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The Effect of the Minimum Wage on Hours of Work

Abstract: Recent studies of the effects of the minimum wage have focused on employment, but employers may adjust hours as well. This study examines the effect of increases in the minimum wage on teen hours of work and employment using both state- and individual-level panel data from the Current Population Survey. The results indicate that teens who are likely to be affected by minimum wage increases are less likely to remain employed than unaffected teen workers, but experience greater increases in hours conditional on remaining employed. The effect of the minimum wage on hours among workers likely to be affected remains non-negative even when accounting for teens who do not remain employed. The results suggest that aggregate data mask employment shifts among teen workers with different skill levels.

JEL classification: J38

1. Introduction

Recent research finding that increases in the minimum wage appear to not reduce employment has challenged the traditional model of the effects of the minimum wage. A growing set of papers find that increases in federal and state minimum wages do not significantly lower employment among teens or fast-food workers.[1] Earlier studies, as summarized by Charles Brown, Curtis Gilroy and Andrew Kohen (1982), usually concluded that a 10-percent increase in the minimum wage reduced teen employment by 1 to 3 percent.[2] Reasons offered for the surprising new findings include poor data, problematic methodologies and the possibility of monopsony power. This study investigates another possibility: employers may reduce hours instead of employment in response to minimum wage hikes.

The traditional model of labor demand posits a downward-sloping labor demand curve and an upward-sloping labor supply curve. If firms demand less labor as the cost of labor increases, imposition of a binding minimum wage should reduce the quantity of labor demanded. In the usual interpretation of the model, employment falls as the minimum wage rises because employers lay off workers whose marginal revenue product is less than the wage floor. However, if all workers are homogeneous and firms seek to minimize payroll subject to staffing needs, employers may cut workers' hours instead of reducing employment. The model thus can predict that the number of labor hours demanded falls as the hourly wage increases. Changes in the minimum wage may cause either employment or hours per

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[1]See Alison J. Wellington (1991), David Card (1992a and 1992b), Lawrence F. Katz and Alan B. Krueger (1992), Card and Krueger (1994) and David Neumark and William Wascher (1995b). Stephen Machin and Alan Manning (1992) provide similar evidence for Britain.

[2]Janet Currie and Bruce C. Fallick (1996), Taeil Kim and Lowell Taylor (1995), and Donald Deere, Kevin M. Murphy and Finis Welch (1995) find similar results that support the traditional model.

employee, or both, to change.

There are several reasons to believe employers may cut hours instead of the number of workers when the minimum wage increases. First, it may be easier for firms to adjust hours than employment, particularly in the short run. Firing workers may lower the morale of retained workers, and waiting for natural attrition to reduce employment levels takes time. In addition, workers earning a higher hourly wage because of a minimum wage hike are likely to be willing to work fewer hours. Since many low-wage workers are part-time and do not receive benefits, employers may not save a fixed cost per worker by reducing employment instead of hours.[3]

Although the minimum wage literature has focused on employment, a few studies have examined the relationship between the minimum wage and hours. Edward M. Gramlich (1976) finds that teens and adult males move from full-time to part-time employment as the minimum wages rises; adult females, however, appear to shift from part-time to full-time jobs. Katz and Krueger (1992) suggest that fast-food restaurants in Texas increased full-time employment and decreased part-time employment when the federal minimum wage rose in 1991. James Cunningham (1981) also finds that higher minimum wages increase full-time while decreasing part-time employment among teens. Neumark and Wascher (1995a) simulate hours changes from the effects of minimum wages on employment and school enrollment and predict that hours worked fall as the minimum wage increases.

This study examines the relationships between the minimum wage and hours and

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[3]Employers may also engage in "labor hoarding" by reducing hours per worker and retaining workers since inflation will eventually erode the real minimum wage. However, most minimum wage jobs are not highly skilled and have high turnover rates.

employment using both state- and individual-level data. State annual averages are first used to examine the effects of the minimum wage on aggregate teen employment and average hours per week. Aggregate data may mask shifts among teen workers of different skill levels, however, so individual data are used to examine the effects on teens' probability remaining employed and hours of work.

The results indicates that teen aggregate employment is not adversely affected by minimum wage increases, and average teen hours either increase or remain constant. Although this implies that teens are better off as the minimum wage increases, the individual-level results show that teens likely to be affected by minimum wage increases are, on average, 2-3 percent less likely to remain employed than unaffected teens. Those teens who do remain employed and are likely to be affected by minimum wage hikes experience a relative increase in hours, suggesting that they substitute work for leisure as their hourly pay increases. The effect of the minimum wage on teen hours is non-negative even when teens who do not remain employed and whose hours of work go to zero are included in the analysis.

2. Methodology and Data

The recent empirical literature on the effects of the minimum wage relies on a differences-in-differences methodology. The effect of a minimum wage increase is estimated by comparing employment among a group likely to be affected by the hike to unaffected workers. Recent studies use either state- or firm-level panel data to compare employment changes among teens, in retail trade or at fast-food restaurants in areas experiencing minimum wage increases to employment changes in other areas. Alternatively, data on individuals can be also used to measure the effect of minimum wage increases on teens' probability of remaining employed. Currie and Fallick (1996) and Neumark and Wascher (1995a) employ this methodology.

This study combines these approaches and extends the analysis to hours. I use panel data at both the state and individual levels to estimate the effect of the minimum wage on teen employment and hours. There are several advantages to using both aggregate and individual data. Using state averages allows for an examination of the aggregate effects of the minimum wage. However, aggregate data may mask shifts among workers. For example, minimum wage increases may draw higher-skilled teens into the labor market, and employers may substitute them for lower-skilled teens. This concern motivates the use of panel data on individuals.

This study uses data from the NBER extracts of the Current Population Survey (CPS) outgoing rotation groups for the years 1979-1993. Participants in the CPS are surveyed for four months, rotate out of the panel for eight months, are surveyed again for four months and then exit the panel. In the last month of each rotation, participants are asked their employment status, hours and earnings. Thus, new individuals are continually entering the CPS, and two observations on employment, wages and hours are available for each individual. Data on teens are used to examine the effect of state and federal minimum wages on employment, hours and earnings.

Both federal and state minimum wages experienced sizable change during the period 1979-1993. The nominal federal wage floor rose from \$2.90 to \$3.10 in January 1980, to

$3.35 in January 1981, to $3.80 in April 1990 and to $4.25 in April 1991. The real value of the federal minimum wage declined by one-third between 1981 and 1990, and 16 states passed state minimum wages above the federal minimum wage during this period.[4] This gives variation in effective minimum wages (the higher of the state and federal minimum wages) not only over time but also across states.

The state-level regressions exploit this variation in effective minimum wages to test the prediction that either aggregate employment or average hours falls as the minimum wage increases. A fixed-effects methodology is used to control for time-invariant unobservable differences across states and business-cycle effects common to all states. In the individual-level regressions, I test the prediction of the traditional model that when a new wage floor is imposed, teens who initially earn less than the new minimum wage ("bound" workers) should have a lower probability of remaining employed or should experience a decline in hours relative to teens who initially earn more than the new minimum. I restrict the sample to teens employed in a given year and compare the effects of minimum wage changes on employment status and hours one year later between "bound" and unaffected teens. Currie and Fallick (1996) also employ this approach but use a different data set and do not examine hours.[5]

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[4]Neumark and Wascher (1992) provide a chronicle of minimum wage changes through 1989. State minimum wages were also obtained from the annual summary of state labor law changes in the January issues of the Bureau of Labor Statistics publication *Monthly Labor Review*.

[5]Currie and Fallick use the National Longitudinal Survey of Youth (NLSY), which offers a longer panel. The NLSY is best used to examine the effects of changes in the federal minimum wage in the early 1980s; the participants were no longer teenagers in the late 1980s, when most of the variation in state minimum wages occurred, and when the federal minimum wage was raised in 1990 and 1991.

3. Aggregate Estimates

State annual averages are used to estimate the aggregate effects of increases in the minimum wage on the teen employment-to-population ratio, hours of work per week, hourly earnings and weekly earnings. The averages are calculated from the individual responses of teens aged 16-19 in the CPS outgoing rotation groups, aggregated using the CPS demographic weights. Teens employed in agriculture, in unpaid jobs or self-employed were dropped from the sample. In the hours and hourly wage regressions, the sample is also restricted to teens paid hourly wages.[6] The data on hours and earnings used here pertain to the primary job if an individual held multiple jobs.

The estimation framework is a basic panel data model and is similar to the model employed by Neumark and Wascher (1992). The model estimated for all of the outcomes is

$$\ln Y_{it} = \alpha + \beta \ln MW_{it} + \gamma URATE_{it} + \delta POP_{it} + \sigma S_i + \theta T_t + \varepsilon_{it} \qquad (1)$$

where Y_{it} is the employment-to-population ratio, weekly hours, real hourly earnings or real weekly earnings for teens. MW_{it} is the effective minimum wage (the higher of the state and federal minimum wages) deflated using the CPI for urban consumers.[7] $URATE_{it}$ is the unemployment rate of males aged 25-64 and is included to capture business-cycle effects. POP_{it} is the teen-to-total population ratio and is included to control for "baby boom" effects.

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[6]Keeping individuals employed in agriculture, in unpaid jobs, self-employed or salaried in the sample does not qualitatively change any of the reported estimates. Results using the unedited data are available from the author upon request.

[7]Several variables can be used as measures of the wage floor. Earlier papers often used the Kaitz index, a coverage-weighted ratio of the minimum wage to the average wage. As Card, Katz and Krueger (1994) discuss, use of the Kaitz index may lead to overestimates of the disemployment effect of the minimum wage. The relative minimum wage (the effective minimum wage divided by the state average manufacturing wage) can also be used, but it should be instrumented using the minimum wage since the manufacturing wage may be correlated with the error term. Instrumenting with the minimum wage usually gives similar results to using the minimum wage as the independent variable, so only the reduced-form estimates are presented here.

S_i and T_t are fixed state and year effects, and the error term ε_{it} is corrected for heteroskedasticity and serial correlation using the Prais-Winston procedure.[8] The lagged real effective minimum wage, MW_{it-1}, is also included in some specifications since Neumark and Wascher (1992) find that lagged effects appear to be larger than contemporaneous effects. The data are annual averages for the 50 states for 1979-1993, giving a total of 750 observations.

[Insert Table 1 here]

The employment-to-population ratio is not significantly correlated with the minimum wage, as shown in the first two columns of Table 1. When the lagged minimum wage is included, its estimated coefficient is negative but not statistically different from zero at conventional levels. These results accord with Neumark and Wascher (1995b), who use state-level data from the CPS for 1977-1989, and much of the recent literature on the employment effects of the minimum wage.

The minimum wage is positively correlated with average weekly hours among employed teens. As shown in the third column of Table 2, a 10 percent increase in the minimum wage is correlated with a 1.7 percent increase in hours per week. This effect is qualitatively small; the sample average is about 26 hours per week, so a 20 percent increase in the minimum wage would raise average hours among employed workers by less than an hour a week. When the lagged minimum wage is included in the regressions, the coefficients on both minimum wage variables are positive, and the coefficient on the lagged minimum

[8]The Prais-Winston method used here corrects for autocorrelation within cross-sectional units. An AR(1) process is assumed, and the serial correlation parameter is constrained to be the same for all states. The panel data used here are not long enough to allow for accurate estimation of a separate serial correlation parameter for each state.

wage is statistically significant at the 10 percent level.

Taken together, the employment and hours results for teens do not appear consistent with the traditional model. Neither employment nor average hours falls as the minimum wage increases. There are several possibilities consistent with these results and the traditional model. First, the increases in the minimum wage might not be binding. If Equation (1) is estimated with either hourly wages or weekly earnings as the dependent variable, however, the results indicate that increases in the minimum wage are indeed correlated with higher teen wages. A 10 percent increase in the minimum wage raises teen hourly wages by 5 percent--or the wages of about one-half of teen workers rise by the full amount of minimum wage increases--and raises weekly earnings by almost 7 percent.[9] The larger effect on weekly earnings than on hourly wages is also consistent with the positive correlation between the minimum wage and hours.

Another explanation for these results is that employers substitute teens for other workers--either other teens or adults--as the minimum wage increases. If employers substitute teens for adult female workers--the other primary low-wage group-- teen employment and/or hours might increase. However, if Equation (1) is estimated for adult women, the results indicate that female employment and hours are not adversely affected by increases in the minimum wage even though female wages are also positively correlated with the minimum wage.

Employers may substitute high-skill teens for low-skill teens, leaving total teen employment unchanged. Cunningham (1981) and Neumark and Wascher (1995a and 1995b)

[9]The coefficients on the minimum wage variable are .490 (s.e.=.064) and .691 (.117) in the hourly wage and weekly earnings regressions, respectively.

suggest that minimum wage increases may cause higher-skilled teens to leave school and work full-time. In this case, employers may also shift employment from part-time to full-time jobs, explaining the surprising positive relationship I find between the minimum wage and teen hours. These possibilities are examined below with panel data on individuals.

4. Individual Estimates

If workers are paid their marginal product of labor, imposition of a binding minimum wage will make it unprofitable for firms to continue to employ workers who initially earn less than the new minimum wage unless these workers raise their productivity. The traditional model therefore predicts that individuals initially earning less than the new minimum wage should be less likely to remain employed or should experience a decrease in hours relative to unaffected workers after a minimum wage hike. These effects should be larger the greater the difference between a worker's initial earnings and the new wage floor, or the larger the "wage gap" for "bound" workers. These predictions are tested by comparing changes in employment status and hours between bound and unaffected teen workers using panel data.

The CPS sample is restricted to individuals who can be matched across two consecutive years and are employed, report wage data and are aged 16-19 in the first rotation. Individuals were matched using the household number, age, sex, race and ethnicity. The resultant data set contains 55,000 matched records, with observations in each of the 50 states for the periods 1979-1980 through 1992-1993.[10] After individuals employed in the

[10]Matches for 1984-1985 could only be made for the months January through June, and matches for 1985-1986 could only be made through the months October through December because the household numbers changed in 1985. An individual is defined as employed if the employment status recode is one.

public sector, agriculture or domestic service (sectors generally not covered by the minimum wage) and 30 individuals whose records did not include an employment status recode in the second year are dropped from the sample, the data set contains 46,915 matched records.

[Insert Table 2 here]

Table 2 provides descriptive statistics for the sample. More than 70 percent of teens employed in the first year were employed twelve months later. About 13 percent of workers were likely to be directly affected by a minimum wage increase between interviews, or were "bound" by a minimum wage increase. To be classified as bound, a worker had to earn at least the first-period minimum wage but less than the second-period minimum wage in the first period, or

$$MW_t \leq Wage_t < MW_{t+1} \tag{2}$$

The average real wage gap, the difference between $Wage_t$ and MW_{t+1}, among bound workers was almost 25 cents.[11] This wage gap is about 6 percent of the average hourly wage for the entire sample. No individuals were bound during the periods 1981-1982 through 1983-1984 because no states experienced minimum wage increases during these intervals.

The definition of bound workers implies that these workers should earn exactly the minimum wage in the second period. The second-period is wage is not used to classify workers as bound or nonbound since the wage at t+1 is not observed for teens who do not remain employed. Teens classified as bound who remain employed are indeed more likely to be paid exactly the minimum wage in the second period than teens who are classified as

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[11]The wage gap was deflated using the CPI for urban consumers in t+1. The regression results are similar if either the nominal wage gap or the difference between the real wage at t and the real minimum wage at t+1 is used instead of the real difference between the wage at t and the minimum wage at t+1.

nonbound and remain employed (20.6 percent v. 8.8 percent).

Those individuals who already earned more than the new minimum wage or who did not experience an increase in the minimum wage during the year between interviews serve as the comparison group. As expected, these workers have higher initial hourly wages than bound workers. Unaffected workers are also more likely to still be employed after a minimum wage increase. These workers tend to be slightly older, more educated and less likely to be female than bound workers. None of the differences between the two groups in the summary statistics presented in Table 2 are statistically significant, however.

A. Employment Effects

As in Currie and Fallick (1996), a linear probability model is used to test the prediction that the larger the wage gap, the lower the probability of remaining employed. Ordinary least squares (OLS) was used to estimate

$$E_{it} = \alpha + \beta \mathrm{Wagegap}_{it} + \delta X_{it} + \eta_{it} \tag{3}$$

where E_{it} equals one if a worker remains employed twelve months later and zero otherwise. As explained above, $Wagegap_{it}$ is the real difference between the hourly wage at t and the minimum wage at t+1 if the wage at t was at least the minimum wage at t and less than minimum wage at t+1. $Wagegap_{it}$ equals zero for unaffected workers. The vector X_{it} contains demographic variables. State and year dummy variables are included in some specifications, and observations are weighted using the CPS demographic weights in the first year. OLS was used for ease in interpreting the estimates; the estimated marginal effects in probit regressions were similar, although smaller in magnitude, to the OLS results reported here.[12]

[Insert Table 3 here]

As shown in the first two columns of Table 3, the wage gap is negatively correlated with the probability of remaining employed. Multiplying the coefficient on the wage gap variable and the average wage gap of bound teens, bound workers' probability of continued employment is about 2.5 percent lower than unaffected workers. This result is similar, although slightly smaller, to that obtained by Currie and Fallick (1996) using a sample from the NLSY in the early 1980s. The coefficient on the wage gap variable is robust to including agricultural, unpaid and domestic workers, classifying workers missing the employment recode in the second period as either still employed or unemployed, and excluding salaried workers. In results not reported here, the estimated effect is also generally the same if subsets of the data for different periods are used, although the estimated coefficient on the wage gap variable becomes smaller toward the end of the period 1979-1993.

B. Hours Effects

A similar model is used to test the prediction that bound workers should experience a decline in hours relative to unaffected workers. OLS was used to estimate

$$\Delta \text{ Hours}_{it,t+1} = \alpha + \beta \text{Wagegap}_{it} + \delta X_{it} + \eta_{it} \tag{4}$$

where Δ $Hours_{it,t+1}$ is the change in an individual's hours between the two interviews. The other variables are the same as above. The sample initially consists only of teens who remained employed, and 40 individuals whose hours changed by more than 50 were dropped

[12]The probit results are available from the author upon request.

because of the likelihood that the observations had sizable measurement error. This reduces the sample size to 33,733 matched records. Both teen who remain employed at t+1 and those who do not are included in a second sample, and second-period hours are set equal to zero for teens who do not remain employed.

When the sample is restricted to teens who remain employed, the wage gap is positively correlated with the change in hours. As shown in the third column of Table 3, the estimate implies that the average bound worker's hours increased by 0.7 relative to unaffected workers when demographic variables are controlled for. The estimate rises to 1.3 when controlling for state and year fixed effects. The average change in hours for both bound and unaffected workers is positive, so the average bound worker experiences an increase in hours both absolutely and relative to unaffected workers. The relationship between the change in hours and the wage gap is robust to changes in the sample.

The above estimates only include workers who are still employed, who may differ substantially from teens who do not remain employed. For example, workers who remain employed may be more skilled than workers who are no longer employed. The employment results indicate that workers with larger wage gaps are less likely to remain employed. To estimate the total effect of minimum wage increases on hours, hours at t+1 were set to zero for individuals no longer employed and equation (4) was re-estimated. As shown in the last two columns of Table 3, the coefficient on the wage gap variable is not significantly different than zero when controlling for demographic variables and is positive when also controlling for state and year fixed effects. The results thus indicate that individuals more affected by minimum wage changes did not experience a decline in hours relative to unaffected workers, even when including teens who are no longer employed and whose hours of work effectively go to zero.

C. Possible Sources of Bias

The inability to match all individuals across successive years may raise concerns about sample selection bias. About 63 percent of the initial sample was successfully matched to a record in the following year, with the match rate declining with age. As discussed in Neumark and Wascher (1995a), regression estimates are inconsistent if the probability of a successful match is correlated with both the employment outcome and the independent variables. Heckman's two-stage method can be used to correct for selection bias if a variable correlated with the probability of matching but not correlated with the error term in the second-stage regression is available. However, no available variable clearly fulfills these requirements.

One reassurance about the possibility of selection bias is that characteristics of teens whose records matched do not significantly differ from non-matching teens. In particular, 13.5 percent of matching teens and 12.4 percent of non-matching teens (both after dropping uncovered industries) were likely to be affected by minimum wage increases. Table 4 gives summary statistics of the characteristics of teens who did and did not match. Most of the differences are likely driven by older teens leaving home and exiting the sample.

[Insert Table 4 here]

Heterogeneity bias is another potential concern. Workers who are bound by minimum wage increases may differ from unaffected workers in unobservable ways. In particular, bound workers may be more likely to leave employment or to experience hours increases regardless of changes in the minimum wage. The usual control for this problem is to use individual fixed effects in a sample with at least three observations per individual, which controls for constant individual-specific heterogeneity. However, only two observations for each individual are available in the CPS outgoing rotation groups. Currie and Fallick (1996) conclude that heterogeneity is not a significant concern based on their fixed effects employment estimates, and there is no reason why heterogeneity should be of greater concern in the CPS than in the NLSY.

Another method of controlling for heterogeneity is to include a dummy variable indicating whether an individual was in a relatively low-wage job in the first year. Teens working in low-wage jobs are likely to have less labor-market experience and fewer skills than higher-wage teens; as these teens accumulate experience and skills, they may be more likely to look for another job (and thus not be observed as employed a year later) or their hours may increase more than those of teens initially earning higher wages. Following Currie and Fallick (1996), a dummy variable for whether the first-period wage was within 15 cents of the current minimum wage was included in Equations 3 and 4 to control for this bias. This effectively gives low-wage and high-wage workers different slopes in the regressions.[13]

The results do not indicate that heterogeneity bias is driving the results in Table 3. Although low-wage teens are less likely to remain employed, the estimated coefficient on the

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[13]This method also exploits the fact that no minimum wage increases occurred during 1981-1984. Low-wage workers during this period are not likely to differ from low-wage workers who were bound by minimum wage increases. This method tests, in part, whether these teens--who were unaffected by minimum wage increases--were more likely to leave employment or experience hours increases than high-wage teens.

wage gap variable does not change in the employment regressions. In the hours regressions, low-wage teens do experience an increase in hours relative to other teen workers. In the hours regressions that include only teens who remain employed, the estimated coefficient on the wage gap variable falls in the hours regressions but remains positive and statistically significant when controlling for state and year fixed effects. In the hours regressions that include both teens who do and do not remain employed, the estimated coefficient on the wage gap variable is negative and statistically significant when only controlling for demographic variables but not statistically significant when also controlling for state and year fixed effects.[14]

Heterogeneity bias may of greater concern in the hours regressions. If bound teens were more likely to change jobs and job switchers experience an increase in hours relative to non-job switchers, the results may just reflect the effect of changing jobs on hours. Since most months of the CPS do not include tenure data, a dummy variable indicating whether an individual's 3-digit industry code in the CPS changed was included in Equation (4) to control for job switching. Equation (4) was then estimated using the sample of teens who remained employed. Switchers's hours do increase relative to non-switchers, but the coefficient on the wage gap variable does not change. The results thus appear robust to controlling for potential heterogeneity bias.

[14]The coefficient on the wage gap variable is -.090 (.030) in the employment regression controlling for demographic characteristics and -.090 (.035) when also controlling for state and year fixed effects. The coefficients in the hours regressions with only teens who remain employed are -.074 (.912) and 2.039 (1.033), respectively, and -2.908 (1.076) and -.377 (1.242) when all teens employed at time t are included.

D. Enrollment Effects

As suggested earlier, teen hours may increase if teens substitute work for school as the minimum wage increases. This may explain the surprising finding that bound teens' hours increase relative to unaffected teens; the substitution effect may dominate the income effect for teens and cause bound teens to leave school more than unaffected teens. To further investigate this possibility, the effect of the wage gap on the probability of being enrolled in school at time t+1 is estimated with OLS. The equation estimated for teens employed at both t and t+1 is

$$\text{Enroll}_{it+1} = \alpha + \beta\text{Wagegap}_{it} + \gamma\text{Enroll}_{it} + \delta X_{it} + \mu_{it} \tag{5}$$

where $Enroll_{it+1}$ equals one if individual i was enrolled at time t+1 and zero otherwise.[15] The other variables are as defined above, and month dummy variables were included in X_{it} because enrollment is considerably lower during the summer. The equation thus estimates the effect of the wage gap on the probability of being enrolled a year later, conditional on being employed at t and t+1.[16]

As shown in Table 5, the results do not indicate that bound teens are less likely to be enrolled in school after minimum wage increases than unaffected teens. The results are similar if the sample is further restricted to teens who are employed in both periods and enrolled in school in the first period. Although the bound teens who remain employed may

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[15]Enrollment status is taken from the major activity question. The employment and hours regressions do not include teens' school enrollment status as a regressor because of the likelihood that enrollment is endogenous. Teens who remain employed or work more hours may base their schooling choice on labor market outcomes.

[16]The effect of minimum wage increases on school enrollment is addressed more broadly by Neumark and Wascher (1995a and 1995b), whose research suggests that teens not initially employed leave school and join the labor force when the minimum wage increases.

substitute work for leisure more than unaffected teens, they do not appear to be more likely to substitute work for school.

5. Conclusion

The state-and individual-level teen employment results I find using the CPS are similar to previous findings. Changes in the minimum wage have no discernable effect on aggregate teen employment rates, and teens likely to be affected by minimum wage increases are, on average, about 2-3 percent less likely to remain employed than unaffected teens. The hours results, which are the new contribution of this paper, indicate that average teen hours rise when the minimum wage increases, and the hours of teens who remain employed and are likely to have been affected by minimum wage increase both relative to unaffected teens who remain employed and absolutely. Even when teens who do not remain employed are included in the sample, the hours of teens likely to be affected by minimum wage increase either rise or remain constant relative to unaffected teens in most specifications.

The aggregate employment and hours results indicate that labor demand for teen workers does not fall when the minimum wage is increased. This appears to contradict the traditional model of labor demand. However, individual-level data shows that teens more likely to be affected by minimum wage increases are indeed less likely to remain employed. Together, the aggregate- and individual-level employment results suggest that employers substitute among teens in response to minimum wage increases. These results accord with Neumark and Wascher (1995a), who find that some teens leave school and enter the labor market when the minimum wage increases. Further research using establishment data is needed to confirm that employers substitute among workers in response to minimum wage changes.

The aggregate- and individual-level hours results can also potentially be reconciled with the traditional model. A plausible explanation for the aggregate result is that high-skill teens who substitute work for school or leisure when the minimum wage rises are more likely to work full-time, and employers substitute these teens for lower-skill, part-time teens. Those teens who remain employed despite probably being affected by the minimum wage increase are likely to be the highest-skilled teens among the group of bound teens. These workers may also substitute work for leisure as the minimum wage increases, explaining the increase in their hours; they do not appear to substitute work for school. Research comparing the characteristics of teens who remain employed after minimum wage hikes and those who do not using individual-level data is a promising area for further work.

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Table 1
Aggregate Estimates of the Effects of the Minimum Wage

Covariate	Employment (1)	(2)	Hours (3)	(4)
Log of Real Minimum Wage	0.003 (0.114)	0.094 (0.143)	0.172 (0.062)	0.091 (0.078)
Lagged Log of Real Minimum Wage	--	-0.154 (0.153)	--	0.139 (0.083)
Unemployment Rate	-2.240 (0.186)	-2.231 (0.187)	-1.033 (0.121)	-1.040 (0.121)
Teen-to-Total Population Ratio	0.425 (0.488)	0.431 (0.488)	0.528 (0.316)	0.519 (0.083)

Notes: The regressions also contain a constant and a full set of state and year dummy variables. The data are annual averages for the 50 states over 1979-1993 for a total of 750 observations. The regressions are estimated using the Prais-Winston method to correct for serial correlation and heteroskedasticity. Standard errors are shown in parentheses.

Table 2
Descriptive Statistics of Individual Sample

	Total Sample	Bound Workers	Nonbound Workers
Number employed in first year and reporting wage data	55,000	--	--
Number after dropping those in uncovered industries or missing employment recode in second year	46,915	--	--
Proportion bound by increases in minimum wage	0.135	--	--
Average real wage gap if bound	$0.24 (0.10)	--	--
Proportion employed in second year	0.728	0.693	0.734
Average real hourly wage in first year	$3.98 (3.02)	$3.57 (0.51)	$4.05 (3.24)
Average change in real hourly wage among workers remaining employed	$0.20 (3.12)	$0.41 (1.31)	$0.17 (3.31)
Average hours per week in first year	24.89 (12.26)	22.49 (10.90)	25.26 (12.42)
Average change in hours per week among workers remaining employed	3.54 (11.57)	4.51 (11.90)	3.30 (11.51)
Percent female	0.48	0.52	0.47
Percent African-American	0.09	0.10	0.09
Percent Hispanic	0.05	0.05	0.05
Average age in first year	17.64	17.33	17.69
Percent completed 12th grade by first year	0.44	0.32	0.46

Notes: Data are based on matched individual records in CPS outgoing rotation groups for teens aged 16-19 for the years 1979-1993. All observations are unweighted. See text for definition of bound workers. Standard deviations are shown in parentheses.

Table 3
Individual Estimates of the Effects of the Minimum Wage

	Remain Employed		Change in Hours			
			Still Employed		All Teens	
	(1)	(2)	(3)	(4)	(5)	(6)
Wage gap	-0.110	-0.114	3.118	5.520	-0.488	2.440
	(0.028)	(0.031)	(0.839)	(0.926)	(0.994)	(1.116)
Age	0.015	0.015	0.231	0.240	-0.288	-0.280
	(0.003)	(0.003)	(0.090)	(0.090)	(0.114)	(0.114)
Male	0.006	0.007	-0.048	-0.041	-0.547	-0.490
	(0.004)	(0.005)	(0.136)	(0.136)	(0.168)	(0.168)
African-American	-0.136	-0.138	-0.642	-0.646	-4.406	-4.466
	(0.009)	(0.009)	(0.261)	(0.271)	(0.313)	(0.325)
Hispanic	-0.031	-0.035	-1.682	-1.683	-3.167	-3.234
	(0.010)	(0.011)	(0.284)	(0.313)	(0.382)	(0.409)
High school graduate	0.091	0.093	-3.175	-3.167	-0.805	-0.748
	(0.006)	(0.006)	(0.201)	(0.201)	(0.257)	(0.257)
State fixed effects	No	Yes	No	Yes	No	Yes
Year fixed effects	No	Yes	No	Yes	No	Yes
Adj. R-squared	0.026	0.029	0.018	0.022	0.010	0.015

Notes: The data are based on matched individual records in CPS outgoing rotation groups for teens aged 16-19 for the years 1979-1993. All regressions include a constant, and observations are weighted using the CPS demographic weights. See text for an explanation of the sample. Standard errors are shown in parentheses.

Table 4
Comparison of Matching and Nonmatching Individuals

	Matching	Nonmatching
Proportion bound by increases in minimum wage	0.135	0.124
Average real wage gap if bound	\$0.24	\$0.23
	(0.10)	(0.10)
Average real hourly wage in first year	\$3.98	\$4.13
	(3.02)	(4.02)
Average hours per week in first year	24.89	26.51
	(12.26)	(12.54)
Percent female	0.48	0.47
Percent African-American	0.09	0.10
Percent Hispanic	0.05	0.06
Average age in first year	17.64	17.96
Percent completed 12th grade by first year	0.44	0.53

Notes: Data are based on individual records in CPS outgoing rotation groups for teens aged 16-19 for the years 1979-1992. All observations are unweighted. See text for definition of bound workers and description of the matching procedure and sample. Standard deviations are shown in parentheses.

Table 5
Individual Estimates of the Enrollment Effect of the Minimum Wage

	(1)	(2)
Wage gap	-0.041	-0.017
	(0.029)	(0.032)
Enrolled first year	0.361	0.354
	(0.007)	(0.007)
Age	-0.091	-0.091
	(0.003)	(0.003)
Male	-0.018	-0.017
	(0.004)	(0.005)
African-American	-0.017	-0.013
	(0.009)	(0.009)
Hispanic	-0.046	-0.053
	(0.009)	(0.010)
High school graduate	0.080	0.081
	(0.006)	(0.006)
Month fixed effects	Yes	Yes
State fixed effects	No	Yes
Year fixed effects	No	Yes
Adj. R-squared	0.298	0.301

Notes: The dependent variable is one if a teen is enrolled in school and zero otherwise. The data are based on matched individual records in CPS outgoing rotation groups for teens aged 16-19 for the years 1979-1993. All regressions include a constant, and observations are weighted using the CPS demographic weights. See text for an explanation of the sample. Standard errors are shown in parentheses.