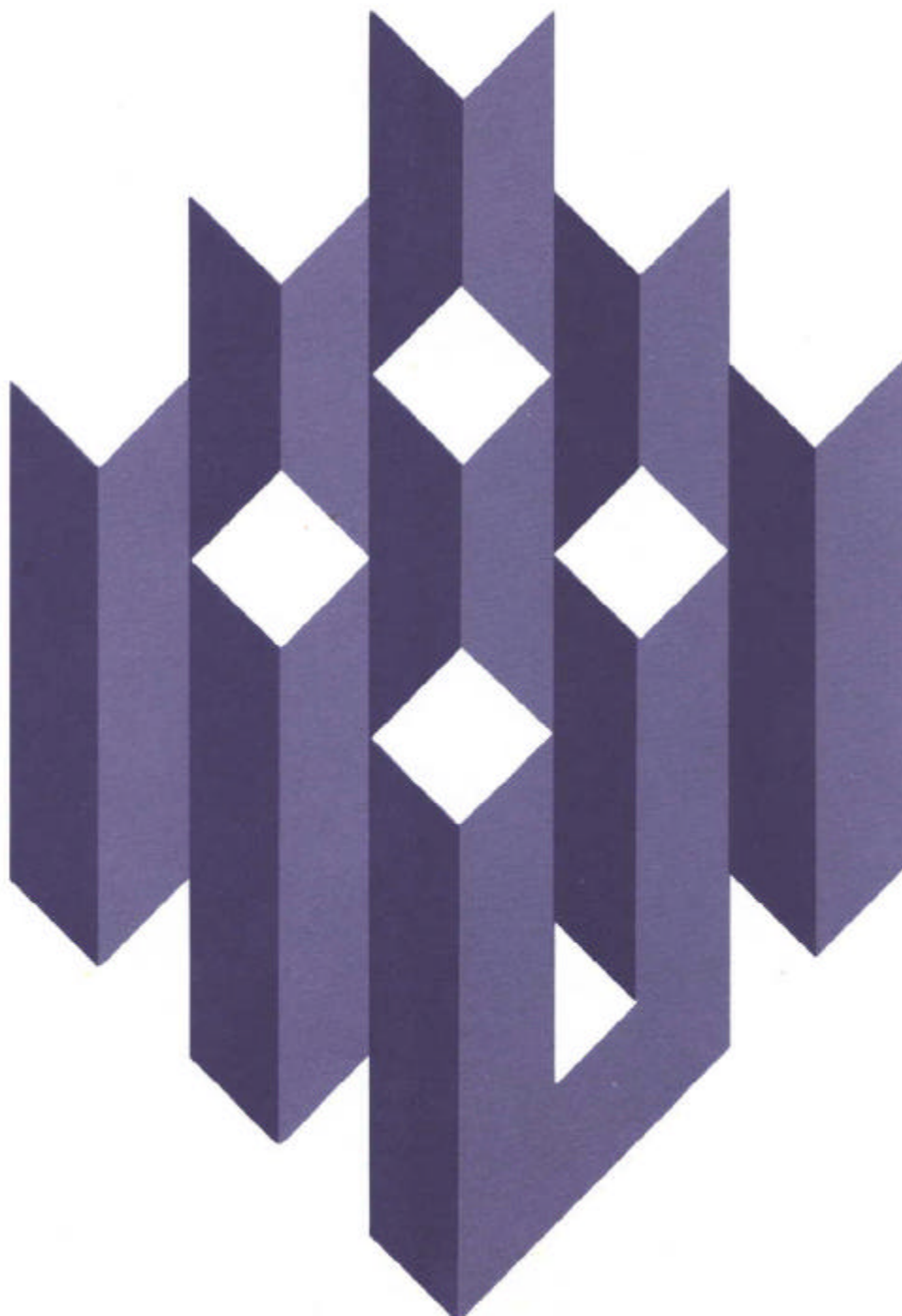# Effect of Unemployment Insurance on Duration of Unemployment: A Study Based on CWBH Data for Florida

Effect of Unemployment Insurance on Duration of Unemployment: A Study Based on CWBH Data for Florida

ACKNOWLEDGEMENTS

ABSTRACT

Previous studies of unemployment insurance have obtained estimates for a reduced form model consisting of two equations: the duration equation and the post unemployment earnings equation. These equations do not, however, allow us to estimate the effects of unemployment insurance accurately. The observed wage distribution is a truncated distribution – it is the wage offer distribution truncated by reservation wages. As such, unless the truncation bias is taken into account, we get misleading results. Also the duration equation itself is derived from more basic behavioral relationships.

To study the effects of unemployment insurance on the duration of unemployment, one needs to go into the basic underlying behavioral equations. In this report we assume that the behavioral equations are: the wage offer equation and the reservation wage equation. An individual is assumed to accept employment if the wage offer is greater than or equal to the reservation wage. The unemployment insurance variables impact on the formation of reservation wages. Past unemployment, of course, has an impact on both wage offers and reservation wages. A stochastic threshold regression model is used to estimate the parameters of the wage offer equation and the reservation wage equation. This model enables us to estimate parameters of both these equations based on data on individual's personal characteristics, employment status, and wages if employed. (See Appendix for an elementary exposition of the model.) The data for this study are taken from the CWBH files for the State of Florida. These data provide a longitudinal summary of individual employment histories.

Once the parameters of the reservation wage equation are estimated, we can get estimates of reservation wages for each individual. These, in conjunction with the estimates of wage offer distribution enable us to estimate the expected duration of unemployment. (This is described in Chapter VI.)

The main aspects of our analysis are:

(i)   Use of individual employment histories based on longitudinal data for 4 years - rather than just a cross-sectional view of the individuals.

(ii)   Use of a structural model of labor market behaviour - rather than a reduced form equation.

(iii)   Estimation of the effects of duration of unemployment using a finite time-horizon model - rather than an infinite horizon model.

In fact, in the model we used, the time horizon of job search is actually estimated (it was found to be 40.87 weeks). Almost all the studies done till now assume an infinite time horizon for job search. We show in this paper that the results are quite different depending on the horizon assumed for job search. For instance, if potential weekly benefit amount is increased by $10.00 the increase in the expected duration of unemployment is estimated as 2.78 weeks in the infinite horizon case and 1.37 weeks in the finite horizon case. Similar results are obtained in changes in potential benefit weeks. The effect of raising potential benefit weeks by one week is to increase the duration of unemployment by 1.4 weeks in the infinite horizon case and 0.8 in the finite horizon case. The study thus suggests that some of the high estimated effects on the duration of unemployment are due to the infinite horizon for job search that is (implicitly) assumed.

Finally, there are many data problems we encountered in our study. These are discussed at length in Chapter IV. Though the new CWBH data tapes are better, our discussion would be of help to others analyzing the CWBH data tapes.

TABLE OF CONTENTS

x

CHAPTER I

INTRODUCTION

The unemployment insurance system originated with the Social
Security Act of 1935. It was designed as a joint federal and state program
to compensate covered workers during spells of unemployment. These
payments have grown substantially over the years. In 1939 benefits were
one-half billion dollars and in 1975 they were nearly 17 billion dollars.
The latter figure represents nearly 1.5 percent of U. S. disposable income
and over 13 percent of all governmental transfer payments. This growth
has caused increasing concern among economists and politicians as to the
effects of these benefits on the unemployed, employed, and the economy
as a whole.

The major thrust of this concern has been directed towards the
duration of unemployment and post-unemployment earnings. It has been
argued, both theoretically and empirically, that unemployment compensation
creates incentives for the jobless to remain unemployed longer than they
would in the absence of these payments. If this is true, then the unemploy-
ment insurance program has effectively acted to increase the duration of
unemployment. This effect may not be economically significant if the
additional time in unemployment is spent in productive job search. Produc-
tive would mean better employer-employee job matches. In most studies the
measure of productivity used has been post-unemployment earnings. The
policy issue then is to determine to what extent these propositions describe

labor market behavior and to determine the magnitude of unemployment insurance effects on labor market activity.

In this report we will present some answers to these questions based on our analysis of the CWBH data for the State of Florida. These data cover the period 1970-75. The CWBH data have several deficiencies, some of which have been corrected in the new CWBH data sets now being collected. For instance, in the data set we used, the human capital characteristics of individuals (education, experience, etc.) were not available and we had to use proxies for them. The new CWBH data contain such information. Further, there is now greater uniformity among states in the form in which data are collected under the CWBH program. Unfortunately, the State of Florida does not participate in the new data collection program and hence we could not extend our analysis to cover a longer period or to make a comparative analysis.

Even with all the limitations of the CWBH data files we used (details of which can be found in Chapter IV), we feel that the analysis we present in Chapters V and VI is worthwhile for the following reasons:

(i)   We construct a structural model of unemployment and estimate the effects of increased benefits on duration of unemployment on the basis of the estimates from this structural model and theoretical constructs based on search theory, rather than by the estimation of a reduced form equation.

(ii)   We derive the estimates by assuming (as usual) an infinite horizon and, more realistically, a finite horizon for job search and show how the assumption of a finite horizon for job search is important. We also derive estimates of the time horizon of job search.

(iii)   We use time-series and cross-section data incorporating the

employment histories of individuals.

The methodology in our study would be useful for others working with

the CWBH data. The CWBH data are being collected with the idea that

several questions relating to the UI program can be studied. Some of these

questions pertain to the adequacy of benefit duration and benefit amounts

and the effects of unemployment insurance on job search behaviour of

claimants. The present study gives an idea as to what extent these questions

can be answered. Specifically, we might ask the following questions:

Q(1)   What is the effect of increases in benefit duration and/or benefit

amounts on the duration of unemployment? Here, the answer

theoretically is unambiguous. It increases the duration. The

empirical question is: By how much? Often, the answers to

these questions have been obtained by the estimation of some

reduced form equations, and also assuming an infinite time horizon

for job search (See Chapter III for a review of the studies).

We obtain these estimates from a structural model and assuming

a finite time horizon for job search.

Q(2)   What is the effect of increases in benefit duration and/or benefit

amounts on post re-employment wages relative to:

(a) post re-employment wages without the increases and

(b) pre un-employment wages.

Again, the answer to question (a) is theoretically unambiguous, i.e, there

will be an increase. As for the answer to question (b), it is theoretically

ambiguous if reservation wages fall over time (which is a reasonable

assumption). Unfortunately, question (b) cannot be answered from the data

3

that are available though several studies have indeed tried to answer this question (See Chapter III for a review) and arrived at unjustified conclusions that the effect of increases in unemployment insurance benefits on post-re-employment wages is negligible and hence that unemployment insurance does not result in productive job search.

The invalidity of such conclusions can be seen from the following hypothetical example: Consider two individuals A and B. A obtains a job that gives higher post-re-employment wages than pre-unemployment wages, but he loses his job soon after because he is unsuited for it. B obtains a job that may not give him higher pre-re-employment wages but can keep the job for a long time. He has spent time in searching for a job suited to his tastes and abilities. In this case B has engaged in productive job search rather than A but a comparison of just post re-employment wages with pre-unemployment wages does not reveal this. We also have to look at the duration of subsequent employment.

Since question 2(b) is an unanswerable question based on the data we have, we decided not to give any spurious answers to this question. Answers to question 2(a) are not very interesting from the policy point of view and hence we did not attempt to answer it, though it can be obtained from the structural model we have estimated. Thus, only answers to question (1) are given and these are presented in Chapter VI.

The scheme of the following chapters is as follows:

---- In Chapter II we provide an overview of the unemployment insurance program in Florida.

---- In Chapter III we give a survey of the empirical literature on the effects of unemployment insurance. This survey forms a background for our

4

criticism of the methods used by others, as well as a comparison with the empirical results we obtain.

---- In Chapter IV we discuss in detail the data files we have used as well as their limitations. This detailed discussion would be useful for those making use of the new CWBH data tapes and encourage them to look for similar data anomalies.

---- In Chapter V we present the empirical results from the structural model we have suggested and estimated.

---- In Chapter VI we estimate the effects of changes in different policies on the duration of unemployment - both in the case of finite time horizon and infinite time-horizon for job search.

The final chapter presents the conclusions of the study. The technical details underlying the models used are explained in the Appendix.

CHAPTER II

OVERVIEW OF THE UNEMPLOYMENT INSURANCE SYSTEM

Structural Framework in Florida

The unemployment insurance system in Florida has been in existence since 1938 and continues to operate today, albeit on a much broader scale. Over the years, there have been many changes in the structure of the unemployment insurance system. These changes have occurred as a result of amendments to, or enactment of, Federal or State laws. In Florida, as well as in the other States, the changes have been made by the State legislature because of changes in the Federal law. The most important changes in provisions that are not subject to Federal law are changes in the duration of benefits and the weekly benefit amount. The only recent major change made at the federal level has been the creation of an extended benefit program which increases the maximum allowable benefits paid during unemployment.

Under the extended benefit program, the United States Secretary of Labor determines whether the statewide or national unemployment situation is severe enough to warrant the extention of benefits beyond the state's current maximum. An extended benefit period begins with the third week after either a national or state "on" indicator has been declared. The extended benefit period will last for at least 13 weeks, and will terminate if both a national and state "off" indicator is declared but not before the end of 13 weeks. As of July, 1975, a national "on" indicator is declared whenever

the rate of insured unemployment (seasonally adjusted) for all states

exceeded 4.5% for three consecutive months. In Florida, a state "on"

indicator is declared if both of the following occur: (a) the rate of

insured unemployment (not seasonally adjusted) for the preceding 13 weeks

exceeded 4.0%, and (b) the rate of insured unemployment (not seasonally

adjusted) equaled or exceeded 120% of the average of such rates for the

corresponding 13-week period ending in each of the preceding two years.

These extended benefit periods may last longer than 13 weeks so long as

the trigger goes on. However, an individual claimant can receive a total

of 39 weeks of regular and Federal-State extended benefits combined.

During the recessionary period 1975-76 some claimants were eligible for

up to 65 weeks of benefits under a special emergency program that was

enacted at the time, and that has now expired.

Generally, an individual is eligible for unemployment insurance

benefits in the State of Florida if (a) a formal claim has been made in

accordance with the rules of the unemployment insurance commission;

(b) the worker has been registered for unemployment compensation at a

claims office; (c) the individual is available for work and able to engage

in work activities; (d) the individual has been unemployed for a waiting

period of one week; (e) the individual has been paid wages for insured

work equal to 20 times his average weekly wages during his base period,

provided that his average weekly wage is at least $20. The base period is

defined in Florida as the first four of the last five completed calendar

quarters immediately preceding the first day of an individual's benefit

year. The amount of weekly compensation an individual is eligible for is

computed as one-half the average weekly wage in the base period, but not less than $10, nor exceeding the current maximum. This is called the weekly benefit amount. Each individual who is eligible for the program and is totally unemployed in any week is allowed to receive a payment equal to his weekly benefit amount for that week, while an individual who is partially unemployed for a given week can receive an amount equal to the weekly benefit amount minus that part of the wages payable to him for that week.

An individual applying for unemployment compensation begins his benefit year on the day in which the claim was filed. It ends exactly one year later. Many spells of unemployment may occur during this benefit year. In each spell, the individual is allowed to collect weekly benefits, beginning after a one-week waiting period, provided he has not exhausted his potential benefits. An individual has exhausted his benefits with respect to any week of unemployment in his benefit year if he has received all of the benefits that were available to him by Florida law, or if his benefit year has expired prior to that week and he does not satisfy the eligibility requirements upon reapplying for compensation.

Disqualification for benefits, under Florida law, may occur when an individual has voluntarily left his job, was discharged for misconduct connected to the work, failed to apply for available suitable work, or failed to accept suitable work when offered to him. If the individual was disqualified for misconduct connected to the work, then he may be ineligible for unemployment compensation for up to 12 weeks after the discharge and until he has earned 10 times his weekly benefit amount. The specific time period is determined for each case, individually. If the individual

fails to work or accept suitable employment, then he may be disqualified for up to five weeks and until he has earned 10 times his weekly benefit amount, his potential duration of benefits may be reduced by up to three weeks. Again, the exact penalty is determined by the local unemployment insurance office on a case by case basis. An individual who has been disqualified for the full period of unemployment must become reemployed and earn wages equal to or greater than 10 times his previous weekly benefit amount to regain eligibility.

In Florida, when a change in the maximum weekly benefit amount is affectuated, all of those claimants beginning a new benefit year would be eligible to receive benefits up to the new maximum. All of those who began their benefit year prior to the effective date would earn benefits subject to' the previous maximum. Therefore, at any point in time, there could be different maximum weekly benefit amounts for different claimants. The changes that have occurred in this variable are as follows:

Effective Date	Maximum
7/1/67	$40
1/1/71	47
7/1/71	54
7/1/72	64
1/1/74	74
7/1/75	82

As can be seen, there have been substantial increases in the maximum weekly benefit amount in Florida.

In addition to the changes in the maximum weekly benefit amount, the Florida legislature has greatly extended the coverage of the

unemployment insurance system in recent years. Since January 1956, all

nonfarm, profit employers with four or more employees payrolled for 20

weeks or more in a calendar year were required to provide coverage. Also,

any employer with four or more employees for eight or more weeks and

who paid at least $6000 in wages in any quarter were required to provide

coverage. Effective January 1, 1972, the unemployment insurance laws were

amended and coverage was expanded to include:

a) Nonfarm, profit employers with one or more employees payrolled for 20 weeks or more in a calendar year. Also, employers that paid $1500 or more in wages in any quarter were required to provide coverage.

b) State government employees.

c) All nonprofit employers with four or more employees, except churches, nonprofit schools and universities, and rehabilitative organizations, were required to provide coverage.

The principal individuals not covered by unemployment insurance are

farm workers and anyone working for commissions only. The coverage was

extended in January 1974, to include local government employees. Also,

up until 1975, an individual receiving social security benefits was not

eligible to receive unemployment insurance compensation. This law was

revised twice and in its present form these individuals may receive

benfits provided they meet the minimum requirements, although these

benefits may be reduced if the individual is receiving a pension from a

base period employer.

Florida Compared With Other States

In comparison with most other states Florida's eligibility require-

ments are slightly more lenient than average, although regaining eligibility

when disqualified is more difficult. Florida requires that the claimant earn at least $400 and work at least 20 weeks in the base period to meet the minimum requirements. A wage of about $650 and at least 20 weeks of employment in the base period is a good approximation to the minimum requirements of the other states. Where Florida appears to be more parsimonious is in the level of benefits paid to claimants. But, oddly enough, Florida has a higher than average ceiling on employer taxes. These data are summarized in Table 2-1. There are eight states that do not require a waiting week before compensation can be paid. In addition, there are all states that provide for compensation during the waiting period under certain restrictions. Florida is not one of these states and requires a one week waiting period.

All things considered, Florida might be regarded as an average state when it comes to its unemployment compensation system. This is because the wage differentials across states tend to make Florida look more restrictive and the other states more lenient in the aggregate.

The main characteristics of the Florida unemployment insurance system reviewed in this chapter pertain to the period covered by our analysis. They need not describe the conditions of the system as they currently exist.

TABLE 2-1

COMPARISON OF UNEMPLOYMENT INSURANCE
PROVISIONS

Provisions	Florida		All Other States		All Other States (Average)	
	Min.	Max.	Min.	Max.	Min.	Max.
Taxes	0.1	4.5	0	6.6	.75	3.76
Potential Duration	10	26	1	36	13.6	26.8
Weekly Benefits ($)	10	74	5	127	14.7	82.5

Source: U. S. Department of Labor, Significant Provisions of State
Unemployment Insurance Laws, January 6, 1975, Manpower
Administration, Unemployment Insurance Service.

CHAPTER III

EFFECTS OF UNEMPLOYMENT INSURANCE:
A SURVEY OF THE EMPIRICAL LITERATURE

A number of studies have been conducted in the unemployment
insurance area. They have tried to explore the theoretical conclusions
of the job search literature and to describe the multitudinous nature of
nonworking activities. Each study has some proportion of virtues and vices,
and these characteristics will be revealed whenever possible. This
literature will be reviewed in chronological order; however, deviations
from this sequence will be entertained if this facilitates a more
contiguous understanding. There will be at least four points examined in
each paper: (1) Is the hypothesis consistent with the theoretical
literature, (2) Does the model describe the hypothesis accurately, (3) Will
the data facilitate the methodology described, (4) Are the results
meaningful in terms of the hypothesis tested.

Gene Chapin [1971] explores "the effects of unemployment benefits
on the incentive to seek work rather than the incentive to work." The
key word in context is 'seek'. This encompasses the question of what
non-working activities will predominate the spell of unemployment. Put
specifically, Chapin is testing the hypothesis that the duration of
unemployment is determined by the probability of finding employment, the
level of unemployment insurance benefits, and the maximum allowable
compensated duration. His regression form is linear with different proxies
used for each of the above:

$$D = \alpha + \beta U + \gamma B/W + \delta M + u$$

Where D is the average duration of unemployment insurance claims, U is the insured unemployment rate, B/W is the average benefit divided by the average wage from previous covered employment, and M is the maximum available compensated duration. This regression equation was estimated from cross-sectional state data for each year from 1962 to 1967. The regression equation was also estimated using pooled data, but a test for structural shifts was not performed. These regression results are presented in Table 3-1.

The most obvious feature of these results is the significance of the coefficient of U. This is to be expected given the statewide nature of the data. The insured unemployment rate is a cross-sectional sample in time that has peculiar problems. These have been discussed in Kaitz [1970], Salant [1977], and Frank [1978]. Basically, there is a counting problem with this variable: Spells of longer duration have a higher probability of being surveyed than those of shorter duration. Hence, the overall rate of unemployment is underestimated, but more importantly, U tends to vary directly with the average duration of unemployment in each state. Therefore, it is not very surprising that the estimates of β are positive and significant in all of the regressions. A second point of note, and a curious one indeed, is that the coefficients of M are all significant. This is a puzzle because M does not vary appreciably across states and this would tend to increase the standard error of the estimate. One possible explanation may be that states which tend to deviate from the average, do so in an upward direction because of some irregular labor force character-istics [1]. A chronically higher incidence of unemployment might lead

14

TABLE 3-1

CHAPIN'S RESULTS

ANNUAL:	$\hat{\alpha}$	$\hat{\beta}$	$\hat{\gamma}$	$\hat{\delta}$	R^2
1962	.462	.788 (5.12)	.041 (0.79)	.122 (2.00)	.37
1963	.194	.929 (4.06)	.088 (1.72)	.193 (3.22)	.34
1964	.219	.589 (4.33)	.049 (1.17)	.240 (0.48)	.45
1965	.199	.755 (5.13)	.070 (1.63)	.224 (4.07)	.41
1966	.196	.832 (5.82)	.030 (0.07)	.224 (4.15)	.49
1967	.933	.892 (5.22)	.076 (1.65)	.269 (4.47)	.44
POOLED:	.313	.841 (15.8)	.056 (3.11)	.153 (5.46)	.49

Source: Chapin [1971].

Note: t-values are given in parentheses.

15

officials to increase the maximum duration [2]. If this is the case, then the positive relationship follows by tautology. Another observation about Chapin's variables is that B/W is probably the most relevant variable so far as economic theory is concerned, acting as a proxy for the opportunity cost of unemployment, but is the least significant in terms of these results. This is undoubtedly caused by the anomalous nature of the other two "independent" variables. As a historical note, if B/W was increased by one percent, then D would increase by approximately one-twentieth of a week [3].

Kathleen Classen [1977a, 1977b] utilizes a micro-data set obtained from the Continuous Wage and Benefit History (CWBH) program for two states: Arizona and Pennsylvania. These data cover insured workers for two years, 1967 and 1968. This two-year period was chosen because an increase in benefit entitlement occurred in 1968 for each of these states. This provides for an increase in the level of benefits that is independent of previous earnings and seasonality. Classen uses this structural change to test the effects unemployment insurance has on the duration of unemployment and post-unemployment earnings.

The duration of unemployment equations are of the following general form:

$$D = f(WBA, MAX\ WEEKS, Z)$$

Where D is the weeks of compensated unemployment, WBA is the weekly benefit amount paid to the claimant, MAX WEEKS is the potential duration of benefits, and Z is a vector of socioeconomic and labor market characteristics. This vector includes age, sex, occupation, industry and cyclical dummies, and high

16

quarter earnings. These last two variables represent a measure of the

level and stability of previous earnings, respectively.

The CWBH data presented a problem in the estimation of the duration

equation. The dependent variable is truncated with respect to the actual

duration of unemployment, and it is this latter variable that is of policy

interest. To circumvent this problem, Classen uses the Maximum Likelihood

technique of TOBIT analysis. A caveat is in order here; whereas the TOBIT

method is applicable to truncated dependent variable equations, it may

vitiate the estimates if the dependent variable is dimensionless. In this

case, there is good reason to believe that Classen might have this problem.

The CWBH data provide information on filing and separation dates. These

are the dates when an unemployment claim was recorded and employment was

terminated, respectively. If these two are not equal or they are not a

constant multiple of each other for all individuals, then the variable

D may lie anywhere on the actual duration of unemployment line. If this

is the case, then it is not clear what effects unemployment insurance

parameters will have on the actual duration of unemployment. Whether

Classen's data have this problem is unknown. Nonetheless, she has

estimated the duration equations using TOBIT for a linear form, and OLS

for both linear and double logarithmic forms. The results for Pennsylvania

are presented in Table 3-2. The Arizona results are similar in sign and

magnitude for the key variables and thus need not be presented.

The only noticeable difference between the OLS and TOBIT estimates

is that the MAX WEEKS coefficient is much lower for the latter technique.

This may be accounted for by the fact that the total duration of

unemployment is no longer constrained and because the TOBIT technique

17

TABLE 3-2

CLASSEN'S DURATION RESULTS

Independent Variables	Linear OLS		LN-LN OLS		TOBIT	
WBA	.11	(5.68)	1.03	(8.41)	.12	(5.79)
Max Weeks	.19	(2.99)	1.09	(4.61)	.07	(0.94)
Sex	.32	(0.88)	.01	(6.28)	.44	(1.11)
Age	.10	(10.69)	.46	(8.82)	.12	(11.10)
Prev. Earnings	-.002	(-5.94)	- .60	(-7.82)	-.002	(-5.76)
Earnings Stability	-3.07	(-10.83)	-1.32	(-12.36)	-3.13	(-10.05)
Industry Dummies:						
Agriculture	1.55	(1.05)	.19	(0.94)	1.66	(1.02)
Mining	-2.19	(-1.99)	- .25	(-1.68)	-2.45	(-2.04)
Construction	1.03	(1.92)	.26	(3.61)	.94	(1.61)
Non-Durables	-2.68	(-6.45)	- .33	(-5.83)	-2.97	(-6.55)
Durables	- .53	(-1.18)	- .09	(-1.55)	- .61	(-1.25)
Transport.	- .79	(-.92)	- .14	(-1.23)	-1.00	(-1.08)
Finance	3.50	(3.29)	.31	(2.16)	4.14	(3.51)
Services	2.60	(4.18)	.31	(3.61)	3.08	(4.49)
Year (1968=1)	- .32	(-.58)	1.07	(2.15)	- .36	(-.60)
Year*Prev. Earn	- .000	(-.70)	- .16	(2.34)	- .000	(-.77)
Constant	5.76		-1.97		8.94	

$$R^2 = .09 \qquad R^2 = .10 \qquad \chi^2 = 2534$$

$$F = 27.8 \qquad F = 29.3$$

Source: Classen [1977b].

Note: Dependent variable equals weeks of compensated unemployment, SEX = 1 for females, and figures in parentheses are ordinary t-values for OLS and asymptotic t-values for TOBIT.

18

recognizes the functional relationship between the dependent variable and
MAX WEEKS. All of these results tend to support the hypothesis that an
increase in potential benefits and/or an increase in WBA will cause a
significant increase in the duration of unemployment.

The post-unemployment earnings equations were estimated using high
quarter earnings in the year following unemployment. Claimants who had no
earnings record in this year were excluded from these regressions. This
introduces a selection bias in the estimates. The importance of this bias
can be tested, but it has not been done in this paper. The independent
variables in these equations are the same as those in the duration equations.
Table 3-3 contains the results of these regressions for Pennsylvania. The
results differ widely between the two functional forms for unemployment
insurance variables, and the significance level of these variables is
relatively low.

In an attempt to improve upon these results, Classen has used the
recommendations of Welch [1977] to estimate the duration and subsequent
earnings equations as a spline function. For the duration equation she uses
a spline function to provide variation in the WBA that is independent
of previous earnings within a single year [4]. These results show a
negative relationship between duration and previous earnings when WBA stops
increasing with previous earnings. This relationship becomes positive when
the WBA and previous earnings increase together. Thus, what could have been
a nonlinear relationship in various directions tends to indicate that the
WBA has a positive effect upon the duration of unemployment.

19

TABLE 3-3

CLASSEN'S POST-UNEMPLOYMENT EARNINGS RESULTS

Independent Variables	Linear OLS		LN-LN OLS	
WBA	1.67	(1.28)	-.10	(-1.59)
Max. Weeks	-8.28	(-1.83)	-.13	(-1.03)
Sex (F=1)	-216.94	(-8.47)	-.13	(-4.70)
Age	-3.15	(-4.49)	-.09	(-3.31)
Prev. Earnings	.66	(31.60)	.77	(17.93)
Earnings Stab.	81.35	(4.15)	.17	(3.02)
Industry Dummies				
Agriculture	92.85	(.88)	.22	(1.97)
Mining	182.0	(2.43)	.16	(2.0)
Construction	273.87	(7.40)	.19	(4.95)
Non-Durables	39.89	(1.38)	.13	(4.40)
Trans.	277.84	(4.73)	.19	(3.13)
Finance	-89.39	(-1.12)	.05	(.63)
Services	-89.78	(-2.02)	-.02	(-.42)
Durables	141.44	(4.55)	.12	(3.70)
CYC. Dummy	-4.86	(-.13)	-.42	(-1.57)
CYC-P.E. Interaction	.05	(2.13)	.07	(1.80)
Constant	620.317		2.52	
	$R^2 = .60$		$R^2 = .39$	

Source: Classen [1977b]

Note: Dependent variable equals earnings in the best quarter of the year following unemployment and t-values are given in parentheses.

20

The spline function is used in the subsequent earnings equations in an effort to get information on the relationship between previous and post-unemployment earnings, while separating out WBA effects. Except for the dependent variable, the spline function has the same specification as that used for the duration equations. The resulting estimates tend to be inconclusive about the effects of WBA on subsequent earnings.

Steven Marston [1975] conducts a study comparing the duration of unemployment between the insured and uninsured workers. He is seeking to resolve three problems that have plagued previous studies: (1) The data on the duration of compensated unemployment obtained from unemployment insurance files often includes more than one spell of unemployment, (2) There is a sample selection problem in the Bureau of Labor Statistics figures on the duration of unemployment, (3) The behavior of the insured and uninsured unemployed may differ for reasons other than unemployment insurace, i.e., demographic and mobility factors.

First, Marston corrects for the sample measurement problem using a procedure similar to that found in Salant [1977]. This procedure gives an estimate of the expected duration of total unemployment. Based upon the 1969 Current Population Survey this estimate was 5.0 weeks with a standard error of 0.0063 weeks.

Second, an estimate of the expected duration of a single spell of insured unemployment is calculated using data from unemployment insurance administrative files for Detroit, during the period 1966-71. The procedure used to develop this estimate is somewhat cumbersome, but the essential ingredient is an equation designed to estimate the unemployment continuation rates. This equation has an unrestricted nonlinear form, but its estimates

21

remain within the acceptable bounds for a probability. These results imply that the probability of finding employment decreases with the duration of unemployment. Using these figures Marson controls for the demographic element by deriving expected duration estimates for twelve groups: male and female, and six age groups. From each of these divisions an estimate of the expected duration of insured unemployment is calculated. This overall estimate for the Detroit area is 5.62 weeks with a standard error of 0.314 weeks.

Finally, the estimates of the total and insured expected duration are used in a weighting scheme to calculate the expected duration of uninsured unemployment.

The main criticism leveled at this study is that it uses both national and local statistics to make nationwide inferences. This would be somewhat acceptable if Detroit was 'typical', but the employment cycle there is tied directly to automobile production. As such, the pattern of job accession and continuation rates are largely employer determined. The unemployed labor force will be composed mostly of temporary layoffs and their job search behavior may deviate from the national norm [5].

In a study on post-unemployment earnings, Burgess and Kingston [1976] have found a positive relationship between unemployment insurance benefits and subsequent earnings. Using data from the Service To Claimants (STC) program, they have studied workers who return to work before potential benefits are exhausted. They have estimated that a $1 increase in WBA increases subsequent earnings by $25. This is an unusual result compared with the ambiguity of Classen's and requires further examination.

The dependent variable in their regressions is post-unemployment minus pre-unemployment earnings. This turns out to be a meaningless

22

distinction because pre-unemployment earnings is also an independent variable. Thus, these regressions might just as well be interpreted as if post-unemployment earnings was the dependent variable. What is meaningful here is how these earnings are obtained from the data. It appears as if they are calculated based upon social security contributions, and therefore, they have biased their sample towards low wage earners [6]. This sample censoring would negatively bias their estimates of unemployment insurance effects.

A slight reduction in the magnitude of the WBA estimate can be derived by using the total effect of a change in WBA on subsequent earnings. This total effect would include the change in duration due to a change in WBA. Using Classen's estimate of 0.14 for $\partial D/\partial WBA$, the $1 increase in WBA will increase subsequent annual earnings by $18.

Another censoring problem that effects the interpretation of these estimates is the method in which the sample was selected. They have chosen only those covered workers who have not exhausted their benefits. This probably introduces an upward bias in the earnings sampled, because those with desirable skills will be the first to find work in the labor market. In light of these considerations, Burgess and Kingston's results seem difficult to interpret.

Arlene Holen [1977] uses the same interstate data base as Burgess and Kingston to estimate the following model

$$D = g(WBA,\ \text{potential duration, age, race, sex, occupation,}$$
$$\text{city dummies, education, city employment rates,}$$
$$\text{previous earnings level and previous earnings stability)}$$

Again, the dependent variable D is the number of weeks of compensated unemployment and g is a linear function. The various city specific variables

capture characteristics that are unique to the five states in the STC file. Since the unemployment insurance system and the method of calculating benefits differs from state to state, this data set is useful in that it allows for independent variation in the different unemployment insurance parameters.

Ordinary least squares was used on two types of regression specification. One of these regressions includes city dummies, while the other uses seasonally unadjusted labor market unemployment rates for the months when the unemployment spell began as control variables. The results of these equations are given in Table 3-4. For both equations the estimated coefficients of the unemployment insurance paremters are very similar and quite significant. Since the compensated duration is used rather than the actual duration of unemployment, there is the same truncation problem present here that Classen encountered in her data set. Holen is aware of this and double checked the effect a change in potential benefits would have on duration by examining whether fewer workers have short spells of unemployment when the potential duration is increased. The nature of the data set allows this sort of comparison. The results tended to verify that the weeks of compensated duration are positively related to potential benefits, although the TOBIT method would still be more appropriate than OLS.

The subsequent earnings equation is also presented in Table 3-4. The dependent variable is defined as the average quarterly post-unemployment earnings for the number of quarters in which earnings were reported. This is the best measure of annual wages available from the sample data. The problem is the same as that of Burgess and Kingston, the earnings history is based on FICA income that is bounded from above. The empirical significance of this will depend upon the actual number of individuals in the sample at the

24

TABLE 3-4

HOLEN'S DURATION AND SUBSEQUENT EARNINGS RESULTS

Dependent Variable:

Independent Variables:	Compensated Duration		Compensation Duration		Subsequent earnings per quarter with earnings	
WBA	.09	(13.4)	.07	(11.4)	8.9	(14.6)
Max Weeks	.77	(66.4)	.81	(73.2)	2.5	(5.7)
Service to Claimants	-1.15	(4.4)	-2.06	(8.1)	32.3	(1.6)
Male	-1.80	(10.3)	-1.81	(10.3)	138.3	(10.1)
Age	.11	(19.6)	.11	(20.0)	- 2.6	(5.6)
Non-white	.85	(4.1)	1.35	(6.8)	8.4	(.5)
Base period wage	- .0004	(22.3)	- .0004	(20.9)		
Number of quarters of previous earnings	.91	(12.2)	- .86	(11.5)	13.5	(1.9)
Previous average quarterly earnings					.2	(14.4)
Education:						
Under 8 years	.54	(1.6)	.07	(.20)	-83.3	(3.0)
8-11 years	.74	(3.9)	.45	(2.3)	-69.4	(4.6)
13-15 years	-1.08	(5.0)	-1.09	(5.0)	14.3	(.8)
College year	- .60	(1.8)	- .13	(.9)	80.0	(2.3)
Missing	-2.42	(7.6)	-4.15	(15.1)	-27.6	(1.2)
Occupation (8 dummies)						
Boston	-2.13	(0.8)			1.0	(.1)
Phoenix	-3.97	(13.8)			61.7	(2.9)
Seattle	4.60	(18.2)			-22.7	(1.1)
Minneapolis (St. Paul)	.17	(.5)			-16.4	(.6)
Unemployment rate			1.20	(35.0)		
Constant	-5.82		-12.25		298.1	
Adjusted R^2	.326		.319		.174	
Standard Error	9.67		9.72		565.8	

Source: Holen [1977].

Note: t-statistics are given in parentheses.

25

upper bound. Holen reports that more than 45% of the sample is at the FICA maximum or has zero earnings. A further breakdown is not given, but this represents a substantial amount even if the breakdown is split down the middle.

This paper also analyzes the effects of other unemployment insurance parameters besides WBA and potential duration. These results may be summarized along with those of Table 3-4 as follows: (1) Higher benefit levels lead to greater duration, (2) Higher benefit levels lead to greater post-unemployment earnings, (3) Longer periods of entitlement leads to increased duration of unemployment spells, (4) Increases in the level of work-test enforcement reduce both duration and benefit exhaustion; although, there is a selection bias in the previous results due to this, (5) Increased eligibility screening leads to higher subsequent earnings, and (6) Job search assistance reduces the duration of compensated unemployment.

Perhaps the most exhaustive study of unemployment insurance effects is that of Ehrenberg and Oaxaca [1976]. In the appendix of their paper they describe analytically the links between job search theory and the empirical literature. They have recognized the need for a structural approach because of the simultaneity aspects of duration of unemployment and subsequent earnings. Nevertheless, they use OLS to estimate their reduced form equations. The specification of the duration and subsequent earnings equations are:

$$\ln(W_t/W_{t-1}) = a_1 F + \sum_{j=2} a_j X_j + u_1 \ ,$$

$$\ln(D) = b_1 F + \sum_{j=2} b_j X_j + u_2 \ ,$$

26

where $\qquad W_t$ = the post-unemployment wage,

W_{t-1} = the pre-unemployment wage,

D = the average duration of unemployment, if the number of spells is one, then this is the actual duration of unemployment.

F = the ratio of unemployment insurance benefits to previous wages,

X_j = the jth control variable,

u_1, u_2 = the random disturbance terms.

The data set is from the National Longitudinal Surveys (NLS). The duration and wage equations were estimated for four different groups: males 45-59 years of age in 1966, Table 3-5; females 30-44 years of age in 1967, Table 3-6; males 14-24 years of age in 1966, Table 307; females 14-24 years of age in 1968, Table 3-8. The observations in each of these groups were selected on the basis of certain criterion. This criterion varied among the different groups, but it still introduces sample censoring problems similar to previous studies, so the generalization of their results is justifiably constrained.

The important variable in these regressions is F, the replacement ratio. This is a proxy for the opportunity cost of unemployment, and is the same varaible that Chapin used in his regressions. If Chapin used the actual percentage figure in his regressions, then $\partial D/\partial F = 5.6$; however, if he did not use percentages, then the results are quite different. Ehrenberg and Oaxaca have a number of different estimates for the coefficient of F. These range from 0.371 to 1.623. If a geometric mean of 6.0 weeks can be assumed for D, the two papers do not given an estimate of the mean duration, then this implies that $2.22 \leq \partial D/\partial F \leq 9.78$. These bounds include Chapin's

27

TABLE 3-5

NLS OLDER MALE RESULTS:
DURATION AND WAGE EQUATIONS

Independent Variable	All Spells			
	Duration		Wage	
F	1.110	(2.0)	44.168	(2.5)
Race	.230	(1.3)	-0.146	(0.0)
Married	.316	(0.9)	-1.189	(0.1)
Own	-.014	(0.1)	3.965	(0.5)
Depend.	.012	(0.2)	0.356	(0.2)
Horizon	.006	(.03)	1.28	(1.8)
Assets	.012	(1.5)	.015	(0.0)
Tenure	---	---	.191	(0.5)
Constant	.598	(1.1)	22.494	(1.3)
R^2	.24		.313	
d.f.	54		53	

Source: Ehrenberg and Oaxaca [1976]

Note: F = weekly benefits/weekly pre-unemployment wage,
        Race = 1 = white; 0 = non-white
     Married = 1 = married and spouse present; 0 = othewise,
         Own = 1 = home owner; 0 = renter,
      Depend = number of dependents excluding wife,
     Horizon = expected number of years to retirement,
      Assets = family net assets/1000,
      Tenure = number of years at previous job
        Educ = years of school completed.
This is a partial list of results.

28

TABLE 3-6

NLS FEMALE RESULTS:
DURATION (D) AND WAGE (W) EQUATIONS

Independent Variables	1968-71 Sample				1968-71 Employer Change Sample			
	D		W		D		W	
F	.3	(1.7)	.12	(4.4)	.43	(2.0)	.14	(4.2)
Age	.006	(0.3)	-.006	(2.3)	-.009	(0.4)	-.009	(2.6)
Race	-.003	(1.7)	-.013	(0.5)	-.036	(0.2)	-.006	(0.2)
Married	-.29	(1.7)	.015	(0.6)	-.41	(1.8)	.031	(0.8)
Depend.	.074	(1.7)	-.01	(1.4)	.066	(1.0)	-.011	(1.1)
Educ.	-.052	(1.7)	.009	(2.0)	-.062	(1.5)	.009	(1.3)
Assets	.001	(0.1)	.004	(2.0)	-.004	(0.2)	.006	(2.1)
Constant	.743	(0.8)	.346	(2.2)	1.375	(1.1)	.484	(2.2)
R^2	.15		.37		.16		.44	
d.f.	234		234		137		137	

Source: Ehrenberg and Oaxaca [1976]

Note: This is a partial list of variables, and t-statistics are given in parentheses.

TABLE 3-7

NLS YOUNGER MALE RESULTS:
DURATION AND WAGE EQUATIONS

Independent Variables	Change Employers Sample			
	Duration		Wage	
F	.538	(2.1)	.093	(0.9)
Age	.003	(0.2)	.021	(2.6)
Race	-.074	(0.8)	.04	(1.1)
Married	-.252	(2.1)	.064	(1.3)
Tenure	-.036	(0.9)	-.002	(0.1)
Educ.	-.057	(2.4)	.029	(3.1)
Assets	-.022	(0.7)	.009	(0.8)
Constant	1.625	(3.8)	-.326	(1.9)
R^2	.05		.46	
d.f.	448		448	

Source: Ehrenberg and Oaxaca [1976].

Note: This is a partial list of regression variables, and t-statistics are given in parentheses.

TABLE 3-8

NLS YOUNGER FEMALE RESULTS:
DURATION, WAGE, AND OUT-OF-LABOR
FORCE EQUATIONS

Independent Variables	Duration		Wage		Out-of-Labor Force	
F	1.222	(3.8)	0.041	(0.4)	-8.002	(2.1)
Age	.027	(1.3)	.012	(1.9)	-1.046	(4.5)
Race	.206	(2.2)	.034	(1.1)	.702	(0.7)
Married	.036	(0.3)	-.007	(0.1)	.064	(0.0)
Depend.	.010	(1.8)	-.002	(0.1)	.857	(1.4)
Educ.	.007	(0.3)	.054	(6.6)	.434	(1.5)
Assets	-.092	(0.5)	.014	(0.2)	-.746	(0.3)
Constant	.521	(1.0)	-.249	(1.4)	12.501	(2.0)
R^2	.17		.60		.19	
d.f.	595		595		595	

Source: Ehrenberg and Oaxaca [1976].

Note: This is a partial list of regression variables, and the figures in parentheses are t-statistics.

31

result, which is surprising given the data problems in his paper. Because this result is positive it tends to reaffirm the hypothesis that unemployment insurance benefits increase the duration of unemployment.

The regression estimates for the earnings equations tend to vary widely between the data groups. Unemployment insurance benefits had a significantly positive effect only among males aged 45-59 and females aged 30-44. Thus, the job search theories seem to be supported among the older age groups, but for the younger groups, an increase in unemployment insurance benefits is not expected to increase post-unemployment earnings.

Recognizing the simultaneity problem, Holen and Horowitz [1976] develop a system of equations to model unemployment insurance effects. They do this primarily to disentangle the interactions between the unemployment rate and benefit liberality at the state level. The basis for an interaction is mentioned in the discussion of Chapin's paper: States with a higher incidence of unemployment, for whatever reason, may have a more liberal system of benefits. Holen and Horowitz have used five measures to describe the benefit liberality in a state: The fraction of workers whose jobs are covered under unemployment insurance; the fraction of claimants with recent work and earnings; denials for voluntary job separation or non-search as a fraction of office visits by claimants who were otherwise eligible; the average weekly benefit amount; and claims exhaustions as a fraction of first payment. Each of these components are used as endogenous variables in a system of eight equations. The functional specification of these equations is as follows:

$$u = f^1 \; (\overline{\overline{UI}}, \overline{W}, \overline{X}, \overline{L}),$$

$$E = g^2 \; (\overline{\overline{UI}}, W, \overline{X}, \overline{P}),$$

$$LF = g^3 \; (\overline{\overline{UI}}, W, \overline{X}, \overline{P}),$$

$$\overline{\overline{UI}} = g^{4-8} (u, W, \overline{X}, PI, R, L, J_i),$$

where

u = the state unemployment rate,

W = wages,

$\overline{X}$ = a vector of industrial mix variables,

$\overline{L}$ = a vector of labor force characteristics,

PI = a political index,

E = the unemployment rate relative to the state population,

LF = the labor force participation rate,

$\overline{P}$ = a vector of population characteristics,

$\overline{\overline{UI}}$ = the vector of benefit liberality components,

$\overline{R}$ = a vector of unemployment insurance administrative resources including quantity and allocation,

$\overline{J}_i$ = a vector of legal variables appropriate to the liberality component being used as a dependent variable.

The equations have been estimated by 2SLS using statewide data from various sources [7]. These results are presented in Table 3-9.

The interesting thing about these results is that the usual measures of benevolence (coverage, eligibility, replacement ratio) are insignificant, but the denial rate and exhaustion variable have a significant impact on the unemployment rate equation. This is puzzling given that the replacement ratio has been quite important in previous studies. They conclude from this evidence that stricter enforcement of the work-test would lead to a reduction in the overall unemployment rate. Specifically, if the denial rate is doubled, then their estimates imply that the unemployment rate would be reduced by 1.2

33

TABLE 3-9

Independent Variable	Equation No.				
	(1)	(2)	(3)	(4)	(5)
Constant	-0.28 (-0.03)	2.40 (0.07)	6.98 (0.22)	20.20 (0.62)	-19.43 (-0.43)
U-E Rate				-1.05 (-1.19)	0.08 (0.10)
Denial Rate	-0.056 (-2.34)	0.038 (0.48)	-0.006 (-0.08)	-0.080 (-0.84)	
Franction Covered	8.14 (1.34)	17.69 (1.03)	18.18 (1.10)	17.40 (1.10)	
% w/Wage Credits	0.049 (0.99)	0.049 (0.28)	0.041 (0.24)	0.029 (0.18)	
WBA:Weekly Wage	-0.80 (-0.09)	38.54 (1.26)	38.09 (1.29)	33.26 (1.17)	
% Exhausting	0.0970 (2.30)	-0.12 (-0.92)	-0.093 (-0.73)	-0.032 (-0.24)	

Independent Variable	(6)	(7)	(8)	(9)
Constant	0.360 (2.63)	73.87 (2.10)	0.270 (3.31)	26.22 (1.10)
U-E Rate		-3.61 (2.00)	-0.0013 (-0.26)	2.74 (2.87)

Source: Holen and Horowitz [1976].

Note: This is a partial list of the results. The dependent variable
for each equation are identified as follows: (1)--Unemployment
Rate, (2)--Employment Rate, (3)--Labor Force Participation Rate,
(4)--Same as in (3), (5)--Percent with Sufficient Wage Credits,
(6)--Fraction Covered, (7)--Denial Rate, (8)--WBA:Weekly Wages,
t-values are given in parentheses.

34

percentage points, ceteris paribus. Unfortunately, this conclusion may be biased by the occurrence of low denial rates in states with relatively higher unemployment rates because of low quit rates and less frequent job opportunities.

In a separate but related paper, Horowitz [1977] examines the effect of unemployment insurance on two groups of people: The unemployed, and those who say they are unemployed but are actually out of the active labor force. To accomplish these ends, he posits two theoretically consistent equations as a method of distinguishing between the two groups. These equations are as follows:

$$P_s = C \; l \; d \; B, \qquad \text{(a product equation)}$$

$$P_n = w \; P_s,$$

where

P_s = the average unemployment insurance payment to a searcher,

P_n = the average unemployment insurance payment to a non-searcher,

C = the fraction of employment covered by unemployment insurance,

l = the fraction of covered unemployed who have enough work experience to qualify for benefits,

d = the fraction of beneficiaries who have not exhausted their benefits,

B = the average benefit amount,

and W = the probability of collecting unemployment insurance when one is not looking for work, estimated by one minus the denial rate.

Horowitz specifies an unemployment rate equation depending upon various supply and demand characteristics, i.e., P_s, P_n, and a vector of $\bar{X}$ of demographic and industrial environment factors. Using data from the 1970 census he estimates a linear form for the unemployment rate equation:

$$u = b_0 + b_1 P_s + b_2 P_n + b_3' \bar{X},$$

or equivalently,

$$u = b_0 + CldB(b_1 + wb_2) + b_3' \bar{X}.$$

Both OLS and 2SLS are used to estimate these equations because of the simultaneity problems with u and the variables 1, d, and B. The results of this exercise seem to indicate that $b_1 = -b_2$. Thus, if w is approximately 0.95, then $\partial u / \partial B = cld(1-0.95)abs(b_1)$. Since Horowitz does not give estimates for the mean of C, 1, or d, a statistical value cannot be calculated, but its bounds would range from 0 to 0.05 given that $0 \leq b_1 \leq 1$. The upper end of this range would mean that a \$1 increase in B would increase u to 0.05. Since u is a fraction measuring the adult unemployment rate, this effect seems unrealistic. A point related to the fractional nature of u has been made by Orley Ashenfelter [1977] in a comment on Horowitz's paper. Ashenfelter wisely notes that because u is a probability, for all practical purposes, the estimation technique should constrain the estimates to the zero-one interval. Because this is not done by OLS or 2SLS probably explains why the comparative statics seem out of proportion.

In the most recent paper analyzing the effects of unemployment compensation, Kiefer and Newmann [1979] develop a reservation wage model similar to that suggested in Chapter V. Their model is described by the following set of equations:

$$W_{it}^0 = X_i \beta + \varepsilon_{it}^0 ,$$

$$W_{it}^r = Z_{it} \gamma + \varepsilon_{it}^r ,$$

where
$$\varepsilon^0_{it} \sim \text{i.i.d. } N(0, \sigma^2_0),$$

$$\varepsilon^r_{it} \sim \text{i.i.d. } N(0, \sigma^2_r),$$

$$E(\varepsilon^0_{it} \, \varepsilon^r_{it}) = \sigma_{0r}.$$

After a detailed discussion of the theoretical merits of this approach, Kiefer and Neumann outline a maximum likelihood procedure to estimate this model. The likelihood function depends upon the joint probability of D_i, the duration of unemployment, and W, the observed reemployment wage. The correct expression for the joint probability is as follows:

$$P_r(D_i, W_i) = P_r(D_i) \, P_r(W_i | D_i)$$

$$= \prod_{t=1}^{D_i - 1} [1 - \alpha(t)] \, \alpha(D) \, \theta(D),$$

where

$$\alpha(t) = \Phi \left[\frac{X_i \beta - Z_{it} \gamma}{\sigma} \right];$$

$$\theta(D) = \int_{-\infty}^{W_i - Z_{iD} \gamma} h(W_i - X_i \beta, \varepsilon^r_{iD}) \, d\varepsilon^r_{iD} / \alpha(D)$$

and $\Phi(.)$ is the standard normal distribution function, $h(.,.)$ is the bivariate normal density function, and $\sigma^2 = \sigma^2_0 + \sigma^2_r - 2\sigma_{0r}$. This expression is not equivalent to the equation given in their paper [8]. More importantly, the joint probability approach is a special case of the employment history probability procedure outlined in Chapter V. It arises when the sample information is restricted to the period of unemployment and reemployment wage information. As such, it tends to

37

emphasize the unemployment spell more than the period of employment relative to the employment history probability approach.

The sample data are from a survey conducted by the Institute for Research on Human Resources of the Pennsylvania State University to study the effects of the Trade Adjustment Assistance Program. These data represent individuals who were initially unemployed due to a plant closing, and thus had to engage in search activity to find another job or drop out of the labor force. The most notable feature of these data is that the average duration of unemployment is quite long --- 61.4 weeks for females and 39.4 weeks for males. In addition, nearly 40 percent of the sample had not found employment at the time of the survey, where the minimum time period between plant closing and interview was two years. This suggests that a large percentage of the individuals involved withdrew from the active labor force.

The estimates that Kiefer and Neumann obtained are given in Table 3-10 and 3-11. Table 3-10 contains the estimates of a "constant reservation-wage model" and Table 3-11 contains the estimates of a "changing reservation-wage model." They differ only in specification. The changing reservation wage model contains a variable, g, that measures the length of time unemployed to date. Its negative coefficient implies that the reservation wage falls about 0.6 percent per week. Both models use logarithms of the wage variables in the estimated equations. These results suggest that as unemployment benefits increase (decrease), reservation wages increase (decrease). If benefits increased by \$1, ceteris paribus, then the constant reservation wage model implies that W_i^r would increase by .043% and the changing reservation wage model implies that W_i^r would increase by .037% [9]. These represent very small

38

TABLE 3-10

CONSTANT RESERVATION WAGE RESULTS

Independent Variables	Wage Offer Equation		Reservation Wage Equation (Normalized)	
Constant	2.94	(7.36)	4.68	(3.59)
Education	.056	(2.73)	.067	(2.46)
Dependents	---		.0079	(.44)
Tenure	.007	(3.11)	---	
Married	---		-.162	(2.12)
U/E Rate	.029	(1.5)	.09	(4.06)
Age	.011	(2.0)	-.046	(3.47)
$(Age)^2$	0	(.33)	.0009	(5.5)
Education x Age	-.0005	(1.35)	-.0008	(1.56)
U/E Benefits	---		.0038	(2.79)
Max. Duration	---		.0021	(1.02)
$\ln W_{t-1}$	.2425	(4.67)	---	
$\hat{w}$			.6346	(2.27)

Source: Kiefer and Neumann [1979].

Note: $\hat{w}$ represents the effect of tenure and $\ln W_{t-1}$ on the mean of the wage offer distribution and the numbers in parentheses are asymptotic t-statistics.

39

TABLE 3-11

CHANGING RESERVATION WAGE RESULTS

Independent Variable	Wage Offer Equation		Reservation Wage Equation (Normalized)	
Constant	2.8	(18.1)	3.46	(3.31)
Education	.081	(3.15)	.031	(1.13)
Dependents	---		-.012	(.67)
Tenure	-.007	(3.03)	---	
Married	---		-.246	(3.63)
U/E Rate	.031	(1.76)	.071	(3.66)
Age	.025	(1.9)	-.036	(3.15)
$(Age)^2$	-.0001	(.41)	.0006	(4.47)
Education x Age	-.001	(1.96)	-.0004	(.65)
U/E Benefits	---		.0033	(2.67)
Max. Duration	---		.0013	(.63)
lnW_{t-1}	.254	(4.89)	---	
$\hat{w}$	---		.273	(1.31)
g	---		-.0173	(12.59)

Source: Kiefer and Neumann [1979].

Note: The figures in parentheses are asymptotic t-statistics.

40

adjustments when compared to a mean employment wage of approximately $149.0 [10]. Also, the estimates imply that if the maximum duration of benefits is increased by one week, then reservation wages will decline by .024% for the constant reservation wage model and increase by .014% for the changing reservation wage model. An ambiguous result, but not very damaging in light of the significance of the maximum duration coefficients.

In summary, it could be said that the Kiefer and Neumann study supports the results of the previous literature in their implications on the effects of unemployment insurance on the duration of search. More about this in Chapter VI. But the falling reservation wage implication is in slight contrast to the work of Holen, who finds that the maximum duration of benefits has a significant and positive effect on reemployment wages, and the work of Marston, who finds that the probability of finding employment decreases over the duration of unemployment.

the search theory as formulated by McCall [1970], Mortenson [1977], and others, implies that in the absence of the above factors, the reservation wage will be constant in the infinite time horizon case and will be declining in the finite time horizon case. What we are interested in is to see what difference the assumption of a finite time horizon makes.

Infinite Time Horizon

When an individual assumes that there is an infinite number of periods in which job search may occur, the duration of job search will have a geometric distribution. This distribution is derived as follows:

Let W_r be the reservation wage and $F(.)$ the distribution function of wage offers. Then the probability that an individual does not find employment in any period = Prob (wage offer $< W_r$) = $F(W_r)$. The probability that the individual finds employment = $1 - F(W_r)$.

Assume that reservation wage W_r is constant over time. Given that an individual follows a reservation wage strategy of job search, what is the probability of stopping after $D = t$ searches? We have

$$\text{Prob } (D = 1) = [1 - F(W_r)] = P_1$$
$$\text{Prob } (D = 2) = F(W_r) [1 - F(W_r)] = P_2$$
$$\text{Prob } (D = 3) = [F(W_r)]^2 [1 - F(W_r)] = P_3$$

etc.

$$\text{Prob } (D = t) = [F(W_r)]^{t-1} [1 - F(W_r)] = P_t.$$

77

This expression can be computed from the estimates of the wage offer equation and the reservation wage equation described in the previous section. With constant reservation wage, if we consider a finite time horizon T for search, equation (6.1) would be

$$E(D) = [\sum_{t=0}^{T} t\,P_t] \Big/ [\sum_{t=0}^{T} P_t]$$

$$= \frac{\sum_{t=0}^{T} tx^{t-1}(1-x)}{\sum_{t=0}^{T} x^{t-1}(1-x)} \qquad (6.3)$$

Now $\quad \sum_{t=0}^{T} x^{t-1} = \dfrac{1-x^T}{1-x} \qquad (6.4)$

To find $\quad \sum_{t=0}^{T} tx^{t-1} \quad$ note that $\quad \sum_{t=0}^{T} x^t = \dfrac{1-x^{T+1}}{1-x}$

Differentiating both sides with respect to x we get

$$\sum_{t=0}^{T} tx^{t-1} = \frac{1-x^T - Tx^T(1-x)}{(1-x)^2} \qquad (6.5)$$

Hence, substituting (6.4) and (6.5) in (6.3) we get

$$E(D) = \frac{1}{1-x^T} \cdot [\frac{1-x^T - Tx^T(1-x)}{1-x}] \qquad (6.6)$$

where, as defined earlier, $x = F(W_r)$. We will use this expression later to get a bound for E(D) in the finite horizon case. Note that in the finite, horizon case W_r will not be constant. But the expression (6.6) can be used to get an upper bound on the value of E(D) when W_r is declining.

79

Thus, $\qquad$ Prob $(D = t) = P_t^*$ (say)

$$= \prod_{j=1}^{t-1} F(W_{rj}) [1-F(W_{rt})] \qquad (6.7)$$

The mean duration of search is given by:

$$E(D) = \sum_{t=1}^{T} tP_t^* \bigg/ \sum_{t=1}^{T} P_t^* \qquad (6.8)$$

where T is the search horizon and P_t^* is given by equation (6.7). What we have here is the mean of a truncated geometric distribution.

Now, this expression does not have a closed form solution. This creates an empirical estimation problem because, in order to estimate (6.8) the reservation wage must be estimated from the initial period of separation to the time horizon T. Kiefer and Neumann [1979] estimate (6.4) numerically assuming $T = \infty$ and "with outliers removed." If we make the assumption of a finite time horizon, then T itself is a parameter that needs to be estimated. For this purpose we proceeded as follows:

We first ask the question: though we cannot get a closed form expression for D as given in (6.8), is it possible to get a bound for it? The answer is yes. Under the assumption that the reservation wages decline over time

$$\text{i.e., } W_{r1} > W_{r2} > W_{r3} > \ldots \ldots > W_{rT}$$

We can show that E(D) as given by equation (6.8) is less than the corresponding expression calculated with the reservation wages held constant at the initial level i.e., $W_{rj} = W_{r1}$.

Note that $= F(W_r) = Prob(W_o < W_r)$ can be computed from the estimates of the wage offer equation and reservation wage equation obtained in the previous chapter. In our computations we used the estimates in Tables 5-2 and 5-3 to predict W_r during periods of unemployment and to construct the mean and variance estimates of the wage offer distribution.

Before we go through any further analysis, it would be interesting to compare the expected mean duration of unemployment as computed from the structural model with the actual mean duration of unemployment from the CWBH data. The mean duration for the CWBH sample that was used in this study was 12.43 weeks. This compares well with the national figures published by the Bureau of Labor Statistics where the means are between 10 and 13 weeks for the same period.

The numbers generated by the analytical procedure are in Table 6-2. The values of E(D) for $T = \infty$ and $T = 40.87$ were 12.99 and 11.17 weeks respectively. Thus, the estimated value of E(D) is remarkably close to the actual value thus suggesting that the results of the policy simulations would be credible.

Empirical Results: Duration Equation Estimates

An estimate of the expected duration of search equation was calculated using the reservation wage estimates in Table 5-3 to calculate reservation wages during unemployment spells and the variance and mean estimates of the wage offer equation to define F(.). For the infinite time horizon case, the residual sum of squares were calculated using the initial reservation estimate and an average of the estimated reservation wages over the period of

83

TABLE 6-1

DURATION EQUATION ESTIMATES

FINITE TIME HORIZON:

$\hat{T}$ 40.87

Standard Error .... .0023

$\hat{\alpha}$ .... -.2069

Standard Error .... .00017

Residual Sum of Squares 171832.60

INFINITE TIME HORIZON:

Residual Sum of Squares[1] 387059.5

Residual Sum of Squares[2] 396211.3

[1]Using the reservation wage in the initial period

[2]Using an average of the reservation wage during the
spell of unemployment.

interest because they bound from above the results of Classen
[1977a] and Holen [1977] for the same variable. Their estimates indicate a
change in the expected duration of unemployment of 1.1 and .9 weeks,
respectively. This also occurs for the estimates generated by changing the
PWEEKS variable, suggesting perhaps that the comparative statics generated
using $T = \infty$ and $T = 40.87$ provide an absolute upper bound on the effects of
unemployment insurance policy variables.

The estimates of the exhaustion effects are a bit disappointing. The
infinite time horizon is particularly difficult to accept being almost two
weeks larger than the finite time horizon change. The reason for these
results is unclear, but the magnitudes can be explained by noting that the
exhaustion effect decreased reservation wages approximately 15% and this
decline impacts upon the expected duration of search equation in a highly
nonlinear way.

In summary, it can be said that the comparative statics for PWBA
and PWEEK's support the hypothesis that unemployment insurance reduces the
costs of search, and thus increases the expected duration of search. Also,
if the finite time horizon concept is valid, then it appears to provide a
dampener on the effect of unemployment insurance variables in the labor
market; particularly in the effect of PWBA on the duration estimates. This
variable has been the major policy tool of the unemployment insurance
administration and these results indicate that its influence will depend
directly upon the time horizon of the unemployed.

Implications

On a theoretical plane the implications of this study are twofold.
First, the empirical methodology conforms to the specification imposed by
reservation wage job search and the estimates of the censored regression model

CHAPTER VII

SUMMARY AND POLICY RECOMMENDATIONS

This study has analyzed the effects of unemployment insurance on labor market behavior using the CWBH data for the State of Florida. The theoretical models of the effects of unemployment insurance have shown that unemployment compensation has unambiguous implications on the duration of unemployment and reemployment wages. These models suggest that both the duration of unemployment and reemployment wages are expected to be higher with unemployment compensation than they would be in its absence. However, one cannot argue on theoretical grounds that reemployment wages with unemployment insurance will be greater than pre-unemployment wages. This has been a central issue in many studies (See Chapter III), but the falling reservation wage creates an ambiguity here.

The empirical results in Chapter V support the falling reservation wage hypothesis. Moreover, these results suggest that the reservation wage falls by approximately $2.30 for each week of compensation. This translates into an increase in the probability of finding a job by approximately 0.005 each week, i.e., a decrease in the expected duration of search by 1.15 weeks for the infinite time horizon case or 0.64 weeks for the finite time horizon case. In addition, the results imply that if the potential weekly benefit amount were decreased by $10, then the expected duration of search would decrease by 2.78 weeks in the infinite time horizon case or 1.37 weeks in the finite time horizon case.

studied this problem extensively with our data set because we feel that the data do not permit any meaningful answers to this question. We did estimate a reduced form equation, as others had done (see e.g,. Classen in Chapter III) and obtained similar results and thus we could have argued that increased benefits do not lead to more productive job search. However, as we have argued repeatedly, the comparison of post re-employment wages with pre-unemployment wages is not a valid indicator of productive job search. One has also to look at the job stability (and job satisfaction) in the new job.

In any case, our conclusion is that the positive effects (of increase in productive job search) are not estimable from the data we have but our feeling is that these effects are rather small. As for the negative effects (viz., increase in the duration of unemployment) these are also small, though not smaller than the positive effects. Overall, these results are not surprising because the increases in unemployment insurance benefits have barely kept pace with inflation.

In terms of policy decisions we do not feel, as Hamermesh [1977] does, that the unemployment insurance program should be enlarged by extending coverage and increasing benefits; although, this assessment comes too late because of the recent extension of coverage to state and local government workers. An extension of coverage and/or benefits would only increase the probability of bankrupting the already beseiged unemployment insurance trust fund. Any case for extending the coverage and increasing the benefits has to be made on grounds of equity - not on grounds of resource allocation.

91

Johnson and Kotz [1970, pp. 81-83] give the mean and variance of the truncated normal distribution.

$$E(X|X > C) = \frac{f(C)}{1-F(C)} = \frac{\text{ordinate at C}}{\text{right-hand tail area}} \qquad (3)$$

where $f(C)$ is the density function at $X = C$ and $1-F(C) = \text{Prob}(X \geq C) =$ the area of the shaded portion in Figure 1.

Let us denote this mean of the truncated normal by M.

$$\text{Var}(X|X > C) = 1-M(M - C) \qquad (4)$$

From the expressions (3) and (4) for the standard normal we can easily convert to the mean and variance of the observed wage distribution, e.g.,

$$E(W_o|W_o > W_r) = \bar{W}_o + \sigma_o \frac{f(\frac{W_r - \bar{W}_o}{\sigma_o})}{1-F(\frac{W_r - \bar{W}_o}{\sigma_c})}$$

Note that the mean of the observed wages will be higher than the mean of the distribution of wage offers.

The Probit Regression Model

Consider the usual regression model

$$y^* = X\beta + u \text{ where } u \sim IN(0,\sigma^2).$$

Suppose we do not observe y^* but instead we observe only a dummy variable y defined as:

$$y = 1 \text{ if } y^* > 0$$
$$= 0 \text{ if } y^* \leq 0$$

93

limitations of the linear probability model can be found in textbooks in Econometrics (see e.g., Goldberger (1964)). Nowadays the probit regression programs are readily available and are also very fast.

Tobin's Model: Threshold Regression Model

Suppose that in the regression model in the previous section we observe y* only if it is greater than a threshold value C. Otherwise we do not observe y*. The observed y is related to the unobserved y as follows:

$$y = y* = X\beta + u \quad \text{if } y* > C \tag{4}$$

$$= 0 \qquad\qquad \text{if } y* \leq C$$

This is called a threshold regression model. This is the model considered by Tobin [1958]. The method of estimation is called the tobit (Tobin's probit) method.

Again, divide the observations y_1, y_2, $\ldots$ y_n into two sets:

S_1 : observations for which $y_i > C$

S_0 : observations for which $y_i = 0$

for any y in S_1, we have $y = y*$

Hence, $P(y) = \dfrac{1}{\sigma\sqrt{2\pi}} \exp[-\dfrac{1}{2\sigma^2} (y - X\beta)^2]$

$$= \dfrac{1}{\sigma} f\left(\dfrac{y-X\beta}{\sigma}\right)$$

95

Let us go back to the model in Chapter V. We have

$$W_o = X\beta + \varepsilon_o \qquad \text{wage offer function}$$

$$W_r = Z\gamma + \varepsilon_r \qquad \text{reservation wage function}$$

and the observed wage W is defined as:

$$W = W_o \text{ if } W_o \geq W_r$$

$$= 0 \text{ otherwise.}$$

Since we have two random variables ε_o, ε_r we now consider their joint distribution, which we assume to be normal with zero means and covariance matrix

$$\begin{bmatrix} \sigma_o^2 & \sigma_{or} \\ & \sigma_r^2 \end{bmatrix}$$

Denote this distribution by $f(\varepsilon_o, \varepsilon_r)$.

Again, as before, partition the observed sample into two sets.

$$S_1 : \text{working group for which } W > 0$$

$$S_0 : \text{non-working group for which } W = 0$$

For the working group we know that

$$W_o = W \quad \text{i.e., } \varepsilon_o = W - X\beta$$

and $\qquad W_r \leq W \quad \text{i.e., } \varepsilon_r < W - Z\gamma$

Hence, substituting these values in the function $f(\varepsilon_o, \varepsilon_r)$ we get the density function for W as

97

then for the probit regression model we have the joint density of y_1, y_2, ...
y_n as

$$L = \prod_{S_1} \left[1 - F\left(\frac{Z_i\gamma - X_i\beta}{\sigma} \right) \right] \prod_{S_0} F\left(\frac{Z_i\gamma - X_i\beta}{\sigma} \right)$$

Note that we cannot get separate estimates of γ and β from this. What we
get is estimates of:

(i) $\dfrac{\gamma_j - \beta_j}{\sigma}$ for those variables which are common to both Z and X.

(ii) $\dfrac{\gamma_j}{\sigma}$ for those variables in Z but not in X.

(iii) $\dfrac{\beta_j}{\sigma}$ for those variables in X but not in Z.

Further discussion of the estimation problems can be found in Chapter V.

Why Consider This Model?

The reason why we go through this involved estimation procedure is
that the observed distribution of wages is a truncated distribution. It is
the distribution of wage offers truncated by reservation wages - which are
themselves stochastic and unobserved. The stochastic threshold regression
model permits us to estimate the parameters of both the reservation wage
equation and the wage offer equation just from the observations on the
individual's personal characteristics, the employment status, and the market
wage if employed. From these parameter estimates we can estimate the
reservation wages for each individual. The expected duration of unemployment
depends on these reservation wages and the wage distribution. This is
explained in Chapter VI. See

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103

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