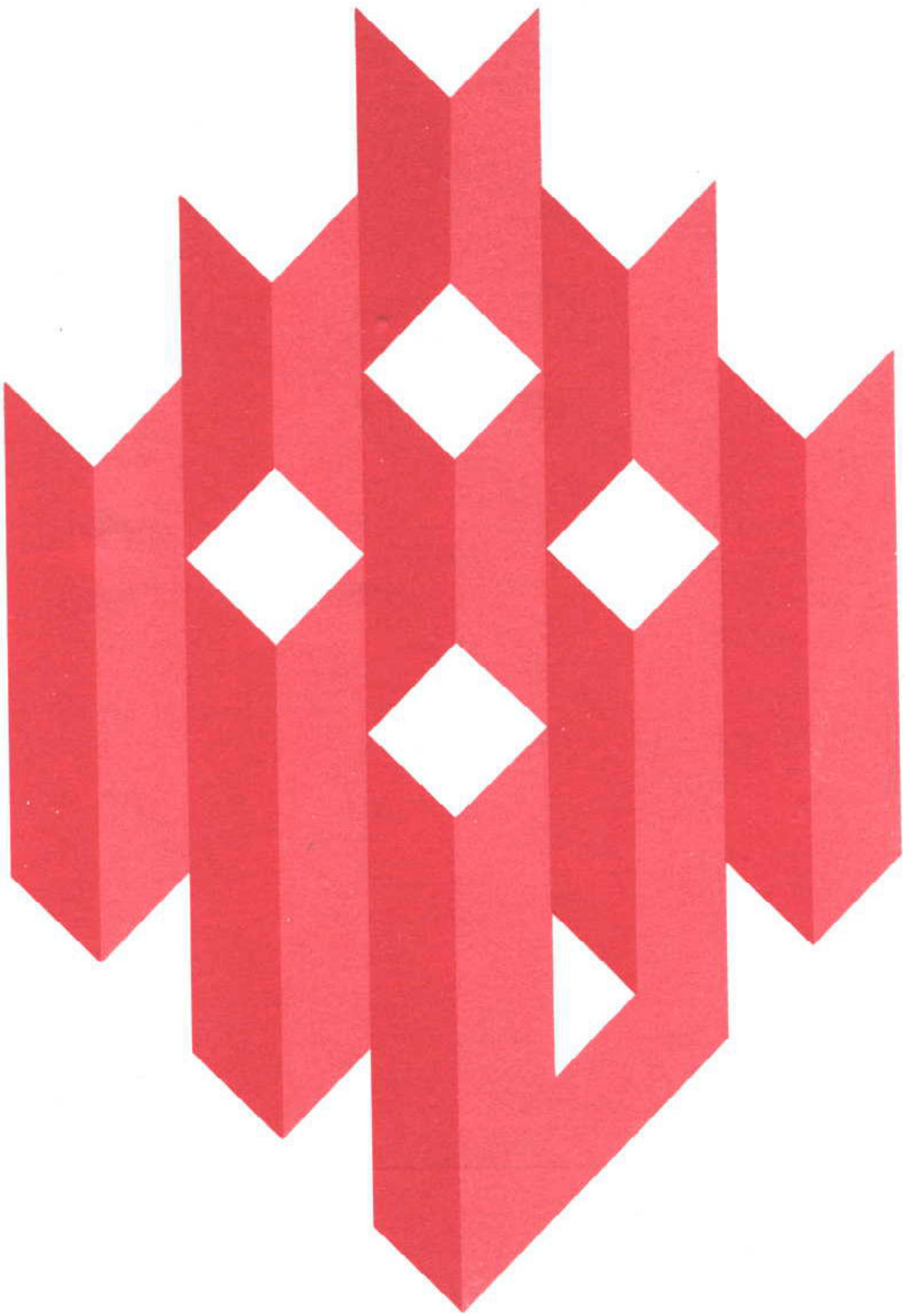
The Effect of State Laws and Economic Factors on Exhaustion Rates for Regular UI Benefits: A Statistical Model

U. S. Department of Labor
Employment and Training Administration



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U.S. Department of Labor
Ray Marshall, Secretary
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CONTENTS

TABLES

FIGURE

ABSTRACT

This report presents a statistical model of the effects of state laws and economic factors on exhaustion rates for regular Unemployment Insurance (UI) benefits. The model was estimated from quarterly pooled time-series, cross-sectional data for all states over the period 1965 to 1974. Results from the model were used to analyze various UI policy issues and to construct a simulation model that permitted a detailed examination of UI policies over the business cycle.

The report is divided into five chapters. In Chapter I, policy issues related to UI exhaustion rates are briefly summarized. Chapter II presents a theoretical model of the exhaustion rate of regular UI benefits that is suitable for empirical estimation. Results of estimating the model are presented in Chapter III. These results were, generally, highly statistically significant and were little affected by using various econometric techniques. A simulation model based on these statistical results and the predictive accuracy of that model are discussed in Chapter IV. Finally, in Chapter V, our models are used to derive a number of conclusions about the determinants of exhaustion rates and the implications of UI policies. The most important of our findings were:

1. Exhaustion rates were positively correlated with Insured Unemployment Rates (IURs). A cyclical, 1 percentage point increase in the IUR was calculated to result in a 4 to 5 percentage point increase in the exhaustion rate. That effect was offset somewhat by the tendency of average potential UI durations to rise during cyclical downturns.

2. Higher average potential UI durations tended to reduce exhaustion rates. We calculated that average potential durations would have to be increased by 3 to 4 weeks to offset the impact of a 1 percentage point increase in the IUR.

3. Wage-replacement ratios on UI tended to have a positive effect on exhaustion rates. The quantitative size of that effect was roughly similar to that derived from other studies that attempt to measure the "disincentive effect" of UI benefits. There was also some evidence that UI enforcement procedures had a negative effect on exhaustions.

4. Uniform-duration states were found to have exhaustion rates that were, on average, 8 to 9 percentage points lower than variable-duration states. About one-half of that difference could be accounted for by the higher average potential durations that occurred in uniform-duration states.

5. Increasing states' maximum potential duration by one week was estimated to reduce average exhaustion rates by one-half of a percentage point. An increase in the fraction of base-period wages to which UI recipients were entitled was also estimated to reduce the exhaustion rates somewhat through its positive effect on average potential durations.

6. Existence of the Extended Benefits (EB) program was estimated to increase exhaustion rates for regular UI by 4 to 5 percentage points. The combined effect of the EB and FSB (Federal Supplemental Benefits) programs was estimated to add nearly 7 percentage points to the exhaustion rate.

7. Our simulation model was used to examine issues related to the triggering of extended benefits. Results of those simulations suggested that current trigger mechanisms operated with a great lag (both for triggering-on and for triggering-off benefits) and were susceptible to seasonal influences. The use of a trigger formula that took exhaustions into account did not significantly affect these problems, but the use of a seasonally adjusted IUR in the state trigger formulas did moderate the sharp seasonal fluctuations.

I. BACKGROUND OF THE STUDY

INTRODUCTION

This report presents a model of the effects of state laws and economic factors on exhaustion rates[1/] of regular Unemployment Insurance (UI) benefits. The model is used to discuss and simulate various policy alternatives, and was estimated from pooled quarterly time-series, cross-sectional data for each state UI system over the period 1965 to 1974. Particular care was taken in constructing the model to specify various UI statutes in ways in which the results would be most directly interpretable to UI policymakers. In this chapter we discuss how the study of exhaustion rates is related to various major policy issues and summarize the results of prior examinations of this relationship. We conclude the chapter with a general overview of the report.

A. EXHAUSTION RATES AND POLICY INITIATIVES

In this section we examine the way in which data on exhaustion rates of regular UI benefits have been used in various policy debates. Our discussion is divided into four subsections. The first three address the relationship between exhaustion rates and UI policy. They also contain a discussion of the maximum potential duration of UI benefits,

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[1/] Theoretically, the exhaustion rate is the probability that a UI recipient will exhaust his or her entitlement to benefits. Since this concept is not directly observable, however, it is often approximated by the ratio of the number of individuals who receive a final UI payment during some period to the number who received a first UI payment in some prior period (usually six months previously to account for the typical twenty-six weeks maximum UI duration).

the variability of individuals' potential durations, and federal extended benefits initiatives, respectively. The final subsection briefly examines the impact of UI exhaustions on other transfer programs.

1. Exhaustion Rates and Maximum Potential UI Durations

The maximum number of weeks for which an individual may collect UI benefits has long been a crucial policy issue. Current duration standards represent a compromise between two opposing concerns. The first is the widely held belief that UI benefits should be available for a sufficient number of weeks to cover completely the unemployment spells of "most" individuals. It is also generally believed, however, that unlimited eligibility for UI benefits would be very costly and could pose work disincentives.[1/] Hence, some upper limit on potential durations is required. Early state UI legislation was strongly affected by these second concerns, and maximum potential durations were relatively low. In 1938, for example, only six states provided maximum durations of over sixteen weeks. As a result, the exhaustion rate of regular benefits was relatively high. During the years 1948 to 1950, national exhaustion rates were above 25 percent and, primarily in those states with restrictive maximum-duration policies, sometimes ran as high as 40 percent. In the early 1950s these experiences with exhaustions, combined with financial surpluses in UI accounts, led many states to increase their maximum potential durations. By 1952 more than half of all workers were covered under state programs that provided a maximum duration of at least twenty-six weeks. All states' programs had adopted such maximum duration standards by the 1960s. Since then, maximum-duration provisions have changed only slightly, even though exhaustion rates have varied widely both among states and over time.

[1/] For a brief summary of these concerns in the early debate over UI, see Report to the President of the Committee on Economic Security (Washington, D.C.: U.S. Government Printing Office, 1935), pp. 11, 13-14.

Instead, the focus of UI policy debate has been on reducing the extent to which individuals' potential durations within a state depart from the maximum (twenty-six-week) duration and on the extension of potential durations through federal initiatives. Exhaustion-rate data have been widely used in these debates, as well.

2. Exhaustion Rates and Variation in the Potential Duration of Benefits

Currently, eight states entitle all UI recipients to their maximum potential duration of benefits.[1/] Other states use a variety of entitlement formulas that produce considerable variation in the potential durations of benefits among recipients. The existence of potential durations of less than the state maximum (in some cases, potential durations of shorter than ten weeks) increases the probability that individuals will experience spells of unemployment longer than their UI duration. Therefore, it is expected that exhaustion rates will be high in those states with such variation in durations. For example, Haber and Murray (1966, p. 208) report that during the period from June 1964 to July 1964, most states that

> provided uniform duration of 26 weeks or more . . . had exhaustion rates of less than 20 percent. . . . On the other hand, only three of the 43 states with variable duration of benefits . . . had exhaustion rates of less than 20 percent.

[1/] As of 1975, these "uniform duration" states were Hawaii, Illinois, Maryland, New Hampshire, New York, Pennsylvania, Vermont, and West Virginia.

Similarly, Murray (1974, p. 15) observed that for the year 1969,

> by and large, the larger the proportion of claimants eligible for less than 26 weeks of benefits, the higher the exhaustion ratio for all claimants.[1/]

In later chapters we will develop a theoretical model that shows why such a (partial) correlation should exist and will demonstrate its importance for explaining exhaustion rates over our period of study.

This relationship between entitlement formulas and exhaustion rates has been a major focus of policy debate on Unemployment Insurance. During the early 1950s, for example, differing exhaustion-rate experiences were often used to argue for the adoption of a uniform-duration policy or for the liberalization of entitlement formulas. By 1952, fourteen states had adopted a uniform-duration policy, and the average duration in variable-duration states substantially increased. On the other hand, low exhaustion rates associated with the relatively full employment experienced during the 1960s reduced the pressure for liberalized entitlements. More concern was expressed about possible disincentive effects of UI benefits, and several states returned to variable-duration formulas as a way of testing claimants' labor-force attachment. In recent years, states' duration formulas have been altered only infrequently, and the focus on the relationship between exhaustion rates and duration policy has shifted to the federal level.

[1/] Murray's data show a correlation coefficient of 0.57 between the exhaustion rate and the percent of claimants eligible for less than twenty-six weeks of benefits.

3. Exhaustion Rates and Federal Extended Benefits Programs

Throughout the history of the UI program there has been a markedly positive correlation between exhaustion rates and the level of aggregate economic activity.[1/] That correlation, as is shown in the next chapter, derives primarily from the tendency of the average duration of unemployment spells to increase in economic downturns. Because general macroeconomic stabilization policies are conducted on a national level, it seems natural that exhaustion rates would play an important role as indicators of the need for federal policy initiatives.

In the early 1950s, these initiatives consisted primarily of urging states to adopt less variable and more generous entitlement formulas. The 1958 downturn, however, led to the congressional enactment of the Temporary Unemployment Compensation Act, which provided UI recipients with an increase in potential duration of one additional week for each two weeks to which they were originally entitled. A similar temporary program[2/] was enacted in 1961 in response to recession-induced increases in exhaustion rates during that year. Experiences with the temporary 1958 and 1961 programs clearly showed the importance of federal initiatives during recessions and indicated the need for a

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[1/] For example, aggregate data for 1964 to 1975 yield a simple correlation coefficient of .96 between the annual exhaustion rate and rate of overall unemployment. A simple regression of the exhaustion rate on the unemployment rate implied that a 1 percentage point increase in the unemployment rate led to about a 4 percentage point increase in the exhaustion rate.

[2/] The principal difference between the 1958 and 1961 programs involved financing. The 1958 program was financed through temporary loans from the federal government to state governments, whereas the 1961 act was federally financed through the Unemployment Insurance Trust Fund. Between 1958 and 1961 several states adopted their own extended benefits programs. The financing of these programs was taken over by the federal government under the 1961 act.

permanent policy regarding such recession-induced changes. Although numerous proposals for permanent programs were made during the period 1963 to 1969, it was not until 1970 that a program was finally enacted.

The Federal-State Extended Unemployment Compensation Act of 1970 is in many ways similar to the earlier temporary programs. Under its provisions, extended benefits (EB) are also payable for an additional week for each two weeks of initial UI entitlement up to a maximum of thirty-nine total weeks of benefits. The program is financed on a fifty-fifty basis by federal and state unemployment tax revenues. Contrary to the earlier programs, however, the 1970 act is permanent. Extended benefits become payable whenever state- or national-insured unemployment rates exceed certain specified levels. The actual mechanism by which the program is "triggered-on" has varied considerably since its enactment. Triggering criteria have generally been based on insured unemployment rates. At times, exhaustion rates have also been used as part of the trigger indicator. Specific legislative action has also been used to trigger-on the program.

Since the passage of the permanent EB program in 1970, benefits have been further extended on a temporary basis on two occasions. The Emergency Unemployment Compensation Act of 1971 (TC) provided additional weeks of benefits equal to 50 percent of the regular UI entitlement up to a maximum of fifty-two weeks in all (including regular UI, EB, and the emergency extension). This program ended March 30, 1973, and was followed in December 1974 by the Federal Supplemental Benefits (FSB) program. FSB, one principal part of the Emergency Unemployment Compensation Act of 1974, increased maximum duration by 100 percent of the regular UI entitlement up to a maximum of sixty-five weeks of benefits (that is, twenty-six weeks of regular UI, thirteen weeks of EB, and twenty-six weeks of FSB). Exhaustion rates were widely used as part of the debate surrounding the enactment of each of these emergency policies. Similarly, predictions of the number of exhaustees were an important parameter used to estimate caseloads and costs both for the emergency programs and for the EB program.

4. Exhaustion Rates and Other Transfer Programs

There are two reasons to expect that exhaustions of UI benefits may be positively related to changes in caseloads of other transfer programs. First, when individuals exhaust UI, they will be more likely to apply for benefits from other programs in an effort to replace the lost income from UI. And second, in the case of income-tested programs such as AFDC and Food Stamps, they will be more likely to be eligible once they have exhausted their UI benefits. For these reasons there is a tradeoff between expenditures for UI benefits and expenditures for other programs which should enter into the consideration of UI extensions.

These potential relationships between the exhaustion of UI benefits and its impact on other programs have been less intensively studied than have the relationships between exhaustion and UI-related issues, and empirical evidence about possible sizes of these effects is minimal.[1/]

[1/] Two types of studies can be used to examine this relationship: studies of exhaustee behavior, and macro-models that correlate benefit flows in various programs with, among other variables, exhaustions of UI. Evidence from the first type of study (contained in Nicholson and Corson, 1976, and Corson, Nicholson, and Skidmore, 1975) shows that few UI exhaustees receive income-tested transfers. Evidence from the second type of study does not exist. Models of transfer program caseloads have not, to our knowledge, explicitly included UI exhaustions as an explanatory variable, although they have usually included economic variables such as the unemployment rate. In an attempt to relate Food Stamp caseloads to UI exhaustions, the current

Similarly, changes in exhaustion rates have had little impact on debates about other policies, and there has been little feedback from the experiences of other programs on UI policy initiatives that affect exhaustion rates.

During the recent recession, however, the tradeoff between UI extensions and increased caseloads and costs of other programs was recognized in the policy debate concerning FSB. Two arguments were raised in support of UI extensions. It was argued that few UI exhaustees would be eligible for other programs, and it was also argued by some state and local officials that the lack of UI extensions would lead to a major financial burden for state and local governments through increased caseloads of the unemployed-father portion of the AFDC programs (AFDC-U). While these arguments appear to be contradictory, they are not because the UI caseload is substantially larger than the AFDC-U caseload.

B. PREVIOUS STUDIES OF THE EXHAUSTION OF REGULAR UI BENEFITS

Previous research on the exhaustion of unemployment compensation can be divided into two groups: (1) an examination of the determinants of the number of exhaustions or the exhaustion rate; and (2) examinations of the characteristics and behavior of exhaustees. The first area of research is important from a policy perspective because the effect that unemployment compensation program parameters have on the level or rate of exhaustions can be determined. The effect of economic variables such as the unemployment rate is also examined in these studies. The second area of research provides policymakers with information that allows them to

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authors were led to the conclusion that it was extremely difficult to disentangle the effects of UI exhaustions from those of other variables describing general economic conditions (e.g., the number of unemployed individuals).

decide what an appropriate level of exhaustion might be. For example, if it is judged that most exhaustees remain unemployed and are looking for work for some time after exhaustion, a case can be made for lowering the exhaustion rate by extending benefits. On the other hand, if exhaustees quickly become reemployed or stop looking for work, the case for extensions is weakened. Our summary of previous studies describes the important findings relating to these two areas of research, although this project primarily concerns the first: the determinants of the UI exhaustion rate.

1. Exhaustion Rate Determinants

There are several research studies that explore the relationship between UI program parameters and exhaustions of UI. In his study of the duration of unemployment benefits, Merrill Murray (1974) showed that, as one would expect, states with larger proportions of their caseload entitled to less than twenty-six weeks of regular UI benefits tend to have larger exhaustion rates. He then demonstrated that in variable-duration states[1/] the proportion of the caseload entitled to less than twenty-six weeks is a function of the fraction of base-period earnings (or weeks of employment) used in the duration formula. The larger this fraction, the smaller the proportion of the caseload with entitlements of less than twenty-six weeks. In a low-employment year, such as the year he analyzed, 1969, those states that restricted, at most, one-third of their claimants to fewer than twenty-six weeks tended to have exhaustion rates of less than 20 percent. Almost one-half of the states that restricted 40 percent or more of their claimants to fewer than twenty-six weeks had exhaustion rates greater than

[1/] In uniform-duration states, all recipients have the same number of weeks of eligibility regardless of base-period earnings or weeks worked.

25 percent. Thus, the proportion of claimants who are entitled to less than twenty-six weeks is an important determinant of the exhaustion rate. This proportion is, in turn, a function of program parameters under state control.

Another study of the determinants of exhaustion rates, by Joseph Hight (1975), also showed that longer potential durations of benefits reduce exhaustion rates. Hight's study used quarterly time-series data from two states--Pennsylvania and Georgia--to examine the relationship among insured unemployment rates, extended benefits, and UI exhaustion. For each state, he regressed the log of the exhaustion rate[1/] on the insured employment rate (IUR), two measures of the effect of state laws (potential duration of benefits and the ratio of the average UI benefit to average wages in covered employment), and a measure of the demographic composition of the caseload (the percent female). Quarterly dummies were also used to control for effects of seasonality. He found that, as hypothesized, the exhaustion rate was positively correlated with the insured unemployment rate, negatively correlated with potential duration, and positively correlated with benefit generosity. He concluded, based on these estimates of the coefficients of the IUR and potential duration, that a UI program which provided thirty weeks for an IUR below 4 percent, forty weeks for an IUR between 4 and 6 percent, and fifty weeks above 6 percent would keep exhaustions around 15 percent.

Finally, a study by Arlene Holen and Stanley Horowitz (1974) on the effect of UI and UI eligibility enforcement on unemployment also reported some findings concerning UI exhaustions. One equation in their model of

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[1/] He defined exhaustions to be the exhaustion of all benefits, including extended benefits. In this report we differ from Hight by using exhaustions of regular UI only (i.e., excluding EB).

unemployment and unemployment insurance expressed the exhaustion rate as a function of the unemployment rate, average hourly earnings in manufacturing, the distribution of employment by industry, and maximum UI benefit duration. Their findings were consistent with those of Hight: the unemployment rate was positively correlated with the exhaustion rate, and while the coefficient was not significant, their measure of potential duration (maximum duration) was negatively correlated with the exhaustion rate.

2. Studies of Exhaustee Behavior

While the aim of the unemployment-insurance system is to provide income support to individuals during spells of involuntary unemployment, it has generally been agreed that the duration of benefits must be limited because of the potential work-disincentive effects of the UI program itself. This judgment leads to a consideration of the related issues of UI durations and exhaustion rates. If it is thought that the UI exhaustion rate is too high, duration policy can be altered to lower the rate and vice versa. Since the late 1940s, data have been collected in a number of studies on postexhaustion behavior which allow policymakers to reach a judgment as to whether the UI exhaustion rate is too high, too low, or about right. The most important data collected in these studies concerns the labor-market experiences of exhaustees. Resolutions to such issues as how quickly exhaustees become reemployed, and whether they withdraw from the labor force, are needed by policymakers to reach a judgment about the appropriate size of the exhaustion rate.

Data concerning these two questions are reported in Table I.1 for several exhaustee studies. Despite the fact that some of the studies were conducted when the unemployment rate was low and others when it was high,

TABLE I.1

PERCENT OF UI EXHAUSTEES EMPLOYED AND OUT OF LABOR FORCE
BY LENGTH OF TIME SINCE EXHAUSTION: SELECTED STUDIES

Study	Employed			Out of Labor Force		
	2 months	4 months	12 months	2 months	4 months	12 months
North Carolina 1956	26%	35%	n.a.	n.a.	n.a.	n.a.
West Virginia 1956	34	43	n.a.	n.a.	n.a.	n.a.
Pennsylvania						
1957-58	33	n.a.	20	17	n.a.	25
1966-67	33	37	37	22	36	35
Four Cities 1974-75						
Whites	17	27	39	15	15	26
Blacks and others	10	21	31	12	14	21

n.a = not available

Source: Merrill G. Murray, "The Duration of Unemployment Benefits," W.E. Upjohn Institute for Employment Research, January 1974, pp. 16-26; Walter Nicholson and Walter Corson, "A Longitudinal Study of Unemployment Insurance Exhaustees," MPR Project Report Series, no. 76-01, January 1976; and Walter Nicholson, Walter Corson, and Felicity Skidmore, "Experiences of Unemployment Insurance Recipients During the First Year After Exhausting Benefits," MPR Project Report Series, no. 76-14, August 1976.

they all found that a substantial percentage of exhaustees remain unemployed at various points in time during the year after exhaustion. For example, the 1966-67 Pennsylvania study, which was conducted during a period of low unemployment, found that after two months 45 percent were unemployed, and that after twelve months this figure dropped to only 28 percent. The 1974-75 study conducted in four cities during a period of high unemployment found comparable figures for the white sample--68 percent and 35 percent.

Thus, these data support the hypothesis that exhaustees of regular UI remain unemployed for a substantial period of time following the termination of benefits. They do not, however, offer an unambiguous argument for providing additional weeks of benefits, since it is possible that providing such benefits could create an incentive for an even further extension of unemployment spells. In Chapter V we present some evidence on the size of such an effect and reexamine the results of these exhaustee studies in light of that information.

C. OUTLINE OF THE REPORT

This report is divided into four additional chapters. In Chapter II we develop a theoretical model of the exhaustion of regular UI benefits which is suitable for estimation from aggregate data. We start with an analysis of the probability of an individual exhausting his or her UI entitlement. We focus on the length of the unemployment spell and the potential UI duration as the two principal determinants of that probability. We then turn to an examination of the empirical implementation of this model by specifying appropriate exogenous variables, discussing issues of aggregation, and briefly identifying salient econometric issues. We conclude the chapter with a discussion of the crucial role of the average potential duration of states for regular UI benefits in the model, and of ways in which the effects of policy parameters on that variable might be estimated.

Empirical results of the study are summarized in Chapter III. After a brief description of the data used, econometric estimates of the basic exhaustion-rate equation are presented with several variants on that equation. The general conclusion to be drawn from a comparison of these variants is that most coefficients of the exhaustion-rate equation are relatively robust to alternative specifications. After a summary of further econometric tests on the basic equation (which are presented in more detail in the appendix), Chapter III concludes with econometric equations for predicting the components of average potential duration.

The econometric results of Chapter III are used in Chapter IV to construct a simulation model of exhaustions and final regular UI payments over the business cycle. After first describing the structure of the model, the chapter is concerned mainly with testing its predictive properties. An appendix to Chapter IV provides a detailed listing of the equations used in the simulation model.

Little attention is given in Chapters III and IV to an assessment of the policy significance of the statistical results. Rather, as an aid to the reader, we have gathered all our policy conclusions together in Chapter V. Our discussion of those issues is divided into three sections. First, in section A, we summarize our findings about the effects of economic factors on exhaustion rates. Section B presents a similar summary of the estimated effects of UI policies on exhaustion rates. As in our theoretical model, we make a distinction between those UI policies that affect exhaustion rates by changing the mean length of unemployment spells, and those that have an effect by changing average potential durations. The final section of Chapter V utilizes our simulation model to evaluate specific triggering formulas for extended benefits. Additional ways in which the simulation model might be used for policy purposes are also briefly mentioned.

II. A THEORY OF EXHAUSTION OF REGULAR UI BENEFITS

INTRODUCTION

In this chapter we develop a theory of the exhaustion of regular UI benefits which is suitable for empirical estimation. We begin, in section A, by sketching a model of the probability that an individual will exhaust his or her UI entitlement. That model stresses the interaction between factors that determine the length of an individual's spell of unemployment and factors that determine his or her potential UI duration. Variables that increase the length of unemployment spells would, ceteris paribus, be expected to increase exhaustion rates, whereas variables that increase potential UI durations should have the opposite effect. We outline our a priori expectation about the direction of the most important of these effects.

In section B we discuss issues related to the empirical implementation of the exhaustion model. Our presentation is divided into two parts. First, we examine issues in implementation that would arise if the model were to be estimated by using individual data on exhaustions. These primarily include questions of functional form and variable specification. Because we intend to estimate the model of exhaustion rates from aggregate cross-sectional, time-series data a number of additional issues of empirical implementation must also be addressed. These are taken up in the second part of section B. The questions that are examined concern special independent variables needed for the aggregate model, appropriate econometric techniques for pooled data, and possible simultaneity in the exhaustion relationship.

Section C of this chapter addresses the empirical implementation of the exhaustion-rate model from a perspective somewhat different than the theoretical discussion of section B. In this section we examine how the model can be made most useful for policy purposes. The distinction between policy instruments and other independent variables is made, and the importance of modeling the precise way in which actual policy decisions affect exhaustion rates is stressed. Particular attention is focused on the ways in which provisions of state laws affect UI beneficiaries' potential durations.

A. THEORETICAL MODEL OF EXHAUSTION

The length of a completed spell of unemployment for a laid-off worker (S_i, measured in weeks) can be considered a random variable whose expected value is determined by personal characteristics of the worker (X_i), the nature of the labor-market environment within which the worker must search for a new job (L_i), and governmental policy parameters (U_i) that provide incentives or disincentives to prolong job search:

$$E(S_i) = f(X_i, L_i, U_i). \tag{II.1}$$

For given values of X_i, L_i, and U_i the distribution of S_i might resemble that illustrated in Figure II.1. Changes in X_i, L_i, or U_i shift the entire distribution of S_i by changing its mean.[1/]

[1/] It is possible that X_i, L_i, and U_i might also affect other moments of the distribution of S_i. We do not consider that possibility here but will examine it briefly in our discussion of empirical implementation.

Laid-off workers who file for unemployment insurance benefits and subsequently receive a first-payment are eligible to collect (full) benefits for a maximum number of weeks (D_i). This "potential duration" is determined by statute and will generally depend on the worker's employment history and on UI formulas.[1/] Once a worker has been laid off, he or she has little if any control over D, and that variable can be considered exogenous to S_i. A particular value for D_i is illustrated in Figure II.1. It has historically been true that $D_i > E(S_i)$ for most types of workers and for most economic environments.

Exhaustion of UI benefits occurs when $S_i \geq D_i$. The individual's probability of exhaustion is denoted by:

$$R_i = \text{Prob}(S_i \geq D_i) = \int_{D_i}^{\infty} g(S_i)DS, \qquad \text{(II.2)}$$

where g is the probability density function for S_i. The probability R_i is illustrated by the shaded area in Figure II.1. It is obvious from the figure that

$$\partial R_i/\partial E(S_i) > 0$$

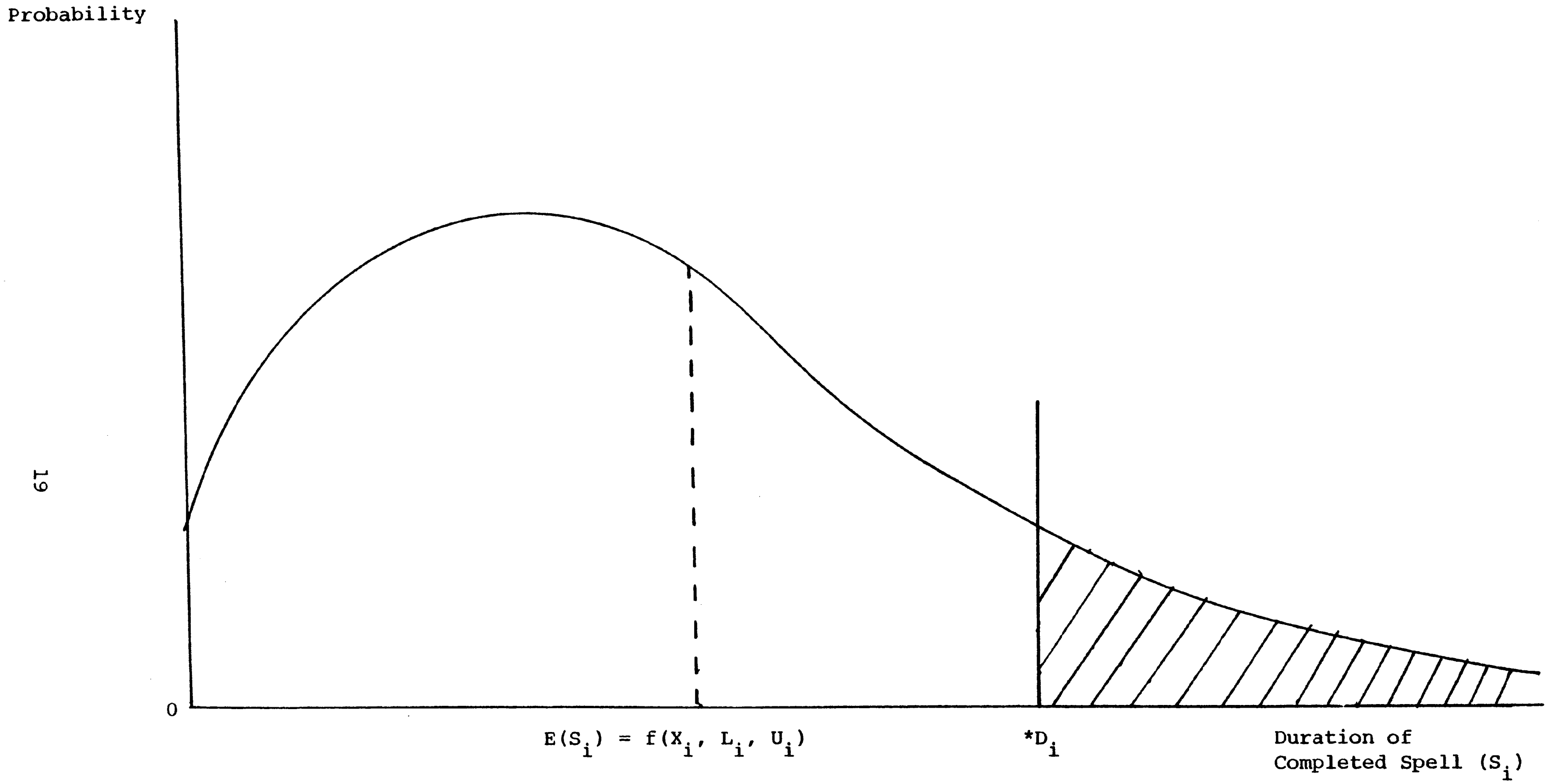
and (II.3)

$$\partial R_i/\partial D_i < 0.$$

[1/] In actuality it is possible for an individual to collect his or her UI benefit entitlement during more than one unemployment spell. We do not treat this complication explicitly in our model. However, the model's conclusions would not be changed much if the random variable S_i were regarded as total weeks unemployed during an individual's benefit year.

FIGURE II.1

THEORETICAL MODEL OF AN INDIVIDUAL'S EXHAUSTION OF UI BENEFITS



$^*D_i$ is the maximum UI duration for which the individual is eligible. The curve shows the probability distribution of completed spells (S_i), and the shaded area represents the probability of the exhaustion of benefits.

It is also likely that for given values of $E(S_i)$ and D_i,

$$\frac{\partial R_i}{\partial \sigma^2_{S_i}} > 0, \qquad (II.4)$$

where $\sigma^2_{S_i}$ is the variance of the distribution of the individual's unemployment spell.

B. EMPIRICAL IMPLEMENTATION

In this section we examine questions relating to the specification of the exhaustion-rate equation (that is, equation II.2) in a way that permits it to be estimated empirically. Three specific issues are addressed:

1. Proper specification of the probability density function, g (and, hence, of the appropriate functional form for equation II.2).
2. Identification of those variables that are believed to affect $E(S_i)$ and the anticipated direction of such effects.
3. The ways in which equation II.2 may be aggregated across individuals together with an analysis of the implications of such aggregation for the estimation strategy employed.

1. Probability Density

The precise shape of the distribution of individual's unemployment spells is not known. A wide variety of functional forces have been postulated, but most of these are intractable for an estimation of equation II.2 by using large quantities of data and sophisticated econometric techniques. Two polar cases that may serve to place bounds on the proper functional form are the uniform distribution and the exponential distribution. Since the actual form of the distribution of unemployment spells across individuals lies "between" these extremes,[1/] and since these functions are easy to handle mathematically, we will briefly examine them here.

If unemployment spells follow the uniform distribution

$$g(S_i) = \alpha_i \quad \text{for } 0 \leq S_i \leq \frac{1}{\alpha_i} \tag{II.5}$$

(where α_i is some constant), then

$$\frac{\partial R_i}{\partial E(S_i)} = - \frac{\partial R_i}{\partial D_i} = \alpha_i . \tag{II.6}$$

That is, R_i is linear in $E(S_i)$ and D_i, and a linear functional form (with $E(S_i)$ and D_i as determinants of R_i) would be appropriate.

If, on the other hand, unemployment spells follow the exponential distribution

$$g(S_i) = \lambda e^{-\lambda S_i}, \tag{II.7}$$

then

$$E(S_i) = \frac{1}{\lambda}$$

[1/] Since the relative frequency of unemployment spells diminishes as longer spells are examined, the uniform distribution is inappropriate. On the other hand, as Kaitz (1970) shows, the exponential distribution declines too rapidly in its upper tail. That distribution implies a constant probability of remaining unemployed one more week across the entire distribution, whereas it appears that, in fact, this probability increases with the length of the unemployment spell.

and

$$\ln R_i = -\lambda D_i = -\frac{D_i}{E(S_i)}. \tag{II.8}$$

In this case, equation (II.8) might be estimated directly, or the functional form in the equation might be relaxed, to allow $\ln R_i$ to depend linearly on $E(S_i)$ and D_i (while dropping the restriction that the coefficient of $D_i = -1$) and, possibly, on interactions among the terms:

$$\ln R_i = \beta_o + \beta_i E(S_i) + \beta_2 D_i, \quad \text{where } \beta_i > 0,\ \beta_2 < 0. \tag{II.9}$$

In Chapter III we follow this strategy. In the appendix to that chapter, however, we present evidence on a variety of functional forms.

2. Determinants of $E(S_i)$

In section A we identified three general influences on the mean length of an individual's unemployment spell: (1) the individual's personal characteristics, (X_i); (2) the economic envircnment, (L_i); and (3) governmental policy parameters (U_i). Table II.1 records several variables in each of these categories and the direction of their expected effect on $E(S_i)$ (and, hence, on the probability of exhaustion of regular UI benefits). Most of the expected effects are based both on strong a priori theoretical arguments and on a considerable body of empirical data. That is true, for example, of the basic demographic (age, sex, race) assumptions. A few of the entries in Table II.1, however, do merit explicit attention. The presumed negative effect of skill and labor-market-attachment variables derives from the assumption that

TABLE II.1

DETERMINANTS OF MEAN LENGTHS OF COMPLETED UNEMPLOYMENT SPELLS

Variable	Direction of Effect in $E(S_i)$
Personal Characteristics	
Age	+
Sex (1 = male, 0 = female)	-
Race (1 = white, 0 = other)	-
Education	-
Other Skills	-
Strength of Labor Market Attachment	-
Seasonal Worker	-
Construction Worker	-
Manufacturing Worker	-
Economic Environment	
Unemployment Rate	+
Extent of Labor Market Information	-
Policy Parameters	
Wage Replacement on UI	+
Enforcement of "Availability for Work" test	-
Benefit eligibility beyond Regular UI	+

there are two a priori reasons to expect them to be an unimportant source of bias. First, state UI laws react to UI outcomes only with a substantial lag. Hence, in a quarterly model, UI policies within a state may reasonably be considered exogenous to the contemporaneous exhaustion rate. Second, if differences in exhaustion rates among states lead to the adoption of differing UI policies, the extent of this type of simultaneity bias can be estimated by using state dummy variables in the pooled regressions. We discuss this point more formally in Chapter III and provide rough quantitative estimates of such biases which indicate they are not substantial.

C. ADAPTATION OF THE EXHAUSTION RATE MODEL FOR POLICY PURPOSES

Thus far, our development of a model of UI exhaustion rates has been restricted by theoretical concerns. Our theory identified two types of UI policy variables that play important roles in the model. The first type, recorded in Table II.1, consists of those variables under the control of UI policymakers that might be expected to affect the length of individuals' unemployment spells. These include measures of the enforcement of UI laws and of the attractiveness of UI over employment. The second type of UI policy variable that is hypothesized to directly affect the exhaustion rate is the average potential duration (APD) of regular UI benefits. This variable differs from those of the first type both in the way it enters into our model and in the extent to which it is subject to direct policy control. Its importance to our model is obvious. Whereas there may be some doubt as to the quantitative magnitude of some of the effects hypothesized in Table II.1, there can be no doubt about the importance of the APD variable. At the same time, however, it should be recognized that APD is not (except in "uniform duration" states) directly determined by state law. Rather, state laws specify various eligibility and entitlement provisions that interact with the characteristics of a state's labor market to determine recipients' APD. To treat APD as a direct policy instrument could obscure the precise mechanisms by which state laws affect APD and, ultimately, exhaustion rates. To investigate these mechanisms directly, our model examines the determinants of APD directly for variable-duration states.[1/] Specifically, APD is assumed to be determined by the interaction of three variables: the state's maximum potential duration (MPD); the fraction of recipients in the state eligible for that maximum potential duration (PMAX); and the average potential duration of those recipients not eligible for MPD (APDO). The precise press of this relationship is given by

$$APD = PMAX * MPD + (1 - PMAX) * APDO. \qquad (II.10)$$

Of the three determinants of APD, only MPD is directly under a state's control. PMAX and APDO are assumed to be determined by MPD, by state eligibility and entitlement provisions (UI), and by variables that characterize the economic environment:

[1/] For uniform-duration states, APD is fixed at maximum potential duration, and there is no variability in that figure. In our regression model in Chapter III we specify uniform-duration states' APDs in this way, but we also add a binary variable for such states to test for the possibility that the effects of uniform-duration laws are not completely captured by our APD and dispersion variables.

$$PMAX = h_1 \ (MPD, UI, E)$$
$$APDO = h_2 \ (MPD, UI, E) \qquad (II.11)$$

Specific, hypothesized determinants of PMAX and APDO are recorded in Table II.2, together with their anticipated direction of effect. Most of these a priori expectations are intuitively clear. MPD has a negative effect on PMAX (given entitlement, fewer individuals qualify for the maximum) and a positive effect on APDO (a higher maximum permits higher potential durations for those not at the maximum). The UI entitlement fraction (as a percent of base-period wages) has a positive effect on both PMAX and APDO, which arises from the mechanical relationship between entitlements and duration. Although it is possible that the nature of a state's UI eligibility and weekly benefit computation formulas (and, possibly, their interaction with entitlement) might have important impacts on PMAX and APDO, the UI literature has not provided any general conclusions about the precise magnitude of such effects.[1/] Hence, although we will investigate that relationship in detail in Chapter III, we are unable to specify an a priori direction.

A number of features of the economic environment should affect PMAX and APDO. Most important, the insured unemployment rate, because it reflects the strength of the labor market, should be positively correlated with both variables. During recessions, UI recipients will tend to have higher base-period earnings and weeks of work than during

[1/] Some estimates of the effects of benefit entitlement formulas on potential durations are reported in "Weekly Benefit Amounts and Normal Weekly Wages of Unemployment Insurance Claimants" (U.S. Department of Labor, Unemployment Insurance Service, August 1977).

TABLE II.2

DETERMINANTS OF PERCENT OF RECIPIENTS
ELIGIBLE FOR MAXIMUM POTENTIAL DURATION
(PMAX), AND AVERAGE DURATION OF THOSE NOT AT THE MAXIMUM (APDO)

	Direction of Effect on PMAX	Direction of Effect on APDO
Maximum Potential Duration (MPD)	-	+
UI Eligibility and Entitlement		
Qualification Provisions	a	a
Entitlement Fraction	+	+
Weekly Benefit Computation	a	a
Economic Environment		
Insured Unemployment Rate	+	+
Percent in Manufacturing	+	+
Percent in Construction	a	a
Percent Male	+	+

a = no direction predicted (see text).

5 Errors follow a components-of-variance structure with unique time and units components.

In addition, the use of state dummy variables (as in Table III.7) can be regarded as a fifth type of generalized least-squares procedure in which error terms are assumed to have differing mean values across states. Our discussion of generalized least squares will focus on comparing coefficient estimates derived from employing these procedures to those derived by ordinary least squares.

Results in Table III.9 tend to support the following generalizations made in section B of this chapter:

1. For most variables, coefficient point estimates were robust to the alternative generalized least-squares procedures.

2. The coefficients of DIS, QUALWK, and WR were relatively more greatly affected by the estimation procedures than were those of the other variables. Even for these variables, however, their estimated direction of influence was unchanged.

3. The coefficient of IUR was raised substantially by all of those procedures that in some way controlled for cross-sectional differences.

4. Although standard errors are not shown in Table III.9, these were found to be roughly similar under the various estimation procedures. Hence, our inferences about the statistical significance of the coefficients in the basic regression were unchanged.

4. Simultaneity

Simultaneous equations bias occurs whenever an equation's dependent variable exhibits "feedback" effects by being a significant determinant of an independent variable in the equation. In the context of estimating the exhaustion-rate equation, there are two ways in which it could be theorized that such a feedback could arise. First, the exhaustion rate might be definitionally related to one of the independent variables.

The most important variable for which such a correlation might exist is the insured unemployment rate (IUR). Because higher exhaustion rates might, ceteris paribus, result in lower IURs (because exhaustees are not counted among the insured unemployed), this correlation would cause the coefficient of IUR, as estimated by ordinary least squares, to be negatively biased. A second way in which coefficients of independent variables might be biased relates to the formulation of state laws. If state UI laws are altered in response to high exhaustion rates, a correlation would exist between the parameters that reflect those laws and exhaustion rates, which would result in the coefficients of the parameters being biased in our equations.

Our statistical investigation of the simultaneity question focused on the first of these sources of bias. There were three reasons why we gave only brief attention to possible simultaneity between state laws and exhaustion rates. First, what simultaneity there is in the relationship probably occurs with a substantial lag attributed to the legislative process. That lag, combined with a two-quarter lag used for many of the independent variables in the model, suggests that contemporaneous simultaneity may be of minor importance. Second, because most coefficients of the model were changed only slightly when state dummy variables were added to the model (see Table III.7), there is little evidence of a long-term correlation between significantly different exhaustion-rate experiences and those parameters of state laws included in the regression. Finally, modeling of state legislative procedures in the necessary detail would have involved data needs and conceptual developments beyond the scope of the present project.[1/]

[1/] Simultaneity between EB and exhaustion rates was explicitly considered, since it is possible (even given the EB triggering mechanisms) that federal extended benefits programs are legislated partially in response to exhaustion rates. In examining several simultaneous specifications of the EB variable, we found little evidence that such an effect was biasing its coefficient of the EB dummy.

In order to investigate possible simultaneity between exhaustion rates (R) and measured IURs, a two-equation model was postulated which explained R and IUR. Exogenous variables used in the exhaustion-rate equation were those used in our basic equation. It was hypothesized that some of these variables (DIS, DURD, QUALWK, and NODIST) did not affect the IUR, but that the other exogenous variables (most notably, the national unemployment rate and the percent of employees covered by UI) did affect the IUR. Results for estimating the exhaustion-rate equation by two- and three-stage least squares as part of such a two-equation simultaneous system are reported in Table III.10. While the coefficient of the IUR in those equations was somewhat higher than in those estimated by ordinary least squares, the change is sufficiently small to suggest that simultaneity between exhaustion rates and the IUR was not a major cause of bias. Other coefficients in Table III.10 were also relatively little changed from those estimated by ordinary least squares in Table III.3.

5. Conclusions

The various econometric tests reported in this section led us to believe that the coefficients of the basic exhaustion-rate equation were relatively robust to alternative specifications and estimation procedures. The major contrary finding related to the substantial increase in the coefficient of the IUR when cross-sectional controls for states were entered into the regression. Such a finding was not unexpected in view of the difficulty in interpreting the coefficient of the IUR in a pooled regression. It suggested that special care be taken in using the coefficient of the IUR in assessing potential

TABLE III.10

SIMULTANEOUS EQUATIONS ESTIMATES OF
THE EXHAUSTION RATE EQUATION
FOR THE PERIOD 1965.II TO 1974.IV

(Dependent Variable lnR)

Independent Variable	Two-stage Least Squares		Three-stage Least Squares	
	Coefficient	t Statistic	Coefficient	t Statistic
APD	-0.0498	-13.3	-0.0491	-13.0
DIS	0.00144	3.8	0.00147	3.9
DURD	-0.167	- 6.0	0.166	- 6.0
PMALE	-0.00680	- 5.9	-0.00595	- 5.2
PMANF	-0.0121	-17.1	-0.0120	-16.8
PCONST	-0.0133	- 8.2	-0.0131	- 8.1
NODIST	-0.130	- 5.7	-0.131	- 5.7
QUALWK	-0.0201	- 3.3	-0.0184	- 2.9
IUR	0.0848	13.2	0.0821	5.8
WR	0.664	4.0	0.650	3.9
NOTAVL	-0.0152	- 7.8	-0.0155	- 7.9
REFUSE	-0.143	-10.3	-0.143	-10.3
EB	0.255	7.3	0.314	9.0
Q1	0.0084	0.3	0.0209	0.7
Q2	0.236	10.4	0.235	10.2
Q3	-0.422	-15.8	-0.436	-16.3
CONSTANT	0.958	7.3	0.908	6.8
	R^2 = .62		R^2 = 61	
	Standard Error: 0.353		Standard Error: 0.357	

policies. When interest is centered on policies that are altered over the business cycle, the results indicated that the higher, "time series" coefficient for IUR should be used for analysis.

Other than these findings that related to the coefficient of the IUR, most of our econometric tests indicated a general stability in the least-squares coefficients. Although some coefficients were changed when structural differences, generalized least squares, and simultaneity bias were investigated, the quantitative size of those changes was usually small. Hence, we decided to use the basic ordinary least-squares regression equation for most of the analysis which follows.

D. AVERAGE POTENTIAL DURATION

In all of our regressions on the exhaustion rate the average potential-duration variable (APD) had a significantly negative coefficient. That finding was consistent with our theory, and it suggested that a principal mechanism by which states may be able to affect exhaustion is policy actions that affect APD. But, as pointed out in Chapter II, APD is not, in general, under the direct control of state policymakers (except in uniform-duration states). Rather, states adopt complex combinations of qualification provisions, entitlement formulas, weekly benefit computations, and maximum-duration provisions that interact with labor-market characteristics in the state to determine APD. To study that interaction we made use of the algebraic identity:

$$APD = PMAX * MPD + (1-PMAX) * APDO, \qquad \text{(III.9)}$$

where

MPD = maximum potential duration,

PMAX = proportion of UI recipients at MPD, and

APDO = average potential duration of those not at MPD.

Of these determinants, only MPD is directly under the control of policymakers. In this section we present results of our examination of the determinants of PMAX and APDO. In Chapter V we use these results together with our exhaustion-rate equation to examine the effect of potential changes in state law on APD and, hence, on exhaustion rates.

Our basic regression equations for APDO and PMAX are reported in Tables III.11 and III.12, respectively. Before discussing these estimates we should describe the dependent variables used. Published data are not available on a quarterly basis for the percent of UI recipients who are eligible for a state's maximum potential duration for the entire period of our study. Instead, data exist on the percent eligible for maximum benefits and duration, and on the percent eligible for the state's maximum weekly benefit amount. Our procedure was to use as a measure of PMAX the proportion eligible for maximum benefits and duration, and to examine the consequences of controlling for the percent at maximum benefits (PMBEN) in the regressions. The variable APDO was then defined from equation III.9 by using our measure of PMAX and published data on MPD and APD. Summary characteristics of the variables calculated in this way were (for the variable duration states only):

Variable	Mean	Standard Deviation
APDO	21.8 weeks	2.9 weeks
PMAX	35.5%	14.1 %

(Based on 1,599 observations)

TABLE III.11

AVERAGE POTENTIAL DURATION OF OTHERS
FOR THE PERIOD 1965.II TO 1974.IV

(Dependent Variable: APDO)

Independent Variable	Coefficient	Standard Error	t Statistic
PCONST	-0.0273	0.0101	-2.7
PMANF	0.0123	0.00452	2.7
PMEN	0.00444	0.00744	0.6
ENTITL	0.0449	0.00354	12.7
IUR	0.365	0.0354	10.3
MPD	0.626	0.0205	30.5
Q1	-0.486	0.186	-2.6
Q2	-0.312	0.150	-2.1
Q3	0.192	0.157	1.2
CONSTANT	2.042	0.772	2.6

R^2: .472 Standard Error: 2.14

R^2 (Adj.): .469 F(9,1589): 157.7

Source: Estimated by OLS

TABLE III.12

PERCENT AT MAXIMUM BENEFIT AND DURATION
FOR THE PERIOD 1965.II TO 1974.IV

(Dependent Variable PMAX)

Independent Variable	Coefficient	Standard Error	t Statistic
PCONST	-0.130	0.0576	- 2.3
PMANF	0.0679	0.0257	2.6
PMEN	0.621	0.0423	14.7
ENTITL	0.0754	0.0201	3.8
IUR	0.384	0.201	1.9
MPD	-1.433	0.117	-12.3
Q1	-4.242	1.056	- 4.0
Q2	-1.142	0.854	- 1.3
Q3	1.644	0.895	1.8
CONSTANT	34.72	4.390	7.9

$R^2 = 0.26$ Standard Error = 12.2

R^2 (Adj.) = 0.26 $F(9,1589) = 63.5$

Source: Estimated by OLS

Because, by definition, uniform-duration states have PMAX = 1 and APD = MPD, no investigation was made of average potential duration in such states. Rather, we treated the policy decision to adopt uniform-duration standards as equivalent to setting APD directly. In Chapter V we examine some of our estimates of the implications of adopting such a policy.

The coefficient estimates reported in Tables III.11 and III.12 generally supported the hypotheses laid out in Chapter II (see Table II.2). Both entitlement fractions (ENTITL, measured in percent) and insured unemployment rates had positive effects on APDO and PMAX, as did the percent of insured unemployed in manufacturing. The percent male variable had a significant positive effect only on PMAX. Also, as hypothesized, MPD had a significant positive effect on APDO and a significant negative effect on PMAX. The overall goodness of fit of the regressions was reasonably good, although, especially in the PMAX equation, a considerable portion of the variance remained unexplained. As shown in Table III.13, the addition of PMBEN to the PMAX regression substantially improved its fit. The coefficient of PMBEN implied that 62 percent of new UI recipients who are eligible for maximum benefits are also eligible for maximum duration.[1/] While the coefficients in Table III.13 were qualitatively similar to those in Table III.12, the quantitative size of some estimates substantially changed. Because we were interested in the "gross" impact of such variables as IUR and MPD on PMAX, we chose to use the estimates from Table III.12 (together with various additions) in most of our subsequent analysis.

[1/] That figure is somewhat below the figure implied by taking the ratio of average PMAX (35.5 percent) to average PMBEN (49.1 percent).

TABLE III.13

PERCENT AT MAXIMUM BENEFIT AND DURATION
(CONTROLLING FOR PERCENT AT MAXIMUM BENEFIT)
FOR THE PERIOD 1965.II TO 1974.IV

(Dependent Variable: PMAX)

Independent Variable	Coefficient	Standard Error	t Statistic
PCONST	-0.129	0.0424	-3.0
PMANF	0.0293	0.0189	1.5
PMEN	0.217	0.0331	6.5
ENTITL	0.0963	0.0148	6.5
IUR	1.154	0.150	7.7
MPD	-0.927	0.0871	-10.6
PMBEN	0.620	0.0170	36.5
Q1	-3.075	0.779	-3.9
Q2	-0.0409	0.631	-0.1
Q3	2.317	0.661	3.5
CONSTANT	12.02	3.297	3.6

R^2: .600 Standard Error: 8.973

R^2 (Adj.): .597 F(10,1588): 238.1

Source: Estimated by OLS

A number of econometric variants of the APDO and PMAX equations were examined. One general conclusion of these investigations was that, contrary to our earlier findings on the exhaustion-rate equation, coefficients were not very stable under alternative specifications. Most important, there appeared to be complex interactions with states' qualification provisions and weekly benefit formulas that were not amenable to modeling in a simplified regression context. To disentangle all of these influences would require a major research agenda--probably using a variety of methodologies. While such an agenda would be beyond the scope of the present study, three conclusions did emerge from our examination which merit reporting:

1. The strong positive effect of entitlements on both APDO and PMAX held up throughout our investigation, and the size of its coefficient changed only slightly. That was true both when adjustments were made for the size of average UI benefits and average wages in a state and when a series of state dummy variables were added to the regressions. Hence, we concluded that the overall effect of changing entitlement fractions could reasonably be modeled by using the equations in Tables III.11 and III.12.

2. Qualification provision interactions within the APDO and PMAX equations appeared to be substantial and remained difficult to interpret. States that use a multiple of the WBA formula for qualification appeared to be somewhat lower in their levels of both APDO and PMAX (and, ceteris paribus, in their level of APD) than were other states.

3. Weekly benefit formulas posed similar complex interactions. There was some evidence that states that use a high-quarter earnings formula obtain a smaller overall impact of entitlement fractions on APD than do other states.

APPENDIX TO CHAPTER III

FURTHER ECONOMETRIC RESULTS

INTRODUCTION

In this appendix we present a few additional econometric results that supplement our basic analysis of Chapter III. Most of these results have already been discussed in Chapter III, and our intent here is primarily to offer a more complete documentation of them. For that reason we provide only a brief interpretive discussion. Our presentation is divided into five sections that parallel the analysis in sections C and D of Chapter III. These sections cover:

A. Functional Form

B. State-by-State Estimates

C. Uniform and Variable Duration Estimates

D. Extended Benefits Effects

E. Average Potential Duration

A. FUNCTIONAL FORM

In Table III.6 we used the Box-Cox Transformation to test for the appropriate functional form of the exhaustion-rate equation. Table IIIA.1 provides some additional detail on those tests. Three regressions are reported: the "basic" log-linear form; a linear equation; and the best Box-Cox transform (λ = .4). Three findings are apparent in the table. First, all of the equations fit the data well. Second, with the exception of the QUALWK variable in the linear regression, each variable's coefficient has the same sign in all the regressions reported, and the statistical significance of the coefficients is quite similar across equations.

TABLE III.A.1

TESTS OF ALTERNATIVE FUNCTIONAL FORMS FOR THE EXHAUSTION RATE EQUATION
FOR THE PERIOD 1965.II TO 1974.IV

	Linear Form		Box-Cox Transform (λ = 0.4)		Log Form	
Independent Variable	Coefficient	t Statistic	Coefficient	t Statistic	Coefficient	t Statistic
APD	-0.0147	-15.9	-0.0150	-17.3	-0.0150	-15.2
DIS	0.00027	2.7	0.003	3.4	0.00157	4.2
DURD	-0.040	- 5.5	-0.041	- 6.2	-0.181	- 6.7
PMALE	-0.0017	- 5.9	-0.0020	- 7.7	-0.00773	- 7.1
PMANF	-0.0030	-16.9	-0.0030	-18.5	-0.0110	-16.5
PCONST	-0.0034	- 5.6	-0.0025	- 6.6	-0.0118	- 7.6
NODIST	-0.0460	- 7.8	-0.041	- 7.6	-0.130	- 6.3
QUALWK	0.00003	0.2	-0.0027	- 1.8	-0.0232	- 3.9
IUR	0.022	13.6	0.021	13.9	0.0778	12.9
WR	0.097	2.3	0.103	2.6	0.834	5.2
NOTAVL	-0.0032	- 6.6	-0.0032	- 7.2	-0.0117	- 6.6
REFUSE	-0.025	- 7.2	-0.028	- 8.6	-0.125	- 9.5
EB	0.033	4.7	0.039	5.9	0.145	5.4
Q 1	0.016	2.3	0.0042	0.7	-0.0586	- 2.3
Q 2	0.065	11.0	0.062	11.5	0.226	10.2
Q 3	-0.069	- 9.9	-0.081	-12.8	-0.415	-16.0
CONSTANT	0.842	25.1	0.058	1.9	0.781	6.2
R^2	0.56		0.62		0.64	
Standard Error	0.092		0.084		0.345	
F (16,1933)	153.5		198.4		210.6	

Finally, the quantitative magnitudes of the estimated effects of the independent variables are also similar once account is taken of the particular way in which the dependent variable was transformed. This may be seen either by working out the derivatives directly, or by using the ratios of some of the coefficients to calculate offsetting changes in the independent variables that are needed to keep the exhaustion rate constant. As an example of the second type of analysis, a 1 percentage point increase in the IUR can be offset by a 1.50 week increase in APD according to the linear equation, by a 1.40 week increase according to the best Box-Cox equation, and by a 1.48 week increase according to the log-linear equation. Similarly, the positive estimated effect of the EB program on the average duration of unemployment spells is 2.24 weeks, 2.60 weeks, or 2.76 weeks according to the linear, Box-Cox, or log-linear equations, respectively. As mentioned in Chapter III, the wage-replacement effect was one of the most unstable of those we estimated, and that result holds true in Table IIIA.1, as well. There, the calculated effect of a 10 percent increase in WR on the duration of unemployment spells ranged from 0.66 weeks (in the linear case) to 1.61 weeks (in the log-linear case). This finding, therefore, again suggests that caution is needed in making precise "disincentive effect" estimates from our equations.

B. STATE-BY-STATE ESTIMATES

Table IIIA.2 reports estimates of the exhaustion-rate equation for five states: California, Florida, Michigan, New York, and Pennsylvania. As the table shows, the results of such estimates were extremely varied. Probably the most consistent coefficient estimate was that for the IUR. Those estimates were all reasonably close to our pooled estimate when

TABLE III.A.2

STATE-BY-STATE ESTIMATES OF THE EXHAUSTION RATE EQUATION--
DEPENDENT VARIABLE lnR--
FOR THE PERIOD 1965.II TO 1974.IV

Independent Variable	California	Florida	Michigan	New York	Pennsylvania
APD	-0.059	-0.0008	0.163	0.0088	-0.046
	(-0.6)	(-0.0)	(1.5)	(0.1)	(-0.7)
PMALE	0.0086	0.0068	-0.0088	0.0028	-0.0012
	(1.2)	(0.6)	(-1.0)	(0.3)	(-0.2)
PMANF	-0.0011	-0.018	-0.012	-0.0164	-0.0116
	(-0.2)	(-2.1)	(-1.1)	(-3.1)	(-2.07)
PCONST	-0.014	0.008	-0.0047	-0.0245	-0.0245
	(-1.8)	(0.8)	(-0.2)	(-1.8)	(-2.0)
IUR	0.091	0.180	0.156	0.229	0.176
	(4.2)	(2.3)	(3.0)	(4.9)	(3.6)
WR	0.477	-6.46	-4.3	-1.59	0.60
	(0.3)	(-2.7)	(-1.7)	(-0.9)	(0.9)
NOTAVL	0.018	0.082	0.064	0.030	0.0471
	(2.4)	(3.1)	(1.2)	(-4.5)	(1.5)
REFUSE	-0.19	-0.582	-0.451	-0.033	-0.324
	(-3.4)	(-3.8)	(-1.1)	(0.2)	(-2.3)
EB	0.028	0.163	-.035	0.049	0.087
	(0.8)	(1.4)	(0.3)	(0.5)	(1.2)
Q 1	0.025	-0.725	-0.774	-0.586	-0.554
	(0.4)	(-4.9)	(-3.7)	(-6.4)	(-7.4)
Q 2	0.154	-0.156	0.433	-0.062	-0.041
	(3.2)	(-2.2)	(2.7)	(-0.9)	(-0.6)
Q 3	-0.354	-0.297	-0.386	-0.337	-0.379
	(-8.4)	(-2.0)	(-2.0)	(-2.1)	(-3.0)
CONSTANT	-0.714	1.016	-3.25	-0.469	-0.163
	(-0.3)	(0.64)	(-1.3)	(-0.1)	(-0.08)
R^2	0.97	0.94	0.90	0.94	0.94
Standard Error	0.056	0.138	0.205	0.121	0.107
F (12,26)	87.0	32.9	21.7	35.2	36.5

NOTE: t Statistics in parentheses.

dummy variables were included (that coefficient was reported in Table III.7 as 0.176--roughly the mean of the coefficient in Table IIIA.2). Other relatively similar estimates in the table were those for the coefficients of PMANF, PCONST, and REFUSE. The variability of many of the other coefficients and their statistical insignificance can probably be attributed to the small within-state variance exhibited by some of the variables (notably, APD, PMALE, and WR), and by collinearity between the variables over the time period examined. While it would undoubtedly be possible to "improve" individual states' exhaustion-rate equations through various alternative specifications, such an examination would probably have had a small payoff for the basic purposes of this study.

C. UNIFORM AND VARIABLE DURATION ESTIMATES

In Table III.8 we showed that the hypothesis that uniform- and variable-duration states had structurally similar exhaustion-rate equations could be rejected at the .01 level by the usual F Test.[1/] Table III.3 reports the results for estimating the equation separately over each type of state. Although there were some major differences in the coefficient estimates for these two equations, the overall impression was one of similarity. Most notably, the coefficients of APD, PMANF, IUR, NOTAVL, REFUSE, and EB were reasonably close in both equations. The lower statistical significance of the coefficients in the uniform-duration equation was probably accounted for by its much smaller sample size and by the small variance of some of the independent variables within that sample. The major differences in the two equations were in the coefficients of NODIST and WR.

[1/] We also showed, however, that this structural difference was resonably well controlled for by allowing differential intercepts in a pooled regression.

TABLE III.A.3

EXHAUSTION RATE EQUATION IN UNIFORM AND VARIABLE DURATION STATES--
DEPENDENT VARIABLE lnR--
FOR THE PERIOD 1965.II TO 1974.IV

	Uniform Duration		Variable Duration	
Independent Variable	Coefficient	t Statistic	Coefficient	t Statistic
APD	-0.0434	-1.7	-0.0560	-15.8
DIS	0.0418	0.2	0.00165	4.3
PMALE	-0.00648	-2.1	-0.00839	- 7.0
PMANF	-0.0111	-4.5	-0.0114	-15.5
PCONST	-0.0077	-1.6	-0.0121	- 7.3
NODIST	-0.482	-4.7	-0.146	- 6.1
QUALWK	-0.0467	-2.0	-0.0220	- 3.3
IUR	0.113	4.2	0.0784	12.5
WR	-1.459	-2.4	0.841	4.5
NOTAVL	-0.0151	-3.3	-0.0116	- 5.7
REFUSE	-0.119	-4.3	-0.134	- 8.3
EB	0.200	2.9	0.126	4.3
Q 1	-0.206	-3.2	-0.0384	- 1.4
Q 2	0.0737	1.3	0.252	10.5
Q 3	-0.445	-6.5	-0.0405	-14.5
R^2	0.68		0.62	
Standard Error	0.300		0.348	
F (d.f.)	32.4 (15,233)		186.6 (15,1685)	

The first of these differences was probably related to peculiarities unique to the one uniform-duration state (West Virginia) without distributional requirements for UI qualification. The statistically significant WR effect in uniform-duration states is more difficult to explain. It may have actually reflected "negative" disincentive effects in uniform-duration states, but it is more likely that the coefficient was a regression "artifact" that resulted from the restriction of the sample to only a few states with a specific type of duration formula. Such a conclusion is, of course, speculative, and it is clear that the estimated differences between uniform- and variable-duration states deserve additional research attention.

D. EXTENDED BENEFITS EFFECTS

The F tests reported in Table III.8 also suggested that the exhaustion-rate equation should be estimated separately over periods for which extended benefits programs were and were not in effect. Table IIIA.4 presents the results of such separate estimates. As the table shows, many of the independent variables had quite similar coefficients during the two periods. Two major differences in the coefficients which might be pointed out concern the IUR and WR. The first difference might have been expected in view of the truncation effect on the relationship between exhaustion rates and the IUR which was introduced by estimating over the subsamples. The increase in the coefficient of WR during the periods when EB is in effect is consistent with the hypothesis that the disincentive effects of UI are increased during such periods. But, as we have pointed out in several other places, the general instability that we observed in our estimates of the wage-replacement effect makes it hazardous to assign major significance to any particular coefficient estimate.

TABLE III.A.4

EXHAUSTION RATE EQUATION FOR EB IN EFFECT, EB NOT IN EFFECT--
DEPENDENT VARIABLE lnR--
FOR THE PERIOD 1965.II TO 1974.IV

Independent Variable	EB Effect Coefficient	EB Effect t Statistic	EB Not Effect Coefficient	EB Not Effect t Statistic
APD	-0.0477	-5.6	-0.0572	-15.5
DIS	-0.00116	-1.0	0.00162	4.2
DURD	-0.152	-2.9	-0.181	- 6.0
PMALE	0.0078	0.3	-0.0913	- 7.6
PMANF	-0.00754	-5.3	-0.0118	-16.0
PCONST	-0.0132	-3.4	-0.0113	- 6.7
NODIST	-0.109	-2.3	-0.162	- 6.6
QUALWK	0.0238	1.6	-0.0256	- 3.9
IUR	0.0489	5.4	0.0862	12.2
WR	1.306	4.2	0.645	3.6
NOTAVL	0.00230	0.5	-0.0129	- 6.7
REFUSE	-0.124	-2.6	-0.119	- 8.6
Q 1	-0.0639	-1.3	-0.0535	- 1.9
Q 2	0.124	2.5	0.241	9.9
Q 3	-0.318	-5.2	0.426	-15.3
CONSTANT	-0.106	-0.3	1.048	7.64
R^2	0.58		0.63	
Standard Error	0.248		0.352	
F	21.8		192.3	
d.f.	(15,238)		(15,1680)	

E. AVERAGE POTENTIAL DURATION

In Chapter III we analyzed the determinants of average potential duration and showed that while our econometric results were in accord with a priori theory, they were subject to complex and difficult-to-interpret interactions. Here we provide some additional detail on these findings, with our principal focus on the relationships among entitlements, weekly benefit formulas, and average potential durations. Table IIIA.5 provides some basic descriptive data as background for our analysis. The data show only slight differences in various duration measures between states that use a high-quarter earnings formula and those that use an average weekly wage formula to compute weekly benefit amounts. On the other hand, states that use an annual wage formula (only Alaska and Oregon use such a formula; they are also variable-duration states) have higher average potential durations, and their data exhibit relatively little variability. For most of our analysis we will therefore give little attention to these states.

Although the mean values for the duration measures were similar for high-quarter and average weekly wage states, regression equations explaining these measures had quite different structures. Regressions on average potential duration of others (APDO), on the percentage of claimants at maximum benefits and duration (PMAX), and on overall average potential durations were found to have differences in coefficients that were statistically significant at the .01 level.[1/] These regressions are

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[1/] Specifically, the F statistic for APDO was 20.9, for PMAX it was 22.1, and for APD it was 21.2. For all of these statistics the degrees of freedom were (20, 1608).

TABLE III.A.5

DESCRIPTIVE STATISTICS ON AVERAGE POTENTIAL DURATION
AND ITS COMPONENTS IN VARIABLE DURATION STATES
FOR THE PERIOD 1965.II TO 1974.IV

	All States		High Quarter States		Average Weekly Wage		Annual Wage States	
	Mean	Standard Deviation	Mean	Standard Deviation	Mean	Standard Deviation	Mean	Standard Deviation
Average Potential Duration (weeks)	23.6	2.5	23.4	2.5	23.7	2.5	26.3	0.9
Maximum Potential Duration (weeks)	27.1	2.8	27.1	2.9	27.0	2.7	26.9	1.0
Average Duration of Others (weeks)	21.8	2.9	21.5	2.9	22.0	2.6	25.6	1.0
Percent at Max. Benefits and Duration (percent)	35.5	14.1	34.7	12.9	37.2	18.0	52.9	8.8
Number of States	42		32		8		2	
Number of Observations	1638		1248		312		78	

reported in Tables IIIA.6 to IIIA.8.[1/] Although, as our statistical tests suggested, these equations had many different coefficients, a few of the differences deserve special mention. First, the coefficient of entitlement (ENTITL) was consistently lower in the high-quarter states' regressions than in those for average weekly wage states. Hence, although high-quarter states have generally higher entitlement percentages (averaging 39 percent in high-quarter states and 34 percent in average weekly wage states), they appeared to receive a smaller "payoff" to entitlements in terms of average potential durations than did average weekly wage states.

A second finding from the subsample regressions was that the coefficient of the IUR was quite erratic. Although the pooled data showed a clear tendency for durations to increase during economic downturns, that result did not hold up under disaggregation. In particular, average weekly wage states seem to have experienced reductions in average durations during economic downturns.

Finally, seasonal patterns in duration appeared to be quite different between high-quarter and average weekly wage states. Average durations were highest in the fourth quarter for high-quarter states, whereas they peaked in the third quarter for average weekly wage states. These differences (and differences in the coefficients of the industry composition variables) suggested that a more careful remodeling of specific labor-market characteristics of the states might help clarify the causes of the observed structural differences.

[1/] In Table IIIA.8 we also present a pooled regression for APD. Qualitative results for that equation were quite similar to those implied by the regressions in the components of APD that were reported in Chapter III.

TABLE III.A.6

AVERAGE POTENTIAL DURATION OF OTHERS
BY TYPE OF WEEKLY BENEFIT FORMULA
DEPENDENT VARIABLE APDO
FOR THE PERIOD 1965.II TO 1974.IV

Independent Variable	High Quarter States	Average Weekly Wage States	Annual Wage States
PCONST	-0.0483	0.139	-0.0764
	(-4.4)	(5.8)	(-3.4)
PMANF	0.00643	0.088	-0.0188
	(1.3)	(7.2)	(-1.2)
PMALE	-0.0086	-0.0258	-0.0016
	(-0.9)	(-1.8)	(-0.2)
ENTITL	0.0575	0.0917	-0.0374
	(15.7)	(3.5)	(-0.3)
IUR	0.161	-0.0432	0.0883
	(3.2)	(-0.6)	(2.0)
MPD	0.626	0.595	0.920
	(27.3)	(16.2)	(6.8)
Q 1	0.113	-0.776	0.131
	(0.5)	(-2.0)	(0.47)
Q 2	-0.448	0.102	-0.0214
	(-2.8)	(0.4)	(-0.11)
Q 3	-0.141	1.22	-0.230
	(-0.83)	(4.77)	(-0.9)
CONSTANT	3.160	-1.953	3.943
	(3.4)	(-1.9)	(0.76)
R^2	0.51	0.69	0.67
Standard Error	2.04	1.49	0.60
F	143.6	73.3	15.6
d.f.	(9,1238)	(9,302)	(9, 68)

NOTE: t Statistics in parentheses.

APPENDIX TO CHAPTER IV

SIMULATION MODEL

This appendix provides a detailed description of the structure of the simulation model of exhaustion rates described in Chapter IV. In its most complete form, this model uses as input the national, seasonally adjusted unemployment rate (SAUR), quarterly dummy variables, and a time trend. The output includes, for each quarter of input data, state and national estimates of the UI exhaustion rate for regular UI, UI first-payments, UI final payments, the insured unemployment rate, and EB status (the national value indicates the number of states on EB). The user of the model can run up to ninety-nine simulations at once, each of which can be run over fifty-five quarters. The model can be used either for prediction or for policy simulation. To use the model for this latter purpose, the user needs to modify one or more of the subroutines.

The model uses the inputs for each quarter to predict the insured unemployment rate and UI first-payments at the state level. Both IURs and first-payments are predicted by using regressions in which these variables were regressed separately on the SAUR, three quarterly dummies, and a time trend.[1/] The estimation period was 1965.II to 1974.IV. The coefficients from these regressions are reported in Table IVA.1 for the IUR regressions and in Table IVA.2 for the first-payments regressions. The national estimate for the IUR is

[1/] The Cochrane-Orcutt method was used to adjust for autocorrelation of the error terms.

TABLE IV.A.1

COEFFICIENTS OF IUR REGRESSIONS USED IN SIMULATION MODEL
(Dependent Variable: IUR)

State	SAUR	Q1	Q2	Q3	Trend	Constant
Alabama	.75	.75	- .05	- .15	- .01	- .52
Alaska	.25	5.9	1.7	3.2	- .001	6.1
Arizona	.80	1.03	.16	.02	- .04	- .32
Arkansas	.72	2.02	.13	- .55	- .03	.27
California	.75	1.45	.44	- .44	- .21	1.69
Colorado	.35	.99	.05	- .22	- .02	- .08
Connecticut	1.17	1.36	.36	.71	- .03	-1.72
Delaware	.52	1.43	.07	.50	.05	- .86
D.C.	.23	.55	.10	.18	.01	.03
Florida	.74	.09	- .14	.47	- .02	-1.22
Georgia	.59	.50	.07	.16	.001	-1.50
Hawaii	.16	.15	- .37	- .53	.04	1.32
Idaho	.47	3.01	.54	.12	.001	.42
Illinois	.56	1.15	.40	.02	- .0004	- .98
Indiana	.66	1.05	.12	- .18	- .004	-1.46
Iowa	.42	1.47	.30	- .07	- .003	- .71
Kansas	.73	1.48	.24	.03	- .03	- .91
Kentucky	.58	1.91	.51	- .12	- .03	.05
Louisiana	.33	1.50	.74	.27	.01	.43
Maine	1.12	1.98	.47	- .21	.01	-2.02
Maryland	.68	1.29	.19	.03	- .01	-1.06
Massachusetts	.96	1.60	.23	.10	.01	-1.16
Michigan	1.34	1.89	.53	.86	.02	-4.21
Minnesota	.77	2.45	.76	- .23	- .01	-1.61
Mississippi	.37	1.48	.42	- .02	- .03	.61
Missouri	.70	1.68	.36	- .02	- .01	- .81
Montana	.48	3.64	.48	- .94	- .001	.65
Nebraska	.39	1.68	.08	- .17	- .01	- .21
Nevada	.69	1.87	- .03	- .67	- .02	1.33
New Hampshire	1.06	.89	.28	.16	- .02	-2.89
New Jersey	.78	1.92	.54	.10	.02	- .48
New Mexico	.61	1.51	.22	- .32	- .01	.27
New York	.83	1.35	.21	- .15	- .01	- .32
North Carolina	.80	1.15	.35	- .08	- .02	-1.76
North Dakota	.57	4.74	1.00	-1.23	- .04	.06
Ohio	.72	1.05	.05	- .19	- .01	-1.43
Oklahoma	.57	.93	.24	- .12	- .03	.58
Oregon	.78	2.33	.12	-1.17	- .01	- .06
Pennsylvania	.84	1.21	.04	- .13	.01	-1.39
Rhode Island	.90	2.07	.33	.63	.03	-1.62
South Carolina	.92	.53	.03	- .05	.01	-2.54
South Dakota	.38	2.35	.19	- .40	- .02	.15
Tennessee	.57	1.69	.33	- .07	- .02	- .09
Texas	.42	.42	.06	- .04	- .03	- .10
Utah	.33	2.23	.10	- .36	- .02	1.78
Vermont	1.19	2.42	.56	- .63	- .01	-2.60
Virginia	.27	.73	.24	.07	- .01	- .36
Washington	.88	2.03	- .38	- .55	.05	- .32
West Virginia	.52	2.32	.13	- .28	- .02	.81
Wisconsin	.70	1.93	.29	- .11	.001	-1.38
Wyoming	.25	2.11	.44	- .45	- .04	1.01

TABLE IV.A.2

COEFFICIENTS OF FIRST PAYMENTS REGRESSIONS USED IN SIMULATION MODEL
(Dependent Variable: FP)

State	SAUR	Q1	Q2	Q3	Trend	Constant
Alabama	6,773	6,724	-2,962	351	192	-19,789
Alaska	339	2,445	-519	-1,853	60	541
Arizona	3,955	2,337	-837	-403	343	-19,901
Arkansas	4,638	5,417	-2,637	-3,029	41	-10,605
California	40,634	71,314	-22,050	-24,205	611	19,409
Colorado	977	4,780	-45	-1,715	11	374
Connecticut	7,580	16,925	-2,156	17,857	543	-21,549
Delaware	3,965	3,824	-847	3,879	39	474
D.C.	609	1,560	-150	907	38	-143
Florida	4,891	1,345	1,542	11,162	-165	-4,977
Georgia	7,015	9,080	168	4,315	174	-22,260
Hawaii	655	190	-1,551	-1,435	89	651
Idaho	507	3,299	-1,568	-58	41	1,245
Illinois	14,174	41,431	1,959	420	256	-22,065
Indiana	9,846	21,498	-2,946	843	205	-24,984
Iowa	1,915	8,546	515	213	95	-4,325
Kansas	2,170	7,046	-851	1,394	25	-3,499
Kentucky	3,905	13,290	-1,533	-1,270	199	-8,822
Louisiana	1,742	10,513	2,587	232	208	1,644
Maine	1,776	2,678	1,867	-86	92	-2,395
Maryland	4,169	13,723	-2,508	5,049	34	-2,917
Massachusetts	9,952	27,853	-7,078	-34	266	-2,891
Michigan	21,917	37,373	-5,927	57,158	877	-66,317
Minnesota	5,353	17,833	-2,550	-5,878	167	-12,077
Mississippi	3,193	6,361	-41	-1,263	178	-13,423
Missouri	4,750	18,308	-327	2,198	196	-1,865
Montana	463	3,891	-937	-2,061	20	1,330
Nebraska	1,190	5,308	-1,220	-997	54	-2,351
Nevada	1,459	1,689	-1,872	-1,749	43	-993
New Hampshire	1,761	1,041	4,276	1,042	81	-6,790
New Jersey	5,981	32,262	586	7,291	1,078	6,515
New Mexico	526	2,228	-653	-988	22	1,32
New York	28,942	100,677	-2,842	12,549	-804	19,042
North Carolina	22,208	23,774	1,734	-1,055	218	-83,514
North Dakota	298	2,965	-624	-1,210	-4	622
Ohio	15,018	32,630	-8,538	2,815	126	-28474
Oklahoma	2,105	3,203	-424	-815	43	-2,235
Oregon	3,660	10,075	-5,461	-7,314	165	2,314
Pennsylvania	18,445	51,583	-10,232	-268	494	-19,491
Rhode Island	2,069	6,054	-1,183	2,684	100	-2,570
South Carolina	6,396	4,558	-405	508	454	-33,771
South Dakota	328	1,884	-236	-314	6	-493
Tennessee	9,684	20,066	-1,241	-2,304	228	-29,457
Texas	9,110	7,376	-1,802	-897	-149	-12,309
Utah	913	4,412	-1,510	-1,086	27	872
Vermont	672	2,281	-303	-1,047	45	-1,262
Virginia	4,420	9,375	1	-604	27	-13,289
Washington	6,104	6,056	-3,888	3,732	328	2,559
West Virginia	1,594	10,726	-2,158	786	37	1,833
Wisconsin	5,635	17,736	-5,625	-2,179	91	-6,308
Wyoming	18	-189	-236	-246	-6	517

calculated by weighting the state estimates by the proportion of insured employment found in each state over the model's estimation period. The national estimate for first-payments is the sum of the state estimates.

Next, the predicted IURs are used to establish EB status by using the EB trigger formula. The actual formula used for EB is slightly modified to fit the quarterly simulation model. Both the national and state triggers are expressed in terms of the quarterly IUR, and the 120 percent criterion is ignored.

The predicted EB status and IURs are then used as inputs to the exhaustion-rate equation by using the regression coefficients reported in Tables III.3 and III.7. The other variables are read into the model as fixed inputs, except the input of average potential duration (APD). That variable can be either read in or predicted by using the predicted IUR and the regression equations reported in Tables III.11 and III.12. The national APD is calculated by weighting the predicted state APDs by the proportion of UI first-payments found in each state.

Finally, the model predicts UI final payments in each state in a given quarter by multiplying the predicted first-payments lagged two quarters by the predicted exhaustion rate. The national first-payment estimate equals the sum of the state final payments, and the national exhaustion rate equals national final payments divided by national first-payments lagged two quarters.

The simulation model has been modified for our report to use an alternative specification of the exhaustion-rate equation and to use alternative specifications of the EB trigger formula. These modifications are discussed in Chapter IV.B and Chapter V.C.

V. POLICY ANALYSIS

INTRODUCTION

In this chapter we use our econometric estimates from Chapter III and the simulation model discussed in Chapter IV to assess various UI policy issues. The chapter is divided into three sections that parallel the presentation provided in prior chapters. In section A we use our basic econometric results to assess in quantitative terms the effects of economic factors on exhaustion rates. Section B reports a similar analysis of the effects of state and federal UI laws on exhaustion rates. Finally, in section C, we make use of the simulation model to examine the implications of adopting alternative trigger mechanisms for the Extended Benefits (EB) program.

A. THE EFFECT OF ECONOMIC FACTORS ON EXHAUSTION RATES

The principal way in which economic factors affect the exhaustion rate is through changes in the strength of labor-market demand. For that reason, this section devotes primary attention to examining the relationship between the insured unemployment rate (IUR) and the exhaustion rate. The section concludes with a brief summary of our findings about the effects of labor-market composition on exhaustion rates.

As we have noted previously in the report, there is an ambiguity in the question "how do IUR's affect the exhaustion rate?" Our econometric results showed that cross-state differences in IURs have relatively slight effects on exhaustion rates, whereas cyclical changes over time in IURs may have a major impact. For the purposes of this section we will focus on the second of these interpretations; that is, we will examine the implications of cyclical changes in IURs on the exhaustion rate. We will do so by using coefficients from Table III.7 in which state-specific, cross-sectional effects are controlled for by the use of dummy variables. The coefficient of the IUR in that table suggests that a 1 percentage point increase in the IUR leads to a 17.6 percent increase in the exhaustion rate. At its mean value (27.4 percent) that would imply a 4.8 percentage point increase in the exhaustion rate (to 32.2 percent). That finding is quite close to the time-series estimates reported by Hight (1975, p. 247) for the states of Pennsylvania and Georgia, but is about 1 percentage point higher than results reported in Chapter I for aggregate data. A possible explanation for this latter discrepancy may be that the simple national time-series regression of exhaustion rates on IURs is capturing offsetting effects on exhaustion rates that come into play when IURs rise. We now turn to examining the most important of such an effect.

In Chapter III (Tables III.1 and III.12) we showed that when the IUR rises, average potential duration (APD) also increases because of the entry of more high-wage workers into the pool of UI recipients. Our econometric estimates suggested that a 1 percentage point increase in the IUR raised APD by about 0.3 weeks.[1/] Because of the strong negative influence of APD on exhaustion rates, such an effect would, in the long run, reduce the impact of the change in the IUR. We estimated the size of this reduction to be about 0.4 percentage points.

1/ This estimate is probably a lower bound because it does not take into account changes in the industrial composition of the insured unemployed (notably, an increase in the percent in manufacturing) that occur when IURs rise.

Hence, our estimate of the impact of a 1 percentage point increase in the IUR is to increase the exhaustion rate by 4.4 percentage points.

One way in which cyclical increases in the exhaustion rate can be offset is by a policy of increasing the average potential duration. Our estimates suggested that the effects of a 1 percentage point increase in the IUR as a result of a cyclical downturn can be offset by a three- to four-week expansion in APD (a 15 percent increase).[1/] Similarly, a recession in which the national unemployment rate rose from 4 to 8 percent would require an increase in APD of ten to twelve weeks (a 50 percent increase), to keep final exhaustion rates constant.[2/] Since the current EB program mandates such a 50 percent increase during recessions, we estimated that the program is sufficient to negate the impact of such a recession on the exhaustion rate for EB.[3/] In a more severe recession the EB program would not be sufficient to prevent total exhaustion rates from rising, whereas in a less severe recession, implementation of EB would actually reduce total exhaustion rates.

Our model documented a number of other ways in which economic factors affect exhaustion rates, but since these economic factors are exogenous to the UI system and are not under the control of policymakers,

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1/ This estimate again agrees with that of Hight (1975, p. 247).

2/ This figure was calculated by using the results from one link of our simulation model, which suggested that such an increase in the unemployment rate would increase the IUR by approximately 3 percentage points.

3/ This computation ignores the positive effect that the existence of the EB program has on the length of unemployment spells (and, hence, on the exhaustion rate). We present a more detailed analysis of the effects of EB (and of FSB) in section B below.

we will only briefly survey these results. We found that the sexual composition of the pool of UI recipients had a significant impact on exhaustions. Other things equal, if state A had 10 percent higher male representation among the insured unemployed than did state B, we estimated that the exhaustion rate in A would be about 2 percentage points lower. That differential would be slightly larger if the fact that males also have higher average potential durations were also taken into account. The finding is consistent with the hypothesis that female UI recipients experience longer spells of unemployment (either because they have more difficulty finding jobs, or because they are more affected by the disincentive effects of UI benefits).[1/]

The industrial composition of the insured unemployed was also found to have a significant influence on exhaustion rates. Variables representing both the percent of the insured unemployed in construction and the percent in manufacturing had significant negative impacts on exhaustion rates. That finding is consistent with the hypothesis that workers in such industries have, on average, shorter unemployment spells than those in other industries. For manufacturing workers the negative effect on exhaustion rates was increased by the fact that such workers also are typically eligible for longer average potential durations. On the other hand, for construction workers, the negative impact on exhaustions was moderated slightly by their typically shorter APDs.

[1/] Age was the only other demographic factor that could be examined with the data we had available. Using data for 1969.II to 1974.IV (the only period for which consistent data were available) we found that the percentage of the insured unemployed younger than twenty-five years old had no impact on exhaustion rates. The percent over fifty-five years of age had (contrary to expectations) a small but statistically significant negative effect on exhaustion rates.

A final "economic" impact on exhaustion rates which was clearly revealed in our estimates was seasonality. Although, as we demonstrated in Chapter III, existence of seasonal influences had little impact on the size of our estimates of the effects of other variables on the exhaustion rate, the size of such influences dwarfed the impact of many of those other variables. Our results showed that exhaustion rates were about 6 percentage points higher than average in the second quarter of a year, and 10 to 11 percentage points lower in the third quarter. This swing in the exhaustion rate between the second and third quarter was, for example, larger than the estimated swing induced by most of the observed cyclical changes in the IUR over the period of estimation. Unfortunately, our data were not appropriate for a detailed examination of the causes of seasonality in the exhaustion rate,[1/] but clearly it is an issue that warrants attention. Later in this chapter we show that seasonality in other UI statistics (most notably, IURs) can have an important effect on the operations of various trigger mechanisms for extended benefits.

B. EFFECTS OF STATE AND FEDERAL LAWS ON EXHAUSTION RATES

In this section we use our econometric model to estimate the effects of various state and federal laws on the exhaustion rate. Our presentation is divided into three subsections. First, we examine state

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[1/] We did find, as might be expected, that the effects of seasonality varied by state. Some of our estimates of these differences were discussed in the appendix to Chapter III. There we showed that states with unusual seasonal unemployment patterns (i.e., Florida) also exhibited unusual seasonal exhaustion patterns.

policies that may affect exhaustion rates through their effects on the lengths of individuals' unemployment spells. Such policy effects include the effects of the wage replacement provided by the amount of weekly benefits, the effects of the extent of UI enforcement procedures, and the effects of UI qualification provisions. Next, we examine state policies that have an influence on the exhaustion rate through their effects on average potential durations. That examination focuses primarily on three parameters that categorize state duration provisions: uniform-duration policies; maximum potential durations; and entitlement fractions. Finally, the third subsection analyzes the effects of federal extended benefits programs on the exhaustion rate for regular UI benefits.

1. State Policies that Affect the Length of Unemployment Spells

Three types of variables in our model measure the effects of state laws on lengths of unemployment spells: the wage replacement ratio; a measure of states' enforcement of UI "availability for work" requirements; and summary measures of UI qualification provisions. The first of these is one of the most commonly used measures of the incentives that individuals who receive UI have to prolong their job search. The wage-replacement ratio (defined as the ratio of average UI benefits to the average wage in manufacturing) provides a measure of the relative benefit of remaining on UI, as opposed to accepting a job at the prevailing wage. Our statistical results suggested that this ratio had a significant positive effect on lengths of unemployment spells and on the exhaustion rate. They indicated that a 10 percentage point increase in the wage-replacement ratio would result in a 0.5- to 1.5-week increase in the average length of unemployment spells.[1/] As we pointed out in Chapter III, that estimate is consistent with a number of studies of the disincentive effects of UI which employ methodologies quite different from those used here. The finding indicates that if states A and B are identical in all respects (except that A has a 10 percent higher wage-replacement ratio than B), on average, then, state A will have an exhaustion rate that is 1 to 2 percentage points higher than B's.

Our estimates indicated that the extent of disqualifications for being unable or unavailable to work and for the refusal of suitable work had a significant negative impact on the lengths of unemployment spells and on the exhaustion rate. Because we have no independent measure of enforcement, these results should be regarded with some skepticism; however, their stability under a wide variety of estimation procedures may at least suggest the influence that UI administrative policies have. However, the quantitative magnitude of the effects, while always significantly different from zero, were relatively small.[2/] They suggested that a 50 percent increase in the average number of disqualifications would reduce the average length of unemployment spells by about one week and would reduce the average exhaustion rate by somewhat more than 1 percentage point.

[1/] This relatively wide range of estimates resulted from the instability in the quantitative magnitude of the wage-replacement effect in response to the variety of estimation procedures employed in Chapter III.

[2/] However, the effects were always considerably larger than those that would result simply from a purely definitional relationship between disqualification and exhaustions.

Finally, our results provided some evidence that state qualification provisions may have a direct effect on the lengths of unemployment spells by affecting the skill levels of individuals in the pool of UI eligibles. In particular, we found that states that require (the equivalent of) more full-time weeks of employment as a condition for UI eligibility had lower exhaustion rates (possibly because workers who did qualify would exert greater job-search efforts). There was also some evidence that states with qualification provisions that make it relatively easier for seasonal workers to qualify for benefits have lower exhaustion rates. That would be consistent with the hypothesis that seasonal workers have shorter (but, perhaps, more frequent) spells of unemployment. Because of the relatively crude ways in which we parameterized qualification provisions, however, the significance of such findings should not be exaggerated.

2. State Policies That Affect Average Potential Durations

Our estimates consistently showed that average potential durations (APD) had a negative effect on exhaustion rates. The size of that effect was practially invariant with respect to the particular estimation procedure used. It indicated that, other things being equal, a one-week increase in APD resulted in a reduction in the exhaustion rate of about 1.5 percentage points.

Although our theoretical model led us to use the APD variable in our exhaustion rate equations, we were also careful to point out that APD is not in fact directly under the control of UI policymakers. Rather, observed average potential durations represent a complex interaction between the following: state entitlement, benefit, and qualification formulas; economic conditions in a state; and characteristics of UI recipients. It was for this reason that we examined the determinants of average potential durations in Chapter III by using a two-equation model that stressed the estimation of the net effects of actual state policy parameters. Although we will not repeat detailed results of that model here, it will be used throughout this section to generate the simplified results reported. Three policy parameters are explicitly examined here: uniform-duration provisions; state maximum potential durations; and entitlement fractions.

Our model showed three reasons why the exhaustion rate was lower in uniform- than in variable-duration states. First, uniform-duration states had higher average potential durations than did variable-duration states. Over our period of observation, that difference averaged 2.6 weeks (26.2 weeks in uniform-duration states versus 23.6 weeks in variable-duration states), and it alone caused exhaustion rates to be about 3.5 percentage points lower in uniform-duration states. Second, there was no dispersion in average potential duration in uniform-duration states, and that factor caused exhaustion rates to be another 0.5 percentage points lower in such states. Finally, our results found a statistically significant effect of uniform-duration policies on the exhaustion rate which persisted even when all the other variables were held constant. Although our investigation did not provide an unambiguous explanation of that effect,[1/] it was clearly of a substantial magnitude.

[1/] Specifically, we found that the hypothesis that uniform- and variable-duration states had identical exhaustion-rate equations could not be rejected at the 95 percent level if the two types of states were permitted to have different intercept terms. An examination of specific coefficients in separate regressions on uniform- and variable-duration states suggested that one possible cause of the intercept differences was a lower wage-replacement effect in the uniform-duration states.

It implied a (ceteris paribus) 4.5 percentage point difference in exhaustion rates between variable- and uniform-duration states. Overall, then, our results indicated that exhaustion rates were 8 to 9 percentage points lower in uniform- than in variable-duration states. About one-half of that difference was accounted for by measurable duration effects, whereas the other half remained a puzzle that required further analysis.

States' maximum-duration provisions were found to have a significant effect on exhaustion rates. However, estimating the quantitative magnitude of that effect required some rather elaborate calculations. For uniform-duration states the effect of a one-week increase in maximum potential duration was easily simulated: average potential duration would be increased by one week for all UI recipients in such states. In variable-duration states, such a simple calculation was not possible, however, because a one-week increase in maximum potential duration would increase only some recipients' potential durations and would not raise APD by a full week. By using our equations that explain the components of APD, we were able to predict that a one-week increase in a variable-duration state's maximum potential duration would increase APD by slightly more than three-tenths of a week. Because other provisions of state UI laws (such as weekly benefit amounts and entitlement fractions) were held constant when we performed the hypothetical experiment of increasing maximum potential duration by one week, only those UI recipients who were actually constrained by prevailing maximums experienced an increase in their potential durations.[1/]

[1/] Of course, if states were to legislate a change in maximum potential durations it is possible they would also change other provisions in their law. Our model did not permit us to estimate what those changes might be, and, hence, they were not included in our estimates.

By combining estimates from uniform- and variable-duration states, our predicted overall effect of an across-the-board increase of one week in maximum potential durations would be to increase national APD by four-tenths of a week and reduce exhaustion rates by about one-half of a percentage point.

A final state policy parameter which was examined by using our model was the fraction of base-period wages to which a UI recipient is entitled. In our sample, entitlements averaged 38 percent of base-period wages, and there was considerable variation in that average.[1/] Although our statistical results showed this variable to have a strong positive effect on average potential durations, they also posed many problems in interpretation because of complex interactions between entitlement fractions and other state-law provisions. As a rough estimate, our findings suggested that a 10 percent increase in the entitlement fraction would have increased average potential duration by about 0.4 weeks. Such a policy would have reduced exhaustion rates by about one-half of a percentage point.

3. <u>The Effects of Extended Benefits Programs</u>

Throughout our statistical investigations we found that the existence of programs that provided for additional weeks of UI receipt after exhaustion

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[1/] Because UI formulas differ among states, there is no commonly reported figure on entitlement fractions. We constructed our data from UI statutes by making some rough conversions in order to obtain comparability across states. These procedures are described in our data documentation. Because of the necessity of adopting such procedures, our entitlement variable may have exhibited substantial errors in measurement, and the estimated effect of the variable on APD would therefore have been biased toward zero.

of regular benefits had a strong positive effect on exhaustion rates.[1/] Such a finding would be consistent with the hypothesis that the availability of extended benefits provided an incentive for unemployed workers to prolong their job-search activities and thereby made it more likely that they will exhaust benefits. Overall, we estimated that the existence of extended benefits increased unemployment spells by an average of 2.75 weeks.[2/] This resulted in an additional 4 to 5 percent of UI recipients exhausting their regular benefits when EB was in effect.

In addition to estimating overall effects of extended benefits programs, we also estimated the effects of three specific programs: regular extended benefits; the Emergency Unemployment Compensation Program (TC); and the Federal Supplemental Benefit Program (FSB). Our estimates of the effects of these programs on lengths of unemployment spells and on exhaustion rates are reported in Table V.1. They showed that the EB program itself had the major impact on exhaustion rates. TC and FSB had smaller (but statistically

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[1/] In Chapter III we discussed the possibility that this correlation may have arisen because exhaustion rates play an important role in the decision to implement extended benefits programs. Although devising a test of that hypothesis involved major statistical problems, we were unable to find any evidence that our estimated effect of EB was biased by this type of simultaneity.

[2/] This figure applied to all regular UI recipients. If extended unemployment spells are experienced only by those who actually go on to collect extended benefits payments, the increased unemployment experienced by that subgroup would be of a considerably longer duration.

TABLE V.1

ESTIMATED EFFECTS OF VARIOUS EXTENDED BENEFITS PROGRAMS ON UNEMPLOYMENT SPELLS AND ON EXHAUSTION RATES

Single Programs	Effect on Length of Unemployment Spells	Effect on Exhaustion Rate
Regular EB	+ 2.75 weeks	+ 4.5%
TC	+ 1.08 weeks	+ 1.7%
FSB	+ 1.20 weeks	+ 1.9%
Combined Programs		
EB and TC	+ 4.1 weeks	+ 6.6%
EB and FSB	+ 4.2 weeks	+ 6.8%

significant) additional effects.[1] These data, therefore, clearly showed that enactment of extended benefits programs can have a significant impact on state UI programs even if the extended benefits program itself is totally federally financed. They provide an additional rationale for federal aid to state programs during cyclical downturns.

C. SIMULATION OF ALTERNATIVE EB TRIGGER POLICIES

One way in which the simulation model described in Chapter IV can be used is to investigate the impact of various potential changes in UI policy. This is done in this section for alternative EB trigger policies. The performance of each alternative is examined for two prototype recessions--a mild recession and a severe recession. Each recession runs over a three-year period and starts with a national, seasonally adjusted unemployment rate of 4 percent. This rate rises in the first year to 6 percent for the mild recession and 8 percent for the severe recession, and returns to 4 percent over the remaining two-year period. Alternative EB triggers are examined by comparing the number and pattern over the recession of states with EB and the percentage of final payments that are eligible for EB. For the most part, the analysis does not indicate that any one trigger mechanism is more desirable than any other; rather, it is intended to provide an example of how the simulation model can be used to provide information for policymakers.

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[1] Notice that the estimated "combined program" effects somewhat exceed the sum of the impacts of the two constituent programs, which is the result of the fact that the effect of EB on the exhaustion rate was slightly greater when the other programs were in effect, perhaps because those programs made EB more attractive.

The current EB trigger mechanism establishes EB benefits in a state if either the national, seasonally adjusted insured unemployment rate (IUR) is 4.5 percent or higher for three consecutive months, or the state IUR (not seasonally adjusted) averages 4 percent or more for any thirteen consecutive-week period and exceeds 120 percent of the average rate for the corresponding thirteen-week period of the two preceeding years. The national trigger goes off if the seasonally adjusted IUR drops below 4.5 percent for three consecutive months, and the state trigger goes off if the IUR for any thirteen-week period drops below 4 percent or below the 120 percent criterion. Since this law is expressed in months or weeks, the criteria were modified slightly to fit the quarterly simulation model. Both the national trigger and the state trigger were expressed in terms of the quarterly IUR. If the appropriate quarterly IUR was above 4.5 or 4 percent, the EB program was triggered-on in the next quarter. In addition, the 120 percent criterion was ignored. The operation of this provision was felt to be unsatisfactory and has been suspended by Congress for most of the history of the EB program.

Three groups of alternative EB trigger mechanisms were simulated. The first group changed the levels of the insured unemployment rate used for the trigger and sets the national and state trigger level equal. The results of these changes are reported in Tables V.2 and V.3, where the triggers were set at 4.0, 4.5, and 5.0. The simulations of these trigger levels were performed in an nonseasonal manner--i.e., there was no seasonal variation in the state IURs and, hence, the state triggers.[1/] The results indicated that, as is obvious, setting the triggers at 4.0 was a more generous mechanism than the current law, while the other two alternatives were less generous. As is reported in Table V.2, the 4.0 trigger increased the percentage of individuals with a final payment who were eligible for EB from 61 percent under current law to 68 percent under the new law in the severe recession. No change occurred in the mild recession, since the national trigger was never operative in that simulation. As can also be seen, there was a cost to this increase in coverage. The presence of EB in more states increased the number of final payments. This occurred in the simulation because of the positive coefficient on the EB variable in the exhaustion-rate equation. For this simulation (and all others) the tradeoff between these two effects was that an increase in EB coverage for each 100 individuals resulted in an increase of thirteen to fourteen final payments.

The 4.0 national trigger also had the effect for the severe recession of triggering-on the national program one-quarter earlier and keeping it on one-quarter later than the current law provided. This earlier operation of the trigger meant that, for the recession that was simulated, the national trigger was on during the quarter with the highest unemployment rate, rather than, as in the current law, being on only during the initial decline of the recession. While this earlier operation of the national trigger is presumably desirable, the simulations clearly show that

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[1/] This was accomplished by setting all the quarterly dummy variables used in the simulation to 1/4.

the fact that since both the national and state triggers used for this set of simulations operate, by necessity, with a lag they will always trigger-on later and stay on later than is probably desirable.

The second type of trigger that was simulated took account of the number of UI exhaustees, as well as the number of insured unemployed. This was done because the insured unemployed includes only those who are receiving unemployment benefits. An "adjusted" IUR was calculated for each state equal to the IUR plus one-quarter of the past year's exhaustions divided by the average monthly number of covered workers in the state.[1/] The implicit assumption was that one-quarter of the past year's exhaustees remained unemployed. This method of triggering-on extended benefits was used for the Temporary Compensation program (TC) in the early 1970s. The trigger level for that program was set at 6.5 percent. This level, plus a 5.5 percent level, was used, and the results are reported in Tables V.2 and V.3. The simulation was again nonseasonal, and no national trigger was used (although such a trigger could easily be incorporated).[2/]

[1/] In the Temporary Compensation program, this average was computed for a given quarter for the first four of the last six completed calendar quarters. For these simulations, the average over the entire period 1965 to 1974 was used. Since covered employment rose during this period, this means that the rate was higher in the simulations than that which would be used if the Temporary Compensation method were reproduced exactly. However, since the timing of when states trigger-on over the recession should be the same, the general conclusions presented here would still hold.

[2/] The TC program itself had only a state trigger, but the program ran concurrent to the regular EB program. A state could have been on both or either one of the programs.

TABLE V.2

FINAL PAYMENTS UNDER ALTERNATIVE
EB TRIGGER MECHANISMS

(1,000's)

	Current Law	State and National Trigger Equal: 4.0	4.5	5.0	Temporary Compensation Method Trigger Equal: 5.5	6.5
Mild Recession						
Total Final Payments	5,410	5,410	5,298	5,236	5,332	5,233
Final Payments in EB States	2,085	2,085	1,260	801	1,515	733
Change Relative to Current Law	--	0	-825	-1,284	- 570	-1,312
Percent of Final Payments in EB States	38.5	38.5	23.8	15.3	28.4	14.8
Severe Recession						
Total Final Payments	6,970	7,034	6,891	6,766	6,878	6,707
Final Payments in EB States	4,278	4,746	3,686	2,744	3,510	2,310
Change Relative to Current Law	--	468	-592	-1,534	- 768	-1,968
Percent of Final Payments in EB States	61.4	67.5	53.5	40.6	51.3	34.4

TABLE V.3

NUMBER OF STATES WITH EB UNDER

ALTERNATIVE EB TRIGGER MECHANISMS

		Current Law	State and National Trigger Equal: 4.0	4.5	5.0	Temporary Compensation Method Trigger Equal: 5.5	6.5
Mild Recession							
SAUR[a]	Quarter						
4.00	1	4	4	2	2	3	2
4.50	2	4	4	2	2	3	2
5.00	3	8	8	4	2	5	2
5.50	4	9	9	6	2	8	2
6.00	5	13	13	9	7	10	4
5.75	6	17	17	12	9	12	8
5.50	7	16	16	9	8	10	6
5.25	8	13	13	9	7	10	6
5.00	9	10	10	9	5	10	6
4.75	10	9	9	6	2	10	5
4.50	11	9	9	4	2	10	5
4.25	12	8	8	4	2	9	5
4.00	13	4	4	2	2	7	3
Severe Recession							
SAUR[a]	Quarter						
4.00	1	4	4	2	2	3	2
5.00	2	4	4	2	2	3	2
6.00	3	9	9	6	2	8	2
7.00	4	17	17	12	9	11	7
8.00	5	24	51	17	14	17	12
7.50	6	51	51	51	51	23	14
7.00	7	51	51	51	14	21	13
6.50	8	24	51	17	14	19	13
6.00	9	19	19	14	11	18	13
5.50	10	17	17	12	9	16	12
5.00	11	13	13	9	7	16	10
4.50	12	9	9	6	2	11	8
4.00	13	8	8	4	2	10	6

[a] Seasonally adjusted national unemployment rate.

The results of these simulations showed that for the trigger levels simulated, the TC method was less generous than the existing EB program. This was true for the mild recession, in which the absence of a national program did not affect the results, as well as for the severe recession. The most interesting finding can be observed by examining the timing of when states trigger-on. Because of the emphasis on exhaustions, the TC trigger went on later during the recession than the current law allowed and stayed on later. For example, for the mild recession and the 5.5 trigger level, fewer states were triggered-on relative to the current law for the first eight quarters. In quarter 9 the same number of states were on extended benefits, and in quarters 10 through 13 more states were collecting than under the current EB program. Thus, this method of triggering-on extended benefits did not remedy the problem observed above--that the trigger operates later during a recession than is desirable. In fact, in this regard the TC trigger was worse than the current EB trigger. A solution to this problem might require a trigger mechanism that takes account of the direction in which the unemployment rate is changing, as well as its level.

The other difference between the EB trigger and the TC trigger is that the states that were triggered-on in any one quarter can be different. In the example discussed above, for the quarter in which both the EB and TC triggers are operative in ten states, nine of the states were the same. For the EB trigger only, California received extended benefits, and for the TC trigger only, Oregon received extended benefits.

The final set of simulations examined the fact that the state IUR used for the trigger is not seasonally adjusted. This means that seasonal fluctuations in the IUR can act to trigger EB on or off independently of underlying economic conditions. For this reason it had been proposed to seasonally adjust the state IUR for use as the trigger. This was done for the simulations reported in Tables V.4 and V.5. For these simulations the recessions were begun separately in each quarter to show the way in which the lack of seasonal adjustment for the trigger affected the results. The data in Table V.4 show that seasonal adjustment reduced the percentage of final payments covered by EB in all of the simulations except one. Overall, for the mild recession, the reduction averaged 1.0 percentage point, and for the severe recession, 4.4 percentage points. This reduction, however, led to a considerable smoothing of the number of states on EB over the two recessions. With no seasonal adjustment, the number of states on EB fluctuated independently of the unemployment rate, while this effect was dampened considerably by the seasonal adjustment. Although this fluctuation was probably overstated by the way in which EB was handled in the simulation model (the actual program uses a sliding thirteen-week period, while the model uses a calendar quarter), the observed dampening in the simulation would still hold under this alternative formula.

TABLE V.4

FINAL PAYMENTS BY BEGINNING QUARTER OF RECESSION AND BY SEASONAL ADJUSTMENT OF STATE EB TRIGGER (1,000's)

Beginning Quarter of Recession

	Mean Of Four Quarters		1		2		3		4	
	No Seasonal Adjustment	Seasonal Adjustment	No Seasonal Adjustment	Seasonal Adjustment	No Seasonal Adjustment	Seasonal Adjustment	No Seasonal Adjustment	Seasonal Adjustment	No Seasonal Adjustment	Seasonal Adjustment
Mild Recession										
Total Final Payments	5,409	5,399	5,789	5,734	5,245	5,255	5,280	5,282	5,320	5,325
Final Payments in EB States	2,168	2,133	2,345	2,176	2,075	2,141	2,118	2,061	2,132	2,072
Change Relative to Current Law	--	-55	--	-169	--	66	--	-57	--	-60
Percent of Final Payments in EB States	40.1	39.1	40.5	37.9	39.6	40.7	40.1	39.0	40.1	38.9
Severe Recession										
Total Final Payments	7,041	7,012	7,432	7,361	6,911	6,884	6,878	6,884	6,943	6,920
Final Payments in EB States	4,691	4,358	4,931	4,515	4,833	4,485	4,404	4,240	4,595	4,191
Change Relative to Current Law	--	-333	--	-416	--	-348	--	-164	--	-404
Percent of Final Payments in EB States	66.6	62.2	66.3	61.3	69.9	65.2	64.0	61.6	66.2	60.6

TABLE V.5

NUMBER OF STATES WITH EB BY BEGINNING QUARTER OF RECESSION AND BY SEASONAL ADJUSTMENT OF STATE EB TRIGGER

Beginning Quarter of Recession

		1		2		3		4	
SUAR[a]	Quarter	No Seasonal Adjustment	Seasonal Adjustment	No Seasonal Adjustment	Seasonal Adjustment	No Seasonal Adjustment	Seasonal Adjustment	No Seasonal Adjustment	Seasonal Adjustment
Mild Recession									
4.00	1	16	3	3	4	1	4	2	4
4.50	2	16	3	3	4	1	4	2	4
5.00	3	5	7	2	7	2	9	20	7
5.50	4	7	9	8	9	25	9	9	9
6.00	5	9	16	28	10	11	14	8	16
5.75	6	28	13	14	17	9	17	13	17
5.50	7	13	16	8	17	10	16	28	13
5.25	8	8	16	9	16	28	10	11	14
5.00	9	8	10	26	10	10	10	7	12
4.75	10	25	9	9	9	7	9	8	9
4.50	11	7	9	5	8	3	9	21	8
4.25	12	2	7	2	9	20	7	5	7
4.00	13	2	4	18	5	3	4	2	4
Severe Recession									
4.00	1	16	3	3	4	1	4	2	4
5.00	2	16	3	3	4	1	4	2	4
6.00	3	9	9	7	9	8	9	25	9
7.00	4	9	17	13	17	28	13	14	17
8.00	5	51	51	41	18	23	26	51	51
7.50	6	43	26	51	51	51	51	51	51
7.00	7	51	51	51	51	51	51	41	21
6.50	8	51	51	51	51	41	18	23	26
6.00	9	15	25	34	17	17	18	14	24
5.50	10	28	13	14	17	9	17	13	17
5.00	11	11	14	8	16	9	16	28	10
4.50	12	7	9	8	9	25	9	9	9
4.00	13	2	9	20	7	5	7	2	7

[a] Seasonally adjusted national unemployment rate.

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