NBER WORKING PAPER SERIES

A NOTE ON THE NEW MINIMUM WAGE RESEARCH

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Working Paper No. 4348

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
April 1993

We are grateful to Nancy Cole for outstanding research assistance. Janet Currie also thanks the NBER for support under the Olin Fellowship program. This paper is part of NBER's research program in Labor Studies. Any opinions expressed are those of the authors and not those of the National Bureau of Economic Research.

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ABSTRACT

Using panel data on individuals from the National Longitudinal Survey of Youth, we find that employed individuals who were affected by the increases in the federal minimum wage in 1979 and 1980 were 3 to 4% less likely to be employed a year later, even after accounting for the fact that workers employed at the minimum wage may differ from their peers in unobserved ways. These results were obtained using a methodology similar in spirit to Card's recent work on the topic, although we use individual rather than state-level data, and an earlier time period.

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The *Industrial and Labor Relations Review* recently published a volume showcasing the "new minimum wage research" which challenges the conventional wisdom regarding the effects of the minimum wage. Time series evidence from the 1970s and early 1980s indicated that a 10% increase in the minimum wage would be associated with a decrease in teenage employment of 1 to 3% (Brown, 1988). Because of the inherent difficulties of drawing inferences from time-series data, much of the new research is based on cross-sections of firms (Katz and Krueger, 1992) or individuals (Card 1992a, 1992b). For example, Card (1992a) exploits the fact that because of regional variation in the wage distributions of teenagers, largely due to variation in state minimum wages, teenagers in different parts of the country had different probabilities of being affected by the increases in the federal minimum which took effect in 1990. He finds that the increase in the federal minimum had no effect on teenage employment.

The volume also presents one paper which exploits state-level panel data. The findings of Neumark and Wascher (1992) are at odds with the rest of the new research in that they find a significant negative effect of increases in the minimum on teenage employment which is similar in size to that found in earlier studies. They suggest that the discrepancy between their results and Card's for example, could be due to the fact that Card's methodology does not allow for lagged minimum wage effects. An alternative possibility suggested by Ehrenberg in his introduction to the volume is that the difference in results is an artifact of the fact that Neumark and Wascher's data span an earlier time period than the other papers: The changes in state minimums which they examine occurred between 1973 and 1989.

Still another possibility is that the reliance on state-level variation in minimum wages confounds the effects of changes in the minimum wage with coincident economic developments which vary in intensity or timing from state to state. For example, Taylor and Kim (1993) use industry and county level data to re-examine the effects of the 1988 increase in California's minimum studied by Card (1992b) and find that "the textbook analysis pertains".

Our contribution to the minimum wage debate is to take the level of disaggregation one step further: We use individual-level panel data from the National Longitudinal Survey of Youth (NLSY). Like Neumark and Wascher, the minimum wage changes we examine span an earlier time period than the Card papers: the increase from $2.90 per hour to $3.10 per hour in January 1980, and from $3.10 to $3.35 one year later. But in an experiment analogous to Card's, we ask whether employed individuals likely to be affected by increases in the federal minimum in these years were less likely than similar individuals to be employed when they were interviewed a year later. We also examine the effects of the minimum on individual year-to-year wage changes. We use the panel aspect of the data to control for possible unobserved differences between the affected workers and other workers, in particular, for the possibility that low-wage workers may also be high-turnover workers for unobserved reasons.

We find that affected workers were 3 to 4% less likely to be employed a year later, even after accounting for unobserved heterogeneity using fixed-effects estimators. We do not find consistent evidence of a positive effect of increases in the minimum on the wages of workers who remained employed a year later. Employment appears to be measured with less error then wages in our data, so the ambiguous findings regarding wages could be due to measurement error, since errors in the data could obscure the correlation between changes in the minimum and changes in wages. Due to data limitations, we cannot say whether differences in the findings of the old and new studies are entirely a function of differences in methodology, or, at least in part, a reflection of real changes in the effects of the minimum wage over time.

Data

The NLSY began in 1979 with 12,686 young people between the ages of 14 and 21. They have been resurveyed every year since. Virtually all respondents were surveyed between February and May of the survey year. At each interview, they are asked whether they are currently employed, and their wage on the current job[1]. Using this information, we can determine whether they were in a job which was affected by the increase in the federal minimum from $2.90 to $3.10 which took affect on January 1980, or by the increase from $3.10 to $3.35 which took affect one year later.

The next year in which the federal minimum wage increased was 1990. The first year in which a significant number of people in our sample would have been affected by state minimums higher than the federal minimum is 1988. Unfortunately, even the youngest people in the NLSY were well out of their teens by this time, which prevents us from examining Card's time period. We do use information for the years after 1981, but only in order to control for individual heterogeneity. Hence, we exclude data collected after 1987 from the sample.

Table 1 describes the extract of data that we work with. The first column shows statistics for all of the cross sections pooled together over the 1979 to 1987 interval. There are 62397 observations with wage data. These pooled observations represent 11607 individuals.

The next two columns show statistics separately for people who had a wage observation in 1979 or 1980. These workers had the potential to be affected by changes in the minimum. The last column describes those that had a wage observation in 1987 for reference purposes.

The first row of the table shows that of the 12,686 people included in the survey, only 30% had valid wage data in 1979. This percentage rose over time: In 1980 it was 46% and by 1987 it had risen to 68%.[2] In what follows we do not attempt to take account of selection into the labor market. We focus only on the effect of changes in the minimum on those who were already employed.

The next row shows that a small fraction of those who were employed in what we will call the "base" year, were missing employment data a year later. On the basis of an examination of the data, we suspect that the majority of these people were not employed. However, a comparison of the third and fourth rows shows that the percent employed a year later is similar whether we delete persons with missing data from the sample, or assume that those with missing data were in fact not employed.

We next show the changes in the minimum between 1979 and 1980 and between 1980 and 1981, as well as the percentage of our respondents who were "bound" by these changes. We consider a person to have been bound by the change if they were working at a wage less than the new minimum but no less than the old minimum in the base year and·if they were not working in the state or local public sectors, in agriculture, or in domestic service. Thus, individuals working in uncovered sectors as well as individuals whose wages were already too high to be directly affected by the increase in the minimum wage serve as a comparison group which controls for changes in the labor market which may have coincided with the increases in the minimum wage. The uncovered workers also act, to some extent, as a control for possible influences of low wages themselves.

Observations with reported hourly wages of less than $1 or greater than $50 were excluded from the sample, as most of these reports appeared to reflect measurement error. By our measure, 20% of the sample were bound by the 1980 increase, and 25% of the sample were bound by the 1981 increase.

Among those who were bound, the average difference between their old hourly wage and the new minimum was $0.15 in 1979 and

$0.18 in 1980. This "wage gap" represents about 4% of the average base wage in each year. The wage gap is meant to measure the extent to which the increase in the minimum affects a worker, and is set to zero for those who were not bound by the new minimum. The gap, which by construction cannot exceed $0.25, may seem small when one considers the fact that most people in this young sample had wages which were growing rapidly from year to year: The average percentage wage increase was 30% between 1979 and 1980 and 27% between 1980 and 1981.

However, Smith and Vavrichek's (1992) contribution to the <u>ILRR</u> minimum wage conference volume suggests that focusing on averages ignores significant heterogeneity among the respondents: They found using the Survey of Income and Program Participation that although most people spent relatively short periods of time working at the minimum wage, a sizeable minority of workers seemed to be "trapped" in minimum wage jobs. These are the workers whom one might expect to be most affected by relatively small changes in the minimum.

The remainder of the table describes characteristics of respondents which have been shown to affect employment and turnover rates, and that we control for in our estimation procedures. The numbers reflect the fact that the NLSY over-sampled poor respondents, African-Americans, and Hispanics.

Estimation and Results

Table 2 shows our estimates of the effect of the minimum wage changes on the probability that individuals employed prior to the change were employed as of the next interview date. In addition to the variables shown in Table 1, we control for possible business cycle effects and for the aging of the sample by including year dummies.

Columns 1 to 3 show Ordinary Least Squares (OLS) estimates of equations of the form

$$E_{i,t} = \alpha Wagegap_{i,t} + X_{i,t,}\beta + \varepsilon_{i,t} \tag{1}$$

where $E_{i,t}$=1 if the individual, who was employed in year t-1, was employed in year t, and =0 otherwise. These are linear probability models. The results are robust to various changes in the definition of the sample, with the coefficient on Wagegap ranging from -0.212 to -0.193. These changes include assuming that those missing employment data a year later were in fact unemployed (column 2), and restricting the sample to those who have at least 4 observations with wage data in the base year and non-missing employment data a year later (column 3). Multiplying the coefficient on the wage gap by the average wage gap of affected persons from Table 1 indicates that the probability of employment for these individuals was reduced by 3 or 4%.

By effectively comparing workers likely to be bound by each increase in the minimum wage with those who are not bound, the OLS estimates control for possible changes in labor market conditions that could affect all workers and that could coincide with increases in the minimum wage. However, the OLS estimates do not take account of the possibility that bound and not bound workers differ from each other in systematic but unobserved ways. In particular, low-wage workers may be more likely to separate from their jobs for the same reasons that they earn low wages, rather than because they are bound by the minimum wage.

This issue can be addressed by exploiting the panel nature of the data, estimating an equation of the form

$$E_{i,t} = \alpha Wagegap_{i,t} + X_{i,t}\beta + \mu_i + u_{i,t} \tag{2}$$

where μ_i represents constant, individual-specific sources of heterogeneity. We estimate this model using only individuals with 4 or more observations to ensure that there are enough observations per person to be able to distinguish the individual-specific component statistically. With this exclusion, and using the years up to 1987, the sample averages 7 annual observations per individual. Column 4 shows fixed effects (FE) estimates of the linear probability model given in (2). The coefficient on Wagegap (-0.184) is similar to the OLS estimate.

An important limitation of the FE estimates are that the important elements of individual heterogeneity are assumed to be constant over time. In particular, it is assumed that the characteristics of an individual which make her likely both to be observed in a low-wage job and to leave that job, remain in force in subsequent years. Alternatively, individual but transient circumstances may be responsible for the fact that someone takes a low-wage job and then, say, leaves in pursuit of better opportunities. In this case, (2) may not adequately control for spurious correlations between the wage gap and the employment history.

An alternative treatment of unobserved heterogeneity is random effects estimation.[3] The random effects estimator has the advantage that it utilizes the "between" individual variation in the data, which the fixed effects estimator ignores, while allowing for correlation between observations of the same individual over time. The disadvantage of a random effects model is that the estimates are inconsistent if the unobserved random variable u_i is correlated with the observed explanatory variables included in the model, which is likely in our case. In any event, the random effects estimate of the coefficient on Wagegap in the linear probability model, shown in column 5 of Table 2, is -0.190 which is very similar to the fixed effects and simple OLS estimates discussed above.

Since the dependent variable in this regression is binary, we would like to estimate a probit or logit model. In such models, however, fixed effects estimators are inconsistent when the number of observations per person is small (Heckman, 1981), and the computational burden is large. An alternative conditional logit model (c.f. Chamberlain, 1982) ignores individuals who do not change employment status at least once over the sample period. This restriction results in a drastic reduction in sample size: All workers who were continuously employed throughout the sample period or who were never employed in two consecutive years (more than twothirds of the sample) are deleted from the sample. It is perhaps unsurprising that when this "control" group is deleted, the coefficients on the variables of interest (not shown) become statistically insignificant. Hence, we confirmed the results discussed above using a random effects probit model (not shown). Unlike the fixed effects model, random effects estimation of a probit model is consistent if the random effects orthogonality assumptions are met.

Because of questions about the reliability of the wage data, we also estimated the employment models using a variable equal to one if the person was bound by the minimum wage increase and zero otherwise instead of the wage gap. We felt that this dichotomous variable might be "cleaner", but this change in specification did not affect our results.

The question of whether the wages of those who remained employed were affected by changes in the minimum wage is addressed in Panel A of Table 3. The first three columns show the results of regressing the change in the log(wage) on Wagegap and the other control variables, although only the coefficient on Wagegap is shown. We do not find any significant effect of the wage gap in the OLS, fixed effect, or random effect models. A less restrictive specification is estimated in columns 4 to 6. Here, the log(wage) is regressed on its one-year lag as well as on the wage gap and other variables. The two specifications are equivalent only if the coefficient on the lagged variable is 1, which does not seem to be the case. The results in columns 3 to 6, show that the wage gap is estimated to have a negative effect on the wages of those that remain employed.

Panel B shows that this counter-intuitive result becomes statistically insignificant if observations with year-to-year wage changes greater than 100% are excluded. We also find that the more restrictive specification in which the change in the log(wage) is regressed on the wage gap yields a significantly positive coefficient on Wage gap in this sample.

These results must be interpreted with caution since, as stated above, we believe that the wage data are subject to more measurement error than the employment data, and the results are not robust to changes in the definition of the sample or to changes in the specification of the estimating equation.

Conclusion

We find that employed individuals who were affected by the new federal minimums in 1979 and 1980 were 3 to 4% less likely to be employed a year later, even after accounting for the fact that workers employed at the minimum wage may differ from their peers in unobservable ways. These results were obtained using a methodology similar in spirit to Card's (1992a), but with individual panel data, which allows us to better control for coincident events which may confound the effects of the minimum wage and for worker heterogeneity. The sample, however, is from an earlier era, more like that studied by Neumark and Wascher (1992). Data limitations prevent us from applying our methodology to the period studied by

Card. Thus we cannot claim to have settled the methodological debate. Whether our results hold for more recent years remains to be seen.

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NOTES

1. These questions are similar to those found in the Current Population Survey.

2. The original 12,686 respondents included 1,280 people enlisted in the armed forces. After 1984, all but 201 of these respondents were dropped from the sample leaving 11,607 people.

3. See Hsiao (1986) for a discussion of random effects models.

Table 1: Means of Key Variables

Base year:	All	1979	1980	1987
# Observations with wage data in base year	62397	3805	4859	7875
# Missing employment in next year	2341	162	86	281
% employed in next year excluding missing	.93	.79	.78	.97
% employed if missing set to zero	.89	.75	.77	.93
change in the minimum base year to next	...	.20	.25	.00
% bound by change minimum	.03	.22	.21	.00
Avg. "wage gap" if bound	.17 (.07)	.15 (.05)	.18 (.08)	...
Avg. wage base year	5.63 (3.34)	3.61 (1.79)	4.13 (2.13)	7.59 (4.46)
Avg. change wage base year to next	.64 (2.90)	.70 (2.04)	.69 (2.31)	.82 (3.81)
Avg. % change in wage base year to next	.24 (.86)	.30 (.73)	.27 (.70)	.23 (.84)
% Poverty sample	.25	.20	.22	.27
% African-American	.22	.18	.19	.24
% Hispanic	.16	.15	.16	.16
% Male	.53	.53	.53	.52
Avg. age in 79	17.92 (2.25)	18.87 (1.72)	18.47 (2.03)	17.62 (2.27)
% High school grad. in 79	.39	.53	.46	.34
# of independent ids	11607	...	...	...

Note: Standard errors in parentheses.

Table 2: Effect of Minimum Wage Changes on Employment Probability

	(1) OLS	(2) OLS	(3) OLS	(4) FE	(5) RE
Sample:	All non-missing only	Impute missing	Persons with >= 4 observations		
Intercept	.949	1.012	.948	...	.791
	(.011)	(.012)	(.010)		(.010)
Wage gap	-.212	-.195	-.193	-.184	-.190
	(.036)	(.043)	(.034)	(.032)	(.033)
Poverty sample	-.030	-.030	-.018	...	-.018
	(.002)	(.003)	(.002)		(.002)
Age in 79	-.001	-.005	.000	...	.000
	(6.113)	(.001)	(.001)		(.001)
High school grad. in 79	.039	.038	.028	...	.028
	(.003)	(.003)	(.003)		(.003)
Male	.041	.025	.027	...	.027
	(.002)	(.002)	(.002)		(.002)
African-American	-.026	-.019	-.016	...	-.016
	(.002)	(.003)	(.002)		(.003)
Hispanic	-.007	-.009	-.002	...	-.002
	(.003)	(.003)	(.003)		(.003)
Year Effects					
1979	-.188	-.181	-.169	-.162	-.167
	(.005)	(.006)	(.005)	(.004)	(.005)
1980	-.189	-.165	-.169	-.178	-.172
	(.005)	(.006)	(.004)	(.004)	(.004)
1981	.007	.025	.010	-.009	.005
	(.005)	(.005)	(.004)	(.004)	(.004)
1982	-.045	-.023	-.028	-.024	-.026
	(.004)	(.005)	(.004)	(.003)	(.004)
1983	-.031	-.020	-.015	-.018	-.016
	(.004)	(.005)	(.004)	(.003)	(.004)
1984	-.010	-.068	.000	-.004	-.001
	(.004)	(.005)	(.004)	(.003)	(.004)
1985	-.008	-.007	.002	.003	.002
	(.004)	(.005)	(.004)	(.003)	(.004)
1986	-.007	-.010	-.003	-.001	-.002
	(.004)	(.005)	(.004)	(.003)	(.004)
R-Squared	.076	.046	.074	.077	.075
# Obs.	62374	64705	57508	57508	57508

Notes: Standard errors in parentheses.

Table 3: Effect of a Change in the Minimum Wage on Log(Wage)
Persons with >= 4 observations only

Panel A: No additional exclusions

Dependent Variable:	Change in log(wage)			Level of log(wage)		
	(1) OLS	(2) FE	(3) RE	(4) OLS	(5) FE	(6) RE
Lag log(wage)	-	-	-	.542 (.004)	.129 (.005)	.644 (.004)
Wage gap	.037 (.077)	.034 (.089)	.040 (.073)	-.251 (.069)	-.173 (.067)	-.249 (.069)
R-squared	.005	.059	.006	.466	.640	.540
# Observations	49864	49864	49864	49864	49864	49864

Panel B: Excluding observations with wage changes > 100%

Dependent Variable:	Change in log(wage)			Level of log(wage)		
	(1) OLS	(2) FE	(3) RE	(4) OLS	(5) FE	(6) RE
Lag log(wage)	-	-	-	.734 (.004)	.383 (.005)	.802 (.004)
Wage gap	.199 (.063)	.213 (.069)	.201 (.061)	-.027 (.060)	-.050 (.060)	-.001 (.060)
R-squared	.006	.163	.007	.594	.719	.652
# Observations	46969	46969	46969	46969	46969	46969

Notes: Standard errors in parentheses.
Variables included but not shown: Intercept, Age in 1979, Male, African-American, Hispanic, Year effects