WORK INCENTIVE EFFECTS OF TAXING UNEMPLOYMENT BENEFITS*

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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 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 limited-dependent-variable technique for the Weibull distribution similar to the Tobit technique for the normal distribution. Despite some variation in the results from different model specifications, the analysis repeatedly rejects the hypothesis of no tax effect. The 1979 policy change is estimated to have reduced average compensated unemployment duration among the sampled high-income claimants by almost a week and a half.

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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 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. Furthermore, the Reagan administration has proposed extending benefit taxation to lower income brackets as one of its "revenue enhancement" measures. Press reports have listed the alleged work disincentive effect of not taxing benefits as one of the reasons for the proposal.(2)

(1) See Feldstein [7], for example.
(2) Cowan [4].
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

Most previous theoretical and empirical research on UI and unemployment duration has 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 theoretical results apply. Whether recipients do, in fact, perceive benefit taxation as a benefit reduction is the main topic of this paper.

Theoretical work has usually analyzed UI and unemployment duration in the framework of job search theory. Mortensen [20] provides a representative example of this approach, and Lippman and McCall [17] 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. 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

(3) Hamermesh [9], for example, analyzes UI’s duration effects in a labor supply framework.
leisure-income preference, and variables associated with wage offer distributions. Disentangling UI effects from wage effects is especially difficult because each state UI program 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 is to analyze data on UI claimants in several states with different benefit formulas; the other is to use single-state data spanning a period when the state's benefit schedule has been substantially changed. With either approach, the sample contains claimants with similar earnings histories but different UI entitlements.

A fair summary of the results of studies using either approach (4) is that, holding earnings variables and other factors constant, raising weekly benefit levels by $10 (in 1980 dollars) increases expected unemployment duration by .3 to .9 weeks. Although each study suffers from methodological problems (5) (some of which are remedied later in this paper), the fact that they yield similar results despite different data bases and model specifications gives some confidence that the results are in the right ballpark. Whether the results can be extrapolated to estimate the duration effects of benefit taxation depends on whether UI claimants

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(4) Classen [3], Ehrenberg and Oaxaca [5], Holen [13], Kieffer and Neumann [15], and Solon [24].

(5) Many of the problems are discussed by Topel and Welch [26] and Welch [28].

- 4 -
respond to benefit taxation as a benefit reduction. This empirical question occupies the remainder of this paper.

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. 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. 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.

(6) See Unemployment Insurance Service [27].
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.(7)

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. The claimant's total entitlement during his "benefit year," the 52-week period following

(7) This sample restriction does not alleviate two other sources of error in ascertaining which claimants were taxable. One is that the CWHH questionnaire data, like other survey data, are subject to considerable income misreporting (see Strouse [25]). Second, the CWHH 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 differences in unemployment duration behavior between taxable and nontaxable claimants.
his initial claim-filing, was the lesser of $1/4 of his base period earnings or 26 times his WBA. Consequently, some 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. As discussions of sample selection bias (8) have made clear, the estimated slope coefficients in the regression analyses below will be biased if the error term in the unemployment duration equation is correlated with the response probability.

This would be the case if the response probability depended on unemployment duration. Fortunately, such dependence seems unlikely because claimants were asked to fill out the questionnaire at the time of initial filing. Nonetheless, one aspect of questionnaire nonresponse deserves mention. A major cause of nonresponse was Georgia's system of employer-filed claims, under which an employer temporarily closing its plant could submit a packet of UI claims for all the employees it had put on temporary layoff. 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 [8] 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 correctly perceived benefit taxation as a reduction in net benefit levels. 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. 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 truncated duration measure. For 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. The trunca-

(9) This is documented for New York claimants in Entes [6].
tion problem 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.

4.1 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.(10)

The basic behavioral equation is posited to take the form

(10) The approach used is similar to Rosen's [23] and Williams' [29] method for estimating the impact of taxes on female labor supply.
DURATION = f{b(1 - pt)WBA + c'X + u};

that is, duration is functionally dependent on the bracketed linear function, in which X is a vector of control variables and u 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 p is a coefficient of tax perception such that \( p = 0 \) if claimants behave as if they are unaware of the tax and \( p = 1 \) if they are fully aware of the tax. 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 suppose that t is approximately a constant \( \bar{t} \) for those claimants whose benefits are taxable, and let D be a dummy variable that equals 1 for those claimants and 0 for the others. Then

\[
DURATION = f\{b(1 - p\bar{t})WBA + c'X + u\}
= f\{b(WBA) - bp\bar{t}(D\cdot WBA) + c'X + u\}.
\]

This last expression allows WBA and D\cdot WBA to be entered as separate regressors in the duration equation. The coefficient of WBA, b, measures WBA's duration effect for claimants whose benefits are not taxable. The estimate of this coefficient should be similar to the estimated WBA coeffi-
cients in earlier studies of UI and duration. The coefficient of D\*WBA, \(-b(pt)\), 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 \(p=0\) (or \(b=0\) 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 \(p>0\) and \(b>0\), in which case the coefficient of D\*WBA should be negative. Moreover, if \(p=1\), the negative of the coefficient of D\*WBA divided by the coefficient of WBA gives an implied value of \(\bar{e}\). The plausibility of this value will serve as a check on the model.

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 and a term interacted with the high-income 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.

The coefficient of WBA is estimated as .049 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 .5 weeks, a result consistent with the findings of previous studies. The coefficient of D·WBA is estimated as -.024, and its difference from zero also is decidedly significant. Hence, the null hypothesis of no tax effect is rejected. If

(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.

<table>
<thead>
<tr>
<th>Explanatory Variables</th>
<th>Level Regression</th>
<th>Log Regression</th>
<th>Weibull Model</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>-6.22 (4.75)</td>
<td>-0.046 (.701)</td>
<td>.476 (.673)</td>
</tr>
<tr>
<td>WBA</td>
<td>.049 (.008)</td>
<td>.0083 (.0012)</td>
<td>.0076 (.0012)</td>
</tr>
<tr>
<td>D-WBA</td>
<td>-.024 (.007)</td>
<td>-.0036 (.0010)</td>
<td>-.0024 (.0010)</td>
</tr>
<tr>
<td>High income</td>
<td>-.21 (.73)</td>
<td>.027 (.108)</td>
<td>-.090 (.104)</td>
</tr>
<tr>
<td>Potential benefit</td>
<td>.43 (.06)</td>
<td>.042 (.009)</td>
<td>.030 (.009)</td>
</tr>
<tr>
<td>duration</td>
<td>.34 (.34)</td>
<td>.083 (.050)</td>
<td>.053 (.048)</td>
</tr>
<tr>
<td>Female</td>
<td>.70 (.37)</td>
<td>.119 (.055)</td>
<td>.117 (.053)</td>
</tr>
<tr>
<td>Black or Hispanic</td>
<td>.82 (.23)</td>
<td>.099 (.034)</td>
<td>.125 (.033)</td>
</tr>
<tr>
<td>Occupation:</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Professional, tech.,</td>
<td>1.85 (.44)</td>
<td>.280 (.065)</td>
<td>.244 (.063)</td>
</tr>
<tr>
<td>managerial</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Clerical, sales</td>
<td>.70 (.37)</td>
<td>.119 (.055)</td>
<td>.117 (.053)</td>
</tr>
<tr>
<td>Service</td>
<td>---</td>
<td>---</td>
<td>---</td>
</tr>
<tr>
<td>Agric., fishery,</td>
<td>1.37 (.105)</td>
<td>.157 (.155)</td>
<td>.275 (.159)</td>
</tr>
<tr>
<td>forestry, related</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Processing</td>
<td>-.38 (.56)</td>
<td>.056 (.082)</td>
<td>-.062 (.077)</td>
</tr>
<tr>
<td>Machine trades</td>
<td>.24 (.43)</td>
<td>.090 (.064)</td>
<td>.035 (.061)</td>
</tr>
<tr>
<td>Benchwork</td>
<td>-.05 (.47)</td>
<td>.079 (.069)</td>
<td>-.006 (.069)</td>
</tr>
<tr>
<td>Structural work</td>
<td>1.11 (.42)</td>
<td>.275 (.062)</td>
<td>.176 (.059)</td>
</tr>
<tr>
<td>Miscellaneous</td>
<td>-.71 (.42)</td>
<td>-.074 (.061)</td>
<td>-.083 (.057)</td>
</tr>
<tr>
<td>High-quarter earnings</td>
<td>-.89 (.25)</td>
<td>-.150 (.037)</td>
<td>-.123 (.035)</td>
</tr>
<tr>
<td>(MQE, in thousands)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>MQE squared</td>
<td>.014 (.014)</td>
<td>.0080 (.0021)</td>
<td>.0021 (.0020)</td>
</tr>
<tr>
<td>MQE X high income</td>
<td>.40 (.13)</td>
<td>.046 (.019)</td>
<td>.049 (.019)</td>
</tr>
<tr>
<td>Base-period earnings/</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>MQE</td>
<td>-.76 (.36)</td>
<td>-.144 (.053)</td>
<td>-.118 (.050)</td>
</tr>
<tr>
<td>Married</td>
<td>-.72 (.35)</td>
<td>-.048 (.051)</td>
<td>-.121 (.048)</td>
</tr>
<tr>
<td>Married X female</td>
<td>-1.12 (.77)</td>
<td>-1.236 (.113)</td>
<td>-1.155 (.105)</td>
</tr>
</tbody>
</table>
Spouse working  -.25  -.024  -.020
(.36)  (.054)  (.050)
Spouse working X female  2.23  .257  .347
(.77)  (.114)  (.107)
Education  .08  .014  .012
(.05)  (.007)  (.007)
Age  .08  .010  .014
(.009)  (.0014)  (.0014)
Expecting recall  -.28  .104  -.039
(.24)  (.035)  (.034)
1979 dummy  .28  .036  .082
(.27)  (.040)  (.039)
Month of filing: January  ---  ---  ---
February  -.05  -.047  .016
(.41)  (.060)  (.058)
March  -.52  -.130  -.041
(.39)  (.057)  (.055)
April  -.65  -.135  -.069
(.38)  (.056)  (.058)
May  -1.09  -.170  -.135
(.40)  (.059)  (.057)
June  -.48  -.100  -.042
(.40)  (.060)  (.057)
Unemployment rate  .59  .079  .054
(.86)  (.127)  (.122)
a  .798  .018
R-Square  .08  .05
Number of observations = 6,610
benefit taxation is fully perceived so that $p=1$, the negative of the ratio of D·WBA's coefficient to WBA's coefficient should yield a plausible value for $\bar{E}$. This ratio turns out to be .49 with an estimated standard error(13) of .17. Since the combined federal and state marginal tax rate for a high-income claimant might typically be about .4, the implied $\bar{E}$ seems a little high, but it is not off by very much.

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 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

\[------------------------\]

(13) The estimated standard error was computed with the formula in Mood, Graybill, and Boes [19], p. 181, for approximating the variance of the ratio of two random variables.
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 net benefits, but to some other factor. As shown in column 3, when D is entered separately, its coefficient is positive and insignificant, and the estimated D·WBA coefficient rises in magnitude to -.033. 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

<table>
<thead>
<tr>
<th></th>
<th>Coefficient</th>
<th>Standard Error</th>
</tr>
</thead>
<tbody>
<tr>
<td>WBA</td>
<td>.049</td>
<td>.009</td>
</tr>
<tr>
<td>D·WBA</td>
<td>-.024</td>
<td>.007</td>
</tr>
<tr>
<td>WBA X high income</td>
<td>- .004</td>
<td>.019</td>
</tr>
<tr>
<td>D</td>
<td>-.911</td>
<td>.169</td>
</tr>
</tbody>
</table>

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. While the addition of 1 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 .0083. 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 about a .8 week increase in duration, at the upper end of the range from previous studies. The estimated D-WBA coefficient of -.0036 is significantly different from zero at any conventional level, rejecting the hypothesis of no tax effect. The ratio of the coefficient estimates implies a plausible tax rate $\tau$ of .44 (with standard error .14).

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 fully perceive benefit taxation as a reduction in net benefits.
4.2 TREATMENT OF THE TRUNCATION PROBLEM

As mentioned before, weeks of regular benefit collection is a truncated measure of unemployment duration for any claimant that continued to be unemployed after exhausting his benefit entitlement. As Welch [28] has pointed out, this problem is likely to cause regression results to understate the duration impact of benefit changes. Classen [3] and Newton and Roson [21] have used Tobit analysis to deal with the truncation 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 truncation 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 [16] 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

$$ h(y) = ay^{a-1} \exp(-b'y) $$

(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, $a$ is a parameter greater than zero, and $b$ is a vector of parameters associated with $X$. If $a=1$, the Weibull degenerates to the special case of the exponential. If $a>1$, the reemployment hazard rises with duration; if $a<1$, it declines.

The probability density function of completed duration $Y$ is then

$$ f(Y) = ay^{a-1} \exp\{-b'X - Y^a \exp(-b'X)\} $$

(2)
and expected duration is

\[ E(Y) = \exp(b'x/a) \Gamma[(a + 1)/a] \]

where \( \Gamma \) is the gamma function. If we differentiate expected duration with respect to \( x_h \), the \( h \)-th variable in \( X \), we obtain

\[ \frac{3E(Y)}{3x_h} = \left( \frac{b_h}{a} \right) E(Y) \]

so that

\[ \frac{3E(Y)/E(Y)}{3x_h} = \frac{b_h}{a} \quad \text{(3)} \]

Thus, estimates of \( b \) and \( a \) 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

\[ y^* = \begin{cases} y & \text{if } y < p + 1 \\ p + 1 & \text{if } y \geq p + 1 \end{cases} \]

where \( y^* \) is compensated duration plus 1, as in the logarithmic regression, and \( p \) is potential benefit duration. To deal with this upper-limit truncation, we now derive a maximum likelihood estimation technique for the Weibull distribution.

(14) See Johnson and Kotz [14] for a detailed discussion of these and other properties of the Weibull distribution.
bution 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

$$
\text{Prob} \left( Y_{i} > P_{i} + 1 \right) = \int_{P_{i}+1}^{\infty} f(Y) \, dY \\
= \exp\left\{ -Y_{i}^{a} \exp\left( -b'X_{i} \right) \right\} .
$$

Hence, the likelihood function for the full sample is

$$
L = \prod_{i} a Y_{i}^{a-1} \exp\left\{ -b'X_{i} - Y_{i}^{a} \exp\left( -b'X_{i} \right) \right\} \prod_{i} \exp\left\{ -Y_{i}^{a} \exp\left( -b'X_{i} \right) \right\}
$$

where $\prod_{i}$ denotes a product taken over the claimants that did not exhaust their benefits and $\prod_{i}$ denotes a product over those that did. It follows that

$$
\log L = n_{1} \log a + (a - 1) \sum_{i} \log Y_{i}^* - \sum_{i} \left\{ b'X_{i} + Y_{i}^{a} \exp\left( -b'X_{i} \right) \right\} \\
- \sum_{i} Y_{i}^{a} \exp\left( -b'X_{i} \right)
$$

where $n_{1}$ is the number of "nonexhaustees," $\sum_{i}$ denotes a sum over nonexhaustees, and $\sum_{i}$ is a sum over exhaustees.

The parameters $a$ and $b$ 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. (15)

The results are reported in the last column of Table 1. The estimated value of .8 for the parameter a is significantly less than 1 and implies that the reemployment hazard declines with duration. As Heckman and Borjas [12] and Lancaster [16] 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 .0076 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 a to obtain the proportional change in expected duration associated with a unit change in untaxed

(15) The maximum likelihood estimation was performed with the Davidson-Fletcher-Powell algorithm in the GQOPT numerical optimization package. The algorithm is discussed in Quandt [22]. The algorithm converged to the same final parameter estimates when started from different initial values.
WBA. The result implies a proportional change of .0095, which should be compared to the least-squares WBA coefficient of .0083 in column 2 of Table 1. As expected, adjustment for truncation of the duration variable results in a larger estimated impact of WBA.

The estimated D·WBA coefficient of -.0024 also is significantly different from zero at conventional levels, again rejecting the hypothesis of no tax effect. Dividing the coefficient by the estimate of a indicates that the proportional duration effect of a dollar change in WBA is reduced by .0030 if benefits are taxed. If claimants fully perceive benefit taxation as a benefit reduction, the ratio of the D·WBA and WBA coefficients implies a tax rate of .32 (with standard error .15), still fairly close to the expected rate.

These results should be taken with at least one grain of salt. Like Tobit analysis, the procedure in this section assumes that the censored observations continue to obey the assumed distribution beyond the point of truncation. This assumption is questionable because follow-up studies of benefit exhaustees have sometimes found an upsurge in reemployment immediately after benefit exhaustion. (16) To put the point another way, this section’s model may overstate the duration impact of benefit level because the assumed hazard function in equation (1) implies that the proportional ef-

(16) See Marston [18], for example.
fect of benefit level on reemployment probability stays the same after the truncation point. This seems extremely unlikely—surely, the effect of benefit level diminishes once the benefits are no longer received. Consequently, while regression results probably tend to understate the full duration impact of benefits, the method to adjust for truncation probably overstates it. More accurate estimation of the total duration impact would require information on length of unemployment following benefit exhaustion.

A less ambitious task, but still a useful one, is to estimate the effect of benefit taxation on mean compensated duration. The expected value of $Y^*$ is

$$E(Y^*) = \text{Prob}(Y < P + 1) E(Y|Y < P + 1) + \text{Prob}(Y \geq P + 1) (P + 1).$$

As shown in the appendix, this expectation can be written as

$$E(Y^*) = \exp(b'X/a) \Gamma(\{a+1\}/a) + \exp\{-\exp(-b'X)\} (P+1)$$

where the subscripted $\Gamma$ term is an incomplete gamma function. (17) This expectation can be estimated for each member of the sample by substituting in his observed $X$ vector, his potential benefit duration $P$, and the maximum likelihood estimates of the parameters a and b.

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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 $D=W_B=0$ 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 9.9 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 11.3 weeks. The implied average effect of benefit taxation on the high-income 1979 claimants is therefore a 1.4 week reduction in their compensated duration.

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 rejected the hypothesis of no tax effect. The 1979 imposition of benefit taxation appears to have reduced average compensated unemployment duration among the sampled high-income claimants by about a week and a half.
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, $\text{DUR}_0$ be his unemployment duration without benefit taxation, and $\Delta \text{DUR}$ be his duration change due to benefit taxation, then the full change in his tax payment induced by benefit taxation is

$$ \Delta \tau = \bar{t} \{ \text{DUR}_0 \cdot \text{WBA} - \Delta \text{DUR} \cdot (w - \text{WBA}) \}.$$ 

The first term in brackets is the benefit income the claimant would have collected in the absence 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, $\text{DUR}_0 \cdot \text{WBA}$ averaged $953$ while the second term, $- \Delta \text{DUR} \cdot (w - \text{WBA})$, averaged $302$. (18) Therefore, a projection

(18) This computation uses $1/13$ of high-quarter earnings in

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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 $\Delta T$ is correspondingly reduced. The empirical evidence on whether UI-induced duration changes are indeed accompanied by wage-rate changes is ambiguous. (19)

Another budgetary implication of benefit taxation’s work incentive effect is the impact on UI program costs. These costs are reduced by $ADUR \cdot WBA$, 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 $120, a 13% reduction from the $953 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. In the case of the base period as an estimate of the weekly wage $w$.

(19) See the studies by Classen [3], Ehrenberg and Oaxaca [5], and Holen [13], and the critical review by Welch [28].
1979 policy change, this distributional effect is not too disturbing because only high-income families suffered a welfare loss. The distributional consequences of extending benefit taxation to lower-income claimants, however, might be less appealing.
APPENDIX: DERIVATION OF THE EXPECTED VALUE OF A RANDOM VARIABLE WITH A CENSORED WEIBULL DISTRIBUTION

Let the random variable $Y$ follow a Weibull distribution with shape parameter $a$ and scale parameter $r$. Then, as given in Johnson and Kotz (14), $Y$ has the probability density function

$$f(Y) = a r^{-a} Y^{a-1} \exp[-(Y/r)^a]$$ \hfill (A1)

and the cumulative distribution function

$$F(Y) = 1 - \exp[-(Y/r)^a] \quad .$$ \hfill (A2)

Now suppose $Y*$ has the censored Weibull distribution described by

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

Then, the expected value of $Y*$ is

$$E(Y^*) = \text{Prob}(Y < B) \ E(Y \mid Y < B) + \text{Prob}(Y \geq B) \ B$$
$$= F(B) \ E(Y \mid Y < B) + \{1 - F(B)\} \ B \quad .$$ \hfill (A3)

The conditional mean is in the first term is

$$E(Y \mid Y < B) = \frac{1}{F(B)} \int_0^B y f(Y) \, dy \quad .$$

Multiplying both sides of this equation by $F(B)$ and substituting the expression in (A1) for $f(Y)$ yields

$$F(B) \ E(Y \mid Y < B) = \int_0^B a (Y/r)^{-a} \exp[-(Y/r)^a] \, dy \quad .$$

20. Similar results on conditional means for the Weibull distribution are derived in Aroian (1).
If we let $Z = (Y/r)^a$, we can rewrite the above equation as

$$P(B) \ E(Y|Y<B) = \int_0^{(B/r)^a} z^{1/a} \exp(-z) \, dz$$

$$= r \int_0^{(B/r)^a} z^{((a+1)/a)-1} \exp(-z) \, dz$$

$$= r \frac{\Gamma((a+1)/a)}{(B/r)^a}$$

(A4)

where $\frac{\Gamma((a+1)/a)}{(B/r)^a}$ is the incomplete gamma function $21^/$ with argument $(a+1)/a$ and upper limit $(B/r)^a$.

Substituting the expressions in (A2) and (A4) into equation (A3) gives

$$E(Y^*) = r \frac{\Gamma((a+1)/a)}{(B/r)^a} + \exp\{-((B/r)^a)\} \ B \ .$$

In the text, the specification of the reemployment hazard function in equation (1) implies a scale parameter of $r = \exp(b'X/a)$. Also, the truncation point is potential benefit duration $P$ plus 1. Hence,

$$E(Y^*) = \exp(b'X/a) \frac{\Gamma((a+1)/a)}{(P+1)^a} \exp(-b'X)$$

$$+ \exp\{-(P+1)^a \exp(-b'X)\} (P+1) \ .$$

21. See Bennett and Franklin (2) for a discussion of the incomplete gamma function.
REFERENCES


