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Unemployment Effects of the German Minimum Wage in an Equilibrium Job Search Model

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We structurally estimate an equilibrium search model using German administrative data and use the model for counterfactual analyses of a uniform minimum wage. The model with worker and firm heterogeneity does not restrict the sign of employment effects a priori; it allows for different job offer arrival rates for the employed and the unemployed and lets firms optimally choose their recruiting intensity. We find that unemployment is a non-monotonic function of the minimum wage level. Effects differ strongly by labor market segment defined by region, skill, and permanent worker ability.

Keywords: monopsony, wages, employment, productivity, structural estimation
JEL-Code: J31; J38; J42; J64

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1. Introduction

A large number of empirical studies have examined the labor market effects of minimum wages (see Dube, 2019 and Neumark, 2019 for recent surveys). Most of these studies have found only small negative effects on employment. The absence of negative employment effects is often used as an argument by proponents of minimum wage increases. However, the ex-post studies of the effects of actually observed minimum wage levels provide only limited guidance about the potential effects of minimum wage increases if these effects are characterized by non-linearities (Bauducco and Janiak, 2018; Christl et al., 2018; Neumark, 2019).

In this paper, we structurally estimate an equilibrium job search model that takes such non-linearities into account. Because of search frictions, employers have market power that allows them to set wages below the marginal productivity of labor. The effects of firms' market power on wages have received increasing attention in the literature, accompanied by a resurgence of interest in monopsony models (e.g., Ashenfelter et al., 2022; Manning, 2021). Azar et al. (2019) explicitly focus on how the employment effects of a minimum wage depend on the degree of firms' monopsony power. While Azar et al. measure monopsony power using regional variation in the concentration of online job postings, our study derives monopsony power from structurally estimated parameters.

Our model is based on the wage-posting model by Bontemps et al. (1999). The model accounts for heterogeneity in both firms' productivity and unemployed workers' reservation wages. It does not restrict the sign of unemployment effects of minimum wages a priori and allows for non-linearities in the effects. Following Shephard (2017), we extend the model to allow for different job offer arrival rates for the employed and the unemployed and let firms optimally choose their recruiting intensity. This labor demand channel – along with the heterogeneity in reservation wages – is important for understanding the unemployment effects of the minimum wage (see, for instance Teulings, 2000).

We study these effects in the context of Germany, which introduced a statutory minimum wage in 2015. Prior to this, minimum wages had existed only at the sectoral level in a small number of industries. The minimum wage introduced in 2015 was set at a uniform level of EUR 8.50.[1] Estimating our model for a country in which a minimum wage was only recently introduced has an important advantage. The shapes of the heterogeneity distributions are important determinants of the magnitude of minimum wage effects. However, data from a market with a minimum wage are not informative on the shapes of the left-hand tails of these distributions, as the minimum wage effectively left-truncates wage outcomes (see e.g. Bontemps et al., 1999). To study counterfactual minimum wage effects, it is useful to have data from periods without minimum wages, as the latter data enable identification of

[1] While a number of transitional measures respected existing collective agreements and those signed in the meanwhile, the uniform minimum wage applied to all industries by 2017 at the latest. A further transitory exemption was given to those industries where industry-specific minimum wages had already been introduced prior to 2015 via the Posting of Workers Act (*Arbeitnehmerentsendegesetz*). The bargaining parties in an industry subject to this legislation may request that the Federal Ministry of Labor declares its (minimum wage) agreement to be generally binding for the entire industry.

the heterogeneity distributions across agents on larger parts of their support. In particular, in the German context, data from before 2015 allow for identification of the effects of minimum wages below the minimum wage that was imposed in 2015. In the paper, we return to the policy change in 2015 at various instances. Furthermore, we use the extensive quasi-experimental literature to validate the out-of-sample predictions of our model.

Our empirical analysis relies on a large administrative data set, the IAB Sample of Integrated Employment Biographies (SIAB). The SIAB is a two per cent random sample of individuals subject to social security contributions during the time period 1975 to 2014. We focus on data from the period 2010–2013 and leave out 2014 because of potential anticipation effects. The SIAB data provide an ideal basis for estimating a structural equilibrium search model for several reasons. First and most importantly, the data allow us to precisely measure the duration of different labor market states and transitions between them, notably job-to-job as well as employment-to-unemployment transitions. These transitions are crucial to the identification of the model's central parameters, such as job arrival and destruction rates.

Second, as the data are based on employers' notifications to the social security authorities, they are less prone to measurement error than comparable information from survey data. Additional advantages over survey data include the larger sample size and absence of panel attrition. We focus on low- and medium-skilled individuals because for these groups the assumption of a wage posting model is more convincing than for high-skilled individuals. The SIAB data do not include information on hours worked. We therefore focus on full-time employment spells and disregard individuals who are employed part-time during the time period under consideration.

Estimating the model based on this sample, we demonstrate that our model does well in replicating unemployment rates and 1-year labor market transitions. We also show that the model's ability to replicate the data varies across different labor market segments. In addition to differences between East and West Germany, we document differences by skill group and permanent worker ability.

Using our structural model, we find that the introduction of the minimum wage of EUR 8.50 in 2015 had a small positive effect on the (full-time) employment of the group that we study, i.e. low- and medium-skilled individuals. According to our simulations, the unemployment rate falls from 8.4% (in the baseline level without a minimum wage) to 8.1%, a decrease of 0.3 percentage points. The positive effect is driven by West Germany, where unemployment falls by 0.4 percentage points (5% of the benchmark level). For East Germany, the introduction of the minimum wage leads to an increase in the unemployment rate of 1.8 percentage points, or 20% of the benchmark level.

Our paper contributes to several strands of the literature on minimum wage effects. A first strand of literature has evaluated the introduction of the minimum wage in 2015 using quasi-experimental variation. These reduced-form studies have found no or at most a small negative effect on employment (see the surveys by Bruttel (2019) and Caliendo et al. (2019) and Section 7.4 below). Connecting our findings to this literature, we document that our model is able to generate unemployment responses that are broadly consistent with these studies. A second strand of literature structurally models unemployment responses to minimum wage policies based on models with perfect competition (Ragnitz and Thum, 2008; Bauer et al., 2009; Knabe and Schöb, 2009; Braun et al., 2020). In such a framework, the effects of a minimum wage on employment can by construction only be zero (if the minimum wage is not binding) or negative. The predictions from these studies stand in sharp contrast to those from the quasi-experimental literature as well as to our findings that the minimum wage leads to a reduction (or at most a slight increase) in unemployment.

Most importantly, our paper is related to the literature quantifying minimum wage effects using structural search and matching models. See, e.g., the papers by Flinn (2006), Ahn et al. (2011) and Engbom and Moser (2022).[2] Our paper is the first that uses the extended wage posting model with reservation wage heterogeneity to study the effect of minimum wage policies on unemployment.[3] Engbom and Moser (2022) use a similar wage-posting model, extended for heterogeneity in workers' ability and search efficiency, to study the role of the minimum wage in the decline of earnings inequality in Brazil. In our framework, worker heterogeneity derives from heterogeneity in reservation wages, i.e. workers differ in their reservation wages by nature. We additionally allow for heterogeneous worker ability by segmenting the sample along observed skill groups and quartiles of worker fixed effects from a decomposition in the spirit of Abowd et al. (1999). Unlike in Engbom and Moser (2022), reservation wages also differ within labor market segments and firms do not use the value of a job applicant's reservation wage when setting their wage offer. This implies that unemployed workers may either accept or reject wage offers.

Our structural approach is informative about the underlying transmission mechanisms and, importantly, allows us to assess the effects of counterfactual policies. We find that, in the German context of 2010–2013 and for full-time workers, unemployment is a non-monotonic function of the minimum wage level, which is in line with recent reduced-form evidence for the United States (Clemens and Strain, 2021). A major contribution of our study is therefore that it allows us to assess the scope of unemployment neutral minimum wage increases. For a relatively wide range of minimum wage levels, the unemployment rate is slightly lower than its benchmark level because a higher share of the unemployed receive acceptable wage offers. This unemployment-reducing effect tapers out at a minimum wage level of about EUR 12.50 because there is little mass left in the reservation wage density beyond this point. Thereafter, unemployment is almost exclusively frictional. At a minimum wage of EUR 13.50, the unemployment rate reaches its baseline level again from below. The search frictions and hence the unemployment rate then continue to grow as firms respond to higher minimum wages by lowering their recruiting intensity. Again, these results and

[2] A related literature analyzes structural models in which imperfect labor market competition arises from imperfect substitutability in workers' preferences across different jobs. The study by Hurst et al. (2022) embeds monopsonistic competition among homogenous firms into a directed search model framework to quantify short- and long-run distributional effects of an increase in the U.S. minimum wage. Berger et al. (2022) develop a similar framework with oligopsonistic labor markets to conduct a normative analysis of the optimal level of the minimum wage.

[3] Flinn (2006) applies a model with wage bargaining to analyze minimum wage effects; see also Breda et al. (2019). Finally, Drechsel-Grau (2022) opts for a model in which workers receive a fixed and exogenous share of output.

conclusions refer to full-time workers only.

Our estimates suggest that the different segments of the German labor market differ in the distribution of reservation wages, firm productivity, search frictions, and the ensuing degree of employers' market power. These differences mean that, while the general mechanisms of the minimum wage effects on unemployment are the same throughout, they operate with different strength and set in at different levels of the minimum wage.

The remainder of the paper is structured as follows: Section 2 gives a brief overview of the model. Section 3 provides a description of the data set and the construction of our main variables of interest, and Section 4 presents descriptive statistics. Section 5 outlines the estimation procedure. Section 6 contains the estimation results and graphical representations of the key steady-state relationships along with an analysis of the model's within-sample fit. Section 7 shows simulation results for the counterfactual introduction of different minimum wage levels and compares our findings with the quasi-experimental literature. Section 8 concludes.

2. Theoretical Model

In this section, we provide a brief description of the model. The framework is based on the wage-posting model by Bontemps et al. (1999), which is extended by allowing the job offer arrival rate to differ across employed and unemployed individuals and by letting firms optimally choose their recruiting intensity and thus the job offer arrival rates, as in Shephard (2017) and Engbom and Moser (2022).[4] We start by describing firms' and individuals' strategies. Individuals maximize their expected steady-state discounted future income and differ by unobserved opportunity costs of employment denoted by b, which may include search costs and unemployment benefits. Later, we will also allow individuals to differ along other dimensions, by segmenting labor markets by skill group and permanent worker ability among others (see Section 2.1). The distribution of b is denoted by H, assumed to be continuous over its support $[\underline{b}, \bar{b}]$. Job offers arrive at endogenous rate $\lambda_u > 0$ ($\lambda_e > 0$) for the unemployed (employed) and are characterized by a draw from a wage offer distribution F with support $[\underline{w}, \overline{w}]$.[5] Note that, in contrast to directed search models, random wage offers rule out that wages drive job seekers' application and search behavior.[6] Layoffs arrive at

[4]Based on a survey of workers in the U.S., Hall and Krueger (2012) show that about two-thirds of new hirings are characterized by wage posting, and that wage bargaining tends to be relevant mostly for high-skilled workers. Brenzel et al. (2011) come to similar results based on German data from the IAB Job Vacancy Survey. They find that wage posting represents 62 percent of all successful hirings in 2011, and that wage bargaining is predominantly relevant for job postings with higher skill requirements, a group that we exclude from our analysis.

[5]By imposing the assumption that each job is characterized by a single, time-invariant wage, the model abstracts from wage growth resulting from human capital accumulation or pure seniority. See, e.g., the model by Bagger et al. (2014) who explicitly incorporate human capital accumulation into a job search model. To address this issue, we conduct various robustness checks with respect to the wage definition in our estimations (see Section 6.4).

[6]The question of whether random or directed search provide a more realistic description of search behavior is difficult to answer. While the study by Banfi and Villena-Roldán (2019) provides some evidence for directed search, the authors emphasize that explicit wage information is rather selective and that most job advertisements at best convey only implicit wage information. We are not aware of any study exploring

constant rate δ. Unemployed individuals searching for a job face an optimal stopping problem, the solution to which consists in accepting any wage offer w such that $w > \phi$. Employed individuals, in contrast, accept any wage offers strictly greater than their present wage contract. By imposing this assumption, our model does not allow for job-to-job transitions that are associated with wage cuts.[7] As in Mortensen and Neumann (1988) and Bontemps et al. (2000), the reservation wage is implicitly defined as

$$\phi = b + (\kappa_u - \kappa_e) \int_{\phi}^{\overline{w}} \frac{\overline{F}(x)}{1 + \frac{\rho}{\delta} + \kappa_e \overline{F}(x)} dx, \tag{1}$$

where ρ denotes individuals' discount rate, $\overline{F}(x) = 1 - F(x)$, and $\kappa_i = \frac{\lambda_i}{\delta}$, $i = u, e$. The distribution of reservation wages, A, is then given by

$$A(\phi) = H\left(\phi - (\kappa_u - \kappa_e) \int_{\phi}^{\overline{w}} \frac{\overline{F}(x)}{1 + \frac{\rho}{\delta} + \kappa_e \overline{F}(x)} dx\right). \tag{2}$$

Equating equilibrium flows into and out of unemployment[8], the number of unemployed individuals with a reservation wage no larger than ϕ for $\phi \leq \underline{w}$ is represented by

$$uA_u(\phi) = \frac{1}{1 + \kappa_u} A(\underline{w}). \tag{3}$$

For $\phi > \underline{w}$, the number is given by

$$uA_u(\phi) = \frac{1}{1 + \kappa_u} A(\underline{w}) + \int_{\underline{w}}^{\phi} \frac{dA(x)}{(1 + \kappa_u \overline{F}(x))}. \tag{4}$$

From this, one can derive the steady-state equilibrium unemployment rate as

$$u = \underbrace{\frac{1}{1 + \kappa_u} A(\underline{w})}_{\substack{\text{(1) unemployed} \\ \text{who accept any job offer}}} + \underbrace{\int_{\underline{w}}^{\overline{w}} \frac{dA(x)}{(1 + \kappa_u \overline{F}(x))}}_{\substack{\text{(2) unemployed} \\ \text{who accept/reject offers}}} + \underbrace{(1 - A(\overline{w}))}_{\substack{\text{(3) unemployed} \\ \text{who accept no offer}}} \tag{5}$$

the prevalence of directed search in Germany.

[7] Postel-Vinay and Robin (2002) allow firms to counter outside offers that their workers receive from competing firms. Workers may accept lower wages than in their current job from more productive firms, by trading off lower wages today for increased chances of higher wages tomorrow. While our framework assumes that firms either cannot observe reservation wages or cannot pay different wages to workers with different reservation wages, the model by Postel-Vinay and Robin (2002) implies that firms offer unemployed workers their reservation wages. As a result, their model cannot capture positive employment effects of a minimum wage that arise from high-reservation wage workers accepting more job offers. Jolivet et al. (2006) allow for wage cuts by modelling reallocation shocks, which are drawn from the wage offer distribution and which workers cannot reject. In their model such reallocation shocks are formally equivalent to a layoff immediately followed by a job offer. While in their model the only alternative to acceptance is to become unemployed, heterogeneous reservation wages in our framework would create incentives for some workers to switch back to unemployment.

[8] For details see Bontemps et al. (1999, equations (2)–(5)).

Moreover, similar to Bontemps et al. (1999) one can show that in steady-state there exists a unique relationship between the unobserved distribution of wage offers and the observed distribution of earnings (i.e., *accepted* wages). Equating the flow of layoffs and upgraded wages of those with a wage lower than or equal to w and the flow of unemployed individuals accepting w, the distribution of earnings $G(w)$ is derived as

$$G(w) = \frac{A(w) - \left[1 + \kappa_u \overline{F}(w)\right] \left[\frac{1}{1+\kappa_u} A(\underline{w}) + \int\limits_{\underline{w}}^{w} \frac{1}{1+\kappa_u \overline{F}(x)} dA(x)\right]}{\left[1 + \kappa_e \overline{F}(w)\right](1-u)}. \tag{6}$$

Each firm offers only one wage and incurs a flow p of marginal revenue per worker. Firms are heterogeneous in their productivity p. The distribution of p across active firms is denoted by $\Gamma(p)$ and is assumed to be continuous over its support $[\underline{p}, \overline{p}]$. Following Shephard (2017), firms choose their optimal level of recruiting intensity ν, which allows them to alter the rate at which they encounter potential employees independent of the offered wage rate. The cost of recruiting effort, c, is a function of ν and p, such that $c(\nu, p)$ may differ across firms. The recruiting cost function takes the form $c(\nu, p) = c(p) \cdot \nu^{\eta}/\eta$, with $c(p) > 0$ and $c(0, p) = 0$ for all p. To ensure convexity of this function, η needs to be greater than one.

The number of workers that a firm attracts at wage w and recruiting intensity ν is denoted by $l = l(w, \nu)$. In what follows, the conditional firm size will be defined as $l(w, \nu) = \bar{l}(w) \cdot \nu / V$, with V representing the aggregate recruiting intensity:

$$V = \int\limits_{\underline{p}}^{\overline{p}} \nu(p) d\Gamma(p),$$

where $\nu(p)$ denotes the recruiting intensity of a firm with productivity p. The number of workers, $\bar{l}$, per unit intensity attracted by a firm that offers wage w solves

$$\bar{l}(w) = \frac{d(1-u)G(w)}{dF(w)},$$

and therefore

$$\bar{l}(w) = \frac{\kappa_e A(w)}{\left(1 + \kappa_e \overline{F}(w)\right)^2} + \frac{\kappa_u - \kappa_e}{\left(1 + \kappa_e \overline{F}(w)\right)^2} \left[\frac{1}{1+\kappa_u} A(\underline{w}) + \int\limits_{\underline{w}}^{w} \frac{1}{1+\kappa_u \overline{F}(x)} dA(x)\right]. \tag{7}$$

It can be shown that $\bar{l}(w)$ is a non-decreasing function of the offered wage. Note that the last term distinguishes $\bar{l}(w)$ from the original model by Bontemps et al. (1999), where $\lambda_u = \lambda_e$ and therefore $\kappa_u = \kappa_e$. The term reflects that if $\lambda_u \neq \lambda_e$, the number of employed and unemployed individuals that are attracted by the firm at a wage w may differ from each other.

Each firm seeks to maximize its steady-state profit flow, by choosing its optimal wage

$w(p)$ and recruiting intensity $\nu(p)$. The latter are determined by

$$(w(p), \nu(p)) = \underset{w(p),\nu(p)}{\operatorname{argmax}} \left[\overline{\pi}(w, p) \cdot \frac{\nu}{V} - c(\nu, p) \right].$$

$\overline{\pi}(p, w) = (p - w) \cdot \bar{l}(w)$ represents the expected profit flow per unit intensity, with $\bar{l}(w)$ denoting the size of a firm's labor force per unit intensity, such that $l(w, \nu) = \bar{l}(w) \cdot \nu / V$. The first-order condition defining the optimal recruiting intensity, ν, is given by

$$\frac{\overline{\pi}(w(p), p)}{V} = c(p) \cdot \nu(p)^{\eta - 1}.$$

Following Shephard (2017), we set $\nu(p) = 1$ in the benchmark, such that $c(p) = \overline{\pi}(w(p), p)$ in the pre-reform setting.[9] With $w = K(p)$ denoting the function that maps the support of the productivity distribution Γ into the support of the wage offer distribution F, we have $F(K(p)) = \int_{\underline{p}}^{p} \nu(y)/V d\Gamma(y)$. With $\nu(p) = 1$ in the benchmark, $F(K(p)) = \Gamma(p) = \Gamma(K^{-1}(w))$. The solution to the optimal wage setting problem of a p-type firm is represented by

$$K(p) = p - \left\{ \frac{\kappa_u(\underline{p} - \underline{w})}{(1 + \kappa_u)(1 + \kappa_e)} A(\underline{w}) + \int_{\underline{p}}^{p} \bar{l}(K(y)) dy \right\} \frac{1}{\bar{l}(K(p))}. \tag{8}$$

To complete the model, the total flow of matches is given by $M(V, S)$, with V denoting the aggregate recruiting intensity as defined above. S is the number of employed and unemployed individuals weighted by their search effort, i.e. $S = s_u \cdot u + s_e \cdot (1 - u)$, with s_u and s_e being the search effort of unemployed and employed individuals, respectively. M is assumed to increase in both, V and S, and to be concave and linearly homogeneous. The model is closed by specifying unemployed and employed individuals' job offer arrival rates, λ_i, with $i = u, e$, as the search effort weighted meeting rates, such that $\lambda_i = s_i \cdot M(V, S)/S$. The matching function is parametrized as Cobb-Douglas, i.e. $M(S, V) = V^{\theta} \cdot S^{(1-\theta)}$. As in Shephard (2017), we set θ equal to 0.5.

2.1. Labor Market Segments

The model assumes that worker productivity is homogeneous and that worker heterogeneity derives solely from unobserved opportunity costs. This implies, for instance, that all workers draw offers from the same wage offer distribution. To relax this assumption, we follow Bontemps et al. (1999) and estimate the model by different labor market segments. In this way, we treat each segment as a separate labor market characterized by its own structural parameters and its own distributions of reservation wages and firms' productivities. Because each segment will feature its own wage offer distribution, this approach assumes that firms are able to observe the characteristics defining the segments. A further underlying assumption is that there is no mobility between segments and no competition among firms across different

[9] As shown by Shephard (2017, Appendix F.1), the worker equilibrium does not depend on the assumptions concerning the recruiting cost function.

segments. As will be motivated in more detail below, we segment the labor market by region, skill group and permanent worker ability (see Section 3).

Apart from imposing homogeneous worker productivity, the model makes further assumptions, such as a uniform wage distribution for the employed and unemployed and exogenous job destruction rates. To assess the plausibility of these assumptions across labor market segments, we will use the separate estimations to explore the in-sample model fit by labor market segment (Section 6.3).

3. Data

Sample Selection and Variables Our empirical analysis uses German register data, the IAB Sample of Integrated Employment Biographies (SIAB). This administrative data set, which is described in more detail by Ganzer et al. (2017), is a two per cent random sample of all individuals who have at least one entry in their social security records between 1975 and 2017 in West Germany and between 1991 and 2017 in East Germany, respectively. The SIAB data cover approximately 80 per cent of the German workforce, providing longitudinal information on the employment biographies of 1,833,313 individuals. Self-employed workers, civil servants, and individuals doing military service are not included in the SIAB.

The data provide an ideal basis for estimating a structural equilibrium search model for several reasons. First and most importantly, the data contain daily information on employment records subject to social security contributions, unemployment records of benefit recipients as well as of registered job seekers. This permits us to precisely measure the duration of different labor market states and the transitions between them, notably job-to-job transitions as well as transitions between employment and unemployment (while receiving or not receiving benefits). Second, due to their administrative nature the data are less prone to measurement error than comparable information from survey data. Additional advantages over survey data include the larger sample size and a much more limited degree of panel attrition.

Sample selection proceeds in several steps. Before restricting the sample to a specific time span and population, we fill in missing values using all the information available in the full dataset (see Appendix A.1). For the estimation of our model, we construct a stock sample by keeping only those employment and unemployment spells[10] including the set date 1 January 2010 and restrict the sample to the period 2010 to 2013, i.e., the years before the introduction of the minimum wage on 1 January 2015. We omit 2014 so that our estimates are not affected by the potential anticipation of the minimum wage. This leads to a sample of 688,869 individuals.

From this sample we select only individuals who are part of the workforce. The data do not make it possible to distinguish between involuntarily unemployed individuals not receiving benefits and individuals who voluntarily left the labor force or who became self-employed or civil servants. To distinguish more precisely between voluntary and involuntary unemploy-

[10] Details on the definition of the different labor market states are given in Appendix A.2.

ment, we follow the assumptions proposed by Lee and Wilke (2009) (see Appendix A.2).

To focus on individuals in the workforce, we restrict the sample to individuals who are at least 20 years old and younger than 63 years. The sample is further restricted to low- and medium-skilled individuals.[11] We exclude highly skilled individuals because this group is less likely to be in a labor market that is characterized by a wage-posting mechanism. We then drop individuals who have missing values in the relevant observables such as daily wages and employment status, or in the variables that we use to define the sub-samples (region, skill level, permanent worker ability). Furthermore, we exclude agricultural jobs because their employment durations are often characterized by seasonality. This leads to a new sample size of 283,180 individuals.

The SIAB data do not include information on hours worked. We therefore focus on full-time employment spells and disregard individuals who are predominately part-time employed during the time period under consideration. Moreover, we exclude individuals if their relevant wage information comes from a part-time spell.[12] By excluding part-time work, we also leave aside marginal employment in the form of "minijobs," i.e. jobs that are exempt from income taxation and social security contributions (cf. Tazhitdinova, 2020). This minijob sector is relatively sizeable in Germany. In 2014, shortly before the introduction of the statutory minimum wage, marginal employment accounted for 14% of total employment. In 2015, right after the imposition of the minimum wage, marginal employment experienced a decline of about 3%. The focus on full-time work is admittedly a limitation of our analysis. We face a trade-off as our administrative dataset is well-suited for our analysis in general, but does not have information on hours, so the dataset would allow an analysis of part-time work only under very strong assumptions. Hours information is available in the Socio-Economic Panel (SOEP), but by using a much smaller survey dataset we would have to forego the advantages of our administrative dataset outlined in the introduction. Independently of the data, our search model is not an appropriate framework for capturing transitions between part-time and full-time work or the conversion of marginal employment into regular employment subject to social security contributions, as these conversions typically take place within the same firm.[13]

To calculate hourly wages for full-time employment spells, we impute the number of hours worked based on information from the German Microcensus. The imputation is done separately by region, sex, sector, job classification, and skill group.[14]

In the model, each job is characterized by a single, time-invariant wage. For individuals

[11]Details on the definition of the different skill groups are given in Appendix A.1.

[12]In our sample, 14.5% of our employees are predominantly part-time employed. For these workers, the median hourly wage amounts to EUR 11.28 (p25: 7.56; p75: 16.58). For predominantly full-time workers the median is EUR 17.09 (p25: 12.62; p75: 22.41).

[13]For instance, vom Berge et al. (2016) show that 50% of the post-minimum wage decline in marginal employment can be explained by such a conversion of marginal into regular employment, mostly within the same employer. Moreover, because marginal employment considerably facilitated a misreporting of hours, the introduction of the minimum wage was accompanied by stricter regulations on documenting working hours. While we are not aware of any causal evidence, this prompts the conjecture that part of the conversion of marginal into regular employment was due to these stricter regulations.

[14]For details, see Appendix A.3.

who were employed on 1 January 2010, we compute this wage as the weighted average of the wages earned over the past year in the same job, where the weights are given by the length of time over which a particular wage was received. Likewise, the wage after an unemployment-to-employment spell is based on the weighted average over the first year after the transition.[15] To reduce the influence of outliers, we discard observations with implausibly low hourly wages (wages below EUR 3 or below the existing sectoral minimum wages). As will be explained below, we split our sample by quartiles of person fixed effects, in order to define our labor market segments based on permanent worker ability. To further mitigate the influence of outliers within quartiles, we truncate the resulting quartile-specific distributions at their first and 99th percentile. Overall, this procedure results in a lowest observed wage of EUR 3.64, which is close to what has been found based on survey data (Burauel et al., 2017). The resulting final sample contains information on 208,626 individuals.

The wage information in the IAB data is censored since there is an upper contribution limit in the social security system. We do not include observations with censored wages.[16]

Definition of Labor Market Segments As spelled out in Section 2.1, estimating the model separately by labor market segments involves the assumption that there is no mobility between segments and no competition among firms across different segments. As individuals of different gender or age are likely to compete within one segment, we define the segments based on two regions (East Germany and West Germany including Berlin), two skill groups (low- and medium-skilled) and individual-specific (or “permanent”) worker ability (see Appendix A.6). The latter dimension is supposed to capture time-invariant productivity characteristics other than the formal educational attainment, which are typically not directly observed in data. To operationalize this “permanent worker ability”, we employ wage decompositions along the lines of Abowd et al. (1999) (henceforth referred to as AKM). Specifically, in a first stage, we use worker-specific fixed effects that were estimated on the universe of the Integrated Employment Biographies (IEB) in Germany (cf. Card et al., 2013, 2015), controlling for establishment-specific fixed effects. The details of this decomposition are in Appendix A.7.[17] We use the worker-specific fixed effects (or “person fixed effects”, or PFE) to break down labor markets by ability types, assigning individuals to quartiles of the PFE distribution. Next, we estimate and simulate the model separately for the labor market segments delineated by these quartiles.[18]

Overall, these dimensions allow us to define fairly well (though not perfectly) segmented labor markets in terms of worker mobility. Regarding region and skill groups, Tables A.1 and A.3

[15] For details, see Appendix A.4.

[16] For details, see Appendix A.5. In a robustness check, we address this issue by replacing censored observations with imputed wages, following Gartner (2005).

[17] The information can be merged for the vast majority of individuals in our sample.

[18] One may consider a closer integration of the fixed effects into the structural model. This would not be trivial. There is an ongoing debate on the compatibility of AKM decompositions with search and matching models. For example, wage decompositions have been criticized for being ill-suited to identify the degree of sorting (see, e.g., Eeckhout and Kircher, 2011). As a more practical issue, the establishment-specific fixed effects that are identified from the German administrative data do not straightforwardly match the concept of an employer or a firm.

show that 95.9% of employment-to-employment transitions remain in the same region, and 98.6% remain in the same skill group. As for unemployment-to-employment transitions, 95.2% occur within the same region and 96.8% within the same skill group (see Tables A.2 and A.4).[19] While a transition from low to medium-skilled may measure a true acquisition of a vocational degree, it may also reflect some measurement error. The latter may arise from employers reporting a job's skill requirements rather than individuals' acquired degree. Finally, defining segments based on permanent worker ability has the advantage that the time-constant nature (over the period considered here) of person fixed effects rules out any transitions between different ability segments.

Recruiting Costs To get an estimate for the recruiting cost function for Germany, we use the IAB Job Vacancy Survey (JVS). The advantage of using these data over indirect information from the firm size distribution is that the JVS contains a direct measure of search costs when filling a vacancy. In general, the JVS provides information on the number of vacancies and information on employers' most recent hiring process. For the latter, employers are asked about the fixed costs of search and the hours spent on search since the year 2014. Similar to Carbonero and Gartner (2022), we assume that total recruiting costs, c, are determined by $c = c_{\text{fix}} + lc \cdot h$, with c_{fix} denoting fixed costs of search, lc hourly labor costs and h denoting hours spent on search. We set hourly labor costs to between EUR 31.80 and EUR 36.10, based on yearly information from the Federal Statistical Office for the period 2014 to 2019. In our model, we impose a recruiting cost function taking the form $c = c(p) \cdot v^{\eta}$, with v denoting the number of vacancies and $c(p)$ being a function of firm productivity. To estimate η, we account for the fact that the JVS provides information for the most recent hiring process, which allows us to interpret the information on recruiting costs as marginal costs. The latter can be expressed as

$$\frac{dc}{dv} = \eta \cdot c(p) \cdot v^{\eta-1}. \tag{9}$$

We present the estimation results and sensitivity tests with regard to the parameter η in Section 6.5.

4. Descriptives

4.1. Transitions

Tables A.5 and A.6 in Appendix A.8 report the type, number, and share of transitions for our stock sample of individuals who were either unemployed (8.1%) or employed (91.9%) on 1 January 2010. Of the 191,710 individuals who were employed on this date, 69% stayed in their job for the next four years while 20% moved to another job and 11% became unemployed. Transitions in the other direction are much more frequent in relative terms:

[19] Note that there are no transitions from medium skill to low skill because of our imputation procedure, which makes sure that individuals cannot lose their degree.

48% of the 16,916 unemployment spells ended with a transition into regular employment during the four-year period after 1 January 2010. At the same time, 52% of individuals who were unemployed on this date remained without a job over the entire period.[20]

The table also breaks down these statistics by region, skill group and permanent worker ability. About 84% of the individuals in the sample worked or searched for a job in West Germany (including Berlin), the remaining 16% in East Germany. On 1 January 2010, the unemployment rate was higher in East Germany (11%) than in West Germany (8%). However, the fraction of unemployed individuals finding a new job over the four-year observation window was almost identical in East and West Germany (49% versus 48%). Looking at transitions of employed individuals, we find that most individuals stayed at their current employer, while around 20% of employed individuals in both West and East Germany changed their employer within the four years. The relative frequency of transitions into unemployment was higher in East Germany (13%) than in West Germany (10%).

Breaking down the sample by skill group, the unemployment rate for low-skilled individuals was twice as high as for the medium-skilled. As for permanent worker ability, the unemployment rate on 1 January 2010 varied between 19% in the bottom PFE quartile and 3% in the upper PFE quartile. The share of unemployment-to-employment transitions was highest in the second and third PFE quartile and lowest in the bottom and upper PFE quartile.

4.2. Durations

Figure A.1 in Appendix A.8 shows non-parametric Kaplan-Meier estimates of the survival function for remaining in the initial state (employment or unemployment) for the whole sample. The survival functions are also shown for the different sub-samples defined by region (Figures A.2), skill group (Figure A.3), and permanent worker ability (Figure A.4). In our estimation sample, the maximum duration of an unemployment spell is nine years.[21] Employment spells can in principle last over the whole observation period: 39 years in West Germany (1975–2013) and 22 years in East Germany (1992–2013).[22]

[20] Left-censoring can occur for the unemployment spells because in some of the data sources for unemployment benefit histories, recording starts at a fixed date which does not necessarily coincide with the beginning of the unemployment spell (see Appendix A.2).

[21] "Unemployment benefit I" (ALG I), a non means-tested transfer which is part of the unemployment insurance system, is typically paid for only one year (two years for older workers). Once ALG I runs out, the unemployed are entitled to the much lower and means-tested "unemployment benefit II" (ALG II), which was introduced on 1 January 2005. Before 2005, ALG I was followed by "Arbeitslosenhilfe" instead of ALG II. This means that individuals receiving "Arbeitslosenhilfe" before 2005 were entitled to ALG II afterwards. However, spells of receiving ALG II are only recorded in the data from 1 January 2007 onwards. This makes 1 January 2005 the earliest starting point for unemployment spells in our estimation sample. These spells refer to those individuals who received ALG I benefits during 2005 and 2006 and who were entitled to ALG II afterwards (starting from 2007). As our sample covers the period 2010–2013, the maximum duration of an unemployment spell is nine years.

[22] 1.16% of the employment spells are left-censored which means employment without interruption at the same firm since 1 January 1975 in West Germany or since 1992 in East Germany. We disregard employment spells recorded in 1991 in East Germany.

Transitions out of Unemployment The chance of transitioning into employment is particularly high within the first year – about 60% of the unemployed were still without a job after twelve months (cf. panel (a) of Figure A.1). By the second year, about 50% of the unemployed had not found employment, and after the third year the survival function flattens out. As can be seen in panel (a) of Figure A.2, the pattern is similar for East and West Germany, but there is substantial variation across skill groups and permanent worker ability (Figure A.3 and Figure A.4). Low-skilled individuals' unemployment durations are higher than those of the medium-skilled. Regarding permanent worker ability, individuals in the bottom and upper quartile of the PFE distribution exhibit longer unemployment durations than those in the middle PFE quartiles.

Transitions out of Employment For individuals who were initially employed, transitions can be either into another job (panel (b) of Figures A.1 to A.4) or into unemployment (panel (c)). The durations of employment spells that end because of unemployment are in general longer than employment spells that end in a job-to-job transition. With regard to employment-to-employment transitions, the probability of still being employed at the current employer is typically around 75% after fifteen years. The durations of employment spells do not differ significantly between East and West Germany. Along the dimension of permanent worker ability, employment stability increases with workers' position in the PFE distribution. Both for transitions into other jobs and for transitions into unemployment, the share of the employed who have left their initial job for one of these destinations is lower in the upper PFE quartiles.

4.3. Wage Distributions

Figure A.5 in Appendix A.8 shows the distribution of wages before and after a labor market transition for the whole sample. As part of the descriptives, we include all three types of transitions (e $\rightarrow$ e, e $\rightarrow$ u, u $\rightarrow$ e) and also document the wage distributions for right- and left-censored spells. In the estimation, only the wages in the initial employment spell or after a transition from unemployment to employment will be used.

As expected, wages of individuals who change their job tend to be higher than wages before a transition into unemployment. Comparing wages before and after a job-to-job transition, we find that wages earned in the new job are on average slightly higher than the wages earned in the old position. Also in line with expectations, wages after an unemployment-to-employment transition tend to be relatively low. A sizeable fraction of the unemployed move to jobs paying less than EUR 8.50 an hour, the statutory minimum wage introduced in 2015, i.e. after our sampling period 2010–2013. This also holds within the different labor market segments defined by region (Figure A.6) and skill group (Figure A.7 in Appendix A.8). As for permanent worker ability, this share is typically very small for those in the upper quartiles of the PFE distribution (Figure A.8).

These figures also confirm the well-known facts that wages tend to be lower in East Germany and tend to increase with workers' skills and their position in the PFE distribution.

5. Estimation

We begin this section by deriving the likelihood contributions of unemployed and employed workers, taking into account stock sampling as well as left- and right-censoring. We then outline the estimation procedure, which combines the likelihood function with a non-parametric estimate of the wage distribution.

Likelihood – Unemployed Workers As seen in Equation (5), the steady-state unemployment rate has three components. For individuals with low enough opportunity costs of employment, unemployment is purely frictional. In a second group, unemployment is driven by both search frictions and the opportunity cost of employment; these individuals will accept some job offers, but reject others. Finally, there is a third group for whom unemployment is permanent given the wage offer distribution F, as any wage offer is below their reservation wage. As a result, the likelihood contribution of an individual who is initially unemployed is a mixture distribution:

$$\lambda_u^{2-d_{ub}-d_{uf}} \cdot e^{-\lambda_u\left(t_{ub}+t_{uf}\right)} \cdot \frac{A(\underline{w})}{1+\kappa_u} \cdot f(w_u)^{1-d_{uf}}$$
$$+ \int_{\underline{w}}^{d_{uf}\overline{w}+(1-d_{uf})w_u} \left(\lambda_u \overline{F}(x)\right)^{2-d_{ub}-d_{uf}} \cdot e^{-\lambda_u \cdot \overline{F}(x)\cdot\left(t_{ub}+t_{uf}\right)} \cdot \frac{f(w_u)}{\overline{F}(x)}^{1-d_{uf}} \cdot \frac{dA(x)}{1+\kappa_u \overline{F}(x)}$$
$$+ [1-A(\overline{w})] \cdot d_{ub} \cdot d_{uf}. \quad (10)$$

The first term of the sum corresponds to purely frictional unemployment. As job offers arrive with Poisson rate λ_u, unemployment durations are exponentially distributed. In a flow sample, where the elapsed ("backward") duration t_{ub} is zero by definition, the density of the residual ("forward") duration t_{uf} is given as $h(t_{uf}) = \lambda_u \exp(-\lambda_u t_{uf})$. In a stock sample, we need to consider the total duration $t_{ub} + t_{uf}$, conditional on the elapsed duration t_{ub}. The latter has the density $h(t_{ub}) = \lambda_u \exp(-\lambda_u t_{ub})$. It can be shown (e.g., Lancaster, 1990) that the conditional density $h(t_{uf}|t_{ub})$ is given as $\lambda_u \exp(-\lambda_u t_{uf})$. For the joint density we then obtain $h(t_{ub}, t_{uf}) = h(t_{uf}|t_{ub})h(t_{ub}) = \lambda_u^2 \exp\left(-\lambda_u(t_{ub} + t_{uf})\right)$, which is the term that figures in the likelihood expression above. The term in front of the exponential function is adjusted if either the elapsed or the residual duration is censored ($d_{ub} = 1$ or $d_{uf} = 1$). $f(w_u)$ is the density function of wage offers evaluated at the offer that we observe as the initially unobserved person transits into employment. If the unemployment duration is right-censored ($d_{uf} = 1$), this term drops out of the likelihood function.

The second term of the sum has the same basic structure, but with some adjustments for the fact that individuals in this group are sometimes faced with wage offers that are below their reservation wage. The unemployment spell hazard rate is therefore given not by λ_u, but by the product $\lambda_u \overline{F}(x)$. The second adjustment concerns the wage offer density, which is now truncated at x, so we have $f(w_u)/\overline{F}(x)$.[23]

[23]Note that as $\overline{F}(x) = 1$ for $x < \underline{w}$, the first term of the sum could be integrated into the second term. We choose to present them separately here to better reflect the conceptual difference between the three

Finally, the third term applied to individuals who, given F, are permanently unemployed. This implies that the observed unemployment spell must be both left- and right-censored, hence the factor $d_{ub} \cdot d_{uf}$.

Likelihood – Employed Workers For individuals who are initially employed, the likelihood contribution is

$$(1-u) \cdot g(w_e) \cdot \left[\delta + \lambda_e \overline{F}(w_e)\right]^{1-d_{eb}} \cdot e^{-\left[\delta+\lambda_e \overline{F}(w_e)\right]\left(t_{eb}+t_{ef}\right)} \left[\delta^v \left(\lambda_e \overline{F}(w_e)\right)^{1-v}\right]^{1-d_{ef}}. \quad (11)$$

In steady state, a fraction $(1-u)$ of all individuals is employed. g is the density of wages in the initial job. Unlike for the unemployed, the reservation wage of a worker is observed and equals his or her current wage, so there is no mixing distribution for the durations. However, there are now two competing reasons for why a spell may end: layoff (at rate δ) or a better job offer (at rate $\lambda_e \overline{F}(w)$). The indicator v equals 1 in the first case and 0 in the second. t_{eb} denotes the elapsed duration, and t_{ef} the residual duration of the current job. d_{eb} equals 1 if the elapsed duration is left-censored, while $d_{ef} = 1$ means that the residual duration is right-censored, i.e. the individual does not change his or her job during the observation period.

Estimation Procedure Maximum likelihood estimation of the model requires functional form assumptions for H and Γ. The estimation is numerically cumbersome as f, g, and $\overline{F}$ are highly non-linear functions of Γ. In particular, optimization involves the numerical computation of the inverse K^{-1}, further complicated by the fact that K contains an integral that has to be evaluated numerically as well. Beyond these numerical concerns, there is the issue that most distributions for Γ imply wage distributions that do not fit the data well.

As an alternative, Bontemps et al. (2000) therefore propose a three-step procedure in which the wage distribution is estimated non-parametrically:

1. In a first step, we estimate G and g (the cdf and pdf of the wage distribution) using a kernel density estimator, and estimate $\underline{w}$ and $\overline{w}$ as the sample minimum and maximum of the wages of workers who are employed on 1 January 2010. Based on these non-parametric estimates and a parametric assumption for the opportunity cost distribution, namely $H \sim \mathcal{N}(\mu_b, \sigma_b^2)$, and setting $\nu(p) = 1$ in the benchmark, we obtain consistent estimates for $\overline{F}$ and f (given $\mu_b, \sigma_b, \lambda_u, \lambda_e, \delta$ and the assumption that $\rho = 0.02$) by numerically solving the following expressions (recall that u is a function of $\overline{F}$):

$$\widehat{\overline{F}}(w) = \frac{A(w) - uA_u(w) - (1-u)\widehat{G}(w)}{\kappa_e \cdot \widehat{G}(w) \cdot (1-u) + \kappa_u \cdot u \cdot A_u(w)} \quad (12)$$

and

$$\widehat{f}(w) = \frac{(1-u) \cdot \widehat{g}(w) \cdot (1 + \kappa_e \widehat{\overline{F}}(w))}{\kappa_u \cdot u \cdot A_u(w) + \kappa_e \cdot (1-u) \cdot \widehat{G}(w)}. \quad (13)$$

components behind the unemployment rate.

2. The estimates from Step 1 are plugged into the likelihood function, which is then maximized with respect to μ_b, σ_b, λ_u, λ_e, and δ.

3. Once these parameters are known, the productivity of a firm can be inferred from the wage that it offers:

$$\begin{aligned} p &= K^{-1}(w) \\ &= w + \left(\frac{\kappa_u \cdot A'(w) \cdot (1 + \kappa_e \cdot \widehat{\overline{F}}(w))}{(1 + \kappa_u \cdot \widehat{\overline{F}}(w)) \cdot (\kappa_e A(w) + (\kappa_u - \kappa_e)) \cdot u \cdot A_u(w)} + \frac{2 \cdot \kappa_e \cdot \widehat{f}(w)}{1 + \kappa_e \cdot \widehat{\overline{F}}(w))} \right)^{-1} \end{aligned} \tag{14}$$

Standard errors are obtained by bootstrapping the three-step procedure.

6. Estimation Results

6.1. Parameter Estimates

Table 1 reports the estimated parameters and the associated bootstrap standard errors.[24] For the whole sample, we estimate a monthly job destruction rate δ of 0.0063. The rate is about 20% higher in East Germany than in the West (0.0073 vs. 0.0061). Regarding permanent worker ability, the job destruction rates decline monotonically with workers' position in the PFE distribution, from 0.0126 in the lowest quartile to 0.0046 in the highest quartile. The same is true for the skill groups, with an estimated job destruction rate of 0.074 for the low-skilled versus 0.062 for the medium-skilled. These orders of magnitude are similar to earlier studies. For France in the 1990s, Bontemps et al. (1999) find a δ between 0.0032 and 0.0069, depending on the sector. Using SIAB data for an earlier period (1995–2000), Nanos and Schluter (2014) estimate the monthly layoff rate to be between 0.0032 and 0.0243 in Germany. Holzner and Launov (2010), who use data from the German Socio-Economic Panel 1984–2001, estimate a δ of 0.0047.

The estimated κ, i.e., the ratio of the job arrival over the job destruction rate, is greater for the unemployed than for the employed. We find κ_u to be 20.06 and κ_e to be 7.10. Holzner and Launov find a κ_e of 2.2, while their three values of κ for the unemployed (they assume that individuals search on skill-specific labor markets) range between 5.6 and 17.1. In their study for France, Bontemps et al. (2000) also estimate a much higher job arrival rate for the unemployed than for the employed. In all cases, this reflects that continental European labor markets are characterized by relatively little job-to-job mobility compared with the United States.

The differences between labor market segments are potentially relevant for the design of the new statutory minimum wage in Germany. After a transition period, the minimum wage became uniform for all workers by 2017 at the latest. Our results suggest that the uniform

[24] While multiple equilibria cannot be ruled out (see van den Berg, 2003), we have not found evidence of this for any of our estimated or simulated equilibria.

Table 1: Estimation results

	N	δ	κ_e	κ_u	μ_ϕ	σ_ϕ
Whole sample	208626	0.0063	7.10	20.06	6.50	3.60
		(0.0000)	(0.10)	(0.32)	(0.02)	(0.05)
Region						
West	176205	0.0061	7.51	20.85	6.79	4.11
		(0.0000)	(0.10)	(0.34)	(0.05)	(0.06)
East	32421	0.0073	5.11	18.23	5.90	3.48
		(0.0000)	(0.16)	(0.67)	(0.05)	(0.09)
Skill Group						
Low-Skilled	15542	0.0074	5.59	13.74	5.41	4.93
		(0.0001)	(0.30)	(0.84)	(0.17)	(0.23)
Medium-Skilled	193084	0.0062	7.33	20.63	6.59	3.44
		(0.0000)	(0.10)	(0.33)	(0.05)	(0.04)
PFE Quartile						
Quartile 1	37049	0.0126	1.97	7.44	4.67	4.09
		(0.0001)	(0.04)	(0.58)	(0.21)	(1.09)
Quartile 2	67970	0.0079	4.76	20.08	7.57	4.56
		(0.0000)	(0.07)	(0.39)	(0.04)	(0.07)
Quartile 3	69438	0.0053	6.28	32.34	10.36	4.46
		(0.0000)	(0.16)	(0.93)	(0.04)	(0.09)
Quartile 4	34169	0.0046	5.24	32.03	12.87	6.22
		(0.0000)	(0.23)	(1.53)	(0.07)	(0.09)

Note: The table reports the estimated parameters for the whole sample, and for subsamples by region, skill group and quartiles of person fixed effects. The monthly job destruction rate is denoted as δ. κ_u and κ_e denote the ratios of the job arrival over the job destruction rate for the unemployed and the employed, respectively. μ_ϕ and σ_ϕ denote the mean and the standard deviation of the reservation wage distribution. We set the discount rate ρ to 0.02. Bootstrapped standard errors in parentheses (250 runs).

rate applies to labor market segments that differ in the extent of search frictions and thus in firms' monopsony power on the labor market.[25]

As we restrict the parameters in the likelihood function to be positive, the distribution of the opportunity costs of employment has a mean μ_b close to EUR 0 per hour.[26] The standard deviation σ_b is estimated to be 3.6 for the whole sample. However, unlike in the

[25] Bachmann and Frings (2017) adopt a different approach to quantify labor market frictions in Germany and estimate labor supply elasticities specific to the individual firm. Using linked employer-employee data from the IAB (LIAB), the authors document substantial differences in employers' market power across industries. Their findings indicate that retailing, hotels and restaurants and agriculture feature a larger degree of monopsonistic power than other services and manufacturing of food products. Bachmann et al. (2022) document that the degree of monopsony power that workers in Germany are exposed to differs by the tasks they perform. Furthermore, Hirsch et al. (2022) show that due to the greater density of labor markets, employers have less wage-setting power in urban than in rural areas.

[26] While this restriction makes sense for the job offer arrival and destruction rates as well as for the variance of the opportunity costs, there is no strong reason to impose it for the mean. When we abolish the restriction, we find a mean that is actually slightly negative for the whole sample and many of the sub-samples. In fact, there are issues with the separate identification of μ_b and σ_b on the one hand and κ_u on the other hand (κ_e is pinned down by observed employment-to-employment transitions), which all enter the expression for the reservation wage distribution $A(\phi)$. The latter matters for the empirical predictions and is hardly affected by the lifting of the restriction.

model of Bontemps et al. (1999), the reservation wages are not identical to the opportunity costs of employment. This is because job offer arrival rates are higher when unemployed, so it is optimal for the unemployed to reject certain wage offers in the hope of getting a higher offer in the future (cf. Equation (1)). Based on the estimated parameters, we find that the distribution of reservation wages is centred around a value of about EUR 6.50 per hour. The reservation wage is higher and slightly more dispersed in the West and increases monotonically with workers' position in the PFE distribution and with the acquired skill level.

Due to the above restriction in our estimation, the differences in the reservation wages between the sub-samples are exclusively driven by differences in the frictional parameters. For instance, the difference between κ_u and κ_e is much larger in the fourth quartile than in the first quartile of the PFE distribution, which is reflected in a much higher μ_ϕ for workers with higher ability. Note that differences in κ_u and κ_e reflect differences in both job offer arrival and layoff rates. The higher the layoff rate, the smaller the expression $\beta \equiv \rho/\delta$ in Equation (1), and thus the smaller the incentive for the unemployed to be picky when accepting a wage offer – after all, accepting a job means giving up a higher job arrival rate. If the job has a higher probability of ending, the costs of accepting it in terms of foregone employment opportunities become smaller.

6.2. Distribution of Wages, Opportunity Costs, Markups, Productivity

Figure 1 shows key plots summarizing the steady-state equilibrium.[27] Panel (a) depicts our non-parametric estimate for g, the pdf of the wage distribution. The cdf G, which is not shown here, is similarly estimated using a kernel density estimator.

The wage offer distribution F (panel (b)) is obtained by combining the estimate for G with the maximum likelihood estimates for the frictional parameters and the opportunity cost distribution, as outlined in Section 5 above. The location and the shape of the wage offer distribution differ from the wage distribution. For instance, more than 70% of the wage offers, but only about 10% of observed wages are below EUR 10.

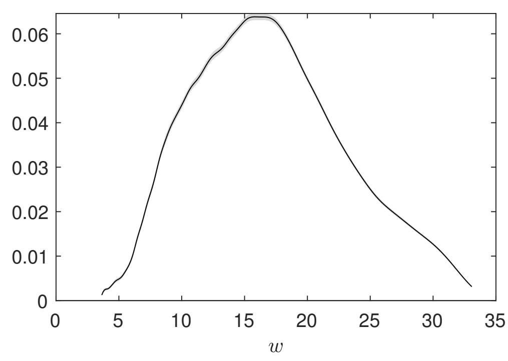
Panel (c) shows the estimated distribution of reservation wages. This is a normal distribution centred around EUR 6.50 and truncated at EUR 3.64, the lowest admissible hourly wage. Note that there is hardly any mass left beyond EUR 12.50.[28] This means that the unemployment-reducing effect of higher minimum wages operating through a lower rate of job offer rejections will be mostly limited to minimum wage levels below this amount.

Panel (d) presents the optimal wage offer as a function of firm productivity p. For example, a firm with a value product of EUR 20 per hour will optimally set a wage of about EUR 15 per hour. The function reveals a monotonically increasing relationship between wages and productivity.[29] As the slope of wage offer function decreases with productivity, high

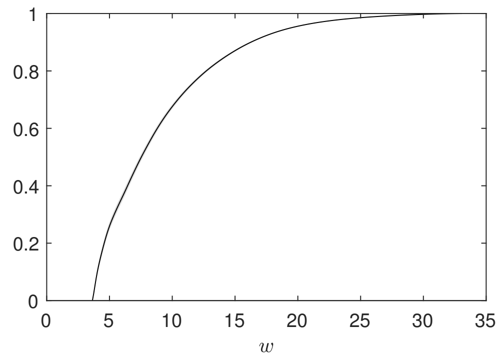
[27]The plots are for the whole sample. Plots by region, skill group, and permanent worker ability can be found in Figures A.9 to A.11 in Appendix A.9.

[28]The distribution of reservation wages is very close to the one estimated by Fedorets and Shupe (2021) based on a survey question from the Socio-Economic Panel.

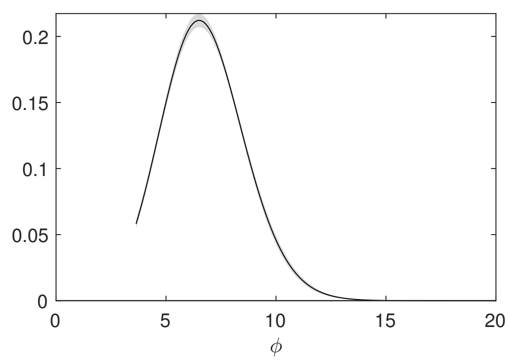
[29]As this pattern is not guaranteed by the non-parametric estimation procedure described in Section 5,



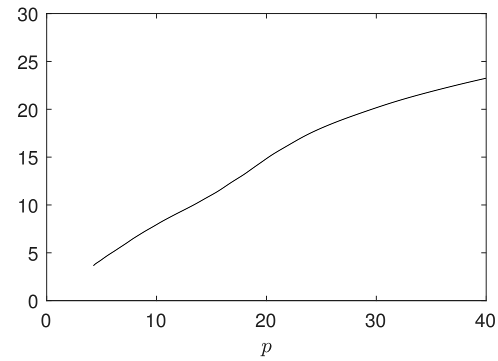
(a) Wage density $g(w)$



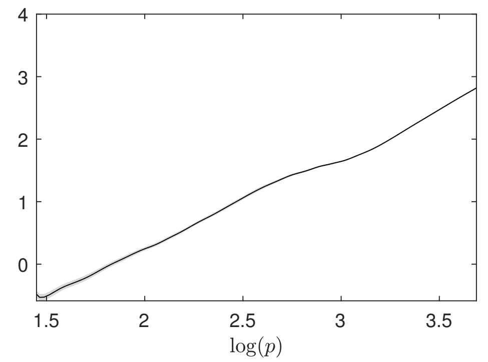
(b) Wage offer distribution $F(w)$



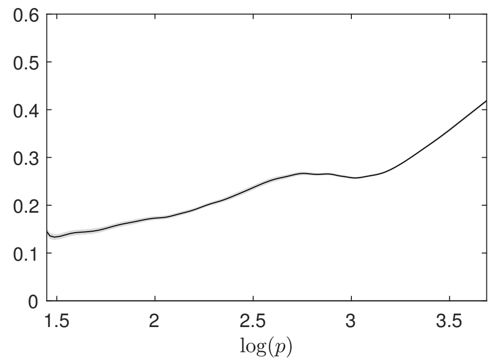
(c) Reservation wage density $dA(w)$



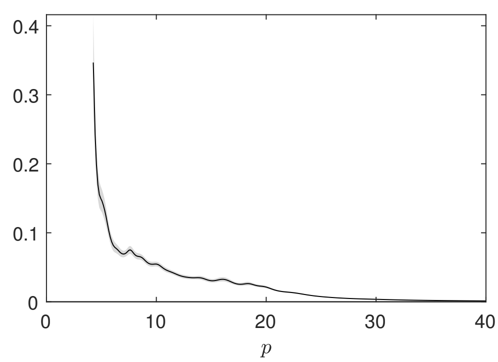
(d) Wage function $K(p)$



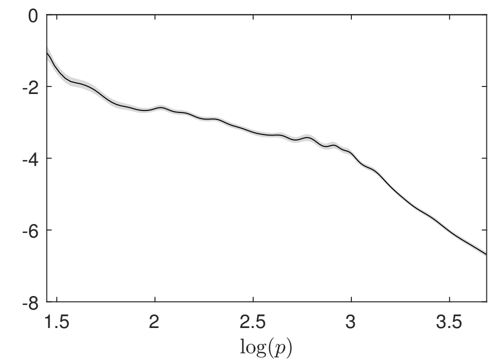
(e) Log absolute markup



(f) Monopsony power $[p - K(p)]/p$



(g) Productivity density $\gamma(p)$



(h) Productivity log density $\log(\gamma(p))$

Figure 1: Main Equilibrium Functions – Whole sample. *Note:* This figure shows the main equilibrium functions summarizing the steady-state for the estimates for the whole sample. Ordinate outcomes can be found in the subfigures' titles. Wages and productivity are measured in EUR. *Key:* Grey area indicates 95% confidence bands based on 250 bootstrap runs.

productivity firms exhibit more monopsony power.

The absolute markup, which is shown in a log-log-scale in panel (e), grows monotonically and at a fairly constant rate with a firm's productivity. Expressed as a percentage of productivity (panel (f)), the relationship is no longer (log-)linear: there is a relatively slow increase first, a plateau at productivity levels around EUR 15–20, and a strong increase thereafter. While the lowest-productivity firm has a markup of about 15%, the markup is over 40% for the firms with the highest productivity. Put differently, workers obtain less than 60% of the value product in these high-productivity firms. However, as the estimate of the productivity density γ in panel (g) makes clear, such cases are fairly rare. Most firms have a value product of less than EUR 20 per hour, and there is hardly any mass left beyond EUR 40 per hour. Finally, panel (h) shows that our three-stage estimate of firm productivity results in a (non-parametric) distribution that is not too dissimilar from a Pareto distribution in that the density γ is a straight line in log-log-coordinates over the relevant range of p.

The main equilibrium functions for the different labor markets defined by region, skill group, and permanent worker ability can be found in Appendix A.9. Both the wage density and the wage offer distribution in West Germany lie to the right of the curves for East Germany. Firms in West Germany tend to be more productive, and they also offer higher wages for any given productivity level above EUR 15-20 per hour. Below this range, the wage offer functions are almost identical in the East and in the West. This difference in wage setting is mirrored in the distribution of markups, which are higher in East Germany for productivity levels of about EUR 15 and more. The relative markup in East Germany grows monotonically with productivity, while in West Germany there is a decline over a short range; these different shapes explain the plateau that is observed for the sample as a whole.

The estimates also reveal considerable heterogeneity across skill groups and permanent worker ability. While the differences across skill groups are not that clear-cut, the differences across PFE quartiles exhibit the expected pattern, with wage and wage offer distributions in the higher permanent ability segments lying to the right to the respective distributions of their lower ability counterparts. Note that the correlation between reservation wages and permanent worker ability indicates that a great deal of unobserved heterogeneity is already captured by heterogeneity in reservation wages. There are also clear differences in wage-setting policies. Monopsony power decreases with the acquired skill level. For the low-skilled segment, the absolute markup exhibits a relatively flat increase at lower productivity levels, which results in a U-shaped pattern of the relative monopsony power. Monopsony power also generally decreases with permanent ability, even though some non-monotonicities in the relationship between productivity and monopsony power for the third and fourth PFE quartile counteract this pattern.

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the monotonicity provides a check of whether the observed distribution of wages may be an equilibrium outcome from the model.

6.3. Within-Sample Model Fit

To explore the within-sample model fit, we compare several simulated model outcomes with their empirical analogues. In a first step, we investigate the model's ability to fit the data for the whole sample. Second, because the plausibility of our model assumptions may vary across labor markets, we also explore the model fit by labor market segment.

We start by computing the steady-state unemployment rate u implied by our estimates (see Equation (5)). For the entire sample, we find a predicted rate of 8.4%, which is close to the rate of 8.1% observed in our stock sample (Table A.10 in Appendix A.10). The variation in the predicted rate across regions, skill groups, and permanent worker ability is consistent with the patterns observed in the sample, i.e. steady-state unemployment is higher in the East than in the West and decreases with skills and permanent worker ability (with the exception of the predicted difference in u between the third and fourth PFE quartile). The model overpredicts unemployment for the medium-skilled and for the upper quartiles of the PFE distribution, while it underpredicts the rate in East Germany, for the low-skilled and the lower PFE quartiles. While the overall fit is less satisfactory across worker ability segments and for the low-skilled, the differences in predicted and empirical unemployment rates for the other segments are small.

In order to explore the mechanisms behind the model's ability to predict u, we next turn to the transitions between labor market states. Table A.9 in Appendix A.10 compares simulated and empirical one-year transition rates. The model does a reasonable job in predicting these transitions and, in line with what was found for the unemployment rate, replicates the variation across segments very well. Keeping in mind that these transition rates are complex functions of underlying determinants, we discuss possible explanations for imperfections in the fit. Predicted job finding rates may be affected by the assumption that employed and unemployed individuals draw from the same wage offer distribution. This could imply that wage offers to the unemployed are overestimated. How well the approximation of an identical distribution works depends on how important the returns to on-the-job experience are in a particular labor market segment. Low-skilled individuals have been found to exhibit lower returns to experience (Dustmann and Meghir, 2005).[30] The assumption of an identical wage offer distribution for the unemployed and the employed is therefore more plausible for this group than for the medium-skilled, which is reflected in a better fit for the unemployment-to-employment transitions among low-skilled workers and the bottom half of the PFE distribution. As job offer arrival rates are determined by λ_u as well as F, we also compare the wage offer distribution F that results from the model with the empirical distribution of wages accepted by individuals who are currently employed at very low wages close to $\underline{w}$.[31] Figure A.12 in Appendix A.10 shows that the closer fit for the

[30] The authors show that low-skilled workers in Germany exhibit significant wage growth in the first two years of work, but virtually no real wage growth afterwards. The wages of (medium) skilled workers, in contrast, exhibit positive returns to experience throughout their working life.

[31] This is based on the idea that these individuals would accept any other job offer, such that the distribution of wages that low-wage individuals receive on their next job provides a direct estimate of the wage offer distribution even in the presence of reservation wage heterogeneity (see also Shephard, 2017).

unemployment-to-employment transitions of the low-skilled and the lower PFE quartiles goes hand in hand with a better fit for F in this direct test.

From Table A.10, we see that despite the good fit of the unemployment-to-employment transitions, the model underpredicts the unemployment rate for the low-skilled. Column (3) and (6) of Table A.9 show that part of this discrepancy may be explained by an underprediction of low-skilled workers' employment-to-unemployment transitions. A similar pattern may be observed for workers in the lower PFE quartiles. This suggests that the assumption of exogenous job destruction rates may be less innocuous for these groups.

We now turn to the fit of the unemployment duration distribution, focusing on the duration dependence of the transition rate from unemployment to employment. Please notice that the structural equilibrium model already predicts negative duration dependence driven by heterogeneity in the workers' values of b, so we look for evidence of duration dependence in the data over and above what is captured by the structural model. Figures A.13 and A.14 in Appendix A.10 show Kaplan-Meier survival curves for the transitions over 48 months. Clearly, at low durations the duration dependence is slightly more positive than what the estimated model predicts whereas at high durations the opposite applies. It seems that the estimated model reflects a compromise between the fit at low durations and the fit at high durations and as a result it slightly underestimates the positive duration dependence at low durations and it underestimates (in absolute sense) the negative duration dependence at high durations.[32]

Finally, the model tends to overpredict employment-to-employment transitions. The fit is better for West Germany, the medium-skilled and the upper PFE quartiles (Columns (2) and (4) in Table A.9). For these groups, the differences between predicted and empirical rates are smaller than one percentage point.

6.4. Robustness Checks

Table A.7 in Appendix A.9 reports results from a number of robustness checks for the whole sample. First, instead of disregarding individuals with wages right-censored at the upper limit for social security contributions (SSC), we use a Tobit regression to impute wages above this limit. Second, we replace the imputation of working hours with the assumption that all full-time employees work 40 hours per week. Third, we experiment with different ways of assigning a single wage to employment spells that last over several years, during which time individuals typically experience wage increases. In the theoretical model, this cannot happen

[32] We formalize this in a test by re-estimating the full model with an ad hoc extension of the likelihood in which we multiply the hazard for the unemployed by a term $\psi \cdot t^{\psi-1}$. Note that this extension is not meaningful in the context of the structural equilibrium model. Rather, it is a convenient way of checking whether the unemployment duration data fit the model. The test confirms the presence of mild positive duration dependence at durations below a year, over and beyond what is captured by the structural model through the heterogeneity in b. The model does not fit well the fraction of unemployed in the data with almost infinitely long unemployment durations. This may be because the normal distribution for b is too restrictive. Alternatively, some individuals who are classified in the data as being unemployed throughout the observation interval may actually not be in the labor force. This may be because in register data, classification into unemployment is not based on the ILO definition of unemployment but instead relies on the receipt of unemployment benefits.

as each job is characterized by a single, time-invariant wage. In our main specification, we use the average wage in the same job over the past year. In further robustness checks, we use the last observed wage only. Additionally, we use the average wage over the whole employment spell instead of the average annual wage. These measures differ from our baseline specification to the extent that individuals experience wage changes within the last year or over the whole employment spell. Fourth, we truncate the wage distribution at a different level. In our main specification, wages below EUR 3.00 per hour are discarded in an initial step. As a robustness check, we change this threshold to EUR 4.00.[33] Moreover, when replacing the right-censoring at the upper limit for SSC with an imputation procedure, we try two variants in which we truncate the imputed wages at the 95th or 99th percentile. Finally, we set ρ, which is assumed to be 0.02 in our main specification, to alternative values (0.01 or 0.04). We also combine the robustness checks along the different dimensions. While the first three dimensions have a negligible impact on the parameter estimates, the assumptions regarding the truncation level and the discount rate ρ matter slightly more. However, the impact is limited to the estimates of the job offer arrival rates, while the other parameters remain almost unchanged. The following comparative statics results remain qualitatively very similar in all these specifications.[34]

6.5. Recruiting Cost Function

Table A.8 in the Appendix provides non-linear least squares estimates of η after taking logs of Equation (9) and including controls for firm size (four categorial dummies) as well as AKM establishment fixed effects pre-estimated for the time period 2007-2013 (see Section A.7 in the Appendix). The AKM effects are included as proxies for firm productivity to capture the term $log(c(p))$. To estimate Equation (9), we use JVS data for the year 2014. Information on recruiting costs is not available before 2014 and this is the last year before the introduction of the minimum wage. For the pooled sample, the estimates for η range between 1.71 and 1.76, depending on whether AKM establishment fixed effects have been included in the specifications. By excluding hiring processes with the highest job requirement level, we obtain estimates between 1.71 and 1.79. Breaking down the sample by region yields estimates of 1.54 for West Germany and 2.06 for East Germany when controlling for AKM establishment fixed effects. Shephard (2017) sets η equal to 2.0 for the UK. However, for Germany and the sample analyzed, an η of 1.75 is closer to other estimates based on German data.[35] For our baseline specification, we therefore take 1.75 as a value for η, but we provide sensitivity checks of the minimum wage simulations for other parameter values of η (see Section 7.2).

[33]The resulting lowest wage observed is a bit higher, e.g. EUR 3.64, because of a further truncation at the first and 99th percentile within each PFE quartile, see Section 3. This further depends on the wage and hours measure used.

[34]For the change in the unemployment rate, this is documented in the last column of Table A.7.

[35]Using a stylized labor demand model, Muehlemann and Pfeifer (2016) estimate an elasticity of recruiting costs with respect to the number of hires of 1.3 to 1.4. In an earlier discussion paper version, the authors also provide estimates of the elasticity of job posting costs with respect to the number of hires, which amounts to about 1.7 to 1.8 (Muehlemann and Pfeifer, 2013).

7. Unemployment Effects of Different Minimum Wage Levels

We begin by discussing and quantifying the different mechanisms through which the minimum wage affects the unemployment rate (Section 7.1). We then present results on the total effect, i.e. incorporating all these pathways, of the minimum wage on the rate and duration of unemployment, first overall (Section 7.2) and then by region, skill group, and permanent worker ability (Section 7.3). Finally, we discuss the out-of-sample performance of our model, in particular by comparing its predictions with the quasi-experimental evidence on the effects of the German minimum wage (Section 7.4).

7.1. Pathways

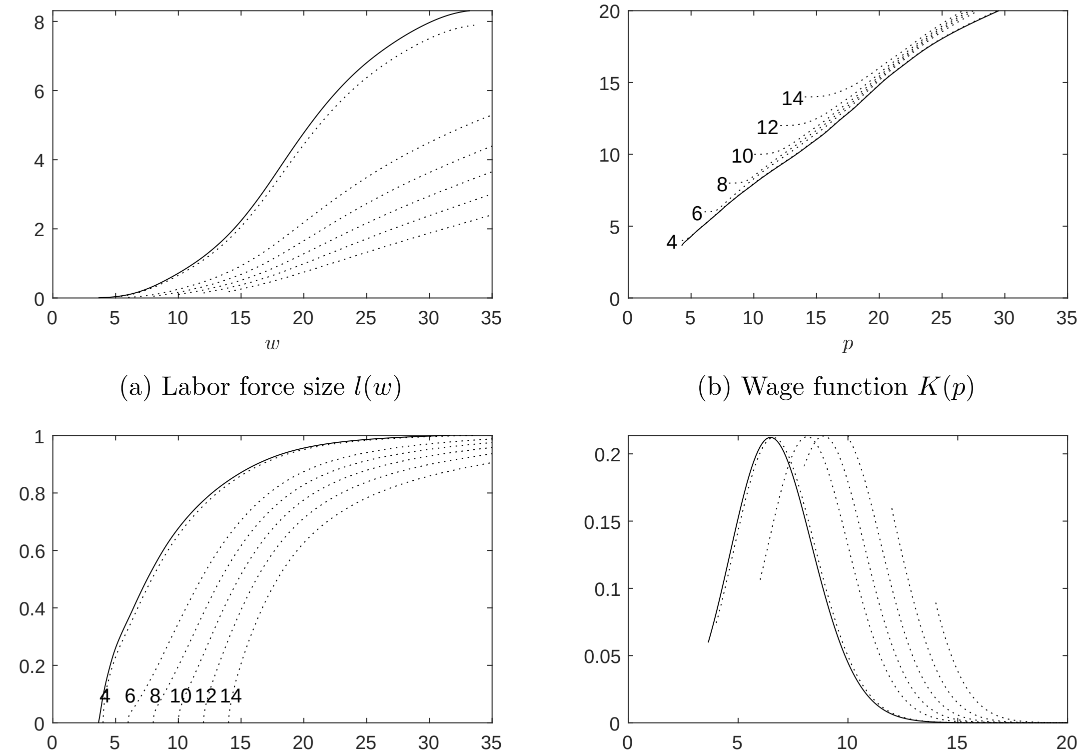
Decomposition of the Unemployment Rate The unemployed fall into three different groups, as shown by the decomposition in Equation (5). Group (1) consists of individuals whose reservation wage is below $\underline{w}$, i.e., who will accept any job offer. This purely frictional unemployment decreases in κ_u, the ratio of the job arrival rate of the unemployed over the job destruction rate. For Group (2), unemployment is partly frictional (through κ_u) and partly driven by the interplay between the reservation wage and the wage offer distribution. Unemployed individuals in this group accept some job offers but reject others, depending on the wage offer. Finally, individuals in Group (3) are permanently unemployed because their reservation wage is higher than the highest wage offer $\overline{w}$.

Effects through the Wage Offer Distribution For minimum wage levels below the lowest productivity level $\underline{p}$, the model predicts that a minimum wage reduces unemployment, as long as the minimum wage shifts up firms' optimal wage offers. The reason is that in this case unemployed individuals are now more likely to receive acceptable wage offers.[36] With $\underline{w}$ = EUR 3.64 and our estimate for the wage offer function, this cutoff level is $\hat{\underline{p}} = \hat{K}^{-1}(3.64)$ = EUR 4.26 for the whole sample. The introduction of a minimum wage of, say, EUR 4.00, slightly above the lowest productivity, limits firms' power to set wages below productivity. The lowest wage is now EUR 4.00 instead of EUR 3.64 and, via Equation (8), this increase has repercussions throughout the wage offer distribution.[37] This is illustrated in Panel (a) of Figure 2 for the whole sample: the higher the minimum wage level, the smaller the workforce l that a firm attracts for a given wage offer w. Moreover, the relationship between l and w becomes less steep for higher minimum wages.

As a result of these interactions operating through $l(w)$, different minimum wage levels lead to different optimal wage offer functions $\hat{K}^{MW}$, and therefore to different wage offer distributions $\hat{F}^{MW}$. Increasing the minimum wage generally shifts $\hat{K}^{MW}$ upwards and $\hat{F}^{MW}$ to the right (cf. panels (b) and (c)). While the biggest changes occur for low wages and

[36]Note that an alternative way in which minimum wages may affect workers' search behavior would be by inducing more search effort in response to the minimum wage. See Christensen et al. (2005) for a model of endogenous search effort and a previous (working paper) version of Ahn et al. (2011), which discusses an extension of their model that allows for endogenous search effort.

[37]Several empirical studies document spillover effects of minimum wages on wages in the upper part of the wage distribution (e.g., Autor et al., 2016 or, in the context of Germany, Gregory and Zierahn, 2022).



(a) Labor force size $l(w)$ (b) Wage function $K(p)$

(c) Wage offer distribution $F(w)$ (d) Reservation wage density $dA(w)$

Figure 2: Equilibrium Functions for Different Minimum Wage Levels (Whole sample). *Note:* This figure illustrates the effect of different minimum wages on equilibrium functions (cf. Figure 1). Panel (a) shows the effect on the labor force size (measured per unit intensity). Different minimum wage levels shifts the optimal wage offer function (panel b) and the wage offer distribution (panel c). Panel (d) shows the effect on the reservation wage density and its support. Wages and productivity are measured in EUR. *Key:* No minimum wage (——); minimum wage levels are EUR 4, 6, 8, 10, 12, and 14 (··········).

productivities, even high-productivity firms adjust their wage offer slightly in response to an increase in the minimum wage.

These changes in the wage offer distribution affect the steady-state unemployment rate. A minimum wage below $\underline{p}$ leads to an increase in $\underline{w}$, which in turn means that some individuals shift from Group (2) to Group (1) in Equation (5). As $1+\kappa_u > 1+\kappa_u\overline{F}(x)$ for all $x \in]\underline{w}, \overline{w}]$, this leads to a reduction in the unemployment rate. For individuals staying in Group (2), unemployment goes down as $\overline{F}(w)$ decreases for all w. Moreover, the highest wage offer $\overline{w}$ increases, which reduces the number of individuals who reject all job offers (Group 3).

Effects through the Job Offer Arrival Rates For minimum wage levels above the lowest productivity level $\underline{p}$, the minimum wage affects the job offer arrival rates, which means that the sign of the minimum wage effect on unemployment becomes ambiguous a priori. In the model by Bontemps et al. (1999, 2000), this mechanism arises because the minimum wage raises the lowest feasible wage offer $\underline{w}^{MW}$ and the productivity level $\underline{p}^{MW}$ that is associated with it. Firms with a productivity below this level leave the market, and

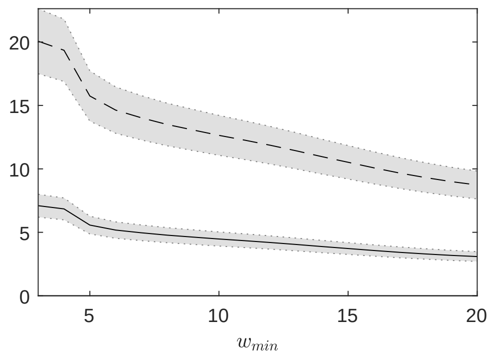


Figure 3: Change in Minimum Wage: Frictional parameters (Whole sample). *Note:* This figure shows κ, the ratio of the job arrival over the job destruction rate, as a function of the minimum wage. *Key:* Employed workers κ_e (——); unemployed κ_u (– – –). Grey area indicates 95% confidence bands based on 250 bootstrap runs.

Bontemps et al. (1999, 2000) assume that κ_u and κ_e are proportional to the fraction $\overline{\Gamma}(\underline{p}^{MW})$ of firms that remain in operation. Because the proportionality assumption is somewhat arbitrary, the present paper models firms' vacancy posting directly. As outlined in Section 2, firms choose the job offer arrival rates such that the return from marginally increasing the rates is just offset by the cost of doing so. An increase in the minimum wage leaves the cost unchanged but reduces the return, which leads to a reduction in the optimal job offer arrival rates. The reduction is particularly pronounced for job offers that are made to the unemployed (cf. Figure 3). Our simulations predict that a minimum wage of EUR 8.50 would bring down the ratio of the job offer arrival rate to the unemployed over the job destruction rate, which is 20.1 in the actual environment without a minimum wage, by almost a third. As a consequence, the unemployment effect of a minimum wage is now the result of two countervailing forces: the reduction in unemployment as higher wage offers lead to less frequent rejections of job offers, and the increase in unemployment arising from the fact that job offers now arrive at a slower rate. Formally, the second effect reduces the denominators in Equation (5), thereby increasing the frictional component of unemployment in Groups 1 and 2.

Effects through Reservation Wages So far, we have discussed the channels operating through the wage offer distribution and the job offer arrival rates. Both channels are already present in the Bontemps et al. (1999) model with homogeneous λ. In the model with $\lambda_u \neq \lambda_e$, there is an additional channel operating through A, the distribution of reservation wages ϕ. This channel is present regardless of whether the minimum wage is below or above $\underline{p}$. As shown in Equation (1), the reservation wage ϕ depends on κ_u, κ_e, F and $\overline{w}$, all of which are functions of the minimum wage. While an increase in $\overline{w}$ raises the reservation wage, a proportional reduction in κ_u and κ_e lowers it. F has a double effect on ϕ, operating both through the numerator and the denominator of the second term in Equation (1). In our simulations, the resulting net influence on A turns out to be positive, i.e. a higher minimum wage shifts the reservation wage distribution to the right (cf. panel (d) of Figure 2).[38] There is also a slight composition effect as a higher minimum wage draws in workers with relatively high reservation wages, while workers with lower reservation wages, who used to work at firms with low productivity that now have to leave the market, find themselves among the unemployed. In the baseline situation without a minimum wage, the distribution H_e of the opportunity costs of employment among the employed has a mean of EUR 0.002, while the distribution H_u for the unemployed has a slightly higher mean of EUR 0.50. (The means of the reservation wage distributions A_e and A_u are EUR 6.57 and EUR 6.86.) With a minimum wage of EUR 8.50, the mean of H_e increases to EUR 0.03, while the mean of H_u decreases to EUR 0.04. The means of A_e and A_u are very similar in this scenario, at EUR 9.17 and EUR 9.18, respectively.

7.2. Whole Sample

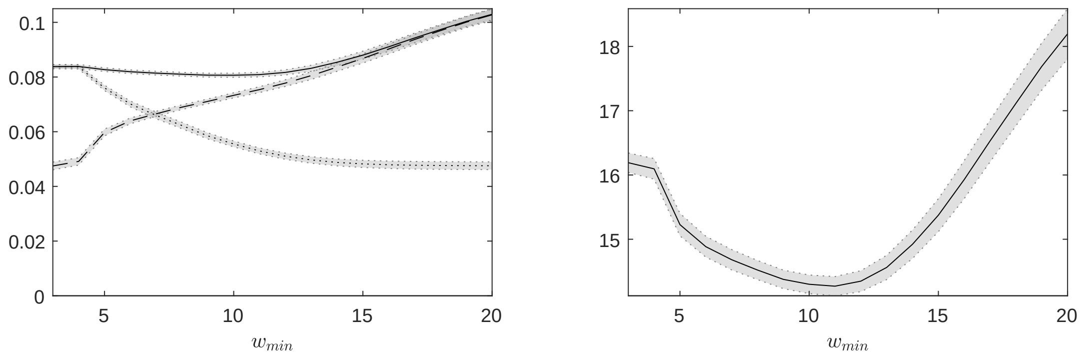
Figure 4 shows the effect of different minimum wage levels on the unemployment rate and the average unemployment duration, based on the estimation results for the whole sample.[39] The solid line is the effect that is actually predicted by the model. In the German context, the introduction of a statutory minimum wage leads to a reduction of the unemployment rate for low levels of the minimum wage. Unemployment is lowest for a minimum wage between EUR 9.00 and 11.00. The minimum wage of EUR 8.50 that was introduced on January 1st, 2015 leads to an unemployment rate of 8.1% in the model, down 0.3 percentage points from the baseline level (with no minimum wage) of 8.4%. At a minimum wage of EUR 13.50, the unemployment rate reaches this baseline level again.

The small effect of the minimum wage on the unemployment rate results from two countervailing forces that happen to almost exactly offset each other over a wide range of minimum wage values. The negative effect (in the sense of reducing unemployment) arises because a higher minimum wage means that unemployed individuals are now more likely to receive acceptable wage offers. This effect is illustrated by a simulation (see the dotted line in Figure 4) in which we allow for heterogeneity in the opportunity cost of employment b, but switch off the channel operating through the reduction in job offer arrival rates; these are held constant at their estimated status-quo levels. The unemployment-reducing effect tapers out beyond a minimum wage level of about EUR 12.50 because, as seen in Figure 1, there is little mass left in the reservation wage density beyond this level. Since the opportunity cost distribution H is unbounded, purely frictional unemployment (corresponding to a situation in which all unemployed individuals are in Group (1)) is reached only asymptotically. As w_{min} approaches infinity, the dotted line converges to an unemployment level of $100 \times 1/(1 + \widehat{\kappa_u}) = 4.8\%$.

This value is the starting point for the dashed line that shows the ratio $1/(1 + \widehat{\kappa_u})$. In this case, the negative effect working through the wage offer distribution is switched off, all

[38] Note that the distribution shown in the figure is truncated at the lowest wage observed in the data because the part to the left of $\underline{w}$ is non-parametrically unidentified.

[39] Selected numerical values for the change in the unemployment rate as a function of the minimum wage level are reported in Table A.13 in the Appendix.



(a) Unemployment rate u (b) Mean unemployment duration

Figure 4: Change in Minimum Wage: Unemployment (Whole sample). *Note:* The figure shows model predicted unemployment rates and unemployment durations for different minimum wage levels, based on the estimation results for the whole sample (see Section 6). Panel (a) shows the unemployment rate (see equation 5 in Section 2). The solid line (——) is the effect that is actually predicted by the model. To illustrate that the mininum wage effects result from two countervailing channels, we also show the unemployment rate for two alternative models: A model without changes in job offer arrival rates (············) and a model which only covers the frictional component of u (−−−). The model represented by the dotted line allows for heterogeneity in the opportunity cost of employment b, but switches off the channel operating through the reduction in job offer arrival rates (these are held constant at their estimated status-quo levels). The model represented by the dashed line shows the unemployment rate as the ratio $1/(1+\widehat{\kappa_u})$. In this case, the model does not allow for heterogeneity in b and all unemployment is purely frictional. Panel (b) shows the mean unemployment duration in months (see equation 15). Grey area indicates 95% confidence bands based on 250 bootstrap runs.

unemployment is purely frictional from the start, and higher minimum wages unambiguously raise unemployment. Such a scenario would lead to a very different conclusion about how the introduction of a minimum wage of EUR 8.50 in 2015 affected the unemployment rate. If all unemployment were always frictional, the rate would have gone up to 7.0%, a sizeable increase from the frictional unemployment rate of 4.8% that we find for the benchmark case without a minimum wage. The fact, however, that this benchmark rate of frictional unemployment is much lower than the predicted unemployment rate of 8.4% shows that, according to the analysis here, a substantial part of unemployment in the benchmark is not frictional but the result of wage offers that fall below the reservation wages of the unemployed. For higher values of the minimum wage, by contrast, the reduction in job offer arrival rates becomes the almost exclusive driver behind the increase in the unemployment rate. In Figure 4, this is reflected by the asymptotic convergence of the solid line and the dashed line.

The mean unemployment duration (panel (b) of Figure 4) is given by[40]

$$\frac{A_u(\underline{w})}{\lambda_u A(\overline{w})} + \int_{\underline{w}}^{\overline{w}} \frac{1}{\lambda_u \overline{F}(x) \cdot A(\overline{w})}\, dA_u(x). \tag{15}$$

[40]Note that the expected value needs to be derived based on the right truncated distribution of reservation wages, as individuals with reservation wages greater than $\overline{w}$ are characterized by infinite unemployment durations. In our application, $A(\overline{w})$ is equal to one. As a result, the mean unemployment durations based on the truncated and untruncated distribution of reservation wages do not differ from each other.

The effects mentioned above in the context of the unemployment rate are again at play here. In fact, each item in the expression depends on the minimum wage level. The effect on the numerator A_u is ambiguous a priori and, given that A changes little, probably fairly small. The main change is likely to take place in the denominator, where λ_u decreases in the minimum wage while $\overline{F}$ increases, again giving an ambiguous effect. The influence of the change in the integral limits $\underline{w}$ and $\overline{w}$ is also an empirical question. Our simulations show that with the introduction of a minimum wage, the mean unemployment duration first decreases from its steady-state level of 16.2 weeks. The minimum of about 14 weeks is reached at a minimum wage level of around EUR 11. For higher levels of the minimum wage, the average unemployment duration increases again.

The last two columns of Table A.7 in Appendix A.9 show the simulated level and the change in the unemployment rate at a minimum wage level of EUR 8.50 for the different sensitivity analyses reported in Section 6.4. We document small decreases in the unemployment rate in all specifications, with the smallest reduction of 0.04 percentage points in the specification in which the lowest wage level $\underline{w}$ is restricted to EUR 4.00, and with the largest reduction of 0.71 percentage points in the specification in which the wage information is based on the average wage of the whole employment spell. An exception is the specification with imputed wages that are truncated at the 99th percentile: Here, the minimum wage effect is positive and the unemployment rate increases by 0.59 percentage points.

In our main specification, the curvature η of the recruiting cost function is set to 1.75. As a robustness check, we use the values of 1.71 and 1.79 that result from our estimation (cf. Table A.8 in the Appendix). For smaller values of η, we simulate slightly smaller reductions in the unemployment rate at EUR 8.50 for the whole sample. For the estimated values of η, 1.71 and 1.79, we find a reduction of 0.24 percentage points and 0.38 percentage points, respectively, while we find a reduction of 0.31 percentage points in our main specification. The finding that unemployment is reduced at the EUR 8.50 minimum wage is robust even when we increase the range of η further to values of 1.65 and 1.85, i.e. ± 0.10 around the value used in the main specification. A graphical representation of these robustness checks for different minimum wages can be found in Figure A.15 in the Appendix.

7.3. Heterogeneity between Labor Markets

The simulation results discussed so far have been based on the estimation for the whole sample. Figures 5, 6 and 7 show the effects when the simulations are based on a separate estimation for each labor market defined by region, skill group, or permanent worker ability. The figures show the change of the unemployment rate compared with its status-quo level in each labor market segment, as a function of the minimum wage level.[41] We find that the

[41] Selected numerical values for these changes are again reported in Table A.13 in the Appendix. The simulations for the different labor market segments can be aggregated in order to derive the overall unemployment rate as a function of the minimum wage level (Table A.13 and Figure A.18 in the Appendix). Overall, the aggregated rates show similar trends compared to the rate that results when the estimation and simulation are directly carried out for the sample as a whole. However, when looking at the aggregation by PFE, we do not find a drop in the unemployment rate for low minimum wage levels and the increase in unemployment sets in earlier.

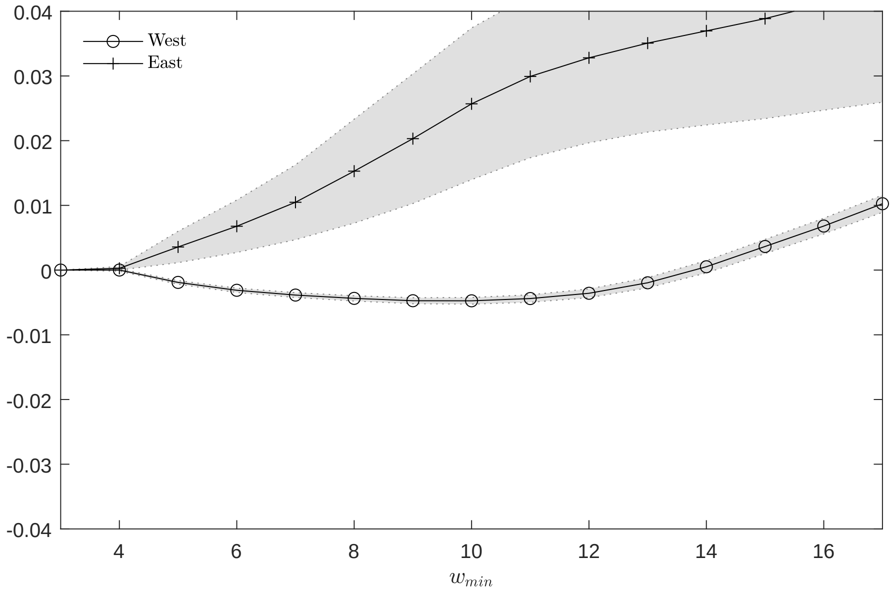


Figure 5: Change in Unemployment Rate by Region. This figure shows the change in the unemployment rate for different minimum wage levels based on separate estimations for East and West Germany relative to the status quo level. Grey area indicates 95% confidence bands based on 250 bootstrap runs.

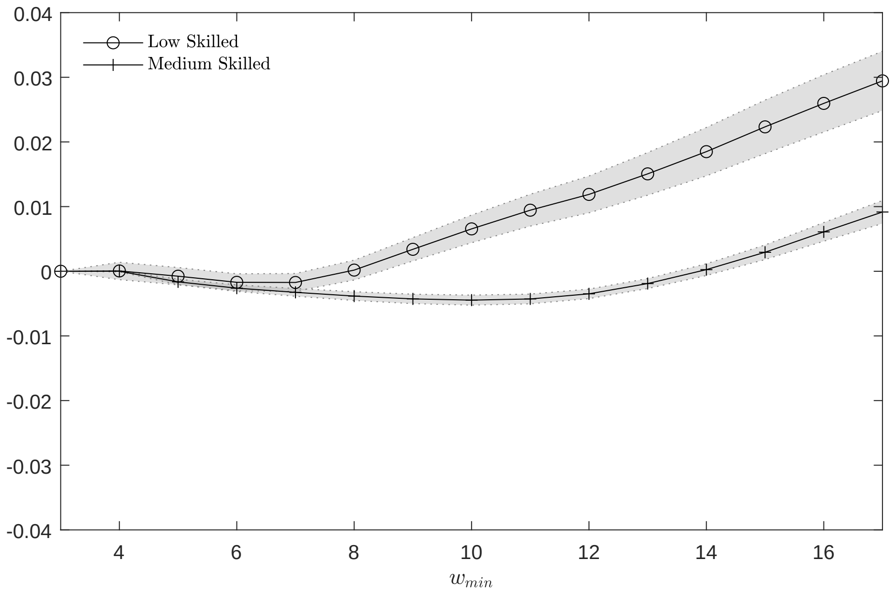


Figure 6: Change in Unemployment Rate by Skill Group. This figure shows the change in the unemployment rate for different minimum wage levels based on separate estimations for low- and medium-skilled individuals relative to the status quo level. Grey area indicates 95% confidence bands based on 250 bootstrap runs.

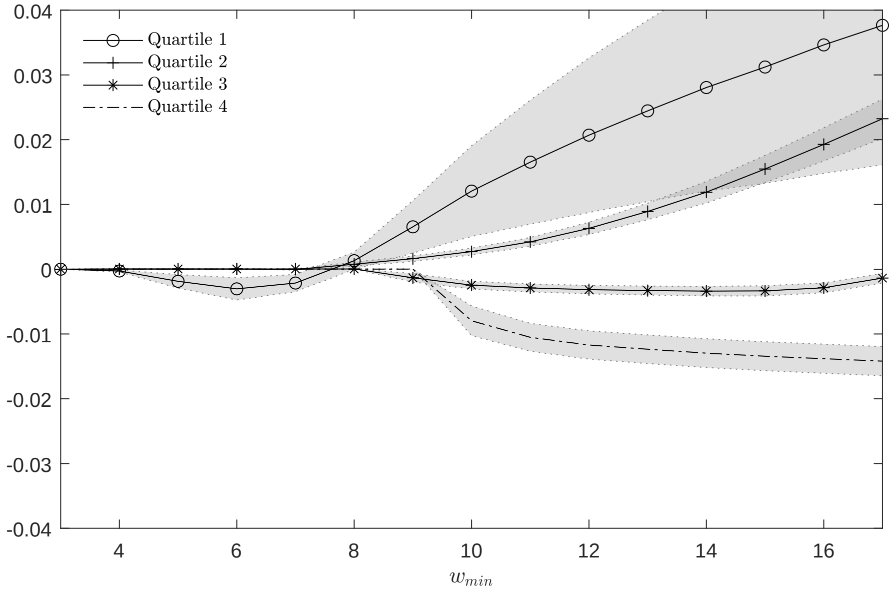


Figure 7: Change in Unemployment Rate by by PFE Quartile. This figure shows the change in the unemployment rate for different minimum wage levels based on separate estimations for each quartile of the person fixed effects relative to the status quo level. Grey area indicates 95% confidence bands based on 250 bootstrap runs.

same minimum wage level can have different effects on unemployment depending on the labor market segment.

By Region East and West Germany differ not only in the level of unemployment (9.0% in the East, 8.0% in the West), but also in how the unemployment rate reacts to the introduction of a minimum wage. According to our simulations, the introduction of any minimum wage – even at low levels – would increase the unemployment rate in East Germany. For the actually implemented level of EUR 8.50, the model predicts an increase of 1.8 percentage points, i.e. 20% of the status-quo level. In West Germany, by contrast, a minimum wage of EUR 8.50 reduces unemployment by 0.4 percentage points, or 5% of the rate observed before the introduction. Only for minimum wage levels of more than EUR 13.50 do we see an increase in unemployment in the West compared with the benchmark case without a minimum wage. These remain fairly moderate over the range of values considered here, while in East Germany a minimum wage of EUR 13.50 or higher would bring up the unemployment rate by about 30% compared with its pre-introduction level.

The effect of the minimum wage on unemployment again results from the different pathways described above. The unemployment-reducing effect – other things equal, unemployed individuals are less likely to reject wage offers after the introduction of a minimum wage – is stronger in the East at lower minimum wage levels because the unemployed have lower reservation wages there. At the same time, the productivity of firms is lower and the decline in the job offer arrival rate is more pronounced there. In the East, this second effect dominates throughout, while in West Germany it leads to an increase in unemployment only for relatively high minimum wage levels.

In our estimates based on the IAB Job Vacancy Survey, we find that the curvature of the recruiting cost function is smaller in the West ($\eta = 1.54$) than in the East ($\eta = 2.06$). When these region-specific estimates are used instead of the value for the whole sample ($\eta = 1.75$), we find that in the West the unemployment rate is slightly higher than in the main specification for all levels of the minimum wage, i.e. there is less scope for minimum-wage increases that lead to a reduction in unemployment (Appendix Figure A.16). In the East, by contrast, unemployment increases less in response to minimum-wage increases than in the main specification (Appendix Figure A.17).

By Skill Group Looking at labor market segments defined by skill group, the unemployment-reducing effect dominates up to a minimum wage level of about EUR 7 for the low-skilled. Afterwards the effect via a lower job offer arrival rate dominates. For the medium-skilled this effect dominates at much higher levels of about EUR 11.

By Permanent Worker Ability When labor market segments are defined via permanent worker ability, the reaction of the unemployment to the introduction of a minimum wage falls into three groups. The segment defined by the first PFE quartile first experiences a reduction in the unemployment rate, which is predicted to increase again beyond a minimum wage level of about EUR 6.00. Workers in the lowest quartile are characterized by low reservation wages, such that the negative effect of the minimum wage on unemployment is already exhausted at quite low minimum wages levels. The positive effect (in the sense of raising unemployment) that operates through a reduction in the job offer arrival rate takes over from then, and the unemployment rate reaches its status-quo level at a minimum wage of about EUR 8.00. In the segment defined by the second PFE quartile, the negative and positive effect on the minimum wage offset each other up to a minimum wage level of about EUR 8.00. Beyond this level the positive effect via the job offer arrival rate dominates. Note that, due to a higher value of $\underline{w}$ in the second PFE quartile, an effect on the unemployment rate may be observed only for minimum wages larger than EUR 6.32. For the third and fourth quartiles this threshold is EUR 8.14 and EUR 9.19, respectively. For these upper quartiles, the unemployment-reducing effect dominates almost completely. For the highest quartile, the effect through the job offer rate sets in so slowly that unemployment is lower than in the status quo even for a minimum wage of EUR 17.

7.4. Post-Reform Comparison

Comparing Predicted and Actual Unemployment We next compare our results with actual changes in unemployment between our pre-reform and post-reform period. Table A.11 shows observed changes in unemployment along with the changes predicted by our model for the whole sample and broken down by labor market segment. Over the pre- and post-reform period, the German labor market experienced a substantial decline in unemployment, despite a major inflow of unskilled migrants in the years 2015 and 2016. For our sample of full-time employed individuals, the unemployment rate fell by 3 percentage points. Thus, our model predicted change in unemployment of 0.03 percentage points accounts only for a small fraction of the observed decline in unemployment. The direction of the predicted change in unemployment across labor market segments is consistent with what is observed in the data for West Germany and the medium-skilled. For East Germany and the low-skilled the model predicts an increase in unemployment, which is counterfactual to what is observed in the data. As to permanent worker ability, workers in the upper quartiles experienced also a decrease in unemployment, which contrasts with the model predicted constant unemployment rates. For the lowest quartiles, the model predicted change in unemployment accounts only for a very small fraction of the observed substantial reduction of about 7 percentage points.

Taken together, from the lens of our model the introduction of the minimum wage can only explain a small fraction of the observed decline in unemployment. This suggests that there have been other important determinants in the German labor market that are relevant for understanding the favorable development of unemployment and that are not captured by our model. For example, a number of studies have documented an increase in matching efficiency after the German Hartz reforms, which involved a considerable change in the German welfare system and a reform of the Federal Employment Agency (Launov and Wälde, 2016; Hochmuth et al., 2021). Other factors include productivity growth coupled with general wage moderation and a considerable increase in labor hoarding due to rising labor shortages, which has led to a substantial decline in separation propensities (Klinger and Weber, 2016).[42] The latter is also visible in our data. Column (3) and (6) in Table A.12 report a substantial decline in separation rates across all segments, which is counterfactual to the predictions by our model.

Note that our model could, in principle, accommodate a change in separation (as well as job finding) rates by allowing for time variant job offer and destruction rates. Simulating a minimum wage of EUR 8.50 after estimating the model using data of the post-reform period would bring the predicted changes in unemployment probably much closer to those observed in the data. However, we wish to note that such a simulation would not reflect a counterfactual that isolates the pure minimum wage effect from other concurrent developments that have taken place during our observation period.

Comparison to Reduced-Form Studies The previous discussion has highlighted that our counterfactual predictions need to be interpreted as resulting from a policy change during which none of the other developments happened. Ex-post reduced-form evaluation studies typically aim at identifying the causal effect of the minimum wage holding everything else constant. Thus, the results from these studies provide a more appropriate reference against which to validate our model predictions.[43] In line with our simulations, much of

[42] See also the study Hutter et al. (2022), who adopt a macro-econometric approach to quantify the extent to which the above determinants contributed to the decline in unemployment during the time period 2012 to 2017.

[43] See the surveys by Bruttel (2019) and Caliendo et al. (2019). There is also a literature that evaluates the pre-2015 industry-specific minimum wages, typically using difference-in-differences designs with industries without a minimum wage as control groups. In what is probably the first quasi-experimental study for

the existing evidence finds that the minimum wage had no or at most a small negative effect on employment.[44] Using the IAB Establishment Panel, Bossler and Gerner (2020) compare establishments with employees affected by the minimum wage with a control group of establishments that are not directly affected. They find that employment remained roughly constant in the treatment group and grew in the control group. As a result, 45,000 to 68,000 jobs were lost (or rather, not created) because of the introduction of the minimum wage. This corresponds to 1.7% of the employment in the affected establishments. In line with our simulations, Bossler and Gerner find that this employment effect is mainly driven by establishments in East Germany. However, the employment effect is no longer statistically significant once the intensity of the treatment is taken into account. Using a similar difference-in-differences strategy at the establishment level, Bonin et al. (2018) find a small negative effect of about the same magnitude. The negative effect is exclusively driven by a reduction in marginal employment (below EUR 450 per month), while regular employment increased after the introduction of the minimum wage.

Several reduced-form studies rely on variation in the regional bite of the minimum wage. Bonin et al. (2020) find a negative effect on total employment of between 0.5% and 0.8%, depending on the specification. Based on the Structure of Earnings Survey, Caliendo et al. (2018) find that overall employment went down by 0.4%, which translates into 140,000 jobs. Using data from the Federal Employment Agency, Schmitz (2019) estimates a slightly larger reduction of up to 260,000 jobs. By contrast, with the same data but a different specification that also exploits variation across gender and age, Garloff (2019) finds some evidence of a positive effect on employment, but the effect is dependent on the specification and in any case very small (11,000 jobs). Adopting a similar approach, Stechert (2018) confirms these results. Based on individual-level data aggregated at the county level and taking commuting flows into account, Ahlfeldt et al. (2018) estimate a small positive effect of the minimum wage on employment (+0.06%). They also find that a one-percentage point increase in the regional bite decreased the unemployment rate by 0.05 percentage points. Holtemöller and Pohle (2020) use variation at the state-industry level. They find a small reduction in overall employment of, depending on their measure of the regional bite, between 20,000 and 50,000 jobs. Dustmann et al. (2022) do not find a significant negative employment effect in their regional analysis.

Friedrich (2020) exploits the variation in the bite of the minimum wage across occupations. He finds a small positive effect on regular employment when the minimum wage was introduced in 2015, but the effect becomes insignificant in 2016 and 2017. The effect on marginal employment is slightly negative, but statistically insignificant in all three years.

Germany, König and Möller (2009) analyze the introduction of a minimum wage in the construction industry. The authors find no significant employment effects in West Germany and small negative effects in the East. In 2011, the German Federal Ministry of Labor commissioned an evaluation of minimum wages in several industries. In general, these studies also tend to find limited employment effects (e.g., Boockmann et al., 2013; Frings, 2013), with the exception of the roofing industry (Aretz et al., 2013).

[44] The small employment effects are in line with firms' expectations and plans as reported in survey data in the months before the minimum wage took effect (Bossler, 2017; Link, 2019). In a more recent survey experiment by Bossler et al. (2020), by contrast, firms do report that they would reduce employment if the minimum wage were to be increased above its current level.

Restricting the analysis to West Germany produces the same pattern, while in East Germany the employment effects are more negative, in line with the predictions of our model.

Finally, using individual-level data, Dustmann et al. (2022) find that the introduction of the minimum wage boosted the wages of low-wage workers, but did not reduce their probability of remaining employed. Umkehrer and vom Berge (2020) study the effect of a minimum-wage exemption for the long-term unemployed and find no effect on transitions out of unemployment at the threshold.

Importantly, even in those studies that do find a negative overall employment effect, the decrease is almost exclusively driven by a reduction in marginal employment. Because we restrict our sample to full-time workers and thereby exclude marginal employment, the results are not directly comparable. The effect on regular employment in the reduced-form studies is at most slightly negative (Caliendo et al., 2018; Schmitz, 2019), while Bonin et al. (2020) find no significant effect and Garloff (2019), Holtemöller and Pohle (2020), and Friedrich (2020) estimate a small positive effect.

Regular employment in most of these studies also includes (non-marginal) part-time work, while our own simulation is carried out for full-time workers only. In addition, we exclude high-skilled individuals, i.e. the group of individuals who are least likely to be affected by the minimum wage. Notice also that our results refer to equilibrium changes whereas reduced-form studies capture short-term adjustments. Finally, we do not consider non-compliance, and our model assumes that prices and the productivity of firms and individuals are unaffected by the minimum wage.[45] Based on the extent to which firms can react to the minimum wage along these margins, the unemployment effects of the minimum wage will be dampened.[46]

While keeping in mind these caveats, we conclude that our simulation results for the introduction of the minimum wage are not in contradiction with the existing evidence, which increases our confidence that the counterfactual simulations are reasonably informative about what might happen to unemployment at other minimum wage levels. Note that this stands in contrast to previous ex-ante studies that relied on the assumption of perfect competition, i.e. on a model that by construction does not allow for positive employment effects of a minimum wage.[47]

[45] See Coviello et al. (2022) for a recent analysis of the effect of the minimum wage on the productivity of workers.

[46] See Clemens (2021) for a discussion of these other margins and a review of the recent empirical evidence. In the German context, Bossler et al. (2020) fail to detect any effects of the minimum wage on establishment-level productivity, based on data from the IAB Establishment Panel. For lack of price data at the micro level, there have been no quasi-experimental studies so far on whether prices were adjusted in response to the minimum wage. Using planned price changes as reported in several business surveys before and after the introduction of the minimum wage, Link (2019) shows that more firms planned to react via the price than via the employment channel.

[47] See, e.g., the studies by Ragnitz and Thum (2008); Bauer et al. (2009); Knabe and Schöb (2009), which predict large negative employment effects of up to one million jobs. Braun et al. (2020) calibrate stylized macro models (both for perfect competition and monopsony) and predict a strong increase in unemployment in their baseline specifications.

8. Conclusion

Based on an equilibrium job search model, this paper argues that the statutory uniform minimum wage of EUR 8.50 that was introduced in Germany in 2015 had a small negative effect on the unemployment rate of full-time workers. The unemployment-reducing effect is driven by West Germany, while in East Germany we do find an increase in unemployment resulting from the introduction of the minimum wage.

We use the model for a series of counterfactual policy experiments and find that unemployment is a non-monotonic function of the minimum wage level. In contrast, simple extrapolations of effects found for actually observed minimum wage levels might be misleading. Our model suggests that there would have been considerable scope for increasing the minimum wage beyond the level of EUR 8.50, at least for full-time workers, i.e. the group that we study in this article. We document substantial heterogeneity not only in the productivity distribution, but also in search frictions and in reservation wages across labor market segments differentiated by region, skill group, and permanent worker ability. To the extent that the minimum wage is motivated by a desire to offset firms' monopsony power, this suggests that a uniform minimum wage is perhaps too blunt a tool. While in East Germany there is basically no scope for unemployment-neutral minimum wages, in West Germany the benchmark level of unemployment is reached again at EUR 13.50. For medium-skilled workers, there would have been scope for unemployment-neutral minimum wages of up to EUR 14.00, while low-skilled workers would have experienced an increase in unemployment with minimum wages of more than EUR 8.00. The heterogeneity in unemployment responses is particularly pronounced across permanent ability segments. While for low-ability workers, the negative effect of the minimum wage on unemployment is already exhausted at quite low minimum wages levels, the unemployment-reducing effect completely dominates for high-ability workers. For these workers, the negative labor demand effect sets in so slowly that unemployment is lower than in the status quo even at quite high minimum wages. Again, these results and conclusions refer to full-time workers only.

Comparing our model predictions with the data suggests that according to our model the introduction of the minimum wage accounts only for a small fraction of the observed decline in unemployment. This indicates that there have been other important determinants of the favorable labor market development in Germany. However, the predictions from our counterfactual simulations are consistent with the results from studies using quasi-experimental variation, most of which have found no or at most small positive effects on unemployment. The counterfactual predictions do not necessarily translate into today's situation, however, because the various changes that have occurred since then (the continuation of the labor market boom and the Covid-19 pandemic in 2020) might affect the external validity of the results. Moreover, while the model allows us to assess the scope for unemployment-neutral minimum wage increases, we do not carry out an explicit welfare analysis and therefore refrain from drawing explicit policy conclusions.

In future research, it will be interesting to study correlates of search frictions and hence firms' market power across labor market segments. For instance, they may be related to differences in workers' characteristics across labor market segments, to firm characteristics, market structure, and union coverage. Finally, it would be interesting to complement the analysis of the present article, which focuses on full-time workers, by a study of the effect of the minimum wage on part-time workers. While part-time work is less important than full-time work in terms of stocks, the lower hourly wages of part-time workers make the latter arguably more sensitive to the minimum wage.

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A. Appendix

A.1. Data Preparation – Data Cleaning and Imputation

Imputation of Missing Information To maximize the available information, we fill in missing values using the full dataset, i.e. prior to imposing our sample selection criteria. When imputing missing information for the variable nationality, we first use information from parallel spells for the same individual, then information from previous spells and, if there are still missing values, with information from later spells. Similarly, we fill in missing information on region, sector, job title, position and employment status with information of previous and following spells but only if individuals stay at the same workplace.

Skill Groups (Educational Status) Missing and inconsistent data on education are corrected according to the imputation procedure IP1 described in Fitzenberger et al. (2006). This procedure relies, roughly speaking, on the assumption that individuals cannot lose their educational degrees. Information on educational status will be aggregated in three values:

- Low-skilled: High school diploma or no qualifications.
- Medium-skilled: Completed vocational training.
- High-skilled: Technical college degree or university degree.

The final sample used in the analysis consists only of low- and medium-skilled individuals.

A.2. Definition of Labor Market States

Employment Employment spells include continuous periods of employment (allowing gaps of up to four weeks) subject to social security contributions and (after 1998) marginal employment. For parallel spells of employment and unemployment (e.g. for those individuals who in addition to their earnings receive supplementary benefits), we treat employment as the dominant labor market state. We disregard employment spells where individuals receive Hartz IV benefits while working (*Aufstocker*), because for this group the wage alone is not a useful metric for work incentives. Furthermore, we disregard individuals in apprenticeships and interns. It is possible that individuals have multiple employment spells at the same time. In this case, only the predominant employment spell is kept. The predominant spell is determined as follows: full-time spells outrank part-time spells. When choosing between two full-time or two part-time spells, the spell with the longest duration is kept. To break any remaining ties, the spell with the highest wage wins.

Unemployment Unemployment spells include periods of registered job searching as well as periods of receiving benefits. Prior to 2005, the latter include benefits such as unemployment insurance and means-tested unemployment assistance benefits. Those (employable) individuals who were not entitled to unemployment insurance or assistance benefits could claim means-tested social assistance benefits. However, prior to 2005, spells of receiving social assistance can only be observed in the data if the job seekers' history records social assistance recipients as searching for a job. After 2004, means-tested unemployment and social assistance benefits were merged into one unified benefit, known as 'unemployment benefit II' (ALG II). Unemployment spells during which individuals receive ALG II are recorded in the data from 2007 onwards. For the period 2010–2013 that is used, the data provides a consistent definition of unemployment.

Distinction between Un- and Non-Employment Extending the procedure proposed by Lee and Wilke (2009), involuntary unemployment is defined as comprising all continuous periods of registered job searching and/or receipt of benefits. Gaps between such unemployment periods or gaps between receiving benefits or job searching and a new employment spell may not exceed four weeks, otherwise these periods are considered as non-employment spells (involving voluntary unemployment or leaving the social security labor force). Similarly, gaps between periods of employment and receiving benefits or job searching are treated as involuntary unemployment as long as the gap does not exceed six weeks, otherwise the gap is treated as non-employment.

A.3. Data Preparation – Weekly Hours of Work

While we observe whether an individual works full-time or part-time (defined as working less than 30 hours per week), the data lack explicit information on the number of hours worked. We only look at full-time employees and assign hours of work in the following way:

Main Specification: Imputation We complement the administrative data using the German Microcensus. To calculate hourly wages for full-time employment spells, we impute hours of work based on information from the German Microcensus. The imputation is done separately by region, sex, sector, job classification, and educational degree.

Alternative Specification: 40 Hours for Everyone In a variant, we assume 40 hours of work per week for all individuals in full-time employment.

A.4. Data preparation – Assignment of Wages

Wages are deflated by the Consumer Price Index of the Federal Statistical Office Germany, normalized to 1 in 2015. In our data, continuous employment spells may consist of a sequence of different spells with time-varying information of daily wages. To address this issue, we adopt two different variants to assign wages to one continuous employment spell. We also assign part- and full-time status consistent with these rules.

Main Specification: Average over one year We assign the duration-weighted average wage confined to the last observed year for employment spell before and without a transition. For subsequent employment spells, the wage information used is an average daily wage in the first year after the transition. An individual is considered mainly full-time employed, if the weighted average duration of full-time spells over one year exceeds 50%.

Alternative Specification: Last and first observations For employment spells before a transition and employment spells without a transition, the last observed wage is assigned. For subsequent employment spells, the first observed wage is assigned. The last part-/full-time status is assigned to the previous employment spell, whereas the first part-/full-time status is assigned to any subsequent employment spell.

A.5. Data Preparation – SSC threshold

Gross daily wages are right-censored at the upper limit for social security contributions.

Main Specification: Exclusion of Censored Observations We do not include observations with censored wages.

Alternative specification: Imputation To analyze this problem, we construct cells based on gender, year, region (East and West Germany), and educational degree. For each cell, a Tobit regression is estimated with log daily wages as the dependent variable and age, age squared, nationality, experience, experience squared, tenure in the current employment, tenure in the current employment squared, two skill dummies, occupational, sectoral as well as regional (Federal State) dummies and dummies for part-time and full-time employment as explanatory variables. As described in Gartner (2005), right-censored observations are replaced by wages randomly drawn from a truncated normal distribution whose moments are constructed by the predicted values from the Tobit regressions and whose (lower) truncation point is given by the contribution limit to the social security system.

A.6. Definition of Sub-Samples

Region

- East Germany: Former GDR, excluding Berlin
- West Germany, including Berlin

The labor market region of an employed individual is given by the location of the workplace. For the unemployed, we use the region where an individual searches for a job. Where this information is missing, we assign the region of the previous workplace.

Skill Groups (Educational Degree)

- Low-skilled: High school diploma or no qualifications
- Medium-skilled: Completed vocational training

Permanent Worker Ability

- Four categories based on quartiles of AKM person fixed effects (PFE) distribution (see Appendix A.7)

Table A.1: Employment-to-Employment Transitions across Region, Percent

	After transition	
Before transition	West	East
West	97.6	2.4
East	14.3	85.7

Note: Of 38,398 individuals experiencing an employment-to-employment transition, 95.9% remain in the same region.

Table A.2: Unemployment-to-Employment Transitions across Region, Percent

	After transition	
Before transition	West	East
West	96.9	3.1
East	11.2	88.8

Note: Of 8,079 individuals experiencing an unemployment-to-employment transition, 95.2% remain in the same region.

Table A.3: Employment-to-Employment Transitions across Skill Group, Percent

	After transition	
Before transition	Low	Medium
Low	79.8	20.2
Medium	0.0	100.0

Note: Of 38,398 individuals experiencing an employment-to-employment transition, 98.6% remain in the same skill group.

Table A.4: Unemployment-to-Employment Transitions across Skill Group, Percent

	After transition	
Before transition	Low	Medium
Low	72.6	27.4
Medium	0.0	100.0

Note: Of 8,079 individuals experiencing an unemployment-to-employment transition, 96.8% remain in the same skill group.

A.7. AKM Fixed Effects

To measure permanent worker ability, we use person fixed effects from a two-way fixed effects model of log wages (see Abowd et al., 1999). This model involves the estimation of the following equation

$$\log w_{it} = \alpha + \mathbf{X}_{it} \cdot \beta + \theta_i + \psi_{J(i,t)} + \epsilon_{it}, \tag{16}$$

where $\mathbf{X}_{it}$ is a vector of time varying observables with coefficient vector β, θ_i is a worker fixed effect that includes time-invariant observables and unobservables of individual i and $\psi_{J(i,t)}$ denotes an establishment fixed effect of establishment J, which employs worker i at time t. ϵ_{it} denotes a time-varying error term.

Estimation of Equation (16) follows the procedure described in Card et al. (2013) (henceforth referred to as CHK). In their two-step procedure, CHK first estimate firm fixed effects from the subgroup of workers who switch employers and control for year effects and age polynomials interacted with educational attainment. Based on these estimates, CHK compute the person fixed effects for stayers in a second step. While the original estimates used data up to 2009, the estimation procedure has been updated at the Research Data Centre of the IAB for the five periods 1985–1992, 1993–1999, 1998–2004, 2003–2010 and 2010–2017, using the universe of the Employment Histories (BeH) of full-time workers who are subject to social security contributions (Bellmann et al., 2020). Unlike in Card et al. (2013) and Card et al. (2015), the estimates for the latter four periods now also cover East Germany. The estimated person and firm fixed effects can be merged to several data sets, such as the Sample of Integrated Labor Market Biographies (SIAB) that is used in our analysis. For our analysis, we use the estimated effects for the years 2003–2010, as this time window defines the last period prior to the introduction of the minimum wage. Note that using an observation period somewhat farther away from the policy change also implies that anticipation effects are less of a concern. Merging the person fixed effect was possible for 94% of individuals in our sample. Unemployed individuals receive their person effect from their last employment spell prior to entering unemployment.

A crucial assumption underlying the AKM decomposition is that the error term in Equation (16) is uncorrelated to the sequence of worker i's employers. This assumption has been frequently criticized in the context of job search theory, as it does not allow for mobility based on the match-specific component of pay. Even though the fixed effects do not have a structural interpretation in our model, we wish to note that CHK demonstrate for the German case that the match-specific log wage components are quantitatively unimportant and unrelated to the direction of worker flows across employers. A further assumption of the AKM model is that firm fixed effects are additive and homogeneous across all workers employed at the same employer. To confirm this additive separability assumption, CHK further show that workers' (log) wage premiums are unrelated to their skill levels.

A.8. Descriptives

Table A.5: Number of Observations

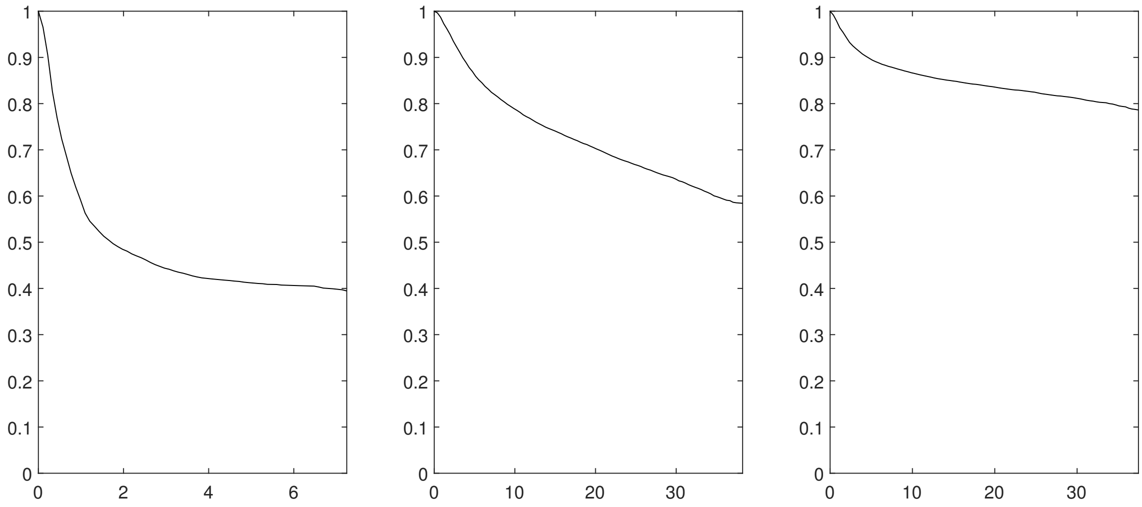
		Unemployment Spells				Employment Spells				
Sample	Total	Total	u → e	rc	lc	Total	e → e	e → u	rc	lc
Whole sample	208,626	16,916	8,079	8,837	27	191,710	38,398	20,818	132,494	2,219
Region										
West	176,205	13,480	6,398	7,082	23	162,725	32,720	17,003	113,002	578
East	32,421	3,436	1,681	1,755	4	28,985	5,678	3,815	19,492	1,641
Skill Group										
Low-Skilled	15,542	2,394	954	1,440	14	13,148	2,701	2,155	8,292	89
Medium-Skilled	193,084	14,522	7,125	7,397	13	178,562	35,697	18,663	124,202	2,130
PFE Quartile										
Quartile 1	37,049	7,003	2,715	4,288	21	30,046	6,391	6,712	16,943	140
Quartile 2	67,970	6,322	3,601	2,721	4	61,648	13,645	7,912	40,091	472
Quartile 3	69,438	2,581	1,431	1,150	1	66,857	12,709	4,399	49,749	902
Quartile 4	34,169	1,010	332	678	1	33,159	5,653	1,795	25,711	705

Note: Arrows (→) indicate that spells end in transitions to another employment spell (e) or to unemployment (u). Spells without an observed transition are right-censored (rc). Additionally, spells might-be left censored (lc). The column entries *Total Unemployment Spells* and *Total Employment Spells* sum up to column *Total* as 100%. The column entries $u \rightarrow e$, *rc* and *lc* sum up to column *Total Unemployment Spells.* The column entries $e \rightarrow e$, $e \rightarrow u$, *rc* and *lc* sum-up to column *Total Employment Spells.*

Table A.6: Percentage of Spell Types

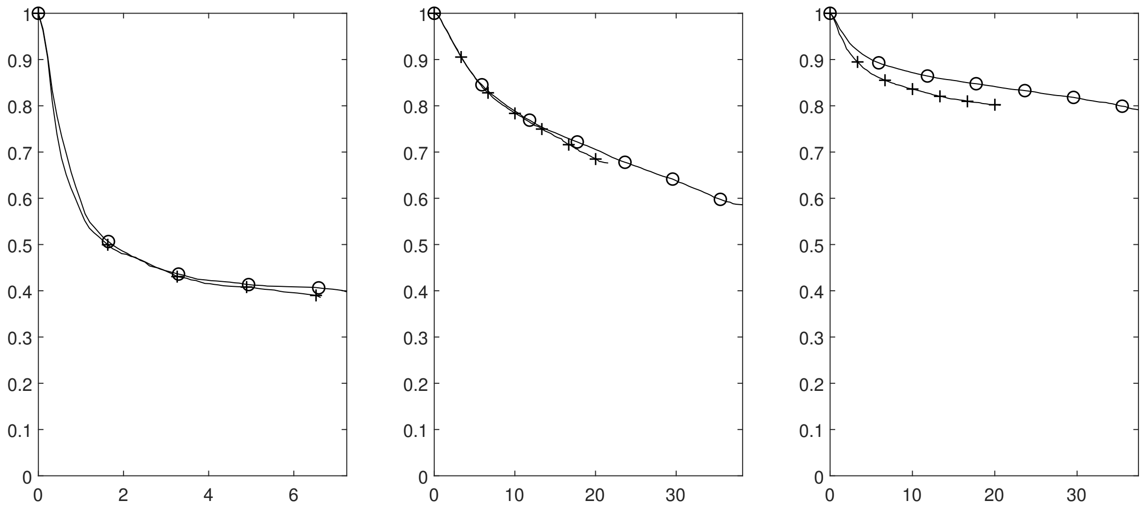
		Unemployment Spells				Employment Spells				
Sample	Total	Total	u → e	rc	lc	Total	e → e	e → u	rc	lc
Whole sample	100.0%	8.1%	47.8%	52.2%	0.2%	91.9%	20.0%	10.9%	69.1%	1.2%
Region										
West	100.0%	7.7%	47.5%	52.5%	0.2%	92.3%	20.1%	10.4%	69.4%	0.4%
East	100.0%	10.6%	48.9%	51.1%	0.1%	89.4%	19.6%	13.2%	67.2%	5.7%
Skill Group										
Low-Skilled	100.0%	15.4%	39.8%	60.2%	0.6%	84.6%	20.5%	16.4%	63.1%	0.7%
Medium-Skilled	100.0%	7.5%	49.1%	50.9%	0.1%	92.5%	20.0%	10.5%	69.6%	1.2%
PFE Quartile										
Quartile 1	100.0%	18.9%	38.8%	61.2%	0.3%	81.1%	21.3%	22.3%	56.4%	0.5%
Quartile 2	100.0%	9.3%	57.0%	43.0%	0.1%	90.7%	22.1%	12.8%	65.0%	0.8%
Quartile 3	100.0%	3.7%	55.4%	44.6%	0.0%	96.3%	19.0%	6.6%	74.4%	1.3%
Quartile 4	100.0%	3.0%	32.9%	67.1%	0.1%	97.0%	17.0%	5.4%	77.5%	2.1%

Note: Arrows (→) indicate that spells end in transitions to another employment spell (e) or to unemployment (u). Spells without an observed transition are right-censored (rc). Additionally, spells might be left-censored (lc). The column entries *Total Unemployment Spells* and *Total Employment Spells* sum up to column *Total* as 100%. The column entries $u \rightarrow e$, *rc* and *lc* sum up to column *Total Unemployment Spells* as 100%. The column entries $e \rightarrow e$, $e \rightarrow u$, *rc* and *lc* sum-up to column *Total Employment Spells* as 100%.



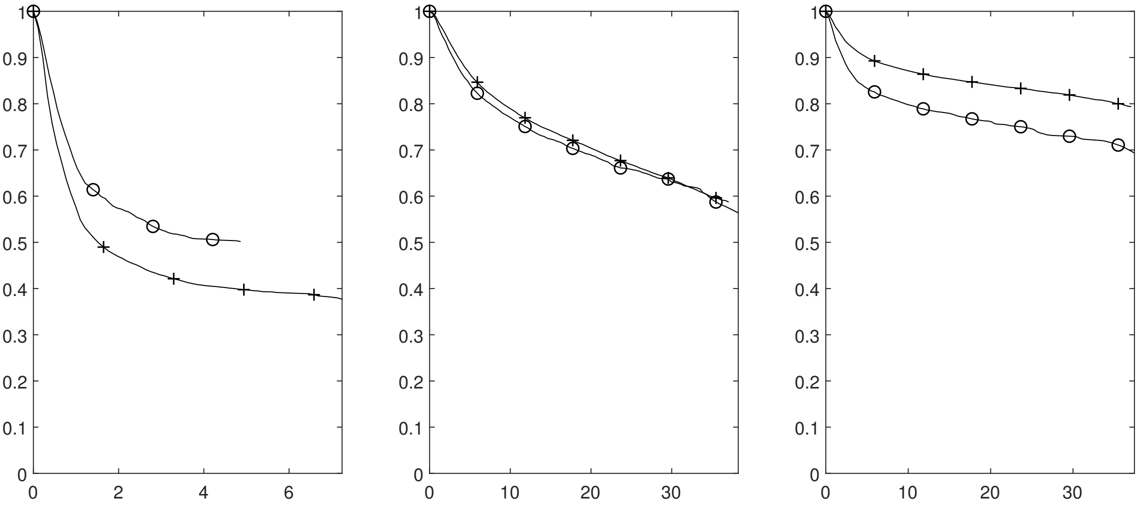
(a) u → e (b) e → e (c) e → u

Figure A.1: Survival Probabilities – Whole sample. *Note:* Plots show Kaplan-Meier survival estimate for durations in years. Arrows (→) indicate that spells end in another employment spell (e) or unemployment (u).



(a) u → e (b) e → e (c) e → u

Figure A.2: Survival Probabilities – by Region. *Note:* Plots show Kaplan-Meier survival estimate for durations in years. Arrows (→) indicate that spells end in another employment spell (e) or unemployment (u). *Key:* West (—o—); East (—+—).

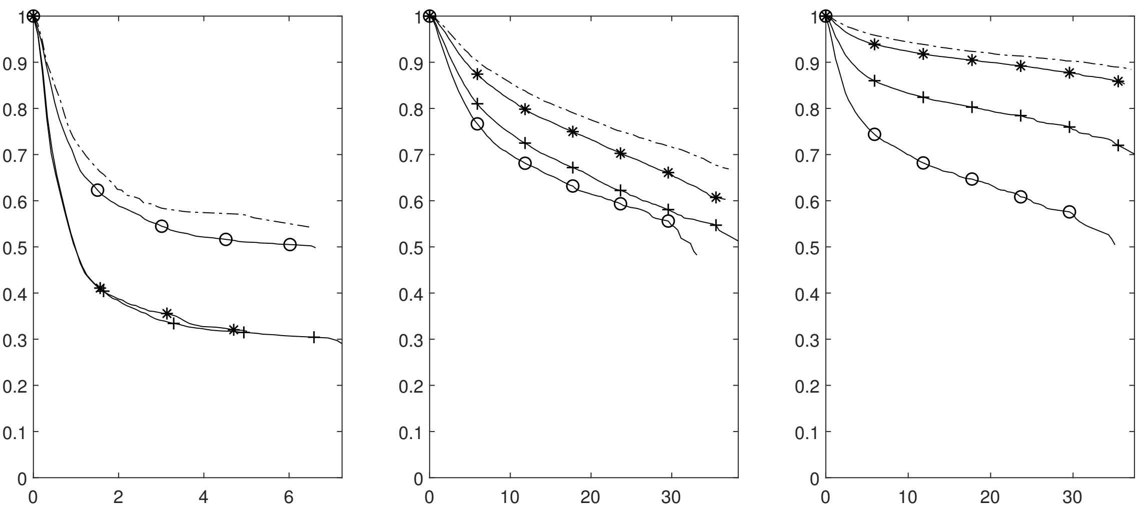


(a) u → e

(b) e → e

(c) e → u

Figure A.3: Survival Probabilities – by Skill Group. *Note:* Plots show Kaplan-Meier survival estimate for durations in years. Arrows (→) indicate that spells end in another employment spell (e) or unemployment (u). *Key:* Low-Skilled (—o—); Medium-Skilled (—+—).

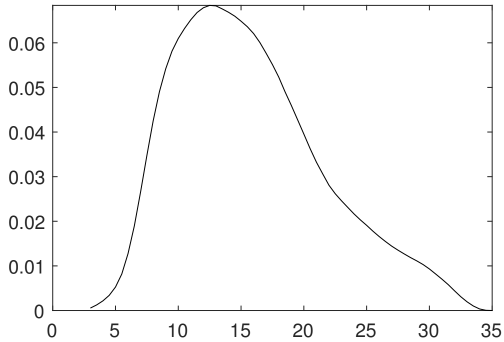


(a) u → e

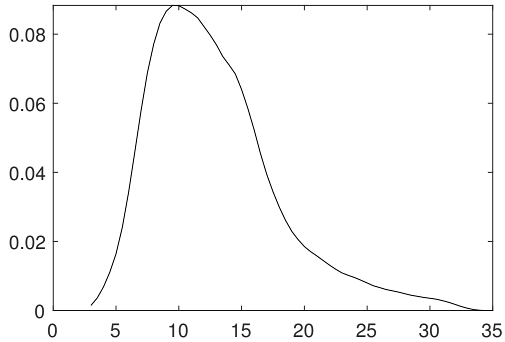
(b) e → e

(c) e → u

Figure A.4: Survival Probabilities – by PFE quartile. *Note:* Plots show Kaplan-Meier survival estimate for durations in years. Arrows (→) indicate that spells end in another employment spell (e) or unemployment (u). *Key:* Quartile 1 (—o—); Quartile 2 (—+—); Quartile 3 (—∗—); Quartile 4 (—·—).

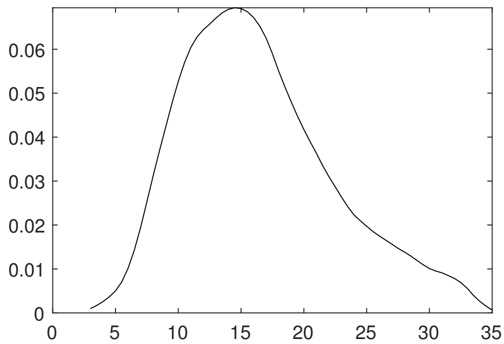


(a) e → e

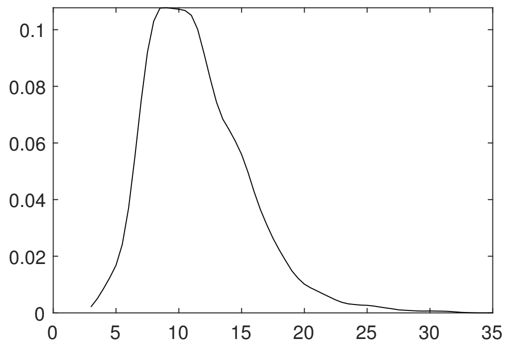


(b) e → u

Panel A – Before Transition

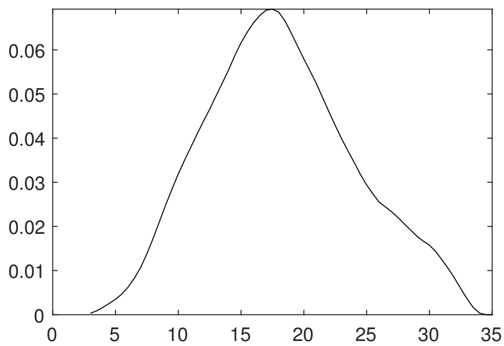


(c) e → e

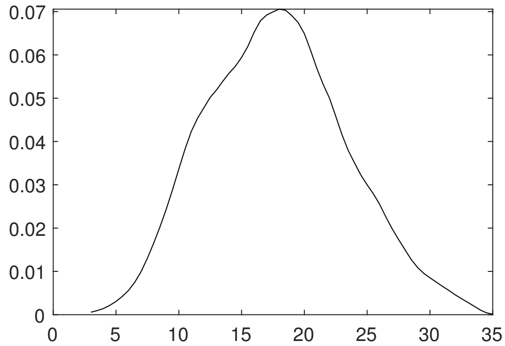


(d) u → e

Panel B – After Transition



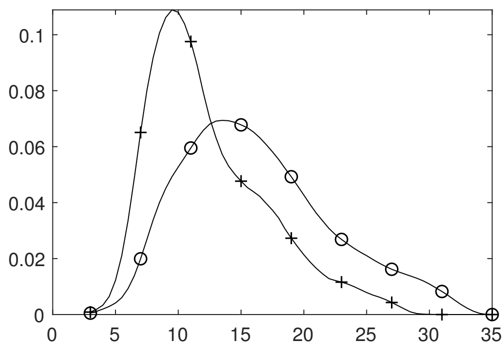
(e) right-censored



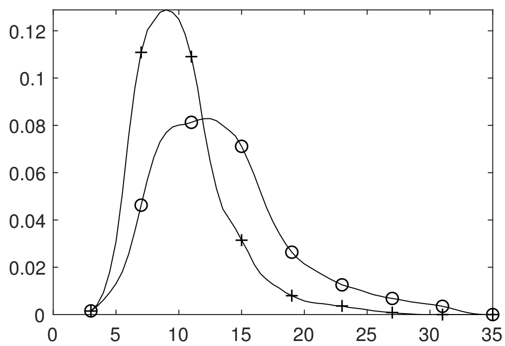
(f) left-censored

Panel C – Censored Spells

Figure A.5: Density of Hourly Wages – Whole sample. *Note:* Epanechnikov kernel density estimate. Arrows (→) indicate that spells end in another employment spell (e) or unemployment (u). Spells without an observed transition are right-censored. Additionally, spells might be left-censored.

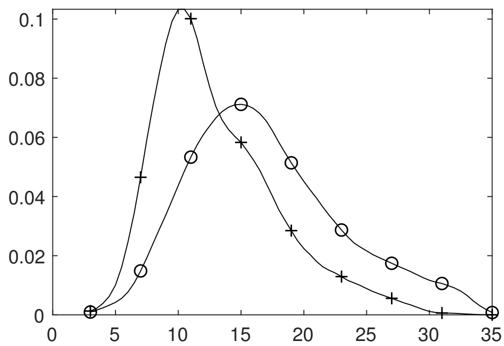


(a) e → e

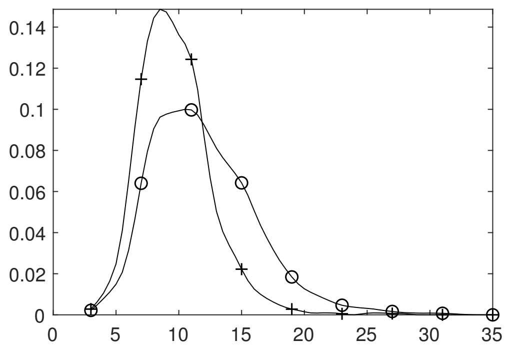


(b) e → u

Panel A – Before Transition

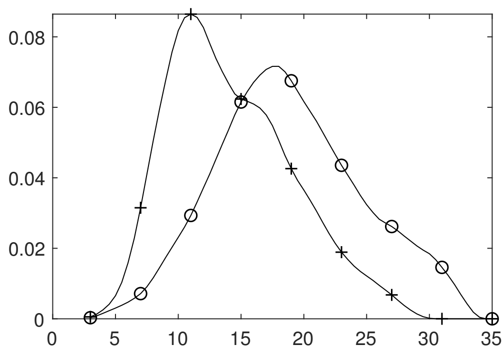


(c) e → e

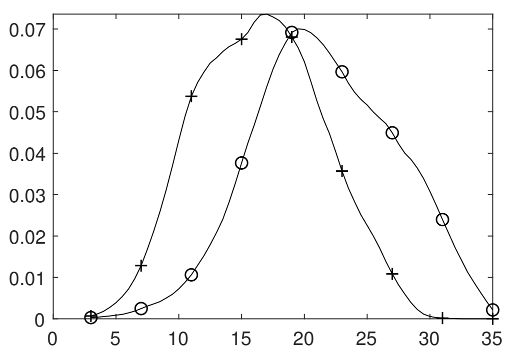


(d) u → e

Panel B – After Transition



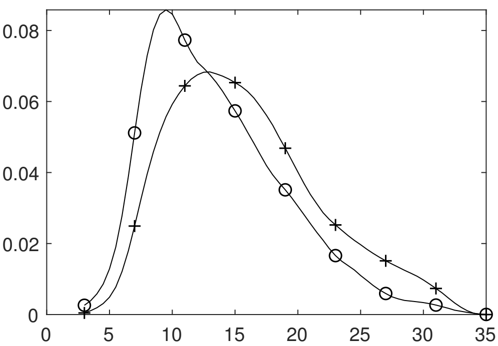
(e) right-censored



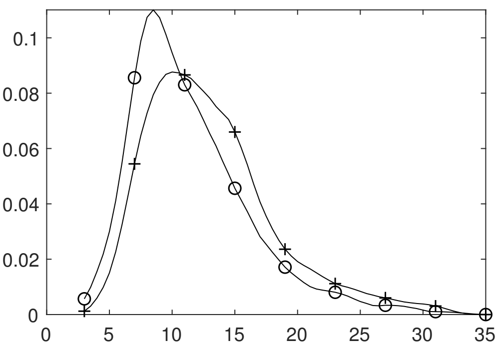
(f) left-censored

Panel C – Censored Spells

Figure A.6: Density of Hourly Wages – by Region. *Note:* Epanechnikov kernel density estimate. Arrows (→) indicate that spells end in another employment spell (e) or unemployment (u). Spells without an observed transition are right-censored. Additionally, spells might be left-censored. *Key:* West (—o—);East (—+—).

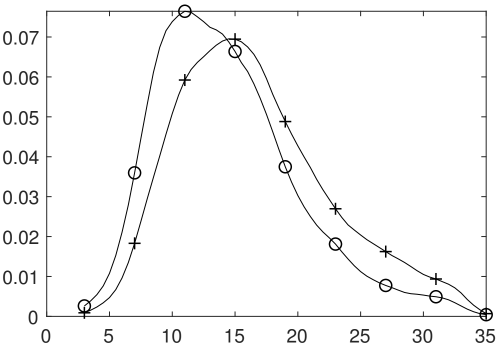


(a) e → e

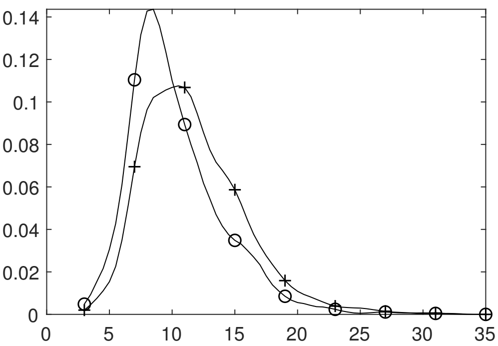


(b) e → u

Panel A – Before Transition

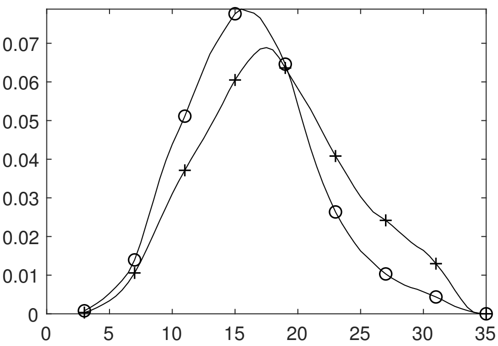


(c) e → e

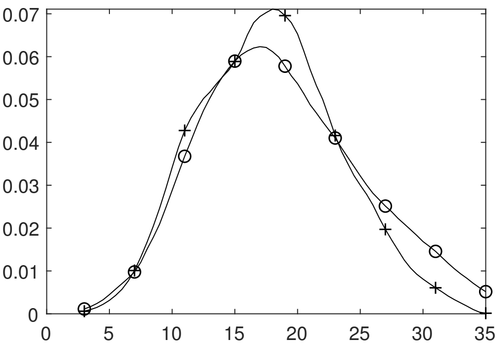


(d) u → e

Panel B – After Transition



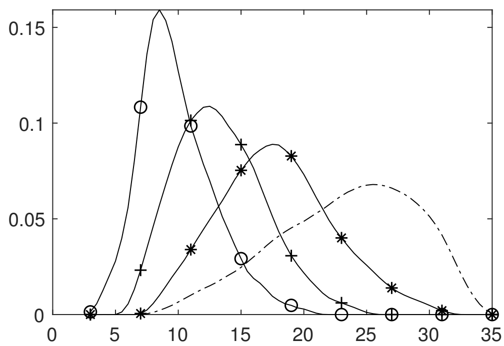
(e) right-censored



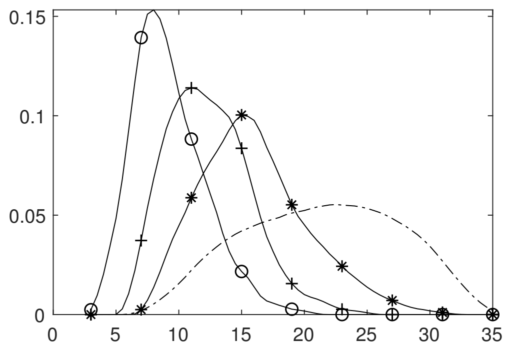
(f) left-censored

Panel C – Censored Spells

Figure A.7: Density of Hourly Wages – by Skill Group. *Note:* Epanechnikov kernel density estimate. Arrows (→) indicate that spells end in another employment spell (e) or unemployment (u). Spells without an observed transition are right-censored. Additionally, spells might be left-censored. *Key:* Low-Skilled (—o—); Medium-Skilled (—+—).

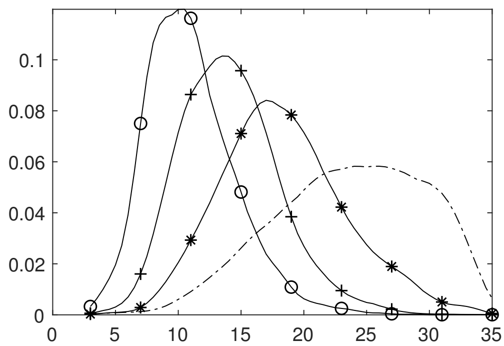


(a) e → e

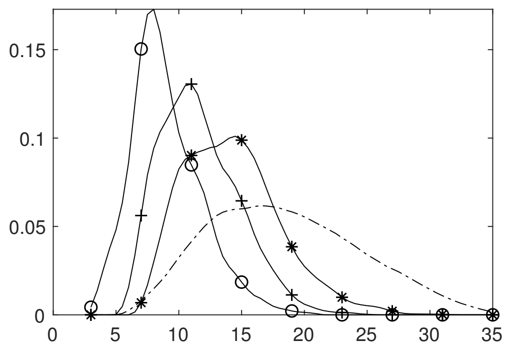


(b) e → u

Panel A – Before Transition

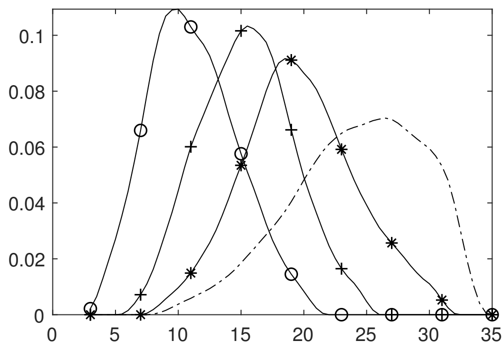


(c) e → e

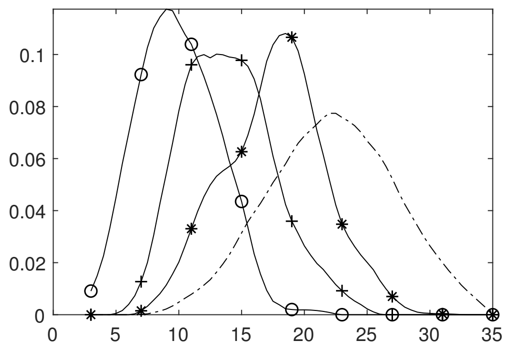


(d) u → e

Panel B – After Transition



(e) right-censored



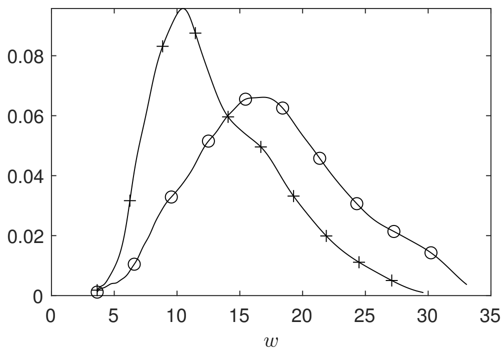
(f) left-censored

Panel C – Censored Spells

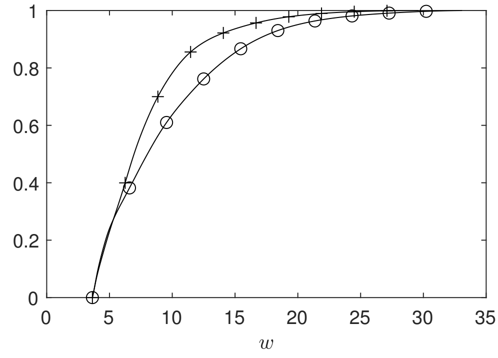
Figure A.8: Density of Hourly Wages – by PFE Quartile. *Note:* Epanechnikov kernel density estimate. Arrows (→) indicate that spells end in another employment spell (e) or unemployment (u). Spells without an observed transition are right-censored. Additionally, spells might be left-censored. *Key:* Quartile 1 (—o—);Quartile 2 (—+—);Quartile 3 (—*—);Quartile 4 (—·—).

A.9. Estimation Results

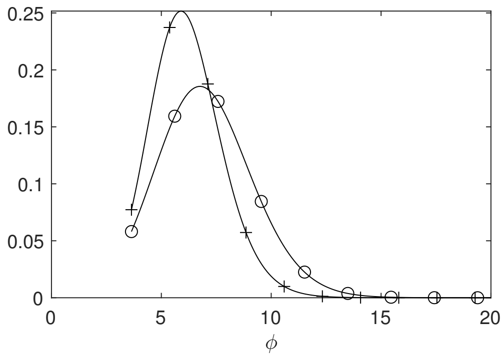
Bootstrapping We report bootstrapped standard errors with 250 draws. In very rare cases we exclude bootstrap runs if the likelihood does not converge: occurs in 1 of 251 bootstrap runs in Whole sample, in 10 of 260 bootstrap runs in East, in 1 of 251 bootstrap runs for Medium-Skilled, in 1 of 251 bootstrap runs in PFE Quartile 1, in 1 of 251 bootstrap runs in PFE Quartile 2, in 1 of 251 bootstrap runs in PFE Quartile 3, in 1 of 251 bootstrap runs in PFE Quartile 4, in 2 of 252 bootstrap runs in the robustness check with 40 hours per week, in 3 of 253 bootstrap runs in the robustness check with truncation of wages at the 99th percentile, in 2 of 252 bootstrap runs in the robustness check that use wages from the last spell and the first spell only, in 2 of 252 bootstrap runs in the robustness check that use wages from average last spell and average first spell, 40 hours per week, in 2 of 252 bootstrap runs in the robustness check with truncation at 4.00 euros, in 16 of 266 bootstrap runs in the robustness check with $\rho = 0.01$.



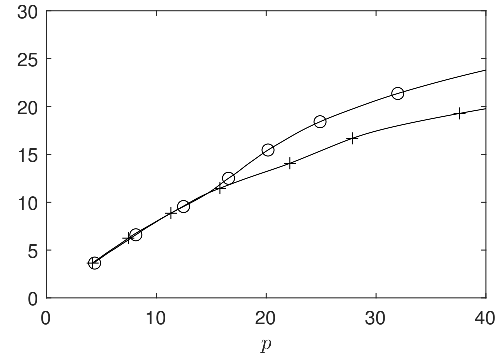
(a) Wage density $g(w)$



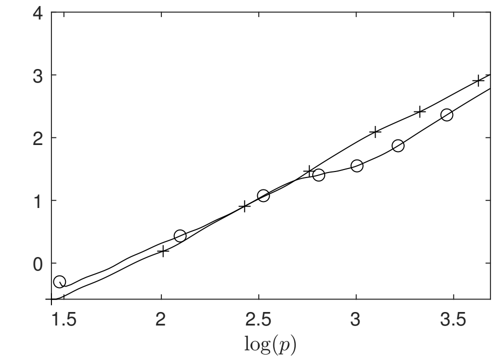
(b) Wage offer distribution $F(w)$



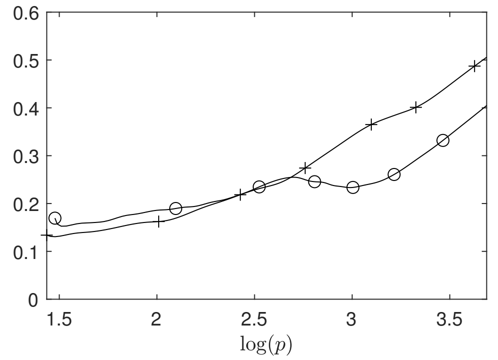
(c) Reservation wage density $dA(w)$



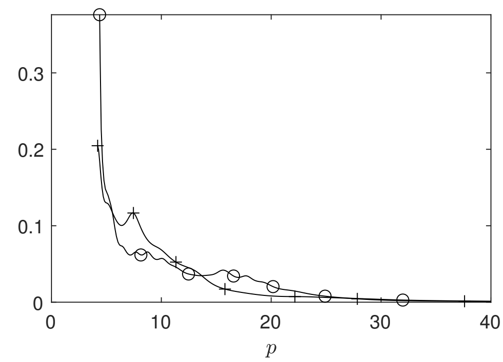
(d) Wage function $K(p)$



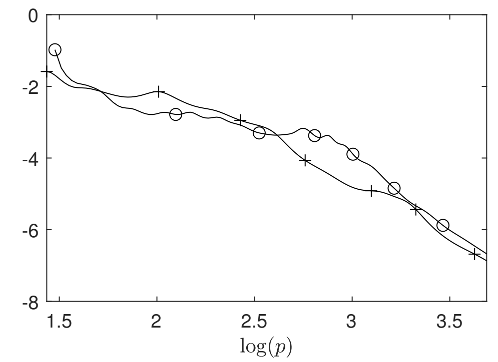
(e) Log absolute markup



(f) Monopsony power $[p - K(p)]/p$

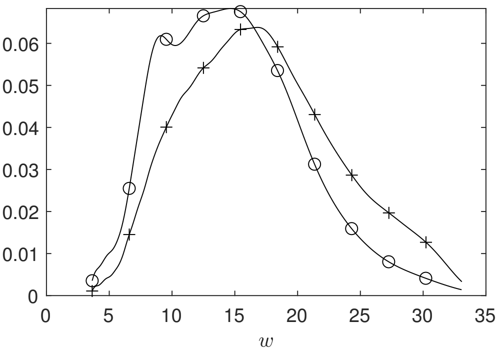


(g) Productivity density $\gamma(p)$

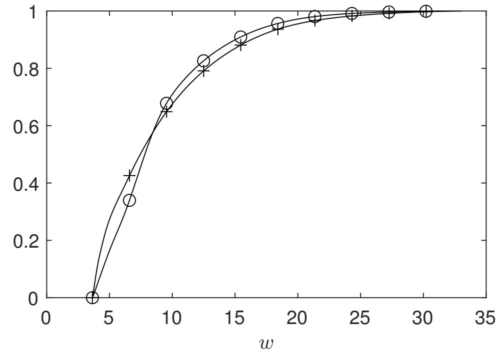


(h) Productivity log density $\log(\gamma(p))$

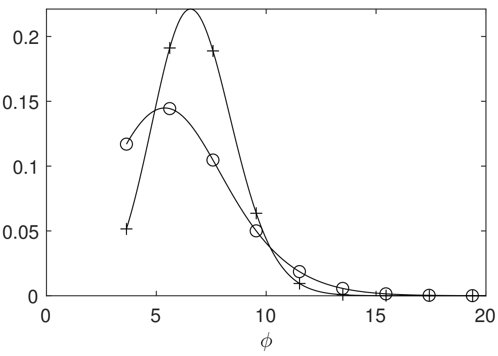
Figure A.9: Main Equilibrium Functions – by Region. *Note:* This figure shows the main equilibrium functions summarizing the steady-state for the estimates by region. Ordinate outcomes can be found in the subfigures' titles. Wages and productivity are measured in EUR. *Key:* West (—○—); East (—+—).



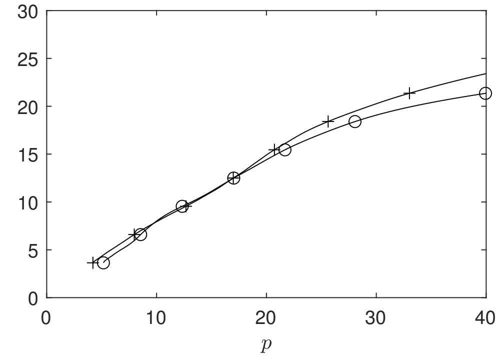
(a) Wage density $g(w)$



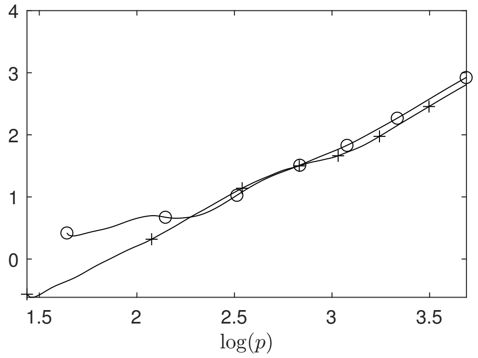
(b) Wage offer distribution $F(w)$



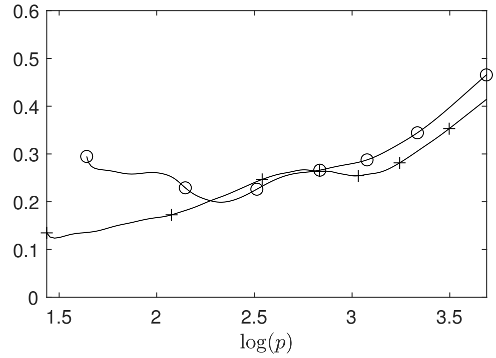
(c) Reservation wage density $dA(w)$



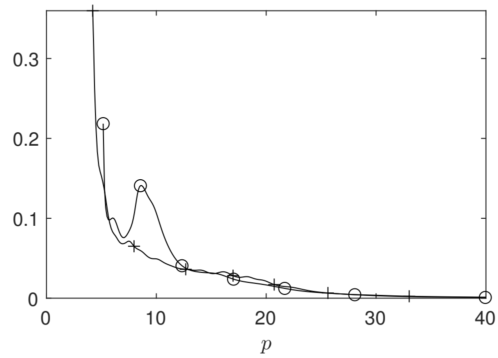
(d) Wage function $K(p)$



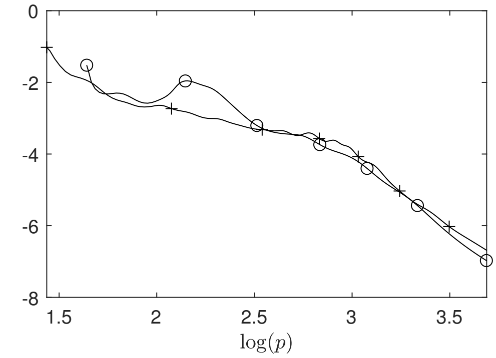
(e) Log absolute markup



(f) Monopsony power $[p - K(p)]/p$

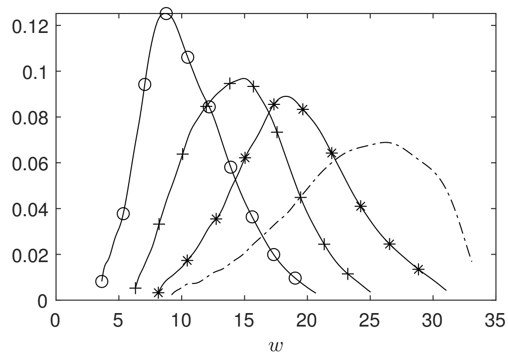


(g) Productivity density $\gamma(p)$

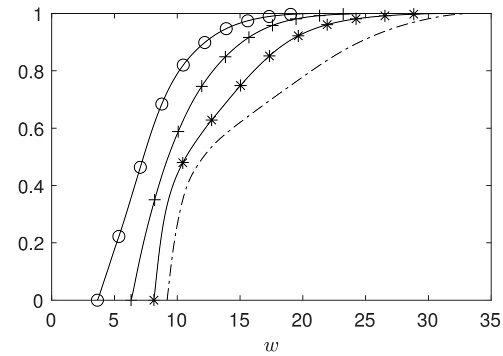


(h) Productivity log density $\log(\gamma(p))$

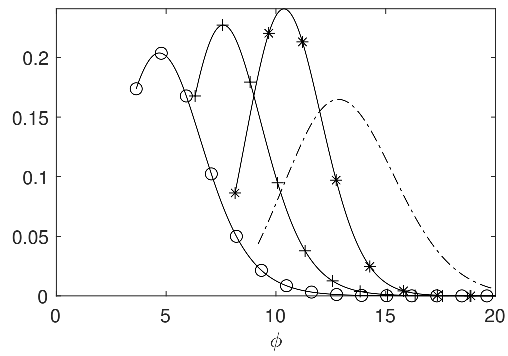
Figure A.10: Main Equilibrium Functions – by Skill Group. *Note:* This figure shows the main equilibrium functions summarizing the steady-state for the estimates by skill group. Ordinate outcomes can be found in the subfigures' titles. Wages and productivity are measured in EUR. *Key:* Low-Skilled (—○—); Medium-Skilled (—+—).



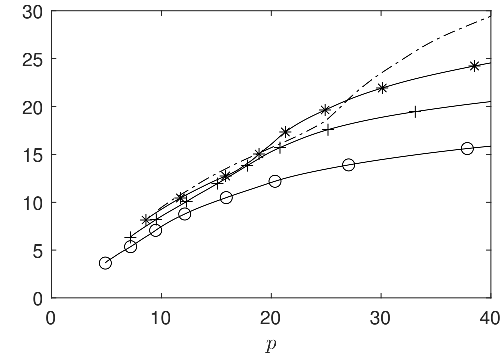
(a) Wage density $g(w)$



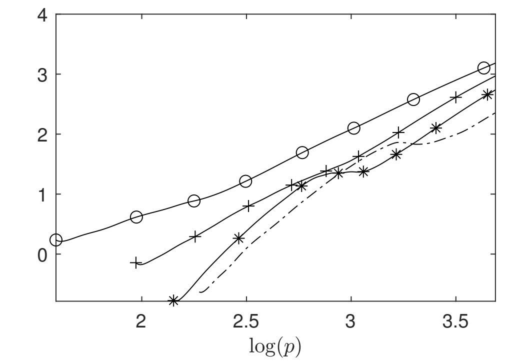
(b) Wage offer distribution $F(w)$



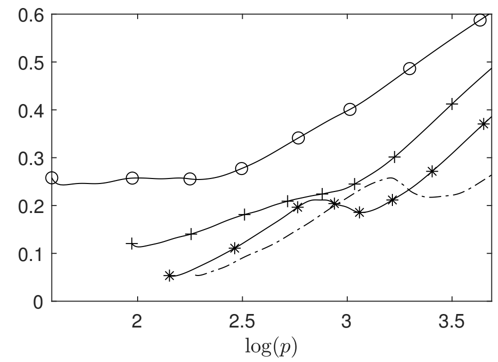
(c) Reservation wage density $dA(w)$



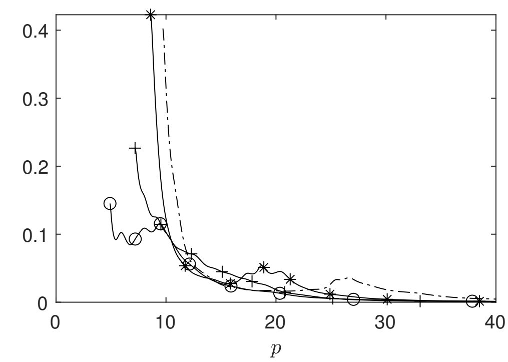
(d) Wage function $K(p)$



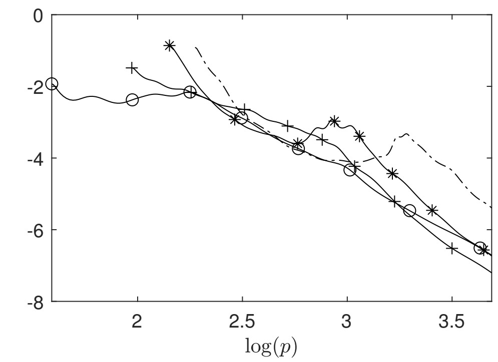
(e) Log absolute markup



(f) Monopsony power $[p - K(p)]/p$



(g) Productivity density $\gamma(p)$



(h) Productivity log density $\log(\gamma(p))$

Figure A.11: Main Equilibrium Functions – by PFE Quartile. *Note:* This figure shows the main equilibrium functions summarizing the steady-state for the estimates by quartile of the person fixed effect. Ordinate outcomes can be found in the subfigures' titles. Wages and productivity are measured in EUR. *Key:* Quartile 1 (—o—); Quartile 2 (—+—); Quartile 3 (—∗—); Quartile 4 (—·—).

Table A.7: Robustness Checks (Whole Sample)

SSC threshold	Hours	Wage measure	Truncation	ρ	η	N	$\underline{w}$	$\overline{w}$	δ	κ_e	κ_u	μ_ϕ	σ_ϕ	u	$u^{8.50}$	$\Delta u^{8.50}$
Censored	Imputed	Avg. one year	3 Euro	0.02	1.75	208626	3.64	33.08	0.0063	7.10	20.06	6.50	3.60	0.0840	0.0808	-0.0031
									(0.0000)	(0.10)	(0.32)	(0.02)	(0.05)	(0.0004)	(0.0003)	(0.0003)
Censored	40	Avg. one year	3 Euro	0.02	1.75	208837	3.61	32.96	0.0063	7.20	20.41	6.55	3.65	0.0842	0.0803	-0.0039
									(0.0000)	(0.09)	(0.30)	(0.03)	(0.05)	(0.0004)	(0.0004)	(0.0004)
Imputed	Imputed	Avg. one year	3 Euro, 95%	0.02	1.75	209012	3.64	33.45	0.0063	7.14	20.15	6.53	3.60	0.0838	0.0809	-0.0029
									(0.0000)	(0.10)	(0.33)	(0.04)	(0.05)	(0.0004)	(0.0004)	(0.0003)
Imputed	Imputed	Avg. one year	3 Euro, 99%	0.02	1.75	218373	3.64	48.27	0.0061	7.58	20.43	6.48	3.42	0.0832	0.0891	0.0059
									(0.0000)	(0.09)	(0.30)	(0.04)	(0.05)	(0.0004)	(0.0008)	(0.0011)
Censored	Imputed	Last and first obs.	3 Euro	0.02	1.75	207103	3.60	33.08	0.0063	7.09	19.92	6.46	3.55	0.0850	0.0810	-0.0040
									(0.0000)	(0.10)	(0.32)	(0.02)	(0.05)	(0.0004)	(0.0003)	(0.0004)
Censored	Imputed	Avg. full spell	3 Euro	0.02	1.75	220257	3.56	33.55	0.0061	6.92	18.76	6.17	3.05	0.0859	0.0801	-0.0058
									(0.0000)	(0.08)	(0.24)	(0.03)	(0.04)	(0.0004)	(0.0003)	(0.0002)
Censored	40	Avg. full spell	3 Euro	0.02	1.75	224006	3.58	32.18	0.0059	6.92	19.63	6.26	3.22	0.0848	0.0777	-0.0071
									(0.0000)	(0.08)	(0.28)	(0.03)	(0.05)	(0.0004)	(0.0004)	(0.0007)
Censored	Imputed	Avg. one year	4 Euro	0.02	1.75	207928	4.51	33.08	0.0063	6.42	19.02	6.71	3.54	0.0821	0.0817	-0.0004
									(0.0000)	(0.07)	(0.25)	(0.05)	(0.06)	(0.0003)	(0.0003)	(0.0002)
Censored	Imputed	Avg. one year	3 Euro	0.01	1.75	208626	3.64	33.08	0.0061	8.34	18.42	6.40	2.86	0.0912	0.0862	-0.0050
									(0.0000)	(0.13)	(0.33)	(0.05)	(0.04)	(0.0004)	(0.0004)	(0.0012)
Censored	Imputed	Avg. one year	3 Euro	0.04	1.75	208626	3.64	33.08	0.0065	5.32	20.50	6.30	5.05	0.0735	0.0716	-0.0019
									(0.0000)	(0.05)	(0.26)	(0.02)	(0.05)	(0.0004)	(0.0004)	(0.0001)
Censored	Imputed	Avg. one year	3 Euro	0.02	1.65	208626	3.64	33.08	0.0063	7.10	20.06	6.50	3.60	0.0840	0.0828	-0.0012
									(0.0000)	(0.10)	(0.32)	(0.02)	(0.05)	(0.0004)	(0.0003)	(0.0003)
Censored	Imputed	Avg. one year	3 Euro	0.02	1.71	208626	3.64	33.08	0.0063	7.10	20.06	6.50	3.60	0.0840	0.0816	-0.0024
									(0.0000)	(0.10)	(0.32)	(0.02)	(0.05)	(0.0004)	(0.0003)	(0.0003)
Censored	Imputed	Avg. one year	3 Euro	0.02	1.79	208626	3.64	33.08	0.0063	7.10	20.06	6.50	3.60	0.0840	0.0802	-0.0038
									(0.0000)	(0.10)	(0.32)	(0.02)	(0.05)	(0.0004)	(0.0003)	(0.0003)
Censored	Imputed	Avg. one year	3 Euro	0.02	1.85	208626	3.64	33.08	0.0063	7.10	20.06	6.50	3.60	0.0840	0.0793	-0.0047
									(0.0000)	(0.10)	(0.32)	(0.02)	(0.05)	(0.0004)	(0.0004)	(0.0004)

Note: The table reports robustness checks with respect to our parameter estimates using either imputed or censored wages, alternative definitions of weekly hours and wages as well as different truncation points. First row: preferred data preparation as reported in Table 1. For alternative data handlings at the SSC threshold, see Appendix A.5. For alternative definitions of weekly hours, see Appendix A.3. For alternative definition of wages, see Appendix A.4. Bootstrapped standard errors in parentheses (250 runs).

Table A.8: Estimated values for η

	Controlling for firm size	Controlling for firm size and AKM FE
Whole sample	1.76	1.71
	(0.05)	(0.05)
Whole sample without highest job requirement level	1.79	1.71
	(0.06)	(0.06)
West	1.56	1.54
	(0.06)	(0.06)
East	2.17	2.06
	(0.10)	(0.10)

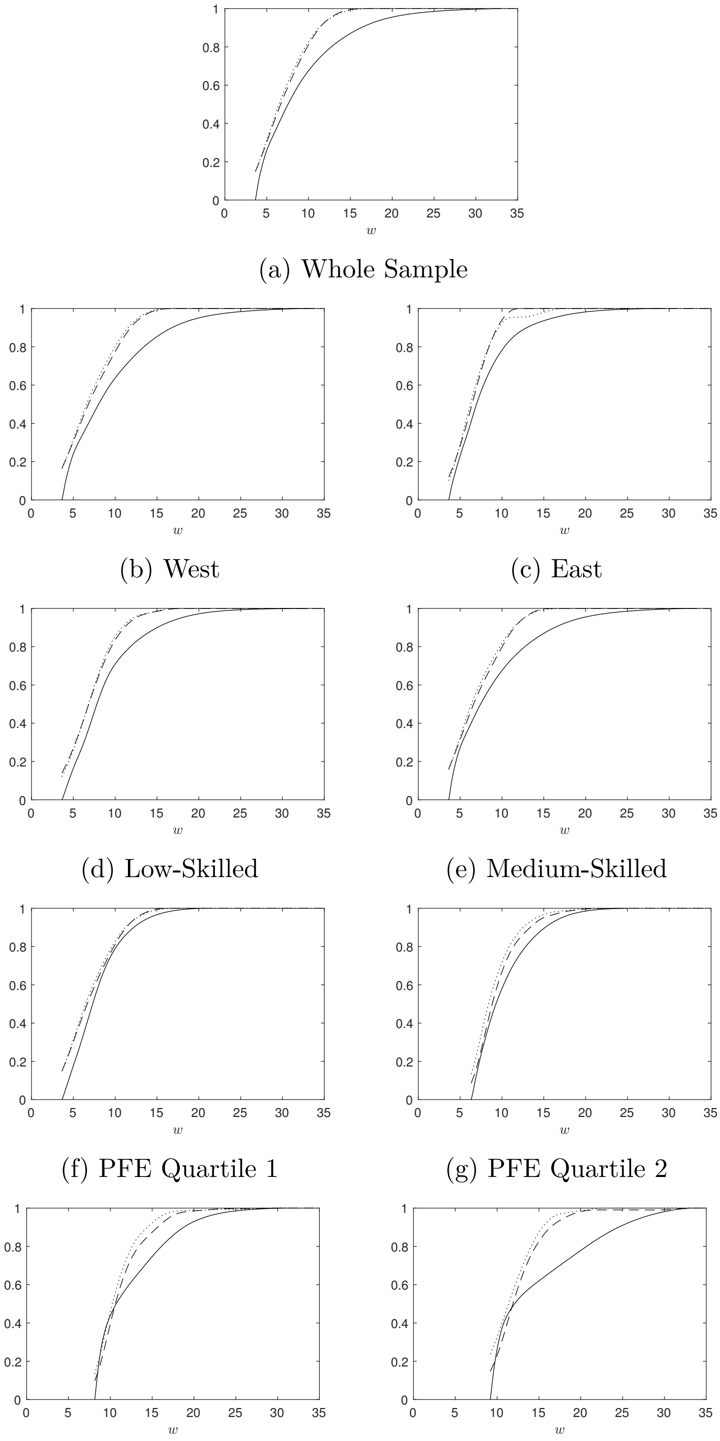
Notes: Lowest job requirement level: no qualification required, medium job requirement level: at least vocational training required, highest job requirement level: at least university degree required. Firm size is measured by four categories (1-9, 10-49, 50-499,500+ employees). AKM establishment fixed effects are measured for the time period 2007-2013. Standard errors are in parentheses. Number of observations: 2,277 - 2,318 (Whole sample), 1,793 - 1,830 (Whole sample without highest job requirement level), 1,454 - 1,484 (West), 823 - 834 (East).

A.10. Model Fit

Table A.9: In-Sample Fit: Transitions After One Year

	(1)	(2)	(3)	(4)	(5)	(6)
	Empirical			Predicted		
Sample	u → e	e → e	e → u	u → e	e → e	e → u
Whole sample	3.6	7.7	6.0	4.5	8.4	4.9
Region						
West	3.4	7.8	5.7	4.4	8.7	4.7
East	4.8	7.1	7.1	5.1	8.8	5.3
Skill Group						
Low-Skilled	5.7	8.1	8.9	5.3	9.3	5.3
Medium-Skilled	3.4	7.6	5.7	4.5	8.5	4.8
PFE Quartile						
Quartile 1	6.7	7.8	11.4	7.8	9.8	8.1
Quartile 2	5.0	8.5	7.1	5.0	10.3	5.2
Quartile 3	2.0	7.3	3.5	3.8	7.1	4.0
Quartile 4	0.9	6.5	2.7	3.5	5.3	3.8

Note: Arrows (→) indicate that spells end in transitions to another employment spell (e) or to unemployment (u). Spells without an observed transition are right-censored. Additionally, spells might be left-censored. Predicted transitions are based on a simulation for 10,000 individuals within each sample. To obtain the complements of the transition rates similar to those reported in Figures A.13 and A.14, the $u \to e$-figures need to be divided by u and the $e \to e$- and $e \to u$-figures by (1-u), respectively. Discrepancies between the figures "Empirical" and the corresponding ones in Figures A.13 and A.14 may arise as the former represent the fraction of individuals with a transition within one year among all individuals observed on January 1, 2010, while the Kaplan-Meier-estimates in Figures A.13 and A.14 are based on the residual duration of unemployed and employed individuals, respectively.



(a) Whole Sample

(b) West

(c) East

(d) Low-Skilled

(e) Medium-Skilled

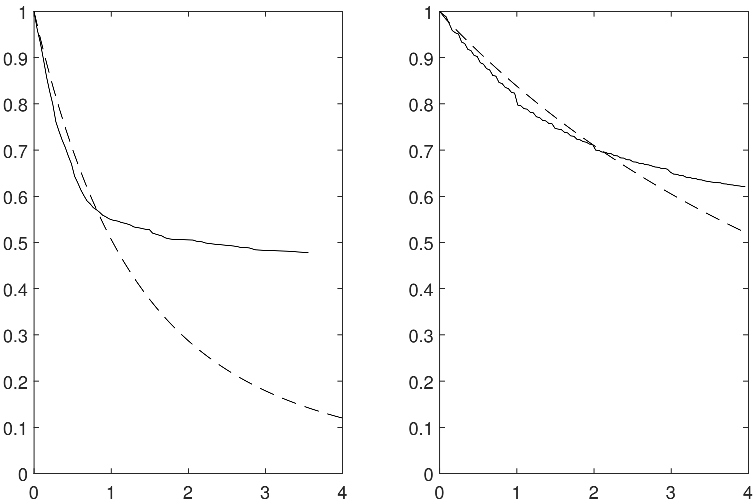
(f) PFE Quartile 1

(g) PFE Quartile 2

(h) PFE Quartile 3

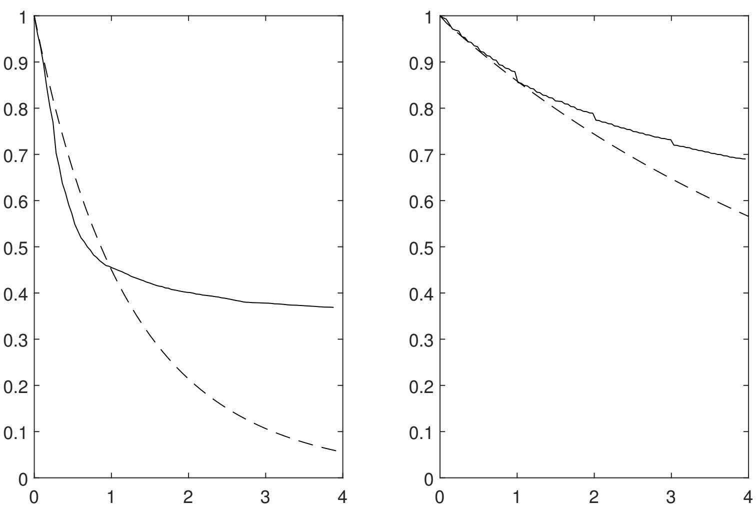
(i) PFE Quartile 4

Figure A.12: In-Sample Fit: Wage Distributions. *Note:* This figure shows the steady-state wage offer distribution $F(w)$ (——) in comparison to empirical wage distributions. The dashed line (----) shows the CDF of accepted wages after a job-to-job transition for wages in the range of $[\underline{w}; 1.25\underline{w}]$. The dotted line (⋯⋯) shows the CDF of accepted wages after a job-to-job transition for wages in the range of $[\underline{w}; \underline{w}+1]$.



(a) Unemployment (b) Employment

Figure A.13: Survival Probabilities (residual duration) – Low-Skilled. *Note:* This figure shows the Kaplan-Meier survival estimates for unemployment and employment residual durations in years (——) in comparison to the predicted survival function (----) which is given by $S_u(t) = A(\underline{w})/(1+\kappa_u) \cdot \exp(-\lambda_u \cdot \overline{F}(\underline{w}) \cdot t)/u + \int_{\underline{w}}^{\overline{w}} dA(x)/(1+\kappa_u \overline{F}(x)) \cdot \exp(-\lambda_u \cdot \overline{F}(x) \cdot t)/u + (1 - A(\overline{w})) \cdot \frac{1}{u}$ and $S_e(t) = \int_{\underline{w}}^{\overline{w}} \exp(-(\delta + \lambda_e \cdot \overline{F}(x)) \cdot t) dG(x)$. The figures' captions refer to the origin labor market state.



(a) Unemployment (b) Employment

Figure A.14: Survival Probabilities (residual duration) – Medium-Skilled. *Note:* This figure shows the Kaplan-Meier survival estimates for unemployment and employment residual durations in years (——) in comparison to the predicted survival function (----) which is given by $S_u(t) = A(\underline{w})/(1+\kappa_u) \cdot \exp(-\lambda_u \cdot \overline{F}(\underline{w}) \cdot t)/u + \int_{\underline{w}}^{\overline{w}} dA(x)/(1+\kappa_u \overline{F}(x)) \cdot \exp(-\lambda_u \cdot \overline{F}(x) \cdot t)/u + (1 - A(\overline{w})) \cdot \frac{1}{u}$ and $S_e(t) = \int_{\underline{w}}^{\overline{w}} \exp(-(\delta + \lambda_e \cdot \overline{F}(x)) \cdot t) dG(x)$. The figures' captions refer to the origin labor market state.

Table A.10: In-Sample Fit: Unemployment Rate

	Empirical	Predicted
Whole sample	0.081	0.084
Region		
West	0.077	0.080
East	0.106	0.090
Skill Group		
Low-Skilled	0.154	0.107
Medium-Skilled	0.075	0.083
PFE Quartile		
Quartile 1	0.189	0.160
Quartile 2	0.093	0.075
Quartile 3	0.037	0.058
Quartile 4	0.030	0.065

Note: For empirical values, also see Table A.5. The predicted unemployment rate is based on equation (5) as implied by our estimates (see also Table 1 and Table A.13).

Table A.11: Out-of-Sample Fit: Unemployment Rate

	Empirical	Predicted
Whole sample	0.051	0.081
Region		
West	0.049	0.076
East	0.060	0.108
Skill Group		
Low-Skilled	0.103	0.109
Medium-Skilled	0.048	0.079
PFE Quartile		
Quartile 1	0.118	0.163
Quartile 2	0.049	0.076
Quartile 3	0.022	0.058
Quartile 4	0.024	0.065

Note: Empirical refers to an out-of sample period using the original data set and the same sample restrictions. The unemployment rate refers to 1 January 2016. Note that the out-of-sample period is affected by shocks not related to the minimum wage as discussed in Section 7.4. The predicted unemployment rate is based on equation (5) as implied by our estimates Table A.13 for a counterfactual of a minimum wage of EUR 8.50.

Table A.12: Out-of-Sample Fit: Transitions After One Year

	(1)	(2)	(3)	(4)	(5)	(6)
	Empirical			Predicted		
Sample	u → e	e → e	e → u	u → e	e → e	e → u
Whole sample	1.8	6.7	3.1	4.5	8.5	4.7
Region						
West	1.7	6.8	3.1	4.4	8.8	4.6
East	2.5	6.6	3.4	5.4	7.4	5.4
Skill Group						
Low-Skilled	2.1	5.8	4.1	5.5	8.8	5.3
Medium-Skilled	1.8	6.8	3.1	4.5	8.5	4.7
PFE Quartile						
Quartile 1	3.3	8.1	5.9	8.2	9.3	8.3
Quartile 2	2.2	7.3	3.3	4.9	9.8	5.2
Quartile 3	1.0	5.9	1.8	3.9	7.1	4.0
Quartile 4	0.8	5.4	1.7	3.6	5.2	3.9

Note: Arrows (→) indicate that spells end in transitions to another employment spell (e) or to unemployment (u). Spells without an observed transition are right-censored. Additionally, spells might be left-censored.

A.11. Minimum Wage Simulations

Table A.13: Unemployment Rate u for Different Minimum Wages

	(1)	(2)	(3)	(4)	(5)	(6)
	Level	Change compared to no MW				
Minimum wage	No MW	7.00 euro	8.50 euro	10.00 euro	11.50 euro	13.00 euro
Whole Sample	0.084	-0.003	-0.003	-0.003	-0.003	-0.001
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
By Region						
West	0.080	-0.004	-0.005	-0.005	-0.004	-0.002
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
East	0.090	0.010	0.018	0.026	0.032	0.035
	(0.001)	(0.003)	(0.005)	(0.006)	(0.007)	(0.007)
Total	0.082	-0.002	-0.001	-0.000	0.001	0.004
	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
By Skill Group						
Low-Skilled	0.107	-0.002	0.001	0.006	0.010	0.015
	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
Medium-Skilled	0.083	-0.003	-0.004	-0.005	-0.004	-0.002
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Total	0.085	-0.003	-0.004	-0.004	-0.003	-0.001
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)
By PFE Quartile						
Quartile 1	0.160	-0.002	0.004	0.012	0.019	0.024
	(0.002)	(0.001)	(0.001)	(0.004)	(0.006)	(0.007)
Quartile 2	0.075	-0.000	0.001	0.003	0.005	0.009
	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)
Quartile 3	0.058	0.000	0.001	-0.002	-0.003	-0.003
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Quartile 4	0.065	0.000	0.000	-0.008	-0.011	-0.012
	(0.001)	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)
Total	0.083	-0.000	0.001	0.001	0.002	0.004
	(0.000)	(0.000)	(0.000)	(0.001)	(0.001)	(0.001)

Note: The table reports the level of u and model predicted changes in u for different minimum wage levels. The first row shows simulations based on the whole sample while the rows *Total* aggregate the minimum wage effects over subsamples. Bootstrapped standard errors in parentheses (250 runs).

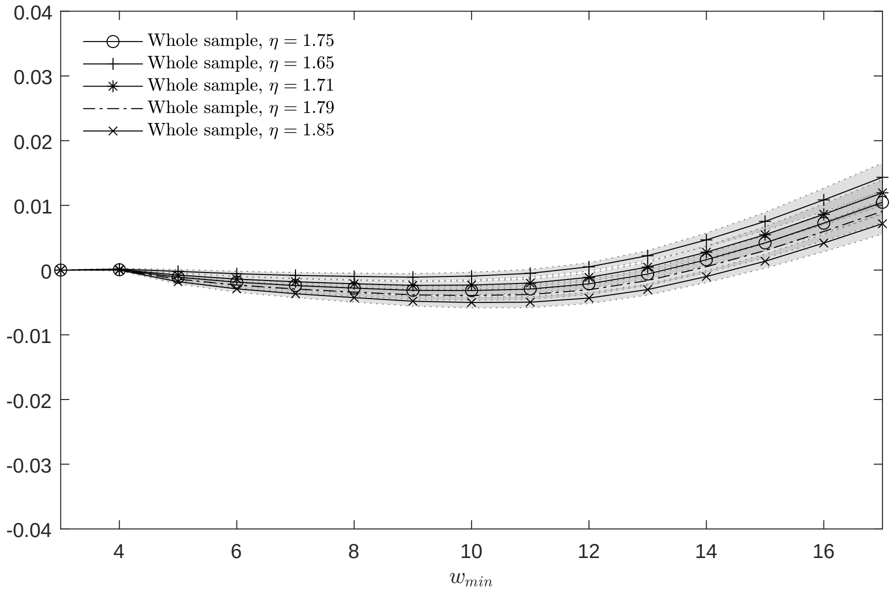


Figure A.15: Change in Unemployment Rate for different calibrations of η (Whole Sample). This figure shows, for the whole sample, the change in the unemployment rate for different values of η relative to the status quo level. Grey area indicates 95% confidence bands based on 250 bootstrap runs.

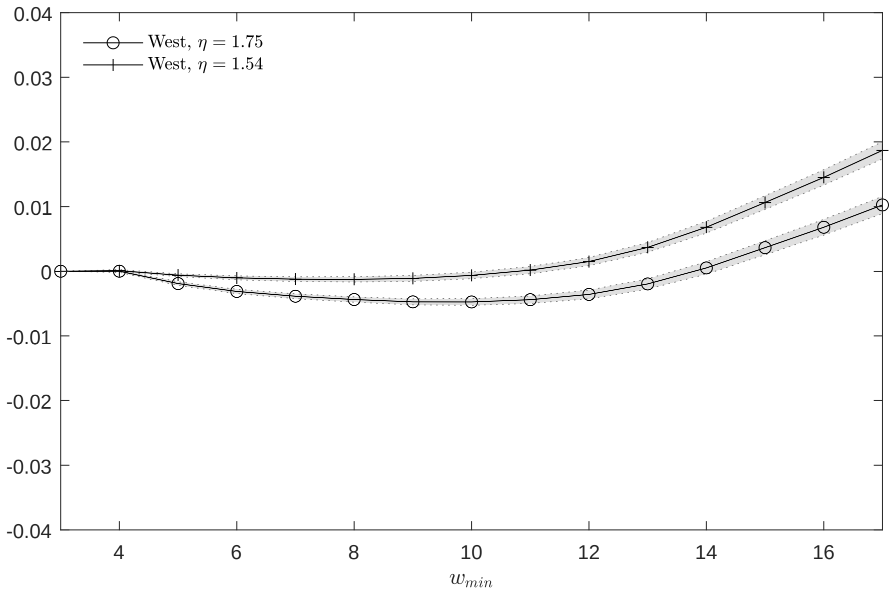


Figure A.16: Change in Unemployment Rate for different Values of η – West Germany. This figure shows the change in the unemployment rate for different minimum wage levels based on separate estimations for West Germany for different values of η relative to the status quo level. Grey area indicates 95% confidence bands based on 250 bootstrap runs.

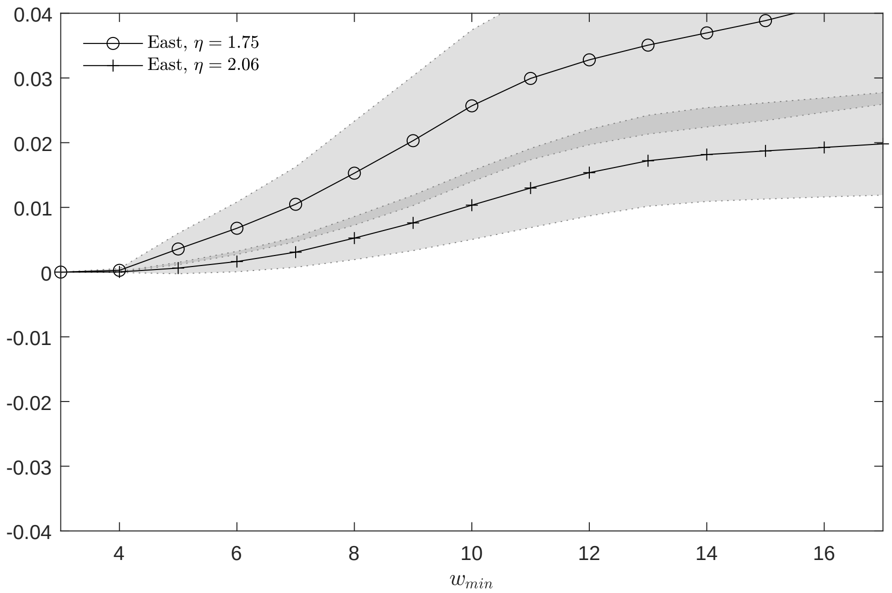


Figure A.17: Change in Unemployment Rate for different Values of η – East Germany. This figure shows the change in the unemployment rate for different minimum wage levels based on separate estimations for East Germany for different values of η relative to the status quo level. Grey area indicates 95% confidence bands based on 250 bootstrap runs.

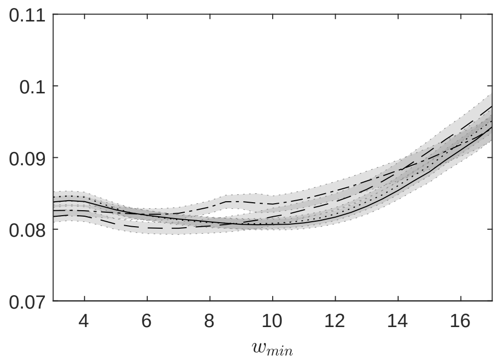


Figure A.18: Unemployment Rate u for Different Minimum Wages. The figure shows the unemployment rate for the whole sample (——) and totals over different subsamples (cf. Table A.13). *Key:* Total over Region (----); Total over Skill Group (··········); Total over PFE Quartile (-·-·-·). Grey area indicates 95% confidence bands based on 250 bootstrap runs.