Leaving teaching in the French-Speaking Community of Belgium: a duration analysis.

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This paper aims at a better understanding of the factors influencing the decision of young graduates who entered teaching to stay in that profession. The field of research covers secondary education teachers in the French-Speaking Community of Belgium. The data analysed comes from an administrative database containing historical records of 50,000 individuals who started teaching between 1973 and 1996. The analysis is carried out assuming a proportional hazard model and using the discrete-time method initiated by Prentice and Gloecker (1978). One of the main results is that the risk of exit is dramatically more important during the first periods of employment. That this risk tends also to increase over time suggesting that drop-out among young recruits is higher now than it was in the past. Location and labour market conditions seem to be of little impact. The risk of exit is the same in rural and urban areas and across provinces wherein unemployment rates vary dramatically. Finally, the significant deterioration of pay conditions (in relative terms) since the mid-80's has had no significant impact on the risk of exit. Of greater importance are supply-side (organisational) elements like the level of centralisation of recruitment decisions or the level of asymmetry between tenure and non-tenure personnel regarding job protection, access to full-time position…

Theme: Public sector labour markets.

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Introduction

Teachers form, with pupils, one of the essential ingredients of an educational system's functioning. Yet, and this might seem a bit paradoxical, very few empirical studies are devoted to them. Our knowledge of the European Labour Market for Teachers is insufficient and our understanding of their career path is limited in scope. This is particularly true of the French-Speaking Community of Belgium. We know with great accuracy the number of teachers in activity. We have a rather good idea of age pyramids or gender distribution, but we know almost nothing about the factors influencing the decision to opt for teaching instead of other sectors of activity (banking, insurance, advertising...). The same is true of the determinants of the propensity to stay in teaching. The point is that these two processes are crucial determinants of the composition and the quality of the teaching work-force at a certain point in history. And both elements are apparently central to good education. Recent work (Hanushek, Kain & Rivkin,1998) on human capital production stresses the importance of 'teachers' effects' and suggests that the quality of teachers could be more important than the quality of schools (i.e. the so-called 'school effects') in maximising pupils' academic prospects.

What follows intends to improve our understanding of teaching careers. The exercise is empirical in essence. The material analysed here comes from administrative databases used to organise the payroll of secondary school teachers in the French-Speaking Community of Belgium. The main interest of this data resides in the fact that it enables us to follow the career of a very large number of individuals. This data does not help us understand the choice made by graduates when entering the teaching profession: it should be clear that all the results presented here are conditional on entry. Similarly, the data set contains no information about the destination of the teachers in case of exit. But despite all these restrictions and limitations, this data is adequate to answer the question of who stays in teaching and why. The results presented here suggest that this question is important and should be addressed by those in charge of education policy.

From a methodological standpoint, results presented and commented here are based on duration analysis (Lancaster, 1990). Duration models are aimed at identifying the determinants of the probability for an individual to move from one status to another (e.g. from life to death, from unemployment to employment...) after a certain period of time. The typical dependent variable of duration models is the time spent by individuals in a certain situation. Variables characterising the individual, his/her context and the policy he/she has benefited from are the most frequent independent variables used in duration analysis (Allison, 1997).

Historically, these models were developed in the medical sector to assess the impact of some molecules or treatments in terms of improved survival prospects. More recently, these models have been applied in labour economics in order to analyse unemployment spells and their determinants (d'Addio, 1998; Dejemeppe, 1997), in particular the efficiency of some training programs. In the field of Economics of Education, their use is even more recent. The first example of their systematic use is by Dolton & van der Klaauw (1995) and Murnane (1991). The focus is similar to ours: these authors are concerned with the determinants of the probability of exit among teachers in the UK and US primary and secondary school systems.

1. Presentation of the educational system of the French-Speaking Community of Belgium

Like many educational systems in Europe, the secondary school system in the French-Speaking Community of Belgium is financed by public money. Teachers get paid by the central administration following uniform salary scales relating teachers' wages to diploma and seniority. This principle is applicable to teachers working in public schools (i.e. those directly ruled by the Ministry of Education, by the Provincial Authority or by the local Municipal Authority) but also for those working for independent Catholic affiliated school boards[2]. Yet, depending on the authority from which schools depend, some aspects of the decision-making process — typically hiring decisions — are more or less centralised. Independent Catholic boards generally delegate the responsibility for hiring teachers to school principals, but that responsibility is much more centralised in public schools, especially those directly operated by the Ministry of Education.

Another central component of the policy towards teachers worth mentioning here relates to *tenure*. Like in most Western educational systems, access to tenure[3] is actually based on seniority. This policy *a priori* limits professional mobility among teachers after a certain number of years. But its logical corollary, *ceteris paribus*, should be a higher exit rate among (young) teachers without tenure. From the employer's side one can say that the last-in-first-out rule prevails. During periods of budgetary austerity or demographic downturn — generally synonymous with job cuts — non-tenure teachers are always the first to be asked to go. In public ministerial schools, non-tenure teachers can also be pushed out by senior colleagues using their right to geographical relocation. Finally, the prevailing quasi-market principle, synonymous with free school choice and per-pupil financing (Vandenberghe, 1998), can also reinforce the asymmetry between tenure and non-tenure staff members. On the quasi-market,

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[2] Representing approximately 50% of the system.
[3] And the right to benefit from the generous public pension system.

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even if the total number of jobs is constant system-wide, some schools can experience dramatic year-to-year enrolment falls. And the cost of those micro-level shifts is likely to be primarily borne by those with a shorter history within the system.

2. Data

Our data set contains information about teachers who worked in secondary schools in the French-Speaking Community of Belgium between 1972 and 1997. It consists of an administrative file containing the list of teachers on the payroll on the 15th of January of each year. We possess payment records for 113,218 teachers who worked in a secondary school for periods (duration) ranging from 1 year or less to 25 years. These periods are measured discretely, by intervals of 1 year. All these teachers have not been retained here. A double restriction has been imposed. Were taken out of the reference set : (1) teachers who were already in the system in 1972 (source of left-hand censoring)[4] (2) observations corresponding to the 1997-1998 interval[5]. This triple restriction led to a reduced total of 50,041 teachers with a majority of women (56.30%).

For each of these individuals we have as many rows (spells) as years in total duration. This means t rows for a person who spent t years in the system. This information helps us build the dependant variable of the model, i.e. a variable indicating if the teacher is still present or not at the end of the year t. Besides, payment records contain all sorts of additional information forming the set of explanatory variables we can use to explain the propensity of individuals to quit the system. There are variables describing each teacher's socio-demographic profile (gender, age...), the context of his/her school (location of the school, the governance regime...) and his/her working conditions (number of schools attended, weekly timetable, wage).

Given the specificity of female career paths, data presentation and analysis carried out here systematically distinguish males and females. The latter (56.30%) slightly outnumber the former (43.70%). For men, the observed exit rate is 50.5% while it reaches 49.9% for their female colleagues.

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[4] One could, for those subjects, measure the length of their stay within the system from the moment our observation starts, but this measure would underestimate the total and effective length of their stay. One exit after a relatively short observed period of time might in fact correspond to a relatively long stay in the system; one that started a long time before 1972 for example.

[5] The 15th of January 1997 being the last point of observation, we *a priori* ignore what occurred beyond that point. It is thus totally impossible to analyse the risk of exit during that interval. To avoid what one calls 'right censoring' we consequently eliminate all observations made on that date, and that includes individuals who entered into the system in 1997.

The first variable describes the age of entering for the first time in teaching (AGE_S). It should be considered as a proxy for the level of qualification in terms of final degree or previous professional experience: the older the new teacher, the higher is his qualification. Table 1 shows that it is on average slightly higher for male teachers (30.89 years old) than for female ones (28.87 years old) therefore suggesting gender-based qualification differentials.

Insert Table 1 about here

The second group of variables reflects the working conditions of the teacher. Table 1 suggests that the average weekly workload (WLOAD) is relatively high for men (94% of the fulltime reference), but also for women (90%). We also measure the dispersion of this timetable between one or several establishments (NSCHOOL). It is superior to unity for both men (1.13) and women (1.15), this last figure suggesting that women get slightly less comfortable working conditions. Then comes the number of provinces in which they work (NPROV) acting as a proxy of the distance between the different schools the teacher works in. Table 1 reveals that most teachers work within the same province (average number of provinces=1).

In the same group of variables we have also introduced a variable capturing the cumulated wage increment teachers have benefited from since their entrance (ISAL). This variable strongly reflects the variability of the wage policy over the last 25 years illustrated by Figure 3. Referring to Table 1, we notice that on average salary progression observed is 65% for men and 69% for women, but that deviation is high especially for women.

Insert Figure 3 about here

Then come variables that are likely to capture local and historical contextual conditions. The unemployment rate in the province (LOGUN) is likely to reflect the impact of the state of the (external) labour market. The dummy variables for the province in which the school is located (PROV) should enable us to capture other influences of the local socio-economic context. In that group of contextual factors, we have also included the level of employment in the educational system (LNEMPL) which is likely to capture the tightness of the (internal) labour market for teachers. One could indeed assume that the larger this market, the more easy it is for young teachers to stay in the system. We introduce that variable also because the labour market for teachers might be very specific and segmented from other sectors of the labour market. We have finally introduced the calendar date (CAL) — not to confuse with the duration variable (DUR). This variable should seize the impact of any other determinant of the risk of exit, at any moment in the history of the system, not captured by the previous variables.

Finally, there is a categorical variable corresponding to the so-called 'réseau' to which the school belongs (RES). As mentioned in section 2, schools can be ruled directly by the Ministry of Education (RES=MIN), by the Provincial Authority (RES=PROV), by the local, Municipal Authority (RES=MUN) or by independent Catholic affiliated boards (RES=CAT).

3. Descriptive analysis: probability (cumulated and instantaneous) of leaving teaching

It is convenient and useful to start with a descriptive analysis of the careers of the 50,041 teachers who entered the system at some point between 1973 and 1996. Figure 1 displays the survival curve (Kaplan-Meier Estimate), which measures the (cumulative) probability of staying within the system until the beginning of interval (year) k or later.

Algebraically, this probability corresponds to the product of conditional survival probabilities up to the beginning of interval $t=1...k$ (equation 1). This can be written as the product over $t=1...k$ of $(1- q_t)$, q_t being the probability of exit during interval t. Following the LIFETABLE method used here, q_t results from the division of d_t (the number of teachers leaving during interval t) by $n_t - w_t/2$ (the number of individuals at risk during that interval)[6].

$$S(k) = \prod_{t=1}^{k} \left[1 - \frac{d_t}{n_t - \dfrac{w_t}{2}} \right]$$ [1]

Looking at the results now (Figure 1), one can see that the (cumulated) survival rate at a 10 year horizon (k=10) is of 55% for men and 53% for women. After 24 years this figure has declined to 39% for men and 42% for women. The well-known tendency of the profession to feminize (Vandenberghe, 1996) does seem to been imputable to the fact that the system cannot retain as many male teachers over the long run.

These rates might appear to be relatively high in absolute terms but their are comparable to those describing the situation in the UK or in the US. Stinebrickner (1997), reporting on his analysis of a sample of primary and secondary education teachers in the US, mentions a

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[6] Note that half the individuals still there on the 15th January 1997 while beginning their year ($w_t/2$) — right-censored individuals — come in deduction of the total number of individuals at risk (n_t). The justification for this is that the method treats censored cases as if they were censored at the midpoint of the interval and consequently out of the 'at-risk' population beyond that midpoint.

survival probability of less that 40% at a 7 year time horizon suggesting a higher cumulated exit rate. By contrast, Dolton & van der Klaauw (1995), analysing the situation in the UK, observed a survival rate of 66% after 5 years of teaching. The corresponding figure we obtain for the French-Speaking Community of Belgium is of 60% for male teachers and 58% for their female colleagues.

Insert Figure 1 about here

Figure 2 usefully complements the survival function. It shows the hazard function which describes the instantaneous probability of exit during interval *k*. Analytically, hazard rate *h(k)* is obtained by dividing the number of teachers who left during that interval (d_k) by the sum of individual-time-units of exposure to the risk of exit (equation 2). As the unit of time used here is unitary (1 year), total exposure time to risk is simply equal to the total number of individuals present at the beginning of interval (n_k), minus half the number of censored observations ($w_k/2$) and half the number of exit cases ($d_k/2$)[7].

$$h(k) = \frac{d_k}{n_k - \frac{w_k}{2} - \frac{d_k}{2}}$$ [2]

One can figure out, by looking at Figure 2, the extent to which the risk of exit is important during the early years of teachers' careers. It is extraordinarily high during the first year, then drops regularly to reach a low and relatively constant level beyond year 7, suggesting that instability is essentially concentrated on newcomers, and that there is almost no career mobility once some seniority has been acquired.

Insert Figure 2 about here

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[7] Still supposing that exit or censoring occurs at the midpoint of the interval

4. Multivariate analysis

The next stage logically consists of trying to identify the factors influencing the hazard rate of teachers: i.e. carrying out a multivariate analysis of the probability of exit. The instrument is a duration model. We have opted here for a semi-parametric continuous time model that we estimate by using per interval (1 year) grouped data. This preference for a continuous time model echoes our concern not to exclude *a priori* the possibility of duration (total time spent in teaching) being a continuous rather than a discrete value.

4.1. Prentice and Gloecker model and complementary log-log transformation

a. Presentation of the model

We assume, following Cockx (1997), a set of *i=1…N* individuals (graduates) that entered the teaching profession at some point between 1973[8] and 1996. Suppose that the information gathered about these individuals is grouped in K intervals of equal length[9]): *[0,1[;[1,2[;[2,3[…[K-1,K[.* For each individual, we know the interval during which he/she left teaching (if relevant): *k-1≤t< k,* with *t* designating the occurrence of the random variable *T*.

If for the underlying continuous time model we note the hazard function[10] in *t* by *h(t),* we have by definition[11] that the survival function[12] is equal to:

$$S(t) = \exp\left[-\int_0^t h(s)ds \right] \qquad [3]$$

If we now consider the probability of exit during an interval P_k — i.e. the discrete version of the hazard rate defined above — we have:

$$P_k \equiv \Pr\left(k-1 \le T < k \,\middle|\, T \ge k-1\right) = \frac{S(k) - S(k-1)}{S(k)} = 1 - \exp\left(-\int_{k-1}^k h(s)ds \right) \qquad [4]$$

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[8] Teachers already present in the system in 1972 are not considered in order to avoid left-censoring problems.
[9] Which is precisely the case in our data set (see above).

[10] The hazard function is defined as: $\lim_{\Delta \to 0} \frac{\Pr\left(t \le T < t+\Delta \,\middle|\, T \ge t\right)}{\Delta} = \lim_{\Delta \to 0} \frac{F(t+\Delta) - F(t)}{\Delta S(t)} = \frac{f(t)}{F(t)}$ with *F(t)*

being the cumulative function and *f(t)* the probability distribution.
[11] See for example Greene (1993), page 717-178 for a demonstration of this.
[12] *S(t)=Pr(T=t)=1-F(t)*

where the right-hand term of the last expression directly stems from equation 3.

If $X_i(t)$ designates a vector of independent variables characterising individual i in t and if we assume that the hazard function is proportional and semi-parametric (Prentice & Gloecker, 1978), we have:

$$h_i(t) = h_0(t)\exp\left(X_i(t)'\beta\right) \tag{5}$$

where β is the vector of unknown parameters to be estimated and $h_0(t)$ the baseline hazard common to all individuals $i=1...N$. Using this specification and assuming that variables $X_i(t)$ take values that remain unchanged within interval k, one can now express the probability that a teacher i leaves the system during interval k as:

$$P_i(k) = 1 - \exp\left[-\exp\left(X_i(k)'\beta\right)\int_{k-1}^{k} h_0(s)ds\right] \tag{6}$$

or, equivalently:

$$P_i(k) = 1 - \exp\left(-\exp\left(X_i(k)'\beta + \gamma(k)\right)\right) \tag{7}$$

where:

$$\gamma(k) = \ln\left[\int_{k-1}^{k} h_0(s)ds\right] \tag{8}$$

It is worth observing that expression 7 links a linear expression to a variable taking values ranging from 0 to 1, meaning that this model belongs to the class of Generalised Linear Models. The inverse function, generally called the complementary log-log function, takes the following form.

$$X_i(k)'\beta + \gamma(k) = \ln\left(-\ln(1 - P_i(k))\right) \tag{9}$$

b. Estimation of the model by maximum of likelihood

The next step is to estimate the parameters of expression 9 by maximum of likelihood. A natural starting point is the standard expression of likelihood commonly used in duration analysis (expression 10). The basic information used for estimation of most duration models consists of the total length of stay within a certain status. Here, the equivalent is the total time

spent by an individual in teaching, measured by the last interval in which he was observed (k_i). The dichotomy variable in expression 10 indicates whether there is right-censoring ($\delta_i = 0$) or not ($\delta_i = 1$) for individual i. The first term of the right-hand part of equation 10 represents the probability of exit during interval k_i and the second term the probability that exit occurs at a later stage.

$$L = \prod_{i=1}^{N} \left[\Pr(T_i = k_i)^{\delta_i} \Pr(T_i > k_i)^{1-\delta_i} \right] \tag{10}$$

But the basic information contained in our data set is not the total length of stay of individual i. Duration is in fact decomposed in sub-periods of equal length. Remember that we have t rows for a person who spent t years in the system. Meyer (1990) and Allison (1982, 1997) propose, in this case, to re-write the likelihood expression 10. Their basic idea is that the probability for individual i to leave the system during interval k_i is the product of the hazard rate (in its discrete version) specific to that interval by the cumulative probability of survival for all preceding intervals:

$$\Pr(T_i = k_i) = P_i(k_i) \prod_{k=1}^{k_i-1} (1 - P_i(k)) \tag{11}$$

and similarly:

$$\Pr(T_i > k_i) = \prod_{k=1}^{k_i} (1 - P_i(k)) \tag{12}$$

Using these two expressions, the likelihood expression 10 can be restated as:

$$L = \prod_{i=1}^{N} \left(\frac{P_i(k_i)}{1 - P_i(k_i)} \right)^{\delta_i} \prod_{i=1}^{N} \prod_{k=1}^{k_i} (1 - P_i(k)) \tag{13}$$

and if one adopts the convention that λ_{ik} equals to 1 if individual i leaves teaching during interval k and 0 otherwise, one finally obtains:

$$L = \prod_{i=1}^{N} \prod_{k=1}^{k_i} \left(\frac{P_i(k)}{1 - P_i(k)} \right)^{\lambda_{ik}} \prod_{i=1}^{N} \prod_{k=1}^{k_i} (1 - P_i(k)) \tag{14}$$

The last thing we need to do is to introduce the specification of the hazard function (in its discrete formulation) $P_i(k)$(equation 7) to get a likelihood expression function of unknown parameters $\gamma(k)$ and β.

It is important to note that this expression 14 is equivalent to the likelihood function of a regression model in which the dependent variable is binary (λ_{ik} to 0 or 1 here); model that can be estimated using common software packages such as SAS, STATA or SPSS[13]. The estimation of these parameters then follows along the usual algorithm of likelihood maximisation. A last point worth noting it that the estimated parameter $\widehat{\beta}$ corresponds to the parameter of the exponential semi-parametric Cox function and should be interpreted in the same way, i.e. in terms of relative hazard rates. For expression 5 the ratio between the hazard levels in case of unit variation of an independent variable equals $\exp(\widehat{\beta})$.

c. Unobserved heterogeneity

The proportional hazard rate model used here is very practical in a sense because it immediately delivers an indication of the dependency to duration (be it with a parametric or non parametric specification of $h_0(t)$ in expression 5). This information is important for policy making. A strong negative dependence to duration suggests that the first months exert a critical impact on young graduates' commitment to the educational sector, while a positive dependence to duration indicates the presence of burnout problems at later stages in the career. Each case requires specific answers and measures. Hence, getting access to unbiased coefficients for duration dependence is essential.

To avoid biases, we inevitably have to envisage the presence of unobserved heterogeneity among individuals. It seems reasonable to assume that some young teachers are highly motivated by their job and want to spend their whole life in the same job. Others, by contrast, are probably unmotivated and consider their current teaching activity as temporary. They stick to that position until something more appealing becomes available in other segments of the labour market.

To understand the sort of bias unobserved heterogeneity potentially creates, let us assume that 50% of entrants are highly motivated and that the other 50% are not motivated at all. The first group has a (unobservable) risk of exit of 0% while the second group is characterised by a risk of 100%. By confounding the two groups as we have done so far, we will measure a risk of exit of about 50% during period 1 and of 0% during the following periods. The (false)

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[13] In SAS in particular, one can use the PROBIT procedure, provided one indicates that the relation between the discrete hazard term ($P_i(k)$) and the linear equation is of the complementary log-log type

conclusion we would draw from these observations would be that every individual surviving the first year is likely to stay indefinitely in the educational sector. The (inadequate) policy recommendation would be to re-deploy human resources policies in favour of entrants and pay more attention to what happens during their first year of activity.

In more analytical terms, the presence of unobserved heterogeneity leads to over-estimated (underestimated) coefficients for negative (positive) duration dependence. Lancaster (1990) demonstrates that the coefficient of the other variables $X_i(k)$ can be biased too.

To remedy this problem, following Liang & Zeger (1986), we introduce into the hazard rate function an additive factor v_i representing the heterogeneity component for which we assume a normal (Gaussian) distribution, $N(0, \sigma_i^2)$. Coefficients to estimate now include the (unique) parameter of the normal function i.e. the variance σ_i^2. This can be achieved in STATA version 6 using the *xtclog* command (STATA, 1999). The higher the estimated variance of v_i, the more serious the presumption of unobserved heterogeneity.

4.2. Results

a. Comparison of the estimated models.

The first explanatory variable examined here is the number of years spent in the system (DUR)[14]. The estimated parameter for this variable corresponds to the term $\gamma(k)$ in equation 9. The other explanatory variables retained here and presented in section 2 correspond to $X_i(k)'$ in the same equation. But before moving to the results *per se*, it is worth discussing the way to model the relationship between hazard and duration (i.e. the specification of the baseline hazard function). It is rather common to assume that the baseline hazard corresponds to Weibull function ($h_0(t) \equiv \alpha \ln(t)$) with a unique parameter α reflecting the sign of duration dependence. The alternative consists of using a non-parametric model (actually a series of dummies for the different values of the duration variable). Both models have been estimated here. Yet, Table 2 and the likelihood ratio test[15] it contains, let us believe that the non parametric model for baseline hazard significantly improves the model fit. Although coefficients have been systematically reported for the Weibull model in Tables 4.1 & 4.2, we shall concentrate hereafter on the results of the non-parametric model (Tables 3.1, 3.2 & 3.3).

Insert Table 2 about here

a. Negative dependence to duration, especially for women.

Table 3.1 (columns 1 & 2) clearly shows that there is a negative dependence of the risk of exit to duration. This result confirms what was already visible in Figure 2 above : the instability is much higher at the very early stage of the career than after a certain number of years. In column 1, the highest (significant) coefficient for the categorical variable DUR corresponds to the first year of the career. Given the proportional hazard model used here (equation 9), an estimated coefficient $exp(\hat{\gamma})$=6.99 for men means that the risk of exit for a male teacher in his first year is almost 7 times higher than that for his 24th year of activity. This ratio decreases sharply up to a point (approximately the 8th year). The situation of women in their first years seems to be even more unstable than when they have a long history within the system. Compared to female teachers it their 24th year, those in their first year show a risk of exit 23.07 times higher. This suggests a higher negative dependence to duration among women.

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[14] Remember that the value taken by DUR represents the number of intervals starting the 15th of January of one year and finishing at the same date one year later.
[15] The likelihood ratio is computed as 2 times the difference of the log-likelihood of the models to be compared. It follows a χ^2 with p degree of freedom, where p is the number of additionally included variables (23 here).

This result of negative dependence to duration is also visible in Table 4.1 (see Appendix) in which we used a *parametric* approach to estimate the baseline hazard. The specification retained is $h_0(t) \equiv \alpha \ln(t)$ generating a Weibull model [$h(t) = t^\alpha \exp(X'\beta)$] in which the sign of the estimated parameter $\hat{\alpha}$ directly measures the negative or positive dependence to duration. In columns 1 & 2, $\exp(\hat{\alpha})$ is clearly inferior to 1 (which means that $\hat{\alpha}$ is negative) and statistically significant.

The robustness of the negative duration dependence is confirmed by the coefficients of the model with unobserved heterogeneity (columns 3 & 4 in Table 3.1). The differences with the first set of coefficients are negligible. This is no surprise if we refer to the statistics displayed at the bottom of Table 3.3. The estimated value for the variance of the Gaussian distributed unobserved heterogeneity term is almost nil ($\ln \sigma_v^2 = 0.00$). The absence of significant unobserved heterogeneity is confirmed by the likelihood ratio tests. Bias caused by unobserved heterogeneity does not seem to be a problem here.

Broadly speaking, the observation of negative dependence to duration is well in line with a central component of the policy in place for many decades now in education: access to tenure solely based on seniority. This policy tends to limit dramatically professional mobility among teachers after a certain number of years, but its corollary is that the turnover is particularly high among young teachers.

b. Increased hazard rates over(calendar) time.

The second interesting result visible in Table 3.2 concerns the impact of the calendar date (CAL) at which exit takes place. The coefficient of the different values taken by this variable represents the ratio between the risk of exit for the particular year considered (e.g. 1973) and the year of reference (1996 here). The general tendency over the last 25 years for both men and women was a rising probability of exit at a certain point of the career. *Ceteris paribus*, the risk was much lower in 1974 et 1975 than it was in 1996, as suggested by the value of $exp(\hat{\beta})$ significantly inferior to 1 in Table 3.2. For both men and women, the hazard rate rose gradually between 1976 and 1986-87. It then decreased a little without recovering its early 1970's values but then rose again to reach its climax during the interval starting on January 15th 1996.

Insert Tables 3.1, 3.2 & 3.3 about here

c. No significant effect of starting age

Referring now to socio-demographic variables in Table 3.3, one observes that starting age (AGE_S) has no significant impact on the risk of exit. This result seems to suggest the absence of difference between individuals who graduated from university or non-university tertiary education.

d. Significant but (very) small effect of wage progression on risk of exit, for men only

One of the most interesting results in Table 3.3 is to observe that the wage policy of the last 15 years (ISAL) — leading to a relative deterioration of teachers' purchasing power as suggested by Figure 3 — has had no significant impact on the probability of exit of women. The coefficient is significant for men, but very small in magnitude: a 10 points additional wage premium[16] would reduce their hazard rate by 1.1 percent.

e. Part-time work increases hazard rates dramatically

By contrast, the variable describing the teachers' weekly timetable (WLOAD) is significant and of higher magnitude, particularly for men. The higher the number of periods taught weekly, the lower the risk of leaving teaching (Table 3.3). A male teacher increasing his weekly workload by 10 points reduces his risk of exit by 7.7%.That percentage if of 7.1% for their female peers.

f. Working in several schools reduces risk of exit

More surprisingly perhaps, working in several schools rather than one (NSCHOOL) is a significant source of reduction of the risk of exit (Table 3.3). The drop is of about 26% for men working in two schools vs. men teaching in a single school. The reduction is of 30% for women. Finally, distance between schools, as approximated here by the number of provinces in which the teacher is active (NPROV), seems to reduce the risk of exit of women by 40%. Yet, this result should be used carefully as descriptive statistics displayed in Table 1 indicate that the average number of provinces of work is very close to unity.

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[16] That premium is, on average, of 60%. See bottom of Table 1.

g. No incidence of outside and inside labour market conditions but lower hazard in some provinces

Although this result might reflex poor quality of data, it is surprising to observe in Table 3.3 that labour market conditions — both external and internal ones — do not help us understand the exit patterns. Teaching employment opportunities estimated by the log of the employment (LNEMPL) level appear insignificant in all regressions. For men, the same is true for the overall unemployment rate (LOGUN) of the province. The positive and significant coefficient of this variable for female teachers (1.29) is puzzling: a higher unemployment rate in the province tends to increase the hazard rate of women, which goes against the idea that a lower 'opportunity cost' should lead to a lower turnover in the teaching sector.

Beyond these dimensions, the province where the school is located (PROV) and all the other contextual elements that are linked to that dummy seem globally insignificant. Some exceptions deserves attention however. First the province of Hainaut. Its coefficients of 0.90 (for men) or 0.86 (for women) suggesting indeed that teachers working there have had a lower propensity to leave teaching, but for reasons that are not strongly related to the unemployment rate. By contrast, but for men only, the province of Luxembourg is synonymous of higher risk of exit.

h. A higher risk of exit when decision-making is more centralised

Finally, the variable describing the so-called 'réseau' (RES) to which the school is affiliated is highly significant. We know that organisational parameters vary along this line. For example, human resources management is more centralised in the schools directly ruled by the Ministry of Education (RES=MIN).The coefficients of Table 3.3 suggest that this might be of some importance in explaining the risk of leaving teaching. That risk is significantly lower in schools ruled by local authorities (RES=PROV or MUN) and subsidised Catholic schools (RES=CAT). Using the most centralised component of the system as a reference (MIN), one observes in Table 2.3 that the risk of departure is up to 46% lower (for women) in schools organised by municipal authorities for example (RES=MUN).

Conclusions

Results presented here strongly support the idea that a good understanding of teachers' propensity to leave teaching is necessary to develop sound educational policies. Previous work on human capital production functions suggests that 'teachers matter' (Hanushek, Kain & Rivkin,1998) and, further, that the characteristics of teachers could be more important than

the quality of schools in maximising pupils' academic prospects. Following this approach most observers and decision-makers usually focus on initial training and recruitment processes to influence the profile of the teaching body. Our analysis of secondary school teachers in the French-Speaking Community of Belgium strongly suggests that this approach might be insufficient. The likelihood of exit among young teachers can indeed be very high.

The actual profile of the teaching body is likely to be strongly influenced by the selection process that takes place at the beginning of the career. This conclusion is well in line with that of other research carried out in the UK for example (Dolton & van der Klaauw, 1995). A better understanding of the determinants of exit could lead to more effective policies aimed at reducing teacher turnover in shortage subjects or schools. In budgetary terms, more investment in teachers' initial training makes little sense if simultaneously turnover among young teachers remains high or keeps rising as suggested by our results.

Contrary to those of Dolton & van der Klaauw (1995), our results do not strongly support the idea that wage policies primarily affect teachers' turnover. Their might be an impact of wage progression rates on the decision of male teachers to stay, but of very small magnitude. Given our results, a more credible assumption would be that non-monetary working conditions matter. Access to a full-time teaching job dramatically reduces the risk of exit. Another central result is simply the strong negative dependence to duration. The risk of exit is dramatically more important during the first years of presence and becomes dramatically low at latter stages of the career, despite potential burnout. We interpret this fact as the logical consequence of both a last-in-first-out principle and an access to tenure solely based on seniority. These elements *a priori* limit professional mobility among older teachers. But its corollary seems to be a higher exit rate among (young) teachers without tenure. Increased hazard rates over (calendar) time observed in this research could simply be the consequence of a growing proportion of tenure teachers in the system, shifting the whole cost of flexibility the system needs on the newcomers.

Finally, the prevailing quasi-market of the French-Speaking Community of Belgium can also reinforce the asymmetry between tenure and non-tenure staff members. On the quasi-market with tenure based on seniority, the cost of inter-school shifts of market shares is primarily borne by those with a shorter history within the system. A tendency of greater use of school choice opportunities by pupils and families, already identified by sociologists in the 70's (Billet, 1977), could thus also explain a rising turnover.

Appendix

Insert Tables 4.1 & 4.2 here

Figure 1 — Survival function in the secondary education system. Breakdown by gender. Observation period 1973-1996
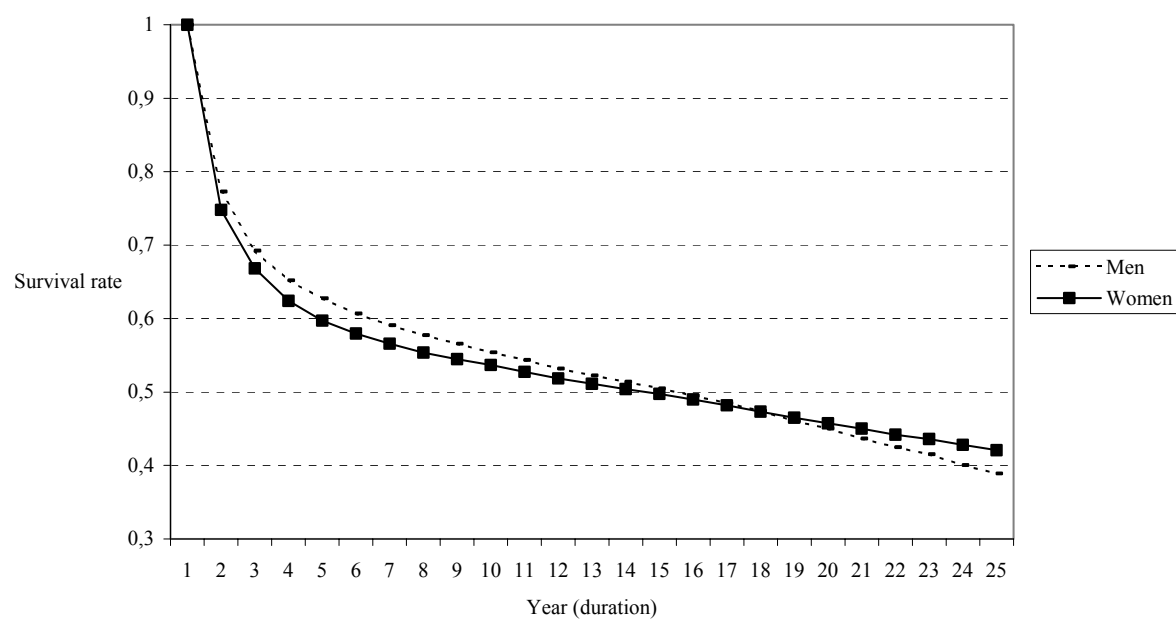
Figure 2 — Hazard rates according to duration in the secondary education system.
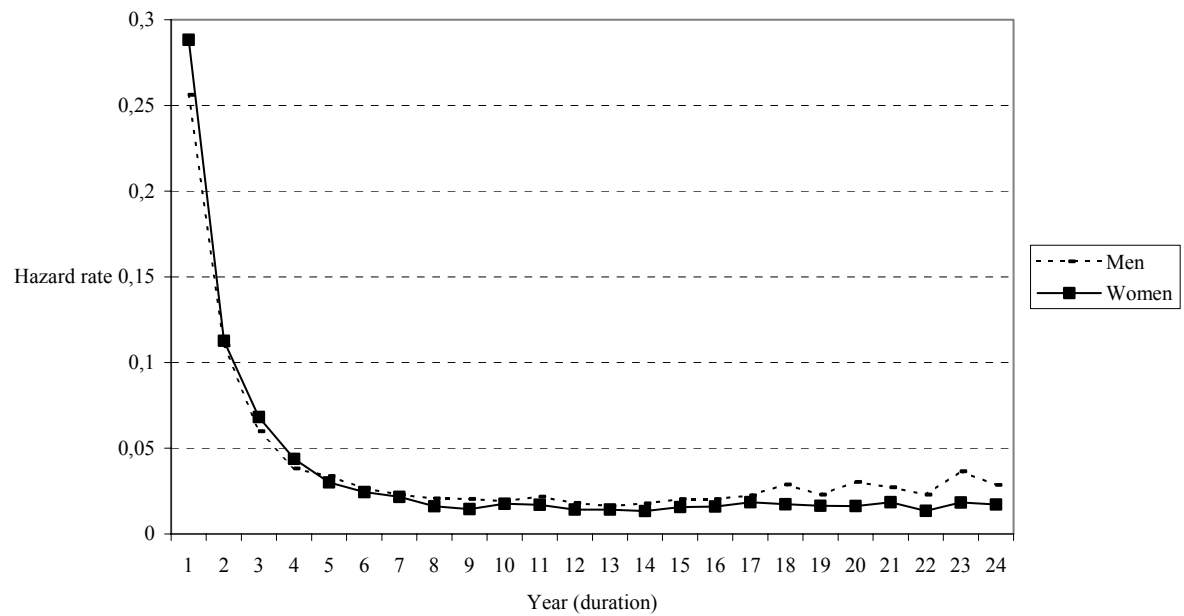*Breakdown by gender. Observation period 1973-1996.*

*Figure 3 — Consumption price index and nominal wage evolution for secondary school
teachers. French-Speaking Community of Belgium. Period 1972-1997. PRICE= consumption
price index SEN=seniority*
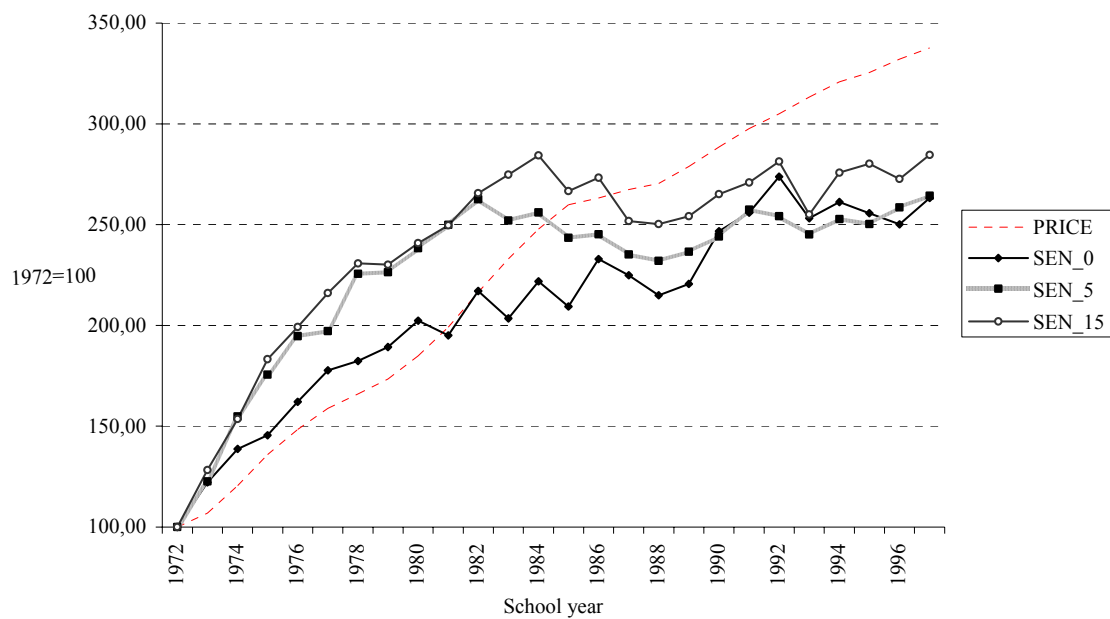
Table 1 – Descriptive statistics about the data set

	Men	*Women*	*Total*
Number of individuals	21,869	28,172	50,041
% individuals	*43.70*	*56.30*	*100.00*
Number of exits (individuals)	11,165	14,191	
Number of right censoring (individuals)	10,946	14,247	
% of exit (individuals)	*50.50*	*49.90*	
Number of spells (years)	191,938	236,963	
Number of exits (spells)	11,165	14,247	
Number of right censoring (spells)	183,758	225,532	
% exit (spells)	*5.73*	*5.94*	
Average number of spells (years)*	8.78	8.41	
Average starting age* : AGE_S	30.89	28.87	
	[7.48]	[7.16]	
Average weekly workload (1=fulltime)* : WLOAD	0.94	0.90	
	[0.23]	[0.24]	
Average number of schools of work* : NSCHOOL	1.13	1.15	
	[0.42]	[0.43]	
Average number of provinces of work* : NPROV	1.00	1.01	
	[0.07]	[0.08]	
Average maximal salary progression : ISAL (1=starting wage level)	1.65	1.69	
	[1.16]	[2.30]	

* Standard deviation in brackets

Table 2 – Likelihood ratio tests (1)

	Men	Women
Non_Para/Weibull $\cong \chi^2_{23}$ (2)	647.67 (p<0.001) H$_0$ rejected	307.06 (p<0.001) H$_0$ rejected
Non_Para/Weibull with Unobs. Hetero. $\cong \chi^2_{23}$ (2)	647.67 (p<0.001) H$_0$ rejected	307.06 (p<0.001) H$_0$ rejected

(1) $[2*(L^{NP}-L^{Weibull})] \cong \chi^2_p$ where p is equal to the number of additionally included variables (23 dummies for duration here); testing H$_0$ i.e. the unconstrained model (i.e. non-parametric baseline hazard) <u>does not</u> improve the fit relative to the Weibull specification,

(2) χ^2_{23} *at* 5% = *35.172*

Table 3.1– Multivariate analysis estimates. Proportional Hazard Model: non-parametric baseline hazard

Variable	exp(estimate) [S.E]			
			Unobserved heterogeneity	
	Men	Women	Men	Women
DUR =1	6.99*[0.24]	23.07* [0.26]	6.99* [0.22]	23.07* [0.25]
2	3.22*[0.24]	10.09* [0.26]	3.22* [0.22]	10.10* [0.25]
3	1.74*[0.24]	6.09* [0.26]	1.74* [0.22]	6.09* [0.25]
4	1.07 [0.24]	3.67* [0.26]	1.07 [0.22]	3.67* [0.25]
5	0.90 [0.24]	2.38* [0.27]	0.90 [0.22]	2.38* [0.25]
6	0.71 [0.24]	1.86* [0.27]	0.71 [0.22]	1.86* [0.26]
7	0.61*[0.24]	1.62 [0.27]	0.61* [0.22]	1.62 [0.26]
8	0.55*[0.25]	1.18 [0.27]	0.55* [0.22]	1.18 [0.26]
9	0.56*[0.25]	1.10 [0.27]	0.56* [0.22]	1.10 [0.26]
10	0.51*[0.25]	1.23 [0.27]	0.51* [0.23]	1.23 [0.26]
11	0.59*[0.25]	1.19 [0.27]	0.59* [0.23]	1.19 [0.26]
12	0.47*[0.25]	1.04 [0.28]	0.47* [0.23]	1.04 [0.26]
13	0.47*[0.25]	1.02 [0.28]	0.47* [0.23]	1.02 [0.27]
14	0.52*[0.25]	0.98 [0.28]	0.52* [0.23]	0.98 [0.27]
15	0.60*[0.25]	1.11 [0.28]	0.60* [0.23]	1.11 [0.27]
16	0.60*[0.25]	1.09 [0.28]	0.60* [0.23]	1.09 [0.27]
17	0.65*[0.25]	1.20 [0.28]	0.65 [0.23]	1.20 [0.27]
18	0.84 [0.25]	1.13 [0.28]	0.84 [0.23]	1.13 [0.27]
19	0.62 [0.25]	1.04 [0.28]	0.62* [0.24]	1.04 [0.27]
20	0.83 [0.25]	0.89 [0.28]	0.83 [0.23]	0.89 [0.27]
21	0.79 [0.25]	1.12 [0.29]	0.79 [0.24]	1.12 [0.28]
22	0.75 [0.27]	0.80 [0.31]	0.75 [0.25]	0.80 [0.30]
23	1.26 [0.27]	1.21 [0.32]	1.26 [0.25]	1.21 [0.31]
Ref=24	*1.00*	*1.00*	*1.00*	*1.00*

*denotes significant difference from zero at 5% level

Table 3.2 – Multivariate analysis estimates. Proportional Hazard Model: non-parametric baseline hazard

Variable	*exp*(estimate) [S.E]			
			Unobserved heterogeneity	
	Men	Women	Men	Women
CAL=1973	0.65 [1.67]	0.10* [0.13]	0.65 [6.23]	-
1974	0.14 [1.40]	0.03* [0.42]	0.14 [5.15]	0.01* [0.41]
1975	0.03*[0.95]	0.02* [0.42]	0.03 [3.30]	0.01* [0.41]
1976	0.03*[0.53]	0.37* [0.26]	0.03 [1.78]	0.01* [0.25]
1977	0.49*[0.12]	0.41* [0.08]	0.49 [0.37]	0.32* [0.06]
1978	0.48*[0.07]	0.41* [0.06]	0.48* [0.14]	0.39* [0.05]
1979	0.52*[0.06]	0.52* [0.06]	0.52* [0.05]	0.41* [0.05]
1980	0.58*[0.07]	0.41* [0.06]	0.58* [0.15]	0.55* [0.05]
1981	0.47*[0.08]	0.55* [0.06]	0.47* [0.21]	0.44* [0.05]
1982	0.67*[0.09]	0.51* [0.05]	0.67 [0.25]	0.61* [0.04]
1983	0.69*[0.06]	0.52* [0.05]	0.69* [0.05]	0.51* [0.05]
1984	0.66*[0.06]	0.51* [0.05]	0.66* [0.05]	0.52* [0.04]
1985	0.69*[0.06]	0.65* [0.05]	0.69* [0.05]	0.52* [0.04]
1986	0.79*[0.06]	0.61* [0.05]	0.79* [0.05]	0.66* [0.04]
1987	0.75*[0.06]	0.54* [0.05]	0.75* [0.12]	0.58* [0.04]
1988	0.64*[0.08]	0.52* [0.06]	0.64* [0.22]	0.50* [0.05]
1989	0.57*[0.07]	0.51* [0.05]	0.57* [0.18]	0.49* [0.04]
1990	0.59*[0.06]	0.50* [0.05]	0.59* [0.09]	0.50* [0.04]
1991	0.60*[0.06]	0.51* [0.05]	0.60* [0.05]	0.50* [0.04]
1992	0.60*[0.06]	0.62* [0.05]	0.60* [0.10]	0.53* [0.04]
1993	0.73*[0.07]	0.61* [0.04]	0.73 [0.17]	0.66* [0.04]
1994	0.75*[0.06]	0.56* [0.05]	0.75* [0.06]	0.62* [0.04]
1995	0.72*[0.05]	0.56* [0.05]	0.72* [0.05]	0.56* [0.04]
Ref=1996	*1.00*	*1.00*	*1.00*	*1.00*

*denotes significant difference from zero at 5% level

Table 3.3 – Multivariate analysis estimates. Proportional Hazard Model: non-parametric baseline hazard

Variable	exp(estimate) [S.E]			
			Unobserved heterogeneity	
	Men	Women	Men	Women
AGE_S	1.01*[0.00]	1.00* [0.00]	1.01* [0.00]	1.00* [0.00]
ISAL	0.89*[0.03]	0.98 [0.01]	0.89* [0.01]	0.98* [0.01]
LOGUN	1.04 [0.03]	1.29* [0.09]	1.04 [0.03]	1.29* [0.07]
LNEMPL	3.23 [0.93]	4.78 [0.00]	3.23 [3.49]	1.20* [0.06]
NSCHOOL	0.74*[0.03]	0.70* [0.03]	0.74* [0.03]	0.70* [0.02]
NPROV	0.98 [0.16]	0.59* [0.17]	0.98 [0.15]	0.59* [0.14]
WLOAD	0.23*[0.03]	0.29* [0.03]	0.23* [0.02]	0.29* [0.02]
PROV=Brabant	1.06 [0.03]	1.06 [0.04]	1.06* [0.03]	1.06 [0.03]
Hainaut	0.90* [0.03]	0.86* [0.04]	0.90* [0.03]	0.86* [0.03]
Liège	1.05 [0.03]	0.96 [0.03]	1.05 [0.03]	0.96 [0.03]
Luxemb	1.10*[0.04]	1.05 [0.04]	1.10* [0.04]	1.05 [0.04]
Ref=Namur	*1.00*	*1.00*	*1.00*	*1.00*
RES=MUN	0.65*[0.02]	0.54* [0.02]	0.65* [0.02]	0.54* [0.02]
CAT	0.73*[0.04]	0.67* [0.03]	0.73* [0.03]	0,67* [0,03]
PROV	0.66*[0.04]	0.57* [0.04]	0.66* [0.03]	0,57* [0,03]
Ref=MIN	*1.00*	*1.00*	*1.00*	*1.00*
Log-Likelihood	*-33,356.84*	*-40,500.13*	*-33,356.836*	*-40,500.13*
$\ln \sigma_v^{\,2}$			0.00 [32,40]	0.00 [28,11]
Likelihood ratio test of rho**=0;Prob > chi2 =			0.9892	0.9862

*denotes significant difference from zero at 5% level

** rho=$\sigma_v^2 / (\sigma_v^2 +1)$; which is the proportion of the total variance contributed by the unobserved heterogeneity component.

Table 4.1 – Multivariate analysis estimates. Proportional Hazard Model: Weibull baseline hazard

| Variable | exp(estimate) [S.E] | | | |
| | | | Unobserved heterogeneity | |
	Men	Women	Men	Women
LNDUR	0.34* [0.01]	0.29* [0.00]	0.34* [0.01]	0.29* [0.01]
CAL=1973	0.62 [1.71]	-	0.62 [5.95]	-
1974	0.13 [1.43]	0.01* [0.00]	0.13 [4.92]	0.01* [0.41]
1975	0.03* [0.97]	0.01* [0.00]	0.03 [3.15]	0.01* [0.41]
1976	0.03* [0.54]	0.01* [0.00]	0.03* [1.70]	0.01* [0.25]
1977	0.43* [0.12]	0.31* [0.00]	0.43* [0.36]	0.31* [0.06]
1978	0.41* [0.07]	0.38* [0.00]	0.41* [0.13]	0.38* [0.05]
1979	0.43* [0.06]	0.39* [0.00]	0.43* [0.05]	0.39* [0.05]
1980	0.47* [0.07]	0.50* [0.00]	0.47* [0.14]	0.50* [0.05]
1981	0.38* [0.08]	0.40* [0.00]	0.38* [0.20]	0.40* [0.05]
1982	0.54* [0.09]	0.54* [0.00]	0.54* [0.24]	0.54* [0.04]
1983	0.55* [0.06]	0.46* [0.00]	0.55* [0.05]	0.46* [0.05]
1984	0.53* [0.06]	0.47* [0.00]	0.53* [0.05]	0.47* [0.04]
1985	0.57* [0.06]	0.47* [0.00]	0.57* [0.05]	0.47* [0.04]
1986	0.67* [0.06]	0.61* [0.00]	0.67* [0.05]	0.61* [0.04]
1987	0.64* [0.07]	0.55* [0.00]	0.64* [0.12]	0.55* [0.04]
1988	0.57* [0.08]	0.48* [0.00]	0.57* [0.21]	0.48* [0.05]
1989	0.52* [0.07]	0.47* [0.00]	0.52* [0.17]	0.47* [0.04]
1990	0.54* [0.06]	0.49* [0.00]	0.54* [0.09]	0.49* [0.04]
1991	0.55* [0.06]	0.49* [0.00]	0.55* [0.05]	0.49* [0.04]
1992	0.56* [0.06]	0.52* [0.00]	0.56* [0.09]	0.52* [0.04]
1993	0.69* [0.07]	0.66* [0.00]	0.69* [0.17]	0.66* [0.04]
1994	0.72* [0.06]	0.61* [0.00]	0.72* [0.06]	0.61* [0.04]
1995	0.70* [0.05]	0.56* [0.00]	0.70* [0.05]	0.56* [0.04]
Ref=1996	*1.00*	*1.00*	*1.00*	*1.00*

*denotes significant difference from zero at 5% level

Table 4.2 – Multivariate analysis estimates. Proportional Hazard Model: Weibull baseline hazard

Variable	*exp*[estimate] [S.E]		Unobserved heterogeneity	
	Men	Women	Men	Women
AGE_S	1.01* [0.00]	0.99* [0.00]	1.01* [0.00]	0.99* [0.00]
ISAL	1.03* [0.01]	1.00 [0.92]	1.03* [0.01]	1.00 [0.01]
LOGUN	1.05 [0.03]	1.29* [0.00]	1.05 [0.03]	1.29* [0.07]
LNEMPL	3.26 [0.95]	1.22* [0.01]	3.26 [3.33]	1.22* [0.06]
NSCHOOL	0.73* [0.03]	0.70* [0.00]	0.73* [0.03]	0.70* [0.02]
NPROV	0.98 [0.16]	0.60* [0.00]	0.98 [0.15]	0.60* [0.14]
WLOAD	0.23* [0.03]	0.29* [0.00]	0.23* [0.02]	0.29* [0.02]
PROV=Brabant	1.05 [0.03]	1.06 [0.14]	1.05 [0.03]	1.06 [0.03]
Hainaut	0.89* [0.03]	0.86* [0.00]	0.89* [0.03]	0.86* [0.03]
Liège	1.05 [0.03]	0.96 [0.20]	1.05 [0.03]	0.96 [0.03]
Luxemb	1.10* [0.04]	1.05 [0.29]	1.10* [0.04]	1.05 [0.04]
Ref=Namur	*1.00*	*1.00*	*1.00*	*1.00*
RES=MUN	0.63* [0.02]	0.53* [0.00]	0.63* [0.02]	0.53* [0.02]
CAT	0.71* [0.04]	0.66* [0.00]	0.71* [0.03]	0.66* [0.03]
PROV	0.63* [0.04]	0.56* [0.00]	0.63* [0.03]	0.56* [0.03]
Ref=MIN	*1.00*	*1.00*	*1.00*	*1.00*
Log-Likelohood	-33,680.675	-40,653.66	-33,680.675	-40,653.659
$\ln \sigma_v^2$			0.00 [34.07]	0.00 [28.77]
Likelihood ratio test of rho**=0;				
Prob > chi2 =			0.9960	0.9894

*denotes significant difference from zero at 5% level

** rho=$\sigma_v^2 \big/ \sigma_v^2 +1$; which is the proportion of the total variance contributed by the unobserved heterogeneity component.

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