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The effect of unemployment insurance on unemployment rate and average duration: evidence from pooled cross-sectional time-series data

PHANINDRA V. WUNNAVA and SYED ALI RAZA MEHDI‡

Department of Economics, Middlebury College, VT, 05753, USA
‡CS First Boston, Park Ave Plaza, 55 East 52nd Street, New York, NY 10055, USA

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This paper empirically analyses the impact of the unemployment insurance system upon the insured unemployment rate and the average duration of unemployment. It employs a simultaneous equation framework because of possible feedback effects between the insured unemployment rate and the average duration of unemployment. Based on a pooled cross-sectional time-series model (covering all the 50 states in the USA for the years 1967–88) that corrects for heteroscedasticity and autocorrelation, and the results show some support for the hypothesis that the unemployment insurance system, by providing workers with a safety net, increases both the insured unemployment rate and the duration period.

I. BACKGROUND

The impact of unemployment insurance on the rate and duration of unemployment has been a controversial one. According to many critics (for example Feldstein and Poterba 1984), unemployment insurance adversely affects unemployment rates and duration by providing workers with a safety net. This acts as a disincentive for workers to be more active in their job search, and so the programme essentially negates its own purpose.

Past studies[1] have analysed the impact of unemployment insurance on labour markets in four different ways. These include looking at the impact of unemployment insurance on the reservation wage-levels, and thus on the average duration of unemployment (Feldstein and Poterba, 1984); the effect of the benefit–income ratio (replacement ratio) on the unemployment rate and the average duration (Chapin, 1971; Maki and Spindler, 1975; Cubin and Foley, 1977; Nickell, 1979; Wunnava and Henley, 1987); the effect of unemployment insurance on the quit rate (Hamermesh, 1979); and the differential impact of unemployment insurance on industrial sectors (Deere, 1991). Some of the studies, such as that of Maki and Spindler (1975) have explored the possibility of potential feedback effects between the replacement ratio and the unemployment rate. The rationale for doing this is that during periods of high unemployment, incomes may be lower which would cause the replacement ratio to be higher, allowing for possible simultaneity between the unemployment rate and replacement ratio. In a similar framework, this paper proposes that the duration and rate of unemployment are jointly determined and thus that substantial feedback effects may exist between the two. The logistics of the proposed simultaneity are explored in the next section.

II. THE SIMULTANEITY HYPOTHESIS

The potential benefits of using such an approach are implicit in the studies done by Chapin (1971), Maki and Spindler (1975), Nickell (1979), Poterba and Feldstein (1984) and Wunnava and Henley (1987). Feldstein and Poterba reported that unemployment compensation raised an individual's reservation wage by providing a safety cushion, and so increased the length of time spent searching for a job, thus increasing the average duration of unemployment. The others have reported the positive correlation

[1]For an excellent review of the literature see Atkinson and Micklewright (1991)

between unemployment compensation and the unemployment rate. So unemployment insurance affects both the insured unemployment rate and the average duration of unemployment; but nobody has tried to determine whether these two have reciprocal effects on each other. An increase in the duration period can cause an increase in the insured unemployment rate if workers with compensation prefer to remain unemployed until they find a job which pays a wage commensurate with their reservation wage, a process that drives up the insured unemployment rate. Conversely, an increase in the insured unemployment rate can affect the average duration of unemployment by changing people's expectations about their probability of finding a job. If these reverse causations are substantial, then empirical estimates ignoring such feedbacks[2] will not even be consistent. Hence in this paper a simultaneous equation framework is employed to account for possible simultaneity between the insured unemployment rate and average duration.

III. THE EMPIRICAL MODEL AND RESULTS

The paper uses a pooled time-series cross-sectional data set for our empirical analysis. The time-series units are years beginning from 1967 until 1988. The cross-sectional observations are all the 50 states (excluding the District of Columbia) in the United States. All data are from publications of the federal government or directly from its agencies. These include the *Social Security Bulletin: Annual Statistical Supplements* and *Statistical Abstracts of the United States*.

The two endogenously determined variables are the insured unemployment rate (*UR*) and the average duration of unemployment (*DUR*). The insured unemployment rate is measured as a percentage, while observations on the average duration of unemployment are in weeks. The following is the system of simultaneous equations:

$$UR_{it}=\beta_0 + \beta_1 DUR_{it} + \beta_2 RR_{it} + \beta_3 PERMANU_{it} + \beta_4 UNION_{it} + \beta_5 LNPINC_{it} + \varepsilon_{it} \quad (1)$$

$$DUR_{it} = \alpha_0 + \alpha_1 UR_{it} + \alpha_2 RR_{it} + \alpha_3 PERMANU_{it} + \alpha_4 UNION_{it} + \alpha_5 EXER_{it} + \gamma_{it} \quad (2)$$

where $i = 1,2, \ldots 50$ (all the states in the USA) and $t = 1967, 1968, \ldots 1988$ (years).

One of the explanatory variables in this model is the level of unemployment insurance (*UI*) compensation. The most appropriate measure is the replacement ratio (*RR*), which is a direct measure of how much of the state's average weekly wage is replaced by benefits under *UI* compensation. The *RR* is normally expected to be positively correlated with both the average duration of unemployment and the level of insured unemployment. The percentage of non-agricultural workers in the manufacturing sector (*PERMANU*) is included to account for differences in industrial composition. States with more workers in the manufacturing sector might have differing average durations and insured unemployment rates due to varying labour–market conditions. Union membership (*UNION*) is taken into account to test whether states with high union membership have higher unemployment rates and longer average duration of unemployment due to the unwillingness of union members to accept non-union jobs, where wage and benefit levels are generally lower.

A fourth explanatory variable in the model is the natural log of real per capita income (*LNPINC*) which would conceivably be negatively correlated with the insured unemployment rate. The last exogenous variable is the percentage of UI claimants who exhaust their benefits (*EXER*). High exhaustion rates could result from both lax enforcement rules and/or poor employment service which does not encourage *UI* claimants to get off the benefit rolls and back to work. Either of these reasons would cause longer average durations of unemployment.

Since this is a pooled cross-section time-series data set, it obviously suffers from the dual problems of heteroscedasticity and autocorrelation. To try and correct for these problems the method[3] specified by Kmenta (1986) is employed. Consistent

[2]To test for possible simultaneity between average duration and insured unemployment rate, a two step procedure of the Hausman (1978) specification test proposed by Spencer and Berk (1981) is performed. In the second stage, the residuals from respective first stage reduced form equations are added as an extra regressor (*R*) to the original structural Equations 1 and 2. The variable *R* is indeed statistically significant in both cases, yielding *t*-values of −6.9014 and −3.3080 respectively. This overwhelmingly supports the theoretical framework of the proposed methodology (i.e. simultaneity between average duration and insured unemployment rate). To conserve space full regression results of the Hausman test are not included in the paper but can be obtained upon a request. Please note that a non-technical setup of this variation of the Hausman test is given in Pindyck and Rubinfeld (1991) pp. 303–5.

[3]This technique, by subjecting the observations to two transformations, one designed to remove autocorrelation and the other to remove heteroscedasticity, comes up with a disturbance term (ε_{it}) that is asymptotically non-autoregressive and homoscedastic. To find consistent estimates, OLS is applied to obtain the regression residuals and then these are used to perform transformations so that the error term is asymptotically non-autoregressive and homoscedastic (for details see Kmenta (1986), pp.618–22). The particular characteristics of this model are as follows:

$E(\varepsilon_{it}^2) = \sigma_i^2$ (heteroscedasticity)

$E(\varepsilon_{it}\,\varepsilon_{jt}) = 0$ (($i \neq j$) — cross-sectional independence)

$\varepsilon_{it} = \rho\varepsilon_{i,t-1} + u_{it}$ (autocorrelation assuming that 'ρ' has the same value for all cross-sectional units and u_{it} is the classical error)

where

$u_{it} \sim N(0, \sigma^2_{ui})$

$\varepsilon_{it} \sim N(0, [\sigma^2_{ui}/1-\rho^2])$

and $E_{(\varepsilon_{i,t-1}\, u_{jt})} = 0$ for all i,j

Table 1. *Second stage pooled regression results*[a]

(A) Dependent variable: *UR*

Variable name	Estimated coefficient	*t*-ratio 1144 DF	Elasticity at means
*DUR***	0.23043	10.590	0.23341
RR	0.21971E–01	2.5368	0.24693
PERMANU	–0.19833E–01	–2.4812	–0.13157
UNION	0.98130E–02	2.8611	0.61432E–01
LNPINC	0.37563E–03	0.50741	0.10973E–02
CONSTANT	1.3584	3.4402	–

R^2 *** = 0.168 Sample size = 1150

(B) Dependent variable: *DUR*

Variable name	Estimated coefficient	*t*-ratio 1144 DF	Elasticity at means
*UR***	–0.25689	–4.0301	–0.18908E–01
RR	0.20527E–01	1.4071	0.57903E–01
UNION	0.16295E–01	3.2611	0.25602E–01
PERMANU	–0.62619E–01	–5.9533	–0.10425
EXER	0.17649	22.684	0.38344
CONSTANT	8.4564	13.350	–

R^2*** = 0.7096 Sample size = 1150

[a]Variable definitions:
UR = insured unemployment rate based on average covered employment in 12-month period by state
DUR = average actual duration (in weeks) by state
RR = the percentage of average weekly wage replaced by average weekly benefit for total unemployment by state
PERMANU = the percentage of non-agricultural labour force in the manufacturing sector by state
UNION = percentage of workers employed in unions (in manufacturing) by state
LNPINC = the natural log of per capita real income (in 1982 dollars) by state
EXER = the percentage of claimants exhausting benefits by state
**Predicted values obtained from the first stage regression (i.e. reduced form equation) run–these regression results can be obtained upon a request
***Computed between the observed and predicted values of the dependent variable.

estimates for Equations 1 and 2, after adjusting for possible simultaneity between insured unemployment (*UR*) and average duration (*DUR*), are reported in Table 1, panels A and B. The first model was run with insured unemployment rate (*UR*) as the dependent variable. All the coefficients, except one, were significant at the 5% level. As hypothesized, an increase in average duration (*DUR*) was positively correlated with the insured unemployment rate, a 10% increase in the predicted value of duration leading to a 2.3% increase in the insured unemployment rate. The disincentive effect of the *UI* scheme as proxied by the replacement ratio *(RR)* variable seemed to be borne out by the results since, holding all other variables constant, a 10% increase in it led to a 0.2% increase in the insured unemployment rate. The result was more striking when we computed the elasticity at the mean for the variable, in which case a 10% increase in *RR* led to a 2.4% increase in the insured unemployment rate. *Ceteris paribus*, an increase in the percentage of non-agricultural labour force employed in manufacturing (*PERMANU*) actually reduced the insured unemployment rate. At the mean value, a 10% increase in *PERMANU* led to 1% decrease in the insured unemployment rate. This result contradicted previous evidence (Wunnava and Henley, 1987) that had suggested that, due to depressed labour-market conditions in the manufacturing sector, states with a higher percentage of the work force employed in manufacturing would have higher unemployment rates. Part of the reason for this might be that we are using more recent data which would incorporate the effects of the revival in the American manufacturing sector that has occurred over the past decade. As expected, those states that had a higher percentage of their labour force in unions (*UNION*) also had higher insured unemployment rates. This is not surprising since unionized labour is most resistant to wage cuts and other such measures. The only variable that was not significant in the equation was the natural log of real per capita income *(LNPINC)*.

In the second model again all coefficients, excepting one, were significant at the 5% level. The insured unemployment (*UR*) was negatively correlated with the average duration of unemployment (*DUR*), with a 10% increase in the former leading to a 0.2% decrease in the latter at the mean value. What this implies is that the increase in the unemployment rate compels workers to increase their efforts to find a job, thus, decreasing the average duration. The sign of the replacement ratio (*RR*) variable is positive as expected, but it is significant only at the 20% level. Thus the replacement ratio seems to create a dual disincentive; it increases the unemployment rate and it also leads to higher duration periods. This finding seems to be supported by the behaviour of another independent variable, the percentage of claimants exhausting their benefits (*EXER*). This has a very significant and positive impact on the duration period, a 10% increase in it leading to a 1.7% increase in the duration period (4% if we use the elasticity computed at the mean value). What this shows is that states which have lax enforcement rules and a poor employment service that does not encourage UI claimants to get off the benefit rolls[4] and back to work find a significant increase in the unemployment period. The union variable (*UNION*) is also positive and significant, lending support to the belief that union workers who get higher wages, and consequently have higher reservation wages, prefer to remain unemployed for longer periods and wait until they find a job that pays a wage rate commensurate with their reservation wage. The unemployment insurance, in effect, acts as a safety net and subsidizes their wait period. Last, the percentage of non-agricultural labour force in the manufacturing sector (*PERMANU*) variable is negatively correlated with the duration period, suggesting that states with relatively larger manufacturing sectors would see shorter periods of unemployment. At the mean value, a 10% increase in *PERMANU* led to a 1% decline in the average duration of unemployment. This is supported by the findings of the first model where the variable was negatively correlated with the insured unemployment (*UR*) rate.

IV. SUMMARY AND CONCLUSION

This paper examined the impacts of the unemployment insurance system on the insured unemployment rates and the average duration of this unemployment period. Results, based on the pooled state level data in the USA for the 1967–88 period, and after adjusting for possible simultaneity between insured unemployment rate (*UR*) and the average duration of unemployment (*DUR*) (as the Hausman (1978) specification test strongly supports such a hypothesis) indicate some evidence of a disincentive effect created by the system. The replacement ratio (i.e. ratio of UI benefits to wages) increases both the rate and the duration of unemployment. This conclusion is consonant with a number of empirical studies surveyed by Burtless (1990) who reports that UI has a positive and statistically significant influence on the unemployment spells of insured workers, although he cautions that evidence exists for both the positive and negative effects of UI and that it is very difficult to make a definitive statement. Second, states which have lax enforcement rules, a higher percentage of their labour force in unions, and a smaller employment base in the manufacturing sector are likely to see longer spells of unemployment. In the last two instances the insured unemployment rate would also increase. In conclusion, this paper finds support for the Feldstein–Poterba (1984) hypothesis that the unemployment insurance scheme has set in motion a disincentive system that negates the scheme's original intention.

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