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Survival and employment growth of Belgian firms with collective layoffs

The impact of relocation, size, age, capital intensity and multinational group membership

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I Survival and employment growth of Belgian firms with collective layoffs, the impact of relocation, size, age, capital intensity and multinational group membership

*This paper studies the impact of collective layoffs with and without relocation on employment growth and survival of industrial firms in Belgium. The effect of size, age, capital intensity and firm membership of a multinational group on its employment growth and probability to survive are also considered. We perform exit and growth regressions for a panel data set of industrial firms with a collective layoff in the period 1990-1996. Econometric methods are applied to handle the problem of sample selection bias caused by using only surviving firms in growth regressions. The setting up and choice of these methods are done in collaboration with Dr. Shadman-Mehta Fatemeh (*UCL*).*

A. Introduction

This paper is part of the SSTC-program on "relocation, innovation and employment". We consider the impact of collective layoffs with and without relocation on employment growth and survival of industrial firms.

In Belgium, firms of at least 20 workers that downsize their workforce by at least 10% are obliged to report this to the regional employment offices (VDAB, ONEM, ORBEM). Information about the firms and workers concerned was reported to the Federal Planning Bureau from 1990 onwards. Data on relocation came from a survey organised with the 3 national labour unions[1]. A relocation is defined as a transfer of (a part of) the activities abroad that is organised by the Belgian firm or its parent.

The aim of the study is to estimate the average direct impact of collective layoffs on firms' employment and survival, and to verify whether this is different if the layoff is caused by relocation or not. We compare the effects of relocations with that of other determinants of firm growth and survival, like firm size, age, capital intensity and a variable indicating that it makes part of a multinational group. This is important, because relocations are more frequent among large and multinational firms and the effects of relocation on employment and firm survival should not be confused with those of size, age, capital intensity or multinational group membership.

1. It concerns a questionnaire sent by the Federal Planning Bureau to the 3 recognised national unions, for every collective layoff in the period 1990-1995. The survey had a high response rate. For a discussion of its results, see Federaal Planbureau (1997).

A major problem, when estimating the impact of any variable on the growth of firms, is that for computing a growth rate, only surviving firms can be used. As we are interested in the total employment effect of relocation, and firm exit[1] is a likely event in the case of a collective layoff, this may lead to a serious bias. For exemple, if small and large firms have the same average growth, but small firms have a higher likelihood to fail, then the growth of small firms is overestimated in a regression using only surviving firms.

Following Doms, Dunne and Roberts (1995) and Hall (1987), we used Heckman's two-step estimation procedure[2] to control for this sample selection problem. This method consists of first estimating, with a probit model, the probability that a firm survives, and then using the results of this regression to correct for the selection bias in the growth regression, performed only on the subset of surviving firms. In both steps we use the same explanatory variables. These are, besides the variables concerning collective layoffs and relocation, the firm's size (measured as its employment in the initial period), its age, and its capital intensity[3].

A reason for including capital intensity in the probit regression is that a higher capital intensity, measured as the stock of tangible assets[4] divided by the average number of workers, implies the existence of relatively high fixed costs. From standard economic theory, it follows that a plant will only shut down if its variable costs are no longer covered. Fixed costs are sunk, and don't matter. Thus we expect firms with a high capital to labour ratio to shut down less easily. In the presence of sunk costs to entry and exit it can also be shown that there exists an option value of remaining in the market even if the producer is incurring losses with respect to variable costs[5]. Another reason for putting the capital intensity of a firm both in the probit regression for survival, as in the growth regression, is that the capital-labour ratio indicates investments done in the recent past. More efficient firms generate higher levels of investment, and larger capital stocks[6]. Also, the simple fact that a firm has invested in the recent past means that it expects to continue or expand production.

In many empirical studies a firm's size and age are found to have a negative impact on its probability to fail[7]. Some of these studies also find a negative impact of size and age on the average growth rate (even after controlling for the sample selection problem mentioned before). This holds particularly when comparing between younger and smaller firms (Evans 1987a, Hall (1987)).

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1. A firm is said to exit if it stops its activities in Belgium, irrespective of having relocated activities.
2. As it was described in Greene (1997), p 977-978.
3. Data on firms employment, capital intensity, age and survival were obtained from annual account data, as gathered by the NBB. (de balanscentrale-centrale des bilans).
4. These include land and buildings, plant, machinery and equipment, furniture and vehicles, leasing and other similar rights, other tangible assets & advanced payments for assets under construction.
5. We used the same argument here as Doms, Dunne and Roberts (1995), who refer to Dixit (1989) for the proof of the option value argument, and found a positive effect of capital intensity both on the probability of surviving, and on growth.
6. This follows both from the theories of passive and active learning of Jovanovic (1982) and Ericson and Pakes (1989).
7. Those studies include Evans (1987a and 1987b), Doms, Dunne and Roberts (1995),Mata, José, Portugal Pedro, Guinaraes Paulo (1995), Sleuwaegen, J. Konings and Mommaerts (1999).

The theoretical justification for entering size and age in the exit and growth regressions is, amongst others, given by the theory of passive learning advanced by Jovanovic (1982) and Ericson and Pakes (1989). Jovanovic's theory of selection (or passive learning) implies that young firms have both a higher growth rate and a higher variability in growth rates, leading to a higher exit probability. This effect is induced by the fact that young firms have to learn about their efficiency as they operate in the industry[1]. The efficient grow and survive; the inefficient decline and fail. As young firms are often small, Jovanovic (1982) predicts small firms to have higher growth and exit rates. Holding age constant, however, his theory does not necessarily imply that small firms have higher growth rates (see also Evans 1987b).

A more direct negative impact of size on growth rates (not on survival!) is provided by the theory of Ghemawat and Nalebuff. They predict that in declining industries the largest firms will downsize first. They do this because they recognize that given the anticipated decline in demand, smaller firms will be able to produce profitably for a longer time (Ghemawat and Nalebuff, 1995). A second reason for the largest firms to cut production (and employment) first, is that they recognize that their production has the largest effect on price levels (Ghemawat and Nalebuff, 1990). Thus the large firm acts as a kind of "Stackelberg leader" that absorbs the general shocks in demand.

Note the difference between the predicted higher likelihood of downsizing for large firms (in declining industries), and that of exit (=the end of all activities). The theory only has implications for downsizing (the likelihood of having a collective layoff), not for that of stopping all activities. Once large firms are as small as the others, there is no higher likelihood for further downsizing. If a higher probability of downsizing for large firms is the basis for the empirically observed negative relation between size and growth in a representative set of firms, then such a negative result no longer has to be expected if one estimates on a subset of firms that have all had a collective layoff in the period for which the growth is computed. For a set of firms that includes both firms with collective layoffs, and firms without collective layoffs, size could have a negative effect on growth again.

We are particularly interested in studying the effect of a firm belonging to a multinational group. Such a firm has a higher probability of collective layoff and relocation, but we show that, when compared with other firms with collective layoffs, it also has a significantly higher growth rate. This is in line with other results indicating that multinational industrial firms in Belgium are more innovative and have a higher capital intensity (Van den Cruyce 1998).

1. Mature firms are aware of their efficiency, and therefore face both a lower probability of failure and a lower average growth rate. The higher average growth for young firms, that should hold even after controlling for selection bias due to the estimation on surviving firms (Jovanovic 1982), could be thought of as a risk premium for their higher variability in growth rates. We derive this interpretation from the fact that Jovanovic (1982) works with a cost function that is convex in production q. With such a cost function, a higher variability of q (with equal means) leads to lower profits. Since expected profits must be positive, this is compensated by a higher average growth rate.

In part B we discuss the methods used for estimating growth and survival, and the econometric problems that arise in that context. In part C we compare the results of different estimators, working only with the variables age, size and capital intensity. In part D we include the variables on collective layoffs and relocation and a dummy for multinational firms. The data used are only those on firms with a collective layoff in the period 1990-1996.

B. Methods used to estimate growth and probability of survival

We first introduce the specification used in the probit and growth regression. In section 2 we discuss briefly some econometric problems that arise when estimating growth with this specification. In section 3 we present the estimation results for the different methods used.

1. Specification of probit and growth regression and correction for sample selection bias.

The specification underlying the estimation for the survival of a firm is:

$$z''_{it} = \alpha_t + \beta y_{it-1} + \gamma' x_{it-1} + e_{it} \tag{1}$$

Here y_{it-1} is the ln of the average employment in firm i in year t-1, x_{it-1} is a vector of the other regressors[1] evaluated at year t-1, and e_{it} is a disturbance term. z''_{it} is a (latent) variable that increases with the probability for firm i to survive in year t. In reality the variable z''_{it} is not observed. We do observe whether a firm has survived or not in year t. The variable z_{it}=1 if the firm has survived, and 0 if it has not. Assuming that e_{ij} is distributed normally, and normalising such that if $z''_{it} > 0$, z_{it}=1, and if $z''_{it} < 0$, z_{it}=0, equation (1) can be estimated with the probit model.

Note that, in contrast to the other parameters, the parameter α has an index t. This is because we do not want to impose that the probability of exit is the same each year. Besides allowing for differences in the business cycle this reduces the problem caused by measurement errors in the exact timing of the exit of a firm.

A firm was declared to fail in year t if it stopped reporting positive employment levels in that year, and had either a special legal status[2], or was no longer present in the central bank's annual account database in the years following the information stop. With this procedure we were able to measure rather well whether a firm eventually disappeared, but not always exactly in which year[3].

The specification for the growth equation is given by:

$$\Delta y_{it} = \zeta_t + \eta y_{it-1} + \theta' x_{it-1} + \iota\lambda(\ldots) + u_{it} \tag{2}$$

1. as described in tabel 13 in the appendix.
2. This includes various statuses, like bankruptcy or liquidation. Mergers, absorptions and scissions were treated differently. If firms with such a legal status stopped reporting information, they were excluded from the sample.
3. This was indicated by data on export and import flows by the same firms that often continued for one year after the firm was declared to have failed based on annual account information.

where the dependent variable is the change in (the ln of) employment with respect to the year t-1 and ζ_t is, like α_t, a year-specific constant term. If one estimates equation (2) without including $\lambda(...)$, the results are biased by the sample selection due to excluding firms that failed during the observation period (since no growth rate can be computed for them).

Following Hall (1987) and Doms, Dunne and Roberts (1995) we resolve this problem by using Heckman's two-step method. This implies that the growth equation also includes the regressor $\lambda(...)=\lambda(\alpha*_t+\beta*y_{it-1}+\gamma*x_{it-1})$, where $\alpha*$, $\beta*$ and $\gamma*$ are the coefficients in (1) as estimated by the probit model and $\lambda(...)$ equals $\phi(...)/\Phi(...)$ (see Greene, 1997). $\phi(.)$ and $\Phi(.)$ are respectively the density and the cumulative distribution of a standard normal distribution[1]. The coefficient ι of $\lambda(...)$ is the covariance between the disturbance terms of equations (1) and (2).

Note that, though the explanatory variables are the same, the coefficients in (2) are not the same as those in (1). The probability of surviving can be considered as a lottery that takes place before a growth figure can be calculated. The growth regression has to be corrected for this, but this does not imply that a variable that has a positive (negative) effect on growth, should always have a positive (negative) effect on survival. From theory, we expect size and age to have a negative effect on growth, but a positive one on the probability to survive.

Since the coefficient of $\lambda(...)$ is the covariance between the disturbance term in (1) and (2) it can be expected to have a positive sign. That is: omitted variables that lead to high growth rates are also likely to increase the probability to survive. Still, a negative covariance cannot be excluded if a high variability in growth is compensated by a higher average growth rate as a kind of risk premium.

2. Some econometric issues

Having solved for the problem of sample selection, some problems, typical of firm-level or panel data, still have to be treated. These are the problem of heteroscedasticity, that of firm specific disturbances, and that of correlation of the disturbance term with the regressor y_{it-1}.

When performing the growth regression in (2) with OLS, White's test for heteroscedasticity rejected the OLS-assumption that the disturbance term u_{it} has the same variance for each observation. This is not surprising, given that the theory of passive learning predicts that the variability of growth in young firms is larger than that in old firms[2]. Heteroscedasticity does not make the coefficients obtained by the OLS-estimator inconsistent, but leads to inconsistent estimates for the standard errors. Like Doms, Dunne and Roberts (1995) and Evans (1987a) we control for this problem by using White 's procedure for computing heteroscedasticity consistent standard errors.

1. If large firms are more likely to survive ($\beta*$ is positive), then $\lambda(\alpha*_t+\beta*y_{it-1}+\gamma*x_{it-1})$ is higher for small firms.
2. Evans (1987a) found that the variability of growth was negatively influenced by firm age and plant size (but not firm size).

Note that measured heteroscedasticity can also be due to misspecification. It has been advanced that the true specification is not linear but of a higher order (see Evans 1987b). Therefore, we also perform estimations with a specification that includes squared terms for age, size and a crossed term for both.

The specification in (1) and (2) is one for panel data, where yearly observations are available for each firm. (2) can be estimated using OLS, with $\lambda(.)$ included as one of the x-regressors to control for sample selection bias. This is not necessarily the most efficient approach, though, since it does not take account of the likely event that growth rates over different years of the same firm are correlated.

Taking account of this, the error term of each firm could be written as $u_{it}=v_i+\varphi_{it}$. φ_{it} is the standard white noise error term, and v_i captures firm specific differences in the growth rate that remain stable over the observation period. This is the approach followed in the "random effects" model. We are able to compute estimates with this structure of the error term, by applying the Fuller-Battese method in SAS. A more radical alternative is to assume that the constant is firm specific. By extending (2) with an additional constant term ε_i, one eliminates all stable firm specific growth effects in the observation period not related to y_{it-1} or the variables in the x_{it-1} vector. This is the approach followed in the "fixed effects"model.

Unfortunately, with the dynamic specification in (2), all three the OLS, random effects and fixed effects estimator are biased. Note that the ln of employment in period t-1 appears both on the right side as on the left side of the equation. This is so because $\Delta y_{it-1}=y_{it}-y_{it-1}$. Thus, equation (2) can be rewritten in the form: $y_{it}=\varepsilon+\zeta_t+(\eta+1)\, y_{it-1}+\theta' x_{it-1} + u_{it}$. This specification entails a problem of correlation between the lagged dependent variable and the disturbance term, causing a negative bias on the coefficient of y_{it-1}. In the random effects model, where u_{it} has a part v_i in the disturbance term, this is because v_i is correlated with y_{it-1}(see also Greene 1997, p 640, and Hsiao, 1988). In the fixed effects model this problem is even worse, because there coefficients are based exclusively on variation in the time dimension within firms, where this problem is caused[1]. To solve this problem, Greene (1997), and Hsiao(1988), propose to estimate:

$$\Delta y_{it} = (\zeta_t - \zeta_{t-1}) + (\eta + 1)\Delta y_{it-1} + \theta'(x_{it-1} - x_{it-2}) + u_{it} - u_{it-1} \qquad (3)$$

To avoid the problem of correlation of the lagged dependent variable with the disturbance term, the change in y_{it-1} is instrumented by either y_{it-2} or $y_{it-2}-y_{it-3}$, using the Generalised Methods of Moments estimation technique. This solution is only valid if the true specification is linear as in (2)[2]. It also uses only a small amount of the variation in the data. We tried it, but without much success[3].

1. Using the fact that estimating the fixed effects model is the same as estimating the model in deviations from the mean, it can be shown that the bias is negative and decreasing in the number of time periods (see Hsiao (1988). If the true η is 0 and T=5 then the asymptotic bias equals -0.188, which is large (see further).
2. In fact, the linear specification in (2) is clearly rejected by the data in favour of the quadratic expression (especially for the crossed term of size and age), see further.
3. Results for these estimates are given in column [3] in table 6. When instrumenting on y_{it-2}, leads to values of $(1-\eta)$ that are highly unstable.

A more straightforward solution is to perform a simple growth equation, where the average yearly growth in a period of n years is regressed on the y and x values of the initial year. Doing this, the within firm differences are not allowed to influence results. This method has the disadvantage that the effect of the other explanatory variables, like the capital intensity (reflecting recent investments) and the occurrence of a collective layoff cannot be estimated in the panel-data dimension. Therefore, we will perform estimations where the size of the firms is kept constant, while the x variables vary over time. To exclude any correlation with the disturbance term we measure size in the year 1990 which precedes the period 1992-1996 for which yearly growth estimations are performed.

C. Comparing the results of the regressions of survival and growth for different estimators

1. Probability to survive

In table 1 and 2, we present the results of the probit estimates for the probability that a firm survives. A description of the explanatory variables and their average values can be found in the appendix. On a total of 369 industrial firms with a collective layoff in the period 1990-1996, 64 have stopped their activities before the end of 1996. Averages of the variables used in the regressions can be found in the appendix.

We estimated both the instantaneous (yearly) probability to survive, and the probability to survive in the period 1992-1996. In the yearly survival estimates a distinction is made between the case where size (lagged) is kept constant at the level of 1990, and the one where it varies over time. In the estimates for the period 1992-1996 size was measured either in 1991 or in 1992. Table 1 presents the results for the linear expression in (1). Table 2 those for the specification with squares and a crossed term for age and size.

From the linear model we conclude that size (as measured by employment) has a significant positive effect on the probability to survive. One should be careful when interpreting this result, though. Consider a firm of 500 workers that downsizes to a size of 10. This firm would still be reported as having survived. If it completely disappears the year after, the specification used in column 1 picks up that “a small firm has failed”. However, the negative effect of size on the probability to fail (or exit) cannot be entirely attributed to such a delayed exit for large firms. The specifications used under [2] and [4] should suffer less from this problem. Although the measured effect of size is smaller, [2] and [4] still lead to a significant positive effect of size on survival.

TABLE 1 - Probit estimation of survival in period 1992-1996, using 369 industrial firms with collective layoffs[a], linear specification

	yearly survival, size of t-1 [1]	yearly survival, size of 1990 [2]	survival in 1992-1996, size of 1991 [3]	survival in 1992-1996, size of 1990 [4]
intercept	0.778 (.481)	0.999 (0.474)	-0.973 (0.615)	-0.886**(0.61)
y_{it-1} or y_{i0}	0.145** (.042)	0.102**(0.046)	0.152**(0.066)	0.119**(0.064)
ln(age)	0.05 (.118)	0.04 (0.119)	0.048 (0.132)	0.049 (0.133)
ln(capital intensity)	0.108**(.046)	0.108**(0.047)	0.157**(0.076)	0.168**(0.075)
dummy 1993	-0.419* (.239)	-0.426*(0.236)		
dummy 1994	-0.619** (0.23)	-0.635**(0.228)		
dummy 1995	-0.804** (0.23)	-0.846 **0.224)		
dummy 1996	-0.485** (0.24)	0.543**(0.24)		
number of firms	369	369	369	369
observations without exit	1678	1678	305	305
observations with exit	64	64	64	64
log likelihood for normal	-255.68	-259.31	-163.28	-164.30

a. Standard errors are given between brackets. A significant difference from zero at a confidence level of 90% and 95% is indicated by * and ** respectively.

TABLE 2 - Probit estimation of survival in period 1992-1996, using 369 industrial firms with collective layoffs[a], specification with second moments for size and age

	panel: yearly survival, size of t-1 [1]	panel: yearly survival, size of 1990 [2]	survival in 1992-1996, size of 1991 [3]	survival in 1992-1996, size of 1990 [4]
intercept	-0.863 (1.831)	-0.231 (1.877)	-0.255 (2.064)	-0.185 (2.043)
y_{it-1} or y_{i0}	-0.162 (0.26)	-0.523* (0.289)	-0.717* (0.352)	-0.733**(0.354)
ln(age)	2.027 (1.3)	2.262* (1.313)	1.39 (1.51)	1.362 (1.513)
$(y_{t-1})^2$ or $(y_{i0})^2$	0.006 (0.018)	0.037 (0.028)	0.043 (0.036)	0.043 (0.034)
$(\ln(age))^2$	-0.457*(0.239)	-0.512**(0.238)	-0.445 (0.292)	-0.439 (0.292)
$(y_{it-1})\ln(age)$ or $(y_{i0})\ln(age)$	0.09(0.074)	0.099 (0.093)	0.17 (0.105)	0.17 (0.107)
ln(capital intensity)	0.109**(0.048)	0.108**(0.048)	0.153**(0.078)	0.160**(0.078)
dummy 1993	-0.471* (0.248)	-0.474*(0.247)		
dummy 1994	-0.679** (0.242)	-0.701**(0.240)		
dummy 1995	-0.864** (0.238)	-0.923**(0.236)		
dummy 1996	-0.528** (0.256)	0.600**(0.252)		
number of firms	369	369	369	369
observations without exit	1678	1678	305	305
observations with exit	64	64	64	64
log likelihood for normal	-253.17	-254.43	-158.98	-159.56

a. Standard errors are given between brackets. A significant difference from zero at a confidence level of 90% and 95% is indicated by * and ** respectively.

In columns [2] to [4] of the non-linear specifications in table 2 there is a significant negative linear effect of size on the probability to survive. However, the crossed term for size and age and the squared term for size are positive. Using the coefficients in column [4] it was computed that for firms with more than 20 workers and the average age of 16 years the impact of an increase in size is always positive.

In table 1 the effect of age on survival is positive, as predicted by Jovanovic's theory, but not significantly different from 0. The results in table 2 show why this is so. There tends to be a positive linear effect of age on survival, and a negative effect of the squared term. Both effects are significant in specification [2], where yearly exit is regressed on 1990 employment levels[1]. For a firm with 100 workers in 1990, using the coefficients of column [2] an increase in age leads to an increase in the probability to survive until the age of 15 years. Further increases lead to a (very) small drop in the probability to survive.

We conclude that, at first, ageing positively affects the probability to survive, but as firms grow older, this effect is weakened. Size has a positive effect on survival. In all the regressions we also find a significant positive effect of capital intensity (measured as capital stock divided by y_{it-1}) on survival. This is in line both with the idea that higher investments (in a recent past) are correlated with a higher likelihood of firm survival[2], and that a firm with more sunk costs exits less easily.

2. Growth estimates

We estimate (2) with OLS, with the random, and the fixed effects model[3]. We use a specification without (table 3 and 4) and one with (table 5 and 6) correction for sample selection bias. Table 3 and 4 show the results of a specification where size (employment lagged with one year) varies, and one where it is kept constant at the 1990 level. Instead of presenting the results for the fixed effects model in the case size is kept constant[4], we included the results of a simple growth regression for the entire period 1992-1996.

1. Because in specification [1] and [2] firms are allowed to get older, and we controlled for business cycle effects by including the yearly dummies, these specifications are more efficient for measuring the effect of ageing than [3] and [4].
2. The causal link here goes in both directions. One will only invest in activities, plants or firms that have a high likelihood of surviving. And secondly, such investments increase the likelihood that the firm survives at least in the near future.
3. See section B.2 for a description of the random and fixed effects model.
4. The results for this model are shown in table 7 in the case of a linear specification

a. Results without correction for sample selection bias

From our discussion in section 2 we know that when estimating the dynamic relation in (2) directly, using panel data, the coefficient for y_{it-1} is biased negatively due to correlation between y_{it-1} and u_{it}. We included these biased estimation results in the first 3 columns of table 3 and 4, besides the (unbiased) results where size is held constant in [4], [5] and [6], to show the importance of this bias. The random and the fixed effects model in [2] and [3] in table 3 show a coefficient of y_{it-1} that is significantly negative, but the estimated coefficients (of -0.03 and -0.573) are very different. This already signals a problem, because if the model is correct, the random and the fixed effects estimator should yield similar results (see Verbeek, 1996)[1].

The results of the random effects model in column [2] are significantly different from those of the same model in column [4], where y_{it-1} is replaced by y_0. (0=the 1990 level). Note that there is not only a difference in the coefficient of size, but also in the other coefficients, which is why this issue is so important. An incorrect estimation of the size effect, affects the estimated impact of the other variables. The OLS-estimator is more robust to size varying over time. There is no significant difference between the coefficients in [1] and [4]. This holds for table 3 and 4. This is probably because the OLS-estimator uses less of the variation in time than the other models. We also find that there is hardly any difference between the coefficients of the OLS and the random effects estimator in [4] and [5], which means that holding size constant at the 1990 level yields more reliable and stable results.

Now consider the specification with a quadratic and crossed term for size and age. In the linear specification the results for a White test of heteroscedasticity, performed in [1], [4] and [6], indicate that the hypothesis of homoscedasticity must be rejected. In the estimations with the quadratic terms, reported in table 4, heteroscedasticity is much less a problem. The hypothesis that the variance of the disturbance is the same for each observation cannot be rejected.

In the robust specifications in [4] to [6] of table 3, we found that size had a positive, but insignificant effect on growth. If one looks at the results in table 4 for the same columns, one sees that the linear term for size has a significantly negative effect on growth, while the crossed term for age and size (in line 6) has a significant positive effect. The squared term for size has no significant effect. For age, we find a significant negative effect on growth in the case of the linear specification in table 3. In the case of the simple growth model in [6], this is decomposed in a significantly negative linear term, and a significantly positive quadratic (and crossed) term in table 4.

Thus we find evidence for a negative effect of age on average growth, and an initial negative effect of size on growth. Due to the positive crossed term for size and age, the latter effect only holds for small and young firms. All this is very much in line with the model of passive learning and the results of the authors mentioned earlier (see Evans 1987ab, Hall 1987). We have to be careful, though, since this result is based on estimates without correction for sample selection bias.

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1. Also note that the R^2 of the fixed effects estimator, where the negative effects are largest, is about 0.2 higher than that of the other estimators. This is not a good sign at all, since it is only the use of the mean over 5 years in the computations (with 1/5=0.2) that leads to the higher R^2.

TABLE 3 - Yearly growth in period 1992-1996, using a panel of 305 surviving industrial firms with collective lay-offs[a], no correction for sample selection bias, linear specification

	panel:OLS size of t-1 [1]	panel: random effects[b] size of t-1 [2]	panel: fixed effects[c], size of t-1 [3]	panel: OLS, size of 1990 [4]	panel: random effects[b] size of 1990 [5]	growth 1992-1996, OLS, size of 1990[d] [6]
intercept	-0.166 (0.115)	-0.266*(0.142)	0.20	-0.170 (0.117)	-0.199 (0.098)	-0.106 (0.126)
y_{it-1} or y_{i0}	0.0037 (0.012)	-0.03**(0.010)	-0.573**(.105)	0.007 (0.008)	0.007 (0.008)	0.009 (0.009)
ln(age)	-0.068**(0.027)	-0.045 (0.032)	0.875 (0.251)	-0.071**(.027)	-0.072**(.023)	-0.069**(0.02)
ln(capital intensity)	0.039**(0.013)	0.065**(0.013)	0.046 (0.069)	0.038**(0.012)	0.038**(0.01)	0.020 (0.014)
dummy 1993	-0.036 (0.029)		-0.148**(.026)	-0.036 (0.029)		
dummy 1994	-0.044 (0.029)		-0.286**(.047)	-0.045 (0.029)		
dummy 1995	0.031 (0.028)		-0.341**(.062)	0.031 (0.029)		
dummy 1996	-0.096**(0.038)		-0.553**(.086)	-0.097**(.038)		
number of observations	1525	1525	1525	1525	1525	305
Adjusted R^2	0.025	0.023	0.242	0.026	0.019	0.058
White Test for heteroscedasticity	X^2=41.8 (0.019) YES			X^2=39.6 (0.032) YES		X^2=19.0 (0.025) YES

a. Standard errors are between brackets. A significant difference from zero at a confidence level of 90% and 95% is indicated by * and ** respectively. Standard errors of OLS are computed from a heteroscedasticity consistent covenantees matrix.

b. The random effects estimator is performed using the Fuller-Batese method in the SAS, TSCREG procedure. This method puts both the stable differences in growth rates between firms as those between years in the disturbance term.

c. The fixed effects coefficients were computed by performing an OLS regression on the variables in deviation from their means. This leads to exactly the same coefficients as the estimation with the ε_i's treated as dummies. (see Verbeek (1996) and Hsiao (1988)).

d. The coefficients of the growth regression are directly comparable to those of the panel regressions because the dependent variable used here was (log (employment 96)-log (employment 1991))/5.

TABLE 4 - Yearly growth in period 1992-1996, using a panel of 305 surviving industrial firms with collective lay-offs[a], no correction for sample selection bias, extended specification

	panel: OLS, size of t-1 [1]	panel: random effects[b], size of t-1 [2]	panel: fixed effects[c], size of t-1 [3]	panel: OLS size of 1990 [4]	panel: random effects size of 1990 [5]	simple growth: OLS, size of 1990[d] [6]
intercept	0.753 (.641	0.816*(0.481)	3.187()	0.522 (0.535)	0.492 (0.387)	0.821**(0.301)
y_{it-1} or y_{i0}	-0.319**(.088)	-0.557**(.066)	-1.509**(.134)	-0.208**(.064)	-0.207**(.046)	-0.161**(.052)
ln(age)	-0.133 (.393)	0.189**(0.323)	2.634**(1.093)	-0.155 (0.368)	-0.154 (0.273)	-0.484**(.208)
$(y_{t-1})^2$ or $(y_{i0})^2$	0.015**(0.007)	0.034**(0.004)	0.072**(0.011)	0.002 (0.004)	0.002 (0.004)	0.003 (0.004)
$(\ln(age))^2$	-0.041 (0.066)	-0.106*(0.059)	-0.944** (.373)	-0.048 (0.068)	-0.049(0.051)	0.032 (0.041)
$(y_{it-1})\ln(age)$ or $(y_{i0})\ln(age)$	0.058**(0.024)	0.064**(0.019)	0.169**(0.035)	0.067**(0.018)	0.067**(0.016)	0.052**(0.013)
ln(capital intensity)	0.038 (0.011)	0.056**(0.013)	0.05*(0.026)	0.037**(0.012)	0.037**(0.010)	0.023*(0.013)
dummy 1993	-0.039 (0.028)		0.012 (0.061)	-0.037 (0.028)		
dummy 1994	-0.049 (0.028)		0.051 (0.115)	-0.047 (0.027)		
dummy 1995	0.024 (0.026)		0.166 (0.171)	0.028 (0.028)		
dummy 1996	-0.102**(.038)		0.121 (0.227)	-0.099**(.038)		
number of observations	1525	1525	1525	1525	1525	305
Adjusted R^2	0.051	0.08	0.269	0.039	0.035	0.128
White Test for heteroscedasticity	X^2=51.8 (0.443) NO			X^2=58.9 (0.21) NO		X^2=29.0 (0.18) NO

a. Standard errors are between brackets. A significant difference from zero at a confidence level of 90% and 95% is indicated by * and ** respectively. Standard errors of OLS are computed from a heteroscedasticity consistent covariance matrix.

b. The random effects estimator is performed using the Fuller-Batese method in the SAS, TSCREG procedure. This method puts both the stable differences in growth rates between firms as those between years in the disturbance term.

c. The fixed effects coefficients were computed by performing an OLS regression on the variables in deviation from their means. This leads to exactly the same coefficients as the estimation with the ε_i's treated as dummies. (see Verbeek (1996) and Hsiao (1988)).

d. The coefficients of the growth regression are directly comparable to those of the panel regressions because the dependent variable used here was (log (employment 96)-log (employment 1991))/5.

b. Results with correction for sample selection bias

Table 5 and 6 report the estimation results with correction for the sample selection problem. The correction is performed by including the term $\lambda(.)$, as defined in part B, in the regressions. Table 5 reports the results where $\lambda(.)$ is computed with the results of the linear probit model in table 1, table 6 those where $\lambda(.)$ is computed with the results of the non-linear probit model in table 2.

In table 5 the coefficient of $\lambda(.)$ tends to be positive, but it is only significant in the linear simple growth regression in column [3]. Column [3] of table 5 can be compared with column [6] of table 3 to see the effect of the correction for sample selection bias. None of the effects of size, age or capital intensity has reversed in sign, but the sizes of the coefficients, and their significance have changed. Size now has a significant positive effect, while the negative effect of age is no longer significant. It is intuitively logic that if one corrects for the bias due to estimating growth on surviving firms only, the effect of size on growth is more positive[1]. In table 1 we found size to have a significant positive impact on survival.

1. This intuition is confirmed by the positive sign of the coefficient of $\lambda(.)$ in column [3], implying that the covariance between the disturbance term in regression (1) and (2) is positive.

It should not be surprising that the coefficient of $\lambda(.)$ is insignificant, and thus that sample selection bias is less important, in the regressions [1],[2],[4] and [5] in table 5 where growth and survival are followed annually. With 64 exits on a total of 369, the probability for a firm to stop its activities in the period 1992-1996 is 17.3%. With the same 64 exits on a total of 1678 observations, the yearly probability of exit is only 0.038.

The insignificance of the correction for sample selection bias in column [6] is more surprising. This column gives the non-linear model with correction for sample selection bias. The $\lambda(.)$ used in this regression is the same as the one used in column [3]. Thus, with a non-linear specification for size and age in regression (2), the correction for sample selection no longer seems to be very important[1]. This implies that the linear specification for growth has to be rejected and the correction for sample selection is less important because the covariance of the disturbance terms in the regression for growth and survival is close to zero[2].

Indeed, as was already the case in table 4, the coefficient of the crossed term for age and size continues to be significantly positive. In that case the results in column [3] of table 5 must be interpreted with care. It may be that the term $\lambda(.)$ in this specification does not measure the impact of sample selection bias, but that, because of its non-linearity in size and age, it takes over the effect of the omitted crossed term for age and size. This reveals a more general identification problem with respect to the effects of sample selection that was already mentioned by Doms, Dunne and Roberts (1995).

If the specification in columns [4] to [6] is preferred over the linear in columns [1] to [3], then the conclusions are quite different. Size has a significantly negative linear effect on growth. Even if the effect of the squared and crossed terms is taken into account it can be computed that the effect of an increase of size on growth is negative, but decreasing for the sizes of the sampled firms[3]. In column [6], age has a significant negative linear effect on growth, even after controlling for sample selection bias. Thus the quadratic specification, as proposed by Evans (1987b), is in accordance with the predictions of the model of passive learning.

Table 6 reports the growth estimates where $\lambda(.)$ is computed with the results of the non-linear probit model in table 2. There is a remarkable difference between the results in table 5 and 6. In contrast to table 5, the coefficient of $\lambda(.)$ now tends to be significantly negative, implying that the disturbance terms in (1) and (2) are negatively correlated. This result is in conflict with the intuition that (omitted) variables that influence growth positively also influence survival positively.

1. As can be verified by comparing the results of the 6th column in table 5 and 4.
2. Implying that in our dataset, the exit of firms is a random process that is uncorrelated with growth.
3. An increase in size from 20 to 30 workers leads to a growth reduction of 1.4%, an increase of 300 to 310 workers to a reduction of 0.2%.

In the linear models, there is now a large and significant effect of $\lambda(.)$ despite the small share of yearly exits. As was the case in table 5, though, the results for the linear specification are suspect, because $\lambda(.)$ may have taken up the effect of the missing crossed term for age and size. In the non-linear regressions for yearly growth (columns [4] and [5]), the selection for sample selection was, as could be expected, of little importance. The results are comparable to those in table 5 and 4. In the non-linear growth model (in column [6]) there is now an important and significant negative effect of $\lambda(.)$ on growth. This is in conflict with the result in table 5, so that a choice has to be made between using the linear or the non-linear probit model for performing the correction for sample selection bias. Further on, we continue with the results of the linear probit model for the following reasons:

1) In the probit estimates of table 2 none of the non-linear terms in size or age has a significant effect in the specifications [3] and [4] where survival is measured for the whole period.

2) the significantly negative coefficient of $\lambda(.)$ implies a negative covariance between the disturbance terms of (1) and (2), for which, with size and age already in the regressions, we have no explanation.

Still, the results for both approaches are less different than it may appear from a comparison of column [6] in tables 5 and 6. Age has in both tables a significant negative linear effect on growth that is weakened either by the squared term for age, or the crossed term with size. Taking the complete quadratic expression, an increase in size also has a negative effect on growth in table 6.

We conclude that even after controlling for the sample selection bias, older and larger firms with collective layoffs tend to have lower growth rates, and that the effects of increasing age and size on growth weaken as firms get bigger and older.

With respect to the effects of capital intensity on growth, the different models presented in table 5 yield different results. In the OLS and random effects estimates for panel data, where capital intensity in year t-1 is allowed to vary over time, this variable has a significantly positive effect on growth. In the simple growth regression, capital intensity has a significant positive effect in the linear model in [3], but an insignificant one in the quadratic model in [6][1].

1. Capital intensity seems to take over a part of the effect of the squared term in size if the latter is not included in the regression.

TABLE 5 - Yearly growth in period 1992-1996, using a panel of 305 surviving industrial firms with collective lay-offs[a], with correction for sample selection bias based on the linear probit results *(table 1)*

	panel: OLS, size of 1990 [1]	panel: random effects, size of 1990 [2]	growth 1992-1996, OLS, size of 1990 [3]	panel: OLS, size of 1990 [4]	panel: random effects size of 1990 [5]	growth 1992-1996, OLS size of 1990 [6]
intercept	-0.302 (0.22)	-0.312 (0.195)	-1.539**(.644)	0.601 (0.54)	0.457 (0.397)	0.98**(1.016)
y_{i0}	0.015 (0.015)	0.012 (0.012)	0.069**(0.029)	-0.218**(.07)	-0.202**(.047)	-0.173*(0.1)
ln(age)	-0.066**(.027)	-0.069**(.023)	-0.034 (0.020)	-0.148 (0.373)	-0.158 (0.273)	-0.495**(.227)
$(y_{t-1})^2$ or $(y_{i0})^2$				0.003 (0.004)	0.002 (0.004)	-0.003 (0.005)
$(\ln(age))^2$				-0.05 (0.070)	-0.049 (0.051)	0.033 (0.04)
$(y_{it-1})\ln(age)$ or $(y_{i0})\ln(age)$				0.067**(0.018)	0.068**(0.017)	0.052**(0.014)
ln(capital intensity)	0.048**(0.018)	0.045**(0.014)	0.112**(0.042)	0.032*(0.018)	0.039**(0.011)	0.014 (0.045)
dummy 1993	-0.057 (0.046)			-0.025 (0.046)		
dummy 1994	-0.083 (0.068)			-0.024 (0.069)		
dummy 1995	-0.029 (0.097)			0.064 (0.1)		
dummy 1996	-0.127 (0.249)			-0.080 (0.063)		
$\lambda(.)$[b]	0.536 (0.917)	0.377 (0.566)	1.358**(0.653)	-0.325 (0.943)	0.136 (0.340)	-0.13 (0.764)
number of observations	1525	1525	305	1525	1525	305
Adjusted R^2	0.025	0.019	0.076	0.039	0.035	0.125
White Test for heteroscedasticity	X^2=44.3 (0.111) NO		22.18 (0.075) YES at 10%	X^2=62.7 (0.45) NO		X^2=33.7 (0.338) NO

a. Standard errors are between brackets. A significant difference from zero at a confidence level of 90% and 95% is indicated by * and ** respectively. Standard errors of OLS are computed from a heteroscedasticity consistent covariance matrix.

b. The adjustment for sample selection comes from the linear Probit model with size held constant at the 1990 level.

TABLE 6 - Yearly growth in period 1992-1996, using a panel of 305 surviving industrial firms with collective lay-offs[a], with correction for sample selection bias based on the nonlinear probit results *(table 2)*

	panel: OLS, size of 1990 [1]	panel: random effects, size of 1990 [2]	growth 1992-1996, OLS, size of 1990 [3]	panel: OLS, size of 1990 [4]	panel: random effects size of 1990 [5]	growth 1992-1996, OLS size of 1990 [6]
intercept	0.138 (0.149)	0.09 (0.138)	0.418**(.212)	0.712 (0.57)	0.484 (0.409)	1.60**(0.423)
y_{i0}	-0.014 (0.011)	-0.009 (0.01)	-0.016(.013)	-0.184**(.071)	-0.208**(.048)	-0.013 (0.059)
ln(age)	-0.072**(.027)	-0.071**(.023)	-0.074**(.020)	-0.288 (0.400)	-0.147 (0.285)	-0.934**(.274)
$(y_{t-1})^2$ or $(y_{i0})^2$				0.001 (0.004)	0.002 (0.004)	-0.003 (0.004)
$(\ln(age))^2$				-0.016 (0.077)	-0.051 (0.055)	0.167** (.063)
$(y_{it-1})\ln(age)$ or $(y_{i0})\ln(age)$				0.06**(0.019)	0.068**(0.017)	0.007 (0.015)
ln(capital intensity)	0.014 (0.013)	0.019*(.011)	-0.014 (0.015)	0.029**(0.013)	0.037**(0.011)	-0.025 (0.017)
dummy 1993	0.017(0.034)			-0.018 (0.031)		
dummy 1994	0.05 (0.043)			-0.013 (0.04)		
dummy 1995	0.183**(.059)			0.083 (0.058)		
dummy 1996	-0.022 (0.043)			-0.072 (0.041)		
$\lambda(.)^b$	-1.382**(0.44)	-1.055**(.354)	-0.53**(0.152)	-0.484 (0.494)	0.020 (0.325)	-0.787**(.268)
number of observations	1525	1525	305	1525	1525	305
Adjusted R²	0.033	0.019	0.113	0.039	0.035	0.145
White Test for heteroscedasticity	χ^2=46.9 (0.069) YES		26.2 (0.025) YES	χ^2=63.4 (0.50) NO		χ^2=39.7 (0.136) NO

a. Standard errors are between brackets. A significant difference from zero at a confidence level of 90% and 95% is indicated by * and ** respectively. Standard errors of OLS are computed from a heteroscedasticity consistent covariance matrix.

b. The adjustment for sample selection comes from the nonlinear probit model.

The explanation for the differences between the simple growth model and the panel data estimations is that the former is based exclusively on the cross-sectional variation in capital intensity. In the panel data estimations in [1], [2], [4] and [5] the cross-sectional variation is mixed with variation over time in capital intensity, due to changes in investment behaviour. To verify this claim one can look at the results of the fixed effects model, which only takes account of variations in the time dimension. In table 7, columns [5] and [6], we report results for the fixed effects model in a regression without size. We find that increasing capital intensity has a large and significantly positive effect on growth[1].

With the problems of the dynamic specification of (2) avoided (since y_{it-1} is left out of the regression), the fixed effects estimator consistently measures the effect of a change in capital intensity within a firm on growth in the next year. Thus, an increase in the capital intensity (as induced by high investments) within a firm leads to higher average growth in the year that follows. We do not find evidence that capital intensive firms have higher growth rates than others. The difference between both disappears if one uses a quadratic specification and corrects for sample selection bias in a simple growth model. Recall that capital intensive firms do have a significantly larger probability to survive, even when considering only the cross sectional variation (see columns [3] and [4] in table 1 and 2).

1. There is thus a discrepancy between the cross-sectional and time-effects of capital intensity. In that case, it is generally preferable to work with the results of the fixed effects model, since in the cross-sectional (or “between”) dimension the variable in question is more likely to be correlated with the disturbance term (due to omitted variables) see Hsiao (1988, and Verbeek(1996).

TABLE 7 - Yearly growth in period 1992-1996, results for some alternative methods[a]

	OLS, panel, size of t-1 [1]	OLS specification in first differences [2]	instrumental variables & first differences Hsiao, Greene [2]	fixed effects[b] with size fixed at 1990 level [5]	fixed effects with size fixed at 1990 level [5]
intercept	-0.166 (0.115)	-0.133 (0.022)	-0.191 (0.101)	_	_
h	0.0037 (0.012)	0.064 (0.114) -1	-0.9735 (1.7) -1	_	_
ln(age)	-0.068**(0.027)	0.634 (0.279)	1.208 (1.026)	0.339 (0.274)	0.372 (0.28)
ln (capital intensity)	0.039**(0.013)	0.129 (0.089)	-0.186 (0.541)	0.229**(0.090)	0.209**(0.093)
dummy 1993	-0.036 (0.029)	-0.028 (0.029)	-0.086 (0.101)	-0.069**(0.028)	-0.029 (0.047)
dummy 1994	-0.044 (0.029)	-0.035 (0.029)	-0.127 (0.163)	-0.111**(0.035)	-0.039 (0.073)
dummy 1995	0.031 (0.028)	0.038 (0.025)	-0.051 (0.153)	-0.073 (0.046)	0.040 (0.107)
dummy 1996	-0.096**(0.038)	-0.084**(0.037)	-0.126 (0.091)	-0.218**(0.063)	-0.167**(0.075)
λ(.)					-1.099 (1.019)
number of observations	1525	1525	1525	1525	1525
Adjusted R²	0.025	0.028	-0.63	0.059	0.059
White Test for heteroscedasticity	X²=41.8 (0.019) YES			X²=21.7 (0.117) NO	X²=41.8 (0.019) NO

a. Standard errors are between brackets. A significant difference from zero at a confidence level of 90% and 95% is indicated by * and ** respectively. Standard errors of OLS and instrumental variables are computed from a heteroscedasticity consistent covariance matrix.

b. The fixed effects coefficients were computed by performing an OLS regression on the model variables in deviation from their means. This leads to exactly the same coefficients as the estimation with the ε_i's treated as dummies. (see Verbeek (1996) and Hsiao (1988)).

D. The direct impact of relocation and collective layoffs on exit and employment growth

In this part we evaluate the direct impact of collective layoff with and without relocation. In section 1 we show that larger firms and foreign multinationals have a higher probability of collective layoff and relocation. This influences the total impact of collective layoff on employment, which is illustrated in section 2. In section 3 we discuss the results of exit and growth estimations that include a relocation and multinational firm dummy variable.

1. The importance of size

In table 8 we compare the size distribution of our sample of 369 industrial employers with collective layoffs (column [2]) with the size distribution of the total population of employers in 1990 (column [1]). Firms employing up to 20 workers are almost unrepresented in the database. This is because firms are only obliged to communicate a collective layoff in the case they have at least 20 workers and the layoff concerns more than 10% of the workers (Federaal Planbureau, 1994).

TABLE 8 - Size distribution in 1990 of firms with collective layoffs in period 1990-1996, compared with total population of employers *(RSZ)*

number of workers in 1990	share of employers in industry (RSZ data) [1]	share in group of industrial employers with collective layoffs in 1990-1996 [2]	share in group of industrial employers with exit [3]	share in group of industrial employers with relocation and collective layoff in 1990-1996 [4]	share in industrial multinational firms with collective layoffs in 1990-1996 [5]
less than 5 workers	0.512	0.011	0.016	0.015	0.004
5-9 workers	0.168	0.011	0.016	0	0.008
10-19 workers	0.119	0.011	0	0	0.008
20-49 workers	0.119	0.106	0.141	0.06	0.061
50-99 workers	0.038	0.179	0.234	0.09	0.126
100-199 workers	0.021	0.247	0.266	0.194	0.24
200-499 workers	0.015	0.244	0.25	0.239	0.29
500-999 workers	0.004	0.095	0.047	0.194	0.118
more than 1000 workers	0.003	0.098	0.031	0.209	0.142
total	25243	369	64	67	246

Of the 369 industrial firms, 67 had a collective layoff that was due to relocation to a foreign country of (a part of) the activities in the period 1990-1996. Relocation is more frequent in the 2 largest size classes, since firms with more than 499 workers make up about 40% of the firms with relocation (see column [4]), while their share in the 369 is only 20%. In contrast to relocation, firm exit is relatively more frequent within smaller size classes, as illustrated by comparing column [2] and [3].

We also made a distinction between uninational Belgian firms and multinational firms[1]. The latter are significantly larger than the uninational Belgian firms. They also had a significantly higher probability of collective layoff[2] as well as of relocation. Of the 246 multinational firms, 58, or 23.6% had a relocation in the period 1990-1996. Among the considered uninational Belgian firms with collective layoffs, only 9, or 7.3% relocated (part of) its activities[3]. For these reasons this is an interesting control variable in the regressions.

1. 78% of these were Belgian firms controlled by foreign groups. 22% were Belgian firms controlling a multinational group.
2. With a total of only about 697 industrial multinational firms in Belgium (Federaal Planbureau, 1997); these are strongly overrepresented within the group of firms with collective layoffs.
3. This difference in share is significant with Fishers Chi-squarevalue=14.6 (prob-value=0.001).

2. The average direct impact of collective layoffs in firms with and without relocation

Table 9 shows the direct impact of relocation on employment in this sample of 369 industrial firms. 13195 workers were fired in collective layoffs of firms with relocation, and 22692 in collective layoffs of firms without relocation. Given the small number of firms that relocated a part of their activities (according to the labour unions), the impact of relocation on employment is quite high. The high impact of relocation on employment is entirely due to the fact that firms with relocation are larger than those without. If one considers the share of the workers hurt by a collective layoff once it occurs, as illustrated in column [6], then there is no difference between firms with and without relocation.

TABLE 9 - The direct impact of collective layoffs[a] on average employment in the case with and without relocation

	Number of firms (only industrial) [1]	number of collective layoffs 1991-1995 [2]	number of workers dismissed [3]	average number of workers hurt [4]=[3]/[2]	average size in year preceding collective layoff [5]	average employment share affected by collective layoff [b] [6]
firms with relocation in 1990-1995	67	87	13195	151.7	892.6	36.1%
firms without relocation in 1990-1995	302	336	22692	67.5	374.8	37.5%
total	369	423	35887	84.8	468.8	0.18

a. The number of workers fired in a collective layoff comes from the regional employment services (VDAB, FOREM, ORBEM).

b. This share was computed using annual account data on average employment. For each firm, the number of persons laid off was compared to the average employment in the year preceding the layoff. If the accounting year and the calendar year were not the same, the end date of the accounting year with respect to that of the collective layoff determined which was the preceding year. Note also that a collective layoff in year t has a direct impact not only on the average employment of year t, but also on that in year t+1. This is because the average employment in year t is still influenced by the employment in the part of the year before the collective layoff. The figures given in column [6] give the total direct impact.

Table 10 illustrates that the average impact and the frequency of a collective layoff depend on the size of a firm. In smaller size classes, collective layoffs are less frequent but if they occur they affect a larger fraction of the workers. For example in the size class of 20 to 49 workers on average 64.5% of the firm’s employment level is hurt in the case without relocation, and 55% in the case with relocation. In the largest size class, a collective layoff only hurts on average 8.4% of the workers in the case without relocation, and 21.2% in the case with relocation.

Except for the smallest size class, the effects of collective layoffs tend to be larger in the case with relocation. This will be tested in the next section.

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TABLE 10 - The direct impact of collective layoffs[a] on average employment, comparison between size classes

	Firms without relocation in 1990-1995		Firms with relocation in 1990-1995	
firms by size class, with and without relocation	frequency of collective layoff in 1991-1995 [1]	average employment share affected by collective layoff[b] [2]	frequency of collective layoff in 1991-1995 [3]	average employment share affected by collective layoff [4]
20-49 workers	0.177	0.645	0.25	0.55
50-99 workers	0.207	0.407	0.2	0.433
100-199 workers	0.221	0.328	0.215	0.524
200-499 workers	0.222	0.354	0.225	0.434
500-999 workers	0.282	0.209	0.323	0.237
more than 1000 workers	0.282	0.084	0.314	0.212
all size classes	0.22	0.375	0.26	0.361

a. The number of workers fired in a collective layoff comes from the regional employment services (VDAB, FOREM, ORBEM).

b. see corresponding footnote in previous table.

3. The impact of relocation on exit and employment growth in firms with collective layoffs

In table 11, we present survival and growth regressions with a dummy variable that is 1 if the firm has relocated activities abroad in the period 1990-1995, and 0 if it has not. In the growth regression we also include a dummy variable that is 1 for firms belonging to a (Belgian or foreign) multinational group. The dummy for multinational firms was not included in the survival regressions, because we only had access to this information for the firms that survived until 1996.

The shaded area gives the results for the new variables. The effects of including these two dummies on the coefficients of the other regressors are small[1]. We find that, compared with other firms with collective layoffs, firms that relocated activities in the period 1990-1995 did not have a significantly different probability of survival (columns [1] to [3]). The dummy for relocation has a negative effect on growth, but the effect is only significant in column [6].

In contrast to the relocation dummy, the dummy that indicates that the firm belongs to a multinational group has a significantly *positive* effect on growth. Thus, while overrepresented in the group of firms with collective layoffs (see earlier), multinational firms are less hurt by collective layoffs in their employment growth than uninational Belgian firms. In earlier work (Van den Cruyce, 1998), using a mixed group of firms (with only 20% firms with collective layoffs) we found a positive effect of belonging to a multinational group on value added growth, but not on employment growth.

1. As can be checked by comparing with the results of the corresponding models in table 1, 2 and 5.

TABLE 11 - Differences between firms with or without relocation and multinational firms and other in survival and employment growth in the period 1992-1996.

	PROBIT (panel) yearly survival, size measured in 1990 [1]	PROBIT (panel) yearly survival, size measured in 1990 [2]	PROBIT survival in period 1992-1996 [3]	OLS (panel) yearly growth, size of 1990 [4]	OLS (panel) yearly growth, size of 1990 [5]	OLS yearly growth in 1992-1996, size of 1990 [6]
intercept	1(0.475)	-0.238 (1.88)	-0.885 (0.611)	-0.301(0.217)	0.545(0.55)	1.191**(1.091)
y_{i0}	0.103**(0.047)	-0.529*(0.290)	0.121*(0.066)	0.008(0.014)	-0.232**(0.07)	-0.205 (0.101)
ln(age)	0.04(0.119)	2.295*(1.315)	0.049 (0.133)	-0.058**(0.026)	-0.063(0.382)	-0.423*(0.234)
$(y_{i0})^2$		0.038 (0.028)			0.005 (0.004)	-0.005 (0.005)
$(\ln(age))^2$		-0.518**(0.239)			-0.06 (0.071)	0.022 (0.043)
$(y_{i0})\ln(age)$		0.099 (0.093)			0.063**(0.019)	0.05**(0.014)
ln(capital intensity in t-1)	0.108**(0.048)	0.104**(0.049)	0.167**(0.076)	0.043**(0.017)	0.025 (0.018)	-0.007(0.044)
relocation in 1990-1995	-0.008(0.16)	-0.07 (0.165)	-0.018 (0.213)	-0.039(0.027)	-0.046 (0.026)	-0.052*(0.027)
(part of) multinational firm in 1996				0.069**(0.032)	0.068**(0.033)	0.066**(0.027)
dummy 1993	-0.426**(0.236)	-0.477** (0.247)		-0.055 (0.045)	-0.017 (0.044)	
dummy 1994	-0.635**(0.228)	-0.703**(0.241)		-0.078 (0.066)	-0.011 (0.068)	
dummy 1995	-0.846**(0.223)	-0.925**(0.237)		-0.022(0.094)	0.086 (0.098)	
dummy 1996	-0.544**(0.240)	-0.604**(0.253)		-0.125**(0.061)	-0.071 (0.061)	
$\lambda(.)^a$				0.458 (0.89)	-0.532 (0.918)	-0.384 (0.758)
number of firms	369	369	369	369	369	369
observations without exit	1678	1678	305	1525	1525	305
observations with exit	64	64	64	64	64	64
log likelihood for normal/ adjusted R^2	-259.3	254.43	-164.3	adj R^2=0.029	adj R^2=0.043	adj R^2=0.151
white test for hetero-scedasticity				X^2=52.6 (0.49) NO	X^2=68.1 (0.94) NO	X^2=62.3 (0.081) YES

a. The adjustment for sample selection comes from the linear probit models in [1] and [3].

Table 11, of course, only gives the results of a comparison between firms with collective layoffs. To measure the impact on survival and growth of the collective layoffs themselves, we performed panel estimates using 4 dummy variables that indicate that there has been a collective layoff with or one without relocation in the year t or t-1. The results are reported in table 12.

The probit estimates show that a collective layoff without relocation has a significant negative impact on the probability for a firm to survive in the same year. A collective layoff also has a negative impact on the survival in the following year, but this effect is not significant. For firms with relocation, collective layoffs had no significant effect on survival. But the absence of significance is more due to the large standard errors than to the smaller coefficients.

In columns [3] and [4] we report results on the growth effects of collective layoffs. These are always significantly negative. There are no significant differences between firms with and without relocation. The total impact on growth of a collective layoff on average employment is given by the weighted[1] sum of its effect in the year of the layoff (as given by the coefficient for the year t), and that in the next year (as given by the coefficient for the year t-1)). This total impact can be compared with the results for the direct effects of collective layoffs computed in table 9, using a different source[2].

Note that the result that multinational firms have a higher growth rate is maintained in these regressions. We also continue to find a positive effect on growth and survival of (an increase in) capital intensity in the year t-1 in these panel data estimates.

TABLE 12 - The direct impact of collective layoffs with and without relocation[a] on survival and employment growth in the period 1992-1996.

	PROBIT (panel) yearly survival, size measured in 1990 [1]	PROBIT (panel) yearly survival, size measured in 1990 [2]	OLS (panel) yearly growth, size of 1990 [3]	OLS (panel) yearly growth, size of 1990 [4]
intercept	1.078**(0.478)	-0.148(1.898)	-0.319 (0.199)	0.459 (0.546)
y_{i0}	0.107**(0.047)	-0.515(0.293)	0.017 (0.015)	-0.206**(0.072)
ln(age)	0.034 (0.119)	2.226 (1.33)	-0.064**(0.025)	-0.112 (0.385)
$(y_{i0})^2$		0.039 (0.028)		0.005 (0.004)
$(\ln(age))^2$		-0.501**(0.242)		-0.045 (0.071)
$(y_{i0})\ln(age)$		0.093 (0.093)		0.056**(0.019)
ln(capital intensity in t-1)	0.109**(0.048)	0.109**(0.049)	0.047**(0.018)	0.035**(0.018)
(part of) multinational firm in 1996			0.070**(0.031)	0.068**(0.32)
layoff without relocation in (book) year t	-0.32**(0.147)	-0.316**(0.149)	-0.103**(0.042)	-0.066**(0.024)
layoff without relocation in (book) year t-1	-0.20(0.147)	-0.191 (0.151)	-0.237**(0.048)	-0.214**(0.049)
layoff with relocation in (book) year t	-0.086(0.325)	-0.193 (0.327)	-0.098**(0.029)	-0.092**(0.030)
layoff with relocation in (book) year t-1	-0.163(0.281)	-0.157(0.294)	-0.250**(0.064)	-0.235**(0.063)
dummy 1993	-0.359 (0.238)	-0.398 (0.248)	-0.019 (0.039)	-0.0005 (0.038)
dummy 1994	-0.577**(0.230)	-0.634**(0.242)	-0.047 (0.061)	-0.006 (0.06)
dummy 1995	-0.819**(0.224)	-0.891**(0.236)	-0.029 (0.089)	0.043 (0.092)
dummy 1996	-0.553**(0.241)	-0.608**(0.253)	-0.131**(0.061)	-0.093 (0.061)
$\lambda(.)$			0.691 (0.874)	0.011 (0.893)
number of firms	369	369	369	369
observations without exit	1678	1678	1525	1525
observations with exit	64	64	64	64
log likelihood for normal/adjusted R^2	-256.6	251.9	adj R^2=0.072	adj R^2=0.084
white test for heteroscedasticity			X^2=66.1 (0.94) NO	X^2=77.0 (0.99) NO

a. For firms with several layoffs in the observation period that relocated at least once, all collective layoffs were considered to be with relocation. This was done to avoid errors in the year 1996, where the information on relocation was not available. Tests with collective layoff dummies where this was not imposed did not lead to significantly different results.

1. The total effect with respect to the year t-1 of a shock in year t is given by: impact in year t+ impact in year t+1 *(1-impact in year t).
2. The effects in table 8 and 9 are based on data on the number of workers fired from the regional employment services. The date used on employment in table 11 are based on annual account data only. The total negative effect on growth of a collective layoff without relocation as implied by the linear specification in table 11 is given by 0.103+0.237(1-0.103)=0.316. That of a collective layoff with relocation by 0.098+0.250(1-0.098)=0.324. These effects are close to the average effects of 0.375 and 0.361 reported in table 8.

E. Conclusions

In this paper, we estimated the effects of collective layoffs on the average employment level of a sample of 369 industrial firms that had a collective layoff in the period 1990-1996.

On average, a collective layoff leads to a reduction of the number of workers by 1/3. This average includes the cases where a collective layoff directly leads to the end of all activities. At the firm level, no difference was found in the effects on employment between collective layoffs with and without relocation. That relocation is responsible for a relatively large share of the job loss due to collective layoffs, is entirely due to the fact that it is more frequent in large firms. Large firms are also hit more frequently by collective layoffs in general. Once it occurs, though, a collective layoff has a relatively higher impact on the employment of a small firm than on that of a large firm.

Among industrial firms, multinational firms had a higher incidence of collective layoffs, and collective layoffs with relocation. However, compared with other firms with collective layoffs, they had significantly higher growth rates of employment. An explanation is that these firms are less reluctant to fall back on a collective layoff (with or without relocation), for reasons of rationalising, labour saving or globalisation. Hence, for smaller and uninational firms, a collective layoff is more often a sign of severe financial and/or managerial problems.

Our estimates for industrial firms with collective layoffs confirm the results of empirical studies in other countries and for other groups of firms, finding that a firm's size, its age and capital intensity have a positive impact on its probability to survive. At the same time size and age are negatively correlated with the average employment growth. This result is found even after controlling for the sample selection problem mentioned before, when using a quadratic specification.

The results provide no convincing evidence that - among firms with collective layoffs - capital intensive firms have higher employment growth rates than others. However, an increase in the capital intensity within the same firm was found to have a positive effect on growth. This attracts the attention to the importance of (recent) investment as a determinant (or at least an indicator) of future growth.

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III Appendix

TABLE 13 - Averages and standard deviations of variables used in regressions, industrial firms with collective layoffs

	description	305 survivors, mean for period 1992-1996	305 survivors, only t=1992	whole group of 369 firms, only t=1992
y_{it}-y_{it-1} or y_{it}-y_{i0}	Change in the ln[a] of employment in year t with respect to t-1 or the base year[b].	-0.115 (0.425)		
y_{it-1} or y_{i0}	ln of employment in year t-1 or the base year	5.081 (1.455)	5.28 (1.366)	5.186 (1.335)
ln(age)	ln of firm age in years in year t	2.919 (0.49)	2.79 (0.562)	2.780 (0.578)
$(y_{it-1})^2$ or $(y_{i0})^2$	square of ln of employment in year t-1 or the base year	27.936 (15.336)	29.735 (15.168)	28.671 (14.56)
$(\ln(age))^2$	square of ln of age in year t	8.760 (2.592)	8.096 (2.778)	8.062 (2.823)
(y_{it-1})ln(age) or (y_{i0})ln(age)	product of ln of employment of firm age and ln of employment in year t-1 or the base year	14.919 (5.187)	14.874 (5.162)	14.534 (5.042)
ln (capital intensity)	ln of the stock of tangible assets divided by employment, evaluated in year t-1 or the base year.	6.696 (1.145)	6.633 (1.096)	6.563 (1.073)
(part of) multinational firm	firm is parent or daughter of multinational firm	0.728	0.728	0.666
relocation in 1990-1995	firm has had collective layoffs with relocation in period 1990-1995 according to unions	0.184	0.184	0.222
$\lambda(.)$[c]	variable that corrects for sample selection bias	0.079 (0.055)	0.29 (0.136)	

a. The ln is the natural logaritm.

b. The base year in the regressions is the year 1991.

c. The mean of $\lambda(.)$ in column 1 results from a linear specification of the probit model for yearly exit. The one given in the second colum is that of the linear probit model for the whole period 1992-1996.