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Changing fortunes during economic transition: Low-wage persistence before and after German unification

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Changing Fortunes During Economic Transition - Low-Wage Persistence before and after German Unification

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Abstract

This paper studies whether the transition from a centrally planned to a market economy offers new perspectives for those who, in economic terms, were relatively deprived under the old regime. Previous empirical research on this question has been limited by the availability of representative longitudinal micro-data that track individuals' labour market careers across different political regimes. Our study seeks to fill this research gap by looking at the transition of Eastern Germany following German Unification. Using a unique large-scale German administrative data set, we measure individuals' relative economic position by exploiting information on whether individuals were in the bottom of the pre-unification wage distribution. We then address the question of how workers' low or high-wage status determines their wage and labour market status within and across different regimes. We document strong evidence of a dynamic selection process into low-wage employment after Unification. Furthermore, consistent with a weak connection between individuals' true productivity and their pre-unification low-wage status, the extent of across regime state dependence is found to be small in magnitude and appears to vanish over time. For males, the small extent of across regime dependence is found to be most pronounced among the medium and high-skilled, suggesting the depreciation of general human capital as a potential explanation.

JEL-Code: J31, J64, P21, C33, C35

Keywords: Low Pay Dynamics, Economic Transition, State Dependence

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1 Introduction

Does the transition from a centrally planned to a market economy offer new perspectives for those who, in economic terms, were relatively deprived under the old regime? Or does the economic transformation bring along an equally unfavourable position under the new system, thereby reflecting some extent of economic state dependence even across completely different political and economic regimes? This question is of considerable relevance against the background of Eastern and Central Europe's recent history of transitions from centrally planned to market economies and is crucial to understand the welfare implications of economic transitions. The issue is also important from a quantitative point of view, as during a large part of the 20th century about one-third of the world's population lived under communist regimes.

In this paper we address these issues in the context of Eastern Germany. The former German Democratic Republic (GDR) is a particularly interesting case because it has experienced a unique pathway of political and economic transformation after 1989 through its Unification with the Federal Republic of Germany (FRG). Compared to other transition countries, the transformation process took place much more quickly, as the political and economic system of Western Germany was immediately transferred to Eastern Germany. Although a large body of research has documented the evolution of the Eastern German labour market after Unification, there is barely any evidence on how one's pre-unification relative economic position determined one's fortunes after Unification. Previous empirical research on this question has been limited by the availability of suitable representative longitudinal micro-data being able to track individuals' labour market careers across different political and economic regimes. Our study seeks to fill this research gap by exploiting a unique large-scale administrative data set (*BASiD*) from the German Pension Register and the German Federal Employment Agency. The *BASiD* data provide an ideal basis for our empirical analysis as it allows us, first, to identify individuals living in Eastern Germany before Unification and, second, it enables us to track

individual employment histories both before and after the fall of the Iron Curtain. To our knowledge, no other study has used administrative data to compare labour market outcomes across different political and economic regimes.

To measure individuals' relative economic positions, we will exploit information on the incidence and duration of individuals' position in the bottom part of the pre-unification wage distribution. Even though earnings inequality in the GDR was considerably lower than in the FRG, one may still identify workers who fared considerably worse than the average. For instance, in 1989 the poorest ten per cent of the working age population earned less than 50 per cent of the overall average of monthly earnings.¹ Having identified individuals' low-wage status prior to Unification, we then proceed to ask as to how workers' low or high-wage status determines their economic fortunes after the transition. In particular, we are interested in the extent of low-wage state dependence across economic regimes by distinguishing persistence in low pay due to observed and unobserved heterogeneity from true state dependence, also referred to as genuine state dependence.

The literature on state dependence acknowledges three main explanations for genuine or true state dependence of low pay (see e.g. Stewart, 2007, Arulampalam et al., 2000). The first one refers to signalling low productivity levels as employers do not observe the true productivity and use the relative wage position as an additional signal for the purpose of evaluation. As to the signalling argument, the heavily regulated pre-unification labour market should have rendered the selection into low-wage jobs based on workers' true productivity very unlikely in the GDR. Given that workers had their wages set according to a centrally determined wage grid, Bird et al. (1994) suspect that "because the human capital model was obviously not relevant under socialism, the observed correlation between wage income and schooling and experience in the GDR must simply exist because the socialist wage regime took factors like these into account in assigning wages". If this was indeed the case, signalling low productivity should not have contributed to genuine low-

¹Own calculations from the 1990 German Socio-Economic Panel retrospective GDR survey.

wage persistence for the pre-unification period. Given the different economic models before and after Unification, the signalling argument would also not serve as an explanation for low-pay persistence across the different regimes.

A further explanation for low-pay persistence may be stigmatization as long as low-wage jobs are systematically associated with characteristics, such as social class affiliation, against which employers might discriminate. In the GDR it is reasonable to think of low-wage jobs as having - at least to some extent - resulted from political discrimination. This in turn might have favoured true persistence due to signalling political opposition. Even though the nature of stigma is likely to considerably differ across the different regimes, such an effect could be present and lead to persistence within each regime. Note, however, that stigmatization is unlikely to contribute to "across regime" genuine state dependence.

A final explanation for low-pay state dependence that may be relevant even across the regimes relates to human capital depreciation. Given that selection into a pre-unification low-wage job should have been unrelated to workers' true productivity, the latter might still have been depreciated due to unfavourable working or job conditions inherent to low-wage jobs. Whether this led to true state dependence largely depends on the extent to which depreciation has affected workers' job-specific or general human capital. Given that general skills have been shown to be transferable to the post-unification labour market (Fuchs-Schündeln and Izem, 2012), a loss in general human capital caused by pre-unification low-wage jobs might contribute to true low-wage persistence even during economic transition.

In exploring the importance of workers' pre-unification wage positions for their post-unification wage outcomes, our analysis contributes to the literature on labour markets of transition economies. A large body of research has examined how returns to human capital have changed during economic transition. The general picture that emerges is that returns to education generally increased, whereas returns to work experience did not change or even decreased during the transition process (see e.g. Rutkowski, 1996, Brainerd, 1998, Munich et al., 2005). For Germany, Bird et al.

(1994) and Krueger and Pischke (1995) show that there was little change in returns to education after Unification. Moreover, these studies document very small returns to age and seniority prior to Unification, which - similar to what has been found for other transition economies - declined during the transition process (see also Gathmann, 2004). Orłowski and Riphahn (2009) show that returns to experience and seniority in Eastern Germany were small compared with Western Germany, even 20 years after Unification. While these results indicate that specific human capital gained during socialist work experience became obsolete in the post-unification labour market, very little is known about the consequences of experience accumulated in low-wage jobs. In addressing the relevance of individuals' pre-unification wage positions for their post-unification outcomes, our analysis thus provides new insights into whether low-wage jobs in a centrally planned economy have also been associated with the depreciation of *general* human capital, thereby leading to adverse long-term consequences for post-unification labour market outcomes.

The remainder of the paper is structured as follows. Section 2 provides institutional background information on the Eastern German labour market prior to and after Unification. Section 3 provides a description of the data set and the sample selection. Section 4 explores the evolution of wages and provides a description of labour market transitions as well as unconditional probability estimates. Section 5 lays out the econometric strategy and presents the estimation results on short and long-run dynamics. Section 6 concludes.

2 Institutional Background

2.1 The Eastern German Labour Market prior to Unification

Following the Soviet example, the GDR introduced tight central economic planning along with subordination of firms to the state administration. Moreover, all citizens of the GDR had the constitutional right and duty to work (where the 'right' included

an unlimited guarantee of employment and the ‘duty’ brought along the threat to be sentenced for antisocial behaviour if one was suspected of remaining voluntarily jobless). The Eastern German labour market before Unification thus was heavily regulated: controlling the supply of and demand for labour was seen as an instrument for efficient use of resources and economic growth (see Grünert, 1997*a* for a detailed overview).

Although enterprises were effectively controlled through centralised redistribution of investment capital, salary funds, and other financial means, they were fairly free in planning and using the labour forces they had at their disposal. Under given general institutional constraints, enterprises could influence, for example, employment policies, regulations pertaining to job transfers, salary ranges, and promotion regimes. At the same time, individuals were – in principle – free to choose their workplace. Once employed, they agreed upon an individual labour contract with their firm, which included far-reaching employment rights (such as the right on employment appropriate to acquired skills or the right to be paid according to the quantity and quality of the work done).

There were clear limits to employer and occupational mobility, though. By the late 1970s, careers had become heavily affected by an increasing influence of the “state-governed labour force allocation”, a system that restricted younger cohorts in the choice of occupational training and their subsequent job (Huinink and Solga, 1994). Since the 1960s, quotas were set for occupations into which individuals were allocated after leaving secondary education. Since the late 1960s, the opportunity to study at a university was strongly restricted through quotas to high-school and university admissions. This implied that many young people could not get the occupation they actually wanted. Very often changing one’s work was then only possible within one’s occupational career via adult education (also see Zühlke and Goedicke, 2000). A restriction on employer mobility was a general tendency among GDR enterprises to keep the fluctuation of their labour force low and to maintain a high level of permanent staff (Stammbelegschaft; see Grünert, 1997*b*, Section 1).

In addition to offering firm-specific fringe benefits like free childcare, holiday arrangements, etc., an important mean of achieving this were bonuses such as ‘loyalty premiums’ (Treueprämie) for long-term employees. Enterprises had more discretion over bonuses than over base wages, where bonuses have been estimated to account for, on average, six per cent of compensation in the GDR (Krueger and Pischke, 1995). Base wages were determined by state-regulated wage grids based on observables (see Stephan and Wiedemann, 1990 for a more detailed account of the wage structure in the GDR). As a consequence, wage dispersion was much lower than in the FRG. While the empirical evidence on the GDR wage structure has established positive returns to education of 4.5 to 7.7 log points for one year of schooling, age-earnings and seniority-earnings profiles - despite the existence of loyalty premiums - have been suggested to be much flatter than in the FRG. Using retrospective information for 1989 from the German Socio-Economic-Panel, Bird et al. (1994) and Krueger and Pischke (1995) estimate returns to experience of about one to two log points for the first year of experience (compared with about 3.4 to 4.1 log points in the FRG).

2.2 The Eastern German Labour Market after Unification

After Unification the Eastern German labour market underwent a period of dramatic structural change. Monetary union between Eastern and Western Germany took place on June 30, 1990. With monetary union, Eastern Germany adopted the legal and economic system from Western Germany, including also its labour market institutions. As a result, Western German trade unions quite rapidly succeeded in transferring the Western German system of collective bargaining to the East. While the first round of wage negotiations, which already took place during summer 1990, mainly resulted in lump-sum wage increases, the second round in winter 1990/91 stipulated wage schedules being tied to a fixed proportion of the western level (Krueger and Pischke, 1995). This gave rise to tremendous wage increases, which were particularly large within the first year following monetary union. According to Hunt (2001), monthly real wages rose on average by 20 log points between

1990 and 1991, with the lower educated benefitting to a significantly larger extent (compare also the similar figures reported by Krueger and Pischke, 1995). For the time period between 1991 and 1996, Hunt (2001) reports an annual monthly wage growth of about nine log points, yielding a cumulative average real monthly wage growth of 78 per cent over the period between 1990 and 1996. Later studies report that real wage growth in Eastern Germany has come to a halt in the mid-1990s (Franz and Steiner, 2000) and even started to decline in the first years of the 21st century (Aretz, 2013). This presumably reflects that since the mid 1990s unions increasingly lost importance, as most employers could not afford the initial wage increases. While union membership rates dropped from about 40 per cent in 1992 to about 18 per cent in 2004 (Addison et al., 2007), the proportion of employees subject to an industry-level contract declined from 56 per cent in 1996 to 41 per cent in 2004.²

At the same time, many Eastern German technologies became obsolete during the transition process. After Unification, Eastern Germany experienced massive inflows of capital and technology from Western Germany. On July 1, 1990 the 'Treuhand' as a holding company for the state-owned sector was set up with the primary purpose to sell all of its holdings. When the Treuhand closed down at the end of 1994, about 860 enterprises had been sold to foreign investors and 3,000 had been acquired through management buy-outs (Kettenacker, 2013). The privatisation process not only resulted in massive worker flows across industries, but was also visible within industries. On the intra-sectoral level, East German firms made intensive use of short-time work ("*Kurzarbeit*") as an instrument provided by the government to adjust to the economic downturn right after Unification. In 1991, more than 2 million workers in East Germany were subject to short-time work and this instrument was especially used in manufacturing (60%), agriculture (15%) and in the water, energy and mining industry (5%), whereas construction, trade, transport and communication were characterised by relatively stable employment during

²Own calculations from the IAB-Establishment Panel. Representative data on collective bargaining coverage in Eastern Germany are available only since 1996.

the first phase of the restructuring process.

The massive structural change brought about by the privatisation process has led some researchers to inquire into whether human capital accumulated during the old regime became obsolete in the post-unification labour market. Using data from the 1990 German Socio-Economic Panel retrospective GDR survey, Gathmann (2004) finds that returns to pre-unification accumulated work experience drop to zero after Unification. The author interprets her results as evidence of a full obsolescence of socialist work experience, suggesting a full depreciation of job-specific human capital. Contrary to that finding, Fuchs–Schündeln and Izem (2012) demonstrate that the low labour productivity in Eastern Germany can mainly be attributed to less favourable job characteristics rather than to individual skills.³ The authors conclude from their findings that a large part of Eastern Germans’ human capital accumulated during the socialist regime was transferable to the post-unification labour market and, therefore, should have been of general nature.

3 Data and Sample

The data used in the empirical analysis are taken from German register data (*BASiD*). The data combine information from the *German Pension Register* with various data sources from the German Federal Employment Agency. The scientific use file of the data (*BASiD-SUF*) is a stratified random 0.25% sample of all birth cohorts from 1940 to 1977, who have at least one entry in their social security records, leading to an overall sample of about 60,000 individuals. The sample has been drawn in a disproportionate manner and can be made representative using a weighting factor that is part of the data set (for a detailed description see Hochfellner et al., 2012 and Bönke, 2009).⁴

³This finding is derived from regional unemployment differences at the inner German border, based on the argument that, if mainly worker characteristics caused the low labour productivity, unemployment rates should jump up discontinuously at the former border.

⁴Note that the representativeness of the data based upon the sample weights that are provided in the data refer only to the calendar year 2007. Later on, in our analysis, we will use administrative population data to construct weights for each gender-year cell.

The data provides longitudinal information on individuals' entire pension-relevant biographies up to the year 2007. Individual work histories cover the period from the year individuals were aged 14 until the age of 67. In Germany, statutory pension insurance is mandatory for all employees in the private and public sector, with the exception of civil servants and self-employed individuals. In addition, contributions to the pension insurance are paid by the unemployment or health insurance during periods of unemployment and prolonged illness.

The *BASiD* data provide an ideal basis for analysing the impact of former GDR citizens' low-pay status on their later career outcomes for several reasons. First, it is the only German administrative data source that encompasses full employment biographies. In particular, the *Pension Register* contains information on all periods for which contributions were paid (employment, long-term illness, unemployment) as well as periods without contributions, which were still creditable for the pension insurance. The latter refers to activities for which an individual receives pension credits, such as periods of school or university attendance after the age of 16, periods of training and apprenticeship and periods of caring.

Second, the *BASiD* data is the only individual level data set that contains employment biographies of former GDR citizens before German Unification. After Unification, former GDR citizens became entitled to transfer their pension-relevant activities to the FRG pension insurance system. For this purpose, the FRG Pension Insurance recorded all periods prior to Unification which were creditable for the pension insurance (see above) as well as earnings up to the GDR social security cap. The pension data therefore allow us to track former GDR workers' entire pre- and post-unification employment histories up to the year 2007. Apart from the individual information on pension relevant activities, the *Pension Register* provides information on age and gender.

Starting from 1975 in Western and from 1992 in Eastern Germany, employment spells subject to social security contributions from the *Pension Register* can be merged with data from the German Federal Employment Agency, the *Integrated*

Labour Market Biographies and the *Establishment History Panel*. The *Integrated Labour Market Biographies* provide further time varying individual information on blue or white-collar status, occupational status, educational status (six categories) and an establishment identifier. The latter allows us to retrieve information on tenure at the current employer. Finally, the *Establishment History Panel* contains information on the establishment’s workforce composition, establishment size as well as sector affiliation. Tables B.1 and B.2 in the appendix provide more detailed descriptions on the variables gained from the *Pension Register* and *Employment Statistics Register*. For former GDR citizens the data lacks explicit information on education prior to 1992. We therefore impute the educational status by using information from the *Pension Register* on individuals’ creditable schooling and apprenticeship periods (for detailed information see Gørtzgen and Nolte, 2016).⁵

In our analysis the main outcome variable of interest is labour earnings, which can be calculated by exploiting information on monthly pension credit points gained from social security employment. Credit points derived from earnings in Eastern Germany before and after Unification are scaled-up to meet the Western pension level according to a factor stipulated in the German Social Act (SGB VI). To obtain the original credit points, Eastern credit points reported in the data have to be divided by this factor. One credit point corresponds to the average of annual earnings of all gainfully employed workers in Germany. This implies that monthly earnings can be obtained by multiplying monthly credit points with average earnings as documented in the German Social Act (SGB VI - see Table B.2). Earnings are top-coded at the social security contribution limit. Compared with the FRG, where the earnings cap increases over time, the GDR threshold remained constant at 600 Mark throughout the entire GDR period. Due to this unchanged earnings cap, the fraction of GDR workers with top-coded earnings increased substantially over time and was much larger than the corresponding fraction in the FRG. Despite the restrictive earnings information, the data is still suited to analyse low-pay transitions as the

⁵More specifically, we adopt Imputation Procedure 2 (*IMP2*), which aims to match three education categories (low, medium, high-skilled), into which the six categories in the *IEB* have been typically summarised in many empirical applications.

earnings information allows us to dichotomise the GDR earnings distribution into a low- and high-wage sector.

For our empirical analysis, we focus on the employment biographies of former GDR citizens. Given that our data covers the cohorts 1940-1977, we confine our sample to the cohorts between 1940 to 1960 and follow their employment histories starting from the year 1980 until 1999. Focussing on these cohorts enables us to track the pre- and post-unification labour market histories of individuals aged between 30 and 50 in 1990. As the employment histories of later cohorts (i.e. those born after 1960) can be observed only after 1980, the restriction to the birth cohorts 1940 to 1960 permits us to observe a reasonable amount of pre-unification labour market years for all cohorts.⁶ This is crucial for our empirical strategy, which will use information on pre-unification labour market histories as a key ingredient in explaining post-unification labour market outcomes. The *BASiD-SUF* file provides monthly information on individuals' pension credit points as well as their main labour market state in a given month. We follow the literature and use the labour market state in July of any given year.⁷ To smooth out seasonal variation in earnings, monthly labour earnings are aggregated to the year level by adding up monthly earnings in a given year and taking the average over the year.⁸ Given that our data lacks explicit information on working time, we are not able to convert monthly into hourly wages. To avoid measuring persistence in working time decisions instead of earnings, we therefore exclude those individuals who based on the information from the *Employment Statistics Register* worked part-time at least once after Unification.⁹

⁶The cohort structure of our data implies that the earliest period in which we observe insured individuals is the year 1954, when those born in 1940 were 14 years old. During the subsequent years younger cohorts successively enter the data set, which gives rise to an increasingly mixed age structure. An overview on the age-year structure of the pension sub-part is given by Bönke et al. (2010). To ensure representativeness within the selected cohorts in terms of the working-age population's age structure, we have constructed weights based upon administrative population data from the German Federal Statistical Office.

⁷In the data set July 1990 is the first month when wages are fully reported.

⁸We exclude individuals from the wage distribution if their monthly wages fall short of 150 Mark, as this is considered as unreasonably low. This causes the exclusion of 35 men and 90 women in total.

⁹The *Employment Statistics Register* is available from 1992 onwards. We do not observe full-time and part-time decisions before 1992. By excluding individuals with at least one part-

Table 1 shows the number of individuals over the whole sample period. Overall,

Table 1: Number of Individuals in the Sample in each Year 1980-1999

| Year | Number of Individuals | West Migration (#) | Retire (#) |
|------|-----------------------|--------------------|------------|
| 1980 | 4818 | | |
| 1981 | 4790 | 28 | 0 |
| 1982 | 4770 | 20 | 0 |
| 1983 | 4761 | 8 | 1 |
| 1984 | 4749 | 9 | 3 |
| 1985 | 4733 | 15 | 1 |
| 1986 | 4723 | 9 | 1 |
| 1987 | 4711 | 10 | 2 |
| 1988 | 4697 | 11 | 3 |
| 1989 | 4683 | 12 | 2 |
| 1990 | 4609 | 71 | 3 |
| 1991 | 4442 | 163 | 4 |
| 1992 | 4284 | 145 | 13 |
| 1993 | 4178 | 105 | 1 |
| 1994 | 4048 | 129 | 1 |
| 1995 | 3949 | 97 | 2 |
| 1996 | 3850 | 94 | 5 |
| 1997 | 3775 | 70 | 5 |
| 1998 | 3694 | 76 | 5 |
| 1999 | 3495 | 192 | 7 |

Source: BASiD 2007.

our sample selection yields an unbalanced panel with 4,818 individuals and 87,671 person-year observations. The main reason for panel attrition is migration from Eastern to Western Germany. While the share of migrants was rather negligible prior to Unification, the fraction of migrants increased to about 3.1 per cent in the first two years after Unification. The observed decline afterwards and the increase in the second half of the 1990s - also referred to as the second wave of migration - is consistent with what has been documented in the literature (for example Fuchs-Schündeln and Schündeln, 2009). The last column refers to early retirement as a

time spell after 1991, we rely on the assumption that people who had a preference for part-time employment after 1991 had also been likely to work part-time prior to Unification. Moreover, it is well established that employment participation rates in communist regimes are generally high also among females. The Socialist Unity Party (SED) defined not the family but rather the socialistic work collectives as the basis of society. This resulted in labour participation rates (year 1989) among females of more than 90%.

reason for panel attrition, with a relatively strong increase in 1990 and 1991 hinting to a potentially selective process.

Table 2 summarises the main variables and provides summary statistics for both time periods prior to the transition (Pre: 1980-1989) and during and after the transition (Post: 1990-1999). Given that we define low-wage workers as those from the first decile of the wage distribution, the fraction of low-wage workers is ten per cent. After economic transition, we are also able to calculate a low-wage threshold based on a fraction of the median wage. By using two-thirds of the median the table shows only a slightly higher fraction of low-wage workers. As regards to qualification, about 14 per cent in the sample did not receive any formal degree, while about two thirds are medium-skilled and obtained some sort of vocational training. As mentioned above, entry to higher levels of qualification was extremely constrained prior to Unification, resulting into a small fraction of ten per cent holding a university degree. The educational information for the remaining eight per cent is missing. Using information on occupational status from the first available year from the *Employment Statistics Register*, about 50 per cent are blue-collar and 40 to 44 per cent are white-collar workers. Differentiating the occupational status into skilled, medium-skilled and simple occupations, about 14-17 per cent of individuals belong to the first, 48-55 per cent to the second and about one third to the final category. Given the increasing fraction of un- and non-employed individuals after Unification, the evolution of experience and age can be observed to diverge after Unification. Finally, the last two rows show the number of employment interruptions and the accumulated length of employment interruptions measured in months, which both increased by construction after Unification.

Table 2: Variable Definition and Description of Basic Variables

| Variables | Description | Mean | | Standard Deviation | |
|--|--|-------|------|--------------------|------|
| | | Pre | Post | Pre | Post |
| Variables over the period 1980-1999 | | | | | |
| Low-wage | Indicator (1=Low-wage), p10 | 0.10 | 0.10 | 0.30 | 0.30 |
| Low-wage | Indicator (1=Low-wage), 2/3 median | - | 0.13 | - | 0.34 |
| Un- and non-employed | Indicator (1=Not working) | 0.03 | 0.16 | 0.22 | 0.42 |
| Migrate | Indicator (1=Migrate) | 0.001 | 0.02 | 0.03 | 0.14 |
| Age | Age in years | 34.7 | 44.5 | 6.58 | 6.54 |
| Female | Indicator (1=Female) | 0.53 | 0.52 | 0.50 | 0.50 |
| Education | | | | | |
| Low-skilled | No formal degree | 0.14 | 0.13 | 0.34 | 0.33 |
| Medium-skilled | Apprenticeship | 0.69 | 0.69 | 0.47 | 0.47 |
| High-skilled | University | 0.09 | 0.10 | 0.31 | 0.31 |
| Occupational status* | | | | | |
| White-collar | White-collar worker | 0.39 | 0.44 | 0.49 | 0.49 |
| Blue-collar | Blue-collar worker | 0.48 | 0.50 | 0.50 | 0.50 |
| Skilled occupation | Engineer, professional, manager | 0.14 | 0.17 | 0.36 | 0.38 |
| Medium-skilled occupation | Qualified manual, service, commercial | 0.55 | 0.48 | 0.36 | 0.38 |
| Simple occupation | Simple manual, service, commercial | 0.31 | 0.35 | 0.34 | 0.47 |
| Labour market characteristics | | | | | |
| Experience | Years worked | 14.5 | 23.4 | 6.96 | 7.23 |
| # Interruptions | Number of interruptions | 3.10 | 5.01 | 3.15 | 4.24 |
| Interruption length | Cum. length of interruptions (in months) | 14.7 | 23.3 | 23.5 | 31.3 |

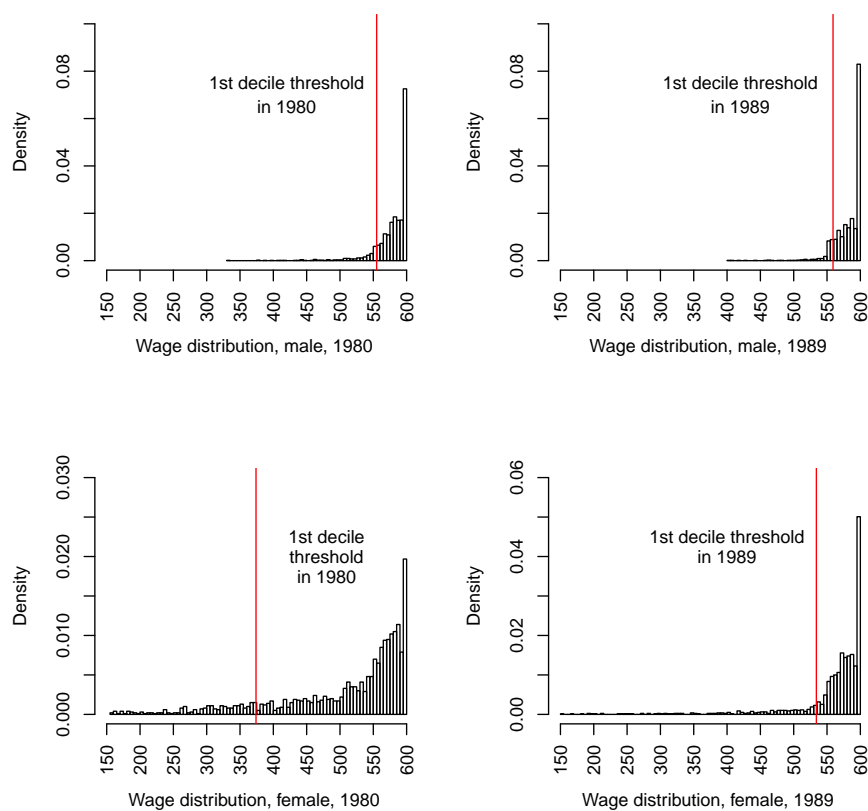
Source: BASiD 2007.

Notes: *Information for occupational status (white- and blue-collar) and the occupational groups are imputed for the pre-unification period based on information from the first available year from the Employment Statistics Register.

4 Descriptive Statistics

4.1 Wage Information before Unification and the Definition of the Low-Wage Threshold

Regarding wage information prior to Unification, wages until the first half of 1990 were censored above at 600 Mark. Figure 1 illustrates the pre-unification wage distribution (separately for male and female workers) for 1980 and 1989, respectively. The figure also marks the first deciles in relation to the censoring limit. It shows that



Source: BASiD 2007, weighted sample.

Figure 1: Distribution of Wages between 1980-1989, by Gender

the fraction of individuals earning monthly wages below 600 Mark decreased over time. The lower panel illustrates that the first decile for males increased only slightly between 1980 and 1989, whereas for females it grew from about 370 Mark in 1980 to

530 Mark in 1989. The figures show that due to the large extent of censoring in our data we are only capable of fully observing the lower part of the wage distribution. To measure individuals' relative economic position, we will exploit this information to measure the incidence and duration of individuals' position in the bottom part of the pre-unification wage distribution. Note that because our analysis focuses on birth cohorts from 1940 to 1960, the relative position needs to be interpreted in relation to this specific sub-population. In accordance with the literature on low pay, individuals are defined as being "low paid" if their wage does not exceed a specified threshold. Previous studies have used different definitions of the low-pay threshold, such as lower quantiles of the wage distribution (Cappellari, 2002 and Cappellari, 2007) or alternatively some fixed proportion of a quantile, such as two thirds of the median wage (Cappellari and Jenkins, 2008 and Uhlenborff, 2006). Given that in 1989, especially among male workers, only a small fraction earned less than the social security contribution limit of 600 Mark, we therefore have to adopt a somewhat more restrictive definition of low pay by fixing the low-pay threshold at the first decile of the wage distribution.¹⁰

Figure 1 shows that shortly before Unification the first deciles approach the censoring limit of 600 Mark. A particular concern is therefore that measurement error due to underreporting might misclassify a certain fraction of individuals as falling below the censoring limit. Even though we cannot fully rule out such kind of measurement error, we argue that there are at least two reasons speaking against it. First, if measurement error due to underreporting played a significant role, this should lead to a downward biased estimate of the first decile of the monthly wage distribution. To check whether this is the case, we compare the decile obtained from our data set with figures from external data sources. An ideal candidate data set is the German Socioeconomic Panel (GSOEP), whose retrospective survey in 1990 provides representative and uncensored information on former GDR workers' monthly labour earnings for the year 1989. According to the GSOEP, the first decile

¹⁰Appendix D in the online publication provides robustness checks based on the wage distribution pooled for males and females. This has the advantage that the first decile threshold becomes farther away from the 600 Mark censoring limit.

of monthly wages was about 560 Mark in 1989, whose order of magnitude is broadly in line with our pooled figure of 550 Mark.¹¹ A second reason speaking against underreporting stems from the administrative nature of the pension data. Central to this argument is the view that earnings dependent pension entitlements should create large incentives to correctly report (or at least not to underreport) earnings. In Appendix C, we demonstrate that even though GDR pension entitlements were only to a limited extent earnings dependent, monthly earnings which fell into a small earnings interval (between 500 and 600 Mark - depending on the number of creditable pension years) effectively raised pension entitlements. This yields an argument against a systematic measurement error due to underreporting, especially within this interval.

Apart from measurement error, another crucial issue concerns the interpretation of our low-pay threshold in terms of real consumption possibilities. Krueger and Pischke (1995) argue that due to the subsidisation of necessities, low paid individuals in a centrally planned economy were relatively better off than comparable individuals in a market economy. Given that individuals at the lower end of the earnings distribution spend a larger fraction of income on subsidised goods, the real earnings distribution should have been therefore more compressed than the nominal one. This raises the question as to whether earnings below our low-pay threshold also reflect economic deprivation in terms of real consumption possibilities. A tentative answer to this question may be provided based on measures of absolute poverty in the GDR. Manz (1992) derives such a measure by defining a minimum consumption level as the poverty threshold. Based on data on the consumption price level from the GDR Statistical Office, the author estimates the nominal value of this minimum consumption basket for a single household to amount to approximately 350 Mark in 1972 and 550 Mark in 1988. However, it needs to be stressed that these results have to be interpreted with caution as the data sources for the underlying consumption basket are lacking in this study.

¹¹Source: German Socioeconomic Panel, own calculations. The figure is obtained by pooling male and female working age individuals with positive earnings, after excluding apprentices, civil servants and the self-employed.

4.2 Unconditional Probabilities

In order to validate the data internally, Table 3 provides unconditional estimates of being low paid based on individual characteristics, separately for males and females. Several results stand out here: First, the raw unconditional probability of being low paid decreases with skills. For males, this decline is somewhat stronger during the post-unification period. Second, while younger individuals are slightly more likely to be low paid than older ones, there appear to be no large differences across both periods. Third, blue-collar workers and those with simple occupations have larger

Table 3: Unconditional probabilities, by Gender and Period

| | First Decile | | | | 2/3 Median | |
|------------------------------|--------------|------|---------|------|------------|---------|
| | Males | | Females | | Males | Females |
| | Pre | Post | Pre | Post | Post | Post |
| Low-wage | 10.0 | 10.0 | 10.0 | 10.0 | 7.1 | 19.3 |
| Education | | | | | | |
| Low-skilled | 12.4 | 15.3 | 11.7 | 11.0 | 10.2 | 21.9 |
| Medium-skilled | 10.4 | 10.1 | 10.4 | 10.9 | 7.3 | 20.1 |
| High-skilled | 5.4 | 2.5 | 3.0 | 2.7 | 2.1 | 4.8 |
| Age groups | | | | | | |
| 20-29 | 12.9 | - | 9.7 | - | - | - |
| 30-39 | 9.7 | 11.4 | 10.3 | 10.7 | 7.2 | 18.5 |
| 40-49 | 8.8 | 9.0 | 9.4 | 9.6 | 6.7 | 17.8 |
| 50-59 | - | 10.6 | - | 9.1 | 8.2 | 19.2 |
| Occupation | | | | | | |
| White-collar | 6.2 | 3.4 | 7.9 | 5.2 | 2.7 | 10.4 |
| Blue-collar | 11.3 | 12.5 | 8.3 | 15.7 | 8.8 | 30.6 |
| Skilled occupation | 5.3 | 2.0 | 4.5 | 1.6 | 1.4 | 4.0 |
| Medium-skilled occupation | 10.9 | 9.8 | 11.4 | 9.5 | 7.2 | 17.0 |
| Simple occupation | 10.8 | 13.3 | 9.7 | 15.2 | 9.4 | 29.8 |
| Labour market history | | | | | | |
| Experience | | | | | | |
| > 20 years | 9.2 | 9.6 | 6.9 | 8.9 | 7.0 | 17.7 |
| ≤ 20 years | 10.4 | 10.4 | 11.0 | 10.8 | 7.3 | 19.2 |
| # Interruptions | | | | | | |
| > 5 times | 20.7 | 14.2 | 9.7 | 9.7 | 10.3 | 20.0 |
| ≤ 5 times | 9.5 | 9.2 | 10.1 | 9.5 | 6.7 | 17.1 |
| Length Interruptions | | | | | | |
| > 20 months | 21.8 | 31.0 | 13.2 | 11.2 | 27.4 | 22.2 |
| ≤ 20 months | 10.0 | 8.7 | 7.3 | 7.3 | 6.1 | 13.7 |

Source: BASiD 2007, weighted sample

Notes: Pooled weighted data for the years 1980-1999. Left numbers (*Prior*) are probabilities prior to Unification. Right numbers (*Post*) refer to probabilities after Unification.

low-pay probabilities, with the differences between occupations becoming slightly more pronounced during the post-unification period. Finally, as to the importance of the labour market history, the probability of being low paid decreases with experience and increases with the number and cumulative length of labour market interruptions.

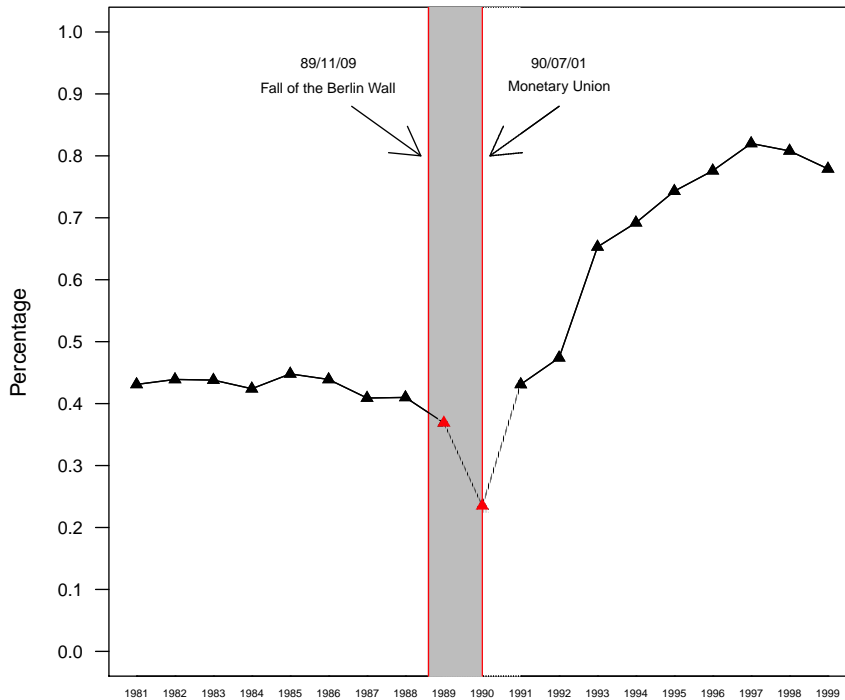
For completeness, the last column shows probability estimates based on the more conventional low-wage threshold of two-thirds of the median. Due to data availability the figures refer to the post-unification period. While the overall pattern of results remains unchanged, the magnitude of the differences becomes somewhat more pronounced.

4.3 Labour Market Transitions

To calculate labour market transitions we compute the extent of aggregate state dependence (*ASD*), defined as the difference in the probabilities of low pay conditional on being initially low paid and highly-paid in period $t - 1$. Thus, *ASD* is defined as:

$$ASD = P(L_t = 1|L_{t-1} = 1) - P(L_t = 1|L_{t-1} = 0), \quad (1)$$

with $L_t = 1$ and $L_t = 0$ meaning low and high pay in year t , respectively. To describe the evolution of low-wage persistence, Figure 2 plots *ASD* against time. Distinguishing the pre-unification (to the left of the grey bar), the transition (between the vertical lines) and the post-unification period (i.e. the time after monetary union), several noteworthy facts emerge from Figure 2. During the pre-transition period, aggregate state dependence varied around 42 per cent. During transition, aggregate state dependence decreased markedly by more than 20 percentage points to 24 per cent in 1990 compared to the pre-unification period. The third part of the figure (to the right of the grey bar) indicates a sharp rise in low-pay persistence, with aggregate state dependence increasing from 40 per cent in 1991 to values above 70 per cent in the late 1990s. Figure 3 shows the evolution of aggregate state dependence by gender. The figure reveals that the sharp decline in aggregate state



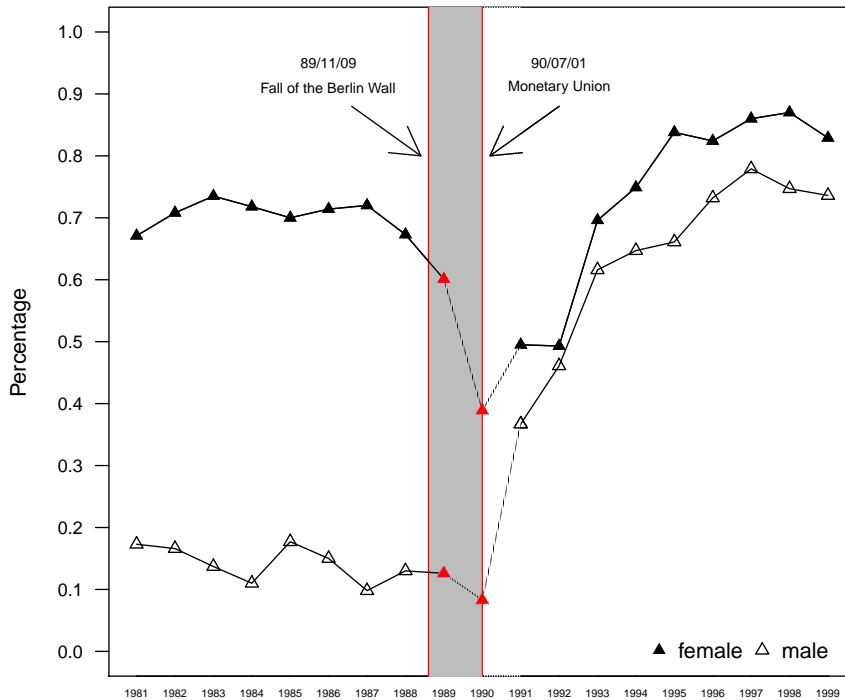
Source: BASiD 2007, weighted sample.

Notes: The figure shows the difference in the probability of being in low-wage employment given low-wage and high-wage employment in $t-1$. The first line marks the fall of the Berlin Wall. The second line marks Monetary Union.

Figure 2: Aggregate State Dependence

dependence during transition is mostly accounted for by female workers. After Unification, there appears to be a strong convergence between male and female workers. Finally, Figure A.1 in the appendix reports the gender-specific evolution of aggregate state dependence based on the pooled wage distribution. Compared to Figure 3, the observed minimum in aggregate state dependence during transition becomes somewhat more pronounced for males.

The overall picture that emerges from Figure 2 along with Figure 3 and Figure A.1 is that aggregate state dependence plummeted with the beginning of a market-orientated economy. The post-unification period is characterised by a steady rise in low-pay persistence during the first years and is facing a level-off at the end



Source: BASiD 2007, weighted sample.

Notes: The figure shows the difference in the probability of being in low-wage employment given low-wage and high-wage employment in $t-1$. The first line marks the fall of the Berlin Wall. The second line marks Monetary Union.

Figure 3: Aggregate State Dependence, by Gender

of the 1990s. Overall, the figures show that the importance of previous low-wage employment for low-pay in the current period reaches its minimum during the time of transition, albeit less pronounced for male workers.

5 Econometric Approach

5.1 Short-run Dynamics within and across Political Regimes

The descriptive figures from the previous sections may hide potential compositional effects and do not allow us to infer any conclusions about the extent of the causal effect of the previous low-pay status. In what follows, we will therefore attempt to

isolate persistence in low-pay due to observed and unobserved heterogeneity from genuine state dependence. The latter basically measures to what extent low-wage employment today causes low-wage employment in the future. We will start the econometric analysis by setting up a joint model of low pay over the whole sample period. This may provide important insights into the relevance of observed and unobserved factors and the difference in the estimated parameters across the different regimes. To quantify the importance of observables and unobservables, we specify the following baseline model of low-pay employment:

$$y_{it}^* = (\gamma + \delta H_{\tau t})y_{it-1} + x'_{it}\beta + \epsilon_i + u_{it} \quad (2)$$

$$y_{it} = I(y_{it}^* \geq 0) \quad (3)$$

where x_{it} are the explanatory variables presented in Table 2, ϵ_i is an individual-specific random effect and y_{it-1} denotes the low-wage status in $t - 1$. We further assume u_{it} to be an idiosyncratic error term with $u_{it} \sim N(0, 1)$.¹² To allow for a change in short-run persistence across the pre- and post-unification period (FRG), we further interact the lagged endogenous variable y_{it-1} with a period dummy variable $H_{\tau t}$, where $H_{\tau t} = 1$ iff $t \geq \tau$ and $\tau = 1990$. The model includes as time-varying explanatory variables the number of labour market interruptions, cumulated labour market interruptions measured in months and a set of experience dummy variables. We additionally include a lagged non-employment indicator variable.

Estimating the dynamic random effects model without modelling the initial condition will bias the estimation results of the coefficients (see for example Heckman, 1981*b*, Stewart and Swaffield, 1999, Honoré and Kyriazidou, 2000, Arulampalam et al., 2000). The dynamic structure of the model implies that individual i 's labour

¹²This approach does not take into account the state of not working in period t . Note, however, that the necessity of modelling selection out of employment differs across the regimes. Prior to Unification, the right and duty to work caused only few individuals to stay non-employed, indicating that selection out of employment is less of a concern. This is true especially for males. In our sample 85 per cent among male workers did never experience a non-employment spell between 1980 and 1989 and 95 per cent experienced no or one year of non-employment. Among females, the numbers are 70 per cent and 85 per cent, respectively.

market position in period 1 depends on status in period 0 (Heckman, 1981*a*). Given the endogeneity of the first period’s outcome due to a correlation with the time-invariant individual effect (selection), state dependence is likely to be overestimated (see Chay and Hyslop, 1998).

In order to treat the initial condition, we follow Wooldridge (2005) and model the density of y_{it} for all $t = 1, \dots, T$ given covariates x_{it} , start at $t = 2$ and condition the density of y_{it} for $t = 2, \dots, T$ on y_{i1} and x_{it} . This approach is comparable to the correlated random-effects model put forward by Chamberlain (1984) (see also Prowse, 2012).¹³ To account for a potential correlation of the unobserved individual effect with observed explanatory variables, we follow Mundlak (1978) and Chamberlain (1984) and model the individual effect, ϵ_i , as:

$$\epsilon_i = a_1 y_{i1} + a_2 \bar{x}_i + \alpha_i \tag{4}$$

where \bar{x} presents individual time averages of all time varying variables, y_{i1} is the initial low-wage status in the first period and $\alpha_i \sim N(0, \sigma_\alpha^2)$.¹⁴

In general, the assumptions made in the specified baseline model (equation 2 to 4) of a constant random effect α_i and β ’s over time might be too strong. To provide an example for α_i , one may think about one’s political attitude (e.g. civil rights campaigner) that might increase the probability of being at the bottom of the wage distribution. This effect might reverse or at least change after Unification. Moreover, the unconditional probability estimates (Table 3) indicate a change in the estimated β -coefficients. Therefore, we estimate a bunch of models that allow sequentially for a change in the coefficients after Unification. We start with the *baseline model* (equation 2) where β and α_i are held constant across time. *Model*

¹³Another approach developed by Heckman (1981*b*) models the joint distribution for the entire sample ($t = 1, \dots, T$) of an individual’s current status. This estimator is, for example, used by Arulampalam and Stewart (2009) and involves the specification of an equation for the initial condition.

¹⁴As shown by Akay (2012) using Monte Carlo results, both estimators (*Heckman* and *Wooldridge*) tend to perform similarly in long panels (above six time periods), with the *Heckman* approach producing less biased results for a small T -dimension.

Table 4: Modelling Approach

| Model | observed heterogeneity | unobserved heterogeneity |
|----------------|------------------------|--------------------------|
| Baseline model | β constant | α_i constant |
| Model 1 | β variable | α_i constant |
| Model 2 | β constant | α_i variable |
| Model 3 | β variable | α_i variable |

Notes: The models are specified over the time periods between 1980 and 1999. All models in the random effects specification account for the initial condition using the *Wooldridge approach* and follow the literature using the Mundlak-Chamberlain device. Model 2 and 3 are estimated by simulation using 100 Halton draws.

1 allows additionally for a change in β after Unification while α_i remains constant. In *model 2*, we allow for a change in α_i holding β constant. *Model 3* sets up the process of low pay over the whole period by allowing both coefficients to change after Unification. The *baseline model* and *model 1* are estimated using quadrature methods whereas the last two models are estimated by simulated maximum likelihood using 100 Halton draws. This sequential estimation approach allows us to learn about regime effects and the importance of confounding factors for the two different regimes.

Results. To quantify the importance of observables and unobservables, we estimate the average partial effect (APE) for each of the time periods considered. The method used is based on counterfactual outcome probabilities based on the lagged endogenous variable. Following Papke and Wooldridge (2008), the specified average structural function (ASF) for period t can be written as:

$$ASF_t = N^{-1} \sum_{i=1}^N \Phi[(\hat{\gamma} + \hat{\delta}H_{\tau t})y_{it-1} + x'_{it}\hat{\beta} + \hat{a}_0 + \hat{a}_1\bar{x}_i + \hat{a}_2y_{i1}](1 + \hat{\sigma}_\alpha^2)^{-1/2}] \quad (5)$$

for $t = 1981 - 1999$. This function corresponds to state dependence accounting for observables and unobservables. By setting $y_{it-1} = 1$ and subtracting the function value for $y_{it-1} = 0$, we get the average partial effect of being in low-wage employment at time t for individuals being low paid in $t-1$ compared to high-paid individuals in $t-1$. Tables 5 and 6 report the results for the defined periods (GDR, FRG). The first

table reports average partial effects without modelling unobserved heterogeneity. The aggregate effects are taken from Figure 3, averaged over the two time periods. We first observe that after accounting for observed heterogeneity (β constant), state

Table 5: Average Partial Effects of Pooled Probit Models, by Gender and Period

| | Aggregate | | β constant | | β variable | |
|---|-----------|---------|------------------|---------|------------------|---------|
| | (1a) | (1b) | (2a) | (2b) | (3a) | (3b) |
| | Males | Females | Males | Females | Males | Females |
| $P(y_t = 1 y_{t-1} = 1) - P(y_t = 1 y_{t-1} = 0)$ | | | | | | |
| APE_{GDR} | 14.1 | 69.3 | 10.8 | 64.7 | 10.2 | 64.9 |
| APE_{FRG} | 58.3 | 70.4 | 48.4 | 61.3 | 46.7 | 59.1 |

Source: BASiD 2007, weighted sample.

Notes: The table reports the results from dynamic discrete choice probit models on the probability of being low paid. $y_t = 1$ denotes the low-wage status. All estimated average partial effects are significant at the 5% level using 100 bootstrapped replications. Each regression for the male (female) sample has 35847 (36178) observations with 2248 (2540) individuals. Detailed results on estimated coefficients are shown in our online appendix.

dependence declines for both males and females during both periods. However, the decline is rather small, suggesting that low-pay persistence is largely independent of differences in observables that are available in our data set. Allowing the β -coefficients to change after Unification reduces the point estimates of the average partial effects only slightly.

Table 6 shows the results based on the random effects specifications. The baseline specification (β and α_i constant) shows a reduction in estimated state dependence for all time periods compared to the pooled model. However, the decline in state dependence in absolute terms compared to the pooled probit model (Table 5, *model 2a-2b*) is most pronounced for the pre-unification period. A concern with the resulting APE_{FRG} is that it does not account for selection out of employment after Unification. This selection process is less of a concern during GDR because of the right and the duty to work. After Unification, state dependence for male workers increases substantially compared to the pre-unification period whereas for female workers the change is not significant. The lower part of the table reports the estimates of the random effects' variances. The variance of the random effect is higher

Table 6: Average Partial Effects of Random Effect Probit Models, by Gender and Period

| | β constant α_i constant | | β variable α_i constant | | β constant α_i variable | | β variable α_i variable | | β variable α_i variable | |
|---|---|---------|---|---------|---|---------|---|---------|---|---------|
| | (1a) | (1b) | (2a) | (2b) | (3a) | (3b) | (4a) | (4b) | (5a) | (5b) |
| | Males | Females | Males | Females | Males | Females | Males | Females | Males | Females |
| $P(y_t = 1 y_{t-1} = 1) - P(y_t = 1 y_{t-1} = 0)$ | | | | | | | | | | |
| APE_{GDR} | 4.7 | 37.8 | 4.2 | 37.4 | 4.9 | 41.6 | 4.9 | 41.9 | 4.9 | 41.6 |
| APE_{FRG} | 39.9 | 38.5 | 37.2 | 36.2 | 37.5 | 46.6 | 34.8 | 45.2 | 34.8 | 39.4 |
| <i>Random Effects</i> | | | | | | | | | | |
| σ^2 | 0.111 | 0.302 | 0.117 | 0.315 | | | | | | |
| σ_{GDR}^2 | | | | | 0.201 | 0.519 | 0.189 | 0.480 | 0.189 | 0.494 |
| σ_{FRG}^2 | | | | | 0.703 | 0.624 | 0.668 | 0.638 | 0.666 | 0.872 |
| $\sigma_{GDR,FRG}$ | | | | | | | | | -0.009 [†] | 0.305 |

Source: BASiD 2007, weighted sample.

Notes: The table reports the results from dynamic discrete choice probit models on the probability of being low paid. $y_t = 1$ denotes the low-wage status. All estimated average partial effects and variance estimates ([†]apart from the covariance estimate for male workers) are significant at the 5% level using 100 bootstrapped replications. Each regression for the male (female) sample has 35847 (36178) observations with 2248 (2540) individuals. Detailed results on estimated coefficients are shown in an online appendix.

for female workers indicating a more heterogeneous group. *Model 2a-2b* allows for a change in all β -coefficients after Unification but assuming again a constant random effect. Similar to the results in Table 5, changing β has only modest effects on the average partial effects. Compared to the first specification of the same table (*model 1a-1b*), the reduction in state dependence after Unification is between about two to three percentage points. The next three models allow for a change in the random effects after Unification. In all specifications, the variance increases stronger for the period after Unification. This might be an indication of offsetting forces within the different economic environments that drive down the variances under the constancy restriction. Put differently, the results suggest that there are unobservable factors in place that increase the probability of being low paid in one regime and change their character after the systemic transformation. While models (3) and (4) restrict the correlation between regime-specific unobservables to be zero, model (5) allows for a covariance term between the two random effects. For male workers the covariance is virtually zero and not significant, whereas for female workers unobservables that determine the stochastic low-wage process in the GDR are partly also of relevance in the new regime.

From the approach so far, we observe that unobserved heterogeneity is the driving force and that it appears to be regime-specific. Compositional effects based on observable characteristics are less important in explaining state dependence. The analysis, however, fails to account for selection out of employment. This is of considerable importance for the post-unification period and might lead to an upward bias of the estimated average partial effect, APE_{FRG} , if low-wage individuals were more likely to experience a transition into non-employment. Moreover, the analysis only sheds light on short-run within-regime dependence. In what follows, we will focus on economic dependence across the different regimes by adopting an empirical strategy that addresses these issues.

5.2 Selection, Long-run Effects and Heterogeneous Short-run Dynamics after Unification

The main purpose now is to model the selection mechanism out of employment after Unification that will otherwise bias the estimation results of the short-run measures of state dependence and to estimate across-regime effects of low-wage employment. The multivariate analysis using a bivariate random effect model has shown that the random effects are regime-specific. Given that the institutional setting (socialism, market-economy) changes at the time of Unification we will start the analysis in 1990, the first year after the fall of the Berlin Wall, and model the stochastic process of low pay until 1999. This enables us to model a constant random effect and to account for selection within the new regime. A further important advantage of this strategy is that it allows us to analyse long-run across-regime state dependence and to use detailed variables of firm characteristics in the market economy which is otherwise not possible due to data availability. Another concern is the chosen low-wage threshold of the first decile. Focussing on the post-unification period enables us to measure the low-pay outcome by adopting the more conventional measure of two-thirds of the median. Our econometric model becomes:

$$y_{ikt}^* = \gamma_k y_{it-1} + x'_{it} \beta_k + \vartheta_{ikt} \quad (6)$$

$$y_{ik1}^* = x_{i1}'\beta_k + \vartheta_{ik1}, \quad (7)$$

where y_{ikt}^* is a latent unobservable variable measuring the propensity of earning a high wage ($k = 0$), a low wage ($k = 1$) or being non-employed ($k = 2$) for individual i in year t for $t > 1$. Similar to the model above, the propensity is a function of an individual's previous state, y_{it-1} , individual characteristics, x_{it} , as well as unobserved characteristics, ϑ_{ikt} . y_{ik1}^* refers to the initial process. We further assume the unobserved part ϑ_{ikt} to consist of an individual and state specific random component, ϵ_{ik} , which is assumed to be time-invariant and uncorrelated with all x -variables at every point in time, and a time-variant idiosyncratic component, u_{ikt} , i.e. $\vartheta_{ikt} = \epsilon_{ik} + u_{ikt}$. By making distributional assumptions about u_{ikt} , we follow the literature and assume a type I extreme value distribution. For identification we set the state of high-wage employment as the reference state.

Our approach to analyse across-regime dependence is twofold. The first issue concerns the modelling of the initial condition. Given that we start directly at the time of Unification, the initial condition is supposed to be determined by conditions in 1990. The motivation of modelling the initial condition is generally to account for selection in the first period. As a result, its significance may help us to assess the randomness of the initial allocation to the low-wage sector during transition. Usually, a (positive) significant value of the respective coefficient would indicate that the start of the modelled stochastic process is correlated with unobservables governing low-pay after Unification. Therefore, if the coefficient on the initial condition is insignificant, the allocation of individuals to the low-wage sector in 1990 may be considered close to random in terms of market-regime unobservables.

Second, the data situation allows us to estimate effects of pre-unification low-wage experience on post-unification low-wage propensities. As a first measure of pre-unification low-wage experience we count the years an individual was below the first decile between 1980 and 1989. In our sample there are 1074 (49 per cent) male and 637 (27 per cent) female workers who spent at least one year below the first wage decile during 1980 and 1989. From those individuals, 90 per cent of males

and 66 per cent of females had between one to three years of low-wage experience. Table 7 reports the descriptive statistics on the probability of being below two-third of the median over the time period between 1990 and 1999 conditional on the across-regime variable *number of years below the first decile*. From the table

Table 7: Percentage of Low-wage Employment Conditional on the Number of Years below the First Decile before Unification

| | 1990 | 1991 | 1992 | 1993 | 1994 | 1995 | 1996 | 1997 | 1998 | 1999 |
|---|------|------|------|------|------|------|------|------|------|------|
| Males | | | | | | | | | | |
| <i>low - wage</i> | 5.3 | 7.3 | 5.4 | 6.0 | 7.0 | 7.3 | 8.6 | 8.9 | 8.8 | 9.3 |
| <i>low - wage</i> # <i>low - wage</i> 1 - 3 | 6.0 | 7.8 | 5.6 | 6.3 | 7.2 | 7.9 | 9.3 | 10.4 | 9.8 | 10.1 |
| <i>low - wage</i> # <i>low - wage</i> > 3 | 16.3 | 19.1 | 16.2 | 16.4 | 11.8 | 9.3 | 9.7 | 11.9 | 16.1 | 17.4 |
| Females | | | | | | | | | | |
| <i>low - wage</i> | 13.9 | 14.4 | 14.8 | 18.6 | 18.6 | 18.9 | 21.4 | 23.0 | 25.1 | 24.8 |
| <i>low - wage</i> # <i>low - wage</i> 1 - 3 | 23.7 | 23.4 | 21.4 | 27.0 | 28.3 | 23.5 | 28.1 | 31.3 | 28.8 | 30.3 |
| <i>low - wage</i> # <i>low - wage</i> > 3 | 48.9 | 42.2 | 31.3 | 40.8 | 35.5 | 34.5 | 40.5 | 41.4 | 44.7 | 41.7 |

Source: BASiD 2007, weighted statistics.

we first observe that the low-wage sector has been increasing over time for both, males and females. Experiencing a low wage episode between one and up to three years during 1980 and 1989 raises the probability of belonging to the low-wage sector after Unification only slightly for male workers, whereas the difference becomes more pronounced for females. The probability of being below two-thirds of the median wage increases further if the number of low-wage years prior to Unification was above three. For both, males and females, the probability has been increasing at the end of the sample period. Thus, individuals who were relatively frequently low paid during the GDR regime are more likely to earn lower wages in a market-based economy. Given the higher uncertainty about one's individual productivity due to weaker signalling potentials, the connection between the low-wage status prior to and after Unification seems to have decreased first and then started to increase again. Additionally to this low-wage measure, the data set allows us to construct employment history measures of labour market interruptions. This includes all kind of interruptions such as child-care and illness. We calculate the number of non-employment months between labour market entry and 1989 (see Mosthaf, 2011 for an empirical strategy to include pre-sample variables into the model).¹⁵

¹⁵See Appendix B for the possible economic states in the data set. To allow for a flexible

Accounting for the pre-unification information, our empirical model becomes:

$$y_{ikt}^* = \gamma_k y_{it-1} + x_{it} \beta_k + \Gamma_i \varphi_k + \epsilon_{ik} + \vartheta_{ikt} \quad (8)$$

where y_{ikt} is the propensity of individual i being in state k at time t with $t = 1991 - 1999$. As shown in the descriptive analysis, the main labour market states k are high-wage, low-wage and out-of-employment (including unemployed and non-employed and referred to as non-employment in the following) with *high-wage* being the reference category.¹⁶ x_{it} represents the explanatory variables and y_{it-1} is a vector of labour market states in $t - 1$. The individual- and state-specific random effect is modeled using the Wooldridge approach:

$$\epsilon_{ik} = a_{1k} y_{i,1990} + a_{2k} \bar{x}_i + \alpha_{ik} \quad (9)$$

where \bar{x} presents individual time averages of all time varying variables, $y_{i,1990}$ is the initial low-wage status in the first period and $\alpha_{ik} \sim N(0, \Sigma_\alpha)$. The cumulative pre-unification low-pay histories are captured by Γ_i . Note that consistent identification of φ_k requires Γ_i to be uncorrelated with α_{ik} . The extent to which this assumption holds depends on the correlation between the regime-specific unobservables. Our results from the previous section suggest that the identifying assumption is likely to be met for males, whereas for females the coefficient may still be upward biased.

Assuming a multivariate normal distribution of the random intercepts leads to a 2-dimensional integral. We estimate this model by simulation (MSL) using 100 Halton draws.¹⁷

dependence we construct labour market interruption intervals that are different for males and females. Due to institutional settings and socio-cultural norms, male individuals typically exhibit rather few interruptions. Therefore, the baseline interval for males is zero interruptions. We further define medium interruptions between one and six months and more than six months to be the high category. The baseline for females is up to 12 months. The medium category comprises 12 to 48 months and more than 48 months is defined as the high category.

¹⁶The out-of-employment state should be interpreted as an absorbing state that might affect low-wage dynamics. We therefore include the migration decision (see Table 2) into this state to control for selective migration.

¹⁷For a detailed description for MSL properties see for example Gourieroux and Monfort (1993). Train (2003) discusses illustratively the properties of Halton draws in terms of coverage and effi-

The importance of firm characteristics for wage mobility is well established in the literature on low-wage persistence (see Gürtzgen and Heinze, 2016). However, estimating multinomial logit models with different sets of covariates is not common in the literature. We include therefore firm size dummies, the share of female workers, the share of low-skilled workers and the median age of the establishment work force as well as detailed occupational information in the low-wage equation to control for firm characteristics in the dynamic process.

5.3 Empirical Results

For expositional purposes, we confine the presentation of the results to the main variables of interest. The upper part of Table 8 shows marginal effects of the lagged low-wage indicator on the probability of belonging to the low-wage sector. For male workers (*model 1a*), the probability of belonging to the low-wage sector is 16.7 percentage points higher if one has been low paid in $t - 1$. For females (*model 1b*), the estimate of true state dependence is with 31.3 percentage points almost twice as high. The interrelation between non-employment and low-wage employment is also found to be statistically significant and is in line with the literature. The middle part of the table shows the results for the initial condition using the Wooldridge approach. Focussing on the first model, the initial condition is not found to be significant. The modelled stochastic process of low-wage employment is independent from the outcome in the first year. This provides first evidence of a random initial allocation to the low-wage sector in 1990. The lower part of the table shows the marginal effects of the pre-unification variables (Γ_i). We do not find any across-regime effects if the number of years below the first decile in the wage distribution between 1980 and 1989 is between one and three. However, having more than three years pre-unification low-pay experience raises the probability of being low paid by 1.1 percentage points for men (albeit insignificantly) and by 3.2 percentage points for females. The second variable measuring across-regime dependence is the number of labour market interruptions.

ciency.

Table 8: Marginal Effects of Multinomial Logit Models with Random Effects, by Gender

| | Males (1a) | | Females (1b) | | Males (2a) | | Females (2b) | |
|-------------------------------------|---------------------|---------------------|---------------------|---------------------|---------------------|---------------------|---------------------|---------------------|
| | <i>low-wage</i> | <i>non-empl.</i> | <i>low-wage</i> | <i>non-empl.</i> | <i>low-wage</i> | <i>non-empl.</i> | <i>low-wage</i> | <i>non-empl.</i> |
| Short-run dynamics | | | | | | | | |
| <i>Low</i> – $wage_{t-1}$ | 0.167*** (0.016) | 0.079*** (0.020) | 0.313*** (0.020) | 0.078*** (0.015) | 0.145*** (0.014) | 0.033 (0.026) | 0.188*** (0.014) | 0.090*** (0.022) |
| <i>Non</i> – $employed_{t-1}$ | 0.096*** (0.013) | 0.137*** (0.009) | 0.117*** (0.012) | 0.174*** (0.008) | 0.088*** (0.012) | 0.102*** (0.008) | 0.115*** (0.012) | 0.128*** (0.017) |
| Initial condition | | | | | | | | |
| <i>Low</i> – $wage_{1990}$ | 0.009 (0.011) | -0.011* (0.006) | 0.003 (0.009) | -0.011 (0.007) | 0.020 (0.014) | -0.000 (0.006) | 0.014 (0.010) | -0.001 (0.007) |
| <i>Non</i> – $employed_{1990}$ | -0.005 (0.011) | 0.000 (0.006) | -0.004 (0.003) | 0.012 (0.008) | 0.013 (0.016) | -0.009 (0.007) | 0.011 (0.017) | 0.006 (0.008) |
| Long-run dynamics | | | | | | | | |
| # <i>Low-wage years</i> | | | | | | | | |
| 1 – 3 | 0.003 (0.005) | -0.004 (0.007) | 0.006 (0.012) | -0.003 (0.008) | 0.004 (0.006) | -0.005 (0.006) | -0.004 (0.007) | -0.002 (0.009) |
| > 3 | 0.011 (0.012) | 0.009 (0.006) | 0.032*** (0.013) | 0.006 (0.006) | -0.002 (0.013) | 0.010 (0.007) | 0.016 (0.012) | 0.007 (0.006) |
| # <i>Interruptions</i> [†] | | | | | | | | |
| <i>Medium</i> | -0.001 (0.013) | 0.003 (0.007) | 0.001 (0.008) | 0.003 (0.008) | -0.001 (0.005) | 0.000 (0.008) | 0.001 (0.007) | 0.003 (0.007) |
| <i>High</i> | -0.004 (0.011) | 0.004 (0.005) | -0.007 (0.009) | 0.006 (0.006) | -0.008 (0.006) | -0.004 (0.005) | 0.013 (0.011) | -0.002 (0.006) |
| <i>Controls</i> | | ✓ | | ✓ | | ✓ | | ✓ |
| <i>Occupation</i> | | ✓ | | ✓ | | ✓ | | ✓ |
| <i>Firm charac.</i> | | x | | x | | ✓ | | ✓ |
| obs. (total) | 17542 | | 18011 | | 16312 | | 16465 | |
| obs. (ind.) | 2181 | | 2289 | | 2165 | | 2267 | |
| LogLik | -6388.2 | | -8236.1 | | -5330.4 | | -5748.2 | |

Source: BASiD 2007, weighted sample.

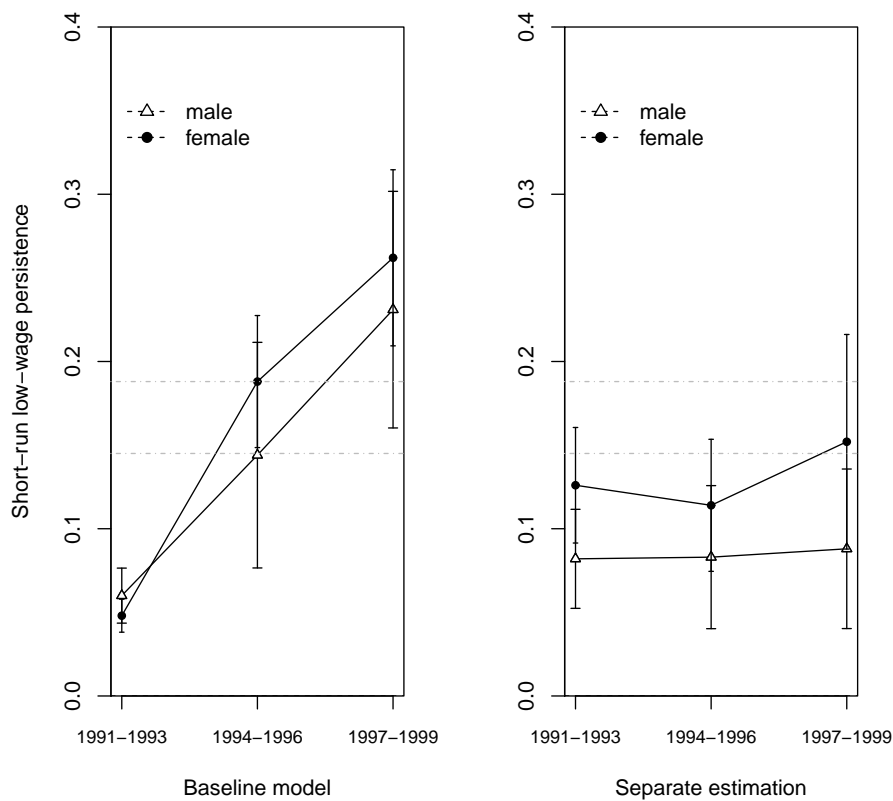
Notes: [†]The reference category differs by gender. For males, the reference is zero labour market interruptions; medium: up to six months; high: more than six months. For females, the reference is zero to 12 months; medium: 12 to 48 months; high: more than 48 months. Robust standard errors are in parentheses. All estimations contain a constant, the specified Mundlak-Chamberlain device, and control for missings in the education variable. Control variables include age dummies, education dummies, year dummies, regional dummies and experience dummies. The out of work equation includes lagged dummy variables of white-collar, skilled and simple occupation. Before Unification, the low-wage threshold is the first decile, while after unification two-thirds of the median is used. The average marginal effects are estimated by averaging over the average marginal effects of 100 Halton draws. Asterisks next to coefficients indicate significance levels as follows: *** 1%, ** 5%, * 10%. Detailed results on estimated coefficients and random effects are shown in the online appendix.

Interestingly, there is no effect of socialist interruptions on the probability of being low paid during the first decade after Unification. The last two models in Table 8 include as additional control variables firm characteristics such as firm size, the share of female and low-skilled workers as well as median age in the low-wage equation. All variables enter the specifications with their expected sign. The short-run measures of state dependence are significantly affected once firm characteristics are accounted for. For male workers, estimated state dependence decreases slightly, whereas for females the coefficient drops by more than ten percentage points. This points to a selection into low-wage firms, especially for female workers and shows how important it is to control for firm characteristics. The coefficients on the variables that measure across-regime dependence are reduced as well. Having experienced more than three years pre-unification low-pay increases the probability of being low paid after Unification for females workers by 1.6 percentage points (with a t -value of 1.45).

Heterogeneity across Time. The results in general do not provide any evidence of across-regime economic dependence and refer to the whole post-unification period. In principle, it might be conceivable that across-regime effects are heterogenous over time. The same is true for short-run state dependence. Given that individual restrictions with respect to qualification and occupational choices decreased after Unification, individuals may have been facing increasing opportunities to acquire skills in the market economy. Therefore, if persistence was the result of human capital depreciation, such a catching up process should have resulted in a decline in state dependence over time. On the other hand, if state dependence was the result of signalling low productivity levels, high initial uncertainty after Unification should have rendered the selection into low-wage jobs based on individual productivity unlikely. To the extent that uncertainty has been declining over time, such a mechanism should have resulted in an increase in state dependence after Unification. According to the short-run measure of state dependence after Unification, an increasing tendency over time would support the signalling mechanisms. Even though we cannot rule out the human capital effect, it is rather unlikely that human capital depreciation strongly changes and cause a rapid increase. Contrarily, catch-

ing up on the qualification level would work in the opposite direction compared to Figure 3.

To distinguish the first restructuring phase from further time periods, we estimate a model similar to *model 2* of Table 8 including interaction effects for the time periods 1994-1996 and 1997-1999 for the short- and long-run dynamic variables and the initial condition. Figure 4 (left) shows the results for the short-run low-pay dynamics ($t-1$). Over time, short-run persistence increased from about six per cent



Source: BASiD 2007, weighted sample.

Notes: The figure shows the marginal effects and the corresponding 90% confidence intervals of the probability of being in low-wage employment conditional on low pay in $t-1$ using a multinomial logit model with random effects and firm characteristics for the low-wage state distinguished by time and gender. The left hand figure shows the marginal effects of being low paid conditional on being low paid in $t-1$ (equation 8). The right hand figure estimates the effects separately for the three time periods. The marginal effects are estimated by averaging over 100 Halton draws. The light grey lines are the estimates from Table 8. Each regression for the male (female) sample (left and middle figure) has 16312 (16465) observations with 2165 (2267) individuals. Detailed results on estimated coefficients are shown in the online appendix.

Figure 4: Effect Heterogeneity - Short-run Dynamics

(three per cent) for males (females) in the first period (1991-1993) after Unification

to 23 per cent (26 per cent) during the last three years of the decade. While all estimates do not significantly differ from the estimates in Table 8 (see light grey lines), the difference between the first period (1991-1993) and the last period (1997-1999) is significant for both groups. However, a major concern with this specification is the potential presence of a dynamic selection process as the pool of low-wage individuals in the last three years might significantly differ from that in the first three years. To deal with this issue, the right hand part of the figure shows the results from running a separate regression for each defined period. It shows for both, males and females, that the probability of staying low paid is higher for workers who have already been low paid and that this probability is rather constant over time. This suggests that the former documented increase in short-run state dependence reflects a strong selection process after Unification.

To assess the randomness of the initial allocation to the low-wage sector, our previous results have shown that the initial condition is not significant. This indicates that the modelled stochastic low-wage process is independent of the low-wage status in 1990. Given that the lagged endogenous variable is interacted with two period dummy variables, we also interact the initial condition variable with the same period dummies. Table 9 shows the results. The point estimate of the initial condition for the first period (1991-1993) is slightly more positive compared to Table 8. It basically supports the findings suggesting that the initial allocation to the low-wage sector is close to random. For female workers the initial condition has a t -value of slightly below 1.6, suggesting a correlation between the observation in 1990 and the individual-specific random effects. The last two columns visualise the importance of the initial condition with separate estimations by sub-period. While the initial condition is insignificant for males and female workers for the first period, its coefficient for the second defined period (1994-1996) is strongly significantly positive. The same holds true for the last sub-period. Note that the strong increase in the initial condition's importance again provides evidence for a dynamic selection process.

Turning to the long-run variables that measure across-regime dependence, we

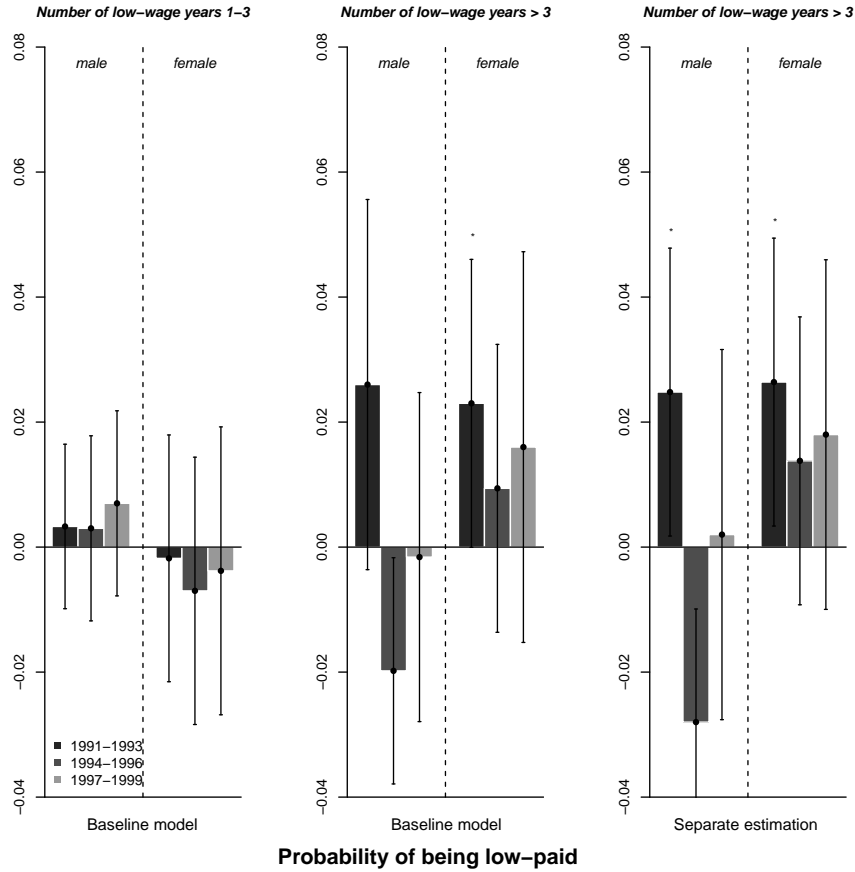
Table 9: Marginal Effects of the Initial Condition for the Low-wage Equation, by Gender

| | Males | Females | Males | Females |
|--------------------------------------|-------------------|--------------------|---------------------|---------------------|
| <i>Initial Condition: low – wage</i> | Baseline Model | | Separate Estimation | |
| 1991 – 1993 | 0.035 (0.022) | 0.022 (0.014) | 0.016 (0.013) | 0.022 (0.014) |
| 1994 – 1996 | -0.009 (0.017) | 0.006 (0.016) | 0.226*** (0.077) | 0.194*** (0.046) |
| 1997 – 1999 | -0.026 (0.017) | -0.025* (0.013) | 0.294*** (0.079) | 0.224*** (0.063) |

Source: BASiD 2007, weighted sample.

Notes: The table shows the marginal effects of the initial condition $low - wage_{1990}$ for the stochastic low-wage process. The initial condition is interacted with two period dummies. The baseline estimate is for the period between 1991-1993. The second period is the time between 1994-1996 and the third period captures the time between 1997-1999. The coefficients for the baseline model report changes relative to the first period (1991-1993). The marginal effects are estimated by averaging over 100 Halton draws. The last two columns report the estimation results separated by the three periods. The initial conditions for the three periods are 1990, 1993 and 1996, respectively. Each regression for the male (female) sample with interactions has 16312 (16465) observations with 2165 (2267) individuals. Asterisks next to coefficients indicate significance levels as follows: *** 1%, ** 5%, * 10%. Detailed results on estimated coefficients are shown in the online appendix.

first focus on the effect of the number of years one was below the first decile between 1980 and 1989 on the probability of being low paid after Unification. Figure 5 illustrates the marginal effects for the variables *1-3 years below the first decile* and *more than 3 years below the first decile*. It shows that there is no heterogeneity with respect to the first variable. For the second one, the figure’s middle part shows that having experienced more than three years pre-unification low-pay raises male workers’ low-pay probability in the first sub-period by about 2.5 percentage points (t -value: 1.5), whereas the effect vanishes after 1993. For females, the point estimates are positive throughout the post-unification period. Even though the estimate is (weakly) significant only for the first sub-period, the results indicate that across-regime dependence appears to be more persistent for female workers. The separate estimations (right part) by sub-period support these findings. Overall, we find (weak) evidence for across-regime dependence for workers who spent many years in low-wage employment before Unification and that this effect disappears for males. Note, however, that according to our discussion above, the estimates for female workers may still be upward biased. Figure 6 shows the impact of pre-unification labour market interruptions on the low-pay probability after Unification. While there is no effect for males, females have a 2.6 percentage points higher low-



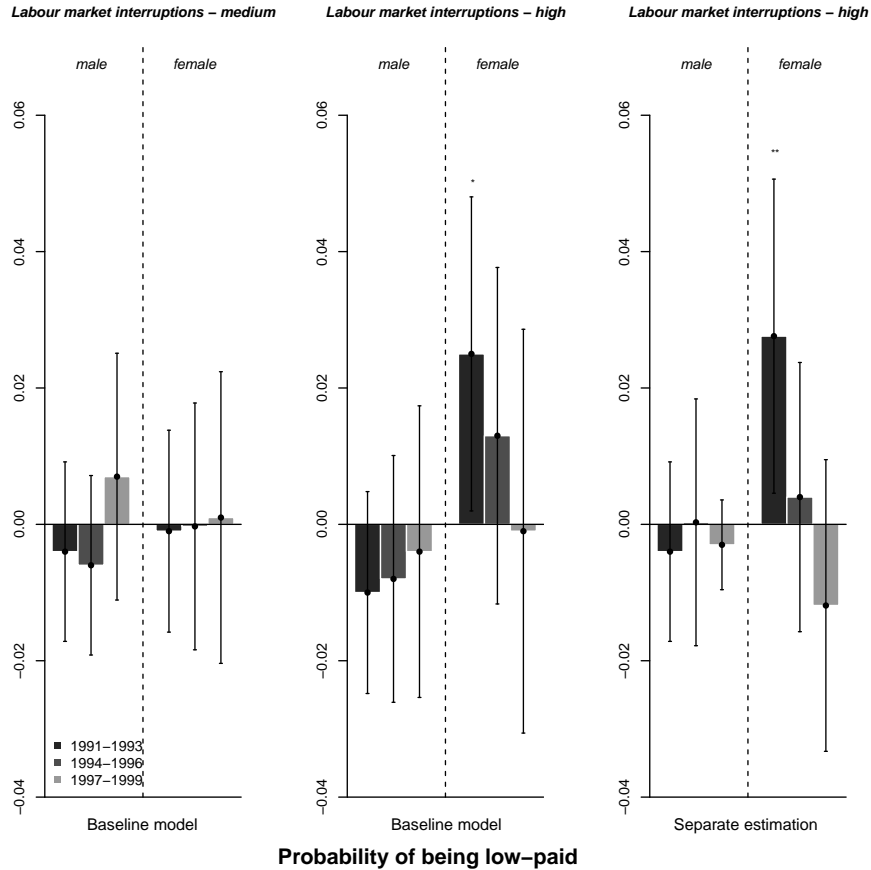
Source: BASiD 2007, weighted sample.

Notes: The figure shows marginal effects and the corresponding 90% confidence intervals of the probability of being low paid given the long-run persistence variables "number of years an individual has been below the first decile between 1980 and 1989" using a multinomial logit model with random effects and firm characteristics for the low-wage state distinguished by time and gender. The marginal effects are estimated by averaging over 100 Halton draws. Each regression for the male (female) sample has 16312 (16465) observations with 2165 (2267) individuals. Detailed results on estimated coefficients are shown in an online appendix.

Figure 5: Long-run Dynamics - Number of Low-wage Years

pay probability during the first three years if they experienced more than four years of cumulated labour market interruptions. The effect vanishes in the second and third sub-period. Repeating the regressions separately (right part of the figure) by sub-period shows similar results with a positive significant impact of 'high' GDR labour market interruptions for female workers for the first sub-period. Overall, there appears to be no effect for male workers and a temporary effect for females.

Heterogeneity across Educational Groups. Thus far, the effects for the long-run dynamics have been assumed to be equal across educational groups. To the extent that the established across-regime dependence in Figure 5 results from hu-



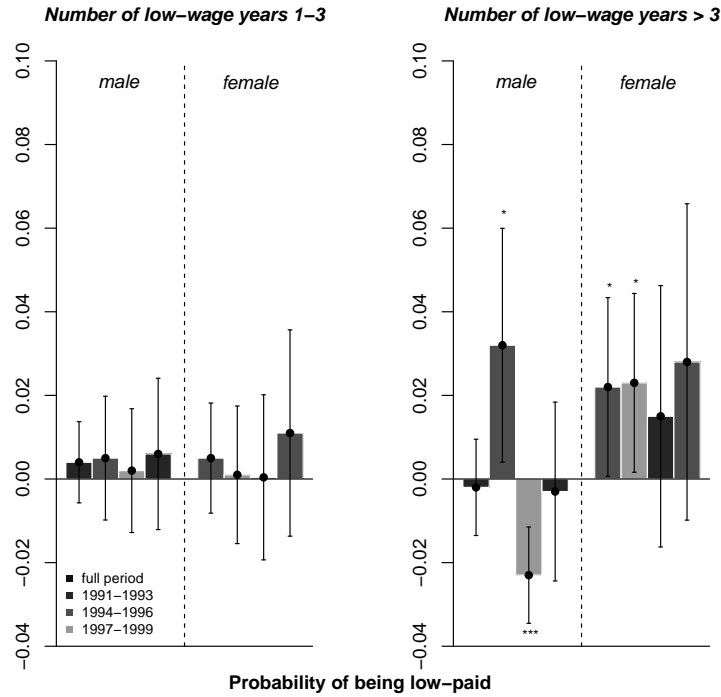
Source: BASiD 2007, weighted sample.

Notes: The figure shows marginal effects and the corresponding 90% confidence intervals of the probability of being out of employment given the long-run persistence variables "number of labour market interruptions between 1980 and 1989" using a multinomial logit model with random effects and firm characteristics for the low-wage state distinguished by time and gender. The marginal effects are estimated by averaging over 100 Halton draws. The reference category differs by gender. For males, the reference is zero labour market interruptions; medium: up to six months; high: more than six months. For females, the reference is zero to 12 months; medium: 12 to 48 months; high: more than 48 months. Each regression for the male (female) sample has 16312 (16465) observations with 2165 (2267) individuals. Detailed results on estimated coefficients are shown in the online appendix.

Figure 6: Long-run Dynamics - Number of Labour Market Interruptions

man capital devaluation, one might expect a stronger effect for high-skilled individuals. To address this issue, we exclude individuals without any level of education and repeat our regressions for the sub-population of medium and high-skilled individuals.¹⁸ Figure 7 illustrates the results for the whole sample period and its sub-periods (based on the baseline model). Compared to the results above, having one to three years of pre-unification low-wage experience has again no effect on the

¹⁸Repeating the regressions 5a and 5b of Table 6 shows that the covariances of the regime-specific random effects are similar once we focus on skilled individuals. For males the covariance is 0.006 and for females 0.285, albeit not significant. The variances of the FRG random effects are higher for both groups.



Source: BASiD 2007, weighted sample.

Notes: The figure shows marginal effects and the corresponding 90% confidence intervals of the probability of being low paid given the long-run persistence variables "number of years an individual has been below the first decile between 1980 and 1989" for skilled individuals using a multinomial logit model with random effects and firm characteristics for the low-wage state distinguished by time and gender. The marginal effects are estimated by averaging over 100 Halton draws. Each regression for the male (female) sample has 14327 (11778) observations with 1895 (1607) individuals. Detailed results on estimated coefficients are shown in the online appendix.

Figure 7: Long-run Dynamics for Skilled Individuals - Number of Low-wage Years

post-unification low-pay probability. However, the effects of having more than three years of pre-unification low-wage experience show the same pattern with different confidence intervals. While the formerly established significant estimates for female workers lose their precision, the picture for male workers becomes more pronounced especially for the first period (1991-1993) after Unification. Their low-pay probability increases by 3.3 per cent if the number of GDR low-wage years exceeds three. The effect, however, is not long lasting and turns out to be significantly negative during the second period. While these findings lend support to the hypothesis that genuine state dependence results from human capital devaluation, they also suggest that formerly low paid men have been able to catch up with their highly-paid counterparts within a relatively short period of time.

Structural Change at the Sectoral Level. The use of short-time work may be viewed as an indicator for a strong restructuring process. The associated mobility patterns along with the fact that short-time work may give rise to systematically lower earnings in our data may affect our estimates of state dependence. To assess the robustness of our results with respect to short-time work, we exclude the sectors that were particularly heavily affected by short-time work (manufacturing, agriculture, and water, energy and mining) from our estimations. The results are similar to the findings above. While there is slightly lower short-run low-pay state dependence after 1990 for male workers, female workers tend to have higher persistence rates. The initial conditions as well as the number of low-wage years prior to 1990 are again not significantly different from zero.

6 Conclusions

Focussing on the economic and political transformation in East Germany, the aim of this study was to investigate how one's pre-unification relative economic position determined one's fortunes after Unification. In answering this question, we attempt to improve our understanding about whether the transition from a centrally planned to a market economy offers fundamentally new perspectives for those who, in economic terms, were relatively deprived under the old regime. To address these issues, we use a novel administrative data set, the *BASiD* data set, focussing on individuals with a sufficiently long period of labour market experience during the pre-unification regime. The data provides an ideal basis for our analysis as it allows us, first, to identify individuals living in Eastern Germany before Unification and, second, as it enables us to track individual employment histories both before and after the fall of the Iron Curtain. To measure individuals' relative economic position, we exploit information on the incidence and duration of their position in the bottom part of the pre-unification wage distribution.

In our empirical analysis, particular emphasis is given to the extent of genuine low-wage state dependence during economic and political transition. We estimate

dynamic discrete choice models taking into account the initial condition problem and different specifications of unobserved heterogeneity and selection effects. First, we show that unobserved factors are the main drivers of low-wage persistence and that these unobservables are regime-dependent. Moreover, true state dependence in low-wage employment after Unification is about nine per cent for males and 15 per cent for females. This measure of persistence is rather constant across time, indicating a strong dynamic selection process after Unification. By measuring across-regime dependencies, our empirical strategy is twofold. We first exploit the modelling of the initial condition in 1990. We show that the initial condition is not significant indicating the exogeneity of the modeled stochastic process in terms of market-regime unobservables. This supports the exogenous initial value assumption indicating that the initial condition is uncorrelated with the unobserved individual-specific effect in the post-unification period. We then proceed by using information on individuals' pre-unification labour market outcomes and show that, consistent with theoretical considerations, economic dependencies across different political and economic regimes are only weakly present. An individual with more than three years of experience below the first decile during 1980 and 1989 has a roughly 2.5 percentage points higher probability of being in low-wage employment during the first three years after Unification. Across all skill groups, this effect is weak and becomes smaller for male and insignificant for female workers.

What drives these findings? As to across-regime dependence, signalling considerations and the stigma-effect channel lead us to expect no relationship between the pre- and post-unification relative wage position, if the heavily regulated pre-unification labour market had precluded any selection into low-wage jobs based on workers' true productivity. The structure of the Eastern German labour market before 1990 characterised by control of labour supply and demand and its central wage and price setting system indeed suggests that individuals' productivity should have been rather unconnected to their low-wage status. Given that general skills have been shown to be transferable to the post-unification labour market, an alternative explanation for the established, albeit weak across-regime persistence might stem

from a loss in general human capital. Especially for men this hypothesis is borne out by the estimates. The empirical results show that spending a reasonable amount of time at the bottom of the wage distribution before Unification is associated with a higher post-unification low-pay probability and that this effect is particularly pronounced for medium and high-skilled workers. However, there is also evidence that this effect was not long lasting. Taken together, especially for males our findings provide good news for those who used to be relatively deprived under the socialist regime. Those who were in the bottom of the socialist wage distribution are not more likely of being low paid after Unification than their highly paid counterparts and if there was any across regime state dependence it was weak and appears to have vanished within a relatively short period of time.

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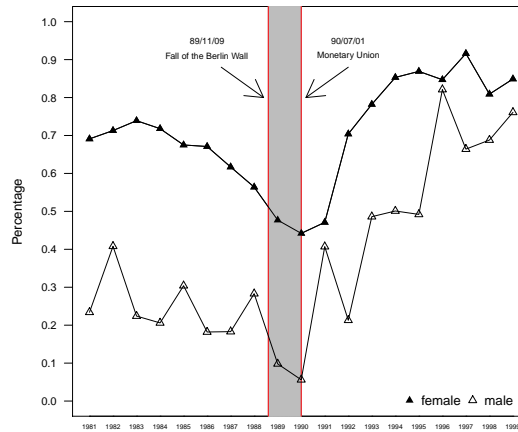
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Appendix

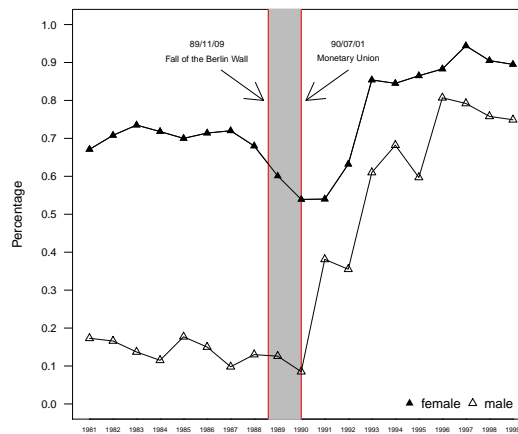
Appendix A: Aggregate State Dependence



Source: BASiD 2007, weighted sample.

Notes: The figure shows the difference in the probability of being in low-wage employment given low-wage and high-wage employment in $t-1$ using the pooled wage distribution. The first line marks the fall of the Berlin Wall. The second one marks Monetary Union.

Figure A.1: Aggregate state dependence, pooled wage distribution, by gender



Source: BASiD 2007, weighted sample.

Notes: The figure shows the difference in the probability of being in low-wage employment given low-wage and high-wage employment in $t-1$ and changing the low-wage threshold from the first decile before Unification to two-thirds of the median after Unification. The first line marks the fall of the Berlin Wall. The second one marks Monetary Union.

Figure A.2: Aggregate state dependence, changing low-wage threshold after Unification, by gender

Appendix B: Data Description

Table B.1: Description of individual employment history variables gained from the *Pension Register*

| Variable | Definition |
|---------------------------------------|---|
| Employment Status¹⁾ | |
| EMPLOYMENT | Employment spells include periods of employment subject to social security contributions and (after 1998) marginal employment. |
| UNEMPLOYMENT | Unemployment spells include periods of unemployment with and without transfer receipt (only FRG). ²⁾ |
| NON-EMPLOYMENT | Non-employment spells include periods of child raising, care giving as well as periods with missing information on the employment status. |
| ILLNESS | Illness spells include periods of long-term illness (FRG > 6 weeks; GDR > 4 weeks before 1984, no minimum restriction afterwards). |
| TRAINING | Training spells include periods of school or university attendance after the age of 16 and periods of training and apprenticeship. |
| # UN(NON)EMPLOYMENT_SPELLS | Number of un- or non-employment spells. An unemployment spell is counted as a new spell if the gap between a preceding unemployment spell exceeds four weeks. |

¹⁾ Note that the recorded pre-unification pension activity histories are less precise than the post-unification histories. The reason is that the transfer of the activities was mainly based on former GDR citizens' social security cards. These cards record the number of months of employment, illness and maternity leave during a particular year, but do not allow for tracking these spells on a monthly basis. As a result, compared to the pension spells after Unification, which provide exact monthly information on all pension relevant activities, information on the incidence of pre-unification employment, illness and maternity leave spells is available only on an annual basis.

²⁾ A spell of unemployment in the *Pension Register* requires individuals to be registered as unemployed *and* to obtain public transfers. The latter include benefits such as unemployment insurance, and - prior to 2005 - the means-tested social assistance and unemployment assistance benefits. After 2004, unemployment and social assistance were merged into one unified benefit, also known as 'unemployment benefit II' (ALG II). As the latter targets only employable individuals, a spell involving the receipt of ALG II automatically fulfills the requirements to be recorded as unemployed in the *Pension Register*. Prior to 2005, spells with social assistance benefits fulfill the above requirements only if individuals were registered as unemployed. Otherwise they are recorded as non-employment spells. As a consequence, the *Pension Register* does not permit a consistent definition of un- and non-employment prior to and after 2005.

Table B.2: Description of individual characteristics
gained from the *Pension and Employment Statistics Register*

| Variable/Categories | Definition |
|----------------------------|--|
| GDR-Spell | GDR spells are identified based on the regional origin (<i>Beitrittsgebiet</i>) of the pension contributions |
| Educational Status | |
| LOW-SKILLED | No degree or highschool degree (Reference category) |
| MEDIUM-SKILLED | Completed vocational training |
| HIGH-SKILLED | Technical college degree or university degree |
| Age | Age in years |
| Occupational Type | |
| WHITE-COLLAR | White-collar worker (Reference: blue-collar) |
| Seniority | |
| TENURE | Number of previous months at current employer. Employment interruptions at the same employer may not exceed 6 months - otherwise tenure is reset to zero after the employment interruption. |
| Earnings | |
| EARNINGS | Gross monthly earnings are retrieved from credit points to the German Pension Insurance. GDR credit points are divided by a factor as specified in Appendix 10 to the German Social Act (SGB VI). One credit point corresponds to the average of yearly earnings of all gainfully employed workers in (Western)Germany. Monthly earnings are thus obtained by multiplying monthly credit points with the average of earnings as documented in the Appendix 1 to the German Social Act (<i>SGB VI</i>). Credit points are reported up to the contribution limit of the German social security system. |

Table B.3: Definition of establishment characteristics
gained from the *Employment Statistics Register*

| Variable | Definition/Categories: |
|------------------------------|--|
| Establishment size | Size ≤ 20 (Reference category) $20 \leq \text{Size} < 50$ $50 \leq \text{Size} < 200$ $200 \leq \text{Size} < 1000$ Size ≥ 1000 |
| Workforce composition | Share of employees younger than 30 years Share of employees older than 50 years Share of low-skilled employees Share of female employees |
| Sector affiliation | Agriculture/Forestry (Reference category) Mining and manufacturing Energy/Water supplies Construction Wholesale and retail trade Transport and communication Financial intermediation Other service activities Public administration |

Appendix C: The GDR Pension Formula

Table C.1: Calculation of GDR Pensions

| Creditable years for pension insurance | Minimum amount | Fixed amount (earnings independent) | Maximum variable amount ¹⁾ | Maximum amount |
|---|-------------------|--|--|-------------------|
| | (1) | (2) | (3) | (4) |
| $t < 15$ | 330 | 170 | 90 | 330 |
| $15 \leq t < 20$ | 340 | 170 | 120 | 340 |
| $20 \leq t < 25$ | 350 | 170 | 150 | 350 |
| $25 \leq t < 30$ | 370 | 180 | 180 | 370 |
| $30 \leq t < 35$ | 390 | 190 | 210 | 400 |
| $35 \leq t < 40$ | 410 | 200 | 240 | 440 |
| $40 \leq t < 45$ | 430 | 210 | 270 | 480 |
| $45 \leq t$ | 470 | 210 | 300 | 510 |

Source: Rosenschon (1990). ¹⁾The variable earnings dependent amount is one per cent of average monthly earnings multiplied by the number of creditable years.

Earnings threshold above which earnings increase pension entitlements calculated as $100 \cdot (\text{Col. (1)} - \text{Col. (2)}) / (\# \text{Creditable years})$.

GDR pension entitlements were only to a limited extent earnings dependent. Column (2) in Table C.1 shows that in 1989, pension rules stipulated an earnings independent amount, which varied between 170 and 210 Mark (depending on the number of contribution years). The additional earnings dependent amount corresponded to one per cent of average monthly earnings multiplied by the number of years creditable for the pension insurance. However, this variable component only applied up to the contribution limit of 600 Mark. Thus, for an individual who e.g. paid pension contributions for 40 years, monthly pensions were capped at a maximum amount of 450 Mark (210 plus 40·6 Mark). The GDR pension system also guaranteed a minimum pension amount that varied between 330 and 470 Mark (again depending on the number of creditable years) shown in Column (1). For a worker who earned on average 300 Mark per months during 40 years, this implied that her pension earnings did not correspond to the amount according to the pension formula ($330 = 210$ plus $40 \cdot 3$ Mark), but was rather fixed at the minimum pension level of 430 Mark. Figure C.1 shows the earnings threshold above which earnings increased pension entitlements. The threshold depends on the number of creditable years for the pension insurance and is calculated as $100 \cdot (\text{Column (1)} - \text{Column (2)}) / (\text{Number of creditable years})$. For example, the minimum pension level implied that for a worker with 40 creditable years, the threshold of monthly earnings above which earnings raised pension entitlements was 550 Mark. I.e., within the range of

550 and 600 Mark additional earnings increased monthly pensions by an amount of up to 20 Mark (corresponding to about five per cent of the minimum pension level). In contrast, for a worker with 44 creditable years, the threshold of monthly earnings above which higher earnings led to higher pensions was already reached at 500 Mark, giving rise to a potential increase in pension entitlements of more than ten per cent of the minimum pension level (44 Mark). According to the German Ministry of Labour and Social Affairs, in 2012 the average number of creditable years was 44.6 for Eastern German males and to 39.1 years for females.

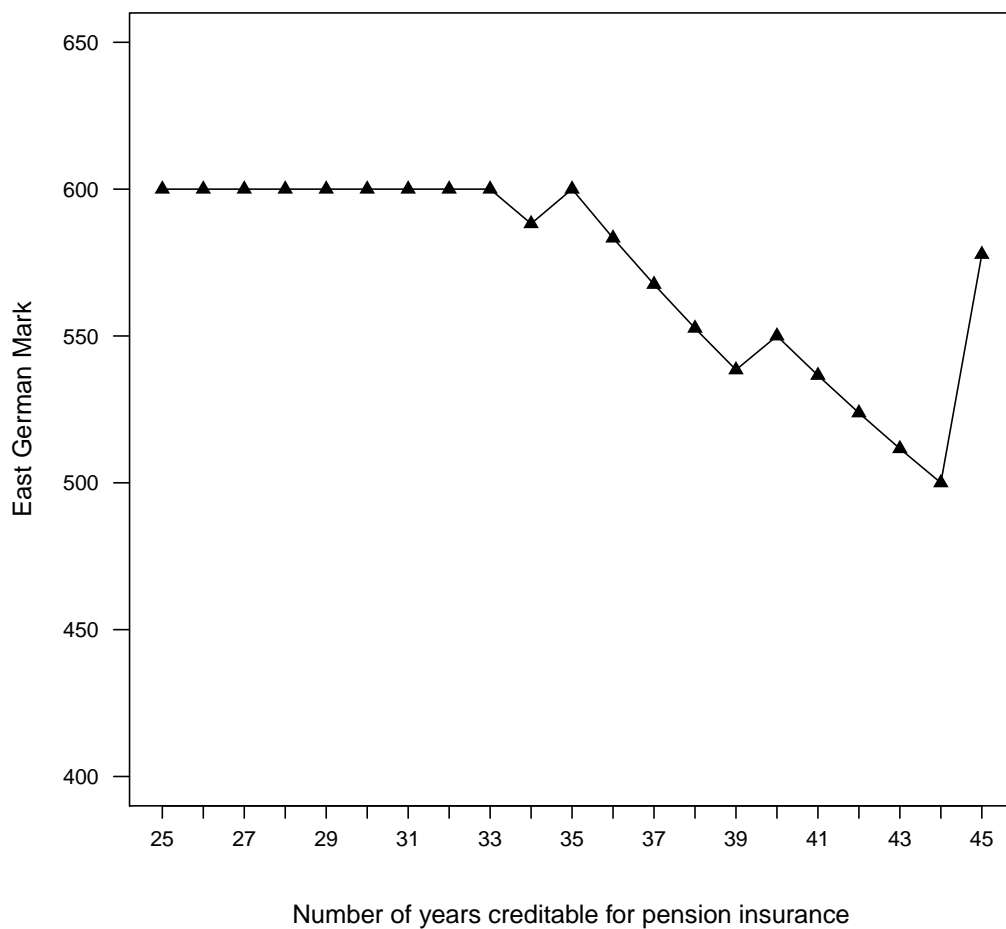


Figure C.1: Earnings threshold above which earnings increase pension entitlements

Appendix D: Robustness Results for Online Publication

Pooled Wage Distribution (Section 5.1)

Given that the male low-wage threshold comes fairly close to the censoring point we perform several robustness checks. To check the robustness of our results with respect to the chosen threshold, we re-ran our specifications based on the first decile of the pooled wage distribution as shown in Figure D.1. Note that from a welfare

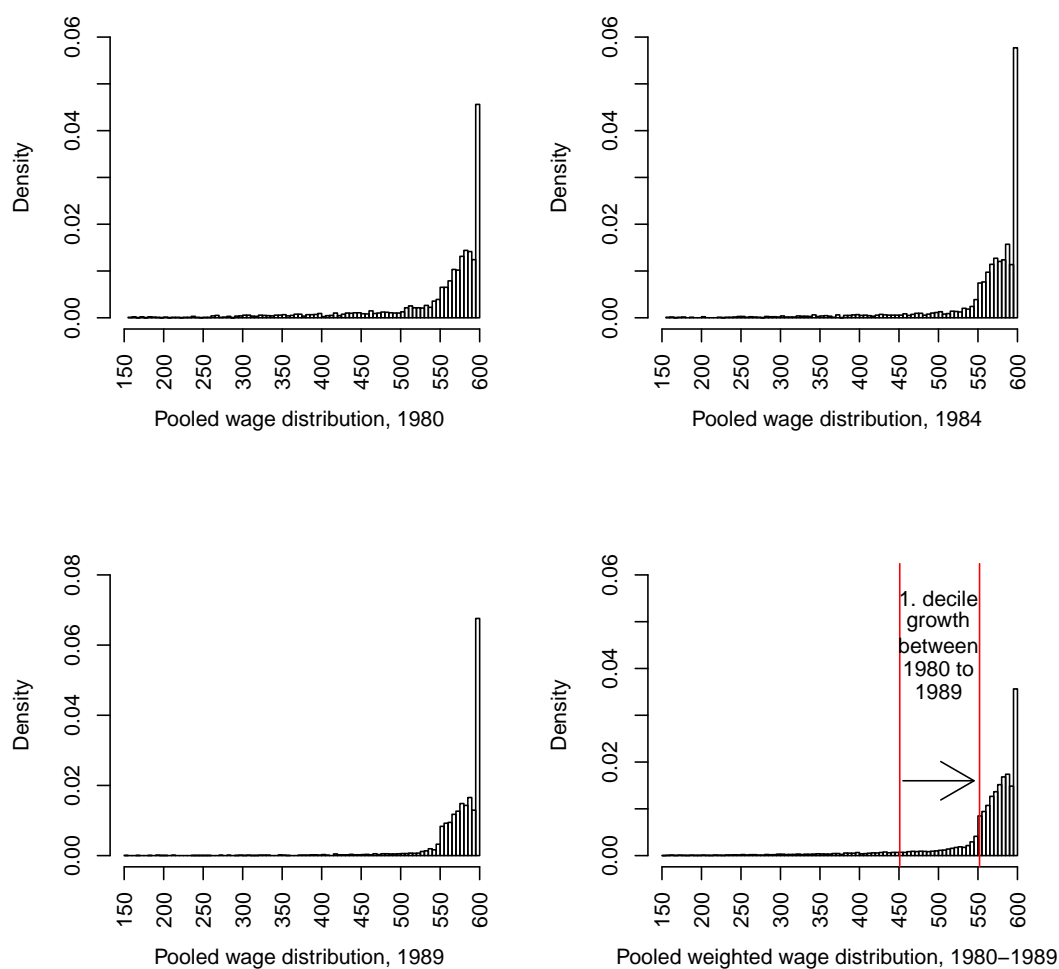


Figure D.1: Distribution of Pooled Wages between 1980-1989

perspective this is a reasonable strategy, as an individual's relative economic position is likely to be determined by the overall working-age population rather than by (arguably narrowly defined) specific peers. While the pooled decile has the advan-

tage that it is farther away from the censoring limit and thereby reduces the concern of misclassifying males as falling below the censoring limit, it comes at the expense of a considerably lower fraction of male workers obtaining a low wage (about five per cent). Table D.1 shows the average partial effects for the pooled wage distribution. Compared to Table 6, the APE for male workers is lower after Unification whereas

Table D.1: Average Partial Effects of Random Effect Probit Models using the Pooled Wage Distribution, by Gender and Period

| | β constant α_i constant | | β variable α_i constant | | β constant α_i variable | | β variable α_i variable | | β variable α_i variable | |
|---|---|---------|---|---------|---|---------|---|---------|---|---------|
| | (1a) | (1b) | (2a) | (2b) | (3a) | (3b) | (4a) | (4b) | (5a) | (5b) |
| | Males | Females | Males | Females | Males | Females | Males | Females | Males | Females |
| $P(y_t = 1 y_{t-1} = 1) - P(y_t = 1 y_{t-1} = 0)$ | | | | | | | | | | |
| APE_{GDR} | 4.3 | 41.7 | 4.0 | 41.3 | 7.2 | 43.6 | 7.3 | 42.6 | 7.5 | 42.4 |
| APE_{FRG} | 20.4 | 51.5 | 19.0 | 47.3 | 30.0 | 54.6 | 28.4 | 50.9 | 23.8 | 48.0 |
| <i>Random Effects</i> | | | | | | | | | | |
| σ^2 | 0.238 | 0.206 | 0.246 | 0.222 | | | | | | |
| σ_{GDR}^2 | | | | | 0.371 | 0.413 | 0.310 | 0.367 | 0.264 | 0.370 |
| σ_{FRG}^2 | | | | | 0.436 | 0.539 | 0.488 | 0.597 | 0.592 | 0.713 |
| $\sigma_{GDR,FRG}$ | | | | | | | | | 0.171 [†] | 0.183 |

Source: BASiD 2007, weighted sample.

Notes: The table reports the results from dynamic discrete choice probit models on the probability of being low paid. $y_t = 1$ denotes the low-wage status. All estimated average partial effects and variance estimates ([†]apart from the covariance estimate for male workers with a significance level of 10%) are significant at the 5% level using 100 bootstrapped replications.

the variance estimate of unobserved heterogeneity increases slightly. For females, both APE's are higher with a slightly lower variance of the unobserved heterogeneity terms. This is due to the fact that a lower fraction of male workers are considered as low-wage employees while the fraction is higher among females. The covariance estimate for the regime-specific random effects are again highly significant for females while for male workers the estimate is weakly significant at the 10% level. However, the overall pattern of results obtained earlier are confirmed.

Moreover, given the fundamental changes the wage distribution underwent during and after economic transition, one might argue that interpreting an escape from out of the first decile into the upper deciles as a transition into the "high-wage sector" might be too restrictive after Unification. For this reason and to adopt a more conventional definition of low pay, we re-run our specifications by defining the low-wage threshold after 1989 as two thirds of the median wage. The quantitative results do not change compared to Table 5 and Table 6.

Low-Wage Threshold: First Decile (Section 5.2)

In section 5.2, we used the conventional threshold of two-thirds of the median. This section provides robustness checks by using the first decile after Unification. The upper panel of Table D.2 shows the main results using the baseline model (full time period with interaction with sub-periods). The lower panel shows the results by estimating the models separately for the sub-periods. The across-regime

Table D.2: Marginal Effects of Multinomial Logit Models with Random Effects, Males, Low-Wage Threshold: 1st Decile

| | 1991-1993 | | 1994-1996 | | 1997-1999 | |
|---------------------------------|---------------------|-------------------|----------------------|--------------------|---------------------|---------------------|
| | <i>low-wage</i> | <i>non-empl.</i> | <i>low-wage</i> | <i>non-empl.</i> | <i>low-wage</i> | <i>non-empl.</i> |
| Baseline Model | | | | | | |
| <i>Low – wage_{t-1}</i> | 0.077*** (0.008) | 0.004 (0.023) | 0.172*** (0.032) | 0.043 (0.029) | 0.275*** (0.039) | 0.018 (0.051) |
| <i>Initial condition</i> | | | | | | |
| <i>Low – wage</i> | 0.015 (0.014) | 0.013 (0.024) | 0.025 (0.016) | 0.003 (0.043) | 0.011 (0.014) | 0.014 (0.023) |
| <i># Low-wage years</i> | | | | | | |
| 1 – 3 | 0.004 (0.008) | -0.002 (0.016) | 0.001 (0.008) | -0.008 (0.018) | 0.009 (0.011) | -0.008 (0.019) |
| > 3 | 0.011 (0.018) | 0.013 (0.014) | -0.030*** (0.011) | 0.008 (0.013) | -0.012 (0.019) | 0.003 (0.007) |
| <i>obs. (total)</i> | | | 16312 | | | |
| <i>LogLik</i> | | | -5665.6 | | | |
| Separate Estimation | | | | | | |
| <i>Low – wage_{t-1}</i> | 0.149*** (0.021) | 0.098 (0.079) | 0.100*** (0.026) | 0.001 (0.021) | 0.096*** (0.028) | -0.047 (0.033) |
| <i>Initial condition</i> | | | | | | |
| <i>Low – wage</i> | 0.000 (0.010) | 0.019* (0.010) | 0.206*** (0.059) | 0.024** (0.015) | 0.278*** (0.074) | 0.094*** (0.033) |
| <i># Low-wage years</i> | | | | | | |
| 1 – 3 | 0.004 (0.007) | -0.002 (0.014) | -0.003 (0.008) | 0.001 (0.012) | 0.002 (0.010) | -0.003 (0.019) |
| > 3 | 0.010 (0.016) | -0.001 (0.012) | -0.033*** (0.013) | 0.027** (0.013) | -0.005 (0.021) | 0.013 (0.019) |
| <i>obs. (total)</i> | 6252 | | 5789 | | 5469 | |
| <i>LogLik</i> | -1788.7 | | -1945.8 | | -2087.1 | |

Source: BASiD 2007, weighted sample.

Notes: †The reference category differs by gender. For males, the reference is zero labour market interruptions; medium: up to six months; high: more than six months. For females, the reference is zero to 12 months; medium: 12 to 48 months; high: more than 48 months. Robust standard errors are in parentheses. All estimations contain a constant, the specified Mundlak-Chamberlain device, and control for missings in the education variable. Control variables include age dummies, education dummies, year dummies, regional dummies and experience dummies. The out of work equation includes lagged dummy variables of white-collar, skilled and simple occupation. The low-wage threshold is the first decile of the wage distribution. The average marginal effects are estimated by averaging over the average marginal effects of 100 Halton draws. Asterisks next to coefficients indicate significance levels as follows: *** 1%, ** 5%, * 10%.

effects are again similar and not significant for male workers during the first three

years after Unification. The short-run results ($Low - wage_{t-1}$) differ strongly which reflects the difference in the chosen threshold. While in the baseline model state dependence is increasing, the results from the separate estimations even point to a slightly decreasing tendency (not significant). Given that the initial condition becomes more important (lower panel), this points to a dynamic selection process in the baseline model. The results for female workers are also in line with the findings

Table D.3: Marginal Effects of Multinomial Logit Models with Random Effects, Females, Low-Wage Threshold: 1st Decile

| | 1991-1993 | | 1994-1996 | | 1997-1999 | |
|---------------------------------|---------------------|---------------------|---------------------|---------------------|---------------------|---------------------|
| | <i>low-wage</i> | <i>non-empl.</i> | <i>low-wage</i> | <i>non-empl.</i> | <i>low-wage</i> | <i>non-empl.</i> |
| Baseline Model | | | | | | |
| <i>Low - wage_{t-1}</i> | 0.037*** (0.006) | 0.029** (0.013) | 0.120*** (0.020) | 0.107*** (0.025) | 0.147*** (0.028) | 0.111** (0.053) |
| <i>Initial condition</i> | | | | | | |
| <i>Low - wage</i> | 0.023** (0.010) | 0.009 (0.019) | 0.009 (0.011) | 0.011 (0.039) | 0.008 (0.012) | -0.038** (0.017) |
| <i># Low-wage years</i> | | | | | | |
| 1 - 3 | 0.006 (0.007) | 0.014 (0.012) | 0.004 (0.009) | 0.007 (0.019) | -0.004 (0.009) | -0.031 (0.020) |
| > 3 | 0.026** (0.012) | -0.004 (0.015) | 0.015 (0.012) | 0.014 (0.018) | 0.016 (0.015) | 0.024*** (0.005) |
| <i>obs. (total)</i> | 6593 | | 5944 | | 5418 | |
| <i>LogLik</i> | -1755.0 | | -2060.5 | | -2023.3 | |
| Separate Estimation | | | | | | |
| <i>Low - wage_{t-1}</i> | 0.074*** (0.016) | 0.039 (0.028) | 0.045** (0.021) | 0.026 (0.019) | 0.041** (0.019) | 0.004 (0.031) |
| <i>Initial condition</i> | | | | | | |
| <i>Low - wage</i> | 0.025* (0.014) | 0.034*** (0.012) | 0.180*** (0.076) | 0.026* (0.015) | 0.190*** (0.057) | 0.119*** (0.029) |
| <i># Low-wage years</i> | | | | | | |
| 1 - 3 | 0.009 (0.008) | 0.029** (0.015) | -0.001 (0.002) | 0.007 (0.014) | -0.001 (0.007) | -0.019 (0.018) |
| > 3 | 0.032*** (0.013) | -0.004 (0.010) | 0.019 (0.014) | 0.019* (0.010) | 0.019 (0.012) | 0.040*** (0.014) |

Source: BASiD 2007, weighted sample.

Notes: †The reference category differs by gender. For males, the reference is zero labour market interruptions; medium: up to six months; high: more than six months. For females, the reference is zero to 12 months; medium: 12 to 48 months; high: more than 48 months. Robust standard errors are in parentheses. All estimations contain a constant, the specified Mundlak-Chamberlain device, and control for missings in the education variable. Control variables include age dummies, education dummies, year dummies, regional dummies and experience dummies. The out of work equation includes lagged dummy variables of white-collar, skilled and simple occupation. The low-wage threshold is the first decile of the wage distribution. The average marginal effects are estimated by averaging over the average marginal effects of 100 Halton draws. Asterisks next to coefficients indicate significance levels as follows: *** 1%, ** 5%, * 10%.

above. However, the initial condition is weakly significant at the 10%-level. The

number of years below the first decile before 1990 is again significant during the first three years after Unification and decreases in magnitude during the subsequent years.