NBER WORKING PAPER SERIES

MINIMUM WAGE EFFECTS AND LOW-WAGE LABOR MARKETS: A DISEQUILIBRIUM APPROACH

David Neumark
William Wascher

Working Paper No. 4617

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
January, 1994

Neumark is Assistant Professor in the Economics Department at the University of Pennsylvania, and a Faculty Research Fellow of the NBER. Wascher is Senior Economist at the Board of Governors of the Federal Reserve System. The views in this paper do not necessarily represent those of the Federal Reserve Board or its staff, or the National Bureau of Economic Research. We thank David Card, Frank Diebold, Chris Hanes, and seminar participants at the NBER Summer Institute, Penn State, Michigan State, Johns Hopkins, Maryland, CUNY and the BLS for helpful discussions. This paper is part of NBER's research program in Labor Studies.

NBER Working Paper #4617
January 1994

MINIMUM WAGE EFFECTS AND LOW-WAGE LABOR MARKETS: A DISEQUILIBRIUM APPROACH

ABSTRACT

We present a new approach to estimating minimum wage effects on employment. In contrast to most previous research, we account for the possibility that the relationship between minimum wages and employment depends on the magnitude of the minimum wage relative to the equilibrium wage in the absence of the legislated minimum. In particular, estimating the employment effects of binding minimum wages requires separation of sample observations into those that are on the labor demand curve but off the labor supply curve, and those that are at labor market equilibria. The paper implements an endogenous switching regression model with unknown sample separation that yields these estimates. The approach also yields estimates of the impact of labor market characteristics on the probability that minimum wages are binding.

We also extend the disequilibrium approach to monopsony, which introduces a third regime, between the equilibrium monopsony wage and the equilibrium competitive wage, in which observations are on the labor supply curve but off the labor demand curve and minimum wages are therefore positively related to employment. Minimum wage effects under monopsony are estimated in a three-regime endogenous switching regression model with unknown regimes, and the monopsony characterization of low-wage labor markets is tested against the competitive characterization.

David Neumark
Department of Economics
University of Pennsylvania
Philadelphia, PA 19104
and NBER

William Wascher
Board of Governors of the
Federal Reserve System
20th and Constitution, NW
Washington, DC 20551

I. Introduction

The competitive view of minimum wages, enshrined in numerous textbooks, is that they reduce employment. This paper presents a new approach to testing this prediction. In contrast to many previous studies, which estimate a single regression of the employment rate on a minimum wage variable (as well as other controls), we use the predictions of standard theoretical models to take explicit account of the possibility that the relationship between minimum wages and employment depends in a discrete fashion on the magnitude of the minimum wage relative to the equilibrium wage in the absence of the legislated minimum. Specifically, we attempt to distinguish observations for which the legislated minimum is below the equilibrium competitive wage--and thus unrelated to employment--from observations for which the minimum wage is above the equilibrium competitive wage and negatively related to employment. To achieve this separation, we implement an endogenous switching regression model with unknown sample separation, with the switch point defined as the intersection of the labor demand and supply curves.

In addition, we extend this disequilibrium approach to account for the possibility of monopsony power in low-wage labor markets.[1] In the monopsony model, as in the competitive model, the minimum wage should be unrelated to employment when the legislated minimum is below the equilibrium monopsony wage. Similarly, in either model, the minimum wage is predicted to be negatively related to employment when the minimum wage is above the equilibrium competitive wage. However, the competitive and monopsony

[1]The plausibility of monopsony power in low-wage labor markets has generally been dismissed in the past (see, e.g., Brown, 1988). But recent evidence of positive effects of minimum wages on employment has renewed interest in the monopsony model (Card, 1992b; Rebitzer and Taylor, 1991).

models differ in their implications for minimum wage effects when the minimum wage is between the equilibrium monopsony wage and the equilibrium competitive wage (Stigler, 1946). In this region, a minimum wage can increase employment by moving the labor market away from the monopsony outcome--where the labor demand curve intersects the marginal cost of labor curve--along the supply curve. In contrast, the competitive model predicts that minimum wages should be unrelated to employment over this range.

To assess the relevance of monopsony power in low-wage labor markets, we set up a model of minimum wage effects on employment with three regimes, with each corresponding to the alternative possibilities implied by the monopsony model. Given specifications of the labor demand and labor supply curves, and given an assumption regarding the relationship between the marginal cost of labor curve (under monopsony) and the labor supply curve, the switch points again arise naturally as functions of the parameters and variables in the labor demand and labor supply curves. We estimate the resulting three-regime endogenous switching regression model with unknown regimes, and test the monopsony characterization of low-wage labor markets against the two-regime competitive characterization.

One rationale for our approach stems from the recent round of studies on the employment effects of minimum wages, some of which challenge the prediction of the standard textbook model. This recent research has greatly expanded the types of data used to estimate minimum wage effects, including: extending time-series data into the 1980s (Wellington, 1991); focusing on minimum wage increases in particular labor markets (Card, 1992b; Spriggs, 1992; Katz and Krueger, 1992; Taylor and Kim, 1993; Card and Krueger, 1993); conducting panel data experiments on state-level observations (Card, 1992a;

Neumark and Wascher, 1992; Williams, 1993), and using individual-level data (Currie and Fallick, 1993). Compared with the earlier time-series studies which, by and large, found similar effects, the findings from this recent research have been quite varied, with some studies finding the conventional negative employment effects of minimum wages, some finding no employment effects, and some even finding positive effects.

A potential explanation for the more disparate findings from recent research may be that the samples used in this research vary with respect to whether minimum wages are binding. For example, even if low-wage labor markets are competitive, it is possible that some studies will reveal no effects of minimum wages on employment simply because they are based on samples in which most of the variation in minimum wages is in the range for which minimum wages are non-binding, while the reverse may be true of studies revealing negative effects. Similarly, if low-wage labor markets are characterized by monopsony, we might nonetheless generally expect to find negative effects of minimum wages on employment, in samples that do not capture minimum wage movements over the perhaps narrow range for which a higher minimum wage raises employment, but instead capture variation along the binding demand regime. On the other hand, we might sometimes expect to find a positive (or non-negative) association between minimum wages and employment.[2]

Thus, the approach developed in this paper offers at least three potential advantages relative to existing single-equation estimates: it leads to estimates of the employment effects of minimum wages when they are binding; it yields a (probabilistic) classification of

[2]This also implies that we cannot test the monopsony model by asking whether minimum wage increases are associated with increases in employment (as claimed, for example, by West and McKee, 1980).

observations based on whether or not minimum wages are binding; and it provides a test of the monopsony characterization of low-wage labor markets.

II. The Data and Single-Equation Results

The data used in this paper are the same as those used in Neumark and Wascher (1992), a panel data set on the 50 states and Washington, DC, covering the period 1977-1989, as well as the subperiod 1973-1976 for those states separately identified in the CPS before 1977. The data set includes information on federal and state minimum wage levels, estimates of coverage by federal minimum wage statutes,[3] and state averages of wages, unemployment rates, employment rates, and age composition variables. In that paper, we followed the literature and estimated regressions of the form

$$(1) \qquad E_{it} = \alpha MW_{it} + R_{it}\beta + \epsilon_{it} \ .$$

E_{it} is an employment-to-population ratio either for teens aged 16-19 or young adults aged 16-24. MW_{it} is a coverage-adjusted relative minimum wage, constructed as federal coverage for the state, multiplied by the higher of the federal or state minimum wage level, divided by the average wage in the state. R_{it} is a vector of control variables, possibly including fixed state and year effects.

Numerous specification issues arise with respect to equation (1), which we explored in our earlier paper, and which have been further discussed by other researchers and ourselves (Card, et al., 1993; Neumark and Wascher, 1993). In this paper we are interested in highlighting an econometric approach that is an alternative to that commonly used in

[3]Published coverage estimates refer to all workers, not the young age groups usually considered in minimum wage studies.

research on minimum wages, not in revisiting these specification issues. Consequently, we focus on a set of specifications and an age group that are, in our view, least controversial.

One issue that has been raised in this context is the appropriate use of supply variables in single-equation estimates. In our case, we explicitly specify the labor demand and labor supply curves in a disequilibrium model. We therefore have a way to introduce supply variables, even when workers are homogeneous and there is no uncovered sector, because supply variables will affect employment when the minimum wage is non-binding.[4] However, we include only the proportion of the population in the age group being studied, and omit the more problematic school enrollment rate, since it may be partly endogenous with respect to labor market conditions. We also work with contemporaneous minimum wage variables, although our earlier paper indicated that specifications using lagged values of the minimum wage variable provided stronger evidence of disemployment effects. In addition, in our earlier paper we allowed for fixed state and year effects. In this paper, though, the econometric approach that we use precludes the inclusion of these effects, although we give some attention to the heterogeneity problem that may arise from their omission. Finally, we study young adults (aged 16-24), for whom the results in our previous paper were much less sensitive to these specification issues than were those for teenagers (aged 16-19).[5]

To see the consequences of these specification choices, Table 1 reports estimates of

---

[4]Brown, et al. (1982) argue that supply variables affect employment in the presence of minimum wages when workers are heterogeneous and there is an uncovered sector.

[5]This may be in part because for the broader young adult age group we have larger cell sizes with which to estimate employment rates and other control variables.

equation (1) for young adults aged 16-24. The top panel reports the basic specifications with which the estimates from the disequilibrium models will be compared. The estimated coefficient of the minimum wage variable is significant and negative. The demand variable --the prime-age male unemployment rate--also has an estimated coefficient that is negative and significant. The supply variable measuring the proportion of the age group in the population has the expected negative coefficient, although the estimate is not significantly different from zero.[6] Note that fixed state and year effects are excluded from these specifications.

The lower panel presents the estimated minimum wage coefficients for specifications that alternatively include fixed state and year effects, and substitute the lagged minimum wage variable for its contemporaneous counterpart. The first row shows, as did our 1992 paper, that including fixed state and year effects leads to some diminution of the minimum wage effects. The second row shows that the estimated disemployment effects of minimum wages are stronger when the lagged minimum wage variable is used. Broadly speaking, however, the estimated minimum wage effects for young adults are not especially sensitive to these specification issues. All of the specifications indicate a significant (or nearly so) negative effect of minimum wages on the employment rate.[7]

---

[6]The expected coefficient is negative because this variable measures relative supply. Higher relative supply reduces the equilibrium wage, resulting in a lower employment rate (but not level).

[7]The estimates for young adults are also insensitive to the inclusion of measures of enrollment rates. In contrast, for teenagers the estimated minimum wage effect is very sensitive to these specification issues (Neumark and Wascher, 1992 and 1993).

III. Disequilibrium Models of Employment Effects

III.a. The Competitive Model

We consider first the two-regime model corresponding to the competitive model of minimum wage effects. Let the labor demand equation be given by

$$(2) \qquad E^D = \alpha W + X\beta + \epsilon^D \ ,$$

and let the (inverse) labor supply equation be given by

$$(3) \qquad W = \gamma E^S - Y\delta - \gamma\epsilon^S \ ,$$

where the "it" subscripts have been dropped.[8] Consistent with most minimum wage research, W is interpreted as a relative wage measure, in particular the wage of young adults relative to the average wage for all workers. Also, following most existing minimum wage research, we assume that the coverage-adjusted relative minimum wage measures the effective minimum wage. Then if the minimum wage (MW) is binding, employment is determined by the demand curve with W set to MW,

$$(4) \qquad E = \alpha MW + X\beta + \epsilon^D \ .$$

If instead the minimum wage is non-binding, then employment is given by the intersection

---

[8]The inverse supply equation is specified with the vector δ having a negative coefficient so that the signs of the coefficients of the supply variables Y in the reduced-form employment equation (given below) agree with those in equation (1). (Note that equation (3) can be rewritten as $E = (1/\gamma)W + (\delta/\gamma)Y + \epsilon^S$.) In addition, throughout the ensuing discussion, we assume that the labor supply schedule is positively sloped. This is because the empirical work focuses on the employment rate, on which, at least in a one-period model, wage increases exert a positive substitution effect, but not a negative income effect.

of the demand and supply curves, or the reduced-form expression for employment

$$(5) \qquad E = \frac{1}{1-\alpha\gamma}(-\alpha Y\delta + X\beta) + \epsilon^{NB} \quad .$$

The error term ϵ^{NB} in equation (5) could be specified as the reduced form error term involving ϵ^{D} and ϵ^{S}. However, we use a less restrictive specification, which allows for the possibility that in the non-binding regime other factors affect employment, and for the possibility that ϵ^{D} and ϵ^{S} do not necessarily reflect demand or supply factors that carry over into the reduced form in the same way as the observable variables. This would be the case, for example, if ϵ^{D} and ϵ^{S} contain measurement error in the employment rate.

Finally, an observation will be classified on the binding regime (equation (4)) if the minimum wage is greater than the competitive equilibrium wage given by the intersection of the labor demand and labor supply curves (i.e., the reduced form for the wage), and vice versa. Thus, an observation is on the binding regime if

$$(6) \qquad MW > \frac{1}{1-\alpha\gamma}(\gamma X\beta - Y\delta) + \epsilon^{I} \quad ,$$

and is on the non-binding regime otherwise. Again, the error term ϵ^{I} is not restricted to have the form implied by the reduced form for the wage, in terms of ϵ^{D} and ϵ^{S}. However, we do allow for ϵ^{I} to be correlated with both ϵ^{D} and ϵ^{S}, and these correlations can reflect the appearance of ϵ^{D} and ϵ^{S} in the reduced form for the wage.

Equations (4), (5), and (6) constitute a disequilibrium model with unknown regimes. The model differs from standard cases in the literature (see Quandt, 1988) in that the switching equation (6) depends on the minimum wage and the variables and parameters in the labor demand and labor supply equations. The availability of the minimum wage as an "indicator" of the regime in which an observation lies makes the application of disequilibrium models to minimum wage effects unique.[9] In addition, because the dependent variable in the switching equation is not a latent variable, the variance of ϵ^I is identified.

Letting f denote density functions, and $Z\psi > \epsilon^I$ be a short-hand for the switching equation (6), we can write the likelihood for an observation as

$$(7) \qquad f(\epsilon^D, \epsilon^I < Z\psi) + f(\epsilon^{NB}, \epsilon^I \geq Z\psi) =$$
$$f(\epsilon^D) f(\epsilon^I < Z\psi \mid \epsilon^D) + f(\epsilon^{NB}) f(\epsilon^I \geq Z\psi \mid \epsilon^{NB}) \quad .$$

Assuming that ϵ^D, ϵ^{NB}, and ϵ^I have a trivariate normal distribution, we can use results for conditional distributions of bivariate normal random variables to obtain the likelihood function

$$(8) \qquad \frac{1}{\sigma_{\epsilon^D}}\phi\left(\frac{\epsilon^D}{\sigma_{\epsilon^D}}\right)\Phi\left(\frac{Z\psi - \frac{\sigma_{\epsilon^D,\epsilon^I}}{\sigma^2_{\epsilon^D}}\epsilon^D}{\left(\sigma^2_{\epsilon^I} - \frac{\sigma^2_{\epsilon^D,\epsilon^I}}{\sigma^2_{\epsilon^D}}\right)^{1/2}}\right) + \frac{1}{\sigma_{\epsilon^{NB}}}\phi\left(\frac{\epsilon^{NB}}{\sigma_{\epsilon^{NB}}}\right)\left[1 - \Phi\left(\frac{Z\psi - \frac{\sigma_{\epsilon^{NB},\epsilon^I}}{\sigma^2_{\epsilon^{NB}}}\epsilon^{NB}}{\left(\sigma^2_{\epsilon^I} - \frac{\sigma^2_{\epsilon^{NB},\epsilon^I}}{\sigma^2_{\epsilon^{NB}}}\right)^{1/2}}\right)\right] \quad .$$

Note that the coefficients of the demand function are identified regardless of which variables appear in the demand and supply functions, since the parameters of the demand function are identified from the binding regime. In contrast, the supply function parameters are not identified without some exclusion restrictions. Variables that appear in both the

---

[9]In most of the standard market disequilibrium literature, the question is whether a market is in excess supply or excess demand, and the "indicator" of these alternatives is whether prices are rising or falling (Quandt, 1988). Bowden (1978) considers the implications for such models of taking an approach closer to ours, by specifying the equilibrium price in terms of the underlying demand and supply curves.

demand and supply equations have coefficients $(-\alpha\delta + \beta)/(1 - \alpha\gamma)$ in the reduced-form employment equation, and $(\gamma\beta - \delta)/(1 - \alpha\gamma)$ in the switching equation. Thus, while α and β are identified from the binding regime, some simple algebra shows that γ and δ are not separately identified unless there is at least one variable that is excluded from the supply function but appears in the demand function; the coefficient of such a variable in the reduced-form employment equation, for example, is $\beta/(1 - \alpha\gamma)$, which identifies γ. This, of course, is the necessary condition for identification in the usual simultaneous equations case. In this particular model, we identify γ and δ by excluding the prime-age male unemployment rate from the supply function.[10]

In addition to obtaining estimates of employment effects of minimum wages (α) in the binding regime, the model can be used to estimate the probability that observations are on the alternative regimes, as functions of the minimum wage variable and the variables that determine the equilibrium wage. For example, the probability that minimum wages are binding is

$$P(\ MW > \frac{1}{1-\alpha\gamma}(\gamma X\beta\ -\ Y\delta)\ +\ \epsilon^{J}\ )\ =\ P(\ Z\psi\ >\ \epsilon^{J}\ )\ =\ \Phi(\frac{Z\psi}{\sigma_{\epsilon^{J}}})\ \ .$$

Such probabilities are informative as to how the overall minimum wage effect is likely to vary with changes in the minimum wage and other labor market variables.[11]

---

[10]We also follow the existing literature in assuming that the variable measuring the proportion of the age group in the population is a supply variable, and does not appear in the demand function, although this is not necessary for identification.

[11]In the multiple-regime approach, minimum wage effects vary depending on labor market characteristics that affect the probability that minimum wages are binding for an observation. In this sense, heterogeneity across workers is introduced. However, this heterogeneity is limited to variation in minimum wages and labor market conditions across

Finally, the two-regime model can be tested against two nested alternatives, one in which all observations are on the demand curve, and one in which all observations are on the reduced form for employment (i.e., at equilibrium employment).[12]

III.b. The Monopsony Model

In contrast to the competitive model, there are three regimes implied by the monopsony model. The first regime is that in which the minimum wage is non-binding, so that minimum wages have no effect on employment. The second regime, which arises uniquely in the monopsony model, is one in which minimum wages increase employment. This regime corresponds to the labor supply curve, between the monopsony wage (the wage along the labor supply curve at the employment level where the marginal cost of labor curve intersects the labor demand curve), and the competitive equilibrium wage. Finally, the third regime is the demand curve above the competitive equilibrium wage, along which minimum wages decrease employment, as in the competitive model.

In addition to the supply regime (in which the minimum wage effect is $1/\gamma$), we need to specify a switching equation between the non-binding regime and the supply regime.

---

states and years; within a state and year labor is treated as homogeneous, with an observation assumed to be on one regime or the other. A richer approach, which is beyond the scope of this paper, would attempt to account for the fact that minimum wages are binding for some individuals but not for others. As it now stands, our approach should not be interpreted as equivalent to one with heterogeneous workers, in which the cumulative probability expressions in equation (8) measure the proportion of workers in a state on one regime or the other.

[12]The coefficients of the reduced form involve all of the coefficients of the labor demand and supply curves. However, if all observations are on this regime these coefficients cannot be identified solely from the employment equation. Thus, in testing this restriction employment is specified simply as a linear function of the demand and supply variables.

(Equation (6) continues to describe switches between the supply regime and the demand regime, at the intersection of the demand and supply curves.) If we assume that under monopsony the inverse labor supply curve facing firms is the aggregate inverse supply curve, then the marginal cost of labor curve is[13]

$$(9) \qquad MC = 2\gamma E - Y\delta \ .$$

Using this, we can solve for the employment level at which marginal cost equals marginal product (the wage on the labor demand curve)

$$(10) \qquad E = -\frac{\alpha}{1-2\alpha\gamma}Y\delta + \frac{1}{1-2\alpha\gamma}X\beta \ .$$

Then we can find the wage that prevails at this level of employment, from the labor supply curve, which is

$$(11) \qquad W = -\frac{(1-\alpha\gamma)}{1-2\alpha\gamma}Y\delta + \frac{\gamma}{1-2\alpha\gamma}X\beta \ .$$

Thus, an observation lies on the demand regime if

$$(12) \qquad MW - \frac{1}{1-\alpha\gamma}(\gamma X\beta - Y\delta) > \epsilon^{J} \ .$$

An observation lies on the supply regime if

$$(13) \qquad MW - \frac{1}{1-\alpha\gamma}(\gamma X\beta - Y\delta) \le \epsilon^{J}$$

[13]In contrast to the standard monopsony model, the Rebitzer and Taylor (1991) efficiency wage model does not suggest any simple relationship between the labor supply and marginal cost of labor curves. In that model, the supervisory input is assumed to be fixed, so that, ceteris paribus, shirking increases with employment. Thus, the marginal cost of labor curve rises above the labor supply curve, because as employment increases firms must raise the wages of all workers to increase the cost of job loss in order to deter shirking.

and

$$(14) \qquad MW + \frac{(1-\alpha\gamma)Y\delta - \gamma X\beta}{1-2\alpha\gamma} > \epsilon^I \ .$$

Finally, an observation lies on the non-binding regime if

$$(15) \qquad MW + \frac{(1-\alpha\gamma)Y\delta - \gamma X\beta}{1-2\alpha\gamma} \leq \epsilon^I \ .$$

It is straightforward to verify that the equilibrium monopsony wage lies below the equilibrium competitive wage as long as $\alpha < 0$ and the equilibrium competitive level of employment is positive. This ensures that equations (12) and (15) cannot hold simultaneously.

Paralleling equation (7), and expressing equations (14) and (15) as $Z\psi' > \epsilon^I$ and $Z\psi' \leq \epsilon^I$, we can write the likelihood for an observation as

$$(16) \qquad f(\epsilon^D, \epsilon^I < Z\psi) + f(\epsilon^{NB}, \epsilon^I \geq Z\psi') + f(\epsilon^S, Z\psi \leq \epsilon^I < Z\psi') =$$
$$f(\epsilon^D) f(\epsilon^I < Z\psi \mid \epsilon^D) + f(\epsilon^{NB}) f(\epsilon^I \geq Z\psi' \mid \epsilon^{NB}) + f(\epsilon^S) f(Z\psi \leq \epsilon^I < Z\psi \mid \epsilon^S) \ .$$

Collecting results for the monopsony model, again assuming that the errors (now ϵ^D, ϵ^S, ϵ^{NB}, and ϵ^I) have a multivariate normal distribution, and using an expression paralleling equation (7), we obtain the likelihood function

$$(17) \qquad \frac{1}{\sigma_{\epsilon^D}}\phi\left(\frac{\epsilon^D}{\sigma_{\epsilon^D}}\right)\Phi\left(\frac{Z\psi - \frac{\sigma_{\epsilon^D,\epsilon^I}}{\sigma^2_{\epsilon^D}}\epsilon^D}{\left(\sigma^2_{\epsilon^I} - \frac{\sigma^2_{\epsilon^D,\epsilon^I}}{\sigma^2_{\epsilon^D}}\right)^{1/2}}\right) + \frac{1}{\sigma_{\epsilon^{NB}}}\phi\left(\frac{\epsilon^{NB}}{\sigma_{\epsilon^{NB}}}\right)\left(1 - \Phi\left(\frac{Z\psi' - \frac{\sigma_{\epsilon^{NB},\epsilon^I}}{\sigma^2_{\epsilon^{NB}}}\epsilon^{NB}}{\left(\sigma^2_{\epsilon^I} - \frac{\sigma^2_{\epsilon^{NB},\epsilon^I}}{\sigma^2_{\epsilon^{NB}}}\right)^{1/2}}\right)\right) +$$

$$\frac{1}{\sigma_{\epsilon^s}}\phi\left(\frac{\epsilon^S}{\sigma_{\epsilon^s}}\right)\left[\Phi\left(\frac{Z\psi' - \frac{\sigma_{\epsilon^s,\epsilon^d}}{\sigma^2_{\epsilon^s}}\epsilon^S}{\left(\sigma^2_{\epsilon^d} - \frac{\sigma^2_{\epsilon^s,\epsilon^d}}{\sigma^2_{\epsilon^s}}\right)^{1/2}}\right) - \Phi\left(\frac{Z\psi - \frac{\sigma_{\epsilon^s,\epsilon^d}}{\sigma^2_{\epsilon^s}}\epsilon^S}{\left(\sigma^2_{\epsilon^d} - \frac{\sigma^2_{\epsilon^s,\epsilon^d}}{\sigma^2_{\epsilon^s}}\right)^{1/2}}\right)\right] .$$

In contrast to the two-regime competitive model, the parameters are identified in the three-regime monopsony model even if all variables appear in both equations, since we have separate observations (with some probability) along the demand and supply regimes. Nonetheless, to maintain comparability we impose the same exclusion restrictions as in the two-regime case.

Because the two-regime and three-regime models are not nested, testing the competitive model against the monopsony model is not as straightforward as testing whether a second regime is needed, relative to a single-regime model.[14] Thus, to compare the models we use the non-nested test procedure developed by Vuong (1989). The applicable test statistic is $n^{-1/2}(L_3 - L_2)/\omega_2$, where L_3 and L_2 are the log-likelihoods for the three- and two-regime models respectively, evaluated at the maximum likelihood estimates. ω_2 is the sample variance of the difference between the log-likelihoods for the three-regime and two-regime models evaluated at each data point. This statistic is distributed asymptotically as N(0,1). Thus, if both models have the same number of parameters, then at the five-percent significance level, if this test statistic exceeds 1.96 we reject the two-regime model in favor

[14]To see that the models are non-nested, note that although the three-regime model involves more parameters than the two-regime model, the three-regime model may be more restrictive than the two-regime model because observations are forced to be on the supply regime, and hence to be unaffected by the demand variables, with some non-zero probability.

of the three-regime model, and vice versa if it is below -1.96. Between these critical values, the models have statistically indistinguishable fits. However, the three-regime monopsony model has two more parameters than the two-regime competitive model, necessitating an upward adjustment of the lower and upper critical values by $2/(\omega_2 \cdot n^{1/2})$.

IV. Results

IV.a. Results from the Disequilibrium Endogenous Switching Models

Results for the two-regime model corresponding to the competitive model of low-wage labor markets are reported in column (1) of Table 2. The estimated minimum wage effect in the binding regime is -.33, considerably stronger than the corresponding single-equation estimate in Table 1. However, it is important to recognize that this effect applies to different observations with probabilities depending on the value of the minimum wage and other labor market variables, and that a sizable proportion of the observations may come from the non-binding regime.[15] Calculations based on the estimated model parameters indicate that the estimated mean probability that an observation is on the binding regime is .48. Another metric, the proportion of observations for which the probability of being on the binding demand regime is higher than the probability of being on the non-binding regime, shows a similar value of .46.

The two-regime model can be tested relative to two nested models, one in which all observations lie on the binding regime, and one in which all observations lie on the non-binding regime (the reduced form for employment). These nested models are based on the

[15]This is similar to what Brown (1988) calls the "fallacy of the inflated denominator," by which he means the possibility that standard minimum wage equations convey small minimum wage effects because many of the workers for whom they are estimated are not affected by minimum wages.

exclusion restrictions discussed previously. Thus, the binding regime includes the minimum wage variable and excludes the supply variable(s), while the non-binding regime includes all variables except the minimum wage variable. The log-likelihoods for these alternative models, and the number of restrictions implied, are also reported in the table. Likelihood-ratio tests indicate that both nested models are rejected at less than the five-percent level.[16]

Finally, we can use the estimated parameters to solve for the mean competitive wage above which minimum wages are binding. The estimate, expressed relative to the average wage for all workers, is .35.

Column (2) of Table 2 reports estimates of the three-regime model, adding the supply regime implied by the monopsony model. The estimated minimum wage effect along the binding demand regime falls to -.21, while remaining significant. The minimum wage effect along the binding supply regime is equal to $1/\gamma$, the estimated value of which is 1.75. The other signs are as expected.[17]

Even if the monopsony characterization of low-wage labor markets is correct, we would not expect a large number of observations to be classified on the supply regime with

[16]We also tested the restriction that the proportion of age group in population variable could be excluded from the demand curve, by estimating the model without imposing this restriction, in which case the parameters are still identified. The estimated coefficient (standard error) of this variable in the demand curve was -.01 (.11).

[17]Partly as a specification check, we also estimated the model including the proportion of population in the age group in the demand function, and the prime-age male unemployment rate in the supply function, in which case the parameters are still identified. The estimated coefficient (standard error) of the population proportion variable was .06 (.11), and that of the unemployment rate was .38 (.24). The log likelihood was 1071.5, compared with a value of 1070.7 in the table, so the exclusion of these variables is not rejected.

a high probability. This turns out to be the case. The mean probability that an observation is on the supply regime is .20, and for only one observation is the probability of being on the supply regime higher than that of being on one of the other regimes. This is also reflected in the mean competitive and monopsony wages implied by the estimates of the structural parameters. At .34 and .31 respectively, these estimates indicate that the region over which minimum wages might increase employment is relatively small. Corresponding to these findings, the likelihood for the model restricting all observations to lie on the supply regime is extremely low compared to the three-regime model. The two other nested models restricting the observations to lie solely on the demand regime or solely on the non-binding regime are also rejected at the five-percent level.

We can also assess the fit of the monopsony model relative to the competitive model by means of the non-nested test outlined above. As the table shows, the likelihood is in fact higher for the three-regime monopsony model than for the two-regime competitive model. However, the Vuong test statistic (1.34), along with the critical values reported in the table, indicate that the statistical fit of the three-regime model is not significantly better. Nonetheless, the data clearly do not reject the restriction that, for some observations, employment is best characterized as determined only by the minimum wage and supply variables.[18]

IV.b. Accounting for State and Year Effects

The results in Table 2 ignore the possibility that heterogeneity across states and years

---

[18]As an additional specification check on the model, we computed the implied equilibrium competitive and monopsony wages for each observation. This is potentially informative since nothing restricts these wages to be positive. For both the two- and three-regime models, no observations had implied equilibrium wages that were negative.

may bias the estimated minimum wage effects. In single-equation estimates this issue has typically been dealt with by including fixed state and year effects in the regression equations (e.g., Neumark and Wascher, 1992). Thus, to assess the importance of heterogeneity bias in these results, we add a set of dummy variables for the nine Census regions to the demand equation, and a time trend to the supply equation. We include the Census dummies in the demand equation, since the standard argument is that demand shocks may generate a negative correlation between employment rates and the minimum wage variable, through a positive correlation between employment rates and wage levels (Freeman, 1982). We put the time trend in the supply equation because we assume that this trend stems from supply factors.[19]

The results are reported in Table 3. For the two-regime model, but not the three-regime model, the negative minimum wage effect falls by about one-half relative to the estimate in Table 2, to -.18, but remains negative and significant. For both models, the proportion of observations classified on the binding demand regime is well above that in Table 2. Finally, for this specification, the Vuong test statistic indicates that the likelihood for the three-regime monopsony model is significantly higher than that for the two-regime

[19]Not surprisingly, we could not get the models to converge with dummy variables for each state or each year. Specifications with Census region dummies introduced into both the demand and supply equation also led to convergence problems. Substituting a time trend for year dummies seems appropriate because the estimated coefficients of year dummies in single employment equations fit a linear trend relatively well. For the purposes of comparison, when the standard minimum wage equation reported in Table 1 (Panel A) was estimated using these region dummy variables and the time trend, the estimated minimum wage effect (standard error) was .01 (.05). Thus, the inclusion of these variables does at least as much as the inclusion of fixed state and year effects to reduce the negative employment effects of minimum wages.

model.[20]

IV.c. Variation in Minimum Wage Effects and the Probability that Minimum Wages are Binding

The disequilibrium framework for estimating minimum wage effects emphasizes that the employment effects of minimum wages may vary in systematic ways depending on whether the minimum wage is binding. This suggests that if we estimate the single employment equation using observations with a relatively higher probability that minimum wages are binding, we should find relatively stronger disemployment effects of minimum wages. On the other hand, when we apply our multiple-regime approach to subsamples of observations with different probabilities that minimum wages are binding, we should find two results: the estimated probability that minimum wages are binding should reflect this difference; and estimated minimum wage effects along the binding regime (α) should differ less across the subsamples than do the single-equation estimates.

We carry out this experiment using subsamples with high and low probabilities that minimum wages are binding, based on estimates of the two-regime model for the whole sample. We restrict attention to the two-regime model, since in the estimates of the three-regime model, there were virtually no observations for which the probability that observations were on the binding supply regime were highest.

Table 4 reports results for subsamples of observations based on whether the estimated probability of being on the binding demand regime was greater or less than .5. As expected, in the subsample with high probabilities that minimum wages are binding the

[20]For these specifications, there was one observation for which the three-regime model implied a negative equilibrium monopsony wage.

single-equation estimate of the minimum wage effect (-.56) is negative and significant. In contrast, in the subsample with low probabilities that minimum wages are binding, this estimate is actually positive and significant. The multiple-regime approach satisfies both of the predictions described above. First, the two-regime estimates for each subsample produce very different estimates of the probability that minimum wages are binding. These estimates exceed .9 for the high probability subsample, and are less than .1 for the low-probability subsample.[21] Second, the estimated minimum wage effects for the two subsamples on the binding regimes are much more consistent than the single-equation estimates, as both are negative and significant.

Finally, it seems worthwhile to assess the validity of the estimated probabilities of lying on alternative regimes. One way to do this is to compute for each state and year the proportion of young adults paid hourly wages equal to or less than the prevailing minimum wage. This proportion should be positively related to the probability that minimum wages are binding. This is confirmed, with an estimated correlation of .31 between the estimated probability from the two-regime model, and the proportion at or below the minimum.[22]

More generally, the estimates from the multiple-regime models can be used to

---

[21]The mean estimated probabilities of being on the binding demand regime based on the estimates for the full sample are .80 and .21 for the two subsamples.

[22]This exercise may also suggest the possibility of estimating the two-regime employment model with the proportion paid at or below the minimum wage replacing the probability that a state's minimum wage is binding. However, this approach is invalid, since it fails to account for the employment at or below the minimum wage that would have occurred in the absence of the minimum. This same argument explains why we cannot make any prediction as to the magnitude of the correlation between the estimated probability that minimum wages are binding, based on the two-regime model, and the proportion at or below the minimum wage; in particular, it need not be near one.

compute minimum wage effects for alternative scenarios regarding minimum wages and labor market conditions, which in turn may be useful in interpreting variation in single-equation estimates of minimum wage effects across different samples. In this context, Table 5 reports the average probabilities that observations are on the binding regime and the average elasticities of employment rates with respect to minimum wages, for each year in our sample period and for each Census division, based on the estimates in Table 2. The elasticity for each observation is estimated as the product of the probability of being on the binding regime, the estimated minimum wage effect, and the ratio of the minimum wage variable to the employment rate.

For the two-regime model, the estimated average probabilities of being on the binding regime, shown in column (2), vary considerably over both years and regions, ranging from .05 to .87. Reflecting this variation, the average elasticities shown in column (3) also vary considerably, ranging from -.01 to -.24. The estimated minimum wage elasticities rise (in absolute value) as the federal minimum wage increases (relative to the mean wage) through 1982, and then fall as the federal minimum wage declines in real terms over the remainder of the 1980s. The variation in the elasticities across regions also corresponds to variation in the coverage-adjusted relative minimum wage (to which it is related by construction). In the low-wage East South Central, West South Central, and South Atlantic states the relative minimum wage variable is high, and the estimated minimum wage elasticity ranges from -.15 to -.24. In contrast, in the high-wage Pacific region the relative minimum wage variable is lowest, and the estimated minimum wage elasticity is only -.05. These calculations, coupled with the single-equation estimates in Table 4 for the subsamples of observations with high and low probabilities of binding minimum wages, suggest that estimated minimum wage effects are quite sensitive to the probability that minimum wages are binding.

As a final exercise, the last two columns of Table 5 report results based on the three-regime monopsony model. Column (4) reports the estimated probabilities that observations are on the binding demand regime, along which minimum wages are estimated to reduce employment. These probabilities are similar to those estimated from the competitive model. Column (5) reports the estimated probabilities of being on the regime that is unique to the monopsony model, namely the binding supply regime along which minimum wages are estimated to increase employment. Except for 1973, the average estimated probability of being on the binding supply regime never exceeds the average estimated probability of being on the binding demand regime. Nonetheless, the probabilities of being on the binding supply regime are clearly non-negligible, and in many cases are above .2. Thus, our results admit the possibility that some minimum wage studies produce positive effects of minimum wages on employment because they use samples for which observations are concentrated along the monopsony regime.

IV.d. A Non-Structural Approach

To gauge the dependence of our results on the assumed structure for the demand, supply, and marginal cost curves, we also consider a non-structural approach that allows the parameters of the standard minimum wage-employment equation (1) to vary depending, in a simple way, on how likely the minimum wage is to be binding. As an analog to the two-regime competitive model, we first allow the coefficient estimates of equation (1) to take on two distinct sets of values, one for observations for which the coverage-adjusted relative minimum wage is relatively low, and one for observations for which it is relatively high.

Rather than specify a priori "high" and "low" values, we use a grid search over the possible range of the coverage-adjusted relative minimum wage, and select the "switch point" for this variable that maximizes the likelihood for the sample.[23]

Estimates in which the equations in each of the two regimes are unconstrained are reported in the first column of Table 6. For the binding regime, which refers to those observations with values of the coverage-adjusted relative minimum wage above the switch point chosen by the grid search, the estimated minimum wage effect is -.41, more than double that from the single-equation estimates, and is significant. For the non-binding regime the estimated minimum wage effect is positive and significant. 58 percent of the observations are classified as being on the binding regime, and the remaining 42 percent as being on the non-binding regime.

In columns (2) and (3) we constrain some of the parameters in each of the regimes in ways suggested by theory, to attempt to increase the correspondence between the theory and the regimes chosen by the grid search procedure. First, in column (2), we constrain the minimum wage effect to be zero in the non-binding regime (and recompute the grid search). The chosen sample separation is the same as in column (1), with, accordingly, the same estimated minimum wage effect in the binding regime. In column (3) we further constrain the model to exclude the supply variable (the proportion of population) from the binding regime, which is supposed to be the demand curve. Since the non-binding regime is the

---

[23]For each iteration of the grid search, we define a dummy variable $B_{it} = 1$ if the minimum wage variable was above the switch point, and 0 otherwise, and then estimate the model

$$E_{it} = (\alpha_B MW_{it} + R_{it}\beta_B)\cdot B_{it} + (\alpha_{NB}MW_{it} + R_{it}\beta_{NB})\cdot(1 - B_{it}) \ .$$

reduced form for employment, it should include both demand and supply variables, so no further constraints are imposed on this regime. The estimates are again little changed with this additional restriction. The likelihood-ratio tests of any of the two-regime models in Table 6 relative to the constrained one-regime model in Table 1 lead to rejection of the latter.

The two-regime estimates from the non-structural approach correspond to those from the structural approach along a number of dimensions. First, the classification of observations on the binding vs. the non-binding regimes is similar, with the non-structural approach assigning .58 of the observations to the binding regime (for all three models in Table 6), while the structural estimates imply that the probability of being on the binding regime exceeds one-half for .46 of the observations. Second, the estimated switch point for all three models in Table 6 is .34. This is close to the mean equilibrium competitive wage of .35 implied by the structural estimates in Table 2. Finally, the parameter estimates for the binding regime from the non-structural approach should correspond to the labor demand curve, and hence can be compared directly to the estimated parameters of the demand curve from the structural approach. These parameter estimates are quite close, as can be seen from comparing the estimated coefficients of the minimum wage variable and the unemployment rate in the top panel of column (3) of Table 6 with the demand curve estimates in column (1) of Table 2. Together, these similarities boost our confidence in the structural approach.

Next, we extend the non-structural approach to three regimes, corresponding to the monopsony model, by allowing two switch points for the coverage-adjusted relative minimum wage. Results for unconstrained equations are reported in the first column of

Table 7. As the monopsony model would lead us to expect, there are positive minimum wage effects in the middle regime, on which 39 percent of the observations are classified. As in the two-regime model, 58 percent of the observations are classified on the binding regime above the higher switch point (i.e., the demand curve), so the estimated minimum wage effect for this region is again -.41. Finally, only three percent of the observations are classified on the non-binding regime (below the lower switch point); the estimated minimum wage effect for this regime is positive but not significant.

In column (2) the model is constrained so that there is no minimum wage effect in the non-binding regime. The estimates are little changed. In column (3) the proportion of population variable is again excluded from the regime that should correspond to the demand curve, and the unemployment rate is excluded from the regime that should correspond to the supply curve. This results in a substantial reduction (to .06) in the proportion of observations classified on the supply curve. In addition, the estimated positive minimum wage effect in this regime is now insignificant, with the standard error rising more than ten-fold.

Again, the estimates from the non-structural approach can be compared with those from the structural approach along a number of dimensions. First, using the column (3) non-structural estimates, the classification of observations on the alternative regimes is similar. The non-structural approach assigns .84 of the observations to the upper binding regime, and .06 to the lower binding regime, while the structural estimates imply that the probability of being on the binding demand regime is highest for .70 of the observations, and the probability of being on the binding supply regime is highest for .001 of the observations. Second, the estimated switch points of .31 and .30 correspond closely to the mean equilibrium competitive and monopsony wages implied by the structural estimates, reported in Table 2, of .34 and .31. Third, the parameter estimates for the upper binding regime from the non-structural approach should correspond to the labor demand curve, and those from the lower binding regime should correspond to the labor supply curve. Comparing the non-structural estimates in column (3) of Table 7 with those in column (2) of Table 2, the correspondence holds in terms of the estimated signs of all of the parameters, and many of the estimated coefficients are quite close.

Finally, the structural and non-structural approaches can be compared with respect to testing the monopsony vs. the competitive model. In the non-structural approach, for the constrained model the likelihood-ratio test of the three-regime model vs. the two-regime model (which has two fewer parameters) leads to a rejection of the two-regime model at the ten-percent significance level, but not the five-percent level. Thus, the non-structural approach provides a bit more support for the monopsony characterization of low-wage labor markets than did the structural approach, although the structural approach also rejected the competitive model in the specifications including region and year effects.

Although we have presented results from the nonstructural specification for purposes of comparison, in our view this approach has several weaknesses as compared with the structural approach on which the endogenous switching model is based. First, the nonstructural approach classifies observations as unambiguously on one of the alternative regimes, whereas the structural approach produces a probabilistic classification of being on the alternative regimes, which captures the inherent uncertainty in classifying observations. Second, the classification indicated by the nonstructural approach depends solely on the value of the minimum wage variable. Although whether or not the minimum wage is binding should depend only on the minimum wage relative to the market-determined wage for affected workers in the absence of minimum wages, the latter is of course unobserved. Therefore, factors affecting the demand for and supply of young workers are likely to provide some information on the likelihood that the minimum wage is binding. The structural approach incorporates this information explicitly by specifying the switching equation in terms of the minimum wage relative to the equilibrium competitive (and monopsony) wage, which are functions of the demand and supply determinants. Finally, while the non-structural approach is intended to identify switch points associated with the structural regimes in the model, we cannot rule out the possibility that it has instead identified switch points associated with nonlinearities in the equation or outliers. The structural approach, in contrast, explicitly specifies the "switch points" in terms of intersections of the demand, supply, and marginal cost of labor curves.

V. Conclusions

In the competitive model of low-wage labor markets, a legislated minimum wage below the equilibrium competitive wage should be unrelated to employment. In contrast, a minimum wage higher than the equilibrium competitive wage should be negatively related to employment. Like the competitive model, the monopsony model implies that there is a region below which minimum wages are non-binding and do not affect employment, and a region above which minimum wages trace out the labor demand curve and reduce employment. In addition, there is a region between these two, along which minimum wages trace out the labor supply curve, and raise employment. We therefore may learn more about minimum wage effects by separating sample observations into these various regimes, and estimating minimum wage effects along the appropriate regimes. This paper develops and implements such an approach, using methods adapted from the market disequilibrium literature.

We claimed at the outset that this approach has three potential advantages relative to existing single-equation estimates: it leads to estimates of the employment effects of minimum wages when they are binding; it yields a (probabilistic) classification of observations based on whether or not minimum wages are binding; and it provides a test of the monopsony characterization of low-wage labor markets. We summarize the results in terms of each of these potential advantages.

First, the two-regime competitive model and the three-regime monopsony model yield significant negative estimated effects of minimum wages on employment in the binding demand regimes, under a variety of alternative specifications. Not surprisingly, these estimated disemployment effects are generally stronger than single-equation estimates.

Second, calculations using the estimated probabilities of lying on alternative regimes implied by the models suggest that the effects of minimum wages on employment may vary considerably depending on the prevailing minimum wage and other labor market conditions. This may help to explain some of the variation in estimated minimum wage effects in recent research.

Third, estimates of the three-regime monopsony model suggest that a small fraction of observations may lie on a supply regime along which minimum wages increase employment. The three-regime monopsony model never fits the data worse than the two-regime competitive model, and sometimes fits the data significantly better.

Finally, the sensitivity of minimum wage effects on employment to the relative magnitude of the minimum wage, and to other labor market conditions, implies that the multiple-regime, disequilibrium approach is useful even if we simply want to predict the employment effects of minimum wages, rather than to uncover the structural estimates of the labor demand and supply curves. In particular, single-equation estimates may be unreliable when the minimum wage and other labor market conditions diverge from those prevailing in the sample from which the estimates are computed. In contrast, the disequilibrium approach can provide a richer set of information about the effects of minimum wages, including whether or not minimum wages are likely to be binding, how changes in other labor market conditions influence minimum wage effects, and how minimum wages are likely to affect employment under a variety of scenarios regarding the minimum wage and other labor market characteristics.

References

Bowden, Roger J. 1978. "Specification, Estimation and Inference for Models of Markets in Disequilibrium." International Economic Review 19(3): 711-26.

Brown, Charles. 1988. "Minimum Wage Laws: Are They Overrated?" Journal of Economic Perspectives 2(3): 133-46.

_____, Curtis Gilroy, and Andrew Kohen. 1982. "The Effect of the Minimum Wage on Employment and Unemployment." Journal of Economic Literature 20: 487-528.

Card, David. 1992a. "Using Regional Variation in Wages to Measure the Effects of the Federal Minimum Wage." Industrial and Labor Relations Review 46(1): 22-37.

_____. 1992b. "Do Minimum Wages Reduce Employment? A Case Study of California, 1987-1989." Industrial and Labor Relations Review 46(1): 38-54.

Card, David, Lawrence F. Katz, and Alan B. Krueger. 1993. "An Evaluation of Recent Evidence on the Employment Effects of Minimum and Subminimum Wages." NBER Working Paper No. 4528.

Card, David, and Alan B. Krueger. 1993. "Minimum Wages and Employment: A Case Study of the Fast Food Industry in New Jersey and Pennsylvania." NBER Working Paper No. 4509.

Currie, Janet, and Bruce Fallick. 1993. "A Note on the New Minimum Wage Research." NBER Working Paper No. 4348.

Freeman, Richard B. 1982. "Economic Determinants of Geographic and Individual Variation in the Labor Market Position of Young Persons." In Richard B. Freeman and David A. Wise, Eds., The Youth Labor Market Problem: Its Nature, Causes, and Consequences (Chicago: University of Chicago Press), pp. 115-54.

Katz, Lawrence F. and Alan B. Krueger. 1992. "The Effect of the Minimum Wage on the Fast-Food Industry." Industrial and Labor Relations Review 46(1): 6-21.

Neumark, David and William Wascher. 1993. "Employment Effects of Minimum and Subminimum Wages: Reply to Card, Katz and Krueger." Economic Activity Section Working Paper No. 144, Federal Reserve Board, Washington, D.C.

_____. 1992. "Employment Effects of Minimum and Subminimum Wages: Panel Data on State Minimum Wage Laws." Industrial and Labor Relations Review 46(1): 55-81.

Quandt, Richard E. 1988. The Econometrics of Disequilibrium (Oxford: Basil Blackwell Ltd.).

Rebitzer, James B. and Lowell J. Taylor. 1991. "The Consequences of Minimum Wage Laws: Some New Theoretical Ideas." NBER Working Paper No. 3877.

Spriggs, William. 1992. "The Effects of Changes in the Federal Minimum Wage: Restaurant

Workers in Mississippi and North Carolina." Unpublished Paper, Economic Policy Institute, Washington, D.C.

Stigler, George J. 1946. "The Economics of Minimum Wage Legislation." American Economic Review 36(3): 358-65.

Taylor, Lowell J., and Taeil Kim. 1993. "The Employment Effect in Retail Trade of California's 1988 Minimum Wage Increase." Heinz School of Public Policy and Management Working Paper No. 93-16.

Vuong, Quang H. 1989. "Likelihood Ratio Tests for Model Selection and Non-Nested Hypotheses." Econometrica 57(2): 307-34.

Wellington, Allison, J. 1991. "Effects of the Minimum Wage on the Employment Status of Youths: An Update." Journal of Human Resources 26(1): 27-46.

West, E. G. and Michael McKee. 1980. "Monopsony and "Shock" Arguments for Minimum Wages." Southern Economic Journal (January): 883-91.

Williams, Nicolas. 1993. "Regional Effects of the Minimum Wage on Teenage Employment." Unpublished Paper, University of Cincinnati.

</ref_section>

Table 1: Estimates of Minimum Wage-Employment Equation for Young Adults (16-24)

Panel A. Minimum Wage-Employment Equation Estimates

Coverage-adjusted relative minimum wage	-.14 (.05)
Proportion of population in age group	-.11 (.08)
Prime-age male unemployment rate	-1.17 (.08)
Log-likelihood	1049.4

Panel B. Estimated Minimum Wage Coefficients from Alternative Specifications

Including fixed state and year effects	-.11 (.07)
Substituting lagged minimum wage variable for contemporaneous variable	-.21 (.05)

There are 751 observations covering the period 1977-1989 for all states, and 1973-1989 for 22 states identified in the CPS beginning in 1973, except in the specification using the lagged minimum wage variable. Standard errors are reported in parentheses. Estimated intercepts are not reported. The ratio of the means of the employment rate to the minimum wage variable, which yields the employment elasticity when multiplied by the minimum wage coefficient, is .64.

Table 2: Estimates of Two- and Three-Regime Disequilibrium Models for Young Adults (16-24)

	Two Regimes (1)		Three Regimes (2)
Labor demand curve:			
Coverage-adjusted relative minimum wage (α)	-.33		-.21
	(.13)		(.08)
Prime-age male unemployment rate (β_1)	-1.28		-1.22
	(.12)		(.09)
Intercept (β_0)	.74		.69
	(.06)		(.03)
Inverse labor supply curve:			
Employment rate (γ)	.71		.57
	(.24)		(.12)
Proportion of population in age group (δ_1)	-.54		-.63
	(.19)		(.17)
Intercept (δ_0)	.16		.11
	(.15)		(.08)
σ_ϵ^D	.06	σ_ϵ^D	.06
	(.005)		(.004)
		σ_ϵ^S	.05
			(.01)
σ_ϵ^{NB}	.06	σ_ϵ^{NB}	.05
	(.004)		(.01)
σ_ϵ^S	.03	σ_ϵ^S	.04
	(.01)		(.01)
$\sigma_{\epsilon^D,\epsilon^1}$	-.0003	$\sigma_{\epsilon^D,\epsilon^1}$	-.001
	(.001)		(.005)
		$\sigma_{\epsilon^S,\epsilon^1}$	-.001
			(.001)
$\sigma_{\epsilon^{NB},\epsilon^1}$	.001	$\sigma_{\epsilon^{NB},\epsilon^1}$	.002
	(.001)		(.001)
Log-likelihood	1065.6		1070.7
Log-likelihood (# restrictions)			
All observations on demand regime	1048.5 (7)		1048.5 (9)
All observations on supply regime	...		951.8 (9)
All observations on non-binding regime	1045.9 (7)		1045.9 (9)
Mean probability on demand regime	.48		.55
Mean probability on supply regime	...		.20
Proportion with probability on demand regime highest	.46		.70
Proportion with probability on supply regime highest	...		.001
Equilibrium competitive wage implied by structural estimates	.35		.34
	(.01)		(.01)
Equilibrium monopsony wage implied by structural estimates	...		.31
			(.01)
Vuong test statistic vs. two-regime model	...		1.34
(lower and upper 5% critical values)			(-1.43,2.49)

Table 2 (continued)

There are 751 observations covering the period 1977-1989 for all states, and 1973-1989 for 22 states identified in the CPS beginning in 1973. Standard errors are reported in parentheses. Probabilities of observations lying on alternative regimes are based on unconditional probabilities. Implied equilibrium competitive and monopsony wages are computed at sample means. Standard errors are computed by linearizing the expression for the equilibrium wage, treating the means as fixed, and applying the estimated variances and covariances of the parameters. The critical values for the Vuong test statistic are not centered around zero because they have been adjusted for the greater number of parameters in the three-regime model.

Table 3: Estimates of Two- and Three-Regime Disequilibrium Models, Including Regional Demand Dummy Variables and Supply Time Trend

	Two Regimes (1)		Three Regimes (2)
Labor demand curve:			
Coverage-adjusted relative minimum wage (α)	-.18		-.20
	(.06)		(.06)
Prime-age male unemployment rate (β_1)	-1.12		-1.27
	(.07)		(.08)
Intercept (β_0)	.67		.69
	(.02)		(.02)
Inverse labor supply curve:			
Employment rate (γ)	.95		1.02
	(.12)		(.22)
Proportion of population in age group (δ_1)	.05		.74
	(.24)		(.36)
Intercept (δ_0)	.21		.04
	(.06)		(.12)
σ_ϵ^D	.05	σ_ϵ^D	.05
	(.002)		(.002)
		σ_ϵ^S	.05
			(.01)
σ_ϵ^{NB}	.08	σ_ϵ^{NB}	.07
	(.01)		(.01)
σ_ϵ^I	.05	σ_ϵ^I	.08
	(.01)		(.02)
$\sigma_{\epsilon^D,\epsilon^I}$	.003	$\sigma_{\epsilon^D,\epsilon^I}$	.004
	(.0005)		(.001)
		$\sigma_{\epsilon^S,\epsilon^I}$	-.003
			(.002)
$\sigma_{\epsilon^{NB},\epsilon^I}$	-.002	$\sigma_{\epsilon^{NB},\epsilon^I}$	-.005
	(.001)		(.001)
Log-likelihood	1220.9		1234.2
Log-likelihood (# restrictions)			
All observations on demand regime	1186.6 (8)		1186.6 (8)
All observations on supply regime	...		962.1 (17)
All observations on non-binding regime	1211.9 (8)		1211.9 (8)
Mean probability on demand regime	.91		.80
Mean probability on supply regime	...		.15
Proportion with probability on demand regime highest	.95		.91
Proportion with probability on supply regime highest	...		.06
Vuong test statistic vs. two-regime model	...		2.32
(lower and upper 5% critical values)			(-1.61,2.31)

See footnotes to Table 2 for details.

Table 4: Estimates of Single-Equation and Two-Regime Models for Observations with High and Low Probabilities that Minimum Wages are Binding, Based on Estimates in Table 2

	P(binding) > .5 (N=343) (1)	P(binding) ≤ .5 (N=408) (2)
Single equation:		
Coverage-adjusted relative minimum wage	-.56 (.12)	.34 (.12)
Log-likelihood	472.7	593.1
Two regimes:		
Coverage-adjusted relative minimum wage (α)	-.52 (.12)	-.28 (.08)
Log-likelihood	473.0	601.2
Mean probability on demand regime	.97	.09
Proportion with probability on demand regime highest	.99	.00

See footnotes to Table 2 for details. The covariance terms for the errors were restricted to zero. Without this restriction the models did not converge for the two subsamples.

Table 5: Estimated Elasticities of Employment Rate with Respect to Minimum Wage, Based on Results in Table 2

	Avg. cov.-adj. relative minimum	Competitive Model (Two Regimes) Avg. prob. on binding regime	Avg. prob.-weighted elasticity	Monopsony Model (Three Regimes) Avg. prob. on binding demand regime	Avg. prob. on binding supply regime
	(1)	(2)	(3)	(4)	(5)
Year					
1973	.31	.05	-.01	.10	.18
1974	.35	.27	-.06	.36	.25
1975	.34	.44	-.10	.49	.22
1976	.35	.44	-.10	.51	.22
1977	.35	.35	-.08	.41	.16
1978	.38	.50	-.12	.57	.20
1979	.38	.52	-.12	.59	.21
1980	.38	.65	-.15	.70	.17
1981	.38	.67	-.16	.72	.15
1982	.35	.61	-.14	.65	.17
1983	.34	.62	-.14	.65	.17
1984	.33	.40	-.09	.48	.22
1985	.35	.55	-.12	.64	.19
1986	.34	.52	-.11	.62	.20
1987	.34	.45	-.10	.56	.22
1988	.33	.38	-.08	.50	.23
1989	.32	.29	-.06	.43	.25
Census Geographic Division (1977-1989)					
New England	.37	.61	-.13	.69	.16
Middle Atlantic	.33	.37	-.08	.46	.22
East North Central	.34	.50	-.10	.58	.22
West North Central	.35	.40	-.08	.50	.25
Mountain	.34	.35	-.07	.44	.25
Pacific	.31	.26	-.05	.34	.23
East South Central	.39	.87	-.24	.90	.07
West South Central	.37	.63	-.15	.69	.18
South Atlantic	.36	.59	-.15	.66	.16

Numbers reported in columns (2) and (3) are based on average effects for the available observations based on probabilities and elasticities evaluated at the state-by-year level. The increased coverage-adjusted relative minimum wage in 1985 is partly attributable to a rise in coverage owing to a court ruling that state and local government workers were subject to the minimum wage. For some of the earlier years, minimum wage level is for workers previously covered under the FLSA, whereas newly covered workers came in at lower minimum wages. All means are unweighted.

Table 6: Estimates of Two-Regime Model, Single Switch Point, Grid Search Results

	(1)	(2)	(3)
Binding regime:			
Coverage-adjusted relative minimum wage	-.41	-.41	-.40
	(.10)	(.10)	(.10)
Proportion of population in age group	.07	.07	0.0
	(.11)	(.11)	
Prime-age male unemployment rate	-1.34	-1.34	-1.35
	(.11)	(.11)	(.11)
Proportion of observations	.58	.58	.58
Non-binding regime:			
Coverage-adjusted relative minimum wage	.35	0.0	0.0
	(.15)		
Proportion of population in age group	-.33	-.33	-.33
	(.11)	(.11)	(.11)
Prime-age male unemployment rate	-.88	-.93	-.93
	(.12)	(.11)	(.11)
Proportion of observations	.42	.42	.42
Estimated switch point	.34	.34	.34
Log-likelihood	1064.0	1061.4	1061.3

Standard errors are reported in parentheses. For each column, the switch point was selected from a grid search over the range of the ratio of the coverage-adjusted minimum wage to the mean wage (.24-.48), in steps of .01. Estimated intercepts are not reported. See Table 2 for additional details.

Table 7: Estimates of Three-Regime Model, Two Switch Points, Grid Search Results

	(1)	(2)	(3)
Binding regime above second switch point:			
Coverage-adjusted relative minimum wage	-.41	-.41	-.31
	(.10)	(.10)	(.07)
Proportion of population in age group	.07	.07	0.0
	(.11)	(.11)	
Prime-age male unemployment rate	-1.34	-1.34	-1.26
	(.11)	(.11)	(.09)
Proportion of observations	.58	.58	.84
Binding regime below second switch point, above first switch point:			
Coverage-adjusted relative minimum wage	.53	.53	1.14
	(.23)	(.23)	(3.07)
Proportion of population in age group	-.49	-.49	-1.11
	(.13)	(.13)	(.33)
Prime-age male unemployment rate	-.99	-.99	0.0
	(.13)	(.13)	
Proportion of observations	.39	.39	.06
Non-binding regime:			
Coverage-adjusted relative minimum wage	.65	0.0	0.0
	(.88)		
Proportion of population in age group	.16	.14	.01
	(.28)	(.27)	(.21)
Prime-age male unemployment rate	-.55	-.56	-.85
	(.29)	(.29)	(.21)
Proportion of observations	.03	.03	.10
Estimated switch points	.28,.34	.28,.34	.30,.31
Log-likelihood	1068.5	1068.2	1064.0

Standard errors are reported in parentheses. For each column, the switch points were selected from a grid search over the range of the lagged ratio of the coverage-adjusted minimum wage to the mean wage (.24-.48), in steps of .01, with the second switch point constrained to be equal to or greater than the first. Estimated intercepts are not reported. See Table 2 for additional details.