The heterogeneous costs of job displacement

The case of firm closures in Norway

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Preface

Working with this thesis seemed at the start a Sisyphean task, but ended much less burdensome. The main reason for this is the motivation and help provided by my supervisors Tarjei Havnes and Knut Røed, to whom I am both heavily indebted. Their vast knowledge of econometric methods and institutional features of the labor market was of great help in writing this thesis. I would also like to thank Espen Moen for valuable input regarding theoretical models of on-the-job search. Furthermore, the constructive feedback from the participants at the Frisch seminar was of great help. I am indebted to the The Ragnar Frisch Centre for Economic Research for generously granting me a scholarship and an office space during the writing of this thesis. This thesis is a part of their project 1172 *Social Insurance and Labor Market Inclusion in Norway*, financed by the Norwegian Research Council's grant no. 202513. I would also like to thank the Centre for Equality, Social Organization and Performance (ESOP) for awarding me a scholarship for working with this thesis. Finally, I would like to thank Øyvind Bruer-Skarsbø, Kristine Wika Haraldsen, Ingrid Hjort and Camilla Karto for being a great study group.

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1 Introduction

Several studies have found that there is in general a negative effect of job displacement on a large set of various labor market outcomes. However, many of these studies only compare the workers who remain in a firm upon closure, and do not consider the downsizing process leading up to this final displacement event. More to the point, the remaining workers could be a selected subset of the original labor force (Bratsberg et al. 2013, p. 141). This is because a downsizing process may be characterized by two opposing effects. The first effect is that the firm wants to retain the most productive workers when downsizing. The second effect is that the most productive workers may fear being laid off at some stage during downsizing, and decide to separate from the firm before this happens (Schwerdt 2011, p. 93). Neglecting these two effects and the associated labor turnover prior to firm closure, and only focusing on the ultimately displaced workers left in the firm upon closure, may therefore lead to a biased estimates of the costs of displacement.

In this thesis, I argue that the differences among workers in the effort exerted in on-the-job search during the downsizing process and the subsequent differences in their post-displacement labor market outcomes are caused by differences in their stock of human capital. To keep the argument simple, I focus only on downsizing processes that culminate in the closure of the firm. Furthermore, I distinguish between only two types of displacing workers in the downsizing process based on when they displace from the firm. The first group are the "early leavers" who depart before closure, and the second group are the "stayers" who remain until closure. The two labor market outcome variables I consider are the post-displacement annual pensionable income and the employment status of the worker. To estimate the difference in the post-displacement labor market outcomes between the early leavers and the stayers, I use matched data on individual workers (administrative register data from Statistics Norway) and individual firms (audited accounting data from the Norwegian bankruptcy court system).

I first identify the early leavers and the stayers by determining when the selective labor turnover directly related to the downsizing process starts by comparing the post-separation outcomes of workers separating from closing and non-closing firms using a difference-in-difference approach. I find that the selective labor turnover related to the firm's downsizing process starts two years prior to the closure of the firm. Therefore, all workers displacing from a firm one or two years prior to the closure of the firm are considered early leavers, and all workers remaining in the firm upon closure are considered stayers.

I then estimate the costs of displacement using a double difference-in-difference approach, where I find that the cost of displacement in terms of annual pensionable income is greater for the stayers compared to the early leavers. Specifically, the annual income loss for the stayers is from 2.2 to 4.5 percent higher in the five years after the displacement year. The immediate cost of displacement in terms of employment opportunities is greater for the early leavers compared to the stayers, with the early leavers' employment rate in the year of displacement being 10.9 percentage points lower than the stayers'. However, this employment rate difference between the early leavers and the stayers is not a persistent effect.

Finally, I construct empirical human capital distributions for the early leavers and the stayers using a three-way fixed effects log-wage model estimated prior to the start of the downsizing process. I find that the general human capital distribution for the early leavers is significantly right-shifted compared to the distribution for the stayers, and that the firm-specific human capital distribution for the stayers is significantly right-shifted compared to the distribution for the early leavers. This lends support to the hypothesis that the stayers are negatively selected in terms of general human capital and positively selected in terms of firm-specific human capital. This could explain the stayers' larger annual income loss due to displacement compared to the income loss for the early leavers due to displacement (measured as percentage loss of the pre-displacement income).

The structure of the thesis is as follows: Part 2 goes through the existing literature on the effect of job displacement on workers' labor market outcomes, as well as the literature on changes in the firms' labor force composition during the downsizing process. Part 3 sketches out a simple on-the-job search model, allowing for the endogenously determined search intensity to depend on the workers' human capital stock. Part 4 describes how the Norwegian labor market institutions affect the incentives, and therefore the behavior, of the workers in the downsizing process and discusses how this may confound the predictions of the theoretical model. Part 5 describes the administrative register data being used, imposes sample restrictions and provides descriptive statistics.[1] Part 6 defines the empirical strategy used to identify the early leavers and the stayers, and presents the estimation results as well as some descriptive statistics for the early leavers and the stayers. Part 7 outlines the empirical strategy for estimating the heterogeneous costs of displacement, and reports the results from the estimation. Part 8 discusses how the empirical human capital distributions are constructed, and illustrates these distributions for the early leavers and the stayers. Part 9 concludes.

2 Literature

There have been several empirical studies on the effects of firm downsizing on various outcomes for individual workers.[2] However, since the focus is on the effect of displacement on the income and employment rate of the displaced workers, attention will be paid to studies concerning the effect on these two outcomes.

Huttunen et al. (2011) define displacement as workers leaving plants that close down or downsize their labor force by 30 percent or more, and use Norwegian administrative register data to investigate the effect of downsizing on the probability of exit from the labor force and post-displacement earnings. The authors find that displacement from a downsizing firm leads to a 31 percent increase in the probability of exit from the labor force for an average male worker working in a manufacturing plant. Displacement is also accompanied by a monotonic reduction in earnings up to the time of exit from the labor force, with workers exiting the labor force when the earnings loss due to displacement is between 10-30 percent of their pre-displacement earnings. In addition, the authors find that the annual earnings loss for those who stay in the labor force reaches its maximum of approximately US$ 2,000 in the second year after displacement, and

[1] All the statistical analyses in this thesis have been performed using Stata/SE 13.1.

[2] See for example Rege et al. (2009) for effect of displacement on crime and Bratsberg et al. (2013) for effect of displacement on disability insurance claims.

that the negative effect persists through the seven years studied after displacement. The novel feature of this study compared to other studies is that the authors also take care to identify within- and between-firm movers, with their results supporting the view that human capital is partly firm-specific and partly job-specific.

Bratberg et al. (2008) use the same definition of displacement as Huttunen et al. (2011), and find that for male workers in Norway displaced from a closing firm there is a negative effect of displacement on annual earnings of about 7 percent five years after displacement, and 9 percent after eight years. In addition, they find that for workers displaced from closed firms the share of unemployed is 14 percentage points higher than for workers who are not displaced at all (measured eight years after displacement), and consistently longer periods of unemployment for workers from closed firms compared to workers who are not displaced at all. The same results hold for those workers who are displaced from downsizing firms that do not close, although the magnitudes of the effects are slightly lower.

Moving on to studies outside Norway, Eliason and Storrie (2006) define displacement as workers separating from a closing firm, and using administrative register data from Sweden the authors find that twelve years after displacement, the annual earnings gap between displaced and non-displaced workers is up to US$ 870 and the unemployment differential is 3.7 percentage points. The authors combine exact matching and propensity score matching to create a control group to identify these effects. The authors also find evidence that job displacement leads to permanent "scars" (in terms of for example lost human capital, both firm-specific and general) or transitory "blemishing" effects (where new employers take into account the previous labor market histories to deduce the worker's productivity). The permanent scars lead to the displaced workers' income and employment opportunities being more vulnerable to business cycle movements.

Using data from Portugal, Carneiro and Portugal (2006) find that three years after displacement from a closing plant, the average earnings of displaced male workers is around 12 percentage points lower than for non-displaced male workers. In contrast to Eliason and Storrie (2006), the authors go further in trying to decompose the observed earnings loss into its main explanatory components. The authors find that for these displaced male workers, 40-46 percent of the earnings loss is due to loss of job tenure, 33-43 percent is due to unemployment spells (where either general human capital depreciation or the blemishing effect takes place), and 14-24 percent is due to change of industry in which they are employed.

For New Zealand, Dixon and Stillman (2009) find that displaced workers from a closing plant have 16 percent lower monthly earnings and 12 percent lower employment rate compared to non-displaced workers four years after displacement. The study differs from most other studies in that the authors identify four categories of firm exits: firm closure without and with employee group transfers, branch closures and probable restructurings. The workers' post-displacement labor market outcome depends on what type of firm exit they are subjected to. The authors find that the effect of displacement on earnings and unemployment rate differs across these four categories, with the most negative effects being for the workers experiencing genuine and complete closures (the first and third category of firm exits). In other words, the "true" firm closure events lead to the most adverse consequences in terms of the workers' post-displacement outcomes.

Considering the role of the firm in the downsizing process, characteristics such as the firm's market share and unionization rates have been found to be important explanatory variables in understanding how a firm responds to negative demand shocks (Abowd et al. 2009, p. 468). However, as the focus is on the on-the-job search behavior of the displacing workers, only studies concerning changes in the firm's labor force composition due to turnover during downsizing are reviewed.

Abowd et al. (2009), using data from the United States, distinguish between workers in terms of human capital levels by estimating a fixed-effects model with a person-specific component (gender, age and education level) and an experience component, and then constructing a human capital distribution for each firm. By doing so, the authors remove firm-specific effects and idiosyncratic residuals from the wage so as to obtain a "cleaner" estimate of the human capital distribution than by only using the wage distribution. From these human capital distributions, the authors subsequently find that firm closures occur substantially more often in firms that employ relatively more workers in the lowest quartile of the human capital distribution and less often in firms that employ relatively more workers in the highest quartile of the human capital distribution.

Lengermann and Vilhuber (2002) describe the competing agendas of the forward-looking workers and the optimizing firm management in the process of downsizing. Based on this the authors study the changes in general human capital quality of the workforce, with the general human capital quality of any worker defined as the worker's position in the distribution of person-fixed effects. Using data from the United States, the authors find that high-skilled workers are more prevalent in the firm up to three quarters before downsizing. This is consistent with the conjecture that high-quality workers leave the firm ("abandoning the ship") to a greater extent than the firm laying off low-quality workers during downsizing.

Fackler et al. (2013) use data from East- and West-Germany and find support for the so-called "shadow of death"-hypothesis, with closing firms shrinking in terms of employment levels in all five years before closure. The authors define the skill level of a worker based on the occupation the worker holds in the firm and find that for West-Germany, in contrast to Lengermann and Vilhuber (2002), that the labor force becomes more skilled in closing firms as closure approaches compared to surviving firms. This indicates that the firm's process of laying off low-skilled workers is a stronger effect than high-skilled workers leaving the firm during the downsizing process. In other words, the firm is "throwing ballast overboard".

Henningsen and Hægeland (2008) find that Norwegian firms use downsizing, defined by the authors as the situation where a firm reduces the number of workers substantially without closing down, as a sorting device to lay off the least profitable workers in the firm. The authors define these least profitable workers as the workers with long records of sickness absence, since these workers are likely to be less productive when at work and more absent from work. These workers are protected against dismissal under normal times of operation by employment protection legislation, but not during firm downsizing processes caused by adverse economic circumstances. In addition, the authors find that it is worker characteristics such as age, firm tenure and education level relative to peers in the firm that matter for the likelihood of a worker keeping his job during the downsizing process.

3 Theory of on-the-job search

In order to impose some structure on the discussion, I construct a simple one-sided on-the-job search model to characterize the behavior of the workers in a downsizing process. The optimal job search effort of the workers is determined by their fixed human capital stock, which consists of both general human capital and firm-specific human capital. Furthermore, I define two types of workers in the downsizing process based on when they displace from the firm ("early leavers" and "stayers"), and consider their respective post-displacement wage levels and employment status.

3.1 The model setup

A worker $i \in [0,1]$ employed in a downsizing firm has a fixed level of general human capital $ghc_i > 0$ and a fixed level of firm-specific human capital $fhc_i > 0$. The defining difference between the two types of human capital is that general human capital is transferable across jobs, whereas firm-specific human capital is nullified upon separation from the current employer (Lazear 2003, p. 1). A worker i earns a period wage $w(ghc_i, fhc_i)$, with $\partial w(ghc_i, fhc_i) / \partial ghc_i > 0$ and $\partial w(ghc_i, fhc_i) / \partial fhc_i > 0$. The former partial derivative could for example reflect the fact that the greater the education level of the worker, the greater the marginal productivity and therefore the greater the wage. The latter partial derivative could for example reflect the fact that the greater the firm-specific skills due to firm-specific on-the-job training, the greater the firm-specific marginal productivity of the worker and therefore the greater the wage of the worker.

The firm faces an exogenous demand for their product at a fixed price, and employs a unit mass of workers. In period $t = 0$, the firm is subjected to an unanticipated negative demand shock known immediately to both the firm and the workers. The demand shock forces the firm to adjust down their production level by initiating a downsizing process of laying off workers. For simplicity, I disregard any adjustment of the firm's capital stock. The firm's downsizing strategy is assumed to be to lay off workers with the lowest level of firm-specific human capital first. One reason could for example be that the firm-worker match value is the highest for workers who have greater firm-specific human capital due to the shared investment from on-the-job training (Borjas 2013, p. 272). Another reason could be trade unions working to protect workers who have invested the most in the firm. The first critical assumption of the model is therefore:

Assumption 1. *In response to the negative demand shock in period $t = 0$, the firm's downsizing strategy is to lay off the workers with the lowest level of firm-specific human capital (fhc_i) first.*

As pointed out by Kuhn (2002, p. 16), this inverse-seniority layoff rule (also known as the "last-in, first-out"-rule) seems to be a nearly universal phenomenon, even when holding age constant. As such, Assumption 1 may be understood not necessarily as an optimal downsizing strategy of the firm, but rather as the firm honoring an implicit contract between the firm and the worker (see Part 4 for more on this). In each period $t > 0$ a worker therefore faces an exogenous probability of being laid off equal to $\lambda (fhc_i) \in (0, 1)$, with $\partial \lambda (fhc_i) / \partial fhc_i < 0$. In response to this, each worker i initiates preemptive on-the-job search in period $t = 0$ and exerts an on-the-job search effort $e_i > 0$. The cost of exerting search effort is $C (e_i)$, with $\partial C (e_i) / \partial e_i > 0$ and $\partial^2 C (e_i) / \partial e_i^2 > 0$. The greater the effort exerted in the on-the-job search, the greater is the probability of obtaining an acceptable wage offer. This probability is denoted as $p (e_i)$, with $\partial p (e_i) / \partial e_i > 0$ and $\partial^2 p (e_i) / \partial e_i^2 < 0$. The offered wage equals $w^n (ghc_i)$, with $\partial w^n (ghc_i) / \partial ghc_i > 0$. I assume that the offered wage, $w^n (ghc_i)$, is strictly greater than the reservation wage of the worker, which equals the expected wage of working in the current firm in the next period, $\lambda (fhc_i) w (ghc_i, fhc_i)$.

3.2 Human capital and optimal search effort

A worker i who learns of the negative demand shock in period $t = 0$ aims to maximize the expected net gain of performing on-the-job search less the cost of doing so, which can be formulated as the following maximization problem:

$$\max_{\{e_i\}} \{p (e_i) [w^n (ghc_i) - \lambda (fhc_i) w (ghc_i, fhc_i)] - C (e_i)\} \tag{1}$$

In Equation (1), I am assuming that the payoff from conducting on-the-job search but not obtaining an acceptable job offer (the probability of this equals $1-p(e_i)$) is zero. The necessary and sufficient first-order condition characterizing the optimal search effort by worker i, denoted e_i^o, tells us that the marginal benefit of exerting search effort must equal the marginal cost of exerting search effort:

$$p'(e_i)\left[w^n(ghc_i) - \lambda(fhc_i)\,w(ghc_i, fhc_i)\right] = C'(e_i) \tag{2}$$

From this first-order condition, the optimal search effort e_i^o is an implicit function of both the general human capital level and the firm-specific human capital level. Taking this into account, the first-order condition can be rewritten as:

$$p'(e_i^o(ghc_i, fhc_i))\left[w^n(ghc_i) - \lambda(fhc_i)\,w(ghc_i, fhc_i)\right] = C'(e_i^o(ghc_i, fhc_i)) \tag{3}$$

Differentiating the rewritten first-order condition above with respect to general human capital level of the worker, we find that the optimal search effort is increasing in the general human capital level (see Appendix A.1 for details):

$$\frac{\partial e_i^o}{\partial ghc_i} = -\frac{p'(e_i^o)\left[\left[\frac{\partial w^n}{\partial ghc_i}\right] - \lambda\left[\frac{\partial w}{\partial ghc_i}\right]\right]}{p''(e_i^o)\left[w^n - \lambda w\right] - C''(e_i^o)} > 0 \tag{4}$$

This results holds under a general assumption, namely that the wage function for the current employer $w(\bullet)$ and a potential new employer $w^n(\bullet)$ respond equally to a marginal increase in the general human capital level. This seems like a defensible assumption given that general human capital is transferable across firms. Furthermore, from the rewritten first-order condition in Equation (3), we find that an increase in the firm-specific human capital level has an ambiguous effect on the optimal search effort (see Appendix A.2 for details):

$$\frac{\partial e_i^o}{\partial fhc_i} = \frac{p'(e_i^o)\left[\left[\frac{\partial \lambda}{\partial fhc_i}\right] w + \lambda\left[\frac{\partial w}{\partial fhc_i}\right]\right]}{p''(e_i^o)\left[w^n - \lambda w\right] - C''(e_i^o)} \tag{5}$$

To establish the sign of this effect, the non-parametric assumption that has to be made is not defensible from an economic point of view. However, if I parametrize the expression using Norwegian data, the condition for the partial effect in Equation (5) to be negative is not wholly implausible (see Appendix A.2 for parametrization). The first proposition of the model is therefore:

Proposition 1. *The optimal on-the-job search effort for worker i (e_i^o) is increasing in the general human capital level of the worker (ghc_i) and decreasing in the firm-specific human capital level of the worker (fhc_i).*

The firm does not only lay off workers in period $t = 1$ as a response to the negative demand shock, but continues to do so in the subsequent periods as well. The reason is not that the firm is subjected to more negative demand shocks, but rather that the firm is recompositioning their labor force as argued by Henningsen and Hægeland (2008). Let us therefore consider the composition of the original labor force as the downsizing process continues by first defining the displacement rate T_i for worker i in period $t > 0$ as:

$$T_i = p\left(e_i^o\left(fhc_i, ghc_i\right)\right) + \lambda\left(fhc_i\right) \tag{6}$$

What I am implicitly assuming in Equation (6) is that there are no overlapping notifications, meaning that a worker who successfully finds an acceptable job offer and sends a quit notification to the firm is not at the same time being sent a layoff notification from the firm. Let us now partition the initial unit mass of workers into two equally large groups based on the human capital dimension, namely a “high-skilled” group with a high level of general human capital and low level of firm-specific human capital, and a “low-skilled” group with a low level of general human capital and high level of firm-specific human capital. Based on this partioning of the initial labor force, I make a second critical model assumption regarding the human capital of a worker (see Part 8.1 for more on this):

Assumption 2. *The greater the level of firm-specific human capital of a worker (fhc_i) the lower is the level of general human capital of the worker (ghc_i), and vice versa.*

Assumption 2 implies that the displacement rate T is higher for the high-skilled group (denoted with subscript HS) than for the low-skilled group (denoted with subscript LS) in each period $t > 0$ following the negative demand shock:

$$T_{HS} > T_{LS} \tag{7}$$

since

$$p\left(e^{o}\left(fhc_{HS}, ghc_{HS}\right)\right) > p\left(e^{o}\left(fhc_{LS}, ghc_{LS}\right)\right)$$

$$\lambda\left(fhc_{HS}\right) > \lambda\left(fhc_{LS}\right)$$

This result stems from both the endogenously determined optimal on-the-job search effort of the workers, and the downsizing strategy of the firm – both of which depend on the human capital stock of the worker. Note that both the job-to-job transition rate and the job-to-unemployment transition rate is higher for the high-skilled group. If we now assume that the finitely-lived firm for some unspecified reason closes down at the end of any period $s > 0$, it follows from Equation (7) that the labor force upon closure is relatively less skilled than the labor force prior to the negative demand shock in period $t = 0$, disregarding any new hires made by the firm. Therefore, the second proposition of the model is:

Proposition 2. *Under Assumption 2, the labor force of a downsizing firm upon closure at the end of any period $s > 0$ is less skilled than it was prior to the negative demand shock in period $t = 0$.*

Following Schwerdt (2008, 2011), I make a further distinction between the workers in the firm based on the temporal dimension. Let us denote all workers displaced prior to firm closure as "early leavers" and all workers still employed in the firm upon closure as "stayers". Both groups contain both high-skilled and low-skilled workers. However, with the greater displacement rate of high-skilled workers compared to low-skilled workers in each period, $T_{HS} > T_{LS}$, it follows directly that the early leavers have on average greater general human capital levels and lower firm-specific human capital levels than the stayers.

3.3 Post-displacement outcomes

Let us first consider the post-displacement wage level conditional on finding employment (either directly after displacement or after an unemployment spell), measured k periods after the firm closure. Keeping the distinction between early leavers and stayers, we have that the post-displacement wage conditional on employment is on average greater for the early leavers than for the stayers. This is because general human capital is the only marketable asset in the labor market after displacement. Since the early leavers have a greater average level of this asset compared to the stayers, they are better off in terms of average wage level since $\partial w^n(ghc_i)/\partial ghc_i > 0$. The third proposition of the model is therefore:

Proposition 3. *Conditional on having found employment in period $s+k$ (k periods after firm closure), the average post-displacement wage level of the early leavers is greater than the average post-displacement wage level of the stayers.*

The second post-displacement outcome we are interested in is the employment rate, which is again measured k periods after the firm closure. When the firm closes down at the end of period s, a fraction 1 of the stayers enter unemployment. Only a fraction $\lambda^s < 1$ of the early leavers have entered unemployment at some point in time between period $t = 1$ and period $t = s$, assuming that upon finding a new job a worker is not displaced again. However, a fraction $\delta \in [0, 1)$ of the early leavers that entered unemployment have during this period managed to find a job, leaving behind only a fraction $(1 - \delta)\lambda^s < 1$ of early leavers who are unemployed in period s. Therefore, at the end of period s, the unemployment rate is 1 for the stayers and $(1 - \delta)\lambda^s$ for the early leavers. Conversely, the employment rate is 0 for the stayers and $1 - (1 - \delta)\lambda^s > 0$ for the early leavers.

In order to analyze the job search of the unemployed stayers and early leavers in the periods following firm closure, we can augment the existing model framework. Once unemployed, I assume worker i receives an unemployment benefit $b(w(ghc_i))$. The unemployment benefit is increasing in the previous wage level, which I assume now only depends on the general human capital level. Furthermore, since unemployment benefits do not fully compensate for the loss of work income, I assume that $\partial b(w(ghc_i))/\partial w(ghc_i) \in (0, 1)$. The remaining model framework is the same as introduced previously. The unemployed worker i now aims to maximize the expected net gain of performing job search less the cost of doing so, which can be formulated as the following maximization problem:

$$\max_{\{e_i\}} \{p(e_i)[w^n(ghc_i) - b(w(ghc_i))] - C(e_i)\} \tag{8}$$

From the necessary and sufficient first-order condition for the maximization problem in Equation (8) we find that the optimal search effort when unemployed, denoted e_i^o, is increasing in the level of general human capital (see Appendix A.3 for details):

$$\frac{\partial e_i^o}{\partial ghc_i} = -\frac{p'\left(e_i^o\right)\left[\left[\frac{\partial w^n}{\partial ghc_i}\right] - \left[\frac{\partial b}{\partial w}\right]\left[\frac{\partial w}{\partial ghc_i}\right]\right]}{p''\left(e_i^o\right)\left[w^n - b\right] - C''\left(e_i^o\right)} > 0 \tag{9}$$

This result holds under the general assumption made earlier in Part 3.2, namely that the wage function for the previous employer $w(\bullet)$ and a potential new employer $w^n(\bullet)$ respond equally to a marginal increase in the general human capital level. It immediately follows that the likelihood of obtaining an acceptable job offer in any period after firm closure is greater for an unemployed early leaver compared to an unemployed stayer. This owes to the early leavers' higher average level of general human capital and therefore greater optimal search effort compared to that of the stayers. Combined with the higher unemployment rate of the stayers in period s and the greater rate of unemployment-to-job transition rate of the early leavers in each of the k periods following firm closure, the fourth and final proposition of the model is:

Proposition 4. *In period $s+k$ (k periods after firm closure), the unemployment rate of the stayers is higher than the unemployment rate of the early leavers. Conversely, the employment rate of the early leavers is higher than the employment rate of the stayers in period $s+k$.*

The key point of this theoretical model is to argue that the downsizing process is best seen as an initiation of on-the-job search by workers in response to the negative demand shock, with the human capital stock of the worker as the underlying variable determining the optimal job search effort and the subsequent post-displacement outcomes of the worker. As such, the model above resembles the traditional neoclassical search-theoretical job-ladder models (Nagypál 2005). Other theoretical approaches include downsizing as a process characterized by information asymmetry (Gibbons and Katz 1991) or entailing two-sided learning (Pfann and Hamermesh 2008). However, neither these alternative modelling approaches with their sequential updating of agents' information sets nor the model sketched out above encourage empirical strategies that are directly compatible with administrative register data.

4 Institutional setting

Having sketched out the optimal job search behavior of the early leavers and the stayers during the downsizing process and after the closure of the firm, we can now consider the Norwegian labor market institutions and the different ways in which these institutions change the incentives for job search. As such, these institutions may confound the predictions from the theoretical model in Part 3. More specifically, let us consider the legal rules and norms regarding firm downsizing and the general labor market institutions.

The Norwegian Working Environment Act states that employment is terminable with one month's notice for workers with tenure less than or equal to five years, two month's notice for workers with tenure between five and ten years, and three months for workers with tenure more than ten years (Arbeidsmiljøloven 2005). However, most employment contracts have a three-month notice requirement for both parties (Huttunen et al. 2011, p. 847). In accordance with this norm, even when having accepted another job offer a worker has (in general) to stay at the firm for three months onwards according to the contractual obligations. The consequence is that the time taken for transition from employment into either unemployment or employment upon separation may vary across different workers depending on their firm tenure and individual contracts. Furthermore, the notice requirement means that all soon-to-be-displaced workers have an incentive to increase their on-the-job search effort to avoid transiting into unemployment upon displacement, so the optimal on-the-job search effort as defined in the theoretical model in Part 3.2 is not time-invariant during the downsizing process.

In addition, the employer is required by the Norwegian Working Environment Act to arrange consultations with workers' representatives in order to jointly decide upon who will be laid off, over what period of time and potential alternatives to layoffs, such as reemployment at other parts of the firm (Arbeidsmiljøloven 2005).

The norm in such consultations is that workers with the lowest firm tenure are laid off first, which is in line with Assumption 1 if we consider the often necessary condition for a worker having firm-specific human capital, namely firm tenure. Therefore, as mentioned in Part 3.1, Assumption 1 of the theoretical model may be best understood as the firm honoring this inverse-seniority layoff norm.

Looking to income-based unemployment insurance (UI) institutions, a worker is entitled to a benefit of 62.4 percent of the previous year's income before tax, or 62.4 percent of the average income over the last three years (Huttunen et al. 2011, p. 847). Benefits may be received for up to 104 weeks if the annual income was at least twice the National Insurance Scheme basic amount (NOK 88 370 per May 2014), and for up to 52 weeks if the annual income was less than twice the National Insurance Scheme basic amount (NAV 2014). In addition, there are certain requirements regarding labor market mobility and active job search that have to be fulfilled to claim unemployment benefits (Duell et al. 2009, p. 70). Although the apparent duration limit for UI was increased in the major reform in January 1997, the absolute duration for UI was in fact decreased when taking into account the "soft" duration constraints prior to the reform. The main content of the reform was to scale down participation in activation programs and instead provide income insurance and encourage job search (Gaure et al. 2012, p. 440). The new "harder" constraints may have led to the exhaustion of benefits more markedly increasing the likelihood of becoming a discouraged worker (an increase in the hazard rate of exiting the labor force altogether) or the likelihood of exiting unemployment (an increase in the hazard rate of exit to employment or other social insurance programs). For example, Gaure et al. (2012, p. 444) find that the employment hazard rises by approximately 50 percent during the last month of the benefit schedule. The consequence of benefit exhaustion is that we must carefully choose the point in time after firm closure where we compare the post-displacement outcomes of the early leavers and the stayers, as the optimal job search effort after displacement may not be time-invariant as in the theoretical model in Part 3.

Moving on to labor market training programs (LMP), we can note that participants in such programs receive payments similar to unemployment benefits in addition to training programs or practical work experience programs. Critically however, the unemployment benefit schedule is not reduced whilst in the program (Gaure et al. 2012, p. 440-441). These LMPs are likely to increase the level or the quality of the general human capital of the participant. For example, Godøy and Røed (2014) find that after completing the LMP the hazard rate to "good" jobs with better earnings increase by 49 percent, albeit with strong lock-in effects during the program. As a result, the availability of LMPs may imply that a disadvantaged worker (in terms of the level of general human capital) in a downsizing firm will conduct little or no on-the-job search. The worker then maximizes the length of wage payments from the firm, and after being laid off enters a LMP to obtain the benefit payments and increase his general human capital stock. The behavior of the stayers in the theoretical model conforms to this, although not explicitly related to the presence of LMPs. However, we may expect that the decision to wait until being laid off will only be taken by the most disadvantaged workers within the stayer group whose outside opportunities are the poorest (due to low levels of general human capital and high levels of firm-specific human capital). Furthermore, their decision will also depend on how the available LMPs complement their existing general human capital stock. As a result, we may find that the stayers are a particularly negatively selected group of workers.

In terms of health-related insurance institutions, Bratsberg et al. (2013) find that job loss increases the likelihood of disability insurance claims, doubling the risk of subsequent program entry for men and raising enrollment for women by approximately 50 percent. The authors note that "there is a strong negative relation between prior earnings and the likelihood of disability benefit uptake" (p. 146). This high substitutability between disability insurance and income can be explained by job loss being a negative shock to the continued value of labor market participation with firm-specific human capital being nullified upon displacement, as in the theoretical model in Part 3. For low-income workers who have mainly firm-specific human capital and a low level of general human capital (such as the stayers), disability insurance may be an attractive alternative as downsizing continues and eventual firm closure seems probable. Indeed, low-income workers receive around 62 percent of last year's earnings in disability pension, which after tax amounts to a compensation rate of up to 83 percent of previous annual earnings (Huttunen et al. 2006, p. 10). This high substitutability between disability pension and labor income for stayers must be taken into account when considering their post-displacement employment status, as many of these workers may transition into disability insurance instead of attempting to obtain a new job.

5 Data and descriptive statistics

Before sketching out the methods and estimation results for identifying the early leavers and the stayers, comparing their post-displacement labor market outcomes and constructing their empirical human capital distributions, I describe the data being used. I then provide descriptive statistics of the workers and the firms for the period around a base year and in the base year itself.

5.1 Data and sample restrictions

The data used comes from Norwegian administrative registers on individuals' labor market and social security histories as well as individual characteristics, which is matched with publicly disclosed financial information of firms and data from bankruptcy court proceedings. The data covers the period 1995 to 2009, and I choose the year 2002 as the base year when the sample firms close down. This allows me to follow each individual worker seven years prior to the base year and seven years after the base year, for a total of fifteen years.

When studying the closing and non-closing firms, the focus is on single-plant firms in the private sector that have more than ten employees within the four years prior to the base year, i.e. 2001, 2000, 1999 or 1998. The reason for focusing on single-plant firms is that transfers to other plants within multi-plant firms after displacement are common, and the displacement costs for within- and between-movers differ (Huttunen et al. 2011, p. 841-842). By focusing on single-plant firms this heterogeneity of costs is disregarded. In addition, accounting and closure data is only available at the firm level, so this data can only be directly matched to workplace data for single-plant firms (Bratsberg et al. 2013, p. 140). Focusing on firms with more than ten employees reduces the volatility of downsizing, since the actions of individual workers in small firms may adversely affect the downsizing process. I follow Rege et al. (2009) in counting the number of workers at a firm by using "full-time equivalents" (FTE), with part-time and minor part-time employment counting as 0.67 and 0.33 FTEs, respectively. I focus on firms in the private sector since downsizing of firms and organizations in the public sector can owe to political restructurings, and not necessarily demand-driven changes. This means that I exclude workers in public administration. Specifically, I only include workers in firms that are organized as sole proprietorships (ENK), joint-stock companies (AS) or public limited companies (ASA).

A firm is defined as closed in year t if the firm identification number was present in the data in year $t-1$ but not present in year t. This is because the data on firms is updated on November 20$^{\text{th}}$ each year, so the time of firm closure was most likely in year t (Huttunen et al. 2011, p. 846). If the firm identification number reappears at some point in time up to the year 2009, the firm is identified as non-closed. Using the number of employees equal to zero as a criteria to identify firm closure is not a valid strategy, as a firm may have several employees left for administrative reasons upon closure. The reason for the firm closing is recorded from the court proceedings, where the reason is either bankruptcy, voluntary liquidation, or takeover by other firms. I define a firm as closed only if it entered bankruptcy or liquidation, as takeovers are likely to be followed by large transitions of workers from the closing firm to the takeover firm. Takeovers cannot therefore be characterized as true closures, but rather as mass transfers (Bratsberg et al. 2013, p. 142). It is worthwhile noting that the dating of the judicial status of the firm upon closure may be significantly lagged compared to the actual closure date due to lengthy bankruptcy court proceedings. For a firm that closes in year t all the available information from the bankruptcy court proceedings in year $t-1$ up to the year 2009 is therefore used.

Moving on to the workers, I define a worker as separating in year t if his firm identification number changes from year $t-1$ to year t. This is again because data on the matching between workers and firms is updated on November 20th each year. Many workers may have several work contracts both within the firm and in other firms (not necessarily single-plant firms), and I define the main work attachment for a worker in year t as the single-plant firm where the largest percentage share of the annual taxable income in year t comes from. This means that some workers who work in single-plant firms as their secondary work attachment are excluded from the sample. As the main work attachment of a worker may change over time, this definition may lead to false separations being recorded. For example, a worker who works 60 percent in a single-plant firm A and 40 percent in another single-plant firm B in year t may in year $t + 1$ work 60 percent at firm B and 40 percent in firm A instead. This will then be recorded as a separation from firm A. This may inflate the measure of total number of job separations and thus overstate the labor market mobility of the workers. In addition, I do not place any restrictions on what labor market state a worker displacing from a closing firm is in two years after displacement and onwards (the first year after displacement I restrict the worker to not transfer into another closing firm). This means that a worker may be an early leaver from a closing firm C, but transfers into another closing firm D two years later and ends up as a stayer. As a result, the number of early leavers and stayers may be inflated and bias the results. Finally, I do not place any restrictions on what type of firms workers transfer to after separation.

Furthermore, I define a worker as employed in year t if he has a firm identification number in year t, and conversely as unemployed in year t if he does not have a firm identification number in year t.[3] Finally, I only consider workers within the age group 25 to 55 in the base year to avoid recording the effect of workers transiting into early retirement programs upon separating from a firm, and to exclude young workers that may still be enrolled in educational institutions in the base year from the sample (Huttunen et al. 2006, p. 10). In contrast to studies such as Bratberg et al. (2008) and Schwerdt (2011), I do not restrict the workers to have any given tenure length in the firm prior to separation,

[3] Note that the workers classified as unemployed will include both workers who have registered for unemployment benefits and workers who are outside the labor force.

which means I include both high-mobility workers and marginal workers. I take the point of view that there are no *a priori* reasons why these workers should not be included in the analysis of the costs of displacement.

It is worth noting that the Norwegian administrative register data is annual data, in contrast to for example quarterly Austrian administrative register data (Schwerdt 2011) and monthly New Zealand administrative register data (Dixon and Stillman 2009). The consequence is that the time resolution of information regarding the downsizing process is in general low. This necessarily makes the analysis of the downsizing process rather crude, as changes in key variables may take place within a calendar year. As a consequence, we may end up studying only the labor turnover in prolonged downsizing processes, with processes that appear to be "sudden death"-events actually containing several stages of downsizing.

Another feature of administrative register data in general is that we cannot identify the reason for a worker-firm separation. To do so, we need survey data containing the stated reason for the separation. This poses a challenge, since it means we cannot determine if a separation was due to successful on-the-job search in response to downsizing, firm downsizing layoff strategies, or due to just cause (such as neglecting one's set tasks at work). More to the point, as noted by Fallick (1996, p. 5), separating workers who are fired for just cause should not be defined as displaced workers, as their separation is not structurally related to the downsizing process. Using administrative data may therefore lead us to pick up these "false" displacements and thus weaken the internal validity of the results.

5.2 Descriptive statistics

In order to characterize the sample of firms and workers in our base year, it is also useful to consider the changes in these stocks around the base year. Using the years $t = 1999, 2002$ and 2005, Table 1 summarizes the main pre-separation variables for the workers in single-plant firms organized as either AS, ASA or ENK. The individual characteristics are measured in year t, and the firm closure indicators are measured over the period $t + 1$ to $t + 3$. I take care to ensure that all firms, both closing and non-closing, had more than ten FTEs at some point in time during the period t to $t+3$. The income measure is the total annual pensionable income, rebased to 2002-NOK. I restrict the age of the workers to be between 25 and 55 for each year to increase comparability with the base year. Finally, the firm closure indicators are not necessarily mutually exclusive, so double-entries of firm closures do occur in the bankruptcy court data due to the unique reason for firm closure not being directly identifiable. The values in square brackets exclude these double-entry firms and the workers associated with these firms.

Table 1: Descriptive statistics for workers in 1999, 2002 and 2005

Year	1999	2002	2005
No. of workers	185,309 [184,545]	166,818 [166,462]	163,674 [163,567]
No. of firms	5,720 [5,697]	5,295 [5,282]	5,807 [5,800]
Age of worker (years)	37.8	38.0	38.4
Education level:			
Compulsory	25.9	23.1	21.3
Secondary	51.4	51.0	49.3
College / university	22.7	25.9	29.4
Income (2002-NOK)	318,746 [318,672]	342,383 [342,401]	356,211 [356,267]
Percent subject to:			
Bankruptcy	10.2 [10.0]	7.9 [7.7]	4.0 [3.9]
Liquidation	7.2 [6.9]	3.9 [3.6]	2.2 [2.1]
Takeover	16.6 [16.4]	12.0 [11.9]	6.9 [6.9]

Age and annual income are mean values of the sample of workers in 1999, 2002 or 2005, with annual income rebased to 2002-NOK using the consumer price index from Statistics Norway. The percentage share of firms subject to closure is measured over a three-year period following 1999, 2002 or 2005. The values reported in square brackets exclude double-entry firms.

From Table 1, we find that the average education level of the workers increases slightly over the period as more workers have college or university education, from 22.7 percent in 1999 to 29.4 percent in 2005. The average age of the workers is around 38 years, and the average annual real income increases over the period from NOK 318,746 in 1999 to NOK 356,211 in 2005. We furthermore find that there is initially a high percentage of the workers who are subject to bankruptcies, liquidations and takeovers during the next three years, but that this percentage declines throughout the period. This reflects the economic downturn in Norway over the period 2001 to 2003, and the subsequent recovery. This economic downturn was characterized by the relative increase in the number of workers registered as unemployed being the greatest among highly educated workers such as managers, engineers and academic professions (Sørbø and Handal 2010, p. 16-17). This differs from both earlier and subsequent economic downturns, where both the absolute and relative increase in the number of workers registering as unemployed was in the manufacturing sector. With 2002 as the base year, the external validity of the estimation results in Part 6.2, Part 7.2 and Part 8.2 may therefore be limited.

If we now consider the sample firms in the year prior to the base year, i.e. in 2001, we find that there are 271 closing firms and 4,272 non-closing firms. The average age of the closing firms is 13.5 years, and 18.9 years for the non-closing firms. The evolution of the number of employees and the annual revenue of the closing and non-closing firms in the period 1995 to 2001 is illustrated in Figure 1. As we see from Figure 1, the number of employees and the firm's annual revenue is consistently higher for the non-closing firms compared to the closing firms throughout the period, with the difference in average employment being 19 in 2001 and the difference in average annual revenue being NOK 42 million in 2001. Furthermore, we see that the employment level and revenue of the closing firms drop markedly compared to the non-closing firms from 2000 to 2001, which may point to the adverse economic circumstances for the closing firms and the associated downsizing process of the firm starting already in 2000.

6 Identifying the early leavers and the stayers

6.1 Empirical strategy

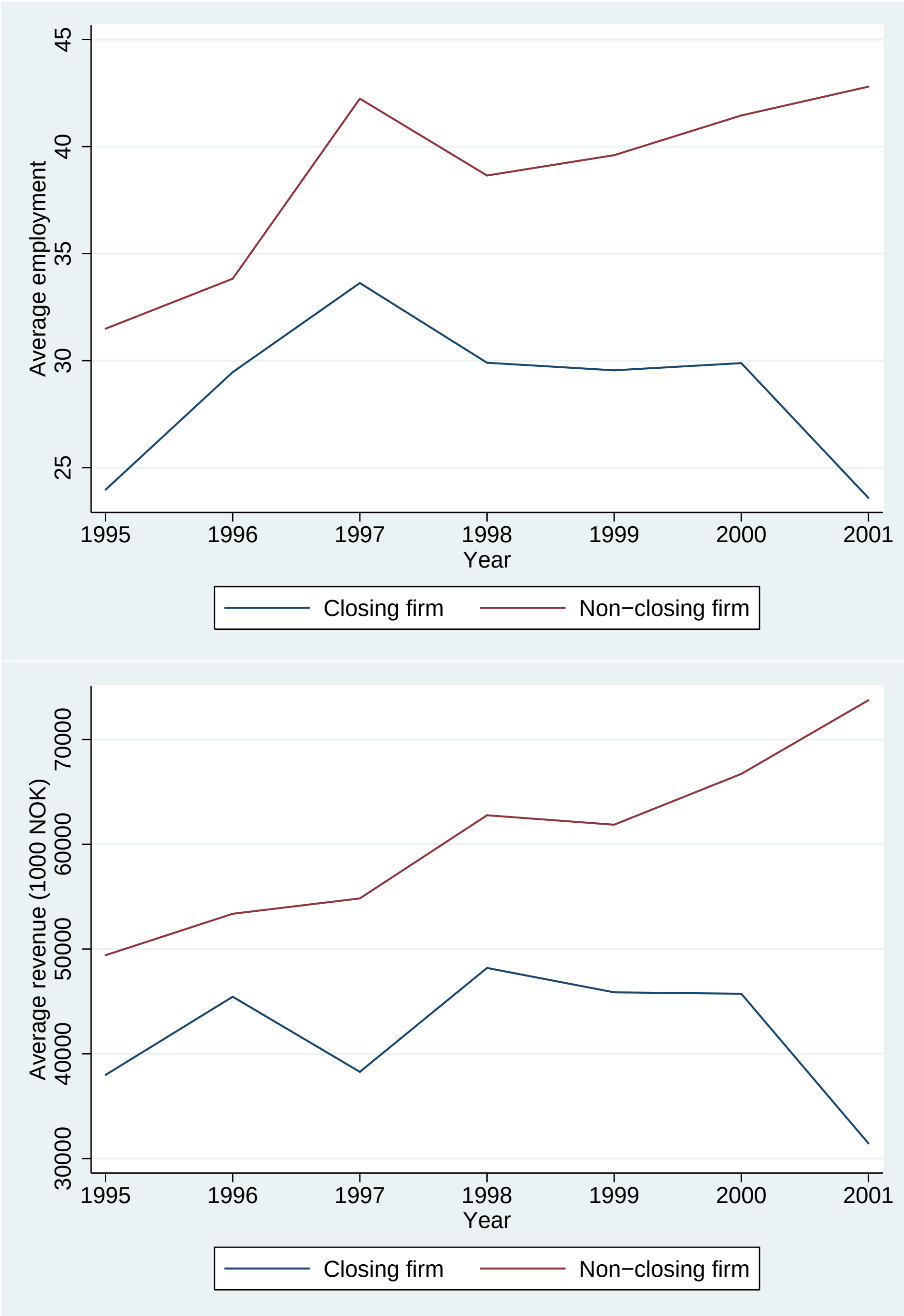
In order to identify the early leavers and the stayers from the theoretical model in Part 3, I follow Schwerdt (2011) in trying to empirically establish a time window prior to firm closure where the separations taking place are directly related to the firm's downsizing process (and in which the separations can be denoted as displacements), and not choose a time window based only on subjective judgement. For example, Huttunen et al. (2011) define workers who separate from a firm that closes within the next two years as early leavers, with no further motivation. Similarly, Henningsen and Hægeland (2008) define workers who separate from a firm that closes the next year as early leavers. Schwerdt (2011) sketches out the idea behind his empirical strategy by pointing out that if the observed separations:

> ... prior to the closure of a plant were due to "normal" labor turnover, which is not related to the upcoming plant closure, then post-separation outcomes [of the displacing workers from the closing plant] should be indistinguishable from post-separation outcomes of separations happening in non-closure plants (p. 96).

In other words, if the post-separation outcomes of the workers separating from closing firms before closure are significantly different from the post-separation outcomes of the workers separating from non-closing firms at the same point in time, there is selective labor turnover in the closing firm. To identify the time window in which selective labor turnover related to firm downsizing occurs, I consider a sample of only separating workers and distinguish between separating workers from closing firms and separating workers from non-closing firms. I require that all workers in the sample separate from the firm in the base year at the latest.

I take advantage of the panel data structure and use both individual- and time-fixed effects to control for individual-fixed characteristics (α_i) and calendar time-varying characteristics (θ_t). By using individual-fixed effects, I control for unobserved heterogeneity among the separating workers, such as for example innate ability. Including time-fixed effects controls for common time-series variation in the outcome variable, such as for example business cycle movements. The following baseline difference-in-difference model for identifying when the selective labor turnover directly related to firm closure starts is estimated over the period $t = 1995$ to $t = 2009$ for every worker i separating from either a closing

Figure 1: Average employment and revenue in closing and non-closing firms



Evolution of the average employment and the average annual revenue in closing and non-closing firms in the period 1995 to 2001. The annual revenue, measured in 1000 NOK, is rebased to 2002-NOK using the consumer price index from Statistics Norway.

or a non-closing firm d years prior to the base year:

$$Y_{it} = \delta T_{it} D_i + \gamma T_{it} + \underline{X}_{it}\underline{\beta} + \underline{Z}_{it}\underline{\chi} + \alpha_i + \theta_t + \varepsilon_{it} \tag{10}$$

The model is estimated for all separations happening up to four years prior to the base year $d = 0$, so Equation (10) is estimated separately for $d = 1$, 2, 3 and 4. Y_{it} is an outcome variable, which is either the employment status of worker i in year t or the total annual pensionable income for worker i in year t. D_i is an indicator variable that takes the value 1 if worker i separates from a closing firm, and 0 if worker i separates from a non-closing firm. T_{it} is an indicator variable that takes the value 1 for worker i in the year of separation from the firm and the three following years, and 0 otherwise. ε_{it} is the error term of the model, which is assumed to have conditional mean zero. Estimation is done by within-estimation, meaning that first dummy variables for each year $t = 1995$ to $t = 2009$ are added to capture the time-fixed effects, and secondly that each variable is then deviated from their individual mean (but not the time mean). Then ordinary least squares estimation is performed on the resulting demeaned version of Equation (10).

In Equation (10), the coefficient γ measures the average change in the outcome variable over the three years after separation from the firm, and therefore measures the effect of separation on the outcome variable. The coefficient δ measures the average effect on the outcome variable of separating from a closing firm compared to separating from a non-closing firm, with the effect being averaged over the three years after separation. This coefficient thus measures the additional average change in the outcome variable if the worker separates from a closing firm.

The baseline model in Equation (10) is modified depending on which outcome variable Y_{it} is being considered. When using the worker's employment status as the outcome variable, I include the vector of observable (and potentially time-varying) individual characteristics for worker i in year t, $\underline{X}_{it}$, to increase the comparability among the separating workers. I include the level of education and the area of residence (on municipality level) in this vector, as well as a constant term.[4] When using the total annual pensionable income for worker i in year t as the outcome

[4] As emphasized by Wooldridge (2012, p. 487), including the age of the worker (whose change is constant across time) in a model with a full set of year dummies as Equation (10) does not

variable, I include both the vector $\underline{X}_{it}$ and a vector of observable (and potentially time-varying) firm characteristics $\underline{Z}_{it}$ to control for potential firm heterogeneity in wage setting (Abowd et al. 2002, p. 12). I include the industry sector of the firm, the firm's annual revenue, and the number of firm employees in this vector.

After having estimated Equation (10) for $d = 1$, 2, 3 and 4, I consider the estimated coefficient $\hat{\delta}$ for each of these estimations to determine a critical time threshold when the selective labor turnover related to firm downsizing starts. If the estimated coefficient is statistically significantly different from zero, this points to the separation of the worker being directly related to the upcoming firm closure. The separation is then due to either being laid off as a result of the firm's downsizing strategy or performing a job-to-job transition in response to the negative demand shock, as argued in the theoretical model in Part 3. All separations from closing firms taking place before the critical time threshold are considered normal labor turnover not related to the downsizing process. Separations from closing firms taking place after the critical time threshold and up to (but not including) the base year $d = 0$ are displaced workers identified as early leavers, and are coded with $EL_i = 1$. The workers who displace from closing firms in the base year $d = 0$ are identified as stayers and are coded with $EL_i = 0$.

The empirical strategy above allows for identification of the separations from the closing firms that are structurally related to the upcoming firm closure, and that are not layoffs due to just cause. In other words, the empirical strategy allows for identification of true displacements despite the lack of survey data, an issue discussed in Part 5.1. A potential problem with this empirical strategy is that the non-closing firms in the control group may downsize their operations substantially in response to the adverse economic circumstances, but without closing. Workers separating from these downsizing but non-closing firms may therefore not serve as good controls to workers separating from closing firms, since worker separations from these firms include displacements as defined by Fallick (1996). I therefore impose a control group restriction where I exclude all workers who separate from

further control for variation in the outcome variable Y_{it}. Therefore, I do not include age in the vector of observable (and potentially time-varying) individual characteristics $\underline{X}_{it}$.

non-closing firms that downsize by more than 30 percent (in terms of FTEs) in the separation year, the same downsizing threshold as used in Bratberg et al. (2008).

6.2 Estimation results

Before estimating Equation (10), it may be interesting to see how the average employment and the average annual pensionable income (rebased to 2002-NOK using the consumer price index from Statistics Norway and not conditional on employment) of the sample of workers separating from closing and non-closing firms evolves over the time period $t = 1995$ to $t = 2009$. This is illustrated in Figure 2 and Figure 3 for the workers who separate in the four years prior to the base year (i.e. for $d = 1$, 2, 3 and 4), displayed in panels (a), (b), (c) and (d) in each figure, respectively. The mean values displayed in Figure 2 and Figure 3 are the raw means, meaning that I have not controlled any observable or non-observable characteristics of either the workers nor the firms. From Figure 2 and Figure 3, we see that the post-displacement employment rate and average annual income for workers separating from non-closing firms three and four years prior to the base year ($d = 3, 4$) is higher than for workers separating from closing firms. However, for separations one year prior to and two years prior to the base year ($d = 1, 2$), no such clear pattern of differences in the post-displacement employment rate or the post-displacement annual income emerges from Figure 2 and Figure 3.

I then proceed by estimating Equation (10) with the employment status of the worker as the outcome variable, with the key results summarized in Table 2. The first and second panel in Table 2 reports the estimation results when the vector of observable (and potentially time-varying) individual characteristics $\underline{X}_{it}$ is excluded and included, respectively, but with the control group restriction described in Part 6.1 not imposed. The third panel reports the estimation results when including $\underline{X}_{it}$ and imposing the control group restriction. The reported Huber-White standard errors in parentheses in Table 2 are clustered on the level of the individual worker, since there is likely to be serial correlation in the demeaned error term $\varepsilon_{it} - \bar{\varepsilon}_i$ when estimating the demeaned version of Equation (10) using ordinary least squares (Angrist and Pischke 2009, p. 319).

Figure 2: Evolution of average employment

(a) Separations in $d = 1$

(b) Separations in $d = 2$

(c) Separations in $d = 3$

(d) Separations in $d = 4$

Figure 3: Evolution of average annual pensionable income

(a) Separations in $d = 1$

(b) Separations in $d = 2$

(c) Separations in $d = 3$

(d) Separations in $d = 4$

The results in Table 2 indicate that the selective labor turnover directly related to the downsizing process starts two years prior to the closure of the firm ($d = 2$) since the post-displacement employment opportunities of separating workers in closing and non-closing firms are significantly different from zero for $d = 1$ and $d = 2$. The preliminary conclusion I draw from the results in Table 2 is that workers displacing from closing firms in 2000 ($d = 2$) and 2001 ($d = 1$) are early leavers, and workers displacing from closing firms in 2002 ($d = 0$) are stayers.

From the first row of all three panels in Table 2, the estimated coefficient $\hat{\delta}$ is positive for $d \in \{1, 2, 3, 4\}$, meaning that the employment rate over the three years after separation is higher for the workers separating from closing firms. This could owe to a greater rate of job-to-job transition as a result of the workers becoming aware of the downsizing process that eventually ends in the closure of the firm, and the subsequent initiation of on-the-job search. The fact that the magnitude of the statistically significant estimated coefficients $\hat{\delta}$ increase the closer the displacement event is to the base year (from 2.0 percentage points in 2000 to 3.0 percentage points in 2001 in the first row of the third panel in Table 2) may point to the on-the-job search intensity of workers in closing firms increasing as closure seems imminent. This leads to a greater job-to-job transition rate for the separating workers in closing firms compared to the job-to-job transition rate for separating workers in non-closing firms, since the latter group does not face downsizing.

The second row of the third panel in Table 2 points to that there is a statistically significant common loss of employment opportunities (ranging from -2.8 percentage points to -12.4 percentage points) for workers separating from both closing and non-closing firms in all four years prior to the base year. This follows from the fact that all the sample workers separate from their job, and within the three years following separation some of these workers will almost inevitably be unemployed and some will stay in unemployment throughout the period.

Moving on, I estimate Equation (10) with the annual pensionable income as the outcome variable to see if this identifies a different critical time threshold than in Table 2, where I now impose the control group restriction from Part 6.1 on all model specifications. The key conclusion of this estimation (the results are deferred

Table 2: Selective labor turnover (employment status)

	$d = 1$	$d = 2$	$d = 3$	$d = 4$
$T_{it}D_i$	0.020***	0.014*	0.013	0.008
	(0.006)	(0.008)	(0.008)	(0.010)
T_{it}	-0.171***	0.030***	-0.019***	-0.019***
	(0.004)	(0.004)	(0.004)	(0.004)
Observations (n)	30,798	25,669	25,673	21,817
Cases ($n \times t$)	461,970	385,035	385,095	327,255
R^2 (within)	0.08	0.09	0.08	0.05
	$d = 1$	$d = 2$	$d = 3$	$d = 4$
$T_{it}D_i$	0.020***	0.013*	0.006	0.008
	(0.006)	(0.008)	(0.008)	(0.010)
T_{it}	-0.016***	-0.110***	-0.064***	-0.083***
	(0.004)	(0.004)	(0.004)	(0.004)
Observations (n)	30,393	25,283	25,348	21,759
Cases ($n \times t$)	433,515	361,785	365,207	317,001
$\underline{X}_{it}$ included	Yes	Yes	Yes	Yes
R^2 (within)	0.09	0.10	0.08	0.05
	$d = 1$	$d = 2$	$d = 3$	$d = 4$
$T_{it}D_i$	0.030***	0.020**	0.009	0.009
	(0.006)	(0.008)	(0.008)	(0.010)
T_{it}	-0.028***	-0.124***	-0.066***	-0.085***
	(0.004)	(0.005)	(0.004)	(0.004)
Observations (n)	23,990	18,880	23,335	20,621
Cases ($n \times t$)	340,945	269,074	336,429	300,382
$\underline{X}_{it}$ included	Yes	Yes	Yes	Yes
Control group restriction	Yes	Yes	Yes	Yes
R^2 (within)	0.09	0.11	0.08	0.05

Results from estimation of Equation (10) over the period $t = 1995$ to $t = 2009$ with the dependent variable Y_{it} being an indicator variable for the employment status of worker i in year t. All regressions control for individual-fixed effects and time-fixed effects. The reported Huber-White standard errors in parentheses are clustered on the individual level.

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$

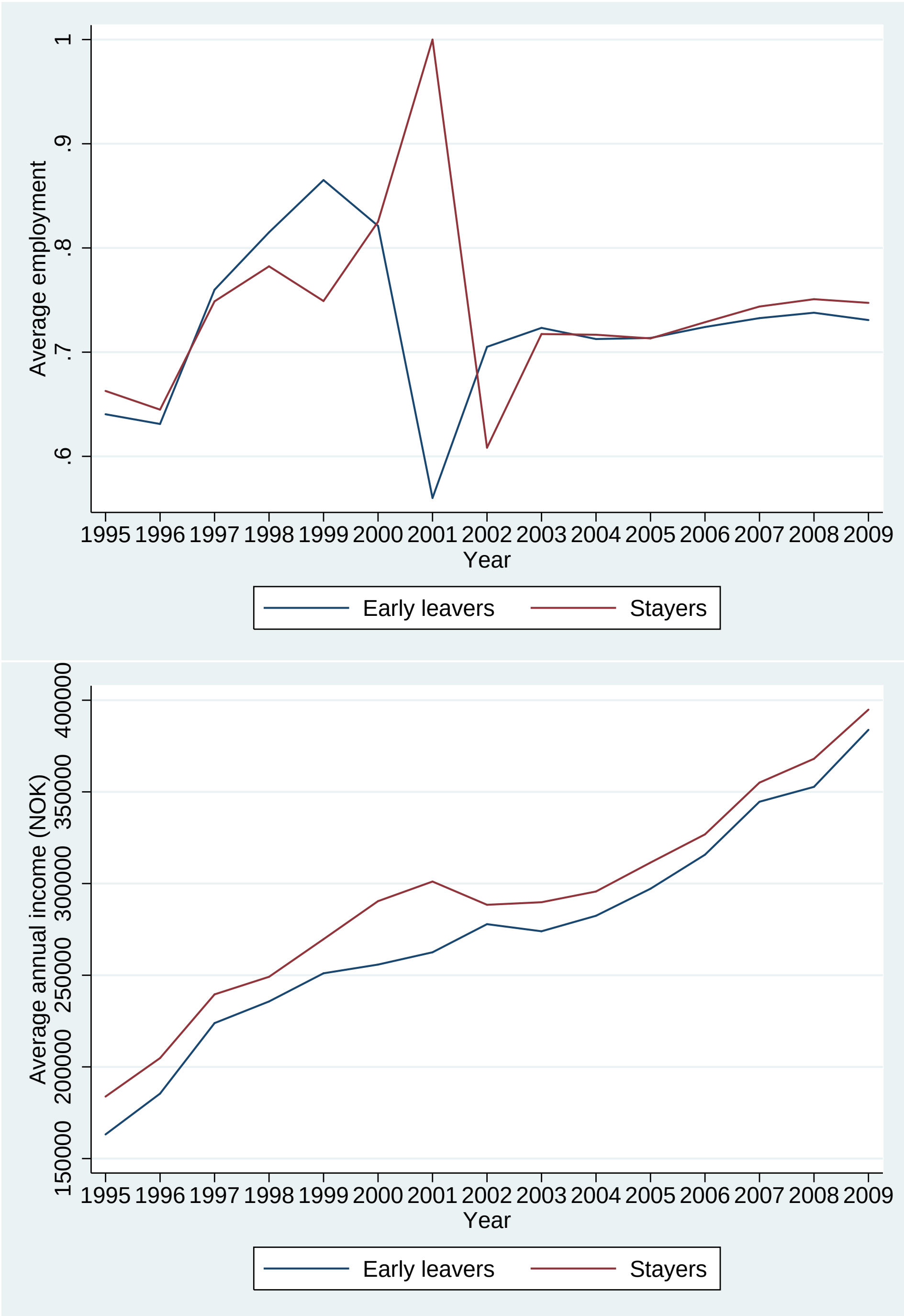
to Table 6 in Appendix B.1) is that it does not identify a different critical time threshold when considering the preferred specification of Equation (10) with $\underline{X}_{it}$ and $\underline{Z}_{it}$ included. This means I do not change the preliminary conclusion from the discussion of Table 2 regarding when the selective labor turnover related to firm closure starts. In other words, I classify all workers displacing from closing firms in 2000 ($d = 2$) and 2001 ($d = 1$) as early leavers and code them with $EL_i = 1$, and classify all workers displacing from closing firms in 2002 ($d = 0$) as stayers and code them with $EL_i = 0$. In the full sample of workers, these definitions identify 12,230 displacing workers (4.1 percent of full sample), with 48.9 percent of these displacing workers being early leavers and 51.1 percent being stayers.

6.3 Descriptive statistics for the early leavers and the stayers

Having identified the early leavers and the stayers, we can investigate the evolution of their employment rates and their average annual pensionable incomes (not conditional on employment) over the period 1995 to 2009. These raw means are illustrated in the first and second panel in Figure 4. As expected, the first panel in Figure 4 shows that there is a negative effect of displacement on the employment rate for both the early leavers and the stayers. The employment rate for both groups stabilizes from 2003 at around 0.72, but from 2005 and onwards the employment rate is greater for the stayers. Interestingly, we see from the second panel in Figure 4 that the annual pensionable income is consistently greater for the stayers than for the early leavers throughout the period 1995 to 2009, and that only the stayers experience a dip in annual real income in the years leading up to 2002. This dip occurs from 2001 to 2002 and therefore plausibly picks up some of the effect of displacement. From 2003 and onwards, the difference in the annual pensionable income between the early leavers and the stayers stabilizes, with the annual incomes following the same real income growth trend.

Furthermore, we can decompose the early leavers into two subgroups: the workers who displace two years prior to the base year, i.e. in 2000, and the workers who displace one year prior to the base year, i.e. in 2001. We can then compare the employment rates and the average annual incomes of these two groups of workers

Figure 4: Average employment and income of displaced workers



Evolution of the employment rate and the average pensionable income for the early leavers and the stayers in the period 1995 to 2009. The annual pensionable income is rebased to 2002-NOK using the consumer price index from Statistics Norway.

against each other, and against the stayers. The evolution of the employment rates and the average annual pensionable incomes over the period 1995 to 2009 for these three groups is illustrated in the first and second panel in Figure 5. From the first panel in Figure 5, we find that the workers who displace in 2000 seem to be consistently worse off in terms of post-displacement employment opportunities compared to workers displacing in 2001 and the stayers. The employment rate of the workers who displace in 2001 returns to a greater level in the year after displacement compared to the stayers, although the magnitude of the drop in employment rate for the workers who displace in 2001 is larger. From the second panel in Figure 5, we see that the stayers consistently have the highest average annual pensionable income throughout the period 1995 to 2009, as in Figure 4. Furthermore, we see that the workers who displace in 2000 are worse off than the other two groups in terms of annual income throughout the period.

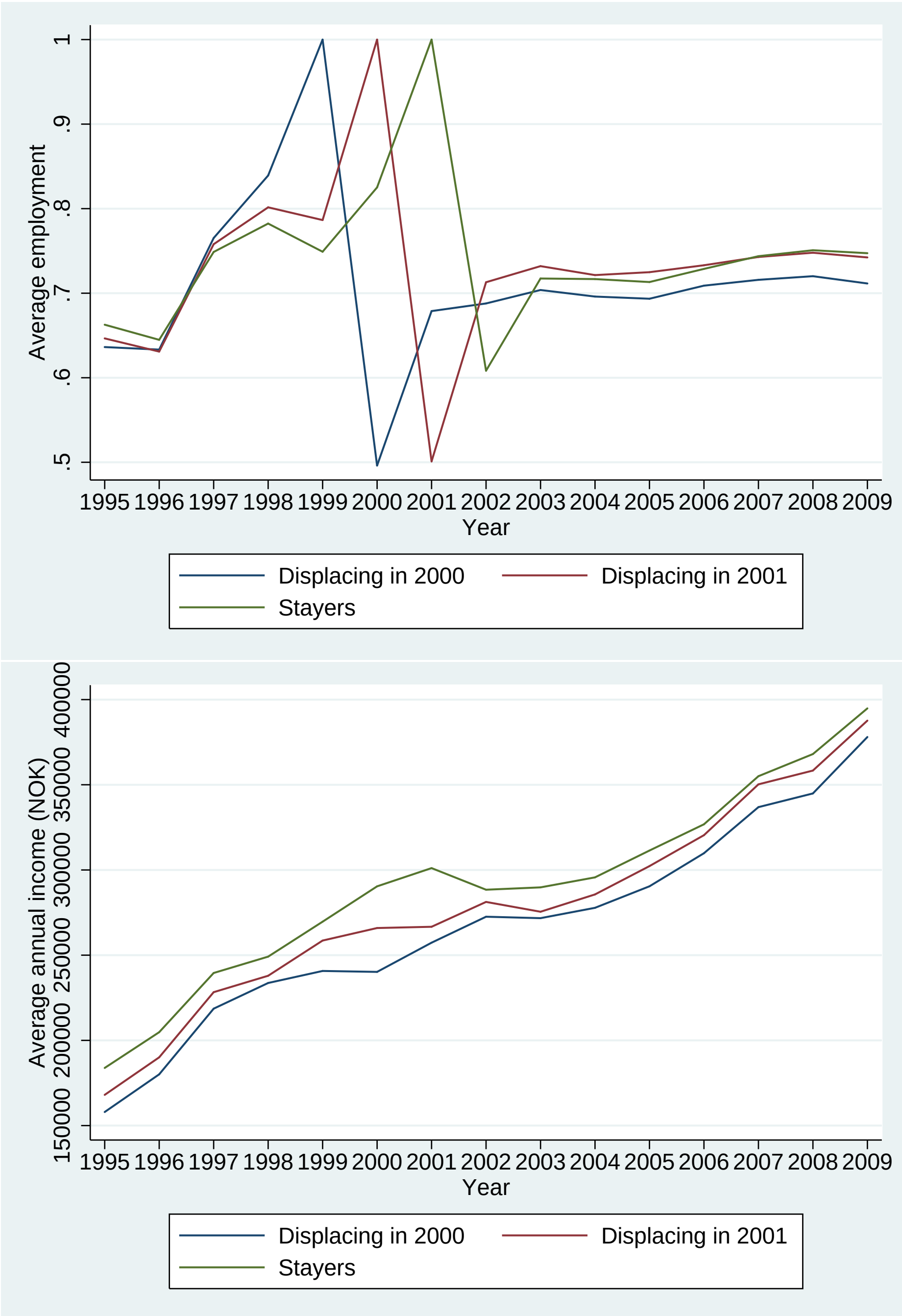
From Figure 4 and Figure 5 we can therefore conclude that there are not only differences between the early leavers and the stayers in terms of their employment opportunities and annual incomes both before and after displacement, but that there are also differences within the group of the early leavers as well.

7 Estimating the heterogeneous displacement costs

7.1 Empirical strategy

Having identified that the early leavers are the workers who displace from closing firms during the two years prior to firm closure and the stayers are the workers who displace when the firm closes down, I move on to estimating the differences in displacement costs between these two groups of workers. I again take advantage of the panel data structure and use both individual- and time-fixed effects to control for individual characteristics (α_i) and calendar time-varying characteristics (θ_t). As explained in Part 6.1, this then controls for unobserved heterogeneity among the workers and common time-series variation in the outcome variable Y_{it}, which is now either the employment status of worker i in year t or the natural logarithm of the total annual pensionable income for worker i in year t. By using the natural

Figure 5: Average employment and income of displaced workers (decomposition)



Evolution of the employment rate and the average pensionable income for the workers displacing in 2000, the workers displacing in 2001, and the stayers in the period 1995 to 2009. The annual pensionable income is rebased to 2002-NOK using the consumer price index from Statistics Norway.

logarithm of the total annual income as an outcome variable, the interpretation of the effect of displacement on income is easier to compare across the early leavers and the stayers, who differ in terms of their pre-displacement annual income.

Figure 4 in Part 6.3 showed large variation in the difference in post-displacement outcomes of the early leavers and the stayers over time. Some of this may owe to the presence of UI institutions, which may imply time-variant post-displacement job search behavior of the displaced workers as discussed in Part 4. Moreover, comparisons of the effects of displacement among the workers may be sensitive to when in the business cycle the comparisons are made. Taking this into account, the following baseline double difference-in-difference model for estimating both the costs of displacement and the heterogeneous costs of displacement over the five years after displacement is estimated over the period $t = 1995$ to $t = 2009$ for the sample of separating and non-separating workers i who are employed in either a closing or a non-closing firm at some stage during the period $t \in [1995, 2001]$:

$$\begin{aligned} Y_{it} = \sum_{k=0}^{5} D_i \mathbf{1}\left\{t - s_i = k\right\} \delta_k + \sum_{k=0}^{5} D_i EL_i \mathbf{1}\left\{t - s_i = k\right\} \beta_k \\ + \underline{X}_{it} \underline{\eta} + \underline{Z}_{it} \underline{\chi} + \alpha_i + \theta_t + \varepsilon_{it} \end{aligned} \tag{11}$$

D_i is an indicator variable equal to 1 if worker i separates from a closing firm at some stage when selective labor turnover has started, i.e. for $t \in [2000, 2002]$, and 0 otherwise. EL_i is an indicator variable equal to 1 if worker i is an early leaver (a worker separating from a closing firm in either 2000 or 2001), and 0 if worker i is a stayer (a worker separating from a closing firm in 2002). The indicator variable $\mathbf{1}\left\{\bullet\right\}$ takes the value 1 if the argument is true, and 0 otherwise. The argument of the indicator variable, s_i, is the year when worker i displaces from the firm. ε_{it} is the error term of the model, which is assumed to have conditional mean zero. As in Part 6.1, estimation of Equation (11) is done by within-estimation.

Note that the displacement year s_i will be different for early leavers and stayers, so the control group of non-displacing workers will be different for the early leavers and the stayers. However, including the calendar time-varying characteristics θ_t ensures that aggregate time-series variation in the outcome variable is controlled for. If we in addition assume that the composition of the control group of non-displacing workers is stable throughout the period $t = 2000$ to $t = 2007$, the comparison of the displacement costs between the early leavers and the stayers will be valid. Note furthermore that since $\mathbf{1}\{\bullet\}$ is not identified for the control group of non-displacing workers, the control group is therefore effectively introduced into the model in Equation (11) through the time-fixed effects θ_t and the covariate vectors $\underline{X}_{it}$ and $\underline{Z}_{it}$ (von Wachter et al. 2009, p. 14).

The baseline model in Equation (11) is modified depending on which outcome variable Y_{it} is being considered. When using the employment status of the worker as the outcome variable, I include only the vector of observable (and potentially time-varying) individual characteristics for worker i in year t, $\underline{X}_{it}$. When using the natural logarithm of the total annual pensionable income for the worker as the outcome variable, I include both the vector $\underline{X}_{it}$ and a vector of observable (and potentially time-varying) firm characteristics $\underline{Z}_{it}$. The contents of these two vectors are described in Part 6.1. The coefficient δ_k measures the average cost due to displacement k years after being displaced compared to not being displaced. The coefficient β_k measures the additional average effect of being an early leaver on the displacement cost k years after being displaced compared to being a stayer. In other words, the coefficient β_k captures the heterogeneous costs of displacement between the early leavers and the stayers k years after being displaced.

7.2 Estimation results

I estimate Equation (11) using the employment status of the worker and the natural logarithm of the annual pensionable income as the outcome variables, with the key estimation results reported in Table 3. The reported Huber-White standard errors in parentheses are again clustered on the level of the individual worker due to potential serial correlation in $\varepsilon_{it} - \bar{\varepsilon}_i$. As a control group restriction, I exclude all workers in the control group who separate from non-closing firms that downsize by more than 30 percent (in terms of FTEs) from one year to the next during the time period when selective labor turnover in the closing firms has started.

From the second column in Table 3, we find that the reduction in the employment rate for the workers who displace from closing firms is statistically

Table 3: Costs of displacement

	Employment status	Log annual pensionable income
$D_i EL_i \mathbf{1}\{k=0\}$	-0.109***	0.032**
	(0.009)	(0.013)
$D_i EL_i \mathbf{1}\{k=1\}$	-0.012	0.022**
	(0.008)	(0.011)
$D_i EL_i \mathbf{1}\{k=2\}$	0.004	0.026**
	(0.008)	(0.010)
$D_i EL_i \mathbf{1}\{k=3\}$	0.003	0.037***
	(0.008)	(0.010)
$D_i EL_i \mathbf{1}\{k=4\}$	0.001	0.045***
	(0.007)	(0.010)
$D_i EL_i \mathbf{1}\{k=5\}$	-0.008	0.022**
	(0.007)	(0.010)
$D_i \mathbf{1}\{k=0\}$	-0.186***	-0.033***
	(0.006)	(0.010)
$D_i \mathbf{1}\{k=1\}$	-0.074***	-0.029***
	(0.005)	(0.007)
$D_i \mathbf{1}\{k=2\}$	-0.062***	-0.046***
	(0.005)	(0.007)
$D_i \mathbf{1}\{k=3\}$	-0.055***	-0.032***
	(0.005)	(0.007)
$D_i \mathbf{1}\{k=4\}$	-0.042***	-0.034***
	(0.005)	(0.007)
$D_i \mathbf{1}\{k=5\}$	-0.027***	-0.010
	(0.005)	(0.007)
Observations (n)	268,101	266,433
Cases ($n \times t$)	3,889,245	2,535,047
$\underline{X}_{it}$ included	Yes	Yes
$\underline{Z}_{it}$ included	No	Yes
R^2 (within)	0.01	0.27

Results from estimation of Equation (11) over the period $t = 1995$ to $t = 2009$. The dependent variable Y_{it} is specified in each column heading, with the annual income rebased to 2002-NOK using the consumer price index from Statistics Norway. All regressions control for individual-fixed effects and time-fixed effects, and all regressions have the control group restriction imposed. The reported Huber-White standard errors in parentheses are clustered on the individual level.
$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$

significant in the displacement year and up to the fifth year after displacement, i.e. for $k \in \{0, 1, 2, 3, 4, 5\}$, ranging from -2.7 percentage points in the fifth year after displacement ($k = 5$) to -18.6 percentage points in the displacement year ($k = 0$). This negative effect on the employment opportunities of the displaced workers decreases in absolute value in the five years following the displacement. This suggests that the negative effect of displacement on the employment opportunities of the workers is a persistent, but not permanent, effect.

Regarding the heterogeneous costs of displacement in terms of employment opportunities, we find from the second column in Table 3 that the employment rate for the early leavers is significantly lower compared to the stayers in the displacement year ($k = 0$) at -10.9 percentage points. In terms of the theoretical model in Part 3, this result may point to the job-to-unemployment effect being stronger than the job-to-job effect for the early leavers during the downsizing process, with the firm laying off workers who then transition directly into unemployment. However, the fact that the effect of displacement on employment opportunities for the early leavers is not significantly worse than the effect for the stayers in the later years, i.e. for $k \in \{1, 2, 3, 4, 5\}$, suggests that the early leavers are in general not worse off than the stayers in terms of employment opportunities.

When considering the natural logarithm of the annual pensionable income as the outcome variable in the third column in Table 3, there is a significant loss due to displacement in the four years following displacement and the displacement year itself, i.e. for $k \in \{0, 1, 2, 3, 4\}$, ranging from -2.9 percent in the first year after displacement ($k = 1$) to -4.6 percent in the second year after displacement ($k = 2$). The fact that the annual pensionable income due to displacement reaches its maximum only in the second year after displacement could owe to compensation agreements the displacing workers may have reached with their previous employer, such as for example severance pay, with such schemes lasting up to two years.[5]

[5] As a robustness check, I estimate Equation (11) with the nominal annual pensionable income as the outcome variable (the results are not reported). The results follow the same pattern as in Table 3, with the displacement costs ranging from NOK 6,248 in the fifth year after displacement ($k = 5$) to NOK 13,698 in the second year after displacement ($k = 2$). However, the only significant difference between the early leavers and the stayers is in the fourth year after

The early leavers are better off in terms of the annual income loss (measured in percent) due to displacement compared to the stayers in all the years after displacement and the displacement year itself, i.e. for $k \in \{0, 1, 2, 3, 4, 5\}$, with the income loss being from to 2.2 percent lower in the fifth year after displacement ($k = 5$) to 4.5 percent lower in the fourth year after displacement ($k = 4$) compared to the stayers. The significant difference of the effect of displacement on the annual pensionable income between the displaced and non-displaced workers and between the early leavers and the stayers in the displacement year ($k = 0$) in Table 3 must however be interpreted with some caution. This is because this difference may pick up both the effect of working at the firm right before closure and the immediate effect of displacement (Huttunen et al. 2006, p. 16). There could for example be agreements between the firm and the workers to enforce pay cuts to avoid downsizing. The estimated coefficients $\hat{\beta}_0$ and $\hat{\delta}_0$ may as such not represent the "pure" effects of displacement on the annual income of the displaced worker.

The effects of displacement on the annual pensionable income reported in the third column in Table 3 are implicitly conditioned on the worker having employment due to the inclusion of the vector of observable (and possibly time-varying) firm characteristics $\underline{Z}_{it}$. Since the employment differential between the displacing and the non-displacing workers is negative (as seen in the second column in Table 3), the displaced workers who succeed in finding work after displacement are likely to be a positively selected subset of the displacing workers. We would therefore expect the point estimates of the annual income loss due to displacement in the third column in Table 3 to understate the cost of displacement in terms of an unconditional income measure (Huttunen et al. 2006, p. 18). However, I consider the post-displacement annual pensionable income as an outcome variable of interest only when conditional on employment. If we were to consider the unconditional income measure, this would include non-employed workers who have transitioned into labor market support institutions, as discussed in Part 4. Although such a transition can also be considered as a displacement cost, I focus rather on how the costs of displacement are manifested in the labor

displacement ($k = 4$), with the displacement cost being NOK 6,821 lower for the early leavers.

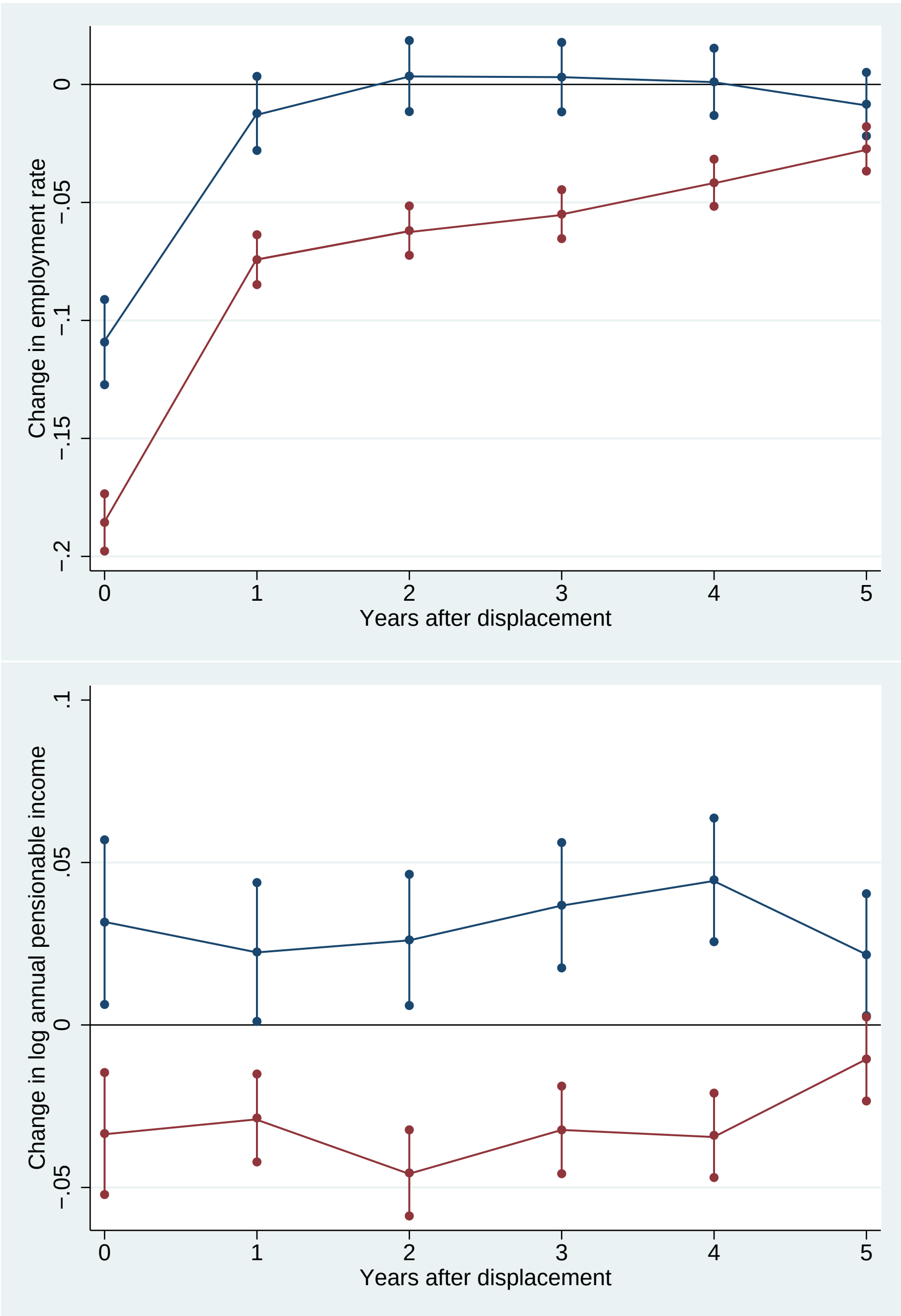
market, and not with regards to transitions to these institutions and non-market states. In this respect, I take the point of view that conditioning on labor market attachment as implicitly done in Table 3 gives a better estimate of these costs.

To summarize the results in Table 3, illustrated in Figure 6 as point estimates with associated 95 percent confidence intervals, displacement leads in general to statistically significant negative effects on both the employment opportunities and the annual pensionable income for the displacing workers compared to the non-displacing workers. The early leavers are better off in terms of the effect of displacement on the annual pensionable income compared to the stayers in the five years following the displacement year and the displacement year itself, i.e. for $k \in \{0, 1, 2, 3, 4, 5\}$. The early leavers are worse off in terms of the effect of displacement on the employment opportunities in the displacement year ($k = 0$) compared to the stayers, but this effect of displacement is not persistent.

Having estimated both the common (δ_k) and group-specific (β_k) marginal effects of displacement on employment opportunities and annual income, these effects must be considered jointly to investigate if the estimation results lend support to Proposition 3 and Proposition 4 of the theoretical model in Part 3.3 (which concerns the levels of the outcome variable, and not the changes as estimated above). To do so, I estimate Equation (11) with the employment status and the annual nominal pensionable income as outcome variables, and graph the average linear predictions for the early leavers and the stayers over the period 1995 to 2009, which is illustrated in the first and second panel in Figure 7.[6] The early leavers and the stayers differ only in terms of being early leavers or stayers, since we have controlled for both observable and non-observable individual characteristics in Equation (11). From the first panel, we see that the predicted employment rate of the early leavers is greater than for the stayers from 2003 to 2008. From the second panel, we see that the predicted annual income of the early leavers is greater than for the stayers from 2003 to 2008. The results illustrated in Figure 7 thus lend support to the propositions, albeit only over the period 2003 to 2008.

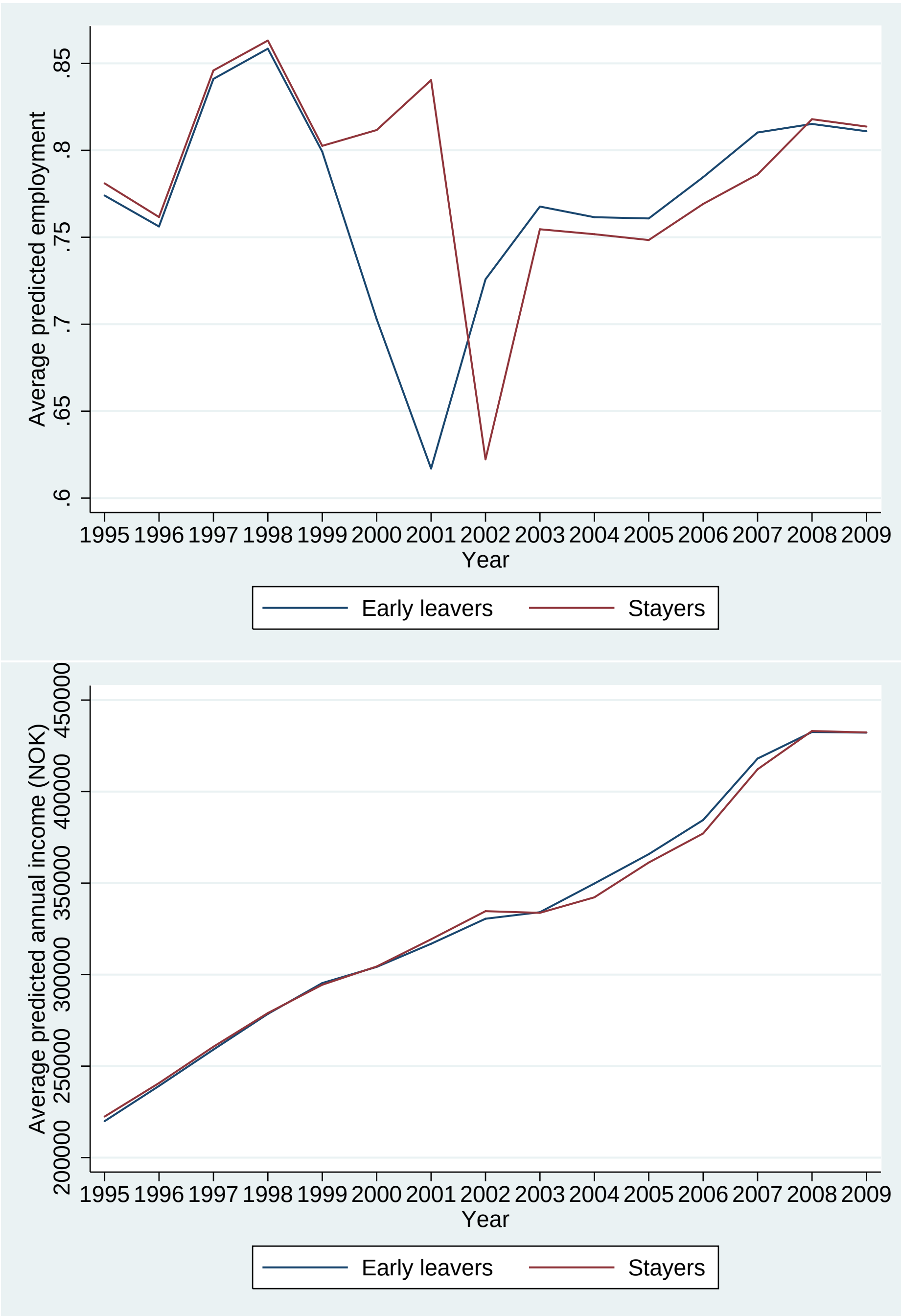
[6] Note that I have not tested the statistical significance of the joint coefficient $\delta_k + \beta_k$, so this exercise is only a visual inspection to see if the results lend support to the theoretical propositions.

Figure 6: Costs of displacement



The point estimates of the effect of displacement on the employment rate and the logarithm of annual pensionable income by year after displacement. The blue line represents the heterogeneous costs of displacement (β_k), and the maroon line represents the common costs of displacement (δ_k). The vertical lines represent the 95 percent confidence intervals of the point estimates.

Figure 7: Predicted employment and annual pensionable income



The linear prediction for the employment rate and the linear prediction for the average annual nominal pensionable income for the early leavers and the stayers from estimating Equation (11) with the control group restriction imposed. The annual income (implicitly conditioned on employment) is rebased to 2002-NOK using the consumer price index from Statistics Norway.

8 Constructing the human capital distributions

8.1 Empirical strategy

Having found that there are statistically significant and persistent heterogeneous costs of displacement in terms of the annual pensionable income, I now investigate if there are systematic differences in the human capital stock of the early leavers and the stayers that could explain this difference, as stipulated in the theoretical model in Part 3. Having thus far attempted to estimate models where the displacing workers differed only in terms of being either early leavers or stayers, I now want to consider how these two groups of workers differ in terms of their labor market assets. It may first be of interest to consider some pre-displacement characteristics of the early leavers and the stayers prior to the start of the selective labor turnover, meaning that we compare the two groups in 1999. These pre-displacement characteristics are summarized in Table 4.

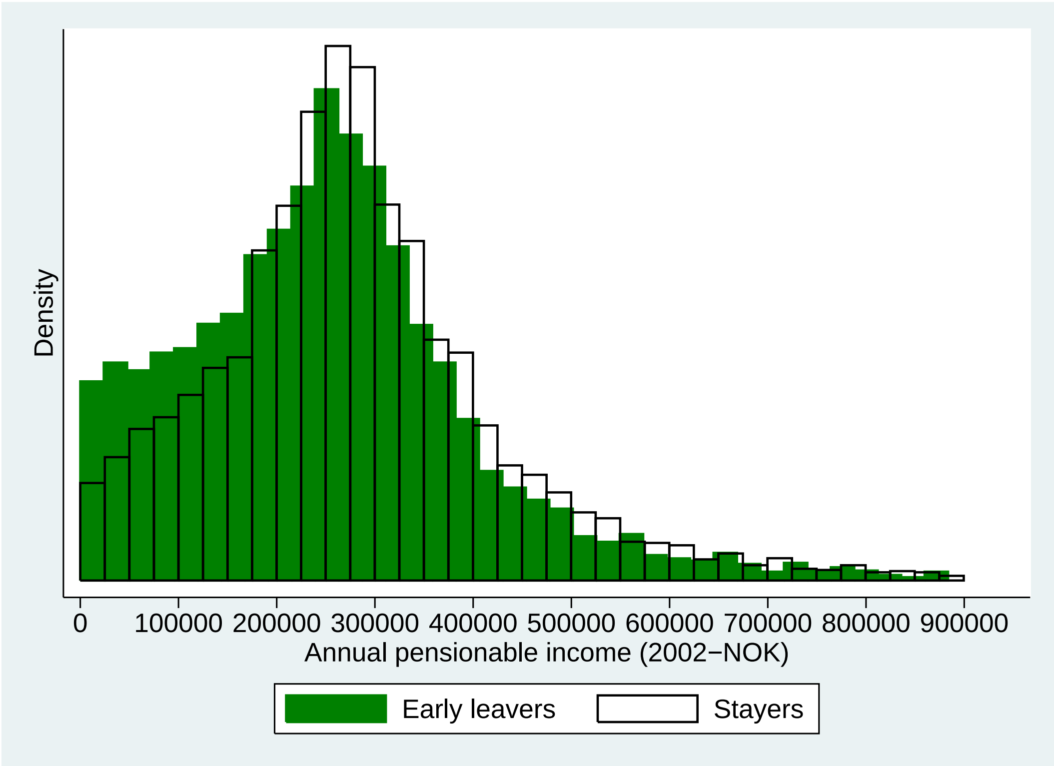
Table 4: Pre-displacement characteristics of displacing workers

	Early leavers	Stayers
No. of workers	5,985	6,245
Age of worker (years)	32.6	33.9
Education level:		
Compulsory	27.8	27.4
Secondary	48.5	44.8
College / university	23.7	27.8
Annual pensionable income (2002-NOK)	251,061	269,660
Unemployment spell in 1995 – 1998	41.7	35.3

Age and annual pensionable income are mean value for the displacing workers in 1999, with the annual pensionable income being rebased to 2002-NOK using the consumer price index from Statistics Norway. The unemployment spell incidence and education level are measured as percentage shares, with the education level measured in 1999.

Interestingly, when considering Table 4 the stayers seem to be better off than the early leavers when considering their annual pensionable income (NOK 269,660 and NOK 251,061, respectively) and incidence of unemployment spells in the period 1995 to 1998 (35.3 percent and 41.7 percent, respectively), and with more having college or university education (27.8 percent and 23.7 percent, respectively). Furthermore, the income distributions of the employed early leavers and the employed stayers are illustrated in the histogram of the annual nominal pensionable income in 1999 (measured in 2002-NOK) in Figure 8.

Figure 8: Pre-displacement income distributions



Histogram of the annual nominal pensionable income (conditional on employment and measured in 2002-NOK) for the early leavers and the stayers in 1999, right-censored at NOK 900,000.

The fact that the pre-displacement annual income of the stayers in general seem to be higher than the pre-displacement annual income of the early leavers in Figure 8 could suggest that the stayers have more human capital than the early leavers. However, whether this reflects actual systematic differences between the early leavers and the stayers in terms of human capital may better be investigated by constructing human capital distributions based on a theoretical framework.

To construct these human capital distributions, let us consider the concept of human capital through the "skills-weight"-framework of Lazear (2003). In this framework, a worker i has a finite set of skills $C_i = \{c_{i1}, c_{i2}, \ldots, c_{iK}\}$ that is considered as general human capital upon starting a job. However, the firm weights the skills of the worker according to the needs of the firm. By allowing the worker to expand upon and add to certain skills through for example firm-specific on-the-job training, the skills set of the worker becomes relatively more firm-specific over time, and the worker's stock of firm-specific human capital therefore increases at the expense of general human capital.[7] Firm-specific human capital is often operationalized as firm tenure, but this operationalization only takes into account the temporal condition that necessarily must be satisfied for a worker's human capital stock to have become relatively more firm-specific through specialization of the skills set C_i. In other words, firm tenure in general often indicates that a worker has firm-specific human capital, but it does not guarantee it. Using the skills-weight approach, let us write the firm's valuation of the finite set of worker skills C_i in period t as the sum of each skill component c_{ikt} multiplied by the skill component rental rate r_{fct} for component c by firm f in period t as:

$$w_{ift} = \sum_{k=1}^{K} r_{fct} c_{ikt} \tag{12}$$

Let now $\sum_{k=1}^{K} r_{fct} = r_{ft}$ in Equation (12) and assume that the skills set of worker i in period t can be captured by a human capital stock scalar hc_{it}. Following Abowd et al. (2002), the income of worker i in firm f in period t, w_{ift}, can then be expressed as the rent obtained from the human capital stock of worker i in period t at the rental rate from the employing firm f in period t, r_{ft}:

$$w_{ift} = r_{ft} hc_{it} \tag{13}$$

[7] The finiteness of the worker's skills set C_i can be seen as the underlying reason why a high level of general human capital is associated with a low level of firm-specific human capital, as stipulated in Assumption 2 of the theoretical model in Part 3.2.

Let us now assume that the human capital of worker i in period t, hc_{it}, is produced by combining a vector containing labor force experience, $\underline{Q}_{it}$, and the person-specific input, α_i, according to the exponential production function:

$$hc_{it} = \exp\left(\alpha_i + \underline{Q_{it}\beta}\right) \tag{14}$$

Combining Equation (13) and Equation (14) and taking logarithms yields the following expression for the relationship between the logarithm of the income of worker i in firm f in period t and the human capital of worker i in period t:

$$\ln\left(w_{ift}\right) = \ln\left(r_{ft}\right) + \alpha_i + \underline{Q_{it}\beta} \tag{15}$$

The right-hand side of Equation (15) captures all the factors which are compensated through the period wage, and therefore encompasses both the general human capital, ghc_{it}, and the firm-specific human capital, fhc_{it}, of worker i in period t (Abowd et al. 2002, p. 12). To consider how the early leavers and the stayers differ in terms of their human capital stocks prior to displacement from the closing firms, I construct empirical human capital distributions for these two groups based on the theoretical framework outlined above. To construct these distributions, I follow Abowd et al. (2002, p. 10) and estimate the following three-way fixed effects model over the sample of both the early leavers and the stayers:

$$\ln\left(Y_{it}\right) = \underline{Q_{it}\beta} + \alpha_i + \phi_f + \theta_t + \varepsilon_{ift} \tag{16}$$

Y_{it} is the total annual pensionable income of worker i in year t. θ_t is the time-fixed effect for year t, ϕ_f is the firm-fixed effect for firm f, and ε_{ift} is the error term of the model. $\underline{Q}_{it}$ is a vector of observable (and potentially time-varying) individual characteristics for worker i in year t that can explain wage compensation in the labor market due to experience. I include age, age squared, and level of education in this vector.[8] Together with the individual-fixed effect α_i for worker i, this then constitutes the explanatory components of the external wage compensated by the

[8] In contrast to the empirical strategy in Part 6.1, I now want to obtain the predicted effect of age on the annual pensionable income to create the human capital distributions, and not merely control for age to increase the comparability among the displacing workers.

labor market (Lengermann and Vilhuber 2002, p. 10). The firm-fixed effect ϕ_f captures the unobserved explanatory component of the internal wage which is compensated by the firm f that hires the worker, meaning that it captures the firm-specific rental rate r_{ft} in Equation (15). The model is estimated over the period $t = 1995$ to $t = 1999$ to compare the workers prior to displacement.

After estimating Equation (16), the predicted core measure of human capital $\hat{h}_{it} \equiv \hat{\alpha}_i + \underline{Q_{it}\hat{\beta}}$ can be used to construct the two empirical human capital distributions (Abowd et al. 2005, p. 156). What this predicted core measure of human capital $\hat{h}_{it}$ measures is not immediately clear from the specification in Equation (16). Age and age squared can be interpreted as proxy variables for the general labor market experience of the worker, which may or may not be a general labor market asset portable across different jobs. The level of education may be both a general labor market asset or an industry-specific labor market asset, depending for example on the specificity of the worker's education.

In light of this and following Lengermann and Vilhuber (2002, p. 10), I interpret the general human capital component of a worker i as being well-proxied by the individual-fixed effect α_i that per definition does not change over time and which is therefore portable across jobs. I therefore construct two empirical general human capital distributions based on the predicted measure $\hat{h}_i = \hat{\alpha}_i$. These individual-fixed effects could include for example ability, social capital and effort (Abowd et al. 2002, p. 12). It is imperative to note that these individual-fixed effects do not necessarily fully capture the general human capital stock of the worker, as there may well be time-variant individual assets that are also considered general human capital. These individual-fixed effects have however been found to be more important in explaining variation in wage levels than the component of observable (and possibly time-varying) individual characteristics $\underline{Q_{it}\hat{\beta}}$ in the predicted core measure of human capital (Abowd et al. 2005, p. 167).

8.2 Estimation results

I first estimate Equation (16) over the sample of early leavers and stayers using the graph-theoretical algorithm proposed by Abowd et al. (2002) and implemented by

Cornelissen (2008). The idea behind this graph-theoretical estimation strategy is best described by Abowd et al. (2002) themselves:

> Start with an arbitrary individual and include all firms that he or she ever worked for. Next, add all individuals who ever worked in any of those firms. Continue adding all additional firms that any of these individuals ever worked for and all additional individuals in any of those firms until no more individuals or firms can be added to the current group (p. 19).

After doing so, the design matrix is reorganized such that each group is placed in an ascending order. However, when using matched administrative data with firms and workers, within-estimation of the normal equations from this estimation strategy entails inversion of a very high-dimensional design matrix. The implementation by Cornelissen (2008) takes advantage of the fact that the high-dimensional fixed-effects matrix for the firm-fixed effects is typically a sparse matrix, since a worker is usually employed in very few firms over time. Therefore, many of the firm dummy coefficients equal zero for this worker and the system of normal equations can therefore be greatly simplified as a result (Cornelissen 2008, p. 172-175).

In the sample of early leavers and stayers, there is sufficient connectedness of workers across firms to identify the individual-fixed effects and firm-fixed effects for 93 percent of the displacing workers. The results of estimating Equation (16) are not reported since we are only interested in the distribution of the point estimates, but a summary of correlations and standard deviations of the variables in Equation (16) is shown in Table 5. Table 5 only reports the results for the largest group of connected firms and workers, as comparison across groups does not readily allow for interpretation since different groups are per definition unconnected.[9]

Notably, in contrast to Abowd et al. (2005, p. 167), we find from the third column in Table 5 that the observable time-varying individual characteristics are more correlated with the logarithm of the annual income than the individual-fixed effects ($r = 0.478$ and $r = 0.280$, respectively). This suggests that the components

[9] The largest group of connected firms and workers contain 75 percent of the sample of displacing workers, with 72 percent of these changing employer in the time period 1995 to 1999.

Table 5: Summary of human capital correlations

Component	Standard deviation	Correlation with $\ln(Y_{it})$	$\underline{Q_{it}\beta}$	α_i	ϕ_f	ε_{ift}
$\ln(Y_{it})$	0.797	1.000	–	–	–	–
$\underline{Q_{it}\beta}$	0.814	0.478	1.000	–	–	–
α_i	0.776	0.280	-0.405	1.000	–	–
ϕ_f	0.580	0.307	0.088	-0.505	1.000	–
ε_{ift}	0.273	0.345	-0.026	0.000	0.000	1.000

Empirical correlations for the largest connected group (n = 9,255) from estimating Equation (16) over the period $t = 1995$ to $t = 1999$ for all displacing workers.

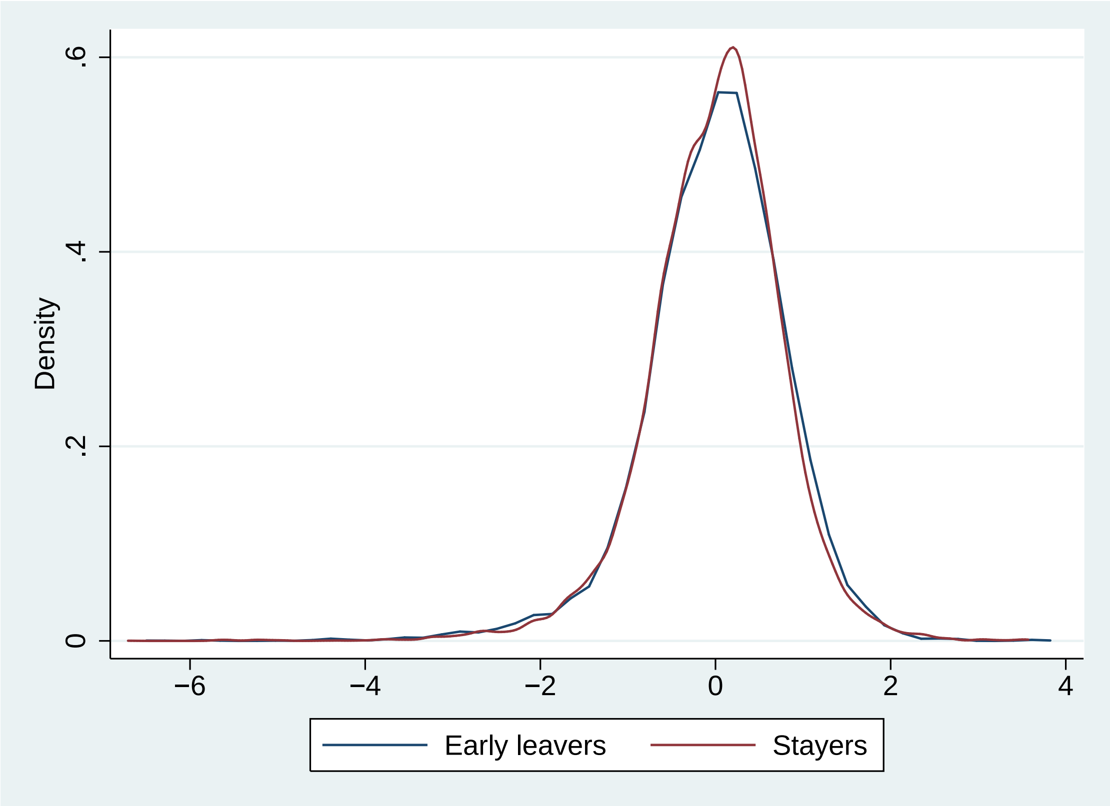
of $\underline{Q}_{it}$ are more important for explaining the variation in annual income than the proxy for general human capital, α_i. Indeed, if we consider the variance decompositions $\text{cov}\left(\underline{Q_{it}\beta}, \ln(Y_{it})\right)/\text{var}(\ln(Y_{it}))$ and $\text{cov}(\alpha_i, \ln(Y_{it}))/\text{var}(\ln(Y_{it}))$, we find that the observable time-varying individual characteristics explain a fraction 0.223 of the variation in the logarithm of annual pensionable income, whereas the individual-fixed effects only explain a fraction 0.157 of the variation. Furthermore, the negative correlation between the firm-fixed effects and the individual-fixed effects ($r = -0.505$) in the fifth column in Table 5 indicate that the high-income workers are not necessarily associated with the firms that pay high salaries, meaning that there is evidence of negative assortative matching.[10]

Using the predicted measure of general human capital for each worker i from the estimation of Equation (16), $\hat{h}_i = \hat{\alpha}_i$, I use semi-parametric density estimation to construct the two empirical general human capital distributions using a Gaussian kernel with a smoothing bandwidth $\hat{b}_{rule} = 0.098$ derived from Silverman's rule-of-thumb (see Appendix C.1 for derivation of the chosen bandwidth). The

[10] However, as noted by Andrews et al. (2008), the estimated correlation between ϕ_f and α_i in Table 5 may be downward biased if there is limited mobility of workers ("limited mobility bias"), which may be a concern for the largest connected group identified from estimation of Equation (16). Therefore, we may have that there is actually positive assortative matching going on.

resulting empirical general human capital distributions for the early leavers and the stayers are illustrated in Figure 9. As we see from Figure 9, the distributions for the early leavers and the stayers are unimodal and centered around zero, where the centering follows by construction from the implementation by Cornelissen (2008). Furthermore, we see that the general human capital distribution of the early leavers is seemingly right-shifted compared to the distribution of the stayers.

Figure 9: General human capital distributions



General human capital distributions constructed from estimating Equation (16) over the sample of early leavers and stayers over the period $t = 1995$ to $t = 1999$, using the predicted measure $\hat{h}_i = \hat{\alpha}_i$. The distributions are constructed semi-parametrically using a Gaussian kernel with a bandwidth $\hat{b}_{rule} = 0.098$.

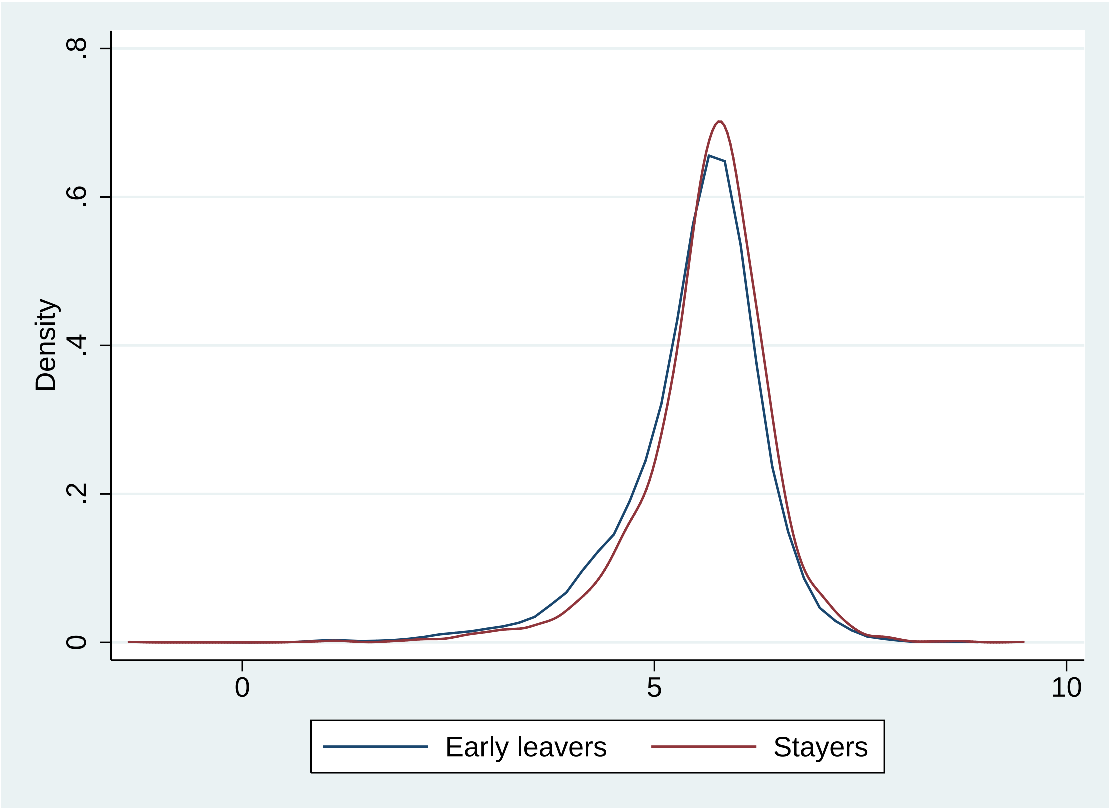
In order to establish whether this visual inspection is correct, I perform a Kolmogorov-Smirnov test of equality of the general human capital distributions in Figure 9. For the non-directional hypothesis that the distributions differ, I obtain a p-value less than 0.001 and thus evidence that there are significant differences in the distributions of general human capital of the early leavers and the stayers. For the directional hypothesis that the cumulative density function of the stayers lies below the cumulative density function of the early leavers, I obtain a p-value of less than 0.001. This provides evidence that the general human capital distribution of the early leavers is right-shifted compared to the distribution of the stayers. This result is in line with the results of Lengermann and Vilhuber (2002).[11]

Given the high correlation between the logarithm of annual pensionable income $\ln(Y_{it})$ and the observable (and potentially time-varying) individual characteristics $\underline{Q}_{it}\underline{\beta}$ in Table 5, it may be interesting to construct the empirical core human capital distributions as well. Using the predicted core measure from the estimation of Equation (16), $\hat{h}_{it} = \hat{\alpha}_i + \underline{Q}_{it}\hat{\underline{\beta}}$ (evaluated at $t = 1999$ to ensure comparability prior to displacements), I use semi-parametric density estimation to construct the two empirical core human capital distributions using a Gaussian kernel with a smoothing bandwidth $\hat{b}_{rule} = 0.138$ derived from Silverman's rule-of-thumb (see Appendix C.1 for derivation of the chosen bandwidth). The resulting distributions for the early leavers and the stayers in 1999 are illustrated in Figure 10.

In contrast to the general human capital distributions in Figure 9, the core human capital distribution of the stayers in Figure 10 seems to be right-shifted compared to the core human capital distribution of the early leavers. A p-value less than 0.001 from the non-directional Kolmogorov-Smirnov test confirms that the distributions are significantly different. For the directional hypothesis that the cumulative density function of the early leavers lies below the cumulative density function of the stayers, the p-value from the Kolmogorov-Smirnov test is less than 0.001. This then provides evidence that the core human capital distribution of the stayers is right-shifted compared to the distribution of the early leavers.

[11] We must however note that the Kolmogorov-Smirnov tests will not take into account that the distributions are estimated distributions, and will therefore reject the null hypothesis too often. Furthermore, since the vertical distance between the two cumulative density functions may be at a maximum when the estimation error in Equation (16) is at its maximum, the results of these tests must be interpreted with some caution.

Figure 10: Core human capital distributions



Core human capital distributions constructed from estimating Equation (16) over the sample of early leavers and stayers over the period $t = 1995$ to $t = 1999$, using the predicted measure $\hat{h}_{it} = \hat{\alpha}_i + \underline{Q_{it}\hat{\beta}}$ evaluated at $t = 1999$. The distributions are constructed semi-parametrically using a Gaussian kernel with a bandwidth $\hat{b}_{rule} = 0.138$.

Despite the general human capital stock of the early leavers being greater than the general human capital stock of the stayers, as illustrated in Figure 9, the right-shifted income distribution of the stayers compared to the early leavers as illustrated in the histogram in Figure 8 indicates that the stayers have a greater stock of some other valuable labor market asset prior to displacement compared to the early leavers, such as for example industry-specific human capital or firm-specific human capital. The fact that the core human capital distribution of the stayers is right-shifted compared to the distribution of the early leavers, as illustrated in Figure 10, supports this hypothesis.

To investigate this, we can perform a decomposition of the vector of observable (and potentially time-varying) individual characteristics into its components. Doing so, we find that the proxies for labor market experience (age and age squared) correlate more with the logarithm of annual income than the level of education does ($r = 0.300$ and $r = 0.142$, respectively). To the extent that the labor market experience of the workers stem from being hired in the same firm, this suggests that $\underline{Q}_{it}$ mainly represents the firm-specific human capital of worker i in year t. Since the stayers have longer average firm tenure compared to the early leavers (757 days and 598 days prior to displacement from the firm, respectively) and taking into account the right-shifted core human capital distribution of the stayers, this points to the firm-specific human capital stock of the stayers being greater than the firm-specific human capital stock of the early leavers.[12]

These result then lend support to Proposition 2 of the theoretical model in Part 3.2, since the stayers have comparably low levels of general human capital and high levels of firm-specific human capital, meaning that the labor force upon closure is less "skilled" than it was prior to downsizing. In the theoretical framework, the comparably low level of general human capital could explain why the negative effect of displacement on the annual pensionable income for the stayers is greater than the effect for the early leavers (as illustrated in the second panel in Figure 6 in Part 7.2), as the stayers then have less of the only post-displacement marketable asset. The fact that the stayers in general have higher pre-displacement annual incomes than the early leavers (as illustrated in the histogram in Figure 8) could on the other hand be explained by the return to their firm-specific human capital prior to displacement. The higher pre-displacement annual income may also point to the fact that firm-worker match value for the stayers is high, and that the downsizing strategy as stipulated in Assumption 1 of the theoretical model in Part 3.1 is an

[12] Going back to the definition from Lazear (2003, p. 1), the firm-specific human capital of a worker is nullified upon displacement from a firm. Ideally, one could therefore identify a shift in the predicted wage distribution ($\widehat{\ln(Y_{it})} = \underline{Q}_{it}\underline{\hat{\beta}} + \hat{\alpha}_i + \hat{\phi}_f$) for early leavers and stayers who separate from a firm and are hired in a new firm the year after during the period 1995 to 1999. However, given the size of the largest connected group identified from estimating Equation (16), the resulting distributions contain more noise than information due to small-sample bias.

intended optimal strategy from the firm's side, and not only the firm honoring the "last-in, first-out" norm as discussed in Part 4 (Borjas 2013, p. 273).

9 Conclusion

It is well-documented that job displacement due to firm closure leads to detrimental effects on the labor market outcomes of the displaced workers (Kuhn 2002). However, many of these studies do not take into account the selection process prior to the closure of the firm. Focusing only on the ultimately displaced workers will therefore lead to biased estimates of the costs of displacement. Using Norwegian data, I identify the selective labor turnover related to firm closure starting two years prior to the closure of the firm, in line with Huttunen et al. (2011).

I find that all workers displacing from closing firms suffer from persistent annual income losses due to displacement ranging from −2.9 percent to −4.6 percent compared to the non-displaced workers, and a reduction in the employment rate ranging from −2.7 percentage points to −18.6 percentage points compared to the non-displaced workers. In terms of the effect of displacement on annual income, I find evidence that the ultimately displaced workers suffer from an annual income loss after displacement 2.2 percent to 4.5 percent greater than the annual income loss of the workers who displace from the firm earlier in the downsizing process. In terms of the effect of displacement on the employment opportunities of the workers, there is no evidence of persistent and significant differences among the two groups of displaced workers, except in the displacement year itself.

The ultimately displaced workers are found to have a greater stock of firm-specific human capital and a lower stock of general human capital compared to the workers who displace earlier in the downsizing process, which may explain the difference between the workers in terms of the effect of displacement on the annual income. In terms of pre-displacement productivity the ultimately displaced workers seem to be positively selected (due to higher pre-displacement annual incomes), but in terms of post-displacement productivity they seem to be negatively selected (due to higher cost of displacement in terms of annual income).

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Appendix A

A.1 Derivation of $\partial e_i^o / \partial ghc_i$ in Equation (4)

Differentiating the expression in Equation (3) by parts and suppressing the arguments of $w(\bullet)$, $w^n(\bullet)$, $\lambda(\bullet)$, $p(\bullet)$ and $C(\bullet)$ to ease notation, we find that:

$$p''\left(e_i^o\right)\left[\frac{\partial e_i^o}{\partial ghc_i}\right]\left[w^n - \lambda w\right] + p'\left(e_i^o\right)\left[\left[\frac{\partial w^n}{\partial ghc_i}\right] - \lambda\left[\frac{\partial w}{\partial ghc_i}\right]\right] = C''\left(e_i^o\right)\left[\frac{\partial e_i^o}{\partial ghc_i}\right]$$

$$\Downarrow$$

$$\left[\frac{\partial e_i^o}{\partial ghc_i}\right]\left[p''\left(e_i^o\right)\left[w^n - \lambda w\right] - C''\left(e_i^o\right)\right] = -p'\left(e_i^o\right)\left[\left[\frac{\partial w^n}{\partial ghc_i}\right] - \lambda\left[\frac{\partial w}{\partial ghc_i}\right]\right]$$

$$\Downarrow$$

$$\frac{\partial e_i^o}{\partial ghc_i} = -\frac{p'\left(e_i^o\right)\left[\left[\frac{\partial w^n}{\partial ghc_i}\right] - \lambda\left[\frac{\partial w}{\partial ghc_i}\right]\right]}{p''\left(e_i^o\right)\left[w^n - \lambda w\right] - C''\left(e_i^o\right)} > 0$$

The sign of this result requires the assumption that the effect of having one more unit of general human capital ghc_i leads to an increase in the wage level of a worker at the new potential employer compared to the current employer satisfying:

$$\frac{\partial w^n}{\partial ghc_i} > \lambda\left[\frac{\partial w}{\partial ghc_i}\right]$$

A.2 Derivation of $\partial e_i^o / \partial fhc_i$ in Equation (5)

Differentiating the expression in Equation (3) by parts and suppressing the arguments of $w(\bullet)$, $w^n(\bullet)$, $\lambda(\bullet)$, $p(\bullet)$ and $C(\bullet)$ to ease notation, we find that:

$$p''\left(e_i^o\right)\left[\frac{\partial e_i^o}{\partial fhc_i}\right]\left[w^n - \lambda w\right] - p'\left(e_i^o\right)\left[\left[\frac{\partial \lambda}{\partial fhc_i}\right] w + \lambda\left[\frac{\partial w}{\partial fhc_i}\right]\right] = C''\left(e_i^o\right)\left[\frac{\partial e_i^o}{\partial fhc_i}\right]$$

$$\Downarrow$$

$$\left[\frac{\partial e_i^o}{\partial fhc_i}\right]\left[p''\left(e_i^o\right)\left[w^n - \lambda w\right] - C''\left(e_i^o\right)\right] = p'\left(e_i^o\right)\left[\left[\frac{\partial \lambda}{\partial fhc_i}\right] w + \lambda\left[\frac{\partial w}{\partial fhc_i}\right]\right]$$

$$\Downarrow$$

$$\frac{\partial e_i^o}{\partial fhc_i} = \frac{p'\left(e_i^o\right)\left[\left[\frac{\partial \lambda}{\partial fhc_i}\right] w + \lambda\left[\frac{\partial w}{\partial fhc_i}\right]\right]}{p''\left(e_i^o\right)\left[w^n - \lambda w\right] - C''\left(e_i^o\right)}$$

The sign of the result above depends critically on the effect of firm-specific human capital on the probability of layoff compared to the effect of firm-specific human capital on the current wage level. More to the point, the sign of the effect above is negative if we assume the effects of firm-specific human capital satisfy:

$$\left[\frac{\partial \lambda}{\partial fhc_i}\right] w + \lambda \left[\frac{\partial w}{\partial fhc_i}\right] > 0$$

$$\Downarrow$$

$$\frac{\partial \lambda}{\partial fhc_i} > -\frac{\lambda}{w}\left[\frac{\partial w}{\partial fhc_i}\right]$$

This does not follow directly from an economic point of view, but let us try to parametrize the expression above to see if it may plausibly hold. For simplicity, let us assume that the marginal return to firm-specific human capital can be proxied by the usual marginal return to general human capital, namely the return to schooling (measured as annuity value). The return to schooling in terms of annual income has from Norwegian data been estimated to be in the interval NOK 4,000 to NOK 9,000 (measured in 2010-NOK) for men in the 1943-1963 cohorts, depending on the estimation strategy (Bhuller et al. 2014, p. 14). The average annual income for the full sample of workers that Bhuller et al. (2014) consider was NOK 402,000. Recalling that $\lambda \in (0,1)$ we have that the value of the right-hand side of the equation above therefore lies in the interval:

$$-\frac{\lambda}{w}\left[\frac{\partial w}{\partial fhc_i}\right] \in (-0.022, 0)$$

If we now apply the necessary condition for increased firm-specific human capital from Part 8.1, namely firm tenure, the expression above tells us that if the probability of a worker being laid off decreases by 2.2 percentage points or less for each year the worker has been in the firm, we will have that $\partial e_i^o / \partial fhc_i < 0$. This conclusion hinges critically on the return to firm-specific human capital being proxied sufficiently well by the return to general human capital and that there is a one-to-one correspondence between firm tenure and firm-specific human capital. I make use of the assumption that $\partial e_i^o / \partial fhc_i < 0$, although the plausibility of the parametrization behind this assumption may of course be questioned.

A.3 Derivation of $\partial e_i^o / \partial ghc_i$ in Equation (9)

Differentiating the first-order condition of the maximization problem in Equation (8) by parts and suppressing the arguments of $w(\bullet)$, $w^n(\bullet)$, $\lambda(\bullet)$, $p(\bullet)$ and $C(\bullet)$ to ease notation, we find that:

$$p''(e_i^o)\left[\frac{\partial e_i^o}{\partial ghc_i}\right][w^n - b] + p'(e_i^o)\left[\left[\frac{\partial w^n}{\partial ghc_i}\right] - \left[\frac{\partial b}{\partial w}\right]\left[\frac{\partial w}{\partial ghc_i}\right]\right] = C''(e_i^o)\left[\frac{\partial e_i^o}{\partial ghc_i}\right]$$

$$\Downarrow$$

$$\left[\frac{\partial e_i^o}{\partial ghc_i}\right][p''(e_i^o)[w^n - b] - C''(e_i^o)] = -p'(e_i^o)\left[\left[\frac{\partial w^n}{\partial ghc_i}\right] - \left[\frac{\partial b}{\partial w}\right]\left[\frac{\partial w}{\partial ghc_i}\right]\right]$$

$$\Downarrow$$

$$\frac{\partial e_i^o}{\partial ghc_i} = -\frac{p'(e_i^o)\left[\left[\frac{\partial w^n}{\partial ghc_i}\right] - \left[\frac{\partial b}{\partial w}\right]\left[\frac{\partial w}{\partial ghc_i}\right]\right]}{p''(e_i^o)[w^n - b] - C''(e_i^o)} > 0$$

The sign of this result requires the assumption that the effect of having one more unit of general human capital ghc_i leads to an increase in the wage level of a worker at the new potential employer compared to the previous employer satisfying:

$$\frac{\partial w^n}{\partial ghc_i} > \left[\frac{\partial b}{\partial w}\right]\left[\frac{\partial w}{\partial ghc_i}\right]$$

Appendix B

B.1 Identifying the early leavers and the stayers

The results of estimating Equation (10) with the nominal annual pensionable income as the outcome variable is summarized in Table 6, where I first estimate the model without either $\underline{X}_{it}$ nor $\underline{Z}_{it}$ (first panel), with only $\underline{X}_{it}$ (second panel) and finally with both $\underline{X}_{it}$ and $\underline{Z}_{it}$ (third panel). The results for the two first specifications (in the first row of the first and second panel in Table 6, respectively) point to the selective labor turnover directly related to the downsizing process starting three years prior to the base year, since the post-displacement annual pensionable income of the separating workers from the closing and the non-closing firms are significantly different for $d \in \{1, 2, 3\}$.

Table 6: Selective labor turnover (annual pensionable income)

	$d = 1$	$d = 2$	$d = 3$	$d = 4$
$T_{it}D_i$	4152.7^{**}	8516.4^{***}	7875.9^{***}	2847.0
	(2109.5)	(2730.3)	(2683.5)	(3316.1)
T_{it}	112313.1^{***}	96547.9^{***}	85686.7^{***}	72212.7^{***}
	(1462.5)	(1413.2)	(1221.2)	(1419.3)
Observations (n)	24,297	19,145	23,625	20,678
Cases ($n \times t$)	346,600	272,834	341,139	302,266
R^2 (within)	0.11	0.12	0.12	0.14
	$d = 1$	$d = 2$	$d = 3$	$d = 4$
$T_{it}D_i$	4918.2^{**}	7770.8^{***}	6318.0^{**}	4168.6
	(1947.2)	(2727.0)	(2677.9)	(3222.8)
T_{it}	91048.0^{***}	81250.7^{***}	72367.5^{***}	52956.6^{***}
	(1483.3)	(1477.4)	(1312.1)	(1423.5)
Observations (n)	23,984	18,875	23,327	20,618
Cases ($n \times t$)	338,120	266,684	334,250	298,956
$\underline{X}_{it}$ included	Yes	Yes	Yes	Yes
R^2 (within)	0.13	0.14	0.13	0.15
	$d = 1$	$d = 2$	$d = 3$	$d = 4$
$T_{it}D_i$	-6132.1^{***}	-222.1	6778.4^{*}	3992.8
	(2283.1)	(5998.6)	(3501.6)	(4372.7)
T_{it}	59383.7^{***}	26436.8^{***}	16351.8^{***}	41635.0^{***}
	(2037.8)	(2277.9)	(2059.5)	(2014.3)
Observations (n)	23,781	18,374	23,119	19,744
Cases ($n \times t$)	169,916	124,038	163,659	140,145
$\underline{X}_{it}$ included	Yes	Yes	Yes	Yes
$\underline{Z}_{it}$ included	Yes	Yes	Yes	Yes
R^2 (within)	0.13	0.17	0.18	0.19

Results from estimation of Equation (10) over the period $t = 1995$ to $t = 2009$ with the dependent variable Y_{it} being the total annual pensionable income for worker i in year t. The annual income is rebased to 2002-NOK using the consumer price index from Statistics Norway. All regressions control for individual-fixed effects and time-fixed effects, and all regressions have the control group restriction from Part 6.1 imposed. The reported Huber-White standard errors in parentheses are clustered on the individual level.

$^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$

However, the preferred specification used to control for firm heterogeneity in wage setting in the third panel in Table 6 does not point to any clear pattern, as there are only significant differences between the separating workers from closing and non-closing firms three years prior to the base year ($d = 3$) and one year prior to the base year ($d = 1$). From this third panel, the workers separating from closing firm three years prior to the base year are better off compared to the workers separating from non-closing firms in the same year by NOK 6,778. However, the workers separating from closing firms the year prior to the base year are worse off compared to the workers separating from non-closing firms in the same year by NOK 6,132. This change in sign of the additional effect of separating from a closing firm on the annual income compared to separating from a non-closing firm leads me to not draw any new conclusions regarding when the selective labor turnover related to firm closure starts compared to the results in Table 2. From the second row of the third panel in Table 6, we see that the annual income for all displacing workers in $d \in \{1, 2, 3, 4\}$ increase after displacement, with the annual income increase ranging from NOK 16,352 to NOK 59,384. This positive effect reflects the growth in real annual income in the Norwegian economy over the time period considered, which is illustrated in the four panels in Figure 3.

As a sensitivity analysis, I estimate Equation (10) with the annual pensionable income being conditional on employment, while also imposing the control group restriction from Part 6.1. As expected, the results of the third panel in Table 6 do not change when performing this sensitivity analysis. However, the conditioning on employment makes the estimated coefficient $\hat{\delta}$ statistically insignificant for $d \in \{2, 3, 4\}$ in the first and second panel in Table 6. As a robustness check, I estimate Equation (10) with the natural logarithm of the annual pensionable income as the outcome variable, both unconditional and conditional on having employment, while also imposing the control group restriction from Part 6.1. Performing this robustness check yields no changes in the qualitative conclusions reached in Table 2 regarding when the selective labor turnover starts when considering the preferred specification including both $\underline{X}_{it}$ and $\underline{Z}_{it}$.

Appendix C

C.1 Derivation of $\hat{b}_{rule}$ for Figure 9 and Figure 10

The choice of bandwidth $\hat{b}_{rule}$ is derived from the following rule-of-thumb formula suggested by Silverman (Hansen 2014, p. 333), where $\hat{\sigma}_{\hat{\alpha}}$ is the standard error of the predicted core measure individual-fixed effects $\hat{\alpha}_i$ from estimation of Equation (16) and n is the number of observations used to estimate $\hat{\alpha}_i$:

$$b_{rule} = 1.06\hat{\sigma}_{\hat{\alpha}} n^{-(1/5)}$$

Inserting the relevant numbers yields the following rule-of-thumb smoothing bandwidth used in the semi-parametric estimation of the empirical general human capital distributions illustrated in Figure 9 in Part 8.2:

$$\hat{b}_{rule} = 1.06\,(0.787)\,(44044)^{-(1/5)} = 0.098$$

For the empirical core human capital distributions illustrated in Figure 10 in Part 8.2, the rule-of-thumb smoothing bandwidth is derived as follows, where the core measure of human capital, $\hat{h}_{it}$, is evaluated at $t = 1999$:

$$b_{rule} = 1.06\hat{\sigma}_{\hat{h}_{it}} n^{-(1/5)}$$

$$\downarrow$$

$$\hat{b}_{rule} = 1.06\,(0.814)\,(9697)^{-(1/5)} = 0.138$$