

Aretz, Bodo; Gürtzgen, Nicole

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The Evolution of Wage Mobility in the German Low-Wage Sector - Is There Evidence for Increasing State Dependence?

Bodo Aretz and Nicole Gürtzgen*

Centre for European Economic Research (ZEW), Mannheim

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Abstract

In this paper, we study how wage mobility in the low-wage sector has changed in Western Germany between 1984 and 2004. Using German individual register data, we document a clear upward trend in the persistence of low-wage employment for both men and women. To explore whether the observed rise in persistence is accounted for by an increase in "genuine" state dependence or by compositional shifts of the low-wage sector, we model low-pay transitions by estimating a series of multivariate probit models. To address the initial conditions problem and the endogeneity of earnings attrition, our estimation approach accounts for the selection into low-wage employment and earnings retention. Using the estimates from the transition model, we determine the evolution of genuine state dependence which is defined as the average difference in low-pay transition probabilities conditional on being initially low and high paid, respectively. For men, our findings strongly argue against an upward trend of genuine state dependence, as the latter exhibits a fairly stationary development. This contrasts with women, whose evolution of genuine state dependence displays a slight upward trend.

Keywords: Wage Mobility, Trivariate Probit, Administrative Data

JEL Code: C23, J31, L13

*Centre for European Economic Research, Department of Labour Markets, Human Resources and Social Policy, L 7.1, 68161 Mannheim, Germany, E-Mail: Aretz@zew.de; Guertzgen@zew.de. We thank Melanie Arntz for providing us with programming routines. Financial support from the German Science Foundation (DFG) is gratefully acknowledged (Grant-No. FR 715/9-1).

1 Introduction

A large body of academic work has documented a sharp increase in earnings inequality particularly in the Anglo-Saxon countries over the last three decades (e.g. Acemoglu 2003, Gosling et al. 2000, Levy and Murnane 1992). However, as long as individuals are able to move up the earnings distribution, a high degree of cross-sectional earnings inequality is likely to exaggerate the extent of inequality over a working life. Thus, any analysis of the evolution of lifetime inequality requires investigating the evolution of both inequality and mobility. For example, a rise in wage mobility could mitigate an increase in cross-sectional earnings inequality, because, in that case, a position at the bottom of the distribution would be of more temporary nature. Conversely, if a rise in earnings inequality were accompanied by a decline in mobility, inequality over a working lifetime would increase due to more persistent positions in the earnings distribution.

Even though there is a large literature that addresses both the evolution of inequality and mobility over time (e.g., Cardoso 2006, Dickens 2000, Gottschalk 1997, Kopczuk et al. 2009), little is known about the determinants underlying the evolution of wage mobility. This is particularly relevant as a number of authors have documented a widening in the distribution of labour earnings that is accompanied by a decline in mobility, giving rise to a larger persistence of low-wage employment (see Buchinsky and Hunt (1999) for the U.S., Cardoso (2006) for Portugal and Dickens (2000) for the U.K.). From a welfare perspective, a high degree of low-pay persistence is of particular concern as it tends to marginalise low-wage workers in the long run. As a result, the determinants of low-pay persistence are of considerable interest to policy-makers. For instance, a decline in low-wage mobility that is accounted for by an increase in the fraction of those without educational attainment, would call for appropriate policy interventions aiming at improving these characteristics. In contrast, if a decline in low-wage mobility were caused by increasing state dependence due to stigmatization effects, policy measures aiming at improving observable attributes of low-wage workers are less likely to succeed.

The purpose of the present paper is therefore to fill in this gap and to investigate the determinants of the evolution of wage mobility. Using German administrative data,

we focus on the western German low-wage sector, which is particularly interesting for several reasons. First, while the German wage structure has long been considered relatively stable at lower percentiles (Prasad 2004), the past two decades have seen a clear tendency towards more earnings inequality at the bottom end of the earnings distribution (Dustmann et al. 2009, Kohn 2006). As a consequence, the low-wage sector has increasingly grown in importance. Second, there is evidence that wage mobility has been declining over the last decades (see Gernandt 2009). However, as mentioned above, this does not necessarily imply a larger degree of persistence in terms of "true" or "genuine" state dependence of low-wage employment, which may occur if low-wage employment today causes low-wage employment in the future for reasons of stigmatization or human capital depreciation. Alternatively, a larger degree of persistence may also be the result of a more unfavorable composition of the low-wage sector. The large extent of selection into low-wage employment has been documented by a number of studies dealing with the determinants of wage mobility. A key finding that emerges from this literature is that the extent of genuine state dependence is often considerably reduced once observable attributes and selection into low-wage employment are accounted for (see Cappellari 2002, 2007, Stewart and Swaffield 1999). The overall aim of our analysis is therefore to quantify to what extent the observed decline in German low-wage mobility is caused by a rise in "true" state dependence by distinguishing the evolution of genuine state dependence from pure composition effects.

The data we use to address these questions stem from the *IAB Employment Subsample 1975-2004* (IABS). This administrative data set is a two per cent subsample of the German *Employment Statistics Register* which is based on reports from employers in compliance with the notifying procedure for the German social security system. The *Employment Statistics Register* offers information on individual wage records and a number of individual characteristics for the whole population of employees who are covered by the German social security system. The IABS provides a useful basis for exploring the evolution of the determinants of wage mobility for several reasons. First, the data allow us to retrieve a great deal of reliable information on workers' previous employment histories, which will be used to model the selection of individuals into the low-wage sector. Second, due to the

administrative nature of the IABS, the problem of panel attrition is considerably reduced as the data track individuals over time as long as they are either employed or, alternatively, unemployed with transfer payments. Even though our data feature less panel attrition than survey data, we still face the problem of non-random earnings retention as individuals may become voluntarily or involuntarily unemployed and may fall out of the earnings distribution. Because this drop out is likely to be non-random, we follow the approach of Cappellari and Jenkins (2006) and Cappellari (2007) by estimating a series of trivariate probit models, which not only account for the selection into low-wage employment, but additionally model non-random earnings retention. To do so, we will take advantage of the precise information on workers' employment histories in order to find appropriate exclusion restrictions that govern both the process of earnings retention and the initial conditions process.

The remainder of the paper proceeds as follows. Section 2 provides a description of the data and some descriptive results. Section 3 presents the empirical analysis. While Sections 3.1. and 3.2. spell out the estimation strategy and define the measure of state dependence, Section 3.3. presents the empirical results. The final Section 4 concludes.

2 Descriptive Empirical Analysis

2.1 Data and Variable Description

The data for our analysis are taken from the regional file of the IAB employment subsample 1975-2004 (*IABS*). This administrative data set, which is described in more detail by Drews (2008), contains a 2 per cent random sample of all social security records spanning the time period 1975 to 2004. It includes employment records subject to social security contributions as well as unemployment records with transfer receipt. The data is representative of all workers subject to the German social security system and covers approximately 80 per cent of the German workforce. Self-employed workers, civil servants and individuals currently doing their military service are excluded from the data set. We restrict our analysis to Western Germany as Eastern Germany experienced profound political and structural changes during our observation period. We further restrict our

sample to the years 1984 to 2004 due to a structural break concerning the wage information, which took place in 1984. The *IABS* provides individual information on (daily) wage records, workers' employment histories and a number of individual characteristics such as age, education, nationality and occupational status. Since the dataset is comprised out of spell data, the data allow us to retrieve a large number of variables used to proxy the stability of workers' employment histories. In particular, the data allow us to measure tenure at the current employer, the number of previous un- and nonemployment spells, the number of previous employers, the number of employment interruptions at the current employer as well as the cumulative duration of previous un- and non-employment spells. A full description of variables used in our analysis can be found in Table B1 in the Appendix.

Although the *IABS* provides a great deal of information on workers' employment histories, the data set has some disadvantages as well. First, the data set provides information on daily – not hourly – wages. While we observe an individual's full time or part time status (defined as working less than 30 hours per week), the data lack explicit information on the number of hours worked. For this reason, we restrict our sample to full-time workers. Second, the data do not allow a distinction between involuntarily unemployed individuals without transfer receipt and individuals who left the labour force or who became self-employed or civil servants. To distinguish more precisely between voluntary and involuntary unemployment, we follow the assumptions proposed by Lee and Wilke (2009) about when the state of unemployment is reached. Further information on the definition of un- and non-employment spells as well as on adjustments of the wage and the education information can be found in the Appendix section B.

2.2 Definition of Low-Pay Status

To study the evolution of wages in the low-wage sector, we define the low-pay status as (full-time) earnings less than two thirds of the median wage. This definition has been used in several other studies (e.g. OECD 1998, Stewart and Swaffield 1999 and European Commission 2004).¹ In order to calculate the low-pay threshold for each year, we consider

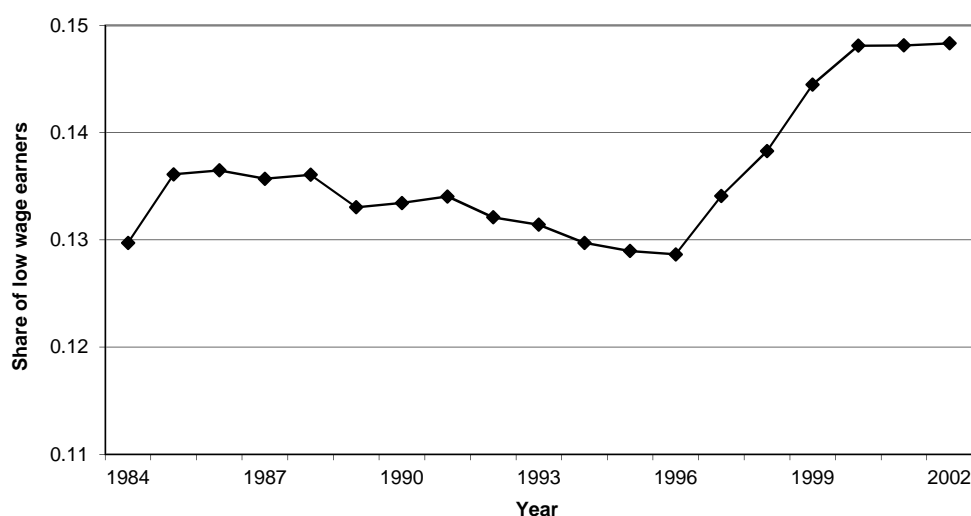
¹Weighting the wage information by the length of the spell did not alter the results substantially.

only spells which include the set date June 30th.² For each year of our observation period, we start from the population of full-time employed workers for whom we define the low-pay status and the respective pay and employment status five years later. This gives rise to 16 relevant transition periods (1984/1989, 1985/1990, ..., 1999/2004). The low or high-pay status is only defined for those who stay full-time employed five years later. We therefore construct a dummy measuring full-time earnings retention, which takes on the value of zero if individuals fall out of the fulltime earnings distribution (i.e. become part-time employed, apprentices, involuntarily unemployed or non-employed).

2.3 The Evolution and Pattern of Low-Pay Transitions

In line with the definition of section 2.2, Figure 1 displays the fraction of low paid individuals among all full-time employees for the years 1984 to 2002.³

Figure 1: Share of low-wage earners in the total population, 1984 - 2002



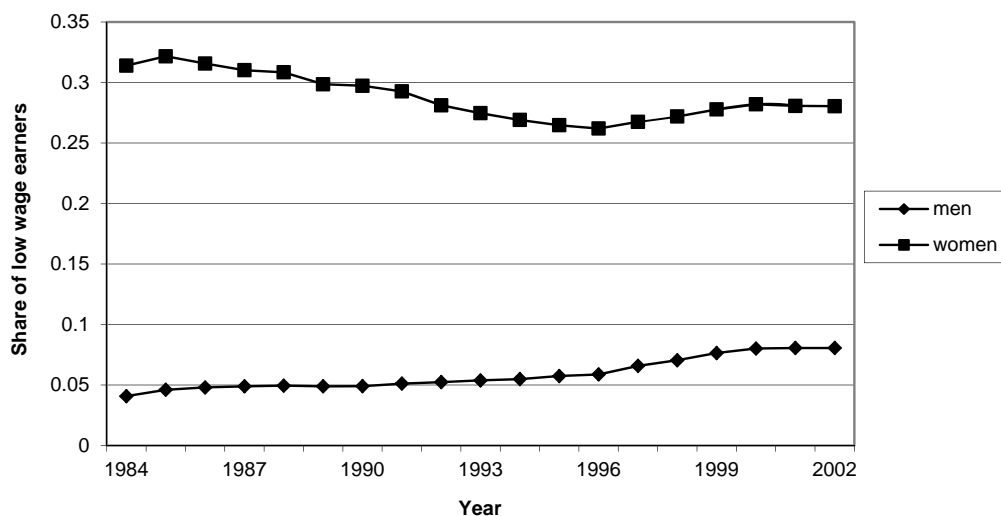
The figure shows that the extent of the low-wage sector has been quite stable in Germany until the mid nineties with a slightly decreasing trend until 1996. After that the low-wage sector experienced a clear upward trend: in 2002, the rate of low-wage earners was 14.9

²Note, however, that we exploit the full spell structure of the data to construct information on individuals' employment histories.

³All figures and tables introduced in section 2.3 can be found in the Appendix sections A and B.

per cent, 14 per cent higher than in 1996.⁴ As shown by Figure 2, the overall increase has been fully made-up by male workers. The share of low paid male workers, which was in the range of 5 to 6 per cent between 1984 and 1996, increased to 8 per cent in 2002. In contrast, the share of low paid women dropped from 33 per cent in 1985 to 28 percent in 2002. Obviously, however, female workers still exhibit a much higher probability of being low paid than male workers. Compared with workers from the high-wage sector, low paid individuals are on average four years younger, less well educated and are more likely to be of foreign nationality (see Table B2). Also, as one would expect, total previous un and non-employment duration periods have on average been longer for individuals in the low-wage sector, whereas their job tenure is on average shorter than that for individuals working in the high-wage sector.

Figure 2: Share of low-wage earners in the total population by sex, 1984 - 2002



In order to describe the overall pattern of the evolution of low-wage persistence, we begin by calculating for each year a measure of aggregate state dependence (*ASD*). This measure is defined as the difference between the probability of staying in the low-wage sector and the probability of entering the low-wage sector five years later, i.e. $ASD = \Pr(L_t = 1 | L_{t-5} = 1) - \Pr(L_t = 1 | L_{t-5} = 0)$. Note that this measure does not account for

⁴This development confirms earlier results by Rhein et al. (2005) who find a very similar evolution of low-wage earners with the same data set for the years between 1990 and 2001. Bosch and Kalina (2007), who use the same data set, arrive at similar results though at a slightly higher low-wage level.

the heterogeneity across formerly low and high paid workers. Later in our analysis, we will therefore attempt to distinguish this measure from the degree of genuine state persistence (*GSD*). This is the measure of "true" persistence we are ultimately interested in and will be defined as the difference in transition probabilities conditional on being initially low and high paid, respectively, therefore providing a measure of differences in price effects of observables across the two groups (see Section 3.2).

Figure 3: Aggregate state dependence by sex, 1984 - 1999

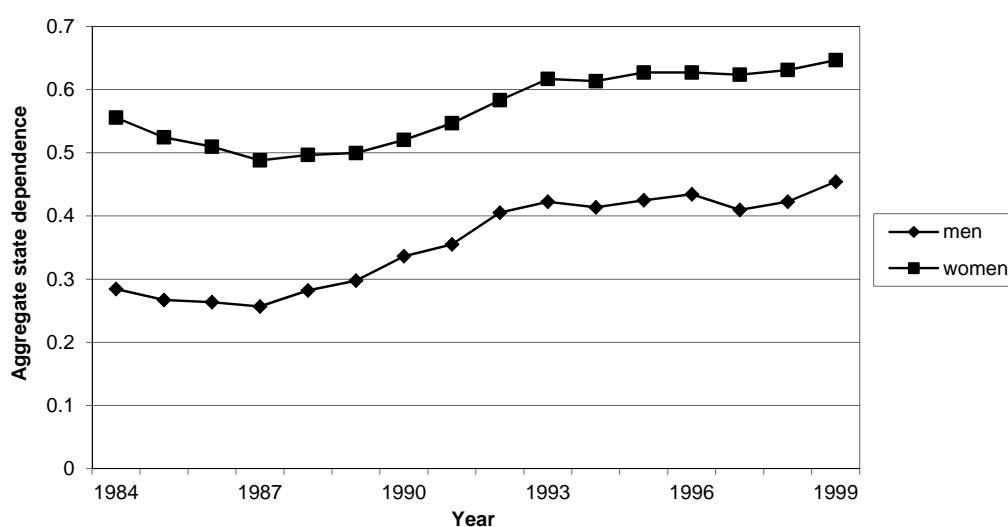


Figure 3 illustrates the evolution of *ASD* between 1984 and 1999 for males and females. As the probability of entering the low-wage sector is typically close to zero, the extent of *ASD* is mainly determined by the persistence probability. In 1984, for example, the overall persistence probability was 50 per cent, i.e. one out of two individuals who earned a low wage and still worked five years later kept his or her low-pay status. The probability of moving from high pay to low pay, however, has been very small in 1984, with only every 20th woman and approximately every 100th man entering the low-wage sector. The first noteworthy fact that emerges from Figure 3 is that the level of *ASD* is about 20 percentage points higher for women than for men. This means that men are considerably less likely to stick to a low-wage job. The development of *ASD* has been quite similar for both sexes, though. Women experienced a larger decrease in *ASD* in the mid 1980's than men, with both sexes reaching their lowest value in the year 1987 with an *ASD* of

26 per cent for males and an *ASD* of 49 per cent for females. Since 1987, *ASD* has been continuously rising with only a slight attenuation period during the mid 1990's. In 1999, the male *ASD* amounted to 45 per cent, whereas women faced an *ASD* value of 65 per cent. Thus, (Western) Germany has experienced a distinct shift towards a higher degree of low-wage persistence in the last 15 years of the 20th century.

What determines the extent of *ASD*? Turning to the association between low-pay persistence and observable attributes, Figures A1 and A2 provide the evolution of *ASD* by different age and job tenure groups. The figures illustrate that there are remarkable differences across groups, with younger and less experienced workers exhibiting considerable less *ASD* than older and more experienced workers. While the difference between the probabilities of remaining low paid and entering low pay was only 29 per cent for workers under the age of 25 in 1984, workers older than 50 years faced a much higher *ASD* (78 per cent). By 1999, this difference had only slightly decreased. However, the share of old workers in the low-wage sector is by far smaller than the share of young workers below 26 years, see Figure A3. Averaged over the whole observation period, only 3 (27) per cent of male (female) workers above 50 years earn a low wage, whereas 18 (41) per cent of the young workers below 26 years are being low paid. In other words, younger workers have a higher probability of earning a low wage, but are at the same time more likely to move up the wage ladder. Once, however, older workers face a low-wage job, it is much harder for them to escape the low-wage sector. Note that this is consistent with the concave shape of age-earnings profiles typically reported in the literature (see e.g. Murphy and Welch, 1990).⁵ With respect to an individual's tenure, the interpretation works similar. The shorter the tenure at the current job is, the higher is the likelihood of earning a low wage (Figure A4) and - on the other hand - the lower will be the extent of *ASD*.

The level of education is also clearly related to the degree of state dependence: the better an individual is educated the lower is his or her *ASD*. Figure A5 reveals for both sexes that while the *ASD* for the high and low-skilled increased only very slightly over our observation period, medium-skilled workers experienced a much larger increase in

⁵For Germany, recent evidence on age-earnings profiles is provided by Bönke et al. 2011

their *ASD*.⁶ Interestingly, the *ASD* level of medium-skilled workers has converged to the *ASD* level of the low-skilled - a development which can be observed for men as well as for women. Put differently, medium-skilled workers, once earning a low wage, seemed to face about the same risk of being caught in the low-wage sector as their low-skilled counterparts in 1999. As Figure A6 points out, however, the share of medium-skilled men (women) who are low paid is with 5 (28) per cent substantially lower than the share of low-skilled individuals in the low-wage sector (13 per cent for men, 38 per cent for women).

There are further observable characteristics that are associated with different levels of *ASD*. With respect to nationality, for example, Figure A7 reveals that foreign individuals have long faced a larger risk of sticking to a low-wage job. By the end of the observation period in 1999, however, the *ASD* for workers of German nationality had converged to the level of their foreign counterparts for both males and females. Such a convergence can not be found when comparing the development of *ASD* of blue and white collar workers (see Figure A8). Over the whole observation period, the *ASD* of blue collar workers has remained about 10 percentage points higher than that of white collar workers. This pattern holds for male as well as for female workers.

Why did the overall aggregate state dependence increase over time? Besides factors like business cycle effects or other macroeconomic developments, one possible explanation could be a change in the composition of the low-wage sector that caused the extent of *ASD* to rise. To investigate such possible composition effects, it is instructive to depict the evolution of some selected observable attributes in the low-wage sector relative to the respective development in the high-wage sector. As can be seen from Figures A9 and A10, the share of young low paid workers below the age of 26 decreased markedly during the last decades for both males and females. While the proportion of young female workers in the low-wage sector dropped from 40 per cent to 19 per cent over the observation period, the decrease was even larger for male workers below the age of 26. For the latter, the corresponding share fell from 52 per cent in 1984 to 24 per cent in 2002. This is a much more pronounced decline compared to the development in the high-wage population

⁶The spikes in the *ASD* of high skilled low paid males can be explained by the small number of observations (<100 per year).

where the share of young male (female) workers dropped from 13 to 6 (24 to 10) per cent between 1984 and 2002.⁷ Not only the demographic change drives this development. The decision to acquire higher education was taken more often in Germany in 2000 (34%) compared to 1985 (20%) so that a smaller share of workers was available for the labor market before the age of 26 (Federal Statistical Office, 2011).⁸ As a consequence, the sharp decline in the share of young workers among the low paid who - as seen above - face a lower *ASD* than the older workforce may have contributed to the overall observed increase in *ASD* over time.

The development of other observed characteristics, however, works against an increase in *ASD* over the observation period. The sex composition in the low-wage sector, for example, has shifted distinctly towards more men being low paid. While the share of males in the high paid population remained roughly constant within the range of 71 to 74 per cent, the share of male workers among the low paid has steadily risen from 21 per cent in 1984 to 36 per cent in 2002, see Figure A11. Since male workers, on average, face a lower *ASD* than female workers, this change in the sex composition of the low-wage sector has favoured a reduction of the overall *ASD* over time. One explanation for the improving situation of women might relate to the decline in the share of low-skilled workers in the female workforce between 1984 and 2002, as can be seen from Figure A12. However, compared to the development in the high-wage sector where the fraction of the low-skilled decreased from 24 to 10 per cent, low paid females experienced a weaker decline in the fraction of low-skilled workers (from 29 to 21 per cent). Male low paid workers exhibit an even less favorable evolution in their skill composition: While in the high wage sector the fraction of male low-skilled workers declined from 17 to 10 per cent, the corresponding share in the low-wage sector even rose by six percentage points (from 27 to 33 per cent, see Figure A13). Moreover, the low-wage sector experiences a less pronounced increase in the share of high-skilled workers than the high-wage sector, where

⁷The figures of the age and tenure evolution of the high wage sector are available from the authors upon request.

⁸Following the definition of the OECD publication "Education at a Glance", the numbers of the Statistical Office are calculated by dividing the absolute number of first-year students by the population of individuals in the typical age range. The number for the year 2000 is including the states of East Germany.

the fraction of high-skilled doubled between 1984 and 2002. Overall, the less favorable evolution of the skill composition among low paid workers should have contributed to a rise in *ASD*. As, however, medium-skilled workers have steadily approached the *ASD* level of their low-skilled counterparts over time, this price effect is likely to have mitigated the effect of the skill composition on *ASD* over time.

There are further developments that support a compositional explanation of the rise in *ASD* over time. As a consequence of the ongoing technological change and the associated structural shift from production to service industries, the share of blue-collar workers (as compared to white-collar workers) has declined by about ten percentage points in the high-wage population for both sexes between 1984 and 2002, see Figure A14. While a reduction in the share of blue collar workers is also observed for low paid women during the observation period, the opposite is true for male workers in the low-wage sector. Despite the overall decreasing importance of blue-collar jobs, the fraction of blue-collar workers among the low paid has increased between 1984 and 2002. As blue-collar workers are more likely to remain low paid (the *ASD* is - as seen above - ten percentage points higher), the male-specific shift towards more low paid blue-collar jobs is likely to have contributed to the rise in *ASD*.

With respect to nationality patterns, Figure A15 illustrates that particularly the low-wage sector has experienced ups and downs over time. While the share of foreign workers in the male high-wage population remained quite stable between 1984 and 1995 (9 to 10 per cent), this fraction more than doubled from 13 to 28 per cent in the male low-wage sector. This development, which may - albeit to a somewhat less extent - also be observed for females, reflects the rising migration flows especially into the low-wage sector after the fall of communism and the opening of the iron curtain (e.g., Bauer and Zimmermann 1999). By 2002, the share of foreigners among low paid male workers declined back to 20 per cent. As foreign workers generally face a higher risk of remaining low paid, these developments may have fostered an increase in *ASD* until the mid-nineties and a slight decrease afterwards. For an overview, Table B3 summarizes the development of the composition of the low-wage sector for males and females by selected characteristics for the years 1984 and 2002.

Taken together, the descriptions shown above provide clear evidence of an increasing degree of persistence in the low-wage sector. Moreover, the degree (and the evolution) of persistence varies considerably across observable attributes. It remains, however, unclear whether the increasing persistence can be fully attributed to a compositional shift towards more unfavorable observable characteristics. In addition to the observed developments fostering a rise in low-wage persistence, such as the increasing fraction of older workers and the less favorable skill composition among the low paid, a low-wage employment per se might have increasingly caused a low-wage employment in the future, regardless of the evolution of observable attributes. Such a development could stem from an increase in "true" or genuine state-dependence. The question to what extent the observed increase in *ASD* is accounted for by a less favorable composition of low-wage workers or an increase in genuine state dependence will therefore be addressed in the next section using a multivariate econometric framework.

3 Econometric Analysis of Low-Pay Transitions

The aim of the multivariate analysis is to quantify the extent to which the observed decline in the probability of escaping low earnings is caused by an increase in the persistence of low-wage employment. To do so, we characterize the determinants of low-pay persistence and exit rates by explicitly distinguishing between observed and unobserved heterogeneity and true state dependence.

3.1 Model specification

To analyze low-pay transitions for each year in our sample period, we estimate the probability of being low paid in period t , conditional on the lagged pay status in $t - 5$, with $t = 1989 - 2004$. An endogeneity issue which is commonly referred to as the 'initial conditions problem' (Heckman 1981b) arises if the starting point of the earnings process cannot be observed in the data and the unobservables affecting this process are correlated. A common solution is to specify an additional equation for the initial condition and to allow for a correlation between the error terms of the initial conditions and the transition

equation. A second endogeneity issue arises since pay transitions are only observable for employees who stay full-time employed five years later. If unobservables affecting the probability of drop out and the initial low-pay status are correlated, the resulting earnings attrition will be endogenous to the pay transition process.

In order to account for these selections mechanisms, we estimate a trivariate probit model. Our model includes the determination of low-pay status in period $t - 5$ (to account for the initial conditions problem), the determination of whether full-time earnings are observed at both points in time, $t - 5$ and t (earnings retention), the determination of pay status in period t , and finally the correlation of unobservables affecting these processes.

We first specify the initial low-pay status. Let l_{it-5}^* denote a latent low-pay propensity for individual i at the start of the observation period and x_{it-5} represents a set of individual-specific characteristics. x_{it-5} includes five different age groups, tenure, tenure squared, a dummy indicating foreign nationality as well as dummies on educational attainment (three categories) and occupational status to capture labour market experience and human capital endowment. We also include information on the previous employment history, such as the employment status prior to entry into the current establishment, the number of previous unemployment and non-employment spells as well as the cumulated duration of previous unemployment and non-employment spells. We further control for the sectoral affiliation as well as regional dummies. u_{it-5} is the sum of an individual-specific effect, μ_i , and an orthogonal white-noise error, δ_{it-5} , and is assumed to follow a standard normal distribution.

$$l_{it-5}^* = \beta' x_{it-5} + u_{it-5}, \quad u_{it-5} \sim N(0, 1) \quad (1)$$

If l_{it-5}^* exceeds some unobservable value (normalized to zero), individual i is observed to be low paid. We define a binary indicator $L_{it-5} = 1$ if $l_{it-5}^* > 0$ and zero otherwise.

The next process to be specified is the earnings retention. We assume that the propensity to observe full-time earnings of individual i in period $t - 5$ and t can be described by

a latent retention index r_{it}^* ,

$$r_{it}^* = \delta' y_{it-5} + \varepsilon_{it}, \quad \varepsilon_{it} \sim N(0, 1), \quad (2)$$

where the error term ε_{it} is standard normally distributed and specified as the sum of an individual-specific effect, η_i , and an orthogonal white-noise error, ξ_{it-5} . y_{it-5} includes factors affecting both earnings and the attachment to paid employment. y_{it-5} contains x_{it-5} , i.e. we assume that all factors affecting earnings levels are also relevant in determining earnings retention. y_{it-5} additionally includes the number of employment interruptions with the current employer as an explanatory factor for earnings retention. If the latent retention propensity of individual i is lower than some critical unobserved value (again normalized to zero), earnings and low-pay status cannot be observed in period t . Let R_{it} be a binary variable of the earnings retention outcome of each individual, where $R_{it} = 1$ if $r_{it}^* > 0$ and zero otherwise.

The third component of the model is the specification of the low-pay status in period t . We assume that the latent propensity of low-pay can be characterized by

$$l_{it}^* = [(L_{it-5}) \gamma'_1 + (1 - L_{it-5}) \gamma'_2] z_{it-5} + v_{it}, \quad v_{it} \sim N(0, 1), \quad (3)$$

with v_{it} denoting the sum of an individual-specific effect, τ_i , and an orthogonal white-noise error, ζ_{it-5} . The column vector z_{it-5} comprises individual-specific attributes affecting the pay status in t . In order to deal with simultaneous changes in covariates and pay status, the individual characteristics pertain to period $t - 5$. The switching specification in (3) allows the impact of the explanatory variables to differ according to the low-pay status in the initial period. Again, L_{it} denotes a binary variable $L_{it} = 1$ if $l_{it}^* > 0$ and zero otherwise, where L_{it} is only observable if $R_{it} = 1$. As a consequence, the sample likelihood will be endogenously truncated.

We assume that the error terms in each of the three equations are jointly distributed as trivariate normal with unrestricted correlations, which can be written as

$$\rho_1 \equiv \text{corr}(u_{it-5}, \varepsilon_{it}) \quad (4)$$

$$\rho_2 \equiv \text{corr}(u_{it-5}, v_{it}) \quad (5)$$

$$\rho_3 \equiv \text{corr}(v_{it}, \varepsilon_{it}). \quad (6)$$

The cross-equation correlations provide a parameterization of unobserved heterogeneity. The correlation ρ_1 describes the relationship between unobservable factors affecting the initial low-pay status and earnings retention. A negative sign suggests that individuals who were more likely to be low paid in the initial period are more likely to drop out of full-time employment compared with highly paid individuals. The correlation ρ_2 summarizes the association between unobservable factors determining the initial and the current low-pay status. Here a positive sign would imply that individuals earning low pay in t are more likely to remain in the low-pay status. The correlation ρ_3 characterizes the relationship between unobservables affecting the retention propensity and the current low-pay status. A negative sign would indicate that individuals employed at both points in time are more likely to escape low pay in t as compared to individuals dropping out of full-time employment. Estimation of unconstrained cross-correlation coefficients provides a test of whether initial conditions and the earnings retention process may be treated as exogenous. In particular, $\rho_1 = \rho_3 = 0$ would imply that the earnings retention process is exogenous and would give rise to a bivariate probit model. Similarly, testing the exogeneity of initial conditions amounts to testing $\rho_1 = \rho_2 = 0$. Finally, if all cross-equation correlations are zero, then γ_1 and γ_2 can be consistently estimated using univariate probit models on sub-samples depending on individuals' initial pay status ($L_{it-5} = 0$ or $L_{it-5} = 1$).

Estimating the model with unrestricted cross-equation correlations requires identifying restrictions, i.e. variables entering x_{it-5} or w_{it-5} but not z_{it-5} . The validity of the exclusion restrictions imposed for identification is tested using functional form as the identifying restriction. As will be shown in section 3.3.1., excluding the number of previous employers until 1989 and the number of previous unemployment spells after 1989 as well as the

number of employment interruptions with the current employer from the equation for low-pay transitions fulfills the requirements of the validity of these restrictions for the men's model. For women, we use the total unemployment duration as well as the number of employment interruptions with the current employer as our exclusion restrictions from the transition equation.

3.2 Measures of State Dependence

To investigate the extent to which the decline in the probability of escaping low earnings is caused by an increase in the persistence of low-wage employment, we distinguish between aggregate state dependence (ASD) and genuine state dependence (GSD). ASD is obtained by computing the difference in average predicted transition probabilities for those who were low paid in $t - 5$ and for those who were initially highly paid:

$$\begin{aligned}
 ASD &= \frac{\sum_{i \in (L_{it-5}=1, R_{it}=1)} \Pr(L_{it} = 1 | L_{it-5} = 1)}{\sum_i L_{it-5} \cdot R_{it}} - \frac{\sum_{i \in (L_{it-5}=0, R_{it}=1)} \Pr(L_{it} = 1 | L_{it-5} = 0)}{\sum_i (1 - L_{it-5}) \cdot R_{it}} \\
 &= \frac{\sum_{i \in (L_{it-5}=1, R_{it}=1)} \frac{\Phi_2(z_{it-5}\hat{\gamma}_1, x_{it-5}\hat{\beta}; \rho_2)}{\Phi(x_{it-5}\hat{\beta})}}{\sum_i L_{it-5} \cdot R_{it}} - \frac{\sum_{i \in (L_{it-5}=0, R_{it}=1)} \frac{\Phi_2(z_{it-5}\hat{\gamma}_2, -x_{it-5}\hat{\beta}; -\rho_2)}{\Phi(-x_{it-5}\hat{\beta})}}{\sum_i (1 - L_{it-5}) \cdot R_{it}},
 \end{aligned} \tag{7}$$

where $\Phi(\cdot)$ and $\Phi_2(\cdot)$ are cumulative density functions of the univariate and bivariate standard normal distributions. This measure does not take into account individual observed or unobserved heterogeneity.

Genuine state dependence arises if initial low pay causes low-pay employment in the future for reasons of stigmatization or human capital depreciation. The absence of GSD can be directly tested by using the endogenous switching structure in (3) and amounts to testing the null hypothesis $H_0 : \gamma_1 = \gamma_2$. To account for individual-specific heterogeneity the GSD measure is based upon individual-specific probability differences. In particular, GSD is derived by computing for each individual the difference between the predicted transition probability conditional on being initially low paid and the predicted transition

probability conditional on being initially high paid, and then averaging the difference over the sample of those with observed earnings in t and $t - 5$:

$$\begin{aligned}
GSD &= \frac{1}{\sum_i R_{it}} \sum_{i \in R_{it}=1} [\Pr(L_{it} = 1 | L_{it-5} = 1) - \Pr(L_{it} = 1 | L_{it-5} = 0)] = \\
&\frac{1}{\sum_i R_{it}} \sum_{i \in R_{it}=1} \left[\frac{\Phi_2(z_{it-5}\widehat{\gamma}_1, x_{it-5}\widehat{\beta}; \rho_2)}{\Phi(x_{it-5}\widehat{\beta})} - \frac{\Phi_2(z_{it-5}\widehat{\gamma}_2, -x_{it-5}\widehat{\beta}; -\rho_2)}{\Phi(-x_{it-5}\widehat{\beta})} \right] \quad (8)
\end{aligned}$$

The log-likelihood contribution for each individual i with earnings information observed in period $t - 5$ is:

$$\begin{aligned}
\log \mathcal{L}_i &= L_{it-5} R_{it} \log [\Phi_3(g_i \gamma'_1 z_{it-5}, h_i \delta' y_{it-5}, d_i \beta' x_{it-5}; g_i h_i \rho_3, g_i d_i \rho_2, h_i d_i \rho_1)] \\
&\quad + (1 - L_{it-5}) R_{it} \log [\Phi_3(g_i \gamma'_2 z_{it-5}, h_i \delta' y_{it-5}, d_i \beta' x_{it-5}; g_i h_i \rho_3, g_i d_i \rho_2, h_i d_i \rho_1)] \\
&\quad + (1 - R_{it}) \log [\Phi_2(h_i \delta' y_{it-5}, d_i \beta' x_{it-5}; h_i d_i \rho_1)] \quad (9)
\end{aligned}$$

where Φ_3 is the cumulative density function of the trivariate standard normal distribution and $g_i \equiv 2L_{it} - 1$, $h_i \equiv 2R_{it} - 1$, $d_i \equiv 2L_{it-5} - 1$. We compute the trivariate standard normal distribution by applying the Geweke-Hajivassiliou-Keane (GHK) simulator, yielding a maximum simulated likelihood (MSL) estimator (see Cappellari and Jenkins 2006).

3.3 Results

3.3.1 Correlation structure and hypothesis tests

Our estimation sample for each year is based on those individuals for whom we observe full-time earnings in our data set. We estimate a trivariate probit model separately for each of the 16 years. While this procedure enables us to obtain the estimation results in

a computationally acceptable way, we face the drawback of not being able to account for feed-back effects as well as correlations of the error terms over time. That the trivariate probit model is necessary to derive consistent estimates of our parameters of interest is ascertained as for all years the hypothesis that $\rho_1 = \rho_3$ and $\rho_1 = \rho_2$ has to be clearly rejected at the 0.1 per cent significance level for both men and women. This provides evidence of the endogeneity of the initial-conditions process and the earnings retention process. The tests also show for men as well as for women that the hypothesis $\gamma_1 = \gamma_2$ and, thus, the hypothesis of no genuine state dependence must be rejected at the 0.1 per cent significance level in each of the 16 years of our observation period.

Table 1: Equation correlation structure

year	Males			Females		
	ρ_1	ρ_2	ρ_3	ρ_1	ρ_2	ρ_3
1984	-.223***	0.008	0.272**	-.157***	0.607	0.213
1985	-.202***	0.054	0.406***	-.131***	0.082	0.477**
1986	-.176***	-.200	0.348***	-.132***	0.285*	0.412**
1987	-.206***	-.183	0.140	-.115***	0.200	0.409*
1988	-.207***	-.299	0.131	-.118***	0.078	0.544***
1989	-.198***	-.046	0.298**	-.113***	0.114	0.459***
1990	-.199***	-.368*	0.199*	-.117***	0.227	0.436***
1991	-.210***	-.206	-.070	-.132***	0.183	0.394**
1992	-.218***	-.261	0.009	-.132***	0.148	-.194**
1993	-.240***	-.229	0.279	-.142***	-.902***	0.003
1994	-.214***	-.061	-.205**	-.145***	-.127	-.239***
1995	-.223***	0.099	-.176**	-.150***	0.108	-.279***
1996	-.235***	-.351*	-.297***	-.154***	-.103	-.048
1997	-.221***	-.098	-.171**	-.150***	0.053	-.146
1998	-.225***	-.356	-.056	-.144***	0.176	-.209**
1999	-.234***	-.204	-.255***	-.160***	-.275	-.181

Note: ρ_1 : Correlation of initial conditions and retention equation;

ρ_2 : correlation of initial conditions and transition equation;

ρ_3 : correlation of retention and transition equation;

Significance levels: ***: $p < 0.01$; **: $p < 0.05$; *: $p < 0.1$.

The cross equation correlation structure is summarized in Table 1 for both male and female workers. As expected, the correlation between unobservables affecting retention and initial conditions is estimated to be negative, indicating that those who were initially low paid are less likely to be employed at both points in time. A significant correla-

tion between the initial condition and the transition equation can not be found, neither for male nor for female workers. This suggests that any bias due to the selection into low-wage employment influences the transition process through its impact on employer retention. Interestingly, the correlation between the retention and the transition equation has changed over time for men as well as for women. In the first years of our observation period, we observe a significant positive relationship between unobservables affecting the retention propensity and the low-pay transition. This indicates that individuals employed at both points in time are less likely to escape low pay in t as compared to individuals dropping out of full-time employment. In the mid 1990s, the sign switches for both sexes and becomes significant in some years. Thus, in recent years individuals employed at $t - 5$ and t seem to have improved their position compared to individuals dropping out of full-time employment.

Table 2: Inclusion of instruments

year	Males		Females	
	Instr. 1 in retention eq.	Instr. 2 in init.cond.eq.	Instr. 1 in retention eq.	Instr. 2 in init.cond.eq.
1984	0.000	0.000	0.029	0.000
1985	0.004	0.000	0.059	0.000
1986	0.001	0.000	0.652	0.000
1987	0.003	0.000	0.463	0.000
1988	0.000	0.000	0.132	0.000
1989	0.000	0.000	0.460	0.000
1990	0.000	0.000	0.006	0.000
1991	0.000	0.000	0.024	0.000
1992	0.000	0.000	0.008	0.000
1993	0.000	0.000	0.017	0.000
1994	0.001	0.000	0.172	0.000
1995	0.006	0.000	0.382	0.000
1996	0.000	0.000	0.401	0.000
1997	0.001	0.000	0.369	0.000
1998	0.023	0.000	0.013	0.000
1999	0.008	0.000	0.000	0.000

Note: All values are p-values. For men, instrument 1 denotes the number of employment interruptions with the current employer, instrument 2. denotes the number of previous employers until 1989; after 1989 instrument 2 refers to the number of previous unemployment spells. For females, instrument 1 denotes the total unemployment duration, instrument 2 is the number of employment interruptions with the current employer.

To be validly identified, the typical identifying conditions must hold for each of the trivariate models, i.e. the excluded variables need to have a significant impact on retention (the initial condition) but not on the low pay transition. Thus, in our case the conditions require the number of employment interruptions with the current employer to significantly affect the men's probability of staying in the sample in a given year, but at the same time to have no significant impact on the probability of staying in the low-wage sector. As can be seen from Table 2, for male workers the two excluded variables have a significant impact on the retention and the initial conditions equation in each of the 16 years. Also for women we observe for all years a significant effect of the number of employment interruptions with the current employer on the probability of initially earning a low wage. The effect of the total unemployment duration on the retention probability is, however, only significant in every second year.

An overview of the tests concerning the second condition - the insignificant impact of the excluded variables in the transition equation - is given in Table 3 for both men and women. While the first two columns show for each sex the significance tests for each excluded variable separately, the third column display the joint significance test of the two excluded variables in the transition equation. We observe for male workers that the number of previous employers (until 1989) and the number of previous unemployment spells (since 1990) seem to be valid instruments for the initial conditions equation. With the exception of 1987, the impact of these variables on the likelihood to stay in the low-wage sector is insignificant. The number of employment interruptions also turns out to be a suitable exclusion restriction as the variable has no significant impact in the transition equation in the majority of years. As a result, the joint significance test accepts the hypothesis (at the 5 per cent significance level) of the two variables having no impact in the transition equation in 12 out of 16 years.

The tests on the insignificance of the excluded variables for women reveal that both variables seem to provide valid exclusion restrictions for the majority of years. This is shown by a joint significance test that accepts the hypothesis (at the 5 per cent significance level) that the two variables have no impact in the transition equation for 11 out of 16 years. However, for both excluded variables there are some periods where the requirements

for a valid identification are not fully met. While the total unemployment duration has a significant impact in the transition equation in the last three years of our observation period (1997-1999), the number of employment interruptions with the current employer affects the likelihood of staying in the low-wage sector significantly between 1987 and 1991. Since - as seen above - the impact of this variable on the initial conditions equation is at the same time not significant, a valid identification of the women's model might not be achieved for some years in the late 1980's so that the results during that time should be interpreted with caution. Taken together, however, the tests show that for the majority of years, the trivariate probit models are well identified for both men and women.

Table 3: Exclusion of instruments from the transition equation

year	Males			Females		
	Instr. 1	Instr. 2	Instr. 1+2	Instr. 1	Instr. 2	Instr. 1+2
1984	0.159	0.328	0.181	0.488	0.611	0.659
1985	0.026	0.316	0.092	0.838	0.724	0.908
1986	0.382	0.495	0.506	0.604	0.444	0.619
1987	0.024	0.021	0.001	0.994	0.008	0.047
1988	0.333	0.376	0.271	0.498	0.030	0.076
1989	0.008	0.388	0.013	0.715	0.001	0.004
1990	0.268	0.301	0.245	0.818	0.000	0.001
1991	0.042	0.156	0.031	0.702	0.001	0.004
1992	0.007	0.742	0.031	0.467	0.570	0.607
1993	0.388	0.435	0.432	0.849	0.256	0.565
1994	0.576	0.213	0.430	0.342	0.975	0.696
1995	0.071	0.314	0.103	0.301	0.244	0.251
1996	0.443	0.751	0.690	0.142	0.427	0.214
1997	0.069	0.312	0.119	0.027	0.603	0.077
1998	0.037	0.442	0.086	0.007	0.961	0.042
1999	0.898	0.252	0.562	0.074	0.189	0.063

Note: All values are p-values. For men, instrument 1 denotes the number of employment interruptions with the current employer, instrument 2 denotes the number of previous employers until 1989; after 1989 instrument 2 refers to the number of previous unemployment spells. For women, instrument 1 denotes the total unemployment duration, instrument 2 is the number of employment interruptions with the current employer.

3.3.2 Regression results

After having clarified the conditions for identification, we summarize the estimation results

over all 16 year for male workers in Table 4. The table shows for all variables their impact on the retention probability (R_t), the initial conditions (L_{t-5}), the likelihood of entering the low-wage sector conditional on being initially high paid ($L_t | L_{t-5} = 0$) and the likelihood of staying in the low-wage sector conditional on being initially low paid ($L_t | L_{t-5} = 1$). We summarize the estimation results for each equation by two different indicators. To get an impression of each variable's robustness over time, the first column not only displays for each equation the direction of the sign which is most often observed over time but also measures the frequency of the signs' appearance of the estimated coefficients on a scale of one (+/-) to three (+++/- ---), with +++/- --- representing a positive/negative effect in each of the 16 years. The second column provides insights into each covariate's consistency over time by summing up the number of years in which the sign of estimated coefficient switched.

The variables quite robustly explain the dependent variables over time, as for the majority of the covariates the estimated coefficients do not change their sign over time. If they do so, they most likely change it only once, indicating a possible structural change in the variables' influence on the different processes. This result applies to men as well as to women, although the estimation results for men seem to show a somewhat higher degree of robustness in terms of both indicators.

Turning first to the earnings retention and the initial conditions process, a comparison of the equations reveals that for male workers almost all covariates exhibit opposite signs with respect to their impact on both processes (see Table 4 for men and Table 5 for women). Characteristics that reflect unstable employment records like longer un and non-employment durations, a higher number of non-employment spells as well as employment interruptions and a change of the employer, favor the likelihood of being initially low paid and reduce the probability of remaining full-time employed five years later. Also, being foreign makes it more likely to be initially low paid and decreases the retention probability.

For other variables, the interpretation works just the other way around: not surprisingly, a higher tenure as well as a higher education reduce the probability of an initial low-wage status and increase the retention probability. Only for blue-collar workers as

Table 4: Regression Summary for men, 1984 - 1999

Variables	R_t		L_{t-5}		$L_t L_{t-5} = 0$		$L_t L_{t-5} = 1$	
	Robustness of direction	number of years in which sign switched	Robustness of direction	number of years in which sign switched	Robustness of direction	number of years in which sign switched	Robustness of direction	number of years in which sign switched
Age 26 - 30 (ref.: age 20 - 25)	+++ ¹	0	---	0	---	0	+++	0
Age 31 - 35	+++	0	---	0	---	0	+++	0
Age 36 - 40	++	1	---	0	---	0	+++	0
Age 41 - 45	++	1	---	0	---	0	+++	0
Age 46 - 50	--	1	---	0	--	2	+++	0
Age 51 - 55	---	0	---	0	-	0	+++	0
Tenure	+++	0	---	0	---	0	+++	0
Tenure squared	---	0	++	1	++	1	--	3
Being foreign (=1)	---	0	+++	0	+++	0	++	4
Medium-skilled (ref.: low-skilled)	+++	0	---	0	---	0	---	0
High-skilled	+++	0	---	0	---	0	---	0
Previously full-time empl. (=1)	+++	0	+++	0	---	0	+	6
Total non-employment duration	---	0	+++	0	+++	0	+++	0
Total unemployment duration	---	0	+++	0	+++	0	+++	0
Number of non-empl. spells	---	0	+++	0	+++	0	++	3
Employment interruption (=1)	---	0	+	7	+	7	++	2
Change of employer (=1)	--	1	+++	0	+++	0	--	4
Blue collar worker (=1)	+++	0	+++	0	+++	0	+++	0
Number of unemployment spells	+++	0	---	0	---	0	---	0
Number of previous employers	---	0	---	0	+++	0	+	3
Number of empl. interruptions	+++	0						

¹⁾ Measure of direction robustness:

+++/- - -: Variable has a positive/negative impact in all 16 years.

++/--: Variable has a positive/negative impact in 12 to 15 of the 16 years.

+/-: Variable has a positive/negative impact in 9 to 11 of the 16 years.

well as for previous full-time employed individuals, we observe that the signs of the two coefficients point into the same direction. The retention probability as well as the probability of being initially low paid is higher for blue-collar (than for white-collar) workers and for those individuals who were previously full-time (rather than not full-time) employed. At first glance, the positive association between a previous full-time employment status and the probability of being initially low paid appears to be somewhat counterintuitive. However, this result may reflect the fact that males exhibiting non-standard employment

relationships reflect a particularly selected group in the labor market. Considering age, we observe that workers below the age of 26 years face the highest risk of entering the sample in a low paid job, whereas we see a u-shaped pattern with respect to the retention probability. Young workers (20 - 25 years) and older workers above 45 years face a lower retention probability than the middle aged.

All in all, the signs of the variables vary little for female workers in comparison to men. Exceptions are skill and occupational status. Other than for men, being high-skilled (versus low-skilled) and a blue-collar (versus a white-collar) worker reduces the retention probability. The estimation results demonstrate that especially females between 26 and 30 years have a lower probability of staying in the sample five years later compared to the young female workers below 26 years. This result might be explained by the fact that women leave the labor market more frequently during that period, for example due to maternity leave. With respect to the initial condition process, a main difference between men and women concerns the previous employment status. In line with what one would expect, a previous full-time employed position now decreases the probability of being initially low paid.

Turning next to the transition equation, the coefficients of all covariates are allowed to differ - in line with our switching regression specification - across those who were initially low paid ($L_t \mid L_{t-5} = 1$) and for those initially high paid ($L_t \mid L_{t-5} = 0$). As one might expect, for most variables the signs of the estimated coefficient point into the same direction as in the initial conditions equation, particularly for those initially high paid. With respect to age, however, we see deviations conditional on the initial low-pay status for male as well as for female workers. Conditional on being initially high paid, younger individuals up to an age of 35 years have a higher probability than those older than 35 years to enter the low-wage sector five years later. In contrast, conditional on being initially low paid, the group of the youngest workers exhibits the lowest probability of sticking to a low-wage job. Thus, once earning a low wage it is much more difficult for older individuals to escape from it than it is for younger ones. This result is valid for both men and women and confirms the result of the descriptive statistics, see section 2.3.

Table 5: Regression Summary for women, 1984 - 1999

Variables	R_t		L_{t-5}		$L_t L_{t-5} = 0$		$L_t L_{t-5} = 1$	
	Robustness of direction	number of years in which sign switched	Robustness of direction	number of years in which sign switched	Robustness of direction	number of years in which sign switched	Robustness of direction	number of years in which sign switched
Age 26 - 30 (ref.: age 20 - 25)	-- ¹⁾	1	---	0	++	4	+++	0
Age 31 - 35	+++	0	--	2	+	3	+++	0
Age 36 - 40	+++	0	--	4	---	0	+++	0
Age 41 - 45	+++	0	---	0	--	2	+++	0
Age 46 - 50	+++	0	--	2	--	2	+++	0
Age 51 - 55	-	2	-	2	--	3	+++	0
Tenure	+++	0	---	0	--	2	+++	0
Tenure squared	--	2	++	1	++	3	---	0
Being foreign (=1)	---	0	—	0	--	4	--	2
Medium-skilled (ref. low-skilled)	+++	0	---	0	o	1	--	2
High-skilled	-	1	---	0	---	0	---	0
Previously full-time empl. (=1)	+++	0	---	0	---	0	-	4
Total non-employment duration	---	0	+++	0	+++	0	+++	0
Number of non-empl. spells	---	0	+++	0	++	2	-	6
Number of previous employers	---	0	---	0	+	1	--	5
Employment interruption (=1)	---	0	+++	0	++	4	++	1
Change of employer (=1)	--	3	+++	0	++	2	+	4
Number of unemployment spells	+	4	++	1	+++	0	++	6
Blue collar worker (=1)	---	0	+++	0	+++	0	+++	0
Total unemployment duration	--	2	---	0				

¹⁾ Measure of direction robustness:

+++/-: Variable has a positive/negative impact in all 16 years.

++/--: Variable has a positive/negative impact in 12 to 15 of the 16 years.

+/-: Variable has a positive/negative impact in 9 to 11 of the 16 years.

o: Variable has a positive and a negative impact in 8 of the 16 years.

To assess the quantitative meaning of a variable's impact, it is necessary to derive the marginal effects as described in Appendix C. Table 6 exemplarily displays the marginal effects of the explanatory variables on the low-pay transition probabilities for male workers in 1999. In line with the switching regression specification, the effects are reported separately for those who were initially low paid and for those initially highly paid. For the former group the effects are to be interpreted in terms of persistence effects, whereas for the latter group the marginal effects refer to the probability of entering low pay. Marginal

effects are to be interpreted as deviations from a reference person which has all dummies set to zero and is defined by setting the continuous covariates equal to their sample median values.⁹

Table 6: Marginal effects for male workers in 1999

	$L_t L_{t-5} = 1$		$L_t L_{t-5} = 0$	
	ME	std. error	ME	std. error
Average prediction		0.476		0.022
Baseline		0.589		0.022
Tenure	0.000	0.572	-.000	3.476
Tenure squared	-.000	0.001	0.000	0.004
Total non-employment duration	0.000	0.072	0.000	0.841
Total unemployment duration	0.000	0.092	0.000	5.608
Number of non-employment spells	-.004	19.967	0.003	754.08
Number of previous employers	0.008	39.628	0.001	188.82
Age 26 - 30	0.063***	0.000	-.009***	0.000
Age 31 - 35	0.099***	0.001	-.010***	0.000
Age 36 - 40	0.121***	0.001	-.011***	0.002
Age 41 - 45	0.172***	0.003	-.011***	0.000
Age 46 - 50	0.185***	0.003	-.008***	0.002
Age 51 - 55	0.231***	0.002	-.004	0.003
Being foreign	0.000	0.008	0.009	0.005
Medium skilled	-.049***	0.009	-.008**	0.003
High skilled	-.140***	0.011	-.021***	0.001
Previously full-time employed	0.018***	0.005	-.012**	0.004
Employment interruption (=1)	0.020***	0.005	-.001	0.003
Change of employer (=1)	0.005	0.012	0.005	0.003
Blue collar worker (=1)	0.012*	0.005	-.041*	0.016
N		14549		175771

Note: All specifications additionally include regional, sectoral, and occupational dummies. Significance levels: ***: $p < 0.001$; **: $p < 0.01$; *: $p < 0.05$.

The first two rows in Table 6 report the average transition probabilities - which represent the two components of the *ASD* derived in equation (7) - and those for the reference individual - which are referred to as the baseline probabilities. The baseline persistence

⁹A reference individual is a German between the age of 20 to 25 years, has a vocational degree, a simple blue-collar occupation and had no regular employment relationship and three previous employers prior to entry into the current establishment. The reference worker has moreover had a median number of 29 spells, worked the median tenure of 2040 days, has been non-employed for 90 days and has not yet been unemployed.

probability of 0.589 is considerably larger than the average transition probability, whereas no difference is observed for entry probabilities. The ME estimates indicate that observable individual attributes significantly affect the probability of both staying and becoming low paid. While a better education reduces the probability of both staying and becoming low paid, other variables have opposite effects on the low-wage persistence and entry probabilities. The marginal effects confirm earlier results from Table 5 which suggest that the likelihood of entering the low-wage sector is highest for the youngest workers below 26 years, whereas the likelihood of remaining low paid is the lowest for this subgroup. An individual who is between 41 and 45 years, for example, has a persistence (entry) probability which is 17.2 (1.1) percentage points higher (lower) than that of an individual aged between 20 and 25 years, *ceteris paribus*. The continuous employment history variables like tenure and total unemployment duration do not have any explanatory power.

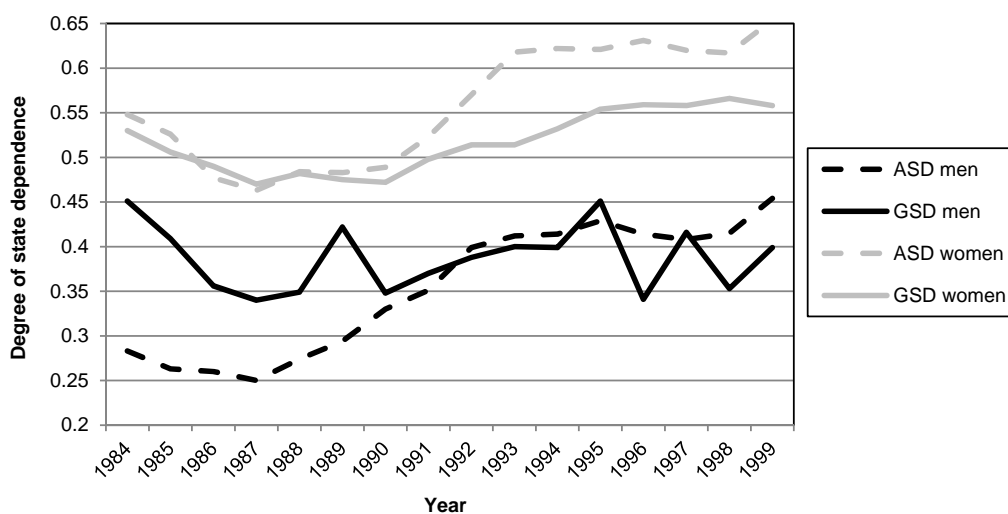
Overall, the estimation results indicate for both males and females that age rather than employment history variables drive the persistence probability. For example, whether an individual has a long or a short record of days in unemployment is not of major importance once an individual earns a low wage. This suggests that the extent of true state dependence is quite substantial.

3.3.3 The evolution of genuine state dependence (*GSD*)

After having estimated the transition equations, we next turn to the evolution of genuine state dependence (*GSD*) as given by eq. (8). In Figure 4, we plot the estimated aggregate state dependence (*ASD*) and genuine state dependence (*GSD*) against time, separately for men and women. The estimated *ASD* values are nearly equal to the descriptive values that we showed earlier in Figure 3. Comparing the evolution of *GSD* and *ASD*, Figure 4 demonstrates that both measures are characterized by a quite divergent development. While the male *ASD* has steadily risen over time from 25 per cent in 1987 to 45 per cent in 2002, the *GSD* measure exhibits a stationary pattern by fluctuating at a rate around 40 per cent over the whole observation period. Thus, once observable characteristics are controlled for, these findings strongly argue against an upward trend in genuine state dependence for male workers. For women, in contrast, there seems to have been a slight

increase in *GSD* between 1990 and 1999 from 47 to 56 per cent. However, the growth of *GSD* has been distinctly lower than that of *ASD* which rose from 46 to 65 per cent between 1987 and 2002. This result highlights the importance of accounting for changes in the composition of the low-wage labor work force and makes clear that the omission of such changing trends might lead to wrong inferences.

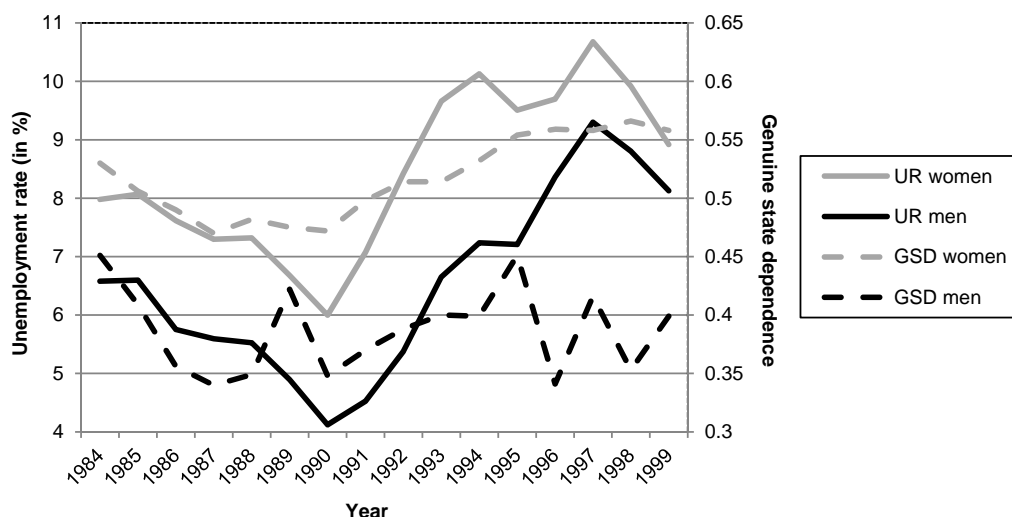
Figure 4: Evolution of the aggregate (*ASD*) and genuine state dependence (*GSD*)



Especially for male workers, we observe that the extent of *GSD* has been substantially higher than *ASD* during the first part of the observation period until 1991. This contrasts with earlier results from other studies (e.g., Cappellari and Jenkins 2004 and Cappellari 2007). Recall, however, that in our *GSD* measure we contrast for each individual - given his or her observed characteristics - the probability of entering the low-wage sector conditional on being initially high paid with the respective probability conditional on being initially low paid. One would typically expect that individuals working in the high wage sector exhibit observed attributes that shelter them from low-wage persistence, even evaluated at the counterfactual persistence probability. However, our strong results with respect to the age structure's impact on transition probabilities lead to the opposite pattern. As individuals from the high-wage sector are on average five years older than those in the low-wage sector, their counterfactual persistence probabilities are considerably higher than their respective entry probabilities. This gives rise to a large *GSD* value,

which even exceeds the (observed) *ASD* especially in the first years of our observation period when the difference in the average age between low and high wage earners was distinctly larger than in later years.

Figure 5: Evolution of the genuine state dependence and the unemployment rate, 1984 - 1999



A closer look at the evolution of *GSD* reveals that the increase during the 1990's for both men and women was preceded by a decline in the mid 1980's. We thus observe some cyclical pattern which differs across men and women only at the end of the observation period. One possible explanation for this pattern could relate to the evolution of the unemployment rate. As e.g. Omori (1997) points out for non-employed individuals with help of US data, the stigmatization of being non-employed is more severe when the local unemployment rate is rather low.¹⁰ Thus, translated to low paid individuals, they might not be afflicted of sticking to a low paid job in times of high unemployment since they at least have a job. This could then correspond to a higher *GSD* than in times with low unemployment rates when the chances of escaping the low-wage sector might be bigger.

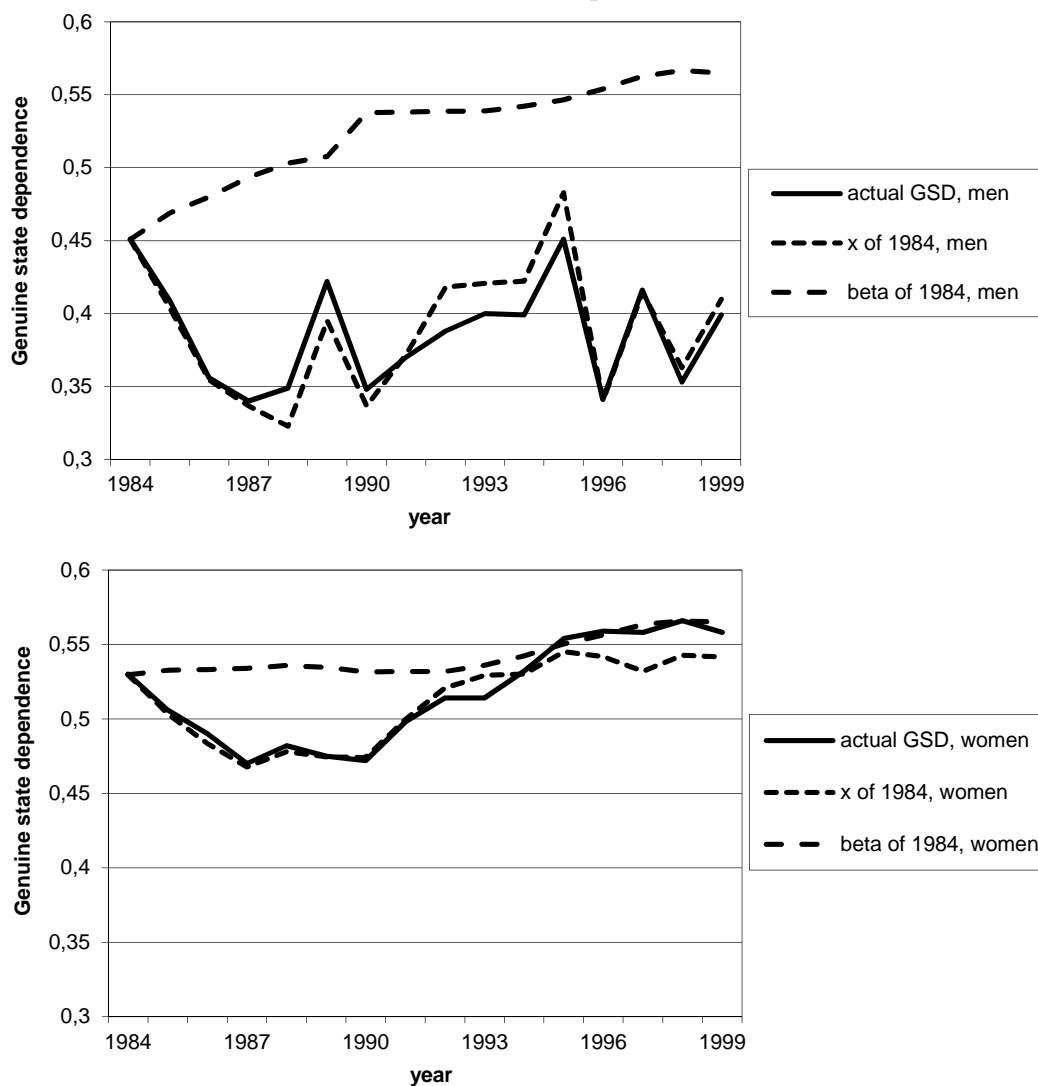
¹⁰Also Lockwood (1991) emphasizes in his matching model in which firms imperfectly test workers prior to hiring them that employers will tolerate longer search durations when general unemployment is high and that they will be stricter when general unemployment is low. Evidence on the relationship between unemployment, state dependence and the business cycle with German household data (GSOEP) is given by Biewen and Steffes (2008). They find that unemployed individuals, who face difficulties finding a new job because employers interpret unemployment as a negative signal, do so especially when it is easier to find jobs, i.e. when unemployment is low.

Although we are not able to make any causal inferences on the relationship between the extent of *GSD* and the unemployment rate, Figure 5 provides some support for the similarity of the evolution of both factors. Especially for women, there is evidence that the evolution of the unemployment rate in year t matches the evolution of *GSD* (measuring the probability of sticking to the low-wage sector in $t + 5$) with some time lag. The decrease in *GSD* from 53 to 47 per cent between 1984 and 1991 was accompanied by a decrease in the female unemployment rate which fell from 8 to 6 per cent during the same period. After 1991, the unemployment rate as well as *GSD* rose distinctly. The positive association between the *GSD* and the first lead of the unemployment rate is confirmed by a high correlation coefficient of 0.53 which is significant at the 5% level ($p=0.036$). A weaker relationship is observed for men, for whom the correlation coefficient is positive (0.21), albeit clearly lower than that for women, and not significant ($p=0.428$). Although we observe patterns for men that are similar to those for women - especially the distinct increase in *GSD* and the unemployment rate at the beginning of the 1990's - the lower correlation can mainly be attributed to the unpronounced pattern of the evolution of *GSD* especially in the last years of the observation period. The results therefore indicate that the unemployment rate fosters an increase in the probability of sticking to a low-wage job especially for women.

At this point it is important to note that the evolution of the *GSD* measure still reflects both changes in the workforce composition as well as changes in the differential impacts of the observed covariates on the transition probabilities. To disentangle these effects, Figure 6 for men and Figure 6 for women show how the *GSD* measure would have evolved, had the composition of the workforce remained constant at its 1984 level. For both sexes, we observe that the counterfactual evolution is similar to the observed development of *GSD*. This means that the composition of the workforce has remained quite stable throughout the observation period and had no profound impact of the evolution of *GSD*. This implies that the evolution is rather driven by changes in price effects, as measured by the differences between the coefficients, γ_1 and γ_2 . Supposed the price effects had remained at their 1984 level, i.e. keeping the coefficients of the starting year constant over time, we observe from the upper part of Figure ?? a continuous positive drift of *GSD* for male

workers. Between 1984 and 1999, *GSD* would thus have risen from 45 to 56 per cent for males (instead of fluctuating around a rate of 40 per cent). On the contrary, keeping the coefficients of the starting year constant for females, the lower part of Figure ?? illustrates that the counterfactual *GSD* approaches the observed *GSD* level at the end of the observation period. This reflects that female low-wage workers indeed witnessed an increase in genuine state dependence.

Figure 6: The *GSD* and counterfactual developments, men and women, 1984 - 1999



4 Conclusions

The purpose of the present paper was to study how individual wage mobility in the low-wage sector has changed over the last two decades of the twentieth century in Western Germany. Using a large administrative data set, we first document that the low-wage sector has increased since the mid 1980's by around 14%. The overall growth of the low-wage sector was accompanied by a distinct rise in the probability of sticking to a low-wage job for both men and women. However, the extent of persistence, as measured by the extent of aggregate state dependence (*ASD*), varies greatly across different groups of individuals. Younger workers below 26 years, for example, face a much smaller risk of sticking to a low-wage job than the oldest age group (50-55 years). As the share of young workers among the low paid has decreased to a much larger extent than the corresponding fraction in the high-wage sector, this compositional shift might have contributed to the observed rise in *ASD* over time.

In order to explore whether the observed decline in low-wage mobility is accounted for by compositional shifts or by an increase in "true" low-wage persistence, our analysis primarily seeks to infer conclusions about the evolution of "genuine state dependence" (*GSD*). Genuine state dependence arises when low-wage employment today causes low-wage employment in the future for reasons of stigmatization or human capital depreciation. We derive a measure of genuine state dependence by contrasting each individual's transition probability of staying in the low-wage sector conditional on being initially low paid with the probability of entering the low-wage sector conditional on being initially high paid. The aggregate *GSD* is computed by averaging these differences over the sample of workers with earnings retention. Hence, to obtain such predicted probabilities, we need to model low-pay transitions depending on a variety of low and high-wage workers' observable characteristics. In order to address the initial conditions problem and the endogeneity of earnings attrition, we estimate a series of annual trivariate probit models that account for the selection into low-wage employment and earnings retention.

Based upon the estimated transitions, our results show that between 1984 and 1999 male workers' *GSD* - opposed to their increasing *ASD* - exhibits a quite stationary

development at a rate of about 40 per cent with some fluctuations over time. For women, there seems to have been a slight increase in *GSD* during the 1990's from 47 to 56 per cent, which is, however, clearly less pronounced than the rise in their *ASD*. The observed fluctuation of *GSD* over time, mainly a decline in the late 1980s followed by an increase in the early 1990s, mirrors the evolution of the unemployment rate. For women, we find a significant positive correlation between the aggregate *GSD* and the first lead of the unemployment rate, suggesting that female low-wage workers are more likely to escape the low-wage sector in times of low unemployment.

As the aggregate *GSD* measure reflects both changes in the workforce composition as well as changes in the differential impacts of the observed covariates on the transition probabilities, we attempt to further disentangle these effects. Our results document that the counterfactual evolution of *GSD* would have been quite close to the actual *GSD* development, had the composition of the low and high-wage sector remained constant at its 1984 level. This holds for men as well as for women. What has changed over time instead, is the differential impact of the observed covariates. Supposed the impact of the observed covariates had remained at its 1984 level, especially male workers would have experienced a continuous increase in their *GSD* from 45 to 56%.

What do these results tell us about the ongoing discussion of decreasing wage mobility in the low-wage sector in Germany? Taken together, our analysis highlights the importance of accounting for possible compositional changes in the low-wage population and makes clear that the omission of such changing trends might lead to wrong inferences about the development of true low-wage persistence. In disentangling compositional shifts from changes in price effects, our findings show that - contrary to common perceptions - especially male workers' decline in low-wage mobility cannot be explained by an increase in "true" state dependence, but rather by compositional changes. Our results therefore lend strong support to the notion that appropriate policy interventions should primarily aim at working against these compositional shifts, by, e.g., improving low-wage earners' skills and intensifying older low paid employees' vocational training opportunities.

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Appendix

Appendix A: Figures

Figure A1: Aggregate state dependence by age group, 1984 - 1999

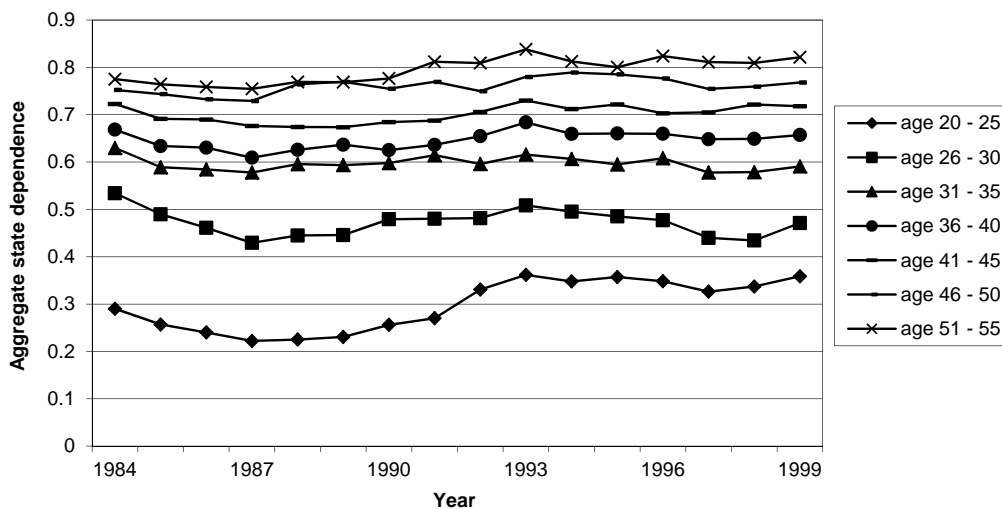


Figure A2: Aggregate state dependence by tenure group, 1984 - 1999

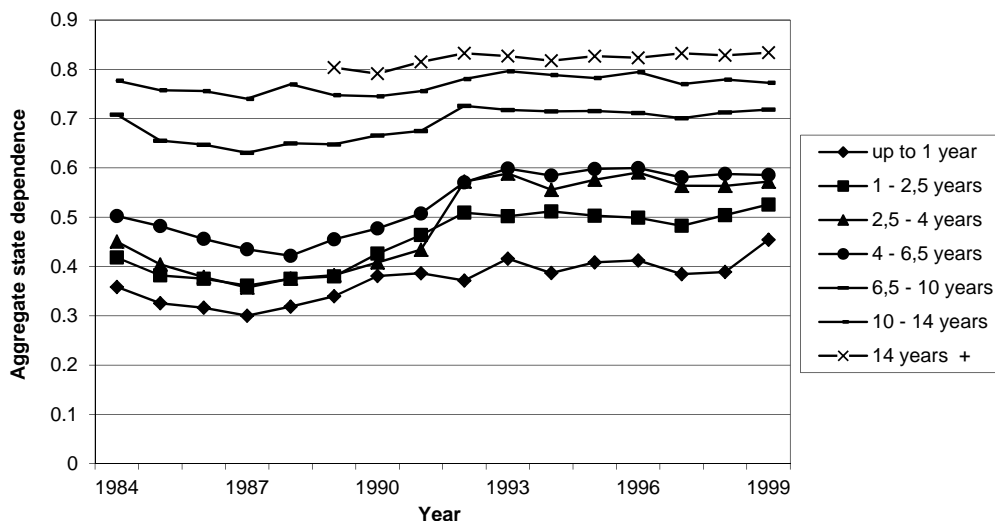


Figure A3: Mean share of low-wage earners by age group and sex, 1984 - 2002

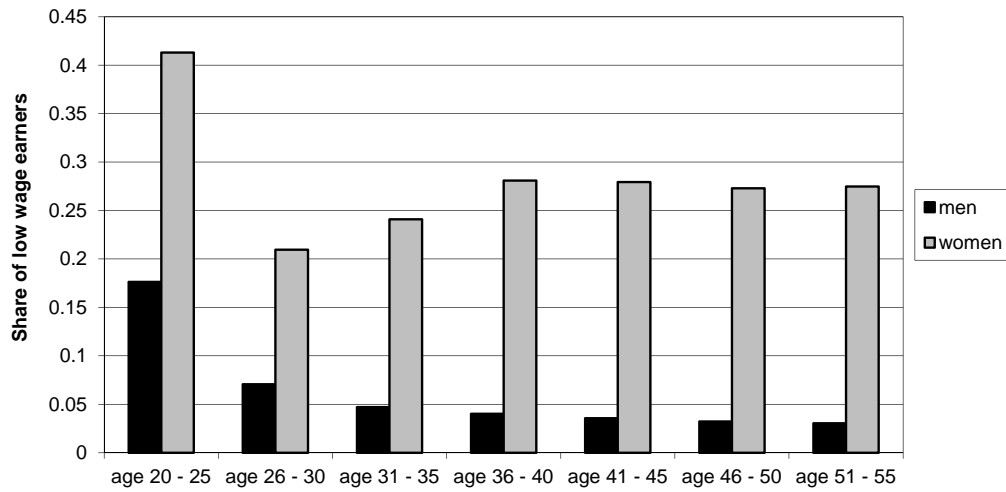


Figure A4: Mean share of low-wage earners by tenure group and sex, 1984 - 2002

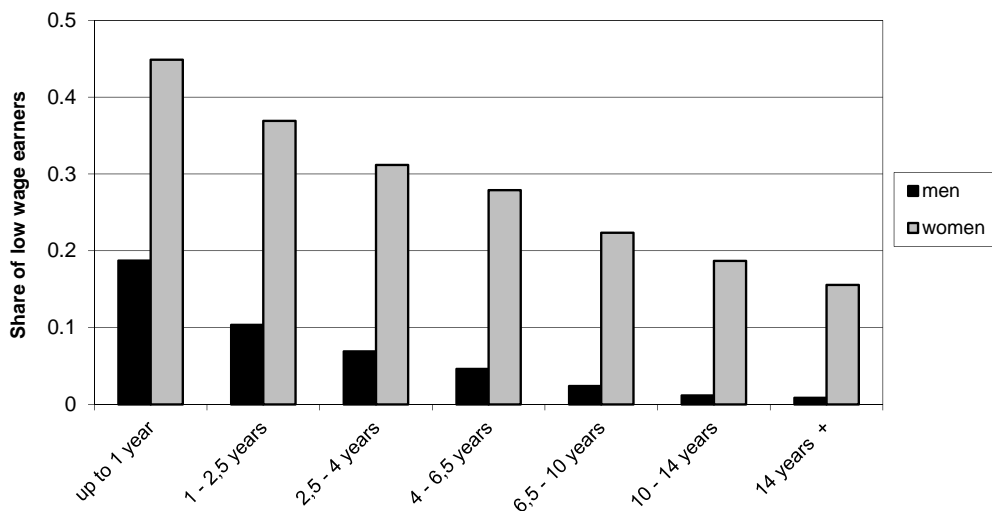


Figure A5: Aggregate state dependence by degree and sex, total population, 1984 - 1999

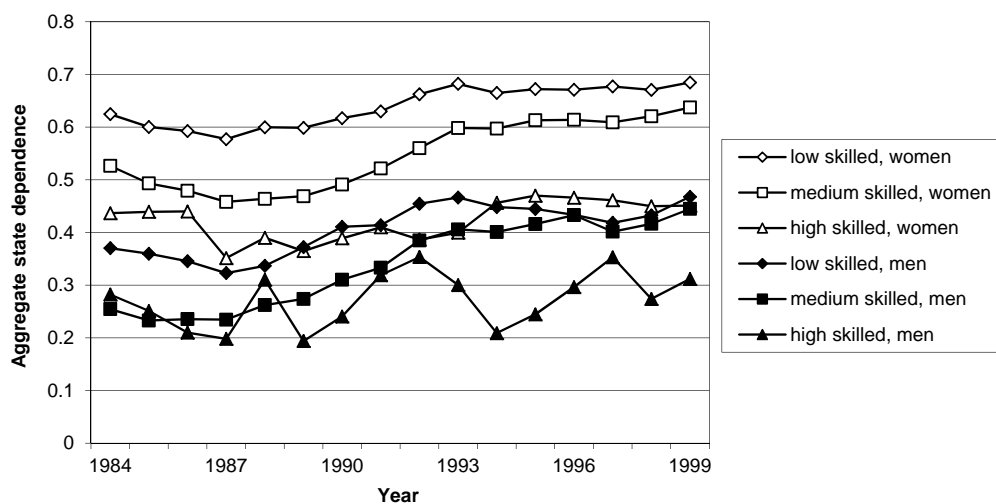


Figure A6: Mean share of low-wage earners by degree and sex, 1984 - 2002

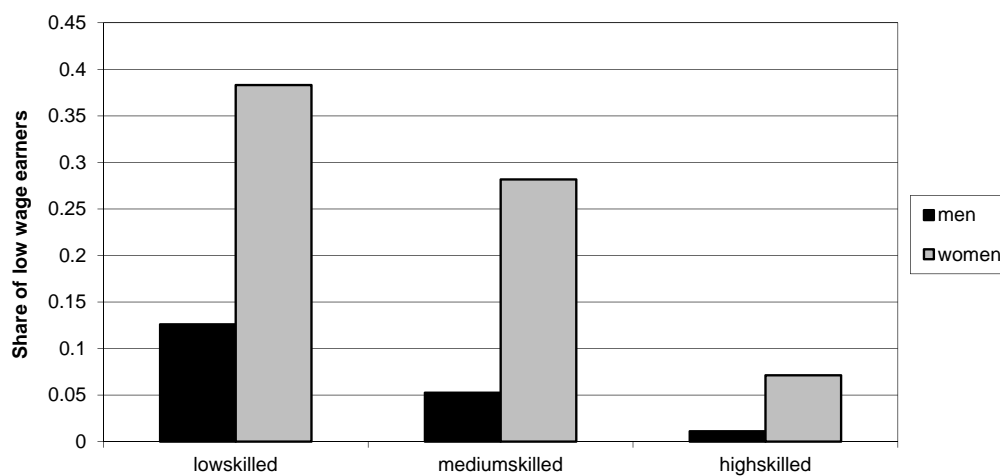


Figure A7: Aggregate state dependence by nationality and sex, total population, 1984 - 1999

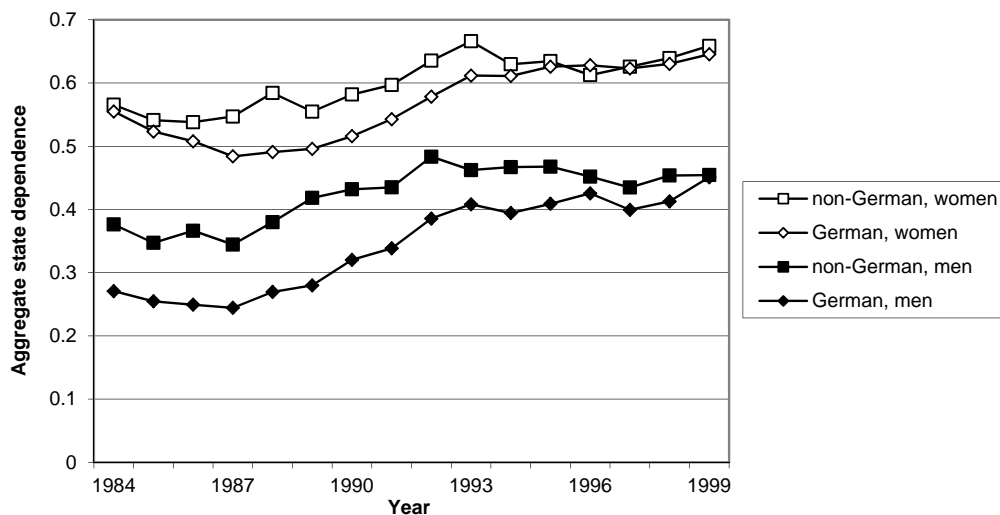


Figure A8: Aggregate state dependence by occupation and sex, total population, 1984 - 1999

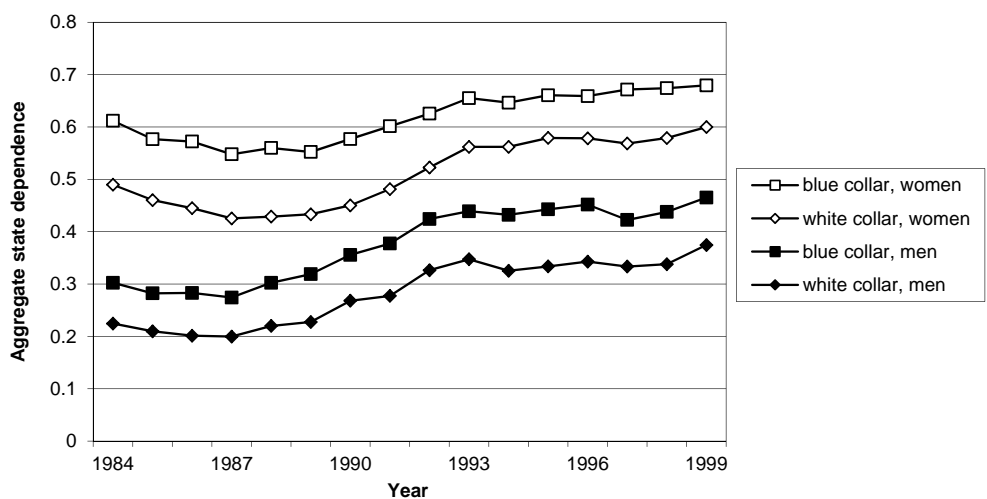


Figure A9: Evolution of the age composition in the low-wage sector 1984 - 2002, men

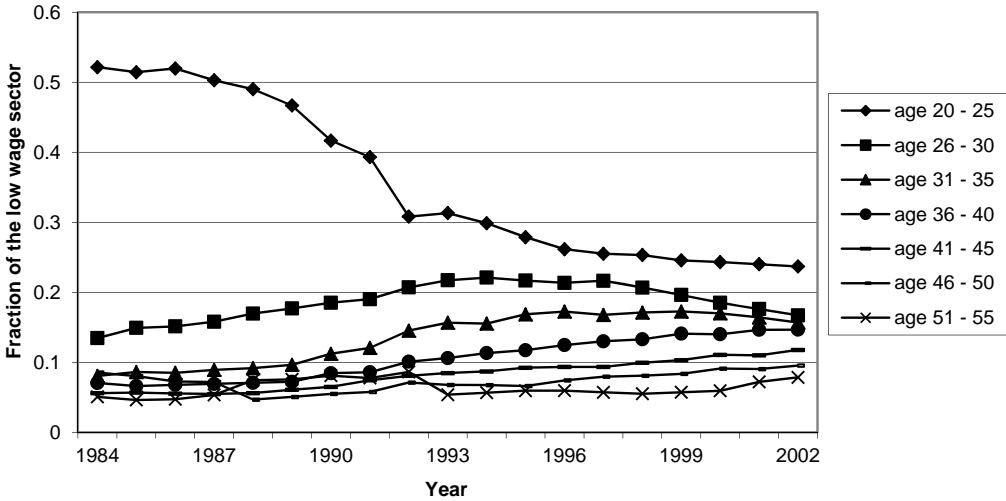


Figure A10: Evolution of the age composition in the low-wage sector 1984 - 2002, women

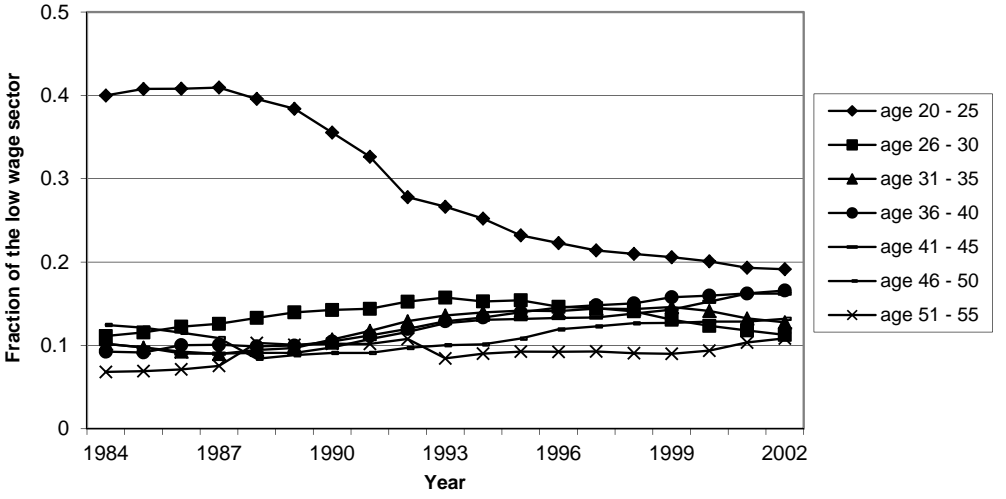


Figure A11: Evolution of the composition of sex in the low and high wage sector, 1984 - 2002

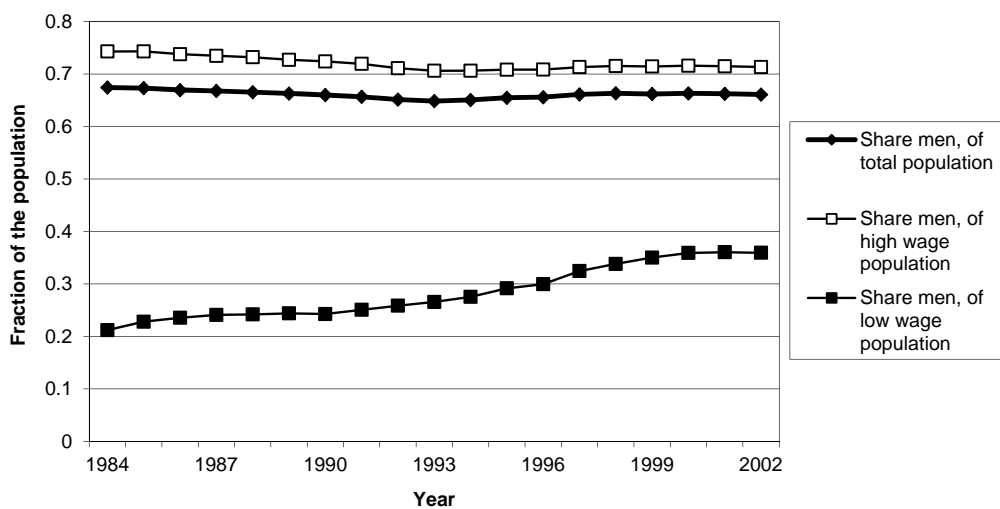


Figure A12: Evolution of the skill composition in the low and high wage sector, 1984 - 2002, women

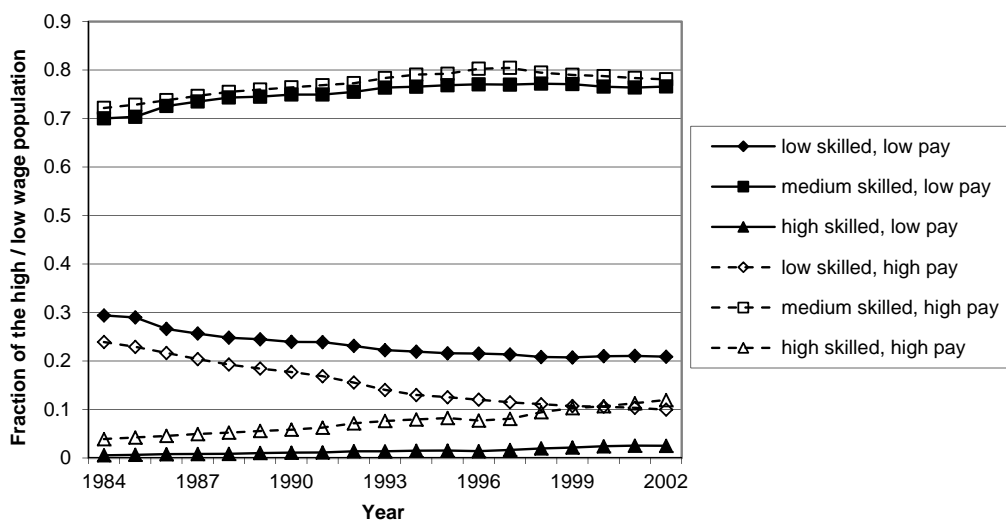


Figure A13: Evolution of the skill composition in the low and high wage sector, 1984 - 2002, men

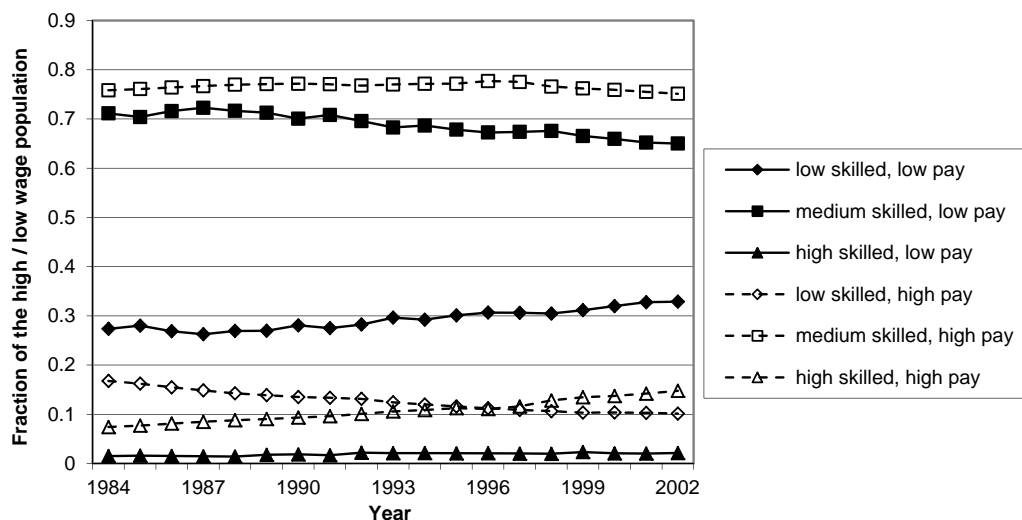


Figure A14: Evolution of the low and high wage sector by collar and sex, 1984 - 2002

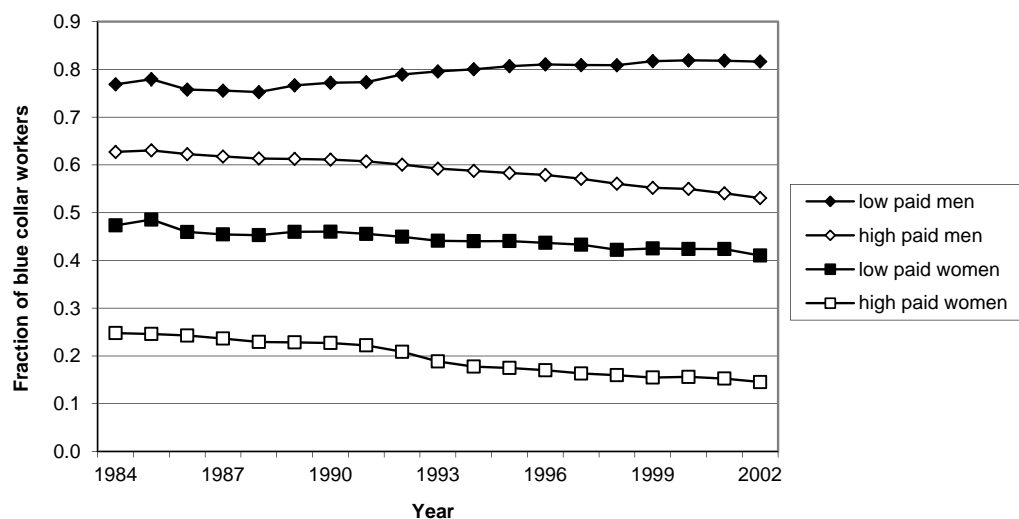
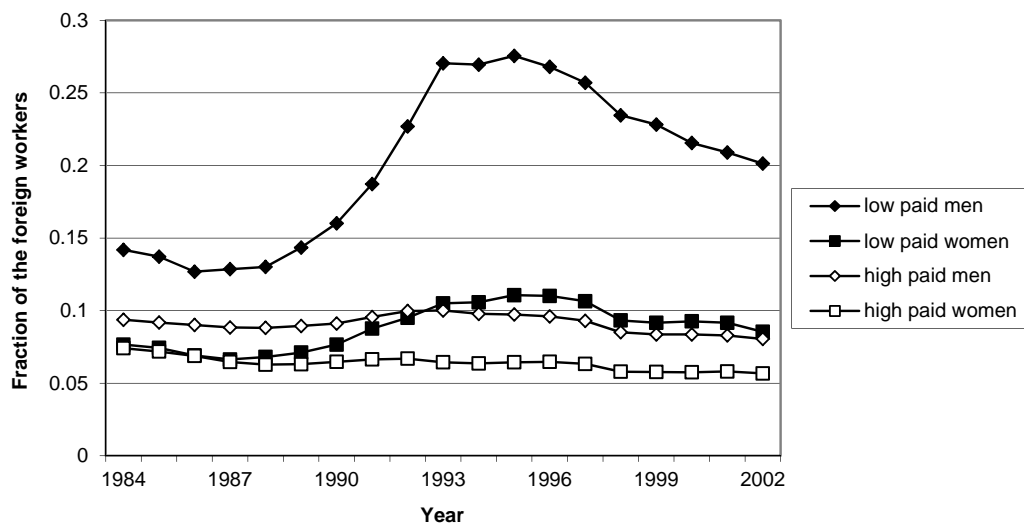


Figure A15: Evolution of the low and high wage sector by nationality and sex, 1984 - 2002



Appendix B: Tables

Table B1: Description of the variables used in the analysis

Variable	Definition
Age	Age (20-55 years) categorized in seven sub-groups
Tenure	Sum of all previous days of employment
Tenure squared	Square of sum of all previous days of employment
Nationality	Dummy=0 if German
Medium skilled ¹⁾	Dummy=1 if completed vocational training but no university degree
High skilled	Dummy=1 if university degree
Previous occupation status	Dummy=1 if previously employed, Dummy=0 if previously un-/non- or part-time employed or in vocational training
Total unemployment duration ²⁾	Sum of all previous days of unemployment
Total non-employment duration	Sum of all previous days of nonemployment
Number of unemployment spells	Sum of all previous unemployment spells
Number of non-employment spells	Sum of all previous nonemployment spells
Number of previous employers	Sum of all previous employers
Previous employment interruption	Dummy=1 if recall from current employer
Number of empl. interruptions	Sum of all previous recalls
Change of employer	Dummy=1 if employer or employment status changes
Profession	Dummy=1 (=0) if blue-collar (white-collar) worker
6 occupation dummies	Either agrarian, salary, sale, clerical, service or production worker
16 sector dummies	Two-digit sectors (for categorization see Drews, 2008, pp. 67)
11 regional dummies	The 10 Western German states plus West-Berlin (until 1990)
Low wage in t (1989-2004)	Dummy=1 if gross daily wage in t < 2/3 of the median wage ³⁾
Low wage in t-5 (1984-1999)	Dummy=1 if gross daily wage in t-5 < 2/3 of the median wage
Retention in t	Dummy=1 if employment status is observed in t

¹⁾ To improve the education variable we use the imputation rules derived by Fitzenberger et al. (2006).

²⁾ Following the procedure proposed by Lee and Wilke (2009), involuntary unemployment is defined as comprising all continuous periods of transfer receipt. Gaps between periods of transfer receipt or gaps between transfer receipt and a new employment spell may not exceed four weeks, otherwise these periods are considered as non-employment spells (involving voluntary unemployment or an exit out of the social security labour force). Similarly, gaps between periods of employment and transfer receipt are treated as involuntary unemployment as long as the gap does not exceed six weeks, otherwise the gap is treated as non-employment.

³⁾ Due to the introduction of the Euro in 1999, all wages before 1999 are transformed from Deutschmark into Euros at a rate of 1€ = 1.95583 Deutschmark. Since the wage variable delivers unrealistic daily wages at the lower end of the wage distribution, we exclude all observations with earnings of less than 16€ per day (in prices of 1995).

Table B2: Mean values of selected characteristics by wage sector

	High wage	Low wage
Females (in %)	28	72
Age (in years)	37.5	33.8
Low skilled (in %)	13	25
High skilled (in %)	10	2
German (in %)	92	88
Previous unemployment duration (in days)	101	163
Previous non-employment duration (in days)	368	727
Tenure in current job (in days)	2797	1571
Retention rate five years later (in %)	66	47
N	4866868	766533

Table B3: Share of low-wage earners in % among the total low-wage population

	males		females	
	1984	2002	1984	2002
High skilled	2	2	1	3
Low skilled	27	33	29	21
Age 20 to 55	52	24	40	19
Age 51 to 55	5	8	7	11
Blue collar	77	82	47	41
Full-time in prior job	16	25	20	25
German	86	80	92	91

Appendix C: Derivation of the marginal effect

In order to be able to interpret the estimation results, we report the marginal effects (ME) showing the impact on the relevant probabilities of a change in the chosen covariate. For a dummy variable, the ME is calculated as a change in the probability resulting from a change in the indicators value from zero to one, holding all other covariates fixed at their sample median values. ME for continuous variables are usually estimated by evaluating the partial derivative, which is equal to the corresponding coefficient multiplied by an evaluation of the normal density function. However, here the computation is not straightforward because the transition probabilities are conditional in nature (e.g. the probability of low pay in t conditional of being low paid in $t - 5$). To clarify this point, the conditional probabilities are given by:

$$e_{it} \equiv \left[\Pr(L_{it} = 1 | L_{it-5} = 1) \right] = \left[\Phi_2 \left(z_{it-5} \hat{\gamma}_1, x_{it-5} \hat{\beta}; \rho_2 \right) \Phi \left(x_{it-5} \hat{\beta} \right) \right] \quad (10)$$

and

$$f_{it} \equiv \left[\Pr(L_{it} = 1 | L_{it-5} = 0) \right] = \Phi_2 \left(z_{it-5} \hat{\gamma}_2, -x_{it-5} \hat{\beta}; -\rho_2 \right) \Phi \left(-x_{it-5} \hat{\beta} \right) \quad (11)$$

As is evident from eqs. 10 and 11, a change in the value of a covariate may affect both the numerator and denominator of the conditional probabilities. In order to deal with this issue, we adopt the procedure suggested by Stewart and Swaffield (1999) (see also Cappellari (2007) and Cappellari and Jenkins (2008)) by keeping the elements of x_{it-5} fixed. To do so, we first predict the low-pay probability in $t - 5$ for all low paid individuals and take the average over these values - denoted as q . By inserting $w = \Phi^{-1}(q)$ into eq. 10, we obtain $\Phi_2(z_{it-5} \hat{\gamma}_1, w; \rho_2) / w$. This expression is used to calculate ME as deviations between the conditional probabilities for a reference person and hypothetical probabilities induced by changing each covariate by an unit. For the reference person, we set continuous covariates to the sample median values and dummy variables to zero. The same procedure is applied to f_{it} .