

## THE ANATOMY OF SORTING—EVIDENCE FROM DANISH DATA

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In this paper, we formulate and estimate a flexible model of job mobility and wages with two-sided heterogeneity. The analysis extends the finite mixture approach of Bonhomme, Lamadon, and Manresa (2019) and Abowd, McKinney, and Schmutte (2019) to develop a new Classification Expectation-Maximization algorithm that ensures both worker and firm latent-type identification using wage and mobility variations in the data. Workers receive job offers in worker-type segmented labor markets. Offers are accepted according to a logit form that compares the value of the current job with that of the new job. In combination with flexibly estimated layoff and job finding rates, the analysis quantifies the four different sources of sorting: job preferences, segmentation, layoffs, and job finding. Job preferences are identified through job-to-job moves in a revealed preference argument. They are in the model structurally independent of the identified job wages, possibly as a reflection of the presence of amenities. We find evidence of a strong pecuniary motive in job preferences. While the correlation between preferences and current job wages is positive, the net present value of the future earnings stream given the current job correlates much more strongly with preferences for it. This is more so for short- than long-tenure workers. In the analysis, we distinguish between type sorting and wage sorting. Type sorting is quantified by means of the mutual information index. Wage sorting is captured through correlation between identified wage types. While layoffs are less important than the other channels, we find all channels to contribute substantially to sorting. As workers age, job arrival processes are the key determinant of wage sorting, whereas the role of job preferences dictate type sorting. Over the life cycle, job preferences intensify, type sorting increases, and pecuniary considerations wane.

**KEYWORDS:** Heterogeneity, wage distributions, employment and job mobility, mutual information, finite mixtures, EM algorithm, classification algorithm, sorting, decomposition of wage inequality.

### 1. INTRODUCTION

TO WHAT EXTENT ARE WORKERS AND FIRMS SORTED IN THE LABOR MARKET? What drives sorting and how does it change over a worker's life? The answers to these ques-

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tions are central to understanding sources of wage inequality and designing labor market policies. To this end, this paper estimates a flexible semiparametric model of wages and mobility with two-sided unobserved and observed heterogeneity on a long matched employer–employee panel data set. The broad-ranging interactions between tenure, experience, and heterogeneity in wages and employment mobility that our model allows for reveal considerable complexity in matching.

We propose a structural job mobility model where workers stochastically move in the direction of higher job preferences, which we identify from job transitions separately from labor earnings. Workers meet firms in labor markets segmented by worker type where both meeting rates and the distribution of vacancies by firm type are segment-specific. In combination with flexibly estimated worker transitions in and out of unemployment, the mobility model allows us to quantify the impact of four different mobility channels on sorting: job preferences, segmentation, layoffs, and job finding. We do so through counterfactuals on the estimated model. While the layoff channel contributes less than the other channels, we find all channels to be significant contributors to sorting.

We find a significant positive correlation between a worker’s valuation of the match and the associated current earnings. But more strikingly, when we calculate the net present value of the expected stream of future earnings given the current match, we find a substantially stronger correlation than that against current earnings. The correlation with net present value is quite strong, in particular at short tenure. Thus, we find convincing evidence of a strong pecuniary component in the variation of worker match valuations across firms. For the sake of emphasis: a worker’s pecuniary compensation varies across firm types and the worker’s revealed preferences show that the worker cares, particularly for short tenure workers. Furthermore, preferences contribute substantially to the sorting patterns we see in the data.

Our paper contributes to the worker-firm sorting literature by offering a rich and flexible framework where sources of sorting can be estimated. The framework allows for wage and nonwage attributes through both worker and firm heterogeneity. One big-picture message is that worker and firm heterogeneity condition mean wages and job preferences differently. Some firms are highly attractive, but they are not necessarily the ones paying the highest wages. Workers, on the other hand, need not share the same global ranking of job types. This is a rather different interpretation from what has been recently pushed forward by several authors (Lindenlaub (2017), Lindenlaub and Postel-Vinay (2020), Lise and Postel-Vinay (2020), Lindenlaub and Postel-Vinay (2021)). In their interpretation, workers and jobs have multiple characteristics that interact in the match production function to generate different dimensions of sorting for each pair of worker-firm characteristics. They emphasize how difficult it then becomes to define a simple notion of sorting and to link wages and mobility to properties of the match surplus. However, they only consider observed attributes (O’NET) and use standard worker panel data (SIPP). Additionally, it is worth noting that the structure of our transition probabilities is perfectly consistent with search-matching theory, yet more flexible. Standard search-matching theory makes both mean wages and match preferences monotone functions of the same input, match productivity. But our model allows mean wages and match preferences to be separate parameters, rendered structurally independent possibly by the existence of mobility costs or job amenities.<sup>1</sup> This story echoes Sorokin’s (2018) recent paper, who first observes on US

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<sup>1</sup>In all fairness, Lindenlaub and Postel-Vinay (2021) only consider employment and job transitions probabilities. Their notion of “worker-job surplus” is therefore independent of wages.

match employer–employee data that close to 50% of job-to-job transitions are not associated with an increase in earnings, and whose revealed preference estimation technique reveals that nonpay characteristics explain 2/3 of the variance of firm-level earnings and 15% of the overall earnings variance. Our paper is in a similar vein as Sorkin’s (2018). An important difference is that his technique identifies commonly held firm rankings across workers, whereas the discrete-mixture approach we adopt identifies mean wages and job preferences that are both firm- *and* worker-specific. With this less restrictive setup, we can analyze and understand preference-driven sorting outcomes. We can furthermore distinguish between sorting on wages and sorting on nonwage characteristics.

Since the seminal contribution of Abowd, Kramarz, and Margolis (1999) (hereafter AKM), the literature studying how individual wages vary across workers and firms has focused on estimating, by ordinary least squares, a linear model with additive worker and firm fixed effects, and sometimes a match-specific effect.<sup>2</sup> As with any panel data model with limited mobility, the OLS estimator of the AKM fixed-effect model is prone to overfitting. In response, finite-sample bias corrections have been proposed (Andrews, Gill, Schank, and Upward (2008), Kline, Saggio, and Solvsten (2020), Azkarate-Askasua and Zerecero (2019)). While they have provided a much needed correction to the bias in the framework’s measurement of wage sorting, the AKM framework does not allow us to study the dynamics of sorting through workers’ transitions between employment and unemployment, nor does it incorporate sorting that may arise from nonwage factors. To advance our understanding on this issue, we therefore depart from the usual fixed-effect estimation by adopting the recent approach of Bonhomme, Lamadon, and Manresa (2019) (hereafter BLM) and Abowd, McKinney, and Schmutte (2019), who regularize estimation by assuming a finite number of worker and firm types.<sup>3</sup>

Our own model is more closely related to BLM in that we impose no restrictions on how wages depend on worker and firm latent types, and we use the Expectation-Maximization (EM) algorithm for estimation. Yet, we extend BLM’s methods along three dimensions. First, BLM classify firms only on wages using a  $k$ -means algorithm and separately from the estimation of the rest of the model, including latent worker types. Instead, we make firm classification a component of the main estimation procedure through a Classification EM algorithm (CEM) inspired by Celeux and Govaert (1992). Monte Carlo simulations show that our CEM algorithm improves over a one-step  $k$ -means. From the perspective of advancing our understanding of sorting and wage inequality, using mobility data to classify employers is crucial as two firms offering an identical wage but nonidentical employment stability would result in different lifetime earnings for workers. Further, if workers have preferences for job stability then employer types should be classified by mobility patterns in addition to wages. We show that it is indeed the case that jobs with low layoff risk are preferred.

Second, while BLM focus on wage residuals, we include observables of both workers (gender, education, tenure, and experience) and firms (sector and industry) in our estimation. Since unobserved heterogeneity may not be orthogonal to observed heterogeneity, for example, some worker latent types may have faster wage growth over tenure or

<sup>2</sup>See Holzer, Lane, and Vilhuber (2004), Martins (2008), Iranzo, Schivardi, and Tosetti (2008), Gruetter and Lalive (2009), Bagger, Sørensen, and Vejlin (2013), Card, Heining, and Kline (2013), Woodcock (2015), Song, Price, Guvenen, Bloom, and von Wachter (2015), Card, Cardoso, and Kline (2016).

<sup>3</sup>BLM put no shape restrictions on the way wages and job-matching depend on worker and firm heterogeneity, whereas Abowd, McKinney, and Schmutte (2019) maintain the AKM additive structure of worker and firm effects in wage means, while adding a stochastic match effect correlated with worker and firm types, and with mobility.

experience than others, we allow for a flexible relationship between latent types and observed heterogeneity. In other words, we let latent types be the channel through which observed *and* unobserved heterogeneity jointly determine wage and mobility outcomes. Thus, the inclusion of observable heterogeneity contributes to the identification of latent types in our population. Specifically, we assume that wages are log normally distributed with means and variances that depend unrestrictedly on  $(x, k, \ell)$ , where  $x$  is time-varying worker heterogeneity (tenure and experience categories),  $k$  is the latent worker type, and  $\ell$  is the latent firm type. Furthermore, we allow wage residuals to be autocorrelated.

Third, we depart from BLM's totally flexible mobility model by using a parametric model for job-to-job transitions that reduces the number of mobility parameters by a factor of 10. This is good for efficiency, but the main reason for constraining the transition probabilities is intelligibility. The transition probabilities are specified as the product of a job sampling probability and a choice probability. The choice probability is an increasing function of the ratio of two job preferences (one for the incumbent employer and one for the poacher) where the job preference parameter can flexibly vary by the  $(x, k, \ell)$  combination. This allows (i) job ranking to differ across both worker types  $k$  and their time-varying characteristics  $x$ , (ii) job ranking to flexibly differ from wage ranking, and (iii) the interpretation of a standard on-the-job search model with random preferences for job types and worker-specific offer arrival rates. Empirically, the model allows us to quantify the importance of job sampling, or "*chance*," relative to worker preferences for job types, or "*choice*," at different career stages. The nonlinear specification does however pose a threat to estimation feasibility. We develop an MM algorithm (Hunter (2004), Hunter and Lange (2004)) nested inside the M-step of the CEM algorithm to overcome this estimation difficulty.

We focus on Denmark in the period 1987–2013 in our empirical application.<sup>4</sup> First, we find a considerable degree of interaction of worker and firm heterogeneity in mean wages. In Denmark, worker and firm types do not determine wages additively—there is a considerable degree of nonlinearity. Second, broadening our measure of sorting, we employ the concept of mutual information (MI) that measures the distance of the estimated match distribution to independent matching, which can flexibly represent the dependence between worker and firm latent types independently of their wage effects. In support of Bagger, Sørensen, and Vejlin (2013), we find that sorting is moderately increasing over time in Denmark, where some is positive assortative on wages. However, the MI reveals that the increased importance of amenities in sorting patterns as workers age and select into long-tenure relationships—this conclusion would have been the opposite had we used the correlation of wage fixed effects as the sorting measure. Consistently, our estimated job preferences revealed from job-to-job moves also confirm that as workers age and tenure rises, preferences are increasingly shaped by nonwage attributes.

Our mobility model allows us to run counterfactual and cohort simulations that comprehensively uncover the key drivers of type sorting measured by MI as well as the classic wage sorting measured by the wage fixed-effect correlation. We find that a cohort achieves a significant part of its sorting pattern, both in types and wages, early on as it enters the labor market. Thenceforth, during a worker's early career, sorting is primarily driven by complementarities between worker and firm types in the offer arrival processes while employed and unemployed. These two channels are apparent particularly in the wage fixed-effects correlation, and thus seem to drive a classical form of sorting via wage effects. It is

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<sup>4</sup>We divide the observation period into five time windows and estimate the model separately for each window. A worker's and a firm's types are constant within each window.

also important to observe that the mobility patterns out of unemployment, like the initial match draws, are a significant contributor to sorting and not the kind of reset that is often found in standard job search models. However, later in life, workers' subjective rankings of jobs become the dominant determinant of matching. The role of job preferences becomes increasingly important as cohorts age, and it is what gives rise to the positive age trend in type sorting. Moreover, the role of job preferences is more apparent in the MI index than in the wage fixed-effects correlation implying that nonwage attributes play a greater role in match preference determination later in life.<sup>5</sup>

Finally, we show that the various channels of sorting are not independent of each other. Our model explicitly considers four drivers of sorting: job preferences, layoffs, market segmentation, and reemployment. We find evidence of substantial interactions between channels. In particular, for both wage and type sorting, there is a strong substitutability between the job preferences and the market segmentation channels (the impact of one channel is bigger in the absence of the other). The job preferences and layoff channels are complementary in particular for type sorting. Preferences select into longer lasting jobs and layoff variation enforces the sorting. The unemployment related channels of layoff and job finding are also complementary, in particular for type sorting. This reemphasizes the importance of classifying employer and employee types using both wage and mobility data.

The layout of the paper is as follows: Section 2 describes the model and the parametric specifications, Section 3 explains the estimation procedure. Then Sections 4 and 5 present the estimation results, Section 6 analyzes sorting, and Section 7 concludes.

## 2. THE DATA AND THE MODEL

### 2.1. *The Data*

We use the Danish matched employer–employee data from 1987–2013. Wages and employment mobility are reported at annual and weekly frequency, respectively. We divide the large panel into five smaller 5-year panels opening different windows on the business cycle: 1989–93, 1994–98, 1999–2003, 2004–08, 2009–13. The first period was one of high unemployment in Denmark (9.5% in 1993). The next period contains a decline in unemployment (to about 5% in 1998). From 1999 to 2008, unemployment stayed low, and went up again after the financial crisis (7.8% in 2012). For each 5-year panel, we use information from the 2 preceding years to distinguish between short- and long-tenure jobs in the stock of jobs at the beginning of each period. For example, we use information from 1987–1988 for the stock of jobs in 1989.

We restrict the sample to employment spells that start after individuals attain their, in retrospect, highest education level. We remove all spells that start after the individuals turn 50-years-old and treat any spell with a positive wage as an employment spell. Time between two jobs is in our analysis referred to as unemployment and it should be understood as a broad definition that includes events such as extended sick leave, but not education or retirement.

There is only one payroll recorded for each employment spell within a year. For example, if a worker transitions from firm 1 to firm 2 after week 12 in a given year, there would

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<sup>5</sup>The divergence between MI and wage fixed-effect correlation could be a result of both non-wage factors and nonlinearity in wage fixed effects. However, we find (not shown) that the contribution of nonlinearity or match effects on wage variance do not vary much by experience. Hence, the divergence between MI and wage fixed-effect correlation over a workers' career is more likely due to nonwage factors.

be one observation for the first 12 weeks in firm 1 and another for the next 40 weeks in firm 2. If the worker remains in firm 2 for the following year, then we have only one observation for the annual pay in that year. We calculate average weekly earnings (total earnings divided by the number of weeks) instead of hourly wages because the data on hours are noisy.

Although most other studies keep only full-time workers, we keep most workers in the analysis and explicitly model wages and mobility. We do this because workers can move between part-time and full-time employment and because treating part-time jobs as unemployment would overstate the share of transitions out of employment. This strategy is unlike those in studies that use the AKM approach, as the latter typically include only full-time workers, neglecting an explicit model of job mobility. Confounding part-time employment with unemployment appears to us more inaccurate than confounding it with full-time employment. Unfortunately, we thus introduce a source of variation in our definition of earnings that is usually absent. Weekly earnings vary either because hourly wages vary or because weekly hours vary.

Workers are indexed by  $i \in \{1, \dots, I\}$  and firms by  $j \in \{0, 1, \dots, J\}$ , where  $j = 0$  marks unemployment. Each worker  $i$  is either drawn from the working-age population in the first year of the panel window, or enters the panel in the first week of the first year following his or her last year of schooling. Individual trajectories  $(w_{it}, j_{it}, x_{it})_{t=1}^T$  are recorded at a weekly frequency, where  $j_{it} := j(i, t) \in \{0, 1, \dots, J\}$  is the employer's ID in week  $t$ ,  $x_{it}$  are time-varying controls, and  $w_{it}$  is the log of the worker's weekly earnings at occurrence  $t$ . Note that although the number of repeated observations  $T$  varies across individuals, we adopt the simplified notation of a balanced panel.

The worker's time-varying characteristics  $x_{it}$  include the short-/long-tenure status and potential experience (time since graduation).<sup>6</sup> Short tenure in a job is defined to be less than 100 weeks of employment (or 2 years). The model also makes a distinction between short- and long-term unemployment and we maintain the language of tenure as a shorthand for duration in the unemployed state also. For unemployment, short tenure is defined to be less than 26 weeks (6 months). We divide experience into four groups: less than 5 years, 5–10 years, 11–15 years, and more than 15 years. Table I displays the distribution of workers' tenure and experience. As expected, most younger workers have short tenure. Notably, older workers are nearly equally split between holding short- and long-tenure jobs. This is important to note as one could think that more experienced workers would be keeping the same job until retirement, a feature that might be expected in less mobile labor markets than the Danish one.

For each worker  $i$ , we observe a set of time-invariant characteristics  $z_i^w$  that include gender and education. Education level is based on the normed number of years of education associated with the worker's highest completed degree. The low education group comprises all degrees normed to less than 12 years of education. The medium education group has a norm of exactly 12 years, and the high education group is any education level with a norm greater than 12 years.

For each firm  $j$ , we observe a set of time-invariant firm characteristics  $z_j^f$  that include the public/mixed/private status and the industry of the firm.<sup>7</sup>

<sup>6</sup>We can calculate actual experience only for workers entering the labor market after 1987. This is why we use potential experience or age.

<sup>7</sup>We produced our own aggregation of the available industry code, but after much work we have not been able to obtain stable and sensible results across panels. We will thus not make a comparison in this aspect across time.

2.2. The Model

We assume that employers (firms) can be clustered into  $L$  latent groups indexed by  $\ell \in \{1, \dots, L\}$  and that workers can be clustered into  $K$  latent groups indexed by  $k \in \{1, \dots, K\}$ . The index  $\ell_j$  is the type of firm  $j$  and  $k_i$  is the type of worker  $i$ . Unemployment is a specific employment type and is denoted by  $\ell = 0$ . Worker and firm type assignments are assumed to be fixed over the duration of the 5-year panel.

We commence by stating the sample likelihood for a given firm classification  $F = (\ell_1, \dots, \ell_J)$  assigning a type  $\ell_j$  to each firm  $j$  in the sample. We denote

$$q(\ell|F) = \frac{\#\{j : \ell_j = \ell\}}{J}$$

as the share of type- $\ell$  firms given a firm classification. For notational simplicity, we write  $\ell_{it} := \ell_{j(i,t)}$  and  $z_{it}^f := z_{j(i,t)}^f$  for the type and the observed characteristics of the firm employing worker  $i$  in period  $t$ . We also let

$$D_{it} = \begin{cases} 1 & \text{if } j_{i,t+1} \neq j_{it}, \\ 0 & \text{if } j_{i,t+1} = j_{it}, \end{cases}$$

indicate an employer change between  $t$  and  $t + 1$ .

For a value  $\beta$  of the parameters (which will be discussed shortly in Section 2.4) and a classification  $F$  of firms, the complete likelihood for one worker  $i$  conditional on the first observation of tenure and experience  $x_{i1}$  is

$$\begin{aligned} \mathcal{L}_i(\beta|k, F) &= \frac{m_0(k, \ell_{i1}|x_{i1})\pi^w(z_i^w|k)\pi^f(z_i^f|\ell_{i1})}{q(\ell_{i1}|F)} f_{\text{static}}(w_{i1}|k, \ell_{i1}, x_{i1}) \\ &\times \prod_{t=1}^{T-1} M(-|k, \ell_{it}, x_{it})^{1-D_{it}} \left( \frac{M(\ell_{i,t+1}|k, \ell_{it}, x_{it})}{q(\ell_{i,t+1}|F)} \right)^{D_{it}} \\ &\times \prod_{t=1}^{T-1} f_{\text{dyn}}(w_{i,t+1}|k, \ell_{i,t+1}, x_{i,t+1}, w_{it}, x_{it})^{1-D_{it}} \\ &\times f_{\text{static}}(w_{i,t+1}|k, \ell_{i,t+1}, x_{i,t+1})^{D_{it}}. \end{aligned} \tag{1}$$

The structure of this likelihood summarizes all of the model’s assumptions and notation, as we now explain.

TABLE I  
SHARE OF EMPLOYMENT BY EXPERIENCE AND TENURE.

Experience (Yrs.):	Short Tenure				Long Tenure			
	0–5	5–10	10–15	15+	0–5	5–10	10–15	15+
1989–93	0.14	0.11	0.08	0.19	0.03	0.08	0.08	0.28
1994–98	0.14	0.12	0.09	0.19	0.02	0.08	0.09	0.27
1999–03	0.16	0.13	0.11	0.23	0.02	0.06	0.07	0.22
2004–08	0.18	0.14	0.11	0.24	0.02	0.06	0.06	0.19
2009–13	0.23	0.11	0.08	0.17	0.03	0.08	0.08	0.21

*Initial Condition.* A worker enters the panel with experience and tenure  $x_{i1}$ , which is conditioned upon. Initial experience and tenure determine a particular distribution of initial matches  $m_0(k, \ell_{i1}|x_{i1})$ . This initial dependence reflects the endogeneity of tenure and matching for the trajectories of workers who are initially drawn from the stock.

Among workers of type  $k$ , there is a particular distribution of gender and education  $\pi^w(z_i^w|k)$ . For simplicity, we assume that gender and education are independent of job tenure and the matching firm, given the worker's type. This conditional independence assumption is in principle innocuous if the number of unobserved types  $K$  is large enough. Similarly, we also assume that the distribution of observed firm characteristics  $\pi^f(z_i^f|\ell_{i1})$  is only a function of firm types. This is exactly like incorporating into the likelihood the usual ex post tabulations of observed individual characteristics by predicted latent type.

Lastly, the initial employer's ID  $j_{i1}$  is drawn given its type with a probability proportional to the relative frequency of each type in the population of firms,  $1/q(\ell_{i1}|F)$ . We thus assume that each firm within a group is equally likely to be selected. Given that inflows to firms are proportional to firm size in our data, this implies in particular that firm sizes within a group should be similar, as we expect large firms to have a higher sampling probability than small firms. Our firm classification algorithm is based on this likelihood. The estimated firm classification will thus be performed subject to the uniform-sampling assumption. Again, there is no loss of generality because, in principle, the algorithm can choose to group firms of different sizes in different groups if size matters for sampling. This is a first difference to be emphasized with BLM. In BLM, the firm classification is operated ex ante and the sampling mechanism of firms given types can be neglected as it contains no parameter left to be estimated. On the contrary, our algorithm will eventually classify firms based on the same likelihood as the workers'. Firm types  $\ell_j$  and the corresponding sampling probabilities  $1/q(\ell_j|F)$  are therefore required in the construction of the likelihood.

*Job Mobility.* In every period, the worker changes employment status (employed or unemployed) and employer type with probability  $M(\ell_{i,t+1}|k, \ell_{it}, x_{it})$ . This probability is conditional on worker type  $k$ , employer type  $\ell_{it}$ , tenure and experience  $x_{it}$ . We do not specifically condition on observed worker and firm characteristics. Since this is a clustering model, the group ( $k$  or  $\ell$ ) is a sufficient statistic for all socioeconomic determination. For instance, if gender per se were to determine mobility and wages, then the proportion of females would vary by group and show in probabilities  $\pi^w(z_i^w|k)$ . The specific employer  $j_{i,t+1}$  is drawn with probability  $1/q(\ell_{i,t+1}|F)$ . The worker stays with the same employer with probability

$$M(-|k, \ell_{it}, x_{it}) = 1 - \sum_{\ell'=0}^L M(\ell'|k, \ell_{it}, x_{it}).$$

Finally, we assume that we do not know whether a mobility occurs for the last observation period.

*Wages.* The initial wage  $w_{i1}$  is drawn from a static distribution  $f_{\text{static}}(w_{i1}|k, \ell_{i1}, x_{i1})$ . We also model the wage distribution after a job-to-job transition or out of non-employment using the same static distribution. However, we allow for autocorrelation in the case of stayers. This assumption is motivated by tests of autocorrelation of wages within and between spells in the Danish data. After estimating a model without wage autocorrelation, we observed that the residual autocorrelation within job was substantial, while it



was much lower between jobs. Moreover, we could make the between-job distribution of wages a function of both origin and destination employer types, as predicted by the sequential auction model (Postel-Vinay and Robin (2002)). However, in doing so, we lose the connection to the AKM framework. Finally, note that removing state dependence from job-to-job transitions greatly simplifies the identification proof. For these reasons, we thus leave it to further research to investigate such extensions.

Due to the annual frequency of observations on job payrolls, we incorporate just one wage likelihood contribution for each job spell in a year. For example, in

$$f_{\text{dyn}}(w_{i,t+1} | k, \ell_{i,t+1}, x_{i,t+1}, w_{it}, x_{it}),$$

$w_{i,t+1}$  refers to the unique wage observation for the first week of the current spell-year and  $w_{it}$  refers to the unique wage observation for the first week of the preceding spell-year. For ease of notation, we use the same time index  $t$ . Firm types are static within a given estimation window. Thus, within a job spell it follows that  $\ell_{i,t} = \ell_{i,t+1}$ .

### 2.3. Identification

BLM extend results in Hu and Schennach (2008) and Hu and Shum (2012) and show that at least four wage observations are sufficient to identify mixtures of dynamic models for matched employer–employee data. In Appendix A, we adapt and simplify BLM’s proof in the context of our specific assumptions. Because each employer change resets the wage dynamic process, three wage observations, one before the change and two after, are sufficient for identification. We use 5-year panels in estimation. So, in theory, we should be on the safe side.

The proof uses the same assumptions as BLM’s. First, there must be enough observed differences across firms—statistics of wage distributions, firm size, entry and exit flows, etc.—to identify the firm classification irrespective of specific worker trajectories. Second, all transitions are realized with positive probability. For example, all workers can go from a type-1 firm to a type-2 firm and move back. It is possible to weaken this assumption, like BLM do with the “alternating cycles,” but the identification argument is essentially the same. We also assume that the transition probabilities, and ratios of transition probabilities, and products of such ratios differ by type. Basically, this assumes that the bipartite graph of worker–firm matches suffices to identify the latent worker types independently of wage observations. Finally, wage densities should be linearly independent with respect to worker types. That is to say, the wage density for type 1 cannot be obtained as a linear combination of the densities for types 2 and 3. This is the nonparametric equivalent of the full column rank condition of the regressor matrix in ordinary least squares.<sup>8</sup>

### 2.4. Parametric Specification

Although the identification proof is nonparametric, in practice we estimate a parametric model. The components of the parameter vector  $\beta$  are detailed herein.

<sup>8</sup>This is the standard identifying restriction for finite mixtures. See Teicher (1963), Yakowitz and Spragins (1968).

2.4.1. *Wage Distributions*

Static wage distributions (initial, out of unemployment, and upon employer change) are assumed log normal given the match type. Specifically,

$$f_{\text{static}}(w|k, \ell, x) = \frac{1}{\omega_{k\ell}(x)} \varphi\left(\frac{w - \mu_{k\ell}(x)}{\omega_{k\ell}(x)}\right), \tag{2}$$

with  $\varphi(u) = (2\pi)^{-1/2} e^{-u^2/2}$ . This specification allows for a match-specific mean  $\mu_{k\ell}$  and variance  $\omega_{k\ell}^2$ .

Within job spells (i.e., while  $D_{it} = 0$ ), we assume serial residual correlation,

$$f_{\text{dyn}}(w'|k, \ell, x', w, x) = \frac{1}{\sigma_{k\ell}(x')} \varphi\left(\frac{w' - \mu_{k\ell}(x') - \rho[w - \mu_{k\ell}(x)]}{\sigma_{k\ell}(x')}\right), \tag{3}$$

where  $(w', x')$  are one period forward relative to  $(w, x)$ . This dynamic specification has the advantage that parameters  $\mu_{k\ell}$  are mean wages. So, we can still perform a variance decomposition as in the AKM literature. Note that  $f_{\text{dyn}}$  and  $f_{\text{static}}$  share the same mean  $\mu_{k\ell}$  but have different variances  $\omega_{k\ell}^2$  and  $\sigma_{k\ell}^2$ .

2.4.2. *Mobility and Preferences*

The probability that a type  $k$  worker at time  $t$  transitions from a firm of type  $\ell = 1, \dots, L$  to a firm of type  $\ell' = 1, \dots, L$ , is specified as

$$M(\ell'|k, \ell, x) = \lambda_{k\ell'}(x) P_{k\ell\ell'}(x). \tag{4}$$

Parameter  $\lambda_{k\ell'}(x)$  is the worker-type  $k$  conditional probability of meeting with a different employer of type  $\ell'$ . This parameter governs *market segmentation*. It generalizes a multiplicative specification  $\lambda_{k\ell'} = \bar{\lambda}_k \nu_{\ell'}$ , where  $\bar{\lambda}_k$  would be derived from a search-matching function for skill  $k$ , and where  $\nu_{\ell'}$  would be the probability of sampling a firm of type  $\ell'$ .

Parameter  $P_{k\ell\ell'}(x)$  is the probability that the meeting results in a transition from  $\ell$  to  $\ell'$ . We assume a Bradley–Terry specification for  $P_{k\ell\ell'}(x)$  (Agresti (2003), Hunter (2004)). That is,

$$P_{k\ell\ell'}(x) = \frac{\gamma_{k\ell'}(x)}{\gamma_{k\ell}(x) + \gamma_{k\ell'}(x)}. \tag{5}$$

Parameter  $\gamma_{k\ell}(x)$ , with  $\sum_{\ell=1}^L \gamma_{k\ell}(x) = 1$ , measures the perceived value of the match  $(k, \ell, x)$ .<sup>9</sup> If the worker draws a same type job, with no loss of generality, since  $\lambda_{k\ell'}(x)$  is unrestricted, we assume that the worker moves with probability  $1/2$ .<sup>10</sup>

Sorkin (2018) is an example of such a view of job-to-job mobility. Search models such as Shimer and Smith (2000), Eeckhout and Kircher (2011), Lise, Meghir, and Robin (2016),

<sup>9</sup>Suppose that the worker moves from a firm  $\ell$  to a firm  $\ell'$  if  $\ln \gamma_{k\ell'} + u > \ln \gamma_{k\ell}$ , where  $u$  is a logistic error (a transitory cost or amenity shock). Then the moving probability is  $P_{k\ell\ell'} = \frac{\gamma_{k\ell'}}{\gamma_{k\ell} + \gamma_{k\ell'}}$ . We will return on this interpretation in Section 5.

<sup>10</sup>Parameters  $\lambda_{k\ell'}$ ,  $\gamma_{k\ell}$  are identified given knowledge of the unrestricted transition probabilities  $M(\ell'|k, \ell)$ . To wit, observe that  $M(\ell|k, \ell) = \lambda_{k\ell}/2$  trivially identifies  $\lambda_{k\ell}$  for all  $k, \ell$  (and  $x$ ), assuming no empty case. Then choice probabilities  $P_{k\ell\ell'}$  follow, and ratios  $\gamma_{k\ell'}/\gamma_{k\ell}$  are identified by odds ratios given the normalization  $\sum_{\ell=1}^L \gamma_{k\ell} = 1$ .

and [Bagger and Lentz \(2019\)](#) all offer variations on this theme. The virtue of our specification is to link the relative strength of preferences over the two matches,  $\gamma_{k\ell}(x)$  and  $\gamma_{k\ell'}(x)$ , to the observed propensity of a worker type  $k$ 's realization of a move from a type  $\ell$  firm to a type  $\ell'$  firm. That is, we assume that job-to-job mobility is a revelation of preferences over the two jobs involved while allowing for differences in chances to move,  $\lambda_{k\ell'}(x)$ . By not restricting the match value to be a function of mean wages  $\mu_{k\ell}(x)$ , we allow for the possibility that worker  $(k, x)$  may value more than just the wage in  $\ell$ .

We also model unemployment–employment transitions in a completely unrestricted way:

$$M(\ell'|k, 0, x) = \psi_{k\ell'}(x), \quad M(0|k, \ell, x) = \delta_{k\ell}(x).$$

By convention,  $M(0|k, 0, x) = 0$  since there is no transition from unemployment to unemployment. It follows that the probability of staying unemployed and the probability of staying with the same employer are

$$M(-|k, 0, x) = 1 - \sum_{\ell'=1}^L M(\ell'|k, 0, x) = 1 - \sum_{\ell'=1}^L \psi_{k\ell'}(x), \quad \text{and}$$

$$M(-|k, \ell, x) = 1 - \sum_{\ell'=0}^L M(\ell'|k, \ell, x) = 1 - \delta_{k\ell}(x) - \sum_{\ell'=1}^L \lambda_{k\ell'}(x)P_{k\ell\ell'}(x), \quad \ell > 0.$$

### 3. ESTIMATION METHOD

The firm classification in the data is unobserved. It is infeasible to evaluate the likelihood function for the formulation of the model where a firm's type is a random effect. Worker mobility across different firm types makes it impossible to separate the complete log likelihood (i.e.,  $\sum_i \ln \mathcal{L}_i(\beta|k, F)$ ) across firm types. Consequently, the estimation delivers a point estimate for each firm type instead of a posterior probability distribution over it as is the case for worker types. We therefore follow the approach in BLM of estimating a random effect model of wages and mobility given the firm classification by the EM algorithm. However, while BLM pre-estimates the firm classification using  $k$ -means clustering, we nest the worker EM algorithm inside a firm classification algorithm that we believe uses all the available information.

In this section, we first explain how we estimate  $\beta$  and posterior probabilities of worker types for a given firm classification  $F$ . We then explain how we set and update  $F$ . Lastly, we address the issue of the calibration of the number of groups  $K$  and  $L$ .

#### 3.1. The EM Algorithm for a Given Firm Classification

For given likelihood parameters,  $\beta$ , and a firm classification  $F$ , the posterior probability that worker  $i$  is of type  $k$  is

$$p_i(k|\beta, F) = \frac{\mathcal{L}_i(\beta|k, F)}{\sum_{k=1}^K \mathcal{L}_i(\beta|k, F)}, \tag{6}$$

which is worker  $i$ 's type- $k$ -complete likelihood relative to the likelihood of the observed data. Note that the factors  $1/q(\ell|F)$  in the definition of  $\mathcal{L}_i(\beta|k, F)$  in equation (1) appear

in the numerator and the denominator of the posterior probabilities in the same way and can be simplified out. This means that given  $F$ , the posterior probability is unaffected by the firm sampling assumption.

The EM algorithm iterates the following steps:

E-step For  $\beta^{(m)}$  and  $F$ , calculate posterior probabilities  $p_i(k|\beta^{(m)}, F)$ .

M-step Determine the update  $\beta^{(m+1)}$  as the  $\beta$  that maximizes the expected log likelihood  $\sum_i \sum_k p_i(k|\beta^{(m)}, F) \ln \mathcal{L}_i(\beta|k, F)$ .

The M-step updating formulas for wage distributions are the usual posterior probability-weighted mean, variance, and autocorrelation for Gaussian mixtures. For the dynamic specification, we can still exploit the model’s linearity to derive analytical formulas.

M-step updating formulas for transition probability are simple frequencies in the unrestricted case. For job-to-job transitions, transition probabilities are nonlinear parametric specifications, and this poses an additional challenge. In Appendix B, we develop an MM algorithm (Hunter (2004), Hunter and Lange (2004)) to maximize the expected log likelihood subject to the parametric restriction on job-to-job transition probabilities  $M(\ell|k, \ell, x)$ , which avoids the calculation of gradients.<sup>11</sup>

### 3.2. Firm Reclassification Given Other Parameters

Given an initial value  $(\widehat{\beta}^{(s)}, F^{(s)})$ , where  $\widehat{\beta}^{(s)}$  is obtained given  $F^{(s)}$  using the previous EM algorithm, we update  $F^{(s)}$  as

$$F^{(s+1)} = \arg \max_F \sum_{i=1}^I \sum_{k=1}^K p_i(k|\widehat{\beta}^{(s)}, F^{(s)}) \ln \mathcal{L}_i(\widehat{\beta}^{(s)}|k, F). \tag{7}$$

In practice, we only search for a firm reclassification that increases the likelihood, that is, some firms may keep the same types from the previous iteration if their types are already maximizing the likelihood. We first order firms by size. Then starting from the largest firm, say  $j = 1$ , we find  $\ell_1^{(s+1)}$  such that it maximizes the criterion in equation (7), keeping all other firm types equal to their values in  $F^{(s)}$ . Then we move to the second largest firm, say  $j = 2$ , and find  $\ell_2^{(s+1)}$  given  $\ell_1^{(s+1)}, \ell_3^{(s)}, \dots, \ell_j^{(s)}$ , and so on until the smallest firm,  $\ell_j^{(s+1)}$ . Thereafter, we return to the EM iterations with the updated  $F^{(s+1)}$  as well as the updated firm types’ sampling probabilities in the likelihood function.

We call this algorithm a Classification EM algorithm, as it resembles the eponym algorithm proposed by Celeux and Govaert (1992) as a variant of the EM algorithm of Dempster, Laird, and Rubin (1977). Gibbs’ inequality guarantees that the likelihood increases at each parameter update provided that one proceeds sequentially (sequential EM algorithm). Notice that, given  $L$ , and as shown by Celeux and Govaert (1992), the discrete

<sup>11</sup>The MM algorithm works by finding a function that minorizes the objective function and that is more easily maximized. Let  $f(\theta)$  be the objective concave function to be maximized. At the  $M$ -step of the algorithm, the constructed function  $g(\theta|\theta_m)$  will be called the minorized version of the objective function at  $\theta_m$  if

$$g(\theta|\theta_m) \leq f(\theta), \quad \forall \theta, \quad \text{and} \quad g(\theta_m|\theta_m) = f(\theta_m).$$

Then we maximize  $g(\theta|\theta_m)$  instead of  $f(\theta)$ , and let  $\theta_{m+1} = \arg \max_{\theta} g(\theta|\theta_m)$ . The above iterative method guarantees that  $f(\theta_m)$  converges to a local optimum or a saddle point as  $m$  goes to infinity because

$$f(\theta_{m+1}) \geq g(\theta_{m+1}|\theta_m) \geq g(\theta_m|\theta_m) = f(\theta_m).$$

classification should settle after a finite number of iterations. It should also be unaffected by very small changes in the other parameters. Hence, the asymptotic standard errors for  $\beta$  calculated given  $F$  should remain valid for the estimated classification. Of course, with administrative data, standard errors will be small thanks to the huge number of degrees-of-freedom.<sup>12</sup>

### 3.3. Starting Values and Choice of Group Numbers $K, L$

The estimation uses twenty different starting values of parameters, and we choose the set of results with the highest likelihood value.<sup>13</sup> For the number of groups, we first cluster firms using  $k$ -means clustering: taking firms' wage quintiles, average size, inflow rate, and outflow rate as inputs of characteristics. We select the value of  $L$  associated with the highest Calinski–Harabasz index, which is the ratio of the between-cluster and the within-cluster sum-of-squares. The idea is that we want to choose an  $L$  that optimally represents distinctive groups of similar firms. We find the optimal  $L$  to be 14 for periods 1–2 and 22 for periods 3–5.

In theory, one could apply the elbow method to pin down the number of worker types  $K$ . However, in practice, it is difficult to find an elbow in worker-type clustering. Using AIC/BIC with the likelihood expression is also not possible when we have double-sided heterogeneity. Therefore, we choose the maximum number of  $K$  that we could handle computationally, specifically  $K = 24$ , and leave a theoretical exercise of pinning down  $K$  as future research.

## 4. WHAT ARE THESE LATENT TYPES?

In this section, we describe the characteristics of latent types and the distribution of matches. The estimation essentially gives us classifications of workers and firms. We assign labels to these groups in a way that we can relate our results to the literature and to understand how sorting may manifest through wage effects that are common across types. We label  $k$  and  $\ell$  based on the global ordering of mean wages using a two-way fixed-effects projection of the estimated mean wages  $\mu_{k\ell}(x)$  with respect to the empirical distribution of match characteristics  $(k, \ell, x)$  in the sample, that is,

$$p(k, \ell, x) \propto \sum_{i,t} p_i(k) \sum_{t=1}^T \mathbf{1}\{\ell_{it} = \ell, x_{it} = x\}, \quad (8)$$

<sup>12</sup>Note that information-based standard errors are usually not thought to be widely useful for evaluating the precision of estimates of finite mixture models, unless the sample size is large (in relation to identification requirements—here, mobility) and groups are well separated. This is why bootstrap has become the preferred method for assessing the variability of EM-based estimates (see, e.g., O'Hagan, Murphy, Scrucca, and Gormley (2019) for a recent comparison for the case of simple Gaussian mixtures). Weighted Likelihood Bootstrap (Newton and Raftery (1994)), in particular, makes sure that all groups are simulated with nonzero probability. Unfortunately, more work is needed to evaluate the properties of information- and simulation-based validation methods for two-way mixtures such as the one we consider in this paper. Nevertheless, BLM develop a simulation method that they use to calculate bootstrap CIs for their wage variance decomposition exercise (see Table S6 in BLM's Online Appendix). Thanks to the large sample, and with only 2 years to estimate their model, they find very small 95%-CIs. This lack of theoretical work on two-way fixed-effect models, the higher computational cost induced by the nonlinearity of our transition probabilities and by the firm reclassification algorithm, and the fact that we use more years of data to improve identification explain why we do not calculate confidence intervals in this paper.

<sup>13</sup>This is standard practice with the EM algorithm, as it only delivers a local maximum.

where the estimated posterior probability  $p_i(k)$  is calculated using equation (6) at the estimated parameters  $\beta$  and  $F$ .<sup>14</sup> We will also denote as  $p(k)$  and  $p(\ell)$ , with a slight abuse of notation, the marginal distributions. This projection takes the standard form:

$$\mu_{k\ell}(x) = \bar{\mu}(x) + a_k + b_\ell + \tilde{\mu}_{k\ell}(x), \tag{9}$$

where  $\bar{\mu}(x)$  represents the marginal effect of tenure and experience interactions,  $a_k$  is the worker effect, and  $b_\ell$  is the firm effect.<sup>15</sup> The last term  $\tilde{\mu}_{k\ell}(x)$  is a residual capturing all remaining interactions. This “match effect” guarantees that the decomposition in (9) is an interpretation, not a restriction.<sup>16</sup> We relabel  $k$  and  $\ell$  after estimation so that  $a_k$  and  $b_\ell$  are now increasing in  $k$  and  $\ell$ . Hence, by construction,  $\mu_{k\ell}(x)$  tends to increase with respect to indices  $k$  and  $\ell$ . This relabeling renders an interpretation of higher- $k$  and higher- $\ell$  as higher wage types.

Note that  $\mu_{k\ell}(x)$  corresponds to the expected current wage *unconditional* on past wages. The autoregressive coefficient  $\rho$  is estimated to be around 0.4–0.6 across the five periods. This is not negligible but not huge. It therefore may mitigate a possible misspecification error in the assumed resetting property of employment and unemployment transitions.

#### 4.1. Worker Effects Explain Most of Conditional Mean Wages

Figure 1 shows the “fixed effects”  $a_k$  and  $b_\ell$ . Note that firm type  $\ell$  has a different set of values in periods 1–2 ( $L = 14$ ) than in periods 3–5 ( $L = 22$ ). Each type’s wage effect is shown with the rows of dots for each time period, indicated on the right vertical axes. By the wage effect ordering, types are arranged in ascending order from left to right. The figures also show the cumulative distributions,  $\sum_{k'=1}^k p(k')$  on  $a_k$  in the left panel

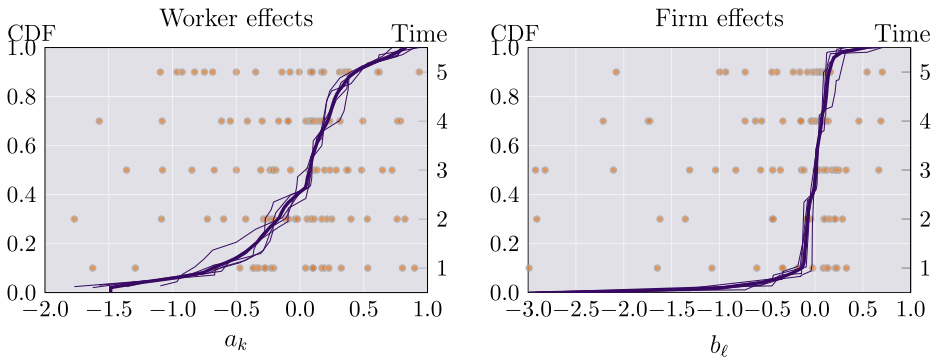


FIGURE 1.—Wage effect types. Notes: Thin curves show the cumulative distributions of fixed effects  $a_k$  and  $b_\ell$  for each time period on the left vertical axes. The thick line is the average over the 5 periods. The right vertical axes show the wage effects by time periods with circle markers. There are 24 worker effects in each time period. By the ordering of wage effects they are ascending from left to right. There are 14 firm types in periods 1 and 2, and 22 firm types in periods 3–5.

<sup>14</sup>The proportionality symbol  $\propto$  means that the right-hand side of the “equality” needs to be normalized for the left-hand side to be a proper probability.

<sup>15</sup>The fixed effects are normalized to average to zero over observations.

<sup>16</sup>Note that it is formally inappropriate to call it a “match effect.” A proper match effect would require classifying matches  $(i, j)$  and not workers  $i$  and firms  $j$  separately.

and  $\sum_{\ell'=1}^{\ell} p(\ell')$  on  $b_{\ell}$  in the right panel. The average over time is shown with a thick line and each time period is shown with a thin line. While there is substantial variation in the estimated wage effects by type over time, in particular at the extremes, the CDFs reveal that the actual distribution of matches over fixed effects is quite stable, possibly suggesting a discrete approximation of types to an underlying continuum.

Worker effects,  $a_k$ , show greater overall dispersion. Although firm effects have a wider negative support, most firm types have values of  $b_{\ell}$  concentrated closer to zero. This impression is reinforced by the cumulative match distribution that shows much greater dispersion in worker-type wage effects. This explains why firm heterogeneity contributes little to the log-wage variance. However, while the firm-type specific wage premiums appear to have little dispersion, we will show in the later section that mobility heterogeneity leads to variation in net present values of future earnings associated with each  $(k, \ell, x)$  match, and this plays an important role in determining workers' job preferences.

Figure 2 provides a graphical representation of the wage projection residuals  $\tilde{\mu}_{k\ell}(x)$  pooled over time by experience (rows) and tenure (columns). Worker effects are on the horizontal axes and firm effects on the vertical ones. There is a broad pattern that the projection misses nonlinearities in the worker dimension: Residuals are positive for low-wage workers and negative for high-wage workers. This is broadly true for low-wage workers regardless of the firm-type pairing, whereas the pattern overestimating the wage effect of high-wage workers is more pronounced in low-type firms. The pattern strengthens in tenure where the variance of the residuals increase in combination with more local nonlinearities. There is no significant variation in these observations over experience.

Much of the literature on matched employer–employee data since AKM was concerned with log-wage variance decomposition. Because this is not the main concern of this paper, we only here summarize the main results (see Appendix C for more details). The between-group variance—that explained by conditional means  $\mu_{k\ell}(x)$ —explains about 50–60%, depending on the time period, of the total log-wage variance. Interactions between characteristics, absorbed in  $\tilde{\mu}_{k\ell}(x)$  account for about 25% of the between-group variance. Worker effects explain more than 40% of mean wages. Firm heterogeneity contributes a lot less to the log-wage variance than worker heterogeneity. Specifically, its contribution decreases over time from 14% in period 1 to 9% in the last period. Lastly, the covariance between worker and firm effects, a classic measure of sorting, explains a similar, modest portion of the between variance (about 10%).

#### 4.2. *The Worker Index Is Monotone in Education and Tenure*

Figure 3a shows the distribution of worker types conditional on gender and education  $\pi^w(k|z^w)$  for the average over the 5 time periods. There is little variation in the patterns over time. The horizontal axis is worker-type rank based on the marginal worker-type distribution,  $p(k)$ . Groups are clearly ordered by education and gender. The distributions of worker types by education and gender exhibit clear stochastic ordering with, for example, more low-educated females in the lower-worker types than in the higher ones. So, worker types partially incorporate gender and education differences.

Figure 3b uses the estimated cross-sectional distribution  $p(k, \ell, x)$  to partial out the distribution of worker types across experience. All CDFs across experience groups are close to each other. A careful inspection reveals some tendency that middle experience workers (5–15 years of experience) represent a greater proportion of higher wage-type workers. Lower experience workers (less than 5 years) are particularly overrepresented in the lower worker types. That said, the overall impression is that of little variation across experience groups.

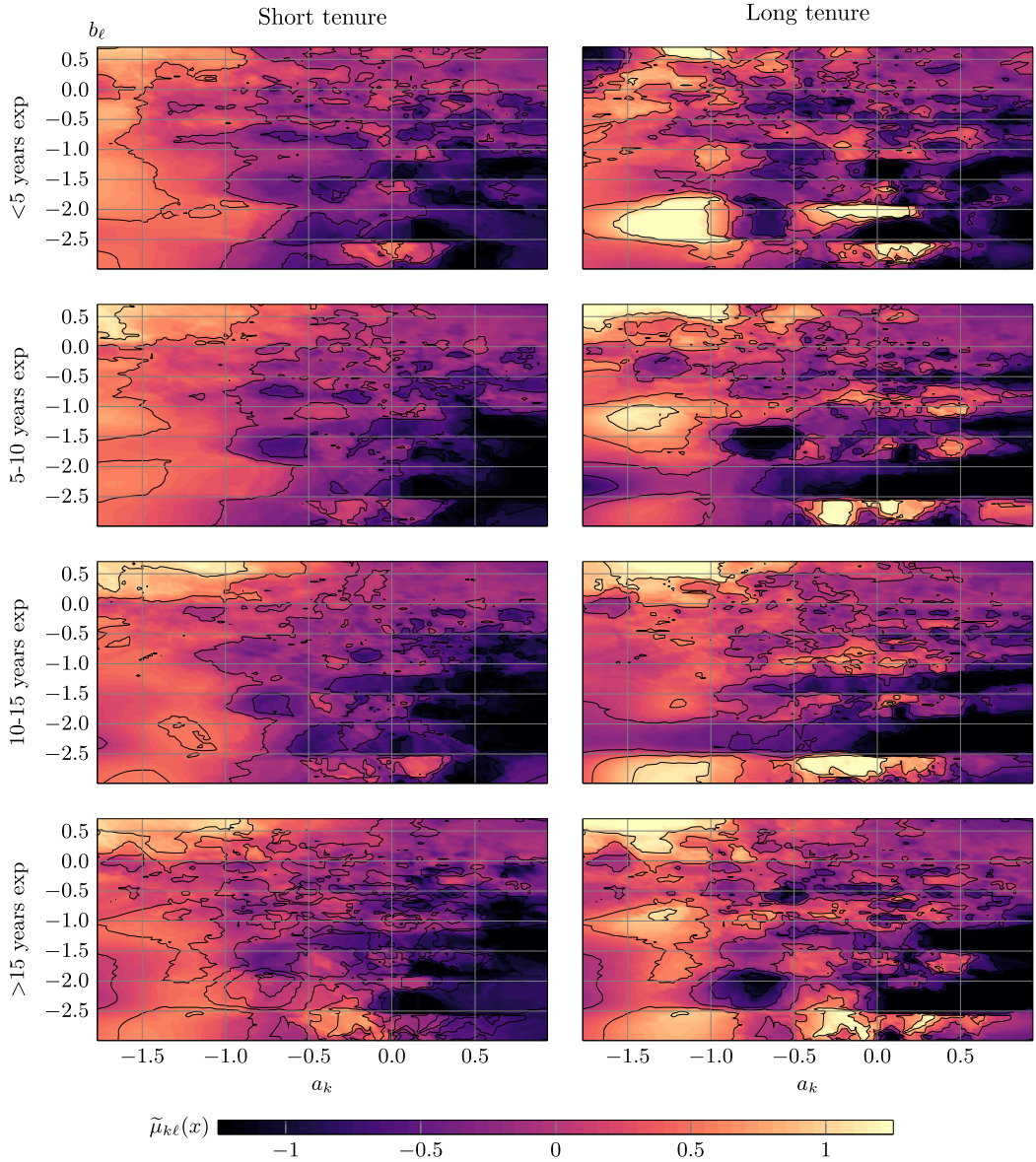


FIGURE 2.—Wage residuals by experience and tenure. Notes: Wage projection residuals—equation (9)—averaged over time by experience (rows) and tenure (columns). Average is done by kernel smoothing on the sample pooled over time. The kernel is Gaussian with bandwidth 0.25 done on the 10 nearest neighbors.

Figures 3c,d show the CDFs across tenure status while employed and unemployed. There is a clear stochastic ordering. Low-type workers tend to separate more from their employers than high-type workers. Low-type workers are also more likely to have long-term unemployment (greater than 26 weeks) than high-type workers. The figure shows the averages over the 5-time periods. The patterns are robust across all five periods. So, we conclude that tenure and worker types are correlated. Tenure is endogenous, and so



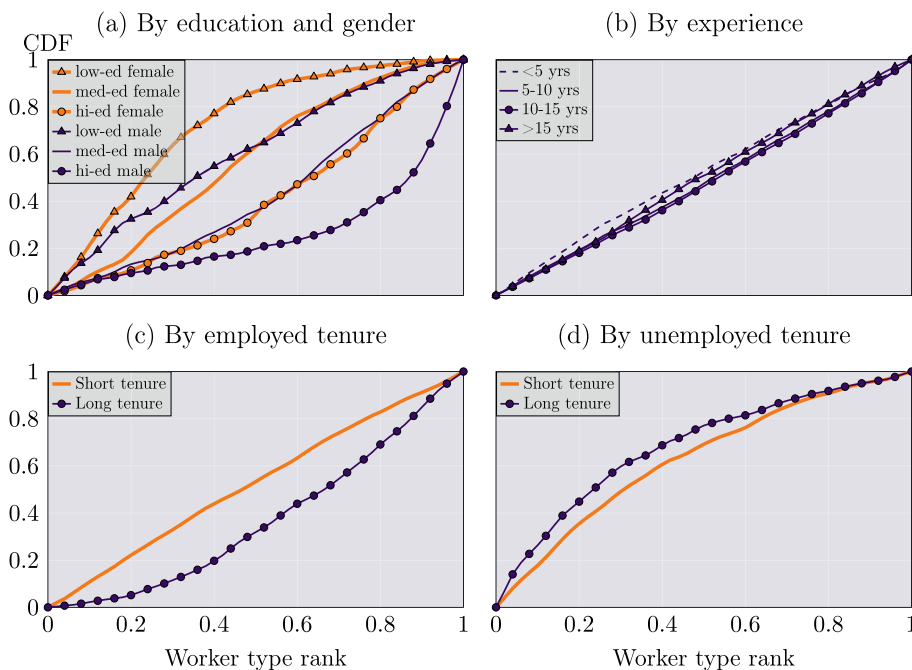


FIGURE 3.—Characteristics of worker types. Notes: CDF of worker type by education, gender and experience. Short tenure in a job is defined to be less than 100 weeks of employment. For unemployment, short tenure is defined to be less than 26 weeks. Averages over the 5 time periods.

we must allow transition probabilities to be tenure-dependent as how we have modeled it.

#### 4.3. Workers Accumulate in Firm Clusters With Lowest Turnover

Table II contains descriptive statistics for firms in period 1, which is representative of all periods (see the Appendix C for periods 2–5). Contrary to workers, the distribution of firm types is far from uniform. Firm groups starkly differ in firm sizes; the largest firms being clustered into smaller groups and the smallest in larger groups. Firm groups also vary in their turnover and age. Lower wage firms tend to have higher inflow and outflow rates, lower paid jobs tend to be less stable (e.g., [Bagger and Lentz \(2019\)](#), and [Jarosch \(2021\)](#)). In terms of industry shares by firm type (not shown), public administration and public services are more prevalent in lower firm groups. The higher firm groups (offering higher wages) are dominant in the construction, transportation, and communications industries in later periods. Finally, looking at all time periods (Appendix C), we see a (weak) declining trend of inflow and outflow rates by firm types over the whole estimation period. The share of firms with mixed public–private status has also increased.

For later reference, it is important to notice that two groups hire a majority of the workers (56%). Group 8 contains some very large firms from the public and mixed sectors. Group 9 also contains larger firms (more than 100 workers on average), mostly in the private sector. Both groups have relatively low exit rates, implying that their employees tend to stay employed longer. Firm clusters 10, 12, and 13 add another 23%. They are smaller private firms, also characterized by a small turnover. Similar patterns can be observed in all time periods.

TABLE II  
FIRM CHARACTERISTICS BY TYPE (PERIOD 1: 1989–1993).

$\ell$	No. Firms	No. Workers	Avg. Size	Legal Status			Avg. Inflow/Yr.	Avg. Outflow/Yr.	Avg. Age
				Private	Public	Mixed			
1	8094	12,127	1.50	0.62	0.34	0.04	0.64	0.62	3.69
2	5836	26,324	4.51	0.66	0.32	0.02	0.54	0.51	6.01
3	15,701	24,080	1.53	0.86	0.10	0.04	0.71	0.70	3.04
4	2558	57,614	22.52	0.74	0.23	0.03	0.54	0.44	6.51
5	26,576	40,283	1.52	0.90	0.08	0.02	0.46	0.45	4.74
6	12,729	51,080	4.01	0.84	0.16	0.00	0.43	0.41	7.49
7	12,282	84,309	6.86	0.93	0.02	0.04	0.53	0.49	4.02
8	91	371,375	4081.04	0.15	0.57	0.27	0.33	0.02	8.80
9	4801	613,119	127.71	0.81	0.16	0.03	0.37	0.22	7.79
10	21,333	143,988	6.75	0.87	0.12	0.00	0.34	0.31	8.05
11	18,384	60,604	3.30	0.97	0.01	0.02	0.56	0.53	5.66
12	9823	186,200	18.96	0.95	0.01	0.04	0.44	0.36	5.15
13	32,614	75,256	2.31	0.91	0.07	0.02	0.36	0.35	5.24
14	27,058	34,746	1.28	0.92	0.04	0.04	0.68	0.67	3.32

Note that it is generally the case that entry and exit rates are close for all firm types. Firm clusters do not generally grow in size. Except, indeed, for the biggest groups, which have inflows significantly greater than outflows. The approach in this paper is not adapted to push this analysis further, because in particular we have chosen to estimate different models for different time periods. This modeling choice allows us to emphasize the features of labor markets that remain stable over the long term. However, it is clear that there is some significant nonstationarity in group sizes that will need to be properly addressed in future work.

#### 4.4. *Instability of Latent Groups Across Periods*

In this section, we emphasize another type of nonstationarity. By estimating different models for different periods, worker and firm group labels may change arbitrarily across time. This may indicate a true change in the worker or firm types, or some statistical variance in the independent labelings. However, the amount of type change that we measure is suggestive of the existence of hidden dynamics in latent types.

In order to measure the stability of workers' labels across the different estimations, we assign to each worker in every period a type  $k_i$  corresponding to the highest posterior probability  $p_i(k)$ . Then we calculate the correlation between worker type in period  $n$  and worker type in period  $n + 1$  (see Table III). We do so both in terms of the ordered groups  $k$ , as well as the associated  $a_k$ . These correlations are substantial (around 55 to 60%) while also demonstrating substantial individual-type dynamics.

Table III also quantifies the stability of firm groups across periods. The correlation of firm types across periods is sizable (around 35 and 50%) but less than that of workers. We also report firm group consistency in terms of the wage effects  $b_\ell$ , where correlations range between 45–50%. Note that the grouping stability analysis between periods 2 and 3 involves comparing a classification with 14 groups in period 2 with one that has 22 groups in period 3. Interestingly, it does not seem to affect the correlation between ordered group labels  $\ell$  much.

TABLE III  
CLASSIFICATION STABILITY OF WORKER AND FIRM LATENT TYPES ACROSS PERIODS.

Period $n$ to $n + 1$	Workers		Firms	
	$\text{corr}(k, k')$	$\text{corr}(a, a')$	$\text{corr}(\ell, \ell')$	$\text{corr}(b, b')$
1 to 2	0.570	0.520	0.357	0.442
2 to 3	0.565	0.525	0.396	0.492
3 to 4	0.561	0.550	0.427	0.491
4 to 5	0.590	0.568	0.460	0.476

#### 4.5. Evidence of Weak PAM

To illustrate the matching of workers to firms, in Figure 4 we show how the worker-type composition varies across firm types  $p(k|\ell, x)$ . Experience is shown in rows and tenure in columns. Firm type is expressed on the horizontal axis by the type’s rank,  $r_\ell = \sum_{\ell'=1}^\ell p(\ell')$ . Worker type is also expressed in the type rank,  $r_k = \sum_{k'=1}^k p(k')$ . For a given pooled sample of  $\{r_\ell, r_k, \sum_{k'=1}^k p(k'|\ell, x)\}_{k,\ell}$  (over the five estimation periods), we evaluate worker-type composition by firm type and experience/tenure. For legibility, we illustrate the match distribution by 6 worker rank groups,  $q_1$  through  $q_6$ , each representing 1/6th of the worker mass. For example,  $q_1$  contains the lowest rank worker types that make up 1/6th of overall employed workers. Given the roughly uniform distribution of workers, this can be thought to represent roughly the 4 lowest wage types. Specifically, for rank group  $q_n$  we evaluate a kernel smoothed average of the pooled sample on the points  $\{\bar{r}_i, n/6\}_i$  where  $\bar{r}_i$  is an evaluation grid on  $[0, 1]$ .

It is clear by the relative increase in high-type worker frequency in high-type firms that there is positive sorting (Positive Assortative Matching, PAM) on worker and firm wage types. For example, by long-tenure workers, the lower ranking half of workers (groups  $q_1 - q_3$ ) represent 60–70% of the workers in the lowest firm types. In the highest firm types, they represent only about 30%. Sorting is nevertheless imperfect, and many workers—as it has already been emphasized—concentrate in a few firm types. This brings us to the next question of what determines job preferences.<sup>17</sup>

#### 4.6. Variance Decompositions

In this section, we extend the above two-way, fixed-effect, latent-type projection to the estimated model’s mobility parameters. For a given measure  $y_{k\ell}(x)$ , we perform the same kind of projection as in equation (9),

$$y_{k\ell}(x) = \bar{y}(x) + a_k^y + b_\ell^y + \tilde{y}_{k\ell}(x) \tag{10}$$

with respect to the same empirical distribution  $p(k, \ell, x)$  as stated in equation (8). This is done separately for each of our five time windows. Given the equation (10) projections, Table IV presents decompositions of  $V(y_{k\ell}(x))$  that are simple averages over the five time window variance decompositions for the model’s estimated wages, net present value, and mobility parameters.

<sup>17</sup>The discrepancy is there because the number of spells per firm type shown in Table II includes all workers in our age range and was computed by counting all spells at each firm. On the other hand, the pdf in Figure 4 was computed using posterior probabilities across  $k$  in each firm—workers who only have observations in the initial 2-year periods will have zero posterior probability, and hence, are not counted  $p(k, \ell)$  in Figure 4.

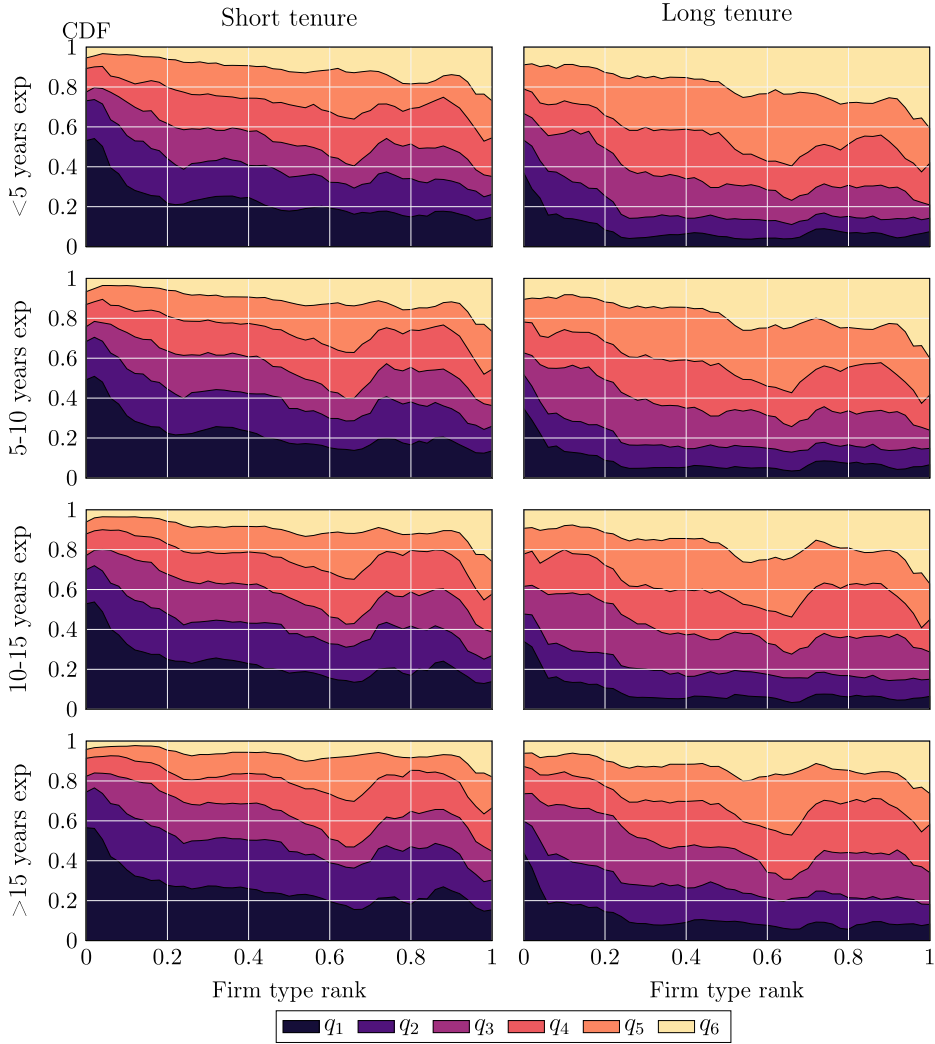


FIGURE 4.—The distribution of workers by firms,  $p(k|\ell, x)$ . Notes: Distribution of workers by firms and experience in rows and tenure in columns. Denote worker and firm type ranks by  $r_k = \sum_{k'=1}^k p(k')$  and  $r_\ell = \sum_{\ell'=1}^\ell p(\ell')$ . Each graph is based on the pooled samples of  $\{r_\ell, r_k, \sum_{k'=1}^k p(k'|\ell, x)\}_{k,\ell}$  over time periods. The graph shows kernel smoothed averages on evaluation points  $\{\bar{r}_l, q/6\}_{l,q=1,\dots,6}$ , where  $\bar{r}_l$  is a uniform grid on the  $[0, 1]$  interval. The kernel is Gaussian with bandwidth 0.05 done on 20 nearest neighbors.

Table IV documents that the worker and the tenure-experience effects combine to contribute 60% of the layoff rate variation. Furthermore, the correlation between worker wage fixed effects and worker layoff fixed effects across the 5-time periods is  $E[\text{Corr}(a^\mu, a^\delta)] = -0.74$  (calculation available upon request). This paints a simple and convincing picture of worker-type variation where high-wage workers are also low layoff rate workers.

The large worker side related contribution to layoff variance is striking. It is a well-known result that layoff rates are high- in low-wage matches. This is for example emphasized in Bagger and Lentz (2019) and Jarosch (2021). However, this relationship can

TABLE IV  
VARIANCE DECOMPOSITIONS.

	$\mu_{k\ell}(x)$	$\delta_{k\ell}(x)$	$\lambda_{k\ell}(x)$	$\psi_{k\ell}(x)$	$\gamma_{k\ell}(x)$
Coefficient of variation	0.08	1.43	1.36	1.08	0.96
Variance decomposition (%)					
worker effect, $a_k^y$	42.02	45.93	36.57	16.47	13.51
match effect, $\tilde{y}_{k\ell}(x)$	23.98	29.14	40.42	43.32	52.60
firm effect, $b_\ell^y$	12.76	6.51	12.24	22.20	23.56
sorting effect, $2\text{Cov}(a_k, b_\ell)$	10.57	4.09	-2.27	1.21	2.64
tenure and experience, $x$	10.67	14.33	13.05	16.80	7.69
$\bar{y}(x)$	4.42	6.11	6.32	18.95	2.99
$2\text{Cov}(a_k, \bar{y}(x))$	5.10	7.02	7.78	2.32	2.84
$2\text{Cov}(b_\ell, \bar{y}(x))$	1.15	1.20	-1.06	-4.47	1.86
$\text{Corr}(a_k, b_\ell)$	0.23	0.11	-0.04	0.05	0.10

be generated in a multitude of ways. The  $\delta$ -decomposition points to worker side heterogeneity as the dominant explanation with match specific variation an important second. A purely common firm side heterogeneity is a relatively weak force in overall layoff rate variation, but the next section emphasizes that it nevertheless shapes preferences and job mobility in important ways.

The  $\lambda$  variance decomposition demonstrates the importance of labor market segmentation in the mobility model. There is significant variation across worker- type labor market segments not just in the level of offer arrival rates, but also in the distribution over firm types that the offers come from. If job type offer distributions were common across worker types, firm and worker effects would dominate and match effects would be irrelevant.

The importance of the match effect variation is even more pronounced in the unemployed job finding rate distribution,  $\psi$ , variation. It foreshadows conclusions in Section 6 that emphasize the first-order impact on sorting in the Danish labor market due to cross-worker variation in the firm-type destination distribution out of unemployment.

To summarize, we see a broad picture where latent worker types can be reasonably distinguished by wages and layoff rates in an inverse relationship. Along both of these dimensions, worker side variation is the dominant contributor to variation. Furthermore, match effects are of first-order importance in both mobility and earnings parameter variances.

The analysis has to this point neglected a serious discussion of the model's estimated job preferences. It is a seemingly striking omission since they are if not *the*, then certainly among the most central measures in the estimation. They are a direct reflection of the worker's valuation of a job. However, preferences are identified through a revealed preference argument within worker type. Thus, the normalization for any  $(k, x)$  that  $\sum_\ell \gamma_{k\ell}(x) = 1$ , and consequently, the variance decomposition must be interpreted with care. The next section employs a within worker-type variation framework to study the role of job preferences. The variance decomposition does allow a comparison of the relative importance of the match effect variation relative to the firm effect variation. About one-third of the variation in preferences across firms is commonly agreed upon by workers (the firm effect). The remaining two-thirds reflect preference variation over firms that is specific to each worker type, a broad measure of job type preference disagreement across workers.

## 5. JOB PREFERENCES

A worker's move from one job to another can be seen as an expression of revealed preference of the new match over the old. A number of papers have studied revealed preferences through job-to-job mobility such as Sorkin (2018), Bagger and Lentz (2019), Taber and Vejlín (2020), and Lentz, Maibom, and Moen (2022).<sup>18</sup> Wages could indeed be a bad measure of match value in presence of compensating differentials. This is why, in our model, we have adopted the assumption that preferences (parameter  $\gamma$ ) are independent of wage levels (parameter  $\mu$ ), and identification manifests itself in workers' mobility patterns. In addition, our model allows for the existence of heterogeneity in the probability of drawing a job offer (parameter  $\lambda$ ). Finally, workers face idiosyncratic layoff ( $\delta$ ) and reemployment ( $\psi$ ) probabilities. We shall study the role of unemployment shocks in the next section. We focus in this section on job-to-job mobility.

It is appealing to adopt the interpretation that  $\gamma_{k\ell}(x)$  directly reflects a type  $(k, x)$  worker's valuation of a job with a type  $\ell$  firm,  $V_{k\ell}(x)$ . It is possible to interpret our wage and mobility model as a dynamic random utility model with independently distributed Gumbel utility shocks attached to job opportunities, or isomorphically a model with logistically distributed mobility costs. In such a model, we have that  $\gamma_{k\ell}(x) = \exp(V_{k\ell}(x)/\xi_{kx})$ , where  $\xi_{kx}$  is the variance parameter of the utility shock or mobility cost realization process. The revealed preference argument allows us to reveal workers' job valuations independently of observed job characteristics like wages.

Our model is not an equilibrium model. A few lessons can nevertheless be drawn from the theoretical matching literature. In the frictional descendants of the partnership model of sorting in Becker (1973) such as Shimer and Smith (2000), Gautier, Teulings, and van Vuuren (2010), Eeckhout and Kircher (2011) and Lise, Meghir, and Robin (2016), an equilibrium with sorting is characterized by preference rank variation, and often disagreement across workers. In these models, workers climb and fall down their respective firm ladders at the same pace independent of the value differences across different rungs on the ladder. The impact of job-to-job mobility on match allocation depends purely on the ordinal properties of preferences.<sup>19</sup> Sorting in these models is a result of different workers ranking firms differently in equilibrium (the surplus function is not positive everywhere the same). Two worker types that have the same ordinal preferences will have the same equilibrium match allocation regardless of any cardinal differences.<sup>20</sup> These models are distinct in that fundamental complementarities in match values manifest in ordinal preference variation across agent types.

In contrast, the sorting in Lentz (2010) and Bagger and Lentz (2019) is a sorting equilibrium where worker types agree on the ranking of firms, but preference/match value intensity varies across workers in relation to expected search gains. So, preferences vary across types only in a cardinal sense. Sorting results from the worker's search intensity

<sup>18</sup>For earlier references, see also Villanueva (2007) and Usui (2008).

<sup>19</sup>Hornstein, Krusell, and Violante (2011) emphasize this point in their discussion of job-to-job mobility rates.

<sup>20</sup>For example, in a positively assorted equilibrium, a low-type worker may reject a match with the most productive firm even though, possibly, this could be the most productive the worker would be. However, in order to make the high-type firm willing to forego the opportunity to match with a higher-type worker, the low-type worker would have to compensate the high-type firm with an unacceptably low wage. The assumption in the model that the firm forfeits the opportunity to match with another worker for the duration of the current match is in this context essential for sorting. Postel-Vinay and Robin (2002) is an example where the elimination of this assumption in an otherwise similar mobility environment results in no sorting even if there are complementarities in joint match values.

choice: the greater the value difference between origin and destination firms, the more intensely the worker searches away from the origin firm, and as a result generates a greater propensity to move.

Our model is a reduced form model insofar as it is agnostic as to how job preferences  $\gamma$  relate to match production and wages  $\mu$ . It is also agnostic as to how job offer probabilities  $\lambda$  (or market segmentation) relate to preferences.

We end this discussion by showing how sorting in our model relates to the standard supermodularity property of match values. Think of  $\gamma_{k\ell}(x)$  as a monotone transformation of the value of a  $(k, \ell, x)$  match. Different worker types may face different ladders, that is, different orderings of job types. Consider a worker in match  $(k, \ell, x)$  receiving a job offer of type  $\ell'$ . The probability of accepting the new job is  $P_{k\ell\ell'}(x) = \frac{\gamma_{k\ell'}(x)}{\gamma_{k\ell'}(x) + \gamma_{k\ell}(x)}$  if mobility incurs a logistic mobility cost or amenity. Now pick two worker types  $k' > k$  and two firm types  $\ell' > \ell$  such that  $\frac{\gamma_{k'\ell'}(x)}{\gamma_{k'\ell}(x)} > \frac{\gamma_{k\ell'}(x)}{\gamma_{k\ell}(x)}$ . Then  $P_{k'\ell\ell'}(x) > P_{k\ell\ell'}(x)$ . Even if the two worker types agree on the ranking of  $\ell$  versus  $\ell'$ , worker  $k'$  will climb her job ladder faster than worker  $k$ . Such ordering of the odds ratios  $\frac{\gamma_{k\ell'}(x)}{\gamma_{k\ell}(x)}$  holds when the function  $\ln \gamma_{k\ell}(x)$  is supermodular in  $(k, \ell)$ .

Of course, if worker  $k$  receives job offers at a lower frequency than worker  $k'$ , this conclusion can be reversed. Understanding sorting therefore involves a complete analysis of the mobility parameters and their interactions. A particularly important feature of our model is that the “chance” component of transition probabilities,  $\lambda_{k\ell'}(x)$ —that we call market segmentation—is separately identified from the “choice” component,  $\gamma_{k\ell'}(x)$ .

Absent of considerations regarding job offers, and assuming that workers start at the bottom rung, log supermodularity in  $\gamma$  will generate positive assortative matching. Note that this form of sorting implies a natural ordering of firm types that every worker recognizes and accepts. In reality, things could be a lot more complex. We could have different groups of workers with different ladders, for example. This is why, in the next subsection, we focus on detecting what we call “preference intensity,” which measures how far from indifference to firm types  $\ell = 1, \dots, L$  a worker  $(k, x)$  can be.

### 5.1. Job Preferences Intensify With Ability and Tenure

Indifference is maximal when  $\gamma_{k\ell}(x) = 1/L$  for all  $\ell$ . Note also that  $\gamma_{k\ell}(x) \geq 0$  and normalization  $\sum_{\ell} \gamma_{k\ell}(x) = 1$  allow to treat  $\boldsymbol{\gamma}_k(x) := (\gamma_{k\ell}(x))_{\ell}$  as a discrete probability distribution over  $\ell$ . Therefore, for each  $(k, x)$ , we calculate the Kullback–Leibler (KL) divergence from uniformity,

$$d_{\text{KL}}\left(\boldsymbol{\gamma}_k(x) \parallel \frac{1}{L}\right) = \sum_{\ell=1}^L \gamma_{k\ell}(x) \ln\left(\frac{\gamma_{k\ell}(x)}{1/L}\right) = \ln L + \sum_{\ell=1}^L \gamma_{k\ell}(x) \ln \gamma_{k\ell}(x),$$

which, by Gibb’s inequality, is nonnegative, and is equal to zero when  $\gamma_{k\ell}(x) = 1/L$  is uniform. Another interpretation is that  $\ln L - d_{\text{KL}}(\boldsymbol{\gamma}_k(x) \parallel \frac{1}{L})$  is the Shannon entropy of distribution  $\boldsymbol{\gamma}_k(x)$  given  $(k, x)$ . The maximum entropy, or maximum surprise, is attained for the uniform distribution. On the other hand, the greater the distance  $d_{\text{KL}}(\boldsymbol{\gamma}_k(x) \parallel \frac{1}{L})$ , the more intense the preference for certain firm types over others. We here use “intensity” as an antonym for “indifference.”

Table V shows  $d_{\text{KL}}(\boldsymbol{\gamma}_k(x) \parallel \frac{1}{L})$ , averaging over worker types  $k$  within three groups (low, medium, and high). If  $d_{\text{KL}}(\boldsymbol{\gamma}_k(x) \parallel \frac{1}{L}) = 0$ , then we have complete indifference and in the

TABLE V  
PREFERENCE INFORMATION BY WORKER TYPE,  $d_{KL}(\gamma_k(x) \parallel \frac{1}{L})$ .

Experience (Yrs.):		Short Tenure				Long Tenure			
		0–5	5–10	10–15	15+	0–5	5–10	10–15	15+
1989–93	Low	0.38	0.47	0.60	0.63	0.44	0.62	0.99	0.97
	Med	0.75	0.67	0.73	0.76	1.05	1.16	0.83	1.00
	High	1.03	0.89	0.91	1.12	1.06	0.99	1.01	1.17
1994–98	Low	0.37	0.34	0.48	0.51	0.84	0.65	0.85	0.77
	Med	0.63	0.59	0.65	0.85	0.81	0.67	0.73	0.92
	High	0.69	0.74	0.77	0.91	0.80	0.68	0.79	0.88
1999–03	Low	0.52	0.59	0.62	0.65	1.10	0.75	0.82	0.79
	Med	0.76	0.64	0.70	0.76	0.95	0.63	0.69	0.83
	High	0.84	0.86	0.95	1.09	0.94	0.95	0.88	1.04
2004–08	Low	0.51	0.56	0.55	0.62	1.18	0.98	0.87	0.70
	Med	0.81	0.79	0.92	0.95	0.98	0.94	1.09	0.95
	High	0.79	0.77	0.90	0.95	0.90	0.80	0.94	0.96
2009–13	Low	0.73	0.53	0.60	0.70	0.84	1.08	0.95	1.00
	Med	0.82	0.76	0.75	0.83	0.87	0.98	1.01	1.11
	High	0.99	0.93	0.98	1.01	1.06	1.09	1.14	1.18

language of labor search models we say that in this case there is no firm ladder for worker type  $k$  to climb. Evidently, workers face a ladder in the sense that they are not indifferent. We furthermore see a robust pattern that preferences for particular firm types strengthen in tenure. We also see such strengthening by experience for short tenure. The link with experience at long tenure is less clear.

There is also a strongly increasing relationship between job preferences and worker wage types, indicating that high-ability workers see greater value loss from mismatch. As explained above, the greater the expected gains, the more likely the move. This creates a basis for sorting whereby higher-type workers are more likely to move to their preferred firm types, all else equal. If those preferred firm types also tend to be characterized by higher wages, then this results in positive wage sorting driven by the intensity variation of job preferences across worker types, regardless of whether worker types agree on the ranking of firms.

### 5.2. Strong Pecuniary Motive at Short Tenure, Less so at Long Tenure

We proceed to ask how workers' preferences over firm types are related to the estimated characteristics of a  $(k, \ell, x)$  match. Table VI shows correlations between model parameters across firm types, averaged across worker types and time periods—but these correlations do not vary much over time—by experience and tenure. As one would expect, matches with higher wages ( $\mu$ ) and lower layoff risks ( $\delta$ ) are more preferred. We also see that short-tenure preferences are more closely aligned (i.e., have higher absolute correlations) with wages and layoff than long-tenure ones. Yet, the KL distance to indifference of job preferences was shown to increase with tenure. This implies that job preferences



TABLE VI  
CORRELATIONS BETWEEN PARAMETERS GIVEN  $(k, x)$ .

Experience (Yrs.):	Short Tenure				Long Tenure			
	0-5	5-10	10-15	15+	0-5	5-10	10-15	15+
$\ln \gamma_{k\ell}(x), NPV_{k\ell}(x)$	0.75	0.75	0.75	0.74	0.48	0.58	0.55	0.59
$\mu_{k\ell}(x), \gamma_{k\ell}(x)$	0.41	0.43	0.42	0.40	0.22	0.25	0.20	0.20
$\mu_{k\ell}(x), \delta_{k\ell}(x)$	-0.25	-0.31	-0.32	-0.27	-0.12	-0.22	-0.23	-0.14
$\mu_{k\ell}(x), \lambda_{k\ell}(x)$	0.14	0.14	0.12	0.10	0.18	0.21	0.22	0.20
$\mu_{k\ell}(x), \psi_{k\ell}(x)$	0.23	0.24	0.24	0.22	0.19	0.23	0.23	0.22
$\gamma_{k\ell}(x), \delta_{k\ell}(x)$	-0.45	-0.46	-0.42	-0.42	-0.14	-0.23	-0.24	-0.26
$\gamma_{k\ell}(x), \lambda_{k\ell}(x)$	0.10	0.06	0.03	-0.01	0.04	0.10	0.11	0.07
$\gamma_{k\ell}(x), \psi_{k\ell}(x)$	0.31	0.31	0.27	0.24	0.25	0.29	0.31	0.29
$\delta_{k\ell}(x), \lambda_{k\ell}(x)$	-0.14	-0.15	-0.12	-0.11	-0.11	-0.14	-0.14	-0.08
$\delta_{k\ell}(x), \psi_{k\ell}(x)$	-0.19	-0.21	-0.17	-0.12	-0.11	-0.15	-0.15	-0.09
$\lambda_{k\ell}(x), \psi_{k\ell}(x)$	0.85	0.83	0.81	0.76	0.75	0.80	0.75	0.70

Note: The correlations are calculated over firm types  $\ell$ , then averaged over worker types  $k$  and time periods, for given values of tenure and experience  $x$ . We use uniform weights in the calculation of the correlations across firm types  $\ell$ . Note that parameters  $\theta_{k\ell}(x)$ , where  $\theta$  stands for any of the parameters in the table, and normalized parameters  $\frac{\theta_{k\ell}(x)}{\sum_{\ell>0} \theta_{k\ell}(x)}$  produce identical correlations across firm types.

become more intense as tenure increases but pecuniary considerations become less important.<sup>21</sup>

Job preferences  $\gamma$  and reemployment rates  $\psi$  are weakly correlated (30%), at both short and long tenure. Hence, layoff shocks are not sending the workers back to the first rung of their ladders, but they are nevertheless strongly mixing. Job preferences and the rate at which offers arrive  $\lambda$  are orthogonal, indicating little ability of workers to generate more preferred offers. Market segmentation thus tends to slow down the speed at which workers move up their value ladder. Interestingly, job offers during employment  $\lambda$  and reemployment rates  $\psi$  are nearly perfectly correlated (80%). This indicates that unemployment and market segmentation work together in ways that are not aligned with workers' preferences for the jobs. Then choice incorporated in  $\gamma$  mixes things up, as  $\gamma$  is orthogonal to  $\lambda$ . One possible interpretation is that unemployed workers do not exert their choice as much as currently employed ones. Finally, layoff rates  $\delta$  are orthogonal to  $\lambda$  and  $\psi$ .

To conclude, the more preferred jobs are broadly the more remunerating and long-lasting ones. Unemployment and market segmentation are mixing. There is supporting evidence for partially directed search, but a large measure of search remains undirected.

### 5.3. Job Preferences Are Well Aligned With Net Present Values, Particularly at Short Tenure

The preceding subsection showed that job preferences  $\gamma$  correlate with wage levels  $\mu$ , job destruction  $\delta$  and reemployment rates  $\psi$ , and that  $\gamma$  and job-to-job offer rates  $\lambda$  are orthogonal. We end this section by considering the relationship between job preferences

<sup>21</sup>It is possible that job destruction is endogenous to the value of a match in which case the negative correlation between preferences and layoff risk reflects the reverse causality; that higher valued jobs are less likely to terminate. We consider both interpretations reasonable.

and the net present value of earnings,  $\text{NPV}_{k\ell}(x)$ . The NPV is a simple way of aggregating wages and employment shocks in a single index and it represents a fundamental measure of interest to most dynamic analyses of earnings.

For this, we simulate 20-year histories forward starting in a  $(k, \ell, x)$  match using a 5% annual discount rate. We prioritize simplicity in the NPV calculation:

$$\text{NPV}_{k\ell}(x) = \mathbb{E} \left[ \sum_{t=0}^T \beta^t \mu_{k\ell_t}(x_t) \right],$$

where the period  $t$  firm-type match and tenure-experience realization  $(\ell_t, x_t)$  follows the estimated laws of motion in the model and initial condition,  $\ell_0 = \ell$  and  $x_0 = x$ . Furthermore, we set  $\mu_{k0}(x) = 0$  for weeks of unemployment. This corresponds to a hand-to-mouth worker with log utility and a close to zero unemployment benefit replacement rate. It has the virtue of excluding wage variance  $\sigma_{k\ell}(x)$  as a confounding factor in the calculation. The NPV reflects purely the estimated log-wage averages and the mobility model parameters. We artificially maintain calendar time fixed in the sense that the NPV calculation for a particular time window panel uses its own mobility model for the entire 20-year horizon.

Table VI, first row, shows the average worker-type correlation between job preferences and net present value of future earnings. We see a strong positive relationship between a worker's job match preferences and the net present value of future earnings associated with the match—and this relationship is stronger than that between preferences and wages or layoff risks. The NPV relationship is also stronger for short tenures than for long tenures, and the difference between short- and long-tenure groups is decreasing in experience. We take this as evidence that pecuniary considerations (labor earnings and their loss) are stronger determinants of job preferences for younger workers and become less important as workers age, although they remain substantial. Additionally, we find an increasing correlation between preferences and NPVs over time (not shown).

It is worth noting that our estimation imposes no particular relationship between job preferences, wages, and layoff risk. The correlations in Table VI demonstrate that workers make job-to-job moves that reveal a preference for matches with firm types where the net present value of future earnings is greater, both because of higher wage and lower layoff risk. We also find that preferences are more substantially shaped by pecuniary considerations for short-tenure workers. As worker tenure increases, preferences become more intense and nonwage attributes increase in importance. While our results are supportive that workers' preferences for job types are closely aligned with pecuniary considerations, the correlations are sufficiently far from perfect that a substantial remaining residual can be attributed to amenities as emphasized in both Sorkin (2018) and Hall and Mueller (2018) (in addition to the idiosyncratic mobility cost/amenity responsible for the stochastic nature of the mobility decision).

## 6. SORTING

In this section, we measure sorting between latent worker and firm types. First, we do so by means of the correlation between the fixed effects obtained from the wage projection as typically done in the literature. Given our estimated importance of nonwage attributes in preferences as workers age and tenure rises, we then use a new measure of sorting, the mutual information, which allows us to flexibly account for both wage and nonwage sorting. Through the mutual information, we find a significant measure of sorting that

is not reflected in the wage correlation measure. Furthermore, the estimated mobility model allows us to understand how sorting arises. In the last two subsections, we quantify the importance of different sorting channels and their interactions through counterfactual analyses.

6.1. *Two Measures of Sorting*

*Wage Sorting.* The standard measure of sorting in the AKM literature is the correlation between worker and firm wage fixed effects. We show the correlation of  $a_k$  and  $b_\ell$  conditional on tenure and experience in Figure 5 for matching from the cross-sectional distribution  $p(k, \ell, x)$  averaged over the 5-time periods.

First, correlations increase with calendar time. This confirms results in [Bagger, Sørensen, and Vejlin \(2013\)](#) and [Card, Heining, and Kline \(2013\)](#) (on German data) although the increases we see are relatively modest. Second, comparing sorting by short and long tenure, the selection in jobs that last into long tenure reduces wage sorting. This is in line with our results in Section 5, where we find a stronger dependence between wages and job preferences at short tenure than at long tenure.

However, it is important to emphasize that the correlation of  $a_k$  and  $b_\ell$  measures sorting on wage effects that are common across types. Sorting on wages that is not common across worker types will not be detected in the classic wage fixed-effects correlation. Thus, the mapping of match preferences into wage fixed effects sorting is confounded by issues of search intensity, their commonality across worker types, and the importance of nonwage attributes.

*Type Sorting.* There may be dependencies between worker and firm types that the correlations of fixed effects are missing. We already saw that worker and firm effects are

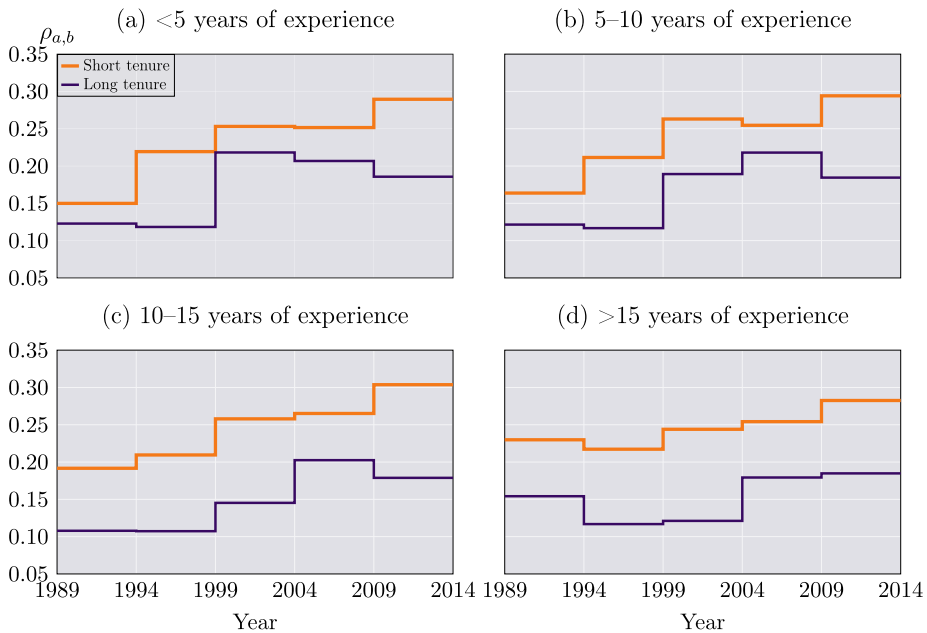


FIGURE 5.—Correlation by tenure and experience. Notes: Correlations between worker and firm wage fixed effects by tenure and experience, averaged over the five periods.

not exactly additive in mean wages. Even if they were, there may be sorting patterns that are not aligned with wage attributes. The mutual information (MI), without imposing any structure, measures the dependence between two variables, say  $X$  and  $Y$ . It quantifies the information about one random variable through knowledge about the other. Specifically, it is the Kullback–Leibler distance between a bidimensional distribution and the product of its margins (forced independence):

$$I(X, Y) = d_{\text{KL}}(p(X, Y) \parallel p(X)p(Y)) = \sum_{x,y} p(x, y) \ln\left(\frac{p(x, y)}{p(x)p(y)}\right).$$

It thus measures, in our case, the distance between observed and independent matching.<sup>22</sup> A drawback of the standard MI index is that it does not say if matching is positive or negative assortative (PAM or NAM), but the direction of matching can be easily inferred from a graphical illustration of worker-type composition across firm types as shown in Figure 4.

Another drawback of the standard MI is that its range is not bounded (just like the covariance). We shall thus consider a normalized version of the MI (like the correlation). If  $X$  is a deterministic function of  $Y$ , that is, perfect dependence, then all information conveyed by  $X$  is shared with  $Y$ . In this case, the mutual information is the same as the uncertainty contained in  $X$  (or  $Y$ ) alone, namely the entropy of  $X$  (or  $Y$ ). We therefore use the following normalized MI:

$$\bar{I}(X, Y) = \frac{I(X, Y)}{\min[H(X), H(Y)]},$$

where  $H(X) = -\sum_x p(x) \ln(p(x))$  is the entropy of  $X$ , and  $H(Y) = -\sum_y p(y) \ln(p(y))$  is the entropy of  $Y$ . In the extreme case of perfect dependence where  $X$  and  $Y$  are the same random variable, then  $\bar{I} = 1$ , and in the other extreme case of independence,  $\bar{I} = 0$ .

Figure 6 shows the normalized MI for the cross-sectional distribution  $p(k, \ell, x)$  given  $x$  and averaged over the 5-time periods. The results are starkly different from the wage fixed-effect correlations, wherein sorting is increasing in tenure. In Section 5, we emphasized that preference intensity is increasing in tenure, which is consistent with MI increasing in tenure. Thus, the selection into long-tenure jobs (presumably revealed to be preferred) is associated with higher dependence between worker types with particular firm types—resulting in increased type sorting. Significantly, correlations between wage effects do not reveal this sorting, which means that good long-term employment relationships are not necessarily the ones with higher wages. Instead, job preferences may depend upon *nonwage* factors (amenities).

## 6.2. Type Sorting Increases With Experience

We have already seen that job preferences are more intense when tenure increases, although not when one gets older for a given tenure. However, as a person ages, better opportunities present themselves, employment retention, and subsequent job transitions are less likely to be driven by wages. To better understand the role of mobility on sorting, we simulate cohorts of individuals drawn from the initial distribution. We expect cohorts to become more sorted as they age, but the sorting on wages should be weakening.

<sup>22</sup>The assortative marriage literature has also measured sorting by comparing the observed and independent matching probabilities between partners. See Greenwood, Guner, Kocharkov, and Santos (2016), for example.

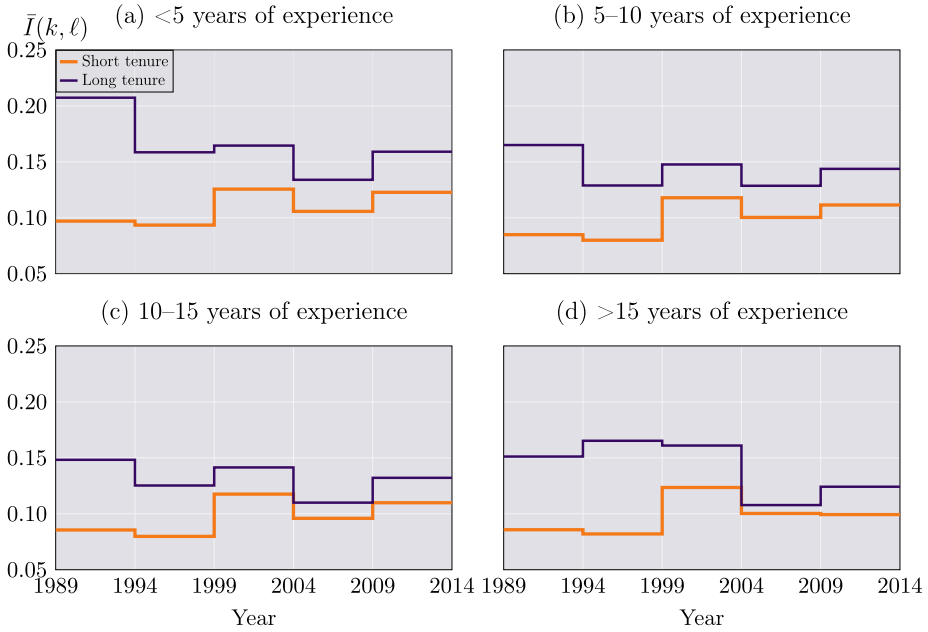


FIGURE 6.—Normalized mutual information by tenure and experience. Notes: Normalized mutual information by tenure and experience, averaged across the five periods.

Figure 7 displays the average profiles of sorting measures across the five periods where the profiles of both mutual information and wage fixed-effects correlations are qualitatively similar in each period. As expected, the mutual information picks an increasing trend with experience that the wage correlation ignores. Next, we investigate what contributes to the rising profile of type sorting.

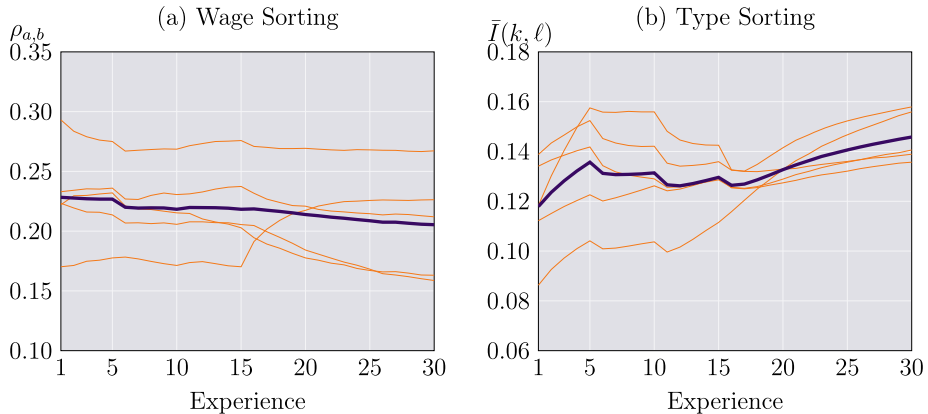


FIGURE 7.—Lifecycle sorting. Notes: Panel a shows the correlation between worker and firm wage fixed-effect over experience, while panel b displays the mutual information by experience. Both measures are averaged across the 5-time periods and drawn with thick line. The thin lines show individual time period sorting levels.

6.3. *With Age, Type Sorting Is by Choice and Wage Sorting Is by Chance*

The model has four channels of sorting: (1) job preferences  $\gamma$ ; (2) layoff  $\delta$ ; (3) market segmentation  $\lambda$ ; (4) reemployment  $\psi$ . To gauge the assortative matching importance of a channel over workers' age, we simulate cohorts of individuals drawn from the initial distribution where we counterfactually remove variation in each channel either across  $k$ , or across  $\ell$ , or both.<sup>23</sup> Mechanically, the model can arrive at no sorting between worker and firm types if all the channels have either (i) no variation across  $k$ , (ii) no variation across  $\ell$ , or (iii) both. Results are qualitatively similar in all three cases; hence, we only present case (iii) in what follows.

We present the counterfactuals to measure a channel's contribution to sorting. Specifically, denote by  $S_{-x}$  for  $x \in \{\gamma, \delta, \lambda, \psi\}$  the level of sorting in the counterfactual where variation in channel  $x$  has been removed. Denote by  $S$  the estimated benchmark from Figure 7. Channel  $x$ 's contribution to sorting is then  $(S - S_{-x})/S$ . We show these contributions in Figure 8.

*Transitions to and From Unemployment.* Figure 8, top row, displays the sorting contributions of the layoff ( $\delta$ ) and reemployment ( $\psi$ ) channels, for the correlation and the mutual information, respectively.

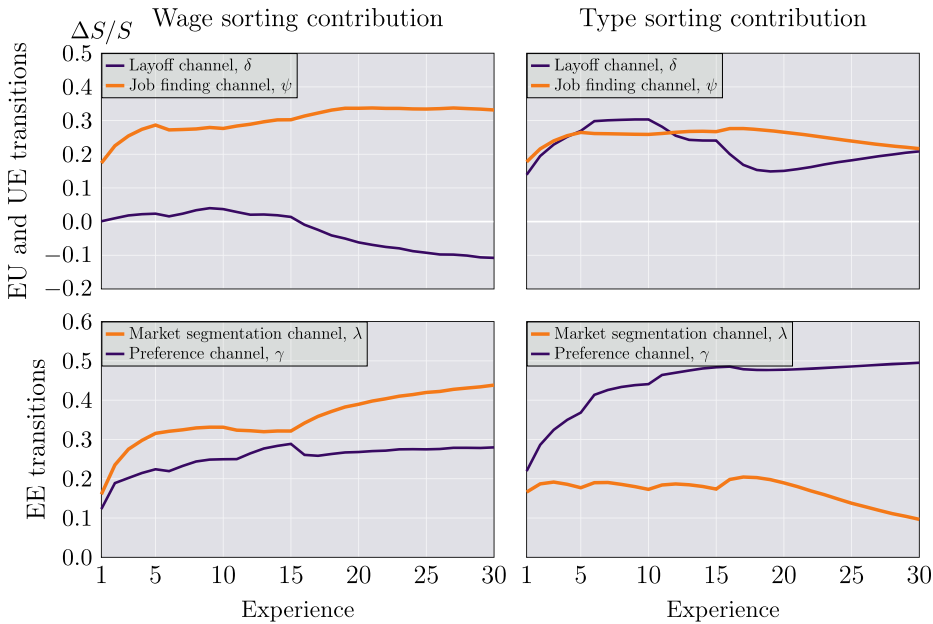


FIGURE 8.—Counterfactual transitions. Notes: Top row shows counterfactuals related to transitions into and out of unemployment. The bottom row shows counterfactuals related to job-to-job transitions. The columns organize counterfactuals into wage and type sorting. Denote by  $S_{-x}$  the counterfactual sorting when variation in  $x \in \{\delta, \psi, \gamma, \lambda\}$  across  $k$  and  $\ell$  are eliminated. The layoff contribution is measured by  $(S - S_{-x})/S$  where  $S$  is the relevant sorting measure shown in Figure 7, either wage correlation  $\rho_{a,b}$  or the type sorting,  $\bar{I}(k, \ell)$ .

<sup>23</sup>For example, for the layoff channel we remove  $k$  variation by the counterfactual  $\hat{\delta}_{k\ell} = \frac{1}{K} \sum_k \delta_{k\ell}$ . The preference channel is special in that eliminating  $\ell$  variation cannot be done without also eliminating  $k$  variation due to the normalization that  $\sum_{\ell} \gamma_{k\ell} = 1$ .

Removing heterogeneity from reemployment probabilities  $\psi$  results in a significant drop in both sorting indexes and, therefore, a substantial contribution to sorting. The impact on wage sorting is increasing with experience and mildly bells shaped in type sorting. Removing heterogeneity from layoff probabilities  $\delta$  has overall little impact on wage sorting (if anything it impacts sorting negatively for experienced workers). Their impact on type sorting is on the other hand substantial and peaking 10 years into careers. The experience effect on type sorting has the effect of making the two channels roughly equally important for high experience workers.

In summary, the parameters governing transitions to and from unemployment contribute to overall sorting with the latter being a relatively notable channel. Reemployment probabilities have an increasing impact in experience. Layoffs primarily impact type sorting.

*Job-to-Job Transitions.* Next, we compare and contrast the relative roles of job preferences  $\gamma$  (choice) and market segmentation  $\lambda$  (chance) in job-to-job transition probabilities. Figure 8, bottom row displays the results, again averaging over five periods where results in each period are qualitatively similar. Job preferences,  $\gamma$ , play a significant role in type sorting, measured by the normalized MI index. It is also a stronger source of sorting for more experienced and tenured workers as the cohort ages, reflecting the greater intensity of preferences. Job preferences appear to be the driving force for increased sorting over worker careers. The impact of preferences on wage sorting is also significant and is also increasing with experience, but the increase is less pronounced.

On the other hand, the effect of  $\lambda$ , which models the segmented offer arrival process, is a lot stronger in the correlation than in the MI index. While a relatively weak force, it does contribute to type sorting early in careers but then decline in importance as workers age to the point of irrelevance in late careers where preferences are strongly dominant. In terms of wage sorting  $\lambda$ , is a strong contributor, and in contrast to its impact on type sorting, the offer arrival process increases in importance in terms of wage sorting.

To sum up, we find that complementarities between worker and firm types in the offer arrival process while employed  $\lambda$  and in the reemployment probabilities  $\psi$  are key drivers of wage sorting during early career. The roles of  $\lambda$  and  $\psi$  become apparent in the correlation, and thus, seem to be driving a classical form of sorting via wage effects. Second, later in life, the way workers rank choices of jobs ( $\gamma$ ) dictates their matches and dominates differences in chances to move on-the-job ( $\lambda$ ). This is more apparent in the MI index than in the wage fixed-effects correlation, and it is what gives rise to the increasing sorting with age. We conclude that choice ( $\gamma$ ) is the key driver of type sorting, a dominance that only increases with age. Chance ( $\lambda$ ) becomes an increasingly dominant driver of wage sorting as workers age.

These four channels, however, may interact and drive sorting jointly. We investigate this in the next section.

#### 6.4. *The Interplay of Sorting Channels*

We end this study by showing how (1) job preferences  $\gamma$ , (2) layoff  $\delta$ , (3) market segmentation  $\lambda$ , and (4) reemployment  $\psi$  may interact and jointly affect wage and type sorting. It is natural that in a more restricted mobility model, multiple channels may work in concert. For example, an endogenous job destruction model would view layoff rates to be a function of the preferences for a match and, therefore, eliminating preference variation would also impact layoff variation.

Specifically, denote by  $S_{at}$  the sorting measure (either mutual information or wage effects correlation) for a given simulation at cohort age  $a$  and time period  $t$ . Denote by  $\bar{S} = \sum_{a,t} S_{at}$  the aggregate measure. To measure the relative importance of each channel to overall sorting, we decompose how much each channel contributes to  $\bar{S}$  through the sequential elimination of channels. Since the channels interact with each other, the order of elimination matters.<sup>24</sup> For example, consider the channel elimination order 4321, where channel 4 is first eliminated, then 3, 2, and 1. In this case, the marginal contribution of channel 4 is measured by  $(\bar{S} - \bar{S}_{-4})/\bar{S}$ . The marginal contribution of channel 3 is measured by  $(\bar{S}_{-4} - \bar{S}_{-43})/\bar{S}$ , and so forth, where  $\bar{S}_{-(\cdot)}$  denotes the aggregate counterfactual measure of sorting. There are 24 different elimination orders. We do channel elimination in 3 ways: no variation across  $k$ , no variation across  $\ell$ , and finally no variation across both  $k$  and  $\ell$ .<sup>25</sup>

Results are similar across periods and whether we eliminate  $k$ -variation,  $\ell$ -variation, or both. While the marginal contribution of each channel does vary by age, to focus attention on the overall interactions between the channels, we summarize the detailed picture through an aggregation over age and over the 5- time windows. As before, we present our results from cases where we eliminate both  $k$  and  $\ell$  variations. In Table VII, we calculate the importance of each channel considering all possible orders of elimination. We emphasize in bold the relative reduction in sorting of removing each of the four channels separately (direct channel effects).

In line with our previous section, the direct effect of job preferences  $\gamma$  is important both in type and wage sorting, in particular in type sorting. The biggest additional sorting reduction after removing the heterogeneity in job preferences is market segmentation  $\lambda$  (another 56% in the case of wage sorting). This is also true when one removes heterogeneity in job offer rates before job preferences. Job preferences and segmentation together explain 75% of sorting both in wages effects and latent types. Furthermore, for both components, their marginal impact is bigger in the absence of the other. This is true for both wage and type sorting. Thus, we find a substitutability between the two channels (each channel is less important in the presence of the other).

The direct effect of layoff rates  $\delta$  is negligible in wage sorting but sizable (22%) in type sorting. Eliminating the layoff channel implies that job preferences  $\gamma$  impact type sorting far less. And after the elimination of the job preference channel, there is no room left for the layoff channel. This indicates a strong complementarity of the job preferences and layoff channels, particularly for type sorting. We have already shown evidence that the most preferred jobs tend to be those that last longer. If these most preferred jobs stop lasting longer, then they become less prevalent in the population distribution, and further removing job preference heterogeneity matters less.

The second biggest drop after removing layoff heterogeneity ( $\delta$ ) is with reemployment probabilities  $\psi$ , and vice versa. Together they explain about 80–85% of type and wage sorting. Removing one of these two channels boosts the other one, but significantly more so for wage sorting than for type sorting. Just as we found a substitutability between the channels involved in the direct mobility between jobs ( $\gamma$  and  $\lambda$ ), we are finding a substitutability between the channels involved with mobility in and out of unemployment, ( $\delta$  and  $\psi$ ).

<sup>24</sup>Taber and Vejlín (2020) face a similar issue.

<sup>25</sup>The preference channel is treated as an exception in that we simply adopt indifference as the elimination of the channel under all three regimes,  $\gamma_{k\ell} = 1/L$ .



We note also that removing the heterogeneity in alternative job offer arrival rates  $\lambda$  affects wage sorting more than type sorting (direct effect of 45% vs. 25%). So, it seems that market segmentation, structuring the job offer rates  $\lambda$  as well as the reemployment probabilities  $\psi$ , are more important for wage sorting, implying that job offers are more likely to be directed toward matches that increase pay.

Overall, we find considerable interaction between the sorting components and that those interactions also vary across wage and type sorting. Our results highlight that sorting is indeed convoluted and multifaceted. Measuring and analyzing sorting sources requires a rich and flexible framework. With this nod to the considerable variation in marginal contributions depending on the interaction with other sorting components, the grand averages in Table VII show the reemployment channel,  $\psi$ , to be a forceful driver of both wage and type sorting. Job preferences,  $\gamma$ , are another major source of sorting, in particular, in sorting on latent types. Layoffs are a relatively weak source sorting, especially wage sorting.

TABLE VII  
DECOMPOSING WAGE AND TYPE SORTING.

Elimination Order	Wage Sorting				Type Sorting			
	$\gamma$	$\delta$	$\lambda$	$\psi$	$\gamma$	$\delta$	$\lambda$	$\psi$
1234	<b>0.252</b>	-0.068	0.450	0.366	<b>0.446</b>	-0.084	0.342	0.296
1243	<b>0.252</b>	-0.068	0.119	0.697	<b>0.446</b>	-0.084	0.180	0.458
1324	<b>0.252</b>	-0.173	0.555	0.366	<b>0.446</b>	-0.037	0.294	0.296
1342	<b>0.252</b>	0.207	0.555	-0.014	<b>0.446</b>	0.144	0.294	0.115
1423	<b>0.252</b>	0.461	0.119	0.169	<b>0.446</b>	0.156	0.180	0.218
1432	<b>0.252</b>	0.207	0.372	0.169	<b>0.446</b>	0.144	0.192	0.218
2134	0.209	<b>-0.024</b>	0.450	0.366	0.145	<b>0.217</b>	0.342	0.296
2143	0.209	<b>-0.024</b>	0.119	0.697	0.145	<b>0.217</b>	0.180	0.458
2314	0.457	<b>-0.024</b>	0.202	0.366	0.281	<b>0.217</b>	0.206	0.296
2341	0.179	<b>-0.024</b>	0.202	0.644	0.089	<b>0.217</b>	0.206	0.488
2413	0.031	<b>-0.024</b>	0.119	0.874	0.005	<b>0.217</b>	0.180	0.598
2431	0.179	<b>-0.024</b>	-0.028	0.874	0.089	<b>0.217</b>	0.096	0.598
3124	0.456	-0.173	<b>0.351</b>	0.366	0.573	-0.037	<b>0.168</b>	0.296
3142	0.456	0.207	<b>0.351</b>	-0.014	0.573	0.144	<b>0.168</b>	0.115
3214	0.457	-0.173	<b>0.351</b>	0.366	0.281	0.254	<b>0.168</b>	0.296
3241	0.179	-0.173	<b>0.351</b>	0.644	0.089	0.254	<b>0.168</b>	0.488
3412	0.381	0.207	<b>0.351</b>	0.062	0.580	0.144	<b>0.168</b>	0.108
3421	0.179	0.409	<b>0.351</b>	0.062	0.089	0.634	<b>0.168</b>	0.108
4123	0.120	0.461	0.119	<b>0.301</b>	0.414	0.156	0.180	<b>0.250</b>
4132	0.120	0.207	0.372	<b>0.301</b>	0.414	0.144	0.192	<b>0.250</b>
4213	0.031	0.549	0.119	<b>0.301</b>	0.005	0.565	0.180	<b>0.250</b>
4231	0.179	0.549	-0.028	<b>0.301</b>	0.089	0.565	0.096	<b>0.250</b>
4312	0.381	0.207	0.111	<b>0.301</b>	0.580	0.144	0.026	<b>0.250</b>
4321	0.179	0.409	0.111	<b>0.301</b>	0.089	0.634	0.026	<b>0.250</b>
<b>Average</b>	<b>0.245</b>	<b>0.129</b>	<b>0.256</b>	<b>0.369</b>	<b>0.300</b>	<b>0.214</b>	<b>0.183</b>	<b>0.302</b>
<b>Std. dev.</b>	<b>0.124</b>	<b>0.237</b>	<b>0.165</b>	<b>0.247</b>	<b>0.199</b>	<b>0.198</b>	<b>0.077</b>	<b>0.140</b>

Note: Marginal contributions of the four channels in the first column include (1) job preferences  $\gamma$ , (2) layoff  $\delta$ , (3) market segmentation  $\lambda$ , and (4) reemployment  $\psi$ . For example, 4123 means removing  $\psi$  first,  $\gamma$  second, then  $\delta$  and  $\lambda$ . The number in the table is the relative change in the correlation or the mutual information index.

## 7. CONCLUSION

In this paper, we extend the finite mixture framework of [Bonhomme, Lamadon, and Manresa \(2019\)](#) (BLM) to estimate a model of wages and employment mobility with two-sided heterogeneity. We propose a new parameterization for job-to-job transition probabilities, which allows us to meaningfully quantify the relative importance of job arrival process versus job preferences over different stages of workers' careers. However, this parameterization is highly nonlinear. We provide a strategy to overcome this estimation challenge by nesting an MM algorithm ([Hunter \(2004\)](#), [Hunter and Lange \(2004\)](#)) inside the M-step of BLM's EM algorithm. We also nest the whole EM algorithm (with the nested MM algorithm) inside a new firm classification algorithm, aimed at producing a classification of firms based on the full information of the wage-and-mobility likelihood instead of BLM's k-means prestep.

Our application focuses on Denmark in the period 1989–2013. Job-to-job transitions reveal that job preferences vary substantially more across worker types than across employer types. We show that the variation has a strong pecuniary component, in particular in terms of the net present value of future earnings associated with the current job. The relationship to current job wages is weaker. The pecuniary motive is particularly pronounced at short tenure and weakens with long tenure. Job preferences intensify both in age, experience, and tenure. We also show that higher wage-type workers have more intense preferences.

We introduce a new measure of sorting, the mutual information (MI). We employ MI to accurately represent the dependence between worker and firm types when sorting is convoluted by nonwage factors or nonlinearity in wage fixed effects. We find that type sorting, measured by the MI index, increases as workers settle into long-tenure relationships. We find the opposite for wage sorting, measured by wage fixed-effect correlations. This implies the increased importance of nonwage job characteristics in sorting patterns as workers age and select into long-tenure relationships.

In addition, we observe that the offer arrival process while employed and the reemployment probabilities are key drivers of sorting during early career. These two channels are apparent particularly in the wage correlation, and thus seem to be driving a classical form of sorting via wage effects. Finally, we show that these various channels of sorting interact with each other.

Our results highlight that sorting is convoluted and one cannot understand its sources by studying wages alone. There is not a global firm ladder across workers, nor is the ladder purely determined by wages. One general conclusion for future studies on sorting is that it is crucial to use both a flexible estimator taking into account the interaction between wages and mobility as well as a flexible measure of sorting that can accommodate nonlinearity and multidimensionality of matching. Our extension of BLM and the mutual information index offers a tractable and flexible way to do that, and goes beyond the wage model of AKM.

## APPENDIX A: IDENTIFICATION

Identification is essentially covered in [Bonhomme, Lamadon, and Manresa \(2019\)](#). We here discuss the argument given our model assumptions, insisting on the main steps.

*Firm Classification.* First, we assume that a vector of firm characteristics (wage moments, size, worker entry/exit statistics, etc.) exists that provides enough information to classify firms into  $L$  groups  $\ell \in \{1, \dots, L\}$ .

*Job-to-Job Moves.* Second, consider all possible worker trajectories with three wages  $(w_1, w_2, w_3)$ , where the first wage corresponds to a job of type  $\ell_1$  and the last couple to a job of type  $\ell_2$ , that is there is a job mobility between periods 1 and 2. The likelihood of this event (conditional on  $\ell_1$ -employment) is

$$p_{\ell_1\ell_2}(w_1, w_2, w_3) = \sum_k \pi_{\ell_1\ell_2}(k) f_{\ell_1}(w_1|k) g_{\ell_2}(w_2, w_3|k), \tag{11}$$

where  $g$  is the joint distribution of  $(w_2, w_3)$  for a short-tenure worker and  $f$  is the cross-sectional distribution of any wage and  $\pi$  is the overall employment probability:

$$\pi_{\ell_1\ell_2}(k) = M_{\ell_1\ell_2}(k) M_{\ell_1}^{-1}(k). \tag{12}$$

Assuming that  $M_{\ell_1\ell_2}(k) > 0$  and  $\sum_{\ell_2} M_{\ell_1\ell_2}(k) < 1$ , all worker groups are represented in all likelihoods. Moreover, given that  $M_{\ell_1}^{-1}(k) = 1 - \sum_{\ell_2} M_{\ell_1\ell_2}(k)$ , the probability of not moving is a function of  $M_{\ell_1\ell_2}(k)$ , we can solve for  $M_{\ell_1\ell_2}(k)$  given  $\pi_{\ell_1\ell_2}(k)$  (quadratic equation).

The idea is to store these probabilities in a matrix

$$P_{\ell_1\ell_2} = [p_{\ell_1\ell_2}(w_1, w_2, w_3)]_{w_1, (w_2, w_3)},$$

where we index rows by the values of  $w_1$  and the columns by the values of  $(w_2, w_3)$ . This is assuming discrete wages. If wages are continuous, we can work with CDFs and a wage grid as in BLM. This argument is essentially the same.

It follows from equation (11) that

$$P_{\ell_1\ell_2} = F_{\ell_1} D_{\ell_1\ell_2} G_{\ell_2}^T,$$

where

$$F_{\ell_1} = [f_{\ell_1}(w_1)]_{w_1, k}, \quad D_{\ell_1\ell_2} = \text{diag}[\pi_{\ell_1\ell_2}(k)]_k, \quad G_{\ell_2} = [g_{\ell_2}(w_2, w_3)]_{(w_2, w_3), k}.$$

BLM use a more sophisticated argument involving general mobility cycles. Let us simply assume that all transitions  $(\ell_1, \ell_2)$  have positive probability for any worker type. We also assume that the matrices  $F_\ell, G_\ell$  are full column rank and that the diagonal matrices  $D_\ell$  have all their  $K$  diagonal entries nonnil.

Fix  $\ell_1 = \ell$ . There exists a singular value decomposition:  $P_{\ell\ell} = U_\ell \Lambda_\ell V_\ell^T$ , where  $U_\ell$  and  $V_\ell$  are orthogonal matrices with  $U_\ell^T U_\ell = I_N, V_\ell^T V_\ell = I_M$  (say  $M = 2N$ ) and  $\Lambda_\ell$  is a rectangular diagonal matrix containing the singular values. The number of nonzero diagonal entries in  $\Lambda_\ell$  is equal to the number of groups  $K$ . Let  $\Lambda_{\ell 11}$  be the  $(K, K)$  diagonal matrix containing the nonzero singular values, and let  $U_\ell = (U_{\ell 1}, U_{\ell 2})$  and  $V_\ell = (V_{\ell 1}, V_{\ell 2})$  partition the columns of  $U_\ell, V_\ell$  so that  $P_{\ell\ell} = U_{\ell 1} \Lambda_{\ell 11} V_{\ell 1}^T$ .

Note also that since the columns of  $U_\ell$  are orthogonal vectors,

$$U_{\ell 2}^T P_{\ell\ell} = U_{\ell 2}^T U_{\ell 1} \Lambda_{\ell 11} V_{\ell 1}^T = 0_{(N-K) \times M}.$$

Hence,

$$U_{\ell 2}^T P_{\ell\ell} = U_{\ell 2}^T F_\ell D_{\ell\ell} G_\ell^T = 0_{(N-K) \times M}.$$

As  $D_{\ell\ell} G_\ell^T$  is a full row-rank  $(K, M)$  matrix, it follows that  $U_{\ell 2}^T F_\ell = 0_{(N-K) \times K}$ . Symmetrically,  $P_{\ell\ell} V_{\ell 2} = 0$  since  $V_{\ell 1}^T V_{\ell 2} = 0$ . Now, since  $F_\ell D_{\ell\ell}$  has rank  $K$ , it follows that  $G_\ell^T V_{\ell 2} = 0_{K \times (M-K)}$ .

Next, we deduce from  $P_{\ell\ell} = U_{\ell 1}\Lambda_{\ell 11}V_{\ell 1}^\top$  that

$$\Lambda_{\ell 11}^{-1}U_{\ell 1}^\top F_\ell D_{\ell\ell} G_\ell^\top V_{\ell 1} = I_K.$$

Given this result, if we define  $W_\ell = \Lambda_{\ell 11}^{-1}U_{\ell 1}^\top F_\ell$ , then  $W_\ell^{-1} = D_{\ell\ell} G_\ell^\top V_{\ell 1}$ .

It is important to note that all these  $W_\ell$  matrices have been calculated independently for all  $\ell$ . Their columns correspond to different  $k$ 's but there is no way, yet, to make sure that the columns of  $W_\ell$  for different  $\ell$ 's are consistently labeled.

Consider any  $\ell_2$ . If  $W_\ell$  and  $W_{\ell_2}$  have been constructed with a compatible ordering of rows and columns, then

$$Q_{\ell\ell_2} := \Lambda_{\ell 11}^{-1}U_{\ell 1}^\top P_{\ell\ell_2} V_{\ell_2 1} = \Lambda_{\ell 11}^{-1}U_{\ell 1}^\top F_\ell D_{\ell\ell_2} G_{\ell_2}^\top V_{\ell_2 1} = W_\ell D_{\ell\ell_2} D_{\ell_2\ell_2}^{-1} W_{\ell_2}^{-1}.$$

Therefore, the first thing to do is to find the permutation of the rows and columns of  $W_{\ell_2}$ , and of the columns of  $U_{\ell_2 1}$  and  $V_{\ell_2 1}$ , that makes  $W_\ell^{-1}Q_{\ell\ell_2}W_{\ell_2}$  diagonal.

Next, consider  $Q_{\ell\ell_2}$  and  $Q_{\ell_2\ell}$ . We have

$$Q_{\ell\ell_2}Q_{\ell_2\ell} = W_\ell D_{\ell\ell_2} D_{\ell_2\ell_2}^{-1} D_{\ell_2\ell} D_{\ell\ell}^{-1} W_\ell^{-1}.$$

Assuming that the diagonal entries of  $D_{\ell\ell_2} D_{\ell_2\ell_2}^{-1} D_{\ell_2\ell} D_{\ell\ell}^{-1}$  are all distinct, they are thus identifiable as the eigenvalues of  $Q_{\ell\ell_2}Q_{\ell_2\ell}$  and the eigenvectors are identified up to a multiplicative constant. Let  $\widehat{W}_\ell$  be the matrix of orthonormal eigenvectors, then

$$\widehat{W}_\ell = W_\ell \Delta_\ell = \Lambda_{\ell 11}^{-1}U_{\ell 1}^\top F_\ell \Delta_\ell$$

for some diagonal matrix  $\Delta_\ell$ .

We cannot immediately deduce that  $U_{\ell 1}\Lambda_{\ell 11}\widehat{W}_\ell = U_{\ell 1}U_{\ell 1}^\top F_\ell \Delta_\ell = F_\ell \Delta_\ell$  because  $U_\ell U_\ell^\top = I_N$  does not imply that  $U_{\ell 1}U_{\ell 1}^\top = I_N$ . Still,  $F_\ell \Delta_\ell = U_{\ell 1}\Lambda_{\ell 11}\widehat{W}_\ell$  because as  $U_{\ell 2}^\top F_{\ell_2} \Delta = 0_{(N-K)\times K}$ , we also have

$$\begin{pmatrix} \Lambda_{\ell 11}\widehat{W}_\ell \\ 0_{(N-K)\times K} \end{pmatrix} = U_\ell^\top F_\ell \Delta_\ell.$$

Hence,

$$U_{\ell 1}\Lambda_{\ell 11}\widehat{W}_\ell = U_\ell \begin{pmatrix} \Lambda_{\ell 11}\widehat{W}_\ell \\ 0_{(N-K)\times K} \end{pmatrix} = U_\ell U_\ell^\top F_\ell \Delta_\ell = F_\ell \Delta_\ell.$$

From  $F_\ell \Delta_\ell = U_{\ell 1}\Lambda_{\ell 11}\widehat{W}_\ell$ , we deduce  $F_\ell = U_{\ell 1}\Lambda_{\ell 11}\widehat{W}_\ell \Delta_\ell^{-1}$ . Since the rows of  $F_\ell$  sum to one (each column being a probability distribution), then  $\Delta_\ell$  is identified (since all its diagonal terms  $\Delta_k$  are identified). Hence,  $F_\ell$  is identified.

Finally, knowing  $W_\ell$  we know  $W_\ell^{-1} = D_{\ell\ell} G_\ell^\top V_{\ell 1}$ . A similar argument allows to separately identify the diagonal matrix  $D_{\ell\ell}$  and  $G_\ell$ . (The rows of  $G_\ell$  sum to one.)

*Long Tenure, No Mobility.* Consider the continuation of the preceding trajectories in the same firm  $\ell_2$  with two additional wages ( $w_4, w_5$ ). Call this new spell long tenure. The likelihood is

$$p_{\ell_1\ell_2}(w_1, w_2, w_3, w_4, w_5) = \sum_k \widetilde{\pi}_{\ell_1\ell_2}(k) f_{\ell_1}(w_1|k) g_{\ell_2}^{\text{ST}}(w_2, w_3|k) g_{\ell_2}^{\text{LT}}(w_4|k, w_3)$$

where the transition probability  $\widetilde{\pi}$  incorporates the additional no-move event. This likelihood easily identifies  $g_{\ell_2}^{\text{LT}}(w_4|k, w_3)$ .

APPENDIX B: AN MM ALGORITHM FOR THE M-STEP UPDATE OF WAGE DENSITIES AND TRANSITION PROBABILITIES

In the M-step of the standard EM algorithm, model parameters are updated to maximize the expected log likelihood of the data, which up to a constant is a minorizing function of the actual likelihood of the data; see, for example, [Hunter and Lange \(2004\)](#). In so doing, the M-step finds parameters that necessarily represent a likelihood function improvement relative to the minorizing function’s defining point. Unlike the standard EM algorithm applications, our expected log likelihood is not linear in some of our parameters (specifically, the ones related to the mobility model). Thus, we modify the expected log likelihood to obtain a minorization of the data likelihood that is indeed linear in our parameters and, therefore, allow an easy maximization of the minorizing function. In this Appendix, we document the first-order conditions associated with the maximization of the minorizing function as well as the modifications of it relative to the standard expected log likelihood.

B.1. Wage Parameters

The first wage in a job spell is drawn from the static wage distribution,

$$f_{\text{static}}(w|k, \ell, x) = \frac{1}{\omega_{k\ell}(x)} \varphi\left(\frac{w - \mu_{k\ell}(x)}{\omega_{k\ell}(x)}\right),$$

and subsequent wages within a job evolve according to the density,

$$f_{\text{dyn}}(w'|k, \ell', x', w, \ell, x) = \frac{1}{\sigma_{k\ell}(x')} \varphi\left(\frac{w' - \mu_{k\ell'}(x') - \rho[w - \mu_{k\ell}(x)]}{\sigma_{k\ell}(x')}\right),$$

where  $(w', \ell', x')$  refer one period forward relative to  $(w, \ell, x)$ .

The wage part of the expected log likelihood is given by

$$W = \sum_{i,k} p_i(k|\beta, F) \sum_{t=1}^{T_i} [D_{it-1} \ln f_{\text{static}}(w_{it}|k, \ell_{it}, x_{it}) + (1 - D_{it-1}) \ln f_{\text{dyn}}(w_{it}|k, \ell_{it}, x_{it}, w_{it-1}, \ell_{it-1}, x_{it-1})],$$

where for convenience,  $D_{i0} \equiv 1$ . While the types  $(k, \ell)$  remain the same within a job, wage parameters still vary across experience and tenure status. Let  $\bar{\mu}$  denote the mean wage parameters of  $(k, \ell, x)$  cell that we are trying to estimate,

$$\begin{aligned} \frac{dW}{d\bar{\mu}} &= \sum_{i,k} p_i(k|\beta, F) \sum_{t=1}^{T_i} \frac{1}{\sigma_{k\ell}(x_{it})} \left[ D_{it-1} \frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}} (w_{it} - \mu_{k\ell_t}(x_{it})) \right. \\ &\quad + (1 - D_{it-1}) \left( \frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}} - \rho \frac{d\mu_{k\ell_{t-1}}(x_{it-1})}{d\bar{\mu}} \right) \\ &\quad \left. \times (w_{it} - \mu_{k\ell_t}(x_{it}) - \rho(w_{it-1} - \mu_{k\ell_{t-1}}(x_{it-1}))) \right] = 0. \end{aligned}$$

For  $D_{it-1} = 0$ , there are three cases for the summand:

$$\begin{aligned}
 S_{ikt} &= D_{it-1} \frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}} [w_{it} - \mu_{k\ell_t}(x_{it})] \\
 &\quad + (1 - D_{it-1}) \left[ \frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}} - \rho \frac{d\mu_{k\ell_{t-1}}(x_{it-1})}{d\bar{\mu}} \right] \\
 &\quad \times [w_{it} - \mu_{k\ell_t}(x_{it}) - \rho(w_{it-1} - \mu_{k\ell_{t-1}}(x_{it-1}))],
 \end{aligned}$$

1.  $\frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}} = 0, \frac{d\mu_{k\ell_{t-1}}(x_{it-1})}{d\bar{\mu}} = 1$  then  $S_{ikt} = -\rho[w_{it} - \mu_{k\ell_t}(x_{it}) - \rho(w_{it-1} - \bar{\mu})]$ .
2.  $\frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}} = 1, \frac{d\mu_{k\ell_{t-1}}(x_{it-1})}{d\bar{\mu}} = 0$  then  $S_{ikt} = w_{it} - \bar{\mu} - \rho(w_{it-1} - \mu_{k\ell_{t-1}}(x_{it-1}))$ .
3.  $\frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}} = 1, \frac{d\mu_{k\ell_{t-1}}(x_{it-1})}{d\bar{\mu}} = 1$  then  $S_{ikt} = [1 - \rho][w_{it} - \rho w_{it-1} - (1 - \rho)\bar{\mu}] = 0$ .

We solve for a set of  $\mu_{k\ell}(x)$  for a given  $k$  using the following procedure. For a given  $k$ , let  $G$  denote the total possible combinations of  $\ell$  and  $x$ , and let  $g$  index the order of this combination. Construct  $A$ , a square matrix  $G \times G$  with each element in row  $g$ , column  $g'$  containing  $A_{gg'}$  where

$$\begin{aligned}
 A_{gg'} &= \sum_{i,k} p_i(k|\beta, F) \sum_{t=1}^{T_i} \frac{1}{\sigma_{k\ell}(x_{it})} \left[ D_{it-1} \frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}_g} \frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}_{g'}} \right. \\
 &\quad \left. + (1 - D_{it-1}) \left( \frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}_g} - \rho \frac{d\mu_{k\ell_{t-1}}(x_{it-1})}{d\bar{\mu}_g} \right) \left( \frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}_{g'}} - \rho \frac{d\mu_{k\ell_{t-1}}(x_{it-1})}{d\bar{\mu}_{g'}} \right) \right].
 \end{aligned}$$

Let  $b$  be a column vector  $1 \times G$  with each element

$$\begin{aligned}
 b_g &= \sum_{i,k} p_i(k|\beta, F) \sum_{t=1}^{T_i} \frac{1}{\sigma_{k\ell}(x_{it})} \left[ D_{it-1} w_{it} \right. \\
 &\quad \left. + (1 - D_{it-1}) \left( \frac{d\mu_{k\ell_t}(x_{it})}{d\bar{\mu}_g} - \rho \frac{d\mu_{k\ell_{t-1}}(x_{it-1})}{d\bar{\mu}_g} \right) [w_{it} - \rho w_{it-1}] \right].
 \end{aligned}$$

We have

$$A \begin{pmatrix} \mu_1 \\ \vdots \\ \mu_G \end{pmatrix} = b,$$

which by inversion of  $A$  delivers  $\mu_{k\ell}(x)$  for a given  $k$ .

The first-order conditions with respect to  $\sigma, \omega$ , and  $\rho$  deliver

$$\begin{aligned}
 \sigma_{k\ell}^2(x) &= \left( \sum_i p_i(k|\beta, F) \sum_t \mathbf{1}\{\ell_{it} = \ell, x_{it} = x, D_{it-1} = 0\} \right. \\
 &\quad \left. \times [w_{it} - \mu_{k\ell_t}(x_{it}) - \rho(w_{it-1} - \mu_{k\ell_{t-1}}(x_{it-1}))]^2 \right)
 \end{aligned}$$

$$\omega_{k\ell}^2(x) = \frac{\left( \sum_i p_i(k|\beta, F) \sum_{t=1} \mathbf{1}\{\ell_{it} = \ell, x_{it} = x, D_{it-1} = 0\} \right) \sum_i p_i(k|\beta, F) \sum_t \mathbf{1}\{\ell_{it} = \ell, x_{it} = x, D_{it-1} = 1\} [w_{it} - \mu_{k\ell_t}(x_{it})]^2}{\sum_i p_i(k|\beta, F) \sum_{t=1} \mathbf{1}\{\ell_{it} = \ell, x_{it} = x, D_{it-1} = 1\}}$$

$$\rho = \frac{\sum_i p_i(k|\beta, F) \sum_t \frac{1 - D_{it-1}}{\sigma_{k\ell}(x_{it})} (w_{it} - \mu_{k\ell_t}(x_{it})) (w_{it-1} - \mu_{k\ell_{t-1}}(x_{it-1}))}{\sum_i p_i(k; |\beta, F) \sum_{t=1}^{T_i} \frac{1 - D_{it-1}}{\sigma_{k\ell}(x_{it})} (w_{it-1} - \mu_{k\ell_{t-1}}(x_{it-1}))^2}$$

### B.2. Mobility Parameters

We maximize the part of the expected likelihood that refers to transitions, that is,

$$H(M|\beta^{(m)}) \equiv \sum_{k=1}^K \sum_{\ell=0}^L \left\{ n_{k\ell-}(\beta^{(m)}) \ln M_{k\ell-} + \sum_{\ell'=0}^L n_{k\ell\ell'}(\beta^{(m)}) \ln M_{k\ell\ell'} \right\},$$

where parameters are

$$M_{k0-} = 1 - \sum_{\ell'=1}^L M_{k0\ell'}, \quad M_{k0\ell'} = \psi_{k\ell'},$$

$$M_{k\ell-} = 1 - \sum_{\ell'=0}^L M_{k\ell\ell'}, \quad \ell \geq 1,$$

$$M_{k\ell\ell'} = \lambda_{k\ell'} P_{k\ell\ell'}, \quad P_{k\ell\ell'} = \frac{\gamma_{k\ell'}}{\gamma_{k\ell} + \gamma_{k\ell'}}, \quad \ell, \ell' \geq 1,$$

$$M_{k\ell 0} = \delta_{k\ell}, \quad \ell \geq 1,$$

and data are

$$n_{k\ell-}(\beta^{(m)}) = \sum_i p_i(k|\beta^{(m)}) \#\{t : D_{it} = 0, \ell_{it} = \ell, x_{it} = x\},$$

$$n_{k\ell\ell'}(\beta^{(m)}) = \sum_i p_i(k|\beta^{(m)}) \#\{t : D_{it} = 1, \ell_{it} = \ell, \ell_{i,t+1} = \ell', x_{it} = x\},$$

where  $\#\{\}$  denotes the cardinality of a set and where we reintroduce the control  $x_{it} = x$  to remind that we are estimating different parameters for all different control values  $x$ . In the rest of this section, we omit the reference to controls  $x$  and to  $\beta^{(m)}$ .

Parameters  $\psi_{k\ell}$  (job finding rate for unemployed) are thus updated as

$$\psi_{k\ell}^{(m+1)} = \frac{n_{k0\ell}}{n_{k0-} + \sum_{\ell'=1}^L n_{k0\ell'}}.$$

The rest of the likelihood is similar to the likelihood of a Bradley–Terry model except that when the incumbent firm  $\ell$  wins we do not know against which  $\ell'$ . The likelihood is thus rendered more nonlinear by the presence of the term in  $\ln M_{k\ell\bar{\cdot}}$ . An MM algorithm can still be developed as follows.<sup>26</sup>

With obvious notation, for  $\ell = 1, \dots, L$ , we can write

$$\begin{aligned} M_{k\ell\bar{\cdot}} &= 1 - \delta_{k\ell} - \sum_{\ell'=1}^L \lambda_{k\ell'} + \sum_{\ell'=1}^L \lambda_{k\ell'}(1 - P_{k\ell\ell'}) \\ &= \frac{1 - \delta_{k\ell}^{(s)} - \sum_{\ell'=1}^L \lambda_{k\ell'}^{(s)}}{M_{k\ell\bar{\cdot}}^{(s)}} \left( \frac{M_{k\ell\bar{\cdot}}^{(s)}}{1 - \delta_{k\ell}^{(s)} - \sum_{\ell'=1}^L \lambda_{k\ell'}^{(s)}} \right) \left( 1 - \delta_{k\ell} - \sum_{\ell'=1}^L \lambda_{k\ell'} \right) \\ &\quad + \sum_{\ell'=1}^L \frac{\lambda_{k\ell'}^{(s)}(1 - P_{k\ell\ell'}^{(s)})}{M_{k\ell\bar{\cdot}}^{(s)}} \left( \frac{M_{k\ell\bar{\cdot}}^{(s)}}{\lambda_{k\ell'}^{(s)}(1 - P_{k\ell\ell'}^{(s)})} \right) \lambda_{k\ell'}(1 - P_{k\ell\ell'}). \end{aligned}$$

Because the logarithm is concave, we can therefore minorize  $M(\bar{\cdot}|k, \ell) \equiv M_{k\ell\bar{\cdot}}$  as follows:

$$\begin{aligned} \ln M_{k\ell\bar{\cdot}} &= \ln \left( 1 - \delta_{k\ell} - \sum_{\ell'=1}^L \lambda_{k\ell'} + \sum_{\ell'=1}^L \lambda_{k\ell'}(1 - P_{k\ell\ell'}) \right) \\ &\geq \ln \underline{M}_{k\ell\bar{\cdot}} \\ &\equiv \frac{1 - \delta_{k\ell}^{(s)} - \sum_{\ell'=1}^L \lambda_{k\ell'}^{(s)}}{M_{k\ell\bar{\cdot}}^{(s)}} \ln \left( \frac{1 - \delta_{k\ell} - \sum_{\ell'=1}^L \lambda_{k\ell'}}{\frac{1 - \delta_{k\ell}^{(s)} - \sum_{\ell'=1}^L \lambda_{k\ell'}^{(s)}}{M_{k\ell\bar{\cdot}}^{(s)}}} M_{k\ell\bar{\cdot}}^{(s)} \right) \\ &\quad + \sum_{\ell'=1}^L \frac{\lambda_{k\ell'}^{(s)}(1 - P_{k\ell\ell'}^{(s)})}{M_{k\ell\bar{\cdot}}^{(s)}} \ln \left( \frac{\lambda_{k\ell'}(1 - P_{k\ell\ell'})}{\frac{\lambda_{k\ell'}^{(s)}(1 - P_{k\ell\ell'}^{(s)})}{M_{k\ell\bar{\cdot}}^{(s)}}} M_{k\ell\bar{\cdot}}^{(s)} \right). \end{aligned} \tag{13}$$

<sup>26</sup>The MM algorithm works by finding a function that minorizes the objective function and that is more easily maximized. Let  $f(\theta)$  be the objective concave function to be maximized. At the  $m$ th step of the algorithm, the constructed function  $g(\theta|\theta_m)$  will be called the minorized version of the objective function at  $\theta_m$  if

$$g(\theta|\theta_m) \leq f(\theta), \quad \forall \theta, \quad \text{and} \quad g(\theta_m|\theta_m) = f(\theta_m).$$

Then maximize  $g(\theta|\theta_m)$  instead of  $f(\theta)$ , and let  $\theta_{m+1} = \arg \max_{\theta} g(\theta|\theta_m)$ . The above iterative method guarantees that  $f(\theta_m)$  converges to a local optimum or a saddle point as  $m$  goes to infinity because

$$f(\theta_{m+1}) \geq g(\theta_{m+1}|\theta_m) \geq g(\theta_m|\theta_m) = f(\theta_m).$$



Note that both sides of the inequality are equal if  $(\lambda_{k\ell'}, \gamma_{k\ell}) = (\lambda_{k\ell'}^{(s)}, \gamma_{k\ell}^{(s)})$  (no parameter change). The MM algorithm maximizes

$$\sum_{k=1}^K \sum_{\ell=0}^L \left\{ n_{k\ell} \ln M_{k\ell} + \sum_{\ell'=0}^L n_{k\ell\ell'} \ln M_{k\ell\ell'} \right\}$$

instead of the initial objective.

Let

$$\tilde{n}_{k\ell\ell'}^{(s)} = n_{k\ell} \frac{\lambda_{k\ell'}^{(s)} (1 - P_{k\ell\ell'}^{(s)})}{M_{k\ell}^{(s)}}.$$

This is the predicted fraction of stayers such as home beats visitor  $\ell'$ .

One can update  $\gamma^{(s)}$  so as to maximize

$$\sum_{\ell=1}^L \sum_{\ell'=1}^L \left\{ \tilde{n}_{k\ell\ell'}^{(s)} \ln \frac{\gamma_{k\ell}}{\gamma_{k\ell} + \gamma_{k\ell'}} + n_{k\ell\ell'} \ln \frac{\gamma_{k\ell'}}{\gamma_{k\ell} + \gamma_{k\ell'}} \right\},$$

subject to the normalization  $\sum_{\ell=1}^L \gamma_{k\ell} = 1$ . Now, because

$$-\ln(\gamma_{k\ell} + \gamma_{k\ell'}) \geq 1 - \ln(\gamma_{k\ell}^{(s)} + \gamma_{k\ell'}^{(s)}) - \frac{\gamma_{k\ell} + \gamma_{k\ell'}}{\gamma_{k\ell}^{(s)} + \gamma_{k\ell'}^{(s)}}$$

with equality when  $\gamma = \gamma^{(s)}$  (see [Hunter \(2004\)](#)), we can instead maximize

$$\sum_{\ell=1}^L \left( \sum_{\ell'=1}^L (\tilde{n}_{k\ell\ell'}^{(s)} + n_{k\ell\ell'}) \right) \ln \gamma_{k\ell} - \sum_{\ell=1}^L \sum_{\ell'=1}^L \left( (\tilde{n}_{k\ell\ell'}^{(s)} + n_{k\ell\ell'}) \frac{\gamma_{k\ell} + \gamma_{k\ell'}}{\gamma_{k\ell}^{(s)} + \gamma_{k\ell'}^{(s)}} \right).$$

That is (taking special care of indices), for  $\ell = 1, \dots, L$ ,

$$\gamma_{k\ell}^{(s+1)} \propto \sum_{\ell'=1}^L (\tilde{n}_{k\ell\ell'}^{(s)} + n_{k\ell\ell'}) / \sum_{\ell'=1}^L \frac{\tilde{n}_{k\ell\ell'}^{(s)} + n_{k\ell\ell'} + \tilde{n}_{k\ell'\ell}^{(s)} + n_{k\ell'\ell}}{\gamma_{k\ell}^{(s)} + \gamma_{k\ell'}^{(s)}},$$

where the proportionality symbol means that the  $\gamma_{k\ell}^{(s+1)}$ 's should add up to one.

One can update  $\delta_{k\ell}^{(s)}, \lambda_{k\ell'}^{(s)}$  by maximizing

$$\sum_{\ell=1}^L A_{k\ell} \ln \left( 1 - \delta_{k\ell} - \sum_{\ell'=1}^L \lambda_{k\ell'} \right) + \sum_{\ell=1}^L n_{k\ell 0} \ln \delta_{k\ell} + \sum_{\ell'=1}^L (B_{k\ell'} \ln \lambda_{k\ell'}),$$

where

$$A_{k\ell} = n_{k\ell} \frac{1 - \delta_{k\ell}^{(s)} - \sum_{\ell'=1}^L \lambda_{k\ell'}^{(s)}}{M_{k\ell}^{(s)}}, \quad B_{k\ell'} = \sum_{\ell=1}^L (\tilde{n}_{k\ell\ell'}^{(s)} + n_{k\ell\ell'}).$$

The FOC for  $\delta_{k\ell}$  is

$$-\frac{A_{k\ell}}{1 - \delta_{k\ell} - \sum_{\ell'=1}^L \lambda_{k\ell'}} + \frac{n_{k\ell 0}}{\delta_{k\ell}} = 0 \Rightarrow \delta_{k\ell} = \frac{n_{k\ell 0}}{A_{k\ell} + n_{k\ell 0}} \left( 1 - \sum_{\ell'=1}^L \lambda_{k\ell'} \right)$$

and

$$1 - \delta_{k\ell} - \sum_{\ell'=1}^L \lambda_{k\ell'} = \frac{A_{k\ell}}{A_{k\ell} + n_{k\ell 0}} \left( 1 - \sum_{\ell'=1}^L \lambda_{k\ell'} \right).$$

The FOC for  $\lambda_{k\ell'}$  is

$$-\sum_{\ell=1}^L \frac{A_{k\ell}}{1 - \delta_{k\ell} - \sum_{\ell'=1}^L \lambda_{k\ell'}} + \frac{B_{k\ell'}}{\lambda_{k\ell'}} = 0 \Leftrightarrow -\frac{\sum_{\ell=1}^L (A_{k\ell} + n_{k\ell 0})}{1 - \sum_{\ell'=1}^L \lambda_{k\ell'}} + \frac{B_{k\ell'}}{\lambda_{k\ell'}} = 0,$$

after substituting out  $\delta_{k\ell}$ .

Let

$$C_{k\ell'} = \frac{B_{k\ell'}}{\sum_{\ell=1}^L (A_{k\ell} + n_{k\ell 0})}, \quad \bar{C}_k = \sum_{\ell'=1}^L C_{k\ell'}.$$

We finally obtain the following updating formulas:

$$1 - \sum_{\ell'=1}^L \lambda_{k\ell'}^{(s+1)} = \frac{1}{1 + \bar{C}_k} = \frac{\sum_{\ell=1}^L (A_{k\ell} + n_{k\ell 0})}{\sum_{\ell=1}^L (A_{k\ell} + n_{k\ell 0}) + \sum_{\ell'=1}^L B_{k\ell'}},$$

$$\lambda_{k\ell'}^{(s+1)} = \frac{C_{k\ell'}}{1 + \bar{C}_k} = \frac{B_{k\ell'}}{\sum_{\ell=1}^L (A_{k\ell} + n_{k\ell 0}) + \sum_{\ell'=1}^L B_{k\ell'}}$$

and

$$\delta_{k\ell}^{(s+1)} = \frac{n_{k\ell 0}}{A_{k\ell} + n_{k\ell 0}} \frac{1}{1 + \bar{C}_k} = \frac{n_{k\ell 0}}{A_{k\ell} + n_{k\ell 0}} \frac{\sum_{\ell=1}^L (A_{k\ell} + n_{k\ell 0})}{\sum_{\ell=1}^L (A_{k\ell} + n_{k\ell 0}) + \sum_{\ell'=1}^L B_{k\ell'}}.$$

For a given value of  $\beta^{(m)}$ , the sequence  $(H(M^{(s)}|\beta^{(m)}))$  is increasing. The MM algorithm can thus be stopped at any time, not only after convergence, to deliver the updated values of transition parameters,  $(\psi^{(m+1)}, \delta^{(m+1)}, \lambda^{(m+1)}, \gamma^{(m+1)})$ .

## APPENDIX C: VARIANCE DECOMPOSITION

As discussed in Section 2.4, both stayers and movers share the same mean wage  $\mu_{k\ell}$  in our model; this allows us to perform a variance decomposition as in the AKM literature. We decompose the log-wage variance following the projection in (9). Specifically, we expand the initial sample so that each individual observation  $i$  is repeated  $K$  times, one for each  $k$  with an associated weight equal to the estimated posterior probability  $p_i(k)$ . Then we decompose the cross-sectional log-wage variance into the effects of experience and tenure ( $\bar{\mu}$ ), worker types ( $a_k$ ), firm types ( $b_\ell$ ), match effects ( $\tilde{\mu}_{k\ell} = \mu_{k\ell} - a_k - b_\ell$ ), and the residual ( $w - \mu_{k\ell}$ ).

Table C.I shows the log-wage variance decomposition over time. The between ( $x, k, \ell$ )-groups and the within-group variances have rather similar contributions, although when the economy is depressed (as in periods 1 and 5), the residual variance contributes much less (up to 50% and down to 40%). The main between-variance contributor is the worker effect  $a_k$ , at just around 25%. After the residual and the worker effect comes the match effect  $\tilde{\mu}_{k\ell}$  that measures the degree of nonlinear interaction between worker and firm effects, whose contribution is decreasing over time (from 15% to 10%). Next, we have the firm effect  $b_\ell$  (from 8% to 5%) and the sorting effect  $2\text{cov}(a, b)$  (from 5% to 7%). We thus confirm an increasing sorting trend that has been previously observed (Bagger, Sørensen, and Vejlin (2013), Card, Heining, and Kline (2013), Song et al. (2015)). In terms of the wage correlation, the increasing sorting trend is monotone over our quarter-century study period with an initial correlation of about 0.2 increasing to about 0.3. Finally and interestingly, the contribution of tenure and experience  $\bar{\mu}(x)$ , which used to be small (around 4% in total, including covariances), has increased in the last period. We lack data to decide whether this is significant or not.

We thus find that the contribution of worker heterogeneity is considerably less than the usual AKM estimates reported in Table C.II, and the residual variance is considerably higher. Furthermore, the correlation between worker and firm effects in this literature is all over the place—sometimes negative, often zero, and sometimes positive, in any case seldom as large as 0.2 or 0.3. However, it is well known that the OLS fixed-effect estimator overfits and induces a negative bias on the correlation of fixed effects. Hence, asymptotic small-sample bias correction techniques have been developed. Andrews et al. (2008) were the first to address this issue. In recent work, Kline, Saggio, and Solvsten (2020)

TABLE C.I

LOG-WAGE VARIANCE DECOMPOSITION (PERCENTS OF TOTAL VARIANCE) AND WAGE CORRELATION.

	89–93	94–98	99–03	04–08	09–13
Residual	42.49	46.21	48.86	50.50	39.19
Total between	57.51	53.79	51.14	49.50	60.81
worker effect, $a_k$	23.67	22.78	19.32	22.86	25.91
match effect, $\tilde{\mu}_{k\ell}(x)$	15.65	14.20	13.13	11.57	10.46
firm effect, $b_\ell$	8.29	6.98	8.87	4.98	5.46
sorting effect, $2\text{Cov}(a_k, b_\ell)$	5.46	5.04	6.09	5.34	6.88
tenure and experience, $x$	4.44	4.79	3.72	4.74	12.09
$\bar{\mu}(x)$	1.29	2.37	1.54	2.52	4.48
$2\text{Cov}(a_k, \bar{\mu}(x))$	2.66	2.01	1.88	1.76	6.01
$2\text{Cov}(b_\ell, \bar{\mu}(x))$	0.50	0.40	0.30	0.46	1.60
Wage correlation, $\text{Corr}(a_k, b_\ell)$	0.19	0.20	0.23	0.25	0.29

TABLE C.II  
LOG-WAGE VARIANCE DECOMPOSITIONS (IN THE LITERATURE).

	AKM												BLM	
	FR 76-87	US1 84-93	AU 90-97	US2 90-99	IT 81-97	BR 95-05	DE 85-91	DE 02-09	US3 80-86	US3 07-13	DK 87-13	VE 99-01	FR2 02-14	SW 02-04
Worker effect	77	82	66	64	44	60	61	51	48	53	43	61	56	60
Firm effect	30	19	37	15	13	27	19	21	16	12	12	13	15	3
Residual	16	9	5	9	15	7	8	5	20	15	39	10	16	25
Sorting effect	-27	-2	-22	1	2	3	2	16	2	7	3	16	-6	12
Match effect				5			3	2						
Observed	7	52	3	4	8	3	11	3	8	7	2		7	
heterogeneity	-3	-69	9	1	15	-3	1	1			1		10	
Sorting	1	9	2	1	3	2	2	2			1		1	
	-0.28	-0.025	-0.23	0.010	0.044	0.040	0.034	0.25	0.029	0.14	0.074	0.283	-0.11	0.49

Note: All decompositions under the label AKM refer to estimations of the AKM model with or without a match-specific component on various matched employer-employee data using the method of Abowd, Creedy, and Kramarz (2002). Columns FR (France) and US1 come from Table 2 in Abowd, Creedy, and Kramarz (2002). Column AU (Austria) comes from Table 4 in Grunetter and Lalive (2009). Column US2 results from Tables 5 and 6 in Woodcock (2015). They use the LEHD data discussed in Abowd et al. (2009). Column IT (Italy) refers to Table 4 in Iranzo, Schivardi, and Tosetti (2008). Column BR (Brazil) refers to Table 1 in de Melo (2018). Columns DE (Germany) are calculated from Table III in Card, Heining, and Kline (2013). Columns US3 result from our own calculations from Table 2 in Song, Price, Guvenen, Bloom, and von Wachter (2018). The worker effect is the sum of the “mean worker effect across firms” and the “difference of worker effect from mean worker effect across firms”, and the effect of observed heterogeneity is the sum of the “mean xb across firms” and the “difference of xb from mean xb across firms”. Column VE is our own calculations based on Table II of Kline, Saggio, and Solysten (2018). They apply their “leave-one-out” estimator to Italian data from the Veneto region. Column FR2 is from Table 2 of Azkarate-Askasia and Zerecero (2019) and uses French data. Column BLM refers to Bonhomme, Lamadon, and Manresa (2017)’s estimation for Sweden, 2002–2004 (Table II).

make the bias correction technique more practical by introducing a Jackknife technique and Azkarate-Askasua and Zerecero (2019) develop a bootstrap approach. In all cases, the bias correction is quite small. The residual explains less than 20% of the log-wage variance and the worker effect is nearly always explaining more than 50%. Finite-sample corrections presumably reduce the amount of overfitting but cannot eliminate it.<sup>27</sup>

The only estimation of an AKM model yielding a large residual component is the one in Bagger, Sørensen, and Vejlin (2013) on the same Danish data as ours. We thus performed Monte Carlo simulations of our estimated discrete mixture model. The artificial OLS estimates are in line with the variance decomposition estimated by Bagger, Sørensen, and Vejlin (2013). The residual contribution was reduced by about 10 percentage points, and the worker effect increased by 10–15 points. The increased firm effect and the reduced contribution of the covariance vary in differing proportions depending on the period. So, whether by estimating an AKM model on the actual or simulated data, it seems that the Danish data are special in that they display a greater idiosyncratic variance and a smaller contribution of worker types.

Finally, unless still biased downward, our estimation of the fixed-effect correlation (between 0.2 and 0.3) is not a large one. The interpretation that we propose in Section 6 is that the correlation between worker and firm effects is a measure of sorting only if workers and firms can be classified along one single dimension, which would then, in turn, correlate with wages. If firms differ along several dimensions, with loading factors that also differ across countries, then fixed-effect correlations may just become meaningless.

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<sup>27</sup>The discrete mixture model is similar to a LASSO penalization method designed to solve the ill-posedness of the latent-variable estimation problem. It can be viewed as a restriction on the number of different values that fixed effects can take. In our framework, it is a lot more than that because it is also a model of employment mobility that allows for more flexible interactions between heterogeneity and time-varying observables.

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