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THE EFFECTS OF UNEMPLOYMENT INSURANCE ELIGIBILITY RULES
ON JOB QUITTING BEHAVIOR*

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Since the beginning of the U.S. unemployment insurance (UI) system in the late 1930s, every state UI program has disqualified from benefit eligibility those persons that voluntarily left their jobs "without good cause." The rationale is that insurance should not be provided to individuals for contingencies resulting from their own voluntary actions. Nevertheless, at the UI system's inception, every state but one disqualified job quitters for only several weeks, after which they could begin collecting benefits if still unemployed. The underlying concept was that even job quitters should be insured against the risk of unusually long unemployment spells. Furthermore, most states applied no disqualification whatsoever when the individual had quit for "good personal cause."¹

Since the 1930s, however, the states have adopted much more stringent disqualification policies. By January 1983, 44 states disqualified job quitters for the full duration of their unemployment. In addition, 31 states restricted the "good cause" exception from disqualification to reasons attributable to the employer or connected with the work.²

The reduced availability of UI benefits to job quitters might be expected to reduce the frequency of quitting because it increases the expected costs of leaving employment. If denial of benefits does in fact reduce quitting, this would underscore the voluntary aspect of quitters' unemployment and strengthen the argument that such unemployment should not be compensated. Indeed, concern about incentive effects on quitting

appears to have been a factor in the trend toward stricter eligibility rules. According to Adams (1971, p. 72), "state lawmakers have responded to the adverse publicity" generated by popular press articles such as the one that cited benefit payments to quitters as a prime example of how, "in many of our states, the unemployment-compensation laws . . . have been riddled with loopholes which encourage all who want something for nothing."³

This study investigates whether the numerous recent changes in state disqualification rules have actually had any discernible impact on quit rates in manufacturing industries. The basic strategy is to perform a before-and-after comparison of quit rates in states that changed disqualification policy, while using quit rate movements in states that did not change policy to control for time patterns due to other factors.

The Empirical Model

The simplest model used in this study assumes that quit rate behavior can be approximated by the equation

$$Q_{it} = f[\alpha + \beta D_{it} + \gamma R_{it} + \phi_t + \delta_i + \varepsilon_{it}] \quad (1)$$

The dependent variable Q_{it} is the quit rate in state i in year t . The function f will be assumed to be either an identity function, so that Q_{it} equals the bracketed expression, or an exponential function, so that the natural logarithm of Q_{it} equals the bracketed expression.

The parameter α is a constant intercept, and ε_{it} is a random error

term. The dummy variable D_{it} equals 1 if, in year t , state i disqualified job quitters from collecting UI benefits throughout the full duration of their unemployment. If, in year t , state i paid quitters benefits after a temporary disqualification period, D_{it} equals 0. Similarly, R_{it} equals 1 if, in year t , state i restricted good cause to reasons attributable to the employer or connected with the work; otherwise, R_{it} equals 0. The coefficient β measures the effect of full-duration disqualification on the quit rate (or its logarithm), and γ represents the effect of restricting good cause. If more stringent disqualification policies do reduce job quitting, these coefficients should be negative.⁴

Of course, UI policy is hardly the only factor affecting the quit rate. Theoretical analyses of job quitting have postulated that the probability that a worker quits depends on his current wage, the form and parameters of his alternative wage distribution, and his job search costs.⁵ Although current wage data are readily available, data on alternative wage distributions and search costs are not. Because these unobservables, however, are thought to be correlated with observable variables such as indicators of the business cycle, previous empirical work has related the quit rate directly to these observables. Time series studies by Pencavel (1970), Parsons (1973), and Hamermesh (1969), for example, have found that quit rates are strongly procyclical. In addition, cross-section analyses by Pencavel (1970), Parsons (1972), and others have found that quit rates vary from industry to industry because of differences in wage levels, occupational composition, urbanization, unionization, and other observable

factors.

It seems reasonable to expect quit rates to vary similarly among states because of state-to-state differences in these same factors as well as in industrial composition itself. If the effects of these factors are approximately stable over time and if the state-to-state differences in these factors also are stable, the effects of these factors on the i^{th} state's quit rate can be simply represented in equation (1) by the constant state-specific term δ_i . This specification has two important advantages. First, it accounts for the effects of unobservable, as well as observable, state-specific factors. Second, it requires no assumptions about the functional form in which the separate factors would enter equation (1)'s bracketed expression. The assumption of stability in the state-to-state configuration of these factors may seem unduly restrictive, but it will be relaxed later in the paper.

The year-specific term ϕ_t is included in the equation to represent general cyclical and trend effects on the quit rate. The implicit assumption that these time effects are common across states also will be relaxed later in a more general version of equation (1).

Estimation

The dependent variable used in the empirical estimation of equation (1) is the annual average quit rate in manufacturing by state or area. The annual average state and area quit rate data, generated by the U.S.

Department of Labor's labor turnover survey, appeared each year in the Labor Department's Employment and Earnings publication (usually in the May issue) until the labor turnover survey was terminated in 1981. These annual rates were averages of monthly quit rates per 100 employees. This study analyzes quit rates in the period 1970-80. During this time, 22 states changed policy on full versus temporary disqualification for job quitting, and five states changed policy on restricting the good cause exception. These changes in state policy provide an opportunity to investigate whether changes in UI eligibility rules for quitters have a discernible influence on job quitting behavior.

Unfortunately, not every state could be included in the study, first because three states did not participate in the labor turnover survey at all. Many other states did not collect statewide data, but reported quit rates for one or more metropolitan areas. Whenever possible, this study uses quit rates for areas in states where statewide data are unavailable (e.g., Chicago in Illinois, Mobile in Alabama), but some area series are unusable either because the area spills over a state line (e.g., the Memphis data pertain to counties in Arkansas as well as Tennessee) or because the definition of the area changed during the 1970-80 period.

After these exclusions, the sample contains 47 units of observation from 45 states. Of the missing states--California, New Mexico, Tennessee, Texas, and West Virginia--only New Mexico had changed D or R during the sample period. The sample therefore contains 21 states that changed D during 1970-80 and four states that changed R. All four that changed R did

so in 1979, with Maryland and both Dakotas changing to $R=1$ and Montana changing to $R=0$. Of course, the small number of R changes and their concentration in one year will prevent very precise estimation of the effect of restricting the good cause exception.

For D , however, the changes are more numerous and better distributed through the decade. Four states (Colorado, Indiana, North Carolina, and Wyoming) changed to $D=0$ in 1971. Changes to $D=1$ were adopted by Wisconsin in 1971; by Arizona, Indiana, and Vermont in 1974; by Georgia, Hawaii, Massachusetts, and South Carolina in 1976; by Connecticut, Maine, Minnesota, Nevada, North Carolina, Oklahoma, and Washington in 1977; and by Montana, North Dakota, South Dakota, and Utah in 1979. The numerous changes in D and their dispersion across years will enable quite accurate estimation of the impact of full-duration versus temporary disqualification.

Equation (1) is estimated by ordinary least squares regression with the standard fixed-effects technique⁶ (also known as the covariance or dummy-variable technique) for analyzing pooled cross-section and time series data. The estimated effects of UI eligibility rules, on both the quit rate and its logarithm, are shown in the first column of table 1. For brevity, the estimated state and time effects are omitted from the table.

TABLE 1

Estimated Coefficients (and Standard Errors) from Quit Rate Regressions

	<u>Fixed-Effects Model</u>	<u>Model with State-Specific Trends and Cycles</u>
Level:		
D	- .080 (.078)	.067 (.094)
R	- .238 (.242)	.292 (.297)
Log:		
D	- .014 (.027)	.020 (.031)
R	- .110 (.084)	.072 (.093)
Number of observations	505*	495*
Degrees of freedom	400	297
R ² (level)	.883	.976
R ² (log)	.913	.983

*In column 1, the sample size is 47 areas times 11 years minus 12 missing observations on the quit rate. In column 2, an additional 10 observations are missing employment information.

In both the level and log specifications, the estimated coefficient of D, which measures the impact of imposing full-duration disqualifications, is nearly equal to zero. Despite the small standard errors, the differences of the coefficient estimates from zero are not statistically significant at any conventional level. These results imply that quit rates are not noticeably affected by whether states impose full-duration or temporary disqualifications. Perhaps this finding should be unsurprising. Even in temporary-disqualification states, quitters can collect benefits only after they have been unemployed for many weeks--apparently not a powerful inducement to quit work.

The coefficient estimates for R, which measure the impact of restricting the good cause exception from disqualification, are much larger in magnitude, but so are their estimated standard errors. In the level regression, the point estimate of the R coefficient implies that restricting good cause reduces the quit rate by almost a quarter of a percentage point (as compared to a sample mean quit rate of 2.4 percent), but the difference of this estimate from zero is not statistically significant at any conventional level. In the log regression, restricting good cause is estimated to reduce the quit rate proportionally by about 10 percent. By a one-sided t-test, the difference of this estimate from zero is just barely significant at the .10 level.

The standard error estimates and associated hypothesis tests assume that the inclusion of time effects in the regressions has purged the quit rate data of serial correlation. First-order autoregressions of the resi-

duals, however, yield estimated autocorrelations of .57 in the level regression and .56 in the log regression. Moreover, as Nickell (1981) has pointed out, these estimates are downward biased. Nickell's (1980) bias correction method revises these estimates upward, respectively, to .78 and .77.⁷

What causes the large serial correlations? One obvious explanation is that the time effects are not actually common to all states. If the response of quit rates to national business cycles varies among states, then during an expansion, for example, some states' quit rates will be persistently above the fitted regression line and others' persistently below. Furthermore, states may differ in their quit rate trends. In other words, the state-to-state differences in the factors affecting quit rates may not be stable, as assumed thus far, but may change over time.⁸ If so, then a state with an especially strong upward trend, for instance, would tend to show persistently negative residuals in the early years of the sample period and positive residuals in the later years. To allow for state-specific trends and cyclical patterns, equation (1) is modified as follows:

$$Q_{it} = f[\alpha + \beta D_{it} + \gamma R_{it} + \phi_t + \delta_i + \lambda_i(t-1970) + \eta_i(t-1970)^2 + \pi_i E_{it} + \epsilon_{it}] \quad (2)$$

The new equation allows the state effects to change gradually according to a state-specific quadratic time trend. It also adds the variable E_{it} , the i^{th} state's level of employment (or log of employment in the log

regression) in year t .⁹ The inclusion of this variable with state-specific coefficients allows states to vary in both the amplitude and timing of cyclical effects. It should be noted that, with state-specific quadratic time trends already included in the equation, including E_{it} is equivalent to including the deviation of E_{it} from its own state-specific quadratic time trend.

When level and log regressions are fitted for equation (2), first-order autoregressions of the residuals yield estimated autocorrelations of $-.09$ and $-.14$. Again, as in the simple fixed-effects model, these autocorrelation estimates are biased, but even so, their closeness to zero suggests that the true autocorrelation is unlikely to be very large.

The estimated effects of UI policy are shown in the second column of the table. As before, in both the level and log regressions, the coefficients of D are estimated with good precision and found to be close to zero, implying no major impact of full-duration versus temporary disqualification. The estimated coefficients of R , which are no longer negative, also differ from zero by less than one estimated standard error. It is worth reemphasizing, however, that the large standard errors of the coefficient estimates for R preclude any precise statements about the true impact of restricting good cause.

Summary and Discussion

State unemployment insurance programs have differed in two main ways in their rules concerning benefit eligibility of job quitters. One is

whether they disqualify quitters temporarily or for the full duration of their unemployment spells. The other is whether they except from disqualification persons that quit with "good personal cause." In both respects, more and more states have been adopting the stricter eligibility rules. This study has examined whether the reduced availability of unemployment benefits to quitters has been associated with an observable reduction in job quitting.

The study has found no evidence that the choice between full-duration and temporary disqualification of quitters has any discernible impact on quit rates. This apparently negligible impact suggests that state legislators choosing between full-duration and temporary disqualification policies need not be concerned about incentive effects on job quitting. It does not necessarily follow, though, that the trend toward full-duration disqualification is misguided. Even in the absence of incentive effects, legislators may reasonably view quit-initiated unemployment as the consequence of voluntary action and, hence, as an unsuitable target for social insurance. To stretch a frequently used analogy, a finding that fire insurance does not increase the frequency of arson would not constitute a compelling case for public provision of fire insurance to persons that burn down their own houses.

The results concerning the impact of restricting good cause to reasons attributable to the employer or connected with the work have been blurrier. The estimates from a simple fixed-effects model seem to provide some very weak evidence that restricted good cause is associated with lower

quit rates. This result is obliterated, though, when the simple model is modified to allow for state-specific trends and cyclical patterns in quit rates. In any case, because of the small number of states changing good cause definitions during the sample period, it has been impossible to obtain precise estimates of the impact of restricting good cause.

Footnotes

1. See Haber and Murray (1966) for a detailed historical discussion of UI eligibility rules.
2. See U.S. Department of Labor (1983) for tabulations of current UI eligibility rules by state.
3. Selby and Selby (1965). See Gilmore (1960) and Anonymous (1964) for other examples.
4. Although D and R describe the most salient features of states' rules for disqualifying quitters, they overlook some subtleties. For example, states that do pay quitters after a temporary disqualification vary with respect to length of disqualification and to whether the disqualification represents merely a delay in receiving benefits or also a reduction in total benefit entitlement. In a fixed-effects model, effects of variations in these policies can be identified only if individual states change policies over time. In fact, these policies changed very rarely in the sample period and hence cannot be separated from the state-specific fixed effects.
5. See Parsons (1977) for a survey of the relevant literature.
6. See Kmenta (1971), pp. 516-17, for a brief discussion of this method. As Mundlak (1978) points out, the fixed-effects approach, unlike the alternative random-effects approach, yields a consistent estimator even when the variables underlying the individual or time effects are correlated with other explanatory variables.
7. For further discussion of the estimation of serial correlation in fixed-effects models, see Solon (1983).
8. I am indebted to David Ashmore for raising the issue of temporal variation in state effects.
9. The employment variable is nonagricultural employment in the relevant area and year, as measured by the U.S. Department of Labor's establishment survey. The data were obtained from U.S. Department of Labor (1979) and monthly issues of Employment and Earnings.

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