Recent Trends in Insured and Uninsured Unemployment: Is There an Explanation?

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RECENT TRENDS IN INSURED AND UNINSURED UNEMPLOYMENT: IS THERE AN EXPLANATION?*

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This paper explores the recent decline in the fraction of unemployed workers who receive unemployment insurance benefits. Using March Current Population Surveys, we compare the fraction who are potentially eligible for benefits with the fraction who receive them. The decline in insured unemployment is almost entirely due to a decline in the early 1980s in the takeup rate for benefits. We analyze the determinants of the takeup rate, using both aggregated state-level data and micro-data. At least half the decline is due to an increasing share of unemployment in states with lower takeup rates.

I. INTRODUCTION

Although over 90 percent of employed workers hold jobs that are covered by the unemployment insurance system, less than 30 percent of unemployed workers currently receive unemployment insurance (UI) benefits. This fraction has fallen over the postwar era, with notable declines in the early 1960s and the early 1980s. The decline in the past decade is especially puzzling since it has occurred at the same time that the fraction of women in the labor force has stabilized and the baby boom has matured. In this paper we analyze recent changes in the fraction of insured unemployment and provide new evidence on the extent to which these changes can be attributed to changes in eligibility for benefits and changes in participation rates among the eligible. A major obstacle to this apparently simple task is the absence of individual data on unemployment insurance eligibility or receipt.[1] We sidestep this problem by using data for unemployed workers in the March Current Population Survey (CPS) to estimate the fraction of unemployed workers who are eligible for benefits in each state. We then compare changes over time in estimated eligibility to changes

1. As far as we know, there are no major data sets that regularly ask currently unemployed workers about their receipt of unemployment benefits, and no data at all on eligibility.

in the actual fraction of workers receiving benefits, as reported by state administrative records.

While far from perfect, this procedure allows us to decompose changes in the fraction of insured unemployment into three components, showing the change due to (1) changes in the state UI laws; (2) changes in the eligibility-determining characteristics of unemployed workers; and (3) changes in the takeup rate for benefits. Perhaps surprisingly, our results suggest that it is mainly changes in the takeup rate that account for the decline in insured unemployment in the past decade.[2] This finding is confirmed on a subset of eight larger states that can be identified in the CPS over the 1968–1987 period. Takeup rates in the nation as a whole and in the subset of larger states show an abrupt decline between 1980 and 1982. By comparison, takeup rates were relatively constant between 1968 and 1979, and have remained relatively stable since 1982.

We go on to analyze the determinants of state-level takeup rates and characterize the extent to which changes in takeup rates can be explained by changes in the unemployment insurance system and the political climate of states, changes in the characteristics of unemployed workers, and changes in the geographic distribution of unemployment. Takeup rates vary widely across states and changes in the regional distribution of unemployment have contributed to the decline in takeup. But even after accounting for these factors, we find that at least one fourth of the decline in takeup rates remains unexplained.

The final part of our paper analyzes takeup rates using micro-data from 1980 to 1982 collected by the Panel Study of Income Dynamics (PSID). Self-reported eligibility in these data provides a validity check on the eligibility calculations in our earlier analysis; summing the micro-data by state yields very similar takeup patterns to those obtained in our aggregate analysis. While the PSID data do not show a decline in takeup rates between 1980 and 1982, the selective nature of the PSID sample may account for this difference from the CPS data. We also use the PSID data to estimate the determinants of takeup rates among individuals. A number of variables that we cannot measure or that appear insignificant in the aggregate analysis of state takeup rates appear to influence individual participation. It seems unlikely, however,

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2. This result is consistent with recent work by Kane [1988] who identifies the major source of decline in UI receipt over the past decade as a decrease in applications by new job losers.
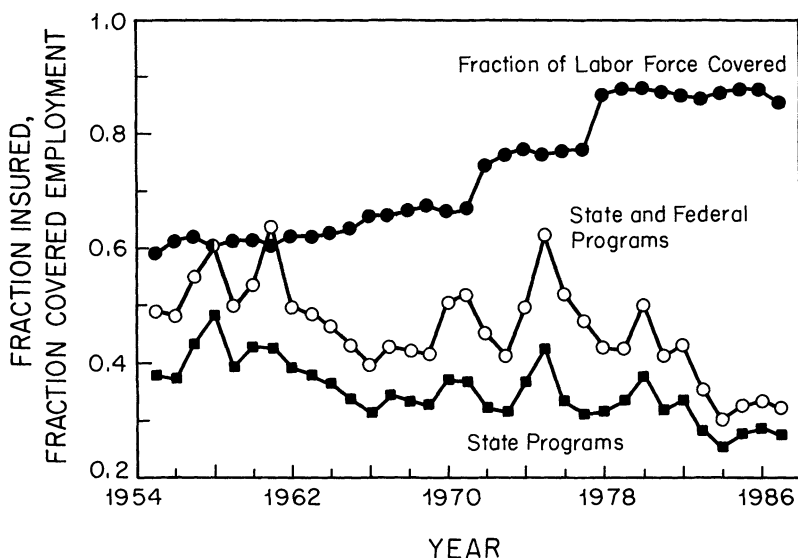
FIGURE I
Characteristics of UI System: 1955–1987

that any of these variables can explain the abrupt decline in takeup rates in the early 1980s.

II. TRENDS IN INSURED UNEMPLOYMENT: 1955–1987

A key indicator of the unemployment insurance system that we use throughout this paper is a ratio that we call the fraction of insured unemployment or FIU. FIU is the fraction of unemployed workers who receive regular UI benefits, calculated by dividing the annual average of weekly insured unemployment in regular state programs by the annual average number of unemployed workers.[3] The fraction of insured unemployment is plotted in Figure I for the period from 1955 to 1987. For comparison we also plot the fraction of unemployed receiving income from all UI programs.[4] In addition

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3. We include in the count of insured unemployed only those workers who are actually receiving UI benefits. As conventionally defined, average weekly insured unemployment also includes workers who are serving out a waiting period for benefits.

4. Sources for the data used in this paper are available from the authors in a Data Appendix. For an excellent discussion of the collection and reporting of UI data, see Blaustein [1980].

to regular state UI benefits, this second fraction includes individuals receiving extended benefits (EB) and supplemental benefits,[5] as well as workers who are covered by special UI programs.[6] Finally, Figure I also presents the estimated fraction of the employed labor force covered by the unemployment insurance system in each year.[7]

The fraction of unemployed workers receiving regular UI payments is countercyclical and shows a decreasing trend, despite the steadily increasing coverage of the UI system among employed workers. Whereas roughly 40 percent of unemployed workers received regular state benefits during the 1950s, the fraction receiving benefits in the late 1980s had fallen to less than 30 percent. The decline in the fraction of unemployed workers receiving payments from all programs has been even greater: from roughly 50 percent in the late 1950s to just over 30 percent in the past few years.

A frequently cited measure of the coverage of the UI system among unemployed workers is the ratio of the insured unemployment rate (IUR) to the total unemployment rate (UR).[8] The insured unemployment rate, which is reported weekly by the states, is the ratio of the number of weekly UI claims (CLAIMS) to the number of employed workers covered by the UI system (COVEMP). The total unemployment rate is the ratio of the number of unemployed workers (UN) to the total labor force (LF). To see the relation between FIU and the ratio of the insured to the overall unemployment rate, write FIU as RECIP/UN, where RECIP is the weekly number of UI recipients, and UN is the number of unemployed. (UN = LF − EMP, where EMP is the number of employed workers). Note that the insured unemploy-

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5. The extended benefit program is a permanent Federal program that provides additional UI payments to workers in high-unemployment states after their regular UI payments have been exhausted. Supplemental benefit programs are similar programs enacted during periods of high unemployment.

6. These special programs are Unemployment Compensation for Federal Employees (UCFE), Unemployment Compensation for Ex-Servicemembers (UCX), and Railroad Unemployment Insurance (RRUI). Workers eligible for these programs are not eligible for regular UI.

7. The number of employed workers covered by unemployment insurance is estimated from payroll (establishment) data; whereas the labor force is typically estimated from individual survey data. If the coverage ratio were calculated using total employment numbers from the establishment survey, it would be approximately 6 to 10 percent higher than in Figure I.

8. The widening gap between the IUR and total unemployment is the focus of Burtless' [1983] paper. The IUR is an important policy concern because extended benefit payments are triggered by state IUR rates.

ment rate is defined in terms of the number of UI claimants, whereas we have defined the fraction of insured unemployment in terms of the number of recipients.[9] Simple algebra then shows that

$$\text{(1)} \quad \frac{\text{IUR}}{\text{UR}} = \frac{\text{CLAIMS}}{\text{RECIP}} * \frac{1}{\text{COVEMP/EMP}} * \frac{1}{1 - \text{UR}} * \text{FIU}.$$

In the past 35 years the ratio of the insured unemployment rate to the total unemployment rate has fallen much more rapidly than the FIU, from an average of 0.75 in the early 1950s to an average of 0.35 in the late 1980s. This is largely attributable to the increase in the coverage rate of the UI system among employed workers (COVEMP/EMP). The ratio of UI claims to UI recipients has also fallen, while the unemployment rate has risen. Thus, the decrease in the ratio of the insured to total unemployment rate significantly overstates the fall in the fraction of unemployed workers who receive UI benefits. For this reason, we focus in this paper on changes in the fraction of UI recipients among the unemployed. Nevertheless, to the extent that government programs are linked to the insured rather than to the total unemployment rate, the divergence of insured and total unemployment rates has important policy effects. In particular, the decline in the IUR has meant that fewer states are eligible for extended benefits.[10]

The failure of the fraction of insured unemployment to increase point-for-point with increases in the coverage of the UI system during the 1970s may at first seem puzzling. However, unless all unemployed workers are previous job losers, and unless the risks of unemployment are the same in the covered and uncovered sectors, increases in the UI system's coverage of employed workers will not generate proportional increases in its coverage of unemployed workers. In fact, many unemployed workers are just entering the labor force and are therefore ineligible for UI benefits regardless of the fraction of employed workers covered by the UI system. In addition, the expansion of the UI system in the 1970s occurred mainly through the addition of state and local

9. In contrast to actual compensation, claims include all UI applications (some of which will be denied), and all workers deemed eligible for UI, some of whom have not yet completed the state-specified "waiting period" which precedes the start of benefits. Individuals receiving extended or supplemental benefits are excluded from the count of UI claims.

10. We estimate elsewhere [Blank and Card, 1989] that three more states would have received extended benefits in 1987 if the gap between insured and total unemployment had not widened.

government employees to the insured rolls.[11] Since these workers have much lower risks of unemployment than private sector workers, their addition to the covered sector generated less than proportional increases in the share of unemployed workers originating from covered employment.[12]

In summary, the time series in Figure I suggest that there has been a steady downward trend in the fraction of unemployed workers collecting UI benefits. This trend was interrupted during the 1970s by expanded coverage of the UI system among employed workers. During the 1960s and 1980s, however, coverage was relatively constant, and the fraction of insured unemployment fell by 6–8 percent per decade.

The decline in the fraction of insured unemployment in the 1960s is often attributed to changes in the demographic structure of the labor force that decreased eligibility. In the past ten years, however, changes in the characteristics of unemployed workers should have slightly increased eligibility. Nevertheless, the fraction of insured unemployment has fallen significantly. In the next section we present a detailed analysis of the effects of recent changes in state UI laws and worker characteristics, in an effort to determine what caused the decrease in the fraction of insured unemployment between 1977 and 1987.

III. CHANGES IN ELIGIBILITY AND IN THE FRACTION OF INSURED UNEMPLOYMENT: 1977–1987

A. Estimating Eligibility and Its Components

There are very few micro-data surveys that contain detailed information on the receipt of UI benefits by unemployed workers, and no aggregate data on UI eligibility. For these reasons, no previous study of UI recipiency has separately analyzed eligibility.[13] In an effort to fill this gap, we estimate UI eligibility for the sample of unemployed workers interviewed in the March CPS between 1977 and 1987 by combining information on state-level UI laws

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11. Employees of nonprofit firms and employees in establishments with fewer than eight workers were also added to the insured rolls in 1972. See U. S. Department of Labor, Employment and Training Administration [1986].
12. For example, in the 1977–1987 March Current Population Surveys, government workers represent on average 16–17 percent of employed workers, but only 7–8 percent of unemployed workers.
13. For example, Corson and Nicholson's [1988] recent analysis of changes in the fraction of insured unemployment makes no attempt to separate the effects of changes in eligibility from changes in takeup rates for benefits.

with individual data on earnings and weeks worked in the previous year. Our choice of sample periods is dictated by the fact that the CPS did not report state of residence for individuals in most states prior to 1977. We present below a longer term analysis based on individuals in the eight states that are identified in the 1968–1976 CPS.

To qualify for unemployment benefits, an unemployed worker must satisfy three sets of eligibility requirements. First, she or he must have lost a job in the covered sector and be currently available for and actively searching for work. Individuals who are fired for cause and job quitters are disqualified for benefits in all but a handful of states. Second, the individual must have been unemployed for a period of time greater than a minimum waiting period (usually one week), and less than a maximum duration (usually 26 weeks, although individuals who received UI during a recent spell of unemployment may be eligible for less). In states that pay benefits to quitters, the waiting period is usually extended by 10–12 weeks for quitters. Finally, the individual must have received a minimum level of earnings, or worked a minimum number of weeks or hours in the twelve-month "base period" prior to the start of the spell of unemployment. In most states the base period is the first four of the five completed calendar quarters prior to the quarter in which the unemployment spell started. Some states also require that the individual have earnings or weeks of employment in at least two quarters during the base period.

To determine whether a worker meets the first set of requirements, we use information on whether an individual quit her or his last job, together with state-specific information on the eligibility of quitters. We use information on an individual's previous industry to check that they were employed in the covered sector.[14] We are unable to eliminate workers fired for cause, or to make any adjustment for those workers who are counted as unemployed by the CPS but are excluded from UI recipiency for failure to actively search for work.

We then use information on the duration of the spell in progress together with state-specific information on the waiting period and maximum duration of benefits to determine whether or

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14. After 1977 we exclude only postal workers, federal public administration workers, and ex-servicepersons from coverage. Workers who do not report a previous industry are included. We are not able to identify and exclude federal workers in nonpublic administration industries.

not the individual satisfies the second set of eligibility require-
ments. Finally, we compare information on reported earnings,
weeks worked, and hours per week in the previous calendar year to
the base period earnings and hours requirements of the state to
determine eligibility for benefits. Given the limited nature of the
earnings and labor supply data in the CPS, we make no attempt to
adjust for special provisions requiring earnings or hours in two or
more quarters of the base period, or to account for individual
variation in benefit duration.[15]

The most serious limitation of our imputation procedure is the
assumption that reported earnings for the previous calendar year
represent base period earnings for the in-progress unemployment
spell. There are two sources of error in this assumption. First, for
workers who have had an earlier spell of unemployment, the base
period may refer to the period of employment prior to the start of
the previous spell. Second, even for unemployed workers in their
first spell of unemployment, there is a significant lag in the
definition of the base period. For the roughly one half of spells in
progress in March that began in the same calendar year, the true
base period is the four-quarter period ending in the previous
September.[16] For longer spells, however, the base period may
represent the four-quarter period ending either in March or June
of the previous calendar year. Reported earnings for the previous
calendar year are therefore likely to understate base period earn-
ings for workers in longer spells, leading us to underestimate the
probability of eligibility for these spells. On the other hand, by
ignoring job search and extra earnings requirements, and individ-
ual variation in the maximum duration of benefits, we overstate
the extent of eligibility for some individuals in some states.[17]

Table I presents the fraction of insured unemployment in the
years 1977–1987, together with our estimate of the fraction of
workers eligible for unemployment benefits, which we shall call
FEU—the fraction of eligible unemployment. Over the 1977–1987

15. Many states have reduced benefit durations for individuals with a concen-
tration of base period earnings or weeks in a single quarter.
16. The median duration of in-progress spells ranges from 8–11 weeks over the
1977–1987 period. Of course, if a worker delays filing for UI or files after multiple
spells of unemployment, the duration of the current unemployment spell will be an
inaccurate indicator of the base period.
17. We have also calculated eligibility rates by assuming that base period
earnings are 52 times average weekly earnings last year. This assumes that the
individual worked at a constant weekly wage with no interruptions during the base
period. This produces virtually identical eligibility rates.

TABLE I

FRACTION OF INSURED UNEMPLOYMENT, FRACTION OF ELIGIBLE UNEMPLOYMENT, AND REASONS FOR INELIGIBILITY: 1977–1987

Year	Fraction of insured unemployment (× 100)	Fraction of eligible unemployment (× 100)	Estimated takeup rate (× 100)	Fraction of eligible unemployment under 1977 laws (× 100)	Reasons for ineligibility (percent of ineligible)			
					< Waiting period	> Maximum duration	Quit	Base period earnings or weeks
Average (1977–1987)	30.4	42.9	70.7	43.5	9.5	26.1	11.7	52.1
1977	31.2	41.7	74.8	41.7	9.5	27.5	12.2	50.0
1978	31.3	42.0	74.5	41.5	11.2	19.0	13.5	55.4
1979	33.2	44.3	74.9	44.0	11.4	17.5	15.8	54.3
1980	37.5	49.3	76.1	48.9	12.6	17.9	14.8	53.9
1981	31.6	44.0	71.8	44.3	9.6	26.8	10.8	52.1
1982	33.4	48.6	68.7	49.5	9.2	27.1	10.4	52.2
1983	27.9	41.2	67.7	42.3	6.4	41.1	7.0	45.2
1984	25.1	37.4	67.1	38.6	7.4	33.2	7.7	51.5
1985	27.6	40.0	69.0	41.3	7.8	27.5	10.2	54.1
1986	28.4	42.2	67.3	43.5	9.6	24.5	13.0	52.7
1987	27.3	41.5	65.8	43.0	10.3	24.7	12.9	51.9

Note. The fraction of insured unemployment is estimated from administrative data on UI recipients and aggregate labor force statistics. The fraction of eligible unemployment and reasons for ineligibility are estimated from unemployed workers in the March CPS. The estimated takeup rate is the ratio of the fraction of insured unemployment to the fraction of eligible unemployment. The omitted category in the set of reasons for ineligibility is "previous employment not in covered sector," which is 1 percent or less in all years. Estimated eligibility under 1977 laws is calculated by assuming that nominal provisions of 1977 laws are inflated at the rate of growth of average hourly earnings for private nonagricultural workers.

period we estimate that 42.9 percent of unemployed workers were eligible for UI. By comparison, only 30.4 percent of unemployed workers actually received benefits during this period. The third column of the table presents our estimate of the takeup rate for benefits, which is simply the ratio of the fraction of insured unemployment to the fraction of eligible unemployment. The reasons for ineligibility are presented in the right-hand columns of Table I.

The fractions of insured and eligible unemployment are plotted in Figure II. The two series have a similar cyclical pattern, although the gap between them diverges after 1981. There is a sharp drop in takeup rates in the early 1980s, from an average of 75 percent pre-1980 to an average of 67 percent post-1982. This decline in takeup rates explains all of the decline in the fraction of insured unemployment over the past decade: our estimated eligibility rates for UI benefits are virtually unchanged between 1977 and 1987.

The reasons for benefit ineligibility have a predictable cyclical pattern, with disqualifications for long durations peaking in the year after a business cycle trough. On average over the eleven-year period, 12 percent of ineligible workers were disqualified for quitting, 10 percent for having a spell duration shorter then the waiting period, and 26 percent for having a spell duration longer than the maximum duration of benefits. The remainder were disqualified for insufficient earnings or weeks worked in the base period or for having previous employment in the uncovered sector.

Although our estimate of the fraction of unemployed workers eligible for benefits is unchanged between 1977 and 1987, changes in state laws and changes in the composition of the unemployed labor force have actually had small but offsetting effects on eligibility. This is shown in the fourth column of Table I, where we estimate the fraction eligible for UI using the 1977 state laws to determine eligibility in each year.[18] If the 1977 state laws had remained in effect, with full indexation of base period earnings requirements, the estimated fraction of workers eligible for benefits would have been one percentage point higher in 1987 than 1977. Thus, holding constant the state laws, the pool of unem-

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18. To compensate for inflation over the past decade, we have inflated the nominal provisions of the 1977 state laws by the annual percent increase in average hourly earnings for nonagricultural production and nonsupervisory workers in private-sector employment.
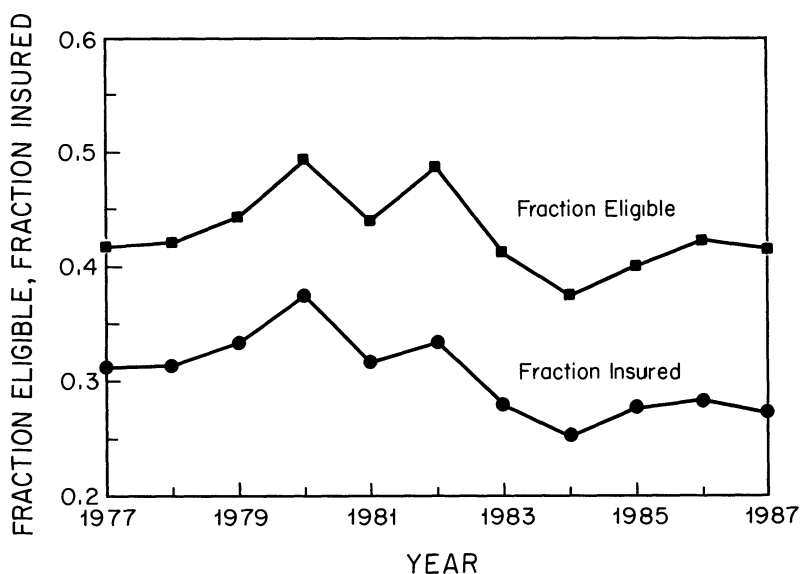
FIGURE II
Fraction Eligible and Fraction Insured

ployed workers was increasingly likely to be eligible for UI benefits. This trend was counteracted by tightening eligibility requirements in the state laws, however, which reduced the fraction of workers eligible for benefits by just enough to leave overall eligibility unchanged.

Further information on the changing composition of the unemployed labor force is presented in Table II. This table summarizes the characteristics of unemployed workers at the beginning and end of our sample period. The comparison is particularly useful because aggregate unemployment conditions were very similar in 1977 and 1987 (the civilian unemployment rate was 6.9 in 1977 and 6.1 in 1987). Between 1977 and 1987 the fraction of women in the unemployed labor force fell 2.8 percentage points, while the fraction of nonwhites increased by 5.6 points. Over the same period, the fraction of unemployed workers in the 16–24 year old range fell from 46 to 36 percent. In terms of previous employment history, the characteristics of unemployed workers at the beginning and end of our sample are very similar, although mean weeks worked in the previous year are slightly

TABLE II
CHARACTERISTICS OF UNEMPLOYED WORKERS: 1977–1987

Variable	1977	1987
1. Percent female	45.2	42.4
2. Percent nonwhite	19.6	25.2
3. Age		
Mean age	30.9	31.9
Percent age 16–24	46.3	35.7
4. Mean education	11.3	11.6
5. Weeks worked last year		
Mean weeks worked	21.7	23.4
Percent 0 weeks	25.9	24.8
Percent 1–23 weeks	18.0	15.6
Percent 14–26 weeks	17.5	17.0
Percent > 26 weeks	38.6	42.6
6. Unemployment spell		
Mean duration (weeks)	8.0	8.0
Percent < 4 weeks	35.0	37.5
Percent 5–13 weeks	31.6	31.8
Percent 14–26 weeks	16.8	16.1
Percent > 26 weeks	16.7	14.6
7. Region		
Northeast	27.1	15.0
North Central	24.3	27.0
South	27.8	36.3
West	20.8	21.7

Note. Estimated from unemployed workers in March CPS, weighted by CPS sampling weights.

higher in 1987. The mean duration of in-progress spells of unemployment is identical in 1987 and 1977.[19]

One of the most significant changes among unemployed workers over the past decade has been in their geographic distribution. Between 1977 and 1987 the fraction of unemployed workers in the Northeast region fell more than ten percentage points, while the fraction in the South rose by almost as much. Not shown in Table II is the distribution of unemployed workers across indus-

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19. At first glance these results may seem to contradict the recent conclusions of Murphy and Topel [1987], who emphasize the shift toward longer spells of unemployment over the past decade. There are two differences between our analysis and theirs, however. First, our analysis is based on the in-progress spells of unemployed workers, whereas theirs is based on the weeks of unemployment in the previous year. Second, Murphy and Topel end their analysis in 1985. During the last two years of our sample period the duration of in-progress spells of unemployment dropped significantly.

tries and occupations, which has remained relatively stable over these eleven years.

B. Regional Differences in FIU and FEU

In light of the significant regional shifts in unemployment during the past decade, it is useful to look more closely at regional differences in the use of UI. Table III provides information on the fraction of insured unemployment (FIU) in each of the nine Census regions in 1977 (column (1)), the change in FIU between 1977 and 1987 (column (2)), the fraction of eligible unemployment (FEU) in each region in 1977 (column (3)), the change in FEU (column (4)), and the change in FEU that would have occurred if the 1977 laws had remained in effect over the entire sample period (column (5)). Columns (6) and (7) show the resulting takeup rates in each region in 1977 and 1987.

A striking feature of Table III is the extent of regional differences in recipiency, eligibility, and takeup rates for UI. The percentage of unemployed workers receiving UI benefits in 1977 varies from 21 percent in the West South Central region to 40 percent in the Mid-Atlantic region. Similarly, the percent of the unemployed workers eligible for UI in 1977 varies from 36 percent in the East North Central region to 50 percent in New England. The resulting takeup rates range from 99 percent in the Mid-Atlantic region to 51 percent in the Mountain region.

There are also significant regional differences in the *changes* in recipiency and eligibility rates for UI. The New England and Pacific regions actually show increases in the fraction of insured unemployment over this period. In the East South Central region, by comparison, FIU dropped eight percentage points, more than twice the national average decline. Similarly, although the fraction of eligible unemployment shows little change in the aggregate, it has increased in some regions and fallen in others. In most regions it is clear that changes in eligibility provide only a partial explanation for changes in overall UI usage. Notable changes in takeup rates occurred in seven of the nine census regions.

A comparison of the numbers in columns (4) and (5) of the table shows that the changes in eligibility that have occurred within the regions are the result of changes in laws as well as changes in the eligibility characteristics of unemployed workers. If UI laws had remained unchanged, changes in the characteristics of unemployed workers would have had varying effects on UI eligibility, with some regions experiencing increases in FEU (the Middle

TABLE III
REGIONAL ESTIMATES OF FRACTION OF INSURED UNEMPLOYMENT AND FRACTION OF ELIGIBLE UNEMPLOYMENT UNDER ACTUAL AND 1977
LAWS, AND TAKEUP RATES: 1977–1987

| Region | Fraction of insured unemployment (× 100) | | Fraction of eligible unemployment (× 100): | | | Takeup rates | |
	1977 (1)	1987–1977 (2)	1977 (3)	1987–1977 (4)	1987–1977 1977 laws (5)	1977 (6)	1987 (7)
All U. S.	31.3	−4.4	41.8	−0.4	+1.2	0.76	0.65
New England	37.3	+2.6	50.3	−4.2	−3.3	0.74	0.85
Middle Atlantic	39.7	−0.8	40.2	+5.7	+7.6	0.99	0.85
East North Central	31.6	−6.1	36.0	−0.6	+2.3	0.87	0.73
West North Central	31.9	−5.9	50.0	−2.7	−1.6	0.64	0.56
South Atlantic	23.2	−3.9	40.2	−1.8	+1.4	0.58	0.52
East South Central	27.8	−8.0	47.9	−10.2	−7.3	0.58	0.53
West South Central	20.5	0.0	46.6	−4.5	−3.3	0.44	0.49
Mountain	22.6	−0.5	44.5	+2.1	+3.2	0.51	0.48
Pacific	33.9	+2.5	40.2	+4.0	+2.3	0.85	0.83

Note. All data are based on weighted sums of state variables, where the weights are the fractions of total unemployment in the state. Thus, column (6) does not equal column (1) divided by column (3), since the weighted sum of ratios is not the equal of the ratio of the weighted sums.

Atlantic, South Atlantic, East North Central, Mountain, and Pacific regions), and some regions experiencing decreases (the New England, West North Central, East South Central, and West South Central regions). With the exception of the Pacific region, however, all regions show smaller increases (or larger decreases) than would have occurred without any changes in the laws. These comparisons suggest that legislative changes tightened UI eligibility in all but one census region.

The evidence in Table III confirms the importance of regional variation in eligibility and usage of UI. Differences in the fraction of insured unemployment arise from differences in eligibility for UI and from differences in takeup rates. The dispersion in takeup rates is particularly large, ranging from 48 to 85 percent in 1987. This suggests one explanation for the nationwide decline in takeup rates over the past decade: as unemployment has shifted from Northeastern states (with higher than average takeup rates), to Southern states (with lower than average takeup rates), the aggregate takeup rate has declined. We explore this explanation more thoroughly in Section IV below.

C. Changes in UI Eligibility and Receipt Among the Larger States: 1968–1987

Our analysis of UI eligibility and receipt at the national level is limited by the absence of state identifiers in the CPS data prior to 1977. Starting in 1968, however, individual state identifiers are available for eight larger states: California, Florida, Illinois, New Jersey, New York, Ohio, Pennsylvania, and Texas. On average during the twenty-year period from 1968 to 1987, these states accounted for 48 percent of aggregate unemployment and 51 percent of UI recipients. An examination of the longer-term trends in eligibility and UI recipiency in these states therefore provides a check on our national results, and allows us to view recent changes in the fraction of insured unemployment from a longer perspective.

Table IV provides information similar to the data in Table I for the eight larger states over the period from 1968 through 1987. The first and second columns show the fraction of insured unemployment and the estimated fraction of eligible unemployment in each year. The third column of the table presents the estimated fraction of eligible unemployment in each year under the assumption that the 1977 state laws were in effect. Finally, the fourth column shows the estimated takeup rate for UI benefits among eligible workers.

TABLE IV
FRACTION OF INSURED UNEMPLOYMENT AND FRACTION OF ELIGIBLE UNEMPLOYMENT
UNDER ACTUAL AND 1977 LAWS IN EIGHT STATES: 1968–1987

Year	Fraction of insured unemployment (× 100)	Fraction of eligible unemployment (× 100): Actual	Under 1977 laws	Estimated takeup rate (× 100)
Average (1968–1987)	33.7	42.2	42.8	81.2
1968	36.4	40.2	42.9	90.5
1969	35.2	40.5	42.5	86.9
1970	38.5	43.4	46.6	88.7
1971	37.4	47.6	49.2	78.6
1972	34.0	40.5	42.3	84.0
1973	33.4	40.1	41.1	83.3
1974	38.3	42.6	43.4	89.9
1975	42.6	50.0	50.6	85.2
1976	34.3	38.3	38.3	89.6
1977	33.2	39.2	39.2	84.7
1978	34.1	40.8	39.7	83.6
1979	36.3	41.0	40.1	88.5
1980	38.8	47.0	46.2	82.6
1981	32.8	42.5	42.6	77.2
1982	34.3	47.7	47.9	71.9
1983	29.7	40.1	40.4	74.1
1984	26.8	38.5	38.8	69.9
1985	28.7	38.6	39.5	74.4
1986	30.0	42.3	42.4	70.9
1987	29.2	42.4	42.5	68.9

Note. Eight states included in calculations are California, Florida, Illinois, New Jersey, New York, Ohio, Pennsylvania, and Texas.

Between the late 1960s and the late 1970s, the fraction of insured unemployment in the eight larger states fell by one half of a percentage point, from 35.8 percent in 1968–1969 to 35.2 percent in 1978–1979. This decline came about through a slight decline in takeup rates, from 88.7 percent in 1968–1969 to 86.1 percent in 1978–1979. The fraction of eligible unemployment, on the other hand, actually increased slightly. Comparisons of the second and third columns of Table IV suggest that this small net increase in eligibility occurred as a result of liberalized eligibility laws, offset by a decline in the eligibility characteristics of unemployed workers.

Between the late 1970s and 1987, however, there was a five-percentage point drop in the fraction of insured unemploy-

ment in the eight larger states. As in the nation as a whole, this decline occurred through a sharp drop in takeup rates in the early 1980s, from 86 percent in 1978–1979 to 71 percent in the post-1982 period. The decline in takeup rates was partly offset by an increase in the fraction of eligible unemployment, due mainly to an improvement in the eligibility-determining characteristics of unemployed workers.

Data for the eight larger states suggest that the decline in takeup rates that occurred in the early 1980s was a marked departure from the experience of the previous decade. If earlier trends had continued, in fact, takeup rates would have been ten to fifteen points higher post-1983, and the fraction of insured unemployment would have been four to six points higher. This finding underscores the critical role that changes in takeup rates have played in the declining fraction of insured unemployment.

IV. The Determinants of Takeup Rates:
A State-Level Analysis

The preceding analysis suggests that changes in takeup rates among eligible workers are responsible for virtually all of the recent decline in the nationwide fraction of insured unemployment. In this section we present a statistical analysis of the determinants of state-level takeup rates, using observations on the 50 states over the eleven-year period from 1977 to 1987. We estimate the correlates of takeup rates across states and over time, using data on the institutional structure of the state's UI system, the political climate of the state, and the demographic and work-related characteristics of the unemployed labor force.[20]

The first set of explanatory variables that we consider are characteristics of the state UI system: the benefit replacement ratio (the average weekly benefit amount divided by the average weekly wage in covered employment); the fraction of employed workers in covered employment; the presence of eligibility requirements for two or more quarters of earnings in the base period; the presence of other miscellaneous eligibility requirements; and the disqualification rate for UI recipients. The latter is defined as the percent of claims that are disqualified for any of three sets of

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20. Available upon request from the authors are appendix tables presenting the means of the explanatory variables as well as the fraction of insured and eligible unemployment and takeup rates for each of the 50 states in 1977–1987.

reasons: refusal of suitable work; not being available for or actively seeking work; or a determination that the previous job was lost due to misconduct.[21]

The coverage of the UI system and the presence of various eligibility requirements presumably affects the fraction of eligible workers, but should not necessarily influence the takeup rate among eligible workers. Correlations between these variables and our estimated takeup rates therefore suggest systematic biases in our eligibility imputations. The benefit replacement rate, by comparison, should be uncorrelated with errors or omissions in our eligibility imputations. This variable measures the generosity of the UI program and is hypothesized to directly influence takeup behavior. Takeup rates may also be affected if states decide to "crack down" on claimants by enforcing rules more strictly or monitoring claimants more carefully. Such changes are proxied by the disqualification rate.

The next variable that we consider is the unionization rate of employed workers in the state. Unions may affect takeup behavior by increasing their members' awareness of UI programs or expediting their applications.

Unionization rates by state for 1977–1981 are taken from tabulations of the May CPS reported by Kokkelenberg and Sockell [1985, Table 5]. Rates for 1982–1987 are taken from unpublished tabulations of the CPS provided to us by Alan Krueger.[22]

Next we consider the political climate of the state, as measured by the fraction of the state's Congressional delegation who are Democrats. Given the timing of changes in takeup rates, which coincided with the first term of the Reagan presidency, they may be part of a national shift toward a more conservative political mood. If increasing conservatism leads to stricter control over unemployment assistance by the states, this would be measured by our institutional UI variables (for example, the disqualification rate). But there may be additional effects on individual behavior if people are generally more reluctant to use government assistance in more conservative political climates.

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21. We omit from the overall disqualification rate those claims that are denied because the claimant quit his or her previous job. As noted above, our eligibility calculations explicitly exclude individuals who quit from the count of eligible unemployed (in states that exclude quitters). Adding the disqualification rate due to quits as a variable in the takeup regressions results in an insignificant coefficient.

22. Krueger's tabulations for 1981 match those of Kokkelenberg and Sockell almost exactly. Because the 1982 CPS does not report union status, we estimate it as the average of 1981 and 1983 state unionization rates.

The final set of variables that we consider are characteristics of the unemployed workers in each state. These variables are meant to capture differences in takeup propensities across different groups. We use samples from the March CPS to estimate the fraction of unemployed workers who were previously employed in eight major industries and nine major occupations,[23] the fraction who were previously self-employed, average weeks worked in the previous year, and the average duration of the current in-progress unemployment spell. In addition, we tabulate the demographic characteristics of unemployed workers: the fraction aged 16–24, the fraction female, and the fraction nonwhite.

Table V presents a series of estimated models for the logarithm of the takeup rate in a given state and year.[24] In interpreting the results in this table, it is useful to keep in mind that the dependent variable, log(FIU/FEU), is based on a noisy estimate of eligibility in each state and year. Provided that the measurement error in FEU is uncorrelated with our explanatory variables, however, this will not lead to any systematic bias in our estimated coefficients.[25]

The first column of Table V presents a model that includes only year effects, with no other variables included.[26] The estimated coefficient in row 15 indicates that takeup rates declined by twelve points between the first three years and the last three years of the sample. The question we seek to answer is how much of this decline can be explained by controlling for other variables in the regression.

Column (2) includes state fixed effects in the regression. This

23. The industries are agriculture, mining, construction, manufacturing, transportation communications and public utilities, trade, services, and public administration. The occupation groups are professionals, managers, sales workers, clerical workers, craftsmen, operatives, laborers, service workers, and private household workers. Since the CPS occupation codes were changed in 1981, we were forced to match the two coding systems. These data are computed among unemployed workers who report industry and occupation on their previous job.

24. The estimates in Table V are obtained by weighted least squares, where the weight for each state-year observation is the fraction of unemployed workers in that state and year. Because disqualification rates are not available in Virginia for all years, the number of state-year observations in the regressions in Table V equals 542.

25. If log(FIU) is used as the dependent variable, placing log(FEU) on the right-hand side of the equation, the results do not support the proportionality hypothesis, suggesting that the measurement errors in FEU are large enough to create a serious downward bias in its coefficient.

26. Eleven independent year effects are entered into each model in Table V, with the sum of the coefficients between 1977 and 1979 normalized to zero. (A test of equality between the year effects in these three years cannot reject the hypothesis in any of the models.) We report in the table the mean of the estimated year effects in 1985–1987. (A test of equality between the year effects in these three years cannot reject the hypothesis in any of the models.)

TABLE V
REGRESSION ANALYSIS OF STATE TAKEUP RATES
(STANDARD ERRORS IN PARENTHESES)

Explanatory variable:	Dependent variable: Log of state takeup rate				
	(1)	(2)	(3)	(4)	(5)
State effects:	No	Yes	Yes	Yes	No
1. Replacement rate	—	—	1.62	1.55	0.90
			(0.31)	(0.31)	(0.18)
2. Coverage rate	—	—	0.56	0.54	0.91
			(0.33)	(0.33)	(0.19)
3. Earnings required in 2 quarters	—	—	−0.02	−0.03	−0.00
			(0.04)	(0.04)	(0.02)
4. Other eligibility requirements	—	—	−0.03	−0.04	−0.12
			(0.05)	(0.05)	(0.03)
5. Disqualification Rate (× 1,000)			−0.34	−0.34	−0.53
			(0.08)	(0.08)	(0.05)
6. Unionization rate (share)	—	—	0.67	0.73	1.17
			(0.27)	(0.28)	(0.16)
7. Democrat's share of house delegation	—	—	−0.06	−0.06	−0.07
			(0.07)	(0.07)	(0.05)
Characteristics of unemployed pool:					
8. Fraction age 16–24	—	—	0.00	0.04	0.08
			(0.13)	(0.14)	(0.16)
9. Fraction nonwhite	—	—	0.03	−0.05	0.01
			(0.13)	(0.14)	(0.08)
10. Fraction female	—	—	0.13	0.03	0.21
			(0.15)	(0.17)	(0.19)
11. Fraction self-employed	—	—	0.62	0.84	1.01
			(0.41)	(0.43)	(0.51)
12. Average weeks worked last year (÷10)	—	—	−0.25	−0.24	−0.12
			(0.04)	(0.04)	(0.04)
13. Average duration of current spell (÷10)	—	—	−0.05	−0.06	−0.04
			(0.03)	(0.03)	(0.03)
14. Industry & occupation shares among unemployed	No	No	No	Yes	Yes
15. Mean of year effects for 1985–1987	−0.12	−0.06	−0.03	−0.05	−0.03
	(0.04)	(0.02)	(0.03)	(0.03)	(0.04)
16. R^2	0.03	0.75	0.82	0.83	0.69

Notes. Sample consists of 542 state and year observations on 50 states between 1977 and 1987. Due to the absence of data on disqualification rates, eight observations for Virginia are omitted. The mean and standard deviation of the dependent variable are −0.3856 and 0.2922, respectively. All equations include eleven unrestricted year effects, normalized by setting the average year effects for 1977, 1978, and 1979 equal to zero. Equations are estimated by weighted least squares, using as weights the number of unemployed workers in the state and year.

greatly increases the explanatory power of the regression and reduces the unexplained decline in takeup rates from twelve to six percentage points. The implication is that interstate variation in takeup rates, together with changes in the distribution of unemployment, account for half of the decline in takeup rates over this time period.

In column (3) we include the main explanatory variables in the regression, and in column (4) we add control variables for the shares of unemployment in major industries and occupations. With fixed effects controlling for permanent differences in takeup rates across states, these regressions indicate the ability of the explanatory variables to explain changes in takeup rates within states over time. The results show a further increase in the explanatory power of the regression, and a slight reduction in the unexplained time trend.

In column (5) we omit state fixed effects, to see how effectively our explanatory variables explain changes in takeup rates both over time and across states. Variables with large cross-state variation have larger and more significant coefficients in column (5) than in column (4). Although the removal of state fixed effects decreases the explanatory power of the regression, it also decreases the unexplained time trend in takeup rates. Three quarters of the decline in state takeup rates over time can be explained by the variables in this model, leaving one quarter of the decline still unexplained.

With respect to the individual determinants of takeup rates, a number of conclusions emerge. First, estimated takeup rates are strongly related to the state's benefit replacement rate. This correlation is robust to the presence or absence of state effects.[27] Second, increases in the coverage of the UI system among employed workers lead to increases in estimated takeup rates. This finding suggests that we have not been entirely successful in measuring the coverage status of previous employment. By comparison, quarterly earnings requirements and other eligibility restrictions appear to have little effect on estimated takeup rates. Third,

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27. We were concerned that the average replacement rate may be an endogenous variable because it is based on average weekly unemployment benefits among those receiving UI. As a check, we estimated the model using maximum replacement rates (based on maximum weekly unemployment benefits.) We also tried instrumental variables, using the maximum weekly unemployment benefit as an instrument for the replacement rate. In either case, the variable remained significant; the instrumental variables procedure actually produced a slightly larger coefficient.

the disqualification rate in a state has a significant negative effect on takeup rates. Note, however, that average disqualification rates have actually *fallen* slightly from 0.047 to 0.045 over this time period, implying that the net effect of this variable would have been a *rise* in takeup rates, all else equal. Thus, there is no evidence that declining takeup is due to increasing administrative strictness by state programs.[28]

The fraction of unionized employees in the state has a strong positive effect on estimated takeup rates, even when state effects are included. Unionization rates over this time period declined by 5.4 points; an estimated coefficient of 0.73 implies that about one third of the 0.12 point decline in log takeup rates can be explained by declining unionization. When state effects are excluded from the model, the aggregate decline in unionization together with the shift of unemployed workers from more to less unionized states accounts for about half of the decline in takeup rates over time.

The share of the state's Congressional delegation that is Democratic has an insignificant effect on takeup rates. The coefficient indicates that takeup rates are slightly lower in states with a stronger Democratic party, not the expected result. To further test this result with alternative data, we attempted to relate changes in presidential voting patterns across states to changes in takeup rates. We correlated the change in the percent of a state's vote going to the Democratic presidential candidate between 1976 and 1980 with the change in the UI takeup rate between the late 1970s and the mid-1980s. The resulting correlation coefficient of -0.05 further demonstrates that there seems to be little relationship between general voting/political patterns in the states and changes in takeup rates.

The race and sex composition of the unemployed labor force, the fraction of unemployed workers in the 16–24 year old age group, and the average duration of in-progress unemployment spells have small and statistically insignificant effects on estimated takeup rates throughout Table V. The fraction of unemployed workers who were previously self-employed shows a strong positive effect on takeup rates.

The average number of weeks worked in the previous year by unemployed workers has a significantly negative effect. We suspect

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28. In other regressions we also included the number of UI offices in the state, another policy variable the states can use to restrict UI availability. The coefficient was uniformly small and insignificant, with an unexpected negative sign.

that this occurs for two reasons. First, this variable is a proxy for the length of the unemployment spell; the more weeks worked last year, the shorter one's current unemployment spell. Indeed, there is a significant negative correlation between average weeks worked last year and the average duration of in-progress spells. When average weeks worked is omitted from the model, the average duration of spells variable has a positive and significant coefficient (the expected result). Thus, the negative coefficient of this variable in Table V may signal that shorter spells are anticipated by some workers who do not bother to apply for UI. Second, because some states have eligibility requirements for minimum weeks of work in the base period, increases in the number of weeks worked in the previous year (our estimated base period) will lead to an increase in the estimated fraction of eligible unemployment. To the extent that weeks last year overstate work experience during the true base period, there will be a negative correlation between reported weeks and the measurement error in estimated takeup rates, further strengthening the negative coefficient.

The industry and occupation shares of unemployed workers are only marginally significant determinants of estimated takeup rates.[29] Corson and Nicholson [1988] suggest that much of the decline in UI receipt may be due to a shift in unemployment across industries, but there is little evidence of this in our analysis. The coefficients on the industry share variables suggest that takeup rates are lower among workers from agriculture, construction, and wholesale and retail trade. The coefficients on the occupation shares indicate lower takeup rates among managers and operatives.

In summary, the analysis in Table V indicates that at least half of the decline in takeup rates over the past decade is due to a shift in unemployment from high to low takeup states. State unionization rates have a particularly strong effect on state takeup rates, accounting for differences between states in takeup levels and for almost half of the decline in takeup rates within states over time. The replacement rate and the disqualification rate within a state also affect takeup rates over time, although trends in these variables cannot explain the recent decline in takeup. Similarly, the characteristics of unemployed workers also affect takeup rates across states, but are less important in explaining changes over time. Overall, the set of explanatory variables in Table V can

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29. For example, the probability value of an exclusion test for the industry and occupation shares of unemployed workers is 0.074.

explain up to three quarters of the aggregate decline in takeup rates between 1977 and 1987.

V. A MICRO-DATA ANALYSIS OF UI TAKEUP RATES

A major shortcoming of the preceding analysis is the absence of accurate information on whether or not a given individual actually receives unemployment compensation. We turn now to an analysis of micro-data that contain information on both eligibility and receipt of UI among a selected group of workers. During the early 1980s the Panel Study of Income Dynamics (PSID) included in its annual survey a series of questions on the receipt of UI among household heads who experienced at least one spell of unemployment in the previous year. In response to these questions, over one fifth of non-UI recipients claimed they were eligible for benefits. Analysis of takeup rates among these PSID respondents complements our aggregate analysis.

In its 1981–1983 surveys, the PSID asked each household head about the last spell of unemployment in the previous year. We have combined information from these questions in the three survey years into a sample of 2,280 "last spells" of unemployment. For each spell, we know when it started, when it ended (if before the 1983 interview),[30] the characteristics of the unemployed worker, and a variety of information about the receipt of UI benefits during the spell.

Unemployment compensation was received in 52 percent of the spells. By comparison, aggregate statistics for the same period suggest that roughly one third of unemployment spells generated a benefit claim.[31] Thus, a spell in our PSID sample was more likely to generate UI benefits than a representative spell from the same time period.

The PSID questionnaire asks non-UI recipients whether or not they thought they were eligible for UI and, if so, the reason for nonreceipt. One third of nonrecipients reported that they thought they were eligible for benefits, or that they were unsure about their

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30. A total of 24 percent of the spells in our sample are censored.
31. An estimate of the aggregate fraction of unemployment spells with UI payments can be obtained by dividing the number of first claims for UI benefits (in a year) by the number of unemployment spells in the year. Unfortunately, there is no accurate count of the number of unemployment spells in a year. An estimate can be obtained by multiplying the average number of unemployed workers per week by 52 and dividing by an estimate of the average completed duration of unemployment spells. The estimate in the text assumes that the average duration of completed spells was thirteen weeks.

eligibility status. Among these, however, one third indicated that they had actually applied for benefits and had been turned down. If UI recipients and nonrecipients who reported themselves as eligible (excluding those turned down for benefits) are counted as eligible, the estimated fraction of PSID spells eligible for UI compensation is 62 percent. The implied takeup rate for benefits among eligible spells is 83 percent. By comparison, aggregate data for the 1980–1982 period show an eligibility rate of 47 percent and a takeup rate of 72 percent.

There are several possible explanations for the differences in eligibility and takeup rates between the PSID sample and the aggregate data. Perhaps most importantly, the PSID is a selective sample of unemployment spells, since it is based on the unemployment experiences of household heads. By the same token, owing to the sampling frame of "last spells," the PSID sample overrepresents longer spells, which may be associated with higher usage of unemployment benefits.

Second, the unit of observation in the PSID sample is a spell, rather than a week, of unemployment. In most cases the fraction of insured weeks in a spell of unemployment is less than unity. Thus, one might expect the fraction of spells eligible for or receiving benefits to exceed the fraction of weeks eligible for or receiving benefits. On the other hand, if most weeks of unemployment are generated by individuals in long spells of unemployment, and if individuals with longer spells are more likely to apply for benefits (as we find below), then the fraction of unemployment weeks with benefits may well exceed the fraction of unemployment spells with benefits.

Finally, the estimated eligibility and takeup rates for the PSID sample are based on self-reported eligibility. Some potentially eligible nonrecipients may have been unaware of their eligibility status and reported themselves as ineligible, leading to a lower estimate of eligibility and a higher estimate of takeup rates in the PSID sample.

In order to compare self-reported eligibility with estimated eligibility as constructed in our aggregate analysis of the CPS data, we have applied the same eligibility imputation methods to the PSID sample.[32] The results of a comparison between estimated

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32. We can construct slightly better eligibility estimates from the PSID data because of its longitudinal nature. Thus, we define the base period as the calendar year prior to the year in which the unemployment spell began, rather than as the previous calendar year. This should improve the estimate of base period earnings for individuals in longer unemployment spells.

TABLE VI
CROSS TABULATIONS OF ESTIMATED ELIGIBILITY FOR UI WITH ACTUAL RECEIPT OF
BENEFITS AND SELF-REPORTED ELIGIBILITY: PSID UNEMPLOYMENT SPELLS,
1980–1982

	Estimated eligibility		
	Eligible	Ineligible	Total
I. Cross tabulation with UI receipt (percent of cases)			
1. Recipient	44.0	7.8	51.8
2. Nonrecipient	26.7	21.5	48.2
3. Total	70.7	29.3	100.0
II. Cross tabulation with self-reported eligibility (percent of cases)			
1. Eligible	50.6	11.5	62.2
and receiving UI	44.0	7.8	51.9
and not receiving UI	6.6	3.7	10.3
2. Ineligible	20.0	17.8	37.8
3. Total	70.7	29.3	100.0

Note. Sample consists of 2,280 unemployment spells reported as "last spells last year" in the PSID survey, 1980–1982. Estimated eligibility constructed by authors using UI laws in the state of residence and income from employment for the year before the spell began.

eligibility and self-reported eligibility are presented in Table VI. The upper panel of the table presents a cross-tabulation of estimated eligibility with reported receipt of UI benefits. The results are quite favorable: in 65 percent of cases, estimated eligibility agrees with the reported receipt of benefits. The discrepancies between estimated eligibility and benefit receipts are of two types. In 8 percent of cases we make a clear error in our eligibility calculations: these are cases where the individual reports UI recipiency but is estimated to have been ineligible. In the remaining 27 percent of cases we estimate that the individual is eligible, but he/she does not report receiving UI.[33] Disagreements in this cell, however, represent a combination of erroneous eligibility imputations and incomplete takeup rates for benefits.

The lower panel of the table presents a cross-tabulation of estimated and self-reported eligibility. Estimated eligibility is

33. Individuals who report themselves eligible for UI but did not receive it are asked their reasons for nonparticipation. A tabulation of these responses is available from the authors upon request. About two thirds indicate either that they wished to avoid contact with administrative red tape or government charity, or that they "didn't need it."

consistent with self-reported eligibility status in 69 percent of cases. In 11 percent of cases we estimate that individuals are ineligible, but they either report receiving UI or report themselves as eligible nonrecipients. Last, in 20 percent of cases we estimate that individuals are eligible for benefits, but they report themselves as ineligible. Again, the latter group may include some individuals who were actually eligible for benefits, but who erroneously reported themselves as ineligible. On balance, these cross-tabulations suggest that our simple imputations of eligibility status are valid in a majority of cases.

Another useful test of our eligibility imputations can be obtained by comparing the average state takeup rates based on self-reported eligibility in the PSID with the average state takeup rates between 1980 and 1982 from our aggregate analysis. The estimated correlation coefficient between the two sets of takeup rates is 0.47, and the implied regression coefficient of the CPS-based takeup rate on the PSID-based rate is 0.68 (with a standard error of 0.20). This finding lends strong support to our eligibility imputations in Section III, and suggests that the interstate variation in estimated takeup rates is a behavioral phenomenon, rather than a systematic measurement problem.

The micro-data from the PSID sample permit us to estimate directly the individual determinants of takeup behavior. Table VII presents a logistic regression model for the probability of receiving UI benefits, estimated over the sample of 1,359 individuals whose last unemployment spell started in 1980, 1981, or 1982, and who reported either receiving UI or being eligible for UI.[34] The dependent variable is a dichotomous indicator for actual receipt of benefits. We have data on a wide variety of individual characteristics in the PSID sample, including age, race, sex, education, family size, and union status and hours of work in the previous survey. We also include in our analysis a dummy variable for short unemployment spells (less than four weeks), regional dummy variables for the nine Census regions, and dummy variables for the individual's occupation and industry (reported in the previous survey).[35] Finally, we match with the PSID data the 1980–1982 state-level data

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34. The results reported in Table VII are based on self-reported eligibility. We have also replicated these results using estimated eligibility. While the estimated coefficients vary slightly, the qualitative results and statistical inferences are similar using the two measures.
    35. These are the same occupational and industrial dummy variables included in Table V.

TABLE VII

MODELS FOR THE PROBABILITY OF UI RECIPIENCY AMONG ELIGIBLE INDIVIDUALS
(STANDARD ERRORS IN PARENTHESES)

	Mean of variable	Logistic probability models			
	(1)	(2)	(3)	(4)	(5)
1. Replacement rate	0.37	—	0.093	−0.165	−0.267
			(0.225)	(0.279)	(0.293)
2. Coverage rate	0.86	—	0.600	0.071	0.022
			(0.231)	(0.314)	(0.322)
3. Earnings required in 2 quarters	0.55	—	0.023	0.013	−0.004
			(0.024)	(0.032)	(0.033)
4. Other eligibility requirements	0.14	—	0.112	0.064	0.048
			(0.043)	(0.053)	(0.055)
5. Disqualification rate (× 1,000)	35.63	—	−0.179	−0.148	−0.141
			(0.042)	(0.059)	(0.062)
6. Democrat's share of house delegation	0.58	—	0.202	0.231	0.247
			(0.054)	(0.061)	(0.065)
Individual characteristics:					
7. Age 16–24	0.16	−0.41	−0.046	−0.055	−0.051
		(0.024)	(0.025)	(0.026)	(0.028)
8. Nonwhite	0.16	−0.058	−0.053	−0.032	−0.031
		(0.025)	(0.027)	(0.028)	(0.031)
9. Female	0.19	−0.029	−0.047	−0.037	−0.050
		(0.023)	(0.024)	(0.024)	(0.027)
10. Hours worked previous year (÷ 1,000)	1.892	0.403	0.428	0.472	0.535
		(0.149)	(0.154)	(0.160)	(0.178)
11. Spell less than 4 weeks	0.36	−0.117	−0.138	−0.131	−0.158
		(0.020)	(0.021)	(0.022)	(0.024)
12. Years of education	11.71	−0.003	−0.001	−0.002	0.004
		(0.005)	(0.005)	(0.005)	(0.005)
13. Family size (number)	2.91	0.020	0.020	0.022	0.022
		(0.008)	(0.008)	(0.008)	(0.008)
14. Union member previous year	0.42	0.126	0.098	0.093	0.060
		(0.024)	(0.025)	(0.026)	(0.029)
15. Dummy for spell starting in 1981	0.33	0.048	0.051	0.053	0.062
		(0.024)	(0.025)	(0.025)	(0.027)
16. Dummy for spell starting in 1982	0.35	0.058	0.044	0.043	0.051
		(0.024)	(0.025)	(0.025)	(0.027)
17. 9 regional controls		No	No	Yes	Yes
18. Dummies for occupa-		No	No	No	Yes
Likelihood value		−480	−456	−446	−424

Notes. Sample consists of 1,359 individuals who report a spell of unemployment starting in 1980, 1981, or 1982, and who report that they either received or were eligible for UI. The dependent variable is an indicator for recipiency of UI. The mean of the dependent variable is 0.8643. Models in columns (2)–(5) are logistic probability models, estimated by maximum likelihood. The estimated coefficients have been multiplied by 0.1173 to give the marginal effects of the explanatory variables on the probability of receiving UI. Observations are weighted by sampling weights in all columns.

on institutional parameters of the UI program and the political climate of the state (as defined for Table V).

The means of the explanatory variables in the regression model are presented in the first column of the table. Estimated coefficients and associated standard errors for a model that includes only individual characteristics are presented in column (2). State-level characteristics are included in the model in column (3), while the model in column (4) includes controls for nine regions, and the model in column (5) includes seventeen dummy variables for past occupation and industry of unemployed workers. In contrast to the state-level takeup regressions in Table V, all of the demographic variables and household characteristics are statistically significant in the micro-level takeup regressions, except race and years of education. Younger workers and female workers are less likely to utilize UI, while union workers, workers from larger families, and workers who had higher hours on their previous job are more likely to utilize UI.

Workers whose unemployment spells are quite short are less likely to utilize UI, suggesting that many of them may have anticipated the short spell. Note that there is a nonlinear relationship between unemployment duration and takeup rates: takeup rates associated with spells longer than one month are very similar.[36]

In contrast to the aggregate data, the effects of the state-level parameters on individual takeup rates are largely insignificant. The aggregate estimates (Table V) show a strong positive effect of replacement rates on takeup rates, while the micro-level estimates indicate that the effect is negative, although poorly determined. We suspect that state-level replacement rates are too crude to provide useful information in a micro-data model. The takeup regressions in the aggregate data also indicate a positive effect of coverage rates on takeup rates (a result we interpreted as indicating bias in our eligibility imputations). The micro-level estimates, however, are unaffected by state coverage rates, suggesting that these data contain fewer errors in eligibility. As in the aggregate estimates, earnings and other eligibility requirements have no effect on individual takeup rates.

36. We have obtained a similar finding by tabulating the likelihood of UI recipiency among workers in the CPS Displaced Worker Supplements. Displaced workers who report one to four weeks of unemployment are much less likely to report receiving UI, but all those with one month or more of unemployment have roughly similar UI recipiency rates.

Similar to the aggregate estimates, a higher state disqualifica-
tion rate for UI claimants shows a negative and significant effect on
individual takeup rates, further verifying that state enforcement
can affect UI participation. Surprisingly, the political orientation of
the state, as measured by the Democrat's share of the Congres-
sional delegation, has a positive effect on takeup rates in Table VII.
Given the insignificant coefficients for this variable in the state
level analysis, it is unclear how to interpret its effect in the
micro-data estimates.

Inclusion of regional controls increases the explanatory value
of the equation only marginally and has little effect on the other
coefficients.[37] The inclusion of all seventeen occupational and
industry controls had a more substantial effect on the fit of the
model,[38] although few of the individual coefficients on these
variables are highly significant and the coefficients on other
variables are unchanged.

We are particularly concerned about the estimated year effects
in the PSID data. The aggregate data in Table I suggest that takeup
rates fell from an average of 75 percent in the 1977–1980 period to
an average of 67 percent in the 1982–1987 period. A similar pattern
is revealed by the estimated year effects from the takeup models in
Table V: relative stability before 1980 and after 1982, with a sharp
decline between 1980 and 1982. The year effects in the micro-level
analysis in Table VII, by comparison, show a slight increase in
takeup rates in 1981 and 1982, relative to the omitted year of 1980.
The same pattern emerges from the raw takeup rates from the
PSID sample, which show a steady increase from 83 percent in
1980 to 86 percent in 1981 and 87 percent in 1982.

We initially speculated that the differences between the aggre-
gate and micro-level data over this period were due to changes over
time in the fraction of weeks with UI benefits in a typical
unemployment spell. As a rough check on this hypothesis, we
reestimated takeup rates for the PSID spells data on a weeks-of-
unemployment basis. For each eligible spell in the sample we
estimated the weeks of unemployment generated by the spell and
the weeks of unemployment benefits received during the spell (if

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37. The probability value for exclusion of the regional controls is 0.018.
Unfortunately, it was not feasible to include a full set of state fixed effects in the
model because of the limited sample size. A significant number of states have very
few spells of unemployment, and attempts to includes state fixed effects lead to
poorly determined coefficients for most of the explanatory variables.
    38. The probability value for the exclusion of these occupational and industry
variables was 0.0003.

any). Using this procedure, we obtained estimated takeup rates of 78 percent, 86 percent, and 87 percent for 1980, 1981, and 1982, respectively. While these findings are based on crude imputations, they suggest that takeup rates on a weeks-of-unemployment basis also increased between 1980 and 1982 in the PSID sample. The discrepancy between the pattern of estimated takeup rates in the aggregate and PSID samples is therefore unresolved.

An alternative explanation for the failure of the PSID data to show a decline in takeup rates is the selective composition of the sample. In comparison with the CPS random sample of all unemployed workers, the PSID sample consists of unemployed household heads. The PSID sample is older and more predominantly male than the aggregate CPS sample. Although we lack exactly comparable data, unemployed workers in the PSID sample also appear to be more highly unionized and to have worked more hours on their previous job. In other words, the PSID sample overrepresents primary workers with high labor market attachment—a group that may be less likely to reduce their takeup rates, particularly during a difficult recession period. It is possible that the decrease in takeup rates observed in the CPS occurred primarily among nonhousehold heads, a group not included in the PSID sample. While we can estimate eligibility among heads and non-heads in the CPS micro-data, the state-level data on actual UI recipiency cannot be broken down between these two groups. Thus, we cannot directly test this hypothesis by estimating separate takeup rates for the two groups.

To summarize, four conclusions emerge from our analysis of eligibility and receipts of UI benefits in the PSID sample. First, a comparison of self-reported eligibility with estimated eligibility using the methodology that we applied in our aggregate analysis suggests that our eligibility calculations are reasonably accurate. Second, average state-level takeup rates from the PSID sample, based on self-reported eligibility, are highly correlated with our estimated aggregate-level takeup rates. This reaffirms the validity of our eligibility calculations, and also underscores the importance of considering state-level variation when investigating changes in national takeup rates. Third, the micro-data analysis suggests that takeup behavior is strongly correlated with various individual characteristics, including age, sex, family size, and length of unemployment spell. Some of these variables are unavailable or are poorly measured in our aggregate analysis. Nevertheless, it is unlikely that their omission can explain the abrupt drop in takeup

rates between 1980 and 1982, since most of these characteristics change slowly over time, if at all. Finally, the PSID data show no evidence of a decline in takeup rates in the 1980–1982 period. Rather, the raw takeup rates from the PSID sample show an increase in takeup rates. We hypothesize that this is due to the selective sample of unemployment spells in the PSID.

VI. Summary and Conclusions

In this paper we have presented and analyzed measures of the fraction of unemployed workers covered by unemployment insurance benefits. We find no evidence that recent declines in the fraction of workers who receive regular UI benefits are due to changes in eligibility. In fact, our estimates suggest that the same fraction of unemployed workers was eligible for benefits in 1987 as in 1977. This finding leads us to focus on the determinants of takeup rates for UI benefits among eligible workers. Nationwide, estimated takeup rates fell sharply between 1980 and 1982, from an average of 75 percent in the 1977–1980 period to an average of 67 percent after 1982. The same pattern appears in a subset of eight larger states for which takeup rates can be calculated from 1968 to 1987. Data for these states suggest that takeup rates were relatively constant between 1968 and 1979 before an abrupt decline in 1980–1982.

Approximately one half of the nationwide decline in average takeup rates for UI benefits appears to be due to changes in the regional distribution of unemployment. States in the Northeast have significantly higher takeup rates for benefits than states in the South. These regional differences are confirmed by micro-data on benefit eligibility and receipt from the PSID. As the distribution of unemployment has shifted over the past decade from Northeastern to Southern states, national takeup rates have fallen accordingly.

The balance of the nationwide decline in takeup rates is due to changes over time within states. About half of this appears due to the decline in unionization over the past decade. However, even after controlling for other variables, about one quarter of the decline in takeup is still unexplained.

Variables that are associated with higher state takeup rates, both within states and across time, include higher unionization rates, higher benefit replacement rates, and higher coverage rates of the UI system. In addition, greater state enforcement of UI

regulations, as measured by disqualification rates, seems to have a discouraging effect on participation. Disqualification rates, however, have actually fallen over the past decade, and therefore cannot explain the aggregate decline in takeup. In our micro-data analysis of UI takeup, individual and family characteristics have a significant effect on individual takeup rates; younger and nonunionized workers, females, and workers with smaller families and shorter spells of unemployment are less likely to apply for available UI benefits.

Several questions deserve further research attention. The wide historical variation in UI participation across different states remains largely unexplained, as does a substantial part of the decline in takeup rates in the early 1980s. It is also unclear why the limited micro-data sample from the PSID does not show the same time trend as the CPS. A full understanding of why individuals decide to receive or not receive unemployment benefits probably awaits the availability of a micro-data survey explicitly designed to investigate UI eligibility, participation, and reasons for nonparticipation among unemployed workers.

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