1 Introduction

Germany[1], as most other European countries, has been plagued by a persistently high level of long–term unemployment since the early 1980's. In contrast, long–term unemployment is much less of a problem in the United States. One potential reason for the different structure of unemployment relates to institutional differences in unemployment compensation systems. The German system is characterized by relatively high income–replacement ratios and extended benefit–entitlement periods which are often followed by the availability of open–ended unemployment or social assistance schemes. In contrast, unemployment insurance payments in the United States are terminated after about six months as a rule, and entitlement to subsequent welfare payments are only available to lone mothers and the disabled. Can these institutional differences explain the higher level of long–term unemployment in Germany relative to the United States?

For the early 1980's Burtless (1987) has calculated the potential unemployment effects of more genereous benefit systems in some European countries relative to the United States. Applying elasticity estimates from two well–known microeconometric U.S. studies [Moffitt and Nicholson (1982), Moffitt (1985)] to the unemployment compensationt systems in these countries, he found that differences in their generosity have contributed only little to the observed differences in unemployment rates between countries. In particular, he estimated that the longer entitlement period in West Germany relative to the U.S. could have added 2.5 to 4 weeks to the average completed spell of unemployment in Germany in the early 1980's [Burtless (1987: 147)]. From this estimate he concluded that only a modest part of the German–U.S. unemployment differential could be attributed to the more generous German unemployment compensation system.

Since then, there have been some important changes in benefit regulations in Germany. In particular, maximum entitlement periods for older workers with long previous employment histories have been extended considerably in several steps. For example, the maximum entitlement period for those aged 54 years and older was successively extended from 12 months before 1985 to 32 months after July 1987. At the same time, long–term unemployment among older workers has increased dramatically, while it has declined among the younger unemployed during the economic upswing in the second half of the eighties. Although it seems likely that the extension of benefit–entitlement periods and the increase of long–term unemployed among older workers are related, the empirical evidence supporting this hypothesis for Germany seems far from conclusive. Recently, there have been some attempts

1 Here, and in the following, "Germany" always refers to the former West Germany before unification.

to estimate the effect of the unemployment compensation system on the duration of unemployment in Germany [Hujer, Löwenbein and Schneider (1990), Wurzel (1990, 1993), Steiner (1994), Hunt (1995), and Hujer and Schneider (1996)]. Although these studies employ similar econometric models and are based on the same data source, they come up with quite different results with respect to the effects of the entitlement to and the level of unemployment compensation on individual unemployment durations. Estimation results range from negative to modestly positive effects, but overall they do not seem to be particularly well determined. The reasons for this divergence in results are not clear yet.

This paper analyzes benefit–entitlement effects within a more general model which encompasses the previous studies' specifications of benefit–entitlement effects on the duration of individual unemployment. The main focus here is on the disincentive effects of the successive extension of benefit–entitlement periods in the 1980's. The next section contains a brief summary of the main economic hypotheses related to the effects of the unemployment insurance system on the duration of unemployment. This section also summarizes the structure of the German unemployment compensaton system and, in particular, the changes in legal regulations since the early 1980's. Section 3 reviews the previous microeconometric studies of benefit–entitlement effects on the duration of unemployment in Germany mentioned above. My own econometric analysis of benefit–entitlement effects is contained in section 4 which also contains a brief data description. The econometric analysis is based on a discrete–time hazard rate model estimated on waves 1 – 12 of the Socio–Economic Panel for West Germany covering the period 1983 to 1994.

To preview the main results of the paper based on my preferred specification of the estimated econometric model, there is strong evidence for entitlement effects for men, but not for women. However, long–term unemployment is much more affected if unemployment assistance, which tends to be open–ended, is available after the exhaustion of unemployment benefits. On the other hand, simulations show that even large reductions of the incomes replacement ratio have very little effect on the duration of unemployment. These results lead to some concluding remarks on the efficiency of the current German system of unemployment compensation and its recent reform in section 5.

2 Theoretical and Institutional Background

2.1 Entitlement Effects in Labor Supply and Job–Search Models

Two types of economic models have been used as theoretical basis for the empirical analysis of the relationship between various aspects of unemployment compensation systems and the duration of individual unemployment spells. The first approach is a straightforward extension of the static neoclassical labor supply model, the second approach relies on the theory of optimal job–search. Although these theories start from very different assumption, they arrive at quite similar conclusions regarding observable effects of benefits on the duration of individual unemployment spells.

To start with the traditional labor supply model [for a summary see Burtless (1989)], Moffitt and Nicholson (1982) describe an unemployed worker's decision to take up a job offer—assumed to be always readily available—as utility maximizing choice between income and leisure subject to a budget constraint. The only modification to the standard set–up is that in the presence of an unemployment insurance system leisure is subsidized and a limited entitlement period results in a kink in the budget constraint at the month of benefit exhaustion: Before this month, remaining unemployed for another month reduces income by the difference between the monthly amount of unemployment benefits received and the offered net wage. The ratio between these two variables is termed the income–replacement ratio. Depending on its size, this segment of the budget constraint may be rather flat while it becomes relatively steep after benefit exhaustion. Given random preferences for income and leisure, a disproportionate number of unemployed people will take up work in the month after the exhaustion of unemployment benefits.

Its static nature and the assumption that all unemployment is simply leisure make this model's implications look rather special. In this respect, job–search theory may provide a somewhat more realistic description of individual unemployment behavior. In its most simple version, unemployed individuals facing a known wage distribution and a constant job offer arrival rate chose a reservation wage that maximizes the present value of expected utility as a function of income and leisure. This reservation wage together with the job offer arrival rate and the distribution of offered wages determines the hazard rate, that is the conditional re–employment probability[2], and hence the duration of unemployment. The hazard rate is decreasing in the reservation wage which positively depends on the level of ALG. Since unemployment benefits act as a search subsidy reducing the costs of continued search, the higher the income–replacement ratio is the longer the duration of unemployment should *ceteris paribus* be in this model.

On the basis of a slightly more general version of this model, also allowing for a varying search intensity, Mortensen (1977) has shown that the hazard rate increases when the unemployed gets closer to the benefit exhaustion point, because search intensity is increased and the reservation wage reduced the shorter the remaining entitlement period becomes. After benefits are exhausted, the hazard rate remains

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[2] That is, the conditional probability of leaving the unemployment state in the next "month", given the spell has already lasted up to this month (see section 4.1 below).

constant because nothing else changes in this model. By implication, the hazard rate for an unemployed without benefit–entitlement should be constant over time. These predictions only hold if benefit exhaustion is the only source of non–stationarity within an otherwise stationary environment. However, there are various extensions to this simple model which introduce non–stationarity of some sort or the other [Lippman and McCall (1976) and Mortenson (1986) provide partial surveys of the literature]. Depending on the source of non–stationarity assumed, this results in negative or positive duration dependence of the hazard rate. Fortunately, this can be taken into account in the econometric model described below.

Benefit–entitlement effects on individual re–employment probabilities will depend on the regulations of the prevailing unemployment compensation system, such as the level of the income–replacement ratio and the length of the entitlement period, as well as on other aspects of the system, such as the degree of monitoring of an individual's search effort by the labor office. In the following, the German unemployment compensation system is therefore briefly described as far as it is relevant for the subsequent empirical analysis.

2.2 The German Unemployment Compensation System

The unemployment compensation system in Germany consists of two parts. The main part is unemployment benefits (*Arbeitslosengeld*, ALG) which is financed from contributions levied equally on employees and employers. The other part, unemployment assistance (*Arbeitslosenhilfe*, ALH), is financed by general government revenues, but administered like ALG by labor offices.

The basic requirements for the entitlement to ALG are that the applicant is registered as unemployed at the local labor office, is available for job placements, has been employed a minimum number of days in a given base period, and is not older than the legal retirement age of 65 years.[3] The *availability criterion* is fulfilled if the unemployed is available on short notice and ready to accept any "suitable" job or take part in training courses offered by the labor office. Until the most recent reform of the Employment Protection Act ("Arbeitsförderungs–Gesetz"), which came into effect on April 1, 1997, criteria for a "suitable job" applied by the labor office depended on the level of formal qualification as well as the required qualification in the previous job, but also on the duration of the on–going unemployment spell. Before 1994, only job offers requiring the same level of occupational qualification, special knowledge and experience as in the previous job were deemed suitable in the first few months of unemployment. In the subsequent months of the unemployment

[3] These and the following facts rely on the official publication of the German Ministry of Labor and Social Affairs [Bundesministerium für Arbeit und Sozialordnung (1995: 99 – 109)].

spell the requirements of a job offer with respect to its level of qualification could be adjusted downwards in a stepwise fashion as layed down in special regulations ("Zumutbarkeits–Anordnung") by the Federal Labor Office. These requirements were tightened by the reform of the Employment Protection Act in 1994, and again by the most recent reform which abolished the regulation with respect to the level of qualification altogether. Since then, the short–term unemployed (less than 3 months) can be obliged to accept any job offering a net wage of up to 20% less than in the previous job, provided the offered wage is not less than the relevant wage layed down in collective bargaining agreements or, in the absence of such agreements, the "common wage" for comparable jobs in the local labor market. For the next three months wage offers of up to 30% less than the wage in the previous job are deemed suitable, while for the subsequent months the level of the unemployment benefits define the lower bound for "suitable" wage offers an unemployed has to accept.

The unemployed may also be required to take up such jobs within a travel–to–works area of up to 2.5 hours a day (3 hours since April 1997), and even longer commuting times are possible under special circumstances layed down in the regulations by the Federal Labor Office. Unemployed persons aged 58 years and older who have decided to take up an early retirement pension because of previous long–term unemployment at the earliest possible age—usually at 60 years—are not expected to be available for any kind of job any longer. In certain seasonal industries, in particular the construction industry, farming and forestry, employees who were layed–off because of seasonal slack and are expected to be re–called by their firm when demand picks up are not required to actively search for a job and are offered no job vacancies by the labor office.

ALG payments can be suspended for up to 12 weeks, and the remaining entitlement period (see below) cut by 25 percent, if the unemployed rejects a suitable job offer or training course offered by the labor office without good reason. In case a suitable offer is rejected twice without good reason, entitlement to ALG may be withdrawn altogether. Although enforcement of requirements look quite stringent in theory, withdrawal of ALG entitlement for the reason of refusing a job offer is very rare in practice.[4] Entitlement to ALG is also suspended for up to 12 weeks, and the remaining entitlement period cut by 25 percent, in case the unemployed has voluntarily quit the previous job without good reason or was fired for reason of misconduct.

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[4] In 1995, when about 2.5 million people were registered as unemployed in west Germany, unemployment benefit payments were suspended in only about 12,500 cases [Bundesanstalt für Arbeit (1996), Übersicht 106).

The required *minimum employment period* is at least 12 months (360 calendar days) or, equivalently, certain other spells covered by the unemployment insurance system within the *base period*. These spells include periods of military service, sickness, maternity leave, and participation in training courses subsidized by the labor offices. The base period has been 3 years since July 1987, before that date it was 4 years. There are special regulations for employees in seasonal industries, where only 180 calendar days of employment within the last three years are required.

Table 1—Changes in benefit–entitlement periods since the mid–1980s.

Months worked in base period	January 1983 – Dec. 1984	January 1985 – Dec. 1985	January 1986 – June 1987	July 1987 – March 1997
12	4	4	4	6
16	4	4	4	8
18	6	6	6	8
20	6	6	6	10
24	8	8	8	12
28	8	8	8	14 (≥42)
30	10	10	10	14 (≥42)
32	10	10	10	16 (≥42)
36	12	12	12	18 (≥42)
40	12	12	12	20 (≥44)
42	12	14 (≥49)	14 (≥44)	20 (≥44)
44	12	14 (≥49)	14 (≥44)	22 (≥44)
48	12	16 (≥49)	16 (≥44)	24 (≥49)
52	12	16 (≥49)	16 (≥44)	26 (≥49)
54	12	18 (≥49)	18 (≥49)	26 (≥49)
56	12	18 (≥49)	18 (≥49)	28 (≥54)
60	12	18 (≥49)	20 (≥49)	30 (≥54)
64	12	18 (≥49)	20 (≥49)	32 (≥54)
66	12	18 (≥49)	22 (≥54)	32 (≥54)
72	12	18 (≥49)	24 (≥54)	32 (≥54)

Notes:
(a) Age groups to whom the various entitlement periods apply are in parantheses.
(b) The changes introduced by the most recent reform of the Employment Protection Act in 1997 are not included; see text.
Source: Hunt (1995, Table 1).

The length of benefit–entitlement periods depends on the cumulated duration of insured employment (or equivalent spells mentioned above) within the relevant base period and on the age of the unemployed person. For benefit–entitlement periods of more than 12 months the relevant base period is 7 years. Entitlement periods have been extended substantially for certain age groups between 1985 and 1987 but, except for the most recent changes, have remained unchanged since then. As summarized in Table 1, they range from 6 months for those aged 41 years and younger with only the minimum employment period to 32 months for those aged 53 and older with employment of at least 64 months in the base period. The recent reform has introduced some changes which will successively be phased in within the next few years. The age limits for extended entitlement periods will be raised in intervals of 3 years, such that only those aged 45 years and older will be entitled to receiving ALG for more than 12 months. When these reform steps are fully phased in, only those aged 57 years and older will be entitled to the maximum period of 32 months.

In contrast to ALG, ALH is means–tested and available to those who fulfill the mentioned requirements for ALG and have either exhausted their entitlement period ("Anschluss–ALH") or have some previous work experience, but cannot prove the minimum employment period required for ALG ("originäre ALH"). There is no definite previous employment period for this latter type of ALH, except that at least 5 months of insured employment (or equivalent spells covered by the unemployment insurance system) within the base period of 3 years (4 years before July 1987) are required. The means–test takes into account the incomes of all household members beyond certain fairly small deductions and also household wealth. Provided this means–test is passed and the mentioned entitlement criteria are met, ALH is granted for one year in the first instance. If household wealth exceeds a certain limit there is a waiting period before ALH becomes available. Those entitled to "Anschluss–ALH" may subsequently apply for a prolongation. Except for the age limit of 65 years, there is no definite time limit for which this type of ALH can be drawn provided the mentioned criteria are met, whereas the other type ("originäre ALH") is restricted to a period of up to twelve months.

Before 1984 the ratio of ALG to net income in the previous job, i.e., the *income–recplacement ratio*, was 68% up to the insured upper threshold amount ("Beitragsbemessungsgrenze"), for ALH the respective income–replacement ratio was 58%. This threshold amount is annually adjusted according to the growth rate of gross earnings of all employed covered by the social security system. For example, for the year 1997 it was set at a level of DM 8,200 of monthly gross earnings in the western German states, which is an increase of 2.5% from the previous year. In 1984 the income–replacement ratio for the unemployed without at least one child was cut to 63% for ALG and to 56% for ALH. In 1994 the ALG income–replacement ratio for the unemployed with (without) children was reduced to 67% (60) percent, for those on ALH it was cut to 57% (53%). The amount from which ALH is calculated ("Beitragsbemessungsgrundlage") is adjusted by a factor of 0.03 per year of unemployment as long as it exceeds 50% of the economy–wide average net wage of all dependently employed persons. However, this adjustment reduces the amount of

ALH only in case the yearly wage increase is below this adjustment factor. These changes affected both ongoing unemployment spells and spells beginning after the legal changes came into effect. The most recent reform of the unemployment compensation system tightened eligibility criteria but left income–replacement ratios unchanged.

3 Previous Empirical Studies for Germany

In his survey on the labor market effects of taxes and transfers covering the period up to the early 1990's Zimmermann (1993) concludes: "*In sum, there is not much evidence that would confirm the hypothesis that the German system of unemployment compensation causes unemployment by creating disincentives to work*". Since then, not much has been published in this area, and the few relevant studies are reviewed below.

Empirical studies on the effects of the entitlement to and the level of unemployment benefits on the duration of unemployment in (west) Germany[5] are based on reduced–form hazard rate models. The hazard rate is the conditional probability of leaving unemployment in the next "month", given unemployment has already lasted until the beginning of that month (for a formal definition see section 4.1 below). By definition, it is inversely related to the completed duration of the inflow into unemployment, which is to be explained in terms of indicators of the unemployment compensation system and other control variables. The various studies differ in the specification of the hazard rate and, in particular, the way potential benefit–entitlement effects are modelled. As to the latter, the studies vary from the simple inclusion of dummy variables for the entitlement to ALG or ALH at the beginning of a spell to the explicit modelling of changes in actual entitlement periods as derived from legal regulations and treated as a time–varying covariate.

Based on the first three waves of the Socio–Economic Panel for west Germany (GSOEP–west), Hujer, Löwenbein and Schneider (1991) find that the level of unemployment compensation and the entitlement–to–ALG dummy do not affect the hazard rate from unemployment significantly, whereas the ALH dummy reduces it substantially. The authors also included a dummy variable with a value of one (and zero otherwise) if entitlement to ALG was at most 2 months until exhaustion and found a very strong negative effect of this variable on the hazard rate from unemployment. Wurzel (1990) also included a dummy for the entitlement to ALG as an explanatory variable in a hazard rate model and found statistically insignificant effects of this variable on the hazard rate from unemployment in all specifications. In

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5 Steiner and Kraus (1995) also analyze entitlement effects in east Germany and find relatively strong effects for east German males.

a more general model, Wurzel (1993) accounted for potential differences in entitlement effects between various demographic groups and also controlled for heterogeneity in offered wages by including the individual expected wage (derived from standard wage equations) as additional explanatory variable. The estimated effects differ somewhat between gender and age groups and also depend on the specification of the particular model, but overall no strong effects seem to emerge from these estimates. In contrast, Steiner (1994) found relatively strong effects from the entitlement to unemployment compensation (benefits or assistance) on the hazard rate, whereas marginal reductions in the level of unemployment compensation had very little effect.

Hunt (1995) makes use of the so–called *difference–in–difference* approach and the various changes in entitlement periods for particular groups in the 1980s to estimate the effect of these changes on the duration of unemployment.[6] In this approach, individuals are divided into two groups: the *treatment* group refers to the unemployed to whom a change in some law or regulation applies, whereas those unaffected by these changes belong to the *control* group. The effect of a law change on some outcome variable, *e.g.* the duration of unemployment, is the difference of the outcomes for the treatment and control groups before and after the law change[7]. In a hazard rate model framework, a dummy for the period to which the new regulation refers and treatment group dummies as well as interaction terms between these and the period dummy are included as explanatory variables together with other control variables. The estimated coefficient on these interaction terms then gives the effect of a law change after controlling for pure period and treatment group effects. Obviously, this methodology can easily be extended to several changes in regulations within the observation period and a varying number of treatment groups by including appropriate dummy variables and interaction terms.

The treatment groups considered by Hunt (1995) consist of the unemployed without children, for whom there was a reduction in unemployment compensation at the beginning of 1984 which remained in effect throughout the observation period (see section 2.2), and three age groups. The control group consists of those aged below 42 years for whom entitlement periods have not changed within the observation pe-

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6 Hunt (1995, Table 6) also presents estimation results for a simple competing risks model (escapes to employment and to out–of–the–labor–force, respectively) with dummies for the receipt of ALG or ALH and the level of unemployment compensation included as exogeneous variables, but without conditioning on the (expected) wage. She finds that entitlement to ALG significantly (insignificantly) increases the hazard rate in employment (to out of the labor force), while entitlement to ALH reduces both hazard rates significantly, and that the level of unemployment compensation has no significant effect on hazard rates.

7 For a survey on this and related approaches to the empirical analysis of policy impact analysis with non–experimental data see, e.g. Meyer (1995).

riod. The age groups and the four period dummies are defined according to the law changes summarized in section 2.2 above, and the interaction terms are defined accordingly. One crucial assumption implicit in the construction of the treatment group dummies is that, given an unemployed belongs to a certain age group, the maximum period of ALG–entitlement is obtained. As mentioned in section 3.1, this does not correspond to the way the German unemployment insurance system works, and is also not compatible with actual or potential entitlement periods as observed in the data (see section 4.2 below).[8]

On the basis of a simple hazard rate model and data from GSOEP for the period January 1983 to December 1988 Hunt (1995) finds significant benefit–entitlement effects. For both males and females in the age group 44 – 48 years there seem to be negative effects on the hazard rate to employment, but they refer to different time periods, January 1986 to June 1987 for females and July 1987 to December 1988 for males, respectively. Since the extension of the entitlement periods occurred at the same time for males and females, it is not quite clear how to interpret these gender differences. Furthermore, it also remains unclear why this effect only shows up for this age group and not for older workers as well, for whom entitlement periods were extended even more. However, for older women (aged 49 – 57 years) the extension of entitlement periods seems to have reduced the out–of–the–labor–force hazard rate substantially.[9]

Hujer and Schneider (1996) extend this approach by including a variable for an individual's potential ALG–entitlement period at the beginning of the spell, dummies for the remaining ALG–entitlement period, a dummy for ALH–entitlement, and the income–replacement ratio. The dummies for the remaining ALG–entitlement period should account for the dynamics of the entitlement effect, where the hypothesis to be tested is that the shorter the remaining period gets the greater the incentive to take up a job offer and, hence, the higher the hazard rate from unemployment becomes (see section 2.1). In contrast to Hunt (1995), the authors derived the potential entitlement period at the beginning of an unemployment spell instead of simply using the maximum entitlement period as given by the legal regulations for all the

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8 Hunt (1995, p. 112) notes that she also ran regressions "with the treatment groups defined as a function of experience as well as age", but it is not clear to me how exactly this was done (see also her footnote 20 on that page).

9 For example, Hunt (1995, Table 10) reports an estimated coefficient on the period3 × aged 49–57 years dummy of –1.36, which would imply that the out–of–the–labor–force hazard rate would be only about a fourth the level of that of the reference person not entitled to an extended benefit period. Note, however, that for a relatively low level of the out–of–the–labor–force hazard rate the effect on the overall hazard rate from unemployment would nevertheless be modest (see section 4.3 below).

unemployed by using information on the actual number of months for which ALG was received from the calendar information in the GSOEP (see the appendix).

As in Hunt (1995), the authors find that the extension of benefit–entitlement periods in the 1980's has only affected the hazard rate of those males aged 44 years and older who became unemployed after July 1987. For females in the age groups 44 – 48 years as well as 49 years and older who became unemployed after Janurary 1986, Hujer and Schneider (1996) also find very strong negative effects on the out–of–the labor–force hazard rate, which they interpret as evidence for the hypothesis that females tend to leave the labor–force just after their entitlement to benefits has been exhausted. On the other hand, for females the income–replacement ratio has a significantly *positive* effect on the employment hazard rate for females and no statistically significant effect on the out–of–the labor–force hazard rate. For males entitled to ALG, the authors find no statistically significant effect of the income–replacement–ratio on the hazard rate from unemployment. In contrast, ALH–entitlement has a relatively strong negative effect on the hazard rate into employment and, for females, out–of–the–labor–force as well.

The length of the ALG–entitlement period has a strong negative effect on the male hazard rate into employment which falls in a stepwise fashion with the decreasing remaining months of entitlement and suddenly jumps up when ALG is exhausted. For females, the length of the potential benefit–entitlement period does not affect the hazard rate into employment, but has a relatively strong *positive* effect on the out–of–the–labor–force hazard rate. On its own, this latter result seems surprising. However, in interpreting it the effects of the dummies for the remaining months of benefit–entitlement have also to be taken into account. These variables show very strong negative effects on the hazard rate out–of–the–labor–force and, to a lesser extent, into employment as well.

To sum up, it seems fair to say that the reviewed empirical studies do not provide a consistent picture of the effects of unemployment compensation on the duration of individual unemployment in Germany. One result on which some agreement seems to exist is that marginal variations of unemployment benefits have only very small effects on the hazard rate, if any. The effects of the entitlement to ALG or ALH tend to be much more important quantitatively, but seem to depend very much on the way these effects are modelled, and also to differ between males and females. The econometric model described in the next section encompasses most of the specifications found in previous research and thus allows to draw more general conclusions.

4 Econometric Analysis

4.1 Hazard Rate Model

Following the standard approach, reduced–form hazard rate models are estimated to quantify benefit–entitlement effects on individual unemployment durations [for general surveys of hazard rate models see, e.g., Kiefer (1988) or Lancaster (1990)]. Given the discrete measurement of unemployment durations derived from the monthly calendar data in the GSOEP and the associated heavy ties of observations, it seems more appropriate to specify a discrete rather than a conventional continuous–time hazard rate model as employed in the studies reviewed in the previous section.

The formal structure of the econometric model is the following. The duration of an individual's *k*–th unemployment spell is described by a non–negative random variable, *T*, which takes on integer values only. If an unemployment spell ends in the interval $\left[I_{t-1}, I_t\right)$ this variable takes on a value of T = t, where the spell can either end in employment or in the out–of–the–labor–force state. The central variable for modelling the transition process from unemployment into any one of these two states is the discrete state–specific hazard rate. For the *i*–th person the hazard rate in spell *k* into state *j* in interval *t*, $\lambda_{ij}^k(t)$, is the conditional probability of a transition into state *j* in this interval, given individual *i* has been unemployed until *t*, i.e.,

$$(1) \qquad \lambda_{ij}^k\left(t | x_i(t), \varepsilon_i^m\right) = P\left[T_{ik} = t, \Omega = j | T_{ik} \geq t, x_i(t), \varepsilon_i^m\right]$$

with i = 1,2,...n; j = 1,2; k =1,2,...K_i;

$x_i(t)$ = vector of covariates of individual i in intervall t

Ω = 1, if transition into employment

= 2, if transition out–of–the–labor–force

ε_i^m = time–invariant individual effect, where

$$E(\varepsilon_i) = \sum_{m=1}^{M} \mathrm{P}(\varepsilon_{\mathrm{i}}^{\mathrm{m}})\varepsilon_{\mathrm{i}}^{\mathrm{m}} = 0; \ \sum_{m=1}^{M} \mathrm{P}(\varepsilon_{\mathrm{i}}^{\mathrm{m}}) = 1; \ E(\varepsilon_i^m x_i(t)) = 0, \ \forall \text{ m (m} = 1,2,\ldots \text{M).}$$

The time–invariant individual effect, ε_i, accounts for unobserved population heterogeneity in the hazard rates and is assumed to come from an arbitrary discrete probability distribution with a small number of mass points, ε_i^m (m=1,2,..M). These mass points and their probabilities, P(ε_i^m), are simultanously estimated with the parameters of the model. This individual effect is assumed to be uncorrelated with the set of explanatory variables in the model, $x_i(t)$. The variables of main interest here are related to the unemployment compensation system. They encompass the various specifications of benefit–entitlement effects found in the previous German studies summarized above. That is, they include the income–replacement ratio and, alternatively, dummies for the entitlement to ALG or ALH at the beginning of the unemployment spell, interaction terms accounting for treatment–group effects, and benefit–entitlement periods as as a time–varying covariate. These variables are briefly described in the next section and, in more detail, in the appendix. In addition to these variables, $x_i(t)$ includes variables that control for differences in individual characteristics and other observed factors affecting individual unemployment behavior (see section 4.2). Note that some of these variables, e.g. the regional unemployment rate, not only depend on process time, i.e. the month of the unemployment spell, but also on historical time.

Conditional on the vector of covariates and the individual effect, transitions into the two states are independent and can thus be modelled as competing risks.[10] The hazard rate from unemployment is therefore given by

$$(2) \qquad \lambda^k\left(t | x_i(t), \varepsilon_i^m\right) = \sum_{j=1}^{2} \lambda_j^k\left(t | x_i(t), \varepsilon_i^m\right)$$

In terms of the hazard rate, the probability of remaining unemployed in period t conditional on having been in that state up to period $t-1$ is simply given by

$$(3) \qquad P\left[T_k > t | T_k \geq t, x_i(t), \varepsilon_i^m\right] = 1 - \lambda^k\left(t | x_i(t), \varepsilon_i^m\right)$$

The survivor function gives the (unconditional) probabilty of remaining unemployed up to period t; in terms of the hazard rate it can be written as

$$(4) \qquad P\left(T_k > t | x_i(t), \varepsilon_i^m\right) = S^k\left(t | x_i(t), \varepsilon_i^m\right) = \prod_{\tau=1}^{t-1}\left(1 - \lambda^k\left(\tau | x_i(t), \varepsilon_i^m\right)\right)$$

The probability of a transition into state j in period t in terms of the respective hazard rates is

$$(5) \qquad P\left(T_k = t, \Omega = j | x_i(t), \varepsilon_i^m\right) = \lambda_j^k\left(t | x_i(t), \varepsilon_i^m\right) \prod_{\tau=1}^{t-1}\left(1 - \lambda^k\left(\tau | x_i(t), \varepsilon_i^m\right)\right)$$

The hazard rates are modelled by means of random–effects logits with three distinct choices (states), namely unemployment, employment and out–of–the–labor–force, the first one being the base category, i.e.,

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10 Of course, without conditioning on the individual effect transitions into the two states will be correlated.

$$(6)\qquad \lambda_{ij}^{k}\left(t \mid x(t),\varepsilon_i^m\right)=\frac{\exp\left(\alpha_j(t)+\beta_j' x_i(t)+\varepsilon_i^m\right)}{1+\sum_{l=1}^{2}\exp\left(\alpha_l(t)+\beta_l' x_i(t)+\varepsilon_i^m\right)}$$

The so–called "baseline" hazard, $\alpha_j(t)$, describes the dependence of the hazard rate on process time ("duration dependence"). To avoid the danger of seriously misspecifying the model, the baseline hazard is modelled in a flexible way by a set of dummy variables. Also note that the specification of the hazard rates in equ. (6) does not imply the rather restrictive proportional hazard assumption on which the empirical models estimated in the studies reviewed in the previous section typically rely. In the present application, this assumption would be rather restrictive because it implies that, say, potential benefit–entitlement effects on the hazard are independent of the duration of unemployment.

The corresponding survivor function for this model is

$$(7)\qquad S^{k}\left(t_i \mid x_i(t),\varepsilon_i^m\right)=\prod_{\tau=1}^{t-1}\frac{1}{1+\sum_{l=1}^{2}\exp\left(\alpha_l(\tau)+\beta_l' x_i(\tau)+\varepsilon_i^m\right)}$$

which will be used to simulate the effects of changes in the unemployment compensation system on the duration of unemployment below.

To derive the sample likelihood function for this model, the indicator variable

$$c_{ik}=\begin{cases}1, & \text{if the } k\text{-th unemployment spell of individual } i \text{ is right-censored}\\ 0, & \text{otherwise}\end{cases}$$

is defined. Right–censored observations include interrupted spells either at the end of the observation period or related to sample attrition. Following usual practice, it is assumed that the censoring mechanism is non–informative (i.e., random). Since there is no operational way to include information on left–censored spells in the likelihood function in a consistent way, they are excluded from the sample.

Defining another indicator variable

$$\delta_{ikj}=\begin{cases}1, & \text{if the } k\text{-th unemployment spell of individual } i \text{ ends in state } j\\ 0, & \text{otherwise,}\end{cases}$$

and assuming that, conditional on the explanatory variables in the model and the individual effect, all observations are independent[11], the sample likelihood function is given by

$$(8) \qquad L = \prod_{i=1}^{n} \sum_{m=1}^{M} P(\varepsilon_i^m) \prod_{k=1}^{K_i} \prod_{j=1}^{2} \left[\lambda_{ij}^k \left(t_i | x_i(t_i), \varepsilon_i^m \right) \right]^{\delta_{ikj}} \prod_{\tau=1}^{t_i-1} \left(1 - \lambda_i^k \left(\tau | x_i(\tau), \varepsilon_i^m \right) \right)^{c_{ik}}$$

For a completed unemployment spell the contribution to the likelihood function is given by the respective state–specific hazard rate, for a censored spell it is given by the survivor function, which is written in terms of the overall hazard rate here. Note that due to the individual effect all observations for a given individual—both within and between spells—are correlated, conditionally on the previous state and the set of explanatroy variables.

The likelihoold function (8) is maximized with respect to the coefficients on the baseline hazard, the coefficients on the explanatory variables and the mass–points together with the corresponding probabilities, $\hat{P}\left(\varepsilon^m\right)$, taking into account the restrictions on the individual effects given in eq. (1) by standard numerical optimization procedures.

4.2 Data

The econometric model of the previous section is estimated on waves 1 – 12 of the Socio–Economic Panel for West Germany (GSOEP) covering the period January 1983 to 1994.[12] In the first wave some 12,000 individuals older than sixteen years of age living in about 6,000 households were interviewed. In the subsequent waves of the panel the same people were followed up and "new" households, of which at least one was member of the households initially included, were added to compensate for sample attrition. In each panel wave detailled information on an individual's labor–force status and various types of income as well as personal and household characteristics and other variables relevant for the explanaton of individual unemployment behavior (see below) is collected. In addition, at the date of interview of each wave, retrospective monthly "calendar" information on an individual's detailed labor force status in each month of the previous calendar year is recorded.

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11 Although the assumption that individual observations are independent in the *sample* is standard in microeconometric models of individual (unemployment) behavior, it need not hold in the population.

12 Details on the GSOEP can be obtained from the webserver of the German Institute of Economic Research (DIW) in Berlin (http://www.diw–berlin.de/soep/).

Depending on the wave, there are between eight and ten different categories for an individual's labor force status, which can be aggregated into the following three labor–force states:

(i) *unemployment*

(ii) *employment*

(iii) *out–of–the–labor–force*.

Since the questionnaire refers to *registered* unemployment, its definition used here is, in principle, the same as in official statistics, and has the same well–known problems of both over– and underreporting. It also corresponds to one of the basic requirement for the entitlement to unemployment benefits mentioned in section 2.2. The employment state includes full–time, part–time and temporary employment as well as vocational training in firms. The out–of–the–labor–force state comprises people in (early) retirement, in full–time education, on military service, working at home, and "others". The unemployed aged 58 years and older are excluded from the sample because of the special regulations for this age group described in section 2.2. Persons previously employed in seasonal industries, like the construction industry, farming and forestry, are also excluded from the analyzed sample because there are special regulations regarding benefit–entitlement periods for them and—since a large proportion of them is on temporary layoff—their labor market behavior is likely to differ substantially from that of the other unemployed. A very small number of civil servants who became unemployed in the observation period was also excluded from the sample.

Completed durations of individual unemployment spells are derived from information on the date of entry into unemployment and the date of the transition from this state into states (ii) or (iii) defined above. Interrupted durations of right–censored spells are calculated from the entry date and the date an unemployed is observed for the last time in the GSOEP, which also includes sample attrition. Unemployment spells which began before January 1983 or for which no calendar information from the previous year is available are treated as left–censored and excluded from the sample. The number of unemployment spells beginning between January 1983 and December 1994 is 2,828 for males and 2,159 for females. After the groups mentioned above were excluded from the sample, 1,736 male and 1,766 female unemployment spells remained in the sample (see Table 2). Of those, 72,9% (58.3%) ended in employment, 10,3% (27.1%) left the labor force, and 16,8% (14.5%) were right–censored at the time they were observed in the sample for the last time. The number of analyzed spells refer to 1,125 males and 1,210 females, respectively. Hence, even though the unemployed previously employed in seasonal industries were excluded from the sample repeat, unemployment spells were quite important within the observation period.

Information on an individual's *actual* entitlement to ALG or ALH is available on a monthly basis from the income calendar data in the GSOEP. Table 2 shows the absolute and relative number of all unemployment spells covered by ALG or ALH, the respective number of spells with censored or exhausted entitlement periods, and the number and ratio of spells with ALH coverage either after the exhaustion of ALG or without prior entitlement to benefits. Of all spells analyzed about 50% received ALG and 10% ALH, where these percentages differ little between males and females. About two–thirds of all spells ended before the exhaustion of ALG–entitlement, which was followed by ALH–entitlement for 12.8% of males and 7.8% of females. ALH without previous ALG–entitlement, i.e. "originäre" ALH, was received by about 50% of all unemployed in the sample.

Table 2—Unemployment spells and benefit–entitlement

	Males		Females	
	#	%	#	%
(1) All spells	2828	100	2159	100
Previously employed in seasonal industries	642	22.7	42	1.9
Previously employed as civil servant	30	1.1	17	0.1
Age > 57 years at spell begin	144	5.1	74	3.4
Left–censored	276	9.8	260	12.0
(2) Spells analyzed	1736	100	1766	100
(3) Received ALG during spell, % of (2)	892	51.4	943	53.4
(4) Received ALH during spell, % of (2)	217	12.5	159	9.0
Spell ended before ALG exhaustion, % of (3)	575	64.5	631	66.9
Spell censored before ALG exhaustion,% of (3)	84	9.4	53	5.6
Spells with exhausted ALG, % of (3)	233	26.1	259	27.5
ALH on exhaustion of ALG, % of (3)	114	12.8	74	7.8
ALH without prior ALG–entitlement, % of (4)	103	47.5	85	53.5

Notes:

(a) ALG = unemployment benefits; ALH = unemployment assistance.

(b) The number of spells analyzed refers to those available for estimation of the hazard rate models in Table 5, i.e. after exlusion of all spells with missing values in any of the explanatory variables in the models estimated.

Source: GSOEP, waves 1 – 12; own calculations.

To model dynamic entitlement effects, information on an individual's *potential* ALG–entitlement period at the beginning of the unemployment spell is required. This information is not directly available in the GSOEP but can be derived by combining monthly calendar data and retrospectively collected employment information available (for the derivation see the appendix). The *remaining* benefit–entitlement period in a particular month of the unemployment spell is then simply calculated by subtracting the respective number of months from the potential entitlement period at the beginning of the unemployment spell. For the sample we analyze below this more involved imputation yields an average potential entitlement period at the beginning of the unemployment spell of 9.0 months for males and 8.5 months for females. These values are only slightly lower than those one obtains by using the simpler imputation method employed by Hujer and Schneider (1996)—9.2 and 8.7 months, respectively—and both measures are considerable shorter than the maximum benefit–entitlement period of 14.3 and 13.8 months implicitely used by Hunt (1995).

There are various ways to model the income–replacement ratio, in particular with respect to the income measure to which the amount of ALG received is to be compared [see Atkinson and Micklerwright (1991)]. The measure used here refers to expected net earnings in the new job: It thus corresponds to the theoretical framework outlined in section 2 but differs from the one used by Steiner (1994) and Hujer and Schneider (1996) calculated on the basis of earnings in the previous job. Expected net earnings are derived from empirical wage equations and a simple tax function estimated on the basis of the GSOEP as described in the appendix. The income–replacement ratio is calculated as the ratio of the amount of monthly unemployment benefits, which is income–tax free, and the net expected monthly wage in employment. This procedure yielded a mean replacement ratio of about 0.5 with a standard deviation of 0.2 for those receiving ALG, where there is very little difference between males and females. The amount of unemployment benefits (ALG or ALH) received in real terms rather than the income–replacement ratio is included in the out–of–the–labor–force hazard rate because the expected wage in employment is obviously not an appropriate measure for the expected household income when leaving the labor–force, which depends in the change of other member's income due to taxes, and social transfers. Lacking such a measure, non–labor household income, the employment status of the spouse, and his or her earnings should proxy the alternative income in the out–of–the–labor–force state.

In addition to the variables describing the unemployment compensation system the set of regressors also includes the usual control variables, such as personal characteristics, indicators of household composition, human capital variables and indicators for the state of the aggregate labor market. Some information on these variables is contained in Table A1 in the data appendix, which refers to the preferred model specification presented in section 4.3.3 below. Spells with missing values on any of the variables included in the estimated hazard rate models were dropped from the sample.

4.3 Estimation Results

The presentation of estimation results starts with the most simple specification which includes two dummy variables for the entitlement to ALG and ALH as well as the level of unemployment compensation. This kind of specification has been used in most of the studies summarized in section 3 and thus serves as the reference model and as the starting point for the more sophisticated analysis reported below. In the following subsection, estimation results for the model with treatment group dummies as used by Hunt (1995) and, in a somewhat extended form, also by Hujer and Schneider (1996) are presented. The final estimates refer to a more general specification of the unemployment compensation system with the remaining months of entitlement as a time–varying explanatory variable. The estimates from this specification are then used to simulate benefit–entitlement effects on the duration of unemployment.

4.3.1 Standard Specification

Estimation results for the "standard" specification of unemployment compensation effects on the hazard rate from unemployment are reported in Table 3A for males and in Table 3B for females. For males, the relatively few out–of–the–labor–force transitions are not explicitely modelled but are treated as right–censored in the estimation. Since the focus of this paper is on the effects of unemployment compensation on the hazard rate, estimation results for the control variables will not be discussed here. Suffice it to note that they do to some extent control for the heterogeneity in the sample, but unobserved heterogeneity remains quantitatively important. Likelihood ratio tests have shown that three (two) heterogeneity groups, i.e. mass points, are sufficient to account for remaining unobserved heterogeneity in the male (female) sample. These mass points and their probabilities are reported at the bottom of Tables 3A and 3B.[13]

Except for the coefficients of the baseline dummies, controlling for unobserved heterogeneity has very little effect on the parameter estimates. It is well known that the negative dependence of the hazard rate into employment on the duration of unemployment tends to disappear if heterogeneity in the sample is adequately controlled for [see, e.g., Steiner (1994)]. This result is also obtained here. Whereas the estimates for the models without unobserved heterogeneity [not shown here] imply negative duration dependence of the hazard rate into employment for both males and females, the estimated baseline hazards in Tables 3A and 3B show that the baseline hazard slightly increases for males and remains constant for females throughout the unemployment spell. For females, the out-of-the-labor-force hazard rate increases significantly after the first few months, again after the ninth month and then remains at that higher level.

In Tables 3A and 3B, the effect of the entitlement to ALG or ALH in any particular month during the unemployment spell on the hazard rate into employment and non–participation is to be interpreted together with the effect of the income–replacement ratio and the amount of unemployment compensation, respectively. Estimated coefficients imply that, for an income–replacement ratio of about 60%, its effect on the hazard rate into employment is slightly negative both for males and females. Marginal changes in the level of benefits have very little effect on individual re–employment probabilities.

[13] These probabilities can be interpreted as the respective proportion of the unemployed in the sample belonging to one of the three (two) heterogeneity groups.

In contrast, entitlement to ALH reduces the re–employment hazard substantially, especially for males, and has a strong negative effect on the out–of–the–labor–force hazard rate for females. Hunt (1995) also obtains significant negative (positive) effects of entitlement to ALH (ALG) on hazard rates. She speculates that the latter variable may act as a proxy for recent labor market experience rather than account for entitlement effects. Note that my specification includes an individual's previous labor force state as a proxy for recent labor market experience, which has a rather strong negative effect on the female re–employment hazard but only a weak effect for males.

4.3.2 Specification with Treatment Groups

Estimation results from models including, in addition to all the control variables in the specification of the previous subsection, period–group interaction terms as explanatory variables are summarized for males and females in Table 4. In this specification, the effects of changes in legal regulations in unemployment compensation rules on hazard rates are measured by the coefficients on the period–group interaction terms, i.e. treatment groups for short. Since estimated coefficients of the control variables change very little compared to those of the previous estimates, they are not reported here.

Estimation results show that only the period3 × aged 44–48 years–dummy has the expected negative effect on the re–employment hazard rate. This effect refers to the same treatment group for which Hunt (1995) and Hujer and Schneider (1996) obtained a significant effect on the re–employment hazard, though their point estimate is only about half the size of the one obtained here. Contrary to these authors, I do not find any significant entitlement effect on the female out–of–the–labor–force hazard rate, however. Nor do I find, as in Hunt (1995), evidence for the hypothesis that the reduction of unemployment compensation for those without children has increased the out–of–the–labor–force hazard rate. Furthermore, for the male re–employment hazard rate the large positive coefficient of the period3 × aged 49–53 years–dummy does not correspond to prior expectations.

There are at least two potential short–comings of the specification relying on treatment groups which may severely bias estimation results. First, as the calculation of potential benefit–entitlement periods reported in section 4.2 has shown, the assumption that people falling into a particular treatment group are in fact entitled to draw ALG for the maximum period is far off the mark. Second, the effects of the treatment group dummies may vary over the duration of the unemployment spell, which the specification in this subsection does not allow for. The next subsection summarizes estimation results for specifications of entitlement effects which take these factors into account.

4.3.3 Specification with Remaining Entitlement Period

The dynamic pattern of benefit–entitlement effects can be modelled by including the remaining number of months of ALG–entitlement as time–varying covariate in the hazard rate model. As mentioned in section 2.1, there are two dynamic effects on the re–employment hazard rate; first, the shorter the remaining entitlement period becomes the higher the hazard rate gets and, second, the hazard increases sharply in the month of benefit exhaustion. To account for non–linearities of entitlement effects on the hazard rate, they are modelled by means of dummy variables here. Estimation results for this specification are contained in Table 5.

As before, estimation results for the control variables are not reported, because they change very little compared to those reported in Tables 3A and 3B, respectively. Of course, since the (remaining) potential entitlement period is just another (and better) proxy for treatment groups, the period–group interaction terms of the previous subsection are not included here. The estimated coefficients of the dummies for the remaining month of benefit–entitlement refer to the base category of 18 or more months. In order to have a sufficient number of observations in each category, several months had to be aggregated. The effect of changing from one entitlement period to the next closer one to the date of benefit exhaustion is thus, *ceteris paribus*, given by the difference in the estimated coefficients of the two respective entitlement categories ("pure entitlement effect"). Since this also implies a corresponding change in the duration of unemployment, this pure entitlement effect is compounded by the pure duration dependence effect which is accounted for by the baseline hazard. These two effects are identified from the variation of individual benefit–entitlement periods at the beginning of the unemployment spell [see Meyer (1990)]. Furthermore, the ALG– and ALH–entitlement dummies as well as the income–replacement ratio or the amount of unemployment compensation also may change at the month when entitlement to ALG is exhausted. Hence, to calculate the effect of moving from one month of remaining entitlement to the exhaustion point one has to add these effects to the "pure entitlement effect" after taking into account changes in the baseline hazard rate.

The dynamic pattern of entitlement effects implied by the estimates in Table 5 is illustrated in Figure 1 for males and Figures 2A and 2B for females. The plots refer to a particular reference group of unemployed who are assumed to have become unemployed between July 1987 and December 1992, to be between 44 – 48 years old and have other characteristics defined by the respective base category for dummy variables and sample means for metric variables [see Table A1 in the appendix]. The hazard rates are calculated for potential ALG–entitlement periods at the spell begin of between 0 and 18 months.

The comparison of the hazard rate for the unemployed without benefit–entitlement (entitlement period = 0 months) and those with the same characteristics but, say, 12 months of potential ALG–entitlement shows the "pure" entitlement effect for each month. To illustrate the effect of entitlement to ALH on the hazard rates, the plots also distinguish between the unemployed with and without ALH–entitlement after the exhaustion of ALG. The plots show "averaged" hazards calculated as the weighted sum of the hazard rates over the three (two) male (female) heterogeneity groups, where the weights are the estimated probabilities of the corresponding mass points (see section 4.1).

For males without entitlement to unemployment benefits, the estimated re–employment hazard rate is about 15% per month at the beginning of the spell; subsequently it declines slowly with an intermittent increase between the 10th and 12th month.[14] For those receiving ALG the evolution of the hazard rate is over time depends on the potential entitlement period at the spell begin and on subsequent ALH–entitlement. Whereas the hypothesis that the hazard rate increases monotonically with a decreasing number of months of ALG–entitlement is not confirmed by the shape of the hazard plots, one effect suggested by the theoretical models referred to in section 2.1 is borne out very clearly for males. If ALH is not available the re–employment hazard rate jumps upwards in the month after exhaustion of ALG–entitlement and subsequently remains at this much higher level. This effect is illustrated in Figure 1 for an initial entitlement period of 6, 12, and 18 months, respectively. In contrast, the re–employment hazard of those unemployed men receiving ALH remains at the level it has reached the month before the exhaustion of ALG–entitlement.

For females, the plots in Figure 2A do not show any systematic entitlement effects on the re–employment hazard rate, which also does not differ significantly between those with and without ALH–entitlement. As Figure 2B shows, there is some evidence for entitlement effects on the out–of–the labor force hazard rate for female who are not entitled to ALH after exhaustion of ALG, but this hazard is too low for this effect to have a noticeable impact on the duration of unemployment (see simulations below).

Estimated entitlement effects differ between age groups and are also likely to depend on general labor market conditions. To quantify these effects, in Table 6 we

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[14] The decline in the hazard is related to the "averaging" of the unobserved heterogeneity components. Within the heterogeneity groups, i.e., for given ε_j, male (female) hazard rates into employment are slightly increasing (constant) in duration. Assuming a constant hazard rate, a monthly value of 15% would imply an average completed unemployment duration of about 7 months (in this case durations are exponentially distributed and the mean of the distribution is equal to the reciprocal of the hazard).

report survival probabilities after 6, 12, 18, and 24 months of unemployment for various age groups in the period July 1987 to December 1992 and January 1993 to December 1994, respectively. Whereas the first period is characterized by a relatively strong employment expansion and a substantial reduction in aggregate unemployment, the second period refers to one of the most severe recessions in post–war German history. These survival probabilities are calcuted by inserting the estimated coefficients from the hazard models in Table 5 into the survival function [equation (7) in section 4.1], where the other observable characteristics are set at the same values as for the reference group and unobserved individual effects are "averaged–out" as described above. More intuitively, these survival probabilities can also be viewed as the proportion of the cohort of people in the reference group entering unemployment in a particular month who are still unemployed after the respective number of months.

Table 6—Estimated age and time effects on survival rates in unemployment

	Males survival rate after ... months				Females survival rate after ... months			
	6	12	18	24	6	12	18	24
	Spell begin between July 1987 and Dec. 1992							
Age<25 years	30.8	13.1	6.3	3.6	18.0	3.2	1.1	0.4
25≤Age≤41 yrs.	42.3	22.4	12.4	7.5	35.4	12.1	6.2	3.7
42≤Age<44 yrs.	47.8	27.8	16.7	10.5	49.3	23.3	14.7	10.2
44≤Age<49 yrs.	66.0	48.9	36.7	28.0	38.7	14.4	7.9	4.8
49≤Age<54 yrs.	70.6	54.8	43.2	34.5	64.3	39.6	29.3	23.0
Age≥54 years	87.1	78.0	70.5	64.2	84.6	69.5	61.1	54.9
	Spell begin between Jan. 1993 and Dec. 1994							
Age<25 years	48.6	28.7	17.4	11.0	33.1	10.4	5.1	2.8
25≤Age≤41 yrs.	59.2	40.5	28.2	20.0	52.3	26.0	16.9	11.9
42≤Age<44 yrs.	63.9	46.2	34.0	25.4	64.8	40.2	30.0	23.6
44≤Age<49 yrs.	78.1	65.0	54.8	46.8	55.5	29.3	19.9	14.5
49≤Age<54 yrs.	81.4	69.7	60.4	52.8	76.3	56.2	46.3	39.5
Age≥54 years	92.5	86.7	81.5	76.9	90.1	79.3	72.8	67.7

Notes:
(a) Survival rates are given in percent.
(b) Dummy–variables are set at the respective base category, metric variables at their values at sample means.

Source: Calculations are based on estimation results in Table 5.

Table 6 shows that, except for females aged 44 to 49 years[15], survival probabilities are strongly increasing in age. Only about 13% of males and 3% of females aged below 25 years who became unemployed in the period July 1987 to December 1992 remained in that state for more than a year. The respective survival probabilities for the oldest age group (≥ 54 years) were almost 80% and 70%, respectively. Survival probabilities in unemployment estimated for the period July 1987 and December 1992 are, *ceteris paribus*, markedly smaller than those estimated for the subsequent period. This holds for all age groups, but cyclical effects seem to have a stronger relative impact on younger workers' survival probabilities in unemployment.

4.4 Simulation of Benefit Entitlement Effects

On the basis of the estimation results in the previous sub–section, simulation results of the quantitative effects of changes in entitlement periods and the level of the income–replacement ratio are presented in Table 7. The simulations again refer to the 44 to 48–years old reference group defined above in the period July 1987 to December 1992. They show how the proportion of unemployed people in this reference group changes ceteris paribus if the parameters of the unemployment benefit system are varied one at a time. For different age groups and the subsequent time period, the simulated effects of variations of the parameters of the unemployment insurance system on survival rates would differ in levels but not in relative size.

Simulation A in Table 7 shows that 43.5% (36.9%) of all males (females) in the reference group with no benefit–entitlement remain unemployed for at least 6 months, and 18.3% (13.6)% for at least a year. Note that the relatively low female survival rates are due to the choice of the reference group which refers to skilled singles aged 44 to 48 years without children, whose work motivation presumably is rather high. If the unemployed are entitled to ALG and subsequently to ALH, simulated survival rates for males are considerabIe higher and also depend on the length of the entitlement period at the beginning of the spell (simulation B). For example, the 12–month survival rate for an unemployed man in the reference group with an ALG–entitlement period of 6 months and subsequent ALH–entitlement is almost 50%, while the respective value for someone without entitlement to ALH is just about half this value (simulation C). Likewise, for the male reference group with longer ALG entitlement periods the share of those subsequently entitled to ALH who remain unemployed longer than 12 months is always considerably higher than those whose benefit–entitlement period has expired.

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15 However, the differences in estimated coefficients of the three age dummies for the 42 to 44 and the 45 to 49 years' old females do not seem to be statistically significant (we have not

formally tested for joint significance of the difference of the respective coefficients in the employmet and out–of–the labor force hazards).

However, there seems to be no monotonic relationship between the length of the entitlement period and the share of long–term unemployed men without ALH–entitlement, as the comparison of the 12–months' survival rate between the different ALG entitlement groups shows. Furthermore, survival rates for those entitled to ALH seem to be negatively related to the length of the ALG–entitlement period.[16] For females, survival rates differ little between simulations A, B and C. Thus, even long ALG entitlement periods raise survival rates only modestly, and entitlement to ALH does not seem to have any effect on long–term unemployment.

The comparison of simulations B and D show that even a rather large reduction of the incomes replacement ratio of 20% (or a corresponding reduction in the amount of unemployment compensation) has hardly any effect on unemployment survival rates. For females, it would even result in higher survival rates, which seems implausible but, given the small effects, is probably not statistically significant. As a comparison of simulations C and E further show, the expected effects from variations of the level of unemployment compensation are also negligible for those not entitled to ALH after exhaustion of ALG.

5 Summary and Conclusions

The results of the econometric analysis have shown that the entitlement to unemployment benefits increases the duration of unemployment for males, but has very little effect for females. The prolongation of entitlement periods and its extension to successively younger age groups in the eighties has thus increased unemployment durations for males. For females, benefit–entitlement in general has little effect on the duration of unemployment. This result contradicts the popular belief that disincentive effects of the unemployment insurance system are especially severe for females by prolonging "wait unemployment" before they withdraw from the labor force. Although females who are not entitled to unemployment assistance have a higher propensity to leave the labor-force after their benefit-entitlement is exhausted, this entitlement effect has little impact on the duration of unemployment due to the very low level of the female out–of–the labor foce hazard rate. Thus, the result of a strong disincentive effect of unemployment assistance on the duration of female unemployment reported in previous econometric studies for Germany does not seem warranted. The estimation results also show that, for both males and females, marginal reductions of the income–replacement ratio have very little effect on individual unemployment behavior.

16 This result may be related to the "averaging–out" of unobserved heterogeneity in the calculation of survival rates.

The recent reform of the unemployment insurance system successively increases the minimum age for prolonged benefit–entitlement. Given the empirical results of this paper hold up in the future, this should lead to some reduction in the duration of unemployment among older men. On the other hand, there is little reason to believe that the reductions in benefit levels already enacted in the past have had substantial effects on individual re–employment probabilities. The stiffer criteria for what is considered a "suitable job" may in effect even increase the individually conceived income–replacement ratio. Furthermore, these criteria may also discourage efficient job search and thus lead to allocative inefficiency in the labor market. On the other hand, unemployment assistance is still open–ended in principle, and, although means–tested, related to previous earnings. Further reform of the unemployment assistance system therefore seems to be required if the reduction of long–term unemployment has a high priority on the economic policy agenda.

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