NBER WORKING PAPER SERIES

WORK INCENTIVE EFFECTS
OF TAXING UNEMPLOYMENT BENEFITS

Gary Solon

Working Paper No. 1260

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
January 1984

This study was partially supported by a grant from the Sloan Foundation to the Department of Economics of Princeton University. The author thanks Orley Ashenfelter, David Card, Angus Deaton, Ronald Ehrenberg, Paul Mackin, Richard Quandt, Michael Ransom, anonymous readers, and seminar participants too numerous to mention for their generous advice. The research reported here is part of the NBER's research program in Labor Studies. Any opinions expressed are those of the author and not those of the National Bureau of Economic Research.

NBER Working Paper # 1260
January 1984

WORK INCENTIVE EFFECTS OF TAXING UNEMPLOYMENT BENEFITS

ABSTRACT

Before 1979, unemployment insurance (UI) benefits were not treated as taxable income in the United States. Several economists criticized this policy on the ground that not taxing UI benefits while taxing earned income allegedly encourages unemployed persons to conduct longer than socially optimal job searches. Since 1979, however, UI benefits received by persons in higher-income families have been subject to income tax. This paper investigates whether the introduction of benefit taxation has had the predicted effect of reducing unemployment duration.

The study uses data on a sample of persons that filed for UI in 1978 or 1979 to examine whether high-income claimants collected benefits for shorter periods after the tax change than they did before benefits became taxable. As part of the empirical analysis, the paper develops a generalization of the Weibull distribution and applies a limited-dependent-variable technique for this distribution similar to the Tobit technique for the normal distribution. Despite some variation in the results from different model specifications, the analysis presents persuasive evidence of a tax effect on unemployment duration. The 1979 policy change is estimated to have reduced average compensated unemployment duration among the sampled high-income claimants by about one week.

Gary Solon
Department of Economics
The University of Michigan
Ann Arbor, MI 48109
(313) 763-1306

WORK INCENTIVE EFFECTS OF TAXING UNEMPLOYMENT BENEFITS

1. Introduction

Before 1979, unemployment insurance (UI) benefits were not treated as taxable income in the United States. Several economists[1] criticized this policy on the ground that not taxing unemployment benefits while taxing earned income produces perverse economic incentives, one of which is allegedly to encourage unemployed persons to conduct longer than socially optimal job searches. Perhaps as a result of this criticism, UI benefits received by persons in higher-income families were subjected to income tax in 1979. Specifically, benefits became taxable on joint tax returns reporting at least $25,000 of adjusted gross income (counting UI benefits) and on single returns reporting at least $20,000. In 1982, these income thresholds were lowered respectively to $18,000 and $12,000. A recent proposal within the Reagan administration to extend benefit taxation still further was motivated by the policy's supposed work incentive effects.[2]

This paper presents an empirical analysis of the work incentive effects of the 1979 policy change. It uses data on a sample of persons that filed for UI in 1978 or 1979 to examine whether claimants collected benefits for shorter periods after the tax change than they did before benefits became taxable. Section 2 of the paper briefly reviews previous theoretical and empirical work on related issues. Section 3 describes the study's data base. Section 4 presents analyses of the data, and Section 5 summarizes and discusses the results.

2. Previous Research

Theoretical analyses of UI and unemployment duration typically have examined the duration effects of changes in weekly benefit level, not changes in benefit taxation. But, if UI recipients do not suffer from "tax illusion," they should respond to benefit taxation as a reduction in their net benefit level, so that the same results apply. Most of the theoretical work has analyzed UI and unemployment duration in the framework of job search theory. Mortensen (1970) provides a representative example of this approach, and Lippman and McCall (1976) survey the search theory literature. Despite variations among models, several general conclusions emerge. One is that unemployment duration of UI recipients is negatively related to cost of unemployment and hence positively related to benefit level. By similar reasoning, individuals that assign a large value to the leisure component of unemployment will impute a smaller opportunity cost to unemployment and will tend to stay unemployed longer. Another conclusion is that expected unemployment duration depends in complicated ways on the individual's wage offer distribution. On one hand, at a given benefit level, a higher-wage worker faces a greater opportunity cost of unemployment so that he might return to work more quickly. Higher-skill workers may have shorter unemployment duration also because they are qualified for a larger proportion of job openings. On the other hand, higher-skill workers may face wage offer distributions shaped in such a way that they set reservation wages high enough to give them longer expected unemployment duration.

Labor supply theory, as well as job search theory, can be used to generate similar conclusions.[3] Indeed, almost any applicable economic theory should reproduce the first conclusion -- that paying people more to be unemployed tends to increase how much they are unemployed.

Drawing from the theoretical conclusions above, numerous empirical studies have investigated the dependence of unemployment duration on UI benefits, variables associated with leisure-income preference, and variables associated with wage offer distributions. Disentangling UI effects from wage effects is especially difficult because each state UI program in the United States computes individuals' benefit levels on the basis of their prior earnings. Researchers have adopted two strategies for obtaining independent variation in UI benefits and wages. One approach, exemplified by Ehrenberg and Oaxaca (1976), is to use a sample of UI claimants from several states with different benefit formulas. The sample then contains individuals with similar earnings histories but different UI entitlements. As Ehrenberg and Oaxaca acknowledge, however, there remains a possibility of simultaneity bias. It is unclear whether the observed positive correlation between duration and benefit level arises because states with more liberal benefits induce their claimants to stay unemployed longer or because states whose unemployed experience longer unemployment spells adopt more liberal benefit formulas.

The other approach, exemplified by Classen (1979), is to analyze within-state data spanning a period when the state's benefit schedule has been substantially changed. Using data from Pennsylvania and Arizona, Classen finds that benefit increases in both states were accompanied by significant increases in unemployment duration. As Welch (1977) discusses in detail, both the Ehrenberg-Oaxaca and Classen studies, like most others, are subject to censorship or truncation biases. More recent analyses of British data by Lancaster (1979) and Nickell (1979) have developed more appropriate econometric methods to treat these problems, but because Britain has a uniform national UI program, these analyses require strong identifying restrictions to disentangle UI effects from the effects of wages and other characteristics.

The present study is similar in spirit to the Classen study in that it analyzes unemployment duration during a period containing a major policy change. This study, however, applies more appropriate econometric techniques and is the first to present direct evidence on the duration effects of taxing unemployment benefits.

3. Data Description

This study analyzes data on several thousand persons that filed valid UI claims in Georgia in 1978 or 1979. Because benefit taxation was initiated in 1979, these data afford the opportunity to compare the unemployment duration of claimants before benefits were taxed with the duration of those that claimed benefits after the tax change. The data were collected as part of the Continuous Wage and Benefit History (CWBH) program, a pilot effort by the U.S. Department of Labor and state employment security agencies to develop data banks on samples of workers covered by the UI program.[4] The CWBH files combine administrative data from the sampled individuals' claims records with questionnaire data on their personal characteristics. The administrative information includes data on claimants' prior earnings, benefit entitlements, and how long they collected benefits. The questionnaire information includes, among other things, income data that enable imputation of which claimants had high enough income to be subject to benefit taxation. (Fourteen percent of the 1979 claimants in the sample were above the relevant income thresholds.) Only Georgia's CWBH data were used because Georgia is the only state with extensive questionnaire data from as early as the beginning of 1978.

This study's sample includes claimants that initiated valid claims between January and June 1978 or January and June 1979. Persons that initiated claims in July-December 1978 are excluded from the study's sample because of the likelihood that they collected part of their benefit entitlement in 1979, in which case that part might have been subject to income tax. The 1978 sample is therefore restricted to early-in-the-year filers to achieve a cleaner separation between the pre-tax and post-tax groups.[5]

A description of some of the features of Georgia's UI program in 1978-79 will clarify the empirical work below. A claimant's benefit entitlement depended on his earnings in the "base period," the first four of the last five completed calendar quarters prior to his filing the claim. His weekly benefit amount (WBA) was set at 1/25 of his highest-quarter earnings in the base period, except that the minimum WBA was \$27 and the maximum was \$90. Forty-five percent of the sample claimants (and 66% of the taxable group[6]) qualified for the maximum WBA. A claimant's total entitlement during his "benefit year," the 52-week period following his initial claim-filing, was the lesser of 1/4 of his base period earnings or 26 times his WBA. Consequently, about a quarter of the claimants qualified for the maximum 26 weeks of potential benefit duration, but most were entitled to fewer weeks.

Although Georgia's weekly benefit schedule was nominally unchanged during the sample period, the high inflation rates of the period meant that the schedule changed substantially in real terms. For example, a January 1978 claimant with high-quarter earnings of \$2500 received the maximum \$90 WBA. A June 1979 claimant with the same real prior earnings had nominal high-quarter earnings of over \$2700. He too received a nominal WBA of \$90, which by then was worth less than \$80 in January 1978 dollars. Thus, compared to his January 1978 twin, the June 1979 claimant experienced well over a 10% reduction in real benefits. This change in the real benefit schedule facilitates the separation of UI's effects on unemployment duration from the effects of wage levels.

This study uses data on only those sample claimants that responded to the CWBH questionnaire. The nonresponse rate of about 2/3 raises the issue of nonresponse bias. By far the main cause of nonresponse was Georgia's system of employer-filed claims, under which an employer temporarily laying off part or all of its work force could submit a packet of UI claims for all its laid-off employees.[7] Because the employees themselves did not appear at a claims office, they had no opportunity to fill out the questionnaire. As a result, this study's sample consists mainly of persons permanently separated from their former employers. This exclusion of employer-filed claims may actually be desirable. Feldstein (1978) has argued that studies of UI's effect on unemployment duration should exclude persons on temporary layoff to avoid confounding UI's duration effects with its effects on frequency of temporary layoffs.

4. Data Analysis

The 1979 institution of benefit taxation applied only to claimants with family income above the thresholds described in the introduction of this paper. The basic empirical strategy of this study is to compare the unemployment duration of high-income claimants before and after the tax change, using duration data on low-income claimants (for whom there was no policy change) to adjust for 1978-79 duration trends not attributable to the tax change. It is conceivable that benefit taxation had no work incentive effect, especially since taxes were not withheld from the benefit checks. If it did not, high-income claimants in 1979 should show no relative reduction in unemployment duration. On the other hand, claimants were formally notified of the tax change and may have responded to the resulting reduction in net benefit levels by altering their job-seeking behavior. If so, high-income claimants in 1979 should show a duration reduction not attributable to other factors.

The duration measure used throughout the analysis is the number of weeks that the claimant collected UI during his benefit year. It should be understood that this is not a pure measure of duration per spell because many claimants collect benefits in more than one spell during the benefit year.[8] The duration measure used therefore does not accord perfectly with job search theory. Nevertheless, the effect of benefit taxation on total weeks unemployed, rather than weeks per spell, is probably of greater policy interest. It should also be noted that number of weeks of benefit collection is a censored duration measure. For the 24% of the sample claimants that used up their entire benefit entitlement, weeks collected measures only their compensated duration and not the weeks they were unemployed after benefit exhaustion. This censorship issue is treated in detail later in the paper.

The results from the more elaborate models presented below can be previewed by a simple comparison of means. Among the low-income claimants in the sample, mean compensated unemployment duration was 8.7 weeks for both the 1978 and 1979 filers, implying no general decline in duration between the two years. Among the high-income claimants, however, mean duration fell from 10.8 weeks in 1978, when their benefits were not taxable, to 8.4 weeks in 1979, when their benefits were taxable. The large duration reduction among high-income claimants suggests the possibility that the introduction of benefit taxation did indeed affect unemployment duration.

Regression Analyses

This section presents the results of regression analyses relating unemployment duration to pre- and post-tax benefit levels. These analyses produce estimated duration effects of WBA that can be compared to the results of earlier studies, and directly test whether claimants respond only to pre-tax benefit levels or whether they react also to tax-induced reductions in net benefits.[9]

The basic behavioral equation is posited to take the form

$$\text{DURATION} = f\{\beta(1 - \rho t)\text{WBA} + \gamma'X + \varepsilon\};$$

that is, duration is functionally dependent on the bracketed linear function, in which X is a vector of control variables and ε is a random error term. The variable t is the tax rate on UI benefits so that $t>0$ for high-income claimants in 1979 and $t=0$ otherwise. The parameter ρ is a coefficient of tax perception such that $\rho=0$ if claimants behave as if they are unaware of the tax and $\rho=1$ if they respond to benefit taxation fully as they do to other benefit reductions. The function f will be assumed to be either an identity function, so that DURATION equals the bracketed expression, or an exponential function, so that the natural logarithm of DURATION equals the bracketed expression.

Now let t equal a constant $\bar{t}$ for those claimants whose benefits are taxable, and let $t=0$ for the nontaxable claimants.[10] This dichotomous treatment of t is admittedly a crude approximation, but it does capture the salient aspect of variation in benefit taxation -- most claimants' benefits are not taxable at all, but the high-income 1979 claimants' benefits are subject to positive (and typically high) marginal tax rates. If we also let the dummy variable D equal 1 for taxable claimants and 0 for nontaxable claimants, then

$$\text{DURATION} = f\{\beta(1 - \rho\bar{t}D)\text{WBA} + \gamma'X + \varepsilon\}$$

$$= f\{\beta(\text{WBA}) - \beta\rho\bar{t}(D\cdot\text{WBA}) + \gamma'X + \varepsilon\} \ .$$

This last expression allows WBA and D·WBA to be entered as separate regressors in the duration equation. The coefficient of WBA, β, measures WBA's duration effect for claimants whose benefits are not taxable. The estimate of this coefficient is comparable to the estimated WBA coefficients in earlier studies of UI and duration. The coefficient of D·WBA, $-\beta\rho\bar{t}$, measures how much the duration effect is reduced when, because of benefit taxation, the claimant cannot keep all of his gross benefits. If benefit taxation has no effect, then $\rho=0$ (or $\beta=0$ if UI benefits have no effect at all). In this case, the coefficient of D·WBA should be zero. But if taxes do affect duration, then $\rho>0$ and $\beta>0$, in which case the coefficient of D·WBA should be negative. Moreover, the negative of the coefficient of D·WBA divided by the coefficient of WBA gives an implied value of $\rho\bar{t}$. Combined with extraneous information on $\bar{t}$, an estimate of $\rho\bar{t}$ provides information on the value of ρ. Tabulations by Daniel Feenberg from NBER's 1979 tax files suggest that the typical marginal tax rate (including the Georgia state income tax) on benefits received by high-income claimants might be slightly above .3. Dividing this value into the estimate of $\rho\bar{t}$ yields a rough estimate of ρ. Alternatively, if one wishes to begin with the hypothesis that $\rho=1$, the negative of the ratio of estimated coefficients gives an estimate of $\bar{t}$, which can be compared with the extraneous information on $\bar{t}$. If the comparison is close, one might accept the hypothesis $\rho=1$.

The first column of Table 1 presents the results of a regression of compensated unemployment duration against WBA, D·WBA, and a set of control variables. The variables WBA and D·WBA were converted to October 1980 dollars with the Atlanta Consumer Price Index. The control variables, similar to those used in other studies of UI and duration, were chosen because of their possible relationship with claimants' cost of unemployment and/or distribution of employment opportunities. The variables include potential benefit duration, high-quarter earnings (also converted to October 1980 dollars), the ratio of base-period to high-quarter earnings (a measure of previous employment stability), years of education, age, the average total unemployment rate in Georgia during the claimant's benefit year,[11] and dummy variables for year and month of filing, sex, race, occupation, marital status, expectation of recall to former employer, and whether family income was above the 1979 threshold for benefit taxation. Because of the importance of separating UI effects from nonlinear wage effects,[12] the high-quarter earnings variable is supplemented by a squared term, a term interacted with the high-income dummy, and a term interacted with the 1979 dummy. In addition, the marital and spouse-working dummies are interacted with the female dummy. Squared terms for education and age also were tested, but their estimated coefficients were not statistically significant and their inclusion had almost no effect on the results.

TABLE 1

Estimated Parameters (and Standard Errors) from Models of Unemployment Duration

Explanatory Variables	Level Regression	Log Regression	Weibull Model	Generalized Weibull Model
Constant	-6.92	-.180	.359	.240
	(4.75)	(.702)	(.668)	(.672)
WBA	.045	.0075	.0071	.0053
	(.008)	(.0012)	(.0012)	(.0014)
D·WBA	-.018	-.0025	-.0016	-.0003
	(.007)	(.0011)	(.0010)	(.0013)
High income	-.53	-.034	-.142	-.142
	(.74)	(.109)	(.102)	(.105)
Potential benefit	.42	.041	.030	.026
duration	(.06)	(.009)	(.008)	(.009)
Female	.34	.082	.052	.050
	(.34)	(.050)	(.048)	(.050)
Black or Hispanic	.82	.099	.125	.124
	(.23)	(.034)	(.033)	(.035)
Occupation:				
Professional, tech.,	1.87	.284	.250	.251
managerial	(.44)	(.065)	(.063)	(.066)
Clerical, sales	.73	.123	.120	.120
	(.37)	(.055)	(.052)	(.054)
Service	---	---	---	---
Agric., fishery,	1.33	.148	.269	.267
forestry, related	(1.05)	(.155)	(.154)	(.166)
Processing	-.35	.062	-.057	-.057
	(.56)	(.082)	(.076)	(.080)
Machine trades	.26	.094	.039	.038
	(.43)	(.064)	(.060)	(.062)
Benchwork	-.02	.085	-.0002	-.0009
	(.47)	(.069)	(.065)	(.068)
Structural work	1.10	.273	.176	.174
	(.42)	(.062)	(.059)	(.061)
Miscellaneous	-.66	-.064	-.078	-.079
	(.42)	(.061)	(.057)	(.059)
High-quarter earnings	-.59	-.093	-.088	-.086
(HQE, in thousands)	(.27)	(.040)	(.036)	(.037)
HQE squared	.009	.0029	.0017	.0015
	(.014)	(.0021)	(.0018)	(.0019)
HQE × high income	.39	.043	.050	.051
	(.13)	(.019)	(.018)	(.018)
Base-period earnings/	-.73	-.138	-.118	-.118
HQE	(.36)	(.053)	(.049)	(.051)

Married	-.73	-.049	-.120	-.123
	(.35)	(.051)	(.048)	(.049)
Married × female	-1.09	-.230	-.155	-.153
	(.77)	(.113)	(.106)	(.107)
Spouse working	-.24	-.022	-.020	-.020
	(.36)	(.054)	(.050)	(.051)
Spouse working × female	2.19	.250	.346	.344
	(.77)	(.114)	(.106)	(.109)
Education	.08	.014	.012	.012
	(.05)	(.007)	(.007)	(.007)
Age	.08	.010	.014	.014
	(.009)	(.0014)	(.0014)	(.0014)
Expecting recall	-.25	.108	-.037	-.038
	(.24)	(.035)	(.033)	(.035)
1979 dummy	1.22	.215	.218	.217
	(.44)	(.064)	(.061)	(.061)
1979 × HQE	-.32	-.061	-.045	-.045
	(.12)	(.017)	(.016)	(.016)
Month of filing:				
January	---	---	---	---
February	-.03	-.044	.022	.022
	(.41)	(.060)	(.058)	(.061)
March	-.50	-.127	-.039	-.038
	(.39)	(.057)	(.054)	(.057)
April	-.62	-.128	-.063	-.063
	(.38)	(.056)	(.053)	(.055)
May	-1.08	-.167	-.131	-.131
	(.40)	(.059)	(.056)	(.058)
June	-.49	-.102	-.043	-.043
	(.40)	(.060)	(.056)	(.059)
Unemployment rate	.62	.085	.061	.062
	(.86)	(.127)	(.120)	(.121)
α			.799	.719
			(.010)	(.037)
γ_1				.00100
				(.00042)
γ_2				-.00067
				(.00033)
Log likelihood × 10^{-4}			-1.733748	-1.733309
R^2	.08	.05		

Number of Observations = 6,610

The coefficient of WBA is estimated as .045 and its difference from zero is statistically significant at any conventional level. The estimate implies that, on average, a \$10 increase in benefit level (in October 1980 dollars) increases an untaxed claimant's compensated duration by almost half a week, a result consistent with the findings of previous studies. The coefficient of D·WBA is estimated as -.018, and its difference from zero also is decidedly significant. Hence, the null hypothesis of no tax effect is rejected. If $\rho=1$, the negative of the ratio of D·WBA's coefficient to WBA's coefficient should yield a plausible value for $\bar{t}$. This ratio turns out be be .40 (with an estimated standard error[13] of .18), reasonably close to the expected $\bar{t}$.

Table 2 presents the coefficient estimates of main interest from some variations on the duration regression. The first column reproduces the WBA and D·WBA coefficients from the regression described above. One question that can be raised about that regression is whether the sizable D·WBA coefficient is due not to a tax effect, but to variation by income in the duration impact of WBA. The regression reported in column 2 addresses this question by including an interaction of the WBA variable with the high-income dummy. This interaction variable turns out to have an estimated coefficient close to zero, and its inclusion has virtually no effect on the estimated WBA and D·WBA coefficients or their standard errors. Another question is how the results would be affected by the separate inclusion of the D variable _not_ interacted with WBA. If this variable showed a significant negative coefficient, and particularly if its inclusion caused the D·WBA coefficient to disappear, one would then suspect that the shorter unemployment duration among high-income 1979 claimants was due not to the tax-induced reduction in their benefits, but to some other factor. As shown in column 3, when D is entered separately, its coefficient is positive and small in magnitude, and the estimated D·WBA coefficient stays about the same. Clearly, the high collinearity between D and D·WBA inflates the standard errors of both coefficients and makes precise estimation impossible, but it is at least somewhat reassuring that the magnitude of the D·WBA coefficient estimate does not decline and that the D variable does not separately explain the duration reduction for the taxable claimants.

TABLE 2

Estimated Coefficients (and Standard Errors) of Key Variables in Duration Regressions

WBA	.045 (.008)	.046 (.009)	.045 (.008)
D·WBA	-.018 (.007)	-.018 (.007)	-.020 (.023)
WBA X high income	---	-.007 (.019)	---
D	---	---	.192 (2.184)

The regression reported in the second column of Table 1 duplicates the one in column 1 except that the new dependent variable is log (compensated duration + 1). Duration is incremented by 1 to avoid taking the logarithm of zero. This is necessary because many sample members initiated valid claims but collected no benefits, presumably because they returned to work before completing a full week of unemployment.[14] While adding 1 to weeks of compensated duration is motivated primarily by computational convenience, it also makes sense as a procedure for rounding fractional weeks of unemployment up to the next integer.

The coefficients of the logarithmic regressions can be interpreted as the approximate proportional changes in duration associated with unit changes in the regressors. The WBA coefficient is estimated as .0075. Evaluated at the sample mean duration of 10 weeks (after adding 1 to compensated duration), this estimate implies that a $10 increase in WBA is associated with a duration increase of about three-quarters of a week, toward the upper end of the range from previous studies. The estimated D·WBA coefficient of -.0025 is significantly different from zero at most conventional levels, rejecting the hypothesis of no tax effect. If $\rho=1$, the ratio of the coefficient estimates implies a very plausible tax rate $\bar{t}$ of .33 (with standard error .16).

In summary, the results of the duration regressions vary somewhat with choices of functional form and explanatory variables. But the results consistently reject the hypothesis of no tax effect, and the relative magnitudes of the WBA and D·WBA coefficients are roughly consistent with the assumption that claimants respond to benefit taxation much as they do to other benefit reductions.

Treatment of the Censorship Problem

As mentioned before, weeks of regular benefit collection is a censored measure of unemployment duration for any claimant that continued to be unemployed after exhausting his benefit entitlement. As Welch (1977) has pointed out, this problem is likely to cause regression results to understate the duration impact of benefit changes. Classen (1979) and Newton and Rosen (1979) have used Tobit analysis to deal with the censorship problem, but the Tobit technique assumes that unemployment duration is normally distributed. This assumption could hardly be further from the truth. A frequency plot of compensated duration for a homogeneous subsample of the Georgia claimants shows not a bell-shaped curve, but a modal frequency for zero weeks and progressively smaller frequencies for longer duration (until a spike appears at the censorship point).

At first glance, the frequency plot suggests that the duration data might be fitted by an exponential distribution. But the exponential distribution implies that a claimant's reemployment probability remains constant over the course of his unemployment spell, and there are several reasons to question this restriction. On one hand, numerous variants of the job search model -- incorporating finite lifetime, risk aversion, capital constraints, or finite potential benefit duration -- predict declining reservation wages and hence rising reemployment probabilities during an unemployment spell. On the other hand, potential employers may perceive lengthy unemployment as a signal of low productivity, and in some cases workers' skills may actually atrophy with prolonged unemployment. The resulting deterioration in the individual's employment opportunities could conceivably cause his probability of reemployment to decline with unemployment duration.

To allow for duration dependence in reemployment probabilities, Lancaster (1979) has proposed the use of the Weibull distribution, of which the exponential is a special case. A convenient formulation of the Weibull distribution implies a reemployment hazard (or exit-from-unemployment rate) function of the form

$$h(y) = \alpha y^{\alpha-1}\exp(-\beta'X) \qquad (1)$$

where y is the number of weeks already unemployed, X is a vector of variables (including WBA and D·WBA) that may affect duration, α is a parameter greater than zero, and β is a vector of parameters associated with X. The elasticity of this hazard rate with respect to y is $\alpha-1$. If $\alpha=1$, the Weibull degenerates to the special case of the exponential. If $\alpha>1$, the reemployment hazard rises with duration; if $\alpha<1$, it declines.[15]

The probability density function of completed duration Y is then

$$f(Y) = \alpha Y^{\alpha-1}\exp\{-\beta'X - Y^{\alpha}\exp(-\beta'X)\} \qquad (2)$$

and expected duration is

$$E(Y) = \exp(\beta'X/\alpha)\,\Gamma\{(\alpha+1)/\alpha\}$$

where Γ is the gamma function.[16] If we differentiate the natural logarithm of expected duration with respect to x_h, the h-th variable in X, we obtain

$$\partial \log E(Y)/\partial x_h = \beta_h/\alpha \quad . \qquad (3)$$

Thus, estimates of β and α can be used to estimate the proportional changes in expected duration associated with unit changes in explanatory variables.

The Georgia data do not permit complete observation of duration. Instead, we observe

$$\begin{aligned} Y^* &= Y \text{ if } Y < P + 1 \\ &= P + 1 \text{ if } Y > P + 1 \end{aligned}$$

where Y* is compensated duration plus 1, as in the logarithmic regression, and P is potential benefit duration. To deal with this "right censorship," we now derive a maximum likelihood estimation technique for the Weibull distribution analagous to the Tobit technique for the normal distribution.

If the i-th claimant's compensated duration $Y_i^* - 1$ is less than his potential benefit duration P_i, his contribution to the likelihood function is simply $f(Y_i^*)$, as in equation (2). But if $Y_i^* - 1 = P_i$, his contribution is

$$\begin{aligned} \text{Prob } (Y_i > P_i + 1) &= \int_{P_i+1}^{\infty} f(Y)\, dY \\ &= \exp\{-Y_i^{*\alpha} \exp(-\beta'X_i)\} . \end{aligned}$$

Hence, the likelihood function for the full sample is

$$L = \prod_1 \alpha Y_i^{*\alpha-1} \exp\{-\beta'X_i - Y_i^{*\alpha} \exp(-\beta'X_i)\} \prod_2 \exp\{-Y_i^{*\alpha} \exp(-\beta'X_i)\}$$

where $\prod_1$ denotes a product taken over the claimants that did not exhaust their benefits and $\prod_2$ denotes a product over those that did.[17] It follows that

$$\log L = n_1 \log \alpha + (\alpha - 1) \sum_1 \log Y_i^* - \sum_1 \{\beta' X_i + Y_i^{*\alpha} \exp(-\beta' X_i)\}$$

$$- \sum_2 Y_i^{*\alpha} \exp(-\beta' X_i)$$

where n_1 is the number of "nonexhaustees," $\sum_1$ denotes a sum over non-exhaustees, and $\sum_2$ is a sum over exhaustees.

The parameters α and β can be estimated by maximizing the log likelihood function with respect to the parameters. This procedure was applied to the Georgia data with the same explanatory variables that were used in the Table 1 regressions.[18]

The results are reported in the third column of Table 1. The estimated value of .8 for the parameter α is significantly less than 1 and implies that the reemployment hazard declines with duration. As Heckman and Borjas (1980) and Lancaster (1979) have observed, however, it is unclear how to interpret this finding. While it may be due to true duration dependence, it may also be explained by unobserved heterogeneity in the sample. If some claimants, because of unobserved factors, have lower reemployment probabilities than other seemingly identifical claimants, they will tend to stay unemployed longer. Then, even if individuals' reemployment hazards are constant over time, the data will display spurious duration dependence — among seemingly identical claimants, those unemployed longer will have lower reemployment probabilities.

The estimated WBA coefficient of .0071 is significantly different from zero at any conventional level. As was shown in equation (3), the coefficient estimate must be divided by the estimate of α to obtain the proportional change in expected duration associated with a unit change in untaxed WBA. The result implies a proportional change of .0089.

The estimated D·WBA coefficient of -.0016 also is significantly different from zero at the .10 level, but not quite at the .05 level. Dividing the coefficient by the estimate of α indicates that the proportional duration effect of a dollar change in WBA is reduced by .0019 if benefits are taxed. If $\rho=1$, the ratio of the D·WBA and WBA coefficients implies a tax rate of .22 (with standard error .15), somewhat (but insignificantly) less than the expected rate. Dividing the ratio by an assumed $\bar{t}$ of slightly above .3 would give a point estimate of about .7 for ρ.

While the Weibull framework provides a convenient and easily interpreted model that allows for duration dependence, it is still somewhat restrictive. In particular, as a referee has noted, equation (1) implies that the proportional effects of benefit variables on the hazard rate stay constant throughout an unemployment spell. To allow for the possibility that these effects diminish as an individual draws closer to exhausting his benefit entitlement, the hazard function can be conveniently generalized to the form

$$h(y) = \alpha y^{\alpha-1} \exp\{-\beta' X - [\log(P+1) - \log y][\gamma_1(WBA) + \gamma_2(D \cdot WBA)]\} \quad \text{for } y < P+1$$

$$= \alpha y^{\alpha-1} \exp(-\beta' X) \quad \text{for } y > P+1 \quad .$$

If $\gamma_1 = \gamma_2 = 0$, this hazard function specializes to the Weibull case. Otherwise, the effects of benefit variables on reemployment probability vary with the time remaining until benefit exhaustion.

As shown in the appendix, this generalization of the Weibull distribution leads to the new log likelihood function

$$\log L = \sum_1 \{\log \alpha + (\alpha + G_i - 1) \log Y_i^* - G_i \log (P_i + 1) - \beta' X_i$$

$$- [\alpha/(\alpha + G_i)] \; Y_i^{*\,\alpha+G_i} \; (P_i + 1)^{-G_i} \exp(-\beta' X_i)\}$$

$$- \sum_2 [\alpha/(\alpha+G_i)] \; (P_i + 1)^\alpha \exp(-\beta' X_i)$$

where $G = \gamma_1(WBA) + \gamma_2(D \cdot WBA)$. The results of maximizing this function with respect to its parameters are shown in the last column of Table 1. A likelihood ratio test of the Weibull model versus the generalized Weibull model rejects the Weibull model at the .05 level.

Unfortunately, the additional complexity of the generalized model makes it more difficult to interpret. Nevertheless, three important observations can be made. First, as expected, the estimates of the β and γ coefficients for WBA and D·WBA imply that net benefit level has a negative effect on reemployment probability and that the magnitude of this effect declines as benefits are used up. Second, to test the hypothesis of no tax effect, the generalized model was reestimated with γ_2 and the β coefficient for D·WBA constrained to equal zero. The resulting log likelihood value was -1.733641×10^{-4}, so that a likelihood ratio test rejects the hypothesis of no tax effect at the .05 level.

Third, to clarify the magnitude of the estimated tax effect, a policy simulation was conducted. As shown in the appendix, the expected value of compensated unemployment duration in the generalized model is

$$E(Y^*) = \{[(\alpha+G)/\alpha](P+1)^G \exp(\beta'X)\}^{1/(\alpha+G)} \Gamma_{[\alpha/(\alpha+G)](P+1)^\alpha \exp(-\beta'X)}\{(\alpha+G+1)/(\alpha+G)\}$$

$$+ \exp\{-[\alpha/(\alpha+G)](P+1)^\alpha \exp(-\beta'X)\}(P+1) \qquad (4)$$

where the subscripted Γ term is an incomplete gamma function.[19] This expectation can be estimated for each member of the sample by substituting in his observed X values, his potential benefit duration P, and the maximum likelihood estimates of the parameters α, β, and γ.

The effect of benefit taxation on compensated duration can be estimated by first computing the sample mean of the estimates of E(Y*) among the high-income 1979 claimants. Then, to estimate what their average duration would have been in the absence of benefit taxation, we set D·WBA=0 and recompute the estimates of E(Y*). A comparison of the sample means with and without benefit taxation yields an estimate of the policy's mean impact.

The mean of the estimates of E(Y*) with benefits taxed is 9.6 weeks, reasonably close to 1 plus the mean compensated duration of 8.4 weeks actually observed for the high-income 1979 claimants. The mean of the estimates of E(Y*) without benefits taxed is 10.8 weeks. The implied average effect of benefit taxation on the high-income 1979 claimants is therefore a 1.2 week reduction in their compensated duration.[20]

5. Summary and Discussion

This paper has presented a series of analyses of the effect of taxing unemployment benefits on unemployment duration. Despite some variation in the results from different model specifications, the analyses have repeatedly found that unemployment benefit levels do affect unemployment duration, and have produced persuasive evidence that the duration impact of taxing benefits is similar to that of other benefit reductions. The 1979 imposition of benefit taxation is estimated to have reduced average compensated unemployment duration among the sampled high-income claimants by about one week.

This finding implies that the budgetary effects of benefit taxation extend beyond the direct revenue increases from taxes collected on benefit income. One additional effect is the tax revenue collected from the increased earnings of claimants who return to work more quickly when their benefits are taxed. If we simplify by letting $\bar{t}$ be a constant tax rate and let W be the claimant's weekly wage when working, DUR_0 be his unemployment duration without benefit taxation, and ΔDUR be his duration change due to benefit taxation, then the full change in his tax payment induced by benefit taxation is

$$\Delta T = \bar{t}\{DUR_0 \cdot WBA - \Delta DUR \cdot (W - WBA)\} \quad .$$

The first term in brackets is the benefit income the claimant would have collected in the <u>absence</u> of benefit taxation. The second term is his additional gross income induced by benefit taxation.

This expression makes clear that, if the budgetary impact of some proposed benefit taxation is to be forecasted, a projection based only on the affected claimants' benefit income before the tax would underestimate the total impact by overlooking the second term. For example, the results for the Georgia sample imply that, among the 1979 high-income claimants, $DUR_0 \cdot WBA$ averaged \$1030 while the second term, $-\Delta DUR \cdot (W - WBA)$, averaged \$337.[21] Therefore, a projection that neglected the work incentive effect of benefit taxation might have underestimated the increase in tax revenue by as much as about 25%. All of the above analysis, however, assumes that the claimant's weekly wage W remains constant. If the claimants faced with benefit taxation return to work more quickly by accepting lower-wage jobs, the second term in the equation for ΔT is correspondingly reduced. The empirical evidence on whether UI-induced duration changes are indeed accompanied by wage-rate changes is ambiguous.[22]

Another budgetary implication of benefit taxation's work incentive effect is the impact on UI program costs. These costs are reduced by $\Delta DUR \cdot WBA$, the tax-induced reduction in gross benefit income. The Georgia results imply that, for the 1979 high-income claimants, this reduction in benefit payments averaged \$115, an 11% reduction from the \$1030 average benefit income they would have collected in the absence of benefit taxation.

The work incentive effects of benefit taxation, along with the attendant budgetary effects, do not by themselves prove that benefit taxation is good policy. Like any cutback in an income transfer program, a tax-induced reduction in net unemployment compensation may undercut the income maintenance objectives of the program. If benefit taxation is not accompanied by an increase in pre-tax benefit levels, work incentives may be improved, but the unemployment insurance program also will be less effective in its purpose of insuring job losers against income reductions.

Appendix: Properties of the Generalized Weibull Distribution

The hazard function for the generalized Weibull distribution is

$$h(y) = \alpha y^{\alpha-1} \exp\{-\beta' X - (\log B - \log y)\, G\} \quad \text{for } y < B$$

$$= \alpha y^{\alpha-1} \exp\{-\beta' X\} \quad \text{for } y > B$$

where $B = P + 1$ and $G = \gamma_1(WBA) + \gamma_2(D \cdot WBA)$. The relation between the cumulative distribution and hazard functions

$$F(Y) = 1 - \exp\{-\int_0^Y h(y)dy\}$$

implies that

$$F(Y) = 1 - \exp\{-[\alpha/(\alpha+G)]\; Y^{\alpha+G}\; B^{-G} \exp(-\beta' X)\} \quad \text{for } Y < B$$

$$= 1 - \exp\{-[Y^{\alpha} - (G/(\alpha+G))\; B^{\alpha}] \exp(-\beta' X)\} \quad \text{for } Y > B \;.$$

Then

$$F(B) = 1 - \exp\{-[\alpha/(\alpha+G)]\; B^{\alpha} \exp(-\beta' X)\} \;.$$

Differentiation of F(Y) with respect to Y gives the probability density function

$$f(Y) = \alpha Y^{\alpha+G-1} B^{-G} \exp\{-\beta' X - [\alpha/(\alpha+G)]\, Y^{\alpha+G} B^{-G} \exp(-\beta' X)\} \quad \text{for } Y < B$$

$$= \alpha Y^{\alpha-1} \exp\{-\beta' X - [Y^{\alpha} - (G/(\alpha+G)) B^{\alpha}] \exp(-\beta' X)\} \quad \text{for } Y > B \;.$$

Now consider the censored variable

$$Y^* = Y \quad \text{if } Y < B$$
$$\quad = B \quad \text{if } Y > B \quad .$$

The likelihood function for a sample of Y^* is

$$L = \prod_1 f(Y_i^*) \prod_2 \left[1 - F(Y_i^*)\right]$$

where $\prod_1$ denotes a product taken over observations with $Y_i^* < B_i$ and $\prod_2$ denotes a product taken over observations with $Y_i^* = B_i$. Dropping subscripts and substituting in the appropriate distribution and density functions,

$$L = \prod_1 \alpha Y^{*\alpha+G-1} B^{-G} \exp\{-\beta' X - [\alpha/(\alpha+G)] \; Y^{*\alpha+G} B^{-G} \exp(-\beta' X)\}$$

$$\prod_2 \exp\{-[\alpha/(\alpha+G)] \; Y^{*\alpha} \exp(-\beta' X)\} \quad .$$

Then the log likelihood function is

$$\log L = \sum_1 \{\log \alpha + (\alpha+G-1) \log Y^* - G \log B - \beta' X$$
$$- [\alpha/(\alpha+G)] \; Y^{*\alpha+G} B^{-G} \exp(-\beta' X)\}$$
$$- \sum_2 [\alpha/(\alpha+G)] \; Y^{*\alpha} \exp(-\beta' X) \quad .$$

The expected value of the uncensored variable Y is

$$E(Y) = \int_0^B Y\, f(Y)\, dY + \int_B^\infty Y\, f(Y)\, dY \quad .$$

Denoting the first integral by M and the second by N,

$$M = \int_0^B \alpha Y^{\alpha+G} B^{-G} \exp\{-\beta'X - [\alpha/(\alpha+G)]\, Y^{\alpha+G} B^{-G} \exp(-\beta'X)\}\, dY \quad .$$

If we let $Z = [\alpha/(\alpha+G)]Y^{\alpha+G}B^{-G}\exp(-\beta'X)$, then

$$M = \{[(\alpha+G)/\alpha]B^G\exp(\beta'X)\}^{1/(\alpha+G)} \int_0^{[\alpha/(\alpha+G)]B^\alpha\exp(-\beta'X)} Z^{[(\alpha+G+1)/(\alpha+G)]-1}\, e^{-Z} dZ$$

$$= \{[(\alpha+G)/\alpha]B^G\exp(\beta'X)\}^{1/(\alpha+G)}\, \Gamma_{[\alpha/(\alpha+G)]B^\alpha\exp(-\beta'X)}\{(\alpha+G+1)/(\alpha+G)\}$$

where the subscripted Γ term is the incomplete gamma function with argument $(\alpha+G+1)/(\alpha+G)$ and upper limit $[\alpha/(\alpha+G)]B^\alpha\exp(-\beta'X)$. The second integral is

$$N = \int_B^\infty \alpha Y^\alpha \exp\{-\beta'X - [Y^\alpha - (G/(\alpha+G))\, B^\alpha]\, \exp(-\beta'X)\}\, dY \quad .$$

If we now let $Z = Y^\alpha \exp(-\beta'X)$, then

$$N = \exp\{(\beta'X/\alpha)+[G/(\alpha+G)]B^\alpha\exp(-\beta'X)]\} \int_{B^\alpha\exp(-\beta'X)}^\infty Z^{[(\alpha+1)/\alpha]-1}\, \exp(-Z)\, dZ$$

$$= \exp\{(\beta'X/\alpha)+[G/(\alpha+G)]B^\alpha\exp(-\beta'X)\}\, [\Gamma\{(\alpha+1)/\alpha\} - \Gamma_{B^\alpha\exp(-\beta'X)}\{(\alpha+1)/\alpha\}]$$

where the unsubscripted gamma term denotes a complete gamma function.

Finally, the expected value of the censored variable Y^* is

$$E(Y^*) = M + \left[1 - F(B)\right] B \quad .$$

Substituting in the expressions derived above for M and F(B) produces equation (4) in the text.

Footnotes

1. See Feldstein (1974), for example.

2. Clines (1982).

3. See Moffitt and Nicholson (1982), for example.

4. See Unemployment Insurance Service (1977).

5. This sample restriction does not alleviate two other sources of error in ascertaining which claimants were taxable. One is that the CWBH questionnaire data, like other survey data, are subject to considerable income misreporting (see Strouse (1980)). Second, the CWBH income variable refers to the claimant's family income during the 52 weeks before he filed his claim, whereas the relevant income measure for tax purposes is family income during the calendar year. These problems in income measurement undoubtedly caused errors in determining whether claimants were above or below the income thresholds for benefit taxation. This misclassification of claimants with respect to taxable status might tend to obscure true between-group differences in unemployment duration and bias the estimated impact of benefit taxation toward zero.

6. The reason 34% of the taxable claimants did not qualify for the maximum WBA is that, although their family income was high, their own earnings were low. Their high income was due mainly to the earnings of other family members.

7. The author thanks Sherryl Edge and Joe Woodall for this information, as well as for other advice about the Georgia progam and data.

8. This is documented for New York claimants in Entes (1980).

9. The approach used is similar to Rosen's (1976) and Williams' (1975) method for estimating the impact of taxes on female labor supply.

10. Accurate imputation of each individual's tax rate is precluded by the broad interval form in which the CWBH income data are reported. For example, for a claimant whose income is above $25,000, the only other information available is whether his income lies in the interval $25,000-29,999 or in the interval $30,000 and above.

11. Other unemployment measures -- the average insured unemployment rate for the year and the insured and total rates for the claimant's month of filing and for the first three months of his benefit year -- also were tried with virtually no effect on the results.

12. Welch (1977) discusses this issue in detail. The interaction of high-quarter earnings with the 1979 dummy is included to allow for a time effect that varies with wage level.

13. The estimated standard error was computed with the formula in Mood, Graybill, and Boes (1974), p. 181, for approximating the variance of the ratio of two random variables.

14. Several earlier studies of UI and unemployment duration excluded individuals that filed valid claims but were unemployed too briefly to collect benefits. As Classen (1979) and Welch (1977) point out, such an exclusion truncates the sample on the basis of the dependent variable and therefore biases the estimated effects of benefit variables. Unsurprisingly, rerunning the regression in column 1 of Table 1 with such an exclusion preserves the qualitative results, but reduces the estimated WBA coefficient from .045 to .036 and the estimated D·WBA coefficient from -.018 to -.011.

15. Inspection of empirical hazard rates for a homogeneous subsample shows that the assumption that the hazard rate changes monotonically with y is consistent with the data.

16. See Johnson and Kotz (1970) for a detailed discussion of these and other properties of the Weibull distribution.

17. This likelihood function is correct provided that the <u>conditional</u> unemployment duration distribution (given the explanatory variables X) is independent of the censoring time. See Kalbfleisch and Prentice (1980) for a detailed discussion of this issue. In their parlance, the present case of censoring on the basis of the predetermined explanatory variable P is censorship on the basis of a "fixed covariate."

18. The maximum likelihood estimation was performed with the Davidon-Fletcher-Powell and GRADX algorithms in the GQOPT numerical optimization package. The algorithms are discussed in Quandt (forthcoming). They converged to the same final paramater estimates when started from different initial values.

19. See Bennett and Franklin (1954) for a discussion of the incomplete gamma function.

20. An analagous simulation for the Weibull model estimates a 1.1 week effect. The estimated effect on <u>total</u> unemployment duration in the generalized model is 2.2 weeks, but confidence in this estimate requires strong faith in the model's goodness of fit beyond the point of censorship.

21. This computation uses 1/13 of high-quarter earnings in the base period as an estimate of the weekly wage W.

22. See the studies by Classen (1979), Ehrenberg and Oaxaca (1976), and Holen (1977), and the critical review by Welch (1977).

23. See Bennett and Franklin (1954) for a discussion of the incomplete gamma function.

References

Bennett, Carl A., and Norman L. Franklin. Statistical Analysis in Chemistry and the Chemical Industry. New York: Wiley, 1954.

Classen, Kathleen P. "Unemployment Insurance and Job Search." Studies in the Economics of Search, ed. S. A. Lippman and J. J. McCall, pp. 191-219. Amsterdam: North-Holland, 1979.

Clines, Francis X. "President's Aides Study Plan to Tax Jobless Benefits." New York Times (November 26, 1982): A1.

Ehrenberg, Ronald G., and Ronald L. Oaxaca. "Unemployment Insurance, Duration of Unemployment, and Subsequent Wage Gain." American Economic Review 66 (December 1976): 754-66.

Entes, Ruth. Spells of Unemployment and Other Studies of Unemployment Insurance Beneficiaries, New York State. Albany: New York State Department of Labor, Division of Research and Statistics, 1980.

Feldstein, Martin. "Unemployment Compensation: Adverse Incentives and Distributional Anomalies." National Tax Journal 27 (June 1974): 231-44.

______________. "The Effect of Unemployment Insurance on Temporary Layoff Unemployment." American Economic Review 68 (December 1978): 834-46.

Heckman, James J., and George J. Borjas. "Does Unemployment Cause Future Unemployment? Definitions, Questions, and Answers from a Continuous Time Model of Heterogeneity and State Dependence." Economica 47 (August 1980): 247-83.

Holen, Arlene. "Effects of Unemployment Insurance Entitlement on Duration and Job Search Outcome." Industrial and Labor Relations Review 30 (July 1977): 445-50.

Johnson, Norman L., and Samuel Kotz. Distributions in Statistics: Continuous Univariate Distributions -- 1. New York: Wiley, 1970.

Kalbfleisch, John D., and Ross L. Prentice. The Statistical Analysis of Failure Time Data. New York: Wiley, 1980.

Lancaster, Tony. "Econometric Methods for the Duration of Unemployment." Econometrica 47 (July 1979): 939-56.

Lippman, Steven A., and John J. McCall. "The Economics of Job Search: A Survey." Economic Inquiry 14 (June 1976): 155-89.

Moffitt, Robert, and Walter Nicholson. "The Effect of Unemployment Insurance on Unemployment: The Case of Federal Supplemental Benefits." Review of Economics and Statistics 64 (February 1982): 1-11.

Mood, Alexander M., Franklin A. Graybill, and Duane C. Boes. Introduction to the Theory of Statistics. New York: McGraw-Hill, 1974.

Mortensen, Dale T. "Job Search, the Duration of Unemployment, and the Phillips Curve." American Economic Review 60 (December 1970): 847-62.

Newton, Floyd C., and Harvey S. Rosen. "Unemployment Insurance, Income Taxation, and Duration of Unemployment: Evidence from Georgia." Southern Economic Journal 45 (January 1979): 773-84.

Nickell, Stephen. "Estimating the Probability of Leaving Unemployment." Econometrica 47 (September 1979): 1249-66.

Quandt, Richard. "Computational Problems and Methods." Handbook of Econometrics, ed. Zvi Griliches and Michael D. Intriligator. Amsterdam: North-Holland, forthcoming.

Rosen, Harvey S. "Taxes in a Labor Supply Model with Joint Wage-Hours Determination." Econometrica 44 (May 1976): 485-507.

Strouse, Richard. An Evaluation of the CWBH Questionnaire. Princeton, N.J.: Mathematica Policy Research, 1980.

Unemployment Insurance Service. Manual for Continuous Wage and Benefit History Data Bank. Washington: U.S. Department of Labor, 1977.

Welch, Finis. "What Have We Learned from Empirical Studies of Unemployment Insurance?" Industrial and Labor Relations Review 30 (July 1977): 451-61.

Williams, Robert George. Public Assistance and Work Effort: The Labor Supply of Low-Income Female Heads of Household. Princeton, N.J.: Industrial Relations Section, Princeton University, 1975.