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The Effect of Recent Increases in the U.S. Minimum Wage: Results from Three Data Sources[†]

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Abstract

This paper investigates the impact on earnings and employment of substantive increases in the minimum wage under the Fair Minimum Wage Act of 2007. Against the backdrop of a thin contemporary literature offering mixed results, our study uses three different data sets, and three different estimation strategies for addressing geographically-disparate trends. Despite the concatenation of seemingly large wage increases and a soft labor market, our evidence is generally unsupportive of material disemployment effects among industrial and demographic groups typically associated with low-wage employment. Our results are consistent with minimum wage workers being concentrated in sectors of the economy for which the labor-demand response to wage increases is seemingly modest.

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I. Introduction

The earliest of the modern literature tackling the impact of minimum wages on labor-market outcomes tended to point fairly uniformly to the existence of small, negative effects of higher minimum wages on employment (Wellington, 1991; Brown, 1999). This result was supported in both time-series studies and state-level panel data research focusing on teenagers. By contrast, a series of studies that focused on employment outcomes within the fast-food sector (most notably, Card and Krueger, 1995) failed to reproduce this finding of a negative employment effect. Even though the robustness of the fast-food estimates to data-measurement issues and statistical approach was soon called into question,[1] recent studies using panel data from the restaurant industry (Addison et al., 2010; Dube et al., 2010) corroborate Card and Krueger's lack of evidence of a negative employment effect. Other contemporary studies considering teenagers, rather than restaurants, have also tended to find less evidence of a general negative effect for teenagers.[2]

What might account for this erosion of adverse minimum wage effects in recent studies? Abstracting for the moment from changes in modeling/statistical approach, it may also be the case that the effect of minimum wages on employment has diminished to such an extent over the years that it is difficult, with currently available data, to distinguish the effect statistically. Specifically, if more recent years have been those in which the minimum wage is low in real and relative terms, adding them to the sample period may well imply that decreasing numbers of, say, teenage workers will be affected by the legislation. Accordingly, disemployment effects will likely be smaller and harder to indentify.[3]

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[1] Kennan (1995) and Neumark and Wascher (2000) questioned the data used by Card and Krueger, with Card and Krueger (2000) responding that such issues did not perturb their primary finding. In addition, Donald and Lang (2007) later criticized the necessary statistical assumptions that underlie the inherent small-scale experiment in the Card and Krueger approach.

[2] See in particular Allegretto et al. (2011). Indeed, Neumark and Wascher (2007a) themselves fail to uncover statistically significant evidence of a negative employment effect for teenagers in general, although disemployment effects are found for particular minority/gender groups. Altough Sabia (2009) still finds evidence of a negative employment effect for teenagers, Allegretto et al. have argued that this result is not robust to specification choice.

[3] Analogously, the opposite circumstance was used as justification by Castillo-Freeman and Freeman's (1992) for their finding of major adverse minimum wage effects in Puerto Rico.

Under the most recent federal legislation, the Fair Minimum Wage Act of May 2007, minimum wages were raised by more 40 percent from 2007 to 2009 – in three separate $0.70 increases in each July. To be sure, many states had already raised their minimum wages above $5.15 before the first installment of the federal increase was due, but the hike in the federal minimum still had an effect in the vast majority of states and was fully effective in over 70 percent of states as of August 2009. Indeed, the effect of these changes in the federal minimum may have been particularly significant, given their major impact was felt in states that had previously not increased their state minima (possibly out of concern that disemployment effects would be more severe in their states).[4]

A second issue, of course is the state of the economy. Unlike other minimum wage hikes since the early 1990s that have tended to take place in relatively tight labor markets, the backdrop to the contemporary increase was one of profound recession, as indexed by a doubling of the unemployment rate to over 10 percent at its peak in October 2009. Taken in conjunction with an acceleration of relative minimum wages, the reversal of favorable labor market conditions after 2007 might be expected to provide a more 'favorable' milieu for observing disemployment effects. Indeed, this very conjecture motivated the research of Even and MacPherson (2010). In concentrating on the post-2004 period, the authors reported evidence of somewhat larger disemployment effects than normally reported in recent research that uses data for earlier periods. Specifically, their estimated elasticities for 16-19 year olds are in the range -0.19 to -0.32, and are at the higher levels of each range with the inclusion of lagged effects. At face value, then, there is some suggestion of an uptick in implied disemployment effects.

We, too, focus on recent data in estimating minimum wage effects on employment. The hallmark of the present treatment is two-fold. First, in examining low-wage sectors and groups – namely, the restaurant-and-bar sector and teenagers – we use three different data sets**:** the Quarterly Census of Employment and Wages (QCEW), the Current Population Survey (CPS), and the American Community Survey (ACS). Second, mindful of issues related to heterogeneity in geographic trends in recent years, we

[4] Bazen and Le Gallo (2009) argue that this is the case, though contrary results are reported in Addison et al. (2010).

estimate our models using three recent approaches suggested for handling the possibility of geographic-specific trends in employment.

II. Recent Minimum Wage Increases

Concern over the impact of the recent increases in the federal minimum wage has most probably centered on the conjunction of two fears: first, the perception that the minimum wage was being increased very substantially; and, second, its particular timing when the economy might not be able to easily 'absorb' the increases (see, in particular, the cautionary remarks of Neumark, 2009).

As regards the former issue, most labor economists were aware that infrequent increases in the minimum wage over the 1980-2005 interval had caused it to fall considerably relative to average wages, and assumed that the 40 percent increase in the mandate would substantially change this situation. Set in context, how sizeable was this minimum wage hike? The top line in Figure 1 plots the real value of the federal minimum wage relative to the average real value of the federal minimum over the 1939-2009 period, where values greater than one signify that the minimum wage exceeds the period average. The final increase in July 2009 did lead to a federal minimum that was higher than the historical average, but just barely. Indeed, the real value of the federal minimum was actually higher throughout the period from the late 1950s to the middle 1980s. On this metric, then, the recent episode is not particularly notable.

The bottom line in the figure charts the federal minimum wage relative to the average manufacturing wage. From this perspective we can see that the federal minimum relative to the average manufacturing wage is only slightly higher in 2009 than it was after the last federal increase in 1997. On the other hand, since the increase in the late 2000s was from a base that had not been that low since 1950, in an absolute sense the increases of 2007-2009 were non-trivial.

Of course, the period since 1980 has witnessed a plethora of state legislation adopting state minima that exceed federal levels. Figure 2 repeats the plot of the federal minimum relative to the average manufacturing wage for the last three decades, but now

also traces that same average when the state minimums are also taken into account.[5] Although state minimums had also caused the overall state average to diverge from the federal minimum in the late 1980s and mid 1990s (both periods in which the federal minimum was not increased over several years), the importance of state minimum wage legislation is noticeably different over the last 10 years. That is to say, once we take state minima into account, the average prevailing minimum wage in 2009 was at its *highest* relative level since 1985.

Turning next to the issue of the broader economic backdrop, the notion that minimum wage effects might also depend on the state of the economy seems worthy of examination. After all, minimum wage jobs tend to be unskilled jobs characterized by high turnover that are most prone to disappear during a downturn. In our case, as we have noted, the minimum wage was increased at the very time the economy was entering into recession. Analyses by Neumark and Wascher (2002) and by Holmes et al. (2009) do indicate disemployment effects that are more severe during recessions, while Allegretto et al. (2011) are unable to detect any role for the cycle.[6]

Studies that cover a similar interval to that examined in the present treatment are no more numerous. The study by Evans and Macpherson (2010), mentioned earlier, uses CPS data to consider the 2005-2009 period and reports negative minimum wage effects on employment for teenagers during this interval. Gittings and Schmutte (2010), using both CPS and Quarterly Workforce Indicators data, also find a negative effect on teenage employment for the longer sample period 2000-2009. Another CPS study by Hoffman and Ke (2011) focuses on just the final installment of the latest federal minimum wage increase in 2009. The authors estimate cross-state differenced models of minimum-wage impact, finding no significant disemployment effects among teenagers but a significantly negative effect for younger workers aged 20 to 24 years. On the other hand, Hirsch et al. (2011), who investigate establishment-level data for quick-service restaurants in Georgia

[5] In calculating the average minimum wage to be used in forming the relative state minimum wage, we weight the minimum in each state by that state's population. The relative state minimum wage calculation includes all states, including those paying the federal minimum.

[6] Cross-country evidence provided by Dolton and Bondibene (2011) reports more detrimental effects of minimum wages during downturns for some groups (namely, youths aged 15-24 years) but not for others (teenagers aged 15-19 years, and adults aged 25-64 years).

and Alabama, 2007-2009, fail to uncover any evidence of a negative employment effect. By way of explanation, the authors conjecture from their surveys of managers that minimum wage increases may elevate technical efficiency along a number of dimensions of business operation (including higher performance standards).

Our goal is to add materially to this literature on the most recent period of minimum wage advance by utilizing three different data sets to comprehensively estimate employment effects for groups typically associated with low-wage employment. In particular, we estimate employment-equation parameters using data just from the 2005-2010 period, enabling us to contrast the minimum-wage effects for this period from the already well-documented effects estimated using data from the period before the "Great Recession." By being selective in the time period examined, we are able to isolate effects of minimum wages that arise in the particular context of that period: a large increase in the minimum wage (from a small base) in the midst of a deteriorating national economy. The tradeoff is that this more focused temporal restriction can lead to less precision in the coefficient estimates of interest, and we will note where this appears to limit our study.

III. Econometric Models and Data

A. Models

Our primary empirical modeling uses the QCEW (or other data) to construct measure of employment and earnings across different counties (or states) in the U.S. The panel nature of our data allows us to exploit both cross-sectional and time-series effects in estimating how these outcomes vary with minimum wages. In particular, our "basic model" is specified as

$$y_{it} = \ _1 \log(MW_{it}) + \ _2' x_{it} + ¥_i + ª_t + £_{it}, \qquad (1)$$

where y is either the log of employment or the log of average weekly earnings, MW is the state-level minimum wage, x is a vector of other covariates potentially affecting y, and i refers to county or state, depending on the data set. In this basic model, fixed-effects estimation directly controls for geographic (¥) and time (ª) effects.

Trends in the dependent variable that vary across counties are handled by using "clustered" standard errors robust to the presence of heteroskedasticity and to any correlation in errors between observations in the same state. The use of these standard errors is particularly important in the QCEW given that our primary variable of interest – the minimum wage – is measured at the state level. This "basic model" requires that such trends are uncorrelated with the state minimum **(**and *x***)** in order to ensure consistent estimated for the minimum-wage effect.

As noted in the introduction, estimates of minimum wage elasticities can be sensitive to the particular time period examined, and to the manner in which geographically-disparate trends are handled in the estimation. In extensions to the above basic model, we employ two different methods of directly controlling for county-specific trends. The first is to explicitly incorporate county-level linear trends in the estimation of the model, so that our estimating equation becomes

$$y_{it} = \ _1 \log(MW_{it}) + \ _2' x_{it} + ¥_i + ª_t + ¢_i t + £_{it}. \qquad (2)$$

Given the large number of counties in the QCEW sample, these trends are most easily incorporated by initially detrending all variables at the county-level, and then using these detrended variables in the estimation. Both Addison et al. (2010) and Allegretto et al. (2011) show that incorporating such trends can materially affect minimum wage estimates (for the 1990-early 2000s period). In the present application, the shorter time-frame focus of our study might suggest that such detrending is less important. However, given the fact that the employment effects of the recession were not shared equally across the U.S., controlling for heterogeneity in trends could still be necessary.

An alternative approach to handling geographically-disparate trends is to study border pairs (as in Dube et al., 2010). In this estimation, every pair of counties bordering each other but in different states is designated as a separate county pair (denoted with the subscript *p*), and the model is specified as

$$y_{ipt} = \ _1 \log(MW_{ipt}) + \ _2' x_{ipt} + ¥_i + ¦_{pt} + £_{ipt}. \qquad (3)$$

Estimation of this model by fixed effects essentially becomes the estimation of a model that differences the dependent and independent variables (including the county fixed effect) across the two counties in any given pairing. This approach has the advantage over the linear-trend model in equation (2) in that the county-pair effect (θ_{pt}) is allowed to vary nonparametrically over time. On the other hand, it does require that this nonparametric trend be shared by the two counties. An additional limitation is that the same county will be included in the data more than once if it borders more than one county in the other state, implying the underlying nonparametric trend for any border county can vary each time the county is included in the sample. To avoid overemphasizing counties in multiple pairs, we follow Dube et al. in using two-way clustering for standard errors.[7] We emphasize that the approaches implicit in estimating equations (2) or (3) each have their strengths and weaknesses, and we are partly concerned with the robustness of our results to the particular method used to control for geographically-disparate trends.

In using the CPS and ACS data, reliable measures of employment and earnings are available at the state level, but not at the county level. As a result, in incorporating these data into our analysis, we modify the above border-county approach to consider instead all border-state comparisons, essentially generalizing the kind of two-state comparison considered in Card and Krueger (1995). As with the border-county estimation, any state that borders more than one other state will have its data included more than once in the estimated model, so we use two-way clustering for standard errors here as well. Admittedly, these comparisons are much cruder than the border-county comparisons possible with the QCEW, but it is hopefully still worthwhile to examine the robustness of findings to this alternative trend comparison.[8]

B. Datasets

1. QCEW

[7]The use of two-way clustering of standard errors allows for correlation of any observations from the same state (note that this includes multiple repeats of the same county/quarter observation that necessarily appear in the data in this approach) as well of any observations that are in different states but share a common state-pair border.

[8] In fact, the New Jersey-Pennsylvania comparison in Card and Krueger was restricted to areas that were in the eastern half of Pennsylvania and the western half of New Jersey.

Our first set of models is estimated using data from the Quarterly Census of Employment and Wages (QCEW). These data are collected by the Bureau of Labor Statistics (BLS) from paperwork filed by employers with the unemployment insurance program, and so have excellent coverage. Geographic information down to the level of the county of the establishment is available in the public use data, and we construct measures of employment and earnings at this level. The industry of the establishments is coded using the North American Industrial Coding System (NAICS), allowing us to construct measures of total employment and average weekly earnings for specific sectors of the economy.

One limitation of the QCEW public use data is that the BLS censors observations when the number of establishments in a sector in the geographic area is low, leading us to omit many small counties from the analysis. Usually observations are present for either every quarter or for no quarter for any given county, and so we construct a balanced panel of counties with every quarter available in our analysis of the QCEW. Another limitation of the QCEW data is that no information on hours of work is available, so we are only able to construct a weekly earnings average.[9] For a more detailed description of the QCEW data, see Addison et al. (2010).

Our choice of sectors to analyze with the QCEW data centers on those that tend to pay low wages, but also with a tendency to have a sufficient number of establishments in a county so as to lessen censoring problems. Following both Addison et al. (2010) and Dube et al. (2010), much of our analysis will focus on *Food Services and Drinking Places*, as this is the sector in which minimum wage employment is most prevalent. The QCEW data also allow us to disaggregate employment within this general sector, and so we also estimate models that use *Full-Service Restaurants* and the *Limited-Service Eating Places* as separate subsectors for which we estimate employment and earnings regressions.[10,11]

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[9] This measure includes most wage-like compensation, including tips, bonuses, stock options, and employer contributions to retirement plans.

[10] The balance of the *Food Services and Drinking Places* sector is made up of drinking places and caterers, both of which have a greater likelihood of being censored than the two subsectors we do analyze.

[11] A share of the workers in the *Full-Service Restaurants* sector includes tipped employees, for whom minimum wage legislation still generally applies. State and federal minimum wages for tipped employees do exist, and are frequently tied to the minimum wage in effect. Specifically, while there is variation in how much "tip credit" is allowable, in all states the hourly wage including tips for "exempt" employees must

The independent variables in our models include relatively standard measures of demand/supply factors that could reasonably affect employment within low-wage sectors in a county. Most importantly, our minimum wage variable is the effective minimum wage measured at the state level (discussed above). For quarters in which the state raised the minimum wage in the middle of the quarter, we calculate the minimum wage for that quarter as the weighted average of the older and newer minimum (with weights depending on the percentage of the time in that quarter in which the particular minimum is in effect). Population is included to capture effects of the size of the county on employment, while total employment in the county (from the QCEW) reflects effects from the size of the labor market. Having both employment and population in the model effectively controls for the state of the economy in the county as well. The measure of average weekly earnings across all sectors is also included as a control, capturing whether the county has particularly high or low average wages in that quarter.[12]

Several models estimated in the minimum-wage literature can be seen as restricted versions of the model we estimate, given that all of our variables enter in logarithmic form. For example, our model is equivalent to one in which the percent of the population employed in the *Food Services and Drinking Places* is the dependent variable (in logarithmic form), if the coefficient on population is equal to -1. Also, a version of our model in which the minimum wage variable is entered as the log of the minimum/average wage ratio (similar to the Kaitz index) would follow if the coefficients on the minimum and average log wages were equal in magnitude but of opposite signs. However, we do not impose these restrictions, as there may be other avenues for population or the average

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still equal (or in some cases exceed) the highest minimum wage enforced in that state (whether it be state or federally determined). Moreover, in several states (Alaska, California, Minnesota, Nevada, Oregon, Montana, and Washington) no tip credit is allowed to employers, so the state minimum wage applies to all employees regardless of their tip status. In twenty-six other states a partial-credit is enforced. Therefore, for employees in the *Full-Service Restaurants* sector most employees (tipped or non-tipped) will be affected by changes in minimum wages laws as the mandate will, at the very least, impact the tip-plus-wage rate they must legally be paid. Nevertheless, we shall check the robustness of our results for the *Full-Service Restaurants* sector by alternatively treating the minimum wage as fixed for states where sub-minimum wages for tipped employees have not varied in recent years.

[12] We also experimented with two other controls, namely the county unemployment rate, taken from the Local Area Unemployment Statistics, as an indicator of labor market slack and the school enrollment rate for individuals aged 16-24 years, and measured at the state level using the Current Population Survey, as a potential demand/supply indicator. Their inclusion did not materially influence the earnings or employment effects of minimum wages in any of our estimations.

wage to affect employment. These restrictions are often not supported in many of our estimated models.

Detailed summary statistics for our measures across the three sectors we analyze are provided in Table 1.

2. *The Current Population Survey*

Although our initial analysis focuses on the QCEW data, we also estimate models for employment and earnings at the state level, using monthly data from the Current Population Survey (CPS). The CPS has been the most widely-used source for data in minimum wage studies since the work of Card (1992) and Neumark and Wascher (1992). Most of this research has concerned potential effects of the minimum wage on particular demographic groups, most commonly teenagers aged 16 to 19 years, but also ranging further afield to include young adults aged 20 to 24 years and/or high-school dropouts in some studies (see the surveys by Neumark and Wascher, 2007b, 2008). In common with most of the literature, we focus on teenagers in our analysis of demographic employment and earnings effects, although findings for the two other groups are available from the authors upon request. When using the CPS, the complete set of respondents to the basic monthly survey is used in measuring teenage employment. However, as earnings information is asked of the approximately one quarter of the sample that is part of the Outgoing Rotation Group (ORG), we have considerably fewer observations when measuring average earnings for this group.[13] That said, one advantage of the ORG sample over the QCEW data is that we can measure average hourly earnings: workers are asked their hourly wage if they are paid by the hour, but information on hourly earnings can be constructed from reports on usual hours worked for those workers whose earnings are specified at a weekly (or higher) rate.[14] Also, given the aggregation of the observation

[13]We find it odd that much of the recent research appears to use only the ORG samples to measure both employment and earnings, when more accurate measures of employment are available using the full monthly CPS.

[14] In calculating the state averages for earnings, we drop all observations for which the relevant earnings level was imputed in the CPS. This addresses concerns raised by Hirsch and Schumacher (2004) that regression estimates may be biased when imputed earnings are used for dependent-variable observations when the imputation is not based on values for relevant independent variables. The CPS does not use state as relevant individual information in constructing earnings imputations, so imputed earnings amounts are not appropriately matched to the minimum wage value possibly relevant to the determination of that earnings value.

from the county level to the state level, we are considerably less likely to experience missing observations due to lack of data with the CPS.

Although the CPS data have been used quite heavily in ascertaining minimum wage effects for particular demographic groups, they have not been widely used to address outcomes in specific industrial sectors. Since January 2003, however, the CPS has used the NAICS coding system for industrial classification, thereby allowing us to replicate our industrial-level QCEW analysis using the CPS data. The sole limitation in this regard is that the CPS does not disaggregate *Food Services and Drinking Places*, meaning that we are unable to undertake a separate analysis of full-service and limited-service restaurants.

We should also note that there are a few minor changes in definition for some of the independent variables in our CPS analysis. In particular, the average wage variable is now calculated using prime-age men only. The employment control is also for adults only (those workers age 25 or older). Means and standard deviations for the CPS data are reported in Table 2. Average employment levels are higher with the CPS than the QCEW because the CPS averages are at the state level, not the county.

3. *The American Community Survey*

The U.S. Census Bureau instituted the American Community Survey (ACS) as an ongoing survey that obtains data similar to the former long-form of the decennial census (which it replaced). The ACS surveys individuals every month, with a sample size that is roughly five times larger than that available from the CPS. Given that earnings and labor-supply questions are asked in the ACS, it potentially serves as a useful alternative to CPS data in labor economics research. To be sure, the sampling frame for the first years of the ACS was relatively small but, beginning with the 2005 survey, this increased to a sample of roughly one percent of the population. Nonetheless, we begin our analysis with the 2006 data, as only in that year did the ACS start to include individuals living in group quarters (such as college dorms). Our analysis ends with the 2009 data, the latest year in which an ACS public use sample was available at the time of writing. We are unaware of any previous minimum wage research that has made use of the ACS.

Although the ACS has the considerable advantage of providing more precise measures of the variables of interest, it has one important weakness: it does not identify the month in which the interview takes place, only the year. As a result, current employment status reported by respondents in the ACS could refer to any of the 12 months of the year. If minimum wage changes were initiated at the start of any year – as is most common with state-level changes – this would pose little difficulty. But much of the variation in minimum wages comes from increases in the federal minimum over 2007-2009; and all such increases occurred in July. This timing problem is perhaps not severe in analyzing employment, where we use annual measures and assign a weighted average minimum wage (similar to the quarter-based minimum-wage averages used with the QCEW). However, the fact that *earnings* in the ACS are measured in retrospective fashion – referring to total earnings in the previous twelve months – implies that reported earnings could be related to any month in a two-year period (so that the corresponding minimum wage variable would need to be a weighted average over a two-year period). For this reason, and given that we are less interested in earnings effects than employment effects, we restrict our analysis to employment outcomes in analyzing the ACS data.

As with the CPS, the ACS allows us to analyze employment outcomes for either demographic groups or industrial sectors. The ACS also uses the NAICS coding for industry, but like the CPS does not separately identify limited-service and full-service restaurants. With the exception of wages, our control variables are taken from the ACS. Problems with measuring earnings led us to use the prime-age male hourly earnings for that state and year from the CPS, simply averaging across the 12 months for any state to obtain our annual measure of this variable. Means and standard deviations for the variables in this analysis are reported in Table 3.

IV. Empirical Findings

A. Results from the QCEW

Our initial set of estimates focuses on minimum wage effects on both employment and earnings in the general *food services and drinking place*s sector. Our expectation is that an increase in the minimum wage should lead to an increase in average earnings among those still employed after the minimum increase. This effect would partly stem

from workers whose wages are directly increased by the increased minimum, but it could also reflect spillover effects for higher-wage workers (as well as compositional changes that shift demand towards those workers).[15] Failure to find evidence consistent with this expectation could cast doubt on the particular empirical approach being used. Although recent minimum wage research for *food services and drinking places* has pointed to an absence of employment effects during the 1990s and early 2000s, the different nature and environment of minimum-wage hikes in the late 2000s could lead to different findings for this latter interval.

Table 4 reports estimates for three different specifications for handling geographic-specific trends. Our "basic" specification includes time dummies for each quarter in our sampled period. It thus controls nonparametrically for a general national trend in employment, but allows for no direct controls for county-specific variation in that trend (see equation 1). The "county-trend" specification (detailed in equation 2) allows for a separate linear trend in each county (in addition to the general nonparametric national trend). Finally, the "border-county" specification allows for a shared nonparametric trend between any two border counties (see equation 3).

The estimated minimum wage coefficients in the earnings equations are generally as expected. Incorporating linear county-specific trends has little impact on this result. The border-county analysis reduces the size of the estimated minimum wage impact on earnings. The estimates are of a somewhat smaller magnitude than those provided in Addison et al. (2010) and Dube et al. (2010).[16] These earlier papers failed to find statistically significant effects on employment in the broad restaurant-and-bar sector, and the same is again true here for all three specifications. The coefficient estimates for the control variables are generally as expected and somewhat similar across specifications.

Although the coefficient estimates for the minimum wage variable in the employment equation are of the expected sign across all three specifications, none is statistically significant at conventional levels. The estimates controlling for

[15] In the case of the QCEW, the one mitigating factor is that we study weekly rather than hourly earnings, and weekly earnings could fall if employers lowered worker hours sufficiently. An hours elasticity greater than one would be necessary for this to occur, however, which result seems very unlikely given prior work on this issue (for example, see Zavodny, 2000).

[16] In Addison et al.'s (2010) preferred specification with county-specific trends, the estimated earnings elasticity is 0.171, while in Dube et al.'s (2010) preferred specification using border-county pairs the estimated elasticity is 0.188. Both estimates are statistically significant at conventional levels.

geographically-disparate trends are also quite precisely estimated. For example, using county-specific trends the minimum-wage employment elasticity has a 95 percent confidence interval with a lower range of -0.05, so that the often-claimed elasticity value of -0.2 would still be significant here. The border-county analysis provides somewhat less precise estimates, but the upper range of a similar confidence interval is still roughly -0.1. Even though the time period examined is short, the data are sufficiently informative to rule out even the small but significant elasticities often reported in the earlier minimum-wage literature.

Separate estimates for the *limited-service* and *full-service* restaurant sectors are reported in Table 5. Our expectation is that as the use of minimum wage labor is more prevalent in the limited-service (fast-food) sector, the minimum wage impact on earnings should also be higher. This expectation was supported in both Addison et al. (2010) and Dube et al. (2010), and it is also borne out here. The two earlier studies differed with respect to employment effects across the two subsectors – Addison et al. (2010) reported a negative employment effect of minimum wages in the limited-service subsector but a positive one in full-service restaurants, while Dube et al. (2010) failed to find a statistically significant effect in either subsector. Here, the border-county approach suggests the same conclusion, but the county-trend specification now points to a statistically significant negative effect for the full-service subsector. That said, this latter estimate is quite small – an elasticity of around -0.05 – and is significant due to the very small standard error provided by the county-trend analysis. Even an elasticity of -0.1 would lie outside the 95 percent confidence interval. Interestingly, most of the standard errors are smaller in the current set of results compared with the parallel results in the two earlier studies, despite the smaller time period. This outcome likely stems from the greater variation in the minimum wage variable in the later period compared to the earlier data.[17]

[17] To explore the sensitivity of the results in the *Full-service Restaurant sector* to the treatment of minimum wages for tipped employees (see footnote 11), we constructed a new minimum wage variable that holds the nominal minimum wage constant in states where changes in the regular minimum wage occurred but where the tipped-worker sub-minimum was not changed. While the employment coefficient estimate for the baseline method is now larger in magnitude and statistically significant when we use this alternative measure of the state minimum wage (an estimate of -0.096, and a standard error of 0.025), conclusions from the county- trend and border-county results are not affected, although the border-trend estimates are more indicative of a negative effect.

Given the difference in the employment results for *full-service restaurants*, we also estimated models that control for both kinds of trends in estimating minimum-wage effects. To be specific, we re-estimated our border-county model also including county-level linear trends in the specification (the model already incorporates county-level fixed effects). While this may seem like an excessive number of trend controls, it does allow us to explore the robustness of the results for *full-service restaurants.* In this expanded model, the coefficient estimate (standard error) of -0.055 (0.044) remains statistically insignificant (as in the border-county result), but it does become negative and actually increases in value relative to the county-trend result. Nonetheless, an elasticity of -0.2 would lie outside the confidence interval in this more general specification.

Next, we return to our broad *food services and drinking places* sector to examine the role of lagged minimum wage effects as well as potential differences in effects during the recession years. Past research has often found support for lagged effects from minimum wage changes (see, in particular, the discussion in Neumark and Wascher, 2006). The justification is that adjustments to mandates are not instantaneous, most notably on the employment side where lags of up to a year have commonly been examined. To avoid expanding dramatically the number of parameters to be estimated, we use a five-quarter moving average of the minimum wage with four lags, equally weighting the current quarter and all quarterly values of the minimum wage within the previous year. Results of this exercise are contained in the upper panel of Table 6. Only the estimated employment elasticity for the border county specification is statistically significant. By contrast, the minimum wage coefficient estimates are uniformly significant for the earnings equation.

The lower panel of Table 6 allows the minimum wage effect to differ between the 2005-2007 and 2008-2009 intervals. The first period pinpoints years in which the economy was still growing while the second period isolates the recession years when the unemployment rate was significantly higher.[18] It is also the case that much more of the variation in the minimum wage in the later period was due to federal minimum wage changes, so that this comparison should also inform as to whether it makes sense to draw

[18]The NBER identified the recession as beginning in December 2007, with the trough occurring in June 2009. Of course, unemployment remained high throughout 2009 and 2010, staying above 9 percent in every month. In comparison, it was below 5.5 percent in every month over 2005-2007.

a distinction between the effects of federal and independent state minimum wage initiatives. From the lower panel of the table it is apparent that the earnings results are scarcely impacted, although in the county-trend specification there is some indication of a modestly higher minimum wage effect in the later period. As far as employment is concerned, there is some indication of a negative minimum wage effect that is present only in the recessionary years, but this result disappears once either method is used to incorporate controls for geographic-specific trends.

By way of summary, our estimates using the QCEW do not provide robust evidence of a negative (or positive) impact of recent minimum wages on employment within the sectors analyzed during this period of increasing minimum wages and economic contraction. They also easily reject employment elasticities between -0.2 and -0.3, a level often suggested for minimum-wage effects in low-wage labor markets.

B. Results from the CPS

The CPS allows us to examine minimum wage effects both at the industry and demographic-group level. Observations here can be measured monthly, but only at the state level. We attempt to assimilate the estimation approaches used with the QCEW data, adding a separate state-specific linear trend in the "state trends" specification and comparing all border states (instead of border counties) in the "border-state" specification.[19] Although there has been much recourse to the CPS to estimate minimum-wage effects for demographic groups, especially teenagers, it has been much less used for estimating effects at the industry level.

Table 7 reports estimates for employment and earnings equations for the general *food services and drinking places* sector. As can be seen, in this instance it is necessary to include some kind of state-level trend controls in order to document significant evidence of a minimum wage increase on earnings. As was found for the QCEW data, the estimated elasticities for employment are all negative, but in the case of the border-state

[19] Recall that a state may be entered multiple times in the border-state specification if it borders more than one state. This explains why the sample size increases by more than four-fold in that specification compared with the other CPS specifications. By contrast, in the case of the QCEW analysis, the border-county approach actually reduces the sample size because a large number of counties in the state's interior are lost in that data. Standard errors in the border-state analysis are calculated allowing for two-way clustering for both the state and the state pair.

specification the coefficient estimate for the minimum wage is now statistically significant. While not reported in the table this finding is not robust; incorporating linear state trends in the border-state model leads to an insignificant estimated coefficient (standard error) of -0.054(0.072). Obviously, the more-aggregated geographic focus of the CPS leads to less precise estimates than does the QCEW; an employment elasticity of -0.2 would be just in the 95 percent confidence interval using both the state-trends and the border-state specifications.

Estimates of minimum wage impacts on teenagers are given in Table 8, and provide a major point of contact with the wider literature. One peculiar result here is the failure of the state-trend specification to suggest evidence of a positive impact of minimum wages on earnings over the 2005-2010 interval. Although the minimum-wage point estimates of the basic and state-trend specifications are of similar magnitude, the higher standard error for the state-trend specification renders the coefficient insignificant.[20] The border-state analysis also provides evidence consistent with this expectation of an increasing impact on earnings. However, only the state trends specification provides (marginally) statistically significant evidence of a negative employment elasticity. That said, including a state-level linear trend in the border-state model suggests that the evidence for a negative effect is not robust; in this case, incorporating both border-state dummies and linear state trends leaves the estimated employment elasticity coefficient estimate (standard error) at -0.036 (0.081). Again, the CPS provides less precise estimates than the QCEW, so an employment elasticity of -0.2 would still be in the confidence interval using all but the border-state specification.[21]

Anther seeming peculiarity concerns our basic equation without state-level trends. Here the insignificant employment results are at odds with the findings of Even and MacPherson (2010), who report a statistically significant negative employment elasticity for teenagers when using a somewhat similar specification. When we replicate their specification, we too obtain a negative and marginally significant minimum wage

[20] Nonetheless, the state-specific trend coefficients are jointly statistically significant, with a p-value less than 0.00001, so the specification that includes these trends is preferred to the basic model based on this specification test.

[21] We also estimated similar specifications for young adults (age 20-24) and junior-high dropouts, as these demographic groups have been examined in some previous studies. After controlling for geographically-disperse trends, the employment elasticity for neither group is statistically significant. These results (and any other unreported results referred to) are available from the authors upon request.

coefficient estimate. However, this result no longer applies when we generalize the national-trend controls included in the model. In particular, Even and MacPherson use a set of month-of-the-year dummies, along with annual dummies to control for trend. But in this period monthly effects vary by year, as the recessionary effects worsened throughout 2008 and 2009. The minimum wage was raised in the middle of these years, and therefore picks up the effects of the worsening recession. Our specification includes a separate dummy for each month in our sample, and so does not impose this restriction. It is also the case that simply weighting by state population also reduces the size of the estimated minimum-wage elasticity, leaving it insignificant even under Even and MacPherson's trend specification,

Finally, extensions of our basic model were estimated, with results presented for *food service and drinking places* in Table 9 and for teenagers in Table 10. First, we examine the possibility of lagged effects by using a 13-month moving average of the minimum wage, equally weighting the current and all previous monthly value of the minimum wage within the previous year. The evidence for an earnings effect for *food services and drinking places* using the state trends specification is weakened, perhaps unsurprisingly given that there should be little lagged effect on the earnings side. The estimated employment elasticities are statistically insignificant other than for the border-state specification in Table 9, although this result is again not robust to including state trends in the specification.

Second, we allow minimum wage effects to differ over the cycle in the same manner as in Table 6. The results are given in the bottom panels of Tables 9 and 10. For *food services and drinking places* (bottom of Table 9), none of the specifications provide evidence of a difference in employment effects between the 2005-2007 and 2008-2010 periods. For teenage workers (bottom of Table 10), however, there is evidence of a negative employment effect restricted to the recessionary period, both for the basic specification and for the specification with state-level trends. The sum of the two coefficients estimates is the estimated minimum-wage effect during the 2008-2010 period, and this sum is statistically different from zero.[22] In this case, the state-trends

[22] One statistical limitation of this result is that we cannot reject that the difference in the coefficients before and after 2007 is zero, nor that that the coefficient in the 2005-2007 interval is zero.

result for employment is robust to including border-state dummies, suggesting evidence of a negative employment effect for teenagers concentrated during the recession.[23]

C. Results from the ACS

As noted in section III, our ACS analysis is restricted to measures constructed at the annual level and deals with employment outcomes only. Data considerations also limit the starting date of our analysis to 2006 (and, as with the QCEW, data were not available for 2010 at the time of writing). In all other respects, however, the ACS allows for similar measurements to those taken from the CPS. Although less ideal than the CPS because of these restrictions, the ACS does have the advantage that its larger base sample allows for accurate measurement of the underlying variables. Yet, it has not previously been used in studies of the minimum wages.

As in the case of the CPS, we estimate employment equations both by sector and demographic group. The left panel of Table 11 presents the results for food services and drinking places. Despite the use of annual measures, and a narrower time frame, the precision of the ACS coefficient estimates is generally better than those obtained from the CPS. Nevertheless, the minimum-wage coefficient estimates from all three specifications are statistically insignificant. Results for teenagers are given in in the second panel of the table, and unlike the case of the CPS, none of the minimum wage coefficients for teenagers is statistically significant. In both sets of estimates, an elasticity of -0.2 lies outside the 95 percent confidence interval, with the sole exception of the teenage elasticity when using the state-trend specification. In sum, the ACS provides reasonably precise estimates, but does not provide corroborating evidence of a negative impact on employment for teenagers or for workers in the food services and drinking places sector.

V. Conclusions

Much recent research on the employment effects of minimum wages has failed to provide convincing evidence that increases in the minimum wage are associated with material reductions in employment. One possible explanation for negative findings of this

[23] The test of no minimum-wage effect on employment for teenagers during the 2008-2010 period had a p-value of 0.04 in the specification combining both trends. I need to check this.

type might be that the typical increases in wage minima over 1980-2005 have simply been too small, taking wages to levels that are experientially low in both real and relative terms. Combined with potentially low own-wage elasticities for low-skilled labor in many service-type sectors, such modest increases may have not provided sufficient identifying information or traction for us to detect statistically any debilitating effects of minimum wages on low-wage employment.

In the present paper, we have focused on contemporary economic developments in the U.S., motivated by the perception that recent minimum wage increases have been altogether more substantial than in the past. This perception although not strongly borne out in practice does receive some affirmation in the data. Our time frame also encompasses a period of major recession that might also be expected to provide a more propitious milieu for detecting displacement effects in examining groups and sectors inherently more susceptible to minimum wage hikes. By focusing on this recent period, we are able to evaluate the possibility that recent effects of minimum wages are larger than the typical small elasticities reported in the literature.

Ours is not the first treatment to examine minimum wages over a period in which minimum wages increases have been more pronounced and/or the economy has deteriorated sharply. But studies that focus on recessionary differences are limited, and mixed results have been obtained from studies that focus on the "Great Recession." One important goal of our study has been to evaluate, using three data sets, a wide range of available evidence for the notion that minimum wages may have had particularly harmful effects during this most recent period.

After examining the robustness of estimated minimum-wage effects across different methods used for handling geographically-disparate trends during the period, our findings should be interpreted as producing only limited evidence of a negative minimum wage impact on employment. In sum, our only robust evidence for negative employment elasticities is from the CPS, and suggests that negative minimum-wage effects primarily occurred for teenagers during the actual recessionary years (2008-2010). The estimated elasticities were not substantially larger than those often reported in the literature (of around -0.2). For their part, results for the specific sector and sub-sectors examined here using either the QCEW or the CPS do not point to robust evidence of

either negative (or positive) impacts of minimum wage increases on employment. And, when compared to the CPS, our final dataset – the ACS – does not provide statistically significant evidence in support of a negative impact on employment for teenagers. On the other hand, even when insignificant, the results from the CPS are often insufficiently precise to allow us to say anything conclusively about employment effects. In this regard, we find the more-precise findings from the QCEW as the most informative, providing results that are at most supportive of a very small negative employment elasticity and one that does not appear to have worsened during the recessionary years.

The bottom line of this investigation is that material increases in the minimum wage – even when implemented during adverse market conditions – do not appear to have particularly strong effects in reducing employment among low-wage groups. Ultimately, it is difficult to resist the conclusion that minimum wage workers are concentrated in sectors of the economy for which the labor-demand response to forced wage increases is minimal.

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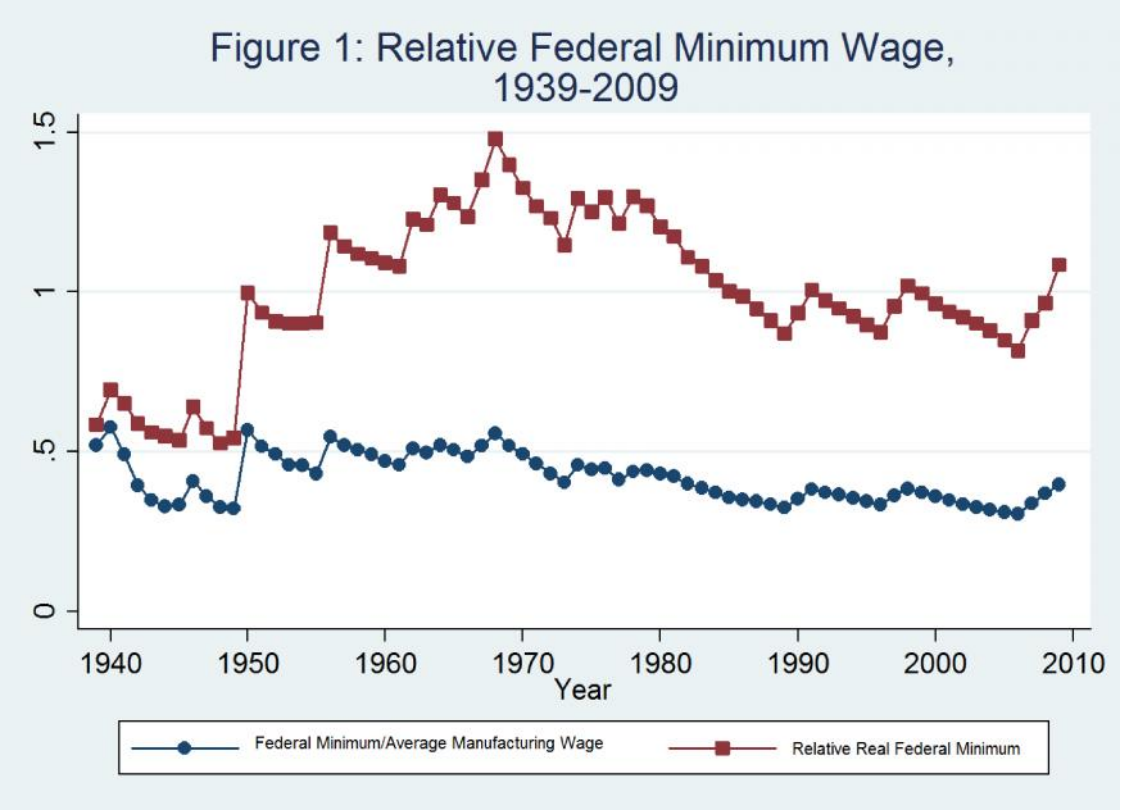
Figure 1: Relative Federal Minimum Wage, 1939-2009
0
.5
1
1.5
1940
1950
1960
1970
1980
1990
2000
2010
Year
Federal Minimum/Average Manufacturing Wage
Relative Real Federal Minimum

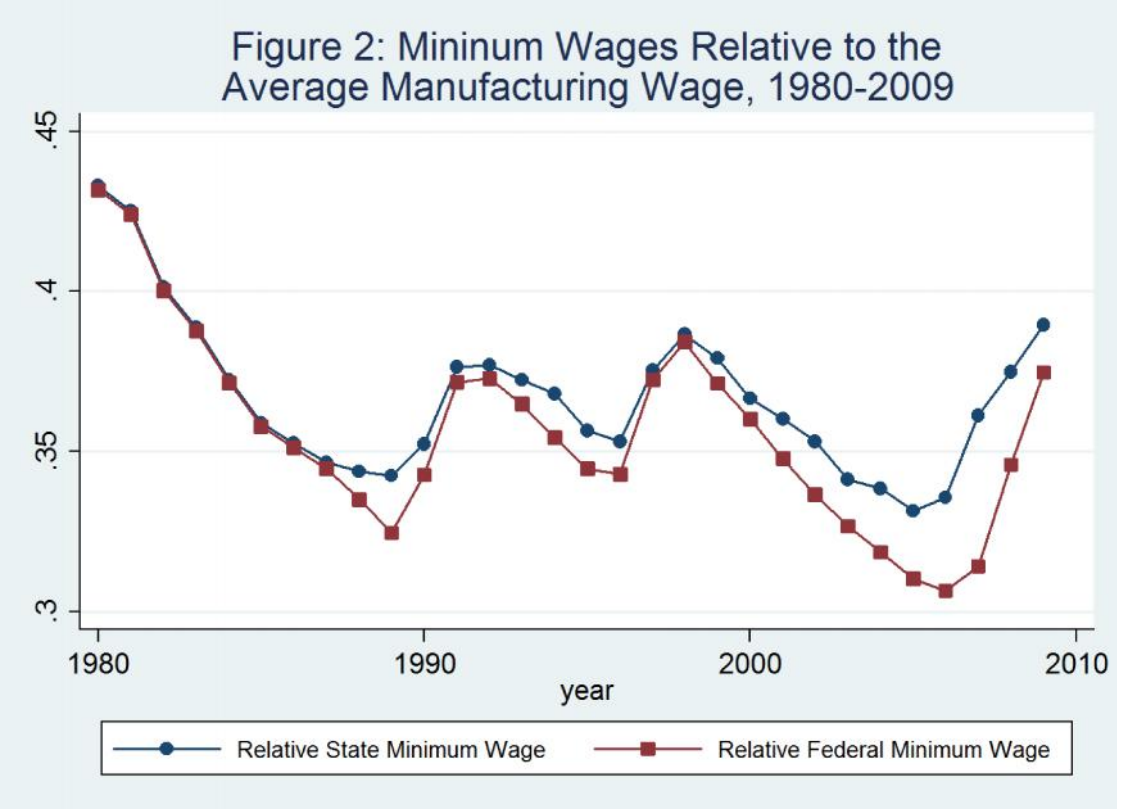
Figure 2: Mininum Wages Relative to the Average Manufacturing Wage, 1980-2009
.3
.35
.4
.45
1980
1990
2000
2010
year
Relative State Minimum Wage
Relative Federal Minimum Wage

Table 1	**Means and Standard Deviations, Quarterly Census of Employment and Wages 2005-2009**	
	Mean	**Standard Deviation**
Food Services and Drinking Places		
Employment	2,302	3,199
Weekly Earnings	224	46
Limited-Service Restaurants		
Employment	1,023	1,360
Weekly Earnings	211	35
Full-Service Restaurants		
Employment	1,157	1,660
Weekly Earnings	228	59
Independent Variables		
Minimum Wage	6.45	0.81
Population	76,067	96,906
Total Private Employment	24,368	33,635
Private Weekly Earnings	629	129
Unemployment Rate (All Participants)	0.060	0.026
Enrollment Rate (State-Level, 16-24 year olds)	0.51	0.08
Notes: These are per-county averages of the variables over the 2005-2009 sample period. The averages for the independent variables are for the sample of counties with complete information for *food services and drinking places* employment. All hourly earnings are expressed in May 2010 dollars, adjusting for inflation using the CPI.		

Table 2	**Means and Standard Deviations, Current Population Survey 2005-2009**	
	Mean	**Standard Deviation**
Food Services and Drinking Places		
Employment	164,030	181,132
Hourly Earnings	$ 8.96	2.36
Teenagers		
Employment	110,643	111,440
Hourly Earnings	8.29	1.77
Independent Variables		
Minimum Wage	$6.51	0.97
Population	5,760,757	6,452,120
Total Adult Employment	2,811,635	3,061,593
Teen Population	330,501	374,835
Prime-Age Male Wage	$22.42	3.76
Notes: These are per-state averages of the variables over the 2005-2010 sample period. The averages for the independent variables are for the sample of state/month observations included in the *food services and drinking places* regressions. All hourly earnings are expressed in May 2010 dollars, adjusting for inflation using the CPI.		

Table 3	**Means and Standard Deviations, American Community Survey 2006-2009**	
	Mean	**Standard Deviation**
Food Services and Drinking Places		
Employment	159,615	178,256
Teenagers		
Employment	114,667	113,871
Independent Variables		
Minimum Wage	$6.71	0.83
Population	5,861,520	6,560,818
Total Adult Employment	2,799,421	3,064,235
Teen Population	344,334	392,554
Prime-Age Male Wage	$22.44	3.02
Notes: These are per-state averages of the variables over the 2005-2010 sample period. All hourly earnings are expressed in May 2010 dollars, adjusting for inflation using the CPI. The prime-age male wage is taken from the CPS data.		

Table 4 **Estimates of Employment and Earnings Equations for the Restaurant-and-Bar Sector, QCEW 2005-2009**

	Employment			*Earnings*		
Independent Variables	Basic	County Trend	Border County	Basic	County Trend	Border County
Minimum Wage	-0.027 (0.021)	-0.019 (0.016)	-0.047 (0.032)	0.126*** (0.016)	0.127*** (0.016)	0.107*** (0.027)
Population	0.390*** (0.081)	0.023 (0.108)	0.275 (0.173)	-0.169*** (0.055)	0.088 (0.096)	-0.311** (0.098)
Total Employment	0.769*** (0.065)	0.955*** (0.106)	0.736*** (0.100)	0.209*** (0.031)	0.280*** (0.033)	0.187*** (0.043)
Total Average Weekly Earnings	-0.272*** (0.037)	-0.165*** (0.022)	-0.218*** (0.035)	0.119*** (0.022)	0.082*** (0.018)	0.103*** (0.031)
N	36,480	36,480	17,960	36,480	36,480	17,960

Notes: All dependent and independent variables are in logarithmic form. Standard errors are in parentheses. See text for details on the estimation procedure for coefficients and clustered standard errors for each specification. All regressions include fixed-effects for county and quarter-year, and are weighted by the average county population over the 2005-2009 period.

***,**,* denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

Table 5	**Estimates of Employment and Earnings Equations by Restaurant Subsector, QCEW 2005-2009**					
Full-Service Restaurants	*Employment*			*Earnings*		
	Basic	County Trend	Border County	Basic	County Trend	Border County
Minimum Wage	-0.044 (0.029)	-0.047** (0.019)	0.006 (0.049)	0.124*** (0.020)	0.112*** (0.016)	0.111*** (0.035)
N	34,820	34,820	17,760	34,820	34,820	17,760
Limited-Service Restaurants	*Employment*			*Earnings*		
	Basic	County Trend	Border County	Basic	County Trend	Border County
Minimum Wage	-0.009 (0.020)	-0.010 (0.021)	-0.013 (0.037)	0.145*** (0.016)	0.146*** (0.020)	0.145*** (0.035)
N	39,020	39,020	19,720	39,020	39,020	19,720

Notes: See Notes to Table 4. Only minimum-wage coefficient estimates are reported, although each regression contains all covariates and fixed effects included in the Table 4 specifications.

***,**,* denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

Table 6	**Alternative Regression Estimates for the Restaurant-and-Bar Sector, QCEW 2005-2009**						
	Employment				*Earnings*		
Specifications Using A Moving Average of the Minimum Wage in Place of the Contemporaneous Minimum Wage							
	Basic	County Trends	Border Counties		Basic	County Trends	Border Counties
Minimum-Wage Moving Average	-0.043 (0.031)	-0.030 (0.030)	-0.093** (0.040)		0.125*** (0.024)	0.147*** (0.025)	0.078** (0.031)
Interaction With Post-2007 Dummy							
Minimum Wage	-0.029 (0.020)	-0.018 (0.017)	-0.045 (0.032)		0.127*** (0.016)	0.123*** (0.016)	0.106*** (0.027)
Minimum Wage*Dummy Interaction	-0.054** (0.026)	-0.004 (0.017)	-0.048 (0.048)		0.023 (0.019)	0.033* (0.018)	0.017 (0.040)
p-value: no minimum wage effect for 2008-2009	0.01	0.16	0.09		0.00	0.00	0.01
Notes: See Notes to Table 5. The upper panel uses an equally weighted 5-quarter moving average of the current minimum wage with 4 lags of the minimum wage. ***,**,* denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively							

Table 7	**Estimates of Employment and Earnings Equations for Food Services and Drinking Places, CPS 2005-2010**					
	Employment			*Earnings*		
Independent Variables	Basic	State Trends	Border State	Basic	State Trends	Border State
Minimum Wage	-0.086 (0.079)	-0.040 (0.082)	-0.130** (0.063)	0.101 (0.096)	0.272** (0.106)	0.085* (0.059)
Population	0.656 (0.407)	0.719 (0.537)	1.012** (0.440)	0.613* (0.349)	1.548*** (0.562)	0.452 (0.342)
Total Adult Employment	-0.176 (0.152)	-0.212 (0.155)	-0.176 (0.146)	0.039 (0.136)	0.014 (0.161)	0.053 (0.120)
Average Prime-Age-Male Wage	-0.057* (0.030)	-0.051* (0.029)	-0.032 (0.028)	0.125** (0.054)	0.124** (0.053)	0.116*** (0.044)
N	3,665	3,665	15,354	3,665	3,665	15,354

Notes: All dependent and independent variables are in logarithmic form. Standard errors are in parentheses. See text for details on the estimation procedure for coefficients and clustered standard errors for each specification. All regressions include fixed-effects for state and month-year, and are weighted by the average state population over the 2005-2010 period.

***,**,* denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

Table 8 **Estimates of Employment and Earnings Equations for Teenagers, CPS 2005-2010**

Independent Variables	*Employment*			*Earnings*		
	Basic	State Trends	Border State	Basic	State Trends	Border State
Minimum Wage	-0.008 (0.124)	-0.178* (0.103)	0.001 (0.090)	0.088** (0.044)	0.078 (0.066)	0.142*** (0.038)
Teen Population	0.895*** (0.069)	0.965*** (0.060)	0.908*** (0.041)	-0.002 (0.034)	0.006 (0.034)	-0.002 (0.027)
Total Adult Employment	0.030 (0.276)	0.092 (0.173)	-0.150 (0.163)	-0.053 (0.064)	-0.050 (0.100)	-0.113 (0.084)
Average Prime-Age Male Wage	0.009 (0.029)	0.027 (0.023)	0.001 (0.022)	-0.057** (0.026)	-0.064** (0.027)	-0.064** (0.019)
N	3,634	3,634	15,138	3,634	3,634	15,138

Notes: See Notes to Table 7.

***,**,* denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

Table 9	**Alternative Regression Estimates for Food Service and Drinking Places, CPS 2005-2010**						
	Employment				*Earnings*		
Specifications Using A Moving Average of the Minimum Wage in Place of the Contemporaneous Minimum Wage							
	Basic	State Trends	Border State		Basic	State Trends	Border State
Minimum-Wage Moving Average	-0.013 (0.111)	0.064 (0.133)	-0.139* (0.074)		0.065 (0.107)	0.206 (0.132)	0.049 (0.062)
Interaction With Post-2007 Dummy							
Minimum Wage	-0.086 (0.079)	-0.048 (0.085)	-0.135** (0.063)		0.101 (0.092)	0.244* (0.122)	0.057 (0.058)
Minimum Wage*Dummy Interaction	-0.020 (0.100)	0.032 (0.116)	0.054 (0.110)		0.245** (0.094)	0.125 (0.155)	0.323** (0.136)
p-value: no minimum wage effect for 2008-2010	0.44	0.90	0.51		0.01	0.01	0.01

Notes: See Notes to Table 7 . The upper panel uses an equally weighted 13-month moving average of the current minimum wage with 12 lags of the minimum wage.

***,**,* denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

Table 10	**Alternative Regression Estimates for Teenage Workers, CPS 2005-2010**						
	Employment				*Earnings*		
Specifications Using A Moving Average of the Minimum Wage in Place of the Contemporaneous Minimum Wage							
	Basic	State Trends	Border State		Basic	State Trends	Border State
Minimum-Wage Moving Average	0.008 (0.163)	-0.166 (0.124)	0.020 (0.084)		0.098* (0.056)	0.085 (0.083)	0.173*** (0.036)
Interaction With Post-2007 Dummy							
Minimum Wage	-0.016 (0.115)	-0.152 (0.106)	0.011 (0.090)		0.092** (0.045)	0.034 (0.070)	0.129*** (0.039)
Minimum Wage*Dummy Interaction	-0.271** (0.134)	-0.118 (0.079)	-0.126 (0.098)		0.127* (0.074)	0.196* (0.099)	0.171*** (0.0560
p-value: no minimum wage effect for 2008-2010	0.03	0.02	0.36		0.03	0.03	0.00

Notes: See Notes to Tables 8 and 9.

***,**,* denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively

Table 11	**Estimates of Employment Equations, ACS 2006-2009**						
	Food Service and Drinking Places				*Teenagers*		
Independent Variables	Basic	State Trends	Border State		Basic	State Trends	Border State
Minimum Wage	0.022	0.115	0.028		0.039	-0.075	-0.005
	(0.057)	(0.108)	(0.054)		(0.113)	(0.083)	(0.042)
Population	1.342	0.868	1.899***		0.869	1.229***	1.028***
	(0.507)	(2.642)	(0.641)		(0.217)	(0.267)	(0.115)
Total Adult Employment	-0.249	0.567	-0.100		0.876	0.468	0.306
	(0.353)	(1.012)	(0.415)		(0.636)	(0.627)	(0.345)
Average Prime-Age Male Wage	-0.154	-0.141	-0.282***		-0.049	0.071	0.211**
	(0.110)	(0.157)	(0.082)		(0.145)	(0.186)	(0.107)
N	204	204	856		204	204	856

Notes: All dependent and independent variables are in logarithmic form, with the exception of the unemployment rate and the enrollment rate. Standard errors are in parentheses. See text for details on the estimation procedure for coefficients and clustered standard errors for each specification. All regressions include fixed-effects for state and month-year, and are weighted by the average state population over the 2006-2009 period. The population variable for the first three columns is total population, and is teen population for the last three columns.

***,**,* denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively