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ABSTRACT

The Impact of Labor Market Entry Conditions on Initial Job Assignment, Human Capital Accumulation, and Wages*

We estimate the effects of labor market entry conditions on wages for male individuals first entering the Austrian labor market between 1978 and 2000. We find a large negative effect of unfavorable entry conditions on starting wages as well as a sizeable negative long-run effect. Specifically, we estimate that a one percentage point increase in the initial local unemployment rate is associated with an approximate shortfall in lifetime earnings of 6.5%. We also show that bad entry conditions are associated with lower quality of a worker's first job and that initial wage shortfalls associated with bad entry conditions only partially evaporate upon involuntary job change. These and additional findings support the view that initial job assignment, in combination with accumulation of occupation or industry-specific human capital while on this first job, plays a key role in generating the observed wage persistencies.

JEL Classification: E3, J2, J3, J6, M5

Keywords: initial labor market conditions, endogenous labor market entry, initial job assignment, specific human capital

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1 Introduction

The recent economic crisis has renewed academic interest in the potential impact of business cycle fluctuations on labor markets (e.g. Elsby *et al.*, 2010). However, while labor economists have studied the short-run association between local labor market conditions and real wages extensively for quite some time (e.g. Blanchflower and Oswald, 1990), longer-run effects of business cycle fluctuations on individuals' wages have only more recently caught the attention of empirical research.[1] Clearly, in the longer run, even small initial wage shortfalls may eventually accrue to substantial overall losses in lifetime earnings if initial wage losses resulting from poor entry conditions persist.[2]

Indeed, recent empirical evidence suggests that substantial losses in lifetime earnings result from entering the labor market during an economic downturn, as opposed to entering during an expansion. Oreopoulos *et al.* (2006, 2008) explore the effects of entering the labor market during a recession on individuals' earnings, using data on Canadian college graduates who entered the labor market between 1982 and 1999. They find a substantial initial wage penalty of about 9% that only fades to zero after the first decade of a worker's career. A similar result is reported in Kahn (2010), who focuses on male college graduates in the United States graduating sometime between 1979 and 1988. She finds that the group graduating in the worst economic situation incurs a wage loss of up to 13% each year, relative to those graduating in the best initial conditions, and that this initial wage loss persists over the first 20 years of workers' labor market career. Similar results are reported in Oyer (2006), who shows that PhD students in economics are considerably more likely to get a position at one of the top universities in the United States if they graduate in times when the demand for economists is high. In a related study, he finds that those MBA students who complete their training during a recession suffer from negative effects on wages (Oyer, 2008). In both studies, the long-term

[1] Most studies estimating the short-run association between fluctuations in local unemployment rates and wages find that wages vary negatively with local unemployment. This negative association is a very robust empirical pattern; it has been shown to exist for a wide range of different countries, using very different sources of data and diverse empirical specifications. See Nijkamp *et al.* (2005) for a comprehensive survey of this literature.

[2] Previous research has shown that the early years in a worker's labor market career are of special importance (Gardecki and Neumark, 1998; Neumark, 2002). In terms of wages, Murphy and Welch (1990) estimate that almost 80% of all (i.e. lifetime) wage increases accrue within the first ten years of labor market experience. Moreover, movements across jobs are considerably more likely at the beginning of a worker's career than later on (Topel and Ward, 1992).

effects on income appear to be caused by the fact that diverse employers hire workers entering the labor market under different conditions, therefore giving them access to distinct jobs. The entry job is thus significant for the future career, and this appears to be of particular importance for highly educated individuals for whom the transition in and out of attractive positions is very low.[3] Mansour (2009) also focuses on college graduates in the US and again finds negative and persistent wage effects from entry into the labor force during a recession. Moreover, he shows that failure to account for endogenous sample composition underestimates both the immediate wage effects as well as the persistence of wage effects resulting from initial labor market shocks.

Empirical evidence for countries outside the US and Canada yields broadly similar findings. Kwon *et al.* (2010) mainly use data from the Swedish labor market.[4] They find that workers who enter the labor market during a boom are not only paid higher wages, but that they are also promoted more quickly to higher ranks than those who enter during an economic downturn. Stevens (2007) finds significant negative, albeit small effects of initial conditions on wages in Germany (much smaller than those found in the US and Canada). In contrast to all other studies, however, she finds that wage losses from poor entry conditions do not fade away, but actually increase over time.[5] The available empirical evidence also underlines the fact that negative wage effects of initial labor market conditions are not confined to highly skilled workers. Genda *et al.* (2010) focus on a separate comparison between more and less educated men in Japan and the United States in the effects of initial conditions. They find negative effects of initial conditions for more highly skilled workers in both countries. However, they only find negative wage effects for less skilled workers in Japan. They argue that the specific

[3] One important concern regarding the validity of these results is that schooling and first entry into the labor force may be endogenous both because individuals may choose to stay in school or continue further training when faced with high unemployment and low starting wages. Indeed, several studies find that enrollment rates are high when unemployment is high and the opportunity costs of schooling are low (e.g. Clark, 2009). In line with these findings, both Kahn (2010) and Oreopoulos *et al.* (2006, 2008) find the duration of schooling to be endogenous. Both tackle the endogeneity problem by instrumenting the unemployment rate at the time of labor market entry with either the prevailing unemployment rate at a lower age or that in the predicted year of graduation. Mansour (2009) presents direct evidence on sample selection over the business cycle based on AFQT scores.

[4] Studies for European countries have mainly focused on the long-run effects of initial conditions on employment rather than wages up until now (e.g. Burgess *et al.*, 2003; Raaum and Roed, 2006).

[5] A similar analysis of wage effects for firm entry cohorts in the German manufacturing sector is given in von Wachter and Bender (2008). However, their analysis is not confined to new labor market entrants but covers workers of all experience levels; their results are therefore not directly comparable to the other studies mentioned.

hiring system and employment protection drive the persistence of the effects for Japanese, while the market for less skilled workers in the United States may indeed be quite close to a competitive market. Consistent with this finding, Kondo (2008) reports that the initial effect of entering the labor market during a recession on wages is less persistent for less skilled workers and for workers with weak labor market attachment in the US.

In this paper, we present estimates of the long-run effects of business cycle fluctuations on young males' wage profiles in the Austrian labor market and derive an empirical estimate of the associated loss in lifetime earnings due to entering the labor force during a recession, as opposed to entry during average aggregate conditions or during a boom. We do so using social security records from Austria that contain detailed individual earnings and employment histories for the universe of private-sector employees from 1972 until 2005. We complement the available empirical evidence on the long-run wage effects of labor market entry conditions with an analysis for Austria, a labor market characterized by a high level of employment protection and a centralized wage bargaining structure. We focus on low and medium-skilled workers, while most of the mentioned studies focus on higher or even highest-skilled workers (in terms of formal education). Note, however, that workers in the Austrian labor market typically have some, potentially very specialized, vocational training.[6] Moreover, and in contrast to Stevens (2007) and Kwon *et al.* (2010) – the only other studies focusing on European labor markets – we take endogenous labor market entry over the business cycle into account in the empirical analysis. Indeed, our analysis shows that there is substantial positive selection of workers in times of high initial unemployment and that, as a consequence, dealing with sample selection is in fact crucial for our conclusions regarding the long-term wage effects of initial labor market conditions.

In the second part of the analysis, we try to discriminate between different mechanisms that can potentially explain persistent cohort effects in wages resulting from differences in labor market entry conditions. Specifically, initial job or task assignment may be important in the longer run if employers assign otherwise identical workers to lower quality jobs or tasks in recession, and if jobs or tasks offer different opportunities for the accrual of human capital (e.g.

[6] A significant part of the initial vocational training in Austria is provided by dual apprenticeship training schemes, i.e. practical training provided by firms coupled with part-time compulsory attendance at a vocational school. Apprenticeships last from two to four years, depending on occupation. Full-time vocational and technical schools provide an important alternative to apprenticeship training, and also last up to four years. Details are available from the report by the Federal Ministry for Education, the Arts and Culture (2008).

Gibbons and Waldman, 2006). Alternatively, workers' initial job or task assignment may have long lasting effects on wages if workers accumulate human capital while on the job that is not fully transferable to other jobs or tasks because it is specific to a worker's task, occupation, or industry. We therefore complement our main analysis with an analysis of the effects of initial labor market conditions on the quality of a worker's first job against this background.[7] Third, we study whether and the extent to which wage changes following involuntary first job moves (i.e. involuntary job changes away from a worker's first employer), as well as wages with the old and the new employers depend on initial labor market conditions. The basic idea in this part of the analysis is that the elasticity of the wage of the new (second) employer with respect to initial labor market conditions informs us about the importance of differences in the accumulation of human capital while on the job and the importance of explanations that do not rest on underlying differences in productivity such as downwardly rigid wage contracts. Finally, we complement our empirical analysis with results for selected subsamples, such as blue and white-collar workers, as well as for additional outcome variables, such as different measures of job mobility.

Our key findings are the following. First, we find a substantial negative effect of unfavorable labor market conditions on starting wages and that this initial negative wage effect is persistent, resulting in sizeable negative effects of poor entry conditions on lifetime earnings. Specifically, our estimates imply that a one percentage point increase in the initial local unemployment rate is associated with a loss in lifetime earnings of about 6.5%. A second main finding is that endogenous labor market entry is a crucial issue, suggesting that there are substantial fluctuations in the recruitment of new workers over the business cycle. Third, we find that labor market entry in times of high unemployment is associated with a lower average quality of a worker's initial job. Fourth, we find that wage losses from having entered the labor market during a recession only partially evaporate upon involuntary termination of first employment. These findings strongly support the view that initial job and/or task assignment is important in explaining the observed wage persistencies resulting from labor market entry conditions, and our additional results for different subsamples and for alternative outcomes are also in line with this view. On the other hand, a share of workers' initial wage losses do vanish upon involuntary

[7]Consistent with this line of argument, Kwon *et al.* (2010) find that workers who enter the labor market during a boom are promoted more quickly and to higher ranks than those who enter during a recession, and Mansour (2009) shows that workers entering in a recession are initially assigned to lower paying jobs.

first job change, and thus mechanisms unrelated to human capital, such as downwardly rigid wage contracts or temporary rents, are also important in explaining the long run wage effects resulting from variation in initial labor market conditions.

The remainder of this paper is organized as follows. The next section presents our data source, details the sample selection process, and discusses the construction of our key measures. Section 3 presents the econometric approach for estimating long-run wage effects of initial labor market conditions along with our main results. We then study the impact of initial labor market conditions on the quality of initial job assignment in section 4. Section 5 studies how changes in earnings resulting from involuntary job loss depend on initial labor market conditions. In section 6 we present complementary additional results for several subsamples and for different outcomes than wages. Section 7 concludes.

2 Data and Sample

2.1 Data Source

We use individual-level social security records from the Austrian Social Security Database (ASSD), a data source described in more detail in Zweimüller *et al.* (2009). The ASSD sample we use basically covers the universe of Austrian private sector workers from January 1972 until December 2005. The data contain complete and precise information about individuals' annual earnings and daily employment histories. The data are therefore ideally suited for studying the impact of labor market shocks on long-run wage profiles because they allow us to construct individual wage profiles for a large number of labor market entrants over a relatively long period of time.

2.2 Key Measures

Our dependent variable for most of the analysis is the real daily wage, i.e. the real wage per actual day of work, adjusted to 2007 prices. Wages are deflated with the consumer price index and include additional/special payments such as a 13th month's salary or holiday pay.[8] Real

[8]The Austrian Central Social Security Administration collects these data with the purpose of administering and calculating entitlements to old-age pension benefits. For this reason, the ASSD includes precise and comprehensive information on annual earnings and daily employment histories. However, contributions to the old-age pension system are capped because old-age pension benefits are limited to a maximum level. As a consequence, annual earnings are only recorded up to the threshold which guarantees the maximum benefit

daily wages are computed as the average earnings over all employers in a given year. That is, we first sum total annual earnings over all employers for any individual. We then divide overall earnings by the total number of days worked in a given year, also summed over all employers for a given individual and taking overlapping employment spells into account.

The regressor of main interest is the annual male unemployment rate, our measure for external labor market conditions at the time individuals first enter into the labor force. We computed annual male unemployment rates from the individual-level employment histories contained in the ASSD raw data. This procedure has the advantage that we can calculate unemployment rates back until 1972 (compared to published statistics, which reach back until 1978 only) and at different levels of cross-sectional aggregation.[9] Unless noted otherwise, we use the male unemployment rate for all workers aged between 16 and 65 at the state level as our main regressor.[10]

2.3 Sample Selection

Mainly for conceptual reasons, but also due to some data limitations, we do not work with the universe of all labor market entrants but only with a specifically selected sample. First, we restrict our attention to male entrants only. On the one hand, female labor supply behavior over the life cycle is much more difficult to model than male labor supply. On the other hand, we believe that the fact that most men work full-time allows us to largely circumvent the problem that the ASSD does not contain information on working hours. Second, we select those workers who start their first regular employment spell sometime between 1978 and 2000, allowing us to observe at least five additional years of earnings for each worker because the data run until the end of 2005 (see also appendix A). As a final restriction, we focus on

level ("Höchstbemessungsgrundlage", HBGr). Similarly, there is a base threshold below which no (otherwise mandatory) social security payments accrue ("Geringfügigkeitsgrenze", GfGr). The two censoring points vary over time in real terms: The lower censoring point increased from about 14 € in 1978 to about 26 € in 2005 (per day worked); the upper censoring point increased from about 78 € to 126 € per workday over the same period of time.

[9] We decided to extract yearly male unemployment rates for the age groups 16 to 65 and 16 to 25, both at the state ("Bundesland") level and at the common classification of territorial units for statistics (NUTS) level. At the NUTS level, we use the most disaggregated level available (NUTS-3), which corresponds to one or more political districts in Austria. There are total of 9 different states and 35 different NUTS-3 regions in Austria. Yearly unemployment rates are within-year averages of monthly unemployment rates.

[10] It is not obvious whether the youth unemployment rate would be preferable to the overall unemployment rate because the youth unemployment rate may suffer from endogeneity bias. Also, because Austria's youth unemployment rate is very low compared to most other countries (e.g. Breen, 2005), the choice between the two is ultimately not important (as we will actually show later on; see table 3).

workers aged between 16 and 21 at the time they first enter the labor force (i.e. start their first regular employment spell).[11] This restriction effectively serves as a restriction on individuals' schooling duration (see also appendix A).[12] Essentially, this restriction excludes individuals with higher education (most importantly, individuals with a university degree), but it should include all or most individuals with an apprenticeship training or an education of similar length and scope, such as full-time vocational school.[13] Our final sample thus consists of male low- and medium-skilled labor market entrants who started their first regular employment between 1978 and 2000. We can observe these workers' full labor market career from the year they first enter into the labor force until the year 2005.

2.4 Sample Description

Because we can follow all individuals from the year of their first regular employment until the end of the data in the year 2005, the resulting data set would have been too large from a practical point of view. In the following, we therefore work with a 30% random sample of all labor market entrants aged between 16 and 21 when first entering into the labor force. This sample contains 220,214 unique individuals and about 3.35 million individual wage observations (i.e. observations at the level of individual×year).

Table 1

Table 1 shows descriptive statistics for our final analysis sample. The first panel shows individual-level characteristics. The average labor market entrant in our sample is about 19

[11]Because the ASSD does not contain a comprehensive measure for schooling, we also use age at entry into the labor force as proxy for education in the regressions below. To mitigate potential collinearity with year of birth and year of entry, we use a slightly different variable as proxy in the regressions. Specifically, we use the smaller of age at start of first regular employment and age at start of first registered unemployment spell.

[12]We focus on a sample of labor market entrants so that each entry cohort is balanced with respect to the potential range of schooling, meaning that the potential range of schooling (i.e. age at entry) is the same for each year of entry considered in the analysis. In order to determine the start of an individual's first job within the full range of "education‘' in each year, we had to restrict the sample period to the years 1978-2000. In the year 1977, for example, we cannot exactly determine the first entry into the labor market for an individual aged 21 because this individual might have already entered the labor market with age 16, i.e. in the year 1971. In this case, his entry is not observed in the data simply because the data do not start before 1972.

[13]Several arguments motivate the restriction on schooling. First, the timing of first labor market entry, and thus the duration of schooling, may be endogenous. However, less skilled workers are presumably less likely to manipulate the duration of schooling. Also, unobserved heterogeneity resulting from, say, unobserved differences in inherent ability, is arguably a more urgent problem for higher-skilled workers. Moreover, we think that our proxy for schooling works best for less-educated workers. Finally, only including less-skilled workers in the sample is an effective way of dealing with right-censored wages (see also appendix A).

years old when starting his first regular employment spell, and he holds his first job for almost three years. The average age at the start of the first job dovetails with the fact that mandatory schooling ends in the year when individuals attain age 15 and that apprenticeships usually last for two to four years. The high fraction of blue-collar occupations is consistent with our intention of only including individuals who received some kind of vocational training. Interestingly, a substantial fraction of the sample (about a third) experiences some unemployment before starting the first regular employment spell (these individuals are registered for unemployment benefits on average for somewhat more than one month). Consistent with this, we find that age at first entry (our proxy for schooling) is about half a year lower than age at the start of the first regular job, reflecting the fact that the transition from education to work often involves short periods of non-employment.

The next panel shows male unemployment rates at the state and NUTS-3 unit levels as well as the aggregate number of labor market entrants at the state level (see footnote 9). The first two unemployment rates refer to the male working-age population, while the other two refer to the population of young workers. The unemployment rate in the year of labor market entry averages about 6.6%, irrespective of the chosen aggregation level. Youth unemployment rates are somewhat lower than overall unemployment rates and equal about 5% on average. Again, there is very little difference across aggregation levels. Finally, about 3,750 male individuals (aged between 16 and 21) enter the labor market in any given state and year.

The bottom of table 1 shows that our sample predominantly consists of individuals working in either manufacturing, construction, or in wholesale and retail trade (representing 41.8%, 20.2%, and 15.9% of the sample, respectively). There are also substantial shares of individuals working in gastronomy and hotel business, transportation, finance, as well as for lobbies and social security agencies. Again, this reflects the fact that we mainly selected lower and medium-skilled individuals for the analysis. The large number of employees of first employer (almost 580 employees) hides the fact that most firms are small and thus that most individuals actually work in small firms. In fact, about 40% of all entrants start their first regular employment in a firm with 25 or fewer employees. Also, the firms in our sample predominantly engage males, probably simply reflecting the overall gender distribution across industries.

3 The Persistence of Initial Labor Market Shocks

We start with a simple graphical depiction of our two key measures (i.e. cohorts' wage profiles and the initial local unemployment rate). First, figure 1 shows wage profiles by entry cohort for all labor market cohorts who first entered the labor force between 1978 and 2000. The black dots therefore represent average starting wages for each entry cohort and the dashed grey line thus shows how starting wages evolve over time. Clearly, real starting wages have increased significantly over the period of analysis, from about 38 € in 1978 to about 50 € in 2000. Also note that there is some cyclical movement in starting wages over time which we expect to be related to economic conditions prevailing in that year.

Figure 1

The solid colored lines, on the other hand, represent long-run wage profiles of cohorts entering the labor market in different years. Cohorts' wages clearly follow an approximate concave path over time, implying that wage growth is highest in earlier working years and then strongly flattens later on. The figure shows, for example, that the 1978 entry cohort starts with a real wage of about 38 € per workday and experiences a raise in real wage up to about 97 € by the year 2005. On average, this cohort's compensation has therefore more than doubled in real terms within the first 27 years of labor market experience. Focusing again on the 1978 entry cohort, we see that this cohort's average wage has grown by approximately 146% ($= [\exp(.9) - 1] \cdot 100\%$) in the first 27 years of experience. Evidently, most of this wage increase happens at the early stage of the labor market career (i.e. the wage increase in the first ten years equals about 86% ($= [\exp(.62) - 1] \cdot 100\%$)).[14]

Figure 2

Panel (a) of figure 2 shows the evolution of state level unemployment rates which we use as our indicator for external labor market conditions at the time individuals first enter the labor force. This figure shows that the period from 1978 to 2000 covers several periods of

[14] Wage profiles of different entry cohorts have somewhat distinct overall shapes. More specifically, the figure shows that returns to experience generally decrease over time, meaning that younger entry cohorts have considerably lower returns to labor market experience than older cohorts. For example, the 1995 cohort only realizes an average wage increase of about 58% ($= [\exp(.46) - 1] \cdot 100\%$) in the first ten years, thus less smaller than that of the corresponding increase of the 1978 entry cohort.

both boom and downturn, and that the identifying variation in initial labor market conditions therefore does not only stem from a few neighboring labor market cohorts. The figure makes it also clear that states not only differ in the level of unemployment but also with respect to cyclical variations around a longer run trend: although all states see an increase over the whole observation period in general, there are marked cyclical differences across states and districts.[15]

Moreover, the lower panel of figure 2 shows that our observation period spans several ups and downs of the business cycle, and that there is considerable differentiation in the strength of these variations across states. We thus have both sufficient cross-sectional and longitudinal variation in our key regressor that we can use to pin down the effect of local business cycle fluctuations on wages.

3.1 Econometric Framework

Because we primarily aim at estimating the long-run impact of economic shocks at the time individuals started their first jobs, we must take care to allow the association between initial conditions and wage to become weaker or stronger as labor market experience increases while also using a generally flexible functional form of wage profiles. Taking these issues into account, our basic econometric model is the following:

$$\begin{aligned}\ln(y_{it}) = {} & ur^0_{j[i]}\alpha_1 + \kappa(\exp_{it})\alpha_2 + [ur^0_{j[i]} \cdot \kappa(\exp_{it})]\alpha_3 + ur_{j[i]t}\delta_1 + [ur_{j[i]t} \cdot \kappa(\exp_{it})]\delta_2 \\ & + \ln(n^0_{j[i]})\beta_1 + x_i\beta_2 + \psi_j + t_i\gamma_j + \epsilon_{it}, \end{aligned} \tag{1}$$

where y_{it} denotes the real daily wage of individual i in calendar year t, $\exp_{it}$ potential labor market experience of i in year t, and $ur^0_{j[i]}$ the unemployment rate as prevailing in region j at the time individual i first entered the labor market. Function $\kappa(\cdot)$ denotes that we allow for a flexible functional form with respect to labor market experience.[16] Note that we allow

[15]For example, and as highlighted in the figure, Burgenland (located in southeastern Austria) experienced a huge increase in the unemployment rate from about 3% in the late 1970s to about 8% in the first half of the 1980s, and then to about 9% in the second half of the 1980s. Vorarlberg (situated in western Austria), in contrast, experienced only a modest increase from about 1% on the 1970s to about 3% in the 1980s. In 1992, however, Vorarlberg underwent a sharp decline in the local labor market conditions, when unemployment jumped from about 3% to about 7–8%.

[16]Specifically, we include the first three polynomial terms of potential labor market experience. We chose the number of polynomial terms on the basis of a non-parametric, and therefore fully flexible, wage-experience model. The first three polynomial terms appear sufficient to reproduce the wage-experience profile predicted

the effect of the initial unemployment rate on current wages to vary as potential labor market experience increases by including interaction terms between the experience polynomials and the initial unemployment rate. We also include the current local unemployment rate, $ur_{j[i]t}$, and its interactions with potential labor market experience $\kappa(\exp)$ to control for the effect of current labor market conditions on wages (while allowing current conditions to have different effects on workers with different experience). We further include the number of labor market entrants aged 16 to 21, $n^0_{j[i]}$, to control for large demographic shifts within the considered birth cohorts. We include additional control variables in some specifications, denoted by vector x_i. Note that the controls variables are predetermined in the sense that they relate to an individual's first regular employment spell or to the time before having started to work (i.e. there is no time index for the controls). Finally, we include a set of state dummies (denoted by ψ_j) and state-specific quadratic time trends (denoted by $t_i\gamma_j$) in most specifications.[17] Parameters α_1 to α_3 describe wage-experience profiles as a function of the initial unemployment rate and are the parameters of main interest. Specifically, α_1 is the elasticity of wages with respect to the initial unemployment rate in the year of first entry (i.e. in the year where labor market experience is equal to 0), while α_3 tells us how the effect of initial conditions changes as labor market experience increases.

One important complication implied by the specification given by equation (1) relates to the fact that the local initial unemployment rate does not vary over time for any given individual. For this reason we cannot use standard panel data estimators such as the fixed-effects or first-differences estimator because these methods not only eliminate all unobserved time-invariant heterogeneity but also all variation in the key regressor. We therefore rely on estimation methods that use the untransformed data. Another critical issue is potential endogeneity regarding the initial unemployment rate due to sample selection over the business cycle. If the composition of entry cohorts changes endogenously over the business cycle, this may lead to inconsistent estimates of the effect of initial conditions on wages.[18] To address

from a corresponding non-parametric specification.

[17]We include a common quadratic time trend in those specifications without state-specific time trends.

[18]This line of argument has been put forth by Bils (1985) and more recently by Solon *et al.* (1994) and Blundell *et al.* (2003). In fact, the timing of labor force entry and thus the composition of labor market cohorts may be endogenous for several distinct reasons. First, some potential labor market entrants may refrain from entering the labor market altogether. Second, both the choice of education as well as the duration of schooling may be endogenous, as both job prospects are weak and opportunity costs of schooling low in times of high unemployment. The most likely reason for endogenous labor market entry is the fact that some workers simply

this issue, we mainly rely on estimates where we instrument unemployment at first entry with the unemployment rate at age 16.[19] In fact, a comparison between estimates that take selection into account and those that do not do so unambiguously shows that this issue is crucial in our context and we thus focus on these instrumental variable estimates (see section 3.3 below). Finally, we have to consider that our key regressor is observed at a higher level of aggregation than the dependent variable, a situation that may lead to grossly misleading statistical inference (Moulton, 1986). All standard errors that we report are therefore clustered by cells defined by year at first entry × state of first entry (there are 9 states and 23 entry years, resulting in 207 distinct cells).[20]

3.2 Main Results: Initial Labor Market Conditions and Wages

We first discuss instrumental variables estimates of variations of the basic regression model given by equation (1). Table 2 shows point estimates of the parameters (i.e. α_1 and α_3) describing the effect of initial conditions on wages at the top and estimated semi-elasticities of wages with respect to the initial local unemployment rate at specific values of potential labor market experience (i.e. potential labor market experience of 0, 5, 10, 15, and 20 years, respectively) in the middle of the table. For example, $\varepsilon^{y}_{ur}(5)$ denotes the estimated semi-elasticity of the real daily wage with respect to the initial unemployment rate at five years of potential labor market experience. It thus corresponds to the estimated relative change in wages resulting from a one percentage point increase in the initial unemployment rate.

Table 2

The model in the first column only includes the log of the number of labor market entrants aged between 15 and 21, the current unemployment rate and the interaction terms with potential labor market experience, a full set of state dummies, and a common quadratic time trend

delay their entry when faced with unfavorable entry conditions, either by registering for unemployment benefits or staying out of the labor force until they find a job. Whatever the underlying reason, if those workers who do not immediately get a job are a selected group of all workers who intend to enter employment in a given year, then the composition of the actual entrants changes along with corresponding changes in the unemployment rate and thus potentially biases the estimated effect of the initial unemployment rate on wages.

[19] Kahn (2010), Kondo (2008) and Oreopoulos *et al.* (2006, 2008) use a similar instrumental variable strategy. OLS and IV estimates are similar in Oreopoulos *et al.* (2006, 2008), but IV are substantially larger in Kondo (2008). Kahn (2010) only reports IV estimates. Kwon *et al.* (2010) and Stevens (2007), the two European studies, only show OLS estimates.

[20] We also computed standard errors that simultaneously account for clustering at both levels for our main estimates. This yields standard errors that are virtually indistinguishable from those actually reported.

across states as additional controls. As expected, there is an immediate large negative effect of the initial local unemployment on wages in the year of entry. Specifically, the semi-elasticity of wages with respect to the initial unemployment rate equals -0.27 in the year of entry. The corresponding standard error equals 0.061, and the immediate wage effect is thus highly significant. Furthermore, the middle panel of the table shows that there is substantial persistence of this negative wage effect. According to the estimates from the first specification, a negative and significant effect of initial labor market conditions remains as much as twenty years after first entry into the labor market, but the wage effect clearly fades away over time.

We include state-specific quadratic time trends instead of imposing a common time trend across states in the second column. This may be rationalized on the observation that the states differ markedly with respect to overall economic conditions (see figure 2). The estimated semi-elasticity in the year of first entry increases and amounts to -0.36 (with a standard error of 0.091). The immediate wage effect is thus about a third larger than the corresponding effect from the first specification. Consistent with this finding, the longer-run wage effects also become more pronounced in this specification, as shown in the middle of the table. For example, the relative wage effect at five years of potential labor market experience increases from -0.13 to -0.17 (a relative increase by almost a third).

The estimates shown in the third column come from a regression specification that includes state-specific time trends and additional control variables at the individual or at the firm level. These additional controls relate to a worker's first regular employment spell or the time before that.[21] Taking these additional individual-level controls into account decreases the short-run effect of the initial unemployment rate markedly (the short run effect shrinks by about a third, compared to the second column). Nonetheless, the estimated immediate semi-elasticity still equals -0.209 with a standard error of 0.073, and it is thus still substantial and highly significant. The longer-run pattern of wage effects also changes substantially, as the impact of initial conditions on wages is substantially reduced at any value of potential labor market experience when controlling for differences in these observable characteristics. In fact, the wage effects at higher values of experience decrease to a larger relative amount.

[21] Specifically, we include our proxy for education, an individual's age at first entry into the labor force, and two indicator variables for blue or white-collar occupation at first regular employment. Concerning the initial employer, we include the number of employees and the fraction of the workforce that is female, as well as location (at the state level) and industry affiliation (15 broad categories).

The fourth column adds a full set of dummies for year of entry, thus picking up any systematic but unobserved differences across entry cohorts such as differences in average educational quality (note that we include an explicit control for cohort size). As expected, the inclusion of entry cohort dummies has a huge effect on the estimated wage effects of initial conditions. The immediate effect drops to -0.111, but remains large and statistically significant nonetheless. It also lowers the effects at lower values of potential experience, but at the same time slightly increases the wage effects at higher levels of labor market experience.

The next specification in column five additionally allows for cohort-specific returns to experience by adding interaction terms between the three included polynomial terms of labor market experience and the year of entry. This accounts for the steady downward shift in returns to experience evident from figure 1. Allowing for varying returns to experience results in only slightly different estimates, however, when compared to the specification that includes dummies for year of entry but no interactions with potential experience.

Finally, the sixth specification includes not only a full set of entry-year dummies but also a full set of calendar-year dummies, thus also accounting for any irregular shifts in average wages over time that are common across states and not already picked up in the parametric time trends. This specification yields essentially the same pattern of wage elasticities as the two preceding specifications.

Table 3

Table 3 shows some simple sensitivity checks with respect to the main regressor, using the same specification as in column 5 of table 2. In the first column, we use (initial and current) unemployment rates at the NUTS-3 rather than at the state level (see also footnote 9). In this case, the cross-sectional dimension of the initial unemployment rate is markedly increased as there are 35 NUTS-3 regions, compared to only nine states. On the one hand, this should decrease standard errors because the relevant cross-sectional dimension for the clustering is now done with 35 instead of 9 regional units. On the other hand, however, using the unemployment rate at a more disaggregated level may run the risk of being endogenous as workers may move to regions with lower unemployment (e.g. Wozniak, 2006). Indeed, there is a notable difference using unemployment rates at different aggregation levels: using the unemployment rate at the district level results in a smaller initial wage effect and in less persistent overall effects. This presumably reflects endogeneity of the local unemployment rate at lower aggregation levels, resulting from workers moving from regions with high unemployment to regions with lower levels of unemployment.

The specification in the second column uses state-level youth unemployment rates instead of overall unemployment rates (again for both initial and current unemployment). The resulting wage effects are somewhat larger, but still reasonably close to our main estimates based on overall unemployment rates.

Quantitative Implications: Effects on Discounted Lifetime Earnings

Even though initial wage differentials between entry cohorts apparently fade away as potential labor market experience increases, our estimates nonetheless imply a non-negligible negative effect on lifetime earnings from entering the labor market when unemployment is high. This is illustrated at the bottom of table 2, which shows an approximate estimate for the loss in lifetime earnings associated with a hypothetical increase in the initial unemployment rate by one percentage point. The loss in lifetime earnings is computed as the average of the accumulated wage losses within the first twenty years of labor market experience.

According to our estimates, an increase in the initial local unemployment rate by one percentage point is associated with an approximate loss in lifetime earnings between 6.4% and 11.7%, depending on the exact specification of the regression model. Therefore, even in the longer run, unfavorable labor market entry conditions have a sizeable negative effect on workers' earnings.

3.3 How Important is Endogenous Labor Market Entry?

A comparison of our instrumental variable estimates with the same regression specifications based on simple OLS estimates is informative about the extent of sample selection over the business cycle. Table 4 therefore compares the estimated semi-elasticities of wages with respect to the initial unemployment rate for different labor market experiences based on IV estimates in panel (a) and based on simple OLS estimates in panel (b), for the same six specifications as in table 2 above (thus panel (a) simply replicates part of table 2).

Table 4

It is immediately evident that selection is important as there are substantial differences between the two estimation methods. Wage elasticities based on simple OLS estimates are much smaller than the corresponding IV elasticities across all specifications and for each value of labor market experience. For example, while the IV estimates of the most extensive specification in column (6) imply a large immediate negative effect on wages of -0.128, the corresponding wage effect based on OLS only amounts to -0.020. Similarly, IV estimates imply a sizeable negative effect on lifetime earnings on the order of 6.5% (specification from column 5), while the corresponding OLS estimates yield a small negative effect of 1.0%. Similar (or even larger) differences between IV and OLS estimates exist for the other specifications as well, notwithstanding the fact that IV estimates are much less precise than the corresponding OLS estimates.

Overall, the comparison of simple OLS and IV estimates makes it clear that there is substantial sample selection of labor market entrants over fluctuations in local unemployment rates in our sample. Cohorts who enter the labor force in times of high unemployment are positively selected, and this selection apparently mutes not only the cyclicality of starting wages with the business cycle, but also weakens the persistence of this wage effect to a considerable extent.

4 Labor Market Entry Conditions and Quality of First Job

One important mechanism that may explain cohort effects in wages starts with the observation that there is cyclical variation in job and/or task assignment within jobs over the business cycle and that high ability workers are assigned to jobs with lower average quality in recessions (e.g. Devereux, 2000).[22] If, moreover, these jobs offer different opportunities for accruing human capital or if, alternatively, a substantial part of human capital accumulated on the job is specific to a worker's task, occupation, and/or industry, then initial job and/or task assignment is probably able to explain a significant part of the observed persistent wage effects resulting from unfavorable initial labor market conditions (e.g. Gibbons and Waldman, 2006).[23] In

[22] Our comparison between IV and OLS estimates in section 3.3 above has already shown that labor market entrants are positively selected in times of high unemployment.

[23] See Gathmann and Schönberg (2010) for evidence on task specific, Kambourov and Manovskii (2009) on occupation specific, and Neal (1995) or Parent (2000) on industry specific human capital. Sullivan (2010) provides evidence that both occupation and industry specific human capital are simultaneously important for the level of wages.

line with these findings, Mansour (2009) finds that the quality of a worker's first job is lower when entry is during high unemployment, even though cohorts entering during recession are positively selected.

Based on this evidence, we provide an additional analysis of the association of initial labor market conditions and the quality of a worker's first job in this section. Our empirical measures for the quality of workers' first job are employment-weighted industry and employer-specific wages for both male and female workers aged 22 to 65 (thus excluding labor market entrants themselves) in any given year.[24] Thus, for example, we compute the average wages paid in industry k and year t as $\overline{y}_{kt} = \sum_{i \in k,t}(y_{ikt} \cdot e_{ikt}) / \sum_{i \in k,t} e_{ikt}$, where y_{ikt} and e_{ikt} are, respectively, the real daily wage and the number of employment days of worker i of industry k in year t. We regress each of these measures on the initial local unemployment rate and additional control variables:

$$\ln(\omega_i) = ur^0_{j[i]}\alpha + x_i\beta + \psi_j + \varphi_t + \epsilon_i, \tag{2}$$

where ω_i denotes the industry or employer-specific wage associated with a worker's first job, and where $ur^0_{j[i]}$ again denotes the initial local unemployment rate. Additional control variables are denoted by x_i, and ψ_j and φ_t denote that we also include a full set of dummy variables indicating state at entry and year of entry into the labor force, respectively. As above, standard errors are clustered by year at entry $\times$ state at entry. We show results with and without additional firm level controls because of potential endogeneity of employers' characteristics. Estimates of α are shown in table 5.

Table 5

The first four columns show the results for industry specific wages.[25] There is a significant and substantial negative effect of initial conditions on the average wage associated with the industry of the first employer in each of the four specifications. The wage effect resulting from a one percentage point increases in the initial unemployment rate ranges from -0.019 to

[24] Mansour (2009), however, uses average occupation-specific wages to measure quality of initial occupational assignment. We compute industry and firm-specific wages because the ASSD does not contain information about workers' occupation (except for the distinction between blue and white-collar jobs).

[25] The ASSD contains two distinct industry classifications and we report estimates for both. The first, older classification (wikl) has been replaced by a common classification (nace) used throughout the European Union. See Zweimüller *et al.* (2009) for details. The correlation between the two wage measures in our sample is 0.75.

-0.008. Workers who enter the labor market during high local unemployment thus start their employment career in industries that pay significantly lower wages on average (again note that labor market entrants' wages do not enter the outcome variables).

The remaining two columns show estimates for employer specific wages as outcomes. We again find a significant and substantial negative effect of the initial local unemployment rate on employer specific average wages. A one percentage point increase in initial unemployment predicts that a labor market entrant starts his employment career with an employer who pays wages that are lower by 2.9% to 3.8%.

Taken together, the results from table 5 point to substantial fluctuations in first job quality over the business cycle. More specifically, and in line with previous evidence, we find that workers who enter the labor force during tight labor markets find better paying jobs than those who enter during times of high unemployment, even though the latter group of workers is positively selected.

5 Initial Labor Market Conditions and Wage Losses from Involuntary First Job Changes

In a recent study, Schmieder and von Wachter (2010) study how wage losses resulting from job displacement are associated with initial conditions at the start of an employment spell as well as with the best conditions in the course of the corresponding job. The key idea is that if wage shortfalls associated with unfavorable entry conditions disappear upon involuntary job loss, this may be viewed as supportive evidence that these wage losses are unrelated to workers' underlying productivity. If workers take these initial wage losses along to the next employer, however, this rather suggests that the initial wage shortfall reflects some difference in underlying productivity. In this section we draw on this basic idea and study the association between initial labor market conditions and, respectively, wages with a worker's new (second) employer and wage changes from involuntary job changes from the first to the second employer. For those workers who move from their initial employer to a new firm, we estimate the association of the initial unemployment rate with both the wage with the old (first) and new (second) employer.

Practically, we first need to determine whether a job change is involuntary or not (while only focusing on first job changes). One possibility is to focus on job displacement following plant closure, which can arguably be seen as involuntary job mobility. However, because only a few first regular employment spells are actually terminated by a plant closure, we cannot rely on this identification strategy. Moreover, the ASSD contains no direct information about the reason for the termination of any employment spell and we thus do not know whether job changes are voluntary or not. Following Gruetter and Lalive (2009), we therefore try to discriminate approximately between voluntary and involuntary job changes based on how the time between the first and second job is spent. Specifically, in what follows, we define as involuntary job mobility if a worker spends at least one day in registered unemployment between the first and the second job.[26] Based on this distinction, we estimate the following regression model for different subsets of first job changers:

$$\omega_i = ur^0_{j[i]}\alpha_1 + \kappa(\exp_i)\alpha_2 + ur_{j[i]}\delta + x_i\beta + \psi_j + \varphi_t + \epsilon_i, \tag{3}$$

where the outcome ω_i is either the log real daily wage in the last year of employment with a worker's old (first) employer, the log real daily wage in the first year with a worker's new (second) employer, or the difference between the two. $ur^0_{j[i]}$ is again the initial local unemployment rate, while $\exp_i$ and $ur_{j[i]}$ now refer to labor market experience and the current local unemployment rate in the year of first job change. Additional control variables do not vary with labor market experience and are denoted by x_i. We also include a full set of dummy variables for year of entry and state at entry.

Table 6

Table 6 shows instrumental variable estimates (using the same strategy as above) of parameter α_1 for the subset of all first job changes in panel (a) and for involuntary first job changes only in panel (b). In the following, we focus on the estimates reported in panel (b) which refer to the subsample of involuntary job moves. Consistent with our main findings (see table

[26]Table B.1 shows some descriptives for workers who move from their first employer to another employer. First, workers who voluntarily move to the second employer have a higher wage both at the end of the first and the beginning of the second employment spell than those who involuntarily change jobs, and thus they also have larger wage gains from switching employers. The next rows report the number of days spent in either non- or unemployment between a worker's first and second job. By definition, the number of registered unemployment days equals zero for the group of voluntary job changers. Involuntary job changers spent about 76% of the time between first and second employer in unemployment. There is a significant difference in the overall gap between first and second job as well. Voluntary (involuntary) job movers spend 185 (593) days out of employment between their first and second jobs.

2), the first column shows that the wage with the old (first) employer is negatively associated with initial labor market conditions. The point estimates implies that a one percentage point increase in the initial local unemployment rate is associated with 19.4% lower wage at the end of the first job.

The second column shows how the wage with the new (second) employer depends on the initial unemployment rate (conditional on current conditions). Interestingly, there is still a substantial and significant negative association between labor market entry conditions and the wage with the new employer. The point estimate is -0.106 (with a standard error of 0.013), thus only about half the size of the corresponding estimate from the first column with respect to the wage with the old employer. This is an interesting result that has two important implications. First, it implies that part of the observed wage differences between entry cohorts reflects underlying differences in human capital. This finding complements and is consistent with our previous findings on cyclical variation in the quality of initial job assignment. Thus workers who enter during a recession are initially assigned to lower quality jobs/tasks, and this initial assignment creates wage persistency because part of the human capital accumulated with the initial job is specific to this job because of occupation or industry specific human capital. Second, part of the wage losses associated with poor entry conditions evaporates upon involuntary job loss and thus part of the observed wage differences appear to be unrelated to differences in human capital. This implies that mechanisms unrelated to productivity, such as downwardly rigid wage contracts (Beaudry and DiNardo, 1991), also play an important role in generating persistent wage differences across entry cohorts.

6 Additional Results

6.1 Subsample and Sensitivity Analysis

This final section presents some additional results that emphasize our interpretation of why labor market entry conditions generate persistent wage effects. Table 7 shows results for some distinct subsamples and some alternative specifications. All columns of table 7 use the same specification as column 5 from table 2.

The first two columns show separate estimates for workers who start their labor market career in either a blue or a white-collar job.[27] Even though the initial wage effect is virtually identical for blue and white-collar workers, the longer-run effect differs substantially between the two groups. Because there is less persistence in wage losses from poor entry conditions, the long-run wage effects for white collar workers are only about half the size of those for blue collar workers (our estimates imply an approximate loss in lifetime earnings of 3.8% and 7.3%, respectively). This result is somewhat surprising because previous research has shown that white collar workers in Austria suffer from much larger and more persistent wage losses from job displacement (Schwerdt *et al.*, 2010). However, in view of our preceding results, this result emphasizes that initial job assignment and occupation specific human capital are important in explaining persistent wage effects especially because occupation specific human capital is arguably more important for blue (e.g. craftsmen) than for white-collar jobs (Sullivan, 2010), at least in our sample, where the most highly skilled workers are excluded.[28]

The next three columns show results for native and immigrant workers separately.[29] The comparison between native and immigrant workers shows that there are only small differences between native and immigrant workers, both in the short and in the long-run: the estimated loss in lifetime earnings amounts to 6.1% for natives and 6.5% for immigrant workers. The fifth column additionally shows estimates for the group of immigrant workers from Turkey and former Yugoslavia, the two largest groups of immigrants in the Austrian labor market (they account for 57% of all immigrants in our sample). We even find a smaller impact of poor entry conditions on wages for these workers than for native workers, both in the short and in the long-run (the estimated loss in lifetime earnings amount to 5.5%, compared to 6.1% for native workers). As for the comparison between blue and white-collar workers, this contrast does probably not confirm prior expectations. However, it can again be reconciled with the argument that immigrant workers probably tend to work in jobs that require less specific skills and more often involve repetitive tasks.

Table 7

[27] Note that these two groups do not exactly add up to the overall sample size reported in table 2 because some employment spells cannot be uniquely identified as either blue or white-collar (see also table 1).

[28] Moreover, we use a totally different sample of workers. Our study focuses on low and medium-skilled labor market entrants who entered the labor force between 1978 and 2000, while Schwerdt *et al.* (2010) focus on prime age workers of any educational level who experienced a plant closure between 1982 and 1988.

[29] Note that an immigrant worker is, unlike in the US, an individual who does not possess the Austrian citizenship, but not necessarily an individual born outside of Austria. Consequently, the ASSD contains information on citizenship but not on country of birth.

We next split the sample with respect to the initial local unemployment rate. Specifically, we first compute the deviations in the initial local unemployment rate from its quadratic time trend (separately for each state). We then run our baseline specification separately for those observations with positive and negative deviations, respectively, as shown in panel (b) of figure 2. This allows us to study whether there is a symmetric impact of the initial unemployment rate for those entering in times of either tight or loose labor markets.[30] The comparison between the estimates for the two subsamples shows that there is an asymmetric effect of the initial local unemployment rate on wages, both in the short and the long-run, and that the wage loss from entry during unfavorable conditions appears to be larger than the gain from entry during good aggregate conditions. More specifically, the immediate wage effect of a one percentage point increase is about twice as large for those observations of workers who enter at an initial local unemployment rate above its trend than for those below its trend (point estimate of -0.183 and -0.088, respectively). The same asymmetry holds in the longer run as well: as shown at the bottom of the table, the estimated loss in lifetime earnings (resulting from a one percentage point increase in the local unemployment rate) for the former group amounts to 9.6%, while the loss for the latter group equals 5% only.

This result again supports our view that initial job assignment and human capital accumulation is probably more important than mechanisms unrelated to productivity because mechanisms such as implicit contracting models with downward rigid nominal wages are more likely to predict larger effects for those who enter during a boom than for those who enter during recession.

6.2 Other Outcomes: Annual Earnings, Unemployment, and Mobility

We next show results for labor market outcomes other than wages in table 8, using exactly the same specification as in column 5 of table 2. The first column of table 8 shows estimates for real annual earnings, i.e. the real daily wages times the number of employment days in any given year. The long run effect is somewhat smaller than the corresponding effect on wages. This implies that there is a small positive effect on employment from initial labor market conditions. Consistent with this finding, the second column shows that there is also a small,

[30] Note that both subsamples contain virtually the full range of both birth and entry cohorts because our observation period spans several ups and downs in local unemployment rates (again see figure 2).

but significant negative effect of the initial unemployment rate on the fraction of the year that is spent in registered unemployment.

Table 8

The remaining three columns show results for different job mobility measures: mobility across employers, mobility across different industries, and regional mobility.[31] In each case, the corresponding dependent variable is a binary indicator taking on the value one if the employer (industry, region) in year t is not the same as in year $t-1$ (and thus the last observation year for any individual drops from the analysis). There is a substantial, and again statistically significant, negative effect of initial conditions on all three mobility measures considered. For example, a one percentage point increase in the initial unemployment rate decreases the probability of moving to another employer by about 7.5% ($= -0.015/0.202 = -0.074$) in the short and by almost the same amount in the long run. Note that these negative effects on mobility may explain the positive (negative) effect on employment (unemployment) because switching employers often involves a period of non or unemployment between two jobs and because stayers accumulate tenure with the current employer.

Besides the negative effect on wages, labor market entrants starting their employment career thus not only face lower wages but also considerably lower mobility prospects later on.[32] Reduced mobility probabilities may explain part of the observed wage persistence, since job mobility is usually associated with wage increases (see also section 5 above). Also, reduced mobility probabilities are line with our previous argument that occupation and/or industry specific human capital, combined with initial assignment to lower quality jobs, are responsible for the persistent wage differentials across labor market entry cohorts.

7 Conclusions

We estimate the long-run impact of initial labor market conditions on wages for young males entering the Austrian labor market between 1978 and 2000. Consistent with previous evidence,

[31] Obviously, mobility across industries and/or across regions also implies mobility across different employers, but not vice versa.

[32] Our mobility results contrast strongly with those from Bachmann *et al.* (2010) for Germany (similar results are reported by Oreopoulos *et al.* (2006, 2008) for the US), who find that workers entering during poor entry conditions switch employers more often later on.

we find a substantial wage penalty from poor entry conditions on starting wages. The estimated semi-elasticity of starting wages with respect to the initial unemployment rate is on the order of -0.12 (thus a one percentage point increase in the initial local unemployment rate is associated with 12% lower starting wages). Moreover, this initial wage loss from first entry into the labor force during high unemployment turns out to be persistent, and sizeable negative effects on lifetime earnings thus result. Our preferred instrumental variable estimates imply that an increase in the initial local unemployment rate by one percentage point is associated with an approximate loss in lifetime earnings of about 6.5%. We also find that endogenous entry into the labor market is a substantial issue in the Austrian labor market, biasing simple OLS estimates towards small initial wage shortfalls and weakly persistent wage effects because of positive selection in times of high initial unemployment.

How do our results compare to previous findings? We find considerably larger wage effects than the two other studies for European labor markets, Stevens (2007) for Germany and Kwon *et al.* (2010) for Sweden. However, neither of these studies controls for endogenous labor market entry that is likely to attenuate the initial wage response as well as to weaken the persistence of the wage loss stemming from aggregate conditions. Our comparison between IV with simple OLS estimates shows that there is considerable positive selection in times of poor entry conditions, and OLS estimates amount to at most 20% of the corresponding IV estimates. Indeed, our OLS estimates are close to the results reported in Stevens (2007) and Kwon *et al.* (2010).[33]

We further show evidence that is consistent with cyclical variation in the quality of a worker's first regular employment, measured by industry and employer specific wages. Thus, even though workers are positively selected in times of high unemployment, they are assigned to industries and/or employers that generally pay lower wages. An additional and complementary analysis of the association of wage losses resulting from involuntary first job changes with initial labor market conditions supports the view that initial job and/or task assignment within jobs is important in explaining the wage persistence we observe. These two findings are also well in line with several findings from our subsample analysis. Specifically, we find that wage losses are substantially larger for blue than for white-collar workers, that wage losses are larger for natives

[33] Our finding of persistent wage effects from poor entry conditions are also in line with Frühwirth-Schnatter *et al.* (2010) who study labor earnings mobility in Austria using the same data source as we do, but a completely different empirical approach.

than for immigrants overall as well as immigrants from Turkey and former Yugoslavia, and that wage losses for those entering during above-trend initial unemployment are twice as large as the corresponding wage gains for those workers entering during below-trend unemployment. We further find that unfavorable entry conditions have a negative effect on workers' mobility as well. Again, this is consistent with the view that a significant part of human capital that accrued on-the-job is occupation and/or industry-specific and that workers are assigned to jobs or tasks of lower average quality during recession. Taken together, these results clearly suggest that initial job assignment is part of the explanation behind cohort wage effects. On the other hand, we have also shown that part of the initial wage gain resulting from labor market entry during good conditions is lost when switching to another employer. This suggests that mechanisms unrelated to workers' productivity, such as downwardly rigid wage contracts or match specific rents, nonetheless also play a role in generating the observed cohort effects in wages.

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Table 1: Summary statistics, male labor market entrants 1978–2000

	Mean	Standard deviation
Individual-level characteristics:		
Real daily wage	48.182	15.448
Age at start of first regular job	19.194	1.007
Age at first entry into the labor force	18.659	1.456
Duration of first regular job (years)	2.836	4.321
Any unemployment before first job	0.307	0.461
Unemployment days before first job	42.703	101.405
Blue-collar	0.744	0.437
White-collar	0.253	0.435
Aggregate-level variables:		
State level unemployment rate	6.597	2.897
State level youth unemployment rate	4.942	1.790
NUTS-3 level unemployment rate	6.655	3.157
NUTS-3 level youth unemployment rate	4.958	2.136
Number of entrants aged 15-21	3,766.782	1,450.335
Firm-level characteristics:		
Number of employees (in 100's)	5.777	24.667
Female share of workforce	0.244	0.219
Region of employer:		
Vienna	0.179	0.383
Lower Austria	0.173	0.378
Burgenland	0.025	0.156
Upper Austria	0.193	0.395
Styria	0.148	0.355
Carinthia	0.067	0.250
Salzburg	0.071	0.256
Tyrol	0.092	0.290
Vorarlberg	0.053	0.223
Industry of employer:		
Agriculture	0.010	0.100
Electricity	0.010	0.100
Mining	0.008	0.087
Manufacturing	0.418	0.493
Construction	0.202	0.401
Wholesale and retail trade	0.159	0.365
Gastronomy, hotel business	0.044	0.205
Transportation	0.038	0.191
Finance	0.059	0.236
Cleaning, body care	0.008	0.088
Arts, entertainment, sports	0.005	0.071
Healthcare, welfare	0.007	0.082
Education, research	0.005	0.069
Lobbies, social security agencies	0.027	0.163
Housekeeping	0.000	0.022
Number of observations	220,214	

Notes: The four unemployment rates are computed from the individual-level raw data of the ASSD. See also notes of table A.1.

Table 2: The long-run wage effects of initial labor market conditions (IV estimates)

	ln(real daily wage)					
Mean	4.257					
Standard deviation	0.380					
ur^0	-0.270^{***}	-0.360^{***}	-0.209^{***}	-0.111^{***}	-0.124^{***}	-0.128^{***}
	(0.061)	(0.091)	(0.073)	(0.011)	(0.012)	(0.012)
$exp \cdot ur^0$	0.039^{***}	0.052^{***}	0.031^{***}	0.014^{***}	0.020^{***}	0.017^{***}
	(0.009)	(0.013)	(0.010)	(0.001)	(0.002)	(0.002)
$exp^2 \cdot ur^0$	-0.002^{***}	-0.003^{***}	-0.002^{***}	-0.001^{***}	-0.002^{***}	-0.001^{***}
	(0.000)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)
$exp^3 \cdot ur^0$	0.000^{***}	0.000^{***}	0.000^{***}	0.000^{***}	0.000^{***}	0.000^{***}
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$\varepsilon^y_{ur}(0)$	-0.270^{***}	-0.360^{***}	-0.209^{***}	-0.111^{***}	-0.124^{***}	-0.128^{***}
	(0.061)	(0.091)	(0.073)	(0.011)	(0.012)	(0.012)
$\varepsilon^y_{ur}(5)$	-0.130^{***}	-0.170^{***}	-0.095^{**}	-0.064^{***}	-0.059^{***}	-0.069^{***}
	(0.030)	(0.043)	(0.037)	(0.007)	(0.007)	(0.007)
$\varepsilon^y_{ur}(10)$	-0.074^{***}	-0.089^{***}	-0.047^{**}	-0.050^{***}	-0.045^{***}	-0.049^{***}
	(0.017)	(0.023)	(0.020)	(0.005)	(0.006)	(0.005)
$\varepsilon^y_{ur}(15)$	-0.069^{***}	-0.077^{***}	-0.038^{**}	-0.051^{***}	-0.053^{***}	-0.048^{***}
	(0.015)	(0.019)	(0.016)	(0.005)	(0.006)	(0.005)
$\varepsilon^y_{ur}(20)$	-0.083^{***}	-0.089^{***}	-0.038^{**}	-0.051^{***}	-0.056^{***}	-0.049^{***}
	(0.015)	(0.019)	(0.015)	(0.005)	(0.006)	(0.005)
Regional controls	Yes	Yes	Yes	Yes	Yes	Yes
ur_t, $ur_t \cdot \kappa$(exp)	Yes	Yes	Yes	Yes	Yes	Yes
State-specific trends	No	Yes	Yes	Yes	Yes	Yes
Individual controls	No	No	Yes	Yes	Yes	Yes
Year of entry	No	No	No	Yes	Yes	Yes
Year of entry $\cdot\ \kappa(exp)$	No	No	No	No	Yes	No
Calendar year	No	No	No	No	No	Yes
Number of observations	3,349,075	3,349,075	3,349,075	3,349,075	3,349,075	3,349,075
Adjusted R-Squared	0.154	0.072	0.299	0.348	0.348	0.348
Loss in lifetime earnings	0.117	0.146	0.079	0.064	0.065	0.067

Notes: ***, **, and * denote statistical significance on the 1%, 5%, and 10% level, respectively. Robust standard errors are given in parentheses and are clustered by state at entry × year of entry. exp and ur^0 denote potential labor market experience (in years) and the initial unemployment rate, respectively. $\varepsilon^y_{ur}(k)$ denotes the estimated semi-elasticity of wages with respect to the initial unemployment rate, evaluated at k years of potential labor market experience. ur_t denotes the current unemployment rate. The initial unemployment rate (and the interactions with labor market experience) are instrumented with the unemployment rate at age 16 (and the corresponding interactions with labor market experience).

Table 3: Sensitivity analysis (IV estimates)

	ln(real daily wage)	
	District-level ur	Youth ur
Mean	4.257	4.257
Standard deviation	0.381	0.380
ur^0	-0.049^{***}	-0.130^{***}
	(0.005)	(0.019)
$exp \cdot ur^0$	0.012^{***}	0.017^{***}
	(0.001)	(0.003)
$exp^2 \cdot ur^0$	-0.001^{***}	-0.001^{***}
	(0.000)	(0.000)
$exp^3 \cdot ur^0$	0.000^{***}	0.000^{***}
	(0.000)	(0.000)
$\varepsilon^y_{ur}(0)$	-0.049^{***}	-0.130^{***}
	(0.005)	(0.019)
$\varepsilon^y_{ur}(5)$	-0.013^{***}	-0.072^{***}
	(0.002)	(0.011)
$\varepsilon^y_{ur}(10)$	-0.010^{***}	-0.057^{***}
	(0.001)	(0.008)
$\varepsilon^y_{ur}(15)$	-0.020^{***}	-0.062^{***}
	(0.001)	(0.008)
$\varepsilon^y_{ur}(20)$	-0.025^{***}	-0.064^{***}
	(0.002)	(0.007)
Number of observations	3,267,615	3,349,075
Adjusted R-Squared	0.354	0.349
Loss in lifetime earnings	0.021	0.075

Notes: ***, **, and * denote statistical significance on the 1%, 5%, and 10% level, respectively. Robust standard errors are given in parentheses and are clustered on state (district) at entry times year of entry. exp and ur^0 denote potential labor market experience (in years) and the initial unemployment rate, respectively. $\varepsilon^y_{ur}(k)$ denotes the estimated semi-elasticity of wages with respect to the initial unemployment rate, evaluated at k years of potential labor market experience. The initial unemployment rate (and the interactions with labor market experience) are instrumented with the unemployment rate at age 16 (and the corresponding interactions with labor market experience). Both columns use the same regression specification as in column 5 of table 2.

Table 4: Comparison between IV and OLS estimates

(a) Semi-elasticities based on IV estimates						
$\varepsilon^{y}_{ur}(0)$	-0.270^{***}	-0.360^{***}	-0.209^{***}	-0.111^{***}	-0.124^{***}	-0.128^{***}
	(0.061)	(0.091)	(0.073)	(0.011)	(0.012)	(0.012)
$\varepsilon^{y}_{ur}(5)$	-0.130^{***}	-0.170^{***}	-0.095^{**}	-0.064^{***}	-0.059^{***}	-0.069^{***}
	(0.030)	(0.043)	(0.037)	(0.007)	(0.007)	(0.007)
$\varepsilon^{y}_{ur}(10)$	-0.074^{***}	-0.089^{***}	-0.047^{**}	-0.050^{***}	-0.045^{***}	-0.049^{***}
	(0.017)	(0.023)	(0.020)	(0.005)	(0.006)	(0.005)
$\varepsilon^{y}_{ur}(15)$	-0.069^{***}	-0.077^{***}	-0.038^{**}	-0.051^{***}	-0.053^{***}	-0.048^{***}
	(0.015)	(0.019)	(0.016)	(0.005)	(0.006)	(0.005)
$\varepsilon^{y}_{ur}(20)$	-0.083^{***}	-0.089^{***}	-0.038^{**}	-0.051^{***}	-0.056^{***}	-0.049^{***}
	(0.015)	(0.019)	(0.015)	(0.005)	(0.006)	(0.005)
Loss in lifetime earnings	0.117	0.146	0.079	0.064	0.065	0.067
(b) Semi-elasticities based on OLS estimates						
$\varepsilon^{y}_{ur}(0)$	-0.011^{***}	-0.011^{***}	-0.014^{***}	-0.022^{***}	-0.021^{***}	-0.020^{***}
	(0.003)	(0.003)	(0.002)	(0.003)	(0.003)	(0.003)
$\varepsilon^{y}_{ur}(5)$	0.003^{***}	0.003^{***}	0.002	-0.006^{***}	-0.004^{**}	-0.008^{***}
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)
$\varepsilon^{y}_{ur}(10)$	0.002^{*}	0.002^{**}	0.001	-0.006^{***}	-0.003^{**}	-0.009^{***}
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)
$\varepsilon^{y}_{ur}(15)$	-0.004^{***}	-0.003^{***}	-0.005^{***}	-0.013^{***}	-0.011^{***}	-0.015^{***}
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)
$\varepsilon^{y}_{ur}(20)$	-0.004^{**}	-0.003^{*}	-0.005^{***}	-0.013^{***}	-0.017^{***}	-0.015^{***}
	(0.002)	(0.002)	(0.002)	(0.003)	(0.002)	(0.002)
Loss in lifetime earnings	0.002	0.001	0.003	0.011	0.010	0.013
Regional controls	Yes	Yes	Yes	Yes	Yes	Yes
ur_t, $ur_t \cdot \kappa(\text{exp})$	Yes	Yes	Yes	Yes	Yes	Yes
State-specific trends	No	Yes	Yes	Yes	Yes	Yes
Individual controls	No	No	Yes	Yes	Yes	Yes
Year of entry	No	No	No	Yes	Yes	Yes
Year of entry $\cdot\ \kappa(exp)$	No	No	No	No	Yes	No
Calendar year	No	No	No	No	No	Yes

Notes: $\varepsilon^{y}_{ur}(k)$ denotes the estimated semi-elasticity of wages with respect to the initial unemployment rate, evaluated at k years of potential labor market experience. See also notes of table 2.

Table 5: Initial labor market conditions and quality of first job (IV estimates)

	$\ln(\overline{y}_t^{\text{wikl}})$		$\ln(\overline{y}_t^{\text{nace}})$		$\ln(\overline{y}_t^{\text{employer}})$	
Mean	4.229	4.229	4.234	4.234	4.182	4.182
Standard deviation	0.158	0.158	0.159	0.159	0.284	0.284
ur^0	-0.019^{***}	-0.012^{***}	-0.014^{***}	-0.008^{***}	-0.038^{***}	-0.029^{***}
	(0.002)	(0.002)	(0.002)	(0.001)	(0.004)	(0.003)
Individual controls	Yes	Yes	Yes	Yes	Yes	Yes
Entry year	Yes	Yes	Yes	Yes	Yes	Yes
Regional controls	Yes	Yes	Yes	Yes	Yes	Yes
Employer characteristics	No	Yes	No	Yes	No	Yes
Number of observations	217,587	217,587	217,587	217,587	215,324	215,324
Adjusted R-Squared	0.190	0.535	0.214	0.519	0.096	0.286

Notes: ***, **, and * denote statistical significance on the 1%, 5%, and 10% level, respectively. Robust standard errors are given in parentheses and are clustered by state at entry $\times$ year of entry. $\overline{y}_t^{\text{wikl}}$ and $\overline{y}_t^{\text{nace}}$ denote average wages paid in different industries, and where wikl and nace denote two different industry classifications available in the ASSD (with, respectively, 59 and 58 distinct categories). Analogously, $\overline{y}_t^{\text{employer}}$ denotes average wages paid by different employers. The initial unemployment rate is instrumented with the unemployment rate at age 16.

Table 6: Initial labor market conditions and wages before and after first job changes (IV estimates)

Dependent variable	$\ln(y^{old})$	$\ln(y^{new})$	$\Delta\ln(y)$
(a) All first job changes			
ur^0	-0.182^{***}	-0.120^{***}	0.062^{***}
	(0.016)	(0.011)	(0.009)
Mean	3.970	4.094	0.124
Standard deviation	0.426	0.376	0.415
Adjusted R-Squared	0.444	0.191	0.152
Number of observations	182,922	182,922	182,922
(b) Involuntary first job changes only			
ur^0	-0.194^{***}	-0.106^{***}	0.088^{***}
	(0.020)	(0.013)	(0.012)
Mean	3.963	4.043	0.080
Standard deviation	0.395	0.358	0.427
Adjusted R-Squared	0.365	0.117	0.128
Number of observations	90,297	90,297	90,297
Regional controls	Yes	Yes	Yes
State-specific trends	Yes	Yes	Yes
Individual and firm level controls	Yes	Yes	Yes
Entry-year	Yes	Yes	Yes

Notes: ***, **, and * denote statistical significance on the 1%, 5%, and 10% level, respectively. Robust standard errors are given in parentheses and are clustered by year of entry × state at entry. y^{old} and y^{new} denote, respectively, the real daily wage with the old (first) and the new (second) employer after first job change. $\Delta\ln(y)$ equals the difference between $\ln(y^{new})$ and $\ln(y^{old})$. ur^0 denotes the initial local unemployment rate.

Table 7: Subsample analysis (IV estimates)

	ln(real daily wage)						
	Blue-collar	White-collar	Natives	Immigrants	Turkish/Yugoslav immigrants	dev(ur^0)>0	dev(ur^0)<0
Mean	4.241	4.367	4.265	4.142	4.095	4.257	4.257
Standard deviation	0.354	0.438	0.379	0.381	0.361	0.377	0.383
ur^0	-0.103^{***}	-0.101^{***}	-0.120^{***}	-0.131^{***}	-0.092^{***}	-0.183^{***}	-0.088^{***}
	(0.013)	(0.014)	(0.012)	(0.023)	(0.021)	(0.020)	(0.013)
$exp \cdot ur^0$	0.010^{***}	0.012^{***}	0.020^{***}	0.016^{***}	0.010^{***}	0.032^{***}	0.014^{***}
	(0.002)	(0.003)	(0.002)	(0.003)	(0.003)	(0.004)	(0.002)
$exp^2 \cdot ur^0$	-0.001^{***}	-0.001^{***}	-0.002^{***}	-0.001^{***}	-0.001^{**}	-0.003^{***}	-0.001^{***}
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$exp^3 \cdot ur^0$	0.000^{***}	0.000^{***}	0.000^{***}	0.000^{***}	0.000^{*}	0.000^{***}	0.000^{***}
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$\varepsilon^y_{ur}(0)$	-0.103^{***}	-0.101^{***}	-0.120^{***}	-0.131^{***}	-0.092^{***}	-0.183^{***}	-0.088^{***}
	(0.013)	(0.014)	(0.012)	(0.023)	(0.021)	(0.020)	(0.013)
$\varepsilon^y_{ur}(5)$	-0.071^{***}	-0.055^{***}	-0.055^{***}	-0.075^{***}	-0.058^{***}	-0.085^{***}	-0.043^{***}
	(0.009)	(0.010)	(0.006)	(0.016)	(0.017)	(0.012)	(0.009)
$\varepsilon^y_{ur}(10)$	-0.062^{***}	-0.030^{***}	-0.042^{***}	-0.051^{***}	-0.045^{***}	-0.070^{***}	-0.036^{***}
	(0.008)	(0.009)	(0.006)	(0.014)	(0.016)	(0.011)	(0.009)
$\varepsilon^y_{ur}(15)$	-0.064^{***}	-0.014	-0.051^{***}	-0.043^{***}	-0.043^{***}	-0.084^{***}	-0.044^{***}
	(0.008)	(0.009)	(0.006)	(0.013)	(0.015)	(0.011)	(0.009)
$\varepsilon^y_{ur}(20)$	-0.064^{***}	0.005	-0.055^{***}	-0.034^{***}	-0.040^{**}	-0.075^{***}	-0.046^{***}
	(0.008)	(0.009)	(0.006)	(0.011)	(0.017)	(0.011)	(0.009)
Number of observations	1,797,908	563,764	3,020,446	203,370	116,999	1,526,797	1,822,278
Adjusted R-Squared	0.275	0.461	0.349	0.298	0.239	0.335	0.358
Loss in lifetime earnings	0.073	0.038	0.061	0.065	0.055	0.096	0.050

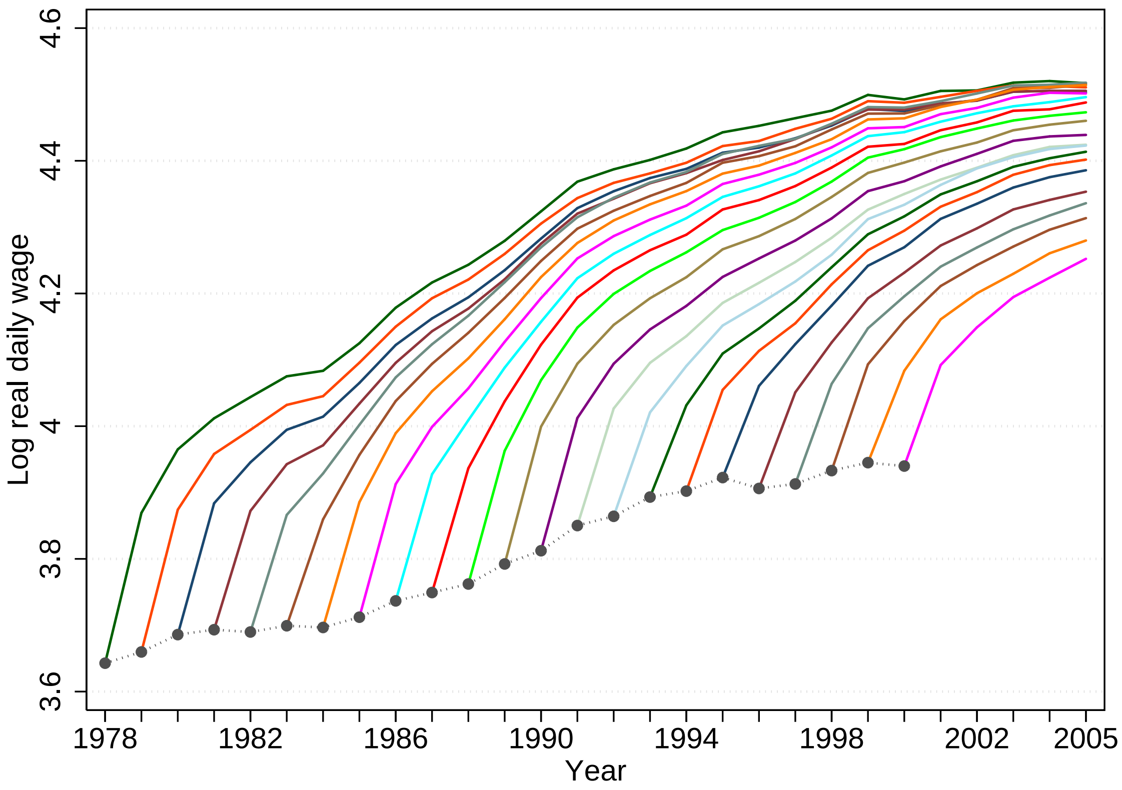
Notes: ***, **, and * denote statistical significance on the 1%, 5%, and 10% level, respectively. Robust standard errors are given in parentheses and are clustered on state (district) at entry times year of entry. exp and ur^0 denote potential labor market experience (in years) and the initial unemployment rate, respectively. dev(ur^0) denotes the deviation of the initial unemployment rate from its state-specific quadratic time trend. $\varepsilon^y_{ur}(k)$ denotes the estimated semi-elasticity of wages with respect to the initial unemployment rate, evaluated at k years of potential labor market experience. The initial unemployment rate (and the interactions with labor market experience) are instrumented with the unemployment rate at age 16 (and the corresponding interactions with labor market experience). All columns use the same regression specification as in column 5 of table 2.

Table 8: Other labor market outcomes (IV estimates)

	ln(annual earnings)	Unemployment	Mobility across		
			Employers	Industries	Regions
Mean	4.091	0.052	0.202	0.110	0.039
Standard deviation	0.656	0.131	0.401	0.312	0.195
ur^0	-0.071^{***}	-0.008^{***}	-0.015^{***}	-0.008^{***}	-0.004^{***}
	(0.008)	(0.002)	(0.003)	(0.002)	(0.001)
$exp \cdot ur^0$	0.006^{***}	0.001^{***}	0.002^{***}	0.001^{***}	0.001^{***}
	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)
$exp^2 \cdot ur^0$	-0.001^{***}	-0.000^{**}	-0.000^{***}	-0.000^{***}	-0.000^{***}
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$exp^3 \cdot ur^0$	0.000^{***}	0.000^{**}	0.000^{***}	0.000^{***}	0.000^{***}
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
$\varepsilon^y_{ur}(0)$	-0.071^{***}	-0.008^{***}	-0.015^{***}	-0.008^{***}	-0.004^{***}
	(0.008)	(0.002)	(0.003)	(0.002)	(0.001)
$\varepsilon^y_{ur}(5)$	-0.056^{***}	-0.006^{***}	-0.010^{***}	-0.005^{***}	-0.002^{*}
	(0.007)	(0.001)	(0.003)	(0.002)	(0.001)
$\varepsilon^y_{ur}(10)$	-0.060^{***}	-0.005^{***}	-0.013^{***}	-0.007^{***}	-0.004^{***}
	(0.007)	(0.001)	(0.003)	(0.002)	(0.001)
$\varepsilon^y_{ur}(15)$	-0.064^{***}	-0.003^{**}	-0.016^{***}	-0.009^{***}	-0.005^{***}
	(0.007)	(0.002)	(0.003)	(0.002)	(0.001)
$\varepsilon^y_{ur}(20)$	-0.050^{***}	-0.001	-0.011^{***}	-0.007^{***}	-0.005^{***}
	(0.008)	(0.002)	(0.003)	(0.002)	(0.002)
Number of observations	3,349,075	3,349,083	3,131,212	3,131,212	3,131,212
Adjusted R-Squared	0.237	0.044	0.032	0.020	0.018
Loss in lifetime outcome	0.063	0.004	0.013	0.007	0.004

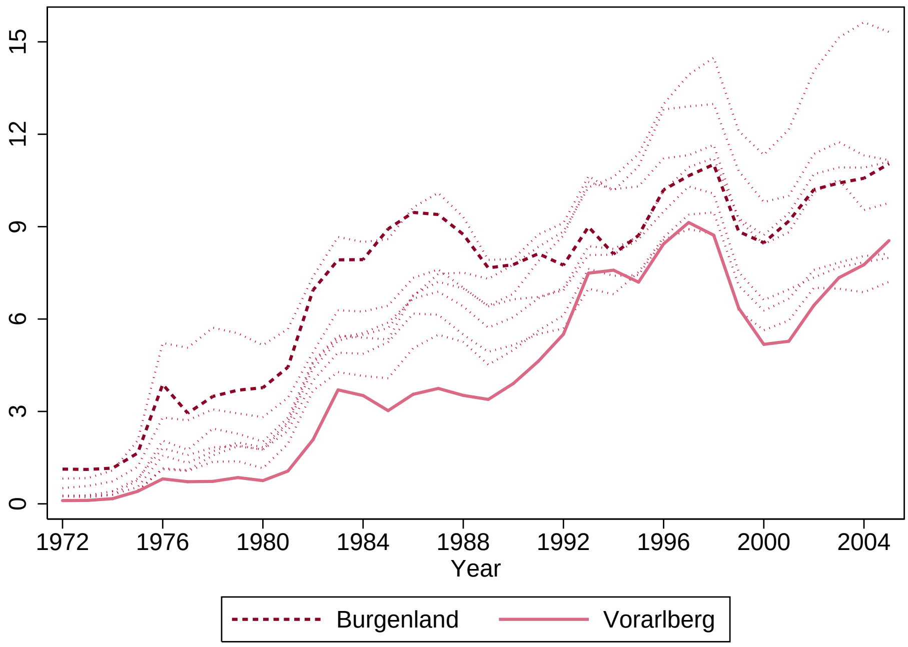
Notes: ***, **, and * denote statistical significance on the 1%, 5%, and 10% level, respectively. Robust standard errors are given in parentheses and are clustered on state at entry times year of entry. exp and ur^0 denote potential labor market experience (in years) and the initial unemployment rate, respectively. $\varepsilon^y_{ur}(k)$ denotes the estimated semi-elasticity of wages with respect to the initial unemployment rate, evaluated at k years of potential labor market experience. The initial unemployment rate (and the interactions with labor market experience) are instrumented with the unemployment rate at age 16 (and the corresponding interactions with labor market experience). All columns use the same regression specification as in column 5 of table 2.

Figure 1: Long-run wage profiles, by labor market entry cohort

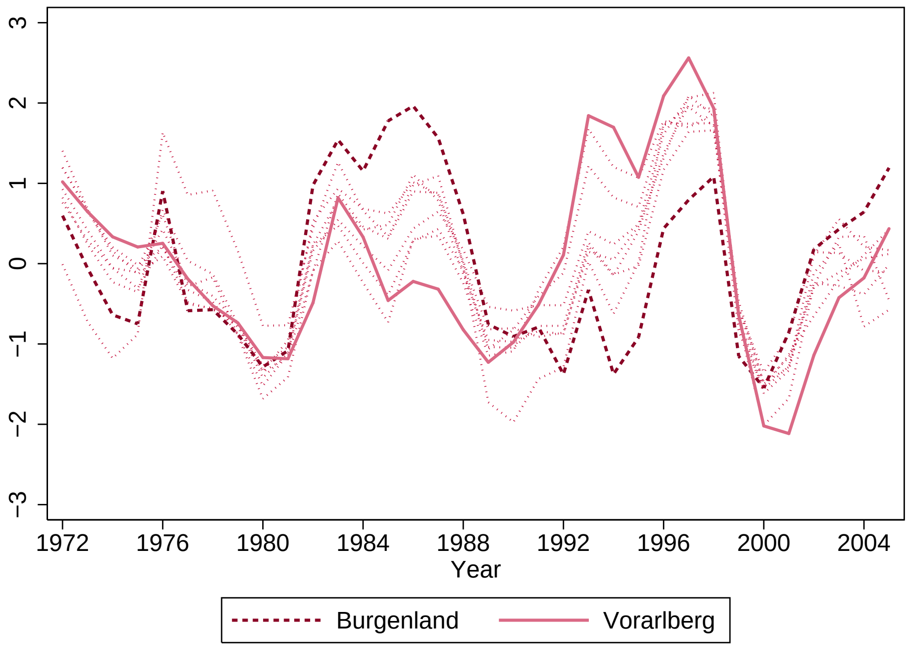


Notes: The figure shows average log real daily wages by calendar year for each labor market entry cohort from 1978 to 2000. Thus the black dots show average log starting wages for each labor-market entry cohort, and the black dotted line highlights the evolution of starting wages. The colored filled lines show how entry cohorts' wages evolve with increasing labor market experience.

Figure 2: Fluctuations in state-level unemployment rates, 1972–2005



(a) Unemployment rate (in %)



(b) Unemployment rate (in %), deviation from state-specific trend

Notes: The figure shows male unemployment rates for workers aged between 15 and 65, aggregated at the level of the state ("Bundesland") and computed from the individual-level raw data of the ASSD. Two of the nine states are graphically highlighted. "Vorarlberg" is situated in the western part of Austria (bordering Germany and Switzerland). "Burgenland" is located at the eastern border of Austria (bordering Hungary in the east and Slovenia in the south). Panel (b) shows deviations from state-specific quadratic time trends.

A Sample Construction

We first determine the start of the first regular employment spell for each male individual born between 1958 and 1985.[34] The restriction with respect to year of birth, in combination with the restriction on age at beginning of one's first regular job that we apply below, ensures that the potential range of age at first entry into the labor force is the same for each entry cohort considered in the analysis (1978–2000). Additionally, we drop all individuals who were self-employed and/or worked as a farmer or civil servant at least once, because the data do not consistently cover these employment spells over the whole period of analysis and/or because earnings are not recorded (in the case of self-employment). We thus cannot fully observe the employment and/or earnings histories of such individuals.

We then determine each individual's age at the start of his first regular employment spell starting between 1978 and 2000. We define regular employment as an employment spell which lasts for at least 180 days.[35] We also focus on individuals aged between 16 and 21 years at the start of their first regular employment spell. This leaves us with 797,846 unique individuals (see table A.1), from which we take a simple 30% random sample. This finally yields a total of 220,214 unique individuals and 3,349,075 observations (= individuals $\times$ years) when following these individuals over time.

Table A.1

Table A.1 shows descriptive statistics for some key variables, by individuals' age at the time of first entry into the labor force. We consistently exclude individuals who start their first regular employment after they attain age 30 because they presumably never enter the labor force at all.[36] The first column of table A.1 shows descriptives for all individuals, the second (third) column shows descriptives for individuals aged 16 to 21 (aged 22 to 30) when starting their first employment. A comparison of the second to the third column shows that our sample restriction with respect to age at the start of first regular employment works as expected. The sample of lower skilled workers, compared to the group of higher skilled workers, contains a higher fraction of blue-collar workers, has considerably lower wages on average, and shorter duration of the first regular employment spell.

Also note that, for the group of individuals aged between 22 and 30 when first entering the labor force, highly skilled workers potentially are mixed up with low-skilled workers: individuals in this group of workers either spent much time in education, were previously unemployed, or had short employment spells not counting as regular employment. This is apparent from the proportion of workers below the lower censoring point or above the higher censoring point. The probability of crossing any of the two points is higher for the sample of older workers. Consequently, the variation in the real daily wage (and thus productivity) is considerably smaller in the sample of younger workers than in the group of older workers.

[34] Obviously, an individual must be covered by the ASSD in order to be included in the sample. An individual is covered by the ASSD if he or she is entitled to future social security benefits (typically old-age benefits) or has already claimed social security benefits before first entering the labor force. Typically individuals "enter" the ASSD once they start working.

[35] Importantly, vocational training such as apprenticeship training is not considered as regular employment (but as formal training).

[36] The group of individuals who enter at a later age probably consists of two very different groups, who are indistinguishable from each other in the data. On the one hand there are truly high-skilled workers who enter the labor market at a later stage because they continued their education until that time. On the other hand, however, there are also low-skilled workers who were never employed or only sporadically employed before starting their first regular employment. Because schooling is not directly observed, we would mix these two groups of workers together if we were to include them.

Table A.1: Sample selection

	Age at start of first regular employment					
	16-30		16-21		22-30	
Real daily wage	50.488	(18.644)	47.536	(15.303)	56.742	(23.027)
Age at start of first job	20.863	(2.948)	19.200	(1.004)	24.386	(2.589)
Duration of first regular job	2.800	(3.995)	2.692	(4.035)	3.030	(3.900)
Blue-collar	0.650	(0.477)	0.726	(0.446)	0.487	(0.500)
White-collar	0.348	(0.476)	0.271	(0.445)	0.510	(0.500)
Below GfGr: Yes = 1	0.059	(0.235)	0.050	(0.219)	0.077	(0.266)
Above HBGr: Yes = 1	0.049	(0.215)	0.019	(0.135)	0.112	(0.315)
Region of employer:						
Vienna	0.239	(0.426)	0.198	(0.398)	0.326	(0.469)
Lower Austria	0.158	(0.365)	0.171	(0.377)	0.129	(0.336)
Burgenland	0.024	(0.152)	0.000	(0.000)	0.021	(0.142)
Upper Austria	0.168	(0.374)	0.183	(0.387)	0.137	(0.344)
Styria	0.139	(0.346)	0.145	(0.352)	0.129	(0.335)
Carinthia	0.065	(0.246)	0.066	(0.249)	0.062	(0.241)
Salzburg	0.070	(0.255)	0.071	(0.257)	0.067	(0.249)
Tyrol	0.090	(0.286)	0.091	(0.288)	0.087	(0.282)
Vorarlberg	0.000	(0.000)	0.049	(0.217)	0.000	(0.000)
Industry of employer:						
Agriculture	0.015	(0.120)	0.012	(0.108)	0.021	(0.142)
Electricity	0.007	(0.085)	0.009	(0.095)	0.003	(0.057)
Mining	0.006	(0.078)	0.007	(0.082)	0.005	(0.068)
Manufacturing	0.344	(0.475)	0.401	(0.490)	0.225	(0.417)
Construction	0.165	(0.371)	0.190	(0.392)	0.114	(0.318)
Wholesale and retail trade	0.150	(0.357)	0.159	(0.366)	0.130	(0.336)
Gastronomy, hotel business	0.060	(0.237)	0.046	(0.209)	0.089	(0.285)
Transportation	0.059	(0.236)	0.060	(0.238)	0.057	(0.233)
Finance	0.082	(0.274)	0.058	(0.235)	0.130	(0.337)
Cleaning, body care	0.009	(0.096)	0.008	(0.089)	0.012	(0.109)
Arts, entertainment, sports	0.010	(0.102)	0.005	(0.072)	0.021	(0.145)
Healthcare, welfare	0.014	(0.116)	0.006	(0.079)	0.029	(0.167)
Education, research	0.015	(0.121)	0.000	(0.000)	0.037	(0.189)
Lobbies, social security agencies	0.063	(0.243)	0.033	(0.180)	0.125	(0.331)
Housekeeping	0.000	(0.000)	0.000	(0.022)	0.000	(0.000)
Number of observations	1,174,523		797,846		376,677	

Notes: Table entries are sample means and standard deviations (in parentheses). Daily wages are given in Euros. Real wage are deflated using the consumer price index with base year 2007. GfGr ("Gerinfügigkeitsgrenze") and HBGr ("Höchstbemessungsgrundlage") denote the upper and the lower censoring point with respect to earnings, respectively. "Below GfGr" is an indicator taking on the value 1 if the nominal daily wage is equal to or below 1.2 times the lower censoring point and 0 otherwise. "Above HBGr" is an indicator taking on the value 1 if the nominal daily wage is equal to or above 0.8 times the upper censoring point and 0 otherwise.

B Additional Tables

Table B.1: Summary statistics, first job changes

	Type of job change			
	All	Involuntary	Voluntary	
			All	Immediate
(a) Wages, first and second employer				
Wage, first employer	57.304	56.158	58.421	66.364
	(21.609)	(18.984)	(23.840)	(25.536)
Wage, second employer	63.762	60.153	67.279	73.903
	(21.344)	(18.604)	(23.177)	(24.464)
Wage change	6.458	3.995	8.859	7.539
	(18.695)	(19.264)	(17.798)	(15.168)
(b) Time between first and second job				
Non-employment days	386.737	593.257	185.397	0.000
	(802.105)	(936.724)	(577.105)	(0.000)
Unemployment days	65.688	133.066	0.000	0.000
	(136.320)	(169.348)	(0.000)	(0.000)
Fraction unemployed	0.379	0.768	0.000	0.000
	(2.218)	(3.109)	(0.000)	(0.000)
Number of observations	182,922	90,297	92,621	37,900

Notes: Table entries are means and standard deviations (in parentheses). Panel (a) shows wages in the last year (first year) of the old (new) employer as well as the difference between the two. Panel (b) shows the number of days spent in nonemployment and unemployment, respectively, between a worker's first and second job since entry into the labor force. Involuntary job changes are those with at least one unemployment day between first and second job. Immediate voluntary job changes are those with zero gap between first and second job.