Institute for Research on Labor and Employment

Institute for Research on Labor and Employment Working Paper Series
(University of California, Berkeley)

Year 2008 *Paper iirwps-166-08*

Do Minimum Wages Really Reduce Teen Employment? Accounting for Heterogeneity and Selectivity in State Panel Data

Sylvia Allegretto
University of California, Berkeley

Arindrajit Dube
University of California, Berkeley

Michael Reich
University of California, Berkeley

Do Minimum Wages Really Reduce Teen Employment? Accounting for Heterogeneity and Selectivity in State Panel Data

Abstract

Traditional estimates of minimum wage effects on employment include controls for state unemployment rates and state and year fixed-effects. Using Current Population Survey data, we show that such estimates often are biased and that the estimates vary ith the source of identifying variation. Without sufficient controls for heterogeneous employment patterns that would occur without minimum wage policies, traditional estimates vary substantially both in sign and magnitude depending on time period and hence with the selectivity of states with minimum wage hikes. Estimates without sufficient controls also vary across demographic groups in a counterintuitive manner that suggests misspecification problems. To account for heterogeneous employment patterns and selectivity among states with minimum wages, we include controls for long-term growth differences among states by using a state-specific linear trend, and controls for heterogeneous responses to economic shocks by including Census division-specific time effects. In the 1990 to 2006 period, including these controls reduces the magnitude of the estimated employment elasticity from -0.168 (significant at the 1 percent level) to -0.024 (not significant). Although the division and state trend controls do not constitute a panacea, they do provide important tools to mitigate the bias that results from unobserved spatial heterogeneities in employment patterns that are correlated with the minimum wage. Since estimates in most previous national-level studies insufficiently address this issue, economists' estimates of minimum wage effects must be revised accordingly.

June 28, 2008

DO MINIMUM WAGES REALLY REDUCE TEEN EMPLOYMENT? ACCOUNTING FOR HETEROGENEITY AND SELECTIVITY IN STATE PANEL DATA

Sylvia Allegretto
IRLE
UC Berkeley, CA 94720
allegretto@berkeley.edu

Arindrajit Dube
IRLE
UC Berkeley, CA 94720
adube@berkeley.edu

Michael Reich
Department of Economics
and IRLE
UC Berkeley, CA 94720
mreich@econ.berkeley.edu

Institute for Research on Labor and Employment
University of California at Berkeley
Berkeley, CA 94720-5555

We are grateful to Lisa Bell for excellent research assistance and to Eric Freeman for helpful suggestions.

DO MINIMUM WAGES REALLY REDUCE TEEN EMPLOYMENT? ACCOUNTING FOR HETEROGENEITY AND SELECTIVITY IN STATE PANEL DATA

ABSTRACT

Traditional estimates of minimum wage effects on employment include controls for state unemployment rates and state and year fixed-effects. We show that such estimates fail to account for heterogeneous employment patterns and selectivity among states with minimum wages. As a result, the estimates are often biased and vary with the source of identifying variation. To correct these problems, we add controls for long-term growth differences among states and for heterogeneous economic shocks. Including these controls reduces the magnitude of the estimated employment elasticity in the 1990 to 2007 period from -0.153 (significant at the 1 percent level) to -0.024 (not significant).

1. Introduction

Despite a steady stream of studies, research on the effects of minimum wage policies on wages and employment continues to arrive at conflicting findings. One set of results—statistically significant disemployment effects with employment elasticities in a "consensus" range of -0.1 to -0.3—are associated with studies that focus on less-skilled groups such as teens, that use national-level household data (usually the Current Population Survey), and that include state and year fixed-effect controls to identify minimum wage effects. Another set of results—employment effects that are close to zero or even positive—are associated with studies that focus on low-wage sectors such as restaurants, that use employer-based data, and that use only local comparisons to identify minimum wage effects

The conflicting findings may arise from differences in the groups being examined and/or differences in the datasets that are used. However, recent evidence suggests other possibilities. Unobserved spatial heterogeneities in employment trends can generate biases toward negative employment elasticities in national minimum wage studies as well as overstate the precision of local studies.

In this study, we seek to address and resolve the conflicting findings by using CPS data on teens over the 1990 to 2007 period and by providing a detailed examination of heterogeneity and selectivity issues. More specifically, we consider whether the source of identifying variation in the minimum wage is coupled with sufficient controls for counterfactual employment growth. With the addition of these controls we are able to reconcile the different findings in the literature, identify the limitations of the previous studies and provide improved estimates.

Our central argument concerns the confounding effects of heterogeneous patterns in low-wage employment that are coupled with the selectivity of states experiencing minimum wage increases. The presence of heterogeneity is suggested by Figure 1 and Table 1, which show that employment rates for teens vary by Census division and differentially so over time. The differences over time are not captured simply by controls for business cycles, school enrollment rates, relative wages of teens, unskilled immigration or by the timing of federal minimum wage increases.[1]

To examine more systematically the importance of spatial and temporal heterogeneity, we begin with the canonical specification of minimum wage effects on teen earnings and employment with national CPS panel data and state and year fixed-effect variables. We then add two controls, separately and together: a) allowing for Census division-specific time effects, which eliminates the variation among the nine divisions and thereby controls for spatial heterogeneity in economic shocks; and b) including a state-specific linear trend that captures long-run growth differences among states. With these geographic controls, the estimates change substantially.

We find that adding division-specific time controls and state-specific time trends as controls reduces the estimated employment elasticity from -0.153 (significant at the 1 percent level) to -0.024 (not significant). We can rule out with 90 percent confidence that the employment elasticity is greater (in absolute value) than 0.147. Our results highlight the importance of estimates that control for underlying growth prospects, even at as coarse a level as Census divisions. These findings suggest that previous studies are compromised by

[1] For detailed analyses that arrives at these conclusions, see Aaronson et al. (2006) and Congressional Budget Office (2004).

insufficient controls for heterogeneity in employment patterns coupled with selectivity of states experiencing minimum wage hikes.

We also analyze the time-path of residual teenage employment around changes in minimum wage, using distributed lags covering 25 quarters around the minimum wage change. When we do not include division-specific time controls and state-specific time trends, the nine quarters *prior* to the actual policy change are all associated with unusually low (and falling) teenage employment, which provides strong evidence regarding the selectivity of states and the timing of minimum wage increases. But when we include these controls, there is no visible reduction in employment following the minimum wage increase.

Although the range of elasticities generated by studies in the literature may seem narrow, they contain important implications for the net benefits of a minimum wage policy for low-wage workers. Whether the net benefits are positive or negative depends upon whether the absolute value of the sum of the estimated employment and hours elasticities is greater than or less than the estimated wage elasticity. The estimates from extant national CPS-based studies (e.g., Neumark and Wascher, 2007) often imply negative net benefits for the low-wage workforce; our estimates reverse this conclusion.

2. Related Literature

For the most part, minimum wage studies using national CPS panel data with state and year fixed-effects find modest but statistically significant negative employment effects on teens, with elasticities that range from -0.1 to -0.3. Sabia (2006) and Neumark and Wascher (2007) are two recent papers in this vein. Sabia (2006) uses grouped CPS data for 1979 to 2004 to study the retail industry only. For teens in retail, Sabia's main specification

includes controls for teen shares in the population and fixed state effects; Sabia also adds year effects in a second specification (Sabia 2006, Table 3, columns 4 and 6). Sabia finds significant disemployment elasticities of -0.217 when year effects are excluded and -0.298 when they are included. Sabia does not, however, control for state-specific time trends. As we shall show below, the robustness of his findings consequently remains open.[2]

Neumark and Wascher (2007) use pooled time-series cross-section individual CPS data for 1997 to 2005. They estimate a negative employment elasticity of -0.136 among teens, significant at the 10 percent level. Neumark and Wascher motivate their selection of the period since 1997 by arguing that welfare reform and expansions of the EITC may have changed the dynamics of the low-wage labor market. A better approach would include the entire period of state minimum wage variation over a longer period with a consistent methodology and test whether minimum wage effects differ in the more recent period. As we report below, their results are driven in large part by the source of the identifying variation in the time period they selected.

As we already mentioned, minimum wage studies that use local restaurant employment data generally do not find disemployment effects.[3] A recent example is the Dube, Naidu and Reich (2007) before-after study of the effects of a citywide San Francisco minimum wage introduced in 2004 and phased in for small firms. To identify the policy's effect, the study exploits differentials in the policy by firm size as well as geographical location, permitting more controls than in previous such studies. The authors do not find a

[2] Orrenius and Zavodny (2008) find similar employment effects among teens, but only when three business cycle controls are included.
[3] Card and Krueger (1994, 2000); Dube, Naidu and Reich (2007). The exception is Neumark and Wascher (2000).

negative employment effect. Moreover, their data contain less measurement error than in previous local studies, so they can reject negative effects with considerably greater confidence than those studies. They cannot, however, address concerns about lags in disemployment effects or common spatial shocks that might lead to overstatement of the precision of their estimates.

Such issues are addressed by Dube, Lester and Reich (2007), who compare all the contiguous counties in the U.S. that lie on state borders. This method uses county-level administrative data on restaurant employment and generalizes the local studies with national data. As previously mentioned, these authors show that previous national minimum wage studies lack adequate controls for spatial heterogeneity in employment growth.[4] Without such controls, Dube, Lester and Reich find significant disemployment effects, within the standard -0.1 to -0.3 range of estimates. In their localized analysis, the economic and labor market conditions within the local area are sufficiently homogeneous to control for spatial heterogeneities in employment growth that are correlated with the minimum wage. Once they add such controls, Dube, Lester and Reich find no significant disemployment effects. An important question, which we consider below, is whether the Dube, Lester and Reich results are replicable beyond the restaurant industry and the QCEW data.

To summarize, a fundamental issue in the minimum wage literature concerns how estimates based upon specifications national panel data and state and year fixed effects compare to estimates from specifications that control for spatial heterogeneity and

[4]In a study of the effect of teen population shares on teen unemployment rates, Foote (2007) also finds that controlling for heterogeneous spatial trends across states generates results quite different from those using national panel data with state fixed effects.

selectivity. To address this question, we use the CPS dataset of the previous literature and we incorporate additional spatial and time controls into the traditional specifications.

3. Data

We construct an individual-level repeated cross-section sample from the CPS Outgoing Rotation Groups for the years 1990 to 2007.[5] The CPS data are merged with data that capture aggregate demand and supply variation—monthly state unemployment rates and population shares for demographically conditioned regressions. Additionally, each observation is merged with a quarterly minimum wage variable—whichever is higher of the federal or state minimum.

Table 2 provides descriptive statistics for the sample of teens aged 16 to 19. Non-Hispanic whites account for nearly 71 percent of the sample, while blacks and Hispanics each account for nearly 15 percent. Hourly pay over the sample period averages $6.44. Although male teens are paid more than female teens—$6.72 versus $6.16—pay differentials by race/ethnicity are considerably smaller. Employment patterns do not correspond to these wage patterns. Over the sample period, 41 percent of all teens 16-19 were employed, with identical percentages for males and females. Black teens had the lowest employment rates—25 percent, followed by Hispanics—34 percent. Employed teens worked an average of 24.9 hours per week, with somewhat higher average hours among males, blacks and Hispanics. Finally, state minimum wages averaged $1.11 above the federal minimum wage.

[5] We use a start year of 1990 to make the clearest comparison with Dube, Lester and Reich's results with QCEW restaurant data, which are available on a consistent basis beginning in 1990. We have also included earlier years in our CPS sample; the results are not affected. See Section 5.3 below.

4. Estimation Strategy

Our focus is to estimate the effect of minimum wage increases upon wages and employment of teenagers. The dependent variables, *y,* are the natural log of hourly earnings, or a dichotomous employment measure that takes on the value one if the person is working, or the natural log of usual hours of work. The baseline fixed-effects specification is then:

$$y_{ist} = \beta MW_{st} + X_{ist}\Gamma + \lambda \cdot unemp_{st} + \phi_s + \tau_t + \varepsilon_{ist} \qquad (1)$$

where *MW* refers to the log of the minimum wage, *i*, *s*, and *t* denote, respectively, individual, state and time indexes, *X* is a vector of individual characteristics, *unemp* is the quarterly (non-seasonally adjusted) unemployment rate in state *s* at time *t,* ϕ_s refers to the state fixed effect and τ_t represents time dummies incremented in quarters.[6] In this canonical specification, including state and time dummies as well as the overall unemployment rate is thought to sufficiently control for local labor market conditions facing teenage workers.

There is, however, growing evidence (Dube, Lester and Reich 2007) that these variables do not fully capture heterogeneity in underlying employment patterns in low-wage employment. To account for this heterogeneity, our second specification allows time effects to vary by Census divisions. Including division-specific time effects (τ_{dt}) eliminates the between-division variation and hence better controls for spatial heterogeneity in differential employment patterns, including region-specific economic shocks:

$$y_{ist} = \beta MW_{st} + X_{ist}\Gamma + \lambda \cdot unemp_{st} + \phi_s + \tau_{dt} + \varepsilon_{ist} \qquad (2)$$

[6] The individual characteristics include 2 gender categories, 4 race/ethnicity categories, 12 education categories and 4 marital status categories. These are the same categories used in Neumark and Wascher (2007).

A state-specific linear trend variable provides a second means of controlling for heterogeneity in the underlying (long term) growth prospects of low-wage employment. Our third specification includes these controls:

$$y_{ist} = \beta MW_{st} + X_{ist}\Gamma + \lambda \cdot unemp_{st} + \phi_s + \psi_s \cdot t + \tau_t + \varepsilon_{ist} \qquad (3)$$

where ψ_s denotes the time trend for each state.

Finally, we add both the division-specific time effect and the state-specific time trend controls for our fourth specification:

$$y_{ist} = \beta MW_{st} + X_{ist}\Gamma + \lambda \cdot unemp_{st} + \phi_s + \psi_s \cdot t + \tau_{dt} + \varepsilon_{ist} \qquad (4)$$

The resulting estimates are less likely to be contaminated with unobservable long term trends and region-specific economic shocks in this final specification.

We estimate these four specifications on all teens 16-19 as well as on sub-samples disaggregated by gender, and race/ethnicity (white (not Hispanic), black separately, Hispanic separately). We report standard errors clustered at the state level to account for the lack of independence among observations within a state at any given time, which might be caused by correlation in employment rates within states over time and across individuals.

5. Main Results

We discuss first our estimated wage, employment and hours effects for all 16-19 year-olds with each of our four specifications. We then check our results by examining the time path of teenage employment around the time of minimum wage changes and by disaggregating our results by different time periods.

5.1 Wage, employment and hours effects for all teens

The estimated wage effects establish the presence of a treatment—that increases in the minimum wage led to increased wages for the teenage population, conditional on employment. Table 3, Panel A presents the estimated effects on wages for all teens, for male teens, and for female teens. The coefficient, which is also the wage elasticity, is positive and significant at the 1 percent level in all the specifications. But the magnitudes vary among the specifications. In Specification 1, the fixed-effects model, the treatment coefficient is 0.120 for all teens. Adding just the division controls (Specification 2) increases the magnitude of the treatment coefficient for all teens to 0.154, similar to that in Specification 1. Adding the state-specific time trends, without division controls (Specification 3) further increases the magnitude of the wage elasticity to 0.205. When both state- and division-specific time trends are included (Specification 4), the treatment effect for all teens is 0.158 and the effect remains highly significant.

These results indicate that the treatment effect of minimum wages remains significant when controls for heterogeneous spatial trends are included. Moreover, the magnitude of the estimated treatment effect is consistent with CPS earnings for teens. In a separate calculation, we found that 30.7 percent of employed teens 16-19 were paid within ten percent of the relevant state or federal minimum wage. Since not all of these teens were earning exactly the minimum wage, the estimated treatment elasticity of 0.158 is consistent with the distribution of pay at or near the minimum wage.

Figure 2, Panel A displays time paths of the wage effects of minimum wage increases.[7] The left-hand column displays results for our Specification 1, while the right-hand column present results for Specification 4, which includes both state-specific time trends and division-specific time effects. Both wage graphs show a clear increase just at the time of the change. However, including the additional controls in the full specification (Specification 4) generates a cleaner "treatment" effect, which we interpret as reinforcing the validity of including additional controls.

We turn next to the employment effects, reported in Table 3, Panel B. Specification 1 shows a significant negative employment coefficient of -0.063. The corresponding employment elasticity is -0.153, which is consistent with the literature that uses the canonical fixed-effects model.[8] In Specification 2, however, allowing for division-specific time effects reduces the elasticity to -0.105 and renders it insignificant. As Specification 3 shows, the addition of a state-specific time trend to the fixed effects model also lessens the effect of minimum wages on employment. Here the elasticity is -0.065 and it is not significant. Finally, in Specification 4, the employment elasticity is -0.024 and remains insignificant. In other words, allowing for variation in employment trends over the 1990 to 2007 period, we obtain minimum wage effects on employment that are indistinguishable from zero. Moreover, estimates using Specification 4 rule out elasticities larger in magnitude than .147 with 90 percent confidence.[9]

[7] To construct the time paths we estimate a distributed lag model with leads and lags in minimum wage covering a 25 quarter window, starting at 8 quarters before the minimum wage change and continuing to 16 quarters after the change.

[8] The elasticity is obtained by dividing the coefficient by the employment-to-population rate of the group in question.

[9] Confidence intervals are reported in Table 9 below.

These results indicate that estimates of minimum wage employment effects using the standard fixed-effects model of Specification 1 are contaminated by heterogeneous employment patterns across states. Controlling only for within-division variation substantially reduces the estimated elasticity substantially. Allowing for long-term differential state trends makes the employment estimates indistinguishable from zero.[10]

The time paths for employment (Figure 2, Panel B) provide strong evidence against the canonical model without controls for heterogeneity across states (i.e., Specification 1). Specification 1 shows negative employment effects throughout the 25 quarter window, including prior to the minimum wage increase. This result provides strong evidence that minimum wage changes occur at times of unusually low teen employment growth. Consistent with this interpretation, Specification 4 (with both added controls) shows relatively more stable coefficients (closer to zero) prior to the minimum wage increase, no clear effect on employment in the subsequent 8 quarters and then a small positive employment effect in quarters 8 to 16 after the minimum wage increase. Overall, the time paths provide further evidence that, in the period under consideration, failure to control for heterogeneity in employment patterns leads to a negative bias in the estimated employment response from minimum wage changes.

Table 3, Panel C provides estimates of the effects of the minimum wage on weekly hours worked. In Specification 1, the elasticity on weekly hours is -0.096 and is significant. It remains significant in Specification 2 but not in Specification 3. In Specification 4, the elasticity is -0.087 and is once again significant. As the time paths for hours in Figure 2,

[10] We discuss in Section 6 below our earnings and employment estimates for gender and race/ethnicity groups.

Panel C indicate, the hours effect with Specification 4 becomes indistinguishable from zero within four quarters of the minimum wage increase and becomes positive after 12 quarters.

This estimated contemporary effect on hours can be put together with the significant effects on wages (and the insignificant effects on employment) to get a combined result as follows. Before a minimum wage increase, the weekly wage averages $160.36 and the average hourly wage is $6.44. Since the estimated minimum wage elasticity is 0.158, the hourly wage increase (6.44x0.158) is $1.02, raising the average to $7.46. Since the weekly hours worked before a minimum wage averages 24.9 hours and the hours elasticity is -0.087, the decrease in average hours is 24.9x0.087, or 2.17 hours per week, and average hours now are 22.73 hours. The weekly wage thus changes from $160.36 to $169.54, or an increase of 5.7 percent, with a decrease in hours of 8.7 percent. Of course, as we have just noted, the dynamic evidence suggests a much smaller hours effect.

5.2 Differences by time period

Our evidence thus far suggests that selectivity of the states with minimum wage increases and heterogeneity of employment patterns across states produce biased estimates for specifications that do not account for such heterogeneity. In this section, we provide further evidence on this issue by looking at different sub-periods in the data—especially the 1990-1998 period versus the 1998-2006 period. The two periods constitute different sources of identifying variation. While most of the changes in the first period occurred through federal changes, in the entire latter period minimum wage increases occurred only at the state level. To the extent that states enacting local minimum wage increases are

systematically different from those that do not, the two periods should provide different estimates when controls for such heterogeneity are not included.

Table 4 reports our estimates for our four specifications for our sample period of 1990-2007 (top row), for 1990-2006 (second row), for a longer sample period, 1987-2006 (third row), and then for various sub-periods: 1990-1998, 1997-2005 (the Neumark and Wascher 2007 period) and 1998-2006.

In 2007 more states implemented minimum wage increases than ever before and the federal minimum wage was increased for the first time in more than a decade. We therefore begin with an assessment of the minimum wage increases in 2007, simply by comparing the first and second row of Table 4. The estimates are extremely close, indicating that the minimum wages increases in 2007 did not lead to disemployment effects.

In the next row, we report results from 1987 to 2006; state minimum wages were first implemented in 1987 in three states (Alaska, Connecticut and Massachusetts). By 1989 fifteen states and the District of Columbia had increased their minimum wage standards above the federal level. Comparing the second and third rows in Table 4 shows that starting the sample in 1987 provides estimates that essentially are identical to starting the sample in 1990.

We next examine the effect of splitting our sample into two time periods of equal duration. For the earlier period (1990-1998), the employment effect is not significant in any of the specifications. For the later period (1998-2006), specifications without both state trends and division-specific time effects (i.e., Specifications 1-3) suggest disemployment

effects. However, the estimates using Specification 4 (with the full set of controls) suggest small and statistically insignificant employment effects.

As we mentioned above, the source of identifying variation is quite different in the two periods. Most minimum wage changes in the earlier period were federal increases, while all of the increases in the later period occurred at the state level. Moreover, the states enacting these policies were not randomly distributed among all states, as they were usually coastal states and slower-growing states.[11]

Equally important, as Figure 1 and Table 1 indicate, the teenage employment rate experienced a large negative shock with the 2001 recession. In every one of the nine Census divisions, teen employment-population ratios showed no long-run trend in the period from 1990 to 1998. The ratios declined subsequently, beginning in 1998 in some divisions and in 2000 in others. Moreover, the magnitude of the decline differed among the divisions and teen employment rates did not rebound in most of the U.S. in the subsequent recovery.

Putting these two facts together suggests that focusing only on the latter period, and doing so *without* sufficient controls for heterogeneity in employment trends, is likely to produce very misleading results. Consistent with this evidence, we find that the results from our fixed-effects specification (Specification 1) vary considerably between the two periods, while the specification with the richest controls (Specification 4) is quite stable, with an employment elasticity close to zero. Moreover, as we have already seen, for the full 1990-2007 period, including either state trends (Specification 2) or focusing on within-division

[11] This pattern began to change by 2006, when a larger set of states implemented minimum wage laws. Dube, Lester and Reich (2007) provide a detailed discussion of the pattern of state minimum wage increases from 1987 to 2007.

variation (Specification 3) produced a small elasticity. In the 1998-2007 period, as Figure 1 suggests, spatial and temporal heterogeneity were much more complex.

The lack of sufficient controls for this heterogeneity consequently is particularly severe if we limit our attention to the 1997 to 2005 time period, as do Neumark and Wascher (2007). As Table 4 shows, even a one year change in the start and ending dates, to 1998 to 2006, generates somewhat different results.

6. Minimum wage effects by gender, race and ethnicity

Figure 3 displays employment rates among teens by gender, race and ethnicity over the period 1990-2007. Three main patterns stand out, each with implications for the effects of minimum wages on specific groups. First, male teen employment rates lost ground relative to female teen employment rates in every race and ethnicity group, indicating that minimum wage policies may have had less negative effects among female teens than among male teens. Second, employment rates are lower among minorities than among whites; since whites, blacks and Hispanics are not equally distributed across states and Census divisions, estimates of minimum wage effects for each group may be affected by inclusion of controls for spatial heterogeneity. Third, minority employment rates are much more affected than are whites by the business cycle, in both prosperity and recession, indicating that spatial heterogeneity of business cycles coupled with selectivity of states with minimum wage increases may be important in estimating minimum wage effects for these minority groups.

Other factors may also be at play. A standard explanation of the lower employment rates among minority teens suggests that they are less skilled and experienced than other teens. Minimum wage increases will then have a greater impact on such groups, especially

insofar as employers adjust to higher minimum wages by substituting toward higher-skilled groups. The prediction is that minority teens will experience higher earnings effects and greater disemployment effects, relative to all teens. An alternative view suggests that barriers to mobility are greater among minorities than among teens as a whole. Higher pay then increases the returns to worker search and overcomes existing barriers to employment that are not based on skill and experience differentials (Raphael and Stoll 2002).

To investigate these issues, we estimate our four different specifications on specific gender and race/ethnicity groups. We begin by discussing minimum wage effects for male and female teens separately. We then examine effects by race/ethnicity.

6.1 Earnings and employment effects by gender

Recent studies of teen wage and employment patterns report that differences between male and female teens of similar educational enrollment status have declined in recent decades and the remaining differences are small (Congressional Budget Office 2004). Figure 3 and the descriptive sample statistics in Table 2 present a similar picture. Average wages in the sample are $6.72 for male teens and $6.16 for female teens and the average employment to population ratio is identical for both. These patterns suggest that minimum wage changes should have similar effects on male and female teens.

Table 3, Panel A reports our estimated wage and employment elasticities by gender. In Specification 1 the wage elasticity is 0.096 for male teens and 0.139 for female teens, indicating a 40 percent larger treatment effect among female teens. In Specification 2, with division-specific time controls included, the male teen wage elasticity increases (to 0.142) and it is much closer to the female wage elasticity (0.154). In Specification 3, with just the

state-specific trend control included, the estimated male teen wage elasticity increases again (to 0.177) and the female teen wage elasticity is 0.231, implying a greater effect on female teens—about 30 percent greater. Finally, for Specification 4, with both controls included, the estimated male teen wage elasticity is 0.148 and the female teen wage elasticity is 0.149.

In summary, without controls (Specification 1), treatment effects are present for both male and female teens, but the estimated treatment effect varies by gender. When both sets of controls are added to account for spatial heterogeneity (Specification 4), the treatment effect on earnings remains substantial and significant for both male and female teens. However, with the full set of controls the gender difference disappears.

We turn next to gender patterns in the estimated employment elasticities that are shown in Table 3, Panel B. In Specification 1, the employment effects for all teens are very similar to those for male and female teens separately.[12] Specification 2 and 3 finds significant disemployment effects among females but not among males. But while Specification 1 estimates significant disemployment effects for both male and female teens, Specification 4 shows no significant employment effects for either male or female teens. These results reinforce our previous finding that controlling for heterogeneity in employment patterns is crucial in estimating minimum wage effects.

6.2 Earnings and employment effects by race/ethnicity

Labor market outcomes for black and Hispanic teens continue to be inferior to those for white, non-Hispanic teens. As Table 2 indicates, during our sample period the employment rate averages 0.25 for black teens and 0.34 for Hispanic teens, compared to

[12] As we indicate below, however, the gender patterns are not similar for specific race and ethnicity groups.

0.46 for non-Hispanic white teens.[13] Moreover, as Figure 3 shows, the employment rates of black and Hispanic teens dropped sharply since the 2001 recession. As we mentioned previously, the poorer outcomes for minority teens may reflect their more limited education or experience, relative to non-Hispanic white teens. Moreover, if minimum wage effects lead to substitution toward more educated and experienced workers, then minimum wage policies may have more harmful effects on the employment on disadvantaged groups.

Structural studies of poorer labor market outcomes for black and Hispanic teens point to a different explanation: the spatial mismatch between urban employment and minority population distributions, as well as other disadvantages that these groups face (Raphael 1998, Raphael and Stoll 2002). In this approach, if minimum wage increases make it more worthwhile for disadvantaged teens to travel greater distances to find employment, then minimum wage increases may create relatively more beneficial employment effects for such groups. The research literature thus does not clearly predict how black teens will be affected by the policies.

Table 5, Panel A reports our estimated treatment effects on wages for separate race/ethnicity groups. For the non-Hispanic white group, the wage elasticities are substantial and significant under all four specifications. These elasticities (and their significance levels) are similar to those in Table 3 for all teens, which is not surprising since non-Hispanic whites accounts for 71 percent of the total teen sample. In summary, whether or not we include controls for spatial heterogeneity, we find a substantial and significant treatment effect for non-Hispanic whites.

[13] Wage rates do not show such a disparity, however. Black teens have the same wage as non-Hispanic white teens ($6.40), while Hispanic teen wages are higher ($6.67).

The treatment effect on wages varies much more among black and Hispanic teens. Among black teens, the wage effect in Specification 1 is positive but not significant. Controlling for division-specific trends (Specification 2), the treatment effect among black teens remains positive but insignificant. Controlling for state-specific time trends (Specification 3), the treatment effect for black teens becomes large and significant and similar in magnitude to that for all teens. The treatment effect on black teens in Specification 4 is significant and it is higher than the effect among all teens (0.227 versus 0.158) and it is more than twice as high in Specification 4 as in Specification 1 (0.227 versus 0.088). Among Hispanic teens, the magnitude and the statistical significance of the wage elasticity fall considerably from Specification 1 to Specification 4 (from 0.101 and 5 percent significance to -0.042 and insignificant). In summary, we find a substantial wage effect for blacks but not for Hispanics.[14]

We turn next to the employment elasticities by race/ethnicity, which are reported in Table 5, Panel B. Among non-Hispanic white teens, the employment effect is negative and significant under Specifications 1 and 2, but not significant under Specification 3. In Specification 4, the estimated effect is also not significant.

The results for black and Hispanic teens reinforce the need for caution in interpreting estimates for disaggregated groups. For both groups, the standard errors are two to three times higher than among non-Hispanic whites. The employment elasticity changes from negative but not significant in Specification 1 to *large and positive* (0.450) and significant at

[14] The increased standard errors among these estimates suggest that caution is called for in interpreting the results.

the one percent level in Specification 4. Among all Hispanic teens, the employment effect is not significant in any of the four specifications.

7. Comparisons with restaurant studies

We examine next whether increases in the minimum wage have similar effects across studies that incorporate analogous controls for spatial heterogeneity. The fixed-effects models without and with controls for division-specific time controls and state-specific time trends in our study are similar to those used in Dube, Lester and Reich (2007), but as already mentioned the data and the group that is studied are different. Table 6 provides results from two similar specifications across the two studies. The Specification 1 results are similar—elasticities in both studies are in the typical range of a 1 to 3 percent disemployment effect from a 10 percent increase in the minimum wage, and both sets of estimates are significant at the 1 percent level. But when we include the division control (Specification 2), in both studies the disemployment effect is reduced substantially and it is not statistically distinguishable from zero. Adding state-specific time trend controls without division controls (Specification 3) also renders the employment outcomes in each study insignificant and (smaller in absolute value). The same is true with division-specific and state-specific time controls included in Specification 4.

While these results are not directly comparable, they support two conclusions. The first concerns the importance of including controls for heterogeneous trends in low-wage employment. In Dube, Lester and Reich, inclusion of division-specific time effects and state-level linear time trends provide imperfect proxies for their local estimators, which also produce employment elasticities indistinguishable from zero. Although CPS data limitations

preclude replicating the analysis at such a local level, the inclusion of these controls attenuates the disemployment effect for teens in the CPS in an analogous manner. The omission of controls for local differences in underlying local labor market conditions induces a serious bias in the teen studies as well.

The second conclusion concerns the similar coefficients for each specification across the two studies. Since the proportions of minimum wage workers who are teens and who are restaurant workers are very similar, it is not surprising that the estimated effects are also similar. Differences in findings are the result of different specifications and identifying assumptions, not different data sets or the groups under investigation.

8. Summary and conclusions

Using the canonical fixed effects specification on the sample of teens, we estimate an employment elasticity of -0.153, similar to the -0.3 to -0.1 percent disemployment consensus of the estimates in other national CPS studies. But including a Census division-specific time control to the specification reduces this effect to -0.105 and it is no longer distinguishable from zero. Adding a state-specific time control also reduces the employment effect and renders it insignificant. The inclusion of division and state-specific time controls thus greatly affects the economic and statistical significance of the estimates. Our plot of the time path of teen employment around the minimum wage change using only the canonical time and state controls indicates that teen employment was unusually low and falling substantially *prior* to the actual increase. Overall, the evidence strongly points to the failure of the canonical fixed-effects specification to control for the heterogeneity and selectivity of states where minimum wages increased over this period.

We also find that traditional fixed-effects estimates are sensitive to the demographic group under study and to the sample time frame, but they become much more similar when richer controls are included in the specifications. Our separate examination of the period from 1998 to 2006 raises serious concerns regarding the source of identifying variation over this period, especially in light of the selectivity of the states and the heterogeneous impacts of the 2001 recession. Absent a rich set of controls for counterfactual employment, we find that such identifying variation can produce misleading results.

Since the proportion of teens and the proportions of restaurant workers who are paid at or near the minimum wage are very similar it is of interest to compare our estimates to those in Dube, Lester and Reich (2007). The estimated minimum wage employment effects are very close in both studies. Moreover, the results in the two studies change in similar ways with the inclusion of controls for spatial heterogeneity. These results suggest that the effects of controlling for such heterogeneity do not result from the focus on any one group, industry or dataset.

Our analysis finds that heterogeneity in employment patterns and selectivity of states constitutes a significant concern for conventional minimum wage studies. Although adding division and state trend controls do not constitute a panacea, they do provide important controls that mitigate the bias that results from unobserved heterogeneities that may be correlated with minimum wage changes. Since estimates in previous national-level studies insufficiently address this issue, the interpretation of the evidence in the existing minimum wage literature must be revised accordingly.

References

Aaronson, Daniel, Kyung-Hong Park and Daniel Sullivan 2006. "The Decline in Teen Labor Force Participation." *Economic Perspectives,* Federal Reserve Bank of Chicago 30, 1: 2-18.

Card, David and Alan Krueger 1994. "Minimum Wages and Employment: a Case Study of the Fast-Food Industry in New Jersey and Pennsylvania." *American Economic Review* 84, 4: 772-93.

_____2000. "Minimum Wages and Employment: a Case Study of the Fast-Food Industry in New Jersey and Pennsylvania: Reply." *American Economic Review* 90, 5: 1397-1420.

Congressional Budget Office 2004. "What is Happening to Youth Employment Rates." CBO Paper. Washington, D.C.: Congressional Budget Office.

Dube, Arindrajit, William Lester and Michael Reich 2007. "Minimum Wage Effects Across State Borders: Estimating Using Contiguous Counties." Working Paper 157-07. Revised November 30, 2007. Berkeley, CA: UC Berkeley Institute for Research on Labor and Employment. www.irle.berkeley.edu/workingpapers/157-07.pdf.

Dube, Arindrajit, Suresh Naidu and Michael Reich 2007. "The Economic Effects of a Citywide Minimum Wage." *Industrial and Labor Relations Review* 60, 4: 522-543.

Foote, Christopher 2007. "Space and Time in Macroeconomic Panel Data: Young Workers and State-Level Unemployment Revisited." Working Paper 07-10. Boston, MA: Federal Reserve Bank of Boston.

Neumark, David and William Wascher 2000. "Minimum Wages and Employment: a Case Study of the Fast-Food Industry in New Jersey and Pennsylvania: Comment." *American Economic Review* 90, 5: 1362-96.

_____2007. "Minimum Wages, the Earned Income Tax Credit and Employment: Evidence from the Post-Welfare Reform Era." Working Paper 12915. Cambridge, MA: National Bureau for Economic Research.

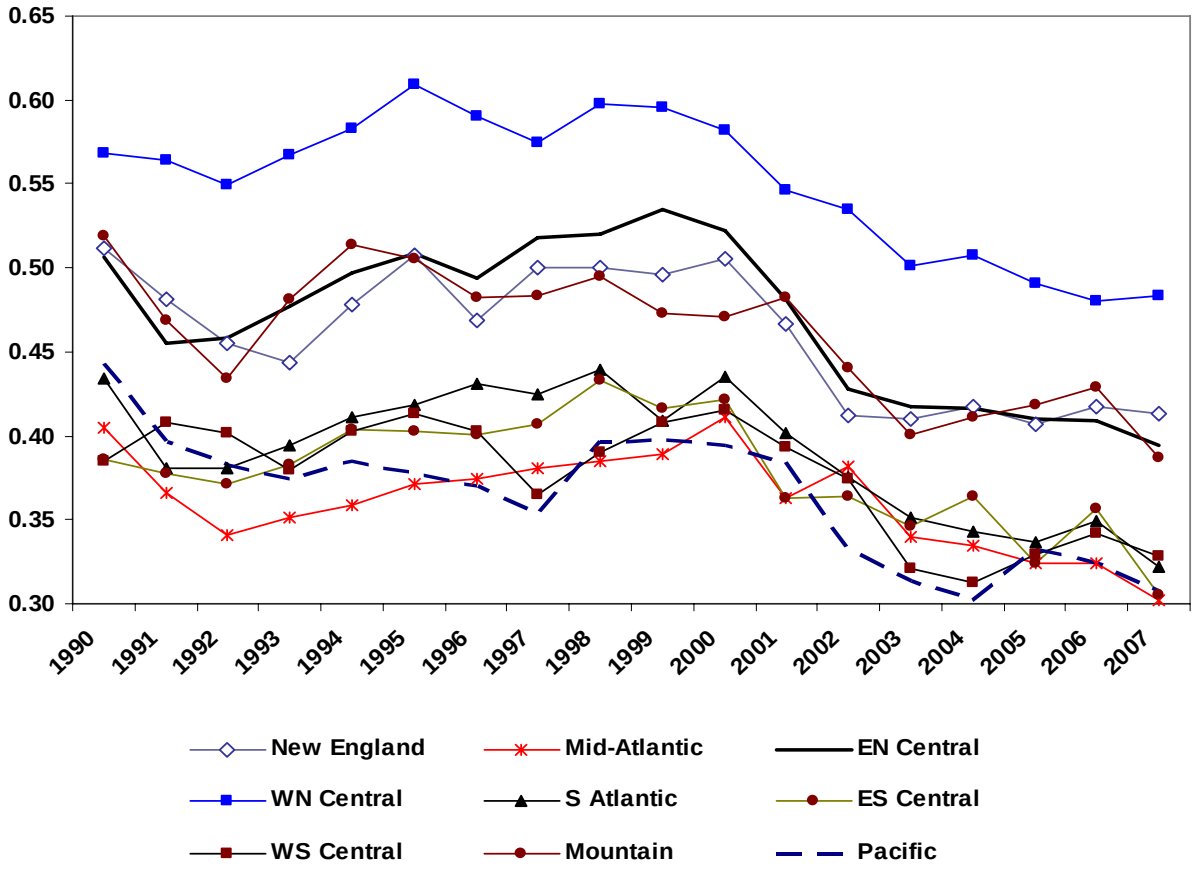
Orrenius, Pia and Madeline Zavodny 2008. "The Effects of Minimum Wages on Immigrants." *Industrial and Labor Relations Review* 61, 4: 544-63.

Raphael, Steven 1998. "The Spatial Mismatch Hypothesis and Black Youth Joblessness." *Journal of Urban Economics* 43, 1: 79-111.

_____and Michael Stoll 2002. "Modest Progress: The Narrowing Spatial Mismatch between Blacks and Jobs in the 1990s." Brookings Institution: Washington, DC.

Sabia, Joseph 2006. “The Effect of Minimum Wage Increases on Retail and Small Business Employment.” Washington, D.C.: Employment Policies Institute.

Figure 1 Employment to population ratio, teens 16-19, by nine Census divisions, 1990-2007

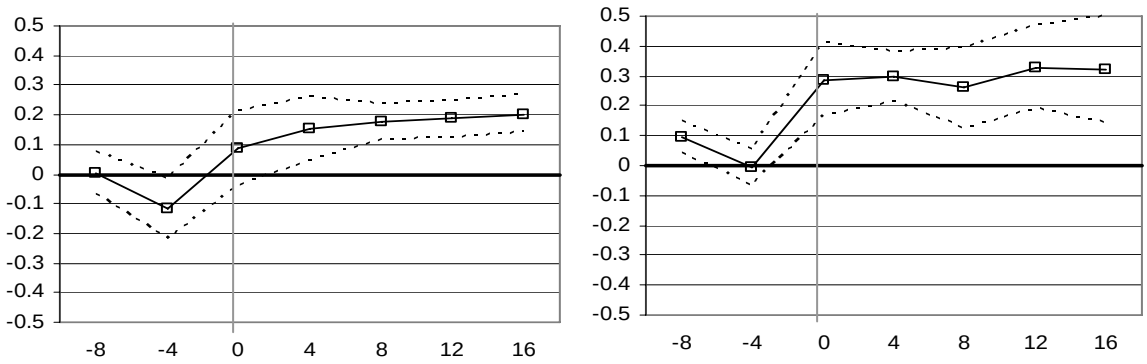


Source: Authors' analysis of Current Population Survey data. See Table 1 for a listing of states within each Census division.

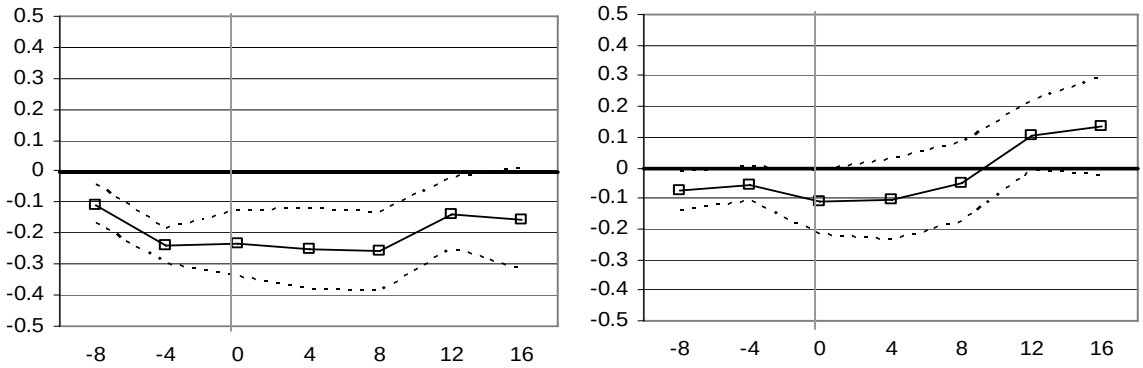
Figure 2 Time paths of wages, employment and hours in response to a minimum wage change

Spec 1 (No additional controls) *Spec 4 (State-linear trends and division-specific time effects)*

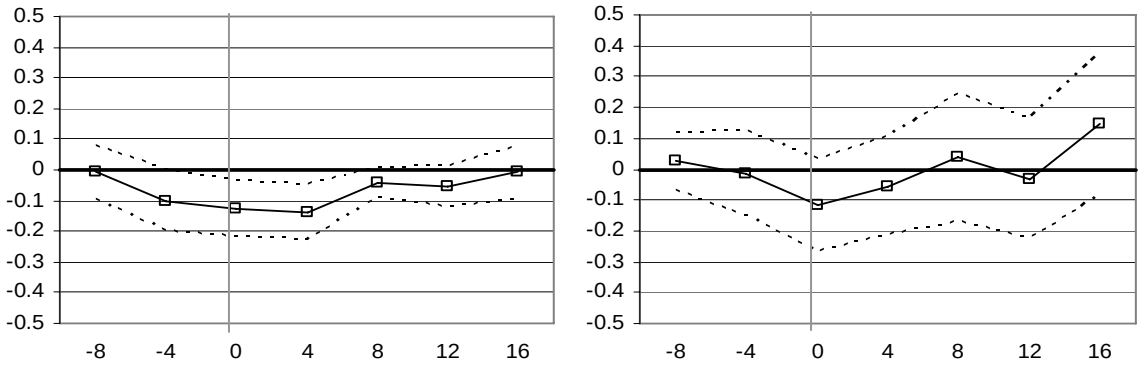
A Log Wages



B Employment

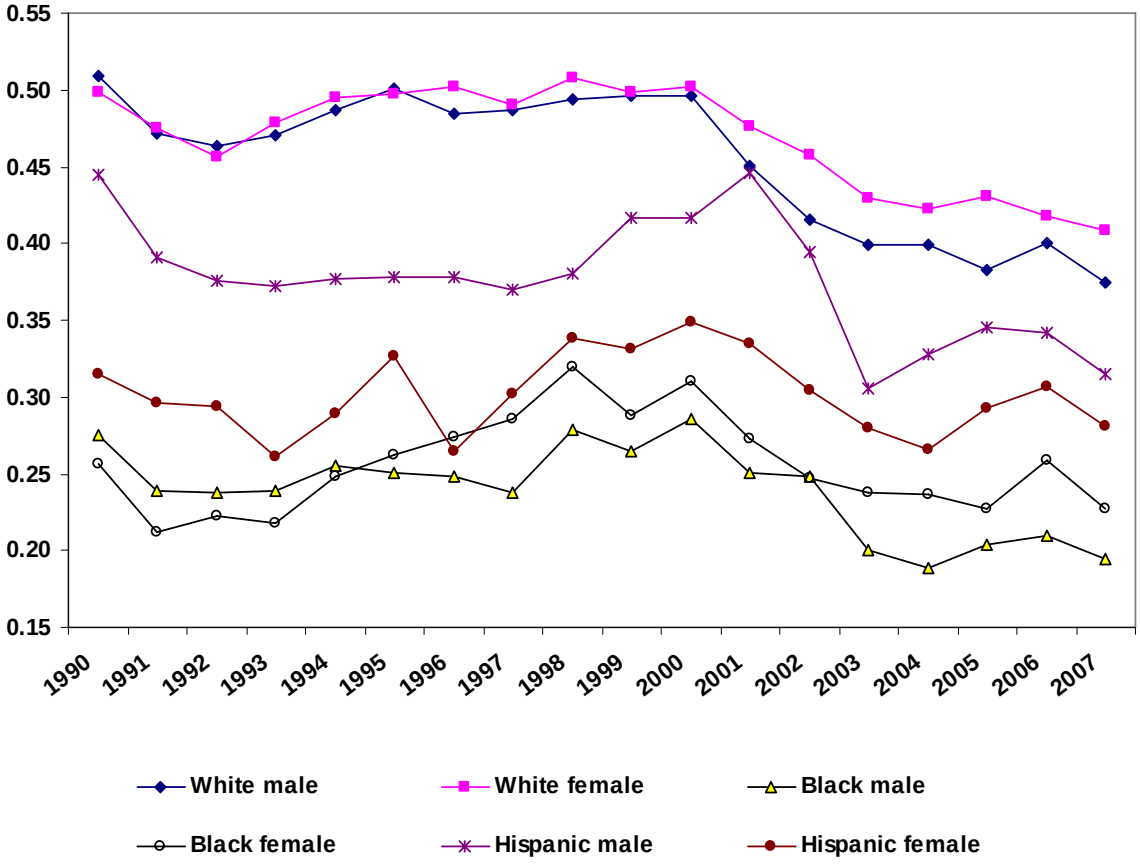


C Log Hours



Notes: Using a distributed lag specification of two leads, four lags and the contemporaneous log minimum wage, the figures above plot the *cumulative response* of log wage, employment and log hours to a minimum wage increase. We consider a 25 quarter window around the minimum wage increase. For employment, coefficients are divided by average teen employment-to-population ratio, so the coefficients represent employment elasticities. Specification 1 includes time and state fixed effects as well as the set of demographic controls reported in the text. Specification 4 additionally includes state-level linear trends and division-specific time effects (hence eliminating the variation among Census divisions). For all specifications we plot the 90% confidence interval around the estimates in dotted lines. The confidence intervals were calculated using robust standard errors clustered at the state level.

Figure 3 Employment to population ratio, teens 16-19, by demographic groups, 1990-2007



Note: Authors' analysis of Current Population Survey data. White refers to non-Hispanic white.

Table 1 Employment to population ratios, teens 16-19, by Census division, selected years

	1990	1998	2007	Change 1990 to 1998	Change 1998 to 2007
United States	0.46	0.45	0.39	0.00	-0.07
New England Maine, New Hampshire, Vermont, Massachusetts, Rhode Island, Connecticut	0.51	0.50	0.41	-0.01	-0.09
Middle Atlantic New York, New Jersey, Pennsylvania	0.41	0.38	0.30	-0.02	-0.08
East North Central Ohio, Indiana, Illinois, Michigan, Wisconsin	0.51	0.52	0.39	0.01	-0.13
West North Central Minnesota, Iowa, Missouri, North Dakota, South Dakota, Nebraska, Kansas	0.57	0.60	0.48	0.03	-0.11
South Atlantic Delaware, Maryland, DC, Virginia, West Virginia, North Carolina, South Carolina, Georgia, Florida	0.43	0.44	0.32	0.01	-0.12
East South Central Kentucky, Tennessee, Alabama, Mississippi	0.39	0.43	0.31	0.05	-0.13
West South Central Arkansas, Louisiana, Oklahoma, Texas	0.39	0.39	0.33	0.01	-0.06
Mountain Montana, Idaho, Wyoming, Colorado, New Mexico, Arizona, Utah, Nevada	0.52	0.50	0.39	-0.02	-0.11
Pacific Washington, Oregon, California, Alaska, Hawaii	0.44	0.40	0.31	-0.05	-0.09

Source: Authors' calculations of Current Population Survey data.

Table 2 Descriptive statistics, teens 16-19, 1990-2007

	Mean	Std dev	N
Sample statistics			401,744
Male	0.51	--	203,320
White, non-Hispanic	*0.36*	--	*153,787*
Black	*0.07*	--	*24,612*
Hispanic	*0.08*	--	*24,921*
Female	0.49	--	198,424
White, non-Hispanic	*0.35*	--	*149,759*
Black	*0.08*	--	*25,303*
Hispanic	*0.07*	--	*23,362*
Labor market outcomes			
Employed	0.41	--	169,783
Male	0.41	--	*85,228*
Female	0.41	--	*84,555*
White, non-Hispanic	0.46	--	*141,264*
Black	0.25	--	*12,192*
Hispanic	0.34	--	*16,327*
Hourly wage	$6.44	$6.53	165,455
Male	$6.72	$7.24	*82,356*
Female	$6.16	$5.72	*83,099*
White, non-Hispanic	$6.41	$6.00	*137,398*
Black	$6.40	$10.91	*12,039*
Hispanic	$6.67	$5.38	*16,018*
Hours worked per week	24.9	12.09	167,814
Male	26.4	12.61	*83,868*
Female	23.3	11.32	*83,946*
White, non-Hispanic	24.1	12.10	*139,490*
Black	25.6	11.10	*12,124*
Hispanic	29.1	11.83	*16,200*
Policy variables			
Minimum wage	$5.06	0.86	--
Minimum wage (state above federal)	$6.17	0.93	--
Unemployment rate	5.43	1.44	--

Source: CPS data. Notes: White includes Asian-American. Weighted statistics. Standard deviations reported for continuous variables. Average hourly wage is calculated for workers who reported a wage and were not self-employed or working without pay. Average hours worked is reported for workers with positive usual hours of work.

Table 3 Minimum wage effects on wages, employment and hours worked, teens 16-19, 1990-2007

Specification		(1)	(2)	(3)	(4)
A Wages					
All	η	0.120***	0.154***	0.205***	0.158***
	se	(0.026)	(0.035)	(0.033)	(0.031)
Males	η	0.096***	0.142***	0.177***	0.148***
	se	(0.026)	(0.033)	(0.037)	(0.034)
Females	η	0.139***	0.154***	0.231***	0.149***
	se	(0.031)	(0.042)	(0.042)	(0.040)
B Employment					
All	coeff	-0.063***	-0.043	-0.027	-0.010
	se	(0.017)	(0.037)	(0.025)	(0.031)
	η	-0.153***	-0.105	-0.065	-0.024
Males	coeff	-0.054***	-0.024	-0.007	0.010
	se	(0.021)	(0.048)	(0.032)	(0.044)
	η	-0.131***	-0.058	-0.018	0.025
Females	coeff	-0.078***	-0.073*	-0.050*	-0.038
	se	(0.022)	(0.037)	(0.029)	(0.042)
	η	-0.190***	-0.176*	-0.122*	-0.103
C Hours					
All	η	-0.096***	-0.091*	-0.038	-0.087**
	se	(0.037)	(0.051)	(0.047)	(0.042)
Males	η	-0.075	-0.105	-0.043	-0.126**
	se	(0.064)	(0.076)	(0.064)	(0.062)
Females	η	-0.116***	-0.078	-0.042	-0.053
	se	(0.037)	(0.059)	(0.050)	(0.056)
Division-specific time controls			Y		Y
State-specific time trends				Y	Y

Notes: η refers to elasticity. Significance levels are ***1%, **5%, *10%. Results are reported for the log of the minimum wage. Each specification includes individual controls for gender, race (4 categories), age (4 categories), education (12 categories), and marital status (4 categories), as well as controls for the non-seasonally adjusted unemployment rate, and the relevant population share for each demographic group. Wage regressions include only those were working and paid between $1 and $100 per hour in 1990 dollars and the log of hourly wage is the dependent variable. Hour regressions are restricted to those who had positive hours and the log of hours is the dependent variable. Standard errors are clustered at the state level.

Table 4 Minimum wage effects on employment, teens 16-19, by time period

Specification	(1)	(2)	(3)	(4)
	Employment elasticities and standard errors			
1990 to 2007	-0.153***	-0.105	-0.065	-0.024
	(0.042)	(0.089)	(0.061)	(0.075)
1990 to 2006	-0.165***	-0.096	-0.020	-0.019
	(0.048)	(0.104)	(0.069)	(0.089)
1987 to 2006	-0.159***	-0.085	-0.030	-0.028
	(0.043)	(0.086)	(0.065)	(0.075)
1990 to 1998	0.072	0.025	0.068	0.057
	(0.068)	(0.125)	(0.069)	(0.125)
1998 to 2006	-0.163***	-0.143*	-0.155**	-0.037
	(0.052)	(0.085)	(0.081)	(0.103)
1997 to 2005	-0.102	-0.196	-0.098	-0.100
	(0.071)	(0.134)	(0.084)	(0.130)
Division-specific time controls		Y		Y
State-specific time trends			Y	Y

Notes: Significance levels are ***1%, **5%, *10%. Results are reported for the log of the minimum wage. Each specification includes individual controls for gender, race (4 categories), age (4 categories), education (12 categories), and marital status (4 categories), as well as controls for the non-seasonally adjusted unemployment rate, and the relevant population share for each demographic group. Standard errors are clustered at the state level.

Table 5 Minimum wage elasticities on wages and employment, teens 16-19, by race/ethnicity, 1990-2007

Specification	(1)	(2)	(3)	(4)
A Wages				
White, non-Hispanic	0.129***	0.179***	0.224***	0.173***
	(0.026)	(0.038)	(0.036)	(0.033)
Black	0.088	0.074	0.206**	0.227***
	(0.054)	(0.083)	(0.081)	(0.086)
Hispanic	0.101**	-0.054	0.107**	-0.042
	(0.044)	(0.062)	(0.053)	(0.070)
B Employment				
White, non-Hispanic	-0.140***	-0.136*	-0.056	-0.045
	(0.042)	(0.080)	(0.050)	(0.067)
Black	-0.116	0.337*	-0.111	0.450***
	(0.153)	(0.198)	(0.187)	(0.167)
Hispanic	-0.101	-0.216	-0.115	-0.187
	(0.095)	(0.211)	(0.112)	(0.183)
Division-specific time controls		Y		Y
State-specific time trends			Y	Y

Note: See notes to Table 3.

Table 6 A comparison of minimum wage employment elasticities

Study	Specification			
	(1)	(2)	(3)	(4)
This study CPS, teens *1990-2007*	-0.153*** (0.042)	-0.105 (0.089)	-0.065 (0.061)	-0.024 (0.075)
90% CI	(-0.222,-0.083)	(-0.252, 0.041)	(-0.166, 0.036)	(-0.147, 0.099)
This study CPS, teens *1990-2006*	-0.168*** (0.020)	-0.103 (0.043)	-0.025 (0.029)	-0.024 (0.038)
90% CI	(-0.247, -0.089)	(-0.274, 0.069)	(-0.140, 0.091)	(-0.173, 0.124)
Dube et al. (2007) QCEW, restaurants *1990-2006*	-0.207*** (0.063)	-0.076 (0.060)	0.055 (0.042)	0.060 (0.041)
90% CI	(-0.312, -0.102)	(-0.176, 0.023)	(-0.014, 0.124)	(-0.007, 0.127)
Division-specific time controls		Y		Y
State-specific time trends			Y	Y

Notes: Elasticities are not directly comparable. They are presented to show the effects of using similar model specifications and controls. Significance levels: ***1%, **5%, *10%.