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Gary Solon


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WORK INCENTIVE EFFECTS OF TAXING UNEMPLOYMENT BENEFITS¹

BY GARY SOLON

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.

1. INTRODUCTION

BEFORE 1979, UNEMPLOYMENT INSURANCE (UI) BENEFITS were not treated as taxable income in the United States. Several economists² 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.³

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 describes the study’s data base. Section 3 presents analyses of the data, and Section 4 summarizes the results.

¹ This study was partially supported by a grant from the Sloan Foundation to the Department of Economics at Princeton University. The author thanks Orley Ashenfelter, David Card, Angus Deaton, Ronald Ehrenberg, Paul Mackin, Richard Quandt, Michael Ransom, the referees, and seminar participants too numerous to mention for their generous advice.
² See Feldstein [6], for example.
³ Clines [3].
2. 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 per cent 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 data analysis 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 \(\frac{2}{5}\) of his highest-quarter earnings in the base period, except that the minimum WBA was $27 and the maximum was $90. Forty-five per cent of the sample claimants (and 66 per cent 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 \(\frac{2}{5}\) of his base period earnings

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\(^4\) See Unemployment Insurance Service [19].

\(^5\) 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 [18]). 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 per cent 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.
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 per cent 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 $\frac{2}{3}$ 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. 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 [7] 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.

The duration measure used throughout the analysis is one plus the number of weeks that the claimant collected UI during his benefit year. Compensated duration is incremented by one because many sample members initiated valid claims but collected no benefits, presumably because they returned to work before completing a full week of unemployment. The incrementation therefore makes sense as a procedure for rounding fractional weeks of unemployment up to the next integer, and later it avoids the problem of taking the logarithm of zero. It should be understood that the variable used does not measure duration per spell because many claimants collect benefits in more than one spell during the benefit year. The duration variable therefore does not accord perfectly with job search theory, but the effect of benefit taxation on total weeks unemployed, rather than weeks per spell, is probably of greater policy interest. Finally, it should be noted that number of weeks of benefit collection is a censored duration measure. For the 24 per cent 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 below.

3. 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 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.

The basic behavioral equation in the empirical analysis is posited to take the form

$$ Y = g\{\delta(1 - \rho t) \text{WBA} + \gamma' Z\}; $$

that is, unemployment duration $Y$ depends on the bracketed linear function, in which $Z$ is a vector of control variables to be discussed later. 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 $\rho$ 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.\(^{10}\)

Now let $t$ equal a constant $\bar{t}$ for those claimants whose benefits are taxable, and let $t = 0$ for the nontaxable claimants.\(^{11}\) This dichotomous treatment of $t$ is

\(^{10}\) This approach is similar to Rosen's [16] and Williams' [21] method for estimating the impact of taxes on female labor supply.

\(^{11}\) 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 $25,000–25,999 or in the interval $30,000 and above.
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

$$Y = g\{\delta(1 - \rho iD)WBA + \gamma'Z\}$$

$$= g\{\delta WBA - \delta \rho i(D \cdot WBA) + \gamma'Z\}.$$ 

This last expression allows $WBA$ and $D \cdot WBA$ to be entered as separate explanatory variables in the duration equation. The coefficient of $WBA$, $\delta$, 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, such as those by Classen [2] and Ehrenberg and Oaxaca [4]. The coefficient of $D \cdot WBA$, $-\delta \rho i$, 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 $\delta = 0$ if UI benefits have no effect at all). In this case, the coefficient of $D \cdot WBA$ should be zero. But if taxes do affect duration, then $\rho > 0$ and $\delta > 0$, in which case the coefficient of $D \cdot WBA$ should be negative. Moreover, the negative of the coefficient of $D \cdot WBA$ divided by the coefficient of $WBA$ gives an implied value of $\rho i$. Combined with extraneous information on $i$, an estimate of $\rho i$ provides information on the value of $\rho$. 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 $\rho i$ yields a rough estimate of $\rho$. Alternatively, if one wishes to begin with the hypothesis that $\rho = 1$, the negative of the ratio of estimated coefficients gives an estimate of $i$, which can be compared with the extraneous information on $i$. If the comparison is close, one might accept the hypothesis $\rho = 1$.

The choice of the control variables $Z$ is guided by theoretical analyses of UI and unemployment duration, such as Mortensen's [14] job search model and Moffitt and Nicholson's [12] labor supply analysis. In accordance with these analyses, the empirical model controls for any observable variables likely to be related to claimants' cost of unemployment, leisure-income preference, or distribution of employment opportunities. The variables include potential benefit duration, high-quarter earnings, 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, 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, the high-quarter earnings variable is supplemented by a squared term, a term interacted with the

\[12\] Welch [20] 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.
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. All benefit and earnings variables were converted to October 1980 dollars with the Atlanta Consumer Price Index.\footnote{Additional analyses reported in Solon [17] check the effects of including alternative measures of the unemployment rate, squared terms for education and age, an interaction of WBA and the high-income dummy, and the \(D\) variable not interacted with WBA. None of these variations produces important changes in the results.}

The main remaining specification issue involves the assumed distribution for unemployment duration. Lancaster [11] has proposed the use of the Weibull distribution, a convenient formulation of which implies a reemployment hazard (or exit-from-unemployment rate) function of the form

\[
\begin{align*}
  h(y) &= \alpha y^{\alpha-1} \exp (-\beta'X) \\
  f(Y) &= \alpha Y^{\alpha-1} \exp (-\beta'X - Y^\alpha \exp (-\beta'X))
\end{align*}
\]

where \(y\) is the number of weeks already unemployed, \(X\) is a vector of variables (including \(WBA\) and \(D \cdot WBA\)) that may affect duration, \(\alpha\) is a parameter greater than zero, and \(\beta\) 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 distribution degenerates to the special case of the exponential distribution. If \(\alpha > 1\), the reemployment hazard rises with duration; if \(\alpha < 1\), it declines.

The probability density function of completed duration \(Y\) is then

\[
F(Y) = \exp (\beta'X / \alpha) \Gamma((\alpha + 1)/\alpha)
\]

where \(\Gamma\) is the gamma function.\footnote{See Johnson and Kotz [9] for a detailed discussion of these and other properties of the Weibull distribution.} If we differentiate the natural logarithm of expected duration with respect to \(x_n\), the \(h\)th variable in \(X\), we obtain

\[
\frac{\partial \log E(Y)}{\partial x_h} = \frac{\beta_h}{\alpha}.
\]

Thus, estimates of \(\beta\) and \(\alpha\) 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

\[
Y^* = Y \text{ if } Y < P + 1 \\
= P + 1 \text{ if } Y \geq P + 1
\]

where \(Y^*\) is compensated duration plus one and \(P\) is potential benefit duration. To deal with this "right censorship," we now derive a maximum likelihood estimation technique for the Weibull distribution analogous 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
\[
\operatorname{Prob} (Y_i \geq P_i + 1) = \int_{P_i + 1}^{\infty} f(Y) \, dY
\]
\[
= \exp \{- Y^*_i \alpha \exp (-\beta'X_i)\}.
\]
Hence, the likelihood function for the full sample is
\[
L = \prod_1 \alpha Y^*_i \alpha^{-1} \exp \{-\beta'X_i - Y^*_i \alpha \exp (-\beta'X_i)\}
\]
\[
\times \prod_2 \exp \{- Y^*_i \alpha \exp (-\beta'X_i)\}
\]
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.\(^{15}\) 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 nonexhaustees, and \( \sum_2 \) is a sum over exhaustees.

The parameters \( \alpha \) and \( \beta \) can be estimated by maximizing the log likelihood function with respect to the parameters. This procedure was applied to the Georgia data with the Davidon-Fletcher-Powell and GRADX algorithms in the GQOPT numerical optimization package.\(^{16}\)

The results are reported in the first column of Table I. The estimated value of .8 for the parameter \( \alpha \) is significantly less than 1 and implies that the reemployment hazard declines with duration. As Heckman and Borjas [8] and Lancaster [11] 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 identical 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

\(^{15}\) This likelihood function is correct provided that the conditional unemployment duration distribution (given the explanatory variables \( X \)) is independent of the censoring time. See Kalbfleisch and Prentice [10] 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.”

\(^{16}\) These algorithms are discussed in Quandt [15]. They converged to the same final parameter estimates when started from different initial values.
TABLE I
Estimates Parameters (and Standard Errors) from Models of Unemployment Duration

<table>
<thead>
<tr>
<th>Explanatory Variables</th>
<th>Weibull Model</th>
<th>Generalized Weibull Model</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>.359</td>
<td>.240</td>
</tr>
<tr>
<td>WBA</td>
<td>(.668)</td>
<td>(.672)</td>
</tr>
<tr>
<td></td>
<td>.0071</td>
<td>.0053</td>
</tr>
<tr>
<td></td>
<td>(.0012)</td>
<td>(.0014)</td>
</tr>
<tr>
<td>$D \cdot WBA$</td>
<td>-.0016</td>
<td>-.0003</td>
</tr>
<tr>
<td></td>
<td>(.0010)</td>
<td>(.0013)</td>
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<tr>
<td>High income</td>
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<td>-.142</td>
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<tr>
<td></td>
<td>(.102)</td>
<td>(.105)</td>
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<tr>
<td>Potential benefit duration</td>
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<td>.026</td>
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<tr>
<td></td>
<td>(.008)</td>
<td>(.009)</td>
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<tr>
<td>Female</td>
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<td>.050</td>
</tr>
<tr>
<td></td>
<td>(.048)</td>
<td>(.050)</td>
</tr>
<tr>
<td>Black or Hispanic</td>
<td>.125</td>
<td>.124</td>
</tr>
<tr>
<td></td>
<td>(.033)</td>
<td>(.035)</td>
</tr>
<tr>
<td>Occupation:</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Professional, tech., managerial</td>
<td>.250</td>
<td>.251</td>
</tr>
<tr>
<td></td>
<td>(.063)</td>
<td>(.066)</td>
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<tr>
<td>Clerical, sales</td>
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<td>.120</td>
</tr>
<tr>
<td></td>
<td>(.052)</td>
<td>(.054)</td>
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<tr>
<td>Service</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Agric., fishery, forestry, related</td>
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<td>.267</td>
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<tr>
<td></td>
<td>(.154)</td>
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<td>Processing</td>
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<td>-.057</td>
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<tr>
<td></td>
<td>(.076)</td>
<td>(.080)</td>
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<tr>
<td>Machine trades</td>
<td>.039</td>
<td>.038</td>
</tr>
<tr>
<td></td>
<td>(.060)</td>
<td>(.062)</td>
</tr>
<tr>
<td>Benchwork</td>
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<td>-.0009</td>
</tr>
<tr>
<td></td>
<td>(.065)</td>
<td>(.068)</td>
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<tr>
<td>Structural work</td>
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<td>.174</td>
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<tr>
<td></td>
<td>(.059)</td>
<td>(.061)</td>
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<tr>
<td>Miscellaneous</td>
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<td>-.079</td>
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<td></td>
<td>(.057)</td>
<td>(.059)</td>
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<tr>
<td>High-quarter earnings</td>
<td>-.088</td>
<td>-.086</td>
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<tr>
<td>(HQE, in thousands)</td>
<td>(.036)</td>
<td>(.037)</td>
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<tr>
<td>HQE squared</td>
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<td>.0015</td>
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<tr>
<td></td>
<td>(.0018)</td>
<td>(.0019)</td>
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<tr>
<td>HQE $\times$ high income</td>
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<td>.051</td>
</tr>
<tr>
<td></td>
<td>(.018)</td>
<td>(.018)</td>
</tr>
<tr>
<td>Base-period earnings/HQE</td>
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<td>-.118</td>
</tr>
<tr>
<td></td>
<td>(.049)</td>
<td>(.051)</td>
</tr>
<tr>
<td>Married</td>
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<td>-.123</td>
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<tr>
<td></td>
<td>(.048)</td>
<td>(.049)</td>
</tr>
<tr>
<td>Married $\times$ female</td>
<td>-.155</td>
<td>-.153</td>
</tr>
<tr>
<td></td>
<td>(.106)</td>
<td>(.107)</td>
</tr>
<tr>
<td>Spouse working</td>
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<td>-.020</td>
</tr>
<tr>
<td></td>
<td>(.050)</td>
<td>(.051)</td>
</tr>
<tr>
<td>Spouse working $\times$ female</td>
<td>.346</td>
<td>.344</td>
</tr>
<tr>
<td></td>
<td>(.106)</td>
<td>(.109)</td>
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<tr>
<td>Education</td>
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<td>.012</td>
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<tr>
<td></td>
<td>(.007)</td>
<td>(.007)</td>
</tr>
<tr>
<td>Age</td>
<td>.014</td>
<td>.014</td>
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<tr>
<td></td>
<td>(.0014)</td>
<td>(.0014)</td>
</tr>
</tbody>
</table>
must be divided by the estimate of $\alpha$ to obtain the proportional change in expected duration associated with a unit change in untaxed $WBA$. The result implies a proportional change of .0089, which is toward the upper end of the range from previous studies.

The estimated $D \cdot WBA$ 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 $\alpha$ 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 $D \cdot WBA$ and $WBA$ coefficients implies a tax rate of .22 (with an estimated standard error of .15), somewhat (but insignificantly) less than the expected $\bar{t}$ of slightly above .3. Dividing the ratio by the expected $\bar{t}$ would give a point estimate of about .7 for $\rho$.

While the Weibull framework provides a convenient and easily interpreted model that allows for duration dependence, it seems somewhat restrictive in two respects. First, it requires the hazard rate to vary monotonically with time already

\[ \text{Log likelihood} \times 10^{-4} = -1.733748 \]
\[ \text{Number of Observations} = 6,610 \]
unemployed. Inspection of empirical hazard rates for a homogeneous subsample, however, reveals that this assumption is consistent with the data. Second, 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

\[ h(y) = ay^{a-1} \exp \{-\beta'X - [\log (P+1) - \log y] \times [\gamma_1(WBA) + \gamma_2(D \cdot WBA)] \} \text{ for } y \leq P+1, \]

\[ = ay^{a-1} \exp (-\beta'X) \text{ for } y \geq P+1. \]

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 to Solon [17], this generalization of the Weibull distribution leads to the new log likelihood function

\[ \log L = \sum \{\log \alpha + (\alpha + G_i - 1) \log Y_i^* - G_i \log (P_i + 1) - \beta'X_i \]

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

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

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 second column of Table I. 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 \( \beta \) and \( \gamma \) coefficients for \( WBA \) and \( D \cdot WBA \) 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 \( \gamma_2 \) and the \( \beta \) coefficient for \( D \cdot WBA \) constrained to equal zero. The resulting log likelihood value was \(-1.733641 \times 10^{-4}\), 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 above-cited Appendix, the expected value of compensated unemployment duration in the generalized model is

\[ (4) \hspace{1cm} E(Y^*) = \left\{\left[\frac{\alpha + G}{\alpha}(P+1)^G \exp (\beta'X)\right]\right\}^{1/(\alpha+G)} \times \Gamma[\alpha/(\alpha+G)](P+1)^\alpha \exp (-\beta'X)\left\{\frac{\alpha + G + 1}{\alpha + G}\right\} \]

\[ + \exp \left\{-\left[\frac{\alpha}{\alpha + G}\right](P+1)^\alpha \exp (-\beta'X)\right\}(P+1) \]
where the subscripted \( f \) term is an incomplete gamma function.\(^{18}\) 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 \( \alpha, \beta, \) and \( \gamma \).

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 \( D \cdot WBA = 0 \) 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.\(^{19}\)

4. SUMMARY

This paper has investigated the effect of taxing unemployment benefits on unemployment duration. The results suggest that unemployment benefit levels do affect unemployment duration and that the duration impact of taxing benefits is qualitatively 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 work incentive effect, however, does not by itself 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.

University of Michigan

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\(^{18}\) See Bennett and Franklin [1] for a discussion of the incomplete gamma function.

\(^{19}\) An analogous simulation for the Weibull model estimates a 1.1 week effect. The estimated effect on total 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.


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