, suggesting that unobserved factors do as well. Moreover, given the existence of spatial clustering of both wage-floor policies, our research needs to adequately address the issue of spatial heterogeneity. At the crux of the issue is the validity of control groups—or the counterfactual for what would have happened in the absence of a change in the minimum or tipped wage. What research design would best account for wage policies that are correlated, but not causal, to growth patterns of low-wage employment?

We start with the traditional two-way fixed-effects model and add controls for time-varying spatial heterogeneity. While estimates of the impact of the tipped wage on earnings are robust across specifications, estimates on employment are sensitive to the inclusion of spatial controls and suggest a strong negative bias that results in an improbably large negative employment effect in the FSR sector in the two-way fixed-effects specification.

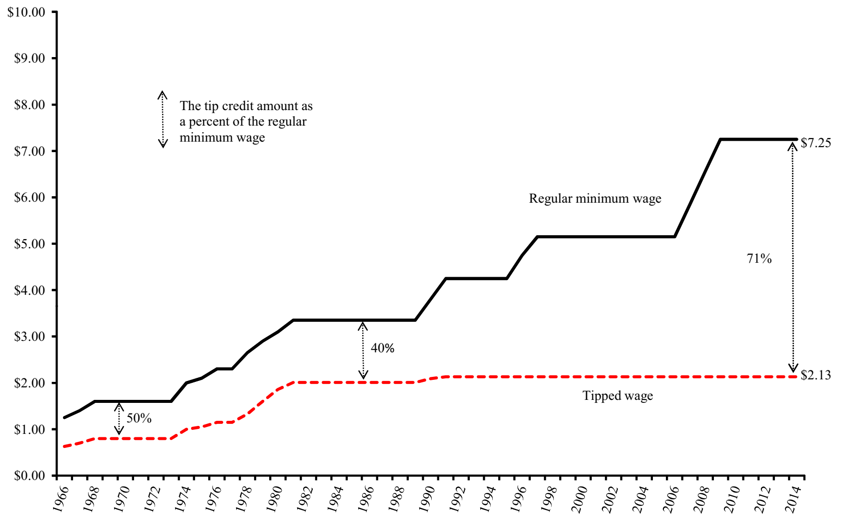
We next extend earlier work by Dube, Lester, and Reich (2010; hereafter, DLR) and compare outcomes between contiguous counties that straddle a state border. Given their geographical proximity, the border counties provide a natural control group when one county's state implements a change in their wage policy and the other county's state does not. Our results from this second analysis support our findings using all counties and including spatial controls. We find a 10-percent increase in the tipped wage increases earnings in the FSR sector by about 0.4 percent. Once we control for spatial heterogeneity, we find small, insignificant effects of the tipped wage on FSR employment.

History of the Tipped Wage

The tip-credit provision. The 1966 FLSA amendments widened the net of labor protections to include coverage for hotel, restaurant, and other service

FIGURE 2

FEDERAL MINIMUM WAGE AND TIPPED WAGE POLICY, NOMINAL VALUE, 1966–2014



SOURCE: Authors' analysis of Fair Labor Standards Act and amendments.

workers. It also introduced a "tip credit" provision that allowed tipped workers to be paid a subminimum or tipped wage that was lower than the regular minimum wage (Elder 1978; Whittaker 2006). The tip credit allows an employer to use tips, provided by customers, as credit toward a tipped worker's wage so long as tips plus the tipped wage paid by the employer equate to at least the regular minimum wage.[2]

Initially, the tipped wage and the tip credit were each 50 percent of the regular minimum wage, as depicted in Figure 2. Over time, the ratio of the tipped minimum to the federal minimum varied—it was as high as 60 percent but didn't fall below 50 percent until 1996. The relatively proportional link between the two wage floors was broken with the passage of the Minimum Wage Increase Act of 1996, which froze the tipped wage at $2.13 into perpetuity. Today the federal tip credit is 71 percent of the regular minimum

[2] Other restrictions apply, such as the worker must make at least $30 per week in tips; for additional information see http://www.dol.gov/whd/regs/compliance/whdfs15.htm. To be in compliance with the FLSA's wage requirements the timing of when to calculate tips plus the tipped wage is assessed on a workweek basis. See 29 U.S.C. 206(a). A workweek is any *fixed and regularly* recurring 168-hour period. A recent (2010–2012) compliance investigation by program analysts at the U.S. Department of Labor reported that 83.8 percent of restaurants had some type of wage and hour violation including 1170, tip-credit violations that resulted in nearly $5.5 million in back wages (email correspondence).

wage, while the federal tipped wage is just 29 percent. Hence, tips are, in part, wages provided by customers via the tip-credit provision.

States act in response to federal inaction. Over the past several decades minimum and tipped wage floors have varied considerably across states. States with minimum wage policies above the federal level ranged from just a few in the mid-1980s to more than thirty in 2008.[3] The number tends to grow considerably when the federal rate is left unchanged for long periods. The situation is a bit different for state tipped wage policies. The number of states with more generous subwage polices has, for the most part, steadily increased over time given the $2.13 federal policy in place since 1991. Seven states do not allow for a tipped wage—in these states all workers are paid at least the regular state minimum wage. In the mid-1980s these seven states, along with five others, had a tipped wage above the federal level—that number increased to twenty-six in 2014.

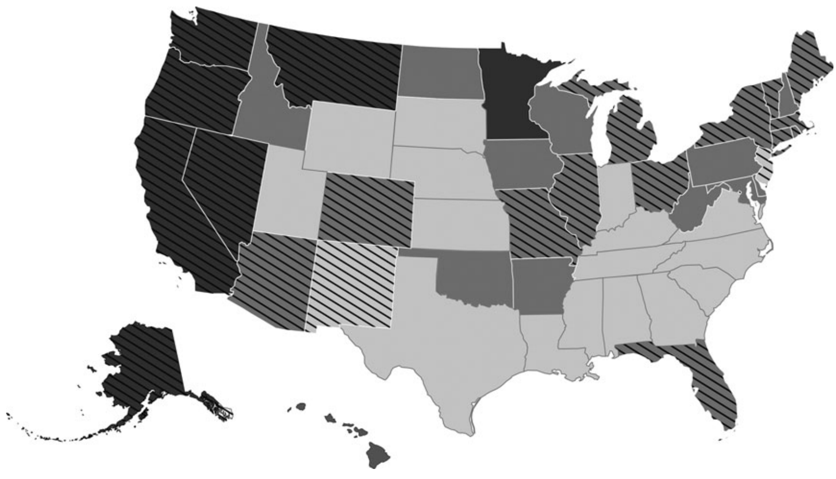
The map depicted in Figure 3 shows state (plus the District of Columbia) minimum wage and tipped wage policies as of January 2014. States with minimum wages above the federal level are marked with black hash marks. The three color codes on the map refer to whether the state tipped wage is set at the $2.13 federal level (light gray), above the federal level but below the regular minimum (medium gray), or if the state does not allow for a tipped wage (dark gray). The three color categories may also be referenced as states with full, partial, and no tip credit, respectively.

The partial tip credit states currently have tipped wages that range from just above the federal level, such as the $2.23 policy in Delaware, to very close to a no tip credit policy, such as Hawaii's tipped wage of $7.00. Of the workforce, 35.5 percent work in full tip credit states, 46.4 percent in partial tip credit states, and 18.1 percent work in the seven no tip credit states (Allegretto and Cooper 2014). It is interesting to note the considerable differences that exist across states at any given time. For example, as of January 2014, the state of Texas followed the federal polices for both wage floors, while Washington State, which does not allow for a tipped wage, had a single wage floor of $9.32. The wage policies at both the federal and state level provide a rich data source with ample variation to examine minimum and tipped wage effects on employment and earnings in the restaurant industry.

[3] Here and elsewhere in the paper the District of Columbia is included as a state. The number of states with higher minimum wages changes depending on changes in state and federal policies. For example, prior to the federal increase in 2007, thirty states had minimum wages higher than the federal minimum, which had been at $5.15 dating back to 1997.

FIGURE 3

STATE MINIMUM WAGE AND TIPPED WAGE POLICIES, JANUARY 2014



NOTES: Hash marks denote states with minimum wage policies above the $7.25 federal rate. Light-gray states follow the federal tipped wage policy of $2.13. Medium-gray states have tipped wages above the federal level but below each state's binding minimum wage. Dark-gray states do not allow for a tipped wage. The state scenarios are always changing as minimum and/or tipped wages change at the federal or state level.

Literature Review

The tipped wage. Unlike the abundant research and debates concerning the regular minimum wage, there has been little interest regarding the tipped wage. A descriptive paper by Allegretto and Cooper (2014) shows that average wages are higher and poverty rates are lower for tipped workers (wait staff in particular) who reside in states that have higher tipped wages. Other descriptive information from Allegretto and Cooper indicates that wait staff are overwhelmingly women—just over 68 percent. And while tipped workers and wait staff are disproportionately young, it is the case that 45 percent and 33 percent, respectively, are at least 30 years old.

The sole published paper that examined tipped wage effects is by Even and Macpherson (2014). They used QCEW data to estimate employment and earnings effects on the restaurant industry. The authors concluded that "results provide fairly convincing evidence that higher cash wages (otherwise tipped wages) increase earnings but reduce employment," but they expressed caution in their degree of confidence and called for additional research (p. 23). In sum, they preferred the canonical two-way fixed-effects estimates and reported that a reduction in the tip credit (in other words, a 10-percent increase in the tipped wage) increased worker earnings by less than 1 percent and reduced employment in full-service restaurants by less than 1 percent.

Even and Macpherson (2014) use two fixed-effects specifications: (1) the traditional two-way specification where they control for time and state fixed effects and (2) the addition of a state-specific time trend to the first specification.[4] Even and Macpherson (2014) prefer the two-way estimator without state-specific time trends. They contend that the inclusion of state-specific time trends, along with the high degree of collinearity between the minimum wage and the time and state fixed effects, overparameterizes the model—washing out the true disemployment effect.

The overparameterization argument is hard to reconcile given the robustness of wage estimates that Even and Macpherson (2014) find for both of their specifications (table 1, p. 643). This result would not follow if the specification with state-specific time trends were overparameterized. Robust minimum wage results are found with even more sophisticated specifications of the fixed-effects model, such as those that use state-linear time trends along with spatial controls (see Dube, Lester, and Reich 2010; Allegretto, Dube, and Reich 2011; Allegretto et al. 2013).

Anderson and Bodvarsson (2005) asked whether states with higher tipped wages boosted server pay. They examined 1999 aggregated data on wait staff and bartenders from the Occupational Employment Statistics from the Bureau of Labor Statistics. Anderson and Bodvarsson (2005) concluded, for the most part, that it does not appear that tipped workers get a boost in total earnings in states with higher tipped wages. The estimate of an earnings effect will be improved by using a panel of data and controls for period and state fixed effects, which are not possible when only a single year of data is analyzed as in the Anderson and Bodvarsson (2005) paper.

A paper by Wessels (1997) theoretically and empirically assessed whether restaurants have monopsony power over wages. Wessels's (1997) tested theoretical model hinged on the fact that tips allow restaurants to pay servers lower wages, and as more servers are hired, each serves fewer customers and consequently earns less in tips—thus restaurants must pay a higher wage to retain workers. Empirically, he concluded that the labor market for tipped wait staff in restaurants is indeed monopsonistic. Wessels (1997) detected the full

[4] They also include additional demographic controls such as the share of the population over 60, the share of the prime-aged (25–60) population, and female labor-force participation rate.

TABLE 1

DESCRIPTIVE STATISTICS FOR THE "ALL COUNTY" AND "BORDER COUNTY" SAMPLES BY FULL-SERVICE (FSR) AND LIMITED-SERVICE (LSR) RESTAURANT SECTORS

	All County Sample				Contiguous Border County Sample			
	FSR		LSR		FSR		LSR	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Population	189,479	434,560	257,058	506,425	222,009	364,542	352,252	451,512
Private-sector average weekly earnings	542	178	570	187	573	207	625	233
FSR average weekly earnings	204	65	218	64	211	70	232	72
LSR average weekly earnings	188	49	188	48	192	52	197	53
Private-sector employment	71,695	179,432	98,886	209,481	89,818	182,474	145,303	230,916
FSR employment	2693	6368	3826	7530	3213	6401	5247	8191
LSR employment	2489	5020	2849	5345	2646	3915	3517	4448
Minimum wage	5.47	1.21	5.46	1.21	5.50	1.23	5.51	1.23
Tipped wage	3.04	1.43	2.93	1.41	3.07	1.31	2.86	1.18
County distribution by Census division								
New England	0.04	—	0.04	—	0.10	—	0.12	—
Middle Atlantic	0.10	—	0.10	—	0.19	—	0.18	—
East North Central	0.22	—	0.17	—	0.22	—	0.15	—
West North Central	0.15	—	0.10	—	0.13	—	0.10	—
South Atlantic	0.16	—	0.22	—	0.15	—	0.22	—
East South Central	0.05	—	0.08	—	0.03	—	0.03	—
West South Central	0.09	—	0.13	—	0.02	—	0.06	—
Mountain	0.10	—	0.08	—	0.12	—	0.09	—
Pacific	0.08	—	0.09	—	0.05	—	0.04	—
Number of counties	1281		890		332		197	
Number of county pairs					281		150	
Number of states (+ DC)	49		49		48		45	
Total county-quarter observations*	119,133		82,770		52,266		27,900	

SOURCE: QCEW 1990Q1–2013Q1 data (except population data) from the U.S. Census Bureau.

NOTE: *Total county-quarter observations indicates the total number of observations within a sample, which has been balanced based on disclosed employment counts within the total private sector and the relevant subsector (either full- or limited-service restaurants). Descriptive statistics for other subsectors have a lower count, because they are based on an unbalanced subset of the sample. In particular, there are 96,803 county-quarter observations with disclosed employment counts of limited-service restaurants within the All Counties sample that has been balanced based on the full-service subsector. Likewise, there are 79,643 county-quarter observations with disclosed employment counts of the full-service subsector within the All Counties sample balanced on limited-service. There are 43,083 county-quarter observations with disclosed counts of limited-service employment within the Border Counties sample balanced on full-service. Lastly, there are 27,149 county-quarter observations with disclosed full-service employment within the Border Counties sample balanced on limited-service.

"reverse C" monopsony employment pattern. He asserted that over some range (not established) a higher wage will increase restaurant employment.

A second paper by Wessels (1993) on minimum wages and tipped employees used the Census of Retail Trade to estimate the effect of allowing a total offset of tips toward the minimum wage requirement. He concluded that restaurant employment would increase by 6.8 percent and those jobs would pay 30 percent or more above the minimum wage (which was $2.01 at the time). Wessels (1993) also found that a 10-percent increase in the tipped wage would result in a 4-percent decrease in employment, and workers who retained their jobs would have their hours cut by 6 percent. In total there would be a loss of 3 percent to 5 percent in total income, coupled with lower employment.

We contribute to this literature in two ways. First, we use more than two decades of variation in state-level tipped wage policy to estimate tip wage effects on earnings and employment. That is, in contrast to earlier approaches relying on cross-sectional data, we use within-county variation for identification. Second, we address potential bias from time-varying spatial heterogeneity using designs from recent research on the minimum wage, focusing on comparisons of counties within the same Census division, or across a state border.

Relevant minimum wage literature. The minimum wage is one of the most studied topics in labor economics. See Brown (1999) and more recently Neumark and Wascher (2008) for an overview of the literature. Recent debate on minimum wages has focused on the importance of research design. Kuehn (2014) gives a broad summary of the two main approaches: (1) those that use the two-way fixed-effects model and (2) those that use "matching criteria." The two-way fixed-effects strategy in its most simplistic form exploits variation in state minimum wages and uses states without minimum wage increases as counterfactuals. We know from looking at the map (see Figure 3) that there is a spatial component to both wage floors, and given that wage policies are not randomly assigned, there is a nontrivial possibility of estimating spurious effects.

As Kuehn (2014) mentioned, different employment trends—for example, the stagnation in the Midwest compared to growth in the South—are due to structural shifts (such as the decline in manufacturing) and not due to minimum wage policies. Allegretto et al. (2013) used four data sets and six approaches—including geographic controls, border discontinuities, synthetic controls, and dynamic panel data models—to show that the two-way fixed-effects estimator for minimum wage studies is biased due to insufficient controls for time-varying heterogeneity.

Research designs such as the case study approach used by Card and Krueger (1994) in their Pennsylvania–New Jersey study and the generalization of that approach by DLR are based on matching criteria—that is, matching or identifying an appropriate comparison group to assess what would have happened to the treatment group in the absence of the treatment. In this case, the treatment is an increase in minimum wage. DLR assert that minimum wage research that relies on the two-way fixed-effects model does not adequately account for unobservable heterogeneity that is correlated, but not causal, to low-wage employment patterns and thus produces spurious negative employment effects.

DLR extended the fixed-effects approach in conjunction with matching criteria by exploiting a research design based on contiguous border county-pairs that assumes counties that are geographically close are better controls than those that are not. In their preferred specification, DLR argued that their research design, which matched a "treatment" county with a neighboring county across a state border as a "comparison" or counterfactual, is preferred to studies that do no such matching. DLR's estimates were essentially a pooled estimate of each contiguous county-pair with a minimum wage differential over a 17-year period. The estimated negative employment effects that resulted from the traditional two-way fixed-effects specification attenuated and became indistinguishable from zero in their preferred specification.

The present research builds upon DLR as we estimate earnings and employment effects for the restaurant industry, but we analyze the full-service and limited-service restaurant sectors separately (DLR pooled them together). As in DLR, we are interested in minimum wage effects, but we extend our analysis to also include the tipped wage.

In relation to the present study it may be that the confounders with the variation in the tipped wage may be similar, but not necessarily identical, to those relating to the minimum wage (as the map in Figure 3 suggests). Hence, spurious effects may differ but heterogeneity remains a potentially serious issue.

The advances of incorporating spatial controls and policy discontinuities to better account for heterogeneity (as presented in Dube, Lester, and Reich 2010; Allegretto, Dube, and Reich 2011; Allegretto et al. 2013) is an often-preferred approach (for example, see Autor 2003; Lee and Lemieux 2010; Magruder 2013), but it is not universally accepted within the discipline. Specifically, research by Neumark, Salas, and Wascher (2014) used a synthetic control approach to argue that areas in close proximity are not more similar. And, as discussed above, Even and Macpherson (2014) preferred the traditional two-way fixed-effects specification.

Data

The QCEW data provide a near census of county-level payroll data on employment and earnings covering approximately 98 percent of all jobs.[5] Importantly, these data are well suited for research on the tipped wage as tipped workers such as wait staff are prevalent in the full-service sector but rare in limited-service restaurants—and the two sectors are separately identified in the QCEW. Furthermore, both restaurant sectors are heavy users of minimum wage workers.[6]

We construct a QCEW panel of quarterly observations of county-level employment and earnings for full-service restaurants and limited-service restaurants from the first quarter of 1990 through the first quarter of 2013.[7] Quarterly employment is the average of the three monthly employment values reported for the corresponding months of each quarter. The earnings variable is the average weekly wage for a given quarter.

The QCEW data include a measure on private-sector employment that we use as a control. Also used as a control is county-level population data for each quarter from the U.S. Census Bureau that was merged with the QCEW data. The dataset is further appended with data on the regular minimum wage and the tipped wage for each state and time period (year, quarter).[8]

We use four subsets of QCEW data. For both restaurant subsectors, we have a sample that includes all counties ("All County" or AC) sample and a subsample of the AC data restricted to contiguous border-county-pairs referred to as the "Border County" (BC) sample. The BC sample is restricted to contiguous county-pairs that straddle a state line and have a minimum or tipped wage differential. Each of the four samples, separately, is restricted to counties that have reported data for all ninety-three quarters. The descriptive statistics are provided in Table 1. The "All County FSR" sample consists of 1281 out of the 3109 counties in the United States with reported data on full-service

[5] QCEW data represent the number of covered workers who worked during, or received pay for, the pay period including the twelfth of the month. Excluded are members of the armed forces, the self-employed, proprietors, domestic workers, unpaid family workers, and railroad workers covered by the railroad unemployment insurance system.

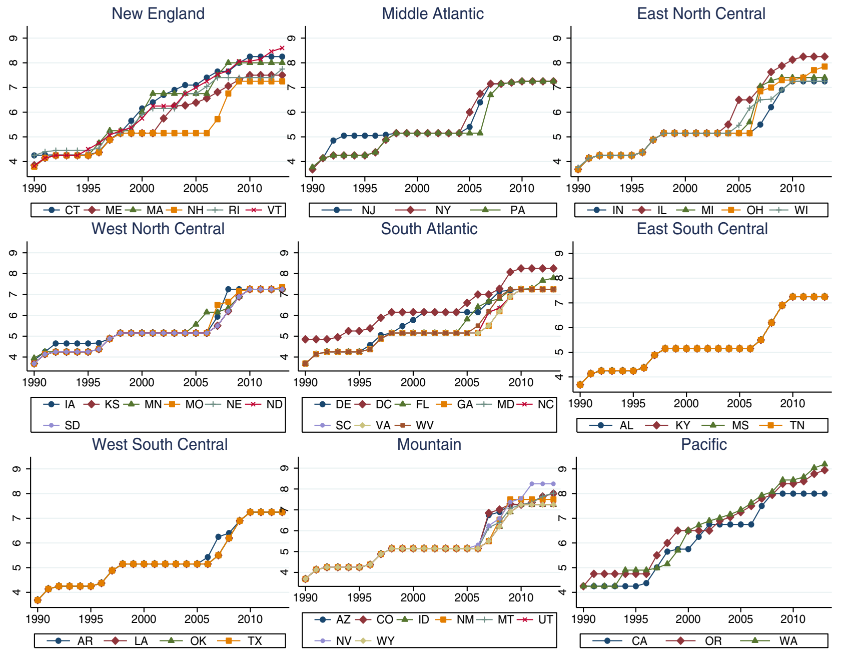
[6] The restaurant industry employs a large share of the minimum wage workforce, and of all workers employed in restaurants, about a third earn wages within 10 percent of the minimum wage (DLR 2010).

[7] NAICS codes from 1990 through 2010: FSR 7221 and LSR 722211; from 2010 onward: FSR 722511 and LSR 722513.

[8] The construction of monthly, state data on the tipped and minimum wage from 1990–2013 were compiled using various sources. Older data were found via annual issues of the *Monthly Labor Review* and its chapter on State Labor Legislation available in January for most years. Other sources were via the Internet and state labor departments.

FIGURE 4

STATE MINIMUM WAGES BY CENSUS DIVISION, 1990–2013



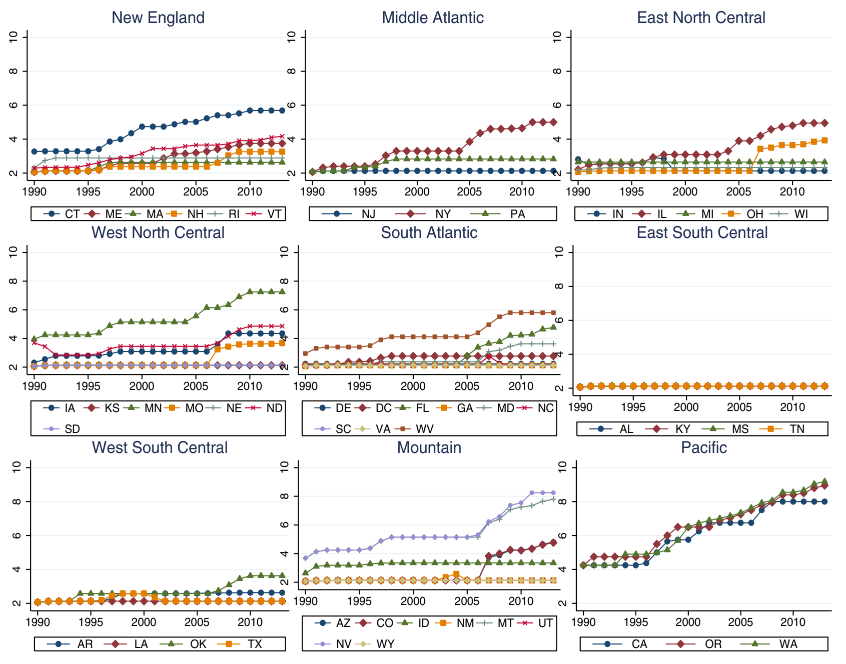
SOURCE: The construction of quarterly state data on the tipped and minimum wage from 1990–2013 were compiled using various sources. Older data were found via annual issues of the *Monthly Labor Review* and its chapter on State Labor Legislation available in January for most years. Other sources were via the Internet and state labor departments.

restaurants. The "All County LSR" sample consists of 890 counties with data on limited-service restaurants. We use this sample to replicate the traditional two-way fixed-effects specification. The spatial depiction of minimum and tipped wages across time is evident in Figures 4 and 5, respectively. The figures illustrate the two wage floors for states within each of the nine Census divisions. Thus, we build upon the canonical model to include spatial controls and state-specific time trends. We discuss the importance of spatial controls further in the Estimation Strategy section.

The BC samples are used for the border-county-pair analysis and consist of all contiguous county-pairs that straddle a state boundary and have a tipped or minimum wage differential. Descriptive statistics are reported in the right-hand panel of Table 1. The BC FSR sample consists of 332 counties and 281 county-pairs, and the BC LSR consists of 197 counties with 150 county-pairs.

FIGURE 5

State Tipped Wages by Census Division, 1990–2013



Source: The construction of quarterly state data on the tipped and minimum wage from 1990–2013 were compiled using various sources. Older data were found via annual issues of the *Monthly Labor Review* and its chapter on State Labor Legislation available in January for most years. Other sources were via the Internet and state labor departments.

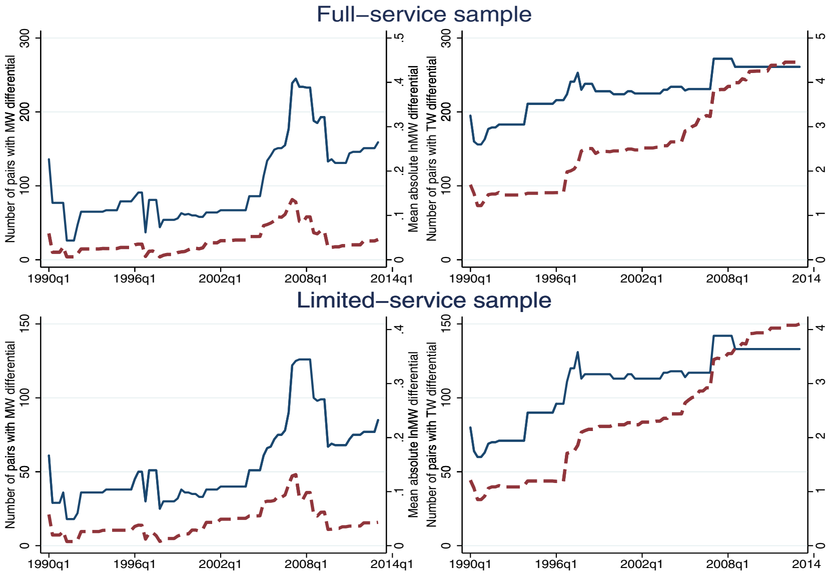
Figure 6 shows the variation in minimum and tipped wages for both restaurant sector datasets along with the mean absolute differential in the (log) minimum wage and the (log) tipped wage. The figures show that there is necessary variation in order to estimate earnings and employment effects. There are substantial pay gaps among these counties, especially regarding the tipped wage that has large increases in later years in the relevant FSR sector.

Estimation Strategy

Specifications using the All County sample. Our first strategy to assess the impact of minimum and tipped wages uses our AC sample of county-level

FIGURE 6

Number of County-Pairs with Minimum or Tipped Wage Differential and Average Absolute Relative Differential



Notes: The solid line plots the number of pairs with a minimum or tipped wage differential in a given quarter. The dashed line plots the average absolute difference in the logged wage between the counties in a pair.

panel data on earnings and employment. We begin by estimating the traditional two-way fixed-effects specification as our baseline model:

$$\ln y_{ct} = \eta_{TW} \ln TW_{s(c)t} + \eta_{MW} \ln MW_{s(c)t} + X_{ct}\Gamma + \phi_c + \tau_t + \varepsilon_{ct} \qquad (1)$$

where c indexes counties and t indexes quarters from 1990Q1 through 2013Q1. $\ln TW_{s(c)t}$ and $\ln MW_{s(c)t}$ are, respectively, the log of the tipped wage and the log of the minimum wage in county c's state, $s(c)$, in quarter t. We define the tipped wage and minimum wage as the maximum of the respective state and federal wage policy. $\ln y_{ct}$ is the log of the labor market outcome of interest, which is either total employment or average weekly earnings in one of two subsectors: full-service or limited-service restaurants. φ_c and τ_t are county and quarter fixed effects, respectively. X_{ct} is a vector of county-level control variables. In all specifications, we control for the log of the county's population, reported on a quarterly basis by the Census Bureau. We also account for economic shocks affecting all industries in the county. In earnings regressions, we include (log) total private sector average earnings, and in employment regressions we include (log) total private sector employment.

Our parameters of interest are η_{TW} and η_{MW}. They represent the elasticities of the dependent variable with respect to the state's tipped wage and minimum wage policies, respectively. Elasticities are estimated without bias under the assumption that $E[\ln TW_{s(c)t}\varepsilon_{ct}] = 0$ and $E[\ln MW_{s(c)t}\varepsilon_{ct}] = 0$. In other words, specification (1) estimates causal impacts of state tipped and minimum wage policies if changes in wage policies are uncorrelated with unobserved changes in the local economy. As previously discussed, there is growing evidence that this condition does not hold. States appear more likely to enact increases in minimum wages during periods in which employment in low-wage sectors is already falling (e.g. Dube, Lester, and Reich 2010; Allegretto et al. 2013). Because changes in tipped wage policies often occur at the same time as changes in minimum wage policies, estimates of the effect of tipped wages on employment may be negatively biased as well.

To address this potential bias, we estimate fixed-effects specifications that additionally control for time-varying regional and state-level economic conditions. To motivate our approach, one may decompose the unobserved within-county variation in a given quarter as follows:

$$\varepsilon_{ct} = \lambda_{d(c)t} + \psi_s t + u_{ct} \tag{2}$$

where $\lambda_{d(c)t}$ is a time-varying, unobserved economic shock that is common across all counties in county c's Census division, $d(c)$; ψ_s is a state-level trend; and u_{ct} is the unobserved heterogeneity that is uncorrelated with $\lambda_{d(c)t}$ and $\psi_s t$. Plugging (2) into (1) we have:

$$\ln y_{ct} = \eta_{TW}\ln TW_{s(c)t} + \eta_{MW}\ln MW_{s(c)t} + X_{ct}\boldsymbol{\Gamma} + \phi_c + \lambda_{d(c)t} + \psi_s t + u_{ct} \tag{3}$$

By including controls for time-varying economic shocks within Census divisions and state trends, the model specified in equation (3) relaxes the previous identifying assumptions. Loosely speaking, identification of η_{TW} and η_{MW} is based on the arguably idiosyncratic timing of changes in state wage policies relative to nearby states presented in Figures 4 and 5, while adjusting for long-term trends in the state economy. In order to gauge the sensitivity of our results, we also estimate intermediate specifications in which we add to equation (1) either division-specific time effects or state trends. In all regressions we cluster the standard errors at the state level, because we are using county-level data to estimate the effect of state-level policies.

Identification using contiguous border-county-pairs. A limitation of our analysis of the AC sample, and of similar studies using state-level panel data, is that we are only able to address the nonrandom timing of state-level changes in tipped and minimum wage policy by using geographically coarse controls or controls for the overall trend within the state. An alternative approach is to focus on local comparisons using pairs of counties that cross a state border. Given their geographical proximity, the border counties provide a natural control group when one county's state implements a change in their wage policy and the other county's state does not. We thus extend the border-county design originally implemented in Dube, Lester, and Reich (2010). Consider a sample of pairs of border counties. To belong in our sample, each pair, indexed by p, must be contiguous and lie across a state border and, at some point in our sample period, have different tipped or minimum wage policies. j indexes the counties that are in a pair: $j = 1, 2$. Because counties can be in multiple pairs, counties appear in the sample for each pair to which they belong.[9] Let $c(j,p)$ denote the jth county in pair p and $s(j,p)$ the county's state. The regression model is:

$$\ln y_{jpt} = \eta_{TW} \ln TW_{s(j,p)t} + \eta_{MW} \ln MW_{s(j,p)t} + X_{c(j,p)t}\Gamma + \phi_{c(j,p)} + \rho_{pt} + \upsilon_{jpt} \qquad (4)$$

where ρ_{pt} is a time-varying economic shock common to both counties in pair $p,$ and v_{jpt} is an unobserved, time-varying, county-level shock uncorrelated with ρ_{pt}. The other variables are as previously defined. η_{TW} and η_{MW} are identified under the assumption that $E[\ln TW_{s(j,p)t}\upsilon_{\mathrm{jpt}}] = 0$ and $E[\ln MW_{\mathrm{s(j,p)t}}\upsilon_{jpt}] = 0$.

This condition requires that local labor-market shocks that would bias our estimate of the effect of the wage policy are also affecting the labor-market outcomes of the county across the state border, adjusting for average differences between the two counties and our time-varying controls. In contrast, the two-way effects model we introduced in equation (1) assumes these confounding labor market shocks are shared across *all counties* in our sample, regardless of geographic proximity.[10,11] Following Dube, Lester, and Reich (2010), we correct our standard errors for clustering within states and border segments of adjacent pairs using multiway clustering (Cameron, Gelbach, and Miller 2006).

Results

Estimates using the All County sample. Table 2 shows the regression results from our analysis on the AC sample. Panel A reports results for

[9] In practice, most counties appear in no more than two pairs.

[10] For comparison with equation (1), we also estimate equation (4) substituting only time effects, τ_t, for the pair-specific time effects, ρ_{pt}.

[11] Because the number of pairs grows with the number of counties, pair-specific time effects are not consistently identified. In practice, we control for this variation by de-meaning at the pair-quarter level.

TABLE 2

ALL COUNTY SAMPLE ESTIMATES OF THE TIPPED AND MINIMUM WAGE ON EARNINGS AND EMPLOYMENT, BY RESTAURANT SECTOR

	Specifications							
	(1)		(2)		(3)		(4)	
Panel A Full-Service Restaurants								
Earnings								
lnTW	—	0.048**	—	0.038**	—	0.034*	—	0.032+
	—	(0.014)	—	(0.010)	—	(0.015)	—	(0.017)
lnMW	0.231**	0.173**	0.204**	0.150**	0.188**	0.165**	0.161**	0.138**
	(0.027)	(0.029)	(0.037)	(0.031)	(0.027)	(0.027)	(0.027)	(0.030)
Employment								
lnTW	—	−0.139*	—	0.006	—	0.014	—	0.011
	—	(0.060)	—	(0.037)	—	(0.072)	—	(0.084)
lnMW	−0.244*	−0.073	−0.017	−0.025	−0.058	−0.068	−0.007	−0.015
	(0.120)	(0.099)	(0.090)	(0.077)	(0.036)	(0.051)	(0.042)	(0.075)
Counties	1,281	1,281	1,281	1,281	1,281	1,281	1,281	1,281
County-pairs								
N	119,133	119,133	119,133	119,133	119,133	119,133	119,133	119,133
Panel B Limited-Service Restaurants								
Earnings								
lnTW	—	0.017	—	0.003	—	−0.021	—	−0.002
	—	(0.015)	—	(0.012)	—	(0.019)	—	(0.020)
lnMW	0.217**	0.196**	0.173**	0.169**	0.182**	0.196**	0.160**	0.162**
	(0.029)	(0.028)	(0.028)	(0.029)	(0.031)	(0.034)	(0.024)	(0.028)
Employment								
lnTW	—	−0.038	—	0.011	—	0.037	—	0.009
	—	(0.051)	—	(0.040)	—	(0.071)	—	(0.064)
lnMW	−0.167*	−0.119+	−0.018	−0.033	−0.073+	−0.098+	−0.015	−0.021
	(0.075)	(0.067)	(0.071)	(0.053)	(0.038)	(0.054)	(0.028)	(0.061)

TABLE 2 (cont.)

	Specifications							
	(1)		(2)		(3)		(4)	
Counties	890	890	890	890	890	890	890	890
County-pairs								
N	82,770	82,770	82,770	82,770	82,770	82,770	82,770	82,770
Controls								
Division-specific period effects			Y	Y			Y	Y
State-specific time trends					Y	Y	Y	Y
County-pair-specific period effects								

NOTES: Results report the estimated elasticities of earnings and employment with respect to the tipped and minimum wage. Standard errors are in parentheses. lnTW refers to the log of the tipped wage policy in a county's state in a given quarter. lnMW refers to the log of the minimum wage. All regressions control for log population and county effects. Earnings and employment models additionally control for log total private sector earnings and employment, respectively. Models (1), (3), and (5) in addition control for quarter effects. See text for construction of the All County sample. In models (1) through (4) standard errors are clustered at the state level. Full results are available upon request. Significance levels are denoted as follows: **1%, *5%, +10%.

earnings and employment in the FSR sector. Panel B shows results for the LSR sector. All models control for log population. In addition, earnings regressions control for log earnings in the private sector. Employment regressions control for log employment in the private sector.[12] In addition to estimating the models we described in the previous section, we also estimate each specification omitting the log tipped wage. We do this for comparison of our results with the related minimum wage literature. We first discuss the AC results, specifications (1) through (4), and wait until the next subsection to discuss the results from our border-county-pair design.

We begin by replicating results from Dube, Lester, and Reich (2010). We exclude the log tipped wage from the regression models and study the effect of raising the minimum wage alone. There are two main differences between our analysis and theirs. First, our QCEW data extend from 1990Q1 through 2013Q1, whereas DLR's QCEW sample ends in 2006Q2. Second, we examine the effects of the minimum wage separately by restaurant subsector. Over all, our results are broadly consistent with theirs (DLR, Table 2). When only county and quarter effects are controlled for, in specification (1), raising the minimum wage 10 percent is estimated to increase earnings a little more than 2 percent. The magnitude of this effect attenuates somewhat as division-specific period effects and state time trends are controlled for in models (2) through (4). Nevertheless, the robustness of this positive earnings effect suggests it is not biased by the nonrandom timing of state changes in wage policy over time. In contrast, estimated effects of raising the minimum wage on employment vary across specifications. Although in specification (1) we find a moderate, negative effect of raising the minimum wage, once division-specific period effects or state trends are controlled for, the effect attenuates dramatically. For example, in full-service restaurants, the effect falls in magnitude over 90 percent from –0.244 to –0.017 once division-specific period effects are controlled for in specification (2).

One difference between our results and those in DLR worth discussing is found in specification (3). Including state trends in the baseline model attenuates the disemployment effects found in model (1), but by a smaller amount than in models (2) or (4) that also control for division-specific period effects. Whereas DLR find a small, statistically imprecise, positive effect of raising the minimum wage once state trends are controlled for,[13] the result in specification (3) predicts raising the minimum wage 10 percent lowers employment –0.58 percent and –0.73 percent in the FSRs and LSRs, respectively. Though both effects are small and estimated imprecisely, the effect on LSRs is significant at

[12] Full regression results are available upon request.

[13] See Dube, Lester, and Reich (2010), Appendix Table A1.

the 10-percent level. Nevertheless, the substantial attenuation of the disemployment effect once division-specific period or state trends are controlled for suggests the two-way effects estimates in specification (1) are negatively biased.

Next, we add the log of the county's state tipped wage to the models in Table 2. Model (1) is our baseline specification, corresponding to equation (1) in the Estimation Strategy section. In addition to the time-varying control variables we discussed earlier, model (1) controls for quarter and county effects. In FSRs, we find that raising the tipped wage 10 percent increases earnings by 0.48 percent. The effect is precisely estimated and statistically significant at the 1-percent level. In model (2), we control for division-specific period effects. The tipped wage effect falls about 20 percent from 0.048 to 0.038, but remains significant at the 1-percent level. Model (3) controls for state trends instead of division-specific period effects. The coefficient on the log tipped wage variable is slightly smaller. When both state trends and division-specific period effects are controlled for in model (4)—corresponding to equation (3) in the Estimation Strategy section—a 10-percent increase in the tipped wage raises earnings 0.32 percent. Though this estimate is similar to those found in models (2) and (3), it is estimated with less precision and is significant at only the 10-percent level.

As we found in models with the tipped wage variable, moving from left to right, the estimated effect of the minimum wage on earnings falls as we add division-specific period effects and state trends but retains statistical significance. In general, comparing specifications with the tipped wage to those without, we find that controlling for the tipped wage causes the estimated impact of the minimum wage to fall in FSRs. Intuitively, since state-level changes in tipped wages are often timed with changes in minimum wages, estimates of earnings impacts of the minimum wage on tipped workers are positively biased when we do not account for changes in the tipped wage policy.

As a falsification test, in panel B, we consider the estimated impacts of the tipped wage on earnings in LSRs. Since workers in LSRs are not generally paid tips, changes in tipped wage policies should not have an impact on their earnings. Models (1) through (4) in Table 2, panel B present the same models we just discussed, except earnings and employment outcomes are specific to LSRs. As expected, across all specifications, we find very small and statistically insignificant impacts of the tipped wage on earnings.

In contrast to earnings, effects of the tipped wage on employment (while also controlling for the minimum wage) are much more sensitive to the inclusion of division-specific period effects or state time trends. Without these controls, the baseline two-way effects estimate in model (1) reports a statistically significant, negative effect of the tipped wage on employment of in the FSR sector. A 10-percent increase is estimated to reduce employment nearly 1.4

percent. A disemployment effect of this magnitude is hard to reconcile with an earnings elasticity of less than 0.5 percent and strongly suggests negative bias. Also surprisingly, once the log tipped wage is controlled for, the impact of the minimum wage in the FSR sector attenuates more than 70 percent and is no longer statistically significant at conventional levels. In models (2) through (4), we find that controlling for either division-specific period effects or state trends or both strongly attenuates the disemployment effect of the tipped wage. In all three specifications, the reported coefficients are very small and positive. Nevertheless, the impacts are imprecisely estimated and cannot rule out small-to-moderate negative impacts on employment. As before, in FSRs, the minimum wage is estimated to have a small negative, though statistically insignificant, impact on employment.

In the LSR sector, in all specifications, increasing the tipped wage is estimated to have an insignificant impact on employment. Although in the FSR sector we find that the tipped wage seemed to explain a large portion of the disemployment effect of the minimum wage in model (1), in LSRs, we find the opposite. In both models (1) and (3), the estimated impact of the minimum wage on employment in LSRs is negative and marginally significant at the 10-percent level. Nevertheless, the sensitivity of this impact to controls for division-specific period effects suggests these impacts may also be somewhat negatively biased.

One concern with including spatial controls such as division-specific period effects in our models is that we may remove valid identifying information (e.g., Neumark, Salas, and Wascher 2014). In our context, we do not feel such a concern is warranted for two reasons. First, the robustness of the estimates of the impacts of the tipped and minimum wage on earnings across specifications demonstrates that state-level variation in wage policies persists even in the presence of division-specific effects and state time trends. Second, if valid identifying information were removed, the inefficiency of our approach would lead us to find larger standard errors once we include these additional controls. This is not generally the case. For example, the standard errors in the employment regression in model (2) are smaller than those in model (1).

Regression results based on the contiguous border-county-pair design. Results from the AC samples strongly suggest that the nonrandom timing of changes in state tipped and minimum wage policies are associated with unobserved regional variation in employment outcomes in the restaurant industry. This correlation negatively biases estimates of the employment effects of these policies in two-way fixed-effects models. So far, we have relied on coarse division-specific period effects and state trends to overcome this bias. In this section, we describe the results from a design that compares pairs of counties that straddle a state border.

TABLE 3

BORDER-COUNTY-PAIR SAMPLE ESTIMATES OF THE TIPPED AND MINIMUM WAGE ON EARNINGS AND EMPLOYMENT, BY RESTAURANT SECTOR

	Specifications			
	(5)		(6)	
Panel A Full-Service Restaurants				
Earnings				
lnTW	—	0.047*	—	0.042*
	—	(0.019)	—	(0.019)
lnMW	0.235**	0.188**	0.187**	0.142*
	(0.037)	(0.039)	(0.056)	(0.059)
Employment				
lnTW	—	−0.075	—	0.070
	—	(0.079)	—	(0.079)
lnMW	−0.096	−0.021	−0.042	−0.116
	(0.103)	(0.090)	(0.079)	(0.097)
Counties	332	332	332	332
County-pairs	281	281	281	281
N	52,266	52,266	52,266	52,266
Panel B Limited-Service Restaurants				
Earnings				
lnTW	—	0.014	—	0.012
	—	(0.027)	—	(0.025)
lnMW	0.213**	0.199**	0.114**	0.099*
	(0.031)	(0.033)	(0.037)	(0.044)
Employment				
lnTW	—	0.012	—	0.051
	—	(0.067)	—	(0.040)
lnMW	−0.143	−0.155*	0.014	−0.044
	(0.092)	(0.072)	(0.083)	(0.081)
Counties	197	197	197	197
County-pairs	150	150	150	150
N	27,900	27,900	27,900	27,900
Controls				
Division-specific period effects				
State-specific time trends				
County-pair-specific period effects			Y	Y

NOTE: Results report the estimated elasticities of earnings and employment with respect to the tipped and minimum wage. Standard errors are in parentheses. lnTW refers to the log of the tipped wage policy in a county's state in a given quarter. lnMW refers to the log of the minimum wage. All regressions control for log population and county effects. Earnings and employment models additionally control for log total private sector earnings and employment, respectively. Models (1), (3), and (5) in addition control for quarter effects. See text for construction of the Border County sample. Standard errors in models (5) and (6) are clustered at the state and border segment level. Full results are available upon request. Significance levels are denoted as follows: **1%, *5%, +10.

Specification (5) in Table 3 estimates the earnings and employment impacts of the tipped and minimum wage controlling only for county and quarter effects, which is similar to the two-way fixed-effects specification in equation (1). That is, we fit models of the form:

$$\ln y_{jpt} = \eta_{TW} \ln TW_{s(j,p)t} + \eta_{MW} \ln MW_{s(j,p)t} + X_{c(j,p)t}\Gamma + \phi_{c(j,p)} + \tau_t + \upsilon_{jpt} \quad (5)$$

The estimated earnings impacts of the tipped wage and minimum wage are remarkably close to what we found earlier (see Table 2, specification 1), despite the fact that we are performing these regressions on a subset of the counties used in the AC analysis.[14] A 10-percent increase in the tipped wage is estimated to raise earnings in FSRs by 0.47 percent. A 10-percent increase in the minimum wage raises FSR earnings 1.9 percent. The log tipped wage and minimum wage coefficients are significant at the 5- and 1-percent levels, respectively. As before, minimum wages are estimated to have a similar impact in LSRs, while tipped wages are estimated to have a small, insignificant impact.

Similar to what we found in specification (1) in Table 2, tipped wages from specification (5) in Table 3 are estimated to have a negative impact on FSR employment. However, the effect in specification (5) is somewhat smaller and not significant at conventional levels. Holding the tipped wage constant, minimum wages are estimated to have an even smaller impact than we found in model (1) in FSRs. Nevertheless, in the LSR sector, a 10-percent increase in the minimum wage is estimated to decrease employment 1.6 percent. This effect is significant at the 5-percent level.

In model (6) we include pair-specific period effects, fitting equation (4) in the Estimation Strategy section. In these models, identification of the effects of the tipped and minimum wage is based on local comparisons between counties and their contiguous neighbors across the state border. The estimated earnings impact of the tipped wage is robust to the inclusion of these effects and maintains significance at the 5-percent level. The estimated elasticity, 0.042, is within the range of earnings estimates we found in the AC sample analysis. Consistent with moderate bias, minimum wage estimates drop about 25 and 50 percent in the FSR and LSR samples, respectively. Both estimates retain significance at the 5-percent level.

Once we control for pair-specific period effects, the estimated employment effect of the tipped wage is now positive. However, the estimate is imprecisely estimated, as a 95 percent confidence interval rules out employment impacts

[14] As noted in Table 2, there are 332 unique counties in the FSR BC sample, compared to 1281 in the AC sample. Likewise, there are 197 unique counties in the LSR BC sample, compared to 890 in the AC sample.

only lower than –0.9 percent and greater than 2.3 percent in response to a 10-percent increase in the tipped wage. Given that we generally find earnings impacts in the range of 0.3 to 0.5 percent, this interval of potential employment responses is consistent with both neoclassical and alternative models of labor-market responses to the increases in the tipped wage. Surprisingly, the tipped wage is also estimated to have a similar positive, though insignificant, impact in LSRs. In contrast, minimum wages are estimated to have more negative, though not significant, employment effects.

Given the imprecision of our estimated employment effects, we interpret our findings from the contiguous border county design as supporting the small, insignificant employment elasticities we found in the AC sample in models (2) through (4).[15]

Summary

The tipped wage has been little researched, especially compared to its better-known and much-investigated minimum wage counterpart. While the federal minimum wage has increased nominally more than 90 percent since the early 1990s, the federal tipped wage remains at $2.13.

We use a panel of QCEW data that spans 23 years to estimate earnings and employment effects of changes in state-level tipped wage policies on FSRs—using LSRs that do not employ tipped workers as a falsification test. We are concerned with spatial heterogeneity given the nonrandom timing of changes in state-level tipped and minimum wage policies. At issue is the validity of control groups—or the counterfactual for what would have happened in the absence of a change in the wage.

We employ two approaches to address this potential bias. First, we estimate earnings and employment effects on a sample of all U.S. counties, controlling for Census division-specific period effects and state trends. Second, we extend earlier work by Dube, Lester, and Reich (2010) and compare outcomes on a sample of contiguous county-pairs that straddle a state border, controlling for time-varying heterogeneity at the pair level.

[15] As a robustness check, in results not shown, we have also estimated models (5) and (6) on an alternative sample balanced on FSRs *and* LSRs. By balancing on both subsectors, we throw out more than half of the counties used in the FSR models, leaving only 165 unique counties. Here, we find similar earnings impacts of the tipped and minimum wage. However, the tipped wage is now estimated to have an employment elasticity of –0.032 in FSRs and 0.064 in LSRs. Both estimates are insignificant at conventional levels. Minimum wage employment elasticities are also small and insignificant: –0.07 in both FSRs and LSRs. Results are available upon request.

Across a range of specifications, we find raising the tipped wage significantly increases earnings of workers in FSRs. Our point estimates suggest a 10-percent increase in the tipped wage raises average earnings about 0.4 percent. In contrast, estimates of the employment elasticity are sensitive to controls for spatial heterogeneity. Models that control for either division-specific period effects, state trends, or border-pair-specific period effects all find small, statistically insignificant employment effects. Although the point estimates of these models are positive, they are also very imprecise. Future research would benefit from richer data on the components of server pay at the worker level. This information would better enable researchers to assess the potential effects of changes in tipped wage policy on those most likely to be affected by policy changes.

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