# The Transition from Welfare to Work and the Role of Potential Labour Income

Hilmar Schneider (IZA, DIW Berlin)
Arne Uhlendorff (DIW Berlin)*

*Very preliminary version*

**Abstract**

It is often argued that the high level of welfare claims in Germany causes little incentive for workers with low productivity to seek for a job. We examine the influence of the ratio between estimated potential labour income and the welfare payment level on the probability of leaving social welfare. Using the GSOEP, we estimate different discrete time hazard rate models with and without unobserved heterogeneity. Controlling for several typical covariates the ratio between potential labour income and the welfare level shows a strong positive effect on the probability of leaving social welfare, independent of the model specification we use.

JEL-Classification: I38, J64, C41

Keywords: Social Welfare; Labour Supply; Duration Analysis

---

* We would like to thank Hans Baumgartner, Peter Haan, Hans-Joachim Rudolph and Katharina Wrohlich, for helpful discussions and comments.

## 1 Introduction

In the year 2002, about 2.8 million people in Germany received social assistance.[1] The number of recipients as well as the amount of income support expenditures have been rising almost continuously in the past. The German social assistance is a means-tested transfer program financed by the municipalities. The receipt of transfer payments requires that the household income including other transfer payments like unemployment benefits does not exceed a certain minimum level. In contrast to the unemployment benefits, everybody is principally eligible, irrespective his or her individual employment history. Although the receipt is in principle unlimited, only a minority of households stays on welfare over a longer period of time. Assistance claims expire as soon as alternative income exceeds a certain threshold. This may be due to labour income but could also be due to changes in household formation or the receipt of alternative transfer payments like educational assistance etc.

In the economic literature as well as in public debate on the German welfare system an incentive argument plays an important role: If the difference between the level of transfers and potential income from a regular job is too small then picking up a job is not attractive for the individual (see for example Ochel 2003). In descriptive analyses, the social assistance levels are generally compared with the average wage of a special group of employees, for example unskilled workers in manufacturing (e.g. Engels 2001 or Boss 2002). Even in more elaborated studies the difference between the potential market wage of an individual and the amount of social assistance has not been considered so far (see for example Voges and Rohwer 1992 or Gangl 1998).

Hazard rate models are an appropriate tool for the analysis of the duration of welfare receipt. For this context, Gangl (1998) has shown that it is important to distinguish transitions to employment from alternative transitions like transitions out of the labour market.

Numerous studies exist on the duration of income support spells. Most of them are referring to North America and dealing with women receiving welfare. A typical result says that the probability of leaving welfare is higher for better educated and white persons and declines with the number of (young) children, disabilities, the amount of benefits and the level of regional unemployment (see e.g. Blank 1989, Stewart and Dooley 1999 or Gittleman 2001, a summary is given by Moffitt 1992).

---

[1] This number of welfare recipients refers to permanent transfers, the so-called *Hilfe zum Lebensunterhalt,* described in detail in section 2.

For Germany, Wilde (2003) examines the difference between social benefits and the average income for unskilled employees on the probability of leaving social welfare using the Low Income Panel. He finds no significant effects. The study of Riphahn (1999) with the German Socio-Economic Panel study (GSOEP) shows no significant influence of a predicted real net income variable for the head of household on the probability of an exit from income support. However, she does not take the amount of social transfers into account. Both, Wilde and Riphahn do not distinguish between different transitions in their regression analysis.

The core questions of this paper are: What determines the transition from social welfare to work? And: How important is the ratio of potential labour income to the amount of transfer payment for this issue?

We use data from the GSOEP. Between 1992 and 2000 retrospective monthly information about social welfare receipt for each month of the previous calendar year is part of the household questionnaire. Spell duration is observed in months. In the literature, the length of welfare spells is often analysed with continuous time hazard rate models (see e.g. Blank 1989, Voges and Rohwer 1992 or Stewart and Dooley 1999) Taking into account the discrete time measurement of the underlying data, we estimate two discrete time models instead: A random effects extreme value model assuming an underlying continuous time proportional hazard rate and a non-proportional multinomial model, which gives us insight to what extent the results vary with the econometric model. Moreover, we analyse the influence of the ratio as well as of the absolute difference between potential labour income and the amount of transfer payment.

Section 2 of this paper gives a short description of the system of social welfare in Germany and its theoretical implications on labour supply. Section 3 provides information on the data and the estimated models. Section 4 presents empirical results and section 5 concludes.

## 2 Incentive Effects of Social Assistance in Germany

The German social assistance (*Sozialhilfe*) is a means-tested transfer program and consists of two main parts: Permanent transfers to households with low income (*Hilfe zum Lebensunterhalt, HLU*) and transfers to persons in special circumstances who need temporary financial support[2] (*Hilfe in besonderen Lebenslagen*). In this study we concentrate on the HLU because these payments are principally unlimited and may act

[2] For example, pregnant women or homeless persons searching for a new apartment.

as a permanent alternative to a labour income. Therefore, HLU could reduce the incentive to search for a market job. In the following, the terms welfare and social assistance are used as synonyms and refer to HLU. The receipt of social assistance requires that the household income including other transfer payments like unemployment benefits (*Arbeitslosengeld* and *Arbeitslosenhilfe,* the latter is also means-tested and principally unlimited) does not exceed a certain minimum level.

In principle, everybody in need may claim for social assistance, while unemployment benefits are only accessible to those who have previously contributed to unemployment insurance for a minimum period within a given time frame. Moreover, the amount of unemployment benefits depends on the income in the previous job, while the amount of social assistance is related to a basic minimum income concept depending on household size and household composition. In addition, the eligibility criteria in case of own income differ between the means-tested unemployment benefits and the social assistance. Therefore, the analysis is restricted to social assistance spells without taking into account spells of the also means-tested *Arbeitslosenhilfe*.[3]

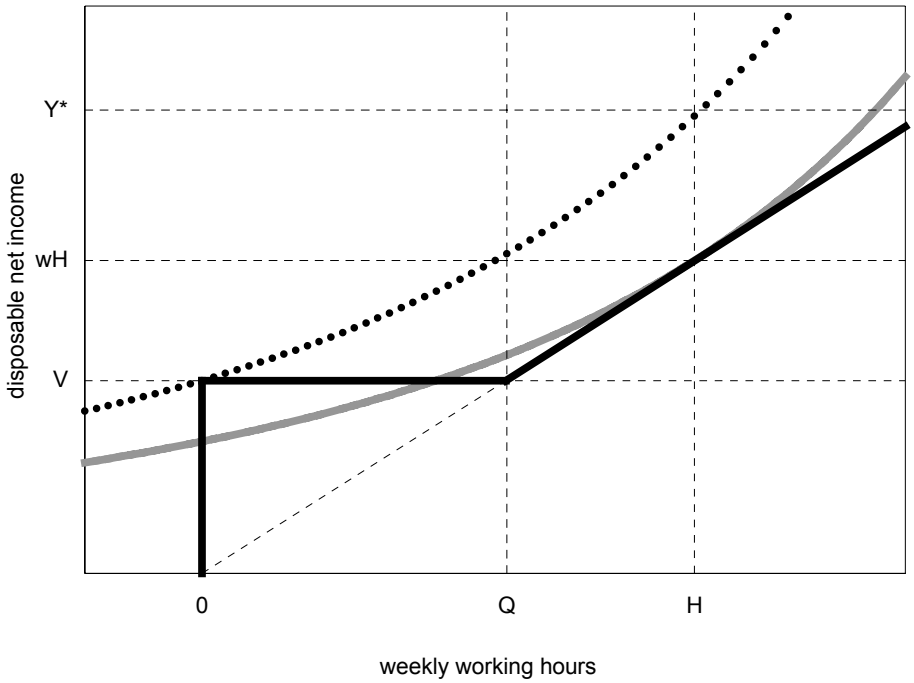
Welfare benefits consist of basic allowances for every adult household member, housing allowances and one-time payments. The amount for basic allowances differs between the federal states depending on the regional minimal costs of living. In 2003, it ranged between 282 and 297 € per month. Children get 50-90 percent depending on age. Expectant mothers, older and disabled persons receive higher basic allowances than "normal" adults. In principle, the amount of social assistance fills the gap between own income and the maximum benefit for the household. Labour income up to 25% of the basic allowance is not taken into account. Additional income is deducted at an implicit marginal tax rate of 85% until the deduction exceeds 50% of the basic allowance. Above this threshold the implicit marginal tax rate is 100%.

The impact of social assistance on work incentives can be described in terms of a very basic utility model for the choice between consumption and leisure (see for example Blundell/MaCurdy 1999). Assume a utility maximizing individual subject to a non convex budget set. A stylised depiction is given in figure 1.

---

[3] Recipients of unemployment benefits are included in our analysis if they are members of households receiving social assistance.

**Fig. 1: The impact of social assistance on work incentives for a stylised budget set**




If a person is not working at all, he or she will receive social assistance at a level of *V*. If this person works at his or her net market wage rate of *w* up to *Q* hours per week, disposable net income will not increase, since earned income is totally deducted from social assistance. Only when the number of hours worked is exceeding *Q*, disposable income will increase with slope *w*. The resulting non convex budget set is expressed in the graph by a thick black line.

If no social assistance existed, it would be optimal to work *H* hours per week with a disposable income of *wH*. Beside the budget set, optimal labour supply is a function of individual preferences that are responsible for the shape of the grey indifference curve tangent to the budget line. The utility level associated with *H* hours of work per week has to be compared to the utility level resulting from not working at all, which is expressed by the dotted indifference curve. In the depicted case, not working generates a higher utility level than working. A utility maximizing person would only work *H* hours per week, if he or she would be able to achieve a disposable income of at least *Y**, which can also be expressed in terms of an implicit minimum wage rate. Note that *Y** in the example given is more than twice as high as *V*. In general, it follows that wage offers are only accepted, if they exceed the implicit minimum wage is. Given that wage offers arrive at a given frequency and given a level of welfare payment, it is more likely to observe exits from social assistance, the higher an individuals market wage. This is what we are going to test in the empirical section of the paper.

## 3 Data, Variables and Methods

This study uses data from the German Socio-Economic Panel study (GSOEP). The yearly repeated GSOEP started 1984 in West Germany and was extended to include East Germany in 1990. In all panel waves, the head of the household provides information about the household and every household member aged 16 or older provides additional individual information (for details on the GSOEP see Schupp and Wagner 2002). Between 1992 and 2000, retrospective monthly information about social welfare receipt for each month of the previous calendar year is part of the household questionnaire. Excluding households with a head and if existing her partner aged 61 years or older at the beginning of the spell we observe 619 uncensored or right-censored social welfare spells between January 1991 and December 1999, distributed on 484 households. The maximum number of spells of each household is five (one household), 379 households experience one spell of social welfare receipt (table 1). These spell data are combined with several time-variant and time-invariant household and individual characteristics.

*Table 1 about here*

In the data there are 432 uncensored and 187 right-censored observations. We are interested in the transition from social welfare to a situation with employment income. Therefore we differentiate between transitions to employment (187 cases) and alternative transitions (245 cases). A transition to employment is defined as a situation with at least one adult household member (head of the household or her partner) working full-time subsequent to benefit receipt, at the latest beginning two month after the spell ending.

Descriptive statistics are documented in table 2. These statistics refer (a) to the status at the beginning of a welfare spell (n=619) and (b) to the monthly status (every month one observation, n=8506). Spell observation mostly ends within one year (71.2%), afterwards the number of spells ending decreases constantly. About 16% of the whole sample end in the second year of observation, 2% in the fifth year of social assistance receipt. These proportions refer to all spells, independent of the censor status. To control for the economic situation we include time dummies for each year of observation. The proportion of spells beginning in different years range from 5% in 1991 up to 15% in 1999.[4] Disproportional numbers of welfare spells start in January or

[4] One has to be careful with interpretations of these descriptive statistics. For example the increase in social assistance spells beginning in 1999 can be at least partly explained by the new sub sample F

end in December. Therefore we include January and December dummies in our analyses. The head of a household or her partner is aged older than 50 years in 12% of the observed spells and in 25% a foreign head or partner is living in the household. Children aged 6 years and younger live in 39% per cent of the households, children between 6 and 18 in 34%. In nearly all households the head or her partner has a school graduation (94%) while only in about one third of all households at least one of these persons has finished vocational training (32%). In every tenth household the head or his partner is handicapped, which means that at least one of these persons answers the question whether he or she is officially registered as having a reduced capacity for work or of being severely disabled with yes. The statistics based on the observed months differ from the reported statistics due to the higher weight of longer spells.

*Table 2 about here*

Before discussing the ratio and the difference between the potential household income in case of one adult person working full-time and the social assistance amount, we describe the estimating and calculating procedures of these two income sources separately in the following.

### 3.1 Estimation of Potential Net-Income

In a first step we estimate potential gross market wages of all heads of the household and as the case may be of their partner. We cannot observe their wages directly because most of the individuals in our data set are not working while receiving social assistance. Therefore we estimate the potential wages using all individuals in working age. Whether or not we observe wages depends on an individual's participation decision. Due to this self-selection we cannot assume the sample of workers to be a random sample of all potential working individuals and we have to account for the sample selection problem.

The sample selection model we apply, also referred to as the type II Tobit model, consists of a log-linear wage equation

$$\log w_i = X_{1i}\beta_1 + \varepsilon_{1i}$$

with $X_{1i}$ as a vector containing exogenous characteristics and $w_i$ as person's $i$ wage and an equation describing the binary choice to work or not to work and therefore determining the sample selection

$$z_i^* = X_{2i}\beta_2 + \varepsilon_{2i}\,.$$

(Innovation Sample) of the GSOEP in 2000. Due to this new sample F the sample size of the GSOEP has increased substantially.

We observe wages according to the rule:

$$w_i = w_i^*, z_i = 1 \text{ if } z_{ii}^* > 0$$
$$w_i \text{ not observed}, z_i = 0 \text{ if } z_{ii}^* \leq 0$$

whereby $z_i$ indicates working or not working and this depends on the characteristics $X_{2i}$. Following Heckman (1979) one can estimate the wage equation consistently assuming that the two error components of the two equations follow a bivariate normal distribution.

We estimate two models, one for East and one for West Germany with a pooled sample using the GSOEP waves from 1991 – 1999. The estimation results are reported in tables 3 and 4. We control for the year and the region. As one can see, the inverse Mills ratio term is positive and statistically significant in both models. Education, measured in years, age and firm specific capital, measured in years being employed at the actual employer, have significantly positive influence on the wage per hour, while the squared age and the squared firm specific capital have a significantly negative impact. Foreigners and women have lower wages in both regions and the absence from the labour market in years, accounting for the previous five years, have a negative impact on the wage. While the squared absence from the labour market influences the wage positively in East Germany, the effect in West Germany is insignificant.

*Table 3 about here*

*Table 4 about here*

Using these estimation results, we calculate a potential monthly full-time gross wage for each head of household and her partner. Calculating the potential net income, we assume that in the case of a partner household the person with the higher income would work and we account for income taxes, social security contributions and child allowance.

### 3.2 Social Assistance

The amount of social assistance was not asked in all waves of the GSOEP. Furthermore, in the years the amount of social assistance was part of the questionnaire, the actual amount but not the monthly amount during the previous year was asked. Therefore we can observe the monthly receipt as a binary variable but not the corresponding amount of social assistance.

Instead of direct observation we calculate the maximum of social assistance. As described above this amount depends on the number and the age of household members and varies with the region and the year of receipt. We use the average yearly individual

basic allowances for East and West Germany to calculate the basic allowance for each household member and add them up. Moreover we consider the one-time payments by using the same method as Breuer and Engels (2003) or Boss (2002): We calculate 16% of the individual basic allowance for the head of household, 17% for the partner and 20% for each child. In addition to that we take an allowance for housing depending on the household size into account.

### 3.3 Ratio and Difference between Employment Income and Social Assistance

We calculate two variables: (a) The difference between the potential household net income in case of one person working fulltime and the amount of transfer payment and (b) the ratio of these two income sources. The nominal differences are deflated with respect to the year 1995. We estimate different models using variables (a) or (b). The empirical distributions of the ratio and the gap between the two income sources in the first month of each spell are plotted in figures 2 and 3. The difference as well as the ratio distribution indicates that the incentives to search for a job may be low for a lot of individuals being on social welfare. Compared to the distributions of the variables corresponding to all observed months (see appendix, figures 6 and 7), the mean of the difference as well as of the ratio is relatively high, which reflects the higher weight of longer spells in the distributions of all month-observations. This indicates that a lower income ratio and income difference may go along with a longer stay in the social assistance.

One could argue that the difference between potential household net income could never be negative because these households would receive supplementary transfer payments (see section 2). Nevertheless we use these negative differences in our analysis. One can observe households who are eligible for social assistance but do non take it up. This (non-) take-up behaviour depends among others on the expected benefit amount (see e.g. Riphahn 2001): The probability of take-up rises with the potential amount of transfer payments. Because we are interested in the leaving processes of social assistance, the difference or ratio between the two separate income sources and not the combination of the different income sources is the relevant variable.

**Fig. 2: Histogram: Monthly difference between potential net income and the amount of social assistance in Deutsche Mark (in prices of 1995), first month of welfare spell (n=619)**

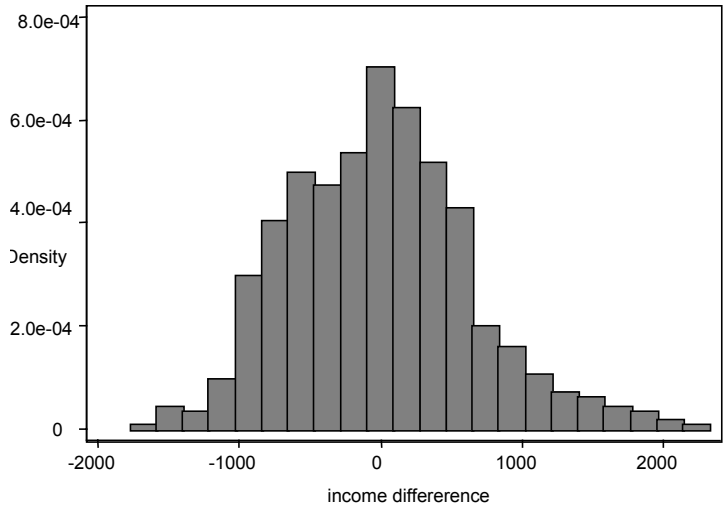
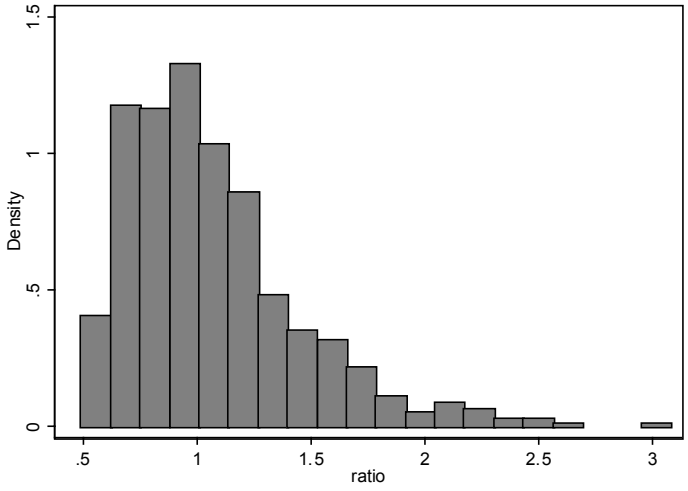



**Fig. 3: Histogram: Ratio between potential net income and the amount of social assistance, first month of welfare spell (n=619)**




## 3.4 Model Specification

The process of leaving welfare in favour of labour income can appropriately be modelled by a transition rate approach. According to the type of data being used here, a discrete hazard rate model has to be applied (see for example Hamerle and Tutz 1989; Han and Hausman 1990; Meyer 1990; Sueyoshi 1992; Narendranathan and Stewart 1993). The hazard rate is defined as the limit of the conditional probability for the ending of a spell in interval $[t;t+\Delta t[$ given that no transition occurred before the start of this interval:

$$\lambda_i(t) = \lim_{\Delta t \to 0} \frac{P(t \le T_i < t+\Delta t \mid T_i \ge t)}{\Delta t}$$

where $T_i$ denotes the length of a spell of the $i$th observation. $T_i$ is assumed to be a continuous, non-negative random variable. The probability that a spell is at least of length $t$ is given by the so-called survivor function:

$$S_i(t) = \exp\left( - \int_0^t \lambda_i(\tau)\, d\tau \right)$$

The hazard rate is depending on individual specific covariates. In the usually applied Cox model (Cox 1972) it is assumed that the influence of these covariates causes proportional shifts of a so-called baseline hazard rate:

$$\lambda_i(t) = \lambda(t \mid \boldsymbol{x}_i(t), \varepsilon_i) = \lambda_o(t) \exp[\, x_i(t)\, \beta + \varepsilon_i \,]$$

with $\lambda_o(t)$ denoting the baseline transition rate, $x_i(t)$ a row vector of covariates, $\beta$ a column vector of parameters and $\varepsilon_i$ an individual specific error term, representing the joint influence of unobserved heterogeneity.

Discrete-time measurement leads to the simplifying consequence that instead of continuous levels of $\lambda_o(t)$ and $x_i(t)$ only their interval specific mean levels have to be taken into account (Kiefer 1988). Assumed that the time axis is divided into intervals of unit length, a given spell consists of a number of $k$ intervals, in the following referred to as subspells. The $j$th subspell covers a range from $t = j\text{-}1$ to $t + 1$, but excluding $t + 1$. The interval specific means of $\lambda_o$ and $x_i(t)$ are then denoted as $\lambda_j$ and $x_{ij}$.

With this notation, the survivor function can be rewritten as:

$$S(\, j \mid \boldsymbol{x}_i, \varepsilon_i) = \exp\left( - \sum_{k=1}^{j} \int_{k\text{-}1}^{k} \lambda(\tau \mid x_{ik}, \varepsilon_i)\, d\tau \right)$$

$$= \exp\left( - \exp(\varepsilon_i) \sum_{k=1}^{j} \exp[\, x_{ik}\, \beta + \gamma_k] \right)$$

$$\text{with } \gamma_k = \ln\left( \int_{k\text{-}1}^{k} \lambda_k\, d\tau \right)$$

In the absence of unobserved heterogeneity, the survivor function ends up in an expression which corresponds to an interval specific binary choice model with an extreme value distributed error term. In the following, we therefore refer to this model as random effects extreme value model.

Here, the parameter $\gamma_k$ has a clear interpretation in terms of an interval specific mean of the baseline hazard rate. Instead of specifying a functional form for the baseline hazard, its time pattern can immediately be taken from the sequence of $\gamma$ parameters.

„Integrating out" the unobserved error term $\varepsilon_i$ may be accomplished in several ways. Heckman and Singer (1984) propose the use of non-parametric methods, to avoid sensitivity of the estimation results with regard to a parametric specification of the distribution of $\varepsilon_i$. However, Trussell and Richards (1985) find that a major part of the sensitivity encountered by Heckman/Singer may also have its origin in the parametric specification of the baseline transition rate. The trade-off between a correct specification of the error distribution and a correct specification of the baseline transition rate is confirmed by Narendranathan and Stewart (1993). Following a common approach, it will here be assumed that $u_i = \exp(\varepsilon_i)$ is gamma distributed with expectation 1 and variance $\sigma^2$. For the survivor function this implies (Lancaster 1979):

$$S(j \mid \boldsymbol{x}_i) = \int_0^{\infty} \exp\left( - u_i \sum_{k=1}^{j} \exp[\, x_{ik} \boldsymbol{\beta} + \gamma_k] \right) f(u_i)\, \mathrm{d}u_i$$

$$= \left[ 1 + \sigma^2 \sum_{k=1}^{j} \exp[\, x_{ik} \boldsymbol{\beta} + \gamma_k] \right]^{-\sigma^{-2}}$$

Following from this, the probability of a transition in the $j$th interval is simply given by the difference of two survivor functions:

$$f(j \mid \boldsymbol{x}_i) = S(j\text{-}1 \mid \boldsymbol{x}_i) - S(j \mid \boldsymbol{x}_i)$$

Together, the probability function and the survivor function allow for the specification of the likelihood function:

$$\mathrm{L}(\boldsymbol{\beta},\sigma^2) = \prod_{i=1}^{n} f(j_i \mid \boldsymbol{x}_i)^{c_i}\, S(j_i \mid \boldsymbol{x}_i)^{1-c_i}$$

$$= \prod_{i=1}^{n} c_i\, S(j_i\text{-}1 \mid \boldsymbol{x}_i) - (2c_i - 1)\, S(j_i \mid \boldsymbol{x}_i)$$

$$\text{with } c_i = \begin{cases} 1, \text{ if a spell ends up in intervall } j_i \text{ with a transition} \\ \quad \text{to employment} \\ 0 \text{ otherwise} \end{cases}$$

It implies that right-censored spells are assumed to be censored at the end of the related interval, but that transitions may occur somewhere between $j$-1 and $j$. This may usually be justified as long as the unit length of intervals is sufficiently small. Competing risks can easily be considered under the same assumption. For a small unit length of intervals, the occurrence of ties - the coincidence of competing risks in the

same interval - can be excluded. If in addition, the individual heterogeneity components $\varepsilon_{ir}$ for different potential transitions are mutually independent, the corresponding likelihood can be written in a factorised form. As a consequence, transitions to different destination states can be estimated separately.

Using a cumulative parameterisation of the baseline hazard function, Han and Hausman (1990) show that the proportional hazard specification may also be interpreted in terms of a specific type of the ordered logit model. In practice, however, especially for competing risks models, multinomial logit specifications are often used instead (see e.g. Allison 1982). The multinomial logit specification does not rely on the proportionality assumption. We also include unobserved heterogeneity in the model which can be interpreted as random effects or random intercepts (Rabe-Hesketh et al. 2001). These random effects are assumed to be risk specific and to be constant for a given episode. A similar approach is presented by Steiner (2001) in an analysis of unemployment duration. We assume the unobserved heterogeneity $\varepsilon_{ir}$ to be normally distributed with expectation 0 and variance $\sigma_r^2$, where $r$ is denoting one of the two competing risks being considered. This corresponds to a discrete hazard rate given by:

$$\lambda_{ir}(t|x_i(t),\varepsilon_{ir}) = \frac{\exp(\alpha_r(t) + \varepsilon_{ir} + \beta_r x_i(t))}{1 + \sum_{r=1}^{2} \exp(\alpha_r(t) + \varepsilon_{ir} + \beta_r x_i(t))}, \quad \varepsilon_{ir} \sim N(0, \sigma_r^2),$$

whereby $\lambda_{i1}$ indicates a transition from welfare to work and $\lambda_{i2}$ an alternative transition. The survivor function is given by:

$$S(t|x_i(k),\varepsilon_{ir}) = \prod_{k=1}^{t-1} \frac{1}{1 + \sum_{r=1}^{2} \exp(\alpha_r(k) + \varepsilon_{ir} + \beta_r x_i(k))}$$

and the likelihood function corresponds to:

$$L = \prod_{i=1}^{n} \int \prod_{r=1}^{2} [\lambda_{ir}(t|x_i(t),\varepsilon_{ir})]^{c_{irt}} \prod_{k=1}^{t-1} [1 - \lambda_{ir}(k|x_i(k),\varepsilon_{ir})] f(\varepsilon_{ir}) d\varepsilon_{ir},$$

whereby $c_{irt} = 1$ indicates an uncensored ending of the welfare spell in period t. The contribution of the episode $i$ to the sample likelihood is given by integration of the transition probability for an uncensored and of the survivor function for a censored spell over the random intercepts distribution[5].

[5] The maximum likelihood function is solved applying Gauss-Hermite quadrature using the routine gllamm (Generalized linear latent and mixed models) for Stata, developed by Rabe-Hesketh et al. (2001, 2002).

## 4 Empirical Results

We estimate two discrete time models, an extreme value model with the assumption of an underlying proportional hazard rate and a non proportional multinomial logit model. For each specification we estimate two versions: model 1 with the ratio and model 2 with the absolute difference between estimated potential labour income and welfare payment level. The results are reported in tables 5 – 8.

The coefficients of the extreme value model can be interpreted with respect to the underlying continuous time proportional hazard rate. In the model without unobserved heterogeneity, the income ratio as well as the absolute income difference between estimated potential labour income and welfare payment level show a significantly positive effect on the probability of a transition from welfare to work (table 5). A household with a 1000 DM higher monthly income difference has a 36% higher probability of leaving social welfare than a household otherwise being equal. The coefficient of the income ratio indicates that a 0.5 higher income ratio goes along with a 36% higher probability of an exit through work. Assuming households with the same welfare level, a difference in the income ratio of 0.5 stands for a difference in estimated labour income by 50% of the social welfare level. Including unobserved heterogeneity in the extreme value model leads to a still significant positive effect of the income ratio (at the 10 percent level) while the effect of the income difference is still positive but not significant (table 6). The estimated variance of the included unobserved heterogeneity is not significant, but the inclusion of unobserved heterogeneity leads to a significant increase in the log likelihood. In addition to a significant increase in the log likelihood, the estimated variance of the gamma distribution is significant in the case of alternative transitions, reported in table 6.

*Table 5 and 6 about here*

One important restriction of the multinomial logit model is the assumption of independence from irrelevant alternatives (IIA), which is fulfilled in our case according to the result of a Hausman test (Hausman and McFadden 1984). In the model without unobserved heterogeneity the income difference shows a positive influence at the 10% level and the income ratio a positive effect at the 5% level (table 7). The marginal effects for the probability of a transition from welfare to work is positive at the same significance levels reported for the coefficients: An increase in the difference of 1000 DM leads to an increase in the exit probability from 1.03% to 1.33% per month, an

increase in the ratio of 0.5 leads to an increase from 1.02% to 1.34%. These effects are calculated at the means of the independent variables. Including unobserved heterogeneity in the model does not substantially change these results. The estimated variances of and the correlation between the two random intercepts are reported in table 8, whereby the inclusion of the unobserved heterogeneity does not change the log likelihood significantly.

*Table 7 and 8 about here*

The other relevant covariates for the transition from welfare to work are quite similar, independent of the model specification and whether or not we account for unobserved heterogeneity. Households with a head being single have a significant lower probability of leaving social welfare via employment than partner households. This effect is especially strong for women. Households with the head or her partner being older than 50 years have a lower exit probability than younger households. The presence of young children extend the duration of welfare receipt, while older children between 6 and 18 show no significant influence on the duration of social welfare receipt. These are the expected results confirming the literature. Households in East Germany exit faster to employment, which is a surprising result because of the relatively bad economic performance of East Germany. One possible (ad-hoc) explanation may be a relatively large number of transitions into public financed jobs for unemployed persons in East Germany, but this has to be checked empirically. The existence of a handicapped adult household member seems to have no influence on the transition probability. Moreover the nationality of adult household members does not affect the exit probability of households. In addition to that the existence of an adult person with vocational or a school graduation has no influence on the probability of exiting social welfare. This result is similar to that of Riphahn (1999)[6] who identifies only a significant effect for a university degree but not for vocational training while Wilde (2003) and Gangl (1998) identify positive effects of a vocational training.

Our results confirm the theoretical predictions: Given a level of social welfare payment it is more likely to observe exits from social assistance to work, the higher an individuals (net) market wage is. In contrast to other studies like Riphahn (1999) or Wilde (2003) we estimate a positive effect of the potential net labour income on the

[6] Riphahn estimates two sorts of models using different covariables: Duration models with continuous time and household as well as individual characteristics on the one hand and duration models with discrete time and household characteristics with an additional estimated income variable on the other hand. We refer to both model categories.

transition probability, especially if we estimate the influence of the ratio between the estimated labour income and the social assistance.[7] The relative distance between the two income sources has more explanation power than the absolute difference.

In contrast to the transitions to work the two income variables have no significant influence on the probability of alternative transitions, reported in the rows "Alternative Transitions" in the tables 5-8. This is an expected result and shows the importance to differentiate between alternative risks when examining the transition from welfare to work and the role of estimated labour income.

## 5 Conclusion

The aim of the study was to estimate the influence of the ratio between estimated potential labour income and the welfare payment level on the probability of a transition from social welfare to work. We use data from the GSOEP waves 1992-2000 including information about spell duration of households receiving social welfare and the monthly employment status of the household members. The potential net labour income is estimated with standard wage equations accounting for sample selection and applying a simple tax function. We estimate two different discrete time hazard rate models with and without heterogeneity.

Independent of the model specification the ratio between potential labour income and the welfare level shows a positive effect on the probability of a transition to work, while the absolute difference between the two income sources shows mostly but not in all models a significantly positive effect.

Our results are contrary to previous studies dealing with the determinants of welfare spell duration in Germany: We identify an effect of the income ratio according to the standard theoretical predictions. This "new" result derives from a simultaneous consideration of both sources of income, the net household labour income and the social welfare level, and additionally from a differentiation between transitions to work and alternative transitions.

[7] These results do not change substantially if we estimate our models with one observation per household and exclude repeated events.

**Table 1: Number of spells per household**

| number of spells | Freq. | Percent | Cumulated Percent |
|---|---|---|---|
| 1 | 484 | 78.19 | 78.19 |
| 2 | 105 | 16.96 | 95.15 |
| 3 | 25 | 4.04 | 99.19 |
| 4 | 4 | 0.65 | 99.84 |
| 5 | 1 | 0.16 | 100.00 |
| Total | 619 | 100.00 | 100.00 |

**Table 2: Descriptive statistics: (a) each spell one observation (household characteristics in the first month), (b) each month one observation**

| variable | (a) Mean / Share (standard deviation) | (b) Mean / Share (standard deviation) |
|---|---|---|
| Time of Observation | | |
| 1 year | 0.71 | 0.53 |
| 2 years | 0.15 | 0.21 |
| 3 years | 0.06 | 0.11 |
| 4 years | 0.03 | 0.06 |
| >4 years | 0.04 | 0.08 |
| | | |
| Year of Observation | | |
| 1991 | 0.05 | 0.01 |
| 1992 | 0.07 | 0.05 |
| 1993 | 0.11 | 0.09 |
| 1994 | 0.15 | 0.11 |
| 1995 | 0.12 | 0.13 |
| 1996 | 0.10 | 0.14 |
| 1997 | 0.13 | 0.14 |
| 1998 | 0.13 | 0.17 |
| 1999 | 0.15 | 0.16 |
| Ratio between potential net income and benefit amount | 1.07 (0.38) | 0.999 |
| Difference between potential net income and benefit amount | -12 (647) | -125 (647) |
| Handicapped adult household member (1/0) | 0.10 | 0.13 |
| Non German adult household member (1/0) | 0.25 | 0.27 |
| At least one adult household member with vocational training (1/0) | 0.32 | 0.39 |
| At least one adult household member with school graduation (1/0) | 0.94 | 0.90 |
| East Germany (1/0) | 0.27 | 0.22 |
| No partner household (female) (1/0) | 0.35 | 0.42 |
| No partner household (male) (1/0) | 0.08 | 0.07 |
| Children aged 6 and younger (1/0) | 0.39 | 0.43 |
| Children aged between 6 and 18 (1/0) | 0.34 | 0.34 |
| December dummy | 0.06 | 0.10 |
| January dummy | 0.34 | 0.07 |
| Adult household member aged > 50 | 0.12 | 0.18 |
| Sample Size | 619 | 8506 |

**Table 3: Probit Selection Equation**

| | East Germany | | West Germany | |
|---|---|---|---|---|
| | Coefficient | t-value | Coefficient | t-value |
| Education in years | 0.0335 | 9.05 | 0.0174 | 7.88 |
| Age | 0.2326 | 47.92 | 0.2119 | 71.12 |
| Age squared | -0.0031 | -55.11 | -0.0028 | -80.11 |
| Woman | -0.1753 | -10.84 | -0.4536 | -42.38 |
| Foreigner | -0.1565 | -2.25 | -0.0305 | -2.05 |
| children under 6 | -0.2467 | -14.71 | -0.3010 | -32.62 |
| partner living in household | 0.1488 | 7.02 | -0.0616 | -4.57 |
| Constant | -4.075 | -4149 | -3.3921 | -57.32 |
| Sample Size | 31074 | | 70248 | |
| Pseudo-$R^2$ | 0.24 | | 0.22 | |

**Table 4: Wage Estimation. Dependent Variable: ln(wage per hour)**

| | East Germany | | West Germany | |
|---|---|---|---|---|
| | Coefficient | t-value | Coefficient | t-value |
| Education in years | 0.0685 | 48.11 | 0.06868 | 85.74 |
| Age | 0.0510 | 7.88 | 0.0678 | 27.34 |
| Age squared | -0.0006 | -7.24 | -0.0008 | -25.32 |
| Woman | -0.1031 | -15.77 | -0.2414 | -43.73 |
| Absence from the labour market | -0.1150 | -12.17 | -0.0631 | -11.36 |
| Absence from the labour market squared | 0.0070 | 2.19 | -0.0006 | -0.37 |
| firm-specific human Capital | 0.0082 | 8.45 | 0.0143 | 21.83 |
| firm-specific human Capital squared | -0.0001 | -3.58 | -0.0002 | -9.04 |
| Foreigner | -0.0484 | -1.91 | -0.0541 | -9.56 |
| Schleswig-Holstein | - | | -0.0397 | -2.81 |
| Hamburg | - | | 0.0496 | 2.83 |
| Niedersachsen | - | | -0.0324 | -2.82 |
| Bremen | - | | -0.0426 | -1.91 |
| Nordrhein-Westfalen | - | | -0.0123 | -1.14 |
| Hessen | - | | 0.0120 | 1.02 |
| Rheinland-Pfalz / Saarland | - | | -0.0247 | -2.02 |
| Baden-Württemberg | - | | 0.0309 | 2.78 |
| Bayern | - | | 0.0089 | 0.80 |
| Sachsen | -0.1942 | -17.86 | - | - |
| Sachsen-Anhalt | -0.1597 | -13.92 | - | - |
| Thüringen | -0.1894 | -16.34 | - | - |
| Mecklenburg-Vorpommern | -0.1179 | -9.35 | - | - |
| Brandenburg | -0.1382 | -11.74 | - | - |
| 1992 | 0.2794 | 26.57 | 0.0634 | 7.58 |
| 1993 | 0.4869 | 44.86 | 0.1119 | 13.41 |
| 1994 | 0.6052 | 55.41 | 0.1236 | 15.01 |
| 1995 | 0.6788 | 62.00 | 0.1656 | 19.75 |
| 1996 | 0.7136 | 64.62 | 0.1876 | 24.04 |
| 1997 | 0.7437 | 66.48 | 0.1936 | 24.54 |
| 1998 | 0.7599 | 66.77 | 0.2094 | 26.31 |
| 1999 | 0.7525 | 67.67 | 0.2002 | 25.86 |
| Constant | 0.4259 | 2.88 | 0.8104 | 14.59 |
| Mill's Ratio | 0.1885 | 4.61 | 0.1668 | 9.84 |
| Sample Size | 15025 | | 31691 | |
| $R^2$ | 0.50 | | 0.43 | |

**Table 5: Discrete Hazard Rates (Extreme Value Model)**

| | Extreme Value | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Model 1 (Income Ratio) | | | | Model 2 (Income Difference) | | | |
| | Transition to Work | | Alternative Transitions | | Transition to Work | | Alternative Transitions | |
| Variable | Coefficient | t-value | Coef. | t-value | Coefficient | t-value | Coef. | t-value |
| 2 years | -0.32 | -1.47 | -0.66*** | -3.36 | -0.32 | -1.48 | -0.67*** | -3.39 |
| 3 years | -0.68* | -1.84 | -0.58** | -2.18 | -0.69* | -1.85 | -0.60** | -2.22 |
| 4 years | -0.54 | -1.17 | -0.70* | -1.90 | -0.56 | -1.20 | -0.71* | -1.92 |
| 5 and more years | -1.73* | -1.71 | -0.74* | -1.73 | -1.75* | -1.72 | -0.76* | -1.77 |
| Year of observation | | | | | | | | |
| 1992 | -0.90* | -1.83 | 0.06 | 0.12. | -0.84* | -1.72 | 0.06 | 0.13 |
| 1993 | -0.56 | -1.25 | 0.14 | 0.30 | -0.51 | -1.14 | 0.15 | 0.32 |
| 1994 | -0.52 | -1.20 | 0.39 | 0.88 | -0.45 | -1.04 | 0.41 | 0.91 |
| 1995 | -0.65 | -1.48 | 0.06 | 0.13 | -0.57 | -1.30 | 0.07 | 0.16 |
| 1996 | -1.14** | -2.47 | 0.27 | 0.59 | -1.07** | -2.34 | 0.30 | 0.65 |
| 1997 | -0.96** | -2.13 | 0.15 | 0.33 | -0.90** | -2.00 | 0.18 | 0.39 |
| 1998 | -0.80* | -1.81 | 0.34 | 0.76 | -0.74* | -1.67 | 0.37 | 0.82 |
| 1999 | -1.61*** | -3.34 | 0.86* | -1.73 | -1.54*** | -3.22 | -0.82 | -1.68 |
| December dummy | 1.95*** | 13.02 | 2.89*** | 21.22 | 1.96*** | 13.06 | 2.89*** | 21.23 |
| January dummy | -1.44** | -2.02 | -0.86 | -1.47 | -1.44** | -2.02 | -0.87 | -1.47 |
| East Germany | 0.49** | 2.57 | 0.09 | 0.50 | 0.48** | 2.50 | 0.07 | 0.38 |
| At least one adult household member with vocational training | 0.14 | 0.67 | 0.09 | 0.57 | 0.14 | 0.65 | 0.07 | 0.44 |
| At least one adult household member with school graduation | 0.32 | 0.71 | 0.19 | 0.73 | 0.35 | 0.79 | 0.21 | 0.79 |
| No partner household (female) | -1.12*** | -5.57 | 0.22 | 1.46 | -1.10*** | -5.48 | 0.22 | 1.48 |
| No partner household (male) | -0.80** | -2.31 | 0.24 | 0.81 | -0.64* | -1.94 | 0.27 | 0.97 |
| Adult household member aged > 50 | -0.80*** | -2.74 | -0.31 | -1.37 | -0.82*** | -2.81 | -0.33 | -1.44 |
| Children aged 6 and younger | -0.33* | -1.80 | -0.13 | -0.81 | -0.35* | -1.89 | -0.16 | -0.99 |
| Children aged between 6 and 18 | 0.08 | 0.45 | -0.04 | -0.22 | 0.10 | 0.53 | -0.07 | -0.40 |
| Non German adult household member | -0.27 | -1.34 | -0.04 | -0.23 | -0.27 | -1.31 | -0.04 | -0.21 |
| Handicapped adult household member | -0.12 | -0.44 | -0.17 | -0.72 | -0.11 | -0.41 | -0.17 | -0.72 |
| Income Ratio | 0.62** | 2.28 | 0.05 | 0.17 | | | | |
| Income Difference ($*10^{-3}$) | | | | | 0.31** | 2.00 | -0.00 | -0.20 |
| Constant | -3.56*** | -5.28 | -4.54*** | -7.70 | -3.00*** | -4.85 | -4.49*** | -8.47 |
| Log Likelihood | -617.24 | | -731.77 | | -617.74 | | -731.66 | |

612 spells, 8373 months, * significant at the 10 percent level, ** at the 5 percent level, *** at the 1 percent level

**Table 6: Discrete Hazard Rates (Random Effects Extreme Value Model)**

| | Extreme Value with unobserved heterogeneity | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Model 1 (Income Ratio) | | | | Model 2 (Income Difference) | | | |
| | Transition to Work | | Alternative Transitions | | Transition to Work | | Alternative Transitions | |
| Variable | Coefficient | t-value | Coef. | t-value | Coefficient | t-value | Coef. | t-value |
| 2 years | 0.25 | 0.81 | 0.28 | 0.81 | 0.25 | 0.78 | 0.28 | 0.81 |
| 3 years | 0.24 | 0.47 | 1.12** | 1.97 | 0.23 | 0.45 | 1.11** | 1.97 |
| 4 years | 0.57 | 0.85 | 1.54* | 1.82 | 0.55 | 0.82 | 1.54* | 1.82 |
| 5 and more years | -0.44 | -0.40 | 2.29** | 2.13 | -0.46 | -0.41 | 2.28** | 2.13 |
| Year of observation | | | | | | | | |
| 1992 | -1.15* | -1.88 | 0.19 | 0.32 | -1.11* | -1.84 | 0.17 | 0.29 |
| 1993 | -0.62 | -1.18 | 0.24 | 0.38 | -0.57 | -1.11 | 0.22 | 0.35 |
| 1994 | -0.56 | -1.09 | 0.75 | 1.22 | -0.49 | -0.98 | 0.72 | 1.18 |
| 1995 | -0.59 | -1.13 | 0.44 | 0.69 | -0.51 | -1.01 | 0.41 | 0.65 |
| 1996 | -1.11** | -1.99 | 0.71 | 1.11 | -1.04* | -1.91 | 0.71 | 1.12 |
| 1997 | -1.03* | -1.90 | 0.33 | 0.53 | -0.97* | -1.83 | 0.33 | 0.53 |
| 1998 | -0.91* | -1.69 | 1.03 | 1.63 | -0.85 | -1.61 | 1.03 | 1.63 |
| 1999 | -1.73*** | -3.15 | -0.51 | -0.78 | -1.66*** | -3.06 | -0.51 | -0.79 |
| December dummy | 2.06*** | 11.91 | 3.27*** | 15.95 | 2.06*** | 11.95 | 3.27*** | 15.96 |
| January dummy | -1.54** | -2.02 | -0.85 | -1.26 | -1.55** | -2.03 | -0.86 | -1.27 |
| East Germany | 0.70*** | 2.63 | 0.31 | 1.02 | 0.68*** | 2.59 | 0.29 | 0.94 |
| At least one adult household member with vocational training | 0.10 | 0.37 | -0.00 | -0.01 | 0.11 | 0.39 | -0.03 | -0.10 |
| At least one adult household member with school graduation | 0.41 | 0.76 | 0.51 | 0.86 | 0.47 | 0.87 | 0.50 | 0.86 |
| No partner household (female) | -1.57*** | -5.12 | 0.57** | 2.01 | -1.55*** | -5.11 | 0.57** | 2.04 |
| No partner household (male) | -1.11** | -2.10 | 0.77 | 1.37 | -0.87* | -1.79 | 0.73 | 1.42 |
| Adult household member aged > 50 | -0.97*** | -2.61 | -0.82* | -1.75 | -1.00*** | -2.68 | -0.83* | -1.78 |
| Children aged 6 and younger | -0.57** | -2.16 | -0.34 | -1.15 | -0.60** | -2.29 | -0.37 | -1.24 |
| Children aged between 6 and 18 | 0.19 | 0.70 | -0.20 | -0.67 | 0.19 | 0.67 | -0.26 | -0.83 |
| Non German adult household member | -0.34 | -1.18 | 0.16 | 0.47 | -0.33 | -1.15 | 0.16 | 0.46 |
| Handicapped adult household member | -0.31 | -0.76 | -0.33 | -0.76 | -0.30 | -0.74 | -0.34 | -0.78 |
| Income Ratio | 0.77* | 1.75 | -0.28 | -0.57 | | | | |
| Income Difference ($*10^{-3}$) | | | | | 0.35 | 1.43 | -0.02 | -0.79 |
| Constant | -3.40*** | -3.99 | -4.43*** | -4.51 | -2.69*** | -3.77 | -4.67*** | -5.42 |
| $Ln(\sigma^2)$ | 0.36 | 0.94 | 1.13*** | 4.11 | 0.35 | 0.92 | 1.13*** | 4.13 |
| Log Likelihood | -613.81 | | -725.98 | | -614.40 | | -725.86 | |

612 spells, 8373 months, * significant at the 10 percent level, ** at the 5 percent level, *** at the 1 percent level

**Table 7: Discrete Hazard Rates (Multinomial Logit Model)**

| | Multinomial Logit | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Model 1 (Income Ratio) | | | | Model 2 (Income Difference) | | | |
| | Transition to Work | | Alternative Transitions | | Transition to Work | | Alternative Transitions | |
| Variable | Coefficient | t-value | Coef. | t-value | Coefficient | t-value | Coef. | t-value |
| 2 years | -0.39* | -1.72 | -0.75*** | -3.62 | -0.40* | -1.74 | -0.76*** | -3.65 |
| 3 years | -0.75** | -1.92 | -0.67*** | -2.31 | -0.75** | -1.93 | -0.68** | -2.34 |
| 4 years | -0.60 | -1.24 | -0.77** | -2.00 | -0.62 | -1.29 | -0.78** | -2.04 |
| 5 and more years | -1.81* | -1.70 | -0.81* | -1.67 | -1.82* | -1.71 | -0.82* | -1.69 |
| Year of observation | | | | | | | | |
| 1992 | -0.94* | -1.88 | -0.04 | -0.08 | -0.90 | -1.78 | -0.04 | -0.07 |
| 1993 | -0.59 | -1.19 | 0.07 | 0.13 | -0.51 | -1.10 | 0.08 | 0.15 |
| 1994 | -0.50 | -1.08 | 0.39 | 0.77 | -0.45 | -0.95 | 0.40 | 0.79 |
| 1995 | -0.70 | -1.45 | -0.00 | -0.01 | -0.63 | -1.30 | 0.01 | 0.02 |
| 1996 | -1.17** | -2.37 | 0.16 | 0.32 | -1.11** | -2.24 | 0.19 | 0.37 |
| 1997 | -1.00* | -2.08 | 0.04 | 0.08 | -0.95* | -1.96 | 0.07 | 0.13 |
| 1998 | -0.84* | -1.79 | 0.29 | 0.57 | -0.78 | -1.66 | 0.32 | 0.62 |
| 1999 | -1.73*** | -3.24 | -1.02* | -1.80 | -1.67*** | -3.13 | -0.99* | -1.76 |
| At least one adult household member with vocational training | 0.14 | 0.64 | -0.09 | -0.52 | 0.14 | 0.65 | -0.73 | -0.42 |
| At least one adult household member with school graduation | 0.35 | 0.76 | 0.21 | 0.80 | 0.39 | 0.85 | 0.23 | 0.87 |
| No partner household (female) | -1.12*** | -5.52 | 0.15 | 0.93 | -1.10*** | -5.24 | 0.15 | 0.95 |
| No partner household (male) | -0.80** | -2.18 | 0.14 | 0.41 | -0.63* | -1.79 | 0.18 | 0.55 |
| Adult household member aged > 50 | -0.87*** | -2.78 | -0.44* | -1.83 | -0.89*** | -2.84 | -0.45* | -1.90 |
| December dummy | 2.37*** | 14.15 | 3.19*** | 21.54 | 2.38*** | 14.17 | 3.19*** | 21.55 |
| January dummy | -1.41** | -1.97* | -0.85 | -1.44 | -1.41** | -1.98 | -0.85 | -1.45 |
| East Germany | 0.48** | 2.39* | 0.16 | 0.79 | 0.47** | 2.28 | 0.14 | 0.69 |
| Children aged 6 and younger | -0.38** | -1.95 | -0.19 | -1.02 | -0.41** | -2.04 | -0.21 | -0.19 |
| Children aged between 6 and 18 | 0.06 | 0.33 | -0.07 | -0.37 | 0.07 | 0.35 | -0.10 | -0.50 |
| Non German adult household member | -0.29 | -1.33 | -0.06 | -0.33 | -0.28 | -1.30 | -0.06 | -0.31 |
| Handicapped adult household member | -0.19 | -0.74 | -0.20 | -0.75 | -0.18 | -0.72 | -0.19 | -0.74 |
| Income Ratio | 0.63** | 2.32 | 0.09 | 0.31 | | | | |
| Income Difference ($*10^{-3}$) | | | | | 0.30* | 1.88 | -0.00 | -0.02 |
| Constant | -3.53*** | -5.29 | -4.35*** | -6.90 | -2.95*** | -4.63 | -4.27*** | -7.32 |
| Log Likelihood | -1553.05 | | | | -1553.62 | | | |

612 spells, 8373 months, * significant at the 10 percent level, ** at the 5 percent level, *** at the 1 percent level

**Table 8: Discrete Hazard Rates (Multinomial Logit Model with random intercepts)**

| | Multinomial Logit with unobserved heterogeneity | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Model 1 (Income Ratio) | | | | Model 2 (Income Difference) | | | |
| | Transition to Work | | Alternative Transitions | | Transition to Work | | Alternative Transitions | |
| Variable | Coefficient | t-value | Coef. | t-value | Coefficient | t-value | Coef. | t-value |
| 2 years | -0.36 | -1.28 | -0.60** | -2.00 | -0.37 | -1.29 | -0.60 | -2.01 |
| 3 years | -0.70 | -1.50 | -0.37 | -0.82 | -0.70 | -1.50 | -0.37 | -0.82 |
| 4 years | -0.56 | -0.95 | -0.40 | -0.67 | -0.57 | -0.97 | -0.40 | -0.67 |
| 5 and more years | -1.78 | -1.61 | -0.29 | -0.40 | -1.78 | -1.62 | -0.30 | -0.40 |
| Year of observation | | | | | | | | |
| 1992 | -0.99* | -1.79 | -0.02 | -0.04 | -0.95 | -1.73 | -0.03 | -0.04 |
| 1993 | -0.61 | -1.22 | 0.07 | 0.12 | -0.57 | -1.14 | 0.07 | 0.12 |
| 1994 | -0.56 | -1.14 | 0.47 | 0.83 | -0.51 | -1.02 | 0.47 | 0.84 |
| 1995 | -0.74 | -1.48 | 0.07 | 0.12 | -0.68 | -1.35 | 0.07 | 0.12 |
| 1996 | -1.20** | -2.32 | 0.26 | 0.44 | -1.14 | -2.22 | 0.27 | 0.47 |
| 1997 | -1.04** | -2.05 | 0.07 | 0.12 | -1.00 | -1.96 | 0.08 | 0.14 |
| 1998 | -0.91* | -1.78 | 0.42 | 0.71 | -0.85 | -1.68 | 0.43 | 0.74 |
| 1999 | -1.78*** | -3.31 | -0.98 | -1.61 | -1.72 | -3.21 | -0.96 | -1.60 |
| At least one adult household member with vocational training | 0.13 | 0.57 | -0.09 | -0.46 | 0.14 | 0.60 | -0.08 | -0.37 |
| At least one adult household member with school graduation | 0.33 | 0.70 | 0.29 | 0.82 | 0.37 | 0.80 | 0.30 | 0.86 |
| No partner household (female) | -1.19*** | -4.40 | 0.27 | 1.08 | -1.17 | -4.32 | 0.26 | 1.07 |
| No partner household (male) | -0.87** | -2.13 | 0.30 | 0.71 | -0.68 | -1.79 | 0.32 | 0.82 |
| Adult household member aged > 50 | -0.90** | -2.80 | -0.54* | -1.74 | -0.92 | -2.88 | -0.56 | -1.79 |
| December dummy | 2.37*** | 13.32 | 3.32*** | 15.25 | 2.38 | 13.32 | 3.33 | 15.30 |
| January dummy | -1.42** | -1.98 | -0.82 | -1.37 | -1.42 | -1.98 | -0.82 | -1.38 |
| East Germany | 0.49** | 2.26 | 0.18 | 0.78 | 0.47 | 2.18 | 0.17 | 0.70 |
| Children aged 6 and younger | -0.41* | -1.93 | -0.21 | -0.95 | -0.44 | -2.05 | -0.23 | -1.09 |
| Children aged between 6 and 18 | 0.06 | 0.28 | -0.10 | -0.46 | 0.62 | 0.29 | -0.13 | -0.57 |
| Non German adult household member | -0.31 | -1.36 | -0.02 | -0.07 | -0.30 | -1.32 | -0.01 | -0.06 |
| Handicapped adult household member | -0.22 | -0.73 | -0.22 | -0.70 | -0.21 | -0.71 | -0.22 | -0.71 |
| Income Ratio | 0.64** | 2.05 | 0.02 | 0.07 | | | | |
| Income Difference ($*10^{-3}$) | | | | | 0.30* | 1.73 | -0.04 | -0.19 |
| Constant | -3.65*** | -4.93 | -4.71*** | -5.71 | -3.06*** | -4.43 | -4.69 | -5.69 |
| $\sigma^2$ | 0.18 | | 0.73 | | 0.17 | | 0.73 | |
| $\rho$ | -0.56 | | | | -0.18 | | | |
| Log Likelihood | -1552.56 | | | | -1553.12 | | | |

612 spells, 8373 months, * significant at the 10 percent level, ** at the 5 percent level, *** at the 1 percent level

**Figure 5: Histogram: Monthly difference between potential net income and the amount of social assistance in Deutsche Mark (in prices of 1995) (n=8373)**

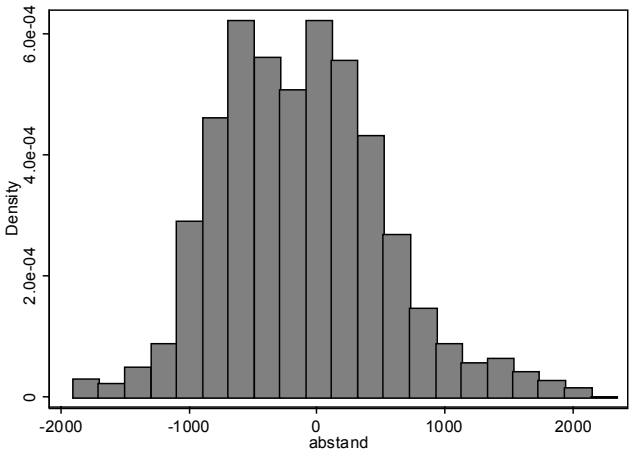
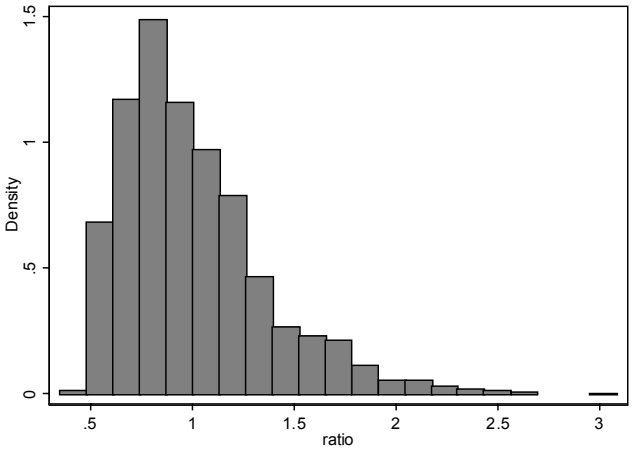



**Figure 6: Histogram: Ratio between potential net income and the amount of social assistance (n=8373)**



## References


Blank, Rebecca M. (1989): Analyzing the Length of Welfare Spells, *Journal of Public Economics*, 39, 245-273.

Boss, Alfred (2002): Sozialhilfe, Lohnabstand und Leistungsanreize. Empirische Analyse für Haushaltstypen und Branchen in West- und Ostdeutschland. Kieler Studien 318. Berlin

Blundell, Richard and Thomas MaCurdy (1999), Labor Supply: A Review of Alternative Approaches. In: Orley Ashenfelter und David Card (Hrsg.), Handbook of Labor Economics, Vol. 3A. Amsterdam, 1559-1591.

Breuer, Wilhelm and Dietrich Engels (2003): Grundinformationen und Daten zur Sozialhilfe. Im Auftrag des Bundesministeriums für Gesundheit und Soziale Sicherung.

Cox, D. R. (1972), Regression Models and Life Tables“, *Journal of the Royal Statistical Society* Series B, 34, 187-202.

Engels, Dietrich (2001): Abstand zwischen Sozialhilfe und unteren Arbeitnehmereinkommen: Neue Ergebnisse zu einer alten Kontroverse, *Sozialer Forschritt*, 50, 56-62.

Gangl, Markus (1998): Sozialhilfebezug und Arbeitsmarktverhalten. Eine Längsschnittanalyse der Übergänge aus der Sozialhilfe in den Arbeitsmarkt, *Zeitschrift für Soziologie*, 27, 212-232.

Gittleman, Maury (2001): Declining Caseloads: What Do the Dynamics of Welfare Participation Reveal? *Industrial Relations*, 40, 537-570.

Hamerle, Alfred, and Gerhard Tutz (1989), Diskrete Modelle zur Analyse von Verweildauern, Frankfurt/Main, Campus.

Han, Aaron, and Jerry A. Hausman (1990), Flexible Parametric Estimation of Duration and Competing Risk Models, *Journal of Applied Econometrics,* 5, 1-28.

Hausman, Jerry and Daniel McFadden (1984): Specification Test for Multinomial Logit Model, *Econometrica*, 52, 1219-1240.

Heckman, James J. (1979): Sample Selection Bias as a Specification Error, *Econometrica*, 47, 153-162.

Heckman, James J., and Burton Singer (1984), Econometric Duration Analysis, *Journal of Econometrics*, 24, 63-132.

Kiefer, Nicholas (1987), Analysis of Grouped Duration Data. In: *Statistical Inference from Stochastic Processes*, edited by N. U. Prabhu. Providence, RI: American Mathematical Society, 107-137.

Lancaster, Tony (1979) Econometric Models for the Duration of Unemployment, *Econometrica*, 47, 939-956.

Meyer, Bruce D. (1990), Unemployment Insurance and Unemployment Spells, *Econometrica*, 58, 757-782.

Moffitt, Robert (1992): Incentive Effects of the U.S. Welfare System: A Review, *Journal of Economic Literature*, 30, 1-61.

Narendranathan, Wiji, and M.B. Stewart, How Does the Benefit Effect Vary as Unemployment Spells Lengthen?, *Journal of Applied Econometrics*, 8, 361-381.

Ochel, Wolfgang (2003): Welfare to Work in the U.S.: A Model for Germany? *Finanzarchiv*, 59, 91-119.

Rabe-Hesketh, Sophia, Andrew Pickels and Anders Skrondal (2001): GLLAMM Manual, Technical Report, Department of Biostatistics and Computing Institute of Psychiatry, Kings`s College, London.

Rabe-Hesketh, Sophia, Anders Skrondal and Andrew Pickels (2002): Reliable Estimation of generalised linear mixed models using adaptive quadrature. *The Stata Journal*, 2, 1-21.

Riphahn, Regina T. (1999): Why Did Social Assistance Dependence Increase? - The Dynamics of Social Assistance Dependence and Unemployment in Germany, 1999, *unpublished Habilitationsschrift* , University of Munich.

Riphahn, Regina T. (2001): Rational Poverty or Poor Rationality? The Take-up of Social Assistance Benefits, *Review of Income and Wealth,* 47, 379-398.

Schupp, Jürgen and Gert G. Wagner (2002): Maintenance of and innovation in long term panel studies: The case of the German Socio-Economic Panel (GSOEP), *Allgemeines Statistisches Archiv*, 86, 163-175.

Steiner, Viktor (2001): Unemployment Persistence in the West German Labour Market: Negative Duration Dependence or Sorting? *Oxford Bulletin of Economics and Statistics*, 63, 91-113.

Stewart, Jennifer and Martin D. Dooley (1999): The Duration of Spells On Welfare and Off Welfare among Lone Mothers in Ontario, *Canadian Public Policy*, 25 Supplement 1, S47-S72.

Sueyoshi, Glenn T. (1992), Semiparametric Proportional Hazards Estimation of Competing Risks Models With Time-Varying Covariates, *Journal of Econometrics*, 51, 25-28.

Trussell, James, and Tony Richards (1985): Correcting for Unmeasured Heterogeneity in Hazards Models Using the Heckman-Singer Procedure. In: *Sociological Methodology 1985*, edited by Nancy Brandon Tuma. San Francisco: Jossey-Bass, 242-276.

Voges, Jürgen and Götz Rohwer (1992): Receiving Social Assistance in Germany: Risk and Duration, *Journal of European Social Policy*, 2, 175-191.

Wilde, Joachim (2003): Why do Recipients of German Social Assistance Opt Out? An Empirical Investigation of Incentives with the Low Income Panel. *Jahrbücher für Nationalökonomie und Statistik*, 223, 719-742.