"REFLECTIONS ON 'WAIT UNEMPLOYMENT' IN EUROPE, II"

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N° 88 / 48

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Printed at INSEAD,
Fontainebleau, France

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April 1988
Revised September 1988

This paper is a revised and expanded version of an earlier paper presented at the London Economic Policy Meetings, April 1988, and has benefitted from comments of David Begg, Richard Layard, John Martin and Charles Wyplosz. I am also grateful to Mme Françoise Coré of the OECD for supplying some essential data and Katrina Maxwell for valuable research assistance.

Abstract

The sustained increase in unemployment rates in Western Europe has provided macroeconomists with an unusual opportunity to test competing theories of unemployed resources and to provide unifying explanations that transcend national boundaries. We find that much of the cross-country variation of long-term unemployment and unfinished spell duration can be predicted by a simple Harris-Todaro model of 'wait unemployment' with an explicit role for unemployment benefits. The model can also account for slower growth of service sector employment in Europe over the past fifteen years. The dynamic response of unemployment compensation to economic conditions may be an important clue explaining in the increase in unemployment rates since the mid-1970s.

The sustained increase in unemployment rates in Western Europe has provided macroeconomists with an unusual opportunity to test competing theories of unemployed resources and to provide a unifying explanation that transcends national boundaries. Yet there appears to be as much confusion now as there was a decade ago, when it became evident that jobless rates were not likely to recover levels of the early 1970s.

All attempts to account for the dramatic rise in unemployment have fallen short. The 'European depression' view that only deficient aggregate demand is to blame has lost considerable credibility after more than a half decade of fiscal conservatism. While Germany and the United Kingdom have pursued restrictive fiscal policy paths, so have Sweden, Finland, Denmark and Norway, with declines in unemployment over the same period.[1] Similarly, the limited employment response to recent declines in product wages relative to productivity has cast doubt on the disequilibrium real wage hypothesis, which maintains that current real wage levels are inconsistent with the elimination of involuntary unemployment. Similarly, a lack of convincing follow-

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[1] In the period 1982-86, the government deficit in Sweden as a fraction of GDP was reduced from 6.3% to 0.3%; in Norway, a surplus of 4.4% of GDP was increased to 5.9%; in Finland, a deficit of 0.4% has been transformed into a surplus of 0.6%. The most astonishing case has been Denmark, which in four years moved from a deficit of 9.1% of GDP to a surplus of 3.4%.

up evidence has discredited strict "hysteresis" accounts of insiders setting wages that effectively exclude outsiders (Coe 1986; OECD 1986b, 1987; Alogoskoufis and Manning, 1987). In general, one can question the premise that the insensitivity of union wages to labor market conditions necessarily raise the natural rate of unemployment, since unemployed workers can presumably always seek employment in the nonunion sector.

An intriguing alternative has been suggested recently by Summers (1986) in a reexamination of issues first raised by Hall (1970). In assessing the modest rise in average unemployment rates in the United States, he finds that much of the increase can be attributable to job losers in high wage sectors such as manufacturing, mining and resource extraction, and construction. While total employment growth is largely orthogonal to changes in unemployment rates for the cross-section of fifty states, he finds a strong negative association between growth in high wage jobs and changes in unemployment. He concludes:

> "These findings on the relationship between changes in employment and unemployment are instructive. They suggest that in analyzing recent changes in unemployment in the United States, it is not enough to focus on the determination of the total level of employment. It is also necessary to examine the composition of employment growth and to consider the incentives individuals may have to remain unemployed." (p.370)

Summers' findings are suggestive of a segmented labor market in which many unemployed choose the strategy of queueing for high paying jobs rather than taking less attractive work in the

secondary sector. This conceptualization of unemployment has its origins in the seminal work of Harris and Todaro (1970), with variations by, for example, Calvo (1978) and McDonald and Solow (1985). Segmented labor market models are currently enjoying a rennaissance in both theoretical and empirical applications (see Dickens and Lang 1988).

While corresponding closely to casual inference of both local and foreign lay observers, these ideas have found little academic interest in the European context.[2] Summers himself dismisses the relevance of "wait unemployment" for Europe, arguing that increasing unemployment there has been associated with a cessation of job creation, distinguishing it sharply from the United States. This judgment may be premature. Zero total employment growth in Europe has masked significant and opposite trends in goods and service producing sectors. While employment in OECD Europe's industrial sectors declined by 10.9% from 1979 to 1986 (14.0% in EEC), employment in services --defined as trade, transport, communications, finance and personal services-- rose by 10.3% (11.2% in the EEC).[3] Second, most of the long-term unemployed are job losers in high paid industrial employment, with higher than average previous experience. Finally, there is

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[2]See, for example, the Wall Street Journal Europe Feb 5, 1988, p.1.

[3]Source: OECD Labor Force Statistics 1965-85, and more recent quarterly reports.

evidence that European labor force growth is more highly correlated with growth in total-- not industrial-- employment.[4]

In this paper we take the position that "wait unemployment," which has both involuntary and voluntary aspects, can make a contribution to explaining both the levels of and the recent increases in normal unemployment that has occurred in some European economies; more importantly, it can explain the absence of this increase in others. Using a variant of Harris and Todaro's (1970) model, we formalize some of the predictions of segmented labor markets for "wait unemployment" when an explict role for unemployment insurance is allowed. The most evident implication is that economies with higher levels of unemployment benefits will have higher average unemployment rates and higher expected duration of unemployment spells. The model predicts that countries with segmented markets and high rates of labor force growth will show more rapid growth of the secondary sector. The model also implies that the correlation between changes in unemployment and the labor force is positively associated with the level of unemployment benefit. By testing these implications on a cross-section of countries, we can better abstract from business cycle influences. Finally the dynamics of unemployment

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[4]Over the period 1975-85, the labor force in OECD Europe grew at an average 0.8% rate (0.7% in EEC), total civilian employment by 0.1% (no growth in EEC), whereas total industrial employment shrank by -1.4% per annum (-1.7% in the EEC).

compensation support the hypothesis that adverse shocks have induced more generous support and thus higher unemployment. The data suggest that cross-country variation in the provision of unemployment insurance may play a central role in explaining diverse labor market outcomes observed in the OECD.

The paper is organized as follows. Section 1 discusses long-term unemployment as a hallmark of the rise in normal unemployment in Europe. In addition, some striking cross-country differences in long term unemployment and average incomplete spell duration are identified. Section 2 sets out the most important predictions of the segmented labor market model for wait unemployment when unemployment benefits are available. In Section 3 we test these predictions on a cross-section of OECD economies heavily weighted towards Europe, and in the following section we discuss and test some alternative explanations of the change in unemployment rates that have occurred in some countries and not others. The final section draws some tentative conclusions.

1. Long-Term Unemployment and Unemployment Duration in Europe

The OECD is often credited with having drawn attention early on to a particularly salient characteristic of high European unemployment, the extent to which it consists of long-term unemployed.[5] Table 1 displays the percentage of unemployed

[5]See the 1983 Employment Outlook and subsequent issues.

individuals with uncompleted spells of twelve months or more, for select years in several OECD economies. It is remarkable that in the boom years of 1973 and 1979, long-term unemployed represented a significant component of total unemployment in Belgium, the Netherlands, Spain, and Ireland. Equally striking is the high correlation over time of these percentages in cross-sections of OECD countries for which data are available: 0.90 between 1979 and 1986 (13 countries) and 0.94 between 1973 and 1979 (7 countries). Not surprisingly, the duration of unemployment exhibits similar variation across economies. Table 2 displays average duration of interrupted unemployment spells in 1973, 1979, and 1982 (recent data are no longer made available by the OECD).[6] While long-term unemployment is in principle consistent with a variety of causes, the wide degree of cross-country dispersion of long-term unemployment and its within country stability shown in Tables 1 and 2 may weaken the case for systematic explanations common to all economies.

There is a strong empirical association between average unemployment rates over the cycle -- corresponding presumably to the natural rate of unemployment-- and average durations of

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[6]It is important to recall Salant's (1977) stricture against hasty conclusions about average duration of completed spells on the basis of interrupted spell data, since shorter spells have less likelihood of being recorded and interrupted spells are by definition incomplete. If the variance of spell length is small relative to the square of the mean length, the sampled average duration of interrupted spells is a good estimate of one half the expected duration of completed spells.

uncompleted unemployment spells. This association is intuitive, since in the steady state, the unemployment rate is a product of a flow and an average duration.[7] The relative constancy of inflow rates thus suggests that the rise in European unemployment reflects primarily a decrease in the escape probability from unemployment, rather than an increase in the rate of job loss. Pissarides (1986) adduces evidence for the United Kingdom that changes in outflows have accounted for nearly all movement in the unemployment rates since the early 1970s. Data from the Bundesanstalt für Arbeit on both quarterly flows into unemployment as well as the number of individuals unemployed for less than one month confirm this phenomenon for the Federal Republic of Germany.[8] In contrast, the unemployment rate in the United States seems driven largely by changes in inflows (Darby, Haltiwanger, and Plant, 1986). This striking difference has been emphasized in a different context by Blanchard and Summers (1986, Table 8) and Flanagan (1987).

In the next section we examine a segmented labor market model which attributes long-term unemployment --thus abstracting from "frictional" components-- to queueing for jobs in the primary

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[7] See Johnson and Layard (1986).

[8] For example, in 1979, the quarterly inflow rate into registered unemployment as a fraction of all employment was 2.7%; in 1983, the rate was 3.7%; in 1986, 3.54%. Over the same period, the unemployment rate rose from 3.8% to 9.0%. Source: Bundesanstalt für Arbeit (1987, p.99) and Sachverständigenrat (1988, p.289).

sector. Such models generate predictions for the relationship between the level and duration of unemployment, conditions in the primary sector and the best alternative available to unemployed workers.

2. A Segmented Labor Market Model

The framework developed here is a hybrid of several approaches to modelling segmented labor markets, including Harris and Todaro (1970), Calvo (1978), and McDonald and Solow (1985). We make several simplifying assumptions that render the model more tractable. We assume the absence of money illusion and risk neutrality on the part of both firms and workers. The latter are in constant total supply, $\bar{L}$, and are equally qualified to work in either the primary or secondary sector. In the primary or industrial sector, a union bargains with an employers' association over wages and employment, following McDonald and Solow's (1981) elaboration of Leontief's (1946) efficient contracts.[9] The secondary sector consists of lower-paid, menial jobs for which the wage is determined by market clearing. Evidence on relative wages in the United States and elsewhere indicate that large segments of

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[9]With the constant elasticity of revenue specification, this could be replaced with a monopoly union, or with bargaining along the labor demand curve à la Carruth and Oswald (1987); none of the results is substantively altered. Although we do not investigate these possibilities, the model might also be consistent with efficiency wages paid in the primary sector, or insider power due to, say, learning curve effects.

the "service sector"-- the broadly defined service producing industries-- possess many of the characteristics of the secondary sector (Bluestone and Harrison 1988). Nevertheless, it should be stressed that some services, such as banking, insurance, and business consulting, are associated with high relative compensation levels.[10]

A key innovation in this paper is an explicit role for unemployment benefits in determining the best available alternative to agents. In the original Harris-Todaro (1970) model, workers compare the option of working in the agricultural sector to that of unemployment and queueing for high-paying industrial jobs. The fallback level of income in the event of not receiving a primary sector job is fixed at zero. The explicit introduction of a variable level of unemployment income, which may be attributable to the drawdown of assets, work in the shadow economy or private assistance as well as unemployment benefits, creates an additional source of variation in predictions of the Harris-Todaro framework.

Let us first describe the outcome of Nash (symmetric) bargaining over the wage and employment in the primary sector. Revenue for the industry is given by $R(\theta,L)$, where L is employment and θ is a shift parameter. We assume $\partial R/\partial\theta>0$, $\partial R/\partial L>0$, and

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[10]This point has been emphasized by US discussants of the "American job machine." See for example, Norwood (1986).

$\partial^2 R/\partial L^2 < 0$. With perfect competition in output markets, R is simply the industry production function multiplied by the market price. The fallback or threat point for the employers' association is revenues covering variable costs (R=wL), and for the union it is b, which is an unemployment benefit plus any monetary value of leisure.[11] With equal weights for both parties, the bargaining process solves

$$\max_{w,L} \ (R(\theta,L)-wL)(w-b)L,$$

where w is the wage. The first order conditions are:

$$w = b + (R-wL)/L \tag{1}$$

$$(w-R')L = R-wL \tag{2}$$

Equation (1) looks suspiciously like profit-sharing, and (2) is the familiar result that the wage in such bargains exceeds the marginal revenue product of labor. These two conditions can be rewritten as

$$R' = b \tag{3}$$

$$w = (b+R/L)/2 \tag{4}$$

In this two-part rule, employment is set to equate the marginal revenue of labor to the worker's best available alternative; the

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[11] This presumes that in case of a collapse of negotiations, the wage in the secondary sector would be bid down to b.

wage paid is the average of b and revenue per worker. A well-known result is that if $R(\theta,L)$ possesses the isoelastic form θL^{α} with $0<\alpha<1$, the bargained wage will be insensitive to shifts in θ. As McDonald and Solow (1985), we also assume this form. It is easy to show that the bargained level of employment and wages are given by $L^m=(b/\alpha\theta)^{1/(\alpha-1)}$ and $w^m=.5(1+\alpha^{-1})b > b$, and that $L^m_b<0$, $L^m_{bb}>0$, $L^m_{\theta}>0$, $L^m_{\theta\theta}>0$, and $L^m_{\theta b}< 0$, where subscripts are shorthand for partial derivatives. The last inequality implies that higher b is associated with less responsiveness of employment to shifts in the revenue function.

In the secondary sector, employment is determined by a labor demand curve $L^s(w^s)$ with $-\infty<L^{s\prime}<0$. Individuals choose either to supply their labor to the secondary sector at a competitively determined wage w^s, or remain unemployed and receive b. In the latter state, one can participate in a lottery for a job in the primary sector. Each instant, a fraction δ of primary sector employment is released -- say through retirement-- and these vacancies are filled from the pool of unemployed workers. The labor force is maintained constant by a flow of new entrants. It is well-known that if secondary sector workers are unable to queue for these jobs while working, there will be unemployment. In their famous paper, Harris and Todaro (1970) characterized urban unemployment in developing countries in this fashion.

It has been noted in many contexts that it is unreasonable to assume that secondary workers do not also search in some way, and

some empirical evidence from the US indicates that on-the-job search is important. On the other hand, such search must be more difficult than when unemployed. At the very least one cannot be out and about during normal working hours, nor is it possible to remain constantly near the telephone. If search on and off the job were equally costly, the secondary sector wage would be bid into equality with b, a seemingly unrealistic prediction. In the Harris-Todaro model, spatial considerations are operative: to work in the city one must migrate from the country. In efficiency wage models where monitoring workers' effort in the primary sector is costly, it has been shown that a separating equilibrium is possible in which firms hire only from the pool of unemployed, effectively using secondary sector employment as a screening device (Jones 1987). That attachment to such secondary jobs is weak (turnover is high) may be indirect evidence of this phenomenon (OECD Employment Outlook 1985).

If workers are mobile between the secondary sector and unemployment, the reservation wage w^s for jobs in the secondary sector will be such that workers are indifferent between secondary work and unemployment. In equilibrium w^s must satisfy:

$$w^s = b + \delta L^m (w^m - b)/U, \qquad (5)$$

where $U \equiv \bar{L} - L^m - L^s$ is the pool of unemployed workers. Note that when

$U=\delta L^m$, $w^s=w^m$.[12]

The model is depicted graphically in Figure 1. The labor force $\bar{L}$ spans the horizontal axis. The demand for labor in the secondary sector is drawn from left to right in the usual way. The bargain struck by the union and management is at point A, with employment L^m such that marginal revenue equals b. The reservation wage for work in the secondary sector is given by the lower branch of the hyperbola defined by (5) to the left of A with $w^s=b$ and $L^s=\bar{L}-L^m$ as asymptotes. Employment in the secondary sector is determined competitively by the intersection of the reservation wage and secondary sector labor demand schedules (point B). Unemployment is simply the difference between the labor force and L^m+L^s.

The implications of this stylized model for the average duration of unemployment are straightforward to derive. If λ is the probability of employment for an unemployed agent, a spell of unemployment will last i periods with probability $\lambda(1-\lambda)^{i-1}$. For given λ, expected duration is $\sum_{i=1}^{\infty} i\lambda(1-\lambda)^{i-1} = -\lambda(d/d\lambda)\sum_{i=1}^{\infty}(1-\lambda)^i = -\lambda(d/d\lambda)(\lambda^{-1}-1) = \lambda^{-1}$. But $\lambda=\delta L^m/U$, so expected duration is $U/\delta L^m$. This is, of course, the link between the unemployment rate and long-term unemployment. Any conclusions drawn with respect to

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[12] A somewhat more elaborate, explicitly intertemporal version of the model would allow for transitions in both directions. In this case (5) would be replaced by the equality of the recursively defined expected value of wait unemployment and work in the secondary sector with an option to wait in subsequent periods.

unemployment in this model will have implications for its *ex ante* duration.

Wait unemployment described here is neither precisely involuntary nor voluntary. Unemployed individuals are indifferent between unemployment and work in the secondary sector, and at the margin will turn down some secondary work. They do however envy those in the high wage primary jobs; in this sense they are involuntarily unemployed. This is very much in the spirit of current rethinking of the nature of unemployment.[13]

We now present some central results, which along with an important Lemma are proved in the Appendix:

Proposition. In the normal case, an increase in b is associated with higher (wait) unemployment.

Proposition. Increases in the labor force $\bar{L}$ will induce increases in secondary sector employment L^s and have no effect on primary sector employment L^m.

Proposition. In the normal case and with $L^{s''}>0$, the effect of an increase in the labor force on (wait) unemployment is increasing

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[13] Lucas (1978) writes:

"[T]he unemployed worker at any time can always find some job at once, and a firm can always fill a vacancy instantaneously. That neither typically does so by choice is not difficult to understand given the quality of the jobs and the employees which are easiest to find. Thus there is an involuntary element in all unemployment, in the sense that no one chooses bad luck over good; there is also a voluntary element in all unemployment in the sense that however miserable one's current work options, one can always choose to accept them." (p. 354).

in b.

Intuitively, the "normal case" requires the probability of employment for unemployed individuals λ not be so high that the negative effect of higher b on the expected gain from waiting dominates the direct positive effect on w^s, as seen in equation (5). In order to get an idea of how restrictive the "normal" case is, consider the following numerical examples. For α=0.75 and δ=0.05, the upper bound λ^* is 0.25; with a turnover δ of 0.10, λ^* rises to 0.35. For a lower elasticity of revenue with respect to labor, α=0.67, we find that λ^*=0.24 when δ=0.05, and increases to 0.31 when δ=0.10. For plausible values of α and δ, the "abnormal case" is only likely to obtain for unrealistically high employment probabilities.

The first proposition is evident, and is by no means unique to the present model. The second is an implication of the bargaining setup and the supply curve to the secondary sector. The last proposition implies that exogenous demographic humps may induce more unemployment in economies with generous unemployment benefits. The strong convexity condition on labor demand is not terribly restrictive, and is satisfied by a constant elasticity specification. In general, as long as the elasticity of demand for secondary sector labor does not decrease too rapidly as the wage falls, the last proposition will hold.[14]

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[14] $L^{s,,}>0$ is only a sufficient condition guaranteeing $\partial^2 U/\partial b \partial \bar{L}>0$. A more general one requires that the product of the

3. Evidence of Wait Unemployment

3.1. UI Benefits and Duration: The Evidence

The most obvious implication of the model is an association between unemployment rates, especially for the long-term unemployed, and the most important component of b, the provision of unemployment insurance (UI) benefits. There is already substantial microeconomic evidence of a relationship between the duration of unemployment and the level of unemployment insurance. Studies with data from the United States have generally found a statistically significant, albeit small, effect of unemployment benefits on unemployment duration.[15] The importance of the duration of UI benefits is supported by evidence that the rate of exit from unemployment rises as the point of benefit exhaustion nears (Moffitt 1985).

This evidence is not limited to the United States. On a cross-section of unemployed men in the United Kingdom, Narendranathan, Nickell and Stern (1985) find elasticities of unemployment duration with respect to unemployment benefits ranging from 0.28 to 0.36. Franz and König (1986) report that the duration of West German unemployment is higher for recipients of Arbeitslosenhilfe (the second stage of unemployment benefits) than

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elasticity of $L^{s\prime}$ with respect to w^s and the elasticity of w^s with respect to b exceed $\alpha/(1-\alpha)$. See the Appendix for details.

[15] The classic study is Ehrenberg and Oaxaca (1976); for a more recent example on a panel of US states, see Wunnava and Henley (1987). More sophisticated studies that explicitly address potential censoring and other biases reach similar conclusions.

for nonqualifiers. In a review of long-term unemployment in France, Cézard (1986) finds the highest concentration of long-term unemployment among older laid-off workers who are more extensively insured (80% of unemployed males age 50 and older). In contrast, younger clerical workers generally experience shorter durations of unemployment. In a recent study of Canadian males, Ham and Rea (1987) estimate a strong influence of UI eligibility on unemployment duration, and a significant effect of declining entitlement on search intensity.

3.2 Measuring the Extent of UI Benefits

Measurement of the extent of unemployment compensation for cross-country comparison poses a thorny problem. Heterogeneity among recipients with respect to marital status, number of dependents, and age can introduce substantial within-country variation of income replacement. Eligibility may be universal or restricted to those unemployed with sufficient work experience; the coverage ratio (the percentage of unemployed receiving insurance or follow-up assistance) varied in 1985 from 34% in the United States to well over 90% in Belgium. In addition, there are other less easily measured components of b which will vary across countries, including ancillary, often nonmonetary benefits, the value of leisure, and returns to economic activity in the underground economy. Finally, variation in the duration of unemployment insurance provision across countries is substantial,

suggesting that the replacement ratio alone is a poor indicator of the extent of unemployment insurance. Consider the example of the Netherlands and Switzerland. That replacement ratios are roughly similar (70% and 75-80% respectively) might lead one to underestimate the difference in the generosity of the two regimes: in Switzerland one may claim unemployment benefits for 150 days, versus about 2½ years in the Netherlands![16]

In light of these issues, we constructed a crude measure of the extent of UI compensation using a metric of (discounted) percent-weeks, which explicitly accounts for duration as well as the intensity of income replacement and corresponds roughly to an annuity valuation of the expected stream of unemployment benefits. Our source of UI system descriptions is Emerson's (1988) survey of eighteen countries, with the exception of Italy, Greece and Portugal, which were dropped.[17] Whenever possible, we considered replacement and duration for an experienced male worker without assets earning average wages in manufacturing. To avoid "contaminating" the data with prior judgments, midpoints of ranges

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[16] As reported by Emerson (1988).

[17] Three substantive modifications were made to Emerson's descriptions: the UK system, which only imposes a means test after the first year, was considered indefinite; in West Germany, insurance and follow-up assistance replacement ratios for childless male experienced workers were used; in Spain, long-term unemployed since 1980 have access to 50% of their last income after eighteen months, indefinitely. Source: Burtless (1987) and Dolado, et al. (1986).

were taken, which in some countries may bias our measure downwards.[18] The interested reader is referred to the Appendix for further details.

3.3. Simple Correlation Evidence of Wait Unemployment

The resulting measure in 1985 is reported for fifteen economies in Table 3. It exhibits substantial cross-country variation which, as suspected, is almost entirely attributable to the statutory duration of benefits and the coverage rate; the initial replacement ratio was virtually orthogonal to the constructed measure (correlation=0.08). This supports the earlier claim that replacement ratios are an insufficient statistic for the generosity of unemployment insurance, and may explain the absence of empirical evidence for macroeconomic effects on unemployment rates (Burtless 1987).

In Figure 2 we plot the percentage of unemployed in 1985 with current spells exceeding six months against the unemployment insurance measure for thirteen countries for which the former were available. The simple correlation is 0.75. When the rate of long-term unemployment as a fraction of the labor force is used instead, the correlation for the sample is 0.61. The measure was also constructed in 1979 for a more limited sample of economies,

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[18] For example, Burtless (1987) reports that the period of UI entitlement in France for an experienced worker of age fifty years or less is thirty months, considerably higher than the midpoint of Emerson's range (21 months).

for which the correlation was 0.58.

The predictions of the wait unemployment model also find support in the behavior of unemployment, sectoral employment and labor force of our sample of OECD economies.[19] First, there is a strong negative correlation between change in unemployment and change in manufacturing employment (sixteen countries, 1975-85, -0.56). In contrast, the correlation between changes in unemployment and those of total employment is insignificant (-0.20). These findings mirror those of Summers (1986) for US state data, and more recently of Glyn and Rowthorn (1988) for nineteen OECD economies. The model predicts positive correlation of $\Delta\bar{L}$ and ΔL^{s}, which for the sixteen countries examined is very close to unity (0.997). The correlation between $\Delta\bar{L}$ and ΔL^{m} is 0.82 for the full sample, and -0.24 when the US and Japan are excluded.

3.4. Competing Explanations: OLS Regression Evidence

Simple correlation evidence lends support to the qualitative predictions of the model, yet is at best only suggestive in an environment where many forces are simultaneously in play. In this section we use simple least squares regression allow the data to discriminate among competing and complementary alternative theories of high unemployment in our cross-section of OECD

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[19] Austria, France, FR Germany, France, Italy, Norway, Sweden, Belgium, Netherlands, Finland, Ireland, Spain, Australia, Canada, Japan, and United States.

economies. Several alternative have already been discussed in the introduction. We discuss each variable and its proxy in turn.

3.4.1. Aggregate Demand. Demand is proxied by the structural budget surplus net of interest payments series constructed by the OECD (1984), with positive values corresponding to surpluses (SDEF). It corrects the usual surplus as a fraction of GNP or GDP for cyclical components of tax revenues and nets out interest payments.

3.4.2. Excessive Real Wages. According to the disequilibrium wage hypothesis, some high unemployment can be attributed to wages exceeding labor market clearing levels. The indicator of the general level of product wages we selected is the OECD's cumulative labor cost gap since 1972, which is computed on a regular basis for a large number of economies (WGAP). As Bruno and Sachs (1985) note, this measure is the cumulative deviation of product wages from actual average productivity rather than that consistent with full employment.

3.4.3. Corporatism. Bruno and Sachs (1985) and Newell and Symons (1987) have stressed the importance of corporatism in determining labor market responses to adverse shocks. Higher degrees of centralization, ceteris paribus, should be associated with improved macroeconomic performance. This argument rests on the

presumption that institutionalized collective bargaining over wages at the national level will respond in a more coordinated fashion to adverse developments than decentralized systems dominated by shop stewards, craft unions, or plant-level bargaining arrangements.[20] Bruno and Sachs (1985, Chapter 11) have derived a numerical index based on several factors, which we employ (CORP).

3.4.4. Centralization of Wage Bargaining. Calmfors and Driffill (1988) have recently claimed that the link between centralization of bargaining and wage outcomes is not monotonic á la Bruno and Sachs, but "hump-shaped." In their view, both highly centralized and decentralized bargaining structures will fare best, since the former induces internalization of effects on the rest of the economy and the latter approximates perfect competition. Calmfors and Driffill publish a cardinal ranking of the relative degree of centralization, and our approach is to let the data speak on the hump-shaped hypothesis by including their index (CLDR), its square ($CLDR^2$) and a constant as regressors. In a regression with an unemployment rate as the dependent variable, one would expect a negative sign on $CLDR^2$ and a positive sign on CLDR, as least squares implicitly selects the coefficient of

[20] "Corporatism is defined as a mode of social organization in which functional groups rather than discrete individuals wield power and transact affairs" (Bruno and Sachs 1985, p.222).

$(CLDR-k)^2=CLDR^2-2k\cdot CLDR+k^2$. The hump will reach an extreme value when the index takes the value k, and this constant can be identified indirectly as minus the coefficient of CLDR divided by twice that of $CLDR^2$.

3.5. Results

In Table 4 we present results of OLS regressions of the rate of long-term unemployment as a percentage of the labor force in 1985 and 1979 on a constant, the unemployment benefit measure (UIEXT), and various combinations of the variables discussed above. The sample size was constrained by the availability of the corporatism index and the long-term unemployment data to eleven economies. Although its estimated magnitude is larger in 1985 than in 1979, the coefficient on the UI measure is consistently and precisely estimated. For our sample, an increase of one hundred percent-weeks of UI compensation is associated with an increase of 0.1% in the long-term unemployment rate.

Aggregate demand as proxied by SDEF, is often significant, although only marginally in the presence of the corporatism measure. This should not be surprising, since cross-sectional regression should in theory estimate long-run responses, which in a world with neoclassical properties should be zero anyway. Aggregate supply conditions as proxied by WGAP are significant, and as theory predicts seems to be a weak substitute for the corporatism measure. While the coefficents of CLDR and its square

are correctly signed, the estimates have very large standard errors. CLDR and $CLDR^2$ are highly correlated, and imprecision of individual coefficients may be misleading; nonetheless we could not reject the joint hypothesis that both coefficients on CLDR and $CLDR^2$ were equal to zero ($F_{2,6}$= 1.32 and 0.51 in 1985 and 1979 respectively). In contrast, the sign of CORP is consistently negative as predicted and its estimated effect is statistically significant. Taken together, these results suggest that the austere administration of unemployment insurance in the United States and Japan rather than their lack of corporatist traits should be credited for good unemployment performance.

In Table 5 we repeated the exercise with actual unemployment rates in 1985 and 1979. Here the aggregate demand and wage gap variables add considerably more explanatory power to the regressions, especially in 1985. Nonetheless, the effect of UIEXT remains significant albeit quantitatively smaller, as one would expect. The data now reject joint irrelevance of the centralization measure and its square ($F_{2,6}$= 6.56 and 6.29 for 1985 and 1979), but the signs of the estimates are incorrect, indicating that the data prefer to give CLDR a negative coefficient, as if it were a corporatism measure. The corporatism index retains statistical significance, with a magnitude slightly larger than in the long-term unemployment regressions.

The estimated effect of SDEF on unemployment in our cross-section may be puzzling, and suggests an important caveat

with respect to omitted variable bias. Our results could be explained if the adjusted structural deficit is positively associated with some left-out variable with a positive effect on the unemployment rate. More generally, this problem extends to unobservables such as the world business cycle, real interest rate disturbances, or even common exchange rate regimes.[21] Even in the most optimistic case of zero correlation with the regressors, such 'spatial' correlation will affect standard errors and the validity of hypothesis tests.

4. Explaining the Rise in Unemployment

4.1 The Problem

Our model can explain the level of long-term unemployment, and the results of the last section suggest that the provision of unemployment benefits, properly measured, explains a large portion of the cross-country variation of long-term and total unemployment. Difficulties arise when one tries to explain why unemployment was so low in Europe fifteen years ago. Although the image of millions of workers made redundant after OPEC I and II 'waiting' for their old jobs is intuitively appealing, the theoretical model yields no clear-cut result with respect to the effect of b on $\partial U/\partial \theta$. Moreover, the change in unemployment per job loss in industry over the period 1975-85 --a statistic recently

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[21] I am grateful to Jose Viñals for this point.

examined by the OECD (1986)-- is more highly correlated with the UI measure in 1985 (-0.61) than in 1979 (-0.44) or 1975 (-0.32). In the next two sections we attempt to explain why this correlation might have arisen in our sample.

4.2. The Impact of UI on the Absorption of Labor Force Entrants

The last proposition of Section 2 implies that under relatively general conditions, the response of unemployment to an exogenous increase in the labor force is increasing in b. If no systematic factors impinge on the countries in our sample, this implies a positive correlation of $\Delta U/\Delta \bar{L}$ with the unemployment insurance measure. The EEC's labor force grew over the period 1975-85 by 7.4% and by 8.1% for OECD Europe as a whole, so the explanation is plausible. At first glance, the correlation is weak for the sample of fourteen countries (0.28). However, when Spain is excluded, for which $\Delta U/\Delta \bar{L}$ over the period 1975-1985 exceeded 9, the correlation between $\Delta U/\Delta \bar{L}$ and b jumps to 0.71. Dolado *et al.* (1986) offer two explanations for the Spanish case: first, a large number of agricultural workers have been displaced in the past ten years, representing an unusual structural shift less characteristic of other industrial economies in the sample, and second, a high rate of fraud with respect to unemployment benefits claims may have distorted official unemployment statistics.

As an additional test, we regressed alternatively the *change* in the rate of unemployment (DRU) and long-term unemployment

(DRLTU) over 1975-85 on: a constant, the percentage change of change of employment in manufacturing (RDMAN), the percentage change in the labor force (RDLF), and an interaction term between the rate of change of the labor force and UIEXT in 1975 (INTER). The results for the sample of the last section are:[22]

$$\underset{}{DRU} = \underset{(1.7)}{2.89} - \underset{(-1.7)}{11.5}\ RDMAN - \underset{(-1.5)}{20.4}\ RDLF + \underset{(2.9)}{0.680}\ INTER$$

$$R^2=0.81,\ s.e.=1.48$$

$$DRLTU = \underset{(-0.2)}{-0.355} - \underset{(-2.0)}{18.3}\ RDMAN + \underset{(0.1)}{1.58}\ RDLF + \underset{(1.9)}{0.503}\ INTER$$

$$R^2=0.72,\ s.e.=1.64$$

The coefficient of RDLF is insignificant and incorrectly signed, but the interaction term is positive and statistically significant. There seems to be some evidence that high UI benefit economies may be more likely to experience a rise in normal unemployment during periods of high labor force growth.

4.3. Dynamics of Unemployment Benefits

In the current model, the rise in (the natural rate of) unemployment could also be explained by a policy response of benefits to adverse labor market conditions. Under this hypothesis, adverse shocks, even if temporary and "aggregate

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[22] For lack of comparable long-term unemployment data, Japan was omitted from the second regression.

demand" in nature may induce an increase in the generosity of UI benefits, which would lead to higher unemployment when equilibrium is restored.

Table 6 shows that the provision of unemployment benefits has changed over the last ten years.This has occurred in two ways. First, statutory provision of benefits can be altered directly. In France, Germany, Austria, and the United Kingdom, replacement ratios were raised over the 1970s. In Spain, this development continued into the current decade, including an extension of the duration. In contrast, UK replacement ratios have fallen sharply since the abolition of the Earnings Related Supplement in 1982. Authorities may also alter, most often through administrative means, the coverage of unemployment compensation programs. In the past decade, for example, Sweden has expanded its UI coverage beyond the union-administered funds through the government KAS assistance payment (Burtless 1987). West Germany, which has by far the most generous system on paper, has reduced the availability of benefits from about 76% of the unemployed in 1975 to 63% in 1985, mostly though a stiffening of the work requirement for insurance and follow-up assistance. Consequently, new labor market entrants have virtually zero prospect of qualifying for benefits until they have acquired sufficient work experience.[23]

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[23]This may in part explain why youth unemployment has remained low in comparison to other OECD countries. In 1986, the unemployment among young adults (20-24 years old) in FR Germany was 9.1%, compared with 22.2% in France, 19.1% in the UK, 16.6% in

To summarize these dynamics I regressed the adjusted unemployment insurance measure on its own lag, lagged total and long-term unemployment rates, the corporatism variable and a constant for the two years in our cross-section. In addition, we repeated the regressions with the percent-weeks measure unadjusted for coverage (RAWUI), which has less variation but reflects more accurately the compensation provided to job losers.[24] The results are displayed in Table 7.

The main conclusion to be drawn from these regressions is that UI benefits are not provided in a vacuum. Past long-term unemployment has a significant positive influence on the both effective and statutory UI benefit levels, especially in recent years. The $UIEXT_{85}$ equation suggests that an additional percentage point of long-term unemployment in 1979 induced 547 additional effective percent-weeks of UI benefits in 1985, with a steady state response of over 1545 percent-weeks. This may be compared with 734 effective percent-weeks in 1979 (2575 percent-weeks in the long run!) The degree of corporatism is also positively associated with UI compensation, with generosity increasing in the 1980s, from a (statistically insignificant) short-run response of 132 percent-weeks per point of corporatism in 1979 to 375

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the Netherlands, and 10.2% in the United States.

[24] Recall that the adjustment for coverage captures the (lower) unconditional probability of benefits for an unemployed worker, whereby the unadjusted measure is conditional on eligibility.

percent-weeks in 1985.

4.4. Synthesis

Our theoretical model suggests a level effect of UI on normal unemployment and an effect of UI on the absorptive capacity of the labor market. Both of these effects are supported by the data. In addition, the cross-country evidence suggests that UI compensation responds to high long-term unemployment. In the most recent years, corporatism may also be associated with more generous benefit levels.

It should not be forgotten that the provision of unemployment benefits is by no means the only explanation of high European unemployment. As the estimates of the last section attest, aggregate demand and real wages also have key roles in determining jobless rates in the 1980s. Indeed recent cuts in effective UI benefits in the FR Germany and the United Kingdom have probably reduced the amount of wait unemployment in recent years, adding additional scope for a "two-handed" demand expansion suggested by Dreze *et al.* (1987).

5. Conclusions

The segmented labor market model contains some testable predictions for the behavior of unemployment across countries. For any given industrial structure, unemployment and the duration of spells will generally be positively correlated with the level of

unemployment insurance. In contrast to analyses of the previous decade, which stressed the effect of UI on frequency of unemployment spells and work attachment, we focus on duration effects, which are better established empirically. Not only do we show that unemployment rates and our extended measure of UI benefits are positively related, but that under plausible conditions, the effects of demographic bulges and, indirectly, adverse real shocks may result in larger increases in unemployment in high b countries. The model is thus consistent with low rates of unemployment in the 1960s and 1970s.

In its current form, our analysis has some evident limitations. First, the association of the secondary sector with service producing industries is clearly at best a crude approximation. Unfortunately, detailed information on wage structure within the service industries is unavailable for a large cross-section of countries. Second, potential labor market distortions in the secondary sector have been ignored, which may affect the reliability of the model's predictions.[25] As with differences in the severity of UI administration and other aspects of b, if such noise is not systematic across countries or nor too variable, the overall covariance structure between the variables should remain unaffected. Third, the difference between real and

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[25]Such factors may be important in countries where labor unions are involved in wage-setting in both sectors, for example in West Germany. See Burda and Sachs (1987).

nominal shocks has been suppressed. Although the economy is by assumption free of money illusion, sectoral price movements can affect the bargaining pie in the primary sector, labor demand in the secondary sector, as well as b, depending on the price index on which the UI benefit is adjusted. Thus a relative price change of sectoral output will affect both wages, both employment levels, and the rate of unemployment. The importance of this type of shock should not be underestimated, given the increasing openness of mature economies to competition from newly industrializing countries. Finally, to the extent UI benefits must be financed by distortionary taxes on employers and employees, the model neglects important general equilibrium effects that have in part already been identified as operative.

Some of the key empirical results of this paper are arguably consistent with other theories admitting an effect of unemployment benefits on the reservation or alternative wage. Models of search unemployment predict that high unemployment benefit levels induce more persistence in the time series response of unemployment to shocks, as agents take more time to find appropriate job matches. Standard accounts of monopoly union behavior predict that higher UI benefits will be associated with greater real wage rigidity, making economies more likely to adjust to adverse supply shocks through an increase in equilibrium unemployment.[26]

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[26]There may be something to this as well. Our UI benefits measure is highly correlated with Coe's (1985) estimate of short run real

It hardly merits repeating that the central message of this paper is not that the unemployed in Europe are lazy. As is discussed in the text, the distinction between voluntary and involuntary unemployment is blurred when labor markets are segmented. Since the palette of vacancies in real world labor markets is not homogeneous, workers may turn down some jobs and covet others. Moreover, the welfare implications of high versus low levels of unemployment benefits have not been addressed, so policy recommendations are probably premature. It might well be the case that unemployed in Europe are better off than their counterparts in the US and Japan, and the case of Sweden demonstrates the viability of low unemployment alternatives to the US 'market solution.' It should however be stressed that variations across countries and over time of unemployment attributable to 'wait unemployment' are unlikely to be susceptible to the usual demand management remedies.

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wage rigidity for 7 countries (0.61). See Bruno and Sachs (1985) or Minford (1985) for formal arguments.

Appendix I

A key result of this model is the following:

Lemma. In the "normal case," an increase in b leads to an increase in the secondary sector wage w^s.

Proof. If $L^m=(b/\alpha\theta)^{1/(\alpha-1)}$ and $w^m= 0.5(1+\alpha^{-1})b$, then the condition for equilibrium (5) can be written as

$$w^s = b + \delta L^m(w^m-b)/U = b\left[1 + 0.5\delta(\alpha^{-1}-1)L^m/U\right]$$

Partially differentiating w^s with respect to b and rewriting yields

$$\partial w^s/\partial b= 1 +0.5(\alpha^{-1}-1)\delta L^m/U + [0.5b(\alpha^{-1}-1)\delta/U]\left[L^m_b -(L^m/U)(\partial U/\partial b)\right]$$

But $L^m_b=-L^m/[b(1-\alpha)]$, so we may write $\partial U/\partial b=-L^m_b-L^{s\prime}(\partial w^s/\partial b)$ $=L^m/[b(1-\alpha)]-L^{s\prime}\partial w^s/\partial b$, so it is straightforward to solve for $\partial w^s/\partial b$ as

$$\partial w^s/\partial b = \left[1-0.5b(\alpha^{-1}-1)L^{s\prime}\delta L^m/U^2\right]^{-1}(1-0.5\delta L^m/U-0.5\delta(L^m/U)^2/\alpha)$$

$$= \left[1-0.5b(\alpha^{-1}-1)L^{s\prime}\delta L^m/U^2\right]^{-1}(1-0.5\lambda-0.5\lambda^2/\delta\alpha),$$

where $\lambda=\delta L^m/U$. Since the bracketed term is always positive, $\partial w^s/\partial b$ and $(1-0.5\lambda-0.5\lambda^2/\delta\alpha)$ will have the same sign, or

$$\partial w^s/\partial b > 0 \quad \Longleftrightarrow \quad \lambda^2+\alpha\delta\lambda-2\alpha\delta < 0.$$

For given α and δ, the condition for a "normal" (positive) response of w^s to b depends on the initial value of λ. Define the

function $f(\lambda) = \lambda^2+\alpha\delta\lambda-2\alpha\delta$. Since $f'>0$, we seek a critical value λ^*, such that $f(\lambda)<0$ for $\lambda<\lambda^*$. Of the two real roots of f, one is always positive; this is our λ^*:

$$\lambda^* = \frac{-\alpha\delta + \sqrt{\alpha^2\delta^2 + 8\alpha\delta}}{2}$$

As long as $\lambda<\lambda^*$, $\partial w^s/\partial b > 0$.■

<u>Proposition</u>. In the normal case, an increase in the unemployment insurance benefit is associated with higher unemployment.

Proof. Differentiating $U \equiv \bar{L}-L^m-L^s$, we find

$$\partial U/\partial b = -L^m_b - L^{s\prime}(\partial w^s/\partial b),$$

which by the Lemma, is positive in the normal case.■

<u>Proposition</u>. Increases in the labor force $\bar{L}$ will induce increases in secondary sector employment L^s and have no effect on primary sector employment L^m.

Proof. $\partial L^s/\partial\bar{L} = L^{s\prime}(\delta w^m L^m/U^2)(1-\partial L^s/\partial\bar{L})$, so solving for $\partial L^s/\partial\bar{L}$, we obtain

$$\partial L^s/\partial\bar{L} = -\left[(1-\delta(w^m-b)L^m L^{s\prime}/U^2\right]^{-1}\left[\delta(w^m-b)L^m L^{s\prime}/U^2\right] > 0,$$

whereas since $L^m = (b/\alpha\theta)^{1/(\alpha-1)}$, $\partial L^m/\partial\bar{L} = 0$.■

<u>Proposition</u>. In the normal case and $L^{s\prime\prime}>0$, the effect of an increase in the labor force on unemployment is positively related to b.

Proof. First note that $\partial U/\partial \bar{L} = 1-(\partial L^s/\partial \bar{L})$, which is positive as long as $\partial L^s/\partial \bar{L}<1$, or from the proof of the last proposition,

$$-[1-\delta(w^m-b)L^m L^{s\prime}/U^2]^{-1}[\delta(w^m-b)L^m L^{s\prime}/U^2] < 1.$$

Rearranged, this condition is

$$1 - \delta(w^m-b)L^m L^{s\prime}/U^2 > -\delta(w^m-b)L^m L^{s\prime}/U^2,$$

which always holds (since $L^{s\prime}>-\infty$), so $\partial U/\partial \bar{L}>0$. Now

$$\partial U/\partial \bar{L} = 1-(\partial L^s/\partial \bar{L})$$

$$= 1 + [1-\delta(w^m-b)L^m L^{s\prime}/U^2]^{-1}[\delta(w^m-b)L^m L^{s\prime}/U^2],$$

$$= 1 + \left[U^2/[\delta(w^m-b)L^m L^{s\prime}] -1\right]^{-1}$$

or given $w^m = 0.5(1+\alpha^{-1})b$,

$$= 1 + \left[U^2/[0.5\delta b(\alpha^{-1}-1)L^m L^{s\prime}] -1\right]^{-1}.$$

Partially differentiating this last expression with respect to b yields

$$= -\left[U^2/[0.5b\delta(\alpha^{-1}-1)L^m L^{s\prime}] -1\right]^{-2} \times \Big[4U[b\delta(\alpha^{-1}-1)L^m L^{s\prime}]^{-1}(\partial U/\partial b)$$

$$-U^2[0.5b\delta(\alpha^{-1}-1)L^m L^{s\prime}]^{-2}\big[0.5\delta(\alpha^{-1}-1)L^m L^{s\prime} + 0.5b\delta(\alpha^{-1}-1)L^m_b L^{s\prime}$$

$$+ 0.5b\delta(\alpha^{-1}-1)L^m L^{s\prime\prime}(\partial w^s/\partial b)\big]\Big]$$

using $L^m_b=-L^m/[b(1-\alpha)]$ and rearranging slightly,

$$= -\left[U^2/[0.5b\delta(\alpha^{-1}-1)L^m L^{s\prime}] -1\right]^{-2} \times \Big[4U[b\delta(\alpha^{-1}-1)L^m L^{s\prime}]^{-1}(\partial U/\partial b)$$

$$-U^2[0.5b\delta(\alpha^{-1}-1)L^m L^{s\prime}]^{-2}(0.5)\delta L^m\big[-L^{s\prime} + b(\alpha^{-1}-1)L^{s\prime\prime}(\partial w^s/\partial b)\big]\Big].$$

Since $\partial w^s/\partial b$ and $\partial U/\partial b$ are both positive in the normal case, one sufficient condition for $\partial^2 U/\partial b \partial \bar{L} > 0$ is $L^{s,\prime\prime} > 0$. A weaker condition (stated in footnote 15) requires

$$\alpha/(1-\alpha) < \left[-L^{s,\prime\prime} w^s/L^{s,\prime}\right]\left[(\partial w^s/\partial b)(b/w^s)\right],$$

or that the product of the elasticity of the slope of L^s with respect to w^s and the elasticity of w^s with respect to b exceed $\alpha/(1-\alpha)$.■

Appendix II

We consider the situation of a single male without assets earning the average manufacturing wage. A discount rate of 20% (weekly: 0.35%) was applied, to the midpoint or upper bounds to ranges of replacement described in Table A.1. A five year truncation was applied to "indefinite" benefit durations, ie for Belgium, FR Germany, UK, and Spain. When necessary, average manufacturing hourly earnings as reported by the Office of Productivity and Technology, US Department of Labor were used to construct replacement ratios (UK, Sweden, Norway, Ireland, and Denmark). An income tax rate of 30% was assumed for the UK and Ireland in 1985. The Belgian minimum wage was BF 32,054 in 1984 and as by law was adjusted by the CPI inflation rate to compute a 1985 figure. The raw percent-weeks measure was then multiplied by the coverage ratio for each country. A more detailed appendix is available from the author on request.

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Table 1
Long-Term Unemployment in OECD Countries†

	1973	1979	1986
Austria	7.4	8.6	12.6
France	21.6	30.3	47.8
FR Germany	8.5	19.9	32.0
UK	26.9	24.8	41.1
Sweden	--	6.8	8.0
Norway	--	3.8	6.7
Netherlands	12.8	27.1	56.3
Belgium	51.0	58.0	68.9
Finland	--	27.0	21.1
Spain	--	27.1	56.6
Japan	--	16.5	17.2
Ireland	--	31.8	44.2
US	3.3	4.2	8.7

†Percentage of all unemployed with current spells of more than 12 months. Source: OECD Employment Outlook, various issues.

Table 2
Average Duration of Incomplete Unemployment Spells for Select OECD Countries (months)†

	1973	1979	1982
US	2.3	2.5	3.6
Austria	5.3	5.3	4.6
France	9.1	9.5	13.4
FR Germany	4.8	7.8	8.2
UK	8.7	9.6	11.2
Sweden	3.9	3.9	4.4
Norway	1.7	2.0	2.3
Netherlands	6.3	9.8	11.3
Belgium	15.6	17.3	17.3
Finland	---	7.5	6.0

†Source: OECD Employment Outlook, 1983.

Table 3
The Extent of Unemployment Insurance in 1985
(in adjusted percent-weeks)†

Austria	1541
France	2118
FR Germany	5857
UK	3856
Sweden	1973
Norway	1979
Netherlands	6380
Belgium	6937
Finland	4101
Spain	4242
Japan	859
Ireland	3658
Denmark	4794
Switzerland	1378
US	481

†Source: Emerson (1988) and author's calculations. See Appendix for details.

Table 4
Long-Term Unemployment Rate Regressions†

	Const.	UIEXT (00s)	SDEF	WGAP	CORP	CLDR	$CLDR^2$	$\bar{R}^2$	s.e
1985:									
1.	-0.047	0.122						0.53	2.5
	(-0.0)	(3.5)							
2.	0.189	0.071	0.870	0.225				0.80	1.6
	(0.2)	(2.6)	(3.0)	(3.2)					
3.	0.880	0.107	0.558		-0.811			0.74	1.8
	(0.7)	(3.7)	(1.9)		(-2.5)				
4.	0.767	0.109	0.384			0.577	-0.202	0.59	2.3
	(0.1)	(2.1)	(1.0)			(0.1)	(-0.3)		
1979:									
1.	0.260	0.044						0.37	1.3
	(0.4)	(2.6)							
2.	-0.645	0.041	0.276	1.408				0.55	1.1
	(-0.9)	(2.9)	(1.4)	(2.4)					
3.	1.119	0.043	0.042		-0.403			0.48	1.1
	(1.5)	(2.8)	(0.3)		(-2.0)				
4.	-2.54	0.034	0.049			2.07	-0.30	0.19	1.4
	(-0.4)	(1.4)	(0.2)			(0.6)	(-0.7)		

†Dependent variable is fraction of labour force which has been unemployed at least six months. Sample: UK, France, FR Germany, US, Norway, Sweden, Belgium, the Netherlands, Finland, Japan, Austria. t-statistics in parentheses.

Table 5
Unemployment Rate Regressions†

Const.	UIEXT (00s)	SDEF	WGAP	CORP	CLDR	$CLDR^2$	RU_{-1}	$\bar{R}^2$	s.e.
1985:									
3.39	0.050	1.242	0.298					0.90	1.30
(4.5)	(2.4)	(5.5)	(5.4)						
4.35	0.098	0.829		-1.10				0.82	1.71
(3.9)	(3.7)	(3.1)		(-3.7)					
16.3	0.142	0.497			-6.06	0.548		0.71	2.16
(0.2)	(2.6)	(3.0)			(-1.1)	(0.8)			
1.08	0.031	1.164	0.192				0.685	0.96	0.84
(1.3)	(2.0)	(7.9)	(4.0)				(3.3)		
1979:									
1.89	0.036	0.506	1.34					0.19	1.81
(1.6)	(1.5)	(1.6)	(1.4)						
4.45	0.038	0.347		-0.805				0.73	1.01
(6.5)	(2.8)	(2.4)		(-4.4)					
6.28	0.040	0.393			-0.602	-0.066		0.43	1.51
(1.0)	(1.6)	(1.8)			(-0.2)	(-0.1)			
0.29	0.030	0.408	1.30				0.533	0.49	1.4
(0.3)	(1.6)	(1.6)	(1.7)				(2.3)		

†Sample of eleven economies in Table 4. RU_{-1} is the unemployment rate in 1979 for the 1985 regressions, and the unemployment rate in 1975 for the 1979 regressions. t-statistics in parentheses.

Table 6
UI Benefits in 1975, 1979, and 1985
(in adjusted percent-weeks)†

	1975	1979	1985
Austria	1543	1674	1541
France	912	2254	2118
FR Germany	7594	6650	5857
UK	7699	5882	3856
Sweden	1427	1652	1973
Norway	1350	1296	1979
Netherlands	5658	5896	6380
Belgium	7123	6974	6937
Finland	2656	3211	4101
Japan	1857	1176	859
US	1089	608	481

†Source: Emerson (1988) and author's calculations. See Appendix for details.

Table 7
The Dynamics of UI Compensation†

$$UIEXT_{85} = \underset{(-1.8)}{-6.62} + \underset{(7.0)}{0.646}\ UIEXT_{79} + \underset{(3.5)}{5.47}\ LTU_{79} + \underset{(3.6)}{3.75}\ CORP$$

$\bar{R}^2=0.95$ s.e.=4.87

$$UIEXT_{79} = \underset{(-0.2)}{-1.12} + \underset{(8.8)}{0.715}\ UIEXT_{75} + \underset{(2.2)}{7.34}\ LTU_{75} + \underset{(1.1)}{1.32}\ CORP$$

$\bar{R}^2=0.94$ s.e.=5.98

$$RAWUI_{85} = \underset{(-0.7)}{-3.67} + \underset{(7.8)}{0.696}\ RAWUI_{79} + \underset{(2.2)}{4.15}\ LTU_{79} + \underset{(2.1)}{3.10}\ CORP$$

$\bar{R}^2=0.93$ s.e.=6.88

$$RAWUI_{79} = \underset{(0.6)}{2.27} + \underset{(15.8)}{0.909}\ RAWUI_{75} + \underset{(1.4)}{3.68}\ LTU_{75} \underset{(-0.2)}{-0.168}\ CORP$$

$\bar{R}^2=0.98$ s.e.=4.50

†Both dependent variables measured in hundreds of percent-weeks. RAWUI is the raw (statutory) percent-weeks measure, without adjustment for effective coverage.

TABLE A.1

Unemployment Benefits in 1985*

United States	50% of earnings (approx.); 26-34 weeks + up to 13 weeks in selected regions. Coverage: 34% (Burtless, 1988)
Japan	80-60% of earnings, 3-12 months depending upon age and other criteria. Coverage: 40%.
France	FF40/day + 42% of earnings, 1-2 ½ years; also special allowance, 3 years. Coverage: 66%.
Germany	63% of earnings, up to 1 year, then 54% indefinitely (means tested). Coverage: 63%.
United Kingdom	£28.50 in first year, means-tested thereafter indefinitely. Coverage: 81%.
Belgium	60% of earnings; 40% 2nd year, then 50% of minimum wage indefinitely. Coverage: 94%.
Netherlands	70% of earnings (262 fl. max. daily) for 26 weeks, then 70% with means-test up to 2 years. Coverage: 89%.
Denmark	90% of earnings (335 kr. max. daily), up to 2 ½ years. Coverage: 79%.
Norway	73% of earnings (310 kr. max. daily), up to 40 weeks. Coverage: 85% (estimate).
Finland	Max. 75% or earnings or 70 mark/day, up to 450 days in 3 years. Coverage: 96%.
Sweden	80-300 kr./day, 1 year. Coverage: 87%.
Switzerland	75-80% earnings, 150 days. Coverage: 87%.
Austria	100-50% earnings, varying inversely with wage up to 30 weeks. Coverage: 73%.
Ireland	£39.5/week + 40-20% of earnings varying inversely with wage, 85% max.; up to 390 days. Coverage: 70%.
Spain	80% earnings, 180 days; 70% to 12th month; 60% to 18th month, 50% afterwards (Dolado et al. 1986). Coverage: 44%.

*Source: Emerson (1988). Coverage defined as % of all unemployed receiving unemployment insurance or assistance; author's calculations except when noted. Thanks to Pekka Hietala, Carmen Matutes, Retsu Sato, Jon Landmark and Tracey Kerntiff for kind assistance with national data sources.

FIGURE 1

Wage and Employment Determination

in the Segmented Labor Market Model

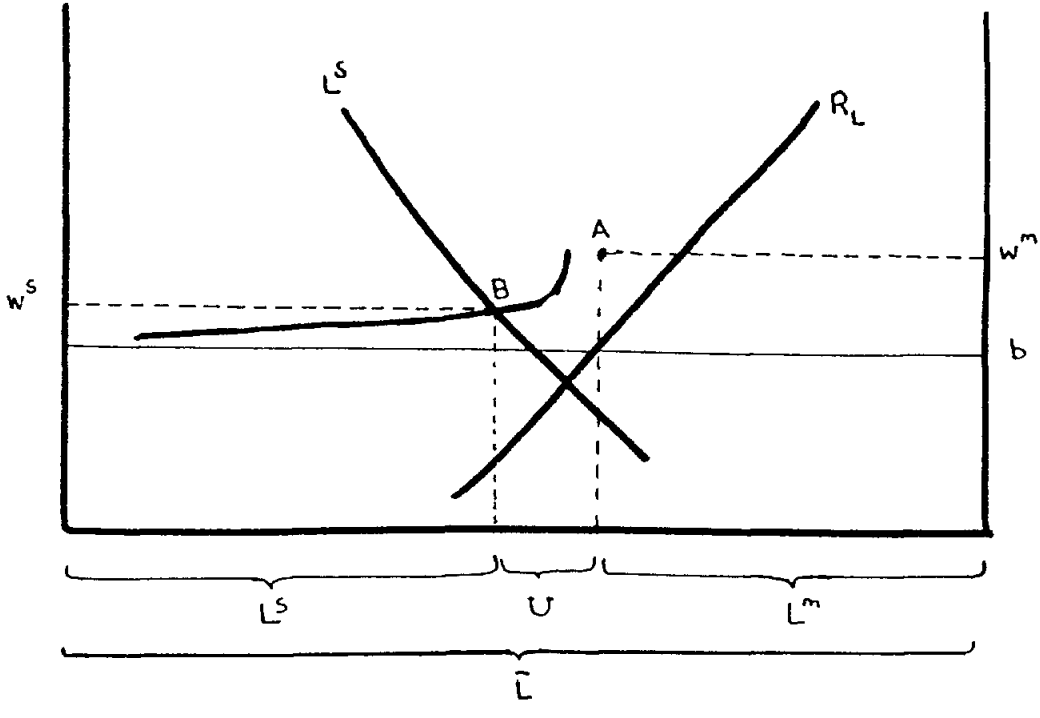
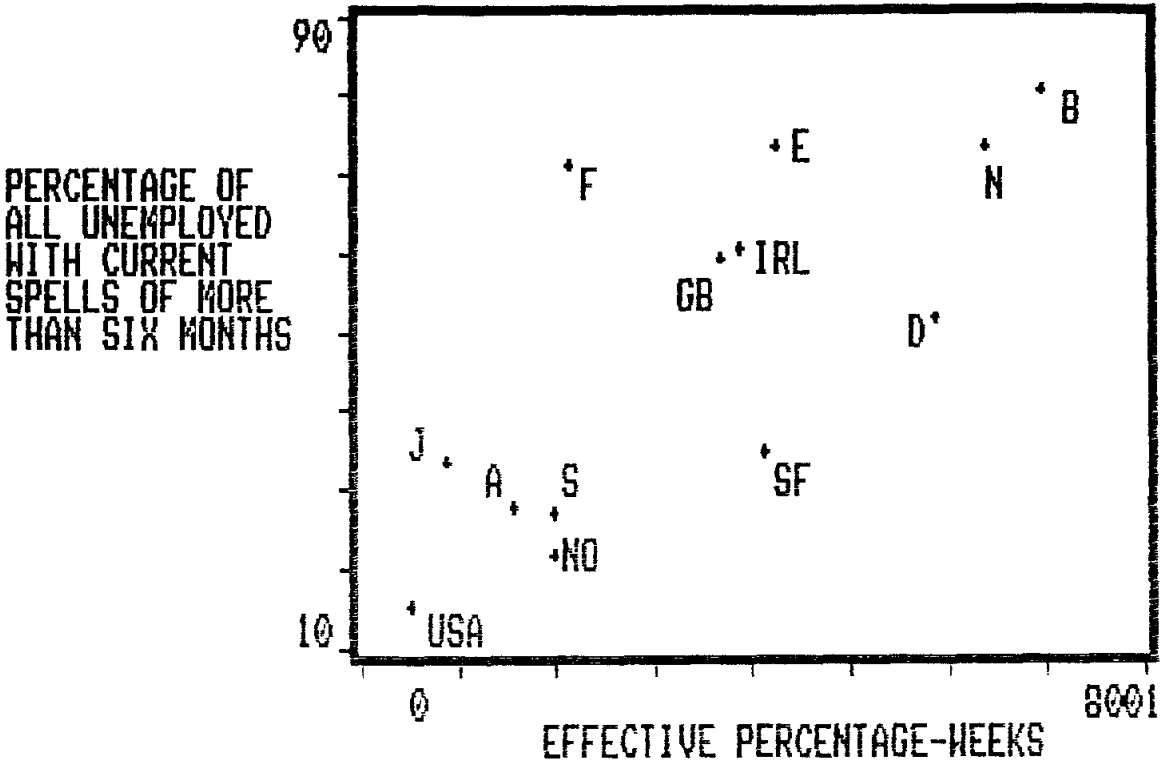


FIGURE 2

LONG-TERM UNEMPLOYMENT AND THE EXTENT OF UNEMPLOYMENT INSURANCE IN 1985



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