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**Interstate Differences in Insured Unemployment:
Some Recent Evidence**

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Abstract: Recent panel data is used to reconsider the determinants of interstate differences in the ratio of insured to total unemployment. We conclude that previous research on the influence of replacement rates, duration of jobless spells and female labor force participation is robust, but find that political affiliations and attitudes could be more important, and unionization rates less important, than once believed.

Keywords: insured unemployment, work norms, replacement rates, unionization

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1. Introduction[*]

The behavior of what is sometimes called the "recipient rate" *RECIP*, the ratio of the rates of insured to total unemployment, is one of the most remarkable but least studied features of American labor markets.[1] With the notable exceptions of Burtless (1983), Blank and Card (1991), Anderson and Meyer (1997) and Rejda (1999), the decline in *RECIP* since the 1960s has received little published attention. For the United States as a whole, the value of *RECIP* peaked in the late 1940s, at close to 100 percent, and reached its nadir of between 40 and 50 percent in the mid-1980s, with sharp declines over the first halves of the 1960s and 1980s, before levelling off around 50 percent.

The absence of comprehensive data on the number of workers who are either eligible for and/or collect UI has proven a serious obstacle for researchers. In an influential paper on the "fraction of insured unemployment" or *FIU*,[2] Blank and Card (1991) infer the number of eligible workers from CPS data and state UI laws for the critical period between 1977 and 1987. Of the three possible explanations for the substantial decline in FIU over this time – changes in state UI laws, changes in the "eligibility determining characteristics" of workers, and changes in "takeup" or collection rates – the last of these is identified as the most important. In particular, their estimated collection rate series exhibits a sharp decline between 1980 and 1982, behavior that Anderson and Meyer (1997) have since attributed to the inclusion of UI benefits in the tax base, which were phased in between 1979 and 1987.

It is not just the behavior of state and national *RECIP* values over time that calls for attention, however. As Table 1 demonstrates, there are also substantial and persistent

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[1] The word "recipient" can be misleading, however, inasmuch as the insured unemployment rate is the ratio of UI *claims* to *covered* employment. Some claims (25 percent, more or less) are disqualified, of course, and about 10 percent of all workers are not covered.

[2] The *FIU* is defined to be the fraction of jobless workers who *collect* regular UI benefits. The decline in *RECIP* has been sharper than that in *FIU*, which Blank and Card (1991: 1161) attribute first and foremost to the increase in the coverage rate over their sample period. The *FIU* is perhaps the more "intuitive" measure, but it is variations in *RECIP* that trigger extended benefit programs.

disparities *across* states, from more than 85 percent (Rhode Island, in 1996) to around 20 percent (South Dakota, in 1994). These differences are mirrored in Blank and Card's (1991) estimated collection rates for 1987, which range from 85 percent for New England as a whole, to less than 50 percent in both the Mountain and West South Central census regions.

[Insert Table 1 Here]

There is, furthermore, an important connection between these interstate differences and the behavior of the national rate over time: some of the decline in the latter has been attributed to the migration of workers from high *RECIP* areas (New England, for example) to low *RECIP* states in the South and West, and to slow(er) labor force growth in the former.

In the penultimate section of their paper, Blank and Card (1991) estimate a cross-sectional model of collection rates, and we undertake a similar exercise here: a simple and distinct model of RECIP differences across states is estimated using pooled annual data over a small, but more recent, sample period, from 1992 to 1996.

2. Data and Model

Our dependent variable is the aforementioned $RECIP_{i,t}$ the recipient rate for state *i* in period *t*, the ratio of the insured unemployment rate, as reported at the ITSC Employment Insurance web site (*www.itsc.state.md.us*), to the overall unemployment rate, available in the *Statistical Abstract of the United States* (hereafter, *SAUS*).

The first of our independent variables is the replacement rate $REPLACE_{i,t}$, the ratio of the average benefit to the average wage in the covered sector. The coefficient is expected to be positive, of course: the more generous the UI program relative to local economic conditions, the greater the claimant rate, other things equal. The second is $DURAT_{i,t}$, the mean duration in weeks of jobless spells, the effects of which are uncertain: if job losers in state *i* have reason to believe more time will be required to find another "match" than those in state *j*, the proportion who submit claims should rise, but as mean duration rises, the proportion of those without work who will exhausted their UI benefits rises, too. Common sense suggests, however, that the first effect will dominate the second in practice, in which case the coefficient should be positive. From a somewhat different perspective, it is well known that state and national *RECIP* rates tend to be countercyclical, and to the extent that duration is an alternative measure of labor market slack, this behavior could manifest itself in our cross-sections. The values of both *REPLACE* and *DURAT* are those reported in the ITSC database.

The third of our "economic" determinants is $UNION_{i,,t}$, the proportion of the workforce in state *i* that was unionized in period *t*, as reported in various issues of *SAUS*. If, as Blank and Card (1991: 1174) find, unions "increase their members' awareness of UI programs or expedite their applications," this coefficient should also be positive and significant.

We wanted to separate unionization effects, however, from possible "industrial" and "urban" influences, and to this end, included both $MANUF_{i,t}$, the proportion of the labor force engaged in manufacturing, and $METROP_{i,t}$, the proportion of the population in metropolitan areas, as regressors, as reported in various issues of Morgan et al's *State Rankings*. On the basis of previous research, we expected a positive, if not significant, coefficient on *MANUF*, but were ambivalent about *METROP*: it seemed to us that the relevant "neighborhood effects" could cut in either direction.

We also consider two other demographic variables, $POV_{i,t}$, the proportion of the population in state *i* classified as poor in period *t*, and $FEMALE_{i,t}$, the female labor force participation rate, as reported in various issues of *State Rankings*. There are at least two reasons to introduce the former, even if these pull *RECIP* in different directions. On the one hand, one would expect that the poorer a state's population and labor force, the higher its claimant and collection rates. On the other, to the extent higher *POV* values reflect the increased presence of workers who have exhausted their UI benefits or differential access to public services, the predicted *RECIP* values could well be smaller. The correlation of POV with other variables we use – in particular, *FEMALE*, *MANUF* and *METROP* – further complicates estimation of its direct influence. The expected sign of the *FEMALE* coefficient is also uncertain. If interstate differences in female labor force participation reflect differences in the proportion of households with multiple wage-earners, and if "second," often female, participant submits fewer UI claims and is disqualified more often, the coefficient should be negative. It will be positive, however, if the differences in *FEMALE* correspond to differences in the proportion of single female-headed households, or if increased participation is associated with more durable labor force attachment. All of this said, previous empirical research suggests that the sign tends to be positive.

Last, Blank and Card (1991) were perhaps surprised to find that their measure of the "political climate" in each state, the fraction of its Congressional delegation that was Democratic, was insignificant. We consider a pair of alternative measures: $GOV_{i,t}$, equal to "1" if state i had a Democratic governor at the start of period *t*, and "0" otherwise; and $RTW_{I,t}$, equal to "1" if *i* was a "right to work" state in period *t*, and "0" otherwise. The rationale for the former is that UI is, and is perceived to be, an ensemble of state programs with federal minimum standards, in which case the relevant political affiliations are perhaps local, not national. The second is premised on the notion that traditional political affiliations will sometimes fail to capture durable social and/or political attitudes vis-à-vis labor markets. In this context, the residents of right to work states might be supposed to hold more "non-interventionist" views. If so, the coefficient on the former should be positive and negative on the latter.

In the absence of fixed effects, then, our benchmark model is:

$$\begin{aligned} RIU_{i,t} = \beta_0 + \beta_1 REPLACE_{i,t} + \beta_2 DURAT_{i,t} + \beta_3 UNION_{i,t} + \beta_4 MANUF_{i,t} \\ \beta_5 METROP_{i,t} + \beta_6 POV_{i,t} + \beta_7 FEMALE_{i,t} + \beta_8 GOV_{i,t} + \beta_9 RTW_{i,t} + e_{i,t} \end{aligned}$$

where $e_{I,t}$ is an error term that is almost certain to be heteroskedastic. For this reason, GLS with cross-section weights are used below. Furthermore, we allowed for period effects, measured relative to 1992, in all of the models we estimated, and considered two sorts of cross-sectional fixed effects, state and regional.[3]

3. Estimation Results

Table 2 reports the estimates for four versions of the model, one without fixed effects, one with regional effects, and two with state effects. We note first that despite the substantial difference in goodness of fit, the introduction of regional effects does not much influence our the estimates of "core" parameters: the coefficients on duration,

[3] The nine census regions are New England, Middle Atlantic, East North Central, West North Central, South Atlantic, East South Central, West South Central, Mountain and Pacific.

unionization, female labor force participation and labor in manufacture are all positive and significant, both in the economic and statistical senses, and the coefficient on metropolitan population is negative and significant. Furthermore, the coefficients on the replacement rate and governorship have the "correct" signs but appear to be insignificant. The signs on two coefficients "flip" between the two models, however: the estimated coefficient on the poverty rate switches, from negative and significant to positive and insignificant, while that on right to work status moves from negative and (almost) significant, to positive and insignificant. The last of these comes as no surprise, however, inasmuch as right to work status exhibit a pronounced regional concentration.[4]

[Insert Table 2 Here]

The substitution of state for regional effects does exert a substantial effect on our results, however. The coefficient on duration remains positive and significant, but much smaller in absolute terms: a one week difference in mean duration is now predicted to push *RECIP* upward about one third of a percentage point. Consistent with intuition, the coefficient on the replacement rate is now positive, significant and much larger (0.62) than before, which implies that the generosity (relative to local economic conditions) of state UI programs does have an important influence the behavior of potential claimants. Surprisingly, perhaps, the estimated influence of unionization is small in both the economic and statistical senses, a result that contradicts Blank and Card's (1991) evidence on collection rates: a one percent difference in unionization rates between states is associated with no more than a 0.21 percent difference in recipient rates.

The coefficients on the poverty rate and proportion of the labor force engaged in manufacture are significant but *negative* in the extended fixed effects model, both provocative results: it is unfortunate, but perhaps not surprising, that holding the replacement rate fixed, jobless workers in poorer states still make fewer UI claims, but the inference that a substantial manufacturing presence exerts a similar effect is unexpected, and more difficult to rationalize. The coefficient on female labor force participation has the "right" sign, on the other hand, but is now small (0.10) and

[4] Three of the four states in the West South Central region (Arkansas, Louisiana and Texas) are right to work, for example, and the current governor of the fourth (Oklahoma) has advocated it. Likewise, three out of four states in the East South Central region, five of eight in each of the Mountain and South Atlantic, and four of eight in the West North Central have passed right to work legislation. There are no right to work states, on the other hand, in the Middle Atlantic, New England, East North Central and Pacific regions, on the other hand.

insignificant. The influence of a metropolitan population is also estimated to be small and negative.

Last, the two "political" variables exert substantial, if not quite significant, effects now: other things equal, the value of *RECIP* is estimated to be a full percentage point lower in right to work states, and three quarters of a point higher in states with Democratic governors. At the least, this suggests that Blank and Card's (1991) conclusions concerning the role of political affiliation merit reconsideration.

The truncated model includes fixed state and period effects, but omits the three "least significant" variables, *UNION*, *FEMALE* and *GOV* . None of the remaining coefficients, except for that on RTW, which increases from –1.00 to –1.42, is much affected.

4. Conclusion

Our results suggest that previous work on the effects of interstate differences in replacement rates, duration of joblessness and female labor force participation on UI claimant behavior is robust with respect to sample period and model specification. We also find, however, that the roles of unions and political attitudes are perhaps less so, and that the effects of a "poor" labor force require further attention.

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Table 1. Descriptive Statistics 1992-1996

	Minimum	Maximum	Mean	Standard Deviation
Recipient Rate	21.21 (NH, 93)	86.27 (RI, 96)	40.41	12.29
Unemployment Duration (Weeks)	8.40 (NC, 95)	27.30 (MA, 92)	15.83	4.31
Replacement Rate	25.10 (CA, 96)	52.80 (HI, 95)	36.85	5.03
Unionization Rate	5.00 (NH, 95)	31.90 (HI, 92)	15.93	5.97
Poverty Rate	5.30 (NH, 95)	26.40 (LA, 93)	13.42	4.07
Female Labor Force Participation	43.70 (WV, 92)	69.80 (MN, 94)	60.18	4.62
Labor Force in Manufacturing	3.13 (HI, 96)	26.58 (NC, 92)	15.37	5.75
Metropolitan Population	23.5 (MT, 96)	100.0 (NJ, 96)	66.98	21.31

Table 2. Regression Analysis of State Ratios of Insured Unemployment, 1992-1996

(Dependent Variable: $RECIP_{i,t}$ x 100, t statistics in parantheses)

	Model 1	Model 2	Model 3	Model 4
Fixed Effects?	No	Yes	Yes	Yes
		Regions	States	States
Duration (Weeks)	1.20	0.67	0.38	0.34
	(6.69)	(3.97)	(4.63)	(3.84)
Replacement (x 100)	0.03	0.07	0.62	0.60
	(0.38)	(0.65)	(3.84)	(4.19)
Union Rate (x 100)	0.88	0.83	0.21	
	(9.71)	(6.95)	(1.03)	
Poverty Rate (x 100)	-0.39	0.04	-0.26	-0.27
	(-2.84)	(0.28)	(-2.13)	(-2.36)
Female LPR (x 100)	0.36	0.96	0.10	
	(2.95)	(6.28)	(0.57)	
Manufacture (x 100)	0.34	0.48	-1.57	-1.50
	(5.12)	(4.04)	(-3.73)	(-3.69)
Metropolitan (x 100)	-0.15	-0.14	-0.12	
	(-7.01)	(-5.45)	(-0.45)	
Right to Work (1=Yes)	-1.54	0.61	-1.00	-1.42
	(-1.62)	(0.60)	(-1.92)	(-2.52)
Governor (1=Dem)	0.92	0.53	0.70	
	(1.20)	(0.73)	(1.53)	
R^2	0.45	0.59	0.91	0.91