The Effect of Unemployment Insurance on Temporary Layoff Unemployment

By Martin Feldstein*

Economists are now beginning to recognize that an understanding of temporary layoffs is crucial for a proper analysis of unemployment. In manufacturing, about 75 percent of those who are laid off return to their original employers. More generally, among all persons classified as "unemployed job losers," temporary layoffs account for about 50 percent of all unemployment spells. Temporary layoffs are an even larger fraction of cyclical changes in the number of job losers. While this group includes some seasonally unemployed, most temporary layoffs are induced by short random or cyclical fluctuations in demand. The conventional model of search unemployment is inappropriate for those on temporary layoff and the modern theory of the Phillips curve requires substantial modification because of the size and cyclical variation of temporary layoff unemployment.1

In a previous paper (1976), I showed analytically that our current system of unemployment insurance (UI) provides a substantial incentive for increased temporary layoff unemployment.2 The present paper provides micro-economic evidence that UI actually has such a powerful effect. The estimates imply that the incentive provided by the current average level of UI benefits is responsible for approximately one-half of temporary layoff unemployment.

It is important to note that the current study shows that UI increases the amount of temporary layoff unemployment, but does not deal with the mean duration per spell. This distinction deserves emphasis because nearly all previous empirical work focused on the potential effect of UI on duration. This focus on duration is both unfortunate and surprising since UI can actually increase total unemployment while decreasing the mean duration per spell. While UI increases the duration of any given spell of unemployment, it may also induce more very short spells of unemployment. This possibility of reduced mean duration is clear in my 1976 theoretical analysis. An additional practical

*Professor of economics, Harvard University. I am grateful to the National Science Foundation for support of this research, to David Ellwood and Joseph Kahan for assistance with the statistical calculations, and to Richard Freeman, Zvi Griliches, Daniel Hamermesh, James Medoff, Melvin Reder, and Jeffrey Sachs for discussions and comments. Earlier versions of this paper were presented at seminars at Chicago, Harvard, and Yale universities.

1In my 1975 paper, pp. 737-42, I discuss the implications of temporary layoffs for the theory of search unemployment, the Phillips curve, and wage inflexibility. Although the standard criterion of unemployment is active job seeking within the past four weeks, individuals are officially classified as unemployed without any inquiry about recent job-seeking activity if they state that they are "on layoff awaiting recall by their employers." Some of those on layoff look for temporary jobs or alternative permanent employment, but the vast majority do return to their original employers. Readers should not be confused by the two quite separate meanings of the term "layoff" in the Department of Labor's lexicon. In manufacturing establishment data, a layoff is a separation initiated by the employer (not a quit) and may be permanent or temporary. In the Current Population Survey (CPS), an individual is on layoff if he is not working but "has a job" to which he is expecting to be recalled by his employer. To emphasize that I am dealing with those layoffs expected to terminate in recall, I use the adjective "temporary." Unfortunately, the CPS uses the word temporary in a different and quite confusing way: persons on layoff are divided into an "indefinite duration" group (in which the individual does not have an expected date of recall within thirty days) and a "temporary" group (when such a date is known). When it is useful to distinguish these groups, I use the terms "indefinite duration" and "fixed duration"; in my usage, the term temporary layoff includes both groups.

2My 1976 paper is really an explicit proof of arguments made more informally in my earlier study for the Joint Economic Committee (1973). For a similar development, see Martin Baily.
matter reinforces this tendency. In the absence of UI, firms might be reluctant to lay off workers for short periods in response to random demand fluctuations, for fear of losing these workers to other firms, or at least of creating costly ill will; UI eliminates these problems and facilitates short-duration layoffs. In contrast, firms might have no choice but to lay off employees for long spells during the less frequent, but more protracted, spells of low demand. Unemployment insurance thus increases the number of spells of temporary layoff unemployment with a relatively greater increase for short spells. Since the duration should increase for any given spell, while the mix should change to add short spells that would otherwise not exist, the net effect of UI on duration is indeterminate. The existing estimates of the effect of UI on the mean duration of unemployment spells should therefore be regarded as an understatement—and, possibly, an extreme understatement—of the effect of UI on total unemployment.

Although the presence of a labor union is not necessary to obtain the effects on temporary layoffs indicated by the theoretical analysis of UI, these predicted effects are likely to be magnified if the employees are unionized. The basic reason is that employers are more willing to lay off workers when they are confident that they will return when recalled, while employees are more willing to be laid off if they can be confident that they will be recalled. Both conditions are more likely to be met in unionized firms where workers often receive compensation that exceeds their market alternative, and have seniority privileges and pensions that are not portable. More directly, union contracts often guarantee that previous workers will be recalled before any new employees are hired (see U.S. Bureau of Labor Statistics, 1972). Finally, unionized firms may have more layoff unemployment because, as Freeman has suggested, unions provide an effective mechanism for expressing workers’ collective preferences to management. All of this implies that temporary layoff unemployment should be higher for union members and suggests that the response of temporary layoff unemployment to UI benefits may also be greater.

The first section of this paper discusses the data and methods used in the present study. The econometric estimates are presented in Section II. The brief concluding section suggests some directions for future analysis and comments on the implications of the research for the optimal redesign of social insurance.

I. Data and Method

The current study uses a sample of nearly 25,000 individual observations collected by the Current Population Survey (CPS) to measure the effect of unemployment insurance benefits on temporary layoff unemployment. The estimated regression equations presented in the next section relate each individual’s temporary layoff unemployment status (a binary variable equal to 1 if the individual is on temporary layoff) to three kinds of variables: 1) his potential UI benefit as a percentage of lost net wages; 2) his basic demographic characteristics; and 3) the basic characteristics of his employment. This section begins by discussing the CPS sample and the method of calculating the potential UI benefit “replacement ratio” for each individual. The measurement of demographic and employment characteristics is then discussed.

A. The CPS Sample

The CPS is the government household survey used by the Department of Labor to calculate official monthly unemployment
rates. About 60,000 households are interviewed each month about the employment activities of their members during the week prior to the survey. The March survey of each year also obtains information about labor force participation, employment, and earnings during the previous year. The current study uses the survey for March 1971, a period of relatively high unemployment.

For this analysis, individuals were eliminated from this CPS sample if they were not in the experienced labor force, were reentrants to the labor force, or were self-employed; none of these groups is at risk of being laid off. Also eliminated because of the atypical character of their employment were employees in the public sector and in agriculture. To avoid the problems associated with those who combine school and work, and with those on the verge of retirement, the sample was restricted to individuals between the ages of 25 and 55. Finally, a few observations were excluded, because data were missing on the individual’s age, sex, color, marital status, industry and occupation of employment, union membership, or previous year’s work experience. The sampling weights indicate that the resulting sample of 24,545 represents a population of 34.2 million persons.

B. Calculation of Potential UI Benefits

The unemployment compensation benefits for which an individual is eligible depend on his previous earnings up to a ceiling of maximum benefits received by about half of all UI benefit recipients. Because unemployment insurance is actually a series of state programs that operate as part of a general federal system, the formulae relating benefits to past earnings and the maximum benefits differ among the states. In addition, dependents’ benefits are also available in states with approximately one-third of covered workers.

The CPS collects no information about the unemployment insurance benefits received by the currently unemployed or the potential benefits of the employed. A special computer program was therefore prepared to evaluate the potential UI benefits for each of the 24,545 individuals in the final CPS sample. The algorithm uses the particular rules for each individual’s state of residence and incorporates information on his industry of employment, previous year’s earnings and work experience, and number of eligible dependents. As a rough test of the accuracy of this method, the program was used to determine the benefit eligibility and to calculate the benefits for all unemployed persons in the full CPS sample (and not the final subsample of 24,545 observations). The implied total benefits for March 1971 was $540 million; this is reasonably close to the total amount actually paid as reported by the individual state UI agencies, $630 million. The accuracy is likely to be greater for temporary layoffs for whom the reporting of previous year’s income is much more reliable.

5The seasonally adjusted unemployment rate in March 1971 was 6.0 percent and had been stable during the previous three months. The March 1971 survey was not “selected,” but was the first CPS tape that became publicly available. The use of that sample indicates the slow gestation of this project.

6Barry Chiswick presents evidence that the recent extension of unemployment insurance to agriculture has substantially increased the seasonality of employment and unemployment in agriculture. It will be important to see if that result is confirmed by data after the 1975 recession year.

7In 1971 there were 50.8 million persons in the labor force between the ages of 25 and 55. The difference between 50.8 million and 34.2 million represents primarily government employees, agricultural workers, and the self-employed.

8Individuals are asked about the total annual value of benefits received during the previous year, but these twelve-month recall data are notoriously bad and, in the aggregate, represent a 50 percent understatement of the amounts paid by the UI program.

9There is no information on “benefit exhaustion,” i.e., on whether an individual has already been unemployed so long that his number of weeks of eligibility for benefits has been exhausted. However, for all types of unemployment, only about 20 percent of spells exhaust available benefits while, for those on temporary layoff, the percentage should be very much lower: in March 1974, only 4 percent of “job losers on layoff” were unemployed for more than twenty-six weeks, while 12 percent of “job losers with no job” were unemployed for that long (see the author, 1975, Table 4).
The central variable of interest is the ratio of the individual's potential unemployment insurance benefit to his foregone earnings net of marginal income and payroll taxes. This UI "benefit replacement ratio" measures the proportion of lost net-of-tax earnings that would be replaced by UI benefits. A 60 percent UI benefit replacement ratio implies that the unemployed individual would lose only 40 percent of his previous net-of-tax wage income. Stated differently, the benefit replacement ratio is analogous to a rate of tax levied on earnings when the alternative is insured unemployment; a 60 percent benefit replacement ratio implies that the individual, by working instead of collecting UI, receives additional income equal to only 40 percent of his total net wage. The computer program evaluated the benefit replacement ratio for each individual, using the federal income tax schedules, to evaluate a marginal tax rate for someone with the individual's family income and dependents who used the standard deduction. The relevant marginal social security tax rate and state income tax rate were added to the federal marginal tax rate.

Although theory predicts that the probability of being on temporary layoff is an increasing function of the benefit replacement ratio, there is no presumption of linearity. A movement in the benefit replacement ratio from 0.70 to 0.80 may increase unemployment by more than a movement from 0.30 to 0.40. To eliminate the restriction of a linear specification, equations are reported in the next section in which the continuous benefit replacement ratio variable (BEN) is replaced by a set of binary variables that classify individuals by their benefit replacement ratios:

- \( B E N = 0 \) (for those not eligible for benefits);
- \( 0 < B E N \leq 0.30 \);
- \( 0.30 < B E N \leq 0.50 \);
- \( 0.50 < B E N \leq 0.70 \);
- \( 0.70 < B E N \leq 0.85 \); and
- \( 0.85 < B E N \).

This method has the further advantage that it can clearly separate those who are ineligible for benefits (\( B E N = 0 \)) from the remaining variation in \( B E N \).

Although I believe that this represents the best method of evaluating the benefit replacement ratio with the available data, there are several problems that should be borne in mind in evaluating the results. First, there is no information on the extent of experience rating that is relevant for each individual's employer. If the extent of experience rating is uncorrelated with the benefit replacement ratio, ignoring experience rating does not bias inferences about the effect of the benefit replacement ratio on temporary layoff unemployment. Second, there are three omissions that are likely to cause an overestimate of the impact of unemployment insurance on temporary layoff unemployment: cash and in-kind transfers that may be available to individuals on temporary layoff, the value of fringe benefits that are lost during unemployment, and the work expenses (transportation, meals, etc.) that are avoided during unemployment. None of these omissions is likely to be large for the quite short duration of unemployment that are relevant here. Moreover, to the extent that a higher probability of layoff is compensated by a higher gross wage (as implied by the firm's budget constraint), there will be an offsetting underestimate of the impact of UI on temporary layoff unemployment. It is difficult to assess the net effect of these countervailing influences, but the resulting bias is likely to be small.

It is much more important to understand that the regression coefficient of the benefit replacement ratio measures the effect of interindividual differences in unemployment benefits and that the effect of such differences is less than the effect of a general change in everyone's benefit replacement

\[ \text{The regression of the unemployment variable on the benefit replacement ratio does, however, understate the effect on unemployment of differences in the net UI subsidy. The net UI subsidy is the difference between the benefits and the additional experience-rated tax payments induced by those benefits. In the notation of my 1976 paper, the net subsidy is } [1 - e(1 - t_p)]b \text{ where } e \text{ is the ratio of induced employer tax to incremental benefits (i.e., the extent of experience rating), by the marginal personal income tax rate, and } b \text{ is the weekly benefit. If } e(1 - t_p) \text{ were constant, the regression coefficient of } b \text{ would understate the effect of changes in the net subsidy by a factor of } [1 - e(1 - t_p)]^{-1}. \]
ratio. As a general rule, it is the employer who makes the decision to lay off and recall a worker, while the employee himself is essentially passive.11 An employer can respond to his employees' benefit replacement ratios only as an average for the group whose layoff he is considering and not individually for each member in the group. It is because the relevant group of employees within a firm has similar benefit replacement ratios12 that the individual benefit replacement information is relevant for understanding what is essentially an employer or employer-employee group decision. Since the benefit replacement ratios are not identical for the relevant group of the firm's employees, some part of the variation of BEN in the sample will not affect layoff unemployment. The effect of this is to make the estimated regression coefficient an underestimate of the effect of a general increase or decrease in all benefits.13

C. Demographic and Employment Characteristics

The demographic characteristics included in the analysis are the standard list of age, sex, marital status, and race.14 As I indicated above, separate equations are also estimated for men only. The sample is limited to individuals between the ages of 25 and 55. To avoid any assumption about the form of the relation between age and temporary layoffs, individuals are divided into four separate age groups and binary variables are used in the regression equation. The age groups included are 25–29, 30–39, and 40–49; the coefficient for persons 50–55 is implicitly zero. The other demographic variables are self-explanatory.

The potential role of unions was discussed briefly in the introduction. In the final CPS sample of 24,545, 6,845 individuals (or 27 percent of the sample) indicated that they were members of labor unions. There is no indication whether the individual's current employment is in a union job. This suggests that the estimated coefficient of the union variable may underestimate the full effect of unionization.

Individuals were classified according to industry group and occupation category and the corresponding binary variables were included in the regression to control for inherent "technological" differences among them in the likelihood of layoffs.15 This procedure entails a danger of "over-controlling" for the exogenous aspect of these variables. Individuals with high potential benefit replacement rates (for example, with low wage rates or high spouse income, or large families in states where dependents' allowances are paid) may seek employment in industries and occupations with high technological probabilities of layoff unemployment. To the extent that this is important, the regression coefficients will overstate the importance of the industry and occupation variables and will underestimate the impact of the benefit replacement ratio. Although it is not possible to model

11 I say "as a general rule" because workers do frequently have "inverse seniority" privileges that permit more senior workers to choose to be laid off before or instead of others. See U.S. Bureau of Labor Statistics (1972) for a description of these privileges.

12 The benefit replacement ratios are similar to the extent that members of the group have similar wages and, being located in the same state, have similar unemployment benefit schedules and state tax rates.

13 This can be stated differently by noting that a firm can only perceive and respond to the mean BEN value for the relevant group of its employees and essentially ignores the within-group variance. A general change in all UI benefits shifts this mean while part of the sample variation includes the within-group variance. In still different language, the coefficient of BEN is biased down but the size of the bias is limited to the extent that the between-group variance is large relative to the within-group variance. This bias can be thought of as a classical "errors-in-variables" bias: the "true" value of BEN required by the model is the mean of the individual BEN values for the relevant employee group, while the actual individual BEN values may be regarded as equal to the "true" value plus an error. This errors-in-variables interpretation also indicates that there is a downward bias that is an increasing function of the within-group variance relative to the between-groups variance.

14 It might be interesting to extend this list to other attributes that reflect differences in tastes for leisure, for example, education, home ownership, age of children of married women, etc. See, however, fn. 13.

15 The twelve industry groups were combinations of two-digit industries. Recall that agricultural workers, the self-employed, and public employees were omitted from the sample. The nine occupation groups were combinations of more detailed two-digit classifications.
this simultaneous relationship, separate results will be presented with and without the industry and occupation variables.

The final variable considered in the analysis is the individual's wage rate. Temporary layoff unemployment is likely to be related to the individual wage in several quite different ways. First, for any given benefit replacement ratio, a higher wage implies both a higher absolute benefit and a greater absolute cost of unemployment; the sign of this effect is therefore indeterminate. Second, if a high wage reflects better pay relative to the individual's market opportunity, the employer will be more likely to lay off workers with a confidence that they will return when recalled; this implies a positive coefficient for the wage variable. Third, a higher wage may ceteris paribus imply greater seniority; greater seniority means fewer involuntary layoffs relative to other employees within the firm, but a group with more seniority on average may have more temporary layoffs because workers are more likely to await recall.

Related to this seniority aspect is the possibility that more senior workers who are laid off perceive themselves (correctly) as only on temporary layoff, while their more junior coworkers who are laid off may regard the separation as permanent because their probability of recall is substantially lower. Finally, jobs with more layoffs may pay higher wage rates ceteris paribus than other jobs, implying that the gross wage is endogenous and positively related to the unemployment probability. While this source of wage variation is likely to be small relative to the wage variation reflecting individual skill differences, etc., some equations without this variable have been estimated to assess the effect of erring in the direction of its omission.

II. The Econometric Evidence

All of the equations that I have estimated imply that the current level of unemployment insurance benefits causes a substantial fraction of the observed temporary layoff unemployment. More specifically, the econometric evidence indicates that the temporary layoff unemployment rates would be reduced by approximately one-half if the adverse incentive provided by the current unemployment insurance were eliminated. This conclusion is not sensitive to the exclusion of questionable regressions or to the restriction of the sample to particular subsamples.

Before looking at the estimated regression coefficients, it is helpful to examine the basic data on temporary layoff unemployment rates and UI benefit replacement ratios. In March 1971 the temporary layoff unemployment rate was 1.6 percent in the population corresponding to the final CPS sample of 24,545 employees; that is, on average, the corresponding population had a probability of 0.016 of being unemployed and on layoff during the sample week in 1971. The mean value of the benefit replacement ratio for this population was 0.55. Only 3 percent of the population was found to be ineligible for benefits; while 60 percent of the sample had benefit replacement rates above one-half and 30 percent had benefit replacement rates about 70 percent.

Table 1 shows the temporary layoff unemployment rates corresponding to six levels of the benefit replacement ratio. This unemployment rate rises monotonically from 0.50 percent among the ineligibles (BEN = 0) to 2.17 percent in the highest benefit group (BEN > 0.85). Taken at face value, these unemployment rates imply that

16Recall that we are "holding constant" the effect of age, sex, color, unionization, industry, and occupation. It might be interesting to add education and other variables.

17The reader should remember the caveats and potential biases discussed in Section I; they will not be repeated here.

18Recall that the sample is restricted to eliminate many groups with no UI benefits, such as new entrants and the self-employed.

19The distribution of benefit replacement ratios for the population should not be confused with the distribution for the unemployed subgroup. The mean benefit replacement ratio of the unemployed was 0.59; if those with zero benefits are excluded, the mean benefit replacement rate for the eligible unemployed exceeds 0.60. This is consistent with the calculation that I presented for a range of hypothetical employees in my 1974 paper.
reducing \textit{BEN} to a maximum of 0.40 would lower the temporary layoff unemployment rate from 1.60 to 1.26, a reduction of 0.34 percentage points. It must, however, be borne in mind that this relation between benefits and temporary layoff unemployment rates is not adjusted for demographic or economic characteristics. We turn therefore to the multiple regression equations.

Table 2 presents the estimated coefficients of four basic regression equations. The dependent variable is binary, taking the value of 1 if the individual is unemployed and on layoff, and the value of 0 otherwise. The regression coefficients have all been multiplied by 100, converting the predicted dependent variable from a probability to a percentage unemployment rate. The sample means and proportions of the explanatory variables are shown in the first column.

Consider first the estimated coefficient of the benefit replacement ratio in equation (1).\textsuperscript{20} The coefficient of 1.345 implies that the mean \textit{BEN} value of 0.55 raises the mean temporary layoff unemployment rate by 1.345 \times (0.55) = 0.74. Since the temporary layoff unemployment rate is 1.60, this equation implies that \textit{BEN} is responsible for 46 percent of the observed temporary layoff unemployment rate. Because the industry and occupation variables may overcorrect for the truly exogenous effects of these variables, the basic specification is repeated without them as equation (2). The coefficient of the benefit replacement ratio rises slightly (to 1.545), implying that the mean benefit replacement ratio of 0.55 is responsible for 53 percent of the observed temporary layoff unemployment rate.\textsuperscript{21}

The coefficient of the binary union variable in equations (1) and (2) provides strong evidence that union members are much more likely to experience temporary

\textsuperscript{20} Note that equation (1) includes all of the variables discussed in Section I; the coefficients of the twenty-seven industry and occupation variables are not shown since they are not of interest in themselves, and would require much extra space in the table.

\textsuperscript{21} Omitting the other potentially endogenous economic characteristic variables (the gross wage rate and unionization) only lowers this coefficient to 1.515. (This equation is not shown in the table.) Other variants cluster around 1.3, rising as high as 1.7 and falling as low as 1.0. Replacing the gross wage rate by a set of six classification variables in gross wages has essentially no effect on the other coefficients.
layoff unemployment than nonunion members. The temporary layoff unemployment rate is 1.15 percentage points higher than the rate for nonmembers even after adjusting for this industry-occupation mix. Without that adjustment, the differential is 2.24 percentage points. I will return below to the evidence that the layoff unemployment rate of union members is also more sensitive to UI benefits.

The coefficients of the other variables are interesting but involve no important economic insights. There is clear evidence that the frequency of temporary layoff unemployment falls quite sharply with age, a reflection of the powerful seniority system. There is no statistically significant difference between either whites and nonwhites or marrieds and singles. Males appear to have a significantly lower temporary layoff unemployment rate when (but only when) the industry and occupation effects are included separately.22

The sex differential is large and surprising to me. It may be an artifact of overadjustment for industry and occupation or it may reflect a real difference between the sexes. Women may be more likely to take seasonal work (within broad industry-occupation groups) or to have relatively long spells of temporary layoff. Nothing is known about these fascinating issues.

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### Table 2—Determinants of Temporary Layoff Unemployment

<table>
<thead>
<tr>
<th>Variable</th>
<th>Sample Means and Proportions</th>
<th>Regression Coefficients</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>(1)</td>
</tr>
<tr>
<td>BEN = 0</td>
<td>0.55</td>
<td>1.345</td>
</tr>
<tr>
<td>0 &lt; BEN ≤ 0.30</td>
<td>0.03</td>
<td></td>
</tr>
<tr>
<td>0.30 &lt; BEN ≤ 0.50</td>
<td>0.28</td>
<td></td>
</tr>
<tr>
<td>0.50 &lt; BEN ≤ 0.70</td>
<td>0.30</td>
<td></td>
</tr>
<tr>
<td>0.70 &lt; BEN ≤ 0.85</td>
<td>0.24</td>
<td></td>
</tr>
<tr>
<td>Union</td>
<td>0.28</td>
<td>1.154</td>
</tr>
<tr>
<td>Age: 25-29</td>
<td>0.18</td>
<td>0.686</td>
</tr>
<tr>
<td>Age: 30-39</td>
<td>0.31</td>
<td>0.220</td>
</tr>
<tr>
<td>Age: 40-49</td>
<td>0.33</td>
<td>-0.196</td>
</tr>
<tr>
<td>Male</td>
<td>0.65</td>
<td>-1.460</td>
</tr>
<tr>
<td>Married</td>
<td>0.91</td>
<td>-0.267</td>
</tr>
<tr>
<td>White</td>
<td>0.89</td>
<td>-0.068</td>
</tr>
<tr>
<td>Gross Wage ($100)</td>
<td>1.64</td>
<td>0.202</td>
</tr>
<tr>
<td>Industry-Occupation</td>
<td>a</td>
<td>a</td>
</tr>
<tr>
<td>Mean of Dependent Variable</td>
<td>1.601</td>
<td>1.601</td>
</tr>
<tr>
<td>N</td>
<td>24,545</td>
<td>24,545</td>
</tr>
</tbody>
</table>

Notes: All coefficients have been multiplied by 100, converting the dependent variable from a probability to a percentage unemployment rate. Standard errors are shown in parentheses. See text for description of sample and definitions of variables.

22The sex differential is large and surprising to me. It may be an artifact of overadjustment for industry and occupation or it may reflect a real difference between the sexes. Women may be more likely to take seasonal work (within broad industry-occupation groups) or to have relatively long spells of temporary layoff. Nothing is known about these fascinating issues.
Equations (3) and (4) replace the continuous benefit replacement ratio variable by a set of six binary classification variables. In each equation, an increase in the benefit replacement ratio always implies an increase in the predicted temporary layoff unemployment rates. Both equations suggest that variations in the benefit replacement ratio below the 30 to 50 percent range have little effect on unemployment but higher benefit replacement ratios have a substantial adverse effect. The coefficients of the $BEN$ variables in equation (3) imply that lowering $BEN$ for everyone to 0.40 (with an implicit coefficient of $-1.53$) would reduce the temporary layoff unemployment rate to 0.46 percentage points. With equation (4), the same calculation implies a reduction of the temporary layoff unemployment rate of 0.49 percentage points (with an implicit baseline coefficient of $-1.40$). It is not clear how much weight should be given to the implications of this more elaborate specification. On purely statistical grounds, there is little basis for choice; the reduction in the residual sum of squares in going from equation (1) to equation (3) is not quite significant at the 5 percent level, while going from equation (2) to equation (4) is not even significant at the 10 percent level. The pattern of the coefficients does correspond to the a priori expectation that variations in benefit replacement ratios will have a weaker effect when UI benefits are "too small to bother taking into account" than when those benefits replace a substantial fraction of lost net wage income. However, the apparently weak effect at low benefit levels may reflect only the small fraction of the sample in this range; since only 11 percent of the sample had $BEN$ values below 0.30, it is difficult to make any inferences about the effects of variations in benefits within the range below 0.30 or between this range and the next higher interval. It is probably best to remain agnostic on this question until more data become available.

Table 3 confirms that union members have a substantially higher temporary layoff unemployment rate and are more sensitive to unemployment insurance benefits than are nonunionized workers. For the sample of 6,845 union members, the temporary layoff unemployment rate was 3.14 percent, twice the rate for the entire sample and thus three times the rate for nonunion members. The coefficient of the benefit replacement ratio variable in equation (1) is 2.72, also about twice the corresponding coefficient for the entire sample. A coefficient of 2.72 implies that the mean benefit replacement ratio of 0.54 (for union members) induces a 1.47 percent temporary layoff unemployment rate.

The coefficients of the $BEN$ variables in equations (3) and (4) correspond quite closely to the conditional "unemployment rates presented in Table 1; for example, an increase in $BEN$ from 0.40 to 0.60 reduces the predicted temporary layoff unemployment rate by 0.50 percentage points in both the multiple regression equation and the unadjusted values of Table 1. It is clear from this comparison that replacing equations (3) and (4) by logit regression instead of ordinary least squares would be very unlikely to change any of the conclusions of the current analysis.
Table 3—Coefficients of Benefit Replacement Ratio Variables for Union Member Subsample and Male Subsample

<table>
<thead>
<tr>
<th>Variable</th>
<th>Subsample (1)</th>
<th>Union Members Only (2)</th>
<th>Subsample (5)</th>
<th>Men Only (8)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Means and Proportions</td>
<td>Means and Proportions</td>
<td>Means and Proportions</td>
<td>Means and Proportions</td>
</tr>
<tr>
<td>BEN</td>
<td>0.54</td>
<td>2.723</td>
<td>2.287</td>
<td>1.419</td>
</tr>
<tr>
<td>BEN = 0</td>
<td>0.01</td>
<td>-2.460</td>
<td>-2.968</td>
<td>-2.349</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(1.499)</td>
<td>(1.495)</td>
<td>(0.708)</td>
</tr>
<tr>
<td>0 &lt; BEN</td>
<td>0.07</td>
<td>-4.382</td>
<td>-3.755</td>
<td>-1.564</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(2.468)</td>
<td>(2.471)</td>
<td>(1.316)</td>
</tr>
<tr>
<td>0.30 &lt; BEN ≤ 0.50</td>
<td>0.36</td>
<td>-3.876</td>
<td>-3.297</td>
<td>-1.470</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(1.518)</td>
<td>(1.516)</td>
<td>(0.869)</td>
</tr>
<tr>
<td>BEN &gt; 0</td>
<td>0.50</td>
<td>-2.834</td>
<td>-2.394</td>
<td>-0.985</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(1.193)</td>
<td>(1.192)</td>
<td>(0.795)</td>
</tr>
<tr>
<td>0.70 &lt; BEN</td>
<td>0.33</td>
<td>-2.508</td>
<td>-2.296</td>
<td>0.019</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(1.216)</td>
<td>(1.128)</td>
<td>(0.878)</td>
</tr>
<tr>
<td>BEN ≤ 0.85</td>
<td>0.19</td>
<td>-2.106</td>
<td>-1.110</td>
<td>0.059</td>
</tr>
<tr>
<td></td>
<td></td>
<td>(1.106)</td>
<td>(1.110)</td>
<td>(0.816)</td>
</tr>
<tr>
<td>Includes Industry-Occupation Variables?</td>
<td>–</td>
<td>Yes</td>
<td>No</td>
<td>Yes</td>
</tr>
<tr>
<td>Mean of Dependent Variable</td>
<td>3.141</td>
<td>3.141</td>
<td>3.141</td>
<td>3.141</td>
</tr>
<tr>
<td>Sample Size</td>
<td>6,845</td>
<td>6,845</td>
<td>6,845</td>
<td>6,845</td>
</tr>
</tbody>
</table>

Each equation also contains age, sex, color, marital status, and wage variables and a constant term (as in Table 1); their coefficients are not shown to save space. The Men Only equations also contain a union variable. The table indicates when industry and occupation variables are included. The omitted BEN category corresponds to BEN > 0.85 and has an implicit coefficient of zero.

Although the mean benefit replacement ratio for union members is almost exactly the same as for the entire sample, the distributions of benefit replacement ratios differ noticeably. The replacement ratios for union members are clustered more closely around the average; 69 percent of union members have BEN values between 0.30 and 0.70, while 58 percent of the entire sample is in this range. Almost no union members appear to be ineligible for benefits. The coefficients of equations (3) and (4) also show that the temporary layoff unemployment rate varies inversely with the benefit replacement ratio. Both equations imply that increasing the benefit replacement ratio from 0.40 to 0.60 raises the temporary layoff unemployment rate by about an entire percentage point.29

The results for the “men only” sample (presented in columns 5–8 of Table 3) are very similar to the estimates for the entire sample and need no detailed comment. The temporary layoff unemployment rate of 1.600 is almost identical to the rate for the entire sample (1.601). The regression coefficients differ substantially from the corresponding numbers of Table 1 only for the BEN = 0 subcategory; since only 0.01 percent of the men and 0.03 percent of the entire population are in this group, the comparison of the regression coefficients is without real substance.

Equations similar to those of Table 1 were also estimated with the “duration of

29Excluding the twenty-seven industry and occupation variables (as in equation (2)) reduces the coefficient slightly but leaves these conclusions essentially unchanged. The industry and occupation variables are themselves statistically significant so that equation (1) would be the clearly preferable specification except for the possible simultaneity problem noted in the text.

30The 1 percent of union members who are ineligible for benefits (BEN = 0) appear to have an unusually high layoff rate. This anomalous behavior also contributes to the relatively high standard error of the BEN variable in equations (1) and (2). The very small sample with BEN = 0 and correspondingly high standard error imply that no weight should be given to this group. For BEN > 0, equations (3) and (4) show a strong monotonic relation.
unemployment to date of survey” as the dependent variable. There was no significant relation between BEN and duration, implying that the effect of UI in inducing more short-duration spells of unemployment offset the effect of UI in lengthening the duration of spells that would otherwise have occurred.

Although I am tempted to compare the estimates presented in this section with the results of other recent studies, I believe that research is too dissimilar to warrant such comparison. There have been no previous econometric studies of the effect of unemployment insurance on temporary layoff unemployment. The recent econometric research has focused on the duration of unemployment spells or on the total unemployment rates for state aggregates.\textsuperscript{31} There are several fascinating problems in the interpretation of these data, but their discussion belongs elsewhere.

III. Conclusion

The evidence presented in this paper implies that unemployment insurance has a powerful effect on temporary layoff unemployment. The average UI benefit replacement ratio implied by the current law can account for about half of temporary layoff unemployment. An increase in the UI benefit replacement ratio from 0.4 to 0.6 raises the predicted temporary layoff unemployment rate by about 0.5 percentage points, or one-third of the current average temporary layoff unemployment rate of 1.6 percent. Temporary layoff unemployment is more than twice as frequent among union members as among others between the ages of 25 and 55 who are in the experienced labor force. Unemployment insurance also has a correspondingly greater effect on that employment rate among union members: an increase in the UI benefit replacement ratio from 0.4 to 0.6 raises the predicted temporary layoff unemployment rate of union members by a full percentage point.

These estimates must be understood as subject to the biases and caveats discussed in Section I. It would clearly be desirable to repeat this research with CPS data for a more recent year. A reanalysis with data from the National Longitudinal Survey would be useful because the temporary layoff character of the unemployment spell could be defined \textit{ex post}. It would be particularly valuable to extend the current data to include information on the experience rated tax of each individual’s employer. More generally, it would be useful to reexamine the effect of UI on temporary layoffs by studying data on a sample of individual firms in a variety of states.

I have refrained throughout this paper from making any normative judgments about the effect of unemployment insurance on layoff unemployment. It is clear, however, that our current UI program does impose an efficiency loss by distorting the behavior of firms to lay off too many workers when demand falls rather than cutting prices or building inventories. The substantial rate of temporary layoff unemployment suggests that this efficiency loss may be quite large.

The redesign of unemployment insurance is a difficult problem because the unemployed include the job losers who must find new jobs as well as those on temporary layoff. For those who are changing jobs, the optimal insurance must balance providing protection from financial loss against the distortion to socially inefficient search.\textsuperscript{32} For those who are on temporary layoff, it is sufficient to eliminate the subsidy element in UI by making each firm repay in taxes the full value of the benefits paid its employees and by making UI benefits subject to the same taxation as other compensa-

\textsuperscript{31}These studies include Kathleen Classen; Ronald Ehrenberg and Ronald Oaxaca; Herbert Grubel and Dennis Maki; Arlene Holen and Stanley Horowitz; Charles Lininger; Marston. It should be clear that the only reliable studies of duration effects exclude temporary layoffs and combine data for individuals in different states or years. See Daniel Hamermesh and Finis Welch for discussions of this research.

\textsuperscript{32}This point is discussed in more detail in my 1973 paper. Baily provides an excellent formal solution of this optimization problem.
The difficult problem arises because the full experience rating that is optimal for temporary layoffs is not optimal for permanent layoffs: it would inappropriately discourage new hiring and desirable layoffs. The problem cannot be solved by a lower tax for those layoffs who are not rehired since that would distort the rehire decision and waste job-specific human capital. The optimum balancing of these considerations is a complex problem that requires more information than is currently available. A formal analysis of the problem would be valuable because it would indicate more precisely the type of information required and might provide new insights about the optimal design even before that information is collected.

As a practical solution, I believe that much could be gained by having full employer experience rating for the benefits paid during the first month of each spell of unemployment (or some other moderately short period). It would also be important to tax individuals on UI benefits in the same way as other compensation is taxed. This combination of reforms would eliminate most of the subsidy currently provided for short spells of temporary layoff unemployment without unduly discouraging either new hiring or permanent separations.

33 See the author (1976).
34 Firms can often assess a worker’s quality only after he has worked for the firm for a period of time. If layoffs of unsuitable workers are made very expensive by experience rating, firms will be reluctant to hire new workers and, when they make a hiring mistake, to discharge those who were inappropriately hired.
35 The bias against new hiring could be reduced further by making the “one-month experience rating” provision apply only to workers with a minimum of, say, six months of experience with the firm.

REFERENCES


F. Welch, “What Have We Learned from

