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EMPLOYMENT POLICIES INSTITUTE

The Effect of Minimum Wages on the Labor Force Participation Rates of Teenagers

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Dr. Wessels' research has appeared in economics and education journals, including *Economic Inquiry* and *The Journal of Cooperative Education*. In 1994, he won the Western Economic Association's Best Article of the Year Award for a paper that appeared in *Economic Inquiry*. Dr. Wessels also wrote "Economics: A Streamlined Course for Students and Business People" for the *Barron's Business Review Series*. He received his Ph.D. from the University of Chicago in 1976.

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Executive Summary

Congress has been considering a hike in the federal minimum wage from $5.15 to $6.15 an hour or higher. It has been estimated that such a raise would affect over 10 million workers, many of whom are teenagers. A considerable body of research shows that while such increases might raise the wages of some workers, it would also eliminate jobs and work opportunities for others. For example, by one consensus view of this effect, a 10 percent increase in the minimum wage would reduce the employment of teenagers overall by anywhere from 1 to 3 percent.

However, this estimate ignores the fact that employers may react in other ways to a minimum wage hike. For example, when employment costs rise, employers may eliminate some fringe benefits such as contributions towards insurance, transportation, or parking, so that the total compensation of workers does not rise even though wages increase. Employers may also raise their expectations of workers, including requiring greater work effort to cover a reduction in total hours worked. Employers may increase the hiring standards for entry-level jobs, such as requiring more education or work experience.

Many of the predictable adjustments by employers to a higher minimum wage reduce the attractiveness of work. If increased requirements are imposed on workers, or less training is available, this may reduce the attractiveness of work itself. Walter Wessels, a professor of economics at North Carolina State University, has studied the effect that higher minimum wages have on the likelihood that teenagers will choose the employment option (i.e., to be employed or look for work). He studies teenagers, who tend to be strongly affected by minimum wage increases because many are in entry-level jobs.

In the first study in 20 years to examine this question, Dr. Wessels concludes that when minimum wages go up, fewer teens on average choose the employment option. This overall outcome is entirely consistent with the findings by others that minimum wage hikes cause teens with greater skills and experience to work more and those with fewer skills and experience to work less. Because work by teenagers has been shown to have beneficial long-term consequences on their subsequent labor force success, Dr. Wessels' study implies that higher minimum wages reduce the future economic well-being of those who are displaced from work and discouraged from seeking work when they are teens.

Study Design

Theoretical Framework

Dr. Wessels argues that the best way to estimate the effect of the minimum wage on the value of being in the labor market is to examine its effect on labor force participation rates. The author recognizes that changes in the minimum wage can affect both the demand for and supply of labor force participants. However, he believes that the main effect of the minimum wage is through the demand side. If supply side effects occur, they are likely to reduce the supply of labor force participants through (1) raising the earnings of other family members (i.e., an income effect); (2) increasing lifetime potential earnings from work (i.e., a wealth effect); or (3) increasing the value of future work relative to current work (a relative wage effect). Dr. Wessels believes these effects will be small because (1) minimum wages have little effect on family income, particularly for families below the poverty line; (2) wealth effects are small because the minimum wage affects teenagers for only a small fraction of their working life; and (3) he finds no evidence that teens shift their labor supply toward the future in response to a minimum wage hike.

If the minimum wage has little effect on the supply of teenage labor force participants, then its primary effect will be through its effect on the demand for labor force participants. The demand for labor force participants (manifested by employers' wage offers and hiring activity) determines the value of being in the labor force.

Data and Estimation

Dr. Wessels uses quarterly data on the labor force participation rates of teenagers from 1978 through 1999 to assess the effects of several rounds of increases in national minimum wage rates. He was able to consider the 1978-1981, 1990-1991, and 1996-1997 increases in the national minimum wage. He finds that from 1978 to 1999, the percentage of teenage workers earning at or less than the minimum wage has generally fluctuated between just above 50 percent to just below 17 percent, and has been under 40 percent since 1991. Historically when this percentage has fallen to about 17, Congress has raised the minimum wage.

After controlling for the effects of the business cycle, per capita income, adult wage rates, and the number of teens affected by a minimum wage hike, Dr. Wessels finds that these minimum wage hikes reduced teenage labor force participation rates. These declines were statistically significant for teenagers overall, and also for whites, males, and females considered separately. Specifically, his research shows the 1978-1981 hikes reduced labor force participation by 6.85 percent (3.615 percentage points), the 1990-1991 hikes, by 4.09 percent (2.07 percentage points), and the 1996-1997 hikes, by 2.78 percent (1.31 percentage points).

— Dr. Richard S. Toikka
Chief Economist

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I. INTRODUCTION

Although the minimum wage's impact on employment is important, its effect on the value of being in the labor market can also be measured by how it affects labor force participation. For example, a minimum wage hike that did not reduce employment could still make workers worse off by causing employers to cut back on non-wage compensation and by increasing the scarcity of job openings. Labor force participation would decline, reflecting the decline in the value of being in the labor force. This paper is the first in 20 years to investigate how the minimum wage affected the labor force participation rate of teenagers. It investigates this effect from 1978 to 1999.

An important issue is how binding the minimum wage is on the market. Most papers use a relative minimum wage measure. This paper uses a more direct measure: the fraction of teenage workers being paid the minimum wage. This allows a more objective measure of how long and to what extent the minimum wage effectively constrains the job market.

Another important issue is how to control for the effects of the business cycle. This paper addresses this issue in several ways. First, it covers three episodes of major minimum wage hikes, each occurring in different economic environments. Second, it uses current time-series modeling to study the effects of minimum wages[1] instead of the current convention in the minimum wage literature of using a series of yearly dummy variables to control for time-series effects. The use of lagged values of the dependent variable and the inclusion of several explanatory variables related to the business cycle are among the steps taken to control for the effects of the business cycle.

It was found that minimum wage hikes between 1978 and 1999 significantly decreased the labor force participation rate of teenagers. This suggests that the minimum wage reduced the value of being in the labor market for many teenagers.

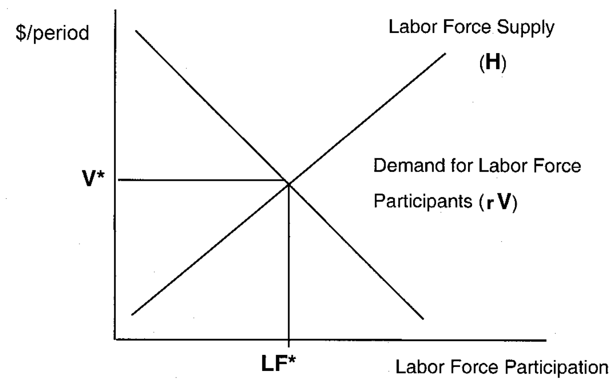
II. LABOR FORCE PARTICIPATION

The effect of the minimum wage on the value of being in the labor market is best estimated by how it affects labor force participation rates. Four articles that estimated the effects of minimum wages on labor force participation are Kaitz (1970), Mincer (1976), Ragan (1977), and Wessels (1980a). All found that the minimum wage decreased (or left unchanged) the labor force participation rate of low-wage workers. Since 1980, there have been no published estimates of the impact of minimum wages on labor force participation rates. This paper updates this research.

The basic assumption in this paper is that if the minimum wage decreases the labor force participation rate, it can be inferred that it has also decreased the value of being in the labor market. A person is in the labor force if they are employed or if they are actively looking for work (that is, "unemployed"). Let V be the value of being in the labor market. It includes the lifetime utility if one joins the labor force in the current period and takes account of the expected cost of searching for a job and the value of future higher earnings from more job experience. A person enters the labor force if the income flow from being in the labor force (rV) exceeds the implicit income of not working (H, where H stands for "home time"); r is the relevant discount rate. The variable H can be thought of as the value of leisure. However, in a more general sense, H is the implicit income flow from the present value of lifetime utility if one did not work in the current period (for most teenagers, it would include the value of being in school full-time). If $rV \geq H$, the worker joins the labor force.

The inference that a lower labor force participation rate implies a lower V comes from the assumption that the minimum wage shifts V more than it does H. Consider the demand and supply for labor force participants as shown in Figure 1.

Figure 1
The Demand and Supply for Labor Force Participants in the Absence of the Minimum Wage

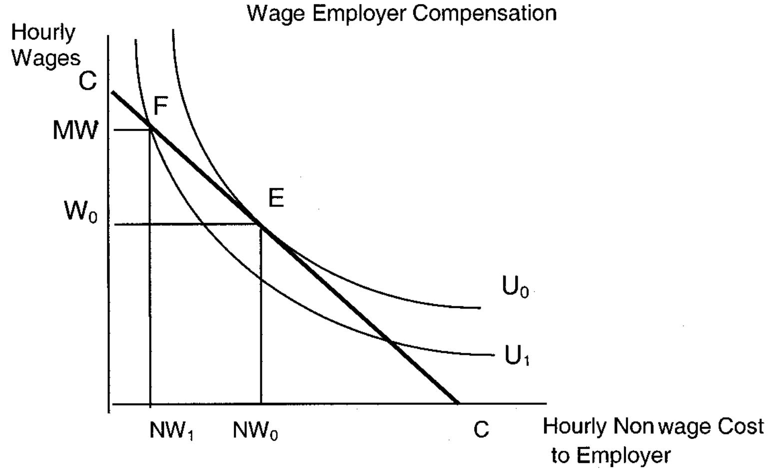


The supply of labor force participants is the schedule of H across the population. The demand curve for labor force participants shows the value of rV. It is negatively sloped in part because a higher labor supply reduces the wage. It is also negatively sloped because, holding employment constant, more labor force participants means more unemployed persons competing for the same number of jobs. This increases the time and thus cost of searching for a job, reducing V.[2] The minimum wage affects the demand for labor force participants because it shifts wages, employment, and the probability of getting a job. The basic assumption here is that increasing the minimum wage does not significantly shift the supply schedule. In this case, if the minimum wage reduces V, it will shift the demand curve down, reducing labor force participation. Although employed persons may be made better off in the current period by the minimum wage, when they reenter the labor market through job turnover, they will then be worse off because of the lower V. In this case, a minimum wage hike lowering V has the potential of making the vast majority of workers worse off.[3]

One must be careful to not confuse the "demand for labor force participants" with the demand for workers by employers. For example, a minimum wage hike could decrease employment, and yet, because it substantially raises wages, increase the demand for labor force participants by making being in the labor force more attractive.

Many different search models have been applied to minimum wages. Wessels (1980a and 1980b) presents a simple version. Thomas Carter (1998) analyzes Rebitzer and Taylor's efficiency wage model (1995) showing that when the minimum wage increases employment in this model, job searchers are made worse off (V goes down). Lang and Kahn (1998) present a bilateral search model with heterogeneous workers in which the minimum wage makes low productivity workers worse off

Figure 2
Effect of Minimum Wages on Employee Utility and the Mix of Wage and Non-Wage Employer Compensation



while not enhancing the welfare of more productive workers. Mortensen and Burdett (1989) present an equilibrium model in which larger firms pay higher wages and have lower turnover whereas smaller firms are, in a sense, more competitive, paying a wage closer to V. Although they definitely do not come to this conclusion, a minimum wage lowering V would make all workers in their model worse off.

The minimum wage has many effects on V other than its obvious effect on wages and employment. First, while it may increase the wage, employers might offset this by cutting back on nonwage compensation. For example, employers might reduce the flexibility of working hours,[4] increase the work pace, demand work "off the clock," reduce work hours, or reduce on-the-job training.[5] If employers fully minimize costs, the offset will reduce the full wage of workers (Wessels, 1980a). Figure 2 shows the offset effect. Given the compensation the firm wants to pay (C), the firm chooses the optimal mix of wages and nonwage compensation to attract workers. This optimal mix at Point E maximizes the worker's full wage (or utility) given C. The worker's full wage is U_0. In Figure 2, the firm is spending W_0 on money wages and NW_0 on nonwage costs (note that a \$1 spent on nonwage items need not be worth \$1 to the worker). A minimum wage of MW that does not change the firm's total compensation cost will push the compensation mix toward wages and, consequently, lower the utility (full wage) of workers. In Figure 2, the firm is shown as being at Point F, with a resulting lower full wage for workers. If firms act to keep the full wage at market levels (or at some fixed percent above market levels), the interaction of labor demand and labor supply will cause the equilibrium full wage to fall between Points E and F. Workers will be worse off.[6]

A second effect the minimum wage has on V is its effect on the availability of jobs and on job turnover. Unfortunately, the actual effect of minimum wages on turnover rates is unknown because the Department of Labor no longer publishes this information.[7] One presumes it lowers turnover, which usually makes new labor force participants worse off (since less turnover means fewer job openings).

It is argued here that a minimum wage hike will not shift the supply curve (the H schedule) significantly. There are two main ways a minimum wage can affect the value of a person's home time (H). First, it can have an income effect, raising the value of leisure or home time, shifting the supply curve inward. The main way this income effect could occur is by the minimum wage increasing the income of other persons in the worker's family. However, the evidence suggests that the minimum wage has little effect on family income, especially families below the poverty line.[8] A related effect is the wealth effect from the minimum wage increasing the value of the worker's future job opportunities. However, since the minimum wage affects teenage workers for only a short span of their working life, this impact is likely to be small.

III. MINIMUM WAGE AND THE SUPPLY OF LABOR

A. Adjustment Costs

The presence of adjustment costs has several consequences for modeling the impact of minimum wages. The reason is that most minimum wage hikes come in sets. The first set in the sample period began in 1978, the second in 1990, and the third in 1996. Congress increased the minimum wage in steps and, for the 1990 and 1996 hikes, the second step in the hike was effectively more binding to employers than the first.

It is reasonable to assume that firms were aware, at least after the first hike, what and when the subsequent hikes were going to be. If there were no adjustment costs, firms would hire and fire according to current wages. On the other hand, in the presence of adjustment costs, firms would be more reluctant to create new job positions if they anticipated still higher minimum wages in the future. In this case, it is appropriate to treat each set of hikes as a unit rather than as separate hikes.

A second consequence is that minimum wage hikes of different sizes are likely to have different effects. A small hike that is transitory may not affect employment at all nor lead the firm to make any nonwage offsets (it can be argued that the

1996-1997 hikes fit this category). The main effect of such a hike will be to reduce job turnover and the availability of new jobs. On the other hand, a larger hike may lead the firm to adjust employment or nonwage compensation, or both. This may lead to a very different outcome. The result is that not all minimum wage hikes will have the same effect. Consequently, I treat each separately. If the differences between hikes are large, this may blunt the criticism by Card and Krueger (1995) that the empirical literature has not found a uniform (or more significant) effect for all hikes over time.

B. Choice of Dependent Variable

A Box-Cox transformation was performed on the labor force participation rate of teenagers (with an AR(9) error structure) in order to choose between a linearized or a log form for the rate. The linear form was preferred (with a log-likelihood of 5974 compared to the log form's 5784). However, the linear form is limited to being between zero and one, violating the assumption of normality. The linear form was then compared to its "normit" transformation[9] (the normit version having a normal cumulative probability equal to the labor force participation rate) as well as its logit form (another variant appropriate for proportions). The normit form's error term was closest to being normally distributed (with a Jarque-Bera normality test statistic[10] of 7.19, compared to the logit's 11.94). In addition, the residuals using the normits (weighted as described below) were homogeneously distributed over all states (the Bartlett test statistic,[11] at a significance level of 11.1%, could not reject the null that states' residuals were homogeneous).

Viewing each teenager in a state as having the same probability of being in the labor force, then labor force participation rate is a binomial process. According to Parzen (1960), reasonable accuracy is achieved by using the normal distribution to approximate a binomial process when $np(1-p) > 10$, where n is the sample size and p is the probability. This is satisfied in the data set. Because of these considerations, the normit form was selected.

C. Choice of Regression Form

Most articles on minimum wages using panel data use the following specification:[12]

$$E_{it} = \alpha_0 + MW_{it}\beta + X_{it}\gamma + T_t\tau + S_i\delta + e_{it} \quad (1)$$

where i is the state index and t is the time index. E is the ratio of teenage employment to teenage population, MW is a variable representing the larger of the state or federal minimum wage, X is a set of explanatory variables, T is a set of yearly dummy variables, and S is a set of identifier variables for the individual states.

A key problem with this equation is its use of yearly dummy variables to control for the business cycle. The use of yearly dummy variables masks the effect of most events (such as increases in the federal minimum wage), making their use sterile and highly nonstandard in the econometric literature for studying the effects of policy interventions. If their use were common in models, for example, the annual dummy variables would appear to drive most recessions.

It is important to control for the effects of the business cycle. The standard method in the macroeconomic intervention literature is to introduce lagged values of the dependent variable as well as lags in the independent variables that are related to the business cycle (for example, the unemployment rate). The regressions in this paper use an autoregressive data generating process of nine lags (AR(9)) in the labor force participation rate. More lags proved to be insignificant. Only one other paper in the minimum wage literature uses more than an AR(1) process (Williams and Mills, 1998). The unemployment rate (with multiple lags) was included to reflect the effects of the business cycle (most minimum wage papers use only the current unemployment rate). The unemployment rate of white males ages 30 to 54 was used, and was significant up to seven lags. In addition, the labor force participation rate of 30- to 39-year-olds was used to reflect other shocks to the labor market. This age group was chosen because its labor force participation rate is not affected by the minimum wage, yet it does vary enough to reflect the state-by-state differences in labor markets and how they change over time. In addition, each state's per capita income was also an explanatory variable.

The other explanatory variables include adult wage rates (ages 30 to 39), the fraction of teens in the working population (ages 15 to 54), and the dummy variable indicating whether the state's effective minimum wage was the federal minimum wage or its own rate (if higher). The age range for the adult wage rate was selected to be close to, but not highly correlated with, the teenage wage rate. As it turns out, any wage measure generally proved to be insignificant, whether chosen from adult wages or teen wages. Much of the minimum wage literature depends on the adult wage to account for the effectiveness of the minimum wage. This is inappropriate given the insignificance of the adult wage.

D. The Choice of Minimum Wage Variables

The usual minimum wage variables in the literature are (1) the minimum wage relative to the adult wage level (for example, Neumark and Wascher (1994) or Card, Katz and Krueger (1994)) or (2) a set of dummy variables for the time periods in which the minimum wage was increased (for example, Deere, Murphy and Welch, 1995).

The relative level of the minimum wage makes sense only if it can be indexed to reflect the minimum wage's effective constraint on firms. Most often the indexing is done by entering both the log of the minimum wage and the log of some adult wage in the regression. A problem with this relative wage measure can be easily illustrated. Suppose the minimum wage is well below market wages for all workers so that it has no effect. An improvement in economic conditions then raises employment and market wages. The relative minimum wage variable would go down in value while employment goes up. What is needed is a way to measure how binding a constraint the minimum wage is.

The second choice is the use of dummy variables to reflect the timing of minimum wage hikes. The key issue here is how long a period one should lag these dummy variables. Certainly not forever, since the effects of the minimum wage wear off. And most likely, one period is too short. One criticism of using lags is that with enough of them, any two variables can be related in any way one wants. Thus, without a way to justify the timing of the minimum wage's effect, the results are subject to question.

E. Measuring the Effective Minimum Wage

To approach the problem of how long the minimum wage represents an effective constraint on a labor market, this paper utilizes the fraction of employed teenagers, ages 15 to 19, who are earning the minimum wage or less.[13] To get an adequate sample, this figure was collected for the United States for each quarter from the outgoing rotations of the Current Population Survey (CPS). This fraction adjusts for state differences in minimum wages when the state has a higher minimum wage.

This variable should be proportional to the fraction of employed teenagers significantly impacted by the minimum wage. Another break point could have been chosen (such as a certain percent higher than the minimum wage) but there is a problem with these other break points: The distribution of wages is not smooth. For example, in the first quarter of 1994, when the minimum wage was $4.25, in the range between $4.25 and $5.25, 78% of the workers were paid exactly $4.25, $4.50, $4.75, $5.00 or $5.25 ($5 being the most popular with 28% of the workers in this range). Most teenagers earn "on the quarter" ($5.50, $5.75, $6.00 and so forth) with few in between. This makes the choice of another break point problematic because the proportions earning below it depend upon how many "quarter" points there are between it and the minimum wage (for example, the minimum wage of $5.15 plus 10 percent includes two quarter points, whereas $3.35 includes only one). In light of this, I chose "the minimum wage or less" as the criterion. From this point on, I will simply refer to this fraction as the "fraction of teenage workers earning the minimum wage" with the "or less" understood.

The monthly fraction of teenage workers earning the minimum wage is shown in Figure 3. Note how this fraction spikes up at each minimum wage hike (1979, 1980, 1981, 1990, 1991, 1996 and 1999).

Figure 3
Percentage of Teenagers at the Minimum Wage or Less

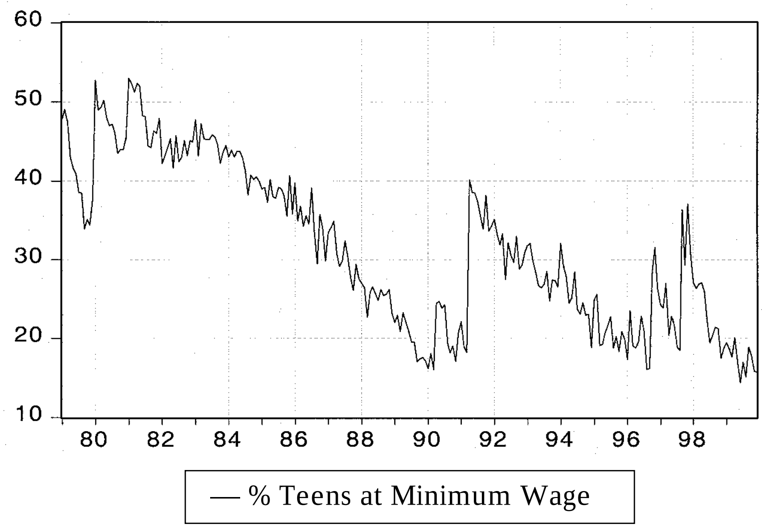


Figure 3 reveals that when approximately 17% of teenagers earn the minimum wage, Congress has passed a new minimum wage law. This includes the 1978 hikes, the 1990 hikes, and the 1996 hikes. The chart does not show the pre-1979 percent, but in the May 1977 CPS survey, 18.77% of employed teenagers were earning the minimum wage (before the new minimum wage was imposed in January 1978). Before the 1990 hike, the fraction bottomed out at 16.1%. Before the 1996 hike, it bottomed out at 17.3%. Currently, 16% of teenagers are earning the minimum wage. Whether a "17% rule" proves to be an economic "law" is an interesting political-economic question not given further consideration here.

A second observation is that between successive sets of hikes, this fraction has fallen at an accelerated rate. Between 1981 and the next hike in 1990, it fell at a rate of 10.2% per year (estimated from a regression on the log of the fraction). Between the hike in 1991 and the next hike in 1996, it fell at a higher rate of 12.8% per year, and between the hike in 1997 and the end of 1999, it fell at an annual rate of 29.1%. The economy has apparently moved faster and faster to "shake off" the effects of the minimum wage.

The fraction of teenage workers earning the minimum wage will be called "Frac_MW". This variable is used in two ways in this paper. First, it is used as a minimum wage variable. Second, it is used to measure the duration and relative strength of a minimum wage hike. The minimum wage variable used in this paper will be of this form:

$$(2) \quad MW_t = \Delta MW \bullet \frac{Frac\_MW_t - Low(Frac\_MW)}{Hi(Frac\_MW) - Low(Frac\_MW)}$$

where the change in the minimum wage is its proportional increase (at the time of the hike); the low Frac_MW is the lowest or baseline fraction after the minimum wage hike (usually near 17%); and the high Frac_MW is its highest value (usually the percent at the time of the hike). Thus, the far right-hand term, which I will refer to as the "weighting" of the minimum wage, is like a decay term: It compares the fraction's current increase over baseline with the total most recent increase over the same baseline. It goes from one to zero over the period after the minimum wage is increased. When it equals one, the minimum wage hike has its full impact; when it equals zero, the minimum wage impact has returned to the level before the minimum wage was increased.

To illustrate, suppose the minimum wage goes up 10%. Before the minimum wage was increased, suppose 25% of the teenage workforce was being paid the minimum wage. Then, at the new higher minimum wage, 45% of the teenagers are paid the minimum wage, with this fraction falling to 35% in the subsequent year and to a low of 25% in 2 years. The weight in equation 2 would be calculated as (Frac_MW–0.25)/0.20. Initially, the minimum wage variable equals 10%. In the subsequent year, it equals 5% (10% x 0.5). After 2 years, it equals zero.

The implicit assumption is that the fraction of workers affected is proportional to this weighting. Continuing the example, the greatest change in the labor force due to a hike occurs when it affects most workers (when 20% more workers are affected compared with the fraction affected before the hike). In 1 year, the change is half of what it was (when only 10% "more" workers are affected), and[14] finally, in 2 years, the change has disappeared (the affected workers return to 25%). Although this assumption can be questioned, it is preferable to assuming a fixed duration of an arbitrary length. Figure 3 shows that the 1978-1981 set of minimum wages affected teenagers throughout the 1980s. On the other hand, the 1996-1997 set of hikes had an effect lasting less than 4 years.

The ΔMW for the 1978-1981 hikes equals the sum of the hikes occurring on or before the current quarter. In 1979, the increase in 1979 was combined with the 1978 increase since the sample begins in 1979. In the tables, it will be referred to as the "1979-1981" set of hikes to make it clear that the sample period began in 1979. The high and low Frac_MW was chosen over the whole period, its highest value occurring in 1981. The other hikes were treated similarly.[15]

The change in the minimum wage was the change in the federal or state minimum wage (using the higher of the minimum wages). In addition, some states increased their minimum wage earlier than did the federal government, so the weighting was shifted forward in these states to reflect this fact. States whose minimum wage did not change in the years of federal hikes were assigned a minimum wage variable of zero. The values of the minimum wage variable for those states where the federal minimum wage was higher are shown in Appendix Table 2.

A potential criticism of the weighting method used is that a growing economy, if its effects are not controlled for, will cause the weighting to decrease (as fewer workers are paid the minimum wage). In this case, it might be found that the minimum wage has a negative effect on labor force participation while what is really discovered is that labor force participation responds to better working conditions. To avoid this problem, numerous steps were taken to control for the effects of the business cycle (using an AR(9) error structure, the inclusion of key variables related to the business cycle and the use of state-specific time trends). In addition, the two main episodes (1990 and 1996) took place in very different business environments, yet the minimum wage had a similar negative effect in both. The fact that the results match those from earlier episodes also is suggestive that the results reflect the true effects of the minimum wage, controlling for the effects of the business cycle.

One of the key variables in the regression is the unemployment rate (for white males, ages 30-54), which is used to control for the effects of the business cycle (up to 7 lags). A desirable minimum wage variable should be exogenous from the key control variables in the equation (Card and Krueger, 1995). A Granger causality test was run over the three minimum wage variables (for each hike, approximately over the period of each variable's effectiveness: 1979:1 to 1989:4 for the 1979-1981 hikes, 1990:1 to 1996:2 for the 1990-1991 hikes, and 1994:4 to 1999:4 for the 1996 hikes). The test was run for seven lags (because unemployment was lagged up to seven periods in the regression). In testing the null hypothesis that unemployment did not cause the minimum wage variable, the highest levels of significance reached by the *F* statistic were, respectively, 77.2%, 8.8% and 13.8%. Unemployment for white males ages 30-54 is not significantly related to the minimum wage variable.

F. The Choice of Period

Most panel studies of minimum wages use monthly data. Unfortunately, monthly data result in small cell sizes. With a small cell size comes measurement error that can bias the results, particularly in dynamic models.[16] It also can be shown to bias the tests for stationarity toward rejecting the nonstationary null hypothesis.[17] One solution to these problems is to use instrument variables, in this case for lagged labor force participation. Because the choice of variables to compose the instruments is somewhat arbitrary, this leaves the results from the use of such instruments open to question. Instead, a longer sampling period (a quarter of a year being one period) was chosen. The result is a larger cell size, which also reduces other biases (see Harris and Tzavalis, 1999).

The use of quarterly data allows for a larger cell size, which is particularly important for investigating the subcategories of teenagers. The variance caused by measurement error can be approximated by the mean of L(1–L)/N, averaged for each quarter over the individual states, where L is the labor force participation rate and N is the number of teenagers in the particular state's quarter cell. The variance equals 0.002131, which is approximately 20% of the variance in L (which equals 0.01073). The figure for the normit of L is 0.01720, which is 24% of the variance of the normit of L (which equals 0.07268). The corresponding percentages for monthly data are much higher, at 35% and 40%.

Further gains could be achieved by using annual data, but this loses too much of the minimum wage's effect because no hike since 1981 has occurred at the beginning of the year. Another solution would be to use national data, but *F* tests show the state fixed effects to be highly significant.

Table 1
Test for Unit Root
Im, Pesaran, and Shin t-bar statistic for Panel Data

Variable	Constant	Constant Plus Trend
LFPRteens	-2.550**	-2.693**
Normit of LFPR teens	-2.548**	-2.646**
Unemployment Rate for Prime Age White Males	-2.472**	-2.945**
LFPR ages 30-39	-1.992**	-2.222
Fraction of Teens in Population	-1.552	-2.015
First Difference of Fraction of Teens in Population	-3.146**	-3.637**
Log Wage ages 30-39	-1.121	-2.716**
First Difference of Log of Wage 30-39	-3.403**	-3.528**
Log Per Capita Personal Income	-0.928	-1.961
First Difference of Log Per Capita Personal Income	-3.189**	-3.128**
Fraction of Employed Teens At Minimum Wage	-0.081	-3.140**

** Significant at 1% level using table from Im, Pesaran and Shin (1997). Unmarked statistics are not significant at the 10% or better level. Labor force and unemployment regressions include dummy variable for break in 1989. The fraction of employed teens at the minimum wage includes dummy variables for breaks starting in 1990:2 and 1996:4 (when federal minimum wage hikes commenced).

Critical values for second column are -1.81 (1%), -1.73 (5%), and -1.68 (10%).
Critical values for third column are -2.44 (1%), -2.36 (5%), and -2.32 (10%).

G. Tests for Stationarity

There is good reason to doubt results derived from using nonstationary variables in time-series regression (see Granger and Newbold (1974) and Phillips (1986)). I tested the variables used in the regression for stationarity, using, in part, the procedures recommended by Enders (1995).

The key variable is the labor force participation rate of teenagers (LFPRteens). Using quarterly data, working down from a high number of lags, nine lags proved to be significant. With nine lags, I could not reject the null hypothesis that the residuals were stationary white noise. Examination of the data indicated a structural break in 1989. This break most likely is related to a change in the CPS coding, when unemployment was changed from "looking" for employment to "looking" or being "laid off."

The Im, Pesaran and Shin *t*-bar test (1997) for panel data was used. This procedure runs the

Table 2
Regression on Normit of Teenage Labor Force Participation
1983:1-1999:4

Variable (lags)	Coefficient	Std. Error	T-Statistic	Probability
ΔMW 1979-81	-0.2213	0.0609	-3.632	0.0003
ΔMW 1990-91	-0.2053	0.0451	-4.547	0.0000
ΔMW 1996-97	-0.1630	0.0798	-2.040	0.0414
Δ Fraction Teens in Population	-0.1643	0.1544	-1.059	0.2901
Δ Log Per Capita Income	0.29349	0.1817	1.614	0.1066
Δ Log Wage 30-39	-0.0343	0.0333	-1.028	0.3040
Unemployment rate	-0.6450	0.1688	-3.8233	0.0001
Unemployment (-1)	-0.1840	0.1699	-1.083	0.2728
Unemployment (-2)	-0.0155	0.1707	-0.091	0.9278
Unemployment (-3)	-0.4086	0.1692	-2.408	0.0161
Unemployment (-4)	-0.5323	0.1689	-3.164	0.0016
Unemployment(-5)	-0.5426	0.1671	-3.254	0.0012
Unemployment(-6)	0.1295	-0.1641	0.756	0.4300
Unemployment(-7)	-0.3707	0.1604	-2.315	0.0208
LFPR 30-39	0.0774	0.1143	0.674	0.5003
LFPR 30-39 (-1)	0.1035	0.1143	0.906	0.3651
LFPR 30-39 (-2)	-0.2104	0.1148	-1.834	0.0667
LFPR 30-39 (-3)	0.1294	0.1152	1.123	0.2612
LFPR 30-39 (-4)	0.0432	0.1143	0.378	0.7053
SHIFT 1989	-0.0263	0.0111	-2.391	0.0169

Other Variables: Seasonal Dummy Variables, State Fixed Effects, State Specific Time Trends, Indicator variable if state had higher minimum wage;
Error Structure: AR to 9 Lags,[1] Weighted Regression
Adjusted R-square 0.7877, Durbin-Watson statistic 2.019

augmented Dickey-Fuller test separately on each panel member (here, states). The change in the key variable is run on its level and past changes (here, nine lags). The *t*-statistic for the variable's level, if it is significant, results in a rejection of the null hypothesis that the process is not stationary (that is, it results in "accepting" the series as stationary). The Im, Pesaran, and Shin *t*-bar statistic is the average of these *t*-statistics over panel members. I report two versions of this statistic in Table 1: The first is for the deterministic term being a constant, the second is for having a constant plus a linear time trend. The labor variables include a dummy variable for the 1989 break (equal to 1 after 1989, 0 before). Table 2 shows the results for all the variables used in the regressions.

To assure that they are stationary, the first difference of per capita income (in log form) of the fraction of teenagers in the population and of the wage rate (in log form) were used in the regressions.

IV. THE IMPACT OF MINIMUM WAGES ON LFPR

A. The Effect on Teenagers

The labor force participation rates were calculated for each state and quarter from 1979 through 1999. For purposes of these data, a teenager is defined as a person 15 to 19 years old. Regressions were run on the normits of the labor force participation rate. These results used a two-step process to weight the secondstage regressions. The weight equals the inverse of an approximation of the variance (see Greene, 2000):

$$\text{(3)} \qquad \frac{\Phi(L^*) \bullet (1 - \Phi(L^*))}{N_{it} \bullet \phi^2(L^*)}$$

where L* is the predicted labor force participation rate from the first step, $\phi(*)$ is the normal density, and $\Phi(*)$ is the cumulative normal distribution. Amemiya (1985) has shown that this procedure has the same asymptotic distribution as the maximum likelihood estimator. The second set of state regressions is estimated using conditional maximum likelihood (given the weighting from the first step).[18] As noted above, the weighting procedure produced homogenous variances across states. Alternative weighting by a feasible generalized least squares (FGLS) estimate of the cross-sectional variances gave very similar results.

Table 2 shows the results of the weighted regressions for teenagers. Appendix Table 1 shows the mean and standard deviations of all variables. The minimum wage variables are those discussed above: The sequentially summed increases in the minimum wage over the period, weighted by the relative impact measure derived from the fraction of teenagers receiving the minimum wage. Appendix Table 2 shows the value of this variable for the states with the federal minimum wage.

All minimum wage hikes had a significantly negative impact on the labor force participation rates of teenagers. I discuss the sensitivity of these results to the form of the regression in subsequent sections.

The fraction of teenagers in the working population (the ratio of teenagers to persons ages 15 to 54) was insignificant. Increases in personal per capita income were modestly significant and positive. Most likely, this reflects an overall income effect on the demand for teenage labor in the economy. The change in the log of wages (of adults 30 to 39 years old) was insignificant. This insignificance was true of other wage variables (including the wage rate of adults ages 20-24, 25-29, and even the teenage wage rate) in alternative versions of this equation.

The set of unemployment rates was highly negative and significant. The minus terms in parentheses indicate how many periods the variable was lagged. Unemployment rates proved to be significant up to seven lags. This may in part reflect the delay between an upturn in the economy and the increase in teenage employment.

The set of labor force participation rates (for ages 30-39) was significant only at the second lag (this varied for other groupings of teenagers). The unemployment variables evidently captured most of the relevant shocks to the labor market.

The shift variable (reflecting the change in the way unemployment was treated in the CPS survey after 1989) was also significant. A Chow break test[19] was run on the data. The results ($F = 1.50$, $p = 0.0002$) indicate that the nature of the equation changed between these periods. Running the regression over the 1979-1988 period alone, the 1979 minimum wage variable became insignificant (–0.3551, with a standard error of 0.2583, $t = -1.41$, $p = 0.157$). This was mainly due to the dropped quarters due to the 7 lags in the unemployment variable. In this subsample, lags 5, 6 and 7 were insignificant (a test failed to reject the null hypothesis that their combined value was zero, with an F of 0.125, $p = 0.772$). Running the same regression with only four lags in the unemployment variable, the 1979-1981 hikes had a significantly negative estimated impact (–0.2862 (0.1441), $t = -1.99$, $p = 0.0471$). Running the regression from 1989 through 1999 did not change the negative effect of the subsequent minimum wage hikes (for the 1990-1991 hikes, –0.2129 (0.0499), $t = -4.26$, $p < 0.0000$; for 1996-1997 hikes, –0.1714 (0.0789), $t = -2.17$, $p = 0.0300$).

Both the constant term and the linear time trend were treated as state-specific effects. The null hypothesis, that the state fixed effects do not make a significant contribution to explaining the variation in the dependent variable, was rejected with an *F* statistic of 412.2 ($p < 0.0000$). The null hypothesis that separate linear time trends (one for each state) do not make a significant contribution (compared to a common time trend) was also rejected, with an *F* statistic of 131.9 ($p < 0.0000$). Apparently economic conditions varying from state to state have a significant impact in the overall relationship between minimum wages and labor force participation rates.

The estimates of the minimum wage's impact were similar for various forms of the estimating equation. For example, removing the labor force participation rates of 30- to 39-year-olds left the effects unchanged (for the respective set of hikes, the coefficients are –0.2276 ($p = 0.0001$), –0.2052 ($p < 0.0000$), –0.1622 ($p = 0.0423$)). Removing the AR(9) error structure also had little effect except for the first set of hikes (for the respective hikes, the coefficients are -0.1079 ($p = .0123$), –0.2138 ($p = 0.0000$), –0.1561 ($p = 0.0464$)). Replacing the state-specific time trends with a common trend also had little effect (for the respective hikes, the coefficients are –0.1615 ($p = 0.0298$), –0.1703 ($p = 0.0004$), –0.1765 ($p = 0.0262$)).

Replacing the fixed effects and the state-specific time trends with a common intercept and time trend reduced the size and significance of only the first set of hikes (for the respective hikes, the coefficients are –0.1305 ($p = 0.1225$), –0.1903 ($p = 0.0002$), –0.1791 ($p = 0.0168$)). Since the fixed effects and state-specific time trends are significant, these results can be discounted. Over a variety of specifications, therefore, minimum wages significantly reduced the teenage labor force participation rate.

On the other hand, removing the unemployment rates reduced the significance of the 1996-1997 hikes (for the respective hikes, the coefficients are –0.2631 ($p = 0.0001$), –0.3159 ($p < 0.0000$), and –0.0976 ($p = 0.2229$)). The 1996-1997 hikes took place in an economic expansion. I interpret this as saying that without controlling for the business cycle with this set of lagged unemployment rates, it is not possible to discern the true impact these hikes had.

The 1996-1997 effects were also smaller in the unweighted regression (–0.1227, $p = 0.1634$). An alternative weighting (using a FGLS estimator of the cross-sectional variances) produced a more significant coefficient for the 1996-1997 hikes (with –0.1401, $p = 0.0594$, with the other sets of hikes also being highly significant and negative: –0.1409, $p = 0.0040$; –0.1899, $p = 0.0002$). In addition, using the White heteroskedasticity-consistent estimator, the 1996-1997 hikes' coefficient was even more significant (–0.1401, $p = 0.0224$). The use of weights (by eliminating heterogeneity and by giving greater weight to the cells with larger sample size) makes the estimations more precise, allowing the 1996-1997 hikes' effects to be discerned. The weighted regression on labor force participation rates (not their normits) gave similar results.[21] The choice to apply weights is important, but the particular set of weights used does not seem to matter.

Running the regression with monthly data showed that the minimum wage had a negative effect. The coefficients were much more varied in size, giving what appear to be unreasonable results (for example, the set of hikes with the smallest increase in the minimum wage (the 1996-1997 hikes) had the most negative impact).[22]

Table 3
Effect of Minimum Wages on Alternative Measure of Teenage Labor Force Participation

Dependent Variable: Normits of Ratio of Teenage LFP to Population 1983:1-1999:4

Employed plus those making contact with employer or employment agency

Variable	Coefficient	Standard Error	T-Statistic	Probability
Δ MW 79-81	-0.1857	0.0531	-3.497	0.0005
Δ MW 90-91	-0.1573	0.0447	-3.520	0.0005
Δ MW 96-97	-0.1113	0.0797	-1.398	0.1622
Regression: 1994:1-1999:4				
Δ MW 96-97	-0.2018	0.0719	-2.805	0.0051

Equation is identical to that shown in Table 2.

Table 4
Effect of Minimum Wages on Teenage Employment
Dependent Variable: Normit of Ratio of Teenage Employment to Population 1983:1-1999:4

Variable	Coefficient	Standard Error	T-Statistic	Probability
Δ MW 79-81	-0.3647	0.0622	-5.861	0.0000
Δ MW 90-91	-0.2652	0.0469	-5.649	0.0000
Δ MW 96-97	-0.0795	0.0823	-0.9646	0.3348

Equation is identical to Table 2's, except for the elimination of the labor force participation rates of 30 to 39 year olds.

Regressions were also run with state-specific Frac_MW.[23] Using this variable in place of the minimum wage variables used above in a regression similar to those in Table 2: the coefficient of the state-specific Frac_MW was –0.0458, with a standard error of 0.0247, a t-statistic of –1.848, and p = 0.0647. The state-specific Frac_MW is estimated from a small sample. As a result, its coefficient may be understated due to the measurement error. To test this hypothesis, a three quarter average of Frac_MW was used, centered on the current period (the average of lead, current and lagged Frac_MW). The coefficient in this case was more negative and significant: –0.1313, with a standard error of 0.0357, a t-statistic of –3.658, and p < 0.0003. These results show that as more teenagers are affected by the minimum wage, fewer of them want to work.

These results are for the normits of the teenage labor force participation rate. The relationship between these coefficients and the change in the actual labor force participation rate (L) is:

$$\frac{dL}{dMW} = \phi(L) \bullet \beta \quad (4)$$

where $\phi(L)$ is the density of the normal curve at L and β is the coefficient of the minimum wage (or whatever variable is of interest). For example, if L = 0.5, $\phi(0.5) = 0.3989$.

Table 9 summarizes the adjusted minimum wage coefficients and the respective elasticities, for all tables in this paper.

Using the results from Table 2, for the 1978-1981 hikes, the teen labor force participation rate went down 3.615 percentage points, or by 6.85% (using as a base the 1979 LFPR of 0.5278). The elasticity of L to MW for this set of hikes was –0.167. For the 1990-1991 set of hikes, the teen labor force participation rate went down by 2.07 percentage points, or by 4.09% (using the 1989 LFPR of 0.5059 as the base). The elasticity of L to MW for this set of hikes was –0.162. For the 1996-1997 hikes, the teen labor force participation rate went down by 1.31 percentage points, or by 2.78% (using the 1995 LFPR of 0.4712 as the base). The elasticity of L to MW for this set of hikes was –0.138. The elasticity of labor force participation rates to the minimum wage is in the range of the elasticity of employment to the minimum wage in Brown et al. (1982).

It is of interest to compare these estimated impacts with those estimated from using Frac_MW. The 1978-1981 hikes increased Frac_MW from 18% to 54%. The Frac_MW's adjusted coefficient, converted to elasticity form, is –0.109, implies that this decreased the LFPR of teenagers by 3.9%. The 1990-1991 hikes increased Frac_MW from 16% to 40%: This implies a 2.6% decrease in the labor force. The 1996-1997 hikes increased Frac_MW from 16% to 36%: this implies a 2.2% decrease in the labor force.

These results (3.9%, 2.6%, and 2.2%) approximate the results derived from Table 2 (6.8%, 4.1%, and 2.8%), Thus, two different forms of estimating the impact of minimum wages yielded similar results.

One criticism of using labor force participation (the sum of employed plus unemployed) is that the number of unemployed is based on the number of positive responses to what can be considered a vague question: "Have you looked for work?" Looking for work can include anything from looking at want ads to visiting employers. To answer this criticism, a regression was run on a more restrictive measure of the labor force participation rate. In addition to the employed, this measure only adds as unemployed those who directly contacted employers or who contacted an employment agency (either public or private). Omitted are those who only contacted friends or relatives, contacted school employment centers, sent out resumes, filled out an application, looked at ads or took job training. Over the sample, the LFPR of teenagers averaged 48.06%; the more restrictive measure averaged 46.99%. Since the employment rate of teens was 40.01%, using the more restrictive measure reduced the measured unemployment rate from 16.8% to 14.9%.

Table 3 shows the results. The minimum wage had a significantly negative effect for all hikes except those in 1996-1997. In 1994, the CPS changed their questions to the unemployed regarding their method of seeking work.[24] To determine if this change could explain the weak 1996 results, the regression was run on data from the period 1994:1 to 1999:4. Taking into account this change, the 1996-1997 set of hikes did have a statistically negative effect on the restricted LFPR of teens.

Although the focus of this study is on labor force participation rates, a similar regression was run over teenage employment (only the labor force participation variables were removed). The results are shown in Table 4.

The effects on employment for the 1978-1981 hikes and the 1990-1991 hikes are large, negative,

Table 5
Effect of Minimum Wages On Various Teen Groups
(Standard Error in Parentheses)
Regression Form: See Table 2

GROUP	Δ MW79-81	Δ MW90-91	Δ MW96-97	State Specific FRAC_MW (separate regression)
Teens 15-17	-0.3052	-0.1923	-0.1815	-0.0702
	(0.0827)	(0.0584)	(0.1009)	(0.0475)
	p = 0.0002	p = 0.0010	p = 0.0722	p = 0.1397
Teens 18-19	-0.0365	-0.1612	-0.1044	-0.1574
	(0.0905)	(0.0684)	(0.1264)	(0.0540)
	p = 0.6863	p = 0.0185	p = 0.4088	p = 0.0036
White Teens	-0.2886	-0.1652	-0.1824	-0.1508
	(0.0637)	(0.0473)	(0.0860)	(0.0391)
	p < 0.0000	p = 0.0005	p = 0.0345	p < 0.0001
Black Teens	+0.1572	-0.2861	-0.2721	-0.1042
	(0.1770)	(0.1225)	(0.2098)	(0.0665)
	p = 0.3745	p = 0.0197	p = 0.1945	p = 0.1176
Male Teens	-0.1486	-0.2120	-0.2746	-0.1747
	(0.0796)	(0.0593)	(0.1069)	(0.0465)
	p = 0.0621	p = 0.0004	p = 0.0102	p < 0.0002
Female Teens	-0.3230	-0.2253	-0.0602	-0.1161
	(0.0823)	(0.0600)	(0.1081)	(0.0482)
	p = 0.0001	p = 0.0002	p = 0.5775	p = 0.0160

Table 6
Effect of Minimum Wages on Young Adults Ages 20-24
Dependent Variable: Normit of Labor Force Participation Rate

Variable	Coefficient	Standard Error	T-Statistic	Probability
Δ MW 79-81	-0.0829	0.0885	-0.9367	0.3490
Δ MW 90-91	-0.0378	0.0486	-0.7739	0.4390
Δ MW 96-97	+0.1231	0.0985	+1.2485	0.2119
Δ Fraction of young adults in population	+0.5182	0.1647	+3.146	0.0017
State-Specific Frac_MW for young adults (separate regression)	-0.0772	0.0884	-0.9137	0.3610

Equation is identical to Table 2's, except fraction of young adults in population is used in place of fraction of teenagers. 1983:1 to 1999:4

and significant. On the other hand, the 1996-1997 hikes had a small and insignificant effect.[25] In the standard model of minimum wages, it would be difficult to explain why the 1996-1997 minimum wage hikes did not increase labor force participation rates. If employment remained unaffected while wages were increased, the standard model would lead one to expect that more people would want to work, especially because this was a period of expansion. The fact that this did not occur suggests that forces other than employment affecting labor force participation played an important role. An interesting possibility is that employers have over time shifted more toward adjusting to the minimum wage by changing nonwage compensation instead of employment. The use of computers for monitoring workers and other changes in labor market conditions may be making this possible.

B. The Impact of the Minimum Wage on Labor Force Participation of Specific Groups of Teenagers

Using the form of the regression in Table 2, the impact of the minimum wage was estimated for different groupings of the teenage population.[26] As teenagers are divided into different groups, the sample size becomes smaller. Consequently, the results are more subject to measurement error. It is not surprising that groups with the fewest numbers in the sample (for example, blacks) did not have significant results.

The form of the regression is that used in Table 2. Table 5 shows the results for two sets of minimum wage variables. The first is for the combined hikes, the second is for the state-specific fraction of employed teenagers at the minimum wage (averaged over three periods, centered on the current period).[27]

The effect of minimum wages on blacks was marginal.[28] Whites, males, and females were all negatively affected. Only the older teenagers appear to have been less negatively affected by the minimum wages. The Frac_MW variable had a significantly negative effect on all groups except blacks and teenagers ages 15 to 17.

C. The Effect of Minimum Wages on Labor Force Participation of Other Groups

The impact of the minimum wage on young adults (ages 20 to 24) was also investigated. The fraction of young adults at or below the minimum wage was used in the weighting of the minimum wage variables. Table 6 presents the results. The minimum wage had no significant effect on this group. This is confirmed by the separate regression run with the state-specific Frac_MW variable (measuring the fraction of young adults earning the minimum wage). Of the other variables, the change in the fraction of young adults in the population had a significant positive impact, corresponding to the results found in Neumark and Korenman (1997).

A similar regression was run to see if the minimum wage variable (without the weights in equation 2) lagged for 1 year, 2 years, 3 years or 4 had any effect on the labor force participation of this group. This is an indirect test of whether the minimum wage induced substitution of leisure time toward one's older years. If this did occur, then the reduction in labor force participation by teenagers could be explained by an inward shift in the supply curve. All lags, together or separately, were highly insignificant.

A regression was also run on high school graduates and, separately, on high school dropouts, ages 18 to 24. To weight the minimum wage variables, the fraction of employed teenagers earning the minimum wage or less was used (results are similar to using the fraction of young adults). Table 7 shows the results, which are mixed overall. Graduates appear to have fared worse (the combined minimum wage variables were negative with a significance level of 0.065, whereas for dropouts, they were positive with a significance level of 0.55). This result, however, may be due to the low number of dropouts (a mean cell size of 42, compared with 175 and 118

Table 7
Effect of Minimum Wages on High School Dropouts and Graduates Ages 18-24
(Standard Error in Parentheses) Regression Form: See Table 2

GROUP	ΔMW79-81	ΔMW90-91	ΔMW96-97	State-Specific FRAC_MW (separate regression)
Dropouts	+0.1218	-0.1631	+0.1616	-0.1472
	(0.1091)	(0.0759)	(0.1450)	(0.0633)
	p= 0.2643	p= 0.0319	p=0.2653	p=0.0201
Graduates	-0.1144	-0.0400	-0.0723	-0.0839
	(0.0665)	(0.0470)	(0.0919)	(0.0377)
	p=0.0855	p=0.3945	p=0.4318	p=0.0261

Table 8
Effect of Minimum Wages on Teenagers Ages 16–18, Labor Force Participation (Normit Form) Based Upon School Attendance
(Standard Error in Parentheses)
Regression Form: See Table 2, Summer Quarter Excluded

GROUP	ΔMW90-91	ΔMW96-97	State-Specific FRAC_MW (separate regression)
Attending School	-0.0751	-0.0650	-0.0715
	(0.0236)	(0.0422)	(0.0517)
	p= 0.0014	p=0.1240	p=0.1670
Not Attending School	-0.2490	-0.0901	-0.0753
	(0.0549)	(0.0999)	(0.1183)
	p=0.0000	p=0.3637	p=0.5350

for graduates), resulting in a large sampling error. Both groups appear to be significantly and negatively affected by the state-specific Frac_MW.

D. The Effect of Schooling Status

Beginning in 1984, the CPS asked if a person 16 to 24 years old was attending or enrolled in school (high school or college) last week. If the person was currently on a holiday or school vacation, then they counted as being enrolled. On the other hand, if they were on summer vacation, they were not counted as being enrolled. Consequently, this question can only be used during the school year to distinguish between those in and those out of school. To account for this, the sample was taken only over the school year (excluding the months of June, July and August). The data are still in "quarterly" form, but the quarters have been redefined to fit the school year (for example, March, April and May are combined into one of the new quarters). The sampling for all variables is for the newly defined quarters.

The labor force participation rate for teenagers ages 16 to 18, both for those in school and those not in school, was calculated for 1984 through 1999. Regressions similar to those above were run, with one important exception: The number of observations proved to be insufficient to estimate an error structure taking into account the absence of the summer quarter. This may not be a serious problem since the absence of autoregression in the error term did not affect the results for teenagers. Because of the 1984 start date, the variable reflecting the 1978 set of minimum wage hikes was not included.

Table 8 reports the results from these regressions. The state-specific FRAC_MW is the average of the forward, current, and previous quarters (results were similarly insignificant for the current FRAC_MW). The results for the 1990-1991 hike show that the minimum wage hike more negatively affected those not attending school. This corresponds to the results for high school dropouts (compared with high school graduates). This set of hikes also decreased employment (especially when compared with the 1996-1997 hikes).

E. A Summary Table

Table 9 summarizes the results from the tables above, making the adjustments shown in equation 3. The first number shown is the adjusted coefficient. The second number is the elasticity of LFPR with respect to the minimum wage. The final number is the significance level. The results for state-specific Frac_MW are from a separate regression. It shows the effect of its three-quarter averaged value, centered on the current quarter.

Two results emerge from the table. First, older groups are less affected by the minimum wage. Second, the minimum wage did not significantly affect any group positively.

V. CONCLUSION

One of the challenges of measuring the effects of the minimum wage is to find a way to determine how many workers it affects and for how long its constraint on wages persists. The fraction of teenagers earning the minimum wage both empirically and logically measures this constraint. A second challenge is to control for the effects of the business cycles. I argue that the many steps taken (including using AR(9) and using multiple lags in the unemployment rate and in other variables related to the business cycle) meet this objective.

The findings in this paper, when combined with past studies, make a strong case that the minimum wage decreases labor force participation. This effect may not be solely due to its negative effect on employment. I suggest that firms also cut back on nonwage compensation. In addition, if the minimum wage also reduced job turnover, the cost of looking for work would be higher. I argue that the decline in labor force participation reflects the minimum wage reducing the value of being in the labor force for most workers. Since the minimum wage apparently decreases labor force participation, those who argue that the minimum wage makes workers better off are left to answer the question, "If that is so, why don't more people want to work?"

Table 9
Summary of Adjusted Coefficients
Adjusted Coefficient
Elasticity
Significance Level

GROUP	1978-1991 Hikes	1990-1991 Hikes	1996-1997 Hikes	Frac_MW
Teens All	–0.088	–0.082	–0.065	–0.052
	–0.167	–0.162	–0.138	–0.109
	0.0003	0.0000	–0.0414	0.0000
Teens 15-17	–0.116	–0.078	–0.069	–0.027
	–0.310	–0.195	–0.184	–0.071
	0.0002	0.0010	0.0722	0.1397
Teens 18-19	–0.013	–0.059	–0.038	–0.058
	–0.020	–0.090	–0.059	–0.088
	0.6863	0.0185	0.0030	0.0036
Young Adults 20-24	+0.024	–0.011	+0.036	–0.023
	+0.031	–0.014	0.046	–0.029
	0.3490	0.4390	0.2119	0.3610
White Teens	–0.115	–0.066	–0.131	–0.060
	–0.228	–0.131	–0.259	–0.119
	0.0000	0.0005	0.0345	0.0001
Black Teens	+0.058	–0.106	–0.100	–0.038
	+0.168	–0.305	–0.290	–0.111
	0.3745	-0.0197	0.1945	0.1176
Male Teens	–0.059	–0.085	–0.110	–0.070
	–0.120	–0.172	–0.222	–0.141
	0.0621	0.0004	0.0102	0.0002
Female Teens	–0.129	–0.089	–0.024	–0.046
	–0.275	–0.191	–0.051	–0.099
	0.0001	0.0002	0.5775	0.0160
High School Dropouts 18-24	+0.046	–0.061	+0.061	–0.056
	+0.074	–0.098	+0.099	–0.090
	0.2643	0.0319	0.2659	0.0201

Continued on next page...

Table 9 (Continued)				
GROUP	**1978-1991 Hikes**	**1990-1991 Hikes**	**1996-1997 Hikes**	**Frac_MW**
High School Graduates				
18-24	–0.034	–0.012	–0.022	–0.025
	–0.044	–0.015	–0.028	–0.032
	0.0855	0.3945	0.4318	0.0261
Teens Attending School				
16-18	N/A	–0.0274	–0.0237	–0.0261
		–0.0641	–0.0554	–0.0610
		0.0014	0.1240	0.1670
Teens Not Attending				
School 16-18	N/A	–0.0779	–0.0282	–0.0230
		–0.1118	–0.0405	–0.0330
		0.0000	0.3674	0.5350

Endnotes

[1] Williams and Mills (1998) also use current time-series methodology. Using vector autoregression to analyze the effects of minimum wages on the employment rates of teen males and females, they found a significantly negative effect. See also Parks and Ratti (1998) who looked at aggregate data for the years 1954 to 1992. They found the minimum wage had no effect on teenage employment. The reasons for this nonsignificance may be the use of aggregate data.

[2] Many different search models have been applied to minimum wages. Wessels (1980a and 1980b) presents a very simple version. Thomas Carter (1998) analyzes Rebitzer and Taylor's efficiency wage model (1995) showing that when the minimum wage increases employment in this model, job searchers are made worse off (V goes down). Lang and Kahn (1998) present a bilateral search model with heterogeneous workers where the minimum wage makes low productivity workers worse off while not enhancing the welfare of more productive workers. Mortensen and Burdett (1989) present an equilibrium model where larger firms pay higher wages and have lower turnover while smaller firms are, in a sense, more competitive, paying a wage closer to V. While they definitely do not come to this conclusion, a minimum wage lowering V would make all workers in their model worse off.

[3] Besides employment, some of the minimum wage literature has focused on unemployment, using the criteria that "if unemployment goes up, workers are worse off." Unfortunately, this is not the case. If unemployment went down, as long as employment does not increase, L must be lower and we can conclude that labor force entrants are definitely worse off. If unemployment went up, only if it goes up enough to increase L can we say that it makes workers better off. Readers of the minimum wage literature will realize that this is exactly opposite of how changes in unemployment have been interpreted. A better variable for detecting the change in V is labor force participation.

[4] The reduction of work hours most likely go hand-in-hand with increased work pace and forcing workers to work off the clock. Neumark, Schweitzer and Wascher (2000) found that the minimum wage reduced work hours of low-wage workers to the extent that earned income declined.

[5] Neumark and Wascher (1998) review the literature on this subject and present evidence that the minimum wage does reduce on-the-job training. Acernoglu and Pischke (1999) found no effect. Both used the presence of a formal training program as the dependent variable.

[6] Wessels (1980a and 1980b) found that the minimum wage increased the quit rate in low-wage industries. It was also found that the quit rate went up more, the lower the wage the industry paid. Unfortunately, turnover data is no longer available to verify if this pattern occurs today.

[7] When I investigated the effects of minimum wages on turnover rates in lower-wage industries in 1981, I could find no consistent patterns (these results were not published). Turnover data is no longer available to see if this pattern persists. The only other evidence I could find is in Light and Ureta (1990) who estimated hazard functions for voluntary job terminations. They estimated these functions for persons, who at the start of their sample, were in their first job and whose ages ranged from 14 to 24. They controlled for industry and many other factors affecting voluntary terminations. They also included a set of yearly dummy variables. The years in which the minimum wage was increased (in this case, 1967-1968 and 1974-1976) either had the same termination rates as other years or, when significant, a lower rate. As the goal of the paper was other than the effects of minimum wages, these results are by no means definitive.

[8] See Neumark, Schweitzer and Wascher (1998) for a discussion of the literature on the impact of minimum wages on family income. The overall result is that the minimum wage has little impact. Even the positive impacts that are found are too small to have a substantial income effect. See also Addison and Blackburn (1999) who found a positive effect for junior high dropouts. They also found some support for the minimum wage increasing the income of teenagers.

[9] A normit is the normal equivalent deviate (Greene 2000, p. 836). If the labor force participation rate is P and the normal distribution is F(z), then the normit of P is z .

[10] See Bera and Jarque (1981). The statistic weights skewness and kurtosis around their normal value. It has a chi-squared distribution with two degrees of freedom under the null hypothesis of normally distributed errors.

[11] Bartlett (1937). The test compares the logarithm of the weighted average variance with the weighted sum of the logarithms of the variances. Under the joint null hypothesis that the sample is normally distributed and that the subgroups have equal variances, the test statistic is approximately distributed with a chi-squared distribution.

[12] Burkhauser, et. al. (1998) extensively review the various regression models used in the minimum wage literature. The form shown here was used in Card, Katz and Krueger (1994), Neumark and Wascher (1992), and Card and Krueger (1995).

[13] An alternative method is explored in Baker et. al. (1999). The method presented here has the advantage of being based on a direct measure of the number of affected workers.

[14] This was not true for the 1978-1981 set of hikes. In May 1978, 54.8% of working teenagers were being paid the minimum or less, compared to 48.2% in January 1979, 50.35% in January 1980, and 52.17% in January 1981. Note that before 1979, wage data was only available in the May CPS.

[15] An alternative is to sum the hikes and use this sum from the start (which may be appropriate, depending on the nature of adjustment costs). This form of the variable gave very similar results to the variable used here (which adds the hikes as they occur).

[16] See, for example, Hsiao (1979) and Collado (1997).

[17] The small sample properties of measurement error and the ADF test have not been dealt with in the econometric literature to my knowledge.

[18] The weighting was so that the unconditional residual of each lag was weighted by weights from its period. In theory, this assures that the residual structure has the same stationary structure across all periods.

[19] See Greene (2000), pp. 287-293. The Chow test compares the restricted residual sums of squares from the regression run over the whole period with the sum from separate regressions run over the subperiods.

[20] The AR coefficients are:

	coef.	std. error	t	p
AR(1)	-0.019581	0.018682	-1.048076	0.2947
AR(2)	-0.050493	0.018739	-2.694491	0.0071
AR(3)	-0.009292	0.018736	-0.495957	0.6200
AR(4)	0.186987	0.018703	9.997599	0.0000
AR(5)	-0.022096	0.018998	-1.163074	0.2449
AR(6)	-0.086536	0.018532	-4.669516	0.0000
AR(7)	-0.037911	0.018644	-2.033391	0.0421
AR(8)	0.013684	0.018701	0.731742	0.4644
AR(9)	-0.083022	0.018540	-4.477960	0.0000

[21] Table 8 gives the converted normit results(-0.088, -0.082, and -0.065) for comparison to these linear coefficients. The results from the regressions on the Labor Force Participation Rate were

1979-1981:	-0.0835 (0.0229)	t=-3.66	p=0.0003
1990-1991:	-0.0780 (0.0173)	t=-4.52	p=0.0000
1996-1997:	-0.0608 (0.0302)	t=-2.01	p=0.0442

[22] The results on monthly data were:

1979-1981:	-0.1193 (0.0370)	t=-3.22	p=0.0013
1990-1991:	-0.1489 (0.0460)	t=-3.23	p=0.0012
1996-1997:	-0.4075 (0.0870)	t=-4.68	p=0.0000

[23] The fraction used for weighting the minimum wage variable was from U.S. aggregate data (modified to reflect the date and whether a higher state minimum wage was imposed). The fraction used here is from each state, reflecting the minimum wage of the state or the federal minimum wage — whichever is higher.

[24] Prior to 1994, the CPS asked of those looking for work " what has ... been doing in the last 4 weeks to find work", with the examiner checking all methods used. Each method was recorded in as a separate variable, with a "yes" or "no" response. In 1994, of those unemployed and looking for work, CPS asked "what are all of the things ... done to find work during the last 4 weeks?" In this case, each variable recorded one of the methods used.

[25] Running the same regression over the 1994:1 to 1999:4 period, the 1996-1997 hikes had a significant negative coefficient (-0.1529 (0.0726), t statistic of -2.105, and p = 0.0355).

[26] The Im, Persaran and Shin t-bar statistic shows that the labor force participation rates of groups are stationary variables. The first t-bar statistic is for a constant term and a break term for the period 1989 forward. The second t-bar statistic adds a time trend. All estimated include 9 lags (except for black teenagers, for whom errors beyond 2 lags were correlated to such a degree as to produce a singular matrix). All reject the null hypothesis of unit root at the 1% level

Group	IMS t-bar statistic (constant term)	IMS t-bar statistic (c plus time trend)
White Teens	-2.504	-2.781
Black Teens	-4.207	-4.257
Male Teens	-2.557	-2.670
Female Teens	-2.601	-2.686
Teens 15-17	-2.488	-2.554
Teens 18-19	-2.773	-2.907
High School		
Dropouts (18-24)	-.2.449	-2.517
HS Grads (18-24)	-2.333	-2.470
Young Adults (20-24)	-2.325	-2.458

For the first column, nonstationarity is rejected at the 1% level as the t-bar statistic < -1.81. For the second column, it is -2.44.

[27] The variable was averaged to reduce measurement error. The nonaveraged fraction was significantly negative for all groups except teens 15-17 (with a coefficient of +0.0824 (standard error of 0.0328), p = 0.0121), black teens (with a coefficient of -0.0273 (0.0669), p = 0.6831), and female teens (with a coefficient of -0.0458 (0.0248), p = 0.0647). Teens 18-19 had a coefficient of -0.1432 (0.0388), p = 0.0002. White teens had a coefficient of -0.0601 (0.0282), p = 0.0332. Male teens had a coefficient of -0.1100 (0.0336), p=0.0011.

[28] For black teens, the state-specific time trend variable had to be dropped and replaced with a common time trend variable to avoid singularity. Singularity was caused by an excessive number of cells with zero or one as the labor force participation rate.

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Appendix Table One

Means and Standard Deviation of Variables Averaged Across States and Years

Variable	Mean	Standard Deviation
LFPR Teenagers	0.4806	0.1036
Cell Size Teenagers	161	127
LFPR 15-17	0.3734	0.1195
Cell Size 15-17	99.7	77.5
LFPR 18-19	0.6564	0.1045
Cell Size 18-19	61.9	50.4
LFPR 20-24	0.7821	0.0589
Cell Size 20-24	154.8	130.3
LFPR White Teens	0.5042	0.1051
Cell Size White Teens	132.9	104.7
LFPR Black Teens	0.3464	0.1990
Cell Size Black Teens	28.7	30.5
LFPR Male Teen	0.4931	0.1148
Cell Size Male Teen	81.3	63.5
LFPR Female Teen	0.4683	0.1151
Cell Size Female Teen	80.3	63.5
LFPR Graduates	0.7747	0.0596
Cell Size Graduates	174.7	142.6
LFPR Drop Outs	0.6230	0.1213
Cell Size Drop Outs	42.0	41.9

Appendix Table One (Continued)

Variable	Mean	Standard Deviation
LFPR 30-39	0.7951	0.0487
Unemployment White Males 30-54	0.0396	0.0235
First Difference: Log Per Capita Income	0.01447	0.0152
First Difference: Fraction of Teenagers in Population (15-54)	-0.000535	0.0170
First Difference: Fraction of 20-24	-0.000672	0.0185
First Difference: Log of Wage 30-39	0.00932	0.0763
"Restricted" LFPR Teens	0.4699	0.1057
Fraction Teens At Minimum Wage: US Aggregate	0.3124	0.1024
Fraction Teens At Minimum Wage: State Specific	0.3209	0.1539
Fraction 20-24 at Minimum Wage: US Aggregate	0.1126	0.0336
Fraction 20-24 at Minimum Wage: State Specific	0.1123	0.0704
Normit of LFPR Teens	-0.5018	0.2696
Employment Rate Teens	0.4001	0.1036

Appendix Table Two
Weighted Minimum Wage Variable

Minimum Wage 1979-1981

1979:1	0.243819	1985:4	0.253410
1979:2	0.194684	1986:1	0.214874
1979:3	0.159342	1986:2	0.197921
1979:4	0.147950	1986:3	0.185361
1980:1	0.329835	1986:4	0.177192
1980:2	0.310629	1987:1	0.183288
1980:3	0.285545	1987:2	0.128411
1980:4	0.272708	1987:3	0.140852
1981:1	0.41048	1987:4	0.112681
1981:2	0.406211	1988:1	0.080852
1981:3	0.354872	1988:2	0.088171
1981:4	0.367068	1988:3	0.079268
1982:1	0.323286	1988:4	0.072561
1982:2	0.336824	1989:1	0.064023
1982:3	0.325117	1989:2	0.064388
1982:4	0.338408	1989:3	0.026706
1983:1	0.358288	1989:4	0.008780
1983:2	0.348896	1990:1	0.000000
1983:3	0.350361		
1983:4	0.327067		
1984:1	0.325117		
1984:2	0.326090		
1984:3	0.285239		
1984:4	0.289384		
1985:1	0.265848		
1985:2	0.267190		
1985:3	0.270117		

Appendix Table Two (Continued)

Minimum Wage: 1990-1991		Minimum Wage: 1996-1997	
1990:1	0.000000	1996:3	0.000000
1990:2	0.040899	1996:4	0.090889
1990:3	0.019491	1997:1	0.063099
1990:4	0.006966	1997:2	0.038331
1991:1	0.012077	1997:3	0.054032
1991:2	0.253952	1997:4	0.201858
1991:3	0.214994	1998:1	0.130271
1991:4	0.209704	1998:2	0.108137
1992:1	0.186375	1998:3	0.048946
1992:2	0.156556	1998:4	0.034528
1992:3	0.156194	1999:1	0.026560
1992:4	0.142126	1999:2	0.007208
1993:1	0.160284	1999:3	0.005945
1993:2	0.110623	1999:4	0.000000
1993:3	0.107257		
1993:4	0.110502		
1994:1	0.142007		
1994:2	0.099080		
1994:3	0.070343		
1994:4	0.044850		
1995:1	0.063367		
1995:2	0.032706		
1995:3	0.033307		
1995:4	0.015271		
1996:1	0.025851		
1996:2	0.032706		
1996:3	0.000000		

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