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Abstract

Job search models offer two complementary predictions about the effects of unemployment benefits on job search outcomes among unemployed workers. By raising workers' reservation wages, unemployment benefits should contribute to both prolonged spell duration and improved post-unemployment job quality. In contrast to many previous empirical studies that have addressed the negative benefit effect on duration only, the current paper jointly addresses the causal effect of unemployment benefits on both unemployment duration and post-unemployment wages. Based on panel data from the Survey of Income and Program Participation and the German Socio-Economic Panel for the 1980s and 1990s, the paper establishes empirical support for both benefit effects in both countries. If anything, there is evidence of a slightly more negative duration effect for the U.S. data, while positive UI effects on post-unemployment wages are stronger in the German data. In any event, the empirical estimates for the positive effects of unemployment benefits on wages substantially exceed those obtained in Addison and Blackburn's recent paper based on Displaced Worker Survey data. In contrast to their findings, the data also provide ample evidence of stronger UI effects in the lower tails of the wage change distribution. At the cost of a fairly small prolongation of unemployment duration, unemployment benefits thus substantially reduce the scar effects of unemployment on workers' future job records.

Keywords: Unemployment insurance, unemployment duration, wage change, displaced workers, search theory

JEL-Codes: J64, J65

Zusammenfassung

Die Suchtheorie macht zwei komplementäre Vorhersagen über den Einfluß der Arbeitslosenversicherung auf die Dynamik von Arbeitslosigkeit. Indem die Anspruchslöhne der Arbeitslosen erhöht werden, sollte die Arbeitslosenversicherung sowohl zu einer verlängerten Arbeitslosigkeitsdauer als auch zu einem verbesserten Matching bei Wiederbeschäftigung beitragen. Im Gegensatz zu anderen empirischen Studien, die nur den negativen Effekt der Dauer der Arbeitslosigkeit betrachten, werden im vorliegenden DP beide Effekte empirisch untersucht. Auf der Basis des sozio-ökonomischen Panels sowie vergleichbarer Paneldaten des Survey of Income and Program Participation können diese Effekte für beide Arbeitsmärkte empirisch nachgewiesen werden. Im Vergleich sind allenfalls die negativen Dauereffekte der Arbeitslosenversicherung in den USA etwas stärker, während in der Bundesrepublik die positiven Effekte auf die Qualität des Matchings vergleichsweise höher ausfallen. Die geschätzten positiven Lohneffekte wohlfahrtsstaatlicher Transfers liegen insgesamt höher als in der Studie von Addison und Blackburn, die auf Daten des Displaced Worker Survey basierte. Im Gegensatz zu dieser Studie lassen sich auch besonders starke Transfereffekte im unteren Bereich der Lohnveränderungsverteilung nachweisen. Auf Kosten einer geringen Verlängerung der Dauer von Arbeitslosigkeit reduzieren wohlfahrtsstaatliche Transferleistungen folglich die negativen Konsequenzen von Arbeitslosigkeit für den weiteren Erwerbsverlauf.

Contents

Introduction

Recently, the microeconomics of unemployment has seen a remarkable shift in emphasis. Whereas the spread of transition data models and longitudinal data had spurred numerous studies of unemployment duration in the late 1980s and up to the mid-1990s (e.g. Devine and Kiefer 1991; Pedersen and Westergård-Nielsen 1993; Machin and Manning 1999 for reviews), attention has increasingly turned towards the economic consequences of unemployment experiences since then. Recently, for example, the Economic Journal featured a series of articles (Arulampalam 2001; Gregory and Jukes 2001; Gregg 2001; cf. Arulampalam et al. 2001 for an introduction) that established substantial scar effects of unemployment, both of youth unemployment on adult unemployment (Gregg 2001), but also for unemployment experiences of prime-age workers on their future work careers. In contrast to popular assumptions among economists, these scar effects are far from transitory: Arulampalam (2001) e.g. gives an initial 6% wage penalty associated with unemployment, that perpetuates into a full 14% penalty over the first three years after leaving unemployment. Gregory and Jukes (2001) estimate a one-year spell of unemployment to imply a permanent wage penalty of 10%. Arulampalam (2001) in particular insists that much of the scarring occurs in workers' first post-unemployment jobs which is carried forward subsequently rather than being gradually eroded over time. Also, Nickell et al. (2002) recently stressed that wage penalties associated with unemployment have grown considerably over the past decades. While still limited in coverage of countries, different aspects of job outcomes, and stringent incorporation of changing macroeconomic conditions, these recent studies nevertheless do strongly suggest that unemployment experiences entail serious negative consequences for individuals' further careers, and that negative effects of unemployment might be much more persistent than commonly perceived.

This shift in emphasis has in fact also prompted new perspectives on the effects of unemployment insurance (UI) on unemployment dynamics, or has at least contributed to revive interest in some long-neglected issues. It is particularly revealing that, while numerous studies had focused on establishing empirical evidence for predictions of negative UI effects on unemployment duration, there was and still is very little empirical research on positive UI effects on post-unemployment wages predicted by basic job search theory. In their recent survey, Addison and Blackburn (2000) cite only a handful of respective studies, and in fact, most of these date back to the late 1970s, and none had appeared since the mid-1980s.[1] Given the increasing popularity of search models, the dearth of respective empirical research is ironic indeed: as Addison

[1] Interestingly, there is a similarly renewed interest in the scar effects of unemployment, and the potential alleviating effects of the welfare state in recent empirical work in sociology (cf. Gallie and Paugam 2000).

and Blackburn (2000:22f.) correctly point out, a negative UI effect on unemployment duration can readily be derived from the simple static labor-leisure framework, without any necessity to introduce search models, dynamics or the notion of a reservation wage at all. In fact, only empirical evidence on UI effects on post-unemployment wages (and job quality more generally) will allow to adjudicate whether UI simply lowers the opportunity cost of leisure or whether UI is better seen as lowering the opportunity cost of job search, which might even increase the efficiency of matching in the labor market.

Along these lines, Addison and Blackburn (2000) offered empirical evidence on UI effects on post-unemployment wages based on data from the 1983-1990 Displaced Worker Surveys. In their analyses, they actually obtained some support for positive UI effects on wages, although the implied magnitudes of the effects were typically small and often statistically insignificant. More recently, Belzil (2000) has conducted a related analysis of a large Canadian administrative database, and found only small positive UI effects on post-unemployment job duration. Against this background, the current paper extends these analyses by presenting empirical evidence on UI effects based on more recent 1984-1995 data from the Survey of Income and Program Participation. Moreover, the paper offers complementary evidence from UI effects in the West German labor market based on the German Socio-Economic Panel study. Like Belzil (2000), yet in contrast to Addison and Blackburn (2000), I will be using a particular variant of hazard rate modeling that allows for joint estimation of the causal effects of unemployment benefits on unemployment duration and post-unemployment wages. In contrast to both recent studies, however, I establish a sizeable positive UI effect on post-unemployment wages in both the U.S. and the West German data.

1. Benefit effects on unemployment dynamics: a brief survey

The relationship between unemployment insurance and labor market dynamics has since long been one of the key issues in modern labor economics, and considerable work has been devoted to clarifying the respective mechanisms at work. Job search models have typically been the work horse in such endeavors, and many of the models' predictions precisely relate to the effects of unemployment benefits, or more accurately non-earned income while unemployed. The basic theory is conventional by now, and there is little need to give an extended account here (cf. the surveys by Mortensen 1986 and Devine and Kiefer 1991; Lippman and McCall 1976a, 1976b). Basically, job search theory represents the job-finding process as a sequential decision process, with job offers randomly arriving and workers having to sequentially decide whether to accept the current offer or to continue job search. According to dynamic programming theory, the optimal search strategy in such contexts is to form a reservation wage as the minimum acceptable wage offer, taking into account features of the labor market environment like job offer arrival rate, the wage distribution, but also available non-earned income and the disutility of work. Having chosen the reservation wage, workers' best strategy is to hold fast for the first offer that exceeds the reservation wage, to accept this offer, yet decline all lower ones received at earlier stages in the job search process.

In the simplest environment of a stationary wage distribution, an infinite time horizon, and inexhaustible unemployment benefits (strictly speaking: non-decreasing non-earned income net of out-of-pocket search costs), the basic job search model yields the constant reservation wage w_r that satisfies the condition

(1) $$w_r - (b+v) = \delta\,[1 - F(w_r)]\,[E(w \mid w > w_r) - w_r]/\,r \quad ,$$

where b gives non-earned income, v the value of leisure, δ the job offer arrival rate, r workers' discount rate, and $F(.)$ the cumulative wage distribution (cf. Mortensen 1986; Devine and Kiefer 1991). With respect to the role of non-earned income b, total differentiation yields the standard result

(2) $$\partial w_r / \partial b = r / [r + 1 - F(w_r)] > 0 \quad ,$$

predicting workers' reservation wages to increase with increasing levels of available income while not working. As unemployment benefits are assumed to raise b for the average worker over the alternative income stream provided by having to rely on family resources and general assistance schemes only, most observers agree on seeing the rise in reservation wages as the main mechanism behind any effects of unemployment insurance on wages (see Atkinson and Micklewright 1991 for a discussion of alternative channels, however). Hence, a most important implication of search theory is the prediction that higher levels of non-earned income (= unemployment benefits) should translate into more favorable post-unemployment wages, and thus smaller scar effects of unemployment. From this starting point, however, most studies are then quick (and correct) to point out that higher reservation wages imply longer spell duration of unemployment, as higher reservation wages imply higher rejection rates of incoming job offers. Indeed, the basic job search model implies the negative relation

(3) $$\partial \tau / \partial b = -r\delta\, f(w_r)/(r + \tau) < 0 \quad ,$$

between job exit rates τ and non-earned income b, so that higher income levels during unemployment spells should prolong unemployment duration.

In sum, job search theory thus offers *two complementary* predictions on the effects of unemployment insurance (UI) on job search outcomes (cf. Mortensen 1986, 1990; Lippman and McCall 1976a, 1976b; Burdett 1979; Atkinson and Micklewright 1991 have a critical discussion of the literature). In essence, the job search argument is that *by raising worker reservation wages, unemployment benefits contribute to both prolonged spell duration and improved post-unemployment wages*. As Addison and Blackburn (2000:23) in a recent paper correctly note, it is the latter prediction that sharply distinguishes the job search interpretation of UI effects from the standard static labor-leisure model. Whereas both models agree in the prediction of negative UI effects on unemployment duration, both models clearly diverge in terms of the alleged source of the effect. To the standard model, the negative UI effect results from the disutility of work, and UI payments thus subsidizing leisure. In the dynamic job search model, however, the negative UI effects on duration are not seen as arising from the disutility of work, but from UI benefits *lowering the opportunity cost of job search*, which by reducing current constraints on search behavior allows workers to trade off some prolongation of current job search (i.e. unemployment) for improved job matches that imply higher utility levels in the longer run. To the extent that both negative UI effects on duration *and* positive UI effects on post-unemployment wages could be established, this evidence would sustain Burdett's (1979) predicament of UI acting as a search subsidy rather than a static work disincentive.

Empirical studies

Against this background, it is in fact surprising to see the bulk of empirical studies focusing on establishing UI effects on unemployment duration. Effectively, most empirical studies seem to conform to expectations, and establish negative effects of unemployment insurance on unemployment duration (e.g. the reviews in Devine and Kiefer 1991; Pedersen and Westergård-Nielsen 1993; Holmlund 1998; Machin and Manning 1999). By now, robust evidence on UI duration effects has been obtained for a number of countries over the past decades, covering the U.S. (e.g. Meyer 1990; Katz and Meyer 1990a, 1990b; Fallick 1991), but also Canada (e.g. Belzil 2000) and European labor markets (e.g. Hunt 1995; Hujer and Schneider 1996; Steiner 1997 for Germany; Carling et al. 2001 for Sweden, or Narendranathan et al. 1985; Narendranathan 1993; Arulampalam et al. 1995 for UK data). However, it seems to be the case that UI effects on duration are typically fairly small in magnitude, except among particular sub-populations like male youth during the school-to-work transition (e.g. Narendranathan 1993). Also, some studies have suggested stronger UI effects from UI eligibility rather than the level of UI replacement rates (e.g. Steiner 1997) – although it might be argued here that this result merely reflects the larger variation in non-earned income *b* inherent in comparing UI recipients to non-recipients relative to the often (especially in

Europe) fairly small cross-sectional variation in actual UI replacement rates. What is important here, however, is the fact that numerous earlier econometric studies almost unequivocally point to some negative UI effect on unemployment duration. Yet as argued above, only an analysis of UI wage effects will be able to tell whether this effect results from UI inducing work disincentives (no wage effects) or from UI subsidizing job search (associated with positive wage effects) and protecting workers' stock of human capital acquired throughout their previous work histories.

Unfortunately, there are few econometric studies to date actually evaluating benefit effects on post-unemployment wages. Cox and Oaxaca (1990), Burtless (1990), and more recently Addison and Blackburn (2000) have surveyed the scant evidence available, and Addison and Blackburn (2000:22) in particular have noted the mixed empirical evidence on positive UI effects on wages that is "varied enough in their approach and conclusions that experienced observers can reach very different interpretations to [the] findings." Addison and Blackburn (2000) then continue to provide their own analyses of UI effects on post-unemployment wages based on Displaced Worker Survey (DWS) data, and conclude that some small positive UI effects on wages might exist, although the evidence does not appear to be statistically robust and the effects are assessed as substantively rather small. Closer inspection of their results suggests, however, that their relatively negative predicament is mainly based on their analyses of replacement rate effects among UI recipients that indeed yield little evidence of any positive effects on wages. In their analyses comparing UI recipients to non-recipients, however, Addison and Blackburn in fact obtain a positive UI effect on post-unemployment earnings in the order of 2%-5% of workers' pre-unemployment earnings levels, which they consider as sizeable themselves (2000:38). Also, they obtain some (weak) evidence of somewhat larger positive UI effects among a sub-sample of more experienced workers, and in the lower tail of the wage change distribution – both features which would seem to fit favorably with the standard search account of UI effects on unemployment dynamics.

Other than this study by Addison and Blackburn, there is very little recent empirical work on UI effects on post-unemployment job outcomes. To the best of my knowledge, there is only one closely related study by Belzil (2000), who addresses UI effects on post-unemployment job stability for a register-based sample of Canadian displaced workers receiving unemployment insurance. Even more clearly than Addison and Blackburn, Belzil (2000) claims evidence of only very small positive UI effects on job stability. His analyses show that a considerable part of UI effects in his models is accounted for by unobserved heterogeneity between workers, so that his preferred estimates of actual UI effects lie in the order of an increase in expected job duration by about 0.5-0.9 days per additional week of benefit duration.

2. Data and econometric methodology

Against these rather negative findings and assessments of UI benefits on unemployment dynamics, the current paper will provide some new analyses based on both new data, somewhat different methodologies, and covering German in addition to U.S. data. Basically, the analytical purpose behind this cross-country study is to use the cross-national comparison as a particular kind of sensitivity analysis that allows to assess the robustness of findings across institutionally and structurally distinct national labor markets. Any U.S.-German comparison indeed provides a veritable array of such differences, be it in terms of the dynamics of labor markets (Schettkat 1992; Garibaldi et al. 1997), the extent of labor market regulation (Grubb and Wells 1993; Abraham and Houseman 1993; OECD 1999), the structure of education and training systems (Müller and Shavit 1998), or the extent of welfare state compensation of income losses (Esping-Andersen 1990; Mitchell 1991). In terms of the latter, Germany as well as many other Continental European countries offers a much more extensive protection against life-course risks like unemployment, ill health, family disruption or old age than is common in the United States. The structure of unemployment insurance is in fact a quite instructive example about the nature of such institutional differences. While UI benefit replacement ratios actually differ relatively little between the United States and Germany (Schömann et al. 2000; Schmid and Reissert 1996; Esping-Andersen 1990), it is benefit eligibility criteria that are considerably stiffer within the U.S. UI system (Grubb 2000). In consequence, actual UI benefit coverage rates among unemployed workers are considerably lower in the United States as compared to the more universal benefit coverage for German prime-age workers (Schmid and Reissert 1996).[2] German unemployment insurance is thus not necessarily more generous in terms of genuine transfer amounts, yet much more encompassing in terms of worker eligibility for unemployment benefits.

Data sources

The current analyses will be based on employment history data drawn from the U.S. Survey of Income and Program Participation (SIPP; U.S. Bureau of the Census 1991) and the German Socio-Economic Panel (GSOEP; cf. Wagner et al. 1994; DIW 1999) study. Both studies are household panel surveys representative of each country's residential population, and both surveys provide rich databases on individual labor market behavior, employment,

[2] Calculated from cross-sectional samples of unemployed workers, Schmid and Reissert (1996:244f.) give UI coverage rates between 70-80% for West Germany in the 1980s and early 1990s. Based on data for the mid-1990s, Schömann et al. (2000: Appendix 1) arrive at UI coverage rate estimates of 40% in the U.S., and 74% for the unified Germany. As these data are based on cross-sectional samples of unemployed workers, the figures will tend to underestimate coverage rates among an inflow sample from employment into unemployment, as will be used in this analysis.

unemployment and job dynamics. Although sharing largely similar interests, both surveys to some extent differ in terms study design. In particular, while the GSOEP design very much follows the design chosen in the Panel Study of Income Dynamics (PSID) in combining annual interval lengths between interviews with extensive retrospective information on both individual life courses and calendar information on labor market events in the year preceding the interview, the SIPP is based on much shorter four-month intervals between interviews. Also, single SIPP panels have been discontinued after eight to ten interviews, whereas the GSOEP sample (including some sample refreshments) has been continuously followed since its original start in 1984. Against these differences in study design, however, both surveys are likely to represent the most appropriate data sources on (short-run) labor market and unemployment dynamics in both countries (cf. also Witte 1989).

For the purpose of this paper, harmonized data from the combined SIPP Panels 1984, 1986, 1988, 1990, 1992, and 1993, and the West German data from GSOEP waves A-M (samples A+B) has been used to generate monthly calendar information in the 12-year observation window between January 1984 and December 1995. To address the effects of UI benefits on unemployment dynamics in the two countries, the subsequent analyses use an inflow sample of all unemployment spells among displaced workers that were begun during this observation period. Throughout this paper, displaced workers are defined rather liberally as workers having entered unemployment from dependent employment immediately preceding an unemployment spell.[3] Hence, the spell sample drawn here excludes any unemployment spells of both first-time entrants to the labor force, but also job search periods of (mostly) women returning to the labor market after career interruptions. The intention behind restricting the analysis to the core work force highly attached to the labor market is to evaluate the effects of UI benefits precisely with respect to those events that UI benefits have been primarily designed to compensate for, namely job losses.

Under these restrictions, the combined SIPP data yield a sample of 24,100 unemployment spells of 21,551 workers that are observed for a total of 98,749 observation months. The smaller GSOEP database still gives a total of 3,251 unemployment spells of 2,264 workers that are observed for a total of 32,498 months. Rates of right-censoring are 17.7% (4,254 spells) in the SIPP, and 11.9% (387 spells) in the GSOEP data. As the spell samples have been drawn conditional on pre-unemployment status, the samples used here by definition exclude any left-censored or left–truncated spells. Added to the core

3 In technical terms, any unemployment spell has been sampled from the two databases if individuals reported to have worked at least up to three months before the start of an unemployment spell. This maximum inactivity gap of two months has been allowed for in order to minimize the impact of late benefit take-up or workers' recall expectations that might result in reporting some time of inactivity rather than active job search behavior.

spell information, the databases include gender, age, ethnicity, workers' education (including completion of vocational training in the German sample), labor force experience, tenure, occupation, industry, earnings and wages with previous employer as main worker-level characteristics, but also a measure of the quarterly vacancy ratio calculated by the quarterly number of hires over the average number of unemployed in any given quarter as an indicator of aggregate labor market dynamics. All earnings and wage data are deflated to 1990 U.S. prices, with German earnings and wage data being adjusted by 1990 purchasing power parities after deflation. Unemployment benefit status is measured time-constant, with benefit receipt being recorded if workers reported receiving UI transfers in any month of the unemployment spell. Compared to properly accounting for the effects of late benefit take-up, temporary benefit disqualification or simple measurement errors, this appeared as the much more robust measure, especially for the purposes of cross-national comparison. The distribution of covariates in the two samples is given in full in Appendix 1.[4]

Econometric modeling

For this spell dataset, two main outcome measures of interest have been recorded: whether exiting unemployment has occurred by taking up a new job, and if so, the wage level of the job entered. The resulting duration data is most conveniently analyzed in a hazard rate framework (e.g. Lancaster 1990; Kiefer 1988; Neumann 1999; Petersen 1995). In line with search theory concerns, the subsequent analyses will primarily address the hazard rate of leaving unemployment into employment as the key dependent variable describing unemployment duration. In this framework, this hazard rate *r(t)* is defined as

$$r(t) = \lim_{\Delta t \to 0} \frac{\Pr(t \leq T < t + \Delta t, T \geq t)}{\Delta t} \quad , \tag{4}$$

representing individuals' instantaneous propensity to leave unemployment at spell time *t*, conditional on the fact that no such event has taken place up to spell time *t*. As the following analyses will apply a discrete-time approach based on monthly spell data (cf. Allison 1982; Meyer 1990), equation (1) becomes the probability of exiting unemployment for paid work within the next monthly interval *t*+1, given that workers have stayed in unemployment until spell time *t*. Also, basic econometric theory on rate models tells that knowledge of *r(t)* is sufficient to deduce several alternative representations of the unemployment duration distribution, including the duration distribution *f(t)* itself, but more importantly also the cumulative duration distribution *F(t)* and the survivor function *G(t)* = 1-*F(t)* (Lancaster 1990).[5]

4 All tables and figures in the appendix page I-X

5 Note at this stage, however, that the following analyses will not address the unemployment duration distribution strictly speaking, as I will refrain from incorporating flows between unemployment and

Modeling duration distributions in terms of hazard rates rather than any other equivalent distribution offers the advantage of easy incorporation of censored cases, i.e. ongoing spells of unemployment by the end of the observation window. Also, if destination states differ in terms of quality, such qualitatively different transitions are straightforward to address in competing-risks frameworks that include several separately estimated rate equations. As this paper is concerned with both reemployment rates and post-unemployment wage levels, the subsequent analyses will use a particular competing-risks specification originally developed by Heckman and Singer (1984) and first applied in Petersen's (1988, 1995) analyses of status attainment processes. More specifically, Petersen (1988, 1995: 500f.) decomposes the destination-specific hazard rates $r_k(t)$ into

$$r_k(t) \equiv r(t) \times \Pr(D = k, T = t) \quad (5)$$

i.e. the product of the overall exit rate $r(t)$ and a destination equation predicting the type of exit k (which might e.g. contrast entering jobs associated with wage losses versus jobs associated with no wage loss). More importantly, Petersen's competing-risks formulation is readily generalized to continuous state space settings (cf. Petersen 1988), i.e. continuous measures of wage levels or other aspects of job quality outcomes. In this case, equation (5) is straightforward to extend into

$$r_k(t) \equiv r(t) \times g(y \mid T = t), \quad (6)$$

where job quality y is measured continuously (Petersen 1988:144). In sum, equations (5) and (6) thus describe a basic, yet quite general and flexible econometric framework to jointly address UI effects on reemployment rates and discrete (equation 5) or continuously measured wage outcomes (equation 6).

In contrast to Petersen's original paper that presented a two-step limited-information maximum likelihood (LIML) estimator of equations (5) and (6), the current paper will implement full-information maximum likelihood estimators (FIML) for these particular competing-risks specifications. Most importantly, conditional on functional form assumptions, the FIML approach developed below allows to test for conditional independence between reemployment rates and wage outcomes, while this assumption has been implicit in Petersen's original LIML estimator. Also, being able to avoid imposing an independence assumption seems indispensable on theoretical grounds, given that job search models clearly imply the prediction of truncation from below in the distribution of accepted wages (equivalent to positive correlation between exit rates and post-unemployment wage levels). In the following, I apply FIML estimators of both (5) and (6), as both continuous and discrete measures of wage outcomes will be

inactivity but rather focus on reemployment rates $r(t)$ only. Implicitly, of course, this assumes conditional independence between U-E and U-N transition rates.

defined in order to address UI effects on different aspects of the post-unemployment wage distribution. In addition to a standard measure of real wage change between pre- and post-unemployment jobs, the paper will also consider discrete measures at different cut-off points in the distribution: given that search models argue about truncation of wage distributions being the main mechanism behind benefit effects, subsequent empirical analyses will address whether UI effects on wages increase in the lower tail of the wage change distribution, i.e. if UI benefits are particularly effective in preventing (severe) wage losses upon reemployment. To that end, empirical results that look at UI effects on the probability of wage losses of various degrees will be presented below.

Assuming joint normality of the latent rate index function and the latent index function underlying the probability of a certain wage change, the FIML estimator in question then becomes a particular variant of a bivariate probit specification of the model. Using Ö to represent the standard cumulative normal distribution, Ö$_2$ to represent the cumulative bivariate normal distribution, and *ä*$_{wk}$ indexing the occurrence of a work exit into jobs of quality k, this results in

$$(7) \qquad \text{Ln}\, L = \sum_{i}^{N} \sum_{k=0,1} \left[\ln \Phi_2 \left(x_w \boldsymbol{b}_w ,\ x_k \boldsymbol{b}_k ,\ \boldsymbol{r}_{wk}\right)\right]^{d_{wk}} + \sum_{t=1}^{T_i-1} \left[\ln \Phi\left(-x_w \boldsymbol{b}_w\right)\right]^{1-d_{wk}}$$

as the log-likelihood function of the model. In this particular setting, the parameter vector *â*$_w$ reflects the effects of covariates x_w on workers' reemployment rates, whereas the second parameter vector *â*$_k$ represents the effects of covariates x_k on the conditional probability of exiting into jobs of type k=1 instead of jobs of type k=0. Moreover, the parameter *ñ*$_{wk}$ reflects any potential correlation between wage levels k and job exit rates r. If wage outcomes are measured continuously instead, the appropriate log-likelihood function becomes

$$(8) \qquad \ln L = \sum_{i}^{N} \left[\begin{array}{l} \left[\ln \Phi\left(\dfrac{x_w \boldsymbol{b}_w + (Y - x_k \boldsymbol{b}_k)\, \boldsymbol{r}_{wk} / \boldsymbol{s}_Y}{\sqrt{1-\boldsymbol{r}_{wk}^2}} \right) - \dfrac{1}{2}\left(\dfrac{(Y - x_k \boldsymbol{b}_k)}{\boldsymbol{s}_Y} \right) - \ln\left(\sqrt{2\boldsymbol{p}}\boldsymbol{s}_Y\right) \right]^{d_w} + \\ \sum\limits_{t=1}^{T_i-1} \ln \Phi\left(-x_w \boldsymbol{b}_w\right)^{1-d_w} \end{array} \right],$$

which is merely a non-standard application of Heckman's selectivity correction estimator (Heckman 1979). In any event, and very much as in more standard analyses, the covariate vectors x_w and x_k will include measures of workers' skills (education, experience, tenure and earnings in previous job, as well as completion of vocational training for German workers) as well as gender and information on non-white, respectively non-German ethnicity. Duration dependence is accounted for by including a third-order polynomial function in both equations. Additional control variables for the rate equation include occupation and industry dummies, a measure of the quarterly aggregate vacancy ratio, year dummies as well as a 'seam' month variable intended to capture the effects of linking several interview waves into a single event history calendar.[6] Excluding these variables from the wage equations effectively also serves as an identification restriction of the model. The core variable of interest to the current paper is of course the effect of individual UI benefit status on both reemployment rates and post-unemployment wages. In order to allow for the possibility of "selection on unobservables" into UI, the analyses will further include the Inverse Mills' Ratio from a first-stage probit model of UI benefit status. The respective estimation results are given in Appendix 2.[7] But before turning to discuss the results from these more involved event history models, the following section will first give some core descriptive information on unemployment dynamics in the United States and West Germany.

3. Unemployment duration and wage outcomes in the United States and West Germay

At fairly comparable levels of aggregate unemployment rates during the mid-1980s to the mid-1990s, the flows underlying aggregate unemployment have differed considerably between the United States and West Germany (cf. Garibaldi et al. 1997; Machin and Manning 1999). As immediately evident from Table 1 below, unemployment duration figures among German workers have well exceeded comparable U.S. figures throughout the period under study. Typically, median spell durations in the German labor market have been about twice the figures common among U.S. workers. Averaging over the 1984-1995 period, median unemployment spell duration has been 2.3 months among U.S. workers, yet amounted to a full 4.8 months among unemployed workers in West Germany. Also, as shown by further disaggregation, while unemployment duration figures have evolved pro-cyclically in both countries, German duration figures have risen particularly strongly during the recession of the mid-1990s.

6 The 'seam' month is the final month of calendar information gained within any single interview. Several papers on both the SIPP and the GSOEP report artificially increased transition rates in these months as individuals having experienced a certain event during the recall period of the subsequent interview are more likely to date the event back to the start of the recall period.

7 Empirically it turns out that there is little evidence of important selection on unobservables, as UI benefit effect estimates are substantively robust to the inclusion of the Inverse Mills' Ratio (respective results are available from the author on request). This indicates that readily observable worker characteristics are suitable predictors for UI eligibility, so that concerns for self-selection are much less of an issue than e.g. in the case of evaluations of training programs. Of course, this is precisely what would be expected from both the nature of unemployment insurance and the institutional regulations concerning UI eligibility.

Unsurprisingly, differences in reemployment rates are the key component behind these cross-national differences in unemployment duration. In both the United States and West Germany, some 70% of all exits from unemployment are into dependent employment. Only relatively few unemployed workers start up their own businesses, a certain proportion of workers facing difficulties in securing reemployment enter training courses or educational programs, and a sizeable minority of workers is - at least intermittently - withdrawing from the work force. Interestingly enough, there is also considerable cross-country similarity in wage change distributions. In both countries, wage change distributions are highly skewed showing average real wage gains for unemployed workers, yet the median worker experiences reemployment at real wages similar to those in her earlier job. A considerable proportion of workers has to face real wage losses, however, which are quite substantial (>20%) for about one quarter to one fifth of all unemployed workers.[8][9]

Disaggregating these figures by individual UI benefit status provides first descriptive evidence on the potential role of UI in accounting for the above findings on unemployment dynamics. There are two relevant pieces of evidence here: first of all, the bottom row of Table 2 clearly shows the expected and substantial differences in UI coverage between Germany and the U.S. While empirically almost 90% of unemployed workers in the German sample have had access to UI benefits, the respective U.S. figure has been as low as 39%. And secondly, there is also some evidence of differences in unemployment behavior between covered and non-covered workers in both countries. At a purely descriptive level, workers receiving UI benefits tend to experience longer durations of unemployment spells, yet at least in Germany, also show somewhat more favorable wage outcomes than workers not eligible to receive benefits. Moreover, the differences between these two groups of unemployed workers are far from trivial empirically. The median spell duration among workers covered by UI benefits in the U.S. is about 1.5 months longer than among workers without access to UI benefits, and the respective differential among German unemployed amounts to even more than two months. At the same time, workers covered by UI benefits in both countries are somewhat more likely to exit unemployment by taking up paid work rather than by – at least intermittently – withdrawing from the labor market. Conditional on leaving unemployment for work, workers who had received UI benefits during their unemployment spell in Germany also tend to show more favorable wage

[8] Note however, that the wage change measure applied here is a less than ideal measure of scar effects proper which would ideally be assessed against the expected wage from continued employment in workers' previous job. In particular, if seniority gradients are steeper in the U.S. labor market, the observations made here are fully consistent with the standard result of smaller scar effects in Germany (e.g. Burda and Mertens 2001). While imperfect, the available wage measure still allows to identify the causal effect of UI benefits on post-unemployment wages, however.

[9] All tables and figures in the appendix page I - X

outcomes, in particular in terms of avoiding fairly sizable real wage losses. Among covered workers in Germany, the probability of experiencing a 10% real wage cut is some 5 percentage points below the corresponding figure among non-covered workers, and in the case of 20% real wage cuts, the differential amounts to even 8 percentage points. In contrast, U.S. figures do not reveal similarly positive effects of unemployment benefit receipt.

As none of these results has been adjusted for group differences in worker characteristics, the above estimates of course provide only a naïve estimate of actual treatment effects. Given the structure of UI eligibility requirements, it is eventually unsurprising to find workers covered by UI exhibiting higher levels of pre-unemployment work experience, higher levels of tenure with former employers, and higher pre-unemployment wages and earnings (cf. Appendix 1). To the extent that any of these worker characteristics affect unemployment processes, systematic group differences between covered and non-covered workers in terms of background characteristics will naturally bias any causal inferences based on simple descriptive statistics. To discuss UI benefit effects on job histories in a more appropriate econometric framework, I now turn to estimation results obtained for the discrete-time hazard rate models that have been developed in Section 2 above.

Unemployment benefits and unemployment dynamics: hazard rate models

Tables 2 and 3 have the estimation results from a series of discrete-time bivariate probit hazard models that simultaneously address job exit rates and post-unemployment wages among U.S. and West German workers. For each country, six different models have been estimated, each addressing a specific aspect of the post-unemployment wage distribution. These models control for a wide range of covariates, including worker characteristics like gender, ethnicity, education and labor force experience, but also aggregate vacancy ratios and potential trends in workers' reemployment rates. As these covariates primarily serve as control variables in the context of this paper, the respective estimation results will be summarized only briefly here. Also, the results obtained for these variables are mostly standard in the empirical literature (Pedersen and Westergård-Nielsen 1993; Devine and Kiefer 1991; Machin and Manning 1999). In general, reemployment rates are found to exhibit negative duration dependence, i.e. reemployment rates tend to fall over the course of unemployment spells. In terms of macroeconomic effects, reemployment rates are also positively related to aggregate labor market dynamics as captured by quarterly vacancy ratios, and relatively more so among German workers. At the individual level, education, labor force experience, previous earnings levels and vocational training among German workers all contribute to higher rates of reemployment, while tenure with workers' previous employer tends to lower workers' chances to find new jobs. Also, women and non-white, respectively non-German workers face lower reemployment rates in both countries (results not shown).

At the same time, these covariates are also found to affect post-unemployment wages. Most importantly, higher pre-unemployment wages imply higher risks of experiencing post-unemployment wage losses among both U.S. and German workers. Effectively, this finding implies that some important part of workers' earnings capacity does carry over to workers' new jobs, yet the fraction of workers' earnings capacity that is carried over apparently declines with pre-unemployment wage levels. There is also some evidence of a negative relationship between labor market conditions and post-unemployment wages, i.e. unemployed workers tend to experience higher wage losses when there are plenty of vacancies. This finding might come somewhat unexpectedly, yet it might simply testify workers' willingness to accept reemployment at lower wages as long as they face a reasonable chance that the deterioration in earnings capacity will be short-lived. Apart from these effects, there is little consistent evidence on further covariate effects, be it in terms of duration dependence or skills. It seems noticeable, however, that all estimated models provide support of a strongly negative structural relation between work exit rates on the one hand, and job quality on the other: job finding rates tend to increase as workers' relative real wage levels fall. This observation is noticeable insofar as job search models would imply a positive correlation resulting from a truncation from below in the distribution of accepted wage offers. Rather to the contrary, and more consistent with dual labor market theories and segmentation models, the empirical evidence points to dominant rationing effects in the primary, high-skill, high-wage sectors of the economy.

More importantly, however, the estimation results of Tables 2 and 3 also provide unequivocal empirical support for both key effects of unemployment benefits on unemployment dynamics. Across all different specifications of the post-unemployment wage distributions, but also if compared across countries, the hazard rate estimates consistently establish a substantial and statistically significant negative effect of UI benefits on reemployment rates among unemployed workers. At the same time, however, UI benefits clearly tend to raise post-unemployment wages, and in particular tend to limit workers' risks of incurring considerable wage losses upon reemployment. Hence, *receiving UI benefits tends to both lower job-finding rates among unemployed workers and, at the same time, raises the quality of jobs taken on by workers leaving unemployment*. Against some prolongation of unemployment spells, UI benefits thus significantly reduce potential scar effects of unemployment, at least in terms of workers' post-unemployment wages. Basically, this conclusion also seems to be robust in the cross-national comparison between the U.S. and the West German labor market. Even though the evaluation of causal UI effects is more involved in the latter case due to both the considerably smaller sample sizes available in the GSOEP and the more encompassing UI coverage among German workers, the results for UI benefit effects in the GSOEP data are remarkably consistent with those established from the SIPP sources. Again, UI benefits imply lower reemployment rates among unemployed workers, although the magnitude of the difference tends to be smaller than established from the U.S. data. In addition, there is also consistent evidence on positive UI effects on post-unemployment wages; if anything, these effects appear even larger than those established with the U.S. data, however.

Given the non-linear nature of the probit model, the implied magnitudes of UI effects are in fact more easily assessed from the marginal benefit effect estimates given in Table 4. As evident from these results, the implied magnitudes of UI effects are far from trivial empirically. To illustrate these, Table 4 contains information on marginal UI benefit effects calculated at four different points of elapsed spell duration (*T*=1, 3, 6, and 12 months). Among the different quantities provided, benefit effects on *r(t)* and *F(t)* obviously describe the negative impact of UI benefits on unemployment duration. Among both U.S. and German workers, unemployment benefits tend to reduce job finding rates *r(t)* among the unemployed, which in consequence also implies lower cumulated probabilities *F(t)* of work exits from unemployment. Apparently, UI effects on unemployment duration are considerably stronger in the U.S. labor market. UI benefits are estimated to lower work exit rates by some 7 percentage points at the beginning of unemployment spells, which is equivalent to a full 30% reduction in outflow rates absent UI benefit coverage. Naturally, this effect translates into respective reductions in the cumulated probability *F(t)* of having exited unemployment by taking up paid work. According to the model estimates, UI benefit effects imply reductions in *F(t)* of some 13-15 percentage points by spell months 3 to 6 (equivalent to some 20-25% reduction in *F(t)*). The comparable German estimates are smaller in magnitudes, with UI benefits lowering work exit rates *r(t)* by about 2.7 percentage points. As reemployment rates are generally much lower in the German labor market, this amounts to reducing exit rates by about 23%, which leads to lowering *F(t)* by 6-9 percentage points by spell months 3 to 6 (equaling a 20% reduction in *F(t)*).

These negative effects of UI benefits on reemployment rates are of course to be set against positive UI effects on post-unemployment wages. As described earlier, UI benefits tend to raise post-unemployment wages, and in particular consistently lower the probability of experiencing wage losses at exiting unemployment. Evaluating the distribution of accepted wages at T=12 months, i.e. when most workers will already have left unemployment, the implied benefit effects on wages are substantial indeed. Among U.S. workers, benefits on average tend to raise post-unemployment wages by 5.3% of workers pre-unemployment real wage levels. In fact, this estimate accords quite well with the most comparable estimate obtained in Addison and Blackburn's

(2000: Table 4, specification 4) recent paper, yet based on different data sources and a somewhat different econometric approach. Moreover, positive UI effects moreover become more pronounced, the more attention focuses on the lower tail of the wage change distribution, i.e. workers' risks of experiencing (severe) wage losses. Across the different cut-points in the wage distribution, benefits tend to reduce the probability of experiencing a given wage loss by 4-6 percentage points. Given a falling risk of experiencing more severe wage losses, this is of course equivalent to an increasingly stronger protection from more severe wage losses. Proportional to the baseline, benefits lower the risk of any post-unemployment real wage loss by 10%, yet the risk of wage cuts of 30% and more by already more than 20%. If anything, these positive UI effects are even stronger among West German workers. On average, the GSOEP data even yield an estimated +9.3% increase in workers' post-unemployment real wage levels. Again, UI effects are particularly strong if it comes to avoiding relatively severe wage cuts. While the probability of experiencing any real wage loss is reduced by 11 percentage points (a proportional risk reduction of about 18%), the likelihood of experiencing wage cuts of at least 10% is lowered by 19 percentage points, and the one for experiencing a 20% wage loss by still a full 14 percentage points. Access to UI benefits in sum achieves a full 40% reduction in workers' risks of experiencing severe wage losses among German workers.

Benefit effects over spell duration and pre-unemployment earnings levels

This evidence of considerable UI effects on both reemployment rates and post-unemployment wages immediately begs further questions. Of many potential concerns, two particular issues will be taken up here. The first of these relates to the relationship between UI effects and spell duration, which embeds concerns like whether UI benefits contribute to perpetuate long-term unemployment as the long-term unemployed may fail to properly adjust wage expectations downward. This may even be reinforced if workers accurately perceive UI as a search subsidy leading to more adequate wage outcomes, yet if these positive UI effects wear off with increasing spell duration. Crosscutting this issue of changing UI effects over spell duration is the nature of the trade-off between prolonged search and improved wage outcomes in different skill groups in the labor market. While the models fitted above might yield the expected trade-off relationship in the aggregate sample, this is in itself no guarantee that this trade-off holds to similar degrees among both low- and high-wage workers. In fact, it well be that UI acts as a search subsidy among high-wage workers who have human capital to preserve through unemployment periods, while pure disincentive effects prevail at the level of low-wage earners. Sustaining the human capital interpretation of the UI effects at any rate requires that UI effects increase among high-wage workers: if anywhere, it is among high-skill workers that UI should unfold its positive effects on wages, and it is also conceivable that search for adequate reemployment should take relatively more time in the high-skill segment.

To address both issues, it is necessary to augment the hazard rate models fitted before by including interaction terms between benefits and spell duration, respectively benefits and pre-unemployment earnings levels. These interaction terms have been included in both the rate and the wage equations of the hazard models, and three different models have been fitted subsequently (with wage change, any wage loss, and wage losses > 20% as the dependent variables in the wage equation). The empirical estimates for these extended models are reported in Table 5 for both countries, with the reported parameter estimates being restricted to the additional interaction terms in the models. For all models, likelihood-ratio tests indicate a statistically significant improvement in model fit by including these interaction terms, although the evidence for the smaller GSOEP database seems less robust than for the U.S. data. Despite imprecise point estimates in the German data and against the resulting reservations about the latter, the results appear mutually consistent in substantive terms for both the U.S. and the West German data, however. For both samples, the parameter estimates show that both negative UI effects on reemployment rates and positive effects on post-unemployment wages decline with spell duration. In turn, and very much in line with the notion of human capital preservation, UI effects increase with pre-unemployment earnings.

The implied effect magnitudes are again more easily assessed from transforming the probit parameters back into the underlying quantities of the models. Figures 1 and 2 provide a graphical illustration of the relationships between marginal benefit effects and spell duration, respectively pre-unemployment earnings. Given that the estimation results are very consistent across different specifications of wage outcomes, both figures illustrate the structure of the interactions for the wage loss model only. With the exception of the model for severe wage losses in Germany, the resulting patterns are qualitatively very similar for the other specifications; the detailed results are of course available from the author on request. Returning to the figures themselves, Figure 1 first clearly shows that whatever effects UI unfolds on unemployment dynamics, these are confined to about the first nine to twelve months in unemployment. This is of course not to say that UI effects are irrelevant empirically: rather to the contrary, given that the overwhelming majority of unemployment spells will be completed by that time, UI does crucially affect unemployment spells and the associated wage outcomes for the bulk of the total inflow into unemployment.[10] The results for the interactions of

[10] Averaging over the 1984-1995 period, some 80% of all unemployment spells among U.S. workers lasted six months or less, some 90% of all spells were completed by nine months, and only about 5% of all unemployment spells continued for 12 months or more. Longer unemployment durations in

benefit effects and spell duration do imply, however, that benefit effects are highly unlikely to account for the persistence of unemployment among the long-term unemployed. Conditional on having spent already one year unemployed, there is little evidence of any remaining UI effects on reemployment rates, i.e. those long-term unemployed supported by benefits are no less likely to take up paid work than those without UI benefits.

At the same time as negative UI effects on reemployment rates decline over spell duration, it is also true that positive UI effects on wages tend to be more pronounced for early job exits out of unemployment. Obviously, UI protection cannot provide ultimate guarantees against wage losses after fairly long duration of unemployment, and it certainly seems to be the case that sustained selective job search does no longer produce favorable job outcomes among the long-term unemployed. The parameter estimates in fact also suggest stronger duration dependence in positive UI effects on wages as compared to duration dependence in UI effects on reemployment rates, i.e. positive UI effects on wages decline relatively more quickly than negative UI effects on reemployment rates. In conjunction with the earlier finding of some negative duration dependence in wage outcomes, this result does suggest the necessity for the unemployed to continuously revise job expectations downwards during the course of unemployment spells. And if anything, there seems to be some evidence that UI benefits indeed slows down this adjustment of expectations as UI wage effects indeed decline more quickly than negative UI effects on reemployment rates. Against the empirical structure of unemployment dynamics, the quantitative impact of this effect is probably minor on the other hand. Given that most workers exit unemployment quickly, the cumulative wage distributions in any event continue to show clear positive benefit effects among both U.S. and German workers.

Against this evidence that UI effects are unlikely to perpetuate long-term unemployment, Figure 2 plots evidence on UI effects across skill groups defined by pre-unemployment earnings levels that is quite consistent with UI protecting and maintaining workers' accumulated human capital through unemployment spells. In general, UI effects on both reemployment rates and wages increase with workers' pre-unemployment earnings. Thus, both positive and negative UI effects occur markedly among high-skill workers, thus clearly supporting the search subsidy interpretation of benefit effects, and also stressing the trade-off relationship between improved wage outcomes on the one hand and prolonged spell duration on the other. According to the results here, it is certainly not the case that UI acts as a work disincentive among low-wage earners, yet has the search subsidy effect among high-wage workers. Rather, UI effects tend to be fairly small among the low-wage group in general. If anything, it seems to be the

West Germany are reflected in the fact that only some 60% of all unemployment spells are completed within six months, about 70% within nine months, and about 80% within 12 months.

case that positive UI effects on wages are particularly pronounced among low- to mid-wage workers in West Germany against a relatively small reduction in reemployment rates.

Recall or Benefit Effects?

There is one final important reservation against the robust evidence of positive UI effects on post-unemployment wages obtained here, and about its correct theoretical reading in particular. While the evidence of positive UI effects seems to be robust econometrically, it is far from clear that these positive effects indeed result from the reservation wage mechanism described by search models. Rather, recalls might act as an important alternative mechanism to the supposed improvement of matching processes achieved by UI. In particular, it might be argued that job exits through recalls potentially imply substantial positive biases in the UI effect estimates reported before, so that true UI effects on matching might be much lower than those reported so far. Ending unemployment spells through recalls is indeed quite common in both the United States (Katz and Meyer 1990b), but also in the West German labor market (Mavromaras and Rudolph 1998). Given that workers' firm-specific human capital is retained, recalls should be associated with relatively positive wage outcomes at exiting unemployment. In addition, if employers are more likely to consider recalls of more experienced workers who might also have better access to UI benefits during unemployment spells, there might be a positive correlation between workers' UI eligibility and workers' probability of being recalled later on. This correlation, in turn, would imply positive biases in estimates of UI effects on post-unemployment wages. Hence, any assertions that UI benefits are effective in protecting workers' current stock of human capital would want to be tested against the effects of recalls.

In contrast to the GSOEP, the SIPP data fortunately provide sufficient information to establish whether given job exits have occurred by returning to workers' former employers or by entering job matches with new employers. According to the SIPP data, slightly less than 30% of all job exits observed in the sample were recalls to workers' former employers, with some anti-cyclical trend in this figure. More importantly, however, the data also do point to a positive correlation between UI benefit receipt and the probability of recall: among workers receiving unemployment benefits, a full 37.7% of all job exits were to workers' former employers. Among workers without benefits, the respective proportion amounted to a mere 24.5% of all job exits.

Yet do these data imply that positive benefit effects on post-unemployment wages are mostly a function of recalls, or are the positive wage effects of UI due to improving matching processes more generally? To test for either possibility, Table 6 contains estimation results for another series of rate models that include recall information in the covariate vector.[11] Given that the full estimation results are substantively quite similar to those presented earlier, Table 6 contains only the estimates for the two key institutional covariates of interest, the effects of recalls and the effects of unemployment benefits. With respect to the former, the results closely conform to the expectations developed above: recalls both do raise reemployment rates among unemployed workers, and tend to considerably improve post-unemployment wages. In particular, recalls show strong effects on limiting the incidence of wage losses upon reemployment; if anything, the associated effect sizes even exceed the positive UI effects established earlier. Against this background, the estimation results also continue to show robust and strong evidence of positive UI effects on wages, however. Controlling for the effects of recall, unemployment benefits continue to have both an important and statistically significant negative effect on reemployment rates and a substantial positive effect on workers' post-unemployment wages. Compared to the parameter estimates of Table 3 above, controlling for the effects of recalls reduces the size of positive UI effects on wages in the order of around 10%. Consequently, the positive UI effects on wages do not mainly result from a positive correlation between firm-specific skills and UI benefit receipt, but from UI achieving a genuine protection of workers' current stock of human capital. As asserted by Burdett (1979), UI thus mainly acts as a search subsidy to workers that results in considerably improved job matches among covered workers.

[11] As the incidence of recall is of course endogenous to the rate equation, recall effects on reemployment rates are controlled by the time-varying control function for the probability of recall. Recall predictions are based on a probit model that includes gender, ethnicity, education, labor force experience, tenure and earnings with previous employer as well as occupation and industry dummies as covariates. The model achieves a Pseudo-R^2 of 4.7%; full estimation results are available from the author.

4. Summary and conclusions

In their recent analysis, Addison and Blackburn (2000) perceived their own econometric evidence as too weak as to allow a definitive judgment about positive effects of unemployment insurance on workers' post-unemployment wages. This does not seem to be the case for the empirical evidence obtained in the current paper: to the contrary, the analyses conducted here do obtain statistically robust evidence of positive UI effects on wages, which is both consistent for both U.S. and West German data and indicative of fairly substantial effect sizes. For the U.S. data, at any rate, the estimated UI effects suggesting an average increase in post-unemployment wages by 5.3% of workers' pre-unemployment wage levels, are certainly at the upper end of those obtained by Addison and Blackburn (2000). Compared to the U.S. data, the evidence for West Germany points to even stronger positive effects, however, with point estimates in the order of a +9% increase in workers' post-unemployment wages. In contrast to Addison and Blackburn (2000), I also obtain consistent evidence that UI effects occur more strongly in the lower tail of the wage change distribution, and among high-skill workers.

In sum, the empirical evidence presented in this paper would strongly seem to support the search theory interpretation of UI effects on unemployment dynamics. Consistent with many earlier studies, there is evidence of a substantial negative UI effect on reemployment rates, yet this effect is counterbalanced by equally strong positive UI effects on post-unemployment wages. Consistent with Burdett's (1979) early notion, unemployment insurance thus mainly acts as a search subsidy to workers implying relatively more favorable job outcomes through sustaining costly job searches for relatively more adequate employment. Acting as a credit slip to unemployed workers, unemployment insurance might thus be seen as an important institutional device for preserving workers' accumulated human capital through unemployment spells. This reading of the evidence is additionally supported by the finding that UI particular acts to avoid (severe) wage losses upon reemployment rather than generating a more uniform upward shift in the wage outcome distribution: very much as implied by basic job search theory, UI raises workers' truncation point of the wage distribution rather than shifting the whole distribution upwards. The human capital preservation aspect stressed here is also consistent with the fact that more pronounced UI effects have been found among high-wage workers. Also, the analyses have stressed that UI indeed tends to preserve workers' general human capital – restoration of firm-specific human capital through recalls has been found to leave UI effect estimates virtually unaffected, and hence the primary effect of UI apparently is to improve actual job matching rather than anything else.

As far as it goes, the results obtained here thus emphasize unemployment benefit provision as one major institutional means of alleviating, or at least dampening the scar effects of unemployment: unemployment benefits do lower the opportunity cost of job search, and hence subsidize workers' job search, and this sustained job search also does show positive payoffs in terms of higher wages. Stated as simple as that, the importance of unemployment insurance for limiting scar effects might be even undervalued: As Arulampalam's (2001) recent results have pointed out, it is the quality of the first post-unemployment job match that is decisive for permanent scar effects of unemployment – empirically, workers apparently have considerable difficulties in making up for initial bad matches in later careers. Effectively, the current analysis has been exclusively confined to wage outcomes in workers' first post-unemployment jobs, and substantial effects of UI on precisely this first post-unemployment match have been documented. Hence, UI might legitimately come to be seen as having considerable positive long-run effects on careers, but potentially even on the macroeconomy more generally (cf. Acemoglu 2001). Certainly, more research to show whether and to which extent initial positive UI effects on wages become dissipated or are maintained over time is direly needed. However, the magnitudes of the effects established in this analysis might even make the positive wage effects that were required to render the associated prolongation of unemployment cost-effective surprisingly small. If UI is associated with an increase of unemployment duration by 0.5-1 month on average, sustaining a 5% gross wage gain for about a single year would be fully sufficient to have a self-financed UI search subsidy at the going replacement rates in the current U.S. or German UI systems.

In addition to these results, there are some apparent differences in UI effects between the U.S. and the West German labor market that would seem to warrant closer further study. If anything, the empirical results of this paper suggest even stronger positive UI effects against smaller negative duration effects in West Germany. In any event, the latter effect is fully consistent with basic job search theory that stipulates benefit effects to increase in more dynamic labor markets. More troubling for the theory, and hence more interesting as a research question, is the finding of stronger positive wage effects in Germany, and in particular the marked effects at relatively low levels of pre-unemployment earnings. Before offering possible conclusions on the issue, replications of the finding with different data would certainly be called for. If substantiated, one then might assess the reasons behind this finding. While there are many possible explanations, it would seem important to control for the effects of a fairly effective public employment service in Germany, which could both offer a decisive advantage to low- to intermediately skilled workers and would certainly be correlated with actual benefit status. Naturally, however, only future research will be able to resolve the issue.

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Table 1
Unemployment duration and post-unemployment wage outcomes, 1984-1995

	United States			West Germany		
	all workers	with UI benefits	without UI benefits	all workers	with UI benefits	without UI benefits
Median spell duration (months)	2.31	3.37	1.78	4.80	5.10	2.80
% work exits	0.695	0.739	0.672	0.708	0.712	0.679
Mean wage change	1.210	1.166	1.236	1.148	1.130	1.356
Median wage change	1.000	1.000	1.000	1.001	0.998	1.088
Prob(wage loss)	0.449	0.472	0.436	0.498	0.506	0.412
Prob(wage loss > 10%)	0.344	0.358	0.335	0.310	0.306	0.360
Prob(wage loss > 20%)	0.257	0.275	0.247	0.232	0.226	0.302
Prob(wage loss > 30%)	0.197	0.211	0.189	0.177	0.172	0.232
Prob(wage loss > 50%)	0.123	0.129	0.120	0.078	0.076	0.103
% workers with UI benefits	0.389			0.892		

Source: Survey of Income and Program Participation, Panels 1984, 1986, 1988, 1990, 1992, and 1993; German Socio-Economic Panel, 1984-1995 data (Waves A-M), weighted data.

Table 2
Hazard rate models estimates, U.S. data

	Job exit rate		Wage change					
			Log wage change	Any wage loss	Wage loss > 10%	Wage loss > 20%	Wage loss > 30%	Wage loss > 50%
Intercept	-1.763	(.123)**	-2.526 (.156)**	-2.881 (.212)**	-3.344 (.212)**	-3.728 (.228)**	-4.000 (.228)**	4.306 (.267)**
Unemployment benefits	-0.250	(.013)**	0.208 (.018)**	-0.246 (.025)**	-0.291 (.024)**	-0.282 (.022)**	-0.304 (.022)**	-0.310 (.026)**
IMR Benefit Status	0.074	(.033)**	0.054 (.041)	-0.177 (.058)**	-0.213 (.059)**	-0.213 (.057)**	-0.201 (.058)**	-0.299 (.070)**
T	-0.033	(.003)**	0.006 (.005)	0.015 (.011)	0.021 (.008)**	0.016 (.009)*	0.010 (.007)	0.005 (.008)
T^2 (x 100)	0.033	(.012)**	0.019 (.026)	-0.127 (.097)	-0.133 (.049)**	-0.135 (.047)**	-0.112 (.044)**	-0.079 (.051)
T^3 (x 10,000)	-0.103	(.078)	-0.311 (.223)	2.350 (2.28)	1.390 (.537)**	1.460 (.557)**	1.260 (.511)**	1.010 (.589)*
Years of education	0.026	(.003)**	-0.006 (.004)	0.023 (.006)	0.035 (.006)**	0.039 (.006)**	0.039 (.006)**	0.049 (.007)**
Labor force experience	0.010	(.002)**	$-3.7e^{-4}$ (.003)	0.002 (.004)	$1.2e^{-4}$ (.004)	$-7.7e^{-5}$ (.004)**	0.002 (.004)	-0.003 (.004)
Labor force experience² (x 100)	-0.027	(.005)**	0.007 (.006)	-0.019 (.009)**	-0.013 (.009)	-0.014 (.009)	-0.016 (.008)*	-0.007 (.010)
Tenure in previous job	-0.001	($3e^{-4}$)**	-0.001 ($4e^{-4}$)**	0.001 ($6e^{-4}$)**	0.001 ($6e^{-4}$)	0.001 ($6e^{-4}$)*	0.001 ($5e^{-4}$)**	0.001 ($6e^{-4}$)
Tenure in previous job² (x 100)	$4.6e^{-5}$	($9e^{-5}$)	$-3.0e^{-4}$ ($1e^{-4}$)**	$-3.0e^{-4}$ ($2e^{-4}$)	$-1.6e^{-4}$ ($2e^{-4}$)	$-2.4e^{-4}$ ($2e^{-4}$)	$-4.0e^{-4}$ ($2e^{-4}$)**	$1.6e^{-4}$ ($2e^{-4}$)
Ln(previous earnings)	0.062	(.012)**	-0.234 (.016)**	0.260 (.024)**	0.267 (.026)**	0.268 (.028)**	0.279 (.026)**	0.296 (.031)**
Quarterly vacancy ratio	0.042	(.005)**	-0.024 (.007)**	0.026 (.010)**	0.033 (.009)**	0.040 (.029)**	0.032 (.009)**	0.031 (.011)**
ρ_{12}			-0.744 (.021)**	0.617 (.070)**	0.675 (.071)**	0.805 (.075)**	0.888 (.043)**	0.838 (.059)**
Log-likelihood			-42,941	-37,874	-37,388	-36,614	-35,774	-34,458
Log-likelihood null model				-39,222	-38,784	-38,057	-37,237	-35,848

Notes: N = 69,480 observation months (11,550 observed work exits). Standard errors in parentheses; statistical significance levels given at **p<.05, and *p<.10. As additional controls, all models include gender, ethnicity, annual dummies as well as a seam month variable in the rate equation. The rate model estimates given are those obtained in the model for any wage losses.

Source: Survey of Income and Program Participation, Panels 1984, 1986, 1988, 1990, 1992, and 1993.

Table 3
Hazard rate models estimates, West German data

	Job exit rate		Wage change					
			Log wage change	Any wage loss	Wage loss > 10%	Wage loss > 20%	Wage loss > 30%	Wage loss > 50%
Intercept	-2.436	(.681)**	4.094 (.585)**	-5.629 (1.28)**	-5.334 (1.42)**	-5.023 (1.34)**	-5.489 (1.13)**	-6.540 (1.74)**
Unemployment benefits	-0.163	(.085)*	0.188 (.071)**	-0.343 (.169)**	-0.560 (.190)**	-0.461 (.181)**	-0.285 (.146)**	-0.374 (.216)*
IMR Benefit Status	-0.266	(.384)	0.491 (.325)	-1.253 (.895)	-1.671 (1.05)	-1.030 (.978)	-0.315 (.780)	-0.271 (1.26)
T	-0.034	(.008)**	0.018 (.008)**	0.029 (.034)	0.028 (.029)	0.012 (.026)	-0.003 (.015)	0.020 (.027)
T^2 (x 100)	0.049	(.034)	-0.064 (.034)*	-0.193 (.210)	0.007 (.100)	0.030 (.087)	0.024 (.061)	-0.058 (.103)
T^3 (x 10,000)	-0.331	(.340)	0.613 (.342)*	3.500 (3.60)	-0.780 (1.13)	-0.752 (.975)	0.343 (.611)	0.392 (1.02)
Years of education	0.028	(.014)**	0.013 (.010)	-0.014 (.023)	0.003 (.025)	0.008 (.023)	0.013 (.020)	0.008 (.029)
Vocational training	0.050	(.047)	-0.053 (.037)	0.037 (.089)	0.015 (.099)	0.030 (.093)	-0.003 (.076)	-0.047 (.123)
Labor force experience	0.006	(.008)	0.010 (.008)	-0.030 (.021)	-0.030 (.023)	-0.017 (.022)	0.007 (.018)	0.015 (.029)
Labor force experience^2 (x 100)	-0.059	(.018)**	-0.002 (.018)	0.046 (.050)	0.046 (.055)	0.009 (.052)	-0.059 (.040)	-0.080 (.070)
Tenure in previous job	-0.004	(.001)**	0.002 (7e^{-4})**	0.002 (.002)	-0.003 (.002)*	-0.002 (.002)	-0.004 (.002)**	-0.001 (.005)
Tenure in previous job^2 (x 100)	5.5e^{-4}	(2e^{-4})**	-3.8e^{-4} (2e^{-4})**	1.4e^{-4} (5e^{-4})	-4.7e^{-4} (6e^{-4})	5.3e^{-4} (.001)	-1.8e^{-4} (8e^{-4})	-0.003 (.004)
Ln(previous earnings)	0.088	(.080)	-0.463 (.067)**	0.682 (.159)**	0.588 (.169)**	0.465 (.158)**	0.426 (.135)**	0.501 (.202)**
Quarterly vacancy ratio	0.634	(.078)**	-0.247 (.071)**	0.227 (.187)**	0.331 (.202)*	0.327 (.200)*	0.513 (.140)**	0.650 (.224)**
ρ_{12}			-0.910 (.022)**	0.702 (.173)**	0.658 (.228)**	0.785 (.208)**	1.000 (3e^{-6})**	0.910 (.154)**
Log-likelihood			-2,707	-2,873	-2,816	-2,766	-2,707	-2,578
Log-likelihood null model				-3,307	-3,251	-3,190	-3,130	-2,987

Notes: N = 15,665 observation months (690 observed work exits). Standard errors in parentheses; statistical significance levels given at $^{**}p<.05$, and $^{*}p<.10$. As additional controls, all models include gender, ethnicity, annual dummies as well as a seam month variable in the rate equation. The rate model estimates given are those obtained in the model for any wage losses.

Source: German Socio-Economic Panel, 1984-1995 data (Waves A-M).

Table 4
Estimated marginal benefit effects, by spell duration

T (months)	1	3	6	12
		United States		
Δ $r_W(t)$	-0.067 (-29.6)	-0.063 (-30.5)	-0.058 (-31.6)	-0.049 (-33.8)
Δ $F_W(t)$	-0.067 (-29.6)	-0.130 (-25.0)	-0.146 (-19.5)	-0.113 (-12.4)
Δ wage change | work exit	+0.061 (+6.7)	+0.060 (+6.6)	+0.057 (+6.4)	+0.053 (-6.0)
Δ Pr(wage loss | work exit)	-0.055 (-12.0)	-0.055 (-11.6)	-0.053 (-11.1)	-0.050 (-10.2)
Δ Pr(wage loss > 10% | work exit)	-0.067 (-18.7)	-0.067 (-18.1)	-0.066 (-17.3)	-0.063 (-15.9)
Δ Pr(wage loss > 20% | work exit)	-0.053 (-19.3)	-0.054 (-18.6)	-0.053 (-17.7)	-0.050 (-16.2)
Δ Pr(wage loss > 30% | work exit)	-0.053 (-24.2)	-0.054 (-23.6)	-0.054 (-22.6)	-0.052 (-21.0)
Δ Pr(wage loss > 50% | work exit)	-0.041 (-28.2)	-0.042 (-27.6)	-0.042 (-26.9)	-0.041 (-25.6)
		West Germany		
Δ $r_W(t)$	-0.027 (-23.2)	-0.025 (-23.8)	-0.023 (-24.6)	-0.018 (-26.0)
Δ $F_W(t)$	-0.027 (-23.2)	-0.064 (-21.4)	-0.092 (-19.1)	-0.108 (-15.8)
Δ wage change | work exit	+0.095 (+10.3)	+0.095 (+10.3)	+0.094 (+10.3)	+0.093 (+10.1)
Δ Pr(wage loss | work exit)	-0.113 (-20.7)	-0.112 (-20.0)	-0.111 (-19.1)	-0.108 (-17.9)
Δ Pr(wage loss > 10% | work exit)	-0.186 (-45.7)	-0.189 (-44.6)	-0.193 (-43.0)	-0.194 (-40.2)
Δ Pr(wage loss > 20% | work exit)	-0.126 (-47.5)	-0.131 (-46.7)	-0.135 (-45.6)	-0.140 (-43.2)
Δ Pr(wage loss > 30% | work exit)	-0.052 (-34.0)	-0.054 (-33.7)	-0.057 (-33.2)	-0.060 (-32.1)
Δ Pr(wage loss > 50% | work exit)	-0.032 (-52.4)	-0.034 (-51.8)	-0.038 (-50.9)	-0.043 (-48.9)

Notes: Average discrete-change effects of UI benefit status on unemployment dynamics in the estimation samples; proportional marginal change effects in parantheses; weighted data.

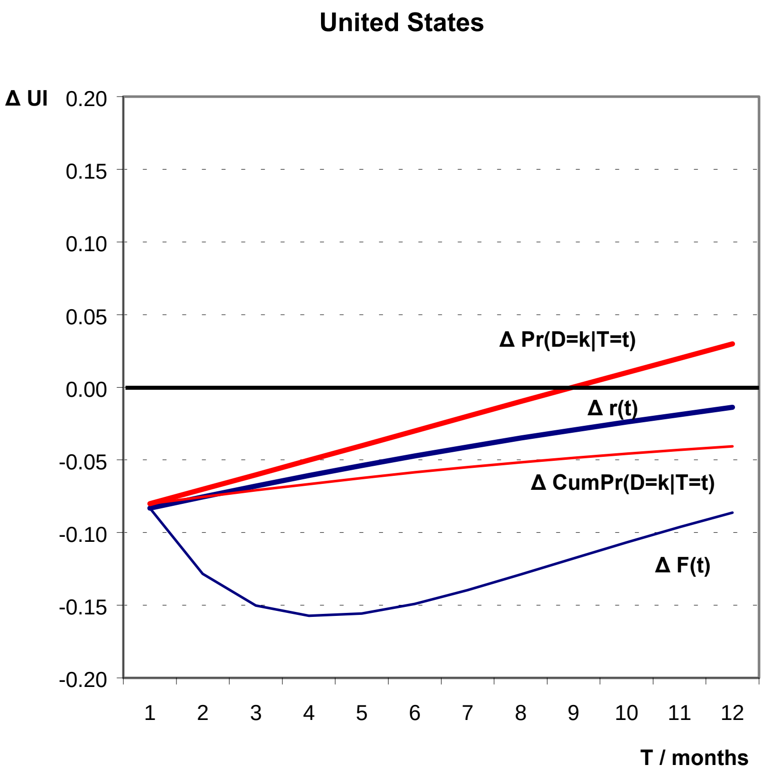
Table 5
Hazard rate model estimates, models including benefit interaction terms

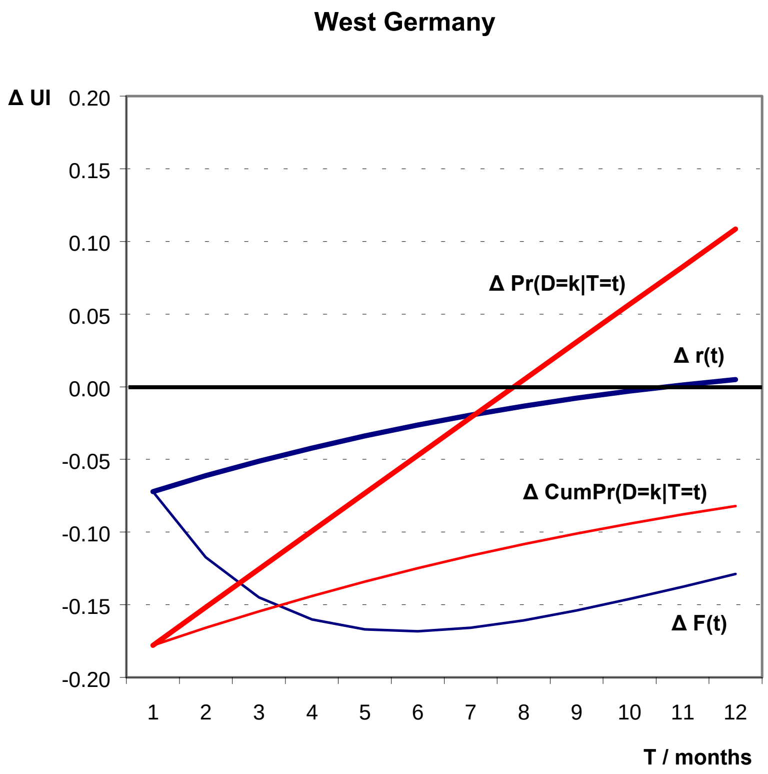
	Job exit rate	Wage change		
		Log wage change	Any wage loss	Wage loss > 20%
		United States		
Unemployment benefits				
x spell duration	0.022 (.003)**	-0.023 (.004)**	0.033 (.006)**	0.035 (.006)**
x log previous earnings	-0.087 (.015)**	0.045 (.020)**	-0.098 (.029)**	-0.128 (.029)**
Log-likelihood		-42,880	-37,810	-36,546
LR-Test χ^2 (df) against baseline		122 (4)**	128 (4)**	136 (4)**
		West Germany		
Unemployment benefits				
x spell duration	0.040 (.021)**	-0.015 (.022)	0.081 (.058)	0.078 (.057)
x log previous earnings	-0.195 (.137)	0.010 (.120)	-0.105 (.370)	0.073 (.349)
Log-likelihood		-2,703	-2,868	-2,761
LR-Test χ^2 (df) against baseline		8.9 (4)*	9.9 (4)**	10.0 (4)**

Notes: Standard errors in parentheses; statistical significance levels given at **p<.05, and *p<.10. Apart from the interaction effects shown, covariate vectors are identical to those given in Table 2 and 3. The rate model estimates given are those obtained in the model for any wage losses.

Source: Survey of Income and Program Participation, Panels 1984, 1986, 1988, 1990, 1992, and 1993; German Socio-Economic Panel, 1984-1995 data (Waves A-M).

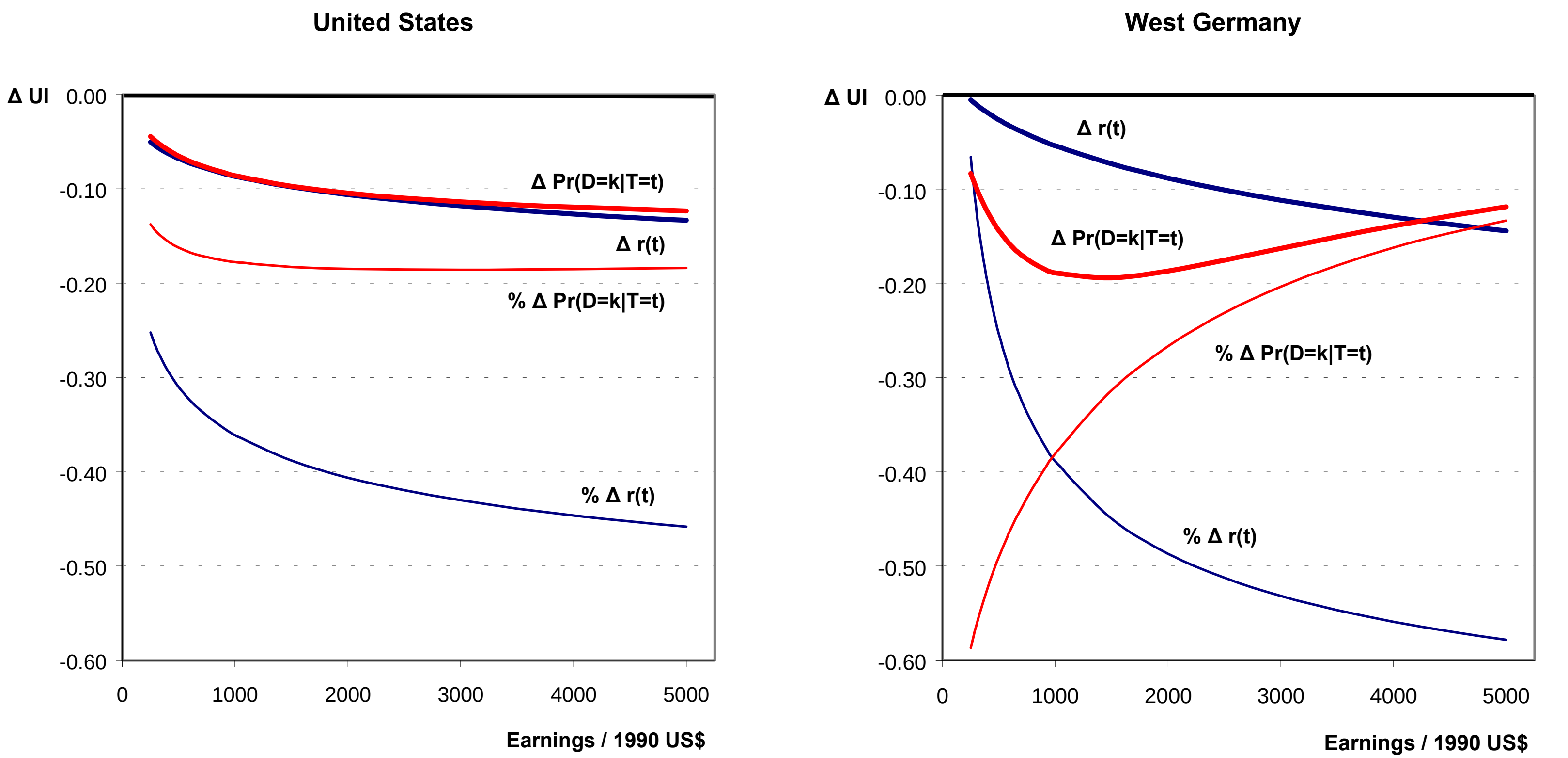
Figure 1
Marginal benefit effects by spell duration, any wage loss model





Notes: Marginal benefit effect calculations based on augmented hazard models given in Table 5.

Figure 2
Marginal benefit effects in first spell months by previous earnings, any wage loss model



Notes: Marginal benefit effect calculations based on augmented hazard models given in Table 5.

Table 6
Benefit effects and recalls, U.S. estimates

	Unemployment Benefits		Recall[1)]	
	b	(s.e.)	b	(s.e.)
Job exit rate	-0.247	(.013)**	0.253	(.047)**
Job quality				
- wage change	0.197	(.018)**	0.091	(.016)**
- any wage losses	-0.185	(.030)**	-0.460	(.032)**
- wage losses > 10%	-0.236	(.030)**	-0.494	(.034)**
- wage losses > 20%	-0.228	(.031)**	-0.478	(.038)**
- wage losses > 30%	-0.275	(.029)**	-0.376	(.041)**
- wage losses > 50%	-0.282	(.034)**	-0.281	(.040)**

Notes: N = 69,480 observation months (11,550 observed work exits). Standard errors in parentheses; statistical significance levels given at **p<.05, and *p<.10. As additional controls, all models include gender, ethnicity, annual dummies as well as a seam month variable in the rate equation. The rate model estimates given are those obtained in the model for any wage losses. [1)] IV estimate in the rate equation.

Source: Survey of Income and Program Participation, Panels 1984, 1986, 1988, 1990, 1992, and 1993.

Appendix 1
Summary statistics for the estimation samples, spell data

	United States			West Germany		
	All workers	with UI benefits	without UI benefits	All workers	with UI benefits	without UI benefits
Women	0.400 (0.490)	0.357 (0.479)	0.422 (0.494)	0.385 (0.487)	0.382 (0.486)	0.412 (0.493)
Non-white / Non-German	0.184 (0.387)	0.152 (0.359)	0.189 (0.391)	0.118 (0.322)	0.116 (0.320)	0.135 (0.342)
Age	33.05 (11.68)	36.96 (11.23)	30.35 (11.11)	34.99 (12.21)	35.62 (12.32)	29.75 (9.84)
Years of education	12.52 (1.89)	12.60 (1.88)	12.45 (1.90)	10.85 (2.12)	10.82 (2.07)	11.11 (2.56)
Vocational training	-	-	-	0.607 (0.489)	0.628 (0.484)	0.435 (0.496)
Labor force experience (years)	12.76 (11.31)	16.60 (11.38)	10.32 (10.56)	16.02 (12.38)	16.82 (12.48)	9.43 (9.26)
Tenure in previous job (months)	20.56 (52.51)	32.79 (65.93)	12.78 (39.86)	51.38 (93.25)	54.25 (95.87)	27.67 (63.06)
Earnings in previous job (1990 US-$, PPP-adjusted)	1141.05 (1131.41)	1510.13 (1280.90)	906.54 (953.69)	1554.83 (758.69)	1585.90 (731.61)	1279.54 (922.96)
Vacancy ratio (quarterly)	3.072 (1.924)	2.966 (1.874)	3.140 (1.953)	0.794 (0.247)	0.796 (0.246)	0.760 (0.204)
Unemployment benefits	0.389 (0.487)			0.892 (0.317)		
N spells (unweighted)	24,100	8,941	15,159	3,251	2,856	395

Notes: Standard deviations in parentheses.
Source: Survey of Income and Program Participation, Panels 1984, 1986, 1988, 1990, 1992, and 1993; German Socio-Economic Panel, 1984-1995 data (Waves A-M); weighted data.

Appendix 2
UI benefit coverage in the United States and West Germany, probit models

	United States	West Germany
Intercept	-4.194 (.052)*	-3.983 (.204)*
Tenure in previous job	0.008 ($3e^{-4}$)*	0.004 (.001)*
Tenure²	$-4.3e^{-4}$ ($3e^{-6}$)*	$-1.5e^{-5}$ ($6e^{-6}$)*
Tenure³	$6.0e^{-8}$ ($5e^{-9}$)*	$1.7e^{-8}$ ($1e^{-8}$)*
Ln(Previous Earnings)	0.420 (.006)*	0.636 (.028)*
Women	-0.063 (.046)	0.088 (.109)
Non-White/German	-0.250 (.012)*	0.070 (.030)*
Age	0.121 (.005)*	0.053 (.017)*
Age²	-0.005 ($3e^{-4}$)*	-0.003 (.001)*
Age³	$5.3e^{-5}$ ($4e^{-6}$)*	$4.4e^{-5}$ ($1e^{-5}$)*
Women x Age	0.008 (.008)	0.022 (.023)
Women x Age²	$-3.9e^{-4}$ ($4e^{-4}$)	-0.001 (.001)
Women x Age³	$8.0e^{-6}$ ($6e^{-6}$)	$2.1e^{-5}$ ($2e^{-5}$)
Education		
- High School / Vocational Training	0.099 (.012)*	0.065 (.031)*
- Some College / *Abitur*	0.073 (.014)*	0.056 (.077)
- Bachelor's degree	-0.089 (.019)*	-
- Master's degree / University	-0.333 (.026)*	-0.355 (.065)*
Labor Force Experience	0.022 (.002)*	0.050 (.006)*
Labor Force Experience²	$-3.2e^{-4}$ ($4e^{-5}$)*	-0.001 ($1e^{-4}$)*
Log-likelihood	-49,360	-5,267
LR-Test (df)	20,821 (32)*	2,419 (31)*
Pseudo R²	0.174	0.187
N	86,915	27,135

Notes: Standard errors in parantheses; statistical significance at *p<.05. The models include seasonal and year dummies as additional controls.
Source: Survey of Income and Program Participation, Panels 1984, 1986, 1988, 1990, 1992, and 1993; German Socio-Economic Panel, 1984-1995 data (Waves A-M).

Bücher

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