Is there monopsonistic discrimination against immigrants?*

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***Abstract*:** This paper investigates immigrants' and natives' labour supply to the firm within a semi-structural approach based on a dynamic monopsony framework. Applying duration models that account for unobserved worker heterogeneity to a large administrative employer–employee data set for Germany, we find that immigrants supply labour less elastically to firms than natives. Under monopsonistic wage setting, the estimated elasticity differential predicts a 7.7 log points wage penalty for immigrants thereby accounting for almost the entire unexplained native–immigrant wage differential of 5.8–8.2 log points. When further distinguishing immigrant groups differing in their time spent in the German labour market, their immigration cohort, and their age at entry, we find that the observed unexplained wage differential is larger for those groups who show a larger elasticity differential relative to natives. These findings are not only suggestive that search frictions are a likely cause of employers' more pronounced monopsony power over their immigrant workers but also imply that employers profit from discriminating against immigrants.

***Keywords*:** monopsony, native–immigrant wage differential, discrimination, Germany

***JEL classification*:** J42, J61, J71

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1 Introduction

Just recently, there have been a growing number of both theoretical and empirical studies investigating the prevalence and causes of monopsony in the labour market (Ashenfelter *et al*., 2010; Manning, 2011). In contrast to the traditional literature on monopsony that saw monopsony power primarily as a consequence of labour markets characterised by demand-side concentration or collusion, this “new” monopsony literature stresses that monopsony power may also be pervasive in labour markets consisting of many competing employers due to search frictions (e.g., Burdett and Mortensen, 1998; Manning, 2003*a*), mobility costs (e.g., Brueckner *et al.*, 2002; Manning, 2003*b*), or heterogeneous preferences over non-wage job characteristics (e.g., Bhaskar and To, 1999; Booth and Coles, 2007).

Empirically, the question whether labour markets should be viewed as monopsonistic rather than competitive or whether the model of perfect competition is a sufficiently good approximation to real-world labour markets boils down to the question whether the labour supply curve to the single firm is horizontal or imperfectly elastic. There is a vivid recent empirical literature that tries to infer the firm-level labour supply elasticity either by exploiting natural experiments (Falch, 2010; Staiger *et al.*, 2010) or by using a semi-structural estimation approach proposed by Manning (2003*a*), which builds on equilibrium search theory with wage posting (Ransom and Sims, 2010; Booth and Katic, 2011; Depew and Sørensen, 2011; Falch, 2011). In summary, this literature finds that the elasticity is far from infinite implying that employers possess substantial monopsony power, so that there is scope for marked deviations from the competitive wage.[1]

Another strand of this “new” monopsony literature has tried to assess whether monopsonistic discrimination, the roots of which originate in Robinson’s (1933) seminal work applying third-degree price discrimination to the labour market, could be a fruitful framework

[1] For a recent overview of this literature, see Manning (2011).

to think about the gender pay gap as there are several reasons to consider women's labour supply to the firm as less elastic than men's. And indeed, a growing literature is emerging which finds that women's labour supply to the firm is less elastic than men's, so that monopsonistic discrimination may explain part of the unexplained gender pay gap in the data (Barth and Dale-Olsen, 2009; Hirsch *et al.*, 2010; Ransom and Oaxaca, 2010).

Up to now, however, no attempt has been made to apply the insights of the "new" monopsony literature to native–immigrant wage differentials, perhaps on account of the immigration literature's focus on questions of wage assimilation and labour market segmentation and less so on discrimination issues. As a consequence, there is no study investigating native–immigrant differences in firm-level labour supply elasticities. This comes at a surprise since there are several reasons to presume that such differences exist and monopsonistic discrimination could contribute to the explanation of native–immigrant wage differentials.

As numerous studies document (see the survey by Dustmann and Glitz, 2011), immigrants experience a considerable wage disadvantage compared to natives upon arrival in the host country and manage to close part of the gap during their stay. Following the seminal work by Chiswick (1978), this partial wage assimilation is typically explained in terms of immigrants gaining host country-specific human capital augmenting their productivity. Yet, from a monopsonistic perspective immigrants' lack of country-specific skills, in particular missing knowledge about labour market institutions, may result in more severe search frictions for immigrants, thereby providing employers with the possibility of engaging in monopsonistic discrimination against them. It may thus be promising to investigate the case for monopsonistic discrimination against immigrants. This paper is intended to fill this gap by estimating labour supply elasticities for native and immigrant males using Manning's (2003*a*) semi-structural estimation approach. In particular, we examine whether a monopsonistic perspective may help understanding the native–immigrant wage differential. The richness and the long time horizon of our administrative linked employer–employee data also enable us to improve on the earlier literature by explicitly controlling for unobserved worker heterogeneity using stratified partial likelihood and conditional logit estimators.

Our main finding is that immigrants show a lower labour supply elasticity than natives and also earn considerably less than natives when controlling for different human capital endowments and workplace characteristics. Remarkably, the wage differential predicted from a simple rule of monopsonistic wage setting is of similar magnitude as the observed unexplained wage differential. Our reading of these results is that monopsonistic discrimination accounts for almost the entire unexplained native–immigrant wage differential. When distinguishing subgroups of immigrants differing in their time spent in the host country, immigration cohort, or age at entry, we also find that the unexplained wage differential is larger for those groups that show a larger elasticity differential relative to natives. These results not only strengthen the case for monopsonistic discrimination, but also are suggestive that search frictions faced by immigrants are a likely source of employers' more pronounced monopsony power over immigrant workers.

The remainder of this paper is organised as follows: Section 2 presents some theoretical considerations on possible reasons why native and immigrant workers may differ in their firm-level labour supply behaviour. Section 3 develops our econometric approach used to infer the firm-level labour supply elasticities from the data, and Section 4 presents our data set. Section 5 includes some descriptive evidence on job mobility and the immigrant wage gap in our sample, while Section 6 presents and discusses our empirical results on native and immigrant workers' firm-level labour supply elasticities. Section 7 concludes.

2 Theoretical Considerations

In most countries, there exist significant native–immigrant wage differentials with immigrants' earnings being considerably lower than natives' (e.g., Dustmann and Glitz, 2011; for

Germany, e.g., Algan *et al.*, 2010; Dustmann *et al.*, 2010). One obvious reason for these differentials may be different endowments in human capital among groups: Immigrants and natives may, for instance, differ in their education levels and their occupations. Furthermore, immigrants may lack country-specific human capital upon arrival in the host country but catch up with natives' earnings over time as they acquire the missing country-specific skills and gain in productivity. Following Chiswick (1978), a vast literature has investigated whether there exists an assimilation process that reduces wage differences between immigrants and natives in the course of time spent in the host country. The general finding of this literature is that there are large initial differentials that tend to decline as immigrants spend time in the host country. Nonetheless, even after a long stay in the host country immigrants still get paid lower wages than comparable natives (cf. Dustmann and Glitz, 2011).

Apart from productivity differences stemming from missing country-specific skills, part of the native–immigrant wage differential may also reflect discrimination against immigrants. While prejudices against immigrants may give rise to Beckerian (1971) taste-based discrimination, employers may also possess more monopsony power over immigrant than over native workers resulting in monopsonistic discrimination against immigrants (Robinson, 1933). Indeed, several studies have found non-trivial unexplained wage differentials likely to reflect discrimination against immigrants (e.g., Kee, 1995; Nielsen *et al.*, 2004; Elliot and Lindley, 2008; Lehmer and Ludsteck, 2011).

Empirically, the scope of monopsonistic discrimination as an explanation of immigrant–native wage differentials leads to the question whether immigrant workers supply labour less elastically to firms than native workers. Put differently, the question is whether immigrants' labour supply to the firm is less driven by wages than natives'. As the "new" monopsony literature stresses, search frictions, mobility costs, and heterogeneous preferences over non-wage job characteristics are the main forces limiting the elasticity of workers' labour supply to employers. If we think of search frictions – as in the Burdett and Mortensen (1998) model – as a source of employers' monopsony power, immigrants lacking country-specific skills may find it harder to climb the job ladder because they face more severe search frictions and thus end up with lower wages than natives. For instance, immigrants may have a lower job finding rate because of limited command of the host country's language, limited knowledge on search channels, application routines and institutional details, or less extensive search networks. Furthermore, there is evidence of considerable workplace segregation of immigrants (Åslund and Skans, 2010; Glitz 2012*b*) likely to limit their choice of employers. What is more, immigrants have been found to be more risk averse and less mobile than natives (Jaeger *et al.*, 2010) and to form enclaves (e.g., Cutler *et al.*, 2008; Granato, 2009), which may also limit their regional mobility. All these factors are likely to further impede immigrants' search activities and therefore should provide employers with additional monopsony power over them.

In line with these considerations, Hotchkiss and Quispe-Agnoli (2012) report that undocumented U.S. workers in Georgia supply labour less elastically to the firm than documented ones and also show that this explains part of the observed within-firm wage differential between these two groups. At first sight, one may argue that undocumented workers face similar restraints in their labour supply behaviour as immigrants. As undocumented workers can rely only on informal search channels, however, search frictions are arguably of different nature for undocumented workers than for the whole group of immigrants. Therefore, comparing firm-level labour supply elasticities for undocumented and documented workers is unlikely to be informative of the native–immigrant differential which is our point of interest.

That said, there are also some reasons to conjecture immigrants to supply labour even more elastically to firms than natives. If we think of horizontal job differentiation as one source of monopsony power, immigrants may have less pronounced preferences over non-wage job characteristics and may thus be more driven by pecuniary considerations than natives. For example, immigrants may be less attached to particular firm cultures and institutional details. Furthermore, immigrants typically experience substantial occupational downgrading when entering the host country's labour market (e.g., Chiswick *et al.*, 2005; Constant and Massey, 2005) and may try to reduce their occupational mismatch by increased job shopping activities.

It is therefore unclear *ex ante* whether the elasticity-inhibiting or the elasticity-enhancing factors dominate and whether monopsonistic discrimination against immigrants is thus viable. To assess whether monopsony may explain part of the native–immigrant wage differential, we will in the following use a semi-structural estimation approach proposed by Manning (2003*a*) to infer the firm-level labour supply elasticity from workers' transition behaviour in the labour market separately for immigrants and natives.

In a second step, we shall investigate whether the firm-level labour supply elasticity varies for different subgroups of immigrants. We consider immigrants differing along three dimensions: (i) their time spent in the host country, (ii) their immigration cohort, and (iii) their age at entry. Since these immigrants are likely to differ in their search frictions (as detailed in Section 6), this not only allows us to further scrutinise the case for monopsonistic discrimination against immigrants but also to shed light on the sources of employers' monopsony power over immigrants.

3 Econometric Approach

The starting point of our econometric approach, which has been pioneered by Manning (2003*a*, pp. 96–104), is a simple dynamic monopsony model for the labour supply to the firm. Consider a firm paying some wage w at time t. We model the change in the labour supply to this firm $L(w)$ as

$$\dot{L}(w) = R(w) - s(w)L(w), \tag{1}$$

where $R(w) > 0$ denotes the number of recruits arriving at the firm at time t with $R^{'} > 0$ and $0 < s(w) < 1$ the separation rate with $s^{'} < 0$. Accordingly, we assume that the firm can increase its labour supply by increasing its wage and that the labour supply adjusts sluggishly over time. Now consider a steady state with $\dot{L}(w) = 0$. Then, we arrive at

$$L(w) = R(w)/s(w) \tag{2}$$

with $L^{'} > 0$.[2] From (2) we get the long-run labour supply elasticity at the level of the firm ε_{Lw} as the difference of the wage elasticity of recruits ε_{Rw} and the wage elasticity of the separation rate ε_{sw}

$$\varepsilon_{Lw} = \varepsilon_{Rw} - \varepsilon_{sw}. \tag{3}$$

Equation (3) further simplifies once we impose more structure on the model. Making use of Burdett and Mortensen's (1998) equilibrium search model with wage posting, which can be thought of as a dynamic general equilibrium model of monopsonistic competition, Manning demonstrates that $\varepsilon_{Rw} = -\varepsilon_{sw}$, so that the supply elasticity becomes

$$\varepsilon_{Lw} = -2\varepsilon_{sw}. \tag{4}$$

Intuitively, this holds because in this model one firm's wage-related hire is another firm's wage-related quit as transitions from and to non-employment are thought of as wage-inelastic. Equation (4) would allow us to identify the firm-level labour supply elasticity by just estimating the separation rate elasticity and is the basis of the studies by Barth and Dale-Olsen (2009), Ransom and Oaxaca (2010), Ransom and Sims (2010), as well as Falch (2011).

Incorporating wage-elastic transitions from and to non-employment, for instance due to voluntary unemployment induced by welfare payments, Manning further shows that the

[2] Note that perfect competition is nested as the case with $L' \to \infty$ due to $s' \to -\infty$ and $R' \to \infty$ at the competitive market wage that equalises labour supply and demand at the market level.

firm-level labour supply elasticity is given by the difference of a weighted average between the recruitment elasticities from employment and non-employment, ε_{Rw}^{e} and ε_{Rw}^{n}, and the separation rate elasticities to employment and non-employment, ε_{sw}^{e} and ε_{sw}^{n},

$$\varepsilon_{Lw} = \theta_R \varepsilon_{Rw}^{e} + (1-\theta_R)\varepsilon_{Rw}^{n} - \theta_s \varepsilon_{sw}^{e} - (1-\theta_s)\varepsilon_{sw}^{n} \tag{5}$$

with the weights being given by the share of recruits from employment θ_R and the share of separations to employment θ_s, respectively. In a steady state, $\theta \equiv \theta_R = \theta_s$ holds.

In Burdett and Mortensen's model, workers change employers whenever offered a wage higher than the on-going one, which is quite restrictive. Allowing for stochastic job-to-job moves and adding more structure to the model, in particular by modelling the separation rate and recruitment functions as iso-elastic in wages, Manning demonstrates that the long-run elasticity of the labour supply to the firm is given by

$$\varepsilon_{Lw} = -(1+\theta)\varepsilon_{sw}^{e} - (1-\theta)\varepsilon_{sw}^{n} - \varepsilon_{\theta w}, \tag{6}$$

where $\varepsilon_{\theta w}$ denotes the wage elasticity of the share of recruits hired from employment and, together with θ, informs us on the recruitment function of the firm.[3] This more general approach is adopted by Hirsch *et al.* (2010) and Booth and Katic (2011). To estimate the supply elasticity as given in equation (6), one has to estimate (i) the separation rate elasticity to employment ε_{sw}^{e}, (ii) the separation rate elasticity to non-employment ε_{sw}^{n}, (iii) the wage elasticity of the share of hires from employment $\varepsilon_{\theta w}$, and (iv) the share of hires from employment θ.

Following Manning's approach, we estimate the separation rate elasticities ε_{sw}^{e} and ε_{sw}^{n} from two hazard models for the instantaneous separation rates to employment and non-employment. We model the respective separation rate of job spell i belonging to worker m as

[3] Note that the more simple approach as given by equation (4) is nested as $\varepsilon_{sw}^{n}, \varepsilon_{\theta w} \to 0$ and $\theta \to 1$.

$$s_i^r\left(t|\mathbf{x}_i^r(t), \upsilon_{m(i)}^r(t)\right) = s_0^r(t)\exp(\mathbf{x}_i^r(t)'\boldsymbol{\beta}^r)\,\upsilon_{m(i)}^r(t) \tag{7}$$

with route $r = e, n$, baseline hazard $s_0^r(t)$, a vector of time-varying covariates $\mathbf{x}_i^r(t)$, a vector of coefficients $\boldsymbol{\beta}^r$, and unobserved worker heterogeneity $\upsilon_{m(i)}^r(t)$, where t corresponds to the time since the start of the job spell, i.e. job tenure. Including the log wage as covariate results in an iso-elastic separation rate as required by Manning's approach, and its coefficient gives the respective separation rate elasticity ε_{sw}^r.[4]

We should make clear that the inclusion of unobserved worker heterogeneity to the separation equations is indispensible for our investigation as we have only few information on natives' and immigrants' socio-economic backgrounds, though these are likely to differ considerably. As a case in point, our data (for the details, see Section 4) do not contain information on immigrants' pre-migration labour market performance or their motivation. We do not know either whether immigrants experienced occupational downgrading when entering the German labour market. All these factors may obviously affect immigrants' transition behaviour and cause it to differ substantially from natives'. However, they are only partly reflected in, though certainly correlated with immigrants' observed characteristics, so that omitting unobserved worker heterogeneity is most likely to cause severe bias in the estimates.

To tackle this problem, we make use of stratified Cox models in which the baseline hazard $s_0^r(t)$ is some arbitrary non-negative function of job tenure as is the unobserved worker heterogeneity $\upsilon_{m(i)}^r(t)$.[5] Put differently, in a stratified Cox model the baseline hazard is assumed to be worker-specific and thus captures time-invariant unobserved worker heterogeneity. It is important to stress that by allowing for worker-specific baseline hazards the proportionality assumption inherent to the class of hazard rate models defined by (7)

[4] Assuming conditional independence of the separation probabilities to employment and non-employment Manning shows that they can be estimated separately by two independent hazard rate models. When estimating the separation rate to non-employment all job spells are used. Yet, when estimating the separation rate to employment only those job spells that do not end with a transition to non-employment are considered.

[5] See Kalbfleisch and Prentice (2002) for details on (stratified) Cox models.

needs to hold only for job spells belonging to the same worker but may very well be violated across workers without invalidating identification (cf. Kalbfleisch and Prentice, 2002, pp. 118/119). As a consequence, our estimations relying on stratified Cox models do not suffer from the widely raised criticism against proportional hazard models.

To estimate stratified Cox models, we adopt the stratified partial likelihood estimator that allows us to sweep out the worker-specific baseline hazard without the need of identifying it and thus to estimate the covariates' coefficients while controlling for unobserved worker heterogeneity in a similarly convenient way as with the within estimator in linear fixed-effects models (cf. Ridder and Tunalı, 1999). The stratified partial likelihood estimator does so by resting the identification of $\boldsymbol{\beta}^r$ on within-variation at the worker level. Hence, it requires multiple job spells per worker and thus long and rich enough data. The use of stratified Cox models therefore allows us to go beyond the existing literature on monopsonistic (gender) discrimination, such as Barth and Dale-Olsen (2009), Hirsch *et al.* (2010), and Ransom and Oaxaca (2010), by controlling for unobserved worker heterogeneity that may be correlated with included covariates. It also considerably mitigates possible concerns on the endogeneity of workers' wages in the separation equations.

That said, estimating stratified Cox models implies that we control for job tenure as the worker-specific baseline hazard $s_0^r(t)v_{m(i)}^r(t)$ in equation (7) drops out of the partial likelihood function without being constrained to be constant over job tenure t. However, from a theoretical perspective, it may seem at first sight plausible to restrict the separation rate not to depend on job tenure. As argued by Manning (2003*a*, p. 103), one of the key results of dynamic monopsony models *à la* Burdett and Mortensen (1998) giving rise to firm-level labour supply as in equation (1) is that firms raise wages in order to increase tenure. Thus, controlling for tenure might take away original variation from the wage variable. Yet, as Manning also notes, the positive relationship between wages and tenure may be spurious under seniority wage scales. Since there is ample evidence suggesting that these are widespread in the German labour market (see, e.g., Zwick, 2011; 2012), we think controlling for tenure is appropriate in our application.

To estimate the wage elasticity of the share of recruits hired from employment $\varepsilon_{\theta w}$, we model the probability that a worker is hired from employment (as opposed to non-employment) as a logit model

$$\Pr\left[y_i = 1 | \mathbf{x}_i, \upsilon_{m(i)}\right] = \Lambda\left(\mathbf{x}_i^{'}\boldsymbol{\beta} + \upsilon_{m(i)}\right), \tag{8}$$

where notation follows the same rules as before, y_i is an indicator variable for a hire from employment, and Λ denotes the c.d.f. of the standard-logistic distribution. To get rid of unobserved worker heterogeneity correlated with observed characteristics, we use the conditional logit (or fixed-effects logit) estimator (see, e.g., Cameron and Trivedi, 2005). Intuitively, this estimator controls for worker fixed effects by conditioning on those workers who are at one point of time hired from employment and from non-employment at another and discarding those always hired from the same labour market status. Identification of $\boldsymbol{\beta}$ thus relies on those workers for whom y_i changes over time, that is workers with at least two job spells and different previous labour market status. As can be easily shown, when including the log wage as a covariate in (8) its coefficient gives the wage elasticity of the share of recruits hired from employment $\varepsilon_{\theta w}$ divided by $1 - \theta$. So multiplying the coefficient by $1 - \theta$ yields the estimate of the wage elasticity we are aiming for.

The last step to get an estimate of the firm-level labour supply elasticity is to estimate the share of hires from employment θ. We do so by simply taking the observed average from the data. Then using (6) yields an estimate for the labour supply elasticity with standard errors for the estimated elasticity being provided by the bootstrap.

To assess the economic relevance of the estimated native–immigrant firm-level supply elasticity differential, we use the fact that in a simple (dynamic) model of monopsony a worker’s wage is given by

$$w = \frac{\varepsilon_{Lw}}{1+\varepsilon_{Lw}} \phi, \tag{9}$$

where ϕ denotes the worker's marginal revenue product.[6] Consider now immigrants and natives with the same marginal revenue product ϕ but different elasticities ε_{Lw}^{I} and ε_{Lw}^{N}, where the superscripts I and N refer to immigrant and native workers, respectively. Assuming the same productivity across workers is plausible as our stratified Cox and conditional logit estimations have controlled for both standard human capital variables and unobserved worker heterogeneity. Then the immigrant–native wage differential following from equation (9) is

$$\ln w^{I} - \ln w^{N} \approx \frac{w^{I} - w^{N}}{w^{N}} = \frac{\varepsilon_{Lw}^{I} - \varepsilon_{Lw}^{N}}{\varepsilon_{Lw}^{N}(\varepsilon_{Lw}^{I} + 1)}, \tag{10}$$

and equation (10) provides us with an estimate of the immigrant–native wage differential under monopsonistic discrimination.

4 Data

To put this approach into practice, we need detailed data on job lengths, preceding and subsequent jobs and periods of non-employment, as well as on workers and employers over a long period of time. Otherwise, correcting for unobserved worker heterogeneity by means of stratified Cox and conditional logit models and multiple-spell data could not be done convincingly. For our purpose we combine two administrative data sets for the period 1985–2008: the Sample of Integrated Labour Market Biographies (SIAB) and the Establishment History Panel (BHP) provided by the Institute for Employment Research (IAB).

The data on job lengths (on a daily basis), transitions, and worker characteristics come from the SIAB. The SIAB comprises a 2 per cent random sample of all wage and salary

[6] To derive equation (9) in a dynamic monopsony model, we would have to assume that firms maximise steady-state profits and discount future profits at a negligible rate.

employees registered with the German social security system during the period 1975–2008, where about 80 per cent of all people employed in Germany are covered by the system (for details, see Dorner *et al.*, 2010). Since the information contained is used to calculate social security contributions, the data set is highly reliable and especially useful for analyses taking wages and job durations into account.

Information on employers comes from the BHP which again consists of data from the German social insurances that are this time aggregated at the plant level as of the 30th of June of a year (for details, see Spengler, 2008). It not only contains information on plants' workforce composition and size but also on downsizing and plant closures. This is particularly important as we aim at identifying the impact of wages on individual workers' separation decisions. Without correcting for downsizing and closing plants, however, part of the measured effect of wages on separations and hirings may be demand-driven rather than a supply-side response and may for this reason not allow us to infer the firm-level labour supply elasticities from separation rate elasticities. To alleviate this problem, we will exclude job spells in downsizing as well as closing plants.

Our data allow us to identify immigrants only on the basis of citizenship. Due to the *jus sanguinis* tradition of the German law, naturalisation rates are traditionally very low, so that second-generation immigrants are still likely to possess foreign citizenship. To mitigate possible effects of naturalisation, we follow Brücker and Jahn (2011) and classify all individuals as immigrants who are reported as foreign citizens in their first observation available. Another important immigrant group, which possesses German citizenship, consists of so-called ethnic Germans or *(Spät-)Aussiedler* (for details on ethnic German immigrants, see Hirsch *et al.*, 2013). Since ethnic Germans' labour market performance in general resembles that of other immigrants (cf. Glitz, 2012*a*; Hirsch *et al.*, 2013), we again follow Brücker and Jahn (2011) and classify them as immigrants.

Although our data contain observations for East German workers from 1992 onwards, restricting our analysis to the post-unification period would markedly reduce our period of observation and thus the scope of our investigation. It would also add only few observations for immigrants as the wage and salary immigrant population in East Germany is small. We will thus focus our analysis throughout on individuals working in West Germany (excluding Berlin) during the period 1985–2008 and further restrict it to males aged 18–55 years.

The merged data set allows us to build up an inflow sample of job spells starting between 1985 and 2008 taking into account workers' previous labour market status, the job length, and – provided the job ended during our period of observation – workers' subsequent labour market status. In the following, we follow our theoretical model and distinguish two labour market status: employment and non-employment. Consequently, a job may end with a transition to employment, which refers to a new job with another employer (i.e. a plant with a different plant identifyer), or with a transition to non-employment, which refers to a subsequent spell in registered unemployment or no spell in the data at all.[7] The latter either implies that the individual has changed to non-employment without receiving unemployment benefits or that he has become, for instance, a self-employed not included in the data set. While our data do not enable us to disaggregate this category of unknown destination, information from other German data sets suggests that the vast majority of employees in this category have indeed moved to non-employment.[8]

Whereas information on job spells and daily gross wages included in the data are highly reliable, the data include no detailed information on the number of hours worked, and wages are also top-coded at the social security contribution ceiling, which affects 10.7 per cent of our observations. To deal with the first drawback, we restrict our analysis to full-time workers for whom daily wages are comparable. To cope with the second, we impute wages

[7] Note that separations are ignored if the employee is recalled by the same plant within three months.

[8] See, for example, Bartelheimer and Wieck (2005) for a transition matrix between employment and non-employment based on the German Socio-Economic Panel that allows stratification of the "unknown" category into detailed categories.

above the ceiling using a heteroscedastic single imputation approach developed by Büttner and Rässler (2008) for this data set.[9] Besides, information on workers' education is provided by employers. Consequently, this information is missing for 5.7 per cent of all observations. To alleviate this problem, we impute the missing information on education by employing a procedure proposed by Fitzenberger *et al.* (2006) that allows inconsistent education information to be corrected. After applying this imputation procedure, only about 1.3 per cent of the job spells are dropped due to missing or inconsistent information on education.

5 Descriptive Evidence

Our final sample consists of 712,466 job spells belonging to 256,373 native workers and 148,013 spells belonging to 57,406 immigrant workers. For descriptive statistics on key variables, see the Appendix Table. As can be seen from Table 1, which gives an overview of our sample of job spells and the transitions between employment and non-employment, the majority of workers (62.2 per cent of natives and 58.7 per cent of immigrants) have at least two job spells in the data, and out of these roughly two thirds have changing previous labour market status. This enables us to implement our estimation strategy relying on stratified Cox and conditional logit models controlling for unobserved time-invariant worker heterogeneity.

- TABLE 1 ABOUT HERE -

Table 1 also makes clear that natives and immigrants show a rather different transition behaviour both with respect to hirings and to separations: First, natives are less often hired from non-employment with a share of hires from non-employment of 52.9 per cent than

[9] Note that including imputed wages as covariates in the stratified Cox and conditional logit models may introduce some bias in estimated wage elasticities. As a check of robustness, we repeated all estimations excluding job spells with top-coded wages, which gave very similar results.

immigrants whose share amounts to 65.4 per cent.[10] This can be seen as a first indication that immigrants face more severe search frictions because the share of hires from non-employment can serve as a simple measure of search frictions in the labour market (see Manning, 2003*a*, pp. 44–49). Intuitively, the share reflects the difficulties to climb the job ladder by wage-increasing job shopping activities. Second, immigrants are relatively more likely to exit existing jobs to non-employment than employment (with separation rates to employment and non-employment of 57.0 and 30.9 per cent, respectively) compared to natives (with rates of 45.4 and 39.7 per cent, respectively). Again, this may reflect differences in search frictions as separations to employment are more likely to reflect voluntary job-to-job moves as a means of achieving better-paying jobs and reducing job mismatch.

Turning to wages, we find a large average raw native–immigrant wage differential of 20.4 log points (see the Appendix Table). In a next step, we run some standard wage regressions controlling for several worker, job, and plant characteristics to arrive at an estimate of the unexplained wage differential in our sample. In these estimations, we include six age and two education dummies as socio-economic variables. Job controls are eleven occupation and seven job tenure dummies. As plant controls we further include four plant size dummies, the shares of immigrant, part-time, high-skilled, low-skilled, and female workers in the plant's workforce, the median age of the plant's workers, and 24 sector dummies. All estimations also include year dummies, dummies for the size of the region the firm is located at (i.e. rural, urban, or metropolitan), and the unemployment rate at the municipality level.

Running these regressions, the results of which are shown in Table 2, yields an average unexplained native–immigrant wage differential of 5.8–15.0 log points depending on specification. Notably, the gap drops from 15.0 to 8.2 log points when adding occupation and

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[10] As detailed in Section 3, our estimation approach assumes steady-state conditions with the share of hires from and the share of separations to employment being equal, i.e. $\theta \equiv \theta_R = \theta_S$. As can be seen from Table 1, our data are in line with this assumption: The share of hires from employment is 47.1 per cent for natives and 34.6 per cent for immigrants, and the share of separations to employment (i.e. the share of non-censored jobs with a subsequent transition to employment) is 46.7 per cent and 35.1 per cent, respectively.

tenure controls, so that part of the pay gap is likely to reflect that immigrants work in low-tenure jobs and low-paying occupations (in line with the literature, e.g., Eckstein and Weiss, 2004; Constant and Massey, 2005). When also controlling for plant characteristics the gap is reduced further to 5.8 log points which indicates that immigrants are employed in low-paying plants (cf., e.g., Pendakur and Woodcock, 2010; Barth *et al.*, 2012). In the following analysis, we will consider the range spanned by these latter two figures as a benchmark of the unexplained native–immigrant wage gap.

- TABLE 2 ABOUT HERE -

6 Results

To estimate labour supply elasticities at the firm level for both immigrants and natives, thereby assessing whether monopsonistic discrimination may provide an explanation of the native–immigrant wage differential in our sample, we use the econometric approach discussed in Section 3. We estimate stratified Cox models for the instantaneous separation rates to employment and non-employment and conditional logit models for the probability that a worker is hired from employment separately for natives and immigrants. Including the job's log wage in these estimations we arrive at estimates of the separation rate elasticities to employment and non-employment as well as the wage elasticity of the share of hires from employment. We can then plug these estimates into equation (6) to arrive at an estimate of the firm-level labour supply elasticity and use equation (10) to calculate the implied native–immigrant pay gap under monopsonistic wage setting.

With the exception of the job tenure dummies (tenure is controlled for by the non-parametric worker-specific baseline hazard in the stratified Cox models), we include in all estimations the same covariates (listed in the notes to Table 3) as in the wage regressions and treat them as time-varying. In particular, we include 11 occupation dummies. Controlling for occupations is important due to occupational downgrading of immigrants likely to be negatively correlated with wages but positively with the separation probability. Omitting occupation controls in the estimations would therefore inflate the wage effect in the separation equations and thus bias the estimated separation elasticity upwards (in absolute value). Furthermore, including variables capturing the plant's workforce composition, like the share of immigrant or high-skilled and low-skilled workers, is important due to possible workplace segregation of immigrants.

As can be seen from Table 3, which reports the estimated parameters required to arrive at the firm-level labour supply elasticity using equation (6), the separation rate elasticities to employment and non-employment as well as the wage elasticity of the share of hires from employment are lower (in absolute value) for immigrants than for natives. As a consequence, the resulting firm-level labour supply elasticity is significantly smaller ($p = 0.018$) for immigrants, the estimates being 1.136 for immigrants and 1.360 for natives.

- TABLE 3 ABOUT HERE -

Applying equation (10) to our results we find that the estimated elasticity differential implies a *ceteris paribus* earnings disadvantage for immigrants of 7.7 log points. Interestingly, this number is very close to the unexplained immigrant–native wage differential in our sample amounting to 5.8–8.2 log points depending on the set of covariates included (see models 2 and 3 in Table 2). Our reading of this result is that almost the entire unexplained differential can be accounted for by monopsonistic discrimination with the less elastic group of immigrant workers receiving lower wages than comparable native workers.

To further scrutinise the robustness of our findings and to shed some light on the likely sources of employers' more pronounced monopsony power over immigrants, we will now repeat our analysis for subgroups of immigrants (i) differing in their time spent in the host country, (ii) belonging to different immigration cohorts, and (iii) migrating at different age. Unfortunately, our data set does not include the date of immigration to Germany. To overcome this missing information, we proxy the date of entry by the date at which immigrants are first registered as either employed or unemployed in our source data that go back until 1975. We are aware that this proxy is rather crude and disregards periods of unregistered non-employment after immigration to Germany, though recent research by Hirsch *et al.* (2013) indicates that periods of unregistered non-employment while migrating to Germany are typically rather short, i.e. only slightly larger than a year on average. For this reason, we restrained ourselves from including immigrants' proxied years since migration, age at entry, and immigration cohort in the standard wage regressions presented in Table 2 and shall only use this information in a back-of-the-envelope manner in the following.

Table 4 presents estimates of the unexplained immigrant–native wage differential for the different subgroups of immigrants, that is the coefficients of the immigrant dummy following from running wage regressions analogous to those from Table 2 for the respective subgroup. First of all, we find that immigrants who have spent at least ten years in the German labour market experience considerably lower unexplained pay gaps relative to natives ranging from 4.6–5.2 log points (models 2 and 3 in Table 4) than those with less experience whose average gap amounts to 6.6–10.5 log points. This is in line with the widely documented finding of partial wage assimilation over time spent in the host country (see, e.g., Dustmann and Glitz, 2011). Following Chiswick (1978), the standard channel discussed in the literature is that during their stay in the host country immigrants acquire missing country-specific skills augmenting their productivity and being reflected in relative wage increases. From a monopsonistic perspective, gaining in country-specific skills is at the same time also likely to mitigate the search frictions faced by immigrants. Therefore, immigrants with more host-country work experience are expected to suffer less from monopsonistic discrimination. Indeed, Table 5 makes clear that immigrants with ten or more years of German working experience supply labour almost as elastically as natives with an insignificant implied pay gap of just 0.8 log points. On the other hand, the supply elasticity is only 1.095 for immigrants entering the German labour market more recently giving rise to a significant implied pay gap of 9.3 log points, which is well in the range of the unexplained differential for this group.

- TABLES 4 AND 5 ABOUT HERE -

In a second step, we distinguish immigrants entering the German labour market before 1990 from those arriving in later years. Following the fall of the Iron Curtain and Germany's reunification in 1990, the West German labour market saw a large inflow both of immigrants coming from Eastern European countries and of East Germans leaving the poorly performing East German labour market. Since this is likely to hamper immigrants' labour market perspectives and search efforts, we expect immigrants arriving in the pre-1990 period to outperform those arriving later in their wage position relative to natives and to show a larger supply elasticity reflecting their better initial position in job search. As Tables 4 and 5 make clear, our data support both these expectations. The unexplained wage disadvantage of immigrants is 2.7–2.8 log points for the pre-1990 cohort but 7.6–11.6 log points for the 1990–2008 cohort. The earlier cohort also shows a larger supply elasticity of 1.246 compared to 1.075 for the later. As a consequence, the implied wage gap is just 2.7 log points for immigrants arriving before 1990 but 10.1 log points for immigrants arriving later, both numbers being within the respective range of estimated unexplained immigrant–native wage gaps.

Finally, we distinguish immigrants entering the German labour market aged 30 years or less from immigrants arriving at older age. Immigrants arriving aged older than 30 years should not only have acquired their education back in the source country but also should have spent part of their employment careers outside of Germany. As a consequence, these immigrants should find it harder to move their way up the wage distribution and should to a higher degree be subject to monopsonistic discrimination. As can be seen from Table 4, the wage disadvantage for immigrants migrating no later than at age 30 amounts to just 3.2–5.2 log points, whereas the wage disadvantage of those migrating later in their lives is 13.1–16.7 log points. In addition, Table 5 makes clear that immigrants moving at young age have a considerably higher elasticity of 1.217 leading to a predicted wage penalty of 4.7 log points than those who move after turning 30 whose elasticity is estimated as 0.881 implying a wage differential of 18.7 log points.

While our findings highlight the robustness of the native–immigrant difference in the elasticity of firm-level labour supply, the heterogeneity of the difference among subgroups of immigrants is also suggestive that search frictions are a likely candidate for explaining why employers possess more monopsony power over their immigrant workers. As we argued in detail above, the lower elasticity for immigrants with few work experience in Germany is likely to reflect immigrants' initial lack of country-specific skills that narrows over the time spent in the host country. Moreover, the higher elasticity for immigrants entering the German labour market at young age is likely to reflect their less pronounced problems in acculturating to the new environment making them more competitive than older immigrants. Finally, the lower elasticity for immigrants arriving in the 1990–2008 period is likely to mirror that these immigrants entered a more crowded labour market giving rise to a worse initial position in job search. On top of that, immigrants' job search activities seem to be further impeded by a lower degree of regional mobility. We find that the average distance between subsequent jobs is 75 kilometres for natives (or 48 minutes commuting time) but just 53 kilometres (or 36 minutes commuting time) for immigrants, which is in line with earlier findings by Jaeger *et al.* (2010).

7 Conclusions

In this paper, we have estimated the firm-level labour supply elasticity for male natives and immigrants in Germany within a semi-structural estimation approach based on a dynamic monopsony model. Addressing unobserved worker heterogeneity by stratified partial likelihood and conditional logit estimators, we find that immigrants supply labour less elastically to the firm compared to natives. Estimated supply elasticities are about 1.36 for natives but only 1.14 for immigrants. Under a simple rule of monopsonistic wage setting this elasticity differential would imply 7.7 log points lower wages for immigrants, *ceteris paribus*. Notably, the unexplained immigrant–native wage differential in our sample amounts to 5.8–8.2 log points (depending on specification) and therefore is very close to the differential arising under monopsonistic discrimination. Put differently, almost the entire observed differential would be due to monopsonistic discrimination if employers were to exploit their different monopsony power over their native and immigrant workers.

When distinguishing immigrant groups differing in their time spent in the German labour market, their immigration cohort, and their age at entry, we find that the unexplained wage differential is larger for those groups who show a larger elasticity differential relative to natives. Immigrants with more German labour market experience, belonging to the pre-1990 cohort, or entering the labour market at young age not only supply their labour more elastically at the level of the firm than those with less experience, migrating in the aftermath of the downfall of the Iron Curtain and Germany's reunification, or entering at old age, respectively, but also experience considerably lower wage penalties. Remarkably, the predicted wage gaps under monopsonistic wage setting are very close to the observed gaps for all subgroups. These findings not only strengthen the case for monopsonistic wage discrimination against immigrants but also are suggestive that search frictions are the likely cause of employers' more pronounced monopsony power over their immigrant workers. Moreover, they complement the extant literature that stresses the productivity effect of gains in host country-specific skills by pointing at a new channel of immigrants' partial wage assimilation.

Following Robinson (1933, p. 224), who argued that "just as we have price discrimination for a monopolist, so we may have price discrimination for a monopsonist," we therefore conclude that differences in firm-level labour supply elasticities between native and immigrant workers can account for a considerable part of the observed unexplained wage differential. Other than Becker's (1971) taste-based discrimination, this sort of monopsonistic discrimination is profit-increasing and thus fostered by market forces in the long run. What is more, monopsonistic discrimination does not force us to introduce (arbitrary) assumptions on individuals' preferences. As Stigler and Becker (1971, p. 89) argue, "no significant behavior has been illuminated by assumptions of differences in tastes. Instead, they ... have been a convenient crutch to lean on when the analysis has bogged down. They give the appearance of considered judgement, yet really have only been *ad hoc* arguments that disguise analytical failures." For this reason, monopsonistic discrimination does not only seem to be a more coherent theoretical framework to think about immigrant wage discrimination, but also offers straightforward policy implications how to fight it. Under monopsonistic discrimination, employers discriminate against immigrant workers because it is profitable to do so. Stopping them from doing so would "just" require the policy-maker to remove the sources of immigrants' lower firm-level labour supply elasticity, for instance by mitigating search frictions through providing better skills on the host country's language, institutions, and labour market to immigrants. On the other hand, removing clichés and racism forming discriminatory preferences against immigrants seems to be a much harder and less obvious job.

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Tables

Table 1: Job spells and transitions

	Natives		Immigrants	
Observations	2,462,340		465,940	
Job spells	712,466		148,013	
Workers	256,373		57,406	
Plants	296,474		79,632	
Average job spells per worker	2.8		2.6	
Percentage of workers with more than one job spell	62.2		58.7	
Percentage of multiple-jobs workers with both employment and non-employment as previous labour market status	64.6		61.5	
Hires from employment	335,570	(47.1)	51,158	(34.6)
Hires from non-employment	376,896	(52.9)	96,855	(65.4)
Separations to employment	282,722	(39.7)	45,739	(30.9)
Separations to non-employment	323,139	(45.4)	84,441	(57.0)
Right-censored job spells	106,605	(15.0)	17,833	(12.0)

Notes: The data sets used are the SIAB, 1985–2008, and the BHP, waves 1985–2008. Percentages are shown in parentheses.

Table 2: Wage regressions

	Model 1		Model 2		Model 3	
Immigrant (dummy)	–0.150	(0.002)	–0.082	(0.002)	–0.058	(0.001)
Medium-skilled (dummy)	0.162	(0.002)	0.119	(0.002)	0.087	(0.001)
High-skilled (dummy)	0.569	(0.003)	0.341	(0.003)	0.258	(0.003)
Age 18–25 years (dummy)	–0.168	(0.001)	–0.105	(0.001)	–0.092	(0.001)
Age 31–35 years (dummy)	0.112	(0.001)	0.079	(0.001)	0.073	(0.001)
Age 36–40 years (dummy)	0.171	(0.001)	0.124	(0.001)	0.113	(0.001)
Age 40–45 years (dummy)	0.193	(0.002)	0.144	(0.001)	0.131	(0.001)
Age 46–50 years (dummy)	0.187	(0.002)	0.146	(0.002)	0.129	(0.002)
Age 51–55 years (dummy)	0.157	(0.003)	0.132	(0.002)	0.114	(0.002)
Tenure 1–4 years (dummy)			0.170	(0.001)	0.116	(0.001)
Tenure 5–9 years (dummy)			0.254	(0.001)	0.168	(0.001)
Tenure 10–14 years (dummy)			0.303	(0.002)	0.198	(0.001)
Tenure 15–19 years (dummy)			0.348	(0.003)	0.222	(0.002)
Tenure ≥ 20 years (dummy)			0.404	(0.005)	0.254	(0.004)
Plant size 11–50 (dummy)					0.074	(0.001)
Plant size 51–200 (dummy)					0.105	(0.002)
Plant size 201–1,000 (dummy)					0.163	(0.002)
Plant size > 1,000 (dummy)					0.210	(0.002)
Share of immigrant workers					–0.094	(0.004)
Share of female workers					–0.126	(0.003)
Share of high-skilled workers					0.280	(0.004)
Share of low-skilled workers					–0.122	(0.002)
Share of part-time workers					–0.048	(0.004)
Median age of workers at the plant					0.005	(0.000)
One-digit occupation			√		√	
Two-digit industry					√	
Observations	2,928,280		2,928,280		2,928,280	
R^2	0.280		0.422		0.526	

Notes: The data sets used are the SIAB, 1985–2008, and the BHP, waves 1985–2008. The regressand is the daily gross log wage. Standard errors clustered at the individual level are given in parentheses. All models include year dummies, two dummies indicating the size of the regional labour market, and the unemployment rate at the municipality level. In addition, models 2 and 3 contain 11 occupation dummies and model 3 also 24 sector dummies.

Table 3: Estimated firm-level labour supply elasticities and the implied immigrant–native pay gap

	Stratified Cox and conditional logit models controlling for unobserved worker heterogeneity	
	Natives	Immigrants
Separation rate elasticity to employment ($\hat{\varepsilon}_{sw}^{e}$)	–1.039 (0.014)	–0.956 (0.035)
Separation rate elasticity to non-employment ($\hat{\varepsilon}_{sw}^{n}$)	–0.793 (0.010)	–0.686 (0.022)
Wage elasticity of the share of recruits hired from employment ($\widehat{\varepsilon_{\theta w}/(1-\theta)}$)	1.110 (0.013)	0.916 (0.030)
Share of hires from employment ($\hat{\theta}$)	0.471	0.346
Firm-level labour supply elasticity ($\hat{\varepsilon}_{Lw}$)	1.360 (0.034)	1.136 (0.086)
Test for equality of firm-level labour supply elasticities between immigrants and natives	p = 0.018	
Implied immigrant–native pay gap ($\widehat{\ln w^{I} - \ln w^{N}}$)	–0.077	

Notes: The data sets used are the SIAB, 1985–2008, and the BHP, waves 1985–2008. Standard errors are shown in parentheses. The standard errors for the firm-level labour supply elasticities are bootstrapped with 200 replications. Covariates included in the estimations are two education, six age, four plant size, 11 occupation, 24 sector, 23 year, and two dummies for the size of the region the plant is located at, the shares of part-time, high-skilled, low-skilled, female, and immigrant workers in the plant's workforce, the median age of its workforce, and the unemployment rate at the municipality level. Detailed results are available on request.

Table 4: Heterogeneity in the immigrant wage gap

	Model 1		Model 2		Model 3	
All immigrants (as in Table 2)	–0.150	(0.002)	–0.082	(0.002)	–0.058	(0.001)
Immigrants with years since migration < 10	–0.186	(0.002)	–0.105	(0.002)	–0.066	(0.001)
Immigrants with years since migration ≥ 10	–0.101	(0.003)	–0.052	(0.002)	–0.046	(0.002)
Immigrants from pre-1990 cohort	–0.066	(0.003)	–0.028	(0.002)	–0.027	(0.002)
Immigrants from 1990–2008 cohort	–0.200	(0.002)	–0.116	(0.002)	–0.076	(0.002)
Immigrants aged ≤ 30 years at entry	–0.106	(0.002)	–0.052	(0.002)	–0.032	(0.001)
Immigrants aged > 30 years at entry	–0.268	(0.003)	–0.167	(0.003)	–0.131	(0.002)

Notes: The data sets used are the SIAB, 1985–2008, and the BHP, waves 1985–2008. The reported parameter is the coefficient of the immigrant dummy following from wage regressions for the respective subgroup of immigrants, where models are specified as in Table 2. Standard errors clustered at the individual level are given in parentheses. Detailed results are available on request.

Table 5: Heterogeneity in immigrants' estimated firm-level labour supply elasticity and the implied immigrant–native wage gap

	Stratified Cox and conditional logit models controlling for unobserved worker heterogeneity						
	Natives	Immigrants with years since migration < 10	Immigrants with years since migration ≥ 10	Immigrants from pre-1990 cohort	Immigrants from 1990–2008 cohort	Immigrants aged ≤ 30 years at entry	Immigrants aged > 30 years at entry
Separation rate elasticity to employment ($\hat{\varepsilon}_{sw}^{e}$)	–1.039 (0.014)	–0.928 (0.046)	–1.066 (0.077)	–0.986 (0.054)	–0.925 (0.046)	–0.960 (0.038)	–0.952 (0.086)
Separation rate elasticity to non-employment ($\hat{\varepsilon}_{sw}^{n}$)	–0.793 (0.010)	–0.587 (0.029)	–0.645 (0.044)	–0.738 (0.036)	–0.657 (0.027)	–0.745 (0.024)	–0.472 (0.046)
Wage elasticity of the share of recruits hired from employment ($\widehat{\varepsilon_{\theta w}/(1-\theta)}$)	1.110 (0.013)	0.777 (0.038)	0.941 (0.065)	0.935 (0.049)	0.857 (0.038)	0.906 (0.033)	0.934 (0.068)
Share of hires from employment ($\hat{\theta}$)	0.471	0.319	0.415	0.412	0.310	0.372	0.277
Firm-level labour supply elasticity ($\hat{\varepsilon}_{Lw}$)	1.360 (0.034)	1.095 (0.108)	1.335 (0.204)	1.246 (0.133)	1.075 (0.110)	1.217 (0.091)	0.881 (0.189)
Test for equality of firm-level labour supply elasticities between immigrants and natives		p = 0.021	p = 0.901	p = 0.546	p = 0.014	p = 0.143	p = 0.012
Implied immigrant–native pay gap ($\widehat{\ln w^{I} - \ln w^{N}}$)		–0.093	–0.008	–0.027	–0.101	–0.047	–0.187

Notes: The data sets used are the SIAB, 1985–2008, and the BHP, waves 1985–2008. Standard errors are shown in parentheses. The standard errors for the firm-level labour supply elasticities are bootstrapped with 200 replications. Covariates included in the estimations are the same as in Table 3. Detailed results are available on request.

Appendix Table: Selected Descriptive Statistics (Means)

	Natives	Immigrants
Log daily gross wage (€)	4.442	4.238
Years since migration (years)		9.217
Years since migration < 10 years (dummy)		0.619
Years since migration ≥ 10 years (dummy)		0.381
Pre-1990 immigration cohort (dummy)		0.617
Immigration cohort 1990–2008 (dummy)		0.383
Age at entry (years)		25.806
Age at entry ≤ 30 years (dummy)		0.734
Age at entry > 30 years (dummy)		0.266
Low-skilled (dummy)	0.084	0.353
Medium-skilled (dummy)	0.800	0.598
High-skilled (dummy)	0.117	0.049
Age (years)	34.431	35.023
Age 18–25 years (dummy)	0.180	0.168
Age 26–30 years (dummy)	0.195	0.185
Age 31–35 years (dummy)	0.193	0.190
Age 36–40 years (dummy)	0.170	0.166
Age 40–45 years (dummy)	0.130	0.139
Age 46–50 years (dummy)	0.088	0.103
Age 51–55 years (dummy)	0.044	0.051
Tenure (years)	2.649	2.180
Tenure < 1 year (dummy)	0.526	0.592
Tenure 1–4 years (dummy)	0.236	0.213
Tenure 5–9 years (dummy)	0.155	0.132
Tenure 10–14 years (dummy)	0.057	0.046
Tenure 15–19 years (dummy)	0.022	0.016
Tenure ≥ 20 years (dummy)	0.004	0.002
Plant size	1043.851	950.498
Plant size ≤ 10 (dummy)	0.118	0.108
Plant size 11–50 (dummy)	0.263	0.250
Plant size 51–200 (dummy)	0.258	0.284
Plant size 201–1,000 (dummy)	0.222	0.227
Plant size > 1,000 (dummy)	0.139	0.131
Share of immigrant workers in workforce	0.079	0.208
Share of female workers in workforce	0.278	0.246
Share of high-skilled workers in workforce	0.078	0.054
Share of low-skilled workers in workforce	0.209	0.316
Share of part-time workers in workforce	0.115	0.100
Median age of workers at the plant	37.299	37.065
Regional unemployment rate	8.656	8.378
Observations	2,462,340	465,940
Job spells	712,466	148,013
Workers	256,373	57,406
Plants	296,474	79,632

Notes: The data sets used are the SIAB, 1985–2008, and the BHP, waves 1985–2008.