Displaced workers, early leavers, and re-employment wages

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Abstract

When receiving information about an imminent plant closure or mass layoffs, workers search for new jobs. This has been the premise of advance notice legislation, but has been difficult to verify using survey data. In this paper, we lay out a search model that takes explicitly into account the information flow prior to a mass layoff. Using universal wage data files that allow us to identify individuals working with healthy and displacing firms both at the time of displacement as well as any other time period, we test the predictions of the model on re-employment wages. Controlling for worker quality and unobservable firm characteristics, workers leaving a "distressed" firm have higher re-employment wages than workers who stay with the distressed firm until displacement.

JEL CLASSIFICATION: J31 - Wage Level and Structure; J65 - Unemployment Insurance; Severance Pay; Plant Closings, J63 - Turnover; Vacancies; Layoffs

KEYWORDS: Displaced workers, search theory, advance notice, linked firm-worker data sets.

1 Introduction

Displaced workers have been the subject of an extensive literature. The basic stylized facts were established by Jacobson, LaLonde & Sullivan (1993): when compared to continuously employed workers, displaced workers suffer an earnings dip prior to displacement, and recovery from displacement is long and persistent, both in terms of work experience and earnings.[1] Other work has studied the effects of advance knowledge of displacement on the outcomes of displaced workers.[2] These studies point to the unemployment-lowering effect of advance notice (mostly through a reduction in the incidence of unemployment as opposed to shorter unemployment spells), but also to the apparent endogeneity of the provision of advance notice (Fallick 1994, Jones & Kuhn 1995, Ruhm 1992). Firms provide advance notice to workers likely to suffer from prolonged periods of unemployment, although this might be due to a correlation of advance notice with other unobserved characteristics of the firm (Ruhm (1994) for US data, Jones & Kuhn (1995) using Canadian data).

Most of these studies suffer from a distinct data problem. Generally, these studies use the Displaced Worker Supplement (DWS) to the Current Population Survey (CPS). All but the 1984 and 1986 DWS have no information on whether workers left before the layoff date specified in the advance notice received, and thus cannot identify and follow early leavers. Furthermore, since the CPS is a cross-sectional survey, it is not possible to follow workers or their firms for prolonged periods of time. For instance, it is not possible to compare displaced workers to continuously employed workers at the *same* firm in other time periods,[3] and to the best of our knowledge, only Abowd & Finer (1997) have contrasted displaced workers and early leavers at the same firm. Even when it is possible to follow workers over longer periods of time (Storer & Van Audenrode 1998, using Canadian panel data) or to observe multiple workers within the same firm (Jones & Kuhn 1995, using Ontario data), the studies involved typically could not distinguish early leavers from workers present at displacement.

The work on advance notice to displaced workers relies implicitly on search models. In fact, the rationale behind mandatory notice laws in Canada and the US is to give workers a chance to search while on the job, rather than being surprised by displacement and searching from the disadvantaged position of unemployment. However, no formal structural model of search that incorporates features of displacement, including the possibility of leaving the displacing firm prior to a mass layoff, has been proposed and estimated in the literature, potentially missing many behavioral pat-

[1] See Fallick (1996) and Kletzer (1998) for overviews and Abe, Higuchi, Kuhn & Sweetman (forthcoming), Abbring, van den Berg, Gautier, Gijsbert, van Lomwel & Ruhm (forthcoming), and Farber (1999) for more recent analyses involving Canadian and US data

[2] Addison & Portugal (1987), Jones & Kuhn (1995), Ruhm (1992, 1994), Swaim & Podgursky (1990). See also Table 5 in Storer & Van Audenrode (1998).

[3] A notable exception is Jacobson et al. (1993).

terns and linkages that one should be looking for around displacement.

In Section 2, we construct a partial equilibrium search-theoretic model that allows for the possibility of surprise announcements of future displacement. This announcement may be formal in the form of mandatory advance notice, or informal, through information diffusion (the *internal grapevine*) within the company, or even through announcement in the public media without any formal notification of workers. Several of the theoretical implications from this model are tested using a unique new American dataset, described in Section 3. In particular, this dataset allows us to compute in which period a mass layoff occurs, where the definition of a mass layoff is very flexible and does not require administrative reporting or survey-based sampling. Although lacking a worker report on the actual receipt of information, we do observe workers leaving (and entering) the firm prior to the displacement period. Lengermann & Vilhuber (2000), using the same data, report significant changes in the distribution of worker skills in the periods prior to displacement. Both based on that study and on the coverage of legislated advance notice, we infer that these movements are due to increased knowledge of impending layoffs. The dataset also allows us to match workers to both pre- and post-displacement firms, and to follow their earnings path for prolonged periods of time. Section 4 outlines the estimation methods used and reports results, and Section 5 concludes.

2 A Search Model of Displacement

2.1 Model assumptions

A starting point for understanding the labor market transitions of workers in the firms in question is the following partial equilibrium model of search with notice or information of impending displacement. Workers search on the job as well as off the job, in line with most other search models (Mortensen 1986). When unemployed, they receive job offers at rate λ_0, when on the job at rate λ_1. Those that receive acceptable job offers leave current employment or unemployment for the new employer. When searching, workers take the wage offer distribution $F(w)$ as given. The value of non-market time while unemployed is b, and jobs are exogenously dissolved at rate δ_1. The discount rate is denoted by r.

In order to introduce some of the features of displacement into this model, the following assumptions are made. First, at rate η_1, employed workers receive information of impending mass layoff. After receipt of this information, which might be formal notice or informal information gathered by other means (the term “notice” is used without implying any formal notice), all participants expect job destruction to occur at rate $\delta_2 > \delta_1$. Mass layoffs are modeled as being stochastic, so workers do not know the precise moment of mass layoff. This is designed to resemble the large variation in actual notice received by workers across firms (Jones & Kuhn 1995). Layoffs (exits) can and do occur at firms that have not issued notice of a mass layoff.

Second, the fraction of firms in the notice state is given by gamma, and it is assumed that these firms do not participate in the hiring process.[4] Although this is probably empirically true for plant closures, it is not quite true for mass layoffs in which the firm continues operations.[5] However, in this partial equilibrium model, this assumption only affects from where workers expect to receive wage offers. Since the proportion of firms in the notice state at any given point in time is small,[6] this assumption is a close approximation of the true distribution of wage offers, and facilitates the analysis.

Third, when a worker receives notice, there is no downgrading of the wage. Again, this might seem to contradict the empirical evidence of a dip in *earnings* prior to displacement,[7] but this finding is not universally upheld in the data.[8] The actual mechanism behind the dip in earnings is not yet fully understood. It may reflect changes in hours of work at a constant wage rate or selection on early leavers. In future work, this can be relaxed.

Fourth, the distressed state is an absorbing state. A firm, once it has given notice, never reverts to a non-distressed state. This is an assumption that would be relaxed in a general equilibrium model, either by specifying the entry of new firms, or a process describing reversion to a non-distressed state, in order to achieve an equilibrium with positive steady-state employment.

No further constraints are imposed on the model at this stage. In particular, the worker's reservation wage strategies for all four possible transitions (employment - unemployment, notice - unemployment, notice - employment, unemployment - employment) are in no way constrained.

[4]In this, our model differs from Burdett & Mortensen (1980). In their model, jobs are characterized by their permanent or temporary layoff probabilities *ex ante*. Here, all jobs have the same *ex ante* probability of becoming notice jobs, and only differ *ex post*.

[5]Lengermann & Vilhuber (2000) provide empirical evidence of increased hiring activities at firms prior to displacement events.

[6]In our data, in any given quarter, approximately 1.3 percent of firms have a displacement event, see Section 3.

[7]This was first established by Jacobson et al. (1993).

[8]See f.i. Schoeni & Dardia (1996) for an example using data similar to ours.

2.2 Value Functions for Employment and Unemployment

The value of employment in a non-notice firm with wage w is given by

$$\begin{aligned} rV_E(w) &= w + \lambda_1 (1-\gamma) \left[\max\{V_E(w'), V_E(w)\} - V_E(w)\right] \\ &\quad + \delta_1 \left(V_U - V_E(w)\right) + \eta_1 \left[\max\{V_U, V_E^n(w)\} - V_E(w)\right]. \end{aligned} \tag{1}$$

While employed the worker receives wage w. With probability $\lambda_1 (1-\gamma)$ the worker receives an outside wage offer w' from a non-notice firm which she can either accept or reject. With probability δ_1 she is laid off and with probability η_1 she receives notice of an impending mass layoff. Upon receipt of notice she has to decide whether to stay employed at the notice firm or go into unemployment.

The value of employment at a notice firm with wage w is given by

$$\begin{aligned} rV_E^n(w) &= w + \lambda_1 (1-\gamma) \left[\max\{V_E(w'), V_E^n(w)\} - V_E^n(w)\right] \\ &\quad + \delta_2 \left(V_U - V_E^n(w)\right). \end{aligned} \tag{2}$$

Here the worker receives outside offers from non-notice firms and must decide to accept or reject them. He now has a higher chance of ending up in unemployment ($\delta_2 > \delta_1$).

The value of unemployment, V_u, is given by

$$rV_U = b + \lambda_0 (1-\gamma) \left[\max\{V_E(w), V_U\} - V_U\right]. \tag{3}$$

While unemployed, workers have non-market time value b and receive wage offers with probability $\lambda_0 (1-\gamma)$, which they can either accept or reject.

2.3 Reservation Wage Strategies

Under the above setup workers will have four state dependent reservation wage strategies.[9] While employed at a non-notice firm and receiving offers from other non-notice firms, it is well known that the current wage w is the reservation wage, i.e. any wage offer above w is accepted. This is still the case in this model. However, the current wage is likely not the reservation wage for those employed at notice firms contemplating non-notice firm offers. Hence, label $r(w)$ the reservation wage function for those employed at notice firms at wage w such that $V_E(r(w)) = V_E^n(w)$. Label w^* the reservation wage of unemployed workers such that $V_E(w^*) = V_U$. Finally, label

[9] Transitions from the non-notice to the notice state of employment do not occur voluntarily, and transitions from unemployment into the notice state have been excluded. If notice firms were to hire, both of these transitions would have an associated reservation wage strategy.

r^* the reservation wage associated with the transition to unemployment when faced with notice such that $V_E^n(r^*) = V_U$.

Given these reservation wage strategies the above value functions can be rewritten as follows. From (1),

$$\begin{aligned} rV_E(w) &= w + \lambda_1 (1-\gamma) \int_w^{\overline{w}} (V_E(w') - V_E(w))\, dF(w') + \delta_1 (V_U - V_E(w)) \\ &\quad + \eta_1 (V_U - V_E(w)) \end{aligned} \tag{4}$$

if $w \leq r$,

$$\begin{aligned} rV_E(w) &= w + \lambda_1 (1-\gamma) \int_w^{\overline{w}} (V_E(w') - V_E(w))\, dF(w') + \delta_1 (V_U - V_E(w)) \\ &\quad + \eta_1 (V_E^n(w) - V_E(w)) \end{aligned} \tag{5}$$

if $w > r$. From (2),

$$\begin{aligned} rV_E^n(w) &= w + \lambda_1 (1-\gamma) \int_{r(w)}^{\overline{w}} (V_E(w') - V_E^n(w))\, dF(w') \\ &\quad + \delta_2 (V_U - V_E^n(w)), \end{aligned} \tag{6}$$

$$\tag{7}$$

and from (3),

$$rV_U = b + \lambda_0 (1-\gamma) \int_{w^*}^{\overline{w}} (V_E(w) - V_U)\, dF(w). \tag{8}$$

where $\overline{w}$ is the highest wage offered.

To solve for w^* and r^* set $V_E(w^*) = V_U$ and $V_E^n(r^*) = V_U$, respectively, under the conjecture that $w^* \leq r^*$. This is the most reasonable conjecture since the expectation is that the non-notice jobs are more attractive than the notice jobs and therefore one is more picky about keeping a notice job, i.e. $V_E(w) \geq V_E^n(w)$. Note that by the definition of $r(w)$, $r(r^*) = w^*$. Solving for w^* and r^* yields.

$$w^* = b + (\lambda_0 - \lambda_1)(1-\gamma) \int_{w^*}^{\overline{w}} (V_E(w') - V_E(w^*))\, dF(w') \tag{9}$$

$$r^* = b + (\lambda_0 - \lambda_1)(1-\gamma) \int_{w^*}^{\overline{w}} (V_E(w') - V_E^n(r^*))\, dF(w') \tag{10}$$

Since $V_E(w^*) = V_E^n(r^*) = V_U$ by definition, the formulas for w^* and r^* are the same and therefore $w^* = r^*$. Thus we have the first result. At the time of notice workers always opt to stay employed; there is no voluntary exit to unemployment to search for another non-notice job.

We now turn to solving for $r(w)$, the reservation wage while employed at a notice firm. The conjecture here is that $r(w)$ will be less than w. That is, workers will accept a lower wage at a non-notice firm in order to escape the higher likelihood of unemployment. To solve for $r(w)$ we set $V_E(r(w)) = V_E^n(w)$. This yields

$$\begin{aligned}
r(w) &= w + (\delta_1 - \delta_2)\left(V_E^n(w) - V_U\right) + \eta_1 \left(V_E(r(w)) - V_E^n(r(w))\right) \\
r(w) &= w + (\delta_1 - \delta_2)\left(V_E^n(w) - V_U\right) + \eta_1 \left(V_E^n(w) - V_E^n(r(w))\right)
\end{aligned}$$

$$\begin{aligned}
r(w) = {} & w + \frac{\delta_1 - \delta_2}{r + \delta_2 + \lambda_1 (1-\gamma)(1 - F(r(w)))} \times \\
& \left[w - w^* - \lambda_1 (1-\gamma) \int_{w^*}^{r(w)} \left(V_E(w') - V_U\right) dF(w') \right] \\
& + \frac{\eta_1}{r + \delta_2 + \lambda_1 (1-\gamma)(1 - F(r(w)))} \times \\
& \left[w - r(w) - \lambda_1 (1-\gamma) \int_{r(r(w))}^{r(w)} \left(V_E(w') - V_E^n(r(w))\right) dF(w') \right]. \qquad (11)
\end{aligned}$$

To show that $r(w) < w$ we rearrange the above expression for $r(w)$.

$$\begin{aligned}
& \left(r + \delta_2 + \lambda_1 (1-\gamma)(1 - F(r(w))) + \eta_1\right)(r(w) - w) \\
= {} & (\delta_1 - \delta_2) \left[w - w^* - \lambda_1 (1-\gamma) \int_{w^*}^{r(w)} \left(V_E(w') - V_U\right) dF(w') \right] \\
& - \eta_1 \lambda_1 (1-\gamma) \int_{r(r(w))}^{r(w)} \left(V_E(w') - V_E^n(r(w))\right) dF(w')
\end{aligned}$$

The term on the left hand side that is multiplied by $(r(w) - w)$ is positive because γ and $F(r(w))$ are less than or equal to 1. The first term on the right hand side is negative because $\delta_1 < \delta_2$ and $V_E^n(w) > V_U$ for all $w > w^*$ (see first line of $r(w)$ expression). The second term on the right hand side is positive because the expression in the integral is positive over the integrated wage range, i.e. $V_E(w') \geq V_E^n(r(w))$ for $r(r(w)) \leq w \leq r(w)$. Therefore $r(w) - w$ must be negative or $r(w) < w$.

The equalization of values at the reservation wage $(V_E(w^*) = V_E^n(w^*))$ is a surprising finding given the intuition about the value of non-notice jobs being higher than the value of notice jobs. However, we will see that this holds only at the reservation wage and otherwise the intuition follows through. To show that $V_E(w) > V_E^n(w)$ for $w > w^*$ we subtract $V_E^n(w)$ from $V_E(w)$. After rearranging we have

$$\begin{aligned}
& \left(r + \eta_1 + \lambda_1 (1-\gamma)(1 - F(r(w))) + \delta_1\right)\left(V_E(w) - V_E^n(w)\right) \\
= {} & \lambda_1 (1-\gamma) \int_{r(w)}^{w} \left(V_E(w) - V_E(w')\right) dF(w') + (\delta_2 - \delta_1)\left(V_E^n(w) - V_U\right).
\end{aligned}$$

The first term on the left hand side is positive because γ and $F(r(w))$ are less than or equal to 1. The first term on the right hand side is positive because $V_E(w) - V_E(w') > 0$ since $r(w) < w$ and therefore $w > w'$. The second term on the right hand side is positive because by assumption $\delta_2 > \delta_1$ and $V_E^n(w) > V_U$ since $w > w^*$. Therefore $V_E(w) > V_E^n(w)$.

2.4 Predictions

The model as outlined generates a number of predictions. First, there are differences in observed re-employment wages among the three groups of workers - workers leaving non-notice firms, workers leaving notice firms, and workers laid off by notice firms. The reservation strategies at notice and non-notice firms immediately imply that accepted wage offers are lower for workers in the second group, conditional on wages at the old firm. Furthermore, once displaced, displaced workers are indistinguishable from other unemployed, who follow a reservation wage strategy defined by w^*. Since $w^* < r(w) < w$ for all $w > w^*$, it then follows that workers at notice firms who separate prior to displacement will on average have higher accepted wages than displaced workers. Thus, conditional on pre-separation wages, the average wage gains observed in the data should decline monotonically across the three groups. This is the primary prediction that will be tested in this paper.

Second, there are no voluntary exits at time of notice, but the quitting likelihood increases at notice firms. Since the layoff decision by the firm is assumed to be exogenous, this implies that the overall separation likelihood also increases at notice firms. Some support for this prediction was found by Lengermann & Vilhuber (2000), who reported that for some skill groups separations increased above the firm-specific mean separation rate up to four quarters before a mass layoff.

3 Data

The data used here were extracted from the Longitudinal Employer and Household Dynamics (LEHD) Program database. The database contains, among other data sources, unemployment insurance (UI) records for several U.S. states covering the 1990s. UI records contain quarterly earnings on all workers covered by the unemployment insurance system[10] in a given state, all linked to their respective employers. One can thus build a precise picture of the sequencing of employment in conjunction with earnings at each job. The LEHD database augments the UI records with basic demographic information (education, age, race, and sex).[11] Experience is calculated as potential experience at observed entry into the data, and updated with actual observed experience at subsequent points in time.

The data set shares a number of advantages as well as a few disadvantages with

[10] Only a small fraction of workers in jobs not subject to state employment taxes are missed. This includes Federal employees, self-employed individuals, and employees of small agricultural enterprises, and philanthropic or religious organizations. Individuals who receive no salary, who are completely dependent on commissions, and who work with no fixed location or home base are also excluded.

[11] Education is known for a subsample of the population, and multiply imputed for the rest. In this paper, only one imputation was used. Age, race, and sex are known for all respondents.

previous work on displaced workers using unemployment insurance records. On the positive side, it provides a very large sample of displaced workers whose earnings can be tracked over long periods of time both before and after displacement. Furthermore, information on firm employment changes as well as individual earnings should be relatively free of measurement error. On the minus side, our analysis is limited to a handful of states, demographic information is not as extensive as in the typical survey, and layoffs cannot be distinguished from quits. Finally, UI records, because they (typically) contain no information on hours worked, do not allow for the distinction between full- and part-time work.

A "firm" in our empirical work refers to the UI reporting unit, i.e. the account attributed to a firm by the state agencies responsible for UI taxes. Such an account number may cover multiple establishments, however, more than 90% of accounts are known to be single-establishment entities.

Crucial to the analysis is the identification of a displacement "event". The data set contains information on all movements in and out of firms, but no administrative or survey reports of displacement (as would be contained in the Mass Layoff Statistics). One of the strengths of the data is that the sensitivity of the result to the definition of displacement can be explored. In the analysis presented here, a "displacement" is deemed to occur when observed job separations surpass 30 % of maximum firm employment (Jacobson et al. 1993). Average employment is required to be larger than 50 workers. In order to properly capture the element of surprise notice postulated by the theoretical model, we restrict our analysis to firms that are observed to have only one displacement event.[12,13] The displacement event for any given worker is identified not from a survey report, but from observed movements out of employment at that worker's firm. Thus, issues of recall bias or multiple displacement that have plagued the Displaced Worker Supplements (DWS) to the Current Population Survey (Farber 1998) are not of relevance here.

Table 1 on page 21 compares statistics based on the displacement measure in our data with the DWS. Farber (2001) tabulates multiple years of data drawn from the DWS, corrected to be consistent as survey questions changed over time. Since the DWS is retrospective survey querying (in this tabulation) about job displacements in the three years before the survey date, we adjusted our data to give a similar picture.

[12] 34.9 percent of all firms having at least one displacement event have multiple displacement events. Inspection of the data reveals that a large fraction look like temporary layoffs of more than one quarter in length; however, very cyclical firms will appear to have multiple "displacements" in the data. The restriction used here is designed to eliminate these cyclical layoff patterns.

[13] Temporary layoffs of less than three months length are difficult to observe in the data, because of low frequency of the data. A worker being laid off sometime in Quarter 1, and recalled sometime in Quarter 2, potentially up to one day less than 6 months later, will nevertheless appear to be continually employed in the UI wage records, albeit with lower earnings, since positive earnings appear in every quarter. The extreme case of a firm laying off its entire workforce on January 2 and rehiring every single one of its former employees on June 29 will be invisible to the algorithm.

Thus, we computed for every worker in our sample whether he or she experienced at least one displacement within a three-year rolling window. This is approximately equivalent to a DWS-like question asked of these respondents at the end of the third calendar year. The higher frequency of sampling possible in the LEHD data allows for a more detailed analysis than the DWS data.[14] When comparing the equivalent three-year reference periods, only three years are common to both data sets. Excluding the 1993 numbers as being unreliable, the numbers in 1995 and 1997 are comparable, though slightly higher than in the LEHD data set, suggesting that our definition of displacement is close to what workers in the DWS understand by displacement. Furthermore, the LEHD data do not show the strong downward trend apparent in the DWS since 1995. The difference in the two measures is presumably due to both definitional issues on the LEHD data, and recall bias on the DWS data, and is being investigated further.

In all, 5 227 firms had displacement events as defined above during the 1990-1998 period, out of a total of 15 560 firms satisfying the size requirement. Although the ratio of firms ever experiencing displacement seems high, note that this corresponds to a probability of less than 1.3 percent of any given company having a displacement in any given period. Within a twelve quarter window leading up to displacement, slightly more than 3 million workers worked for these firms for at least one quarter.

We construct a sample designed to address some of the sample selection and data quality issues. First, only earnings from "full-quarter employment" quarters are used. Under the full quarter assumption, a worker is counted as working for a firm for the entire period t if and only if she appears at the same firm in periods $t-1$ and $t+1$. This is designed to correct for the problem of unobservable hours.

Second, we select individuals who were in "full-quarter employment" four quarters before the displacement event, continually employed until the separation, and who were in "full-quarter employment" four quarters after separating from the displacing firm (which means they found new employment or were recalled to their old job by the third post-separation quarter). This sample will not include the typical displaced worker who experiences a long unemployment spell.[15] Rather, it will include those displaced workers who, like the early leavers, found a job fairly quickly. Other variables (such as experience or the size of the firm) are taken from the beginning of the pre-displacement period or the end of the post-displacement period here under consideration, as appropriate. Unemployment duration is computed between the quarter separation occurred and the first quarter of observed positive earnings with any firm.

[14] A data quality issue seems to be at the root of the very high displacement rates in 1991 and 1992. Observations from years before 1993 are excluded from the analysis. Future updates to the database will hopefully resolve this problem.

[15] Mean unemployment is around 27 weeks in the CPS (Ruhm 1992).

The resulting sample contains data on approximately 30 000 men who are present in all periods 3 to 5 quarters before a firm's displacement event, as well as working in periods 3 to 5 after having left the displacing firm, either as an "early leaver" or as a displaced worker.[16]

4 Estimation and Results

The base wage equation is an expanded version of the generic displaced worker regression (Jacobson et al. 1993). Let $T_{displ,j}$ denote the displacement date of employer j. Let $T_{depart,i,j}$ denote worker i's separation date from employer j. Finally, let $J(i,t)$ be the function identifying worker i's employer at time t.

The effect of displacement on the wages of workers prior to displacement is captured by

$$DJ'_{J(i,t)}\mu_1 = \sum_{-m\leq\tau\leq 0} DJ^{\tau}_{J(i,t)}\mu_\tau, \tag{12}$$

where $DJ^{\tau}_{J(i,t)}$ is unity if displacement will occur in $-\tau$ periods at the worker i's current employing firm $J(i,t)$ (i.e. $t - T_{displ,J(i,t)} = \tau$). m denotes how many periods in advance this vector of dummies is started. For instance, μ_{-1} measures the effect of next period's displacement on the present period's earnings.

The pre-displacement dummies are specific to a firm and likely apply to that firm's entire workforce, whether or not any particular member of that workforce is actually displaced at $T_{displ,j}$. In particular, workers leaving at some time $T_{depart,i,j} < T_{displ,j}$, whom we will call "early leavers", are likely to experience similar wage changes as "displaced workers" in the stricter sense ($T_{depart,i,j} = T_{displ,j}$), up to the time of departure from the firm.

On the other hand, the post-displacement effects on wages are worker specific, independent of the firm that they work at after separation or displacement. The effect of person-specific post-displacement dummies DI_{it} can be constructed in a similar fashion as the pre-displacement dummies:

$$DI'_{it}\mu_2 = \sum_{0<\tau\leq m} DI^{\tau}_{it}\mu_\tau, \tag{13}$$

where DI^{τ}_{i} is unity if a worker left a displacing firm τ periods ago (i.e. for some j, $t - T_{depart,i,j} = \tau$ and $m > T_{displ,j} - T_{depart,i,j}$). For instance, μ_4 measures the effect of having worked at a displacing firm one year ago on this period's earnings. The notation here corresponds to that in Jacobson et al. (1993) for workers with $T_{displ,j} = T_{depart,i,j}$. However, the post-displacement dummies are person-specific, and are a function of the worker's employment history.

[16] More details on the construction of the data set are available in Appendix A on page 28.

Both Equations (12) and (13) assume that the earnings patterns related to displacement are the same for early leavers and displaced workers. This assumption can be relaxed. Let $e_{i,t,J(i,t)} = 1\left(-m \leq T_{depart,i,j} - T_{displ,j} < 0\right)$ flag early leavers from firm j in period t. A more general specification allowing for variation in the earnings patterns is

$$\begin{aligned} DJ'_{J(i,t)}\mu_1 = \quad & DJ'_{J(i,t)}\mu_{11} * \left(1 - e_{i,t,J(i,t)}\right) \quad + \\ & DJ'_{J(i,t)}\mu_{12} * e_{i,t,J(i,t)} \end{aligned} \tag{14}$$

$$\begin{aligned} DI_i{}'\mu_2 = \quad & DI_i{}'\mu_{21} * \left(1 - e_{i,t,J(i,t)}\right) \quad + \\ & DI_i{}'\mu_{22} * e_{i,t,J(i,t)} \end{aligned} \tag{15}$$

Assembling all the elements defined above yields the basic wage specification:

$$w_{it} = X_{it}\beta + DJ'_{J(i,t)}\mu_1 + DI'_{i,t}\mu_2 + \theta_i + \psi_{J(i,t)} + \varepsilon_{it} \tag{16}$$

where w_{it} measures log earnings for individual i at time t, X_{it} are individual characteristics, both time-varying and time-invariant, θ_i measures the effect of time-invariant individual characteristics (“worker quality”), $\psi_{J(i,t)}$ is a firm-specific (productivity) effect on wages. ε_{it} is a statistical residual, uncorrelated with all the right hand side variables.[17] In our data, X_{it} includes a quadratic in experience, education, race, and year.

In this paper, we consider persons who worked for the same firm during the same time period prior to a displacement event, and who either left early (within two quarters of the displacement quarter) or who were displaced. Their pre-displacement earnings a year before the displacement event are then compared to earnings a full year after separation from their previous firm. In the case of early leavers, this is computed not from the date the firm displaced its remaining workers, but from the date they left the firm. In terms of the above defined variables, all these individuals satisfy $DJ^{-4}_{J(i,-4)} = 1$, and we compare post-separation earnings for the quarter in which $DI^{post}_{i,post} := DI^{+4}_{i,+4} = 1$ with pre-separation earnings from the quarter in which $DJ^{pre}_{J(i,pre)} := DJ^{-4}_{J(i)} = 1$. Differencing (16) obtains

$$\begin{aligned} w_{i,post} - w_{i,pre} \quad = \quad & \left(X_{i,post} - X_{i,pre}\right)\beta \\ + \quad & \left(DI^{post}_{i,post}\mu_{21} - DJ^{pre}_{J(i,pre)}\mu_1\right) \\ + \quad & DI^{post}_{i,post}e_{i,t,J(i,t)}\left(\mu_{22} - \mu_{21}\right) \\ + \quad & \psi_{J(i,post)} - \psi_{J(i,pre)} + \Delta\varepsilon_i \end{aligned} \tag{17}$$

where we have assumed that pre-displacement wage paths are identical within the displacing firm for both early leavers and displaced workers ($\mu_{12} = \mu_{11} = \mu_1$ in (14)). Rewriting,

$$\begin{aligned} w_{i,post} - w_{i,pre} \quad = \quad & \alpha + \Delta X_i\beta \\ + \quad & e_{i,t,J(i,t)}\Delta\mu + \psi_{J(i,post)} - \psi_{J(i,pre)} + \tilde{\varepsilon}_i \end{aligned} \tag{18}$$

[17] See Abowd & Kramarz (1999) for a more detailed description of this model.

where $\alpha = \mu_{21} - \mu_1$ is the component for all workers finding employment post-displacement, $\Delta X_i = X_{i,post} - X_{i,pre}$ captures any changes in time-varying observables,[18] and $\Delta\mu = \mu_{22} - \mu_{21}$ is the difference in post-displacement earnings due solely to the fact that some workers left earlier than others, and all the displacement dummies are set to unity due to sample selection. Note in particular that θ no longer plays a role in (18) because of first-differencing, but that ψ still enters for two different firms.

The starting point of this analysis are the first two rows of Table 2. The first row shows the raw earnings differential between the fourth predisplacement quarter and the fourth post-displacement quarter. The second row shows the differential when computed using "full-quarter" earnings, as defined earlier. The difference between the two columns is a first estimate of the parameter of interest $\Delta\mu$, and both difference-in-differences tell the same story: Earnings for early leavers are significantly higher than for workers from the same firms who stay until displaced, by approximately 10 percent if using full-quarter earnings, and by more than thirteen percent when using raw quarterly earnings, consistent with the search model outlined earlier.

The table also reveals marked differences between the groups. Levels of earnings are lower for early leavers, as is their age, experience, and education. Early leavers leave smaller firms, but also move to smaller firms. The racial composition is also more diverse among early leavers. There are small differences in the estimated person fixed effect, a measure of long-term earnings potential, but both groups are quite close to the population average of zero. However, there are larger differences in the fixed effect of firms for which they work before and after separation, a measure of pay policy differences. The more seasoned displaced workers separate from higher-paying firms than the early leavers, and also find new jobs in such firms, but the early leavers experience a larger improvement. This finding is again consistent with the search model.

The search model implies that at least for early leavers, transitions occur directly from one job to the next with no intervening unemployment. Empirically, we observe positive unemployment spells for both groups of workers, as well as substantial recall for displaced workers. To more closely approximate the requirements of the model, we restrict our sample further to those experiencing at most one period of unemployment, and tabulate the characteristics of this subsample in Table 3 on page 23. Most workers in both categories still experience some unemployment. The FQ difference-in-difference increases to nearly 18 percent, and the raw difference-in-difference to over 22 percent. But other characteristics also change. The difference in θ is now more marked, and reversed in favor of early leavers. Both types of separators now work in more similarly sized firms before and after separation. The fraction of temporary layoffs falls dramatically. On the other hand, the fixed effects of pre- and post-separation firms are quite similar to the full sample, although both groups make larger improvements. The fraction of industry stayers remains essentially unchanged

[18] In practice, the data we use has few or no time-varying observables.

for early leavers, but decreases slightly for displaced workers, despite the elimination of most recalls.

Many of the differences noted in Tables 2 and 3 are correlated with wage levels, and in the further analysis, we will use regressions based on Equation (18) to disentangle the determinants of wage levels from the more basic implication of the search model, namely that the wage difference is due to the fact that the early leavers received a better draw than the displaced workers. [19]

Table 4 on page 24 presents results from a series of OLS specifications using the full sample described by Table 2. Column (1) builds on the basic difference-in-difference (DID) comparison done in the first row of Table 2 by controlling for a number of person-specific observables as well as the time difference between the two wage measurements. Although in particular the experience and education variables have a significant impact on the log wage differential, the estimate of $\Delta\mu$ remains virtually unchanged from the naive DID.

Column (2) estimates Equation (18) using independently estimated firm fixed effects. The estimation of these firm fixed effects is described in Appendix A on page 28. In this specification, a firm will have the same fixed effect whether it is a displacing firm, a new firm, or both.[20] Controlling for estimated firm fixed effects reduces the difference due to early leavers, but the latter still remains economically and statistically significant.[21] Column (3) augments Equation (18) with the person-specific characteristics, which slightly increases $\Delta\mu$. The controls included in the base specification do not alter the conclusion from the raw data: Early leavers earn more than displaced workers with similar characteristics.

It could be argued that among early leavers, workers with higher search intensities (represented in the search model by λ_0 and λ_1) are over-represented, and a control for the unobserved ability associated with higher search intensity or success should be controlled for. Since a higher search intensity leads to faster wage growth, omission of such a variable would lead to an overestimate of $\Delta\mu$.

No direct measure of search intensity is available in the data. On the other hand, covariates such as experience may already control for differences in search intensity,

[19]The other major implication, the difference between workers of non-notice firms who change jobs and early leavers from notice firms, will be tested in a later revision of this paper.

[20]Of the 5,343 new firms on the full analysis data set, 1,525 are also displacing firms (51.8 % of all displacing firms). When restricting the sample to workers with no more than one quarter of intervening unemployment, which eliminates a large fraction of recalls, only 566 are also displacing firms (26.7 % of all displacing firms).

[21]Note that the coefficient on $\hat{\psi}_{J(i,pre)}$ according to the Equation (18) should be negative unity, and that on $\hat{\psi}_{J(i,post)}$ should be positive unity. The significant deviation from that value might indicate of a time-varying component to firm pay policies not well captured by the base regression used to estimate these variables.

since younger workers might have higher search intensities. In all columns, experience is negatively linked to earnings growth, consistent with the usual perception of a concave experience profile. Column (4) introduces a different control. We include the estimated person fixed effect $\hat{\theta}$. If higher average earnings are a product of faster wage growth, then $\hat{\theta}$ is correlated with a higher search intensity. Table 4 shows that $\hat{\theta}$ is positively correlated with the wage gap for all workers.[22] However, conditional on $\hat{\theta}$, the gap between early leavers and displaced workers is actually slightly higher than when $\hat{\theta}$ is not included, suggesting if anything that $\hat{\theta}$ is negatively correlated with search intensity.

One feature of Table 4 is the strong impact of temporary layoffs. While the coefficient on the temporary layoff flag allows to evaluate the strength of the observed wage growth of early leavers - early leavers gain 3 to 4 percent more than workers who returned to their old jobs and presumably pay schedules - it also begs the question of whether or not the base comparison group of displaced workers is appropriately chosen. Over 60 percent of displaced workers eventually return to their jobs in the base sample. However, temporary layoffs also experience much longer unemployment periods, and in Table B.1 on page 31 in the Appendix, we have restricted the analysis to the sample described in Table 3. Among those workers having experienced no more than one quarter of intervening unemployment, only 7 percent of displaced workers return to their pre-displacement jobs. This sample, as previously explained, is also closer to what the theory describes. Among these workers, the naive DID estimator of $\Delta\mu$ is 0.178. The level of the estimated $\Delta\mu = 0.113$ is higher, but the conclusion reached based on Table 4 is unaltered.

Tables 4 and B.1 used estimated firm fixed effects. It is however feasible to explicitly re-estimate fixed effects, separately for displacing and receiving firms.[23] Table 5 on page 25 reports results from an OLS regression which explicitly estimated fixed effects for all 2,942 displacing firms and all 5,343 new firms, based on the full sample.[24] Column (1) corresponds to the basic specification of Equation (18), and compares directly to column (2) of Table 4. Explicitly re-estimating the fixed effects actually results in a higher estimated $\Delta\mu$, which again is very close to the naive DID estimator. The (cumulative) addition of person-specific characteristics in column (2) and $\hat{\theta}$ in column (3) do not alter the estimated coefficient significantly.

Column (4) of Table 5 investigates some of the sources of the large wage gains observed for early leavers by distinguishing between those whose new job was found in

[22] $\hat{\theta}$ is estimated, but we are interested only in the significance of the coefficient on $\hat{\theta}$, $\beta_{\hat{\theta}}$, since we have no priors as to what value it should have. For the hypothesis test $\beta_{\hat{\theta}} = 0$, the OLS standard errors are consistent and t-statistics valid (Pagan 1984).

[23] Contrary to the previous regressions, a firm that is at once a displacing firm for some workers and a new firm for others will have two different fixed effects, depending on the role it plays towards each worker.

[24] Table B.2 on page 32 reports results for the same specifications based on the sample with restricted unemployment. The conclusions remain the same.

the same industry as the displacing job, and those who switched industries. Tables 2 and 3 reported that the fraction of early leavers who stayed within the same industry when switching jobs was signficantly lower than for displaced workers, though a majority did not change industries. The coefficients reported in column (4) indicate that much of the wage gain is due to individuals who change industries. Nevertheles, even those who find new jobs in the same industry make higher wage gains than do displaced workers who stay in the same industry.

To further explore the correlation between observable characteristics, $\hat{\theta}$, and the likelihood of being an early leaver, Table 6 on page 26 presents results for the likelihood of leaving a displacing firm early from a probit specification. Column (1) includes demographics and year dummies. Workers who leave the firm early tend to be less educated, and tend to come from smaller firms. Conditional on the other variables, experience is linked to a positive and convex likelihood of leaving early, contrary to the difference in unconditional sample means reported in Table 2 on page 22, and in contrast to the earlier hypothesis that experience is negatively correlated with a higher search intensity.[25] Racial background also seems to matter, with blacks and hispanics more likely to leave than whites and other racial groups.

The addition of $\hat{\psi}_{J(i,pre)}$ in column (2) and of $\hat{\psi}_{J(i,post)}$ does not alter the effect of other variables. Both estimated firm fixed effects are negatively correlated with the likelihood of departure, reflecting the observations from the different sample means. Finally, the addition of the person fixed effect in column (4) does reduce the impact of racial indicators, suggesting different population averages of long-term wage levels across these groups. However, the effect of $\hat{\theta}$ is *negatively* correlated with the likelihood of leaving, contrary to what should be expected if $\hat{\theta}$ were correlated with a higher search intensity, but consistent with the OLS results. The general picture, in particular the effect of experience and education, is not affected.

Even though the probit results reinforce the previous OLS results, they do indicate that the controls included in the specifications from Table 4 may not fully control for heterogeneity in the sample. The primary focus of the search model is on homogeneous workers. One way of introducting heterogeneity into the model is through segmented labor markets. Workers are homogeneous within, but not across labor markets.

Table 7 reports results from separating the restricted analysis sample into four sub-groups, based on the population distribution of $\hat{\theta}$, in order to obtain homogeneous sub-populations. Each column is equivalent to the specification used in column (4) from Table 4, when estimating OLS for each of the subgroups separately. Even though the groups are selected within quartiles of the distribution of $\hat{\theta}$, $\hat{\theta}$ is still included in the regression as a control.

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[25]The sample includes only persons with more than 20 quarters of labor market experience. Over the range of feasible values in the data, the experience terms are positive.

The parameter of interest $\Delta\mu$ is significant in all but the lowest quartile, and is of the same order of magnitude as for the sample as a whole, but varies significantly across the distribution of $\hat{\theta}$. Also, the included $\hat{\theta}$ is significant only for the bottom quartile. The second row of the table shows means and standard errors of $\hat{\theta}$ within each quartile, and not surprisingly, the variation (and range) of $\hat{\theta}$ in the top and bottom quartiles is far larger than in the middle of the distribution. Thus, the workers captured in columns (1) and (4) are far less homogeneous than those workers underlying the regressions in columns (2) and (3).[26]

5 Conclusion

One of the primary concerns of policy makers when faced with mass layoffs is how to quickly return these individuals to work. Mostly, the emphasis has relied on mandatory advance notice laws, but their efficacy has only circumstantially been proven. Firms, on the other hand, might worry about destructive attrition prior to displacement. In particular, if the mass layoff was the result of a plant closure which the firm had deemed avoidable, then attrition might have been detrimental to the rescue attempt.

In this paper, we provide some evidence that a solution to these competing incentives is non-trivial. We lay out a search model that incorporates aspects of displacement, in particular the receipt of information as to the viability of a worker's job (which here is interpreted be related to a mass layoff). Workers endogenously adapt their reservation wages to changed circumstances. The model predicts that workers who have received "notice" of a higher job failure risk, will adjust their reservation wages downwards. This implies that their departure from the firm is more likely than if they had not received this information, and furthermore that their re-employment wages will lie below normal re-employment wages, but above the wages obtained by displaced and other unemployed job seekers.

The data, obtained from US universal wage record data, support these conclusions. The data is used to determine when mass layoffs occur, and then compares those workers who left up to 2 quarters prior to the mass layoff with workers displaced at the time of the mass layoff. The results indicate that within categories of homogenous workers, and controlling for characteristics of workers and displacing firms, early leavers consistently obtain higher re-employment wages than displaced workers, except for workers at the low end of the labor market.

Although the data do not report if these workers had received formal advance

[26] Not reported here, probit estimation by quartiles of $\hat{\theta}$, of Table 6 yielded similar results. Once firm characteristics have been controlled for, $\hat{\theta}$ is no longer significantly related to early exit from the displacing firm except in the lowest quartile.

notice, the results are suggestive of the beneficial effect to workers of advance notice. On the other hand, accelerated attrition is clearly a feature of the model used here.Whether this accelerated attrition is beneficial in a general equilibrium framework, for instance through improved reallocation of workers, remains to be determined in future work.

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6 Tables

Table 1: Comparison of displacement measures

Year	Farber (2001)	LEHD UI records	
	3-year [1]	1-year [2]	3-year [3]
1983	0.128		
1985	0.103		
1987	0.095		
1988			
1989	0.085		
1990		0.043	
1991	0.118	0.120	
1992		0.094	0.252
1993	0.109	0.039	0.252
1994		0.031	0.130
1995	0.115	0.030	0.081
1996		0.034	0.083
1997	0.091	0.032	0.085
1998		0.021	0.076
1999	0.086		

Notes:

1: Source: Farber (2001), Appendix Table 2b, *Total Three-Year Rate of Job Loss*, defined as "At least one displacement in the past three years, Discounted Other Job loss."

2: Source: LEHD data sources. At least one displacement in the past 4 quarters, as of 31 December. For other data restrictions, consult the text.

3: Source: LEHD data sources. At least one displacement in the past 12 quarters, as of 31 December.

Table 2: Person summary statistics

Variable	Displaced Mean	Displaced (Std Dev)	Early leavers Mean	Early leavers (Std Dev)
Raw Earnings difference	0.089	*(0.344)*	0.224	*(0.876)*
FQ Earnings difference	0.083	*(0.449)*	0.189	*(0.979)*
Person characteristics:				
Education	13.217	*(2.503)*	12.897	*(2.672)*
Race: Black	0.101	*(0.301)*	0.142	*(0.349)*
Race: Hispanic	0.038	*(0.192)*	0.069	*(0.254)*
Race: Other	0.062	*(0.242)*	0.064	*(0.245)*
Age first observed	38.709	*(9.592)*	33.605	*(9.586)*
Total experience (Q)	94.827	*(38.937)*	77.332	*(37.269)*
θ person fixed effect	0.0156	*(0.574)*	-0.005	*(0.601)*
Pre-displacement job:				
Log FQ earnings	9.072	*(0.642)*	8.583	*(0.987)*
FQ earnings	10822.50	*(24432.37)*	8339.97	*(20407.22)*
Number of establishments		*( )*		*( )*
Average employment	4312.18	*(6483.07)*	2094.77	*(4750.87)*
ψ firm fixed effect	0.173	*(0.311)*	0.032	*(0.332)*
Post-displacement job:				
Log FQ earnings	9.156	*(0.672)*	8.771	*(0.934)*
FQ earnings	12113.46	*(29064.07)*	9228.63	*(12452.13)*
Number of establishments	21.075	*(56.535)*	8.950	*(37.711)*
Average employment	4425.03	*(6726.57)*	1949.47	*(5449.03)*
ψ firm fixed effect	0.177	*(0.313)*	0.070	*(0.342)*
Temporary layoffs	0.616	*(0.486)*	0.029	*(0.169)*
Unemployment duration	1.722	*(0.614)*	1.455	*(0.923)*
Industry stayers	0.894	*(0.307)*	0.528	*(0.499)*
Observations	26955		2486	

Source: LEHD data sources, 10 percent random sample, authors' computations. For computation of θ and ψ, see Appendix A.

Table 3: Person summary statistics: restricted unemployment

Variable	Displaced		Early leavers	
	Mean	*(Std Dev)*	Mean	*(Std Dev)*
FQ Earnings difference	0.097	*(0.563)*	0.275	*(0.985)*
Raw Earnings difference	0.097	*(0.451)*	0.321	*(0.873)*
Person characteristics:				
Education	12.981	*(2.560)*	12.940	*(2.658)*
Race: Black	0.097	*(0.296)*	0.132	*(0.338)*
Race: Hispanic	0.053	*(0.225)*	0.066	*(0.249)*
Race: Other	0.067	*(0.250)*	0.070	*(0.255)*
Age first observed	36.981	*(9.896)*	33.269	*(9.604)*
Total experience (Q)	94.973	*(40.421)*	75.879	*(37.131)*
θ person fixed effect	-0.016	*(0.582)*	0.040	*(0.589)*
Pre-displacement job:				
Log FQ earnings	9.008	*(0.718)*	8.566	*(1.018)*
FQ earnings	10842.03	*(38202.04)*	8118.37	*(13298.63)*
Number of establishments		*(*		*( )*
Average employment	1353.13	*(3671.38)*	1865.42	*(4412.53)*
ψ firm fixed effect	0.160	*(0.286*	0.018	*(0.329)*
Post-displacement job:				
Log FQ earnings	9.107	*(0.733)*	8.840	*(0.918)*
FQ earnings	12102.69	*(39990.83)*	9703.76	*(13277.08)*
Number of establishments	10.799	*(28.813)*	8.420	*(38.384)*
Average employment	1749.99	*(4981.06)*	1921.45	*(5511.89)*
ψ firm fixed effect	0.172	*(0.284)*	0.082	*(0.332)*
Temporary layoffs	0.071	*(0.256)*	0.008	*(0.089)*
Unemployment duration	0.982	*(0.130)*	0.955	*(0.207)*
Industry stayers	0.785	*(0.410)*	0.510	*(0.500)*
Observations	9021		1734	

Source: LEHD data sources, 10 percent random sample, authors' computations. For computation of θ and ψ, see Appendix A.

Table 4: OLS results: Log wage differential

	(1)		(2)		(3)		(4)	
Variable	Parameter Estimate	S.E.	Parameter Estimate	S.E.	Parameter Estimate	S.E.	Parameter Estimate	S.E.
Early leaver $\Delta\mu$	**0.107**	*(0.023)*	**0.069**	*(0.019)*	**0.081**	*(0.020)*	**0.086**	*(0.020)*
Intercept	**0.251**	*(0.035)*	**0.027**	*(0.016)*	**0.160**	*(0.037)*	**0.164**	*(0.037)*
Temp layoff	**0.067**	*(0.014)*			**0.061**	*(0.014)*	**0.057**	*(0.014)*
Total experience (Q)	**-0.002**	*(0.000)*			**-0.001**	*(0.000)*	**-0.001**	*(0.000)*
Exp.^{2}/10000	**3.947**	*(1.891)*			1.994	*(1.788)*	0.616	*(1.793)*
Education	**-0.003**	*(0.001)*			**-0.003**	*(0.001)*	**-0.007**	*(0.001)*
Race: Black	-0.013	*(0.010)*			-0.014	*(0.009)*	0.005	*(0.009)*
Race: Hispanic	0.010	*(0.019)*			0.009	*(0.017)*	**0.030**	*(0.017)*
Race: Other	0.002	*(0.011)*			-0.001	*(0.010)*	0.000	*(0.011)*
Theta $\hat{\theta}$							**0.063**	*(0.011)*
Unemployment	**-0.049**	*(0.009)*			**-0.036**	*(0.008)*	**-0.035**	*(0.008)*
Same industry					0.011	*(0.018)*	0.009	*(0.018)*
$\hat{\psi}_{J(i,pre)}$			**-1.262**	*(0.051)*	**-1.245**	*(0.051)*	**-1.229**	*(0.051)*
$\hat{\psi}_{J(i,post)}$			**1.258**	*(0.048)*	**1.246**	*(0.047)*	**1.247**	*(0.047)*
Adj. R^2	0.0141		0.1411		0.1476		0.1511	

NOTE: Dependent variable is the the log difference in "full-quarter" wages (mean = 0.092). All regressions include year of displacement dummies and a control for the time difference between earnings measures. Standard errors are corrected for clustering by 2,110 displacing firms. 29,351 observations. Source: LEHD data sources, 10 percent random sample, authors' computations.

Table 5: OLS results: Log wage differential, full fixed effects

	(1)		(2)		(3)		(4)	
Variable	Parameter Estimate	S.E.	Parameter Estimate	S.E.	Parameter Estimate	S.E.	Parameter Estimate	S.E.
Early leaver $\Delta\mu$	**0.108**	*(0.011)*	**0.105**	*(0.012)*	**0.110**	*(0.012)*		
(Same Industry)							**0.050**	*(0.015)*
(Industry leaver)							**0.193**	*(0.018)*
Intercept	**0.058**	*(0.011)*	**0.274**	*(0.027)*	**0.275**	*(0.027)*	**0.249**	*(0.027)*
Temp layoff			**0.055**	*(0.008)*	**0.051**	*(0.008)*	**0.044**	*(0.008)*
Total experience (Q)			**-0.002**	*(0.000)*	**-0.001**	*(0.000)*	**-0.001**	*(0.000)*
Exp.^{2}/10000			**3.998**	*(1.835)*	2.130	*(1.837)*	1.982	*(1.836)*
Education			**-0.003**	*(0.001)*	**-0.007**	*(0.001)*	**-0.007**	*(0.001)*
Race: Black			-0.013	*(0.010)*	0.010	*(0.010)*	0.009	*(0.010)*
Race: Hispanic			0.011	*(0.015)*	**0.035**	*(0.015)*	**0.034**	*(0.015)*
Race: Other			0.003	*(0.012)*	0.005	*(0.012)*	0.005	*(0.012)*
Theta $\hat{\theta}$					**0.074**	*(0.006)*	**0.075**	*(0.006)*
Unemployment			**-0.050**	*(0.006)*	**-0.048**	*(0.006)*	**0.028**	*(0.011)*
Same industry			-0.001	*(0.010)*	-0.003	*(0.010)*	**-0.047**	*(0.006)*
Adj. R^2	0.0034		0.0136		0.0185		0.0198	

NOTE: Dependent variable is the the log difference in "full-quarter" wages (mean = 0.092). All regressions include a control for the time difference between earnings measures; and fixed effects for 2,942 displacing firms and 5,343 new firms. 29,351 observations. Source: LEHD data sources, 10 percent random sample, authors' computations.

Table 6: Probit results: Likelihood of leaving before displacement

	(1)		(2)		(3)		(4)	
Variable	Parameter Estimate	Standard Error	Parameter Estimate	Standard Error	Parameter Estimate	Standard Error	Parameter Estimate	Standard Error
Intercept	**-0.438**	*(0.074)*	-0.004	*(0.159)*	0.004	*(0.239)*	-0.055	*(0.241)*
Total experience (Q)	**-0.013**	*(0.001)*	**-0.010**	*(0.001)*	**-0.007**	*(0.001)*	**-0.009**	*(0.001)*
Exp.^{2}	**0.004**	*(0.001)*	**0.003**	*(0.001)*	**0.002**	*(0.001)*	**0.002**	*(0.001)*
Education	**-0.029**	*(0.004)*	**-0.027**	*(0.005)*	**-0.023**	*(0.005)*	**-0.012**	*(0.005)*
Race: Black	**0.205**	*(0.033)*	**0.206**	*(0.034)*	**0.220**	*(0.037)*	**0.159**	*(0.038)*
Race: Hispanic	**0.180**	*(0.048)*	**0.103**	*(0.049)*	**0.120**	*(0.052)*	0.066	*(0.053)*
Race: Other	0.037	*(0.045)*	0.031	*(0.046)*	0.073	*(0.049)*	0.066	*(0.049)*
Theta $\hat{\theta}$							**-0.176**	*(0.024)*
Same industry					**-0.989**	*(0.029)*	**-0.972**	*(0.029)*
$\hat{\psi}_{J(i,pre)}$			**-0.507**	*(0.045)*	**-0.398**	*(0.064)*	**-0.409**	*(0.064)*
$\hat{\psi}_{J(i,post)}$					**-0.122**	*(0.060)*	**-0.112**	*(0.060)*
Log likelihood	-7954.8284		-7623.5471		-6690.6645		-6664.3343	

NOTE: Probit analysis. Dependent variable is the observed indicator for early leavers, $e_{i,t,J(i,t)}$. All regressions include year of displacement dummies; columns (2), (3), and (4) an additional 9 displacing industry dummies and 4 firm size dummies. Column (3) and (4) also include an additional 9 industry dummies and 4 firm size dummies associated with the new firm. 29,441 observations. Source: LEHD data sources, 10 percent random sample, authors' computations.

Table 7: OLS results: by theta quartiles, restricted unemployment

Variable	(1) Quartile 1		(2) Quartile 2		(3) Quartile 3		(4) Quartile 4	
	Parameter Estimate	Standard Error .	Parameter Estimate	Standard Error .	Parameter Estimate	Standard Error .	Parameter Estimate	Standard Error .
Mean of								
flag early leaver	0.096	*(0.295)*	0.079	*(0.270)*	0.083	*(0.276)*	0.086	*(0.280)*
Theta θ	-0.833	*(0.325)*	-0.262	*(0.150)*	0.202	*(0.134)*	0.833	*(0.378)*
Early leaver $\tilde{\gamma}$	0.072	*(0.079)*	**0.071**	*(0.038)*	**0.208**	*(0.045)*	**0.085**	*(0.044)*
Intercept	0.575	*(0.489)*	0.111	*(0.143)*	-0.120	*(0.144)*	0.158	*(0.230)*
Temp layoff	0.042	*(0.067)*	**0.101**	*(0.045)*	0.026	*(0.049)*	0.048	*(0.074)*
Total experience (Q)	0.003	*(0.003)*	**-0.002**	*(0.001)*	0.001	*(0.001)*	0.002	*(0.002)*
Exp.^{2}/10000	-13.446	*(12.313)*	8.283	*(6.055)*	-7.294	*(6.471)*	**-25.245**	*(14.584)*
Education	-0.001	*(0.006)*	-0.003	*(0.004)*	-0.002	*(0.004)*	**-0.011**	*(0.006)*
Race: Black	-0.044	*(0.047)*	**-0.047**	*(0.026)*	0.019	*(0.036)*	0.154	*(0.100)*
Race: Hispanic	0.006	*(0.057)*	0.017	*(0.036)*	-0.031	*(0.042)*	0.041	*(0.106)*
Race: Other	0.037	*(0.051)*	0.004	*(0.035)*	-0.012	*(0.033)*	**-0.090**	*(0.039)*
Theta θ	**0.253**	*(0.100)*	0.009	*(0.059)*	-0.034	*(0.073)*	0.061	*(0.061)*
Unemployment	0.082	*(0.095)*	0.052	*(0.075)*	0.094	*(0.098)*	-0.105	*(0.137)*
Same industry	**0.131**	*(0.060)*	-0.048	*(0.029)*	-0.032	*(0.033)*	-0.064	*(0.045)*
$\psi_{J(i,pre)}$	**-1.310**	*(0.148)*	**-1.113**	*(0.091)*	**-1.191**	*(0.108)*	**-0.988**	*(0.156)*
$\psi_{J(i,post)}$	**1.387**	*(0.163)*	**1.178**	*(0.090)*	**1.082**	*(0.091)*	**0.966**	*(0.143)*
Adj. R^2	0.1923		0.2340		0.2702		0.1439	
Observations	1717		3844		3073		2068	
Displacing firms	717		1172		1162		924	

NOTE: Dependent variable is the the log difference in "full-quarter" wages. All regressions include year of displacement dummies and a control for the time difference between earnings measures; Standard errors are corrected for clustering by displacing firm. Source: LEHD data sources, 10 percent random sample, restricted unemployment, authors' computations.

A Data construction

The full LEHD data base as used for this research has 219,414,147 observations for 17,381,486 unique workers and z unique firms. To construct the sample of firms, all quarters in which firms satisfied the displacement definition were retained. Only firms having a single displacement were retained. 34.9 percent of all firms having at least one displacement event have multiple displacement events. Inspection of the data reveals that a large fraction look like temporary layoffs of more than one quarter in length; however, very cyclical firms will appear to have multiple "displacements" in the data. The restriction used here is designed to eliminate these cyclical layoff patterns.

All workers having worked within a three year window around the displacement were extracted. Only workers having experienced no more than 4 displacements were retained, eliminating about 0.19 percent of all workers (slightly more than 6 thousand individuals).

To speed up analysis, a random 10 percent sample of people was taken, yielding a data set with 309 706 workers and 8 391 992 quarterly wage observations. We further restrict the sample to the men with more than 5 years of labor market experience, leaving 3 562 101 observations for 133 998 workers. This constitutes our basic analysis sample.

Inspection of the data revealed data quality issues in 1991-1992, generating an artificially high displacement rate in these years. In the analysis, only workers displaced in years 1993-1997 are included.

For some of the analysis, we use person- and firm-specific productivity factors, $\hat{\theta}$ and $\hat{\psi}$. These are computed from the full LEHD data base using OLS based on Equation (16) on page 12 with all displacement dummies set to zero. Dependent variable is full-time-equivalent FQ earnings, where an adjustment has been made based on an individual's imputed full-time or part-time status. Mean $\hat{\theta}$ is normalized to zero in the population of workers, weighted by the number of wage observations for each workers. $\hat{\psi}$ is set to zero for one arbitrary firm. Its mean is restricted to be zero across all wage observations. See Abowd & Kramarz (1999) for a more detailed explanation of the estimation procedures used in this step of the data preparation.

Employment used to select firms is computed as a moving two-period average of all workers appearing during one quarter. This is then averaged over the entire period a firm appears in the data to obtain average employment.

Temporary layoffs are defined as workers whose first job after separation is with the same firm as the displacing firm. Industry status is defined at the division level, and industry stayers are those whose firm is in the same division as the displacing firm.

B Appendix tables

Table B.1: OLS results: Log wage differential, restricted unemployment

Variable	(1) Parameter Estimate	(1) S.E.	(2) Parameter Estimate	(2) S.E.	(3) Parameter Estimate	(3) S.E.	(4) Parameter Estimate	(4) S.E.
Early leaver $\Delta\mu$	**0.171**	*(0.027)*	**0.124**	*(0.022)*	**0.111**	*(0.023)*	**0.113**	*(0.023)*
Intercept	0.120	*(0.075)*	0.021	*(0.027)*	0.076	*(0.072)*	0.080	*(0.072)*
Temp layoff	**0.061**	*(0.031)*			**0.059**	*(0.028)*	**0.061**	*(0.028)*
Total experience (Q)	**-0.002**	*(0.001)*			**-0.001**	*(0.001)*	-0.000	*(0.001)*
Exp.^{2}/10000	4.217	*(3.479)*			0.373	*(3.222)*	-1.194	*(3.218)*
Education	0.001	*(0.003)*			0.001	*(0.003)*	-0.003	*(0.003)*
Race: Black	-0.002	*(0.022)*			-0.027	*(0.021)*	-0.010	*(0.021)*
Race: Hispanic	-0.009	*(0.027)*			-0.002	*(0.025)*	0.016	*(0.025)*
Race: Other	-0.008	*(0.022)*			-0.011	*(0.019)*	-0.008	*(0.019)*
Theta $\hat{\theta}$					**0.054**	*(0.018)*		
Unemployment	0.044	*(0.057)*			0.034	*(0.050)*	0.031	*(0.049)*
Same industry					-0.009	*(0.021)*	-0.010	*(0.021)*
$\hat{\psi}_{J(i,pre)}$			**-1.261**	*(0.063)*	**-1.243**	*(0.063)*	**-1.242**	*(0.062)*
$\hat{\psi}_{J(i,post)}$			**1.230**	*(0.059)*	**1.221**	*(0.058)*	**1.214**	*(0.058)*
Adj. R^2	0.0215		0.1749		0.1793		0.1809	

NOTE: Dependent variable is the the log difference in "full-quarter" wages (mean = 0.1260). All regressions include year of displacement dummies and a control for the time difference between earnings measures. Standard errors are corrected for clustering by 2,110 displacing firms. 10,702 observations. Source: LEHD data sources, 10 percent random sample, restricted to a maximum of one quarter of intervening unemployment, authors' computations.

Table B.2: OLS results: Full fixed effects, restricted unemployment

Variable	(1) Parameter Estimate	(1) S.E.	(2) Parameter Estimate	(2) S.E.	(3) Parameter Estimate	(3) S.E.	(4) Parameter Estimate	(4) S.E.
Early leaver $\Delta\mu$	**0.193**	*(0.017)*	**0.161**	*(0.018)*	**0.165**	*(0.018)*		
(Same Industry)							**0.095**	*(0.024)*
(Industry leaver)							**0.260**	*(0.027)*
Intercept	0.017	*(0.017)*	**0.146**	*(0.064)*	**0.147**	*(0.064)*	**0.123**	*(0.064)*
Temp layoff			**0.065**	*(0.027)*	**0.069**	*(0.027)*	**0.062**	*(0.027)*
Total experience (Q)			**-0.002**	*(0.001)*	**-0.001**	*(0.001)*	**-0.001**	*(0.001)*
Exp.^{2}/10000			4.267	*(3.581)*	2.314	*(3.595)*	1.969	*(3.592)*
Education			0.001	*(0.003)*	-0.003	*(0.003)*	-0.003	*(0.003)*
Race: Black			-0.001	*(0.021)*	0.020	*(0.021)*	0.019	*(0.021)*
Race: Hispanic			-0.008	*(0.028)*	0.015	*(0.028)*	0.014	*(0.028)*
Race: Other			-0.006	*(0.025)*	-0.002	*(0.025)*	-0.002	*(0.025)*
Theta $\hat{\theta}$			**0.070**	*(0.013)*	**0.073**	*(0.013)*		
Unemployment			**-0.026**	*(0.015)*	**-0.028**	*(0.015)*	0.009	*(0.017)*
Same industry			0.043	*(0.044)*	0.037	*(0.044)*	0.028	*(0.044)*
Adj. R^2	0.0211		0.0117		0.0237		0.0256	

NOTE: Dependent variable is the the log difference in "full-quarter" wages (mean = 0.1260). All regressions include a control for the time difference between earnings measures; and fixed effects for 2,118 displacing firms and 3,588 new firms. 10,702 observations. Source: LEHD data sources, 10 percent random sample, restricted to a maximum of one period of intervening unemployment, authors' computations.