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MINIMUM WAGE EFFECTS ON EMPLOYMENT AND SCHOOL ENROLLMENT

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ABSTRACT

We argue in this paper that the focus oo employment effects in recent studies of minimum wages ignores an important interactioo betweeo schooling, employment, and the minimum wage. To study these linkages, we estimate a conditional logit model of employment and enrollment outcomes for teeoagers using state-year observations for the period 1977 to 1989. The results show a negative influence of minimum wages oo school enrollment and a positive effect on the proportion of teens neither employed nor in school. We further suggest that our results are consistent with substitution by employers of higher- for lower-skilled teenagers, with the displaced teens ending up both out of work and out of school.

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Recent increases in minimum wages at the state and federal levels and the debates surrounding these adjustments have renewed interest in the labor market effects of minimum wage laws. Taking advantage of the additional variation provided by new legislation, a number of recent studies have reexamined the long-standing consensus that had emerged from the work of the Minimum Wage Study Commission and others (see Brown et al., 1982) that minimum wage laws reduce employment opportunities for youths (Card, 1992a, 1992b; Katz and Krueger, 1991, 1992; Spriggs, 1992; Card and Krueger, 1993; Neumark and Wascher, 1992; and Taylor and Kim, 1993). A striking feature of most of these studies (including ours) is that simple comparisons, or regressions controlling for exogenous shifts in labor demand, do not reveal disemployment effects of minimum wages for teenagers.

The recent set of studies of the minimum wage have focused on estimating employment effects and have not explored in any detail the potential implications for other aspects of the teenage labor market. In this paper, we argue that the standard employment equations are suggestive of an additional effect on youth labor market behavior. In particular, when the school enrollment rate is controlled for in the employment equation, the estimated disemployment effects of the minimum wage rise markedly. We interpret the sensitivity of the minimum wage effect on employment to the introduction of schooling controls as evidence that the recent literature has missed an important influence of minimum wages that operates through interactions between schooling, labor force participation, and minimum wages. Thus, in this paper, we move beyond the earlier research and examine in more detail the linkages between minimum wages, employment, and enrollment.

In our empirical work, we specify enrollment and employment as jointly determined, and therefore add exogenous determinants of enrollment as variables affecting employment. Our procedure does not significantly alter the reduced-form estimates of the disemployment effect of minimum wages. However, we find that minimum wages also lead to a decline in the school enrollment rate and an increase in the proportion of teenagers who are neither employed nor enrolled. One hypothesis consistent with these findings is that labor demand shifts toward higher-productivity teenagers after a minimum wage increase, and that these teenagers take jobs in lieu of enrolling in school, in the process displacing lower-productivity teens from employment. Alternatively, there may not be substitution away from the lower-productivity workers, but rather a shift from enrollment to queuing for minimum wage jobs. In either case, our results indicate effects of minimum wages that are not revealed in the standard reduced-form estimates of the effects of minimum wages on employment.

I. Reduced-Form Effects of Minimum Wages on Teenage Employment: A Reinterpretation

The basic dataset used in this paper is similar to that used in Neumark and Wascher (1992). In particular, the sample consists of panel data on the 50 states and Washington, D.C. for the period 1977-1989. Included in the dataset are variables measuring federal and state minimum wage levels, coverage by federal minimum wage statutes, and state averages (estimated from May CPS files) of wage rates, unemployment rates, employment rates, school enrollment rates, and the age composition of the state population. In addition, we have augmented the standard set of explanatory variables used in research on employment effects of minimum wages with information on statutory schooling requirements and educational quality: specifically, the upper end of the age range mandated by state compulsory schooling laws and average teacher salaries (deflated by the consumer price index). In both cases, the variables are intended to capture changes engineered by state and local education officials that influence enrollment rates. Summary statistics for each of the variables are shown in the appendix table.

To begin, we review evidence from the standard employment equation estimated in much of the minimum wage literature:

$$(1) \qquad E_{it} = \alpha MW_{it} + X_{it}\beta + \varepsilon_{it} \quad .$$

E_{it} is the employment-to-population ratio for teenagers aged 16-19 in state i at time t. MW_{it} is a measure of the coverage-adjusted relative minimum wage, constructed as the federal coverage rate for the state, multiplied by the higher of the federal or state minimum wage level, divided by the average wage in the state. Finally, X_{it} represents a set of control variables, including fixed state and year effects. We include in the regressions the previous year's minimum wage variable along with the contemporaneous variable in order to pick up the lagged effects of the minimum wage on employment that were documented in Neumark and Wascher (1992).

Some previous research (e.g. Mincer, 1976; Abowd and Killingsworth, 1981; and Tauchen, 1981) has distinguished between minimum wage effects on individuals covered and not covered by minimum wage laws, or has studied the effects on each sector separately, an approach that might be relevant here as well. However, lacking clear identifying information as to the classification of individuals into those not covered by minimum wage laws and those who are covered but whose equilibrium wage is above the minimum, we have chosen here to address aggregated estimates as in equation (1). Because this aggregate equation is then a combination of observations for which the minimum wage is binding and observations for which it is not (and for which wage rates are determined by the intersection of the demand and supply curves), equation (1) should be thought of as a reduced-form model that includes both exogenous demand and supply shifters. Thus, the estimated minimum wage coefficient should not be interpreted as a labor demand elasticity, but rather as the net effect averaged over workers for whom the minimum wage is and is not binding. (In Neumark and Wascher (1994), we provide an alternative approach to this issue.)

The first column of Table 1 shows the estimates of this standard specification using our dataset. The estimates are based on a GLS estimator that permits a block-diagonal covariance matrix, with heteroskedastic errors across states and first-order autocorrelation in the residuals within states, to obtain consistent estimates of the standard errors and efficient estimates of the coefficients. Consistent with other recent studies that fail to find disemployment effects for teenagers, the elasticity of employment with respect to a change in the minimum wage is small (-0.03) and is not statistically significant.

Because this equation is essentially a reduced-form model, it is appropriate to include exogenous supply variables in the specification. A variety of such variables have been used in past studies (for a survey see Brown et al., 1981), although in column (1) of the table, we have used only a cohort size variable. An alternative supply variable that has been used in some previous minimum wage studies (e.g. Al-Salam et al., 1981: Ragan, 1977, 1981: Mattila, 1978, 1981) is school enrollment. Schooling is an important alternative to work for many teenagers, and thus exogenous variation in enrollment can arguably influence employment rates.

The remainder of the table reports estimates of the basic employment equation that control for shifts in the school enrollment rate. Specifications using two alternative measures of enrollment are included to ensure that the results are not overly sensitive to measurement or definitional changes. The first (labeled S1) includes only teenagers who were enrolled in school and were not in the labor force. This variable is designed to capture enrollment decisions that preclude seeking employment. The second definition of enrollment (labeled S2) includes all individuals who reported that their major activity was school. This is a broader measure that includes most individuals who were both working and in school.

As shown in the second column, when the proportion enrolled in school and not in the labor force is included in the regression (and treated as exogenous), the elasticity jumps to -0.23 and is three times its standard error. When the broader measure of enrollment is included instead (column (3)), the minimum wage elasticity is the same (-0.23), but with a larger standard error of 0.11. In both equations, the coefficient on the enrollment variable is negative and significant, consistent with the view that employment and enrollment are alternative activities for many teenagers.

A potential problem with a causal interpretation of the estimates shown in columns (2) and (3) of Table 1 is that the enrollment rate may be endogenous. If exogenous factors that raise employment rates tend to lower enrollment rates by drawing individuals out of school, then this endogeneity transmits a negative bias to the coefficient on the school enrollment variable. (On the other hand, if exogenous factors that raise employment rates also lead individuals to remain in school or draw into school individuals who previously were neither working nor in school, then the endogeneity bias in the school enrollment coefficient is positive.) The coefficient on S1 may be especially prone to this type of bias because, as Card, et al. (1994) point out, employment and enrollment (as defined by S1) are mutually exclusive activities.

Nonetheless, the sensitivity of the estimates of minimum wage effects in standard specifications such as equation (1) to the inclusion of the school enrollment rate, even if overstated, suggests that exploring the linkages between employment, enrollment, and the minimum wage may yield additional insights into the workings of the youth labor market. Indeed, one explanation of the sensitivity of the minimum wage coefficient to the addition of controls for enrollment shifts might be that there is another effect of the minimum wage through its influence on the teenage school enrollment rate. In particular, if a higher minimum wage encourages individuals to leave (or not enroll in) school in order to take a job, then we would expect to see a larger negative effect of minimum wages on employment after taking account of the negative partial correlation between employment and school enrollment. In this case, the negative impact of higher minimum wages on enrollment rates can be interpreted as a labor supply response stemming from a demand shift towards the teenagers who would have otherwise chosen to be enrolled in school.

To see this more clearly, suppose that, at the state level, employment is determined by

$$(2) \qquad E = \alpha MW + \gamma S + \varepsilon \ ,$$

where S is the enrollment rate, $\alpha < 0$ and $\gamma < 0$ (as in the Table 1 estimates), and X has been dropped. γ is assumed to reflect the effect of exogenous variation in S on E. If enrollment is negatively related to the minimum wage through the reduced-form equation

$$(3) \qquad S = \alpha' MW + \varepsilon' \ .$$

where $\alpha' < 0$, then the reduced form for the employment equation is

$$(4) \qquad E = [\alpha + \alpha'\gamma]MW + \varepsilon'' .$$

In this reduced form, the expected sign of the minimum wage coefficient may be close to zero (or even positive) even though the individual minimum wage coefficients α and α' are negative. Note that if the minimum wage had no effect on the school enrollment rate (so that $\alpha'=0$) or if employment and enrollment decisions were unrelated (so that $\gamma=0$), the coefficient on the minimum wage in the employment equation would not be sensitive to the exclusion or inclusion of the enrollment rate.

Previous research has identified correlations between minimum wage changes and the schooling decisions of youths, although the evidence is relatively scant and there is some disagreement about the direction of the correlation. Using time-series data, Mattila (1978) finds a positive effect of the minimum wage on school enrollment, which he argues suggests that raising the minimum wage prompts teenagers to remain in school to increase their likelihood of gaining employment in the covered sector in the future. Ehrenberg and Marcus (1980, 1982) examine the effects of the minimum wage on school enrollment with cross-section data. Using grouped data by state for 1970, they find very little evidence of an effect of the minimum wage on enrollment. In contrast, their estimates derived from the 1966 National Longitudinal Survey suggest that minimum wages reduce enrollment rates for low-income teenagers and increase enrollment rates for teenagers from high-income families. Cunningham (1979) also finds that a higher minimum wage reduces enrollment; we will return to his results in more detail later in the paper. Finally, Card (1992b) finds a negative partial correlation between school enrollment and the minimum wage in his study of the 1988 increase in California's minimum wage. In the next section, we turn to a fuller characterization of minimum wage effects on employment and enrollment of teenagers using our dataset.

II. Minimum Wage Effects on Enrollment and Employment

The Empirical Approach

Our modeling strategy is based on the assumption that youths choose among a set of J alternative activities (work, school, etc.), with the choice influenced by a set of determinants, X. In particular, let

$$(5) \qquad p_{kj} = f(X) + \varepsilon; \quad k=1,\ldots,n; \; j=1,\ldots,J$$

be the probability that individual k chooses activity j. In this framework, X includes variables relating to the costs of participating in each activity (e.g. schooling costs and foregone leisure) and the benefits from participation (e.g. current and future wage rates), while unmeasured individual specific tastes and abilities as well as unmeasured costs and benefits are captured in ε.

Averaging over individuals (for example, for each state and year), the proportion of teenagers choosing a particular activity becomes

$$(6) \qquad P_j = g(X') + \varepsilon'; \; j=1,\ldots,J.$$

Average differences in the unobservables are captured in the error term, ε'. In addition, some elements of X that were exogenous to individuals are endogenous with respect to the aggregate proportions (e.g. the market wage for teenagers), and so X' has been redefined as the exogenous determinants of the costs and benefits of each activity. In this sense, equation (6) can be viewed as a set of reduced-form equations, subject to the constraint that the sum of the proportions is equal to one.

This framework gives rise to the conditional logit model suggested by McFadden (1973). In particular, we specify employment and enrollment as jointly determined by a set of exogenous variables, including the minimum wage. In the results shown below, we employ grouped data to estimate the parameters of the multinomial logit specification. This approach has been used in previous analyses of the youth labor market by Ehrenberg and Marcus (1980, 1982) and Wachter and Kim (1982).

Following Ehrenberg and Marcus, we divide the youth population into four mutually exclusive categories of youth activity, distinguished by employment status and enrollment status. Specifically, SNE is the proportion of individuals in school but not employed: SE is the proportion of individuals in school and employed: ENS is the proportion of individuals employed but not in school: and NSNE is the proportion of individuals not in school and not employed. We consider several groupings of these activities in the results that follow, but in the terminology of the GLS estimates presented in Table 1, E/P = SE + ENS, S1 = SNE, and S2 = SNE + SE. (The inexact correspondence between S1 and SNE reflects the fact that the use of the employment status recode of the CPS for the construction of S1 causes individuals both unemployed and in school to be excluded from this measure, while the use of the major activity variable to distinguish the alternative activities in this section leads to the inclusion of such individuals in SNE.)

Given the categorization just described, the conditional logit model is comprised of a set of equations specifying the logarithms of the odds ratios as functions of a set of independent variables:

$$\text{(7a)}\quad \log(SNE_{it}/NSNE_{it}) = \alpha_{11}MW_{it} + \alpha_{12}MW_{it-1} + X_{it}\beta_1 + \varepsilon_{1it}$$

$$\text{(7b)}\quad \log(SE_{it}/NSNE_{it}) = \alpha_{21}MW_{it} + \alpha_{22}MW_{it-1} + X_{it}\beta_2 + \varepsilon_{2it}$$

$$\text{(7c)}\quad \log(ENS_{it}/NSNE_{it}) = \alpha_{31}MW_{it} + \alpha_{32}MW_{it-1} + X_{it}\beta_3 + \varepsilon_{3it}$$

where i indexes states and t indexes time. This set of equations is estimated with grouped state-year observations as in Table 1. The exogenous variables used in the estimation include the current and lagged relative coverage-adjusted minimum wage (MW_{it} and MW_{it-1}) as defined in the previous section, and other exogenous determinants of employment and enrollment (X_{it}). In addition to the relative minimum wage, we use the prime-age male unemployment rate as an exogenous labor demand indicator, the relative size of the teenage cohort as an exogenous supply indicator, and average teacher salaries and the upper end of the age range of compulsory schooling laws as exogenous indicators of the demand for education. Compulsory schooling laws are split into controls for four distinct categories: less than age 16, age 16, age 17, and age 18. We use compulsory school age of 16 as the omitted reference category in the results described below. Fixed state and year effects are also included to capture state- or year-specific variation in tastes or abilities or in unmeasured determinants of the costs and benefits of these alternative activities. Finally, as before, the model is estimated with a block-diagonal residual covariance matrix allowing for heteroskedasticity and AR(1) errors.

Results

Parameter estimates from the conditional logit model of youth activity are presented in Table 2. In the left two columns, we divide potential youth activities into three distinct groups: (1) employment (E=SE+ENS), (2) in school and not employed (SNE), and (3) neither in school nor employed (NSNE). This grouping corresponds closely to column (2) of Table 2, where we define school enrollment (S1) using the employment status recode.

The results in Panel A indicate that minimum wages reduce both the proportion in school and the proportion employed relative to the proportion neither in school nor employed. The sum of the minimum wage coefficients is statistically significant and negative in both equations. With regard to the remaining variables, the adult male unemployment rate is estimated to have a negative effect on employment, but no effect on enrollment. Higher teacher salaries are estimated to have a positive effect on enrollment and employment, while compulsory schooling laws have essentially no determinable effect.

To get a better sense of how to interpret these results, Panel B of the table shows the estimated implied elasticities of the proportion in each employment-enrollment category with respect to the minimum wage. Consistent with the evidence from the results in Table 1, the elasticity of the proportion employed with respect to the minimum wage is essentially zero

(0.05 and insignificant). However, the results also indicate a significant negative elasticity (-0.34) of enrollments with respect to the minimum wage and a positive and significant elasticity (0.67) of the proportion neither in school nor employed. Thus, the primary net impact of a higher minimum wage is to increase the proportion of teens neither in school nor employed (NSNE), and to reduce the proportion in school (SNE+SE). These findings clearly indicate that the reduced-form employment equation masks shifts in enrollment rates of teenagers, and perhaps also minimum wage effects on employment rates of subgroups of teenagers.

To explore these relationships further, columns (3) to (5) of Panel A present results that additionally split the employed category into two distinct groups: in school and employed (SE), and employed but not in school (ENS). This results in a model with the four mutually exclusive categories as originally defined. Consistent with the earlier results, the estimates show generally statistically significant minimum wage effects that induce individuals away from each of the employed and/or enrolled groups to the not enrolled and not employed category. Panel B again shows the implied elasticities of the proportions of teenagers in each employment-enrollment category with respect to the minimum wage. As in the three-category estimates, the largest increase occurs in the proportion of teenagers neither in school nor working. In this specification, however, the estimates also suggest a sizable decline in the proportion in school and employed. Finally, on net, the employment effect (SE+ENS) is negative, although small.

Robustness Checks

Table 3 reports a variety of estimates of the minimum wage elasticities implied by the conditional logit model, exploring the robustness of the results to variations in model specification and the structure of the covariance matrix. For comparison purposes, Panel A repeats the baseline elasticities for the four-category model from Table 2. In Panel B, we report estimates that exclude the exogenous determinants of schooling (teacher salaries and compulsory schooling laws). The estimates are little affected by this omission, suggesting either that there is room to improve the specification of the determinants of enrollment decisions or that enrollment is little affected by schooling measures. In Panel C. we omit the lagged minimum wage variable to focus on the contemporaneous effects of an increase in the minimum. As expected given the results in our previous paper (Neumark and Wascher. 1992). some of the effects on schooling and employment are smaller when the lagged variable is omitted. but not strikingly so. The estimate of the elasticity of the proportion neither in school nor employed with respect to the minimum wage is 0.44. about a third less than the baseline. and still statistically significant.

In Panels D and E. we examine the effects of changing the assumed structure of the covariance matrix. When a single autocorrelation parameter across all states is estimated (Panel D). the results are nearly identical to the baseline shown in Panel A. although with larger standard errors. Using a scalar residual covariance matrix. as in Panel E. leads to some (offsetting) decline in the minimum wage effects on the two employment categories (SE and ENS) as well as increased standard errors. but overall the results again are qualitatively similar.

Finally. in Panel F. we report the elasticities when the model is estimated excluding the fixed state effects. This should produce results more analogous to the cross-section results of Ehrenberg and Marcus (1980). although that study focused on differences in minimum wage effects across income classes and did not introduce any time-series variation. Nonetheless. comparisons of the estimates in Panel F with those in Panel A are consistent with the criticisms of cross-section studies of minimum wage effects put forth by Freeman (1982). In particular. the estimated disemployment effect (SE+ENS) rises considerably when the fixed state effects are excluded. consistent with Freeman's argument that unmeasured economic conditions across states could give rise to a positive correlation between wages and employment and hence a spurious negative correlation between the relative minimum wage variable and employment. If high wage rates also induce a positive labor supply response. there could be a spurious positive correlation between the relative minimum wage variable and the enrollment rate as well: this is consistent with the finding that the elasticity of the proportion in school and not employed becomes positive when the state effects are excluded. The spurious correlations between the relative minimum wage and employment and enrollment induced by ignoring state effects are roughly offsetting, and the elasticity of the proportion neither in school nor employed is slightly smaller (0.58) in this specification.

Interpreting the Results

The structure of the model is not sufficiently complex nor are the data sufficiently detailed to enable us to infer precisely why the proportion of teenagers neither in school nor employed rises in response to a higher minimum wage while the employment rate is little affected. However, the estimates in Table 2 are consistent with two possibilities.

First, the estimates may be picking up a shift in labor demand toward higher-skilled workers. One effect of a higher minimum wage is to raise the cost of lower-skilled labor relative to the cost of higher-quality substitutes (such as teenagers with more schooling). If high- and low-skilled teenagers enter separately in the production function, this change in the relative wage reduces the demand for lower-productivity workers by moving employers up the demand curve for low-skilled labor. However, the increase in the minimum wage also shifts out the demand curve for higher-productivity labor, inducing higher labor supply and employment for these workers. Thus, if higher-skilled teenagers are more likely to be enrolled in school, and if school is properly viewed as an alternative to working, this increase in labor demand will result in a decline in enrollment rates, as the higher-productivity teenagers choose to enter the labor market rather than enroll (or remain enrolled) in school. Moreover, these teenagers will displace some relatively lesser-skilled teenagers in the work force, mitigating the overall employment loss associated with the minimum wage and, of course, masking the employment declines among the less-skilled teens. The displaced lower-skilled workers would tend to end up neither enrolled nor employed given their presumably lower propensity toward additional schooling. Thus, in the substitution model, increases in the minimum wage do reduce the employment of some workers, presumably those who are the least productive. In this sense, the competitive view of minimum wages applies to this subset of workers. even though reduced-form employment equations reveal little or no disemployment effect of the minimum wage for the teenage group in the aggregate.

A second and related possibility is that teenage workers are homogeneous. but that the higher minimum wage leads some teenagers to leave school and queue for jobs in the covered sector. For this model to explain the results. the higher minimum wage must induce an increase in desired labor supply but have little or no effect on labor demand. Thus. some individuals leave school and end up neither enrolled nor employed. but there is no disemployment effect even for a subset of workers.

Note that in both models. the main result is that a higher minimum wage increases the proportion of teenagers neither enrolled nor employed and decreases the proportion enrolled. The primary difference is that the substitution hypothesis suggests that a subset of workers may be "priced out of the market" relatively permanently by a minimum wage increase. whereas in the queuing model. the time spent neither enrolled nor employed is more equally distributed across the teenage population. especially if some teenagers who queue for minimum wage jobs eventually obtain them.

In our view. there is some evidence suggesting that the substitution model better characterizes the data. The substitution model predicts an exogenous effect of enrollments on employment. because as teenagers leave school to take jobs. the price of their labor should fall. and employment rise. In contrast. the queuing model predicts essentially no exogenous effect of enrollments on employment. because the teenagers who leave school simply queue for minimum wage jobs. The estimates of the coefficient on the enrollment variable in the employment equations of Table 1 are decidedly negative. which seems to lend support to the substitution argument. However. as we indicated. these estimates are potentially contaminated by endogeneity bias. and if this bias is the main source of the negative partial correlation between enrollment and employment. these results cannot be used as evidence in favor of the substitution model. In fact. a greater propensity for teenagers to join the queue when employment conditions are good would generate this negative endogeneity bias.

To assess the importance of endogeneity bias, we reestimated the specifications in columns (2) and (3) of Table 1, instrumenting for the enrollment rate with the school quality and compulsory schooling variables. When the proportion of teenagers in school and not in the labor force (S1) is used as the enrollment variable, the IV procedure produces a slightly larger negative coefficient on the enrollment rate (-0.93 with a standard error of 0.28). When the broader enrollment measure (S2) is used, the coefficient drops slightly to -0.31 with a standard error of 0.16. Thus, the IV results do not suggest that the negative coefficients on the enrollment rate in the employment equation are due to endogeneity bias, and in that sense lend support to the substitution hypothesis.

In addition, the substitution model is, in our view, somewhat more consistent with the findings of other research on the effects of the minimum wage, although this research has not addressed the substitution versus queuing models explicitly. One relevant study is a recent paper by Currie and Fallick (1993), who study the effects of the minimum wage on the employment of low-wage workers using panel data on individuals. As long as there is excess supply at the original minimum wage, an increase in the number of identical workers queuing for minimum wage jobs should not affect the probability of job loss for an existing worker. In contrast, this study finds that the probability of a job loss for a low-wage worker rises 3 to 4 percent with an increase in the minimum wage, a result consistent with the prediction of the substitution hypothesis that minimum wages lead to job losses for some workers.

Another pertinent study is that by Cunningham (1979), who employs a methodology similar to ours, using state data from the 1960 and 1970 decennial Censuses of Population. He finds that a higher minimum wage leads to an increase in full-time employment in the sector covered by minimum wage laws and a decline in part-time employment in that sector; in addition, school enrollment among workers in the covered sector falls. This is consistent with a substitution of higher- for lower-quality workers, as part-time workers in the covered sector who were also in school increase their hours (and leave school) and displace the lower-quality workers (who were not in school). In contrast to our findings. Cunningham's results do not show a sharp increase in the proportion of teenagers neither in school nor employed, but instead show that most of the displaced workers enter the uncovered sector. This difference may simply reflect increases in coverage since 1970 that have reduced the size of the uncovered sector. Finally, using cross-state data from the 1970 Census, Ehrenberg and Marcus (1980) find that for male teenagers, minimum wage increases redistribute jobs from the children of the poor to the children of the nonpoor, although their results using 1966 NLS data suggest the opposite.

On the other hand, Card (1992b) reports enrollment declines associated with the 1988 minimum wage increase in California and argues that these declines "were not directly associated with the relative growth in California employment" (p. 48). He bases his claim on the fact that the employment rate for enrolled teenagers increased by about as much as the employment rate for all teenagers following the minimum wage increase. However, this evidence does not rule out the substitution hypothesis since there may have been increases in employment among those who remained enrolled as well as transitions from employment to enrollment. Thus, on balance, we view the results reported in these previous studies as more supportive of the substitution model than of the queuing model, although, as with our results, the evidence is indirect.

III. Conclusions

The evidence in this paper indicates that an exclusive focus on the effects of minimum wages on employment is inadequate for understanding how minimum wages influence the low-wage labor market. In particular, the sensitivity of the estimated disemployment effect of minimum wages for teenagers to the inclusion or exclusion of the school enrollment rate suggests that there is an important interaction between schooling, employment, and the minimum wage that has been ignored in much of the recent literature.

To study these linkages more carefully, we broaden the analysis of the effects of the minimum wage using a conditional logit model of alternative employment and enrollment outcomes. Estimates from this model consistently show a negative influence of a higher minimum wage on school enrollment. Perhaps most importantly, we find that there is a significant increase in the proportion of teenagers neither in school nor employed in response to an increase in the minimum wage.

The results in this paper are consistent with the hypothesis that a higher minimum wage leads employers to substitute away from low-productivity teenagers toward higher-productivity teenagers. This increase in the demand for higher-quality workers induces a labor supply response among teenagers who were formerly in school. The upshot is that lower-productivity workers are displaced by higher-productivity teens (who also leave school), and become neither employed nor enrolled. Under this interpretation, reduced-form estimates of minimum wage effects on employment mask disemployment effects among the least productive workers. The results are also consistent, although in our view less so, with the hypothesis that, as a result of a minimum wage increase, teenagers leave school to queue for minimum wage jobs, without any disemployment effect among those already employed. Of course, some individuals who queue for minimum wage jobs will eventually find employment so that in either case, some teenagers leave school for work.

Whether a minimum wage-induced transition from school to work is desirable from a policy perspective is ambiguous, depending on, among other factors, the substitutability of experience for education in a human capital production function. However, individuals displaced from both school and work seem unambiguously worse off. Whatever the welfare evaluation of these changes, the results clearly indicate that focusing solely on the reduced-form estimates of minimum wage effects on employment masks changes in school enrollments of teenagers and perhaps also changes in employment that should be of interest to policymakers.

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Table 1: Within-Group Estimates of Minimum Wage Effects on Employment, Teenagers (16-19)[1]

	Teenagers (16-19)		
	(1)	(2)	(3)
Minimum wage	.15	-.13	.01
	(.11)	(.07)	(.10)
Minimum wage, lagged one year	-.18	-.11	-.23
	(.12)	(.08)	(.11)
Proportion of age group in school and not employed (S1)	---	-.78	---
		(.03)	
Proportion of age group with major activity school (S2)	---	---	-.39
			(.03)
Minimum wage elasticity[2]	-.03	-.23	-.23
	(.13)	(.08)	(.11)

1. The sample covers the 50 states and Washington, D.C., for the years 1978-1989. Standard errors are reported in parentheses. The minimum wage variable is the minimum wage level, multiplied by coverage, divided by the average wage in the state. All specifications include fixed state and year effects, the unemployment rate for prime-age males, and the proportion of the population in the age group. The estimates allow for different residual error variances across states, and AR1 errors, with different autocorrelation parameters across states. The estimates are computed by estimating the autocorrelation parameter from the OLS residuals for each state, quasi-differencing (dropping the first observation), and applying GLS.

2. Long-run elasticity, evaluated at sample means. Standard errors treat coefficient estimates, but not means, as random.

Table 2: Conditional Logit Estimates of Minimum Wage Effects on Employment and School Enrollment, Teenagers (16-19)[1]

A. Estimates[2]

	Using Empl. Status Recode (S1)		Using Empl. Status Recode and Major Activity (S2)		
	S/NSNE (1)	E/NSNE (2)	SNE/NSNE (3)	SE/NSNE (4)	ENS/NSNE (5)
Coverage-adjusted relative minimum wage	-2.26 (.60)	-.47 (.65)	-1.40 (.70)	-1.53 (.79)	-.15 (.75)
Coverage-adjusted relative minimum wage, lagged one year	-.65 (.67)	-1.31 (.71)	-.86 (.75)	-1.56 (.88)	-.95 (.83)
Proportion of population aged 16-19	.06 (1.14)	-1.43 (1.28)	-1.14 (1.25)	-2.54 (1.60)	-.61 (1.49)
Prime-age male unemployment rate	.34 (.56)	-2.65 (.62)	1.76 (.61)	-1.69 (.79)	-2.17 (.64)
Compulsory school age<16	-.04 (.21)	-.04 (.21)	.15 (.23)	.13 (.29)	-.11 (.22)
Compulsory school age=17	-.12 (.09)	.01 (.09)	-.10 (.09)	-.08 (.12)	-.07 (.09)
Compulsory school age=18	-.21 (.10)	-.06 (.11)	-.05 (.10)	.07 (.14)	-.05 (.11)
Average teacher salaries/100	3.42 (1.03)	2.41 (1.22)	.57 (1.28)	-1.28 (1.51)	1.88 (1.45)
Sum of minimum wage effects	-2.90 (.65)	-1.78 (.69)	-2.26 (.76)	-3.08 (.89)	-1.11 (.77)

Table 2 (continued)

B. Minimum Wage Elasticities[3]

	Mean proportion (1)	Predicted change (2)		Mean proportion (3)	Predicted change (4)
Proportion in school (S)	.40	-.34 (.17)	Proportion in school, not employed (SNE)	.45	-.13 (.17)
Proportion employed (E)	.43	.05 (.16)	Proportion in school, employed (SE)	.21	-.40 (.27)
Proportion not in school, not employed (NSNE)	.17	.67 (.14)	Proportion employed, not in school (ENS)	.23	.28 (.25)
			Proportion not in school, not employed (NSNE)	.12	.64 (.14)

1. The sample covers the 50 states and Washington, D.C., for the years 1978-1989.

2. Standard errors are reported in parentheses. The minimum wage variable is the minimum wage level, multiplied by coverage, divided by the average wage in the state. All specifications include fixed state and year effects. Compulsory schooling age of 16 is the omitted reference category. The estimates allow for different residual error variances across states, and AR1 errors, with different autocorrelation parameters across states. The estimates are computed by estimating the autocorrelation parameter from the OLS residuals for each state, quasi-differencing (dropping the first observation), and applying GLS equation-by-equation to the quasi-differenced data. (This latter step is equivalent to applying GLS to the system of equations for the quasi-differenced data.)

3. Mean proportions are sample means. Predicted changes are evaluated at sample means. Standard errors of predicted changes are calculated treating sample means as fixed, and coefficient estimates as random, using first-order linear approximations to the nonlinear functions of the parameter estimates. While the estimated minimum-wage elasticities are independent of which category is the reference category in the conditional logit estimation, the standard errors vary slightly based on this choice, since the first-order approximation used in the standard error calculation varies.

Table 3: Sensitivity Analysis for Conditional Logit Estimates, Minimum Wage Elasticities[1]

A. Baseline (Table 3)

SNE	SE	ENS	NSNE
-.13	-.40	.28	.64
(.17)	(.27)	(.25)	(.14)

B. Excluding Exogenous Schooling Determinants

SNE	SE	ENS	NSNE
-.14	-.40	.29	.67
(.17)	(.27)	(.25)	(.14)

C. Omitting Lagged Minimum Wage Effects

SNE	SE	ENS	NSNE
-.14	-.27	.28	.44
(.14)	(.22)	(.20)	(.12)

D. Single Autocorrelation Parameter for Each State

SNE	SE	ENS	NSNE
-.09	-.50	.29	.65
(.19)	(.31)	(.27)	(.16)

E. Homoskedastic, Non-Autocorrelated Errors

SNE	SE	ENS	NSNE
-.15	-.17	.12	.60
(.24)	(.38)	(.36)	(.20)

F. No Fixed State Effects

SNE	SE	ENS	NSNE
.04	-.30	-.11	.58
(.10)	(.20)	(.14)	(.08)

1. Except where otherwise specified, footnotes from Table 2 apply. All specifications use major activity in the survey week to define enrollment (S2).

Appendix Table: Descriptive Statistics

	Characteristics of Teenagers (16-19)			
	Mean	Standard deviation	Minimum	Maximum
Employment rate	.43	.09	.09	.67
Proportion of population aged 16-19	.09	.01	.06	.14
Proportion of population in school and not employed (S1)	.40	.08	.17	.73
Proportion of age group with major activity in school (S2)	.65	.07	.39	.87
	State Characteristics			
	Mean	Standard deviation	Minimum	Maximum
Coverage-adjusted relative minimum wage	.35	.04	.25	.49
Prime-age male (25-64) unemployment rate	.05	.03	.004	.20
Average teacher salaries (1,000 1982 dollars)	20.1	3.7	13.8	35.7
Compulsory school age	Proportion	# changes to	# changes from	
<16	.04	0	2	
16	.67	3	7	
17	.14	5	2	
18	.14	4	1	

1. The sample covers the 50 states and Washington, D.C. for the years 1978-1989. The minimum wage level is the greater of the states or federal minimum wage level. The coverage rate is coverage by federal minimum wage laws for all workers in the state, divided by the average wage in the state. Information on teacher salaries and compulsory schooling ages is taken from various years of Digest of Education Statistics, U.S. Department of Education, National Center for Education Statistics, Estimates of School Statistics, National Education Association, and Statistics of Public Elementary and Secondary Day Schools, U.S. Department of Education, National Center for Education Statistics.