NBER WORKING PAPERS SERIES

EVIDENCE ON EMPLOYMENT EFFECTS OF MINIMUM WAGES AND SUBMINIMUM WAGE PROVISIONS FROM PANEL DATA ON STATE MINIMUM WAGE LAWS

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Working Paper No. 3859

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
October 1991

September 1991. We gratefully acknowledge helpful comments from Charles Brown, Daniel Hamermesh, Alan Krueger, Paul Taubman, Wayne Vroman, and participants at the NBER Labor Studies Summer Institute, and research assistance from Donna Boswell and Richard Johnson. We thank Wayne Vroman for providing minimum wage data for Washington, D.C. This paper is part of NBER's research program in Labor Studies. Any opinions expressed are those of the authors and not those of the National Bureau of Economic Research, the Federal Reserve Board or its staff.

NBER Working Paper #3859
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ABSTRACT

We construct a panel data set on state-level minimum wage laws and economic conditions to reevaluate existing evidence on minimum wage effects on employment, most of which comes from time-series data. Our estimates of the elasticities of teen and young-adult employment-to-population ratios fall primarily in the range -0.1 to -0.2, similar to the consensus range of estimates from time-series studies. We also find evidence that youth subminimum wage provisions enacted by state legislatures have moderated the disemployment effects of minimum wages.

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I. Introduction

A federal minimum wage was first implemented in the United States with the passage of the Fair Labor Standards Act of 1938. Although the proportion of workers covered by the minimum was originally below 50 percent, coverage has expanded to now include more than 90 percent of all workers. Since its enactment, there has been widespread debate about the merits of minimum wage laws, together with numerous efforts to evaluate their economic effects. By the early 1980s, the large body of theoretical and empirical research by economists, including that of the Minimum Wage Study Commission (1981), suggested a rare consensus; the imposition of minimum wages decreases employment opportunities for workers with wages at or near the minimum wage. More explicitly, a 10 percent increase in the minimum wage apparently reduced teenage employment by 1 to 3 percent, with proportionately smaller effects for 20-to-24 year-olds, reflecting their smaller representation in the minimum wage population (Brown, et al., 1982, and Brown, 1988).

Two recent papers have challenged the findings from the earlier time-series studies. Wellington (1991) argues that the addition to the sample period of the 1980s, during which the minimum wage fell sharply in real terms, has the effect of reducing the estimated disemployment effects. For an hypothesized 10 percent rise in the minimum wage she reports disemployment effects of less than 1 percent for teenagers and essentially zero for young adults. Card (1991) compares the employment experiences of Californian workers during the late 1980s, when the state minimum wage level rose sharply, with the experiences of similar workers in labor markets with no changes in minimum wage laws. He finds no evidence of a disemployment effect from the higher minimum wage, which he interprets as casting some doubt on the conventional competitive model of the low-wage labor market.

In this paper, we reevaluate the minimum wage effects reported in the existing literature by considering the influences of state minimum wage laws along with federal legislation. Aside from federal standards, individual states have frequently legislated changes in the level or coverage of the minimum wage in their specific jurisdictions. States cannot enact a minimum wage lower than the federal minimum (for workers covered by the federal minimum). But many states have periodically raised their minimum wage above the federal level (or extended coverage to workers excluded from federal legislation). In addition, states with minimum wages above the federal level have sometimes implemented exemptions from the higher state level for specific subgroups of the labor force, such as teenagers or students.

Despite the widespread existence of state minimum wage laws, empirical studies of the effects of minimum wages on the employment status of teens and young adults have largely ignored the state statutes. Instead, existing research is based almost exclusively on time-series data on the federal minimum wage, on workers covered by the federal minimum, and on the aggregate (i.e., national) labor market. The authors of these time-series minimum wage studies often bemoan the inadequacy of their data for a serious empirical analysis of minimum wage effects (e.g., Wachter and Kim, 1978). Yet out of the vast number of minimum wage studies conducted for the U.S., only about ten do not resort to time-series data.

In this study, we augment the national data with information on geographical differences in the minimum wage. In particular, we construct a panel data set on minimum wage laws and labor market conditions at the state level and use the data set to reevaluate the existing evidence on the consequences of minimum wages for youth employment and unemployment. In our view, the integration of state data into the analysis of minimum wage effects expands upon the existing literature in three directions.

First, the explicit consideration of state minimum wage laws in our construction of a minimum wage variable eliminates a source of measurement error manifest in earlier studies that used only the federal minimum wage. A few states nearly always set minimum wage levels above the federal level, and in some years the percentage of states with higher minimums has been as high as 25 percent. This source of measurement error may be especially worrisome for new studies estimating minimum wage effects using data for the 1980s, because the number of states with minimum wage levels above the federal level rose significantly in this period.

Second, the use of panel data addresses an important criticism leveled at the small number of previous studies using single cross sections. Much of the variation in the minimum wage variables used in these studies arose from variation in average wage levels across states, leading critics to argue that the estimated wage effects largely reflected state "average wage" effects rather than minimum wage differences (Brown, et al., 1982). In contrast, our use of panel data permits explicit estimation of state (or year) effects as distinct from the effects of changes in the minimum wage variable, and so permits a cleaner evaluation of minimum wage effects from the cross-sectional (and time-series) variation in the data.[1]

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[1]A small subset of the existing studies use cross-sectional data, sometimes exploiting information on state minimum wage laws. Most of these studies focus on a single cross-section and include either a dummy variable for the existence of a minimum wage law (Katz, 1973; Freeman, 1982) or data on state levels and coverage (Cotterill and Wadycki, 1976; Schiller, 1991; Welch and Cunningham, 1978). Only two studies for the United States and one for Canada use repeated observation on states to remove the influence of state effects on cross-sectional estimates. For the U.S., Cunningham (1981) and Cogan (1981) use states as the unit of observation, and attempt to identify minimum wage effects from changes in employment rates across decennial Censuses; neither of these studies uses information on youth or student subminimums, and only the

Third, the use of state data permits a direct evaluation of the effects of lower minimum wage levels for subgroups of the population. The latest federal minimum wage legislation--effective in April 1990--attempts to alleviate the disemployment effects on youths by introducing, for the first time at the federal level, a "subminimum" wage for newly employed workers. Because a broad federal youth subminimum has never before been implemented, however, aggregate time-series data for the U.S. economy cannot be used to evaluate the effectiveness of a subminimum in mitigating the adverse effects of the minimum wage on youth employment.[2] In contrast, numerous states have implemented subminimums for young workers or students in years past, so that data on state legislation can be used to examine the effects of subminimum wage provisions. Based on this past experience at the state level, predictions can be made regarding the consequences of the recently enacted federal subminimum.

To summarize our findings, our estimates indicate disemployment effects of

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Cunningham paper uses data on state minimum wage levels. Lester (1946) provides a similar experiment, comparing employment growth in particular industries in the North and the South following the implementation of the FLSA, which had a larger impact in the lower-wage South.

Swidlinsky (1980), studying minimum wage effects in Canada, uses an approach most similar to ours, using data for five regions over a twenty-year period. He reports a disemployment elasticity of -0.17 for Canadian teenagers. Swidlinsky also notes that youth and student subminimum wages vary across regions, but does not carry out a direct analysis of the moderating effects of subminimum wages.

[2]Past legislation sometimes permitted some classes of employers (primarily colleges and universities) to pay subminimum wages to full-time students, but had never before been generalized to all young or new workers. In a study for the Minimum Wage Study Commission, Freeman, et al. (1981) analyzed this specific provision using a data set constructed from records of private-sector employers certified to hire full-time students. Their main focus, however, is on the micro-level variables associated with the employment of workers under this subminimum wage. Brown, et al. (1983) investigated the effect of adjusting the Kaitz index of the minimum wage for coverage exemptions for students on time-series estimates of minimum wage effects, and found no discernible impact.

minimum wages similar to the consensus range of estimates from existing time-series studies (Brown, et al., 1981). In particular, the disemployment elasticities we estimate are concentrated in the range from -0.1 to -0.2. In addition, we find evidence that youth subminimum wage provisions enacted by state legislatures have moderated the disemployment effects of minimum wages.

II. The State Minimum Wage Panel Data Set

To better assess the effects of minimum wage laws on youth employment, we have constructed a panel data set of minimum wages, minimum wage coverage, and local economic conditions, covering the 50 states and the District of Columbia for the years 1977 through 1989, and extending back to 1973 for a subset of 22 larger states for which the CPS identified state of residence from 1973-1976. Reflecting limitations in the availability of state-specific data, the data set is compiled at an annual frequency.

Minimum wage laws

To assemble the state minimum wage levels, we obtained from each state labor department a chronology of applicable minimum wage legislation dating back to the early 1970s. These data were cross-checked against information available from the Bureau of National Affairs' Compensation Primer and the chronologies constructed by Quester (1981).[3] In most cases, this procedure yielded a single time series of the minimum wage level, from which we extracted the value in effect during May of each year (for consistency

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[3]Our principal focus is on legislated minimum wage levels, although state laws extending minimum wage coverage, perhaps at a wage lower than the federal minimum, have often been more prevalent than laws boosting the minimum above the federal level (Cox and Oaxaca, 1981). However, as explained below, estimates of coverage by state laws are hard to come by. In addition, in the few time-series studies that consider the effects of coverage separately from the effects of (relative) minimum wage levels, coverage effects tend to be weaker (Brown, et al., 1982).

with the data on state labor market conditions described below). For a few states, however, the state minimum wage level differed by occupation or labor force group, and additional steps were necessary to obtain a single value that best captured the effective state law. Where the varying levels clearly were subminimums for youths, students, newly covered workers, or very low-skilled occupations (e.g., housekeeping in North Dakota), we used the highest value at each point in time and documented the existence (and type) of a subminimum wage for use as a separate variable. In two other cases, the existence of multiple minimum wage levels was not automatically suggestive of a subminimum. In the District of Columbia there are nine separate minimum wage levels for different industries and occupations, as well as differing minimum wage levels for youths, students, and JTPA workers. In this case, we used the weighted average of the minimum wage levels for adults across the nine categories, weighted by estimated employment in each category in each year.[4] In Minnesota in recent years, the state minimum wage level for workers covered by the Federal law (FLSA) differed from the level for those covered only by state law; in this case, we used the state minimum for workers covered by the FLSA.[5]

Table 1, column (1) shows the states with legislated minimum wage levels above the federal level for each year in our sample, along with the legislated levels.[6] Throughout the

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[4]We are grateful to Wayne Vroman for supplying these estimates.

[5]This is relevant in Minnesota because the state minimum wage levels for both groups were higher than the Federal minimum wage in 1988 and 1989. New legislation effective in 1991 eliminated this two-tier schedule, although the new structure sets a higher minimum wage for large employers than for small employers (who are less likely to be covered by the FLSA).

[6]For some years the number of states with minimum wages above the federal level differs from numbers reported in Vroman (1991). This occurs because we use minimum wage levels as of May of each year, while Vroman uses annual averages.

1970s and much of the 1980s, only a few states set a minimum wage above the federal level.[7] However, in the late 1980s, with legislation to adjust the minimum wage absent from the federal agenda, concerns about the ongoing decline in the real value of the minimum wage led additional states to enact new legislation to raise their minimum above the federal level, and the number of states with higher minimums rose to 13 by 1989.[8] Columns (2) and (3) report the federal minimum wage level for each year, and the (unweighted) average percentage difference between the state and federal minimum wage level, for states with legislated minimum wage levels exceeding the federal level. Perhaps not surprisingly, the average percentage differential is greatest when the number of states with minimums exceeding the federal minimum is largest, rising to a peak of 16 percent in 1989.

In addition to minimum wage levels, the chronologies collected for each state included information on the existence of subminimum wage provisions that permit employers to pay a lower wage to specific subgroups of the labor force. In contrast to federal legislation, which only very recently began to experiment with a subminimum wage, many states have included subminimums for years. Generally, subminimum wage provisions enacted by state legislatures in the past have taken two forms: a subminimum (or

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[7]Legislation in both Connecticut (beginning in 1974) and Alaska (beginning in 1977) automatically keeps the state minimum wage above the federal level. In Alaska, a constant differential of 50 cents per hour is maintained. In Connecticut, the law through 1987 set the state minimum 1/2 percent above the federal level, resulting in a differential of just a few cents.

[8]With the increase in the Federal minimum to $3.80 per hour in 1990 and $4.25 in 1991, the number of states with higher minimum levels has again dropped, to five in 1991.

exemption) based on age; or one based on student classification.[9] As shown in Table 2, columns (1) and (4), all or nearly all states with the minimum wage above the federal level had subminimum wage provisions in their minimum wage legislation in the 1970s and early 1980s. And even as the number of states with high minimum wage statutes rose in the second half of the 1980s, the percentage with subminimum provisions fell off only slightly, to roughly 80 percent. Columns (2) and (5) show that, over the sample period, about half of all states have had youth subminimums, and two-thirds of all states have had student subminimums. Finally, columns (3) and (6) report the number of states that either enacted or repealed a subminimum wage provision in each year; the variation induced by these changes is essential for our estimation of the effects of subminimums, in the presence of fixed state effects.

Comprehensive time-series information on state minimum wage coverage was more difficult to assemble. For the federal law, the Department of Labor has published estimates of the number of wage and salary workers in each state by their coverage status under the minimum wage provisions of the FLSA for most years in our sample. Estimates of coverage are complicated, however, by the treatment of state and local government workers. Prior to 1976, virtually all state and local government employees were covered by the FLSA as a result of the 1966 and 1974 amendments to the Act. In 1976, however, the Supreme Court ruled that the minimum wage and maximum hours provisions of the FLSA could not be constitutionally applied to state and local government activities that are an integral part

[9]A few states in the past have legislated specific minimum wage levels for women. However, these typically were not subminimums but rather existed in the absence of other minimum wage legislation covering men. The last of these separate laws covering women disappeared in 1990 when Utah enacted more general minimum wage legislation.

of traditional government functions (National League of Cities v. Usery). That ruling was reversed in 1985 (Garcia v. San Antonio Transit Authority) so that federal minimum wage laws again became applicable to virtually all state and local government employees. As a result, federal coverage was depressed by the absence of state and local workers from 1977 to 1984, but picked up these workers again in 1985.

For coverage by state laws (above and beyond FLSA coverage), data are available from the U.S. Department of Labor only for the years 1974, 1975, and 1977. Because we currently have no way to capture changes over the remainder of our sample period in coverage by state laws, we simply use the FLSA coverage estimates for each state for the available years.[10] For years with no official estimates (1979-81), we assumed that federal coverage on a state-by-state basis changed in proportion to the change in coverage for the United States as a whole. For the years after 1986 (the latest data available), we assumed that coverage rates held steady at their 1986 level. Because FLSA coverage differs across states, this approach seems preferable to using just data on minimum wage levels, and should help to reduce measurement error in the effective minimum wage.

Using the data on federal minimum wage levels, state minimum wage levels, and coverage, we then construct a coverage-adjusted minimum wage. For each state-year observation, this variable is the product of the greater of the federal or state minimum wage, and federal coverage for the state. To complete the formulation of the minimum wage variable, we computed the ratio of the coverage-adjusted minimum wage prevailing in May of each year to the average hourly wage in each state during May of that year; the

[10]Later in the paper, we report results for our final specifications using estimates of coverage by state minimum wage laws, as well as federal minimum wage laws. The estimates based on these coverage estimates are substantively the same as those based on federal coverage.

resulting variable is the coverage-adjusted relative minimum wage used in the analyses reported below.[11] Table 1, columns (4) and (5) use these coverage-adjusted relative minimum wages to provide more information on the role of state minimum wage laws in influencing effective minimum wages. These columns report the average coverage-adjusted relative minimum wage variable separately for states with minimum wage levels exceeding the federal level, and states in which the federal minimum wage level is binding.

A comparison of columns (4) and (5) reveals that, in our sample period, minimum wage levels higher than the federal minimum wage level generally did not result in higher relative minimum wages; indeed, throughout most of the sample period, relative minimum wages were higher in states without minimum wage levels exceeding the federal level, reflecting the lower average market wage in those states. Over the 1980s, however, the average relative minimum wage in states in which the federal minimum wage is binding declines, and by the end of the sample period, the average relative minimum wage is roughly the same for both sets of states. Thus, when evaluated in terms of changes, the rising incidence of state minimum wage laws did boost relative minimum wages during the

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[11]Because state-specific wage rates are not published outside of the manufacturing industry, the average state wages used are estimated as the mean usual hourly wage from the May Current Population Surveys (CPS). We chose the May CPS, because prior to 1983, the questions pertaining to usual weekly hours and earnings were only asked in May. After 1982, the questions were asked in every month, but only of one-fourth of that month's CPS sample. For all data computed from the CPS, we deleted persons under age 16, self-employed workers, unpaid family workers, and persons indicating agricultural production or agricultural services as their current or most recent detailed industry classification.

The ratio of the coverage-adjusted minimum wage to the average wage for the age group studied may be more informative as to how much the minimum wage cuts into the wage distribution. However, for many states the cell sizes from which we can compute mean wages for teenagers and young adults are quite small (especially after 1982, when wage information was elicited from only one-fourth of the sample in each month).

1980s.

State-level economic data

We constructed data on state labor market conditions over the same period, for consistency also using the May files of the Current Population Survey (CPS). Variables estimated from the CPS include: employment rates for teens (16-19) and young adults (aged 16-24); unemployment rates for prime-age (25-64) males; proportions of the population aged 16-19 or 16-24; and the proportions of individuals aged 16-19 or 16-24 enrolled in school. In all cases, the variables are calculated from the individual survey responses, aggregated to the state level using the CPS demographic weights.[12]

III. Reconsidering the Existing Evidence on Minimum Wage Effects

As noted above, the overwhelming majority of past minimum wage studies have focused on time-series data at the national level. Accordingly, we begin our investigation by fashioning a norm with which to compare the panel data results presented below. In particular, the typical time-series study estimates a regression equation of the form:

(1) $E_t = \alpha_0 + \alpha_1 MW_t + X_t\beta + \epsilon_t$.

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[12]Madden (1991) describes four adjustments to the weighting of households in the CPS in the 1980s; the largest of these was in January 1985, when data from the 1980 U.S. Census were incorporated. Madden reports that only the January 1985 reweighting resulted in substantial changes, shifting the CPS towards more prosperous households (presumably reflecting the shift of the population from urban to suburban areas). While such reweighting is relevant to the estimation of means from the data, we are estimating regression relationships in which most of the variables (employment-to-population ratios, mean wages, unemployment rates, etc.) should be similarly affected. Thus, a priori, reweighting may not be important in our results. Furthermore, the individual-year dummy variables that we include should capture the effects of reweighting on the intercept.

E is the employment-to-population ratio for the age group under study,[13] MW is a coverage-adjusted minimum wage level, and X is a set of control variables to capture aggregate business-cycle effects, the changing age structure of the population, and in some specifications, school enrollment rates; the "t" subscript indicates the year or quarter to which the data apply. Although economists think of employment and wages as being determined by the interaction of labor supply and labor demand, this equation is assumed to represent a simpler analysis in which the effects of exogenous variables on equilibrium employment or unemployment are estimated (i.e., equation (1) is the reduced form).[14]

Table 3 presents estimates of this "typical" time-series model, using annual data from 1955 to 1989 as well as for the 1973-1989 subperiod covered by the data in our state panel data set. The model is estimated in first-difference form, and we use annual data to enhance comparability with the results from the panel data.[15] For teenagers, the estimated employment elasticity from the minimum wage is roughly -0.1 over the entire sample (column (1)), similar to that found by Wellington (1991) and at the bottom of the consensus range suggested by Brown, et al. (1982). When the minimum wage effect is permitted to differ after 1972 but the entire sample is used to estimate the equation (column (2)), the

[13]We focus on the employment-to-population ratio, rather than the unemployment rate, because as pointed out by Mincer (1976), the effects of minimum wages on unemployment rates are ambiguous.

[14]As Brown, et al. (1982) point out, a simple supply and demand model with homogeneous workers implies that in the presence of a (binding) minimum wage, employment is demand determined. Supply variables become important because many workers, including teenagers, earn more than the minimum wage. Thus, overall employment will depend on supply as well as demand variables.

[15]Estimates of the model in level form were suggestive of a unit root in the dependent variable. In addition, the minimum wage variable is lagged one year; the contemporaneous value of the minimum wage variable and longer lags entered with statistically insignificant coefficients.

minimum wage effect in the latter period falls by half, although the difference between the two periods is not statistically significant. Similarly, the minimum wage elasticity is roughly -0.06 when the sample is restricted to the 1973-1989 subperiod, again consistent with Wellington's results. For young adults (aged 16 to 24), the estimated elasticities are smaller than for teenagers alone, over the entire sample period (column (4)). However, there is little change in the estimated elasticity for the 1973-1989 subperiod.

With the time-series results in mind, the first component of our research entails a reevaluation of the evidence on the consequences of minimum wages for youth employment. In particular, we view the existing research using time-series data as potentially flawed for three important reasons. First, the federal minimum wage variable typically used in time-series studies exhibits relatively little variation over time. Indeed, one research paper characterizes the coverage-adjusted time-series minimum wage variable as "largely a spike in 1967" (Wachter and Kim, 1978). Second, some workers have been covered by state minimums set higher than the federal level, so that the federal minimum wage variable used in past time-series studies measures the effective minimum wage with error.[16] Third, minimum wages have risen over time concurrently with the expansion of government social welfare programs. These programs may have had independent effects on

[16]This problem may be particularly severe in time-series estimates that include the latter part of the 1980s, such as those in Table 3, when many states had minimum wage levels above the federal level, and is a potential explanation for the results in Wellington (1991). She compares time-series estimates of minimum wage effects including and excluding the 1980s (through 1986), and concludes that estimates based on data including the 1980s are lower. However, a minimum wage series based solely on the federal minimum overstates the decline in the minimum wage in the 1980s, which would be expected to bias the minimum wage coefficient towards zero, given that teen and young adult employment rates increased over much of this period. This problem would likely be more severe were Wellington's study extended through 1989, by which time more states had raised their minimum wages.

teen and youth employment or unemployment, and collinearity between increases in the minimum wage and social welfare programs makes it difficult to isolate the effects of minimum wages.

In this subsection of the paper, we use our data set to reestimate the impact of minimum wages on employment of teens and young adults, using the specifications standard in the existing literature. In contrast to the existing time-series studies, the data on state minimum wage laws and state-level economic conditions permit a pooled time-series cross-section analysis of minimum wage effects, exploiting the greater variation in relative minimum wages at the state level. Moreover, the measurement error caused by using a uniform federal minimum wage is avoided. Finally, because state minimum wages vary across states in a manner that is relatively more independent of the growth of social welfare programs, this analysis should yield more reliable estimates of the effects of minimum wages on employment and unemployment of teens and young adults.

Specifically, our panel data set permits us to estimate an equation of the form

(2) $E_{it} = \alpha_0 + \alpha_1 MW_{it} + X_{it}\beta + Y_t\gamma + S_i\delta + \epsilon_{it}$,

where i indexes states and t indexes years. Y_t is a set of year effects, and S_i is a set of state effects. For the analysis, 751 observations are available (data for the 50 states and Washington, D.C., multiplied by the 13 years for which the data are available, plus an additional four years of data for the 22 larger states identified in the CPS from 1973-1976).

As the equation specification indicates, the panel data permit an examination of the presence of year or state effects in the regression models. One advantage of this approach is that it addresses a major criticism of studies that use purely cross-sectional data: that the "variation across regions may reflect regional differences in "competitiveness"--the performance of one area versus another--that provide little insight into the possible causes of aggregate problems" (Freeman, 1982, p. 115). That is, individual state effects may be present corresponding to unmeasured economic conditions of state economies, which give rise to persistently tight labor markets and high wages in particular states. In equation (2), this generates a negative correlation between the minimum wage variable (MW) and the employment-to-population ratio (E), and hence an estimate of the disemployment effect that is too large in absolute value. The use of panel data, accounting for the presence of state effects, can resolve this problem, while still exploiting geographic variation. An additional advantage of the panel data set is that we can control for the possibility of year effects in the data, corresponding perhaps to business-cycle or cohort-size effects that are not captured in the variables usually included in minimum wage studies; this problem is insurmountable in a time-series study.

Tables 4 and 5 reports results from alternative specifications of the pooled time-series cross-section standard minimum wage model. Table 4 reports specifications with different combinations of fixed state and year effects for teenagers (16-19), while Table 5 repeats the same analysis for young adults (16-24). In each case, estimates of equation (2) are reported excluding and including the proportion of the age group enrolled in school as an independent variable. Because school and work represent alternative opportunities for many young persons, school enrollment rates are potentially endogenous. This endogeneity presumably imparts a negative bias to the coefficient of the school enrollment variable because factors associated with high employment rates lead to low enrollment rates. The bias transmitted to the coefficient of the minimum wage variable is ambiguous a priori. High minimum wages may lead young persons to stay in school, either reflecting worsened employment opportunities or because schooling increases the probability of employment in the covered sector (Leighton and Mincer, 1981).[17] Alternatively, if job search is more difficult while in school, high minimum wages may reduce enrollment as young persons leave school to queue for minimum wage jobs (Mincer, 1976).[18] Unfortunately, we do not believe that there are valid exclusion restrictions on the basis of which to instrument for the enrollment variable, so instead, throughout the paper we report both sets of results.

Columns (1) and (2) of Tables 4 and 5 report estimates from specifications using fixed state and year effects. In the specifications excluding the school enrollment rate, the estimated minimum wage effect is slightly positive for teenagers, and negative for young adults, and is marginally significant only for young adults. In the specifications including the school enrollment rate, the disemployment effects are negative and statistically significant for both teenagers and young adults, with a larger disemployment effect for teenagers. The corresponding elasticities of the employment-to-population ratio with respect to the minimum wage variable, evaluated at the sample means, are reported in the last row of each table.

The remaining columns of Tables 4 and 5 report estimates from specifications excluding year effects (columns (3) and (4)), excluding state effects (columns (5) and (6)), and excluding both year and state effects (columns (7) and (8)). The estimated minimum wage effects vary considerably as different fixed effects are excluded. The elasticities are more positive when year effects are dropped, compared to columns (1) and (2), and more negative when state effects or state effects and year effects are dropped. The differences

[17]Mattila (1978) finds that, over time, minimum wages are positively associated with enrollment rates, consistent with this view.

[18]This negative relationship could also arise if legislators seek higher minimum wages when a larger proportion of young persons is out of school and looking for work.

between the estimated elasticities with and without the fixed state effects seem to confirm the suspicions of critics of earlier cross-section studies. Omitting state effects apparently imparts a negative bias to the minimum wage estimates, suggesting that unmeasured local economic conditions can complicate the estimation of minimum wage effects in such studies. The other coefficients also change when different fixed effects are excluded, most noticeably the estimated effect of the prime-age male unemployment rate. These changes in coefficients, in addition to the rejection of the restrictions imposed by dropping either the state or year effects, leads us to select the specification with fixed state and year effects as our best specification.[19]

More generally, minimum wage elasticities from our fixed-effects models that include the school enrollment variable are somewhat larger than the time-series elasticities presented in Table 3 and are broadly consistent with the evidence presented by Brown, et al. (1982), if perhaps toward the lower end of their consensus range.[20] However, the positive estimated minimum wage elasticities from the fixed-effects models that exclude the

[19]In results not reported in the tables, we checked the validity of the linear specification of our fixed year- and state-effects specification. We set up a grid search over a specification of the model with a Box-Cox transformation of the dependent and independent variables, with the same transformation applied to all variables; the Box-Cox parameter ranged from 0 to 1 in increments of 0.1. For all four alternatives (teenagers and young adults, with and without the school enrollment rate), the likelihood was monotonically increasing as the Box-Cox parameter increased. This validates the linear specification, and in particular rules out a double-log specification. (A double-log specification seems dubious on a priori grounds, since it implies that, for example, an increase in the minimum wage from 50 cents to one dollar has the same percentage effect on employment as an increase from two dollars to four dollars.)

[20]Al-Salam, et al. (1981) report results for male teenagers for specifications including and excluding three "potentially endogenous" variables: the proportion enrolled in school; the proportion in the military; and the proportion in federal youth programs. They obtain negative minimum wage effects in both cases, but the effects are <u>smaller</u> when these three variables are included.

school enrollment rate are unusual. In the Brown, et al. survey of employment elasticities, only one of the eighteen estimates for teenagers was positive (Iden, 1980).

Fixed-effects vs. first-difference estimates

Our results also contrast with those reported by Card (1991), in his study of the rise in California's minimum wage between 1987 and 1989. Specifically, he found a positive contemporaneous correlation between changes in the minimum wage and in the employment of teens and young adults, which he interprets as evidence against the standard competitive model of the low-wage labor market. There are basically two key differences between Card's specification and the estimates reported in Tables 4 and 5: first, Card focuses on California whereas we use data on all states; second, Card uses a first-difference form of the model rather than the fixed-effects specification used in this paper. Thus, to determine whether Card's results hold more generally at the state level, we also examined first-difference estimates of our four specifications; these are presented in Table 6, both including and excluding year effects.

The estimated minimum wage effects when year effects are included--reported in columns (1), (2), (5) and (6)--are strikingly different from the fixed-state-effects estimates reported in Tables 4 and 5, and strikingly similar to the results reported by Card for California. For both teenagers and young adults, there is a statistically significant positive effect of minimum wages on the employment-to-population ratio in the specifications excluding school enrollment. A positive effect is also estimated for the specifications including school enrollment, although these coefficients are small and not statistically significant. Larger positive elasticities result when the year effects are excluded; these changes and the p-values from the likelihood-ratio tests indicate, however, that the year-effects specifications are preferred. The estimates in Table 6 imply that Card's (1991) finding of a positive contemporaneous correlation between minimum wage changes and employment changes at the state level is not unique to California's increase in the minimum wage between 1987 and 1989, but is a more general finding applicable to the U.S. as a whole.

Dynamic specifications

One plausible explanation of the differences between the fixed-effects and first-difference estimates is that the basic model is misspecified by ignoring lags in the effects of minimum wage changes. The first-difference estimates reflect only the contemporaneous relationship between minimum wage changes and changes in the employment-to-population ratio. The fixed-effects estimates, on the other hand, define changes relative to the state-specific mean, and hence are relatively more weighted towards the relationship between long-run changes in these variables. Thus, the existence of lagged negative minimum wage effects, coupled perhaps with a positive contemporaneous correlation, could account for the difference between the fixed-effects and first-difference estimates.[21]

In Table 7, we report estimates of dynamic specifications of minimum wage effects using both fixed-effects and first-difference estimators. The specifications for which coefficient estimates are reported include the contemporaneous minimum wage variable, as well as the variable lagged once. The estimates of the minimum wage coefficients reveal evidence of significant lags in these effects. In all eight columns the coefficient on the lagged minimum wage variable is more negative than that on the contemporaneous

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[21]Brown, et al. (1982), discuss the arguments for and against the likely existence of significant lags in minimum wage effects. Lagged effects may arise for the standard reasons, either because of hiring and training costs, or inability to adjust other inputs quickly (Nadiri and Rosen, 1969). But strong lags in minimum wage effects are perhaps less likely because of high turnover among these workers, and because minimum wage changes are enacted some time before they actually take effect.

minimum wage variable, and in the young-adult specifications, the lagged coefficient is strongly statistically significant. The importance of the lags is especially pronounced in the first-difference estimates. For three of the four first-difference specifications, the positive contemporaneous correlation persists, although with statistically insignificant coefficients. For all four first-difference specifications, however, there are negative lagged effects and overall negative elasticities. The marked changes in the estimated employment elasticities in these specifications are consistent with the first-difference estimates being more heavily influenced by contemporaneous changes in the minimum wage variable than are the fixed-effects estimates.

The table also reports results of specification tests and robustness checks for these dynamic specifications. The p-values from likelihood-ratio tests indicate that the data reject the specification with the contemporaneous minimum wage variable alone for all of the young-adult models. For teenagers, the p-values range from 0.12 to 0.39, which does not indicate rejection. But in results not reported in the table, we compared the likelihoods for the non-nested models including only the contemporaneous minimum wage variable, or alternatively only the lagged variable. The likelihood was higher for the model with lags in the three cases in which the contemporaneous specification indicated a positive elasticity (the fixed-effects specification excluding school enrollment, and the two first-difference specifications). The last two rows of the table report elasticities for the specification including only the contemporaneous minimum wage variable, and only the lagged minimum wage variable. (The elasticities from the contemporaneous specification differ slightly from Tables 4 and 5 because here they are computed from the same sample for which the lagged model can be estimated.) These elasticities tell much the same story; once lagged effects are allowed, negative elasticities results for all specifications.

The identification of these lagged effects is itself an interesting result.[22] But three additional findings emerge from these estimates. First, the long-run elasticities from the fixed-effects and first-difference estimates are considerably closer, although a relatively large difference still persists for teenagers in the specification excluding school enrollment. Second, the fixed-effects estimator provides more efficient estimates of the minimum wage effect. The standard errors of the long-run elasticities are considerably larger for the first-difference estimates than for the fixed-effects estimates, despite quite similar point estimates.[23] Together, these results argue for preferring the fixed-effects estimates.

Finally, the fixed-effects estimates of the elasticities of the employment-to-population ratio are larger (more negative) than in any of the previous fixed-effects (or first-difference) estimates. The elasticities are bunched in a range from -0.15 to -0.2; the single exception is the specification for teenagers excluding the school enrollment rate. These estimates are close to the midrange of the consensus of past time-series studies, and larger than our time-series estimates based on more recent data.

Robustness checks

In Table 8 we report the elasticities estimated from a number of additional fixed-

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[22]We identified two time-series studies that reported coefficients of contemporaneous and lagged minimum wage variables. Adie (1973) finds that in monthly data, effects lagged two years increase unemployment rates of teenagers more than do immediate effects. Betsey and Dunson (1981), using quarterly data, find substantial lagged effects of up to two years on employment rates of teenagers, with a pronounced effect at two years for non-whites (using quadratic Almon lag specifications). On the other hand, Brown, et al. (1983) find no evidence of lagged minimum wage effects in quarterly data for teen employment.

[23]These relatively larger standard errors are not surprising, since the differences for the first-difference estimator are computed relative to each year's data, while for the fixed-effects estimator the differences are computed relative to the state's mean for all years.

effects specifications, partly as a robustness check, and partly to provide a summary of findings from some of the specifications considered in our exploration of the data. In panels A-C we consider some alternatives to our treatment of minimum wage coverage. In panel A we report estimates of the specifications in Table 7 using no coverage adjustment; the minimum wage variable is simply the minimum wage level divided by the mean wage in the state. Some existing evidence suggests that this unadjusted minimum wage variable and the coverage rate may have different effects (Brown, et al., 1983; Gramlich, 1976). However, we are more interested in a sensitivity analysis of our estimated minimum wage effects, given the possibility of considerable measurement error in the coverage estimates. The estimated elasticities are slightly smaller in absolute value without the coverage adjustment, compared to the elasticities in Table 7; this makes sense since a one-unit increase in the coverage-adjusted minimum wage variable corresponds to a larger increase in the minimum wage level than does a one-unit increase in the unadjusted minimum wage variable. In panel B, we attempt to adjust for coverage by using state minimum wage laws along with the FLSA provisions.[24] In this case, the evidence of negative elasticities for

[24]We used the 1974 estimates from the Department of Labor for 1973, and interpolated the 1975 and 1977 estimates for 1976. As mentioned in the data section, the last available estimates of coverage by state minimum wage laws are for 1977. These figures were extrapolated as constant percentage equal to the value in 1977 for subsequent years. Consistent with the treatment of state and local government workers in the federal coverage estimates, for states that explicitly covered these workers we assumed that state coverage was equal to its 1977 value from 1978-1984, and its 1976 value (interpolated from 1975 and 1977 Department of Labor estimates) for 1985-1989. For other states, we simply used the 1977 value throughout the sample period. Colorado and Kansas did not enact a state minimum wage until 1978, after the last available estimate of state coverage; we currently use state coverage estimates of zero for these states. In cases for which the sum of the federal coverage estimate and the state coverage estimate exceeded unity, we adjusted the state coverage estimate downward to constrain the two to sum to unity.

Given our coverage estimates, the numerator of the relative minimum wage variable is defined as follows: for observations with minimum wage levels greater than teenagers disappears. As discussed above, however, we have very little hard evidence on coverage by state minimum wage laws, above and beyond the FLSA, and therefore discount these estimates. In panel C we examine the extent to which our use of state minimum wage levels, in contrast to the federal minimum wage levels used in time-series studies, influences the results, by substituting the federal minimum wage level for the state level. The estimated elasticities differ little from those using the state minimum wage levels. The implication of this is that Wellington's (1991) and our finding of diminished minimum wage effects in recent time-series data does not appear to be the result of a greater understatement of the effective minimum wage in the 1980s associated with the use of the federal minimum wage level.

In panel D we reestimate the fixed-effects specifications from Table 7 introducing one lag of each of the control variables, to ask whether the lagged minimum wage effect is simply picking up lagged effects of the other variables. Again, the elasticities are little changed. Finally, in panel E we reestimate the same models using state population estimates from the CPS to weight the observations. Again, there is relatively little change in the estimates, although there is some widening of the range of elasticities.

IV. Evidence on Student and Youth Subminimum Wage Provisions

The final issue we explore in this paper is the estimation of the effects of youth or student subminimum wage provisions in reducing the adverse disemployment effects of minimum wages. One existing study (Katz and Krueger, 1991) attempts to assess the likely impact of the new federal subminimum through a small survey of employers in the period

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or equal to the federal level, the state level multiplied by the sum of federal and state coverage; for observations with a minimum wage below the federal level, the federal minimum wage multiplied by federal coverage, plus the state minimum wage multiplied by state coverage.

immediately following the implementation of the new federal legislation, studying in particular the extent of usage of the subminimum. Because many states have had such subminimums in the past, however, our state-level data provide a complementary means of estimating the impact of these subminimums.

In particular, we expand equation (2) to incorporate data on state-level youth and student subminimum wages. The empirical question is whether--controlling for the state minimum wage level as well as state labor market conditions--states with subminimum wages exhibit higher employment rates. The simplest approach to this question is to augment equation (2) to

$$(3)\quad E_{it} = \alpha_0 + \alpha_1 MW_{it} + \alpha_2 SUB_{it} + X_{it}\beta + S_i\delta + \epsilon_{it},$$

where SUB is a dummy variable for the existence of either a youth or student subminimum wage. Estimates of α_2 greater than zero would support the hypothesis that youth or student subminimums reduce the adverse effects of minimum wages on employment of young workers.

An alternative specification interacts the dummy variable with a variable measuring the potential impact of a subminimum. We construct such a variable in two steps. First, for workers covered by a state minimum wage law, but not the FLSA, a subminimum could reduce the wage paid from the level of the state minimum for all workers down to the minimum allowable wage for youths or students. The limited information we have on student and youth subminimum wage provisions suggests that, on average, these provisions permit wage payments equal to about 75 percent of the minimum wage for other workers. Consequently, for each observation we construct a variable equal to 25 percent of the state minimum wage level, multiplied by state (and not federal) coverage, and divided by the mean wage in the state.[25] Second, in states with a minimum wage level above the federal level, a subminimum wage provision can reduce the wage paid to workers covered by the FLSA from the state minimum wage level to the greater of the federal level, or 75 percent of the state level. For these states we add a second term that is the smaller of 25 percent of the state level and the difference between the state and federal levels, all multiplied by federal coverage and divided by the mean wage in the state. One advantage of the addition of this variable to the model is that it has greater within-state variation than the simple dummy variables used to indicate the existence of youth or student subminimums.

Results for the dummy variable and interactive specifications, using alternatively contemporaneous and lagged values of the subminimum wage variables, are reported for student subminimums in Table 9, and for youth subminimums in Table 10. In Table 9 neither the dummy variable specifications--columns (1)-(4)--nor the interactive specifications--columns (5)-(8)--reveal any statistically significant effects, providing no evidence that student subminimum wage provisions moderate the disemployment effects of minimum wages. In Table 10, however, there is evidence that state youth subminimum wage provisions moderate the disemployment effects for teenagers. The contemporaneous and lagged dummy variables are positive and statistically significant in the specifications excluding the school enrollment rate (specifications for which, overall, minimum wage effects are negative). And, for three of the four interactive specifications the coefficients are positive and statistically significant.

To interpret the interactive coefficients, we can write out the complete minimum

[25]There were a couple of cases in which states with minimum wage levels below the federal level had a subminimum wage provision on the books, but our estimate of state coverage was zero. In these cases we treated the observation as if there was no subminimum.

wage effect as ($\alpha_1 \cdot$ MW + $\alpha_2 \cdot$(MW-SMW)$\cdot$SUB), where (MW-SMW) is our minimum wage "gap." As an example, consider a state with its minimum wage level currently set above the federal level, and with a subminimum wage provision. In this case, the second term (assuming for simplicity complete coverage) would be the difference between the state and federal levels, and the disemployment effect of raising the state minimum would be (α_1 + α_2). In column (8), for instance, raising the minimum wage variable by 0.1 (by raising the state minimum wage) would have a long-run effect of (-0.11-0.16) × 0.1 + 0.47 × 0.1, which is actually positive. We might want to discount the possibility of a positive effect, but an increase in teen employment is not entirely implausible given that the minimum wage increase, in the presence of a widespread subminimum, would make teenage workers covered by the subminimum less expensive relative to close substitutes in other age groups.

Given existing skepticism regarding the use of subminimum wage provisions (Katz and Krueger (1991)), the finding of significant effects of youth subminimums in moderating the disemployment effects of minimum wages might be regarded as spurious. For example, the existence of subminimum wage provisions may simply coincide with relatively high employment in particular states and years, perhaps because states with subminimum wage provisions also have relatively lax enforcement of state minimum wage laws. To provide evidence that the relationship between youth subminimum wage provisions and higher teen employment is causal, in Table 11 we report estimates of the same specifications as in Table 10, but only for individuals aged 20-24. Because youth subminimums typically apply to individuals aged 18 or less, if the youth subminimums boost employment of teenagers, they should have little effect on, and perhaps even reduce, the employment of those aged 20-24. On the other hand, if the findings in Table 10 reflect a spurious correlation between youth subminimums and employment rates, we might expect to find a similar positive association for these older youths. In Table 11, the point estimates of the effects of subminimum wages on the employment rates of 20-to-24 year-olds are negative in six out of eight cases. Although the estimated coefficients of the subminimum wage variables in Table 11 are not statistically significantly different from zero, they clearly do not replicate the patterns found for teenagers. This argues against the notion of a spurious correlation generating the results for teenagers, and strengthens the finding that youth subminimums moderate the disemployment effects of minimum wages.[26]

V. Conclusions

Using a specially constructed panel data set on state minimum wage laws and labor market conditions, this paper presents new evidence on the effects of minimum wages on employment of teenagers and young adults, and assesses the extent to which youth or student subminimum wages may reduce any adverse disemployment effects of minimum

[26]One alternative approach would be to use wage distributions to examine directly whether the presence of a youth subminimum alters the wages paid to teenagers. In this vein, we regressed the average teen wage for each state-year observation on the effective minimum wage and the subminimum wage variables, using the same fixed-effects specifications and control variables used in Table 10.

The findings from these regressions appear at odds with the conclusion that youth subminimums boost teen employment; the subminimum wage variable has a positive coefficient. However, the minimum wage variable itself has a negative coefficient, implying that minimum wage increases reduce teen employment as well as teen wages. This is consistent with the presence of an important uncovered sector, in which case an increase in minimum wages could reduce employment while also reducing average wages overall (but not in the covered sector alone), as disemployed workers are crowded into the uncovered sector. Estimates from our sample indicate that between 10 and 20 percent of teenagers are paid less than the state minimum wage, in states that do not have youth subminimums; this provides a lower bound on the size of the uncovered sector, since workers in the uncovered sector may earn more than the minimum. Other estimates (Schiller, 1990) imply a much larger uncovered sector for teenagers.

The presence of a significant uncovered sector for teenagers, coupled with the inability--in our data set--to determine whether workers are in the covered or uncovered sector, makes it impossible to draw conclusions regarding the use of subminimum wages from information on wage distributions.

wages. Our re-examination of the existing evidence provides a range of estimated elasticities of employment-to-population ratios with respect to minimum wages. Based on the evidence, our sense of the best estimate of the range of elasticities for teenagers is from -0.1 to -0.2, with the elasticity falling in the higher range for specifications taking account of (potentially endogenous) enrollment rates. For young adults, our best estimate of the range is from -0.15 to -0.2. In general, our results support the range of negative impacts of minimum wages on employment of teenagers and young adults suggested by the earlier time-series evidence surveyed by Brown, et al. (1982). These estimated disemployment effects contrast with those reported in the more recent studies by Wellington (1991) and Card (1991).

We have no explanation for the differences between our estimates and the smaller negative elasticities reported in Wellington's recent time-series study. We examined the possibility that ignoring the potential effects of higher state minimum wage levels posed a particular problem in the 1980s, which could be expected to exert a downward bias on more recent time-series estimates. However, when we constructed the minimum wage variables using just the federal minimum wage level, there was little change in our results.

With regard to the positive employment effects of minimum wage changes found by Card using data for California, we find evidence of a positive contemporaneous relationship between changes in minimum wages and employment rates in our data set as well. However, when the model is generalized to allow for lags, we find strong negative lagged effects running from minimum wages to employment, with negative long-run elasticities in the range reported above.

Based on our preferred specifications, we also provide estimates of the role of youth or student subminimum wages in mitigating the disemployment effects of minimum wages on teenagers. Our results indicate that youth subminimums, but not student subminimums, moderate the disemployment effects of minimum wages.

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Table 1

State Minimum Wage Levels

States with Minimum Wages Exceeding Federal Minimum (1)	Federal Minimum (2)	Average % Difference Between State and Federal Minimum (3)	Average Coverage-Adjusted Relative Minimum Wages: State Minimum > Federal Minimum (4)	Average Coverage-Adjusted Relative Minimum Wages: State Minimum = Federal Minimum (5)
1973 CA CT DC MA NJ NY 1.65 1.85 2.16 1.85 1.75 1.85	1.60	15.7	.30	.31
1974 CT DC 2.01 2.19	2.00	5.0	.29	.36
1975 CT DC NJ 2.11 2.45 2.20	2.10	7.3	.30	.35
1976 CT DC HI 2.31 2.55 2.40	2.30	5.2	.31	.36
1977 AK CA CT DC HI NJ 2.80 2.50 2.31 2.76 2.40 2.50	2.30	10.7	.33	.35
1978 AK CT DC 3.15 2.66 2.79	2.65	8.1	.33	.36
1979 AK CT DC 3.40 2.91 2.95	2.90	6.4	.33	.38
1980 AK CT DC 3.60 3.12 3.14	3.10	6.0	.32	.38
1981 AK CT DC 3.85 3.37 3.48	3.35	6.5	.34	.38
1982 AK CT DC 3.85 3.37 3.62	3.35	7.9	.30	.35
1983 AK CT DC 3.85 3.37 3.82	3.35	9.9	.32	.34
1984 AK CT DC 3.85 3.37 3.82	3.35	9.9	.32	.33
1985 AK CT DC, ME 3.85 3.37 3.85 3.45	3.35	8.4	.35	.35
1986 AK CT DC ME 3.85 3.37 3.86 3.55	3.35	9.2	.33	.34
1987 AK CT DC MA ME NH RI VT 3.85 3.37 4.16 3.55 3.65 3.45 3.55 3.45	3.35	8.3	.34	.34
1988 AK CT DC HI MA ME MN NH RI VT 3.85 3.75 4.33 3.85 3.65 3.65 3.55 3.55 3.65 3.55	3.35	11.6	.33	.33
1989 AK CA CT DC HI MA ME MN NH PA RI VT WA 3.85 4.25 4.25 4.33 3.85 3.75 3.75 3.85 3.65 3.70 4.00 3.65 3.85	3.35	16.5	.33	.32

Table 2

States with Youth or Student Subminimum Wages

	Youth Subminimum			Student Subminimum		
	States with minimum wages exceeding federal	Proportion of all states	Number of Changers	States with minimum wages exceeding federal	Proportion of all states	Number of Changers
	(1)	(2)	(3)	(4)	(5)	(6)
1973	CA CT DC MA NJ NY	.68	...	CA CT DC NJ	.50	...
1974	CT DC	.64	1	CT DC	.55	1
1975	CT DC NJ	.59	1	CT DC NJ	.55	0
1976	CT DC HI	.59	0	CT DC	.55	0
1977	AK CA CT DC HI NJ	.53	1	AK CT DC NJ	.61	2
1978	AK CT DC	.55	1	AK CT DC	.63	1
1979	AK CT DC	.57	1	AK CT DC	.61	1
1980	AK CT DC	.57	0	AK CT DC	.61	0
1981	AK CT DC	.57	0	AK CT DC	.61	0
1982	AK CT DC	.53	2	AK CT DC	.57	2
1983	AK CT DC	.53	0	AK CT DC	.57	0
1984	AK CT DC	.53	0	AK CT DC	.57	0
1985	AK CT DC	.53	0	AK CT DC ME	.57	0
1986	AK CT DC	.53	0	AK CT ME	.65	6
1987	AK CT DC MA NH RI VT	.53	0	AK CT MA NH RI VT ME	.65	0
1988	AK CT DC MA MN NH RI VT	.53	0	AK CT MA ME MN NH RI VT	.65	0
1989	AK CA CT DC MA MN NH RI VT WA	.55	1	AK CA CT MA ME MN NH PA RI VT WA	.65	0

Table 3
Time-Series Estimates of Minimum Wage Effects on Employment-to-Population Ratio[1]
(All variables are first differences)

	Teenagers (16-19)			Young Adults (16-24)		
	(1)	(2)	(3)	(4)	(5)	(6)
Minimum Wage 1955-89	-.12 (.06)	--	--	-.09 (.04)	--	--
Minimum Wage 1955-72	--	-.14 (.07)	--	--	-.08 (.05)	--
Minimum Wage 1973-89	--	-.06 (.12)	-.08 (.08)	--	-.11 (.09)	-.12 (.06)
Proportion of population in age group	-.10 (.70)	-.15 (.72)	.75 (.66)	-.07 (.29)	-.06 (.30)	-12.80 (2.57)
Prime-age male unemployment rate	-1.16 (.16)	-1.20 (.18)	-1.17 (.14)	-1.12 (.12)	-1.11 (.13)	-1.07 (.12)
$\bar{R}^2$	.64	.64	.91	.76	.75	.91
D.W.	1.43	1.45	2.23	1.20	1.20	1.47
Implied Minimum Wage Elasticity 73-89	-.10	-.05	-.06	-.06	-.06	-.07

1. Sample period is 1955-1989 for columns (1)-(2) and (4)-(5), 1973-89 for columns (3) and (6). Minimum wage variable is the federal minimum adjusted for coverage divided by average hourly earnings; the lagged value of this variable is used in each regression. Prime-age male unemployment rate is for men aged 25-54. The proportion of the population in the armed forces also is included in the model. Standard errors are in parentheses.

Table 4

Fixed-Effects and OLS Estimates of Minimum Wage Effects on Employment-to-Population Ratio, Teenagers (16-19)[1]

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Minimum wage	.07	-.17	.21	-.11	-.33	-.25	-.18	-.23
	(.10)	(.07)	(.07)	(.05)	(.08)	(.04)	(.07)	(.04)
Proportion of population in age group	-.19	-.11	-.24	-.13	.35	-.05	.02	-.16
	(.22)	(.15)	(.15)	(.10)	(.28)	(.15)	(.21)	(.11)
Prime-age male unemployment rate	-.54	-.31	-.86	-.63	-1.53	-.76	-1.39	-.80
	(.11)	(.07)	(.08)	(.06)	(.13)	(.07)	(.11)	(.06)
Proportion of age group in school	...	-.75	...	-.75	...	-.95	...	-.96
		(.03)		(.03)		(.02)		(.02)
Year effects	Y	Y	N	N	Y	Y	N	N
State effects	Y	Y	Y	Y	N	N	N	N
P-value for restricted model[2]	...	...	.00	.00	.00	.00	.00	.00
$\bar{R}^2$	.69	.86	.68	.85	.20	.78	.18	.78
Elasticity[3]	.06	-.14	.17	-.09	-.27	-.20	-.14	-.18

1. The sample consists of 751 observations covering the 50 states and Washington, D.C. For the twenty-two states identified in the CPS as early as 1973, the sample covers the years 1973-1989; for the remaining states it covers the period 1977-1989. Minimum wage variable is state minimum wage level, multiplied by federal minimum wage coverage for state, divided by average wage in state. Prime-age male unemployment rate is for men aged 25-64.
2. Likelihood-ratio test.
3. Evaluated at sample means.

Table 5

Fixed-Effects and OLS Estimates of Minimum Wage Effects on Employment-to-Population Ratio, Young Adults (16-24)[1]

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Minimum wage	-.11 (.07)	-.16 (.06)	.02 (.06)	-.12 (.04)	-.21 (.06)	-.23 (.04)	-.14 (.05)	-.21 (.04)
Proportion of population in age group	.42 (.10)	.08 (.07)	-.19 (.06)	-.24 (.05)	.28 (.11)	-.16 (.08)	-.11 (.08)	-.27 (.05)
Prime-age male unemployment rate	-.53 (.08)	-.47 (.06)	-.70 (.06)	-.69 (.05)	-1.36 (.09)	-1.13 (.07)	-1.17 (.08)	-1.06 (.05)
Proportion of age group in school	...	-.80 (.04)	...	-.82 (.04)	...	-.99 (.03)	...	-1.02 (.03)
Year effects	Y	Y	N	N	Y	Y	N	N
State effects	Y	Y	Y	Y	N	N	N	N
P-value for restricted model[2]	...	...	.00	.00	.00	.00	.00	.00
$\bar{R}^2$	.70	.82	.66	.80	.28	.66	.23	.65
Elasticity[3]	-.07	-.10	.01	-.07	-.13	-.14	-.09	-.13

1. The sample consists of 751 observations covering the 50 states and Washington, D.C. For the twenty-two states identified in the CPS as early as 1973, the sample covers the years 1973-1989; for the remaining states it covers the period 1977-1989. Minimum wage variable is state minimum wage level, multiplied by federal minimum wage coverage for state, divided by average wage in state. Prime-age male unemployment rate is for men aged 25-64.
2. Likelihood-ratio test.
3. Evaluated at sample means.

Table 6

First-Difference Estimates of Minimum Wage Effects on Employment-to-Population Ratio[1]

	Teenagers (16-19)				Young Adults (16-24)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Minimum wage	.23 (.12)	.00 (.09)	.38 (.10)	.10 (.07)	.15 (.08)	.01 (.07)	.27 (.07)	.07 (.06)
Proportion of population in age group	.29 (.24)	-.06 (.18)	.20 (.23)	-.10 (.18)	.20 (.11)	.00 (.09)	.09 (.10)	-.08 (.09)
Prime-age male unemployment rate	-.06 (.13)	.02 (.09)	-.46 (.10)	-.32 (.07)	-.08 (.09)	-.11 (.08)	-.45 (.07)	-.42 (.06)
Proportion of age group in school	...	-.74 (.03)	...	-.74 (.03)	...	-.71 (.04)	...	-.72 (.04)
Year effects	Y	Y	N	N	Y	Y	N	N
P-value for restricted model[2]	...	...	.00	.00	...	...	.00	.00
$\bar{R}^2$	.08	.50	.05	.48	.13	.38	.07	.34
Elasticity[3]	.19	.01	.31	.08	.09	.01	.17	.04

1. The sample consists of 700 observations covering the 50 states and Washington, D.C. For the twenty-two states identified in the CPS as early as 1973, the sample covers the years 1974-1989; for the remaining states it covers the period 1978-1989. Minimum wage variable is state minimum wage level, multiplied by federal minimum wage coverage for state, divided by average wage in state. Prime-age male unemployment rate is for men aged 25-64.
2. Likelihood-ratio test.
3. Evaluated at sample means.

Table 7

Fixed-Effects and First-Difference Estimates of Dynamic Specifications of Minimum Wage Effects on Employment-to-Population Ratio[1]

	Teenagers (16-19)				Young Adults (16-24)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Minimum wage	.10 (.11)	-.12 (.08)	.09 (.14)	.02 (.10)	-.03 (.08)	-.10 (.06)	.01 (.10)	-.04 (.08)
Minimum wage, lagged one year	-.14 (.12)	-.12 (.07)	-.19 (.13)	-.08 (.09)	-.26 (.08)	-.18 (.06)	-.25 (.09)	-.18 (.08)
Proportion of population in age group	-.11 (.23)	-.13 (.16)	.33 (.25)	-.11 (.19)	.42 (.10)	.08 (.08)	.15 (.11)	-.06 (.09)
Prime-age male unemployment rate	-.53 (.11)	-.29 (.08)	-.10 (.12)	-.05 (.09)	-.52 (.09)	-.47 (.06)	-.07 (.09)	-.13 (.07)
Proportion of age group in school	...	-.77 (.03)	...	-.74 (.03)	...	-.81 (.04)	...	-.70 (.04)
Estimator[2]	FE	FE	FD	FD	FE	FE	FD	FD
P-values:[3]								
0-1 lags vs. 0 lags	.20	.12	.13	.39	.00	.00	.01	.02
No year effects	.00	.00	.00	.00	.00	.00	.00	.00
$\bar{R}^2$	.70	.86	.07	.48	.71	.83	.12	.36
Elasticity[4]	-.03 (.11)	-.19 (.07)	-.09 (.18)	-.06 (.13)	-.18 (.06)	-.17 (.04)	-.15 (.09)	-.13 (.08)
Elasticities from alternative specifications:								
0 lags	.04	-.14	.13	.03	-.08	-.11	.06	.02
1 lag only	-.08	-.14	-.18	-.07	-.17	-.14	-.15	-.10

1. The sample covers the 50 states and Washington, D.C. For the twenty-two states identified in the CPS as early as 1973, the sample covers the years 1974-1989 for the fixed-effects estiamtes, and 1975-1989 for the first-difference estimates; for the remaining states the corresponding periods are 1978-1989 and 1979-1989. Minimum wage variable is state minimum wage level, multiplied by federal minimum wage coverage for state, divided by average wage in state. Prime-age male unemployment rate is for men aged 25-64.
2. All specifications include year effects.
3. Likelihood-ratio test.
4. Long-run, evaluated at sample means. Standard errors, reported in parentheses, treats coefficient estimates, but not means, as random.

Table 8

Fixed-Effects Estimates of Alternative Specifications of Dynamic Specifications of Minimum Wage Effects on Employment-to-Population Ratio, from Table 7, Long-run Elasticites Evaluated at Sample Means[1]

	Teenagers (16-19)		Young Adults (16-24)	
	Excluding Enrollment (1)	Including Enrollment (2)	Excluding Enrollment (3)	Including Enrollment (4)
A. No coverage adjustment	-.06	-.13	-.09	-.11
B. Federal and state coverage adjustment	.11	-.02	-.07	-.09
C. Federal minimum wage level in place of state minimum wage level	-.01	-.15	-.17	-.14
D. Adding one lag of other control variables	.01	-.17	-.15	-.15
E. Weighting by state population[2]	.06	-.18	-.07	-.12

1. The sample covers the 50 states and Washington, D.C. For the twenty-two states identified in the CPS as early as 1973, the sample covers the years 1974-1989; for the remaining states it covers the period 1978-1989. Minimum wage variable is state minimum wage level, multiplied by federal minimum wage coverage for state, divided by average wage in state. Prime-age male unemployment rate is for men aged 25-64. All specifications include state and year effects.
2. Elasticities are evaluated at weighted means.
3. Elasticities are sums of coefficient estimates.
4. For the twenty-two states identified in the CPS as early as 1973, the sample covers the years 1973-1989; for the remaining states it covers the period 1977-1989.

Table 9

Fixed-Effects Estimates of Effects of Student Subminimum Wage Provisions on Employment-to-Population Ratio of Teenagers[1]

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Minimum wage	.10	-.13	.10	-.13	.08	-.14	.10	-.12
	(.11)	(.08)	(.11)	(.08)	(.12)	(.08)	(.12)	(.08)
Minimum wage, lagged one year	-.15	-.12	-.14	-.12	-.13	-.11	-.14	-.13
	(.12)	(.08)	(.12)	(.08)	(.12)	(.08)	(.12)	(.08)
Student subminimum	-.01	-.01	...	...	-.01	-.01	...	...
	(.01)	(.01)			(.01)	(.01)		
Student subminimum, lagged one year	...	...	-.01	-.01	...	...	-.00	-.02
			(.01)	(.01)			(.01)	(.01)
Student subminimum x minimum wage gap[2]	...	...	...	...	.21	.16	...	...
					(.28)	(.19)		
Student subminimum x minimum wage gap, lagged one year	...	...	...	...	...	...	-.02	.28
							(.36)	(.24)
Proportion of population in age group	-.12	-.14	-.12	-.15	-.12	-.14	-.12	-.14
	(.23)	(.16)	(.23)	(.16)	(.23)	(.16)	(.23)	(.16)
Prime-age male unemployment rate	-.54	-.30	-.54	-.30	-.54	-.30	-.53	-.31
	(.11)	(.08)	(.11)	(.08)	(.11)	(.08)	(.11)	(.08)
Proportion of age group in school	...	-.77	...	-.77	...	-.77	...	-.77
		(.03)		(.03)		(.03)		(.03)
Joint significance of subminimum wage variables[3]	...	...	...	...	.42	.13	.85	.08
$\bar{R}^2$	.70	.86	.70	.86	.70	.86	.70	.86

1. The sample covers the 50 states and Washington, D.C. For the twenty-two states identified in the CPS as early as 1973, the sample covers the years 1974-1989; for the remaining states it covers the period 1978-1989. Minimum wage variable is state minimum wage level, multiplied by federal minimum wage coverage for state, divided by average wage in state. Prime-age male unemployment rate is for men aged 25-64. All specifications include state and year effects.
2. The minimum wage gap is defined as follows: for all observations, it includes 25 percent of the state minimum wage level, multiplied by state coverage (i.e., the proportion of workers covered by the state minimum wage, but not the FLSA), divided by the mean wage; for observations with minimum wage levels greater than the federal level, this is added to the smaller of either 25 percent of the difference between the state and federal minimum wage level or the difference between the state and federal minimum wage level, multiplied by federal coverage, divided by the mean wage.
3. P-value for likelihood-ratio test.

Table 10

Fixed-Effects Estimates of Effects of Youth Subminimum Wage Provisions on Employment-to-Population Ratio of Teenagers[1]

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Minimum wage	.09 (.11)	-.12 (.08)	.10 (.11)	-.12 (.08)	.02 (.12)	-.15 (.08)	.11 (.11)	-.11 (.08)
Minimum wage, lagged one year	-.13 (.12)	-.12 (.08)	-.14 (.12)	-.12 (.08)	-.11 (.12)	-.11 (.08)	-.20 (.12)	-.16 (.08)
Youth subminimum	.05 (.02)	.00 (.01)	...	...	.04 (.02)	.00 (.01)	...	...
Youth subminimum, lagged one year	...	...	.03 (.01)	.00 (.01)	...	...	.02 (.01)	-.00 (.01)
Youth subminimum x minimum wage gap[2]	...	...	...	...	.50 (.28)	.18 (.19)	...	...
Youth subminimum x minimum wage gap, lagged one year	...	...	...	...	...	...	.61 (.32)	.47 (.22)
Proportion of population in age group	-.07 (.23)	-.12 (.16)	-.08 (.23)	-.12 (.16)	-.06 (.23)	-.12 (.16)	-.08 (.23)	-.13 (.16)
Prime-age male unemployment rate	-.55 (.11)	-.29 (.08)	-.55 (.11)	-.29 (.08)	-.54 (.11)	-.29 (.08)	-.57 (.11)	-.31 (.08)
Proportion of age group in school	...	-.77 (.03)	...	-.77 (.03)	...	-.77 (.03)	...	-.77 (.03)
Joint significance of subminimum wage variables[3]	...	...	...	...	.00	.54	.01	.07
$\bar{R}^2$	.70	.86	.70	.86	.70	.86	.70	.86

1. The sample covers the 50 states and Washington, D.C. For the twenty-two states identified in the CPS as early as 1973, the sample covers the years 1974-1989; for the remaining states it covers the period 1978-1989. Minimum wage variable is state minimum wage level, multiplied by federal minimum wage coverage for state, divided by average wage in state. Prime-age male unemployment rate is for men aged 25-64. All specifications include state and year effects.
2. The minimum wage gap is defined as follows: for all observations, it includes 25 percent of the state minimum wage level, multiplied by state coverage (i.e., the proportion of workers covered by the state minimum wage, but not the FLSA), divided by the mean wage; for observations with minimum wage levels greater than the federal level, this is added to the smaller of either 25 percent of the difference between the state and federal minimum wage level or the difference between the state and federal minimum wage level, multiplied by federal coverage, divided by the mean wage.
3. P-value for likelihood-ratio test.

Table 11

Fixed-Effects Estimates of Effects of Youth Subminimum Wage Provisions on Employment-to-Population Ratio of 20-24 Year-Olds[1]

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Minimum wage	-.05 (.10)	-.08 (.08)	-.05 (.10)	-.08 (.08)	-.04 (.10)	-.05 (.09)	-.05 (.10)	-.09 (.08)
Minimum wage, lagged one year	-.21 (.10)	-.19 (.09)	-.22 (.10)	-.19 (.09)	-.21 (.10)	-.20 (.09)	-.21 (.10)	-.16 (.09)
Youth subminimum	-.00 (.01)	-.01 (.01)	...	...	-.00 (.01)	-.01 (.01)	...	...
Youth subminimum, lagged one year	...	...	.01 (.01)	.01 (.01)	...	...	.01 (.01)	.01 (.01)
Youth subminimum x minimum wage gap[2]	...	...	...	...	-.01 (.29)	-.16 (.21)	...	...
Youth subminimum x minimum wage gap, lagged one year	...	...	...	...	...	...	-.09 (.28)	-.37 (.24)
Proportion of population in age group	.51 (.14)	.36 (.12)	.49 (.14)	.32 (.12)	.51 (.14)	.37 (.12)	.49 (.14)	.34 (.12)
Prime-age male unemployment rate	-.59 (.09)	-.62 (.08)	-.59 (.10)	-.63 (.08)	-.59 (.09)	-.62 (.08)	-.59 (.10)	-.61 (.08)
Proportion of age group in school	...	-.77 (.05)	...	-.77 (.05)	...	-.77 (.05)	...	-.77 (.05)
Joint significance of subminimum wage variables[3]	...	...	...	...	.97	.38	.65	.20
$\bar{R}^2$	.56	.67	.56	.67	.56	.67	.56	.67

1. The sample covers the 50 states and Washington, D.C. For the twenty-two states identified in the CPS as early as 1973, the sample covers the years 1974-1989; for the remaining states it covers the period 1978-1989. Minimum wage variable is state minimum wage level, multiplied by federal minimum wage coverage for state, divided by average wage in state. Prime-age male unemployment rate is for men aged 25-64. All specifications include state and year effects.
2. The minimum wage gap is defined as follows: for all observations, it includes 25 percent of the state minimum wage level, multiplied by state coverage (i.e., the proportion of workers covered by the state minimum wage, but not the FLSA), divided by the mean wage; for observations with minimum wage levels greater than the federal level, this is added to the smaller of either 25 percent of the difference between the state and federal minimum wage level or the difference between the state and federal minimum wage level, multiplied by federal coverage, divided by the mean wage.
3. P-value for likelihood-ratio test.