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ABSTRACT

Unemployment Benefits and Unemployment Rates of Low-Skilled and Elder Workers in West Germany: A Search Equilibrium Approach*

In this paper we investigate whether the extension of the entitlement to unemployment benefits in the mid 80s can explain the increase in the unemployment rates of unskilled and elder workers in western Germany. To answer this question we estimate a version of the Burdett-Mortensen search equilibrium model and analyze how workers' search behaviour responded to these reforms. We try both nonparametric and fully-parametric estimation methods and identify the cases in which the nonparametric approach cannot be applied. We find that the entitlement reforms are largely responsible for the increase of unemployment among unskilled workers.

JEL Classification: J64, J65

Keywords: search equilibrium, unemployment benefit, parametric estimation, Germany

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1 Introduction

Generous unemployment insurance benefit is one potential reason for the high level of unemployment in European economies. The studies of Nickell (1997) and Siebert (1997) provide evidence for this hypothesis. Moreover, Nickell (1997) and Nickell and Layard (1999) demonstrate that to a large extent there exists a positive dependence between long-lasting entitlements to unemployment benefit and long-term unemployment. The German labour market is a typical representative of the above pattern. Evidence for this is presented for instance in Hunt (1995) or Steiner (1997) who in a reduced form estimation of a duration model show that the length of entitlement is associated with an increase in unemployment duration.

The time profile of the West German unemployment rates shows some well-known and interesting features: From the mid-1980s to the mid-1990s the unemployment rate of low-skilled and aged workers was rising faster than that of the other skill or age groups. Relatively high unemployment rates of low-skilled workers are not only a German phenomenon, but they are particularly high in Germany. Nickell and Layard (1999) present figures of the unemployment rates of low and highly educated male workers for ten OECD countries from the 1970s to the early 1990s[1]: From 1983 to 1986, the unemployment rate of low skilled workers relative to the total unemployment rate was 2.2 in West Germany. For the other countries the ratio ranged from 0.6 to about 1.8, and the average was about 1.4. Moreover, until 1991 to 1993, for Germany this ratio rose by about 18 %, while in the other countries the rise was lower[2]. For elder workers, figures from the OECD Employment Outlook (1996) on standardized unemployment rates show that the West German situation differs substantially from that of many large economies[3]. For instance, in 1983 in France, Italy, Spain, and the US the ratio of the unemployment rate of workers aged 55 to 64 years to the total unemployment rate was below one. Until the year 1990 it rose only for Spain. In contrast, for West Germany this ratio was about 1.16 in 1983 and more than doubled by 1990[4] demonstrating again the highest value and the sharpest increase.

There may exist quite a number of reasons why by the mid 90s West Germany has got a leading position in unemployment of unskilled and elder workers. In the present paper we would like to concentrate on the one which we consider especially important. In the mid 80s the government has introduced a series of reforms aimed at raising the length of entitlement to unemployment insurance benefits. Additionally, the increase in the entitlement length was highest for elder unemployed people. We expect that as a result of these reforms work

[1]The countries are France, Germany, Italy, Netherlands, Norway, Spain, Sweden, United Kingdom, Canada, and the United States.

[2]For the United Kingdom and Sweden there was even a decline. Note, however, that the definition of low skilled workers in Nickell and Layard (1999) is not entirely the same for all the countries. Hence, we have to take these comparisons with some caution.

[3]Note that only until the year 1990 such figures are available for West Germany only.

[4]The numbers are computed from Table B. and Table L., OECD Employment Outlook (1996).

disincentives among elder unemployed workers have significantly gone up. Furthermore, the reforms may have had particularly strong adverse effect on the incentives of low-skilled unemployed workers to return to work.

Considerations of this type are not unfamiliar in the literature that documents the German labour market. For instance, Sinn (2002) argues that changes in the unemployment benefit system can potentially have an adverse effect on the incentives of low skilled workers, because the wages they earn are rather low. For elder workers, longer entitlement to unemployment benefit could be interpreted as a de facto reduction of the (early) retirement age.

In the present paper we try to investigate empirically the impact of the extension of entitlement to unemployment benefits on the unemployment rates of low-skilled and elder workers in West Germany. To do so we study the arrival rates of job offers and exit rates from full-time employment into unemployment in the mid 1980s and mid 1990s for different skill and age groups. As a framework for the analysis we choose the Burdett-Mortensen model of search equilibrium. There are two important reasons for this choice. First of all this framework allows a structural econometric estimation of the theoretical model, i.e., the estimation procedure takes into account all the restrictions imposed by economic theory. Examples of such restrictions could be the endogenously derived functional form of the theoretical wage offer distribution, functional dependence between wage offer and earnings distributions etc. Secondly, through the adjustment of individual search behavior one can to establish the link between the entitlement extension and the dynamics of unemployment rates.

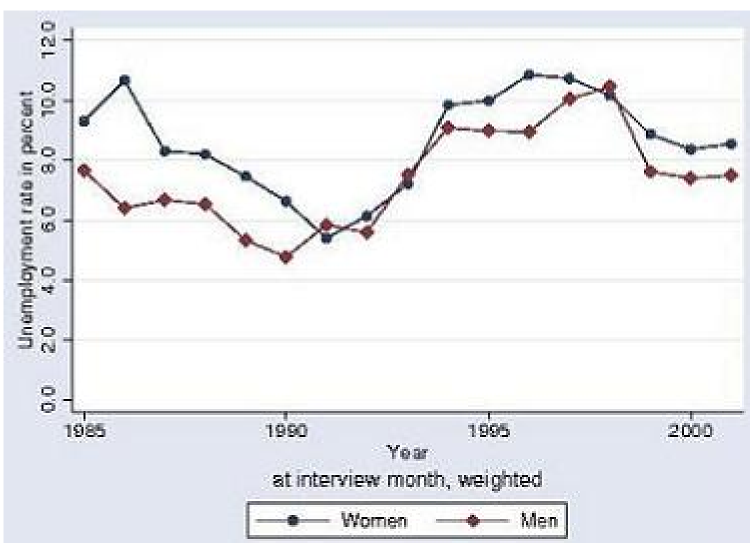
As the models of search equilibrium attract considerable attention in the contemporary labour economics literature we do not present any overview of the theory in this paper. We simply use the existing theoretical results to develop our own argument. For an extensive treatment of the theory interested readers are referred to Burdett and Mortensen (1998), Mortensen (1990) and Bontemps et al. (2000). At the same time we provide the detailed analysis of the two existing structural estimation methods. The primary reason for doing so is that in our analysis we discover that the relatively more attractive nonparametric procedure of Bontemps et al. (2000) may not be always applicable.

The paper is organized as follows. Section 2 motivates our study. Here we describe the evolution of unemployment rates in West Germany for different skill and age groups. We also provide a number of potential explanations of these developments. In Section 3 we present an overview of the necessary theoretical results from the search equilibrium modelling and develop an argument that links entitlement reforms with unemployment rate dynamics. Section 4 discusses the microdata, which we use in our study. Methodological questions on the estimation of empirical search equilibrium models are discussed in Section 5. Here we present an overview of the two existing estimation techniques – nonparametric and parametric, and demonstrate the limitation of the first one. We also discuss some further inference-related issues. Section 6 presents our estimation results and discusses their main economic implications. A summary and some important conclusions are given in Section 7.

2 Motivation

An important feature of the West German unemployment rate is its development for the specific groups of workers. Already in the 1980s unemployment rates of unskilled and elder workers were particularly high relative to overall unemployment rate and they still rose considerably until the mid 1990s. In the figures below we demonstrate this phenomenon. Figure 1 shows the economy wide development of the unemployment rate of men and women from the year 1985 to the year 2001[5]. For both males and females these rates tended to fall from 1985 until, the beginning of the 1990s, reaching levels at around five percent. Thereafter they rose until the mid 1990s, when they ranged form about nine to eleven percent. Figures 2 and 3 show for the same period the unemployment rates of four different skill-groups relative to the overall unemployment for each gender. Figures 4 and 5 repeat this exercise for different age groups. All figures were computed using data from the German Socio-Economic Panel (GSOEP). The samples are limited to workers aged 16 to 64 years. As to qualification, the GSOEP categorizes workers according to "International Standard Classification of Education" (ISCED) code, which takes into account both general schooling and occupational qualifications. We discern four such groups: 1 - inadequately trained or with general elementary schooling, 2 - middle vocational training, 3 - vocational training and college entrance exam or higher vocational training and 4 - higher education.

Figure 1: West German Unemployment Rates by Gender



[5]These figures are based on the Federal Labour Office records. Official unemployment figures are virtually identical to those based on the GSOEP data (see below).

Figure 2: „Relative Unemployment Rates by Skill Groups – Males“

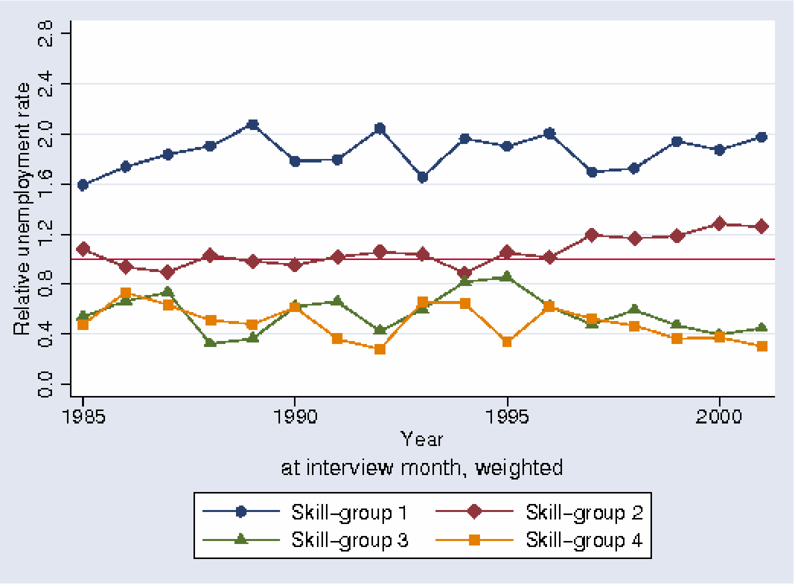


Figure 3: „Relative Unemployment Rates by Skill Groups – Females“

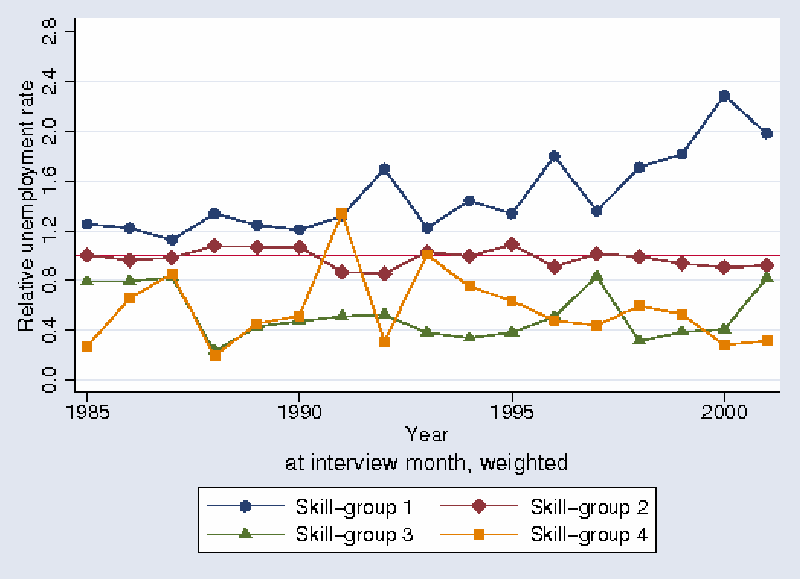


Figure 4: „Relative Unemployment Rates by Age Groups – Males“

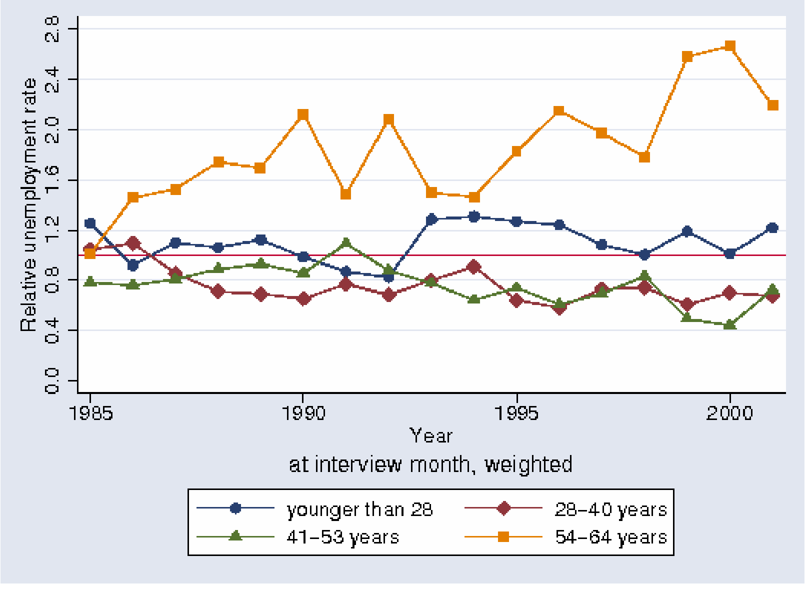


Figure 5: „Relative Unemployment Rates by Age Groups – Females“

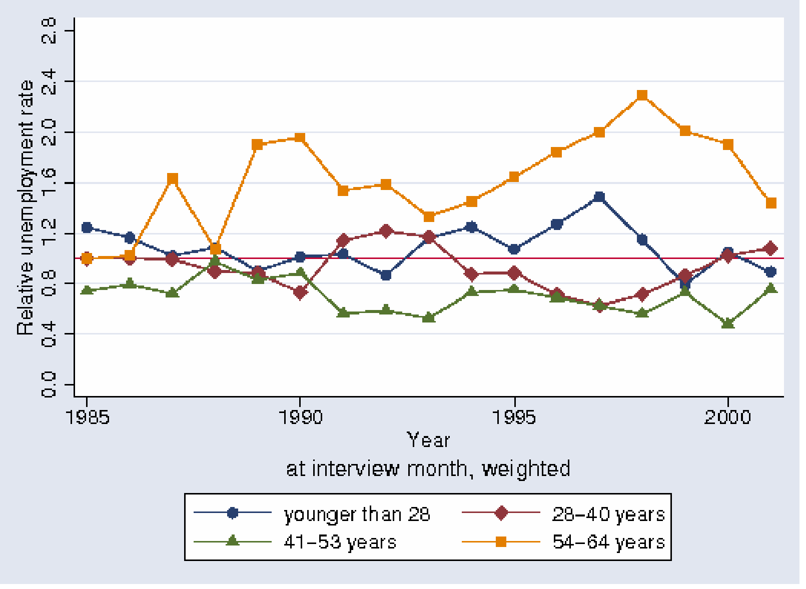


Figure 2 clearly shows that for skill-group 1 the unemployment rate is far above the average unemployment rate in the economy. Since 1988 it is most of the time about twice as high as the average male unemployment rate. Figure 3 displays such relative unemployment rates for women. Its striking feature concerns again the unemployment rate for women in the lowest skill group. From 1985 to 1991 it exceeds the overall unemployment rate by roughly 13 up to 34 %. In contrast, these relative differences are much higher after 1991 ranging from about 22 to 130 %.

For both males and females, there are most of the time no remarkable difference between the economy wide unemployment rate and that of group 2. The unemployment rates of the two highest skill groups are usually somewhat and sometimes considerably lower than those of the entire economy. Taken together, Figures 2-3 demonstrate that unemployment rates of the unskilled workers are the highest among all other skill groups and for women their relative deviation from the economy-wide unemployment rate became particularly high in the 1990s.

The evolution of such differences in the West German unemployment rates was also highlighted by Sinn (2002) who points out that high unemployment rates of the unskilled reflect adverse effects of changes in benefits. Indeed, the standard argument that increased benefit levels may raise the reservation wage and/or decrease job search intensity and therefore induce a higher level of unemployment, may apply. And this can be especially important for low-skilled unemployed workers, whose potential earnings are relatively close to the benefits that they receive. However, in the period under review the replacement rates of the German unemployment benefit system were not increased. So this can hardly explain why unemployment rates of the low-skilled rose considerably from the 1980s to the mid 1990s. At the same time, as we will discuss in more details below, there is one major difference between the mid 1980s and the mid 1990s. In the mid 1990s unemployment insurance benefits have become being paid for a much longer period of time. So it could have been an increase in the entitlement period that may have adversely affected the unemployment rates of low-skilled workers.

Now let's consider the age dimension. Figure 4 displays the development of unemployment rates for several age groups of male workers relative to the total unemployment rate: workers younger than 28, 28 to 40 years, 41 to 53 years and 54 to 64 years old. The most important feature of this figure is the development of the unemployment rate of the eldest age group. In the year 1985 it is still relatively close to the aggregate male unemployment rate. But from 1986 to 1989 it exceeds the aggregate unemployment rate by about 46 to 74 %. From 1995 to 2001, this relative difference ranged even from 79 and 167 %. The unemployment rates of all other age groups deviate much less from the aggregate unemployment rate.

The corresponding relative unemployment rates for women are shown in Figure 5. The evidence on the eldest workers is not exactly the same as for men. Still, the figure shows that the unemployment rate of those aged 54 to 64 years tends to exceed the aggregate unemployment rate in the second half of the

1980s and the first half the 1990s. Its deviation from the overall unemployment rate is remarkable since 1995 and on average higher than in the period before. It is sometimes even more than twice as high as the overall unemployment rate of women.

In Germany two important institutional changes may have contributed to a large extent to the increase in the relative unemployment rates of the aged workers. First of all over the 1980s several benefit reforms tended to raise the potential length of the unemployment insurance (UI) benefits. Table 1 shows the length of UI benefit receipt over several time periods.

Table 1: "Entitlement Length of Unemployment Insurance Benefit"

Work History	Length of UI entitlement during specific periods			
(months)	January 1985 to December 1985	January 1986 to June 1987	July 1987 to March 1997	April 1997 to December 2003
12 − 15	4	4	6	6
16 − 17	4	4	8	8
18 − 19	6	6	8	8
20 − 23	6	6	10	10
24 − 27	8	8	12	12
28 − 29	8	8	14 (age ≥ 42)	14 (age ≥ 45)
30 − 31	10	10	14 (age ≥ 42)	14 (age ≥ 45)
32 − 35	10	10	16 (age ≥ 42)	16 (age ≥ 45)
36 − 39	12	12	18 (age ≥ 42)	18 (age ≥ 45)
40 − 41	12	12	20 (age ≥ 44)	20 (age ≥ 47)
42 − 43	14 (age ≥ 49)	14 (age ≥ 44)	20 (age ≥ 44)	20 (age ≥ 47)
44 − 47	14 (age ≥ 49)	14 (age ≥ 44)	22 (age ≥ 44)	22 (age ≥ 47)
48 − 51	16 (age ≥ 49)	16 (age ≥ 44)	24 (age ≥ 49)	24 (age ≥ 52)
52 − 53	16 (age ≥ 49)	16 (age ≥ 44)	26 (age ≥ 49)	26 (age ≥ 52)
54 − 55	18 (age ≥ 49)	18 (age ≥ 49)	26 (age ≥ 49)	26 (age ≥ 52)
56 − 59	18 (age ≥ 49)	18 (age ≥ 49)	28 (age ≥ 54)	28 (age ≥ 57)
60 − 63	18 (age ≥ 49)	20 (age ≥ 49)	30 (age ≥ 54)	30 (age ≥ 57)
64 − 65	18 (age ≥ 49)	20 (age ≥ 49)	32 (age ≥ 54)	32 (age ≥ 57)
66 − 71	18 (age ≥ 49)	22 (age ≥ 54)	32 (age ≥ 54)	32 (age ≥ 57)
≥ 72	18 (age ≥ 49)	24 (age ≥ 54)	32 (age ≥ 54)	32 (age ≥ 57)

We start with the year 1985, as we will analyze the period from the mid 80s until the year 2000. The length of UI receipt depends positively on work-history in insured employment in the seven years prior to the benefit claim. The first column of Table 1 shows the relevant work-history intervals in months. In how far additional work-history raises the UI entitlement length however also depends on age-limits[6]. These age-limits are shown in brackets next to the entitlement lengths in the other columns of the Table 1. The Table shows the

[6]Note that unemployed people who run out of their UI benefit may still receive unemployment assistance benefit (UA). UA is generally lower than UI benefit and is not time limited. It can be paid until people reach the regular retirement age. Before 1994 the formal replacement rates of the UA benefit were 58 % for parents and 56 % for childless people, while for UI they

rules on the entitlement lengths, which are measured in months, that were in force in the year 1985 (second column), from January 1986 to March 1987 (third column), from April 1987 to March 1997 (fourth column) and from April 1997 to December 2003 (fifth column)[7].

Table 1 demonstrates that except for the last reform, all benefit reforms raised the length of UI entitlement. However, it also shows that this increase was usually limited to some age-groups. The reforms made the benefit system more and more generous for elder workers. With a sufficient work-history, unemployed workers aged older than 54 from July 1987 to March 1997 could be entitled to UI benefits for up to 32 months, while it was only 24 months from January 1986 to June 1987 and 18 months in the year 1985. For workers younger than 42 years instead, the maximum length of UI entitlement was never raised in the 1980s. They could receive UI for no more than 12 months. However, the amount of work-history to achieve this maximum was reduced from 36 in 1985 to 24 by March 1987. Also the maximum entitlement lengths of the people aged 42 to 53 years were raised by the reforms in the 1980s. But the rise for those the aged 54 or older is higher. Hence the incentives to actively search for a job decreased particularly for all those workers aged 54 or older.

Furthermore, UI recipients aged 54 or older faced even less strong incentives to search for a job. The reason is that at the age of 60 they have an option of exit into early retirement. To qualify for early retirement one must have at least 12 months of unemployment in the 18 months prior to reaching this age limit (see Lampert, 1996, p.267). For workers near sixty, this type of early retirement together with the high length of UI entitlement was a major disincentive to search actively for a job.

The second important institutional change is concerned with the availability of elder workers for jobs. Since the reform of the Employment Promotion Act in the year 1986, unemployed workers aged 58 or older could agree with the labour offices to enter early retirement at the earliest possible date (see Steffen, 2003). In turn they need not be (fully) available for the mediation into suitable job offers. This setup further raised the disincentives for elder workers to search for a job. It paved the way into early retirement within the two years prior to reaching the age limit of 60 years. Even though such elder workers are highly protected against dismissal, in practice these rules made their dismissal for both the employer and the employee more attractive. Arnds and Bonin (2002) argue that these reforms enabled employers to change the structure of their staff towards younger workers. And apart from the unemployment benefit, dismissed elder workers could even receive some additional financial support from their last employer.

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were 68 % and 63 %, respectively. In 1994 these replacement rates were cut for UA benefit to 57 % and 53 % and for UI benefit to 67 % and 60 %. However, the UA benefit is means-tested and the benefit level may hence by far lower than the formal replacement rates suggest.

[7] We need to note here that due to some special exemptions the rules displayed by the last column fully affected unemployed workers only two years after their introduction. See Wolff (2003) for details.

Taken these two changes together we should expect a very low incentive for workers aged older than 53 to search for a job.

The discussion above shows that for both unskilled and aged workers we ask one and the same question. We are interested in how far the rise in the UI entitlement length influenced their equilibrium unemployment rate. To answer this question we need a theory that links UI entitlement reforms with equilibrium unemployment rates. We consider such a theory in the next section.

3 Theoretical Results and their Implications for our Analysis

The theoretical Burdett-Mortensen model of search equilibrium formalizes strategic interactions between supply demand sides of the labour market. Representatives of the supply side, i.e. workers, search for better jobs while representatives of the demand side (employers) offer job opportunities. Workers maximize their utility of being employed and employers maximize their profits. The model describes equilibrium flows between the two states of the labour market, namely "employment" and "unemployment" by means of the three key parameters: arrival rate of a job offer to unemployed worker, λ_0, arrival rate of a job offer to employed worker, λ_1 and arrival rate of a match dissolution and return to unemployment, δ. The individual search process in any of these two states is viewed as a repeated drawing of job offers from a certain probability distribution $F(w)$ and acceptance or rejection of the offer after each draw.

Three equations of the model by Burdett and Mortensen (1998) are central to our application. First, Burdett and Mortensen (1998) demonstrate that the steady state level of unemployment is

$$u = \frac{\delta}{\delta + \lambda_0}. \tag{1}$$

Secondly, the model allows calculating the theoretical reservation wages of the agents[8]. Specifically, for any unemployed agent who has an opportunity cost of employment b, which is normally associated with unemployment benefits, the reservation wage becomes

$$R = b + (\lambda_0 - \lambda_1) \int_R^{\overline{w}} \frac{1 - F(w)}{\delta + \lambda_1 (1 - F(w))} \, dw. \tag{2}$$

Additionally, Mortensen and Neumann (1988) argue that considering (2) the arrival rates of job offers, λ_0 and λ_1, can, without loss of generality, be interpreted as search intensities of the participating workers. This interpretation will be quite useful later on.

Finally, Burdett and Mortensen (1998) show that whenever all the employers are homogeneous with respect to their productivity the equilibrium wage offer

[8]This somewhat earlier result is due to Mortensen and Neumann (1988).

distribution takes a form

$$F(w) = \frac{\delta + \lambda_1}{\lambda_1}\left[1 - \sqrt{\frac{p - w}{p - R}}\right] \tag{3}$$

One can further relax the assumption of employer homogeneity which will lead to the wage offer distribution of a form $F(w) = \int F(w|p)d\Gamma(p)$ where $\Gamma(p)$ is a certain productivity distribution that can be also derived endogenously. In the earlier paper Mortensen (1990) derives the theoretical wage offer distribution assuming that $\Gamma(p)$ is discrete. Bontemps et al. (2000) study the case when productivity distribution is continuous. In our application we will estimate the model for both discrete and continuous productivity distributions. Therefore, we reserve the discussion of the issues related to the functional form of the wage offer distribution for Section 5, where we in details deal with the structural econometric estimation of the theoretical model.

Before presenting a mechanism that links the extension of entitlement to UI with equilibrium unemployment rates it will be quite instructive to take a closer look at equations (1) and (2).

Consider first (1). Differentiating u with respect to λ_0 one can see that u is a decreasing function of λ_0. *Ceteris paribus* a reduction in search intensity of unemployed workers leads to an increase in the equilibrium unemployment rate. The opposite is true with respect to δ: A higher incidence of exit into unemployment raises the equilibrium unemployment rate. Equation (1) will be central for our inference.

Now let us look into the dependence of the reservation wage on the adjustment of search behavior. Consider (2). After some algebra (2) can be represented as a function $G(R, \lambda_0, \lambda_1, \delta, b)$, which equals zero. Differentiation of G with respect to its arguments and application of Implicit Function Theorem (see Appendix B) leads to a number of results. First of all, it shows the impact of a rise in b, and hence the impact of increased unemployment benefit levels. Its effect on single parameters, holding everything else constant, is positive for R and δ, negative for λ_0 and ambiguous for λ_1. We expect the effect of increased entitlement length of benefit receipt to go in the same direction. Second, it leads us to the following result:

$$\begin{array}{llll} & \dfrac{\partial R}{\partial \lambda_0} = -\dfrac{\partial G \,/\, \partial \lambda_0}{\partial G \,/\, \partial R} > 0, & & \dfrac{\partial R}{\partial \delta} = -\dfrac{\partial G \,/\, \partial \delta}{\partial G \,/\, \partial R} < 0 \\ \lambda_1 \to \lambda_0 : & \dfrac{\partial R}{\partial \lambda_1} = -\dfrac{\partial G \,/\, \partial \lambda_1}{\partial G \,/\, \partial R} < 0, & \lambda_1 \to \delta : & \dfrac{\partial R}{\partial \lambda_1} = -\dfrac{\partial G \,/\, \partial \lambda_1}{\partial G \,/\, \partial R} > 0 \end{array} \tag{4}$$

for $\lambda_1 \in [\delta, \lambda_0)$.[9] The partial derivatives $\partial R/\partial\lambda_0$ and $\partial R/\partial\lambda_1$ have quite an intuitive interpretation. They establish that unemployed workers who search

[9] Even though the condition $\lambda_1 \in [\delta, \lambda_0)$ might seem to bee too restrictive, indeed it is not so. The reason is that $\lambda_1 \leq \lambda_0$ implies that expected job duration is at least as high as expected unemployment duration. Furthermore $\lambda_1 \geq \delta$ implies that for employed workers with no job-to-job changes so far the probability of finding the next job is at least as high as the probability of being fired. Thus, the values of λ_1 will typically lie in the interval $[\delta, \lambda_0]$.

more actively, i.e have higher λ_0, must have higher reservation wages. Better prospects of promotion on the job reflected by high λ_1 reduce the reservation wage and create an incentive to accept lower wages to get out of unemployment faster (note that each promotion on the job is treated as a job change here). Poor promotion possibilities, i.e. low λ_1 , increase R creating thus an additional incentive to stay longer in unemployment and wait for better times.

The results above make it particularly easy to show how increased entitlement length influences the dynamics of unemployment rates. We would suggest the following argument. Although it is not explicitly stated in (2) which only considers the current benefit level and not its discounted present value, a reasonable interpretation is that an increase in the duration of UI benefit payments increases the value of unemployment. As a result, unemployed workers become more choosy to the arriving wage offers, i.e. the reservation wage of the agents goes up. It should be also true that for any agent the search process is associated with certain disutility generated by search efforts. Therefore, facing the exogenous increase in the value of unemployment, unemployed agents will tend to substitute certain degree of search intensity that brings disutility for some other activities, i.e. search less. Considering (1) we conclude that this will unambiguously rise the equilibrium unemployment rate. This establishes the expected direct effect of the extension on the unemployment rates.

Additionally there may also exist an indirect effect. As we see from (4) the reduction of unemployment search intensity drives the reservation wage down. This counteracts the initial increase in R. As a result of the initial exogenous shock and subsequent unemployed search behavior adjustment we will receive a new equilibrium level of the reservation wage. An interesting (and likely) case arises whenever this new level is higher then the one before the entitlement extension. In this situation the low-productivity firms with limited capacities for productivity enhancement may offer too low a wage to attract any worker. This will result in a higher degree of structural unemployment among lower-skilled workers.

Finally, the contribution to the increase in equilibrium unemployment rates may come from the side of match dissolution parameter δ. Even though in the model this parameter is exogenous and not really related to workers' adjustment behavior, it may still reflect some effects induced by UI extension. In particular, the increased generosity of the UI system may increase the incentives to shirk and as a result increase the match break incidents. From (1) we know that an increase in the frequency of match dissolution incidents leads to the increase in the equilibrium unemployment rate.

The arguments presented above imply that by analyzing empirically the key parameters of the model before and after the reform we will be able to tell whether the entitlement extension indeed contributed to the increase in unemployment rates of unskilled and elder workers as discussed in the previous section. Even though the reservation wage equation in the contemporary formulation of the model does not explicitly include the timing of UI payments, the available econometric procedures are robust to this theoretical shortcoming (see Section 5, page 21 for the discussion). So we will be able to avoid possible specification bias in our structural estimation and find the estimates that are consistent with the most general formalization of UI payment schedules that would consider the duration of benefit payments.

This concludes the summary of main theoretical results and their implication for our paper. After discussing the data used for analysis we proceed with the econometric specification and structural estimation issues. Here the key theoretical results will be revisited.

We also need to notice that the effect of benefit reforms on job search behavior of employed workers (λ_1) is unclear theoretically. In addition, empirical studies by Belzil (1995), (2001) demonstrate that changes in the duration of benefit payments do not significantly alter the length of subsequent reemployment spell. For these reasons, our discussion concentrates on the impact of benefit reforms on the arrival rate of job offers while unemployed and on the match dissolution parameter.

4 The Data

We use data from the German Socio-Economic Panel. It is a longitudinal survey of German households, which was started at 1984 and conducted on the annual basis ever since. We use the information from the 1984 to the 2001 waves. Our analysis is restricted to samples A and B of the GSOEP. Sample A represents households with a household head being a native West German. Sample B represents households whose head belongs to the main groups of foreigners in West Germany. Additionally, we only include respondents aged 16 to 64 years.

4.1 Classification of Workers in the Stock Samples

Estimation of the empirical model of search equilibrium relies on stock sampling. We analyze the stocks of employed and unemployed people from two specific waves: the wave of the year 1986 and the wave of the year 1995. As the extension of entitlements occurred in-between, this should allow us to investigate the reaction in the search behavior of the agents. The choice of years is also influenced by the fact that macroeconomic conditions in these two years were rather similar, i.e. the economy was in roughly the same phase of the cycle. Finally, such a choice minimizes the amount of censored job and unemployment durations in the samples under study.

The samples for these two years were drawn according to the implications of the theoretical model. We analyze agents who are "unemployed" and "full-time employed". We classify workers as "unemployed" if for the modal interview month of the chosen year they reported to be registered as unemployed. For this classification, we use information from the subsequent wave's retrospective labour force status calendarium[10]. In contrast, we classify people as "full-

[10]With the labour force status information in the interview month, the construction of a genuine stock sample at a specific month is not possible, as not all the respondents are

time employed" on the basis of their current labour force status reported at the interview[11].

Due to the restrictions of the theoretical model we did not include part-time employed workers and non-participants in our sample. These should be left out because it is likely that their behavior is different from behavior of the agents represented in the model (see also Koning et al., 1995).

4.2 Unemployment and Job Durations, Exit States

To construct the likelihood function for the model we need to use both wage and duration data. Whenever we observe a change of states, we need to record information about the new state. In the setting of the model, job-to-job changes are also considered as a "change of state".

Unemployment duration is calculated from the retrospective labour force status calendarium of the GSOEP, in which respondents have to provide their labour force status for every month of the previous calendar year.

Apart form completed spells, unemployment spells can also be left-censored, right-censored or both. In our sample, unemployment spells are left-censored mainly because a respondent was already unemployed before he/she first filled in the labour force status calendarium. The main reasons for right-censoring is either that a respondent temporarily did not respond to the GSOEP or due to the fact that the respondent completely dropped out of the panel study. Finally, some of the spells did not terminate before the end of our observation period.

The information on the beginning and end of a job spell is more difficult to obtain. There are various pieces of information on the job history of individuals that the GSOEP collects retrospectively. First of all respondents who state that they are currently employed provide the calendar year and the calendar month of the start of the job. Provided that there is a job change, employed respondents have to state in which calendar month this job event took place and indicate the type of job event: first job, new employer, self-employment, change within the firm, company takeover, or return to work. This information allows us to identify when the jobs of the individuals in the current employment stock started[12].

interviewed in the same month of a year.

[11]The reason is that for employed people we need the wage in the current job, which is only available for the month prior to their interview. There were also cases where people report in the interview month to be full-time employed, while in the subsequent wave their retrospective labour force status for the modal interview month of the previous year is registered unemployment. These people were classified as registered unemployed in our samples.

[12]If a job spell of a respondent in our employment stock was already in progress at the interviews of previous waves we use the related job start information of these previous waves to determine the respondents' start of the job. In case for one and the same job a person reports different job starting dates over different waves of the GSOEP and there was a modal calendar start, the job start was set to this modal value. If there were no such modal calendar start, the job start is taken as reported in the wave, in which the person's current job was first observed. For individuals where we have no information on the calendar month and the type of event that lead to a job start, we used the employer start information.

Table 2: "Descriptive Statistics of Event History Data for the Two Stock Samples[1]"

	1986			1995		
	Full Sample	Elder	Low Skilled	Full Sample	Elder	Low Skilled
Number of Individuals	4873 [1.000]	571 [1.000]	1401 [1.000]	4030 [1.000]	637 [1.000]	933 [1.000]
Employed:	4551 [0.934]	518 [0.907]	1272 [0.908]	3681 [0.913]	533 [0.837]	780 [0.836]
Unemployed:	322 [0.066]	53 [0.093]	129 [0.092]	349 [0.087]	104 [0.163]	153 [0.164]
Employed Agents:	4551 [1.000]	518 [1.000]	1401 [1.000]	3681 [1.000]	533 [1.000]	780 [1.000]
Uncensored observations with:						
job → job transition:	706 [0.155]	6 [0.012]	138 [0.108]	423 [0.114]	7 [0.013]	49 [0.063]
job → unemployment transition:	385 [0.085]	42 [0.081]	157 [0.123]	277 [0.075]	68 [0.128]	101 [0.129]
mean time spell between two states [job duration]:	139.95	248.94	150.53	106.82	248.33	129.35
(std. deviation):	(115.44)	(138.18)	(113.66)	(101.08)	(141.28)	(115.29)
Censored observations[2]						
a) left-censored durations only						
with job → job transition:	97 [0.021]	5 [0.010]	24 [0.019]	22 [0.006]	1 [0.002]	3 [0.004]
with job → unemployment transition:	74 [0.016]	16 [0.031]	28 [0.022]	16 [0.004]	2 [0.004]	1 [0.001]
b) right-censored durations only:	2898 [0.637]	361 [0.697]	784 [0.616]	2857 [0.776]	445 [0.835]	603 [0.773]
c) both left- and right-censored durations:	391 [0.086]	88 [0.170]	141 [0.111]	86 [0.023]	10 [0.019]	23 [0.029]
Mean time spell [both uncensored and censored]:	168.85	236.69	158.99	155.05	263.87	161.88
(std. deviation):	(136.41)	(167.16)	(123.92)	(118.89)	(143.26)	(117.13)
Unemployed Agents:	322 [1.000]	53 [1.000]	129 [1.000]	349 [1.000]	104 [1.000]	153 [1.000]
Uncensored observations (u → j transition):	116 [0.360]	3 [0.057]	42 [0.326]	105 [0.301]	4 [0.038]	38 [0.248]
mean time spell between two states [unempl. duration]:	14.18	11.67	14.91	20.81	14.50	19.92
(std. deviation):	(18.94)	(4.16)	(12.57)	(22.95)	(8.66)	(14.30)
Censored observations						
a) left-censored durations (u → j transition) only:	14 [0.043]	-	11 [0.085]	3 [0.009]	-	1 [0.007]
b) right-censored durations only:	160 [0.497]	33 [0.623]	58 [0.450]	226 [0.648]	96 [0.923]	106 [0.693]
c) both left- and right-censored durations:	32 [0.099]	17 [0.321]	18 [0.140]	15 [0.043]	4 [0.038]	8 [0.052]
Mean time spell [both uncensored and censored]:	29.20	45.51	34.95	35.43	47.25	40.92
(std. deviation):	(33.02)	(37.02)	(36.07)	(33.35)	(36.75)	(36.25)

[1] Duration data in Months. Share of the sample in brackets.
[2] In the framework of the theoretical model a spell with transition to non-participation qualifies as right-censored with unobserved exit state.

To define the calendar end of the jobs we tracked the job start and end information as well as the labour force status information at the interviews over the waves that followed the year of the stock sampling. The calendar end of job spells is set to the first reported job end in subsequent waves or to the first reported job start due to a within firm job change.

Similar to unemployment spells, job spells can be left-censored, right-censored or both and we proceed in similar fashion to the treatment of unemployment spells[13]. For all spells where we could observe the calendar end, we determined the exit state. In case of the unemployment spells, using the retrospective labour force status calendarium information, we determined whether they ended in full-time employment or in any other labour force state. In case of the job spells, we used the labour force status calendarium and job events information to see whether a job ended by transition to unemployment, another job or non-participation.

Table 2 provides a summary statistics for employment and unemployment spells in the resulting stock samples. Additionally it shows the percentage of spells that are completed, left-, right- and both left- and right-censored.

It is important to notice here that we treat spells that terminate by an exit into non-participation as right-censored (see, for instance, Koning et al., 1995 and van den Berg and Ridder, 1998). The reason is that the theoretical model does not have states other than "full-time employment" and "unemployment". Because of this, we observe a rather large share of right-censored durations.

4.3 Wages and Benefits

The final piece of information necessary for the estimation of the model is earnings. We use the data on net wages provided by the GSOEP. Individuals who are employed at their interview provide the monthly net wage in the month prior to the interview. For the stock sample of job spells we use the wage information that the respondents stated at the year for which the sample is drawn. For the stock sample of unemployment spells we use the first reported wage after the end of unemployment, provided that the unemployment spell is not right-censored. All wage are deflated by the West German consumer price index at prices of 1998.

Having once estimated the model we compute the reservation wages predicted by the theory. To do this we need to know either the true benefit receipt

[13] For some spells in our sample we cannot determine the exact calendar start of their job, but only the year of the job start. These were considered as left-censored with calendar start being December of that year. Likewise we cannot always determine the exact end of the job spell. One reason is that for at least one of the subsequent waves a respondent was not interviewed prior to the termination of his/her job. In this case the interview month of the wave before determines the right-censored job end. Of course this rule applies to all jobs that are still in progress by the interview month of the last available wave of the GSOEP. Additionally, right-censoring applies if without providing job end information, some respondents either stated not to be employed or indicated a start of a different job in one of the waves that follow. Again in these cases the right-censored job end is set to the interview date of the wave before, i.e., the last month for which we have a valid observation of the job.

or a potential benefit level of our sample members. We considered three types of benefits: unemployment insurance benefits (UI), unemployment assistance benefits (UA) and welfare benefits (WB). UI and UA benefits are determined by formal replacement rates. Though the UA benefit is means-tested and hence may be much lower than the formal replacement rates suggest. A means-test also applies to the WB.

UI and UA Benefit Levels: For unemployed people, we set the UI or UA benefit at the level that they received at the date, where the stock sample was drawn. These benefit levels are reported retrospectively in the subsequent wave. The respondents provide the monthly average benefit level for the months in which they received the benefit during the previous calendar year. There are also a few unemployed individuals in our sample who receive a training benefit but no unemployment benefit. For all full-time employed individuals, we set their unemployment benefit level to the value of the replacement rate of the UI benefit multiplied by their net wage.

Welfare Benefits: Welfare benefits are means-tested. We did not attempt to simulate the means-test for the households in our sample in order to compute a welfare benefit level. However, we used information on social benefits provided by the household heads for the households in which the respondents live. We took into account receipt of rent subsidy payments (Wohngeld), continuous aid for living expenses (laufende Hilfe zum Lebensunterhalt) as well as social welfare assistance to meet special contingencies in life (Hilfe in besonderen Lebenslagen). For the year 1995 the GSOEP questionnaire provides a variable that records monthly amount of such benefits received by a household in the interview month. We assume that the sum of these amounts divided by the household size represents the potential social benefit that is available to a member of the household.

In the 1986 wave the GSOEP did not collect information on the current level of welfare benefits. However, such benefit levels and months of benefit receipt were collected retrospectively in the wave of 1987. The household questionnaire asked whether people in the household received these social benefits in 1986. Two additional questions also provide the number of months and the average monthly amount of each of these welfare benefits. From this information we computed monthly welfare benefit levels of the respondents in our stock sample of 1986. In Appendix A we describe in details the computation and introduce some related assumptions.

The total benefit level is computed by the sum of the unemployment benefit and per capita welfare benefits. All benefit information is also deflated by the West German consumer price index with price base being the year 1998.

5 Structural Econometric Model of Search Equilibrium

5.1 The Likelihood

A short summary presented in this subsection relies on the distributional properties reviewed by Lancaster (1990) and certain theoretical results developed by Burdett and Mortensen (1998).

The process that governs the arrival of job offers in the theoretical model is Poisson (θ). Therefore, the waiting time between any two adjacent events is distributed exponentially with parameter θ. However, due to the non-randomness of the sample of job and unemployment durations (see Ridder, 1984), this property cannot be applied directly. We follow Ridder (1984) and analyze instead a joint distribution of elapsed (t_e) and residual (t_r) durations of a spell. On the distribution of elapsed duration it is known that certain time t_e ago there was a renewal of states and since then an individual spent at least t_e in a new state. Renewal probability for Poi(θ) is shown to be equal to θ. On the distribution of residual duration our knowledge is that given a certain elapsed time t_e an individual spends in his current state additional time t_r ($t_r > 0$). Therefore the appropriate densities are:

$$\begin{array}{ll} \text{Elapsed:} & f(t_e) = \theta e^{-\theta t_e}, \\ \text{Residual:} & f(t_r|t_e) = \theta e^{-\theta t_r},\ t_r > 0, \end{array}$$

$$\text{Joint:} \quad f(t_e, t_r) = \theta^2 e^{-\theta(t_e+t_r)}, t_r > 0. \tag{5}$$

Denote the arrival rate of job offers to unemployed and employed workers as λ_0 and λ_1 respectively. Then, using the property of the exponential distribution, the exit rate from unemployment is the arrival rate of job offer: $\theta_u = \lambda_0$. For employed individuals the hazard rate from the current job is a sum of the transition intensity to a job that pays a higher wage and the transition intensity to unemployment: $\theta_e = \lambda_1 [1 - F(w)] + \delta$, where $F(w)$ is an unobserved wage offer distribution. Substitution of θ_u and θ_e into (5) will give the correctly specified density of job and unemployment durations.

To complete the formulation of individual contributions to the likelihood we consider separately the cases of employed and unemployed individuals:

1. *For Unemployed:* In equilibrium the probability of encountering an unemployed agent is $\delta (\delta + \lambda_0)^{-1}$. In case the transition to the job is observed we know the offered wage hence record a realization of the wage offer distribution $f(w)$.

2. *For Employed:* In equilibrium the probability of encountering an agent employed at given wage is $\lambda_0 (\delta + \lambda_0)^{-1} g(w)$. In case the transition to the next state is observed we record the destination state. The probabilities of exit to unemployment and to next job are respectively:

$$\pi_{j\to u} = \frac{\delta}{\delta+\lambda_1\bar{F}(w)} \qquad \text{and} \qquad \pi_{j\to j} = \frac{\lambda_1\bar{F}(w)}{\delta+\lambda_1\bar{F}(w)}.$$

Defining for the convenience of notation $\bar{F}(w) = 1 - F(w)$ and for the convenience of subsequent estimation $\kappa_0 = \lambda_0/\delta$, $\kappa_1 = \lambda_1/\delta$ we get the following likelihood contributions of unemployed ($\mathcal{L}_u$) and employed ($\mathcal{L}_e$) individuals:

$$\mathcal{L}_u = \frac{1}{1+\kappa_0}\left[\delta\kappa_0\right]^{2-d_r-d_l} e^{-\delta\kappa_0[t_e+t_r]}\left[f(w)\right]^{1-d_r}, \tag{6}$$

$$\mathcal{L}_e = \frac{\kappa_0 g(w)}{1+\kappa_0}\left[\delta\left(1+\kappa_1\bar{F}(w)\right)\right]^{1-d_l} e^{-\delta\left(1+\kappa_1\bar{F}(w)\right)[t_e+t_r]}\left[\left[\delta\kappa_1\bar{F}(w)\right]^{d_t}\delta^{1-d_t}\right]^{1-d_r} \tag{7}$$

In (6) and (7) $d_l = 1$, if a spell is left-censored, 0 otherwise; $d_r = 1$, if a spell is right-censored, 0 otherwise; $d_t = 1$ if there is a job-to-job transition, 0 otherwise. Since all labor suppliers are assumed to act independently, the total likelihood is a product of all individual contributions.

5.2 Nonparametric Estimation and Its Limitations

Define the observed earnings density and distribution as $g(w)$ and $G(w)$ respectively. Then using the steady state identities

$$\bar{F}(w) = \frac{1-G(w)}{1+\kappa_1 G(w)} \quad \text{and} \quad f(w) = \frac{(1+\kappa_1)}{\left[1+\kappa_1 G(w)\right]^2} g(w) \tag{8}$$

implied by the theoretical Burdett-Mortensen model. Bontemps et al. (2000) propose the following 3-step estimation procedure. In a first step $g(w)$ and $G(w)$ in (8) are estimated nonparametrically. In the second step the expressions in (8) are substituted into (6) and (7) and the likelihood function is maximized with respect to $\{\kappa_0, \kappa_1, \delta\}$. In the third step the equilibrium productivity levels

$$p = K^{-1}(w) = w + \frac{1+\kappa_1 G(w)}{2\kappa_1 g(w)} \tag{9}$$

and productivity density

$$\gamma(p) = \frac{2\kappa_1(1+\kappa_1)g(w)^3}{3\kappa_1 g(w)^2[1+\kappa_1 G(w)]^2 - g'(w)[1+\kappa_1 G(w)]^3} \tag{10}$$

are calculated. Bontemps et al. (2000) notice that the third step is possible only if the model is well specified with respect to the equilibrium productivity distribution, i.e., if $3\kappa_1 g(w)^2 - g'(w)[1+\kappa_1 G(w)] > 0$. In case this condition is not satisfied they suggest to perform the second step of the procedure under this theoretically implied constraint, which can be conveniently rewritten as

$$\kappa_1\left[3g(w)^2 - g'(w)G(w)\right] > g'(w) \quad \{w : g'(w) \geq 0\}.[14] \tag{11}$$

[14] Notice that if $g'(w) < 0$ productivity density $\gamma(p)$ is always positive.

In the applications of the proposed methodology so far (see, for instance, Bontemps et al., 2000) the constraint in (11) was never violated. The present paper, to the contrary, faces the opposite case. Therefore, we follow the suggestion of Bontemps et al. (2000) and on the second step maximize the likelihood with respect to (11). It turns out, however, that the constrained optimization may not always be feasible. To see this notice that for some values of w the term $3g(w)^2 - g'(w)G(w)$ on the l.h.s. of (11) can be negative. This is exactly the case when we observe clusters of those who earn very high wages. Such clustering is represented by a bump far on the right tail of the estimated earnings density. Whenever such bump occurs, $g'(w)$ is greater than zero and at the same time $G(w) \to 1$ and $g(w) \to 0$. So the value of $g(w)$ may be too small to make the whole term $3g(w)^2 - g'(w)G(w)$ positive. In this situation the constraint yields

$$\kappa_1 < \min_{\{w\}} \frac{g'(w)}{3g(w)^2 - g'(w)G(w)} < 0 \quad \{w : g'(w) \geq 0\} \tag{12}$$

As a result there is no κ_1 that can satisfy (11), since κ_1 is always greater than zero. We will refer to this case as to *"constraint inconsistency"*.

In the opposite situation when $3g(w)^2 - g'(w)G(w) > 0$ the constraint is formulated as

$$\kappa_1 > \max_{\{w\}} \frac{g'(w)}{3g(w)^2 - g'(w)G(w)} > 0 \quad \{w : g'(w) \geq 0\} \tag{13}$$

and the second step indeed returns an appropriate estimate of κ_1. A typical example for this case will be the left tail of earnings distribution, where $g(w)$ increases, but its values are high enough to insure that $3g(w)^2 - g'(w)G(w) > 0$ holds true $\forall w : g'(w) \geq 0$.

Since we find that constraint inconsistency is a pure earnings data property we suggest

$$\texttt{sign}\left[3g(w)^2 - g'(w)G(w)\right] \tag{14}$$

as a quick check for applicability of the nonparametric 3-step procedure.

In our application we face the case of an inconsistent constraint, i.e., we cannot apply the nonparametric estimation procedure directly. We also warn from using oversmoothing of the kernel density estimator in order to achieve consistent constraint. By oversmoothing one can indeed get a strictly decreasing right tail with minor changes of the curvature of the rest of estimated density. However, from (13) it can be seen that by manipulating the magnitude of the bandwidth one arbitrarily fixes the value of the constraint. This will generate bias in the estimated κ_1.

5.3 Parametric Estimation of the Model

Facing the situation of constraint inconsistency we cannot perform the nonparametric estimation of the model any longer. So we need to use the alternative parametric procedures. In other words we have to impose certain assumptions concerning the form of either earnings or productivity distribution.

5.3.1 Parametric Assumptions on the Earnings Distribution

The easiest way to avoid an inconsistent constraint is to assume some parametric form of $g(w)$, instead of using its nonparametric estimate. Inspecting the shape of the kernel estimate of earnings distribution the most natural suggestion would be that $g(w)$ is distributed lognormally with parameters μ and σ. We estimate the model under this assumption and find that indeed (11) is never violated. However, calculating (9) we discover that it violates the requirement that offered wage is a monotone increasing function of productivity.[15] This generates an improper estimated productivity density and implies the necessity of imposing parametric assumptions on the productivity density directly.

5.3.2 Parametric Assumptions on the Productivity Distribution

This approach differs from the one in Section 5.2 because now the productivity appears in the likelihood function explicitly. This happens because for $g(w)$, $f(w)$ and $F(w)$ instead of nonparametric estimates we take the theoretical expressions that constitute a part of the equilibrium solution of the model (for derivations of the theoretical earnings and offer distributions see Mortensen, 1990; Burdett and Mortensen, 1998). These expressions depend on both search intensity parameters and productivity parameter p. It is theoretically demonstrated that the dispersion of p leads to a decreasing right tail of the theoretical earnings density, which matches the empirical regularity.

There exist two approaches to estimate the model in which the productivity level is assumed to have a certain probability distribution. The first approach is developed by Koning et al. (1995) and Christensen et al. (2000). Koning et al. (1995) assume that the productivity parameter is distributed lognormally with parameters μ and σ and consider marginal likelihood, where marginalization is made with respect to unknown productivity p. The likelihood function is then maximized with respect to $\{\kappa_0, \kappa_1, \delta, \mu, \sigma\}$. Christensen et al. (2000) rather suggest that the unknown productivity parameter p is multiplied by the term $\exp\{\eta\}$, where $\eta \sim N(0, \sigma^2)$. The likelihood function in their application is maximized with respect to $\{\kappa_0, \kappa_1, \delta, p, \sigma\}$.

The second approach to the specification of the productivity distribution is due to Bowlus et al. (1995), (2001). It assumes that the productivity distribution is discrete rather then continuous. Moreover the exact form of the distribution is *a priori* unknown. So its support points and corresponding probability mass values are to be estimated together with the structural parameters of the model. In this sense the approach minimizes distributional assumptions on p and becomes conceptually equivalent to the semiparametric one. Therefore in the present paper we choose this way to estimate the model.

Mortensen (1990) shows that whenever the productivity distribution is discrete and has Q points of support the theoretical wage offer distribution has

[15] Monotonicity of offered wages as a function of productivity follows from *Proposition 10* of Bontemps et al. (1997), which is a generalization of Burdett and Mortensen (1998) finding that more productive firms pay higher wages.

Q kinks each of them corresponding to the highest wage paid by a p_j-type employer ($j = 1, ..., Q$). Moreover, firms with higher productivity pay higher wages, which implies that the lowest wage paid by p_j-type employer (w_{L_j}) is equal to the highest wage of p_{j-1} -type employer ($w_{H_{j-1}}$). Consequently the ranking $w_{H_{j-1}} < w_{H_j}$ $\forall j = 1, ..., Q$ applies. Mortensen (1990) derives the following expression for the theoretical wage offer distribution with Q distinct productivity types

$$F(w) = \frac{1+\kappa_1}{\kappa_1}\left[1 - \frac{1+\kappa_1(1-\gamma_{j-1})}{1+\kappa_1}\sqrt{\frac{p_j - w}{p_j - w_{H_{j-1}}}}\right], \quad w \in (w_{L_j}, w_{H_j}] \tag{15}$$

$j = 1, ..., Q$, ($w_{L_1} = R$, $w_{H_Q} = \overline{w}$). In the expression above γ_j indicates the probability mass attached to the productivity level p_j ($\gamma_0 = 0$, $\gamma_Q = 1$). Differentiating (15) with respect to w and using (8) we can show that the theoretical wage offer and earnings densities are

$$f(w) = \frac{1+\kappa_1(1-\gamma_{j-1})}{2\kappa_1}\frac{1}{\sqrt{p_j - w}\sqrt{p_j - w_{H_{j-1}}}}, \tag{16}$$

$$g(w) = \frac{1+\kappa_1}{2\kappa_1(1+\kappa_1(1-\gamma_{j-1}))}\frac{1}{p_j - w}\sqrt{\frac{p_j - w_{H_{j-1}}}{p_j - w}} \tag{17}$$

$w \in (w_{L_j}, w_{H_j}]$, $j = 1, ..., Q$. Substitution of (15)-(17) into the expressions for likelihood contributions (6)-(7) gives us the likelihood function with unknown parameters $\left\{\kappa_0, \kappa_1, \delta, \gamma_1, ..., \gamma_{Q-1}, p_1, ..., p_Q, R, w_{H_1}, ..., w_{H_{Q-1}}, \overline{w}\right\}$.

Following Mortensen (1990) it is possible to represent productivity levels as a function of wage cuts w_{H_j}, probability mass points γ_j and structural parameters, namely

$$p_j = \frac{w_{H_j} - B_j w_{H_{j-1}}}{1 - B_j}, \tag{18}$$

where $B_j = \left[\frac{1+\kappa_1(1-\gamma_j)}{1+\kappa_1(1-\gamma_{j-1})}\right]^2$. Moreover he shows that there holds an equality $F(w_{H_j}) = \gamma_j$ $\forall j = 1, ..., Q$. On the basis of this information Bowlus et al. (1995), (2001) develop an iterative procedure for estimation of the unknowns of the model.

The estimation procedure can be represented as follows. Initially we focus on the subsets $\theta_1 = \{R, \overline{w}\}$, $\theta_2 = \left\{w_{H_1}, ..., w_{H_{Q-1}}\right\}$ and $\theta_3 = \{\kappa_0, \kappa_1, \delta\}$ of the parameter space. As an estimator of θ_1 Bowlus et al. (1995), (2001) suggest minimum and maximum of the observed wage sample: $\hat{\theta}_1 = \{w_{\min}, w_{\max}\}$. The authors argue that sample minimum and maximum are asymptotically MLEs of R and $\overline{w}$. This fact is especially useful because the estimator $\hat{R} = w_{\min}$ allows us to get the consistent estimate of R even when the timing of UI benefit payments is not explicitly introduced in the model. More generally, application of this estimator contributes to avoiding the case in which the likelihood function has non-standard properties and cannot be maximized by gradient methods (see also Kiefer and Neumann, 1993).

Given the above $\hat{\theta}_1$ the estimation procedure is stepwise:

1. On the first step given $\hat{\theta}_1$ and starting values for $\{\kappa_0, \kappa_1, \delta\}$ and $\{\gamma_1, ..., \gamma_{Q-1}\}$ we estimate the set of wage cuts θ_2
2. On the second step estimates $\hat{\theta}_2$ and starting values for $\{\kappa_0, \kappa_1, \delta\}$ and $\{\gamma_1, ..., \gamma_{Q-1}\}$ are used to calculate (18), substitute it into (6)-(7) and maximize the likelihood with respect to $\theta_3 = \{\kappa_0, \kappa_1, \delta\}$.
3. Using estimated productivity levels and structural parameters $\hat{\theta}_3$ we use (8) to calculate the implied point mass probabilities γ_j and return to the first step.

With respect to this procedure Bowlus et al. (1995), (2001) notice that since (15) has Q kinks in w_{H_j} the likelihood function is discontinuous in $\theta_2 = \left\{w_{H_1}, ..., w_{H_{Q-1}}\right\}$. To facilitate the estimation of wage cuts they derive the following useful property of θ_2.

Theorem 1 *Let $\{W_N\}$ be a set of observed wages from a sample of size N drawn from the distribution specified in (15). Then the maximum likelihood estimator for $\theta_2 = \left\{w_{H_1}, ..., w_{H_{Q-1}}\right\}$ is a $Q-1$ element of $\{W_N\}$.*

Proof. See Bowlus et al. (2001). ■

To estimate the discontinuity points in the first step Bowlus et al. (1995), (2001) suggest a simulated annealing algorithm as introduced by Kirkpatrick et al. (1983). Useful hints for practical implementation of the algorithm could be found in Goffe et al. (1994). On the smooth second step the likelihood is maximized by standard methods.

The number of mass points in the productivity distribution is treated as unknown. We start from the homogeneous case ($Q = 1$) and add points one by one. The exact distribution of the likelihood ratio in this particular case is also not known. Bowlus et al. (2001) propose a quasi-LRT test $V = -2\left(\log L_{j-1} - \log L_j\right) < \chi^2(1)$. Performing a simulation study they notice, however, that this criterion can be applied for small Q only, because the critical region increases with Q. Therefore we make our choice of the number of mass points on the basis of information criteria (Consistent AIC, SBC; see Appendix C, Tables C1-C2).

Following Bowlus et al. (2001) we finally notice that the asymptotic distribution of the resulting estimates of $\{\kappa_0, \kappa_1, \delta\}$ is too cumbersome to be derived analytically. The authors propose bootstrap to find the correct confidence intervals for the estimated parameters. However, the bootstrap and MLE standard errors in the paper of Bowlus et al. (2001) are not significantly different from each other. At the same time, because of the simulated annealing step, for large Q bootstrap of the suggested stepwise estimation procedure is an extremely computationally intensive task. Therefore we base our inference on the standard ML covariance estimates (i.e. inverted expected negative Hessian).

6 Estimation Results and Discussion

6.1 General Issues

As discussed in Section 2, the main aim of the present paper is to investigate whether the UI reforms of the late 1980s which extended the entitlement to the receipt UI benefits have significantly contributed to the increase in unemployment rates of low skilled and elder workers (see Figures 1-4).

To link the entitlement extension with subsequent dynamics of unemployment rates we use a theoretical search equilibrium framework. In Section 3 we argue that such an extension should negatively effect the search intensity of the unemployed λ_0 and increase the exit rate to unemployment δ. According to the theory both of these factors must result in an upward shift of the equilibrium unemployment rate. We also stipulate that the adjustment dynamics of search intensity induced by the reforms may change the reservation wage level, which can contribute to an increase in structural unemployment.

In order to analyze the effects of the reform quantitatively we use the structural estimation methods described in Section 5 and empirically analyze to what extent the arrival rate of job offers while unemployed or employed as well as the rate of job loss and reservation wages changed from the mid 1980s to the mid 1990s.

With respect to skills we expect that the increased generosity of the benefit system will affect the arrival rates of job offers and reservation wages of the low skilled workers more than those of the high skilled ones. One line of argument to support such a hypothesis is that the value of household production of skilled and unskilled workers is about the same. At the same time the ratio of benefits plus value of household production while unemployed to the potential wages plus the value of household production while employed is much higher for the low-skilled than for the skilled workers. Thus extending the entitlement length may affect the job search behavior of the low-skilled workers more than that of skilled workers[16].

Given the nature of the unemployment benefit system, other factors might lead to a different conclusion. In particular, the increased length of UI entitlement could be also be more important for the search behavior of high-skilled rather than low-skilled workers. The main reason is that low-skilled/low-wage workers are very likely to pass the means-test for UA receipt, once they run out of their UI benefit. In contrast high-skilled/wage workers are unlikely to pass it. As UA receipt is not time limited, low-skilled workers as opposed to high-skilled ones are more likely to receive unemployment benefit without a time limit. Therefore, the effects of the extension of UI receipt may also adversely influence the search incentives of the skilled rather than the unskilled. We will

[16]One should note, though, that such differences may also be caused by other influences in the labour market. For instance it could be a skill-biased technological change that could decrease the relative demand for low-skilled workers. As a result we may expect a reduction in the arrival rate of job offers to unskilled relative to skilled workers as well as an increase in their relative rate of job loss.

have to sort out these different possible effects empirically.

As shown in Section 2, with respect to age, the rise in the potential duration of UI benefit receipt is higher the older the workers are. So we expect that search intensity of elder workers falls faster than that of younger workers. The opposite should be true for the reservations wages of the younger age groups.

We also have to notice that even though the comparison of the reservation wages predicted by the model using (2) may provide us with rather useful results we have only limited possibility to interpret it. The reason is that the reservation wage calculation relies on a formula that contains the opportunity cost of employment b. We set this quantity equal to the benefit level received by the agents. By doing so we do not explicitly take into account other possible contributions to b such as household production, black market work etc. In this way we may underestimate the magnitude of the reservation wage. Furthermore, reservation wages predicted by means of (2) will also ignore the actual change in the entitlement period. This shortcoming hampers the inference about the possible contribution to the unemployment dynamics. Limited possibility to interpret the predicted reservation wages also prevents us from inferring much from the changes in employed search intensity λ_1. The latter is not a big obstacle, though, since we know that the dynamics λ_1 is not caused by the entitlement extension reforms.

Let us now turn to estimation results. These are reported in Appendix C. We did not carry out the estimation separately for men and women. The reason is that when we estimate the models for different skill or age groups, the sample sizes would become very small.

First of all we estimate the search equilibrium model for the whole economy. For the reasons explained in Section 4.3 we base our choice of the number of support points of the productivity distribution on information criteria (Consistent AIC, Schwarz). The specification selection procedure is demonstrated in Tables C1-2. We estimate the model for the whole economy primarily for analyzing its fit to the data. There is only one criterion that tells us about the goodness of fit. It is the discrepancy between the predicted theoretical earnings distribution and the nonparametric estimate of earnings distribution obtained from wage data. From Figures C1-2 we can see that for both 1986 and 1995 samples this fit is very close. This assures sound inference from the obtained estimation results.

We can go a bit further and compare the equilibrium unemployment rates predicted by the model with actual unemployment rates reported in Table 2. Using (1) and the results reported in Table C3 we find that the model predicts unemployment rates of 7.3 % and 9.7 % for 1986 and 1995 samples respectively. From Table 2 we see that the share of unemployed workers in 1986 and 1995 was 6.6 % and 8.7 % respectively. Again, this reflects a fairly good fit of estimated model to the data. Therefore we conclude that the chosen model can provide a reliable information for our subsequent analysis.

In what follows we estimate the model for different skill and age groups. As before we treat the number of points of increase in the productivity distribution as unknown. We start with the homogeneous model and, adding the support points one by one, use information criteria to find the best specification. Estimation results for skill groups are reported in Table C4 and for age groups in Table C5. Using the fact that $\lambda_i = \kappa_i \delta$ $(i = 0, 1)$ in Tables C4-5 we report the results already in the form of arrival rates of job offers. Our attention will be mainly focused on the results for the least-skilled workers (Table C4, group 1) and elder workers (Table C5, group 4).

6.2 Estimation Results by Skill Group

Table C4 of the Appendix presents our estimation results for different skill groups in both the year 1986 and 1994. The least skilled workers of group 1 are those who went through inadequate training or only general elementary education. We find that the predicted equilibrium rate of unemployment in this group of workers goes up from 10.3 % in 1986 to 15.1 % in 1995. These results somewhat underpredict the true rise from 9.2 % to 16.4 % for our sample (see Table 2). However this underprediction is minor. The results also demonstrate a considerable rise of the unemployment rate of the low-skilled relative to those of all other skill groups, which matches the empirical regularity presented in Figures 2-3.

Remembering that the equilibrium unemployment rate is found to be $\delta/(\delta + \lambda_0)$, let us have a look at how λ_0 and δ changed over the observation period. For the unskilled, λ_0, the arrival rate of job offers while unemployed fell over the observation period from 0.0373 to 0.0273, i.e., by roughly 27 %. This change was significant, as a Wald test for the constancy of this parameter in the first row of Table 3 demonstrates. So we may conclude that by significantly slowing down the search intensity of unemployed workers the entitlement extension reform has indeed contributed to the increase in the unemployment rate of unskilled workers.

Table 3: "Test Results for Search Intensities"

	H_0 :	$\chi^2_{(1)}$	p-Value
Skills	$\lambda_0^{(86)} = \lambda_0^{(95)}$	19.5996	0.0000
(Group 1)	$\lambda_1^{(86)} = \lambda_1^{(95)}$	16.4417	0.0000
	$\delta^{(86)} = \delta^{(95)}$	16.9996	0.0000
Age	$\lambda_0^{(86)} = \lambda_0^{(95)}$	0.9781	0.3227
(Group 4)	$\lambda_1^{(86)} = \lambda_1^{(95)}$	0.1451	0.7033
	$\delta^{(86)} = \delta^{(95)}$	15.1216	0.0001

The arrival rate of job offers for the unskilled also fell much more than those of the higher skill groups. For both skill group 2, middle vocational training, and group 3, vocational training and college entrance exam or higher vocational training, it decreased by only about 12 %. This finding goes in line with the argument that the benefit reform may potentially have an adverse negative effect on the unemployment rates of the unskilled.

A remarkable result displayed in Table C4 is that for workers with the highest skills (group 4) λ_0 increased from 0.0659 in 1986 to 0.0864 in 1995 which amounts to more than 30 %. Moreover, their reservation wage rose by about 1,200 D-Mark. In contrast, for skill groups 1 and 3 it nearly did not change and for group 2 it rose by only about 200 D-Mark. A potential explanation for these results can be a skill-biased technological change that raised the productivity and arrival rate of job offers of workers with the highest skills as well as their reservation wages. This hypothesis is supported by the fact that over the observation period, the wage offer density for the high-skilled became less skewed and has considerably shifted to the right (see Figure C3). Such change of curvature is determined by the relative increase in productivity and higher concentration of probability mass on the right tail of the productivity distribution, which is in line with the acceleration of technological progress.

Consider now the arrival rate of job loss, δ. For the least skilled workers of group 1, it rose from 0.0043 in 1986 to 0.0049 in 1995; i.e., by about 14 %. While this percentage rise is similar for group 2, for group 3 it is about 26 % and for group 4 even 33 %. This would be in line with the interpretation that the rise length of UI entitlement had only a small effect on the job loss probability of the least skilled workers. It rather affected the probability of job-loss of the two groups of high skilled-workers; for them extending the length of UI receipt was more important as they have much lower chances than the low-skilled to receive UA benefit once they run out of their UI benefit.

Still, if we take look at the third row of Table 3 we will see that the observed 14% increase of the exit rate to unemployment among the low-skilled workers is statistically significant. This implies that the increase of δ has also significantly contributed to the upward shift of the unemployment rate. Though, as long as δ is entirely exogenous to the model and theoretically absorbs all other possible reasons for match dissolution, it is an open question what share in the observed eventual 14% increase is due to the entitlement extension.

Summarizing all the findings above we conclude that the entitlement extension reforms of the late 1980s have led to a significant slowdown in the search intensity among the unemployed low-skilled workers. Moreover they have quite likely contributed to the significantly increased incentive of shirking among the employed low-skilled workers. Taken together these two effects have led to the leap of unemployment rates of unskilled workers observed in the data. Furthermore, as it could be generally expected (see for instance Nickell and Layard , 1999), this reform has brought about a significantly longer duration of unemployment for the unskilled[17].

We also observe the similar influence for the second and third qualification groups. However, the magnitude is much lower. For the highest skill group, skill biased technological change may have counteracted this effect over the

[17]This conclusion follows automatically, since the expected unemployment duration within the theoretical model is just a reciprocal search intensity parameter of unemployed workers.

period under review since as a result no changes in unemployment rates were predicted.

Now let us proceed with the parameter estimates for λ_1, the arrival rate of job offers while employed. They are also displayed in Table C3. From 1986 to 1995 for skill groups 1, 2 and 4, the estimates reveal a decrease of this arrival rate of about 20 %, 12 % and 34 %, respectively. For group 3 instead, it rose by roughly 20 %. These results are somewhat puzzling. With skill-biased technological change, one would have expected, that the higher is the skill level, the bigger should be the percentage change in the arrival rate of job offers. Instead, however, we observe that after having increased for the third group the arrival rate of job offers has immensely decreased for group 4, workers with the highest skills. As a possible explanation to this phenomenon one may suggest that firms post too high wages for workers of the skill group 4 because the highly skilled personnel becomes increasingly important. However, we regard this interpretation as rather speculative.

Finally, considering the changes in λ_1 for the first two qualification groups we may think that the skill-biased technological progress obscures the promotion prospects of the low-skilled. We observe that the less qualified the worker is, the fewer chances of finding a better paid job he/she has.

Concluding the discussion of this subsection it would be natural to go over the policy measures that our results imply. As we have discovered, the extension of entitlement to UI has significantly affected the search intensity of unemployed low-skilled workers and through this contributed to the increase of equilibrium unemployment rate in this group. Moreover it has also raised the expected duration of being unemployed. Therefore if one pursues the goal of reducing the unemployment rate and tries to enhance incentives to return to work faster, entitlement length could be a valuable instrument.

6.3 Estimation Results by Age Group

Now consider the results for different age groups as displayed in Table C5 of the Appendix. Remember from Section 2 that in 1987 the maximum duration of UI receipt rose from 24 to 32 months for workers older than 53 years. So we should expect an increase of their group-specific equilibrium unemployment rate. Our results indeed reflect such an increase. Table C5 shows that from 1986 to 1995 the predicted equilibrium unemployment rate of workers aged 54 to 64 years went up by 28 % (from 11.4 % to 14.6 %). However here we significantly underpredict the magnitude of the change, since in fact their sample unemployment rate increased from 9.3 % to 16.3 %, i.e. by more than 70 % (see Table 2).

The percentage changes of the predicted unemployment rates of the other age groups 1 (16-27 years), 2 (28-40 years), and 3 (41-53 years) are about 15 %, 44 % and 22 % respectively, i.e., for the 28 to 40 year old the percentage rise of the predicted equilibrium unemployment rate is even higher than for 54 to 64 year old. These results are also somewhat surprising because the maximum duration of UI receipt remained constant at 12 months for workers younger than

42 years. Only for the 42 to 53 year old it rose to some extent.

Let us turn to the estimation results for λ_0 and δ. The results in Table C5 imply that for the oldest group the arrival rate of job offers while unemployed, λ_0, decreased by roughly 10 % over the period under review. *Ceteris paribus*, this change would have raised the equilibrium unemployment rate of the elder workers to 12.6 %. Yet according to the Wald test in Table 3 for the elder workers we cannot reject the hypothesis that the parameter λ_0 is the same in 1986 and 1995. This means that our model does not support the argument that the entitlement extension affects the unemployment rates of aged workers through influencing their search intensity.

As to the other age groups, we could have expected that the arrival rate of job offers would have fallen much more for aged workers than for younger ones. Compared with the 16 to 27 year old the results are in line with our expectation: the search intensity of the youngest fell by only about 2 % (see Table C5). Though for the other two age groups its percentage change is quite similar to that of the eldest workers.

Consider now the arrival rate of employer-employee match dissolution δ. From 1986 to 1995 the incidence of job loss rose by about 18 % for the 54 to 64 year old, while for the three younger groups, it rose by about 14 %, 30 % and 13 % respectively. So, the percentage rise in the probability of job loss of the elder workers at least somewhat exceeds that of 16 to 27 year old and the 41 to 53 year old. Still we cannot see that the exit rate positively depends on age.

To see whether the observed 18% increase in δ has significantly contributed to the increase of unemployment rates among the oldest workers we again test the hypothesis of the constancy of δ over time. The results in Table 3 indicate clear rejection. This proves that the rise in unemployment of aged workers is mainly explained by the by the in employer-worker separation rate.

This is partly in line with our expectations. Over the second half of the 1980s the generosity of the benefit system increased more for older workers than for the younger ones. Additionally due to the Employment Promotion Act of 1986 (see Steffen, 2003) unemployed people of at least 58 years old were granted a possibility to be no longer available for mediation into jobs, provided that they would retire early at the age of 60. This could have made a job loss for elder workers more acceptable and therefore could have increased their incentives to work less intensively. This implies that the likelihood that firms terminate the employment of elder workers should have increased, since such a termination could be done more amicably given the generous (and essentially indefinite) benefit entitlements the elder workers have become able to get. As a result incidence of match dissolution has gone up.

Table C5 also displays the predicted reservation wages of the four age groups for the years 1986 and 1995. The reservation wages of the 54 to 64 year old hardly changed over this period. So we cannot find that the benefit reform had a major impact on their reservation wage. The reservation wages of both the 16 to 28 year old and the 41 to 53 year old rose by somewhat more than 450 D-Mark, while those of the 28 to 40 year old fell by roughly 200 D-Mark.

To summarize, we find that the chosen theoretical model is not rich enough to shed light on the precise mechanisms that shifted up the unemployment rate of the eldest workers. For this group the predicted change of the equilibrium unemployment rate is considerably lower than the actual change in the sample. Hence we would also expect that the changes in the parameter estimates are biased. This may be the reason why our results for elder workers do not generally reflect our expectations about changes of their search intensities, job loss rate and reservation wages. We discover that the dynamics of unemployment rates of the aged workers is mostly determined by the changes in their rate of job loss. But still we see that the institutional influence here is more complex, because now it consists of not only prolonging unemployment insurance benefit entitlement but also the possibility of earlier retirement. As long as under the assumptions of the model the rate of job loss is exogenous we are not able to say definitely which of the two stands behind it. Though in view of statistical insignificance of the changes in search behavior of the eldest group, we would tend to think that the suggested early retirement argument may be an explanation.

7 Summary and Conclusions

In this paper we ask a question whether reforms that extended the entitlement length to UI benefit payments in West Germany had a significant contribution to the increase in unemployment rates among unskilled and aged workers. We try to answer this question by estimating parameters of the theoretical search equilibrium model of Burdett and Mortensen (1998) with heterogenous employers. Our choice of the theoretical framework is determined by the fact that through the individual search behavior the model makes it possible to link the increased UI entitlement length with the subsequent dynamics of unemployment rates. Furthermore we turn to the search equilibrium approach because only this framework allows to give the quantitative description of individual search behavior fully consistent with the solution of the economic-theoretical model.

To estimate the model we firstly use the structural nonparametric approach suggested by Bontemps et al. (2000). However, we discover that this procedure cannot be always applicable and find a data-driven condition, which demonstrates the limitations of this estimation techniques. As long as the applicability condition which we refer to as "constraint inconsistency" is not satisfied in our case, we proceed with the structural estimation method suggested by Bowlus et al. (1995), (2001).

In our study we find that for unskilled workers the extension of the entitlement period has significantly contributed to the changes in their search behavior. Both arrival rates of job offers to unemployed and employed workers went down. Moreover the arrival rate of employee-employer match went up considerably. A slowdown in unemployment search intensity along with increased incentives to shirk induced by the UI system after the reforms has led to the increase of predicted unemployment rates in this group. Unemployment rate for the unskilled predicted by the model shifts from 10.3% to 15.1% which almost completely matches the 9.2 % to 16.4 % increase of the same rate observed in the data.

As to the elder workers, a pure search intensity adjustment argument is unfortunately insufficient to present a satisfactory explanation of unemployment rate dynamics. However, the model mirrors the phenomenon of increased unemployment rates between 1986 and 1995 predicting a higher exit from jobs into unemployment. We know that for this group of workers the entitlement to unemployment benefit payments became particularly long in the second half of the 1980s. Additionally, whenever out of job, under certain conditions elder workers were granted a possibility to retire earlier. Taken together this may have made a job loss more acceptable and give employers an incentive to dismiss aged workers rather than the others. So the benefit and retirement reforms have rather affected the exit rates into unemployment than search behavior and reservation wage of the elder. Still our model is not rich enough to separate these two institutional effects.

In this context, it is interesting to note that recent labour market reforms instituted in Germany are likely to reverse some of the phenomena observed in this paper. In particular the entitlement period to unemployment insurance has been shortened twice recently (in 1997 and in 2003), the levels of UA benefits are being adjusted downwards to the level of social assistance (starting in 2005), a variety of incentive mechanisms (such as increased sanctions) to increase job search have been instituted in 2003, and reforms in the job referral system of employment agencies aimed at lowering the costs of job search were undertaken (in 2003). Given the logic of this paper, one would expect these reforms to result in lower unemployment rates, although it is difficult to precisely predict their quantitative effects.

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8 Appendix

8.1 Appendix A: Welfare Benefit of the Year 1986

As mentioned, we computed the welfare benefits at the month of the stock sample of the year 1986 using retrospective information of the year 1987 on the different types of welfare benefits. For each type of welfare benefit we proceeded as follows. We first of all assumed an equal distribution of the benefits over the household members, so the monthly welfare benefit level was divided by the household size. For individuals living in households which receive the benefit all over the year the monthly per capita welfare benefit was assigned. For individuals who live in households without such a benefit all over the year, we assumed a potential benefit level of zero.

For unemployed individuals who got no unemployment benefit but some welfare benefit receipt was reported, we set their welfare benefit level to the monthly per capita welfare benefit. For unemployed people who at the date of the stock sample received unemployment benefit, we proceeded differently. If their household received welfare benefit for some months, we assumed that the welfare benefit was certainly during the months in which unemployment was not covered by unemployment benefit. But the number of months of welfare benefit can exceed the number of months of unemployed without unemployment benefit. If this difference is positive, we assumed that there was also some welfare benefit receipt while these individuals were entitled to unemployment benefit. Hence we assigned as their welfare benefit the monthly per capita welfare benefit multiplied by the share of months with unemployment benefit receipt in which welfare benefit was available.

8.2 Appendix B: Reservation Wage and Search Intensities

Homogeneous employers

Consider the reservation wage equation in (2). Firstly, for $F(w)$ substitute the wage offer distribution with equally productive employers

$$F(w) = \frac{\delta + \lambda_1}{\lambda_1}\left[1 - \sqrt{\frac{p-w}{p-R}}\right]$$

After some algebra we get

$$G(R, \lambda_0, \lambda_1, \delta, b) = R - b - \frac{\lambda_0 - \lambda_1}{\lambda_1}(\overline{w} - R) - \frac{2\delta(\lambda_0 - \lambda_1)}{\lambda_1(\delta + \lambda_1)}\left[\sqrt{p - \overline{w}}\sqrt{p - R} - (p - R)\right] = 0$$

Generally we have:

$$dG = \frac{\partial G}{\partial R}dR + \frac{\partial G}{\partial \lambda_0}d\lambda_0 + \frac{\partial G}{\partial \lambda_1}d\lambda_1 + \frac{\partial G}{\partial \delta}d\delta + \frac{\partial G}{\partial b}db$$

To apply the Implicit Function Theorem we find the partial derivatives of G with respect to $b, R, \lambda_0, \lambda_1$ and δ (for simplicity we treat p as constant)

1. $$\frac{\partial G}{\partial b} = -1 < 0$$

2.
$$\frac{\partial G}{\partial R} = 1 + \frac{\lambda_0 - \lambda_1}{\lambda_1}\left[1 - \frac{2\delta}{\delta + \lambda_1}\left(1 - \frac{1}{2}\sqrt{\frac{p - \overline{w}}{p - R}}\right)\right]$$
It could be seen that provided $\lambda_0 \geq \lambda_1$ the above expression is positive at least for $\lambda_1 \geq \delta$. Moreover, when p is large enough the above expression is positive for all $\lambda_1 \in (0, \lambda_0]$.

3.
$$\frac{\partial G}{\partial \lambda_0} = -\frac{(\overline{w} - R)}{\lambda_1} - \frac{2\delta}{\lambda_1(\delta + \lambda_1)}\left[\sqrt{p - \overline{w}}\sqrt{p - R} - (p - R)\right]$$
It is easy to show that the inequality $(\overline{w} - R) + \left[\sqrt{p - \overline{w}}\sqrt{p - R} - (p - R)\right] > 0$ is always true. So whenever $\lambda_1 \geq \delta$ the above derivative is necessarily negative.

4.
$$\frac{\partial G}{\partial \lambda_1} = \frac{\lambda_0}{\lambda_1^2}(\overline{w} - R) + \frac{2\delta\left(\lambda_0\delta + 2\lambda_0\lambda_1 - \lambda_1^2\right)}{\lambda_1^2(\delta + \lambda_1)^2}\left[\sqrt{p - \overline{w}}\sqrt{p - R} - (p - R)\right]$$
Using the fact that $(\overline{w} - R) + \left[\sqrt{p - \overline{w}}\sqrt{p - R} - (p - R)\right] > 0$ to say that $\partial G \,/\, \partial \lambda_1$ is positive it is sufficient to show that λ_0 exceeds $\frac{2\delta\left(\lambda_0\delta + 2\lambda_0\lambda_1 - \lambda_1^2\right)}{(\delta + \lambda_1)^2}$. This is indeed true at least for those λ_1 that are close to λ_0. However, if $\lambda_1 \to \delta$ the derivative reverses its' sign and becomes negative. So eventually on the interval $[\delta, \lambda_0]$ we observe that $\partial G \,/\, \partial \lambda_1 \lesseqgtr 0$.

5.
$$\frac{\partial G}{\partial \delta} = -\frac{2\left(\lambda_0 - \lambda_1\right)}{(\delta + \lambda_1)^2}\left[\sqrt{p - \overline{w}}\sqrt{p - R} - (p - R)\right]$$
As long as $\lambda_0 > \lambda_1$ the above derivative is clearly positive.

Therefore, for $\lambda_1 \in [\delta, \lambda_0)$, we conclude the following. When b increases and affects only one of the parameters, then the sign of the impact on the different relevant parameters is represented by the following derivatives:

$$\frac{\partial R}{\partial b} = -\frac{\partial G/\partial b}{\partial G/\partial R} > 0, \qquad \frac{\partial \lambda_0}{\partial b\delta} = -\frac{\partial G/\partial b}{\partial G/\partial \lambda_0} < 0, \qquad \frac{\partial \delta}{\partial b} = -\frac{\partial G/\partial b}{\partial G/\partial \delta} > 0$$
$$\lambda_1 \to \lambda_0 : \quad \frac{\partial \lambda_1}{\partial b} = -\frac{\partial G \,/\, \partial \lambda_1}{\partial G \,/\, \partial R} < 0, \quad \lambda_1 \to \delta : \quad \frac{\partial \lambda_1}{\partial b} = -\frac{\partial G \,/\, \partial \lambda_1}{\partial G \,/\, \partial R} > 0,$$

We emphasize that these derivative do not formally show the effect of an increased length of entitlement to unemployment benefit, but we expect its effect to go in the same direction.

Naturally, we could have an increase in b that affects, R, λ_0, and δ, while we leave all other parameters constant. We argue as if this does not change the sign of the effect on R, but reduces its size! So we would argue that a rise in b raises R. But the total increase in R will be lower, if due to the rise in b there is also a downward adjustment of λ_0 due to a reduction of search intensity while unemployed and/or if there is an increase in δ. The following equations show in which directions changes of the other parameters would affect R:

$$\frac{\partial R(\lambda_0, \lambda_1, \delta)}{\partial \lambda_0} = -\frac{\partial G \,/\, \partial \lambda_0}{\partial G \,/\, \partial R} > 0 \qquad \frac{\partial R(\lambda_0, \lambda_1, \delta)}{\partial \delta} = -\frac{\partial G \,/\, \partial \lambda_0}{\partial G \,/\, \partial R} < 0$$
$$\lambda_1 \to \lambda_0 : \quad \frac{\partial R(\lambda_0, \lambda_1, \delta)}{\partial \lambda_1} = -\frac{\partial G \,/\, \partial \lambda_1}{\partial G \,/\, \partial R} < 0, \quad \lambda_1 \to \delta : \quad \frac{\partial R(\lambda_0, \lambda_1, \delta)}{\partial \lambda_1} = -\frac{\partial G \,/\, \partial \lambda_1}{\partial G \,/\, \partial R} > 0,$$

Heterogeneous employers

Now consider the more general case of different productivity levels. The wage offer distribution in this case is given by

$$F(w)=\frac{\delta+\lambda_1}{\lambda_1}\left[1-\frac{\delta+\lambda_1\left(1-\gamma_{j-1}\right)}{\delta+\lambda_1}\sqrt{\frac{p_j-w}{p_j-w_{H_{j-1}}}}\right],\quad w\in\left(w_{L_j},w_{H_j}\right]$$

In this case equation (2) becomes

$$G=w_{L_1}-b-\frac{\lambda_0-\lambda_1}{\lambda_1}\sum_{i=1}^{Q}\left(w_{H_i}-w_{L_i}\right)-$$
$$\sum_{i=1}^{Q}\frac{2\delta(\lambda_0-\lambda_1)}{\lambda_1(\delta+\lambda_1(1-\gamma_i))}\left[\sqrt{p_{i+1}-w_{H_i}}\sqrt{p_{i+1}-w_{L_i}}-\left(p_{i+1}-w_{L_i}\right)\right]$$

where Q is the number of support points in the productivity distribution, $w_{L_1}=R$ and $w_{HQ}=\overline{w}$.

Partial derivatives of G with respect to the parameters of interest are:

$$\begin{aligned}
\frac{\partial G}{\partial R} &= 1+\frac{\lambda_0-\lambda_1}{\lambda_1}\left[1-\frac{2\delta}{\delta+\lambda_1}\left(1-\frac{1}{2}\sqrt{\frac{p-w_{H_1}}{p-R}}\right)\right]\\
\frac{\partial G}{\partial\lambda_0} &= -\lambda_1^{-1}\sum_{i=1}^{Q}(w_{H_i}-w_{L_i})-\sum_{i=1}^{Q}\frac{2\delta\left[\sqrt{p_{i+1}-w_{H_i}}\sqrt{p_{i+1}-w_{L_i}}-\left(p_{i+1}-w_{L_i}\right)\right]}{\lambda_1(\delta+\lambda_1(1-\gamma_i))}\\
\frac{\partial G}{\partial\lambda_1} &= \frac{\lambda_0}{\lambda_1^2}\sum_{i=1}^{Q}(w_{H_i}-w_{L_i})+\sum_{i=1}^{Q}\left\{\frac{2\delta\left[\lambda_0\delta+\left(2\lambda_0\lambda_1-\lambda_1^2\right)(1-\gamma_i)\right]}{\lambda_1^2(\delta+\lambda_1(1-\gamma_i))^2}\right.\\
&\quad\left.\left[\sqrt{p_{i+1}-w_{H_i}}\sqrt{p_{i+1}-w_{L_i}}-\left(p_{i+1}-w_{L_i}\right)\right]\right\}\\
\frac{\partial G}{\partial\delta} &= -\sum_{i=1}^{Q}\frac{2\left(\lambda_0-\lambda_1\right)(1-\gamma_i)}{(\delta+\lambda_1(1-\gamma_i))^2}\left[\sqrt{p_{i+1}-w_{H_i}}\sqrt{p_{i+1}-w_{L_i}}-\left(p_{i+1}-w_{L_i}\right)\right]
\end{aligned}$$

We observe that the sign pattern of $\partial G/\partial R$ is the same as it was before. The same applies to $\frac{\partial G}{\partial b}$ since it does not depend on p. As to the rest of the derivatives the additional ambiguity is introduced by the term $(1-\gamma_i)$. This term implies that in order to preserve the same sign pattern as we have discovered before the conditions for λ_1 should be stronger with $\gamma_i\to 1$. In other words λ_1 should be closer and closer to λ_0. At the same time importance of each additional i-th term declines since for increasing i the productivity level increases by incomparably much more then the cutoff wage. This implies that for increasing i $\left[\sqrt{p_{i+1}-w_{H_i}}\sqrt{p_{i+1}-w_{L_i}}-\left(p_{i+1}-w_{L_i}\right)\right]\to 0$ and this convergence is very rapid. This makes the value of the complete additional i-th term more and more negligible and so relaxes the assumption on the magnitude of λ_1. Therefore we believe that in the general case of $F(w)$ partial derivatives should behave in the way they do with the homogeneous function.

So we suggest that without loss of generality the derivatives of b and R with respect to λ_0,λ_1 and δ have the signs as in the subsection above.

8.3 Appendix C: Estimation Results

Table C1: "Choice of Specification for $\gamma(p)$ in 1986"

Q		Coefficients	Log(L)	CAIC / SBC	
1 homogen. model	κ_0	9.7780 (0.2549)	-80715.714	161469.4	161465.4
	κ_1	0.1696 (0.0079)			
	δ	0.0058 (0.0001)			
2-6		...	...	...	...
7	κ_0	12.3247 (0.4559)	-74416.835	148928.6	148918.6
	κ_1	3.3833 (0.0847)			
	δ	0.0039 ($6.7 \cdot 10^{-5}$)			
8	κ_0	12.7914 (0.4767)	-74245.072	*148594.6*	*148583.6*
	κ_1	4.8014 (0.1157)			
	δ	0.0036 ($6.3 \cdot 10^{-5}$)			
9	κ_0	12.8191 (0.4838)	-74242.596	148599.1	148587.1
	κ_1	4.8895 (0.1181)			
	δ	0.0036 ($6.3 \cdot 10^{-5}$)			

Table C2: "Choice of Specification for $\gamma(p)$ in 1995"

Q		Coefficients	Log(L)	CAIC / SBC	
1 homogen. model	κ_0	7.1764 (0.2338)	-65566.955	131171.1	131167.1
	κ_1	0.1324 (0.0064)			
	δ	0.0066 (0.0001)			
2 - 8		...	...	...	...
9	κ_0	9.3407 (0.3418)	-61099.327	122310.3	122298.3
	κ_1	3.9950 (0.1120)			
	δ	0.0042 ($7.9 \cdot 10^{-5}$)			
10	κ_0	9.3459 (0.3428)	-61075.378	*122271.7*	*122258.7*
	κ_1	4.0133 (0.1126)			
	δ	0.0041 ($8.0 \cdot 10^{-5}$)			
11	κ_0	9.3544 (0.3416)	-61073.631	122277.5	122263.5
	κ_1	4.0443 (0.1136)			
	δ	0.0041 ($8.0 \cdot 10^{-5}$)			

Table C3: "Estimated Model for the Whole Economy"

1986				1995		
Coefficients (Std.Errors)				Coefficients (Std.Errors)		
κ_0	12.7914	(0.4767)		κ_0	9.3459	(0.3428)
κ_1	4.8014	(0.1157)		κ_1	4.0133	(0.1126)
δ	0.0036	($6.3 \cdot 10^{-5}$)		δ	0.0041	($8.0 \cdot 10^{-5}$)
Estimated Productivity Distribution:				Estimated Productivity Distribution:		
i :	P_i	γ_i		i :	P_i	γ_i
1	2304.6	0.65561		1	2758.0	0.62421
2	2726.6	0.81784		2	3120.4	0.79455
3	3289.8	0.90804		3	3845.8	0.88384
4	4601.5	0.95306		4	4738.5	0.92208
5	7997.2	0.98269		5	6147.2	0.94792
6	18630.5	0.99529		6	8673.3	0.97320
7	62728.1	0.99897		7	13906.5	0.98731
8	437143.1	1		8	24442.1	0.99331
				9	53593.8	0.99769
				10	232585.7	1
Log(Likelihood):		-74245.072		Log(Likelihood):		-61075.378

Table C4: "Estimated Model for Different Qualification Groups"

Qualification Groups	1986 Coefficients (Std.Errors)		1986 u	1986 R	1995 Coefficients (Std.Errors)		1995 u	1995 R
Group I: (inadequately or general elementary)	λ_0	0.0373 (0.0023)	0.103	1768.9	λ_0	0.0273 (0.0017)	0.151	1800.4
	λ_1	0.0110 (0.0006)			λ_1	0.0088 (0.0006)		
	δ	0.0043 $(1.4 \cdot 10^{-4})$			δ	0.0049 $(1.9 \cdot 10^{-4})$		
	Number of mass points in $\gamma(p)$:		7		Number of mass points in $\gamma(p)$:		7	
	Number of observations:		1401		Number of observations:		933	
	Log(Likelihood):		-20331.83		Log(Likelihood):		-13319.29	
Group II: (middle vocational)	λ_0	0.0513 (0.0028)	0.067	2036.0	λ_0	0.0449 (0.0024)	0.087	2238.6
	λ_1	0.0179 (0.0006)			λ_1	0.0157 (0.0006)		
	δ	0.0037 $(9.1 \cdot 10^{-5})$			δ	0.0043 $(1.2 \cdot 10^{-4})$		
	Number of mass points in $\gamma(p)$:		8		Number of mass points in $\gamma(p)$:		7	
	Number of observations:		2381		Number of observations:		1973	
	Log(Likelihood):		-35969.58		Log(Likelihood):		-29618.12	
Group III: (vocational + 'Abitur' or higher vocation)	λ_0	0.0746 (0.0104)	0.040	2852.1	λ_0	0.0659 (0.0080)	0.055	2858.3
	λ_1	0.0184 (0.0012)			λ_1	0.0221 (0.0015)		
	δ	0.0031 $(1.5 \cdot 10^{-4})$			δ	0.0039 $(1.9 \cdot 10^{-4})$		
	Number of mass points in $\gamma(p)$:		8		Number of mass points in $\gamma(p)$:		7	
	Number of observations:		616		Number of observations:		587	
	Log(Likelihood):		-9643.23		Log(Likelihood):		-9078.60	
Group IV: (higher education)	λ_0	0.0659 (0.0103)	0.043	2488.2	λ_0	0.0864 (0.0131)	0.044	3687.5
	λ_1	0.0434 (0.0031)			λ_1	0.0285 (0.0021)		
	δ	0.0030 $(1.8 \cdot 10^{-4})$			δ	0.0040 $(2.2 \cdot 10^{-4})$		
	Number of mass points in $\gamma(p)$:		9		Number of mass points in $\gamma(p)$:		7	
	Number of observations:		450		Number of observations:		488	
	Log(Likelihood):		-7146.4097		Log(Likelihood):		-7828.00	

Table C5: "Estimated Model for Different Age Groups"

Age Groups	1986 Coefficients (Std.Errors)		u	R	1995 Coefficients (Std.Errors)		u	R
Group I: 16-27 years old	λ_0	0.0781 (0.0053)	0.096	1016.1	λ_0	0.0766 (0.0060)	0.111	1479.2
	λ_1	0.0533 (0.0026)			λ_1	0.0357 (0.0022)		
	δ	0.0083 ($3.1 \cdot 10^{-4}$)			δ	0.0095 ($4.2 \cdot 10^{-4}$)		
	Number of mass points in $\gamma(p)$:		7		Number of mass points in $\gamma(p)$:		6	
	Number of observations:		1110		Number of observations:		745	
	Log(Likelihood):		-15799.39		Log(Likelihood):		-10334.69	
Group II: 28-40 years old	λ_0	0.0664 (0.0047)	0.057	2437.3	λ_0	0.0584 (0.0037)	0.081	2244.5
	λ_1	0.0164 (0.0007)			λ_1	0.0238 (0.0010)		
	δ	0.0040 ($1.2 \cdot 10^{-4}$)			δ	0.0052 ($1.6 \cdot 10^{-4}$)		
	Number of mass points in $\gamma(p)$:		7		Number of mass points in $\gamma(p)$:		8	
	Number of observations:		1609		Number of observations:		1527	
	Log(Likelihood):		-25096.58		Log(Likelihood):		-23409.19	
Group III: 41-53 years old	λ_0	0.0408 (0.0028)	0.070	2983.0	λ_0	0.0372 (0.0027)	0.087	3328.2
	λ_1	0.0068 (0.0003)			λ_1	0.0056 (0.0004)		
	δ	0.0031 ($9.1 \cdot 10^{-5}$)			δ	0.0035 ($1.3 \cdot 10^{-4}$)		
	Number of mass points in $\gamma(p)$:		7		Number of mass points in $\gamma(p)$:		4	
	Number of observations:		1583		Number of observations:		1121	
	Log(Likelihood):		-24712.09		Log(Likelihood):		-17411.64	
Group IV: 54-64 years old	λ_0	0.0215 (0.0021)	0.114	3165.2	λ_0	0.0194 (0.0015)	0.146	3114.4
	λ_1	0.0032 (0.0003)			λ_1	0.0030 (0.0003)		
	δ	0.0028 ($1.5 \cdot 10^{-4}$)			δ	0.0033 ($1.5 \cdot 10^{-4}$)		
	Number of mass points in $\gamma(p)$:		5		Number of mass points in $\gamma(p)$:		7	
	Number of observations:		575		Number of observations:		637	
	Log(Likelihood):		-8021.41		Log(Likelihood):		-9344.46	

Figure C1: „Estimated Theoretical Offer and Earnings Distributions for the whole Economy: Sample 1986“

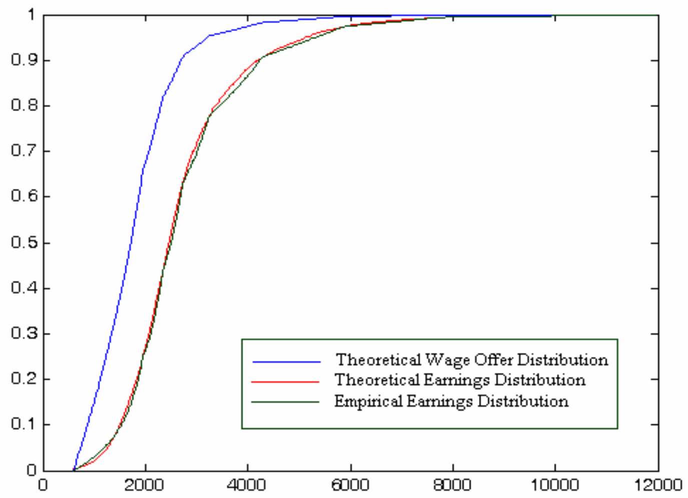

Figure C2: „Estimated Theoretical Offer and Earnings Distributions for the whole Economy: Sample 1995“

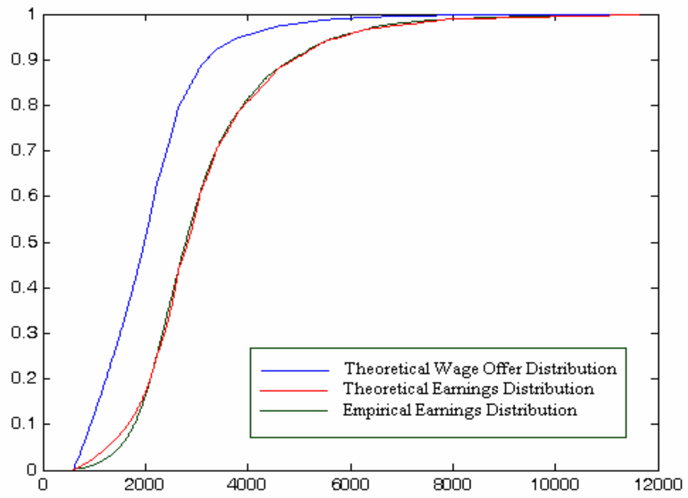

Figure C3: „Estimated Theoretical Wage Offer Densities for the High Skilled Group“