

# **Non parametric Estimation of the Determinants of the Unemployment Duration Distribution: Application to the higher schooled qualified**

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**Non parametric Estimation of the Determinants of the  
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**Abstract** An issue in the analysis of unemployment duration concerns distinguishing genuine duration dependence of the exit rate out of unemployment from unobserved heterogeneity. We present a method for the non parametric estimation of both phenomena, designed to be applicable to time-series data on outflows from different duration classes. The method is applied to French data. We find diverging duration effects among university, commerce/ engineers school and IUT / STS school degree qualified. However except for the men outgoing crowd for commerce/ engineers school , duration dependence is dominated by unobserved heterogeneity.

**Keywords:** duration dependence, unobserved heterogeneity, unemployment dynamics, higher education, time-series data

**Classification JEL :** C14, C32, C41, J64

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## INTRODUCTION

In the past decade, the econometric analysis of unemployment duration has become widespread. One of the major issues in this literature concerns the distinction between duration dependence of the hazard rate (or exit rate out of unemployment). And unobserved heterogeneity (for survey, see, e.g., Lancaster 1990; and Devine and Kiefer 1991).

Often, there is reason to believe that for a given individual the hazard rate decrease as a function of duration. For example, there may be stigma effects reducing the number of job opportunities for the long-term unemployed (see, e.g., Vishawanath 1989; and Van den Berg 1994).

However, the presence of unobserved heterogeneity in the distribution of duration variable causes the hazard rate of the distribution of observed duration to decrease as well. This follows from the fact that, on average, individual with the largest hazard rate leave unemployment first. Obviously, from a policy point of view, it is important to know the relative importance of genuine duration dependence (also called state dependence) on the one hand and unobserved heterogeneity on other hand. For example, if duration dependence is the dominant factor, then efforts may be concentrated on the long-term unemployed, while otherwise it may be useful to screen short-term unemployed and concentrate efforts on those with bad characteristics. However, since both factors affect the hazard rate in a similar way, it seems to be hard to distinguish empirically between them.

It is known that in the class of mixed proportional hazard (MPH) models, both the shape of the duration dependence of the hazard and the distribution of unobserved heterogeneity are non parametrically identified (see Elbers and Ridder 1982). However, it is generally believed that in practice it is next to impossible to distinguish between these elements if non strong prior information is present on the shape of the duration dependence or the heterogeneity distribution. In any case, up to now no nonparametric estimation strategy has been developed.

Therefore, in reduced-form empirical analysis of unemployment duration, it has been common to make functional form assumption on (i) the shape of the duration dependence, (ii) the distribution of unobserved heterogeneity, and, (iii) the way the observed explanatory variables enter the model. For example, typical choices are (i) weibull duration dependence, (ii) gamma-distributed unobserved heterogeneity, and, (iii) log-linear dependence of the hazard on observed

explanatory variables (see the surveys mentioned above and the reference therein). Sometimes more flexible forms are chosen, or semi parametric approaches are followed in which only part of the assumption mentioned above are made. In any case, the results are conditional on the particular parametric parts of the specification. Intuitively, it is clear that the results on the degree of duration dependence and unobserved heterogeneity may be extremely sensitive with respect to misspecification of the corresponding parts of the model. As an example, Ridder (1987) proves that estimates may be heavily biased if the form of the duration dependence is misspecified. Since in general the choice of this specification is based on analytical tractability rather than strong prior information, empirical analyses based on such assumption may be unreliable.

In this article, we present a method for the non parametric estimation of the determinants of the unemployment duration distribution. This method is designed to be applicable to discrete-time time-series data on gross outflows from different unemployment duration classes. Gross (or aggregate, or macro) data have the advantage they provide the exact values of the exit probabilities (or exit rates) out of the duration classes considered (averaged over unobserved heterogeneity).

Section II presents the model and the estimation method. Basically, the model is an MPH model in which calendar time replaces the role of observed explanatory ( $x$ ) variables. The estimation method generalizes the method proposed in Van Ours (1992). It enables one to estimate the quantities of interest from ratios of observed hazards, and we extend the model to allow for seasonal effects in the inflow into unemployment..

Section III describes the data in some detail. Section IV contain the results. In some respects, unemployment dynamics in French higher education degree differ greatly between individuals with different sex and the type of the higher education institution (e.g. university, business school and/or college of engineering, or vocational training certificate (BTS) and/or two-year diploma taken at a technical college (DUT).

It would be too restrictive to assume that the duration dependence and calendar-time dependence patterns are the same for all six groups that can be distinguished. The data are disaggregated over these groups, so the empirical analysis is carried out separately for each group. It turns out that duration dependence patterns and the distributions of unobserved heterogeneity differ between groups with different sex/institution characteristics. Section V concludes.

## THE MODEL AND THE ESTIMATION METHOD

### Model Assumption

In this subsection, we present the unemployment duration model and the underlying assumption. An estimation strategy is proposed that enables us to estimate the quantities of interest. Recall that our inferences are purely nonparametric.. That is, we do not parametrize the model, and, strictly speaking, we estimate (summary measures of) functions rather than parameters.

We use two measures of time, each with a different origin. The variable  $t$  denotes the duration of unemployment for a given individual, as measured from the moment the individual becomes unemployed. The variable  $\mathbf{t}$  denotes calendar time, which has its origin somewhere in the past. For simplicity we take  $t$  and  $\mathbf{t}$  to have the same measurement scale (apart from the difference in origin). Both  $t$  and  $\mathbf{t}$  are discrete variables. As an example, consider an individual who is unemployed in period  $t$ , he will be unemployed for  $t + 1$  periods at calendar time  $\mathbf{t} + 1$ .

We denote the probability that an individual leaves unemployment right after  $t$  periods of unemployment, given that he is unemployed for  $t$  periods at calendar time  $\mathbf{t}$ , and conditional on his unobserved characteristics  $\mathbf{n}$ , by  $\mathbf{q}(t / \mathbf{t}, \mathbf{n})$ .

The unemployment duration density conditional on calendar time and conditional on  $\mathbf{n}$  can be constructed from the individual exit probabilities. For example, the probability that unemployment duration equals  $t$ , when calendar time was  $\mathbf{t} - t$  at the moment of inflow into unemployment, conditional on  $\mathbf{n}$ , equals

$$P(T = t / \mathbf{t} - t; \mathbf{n}) = \mathbf{q}(t / \mathbf{t}, \mathbf{n}) \cdot \prod_{i=1}^t [1 - \mathbf{q}(t - i / \mathbf{t} - i, \mathbf{n})] \quad (1)$$

for all  $t \in \{0, 1, \Delta\}$ . We take the product term to be one if  $t = 0$ .

*ASSUMPTION 1.* MPH :  $\mathbf{q}(t / \mathbf{t}, \mathbf{n})$  has a mixed proportional hazard specification ( Hahn,- Jinyong, 1994); that is, there are functions  $\mathbf{y}_1$  and  $\mathbf{y}_2$  such that

$$\mathbf{q}(t / \mathbf{t}, \mathbf{n}) = \mathbf{y}_1(t) \cdot \mathbf{y}_2(\mathbf{t}) \mathbf{n} \quad (2)$$

with  $y_1$  and  $y_2$  positive and uniformly bounded from above. Further, the distribution of  $\mathbf{n}$  is such that, for every  $t$  and  $\mathbf{t}$ ,

$$P(0 < \mathbf{q}(t/\mathbf{t}, \mathbf{n}) < 1) = 1$$

*ASSUMPTION 2.* Independence of  $\mathbf{n}$  and  $\mathbf{t}$ : the distribution of  $\mathbf{n}$  in the inflow into unemployment does not depend on the moment of inflow. Further, the individual level of  $\mathbf{n}$  does not change during unemployment.

*ASSUMPTION 3.* Variation over calendar time: the function  $y_2$  is not constant.

The functions  $y_1$  and  $y_2$  represent the duration dependence and the calendar time dependence of the individual exit probabilities out of unemployment. As we will see, assumptions 1-3 ensure non parametric identification of the model. In particular, they ensure that duration dependence and unobserved heterogeneity can be distinguished empirically.

Assumption 1 is reminiscent of the standard MPH assumption in reduced-form models for micro duration data (see Lancaster (19890) for an extensive survey of such models). In models for micro duration data, dependence on calendar time is usually ignored, and the role of  $\mathbf{t}$  in the model above is replaced the role of observed explanatory variables  $x$ .

### **Observed exit probabilities**

Let  $T$  denotes the random unemployment duration and  $t$  its realization. In obvious section, it holds that

$$\mathbf{q}(t/\mathbf{t}) = \frac{P(T = t / \text{inf low at } \mathbf{t} - t)}{P(T \geq t / \text{inf low at } \mathbf{t} - t)} \equiv \frac{E_{\mathbf{n}}(\Pr(T = t / \text{inf low at } \mathbf{t} - t; \mathbf{n}))}{E_{\mathbf{n}}(\Pr(T \geq t / \text{inf low at } \mathbf{t} - t; \mathbf{n}))} \quad (3)$$

in which  $\Pr(T = t / \text{inf low at } \mathbf{t} - t; \mathbf{n})$  and  $\Pr(T \geq t / \text{inf low at } \mathbf{t} - t; \mathbf{n})$  can be expressed in terms of  $\mathbf{q}(t/\mathbf{t}; \mathbf{n})$ . (Note that eq. (1) gives  $\Pr(T = t / \text{inf low at } \mathbf{t} - t; \mathbf{n})$ ).

By doing this, and by substituting equation (2), we get

$$\mathbf{q}(t/t) = \frac{\mathbf{y}_1(t)\mathbf{y}_2(t) \cdot E_n \left[ \mathbf{n} \cdot \prod_{i=1}^t \left[ 1 - \prod_{i=1}^i [1 - \mathbf{y}_1(t-i)\mathbf{y}_2(t-i)\mathbf{n}] \right] \right]}{E_n \left[ \prod_{i=1}^t [1 - \mathbf{y}_1(t-i)\mathbf{y}_2(t-i)\mathbf{n}] \right]} \quad (4)$$

Denote  $E_n(\mathbf{n}^i)$  by  $\mathbf{m}_i$ . We have the following result:  $\mathbf{q}(t/t)$  depends on  $\{\mathbf{y}_1(i), \mathbf{y}_2(\mathbf{t}-t+i), \mathbf{m}_{t+1} \text{ with } i = 0, 1, \dots, t\}$ .

However, note that we are primarily interested in estimating the duration dependence and unobserved heterogeneity parameters. For this, the calendar time dependence parameters are nuisance parameters. Thus it would be nice if the parameters of interest could be estimated without the need to deal with the calendar time dependence parameters.

It turns out that the ideas in Van Ours (1992) can be extended and applied to achieve this aim. Basically, these ideas amount to substituting values of past observed exit probabilities into the expressions (4) for  $\mathbf{q}(t/t)$  and examining ratios of the resulting expressions for different  $t$ . We illustrate this in the first appendix.

## THE DATA SET

We use data resulting from the investigation of the center from studies and search into the qualifications (Céreq) carried out in march 1997 to the outgoing of higher education in 1994. The data describes the various situations of the applicants for work month by month of the calendar for the period active of september 1994 until march 1997.

However, we may consider the data as synthetic cohorts, even more so because the sample size is much larger than typically encountered in micro data sets. Still, when we do so, the data generate a small number of inconsistent ( $< 0$ , or very small) monthly exit probabilities. In some cases this led to very large ratios  $\mathbf{q}(t/t)/\mathbf{q}(t-1/t)$ . We therefore skipped observations for which  $\mathbf{q}(t/t)$  was smaller than 0.05. This restriction is arbitrary, but the use of similar restrictions with different boundaries did not lead to substantially different results.

In the analysis we use a time series of monthly exit probabilities out the unemployment for six groups of workers: commerce/engineers males workers, commerce/engineers females workers, university males workers, university females workers, IUT/STS males workers, IUT/STS females workers.

It is clear that the six groups that are distinguished have quite different unemployment dynamics characteristics. The model of section II predicts that there is unobserved heterogeneity in the unemployment duration distribution if ratios of observed exit probabilities for different duration classes change over calendar time.

## THE RESULTS

### **Parameters estimates**

The estimation results are shown in table 1 (see appendix 2). We start by discussing the  $h_t$  estimates. These indicate that for university males workers there is negative duration dependence of the exit probability out of unemployment. For university males, the exit probability in the fourth month is 83 % of the exit probability in the second month, and the exit probability in the fourth month is 79 % of the exit probability in the third month. Thos may be due to a stigma effects of not being short-term unemployed.

For both male and female commerce/engineers workers,  $h_1$  is significantly larger than one, indicating that there is significant positive duration dependence in the second month of unemployment in comparison to the first.

The less negative duration dependence for IUT/STS workers (relative to university workers) during the first few months can be “explained” in a numbers of ways (relatively strong anticipation of unemployment benefits exhaustion; relative importance of particular recall options; relatively large nonpecuniary utility of being short-term unemployed; increase transitions from unemployment to nonparticipation; etc.). However, in the absence of additional information, it is hard to assess the power of such possible explanations.

One potentially interesting explanation for the male difference takes into account that the initial level of the exit probability out of the unemployment (i.e.,  $q(0/t)$ ) for university males is on average about 76 % of the corresponding level for commerce/engineers males. It may well be

that university males carry a stigma from the moment they become unemployed. Now suppose that, in addition to that, in each group, individuals are heavily stigmatized from the moment their spell length is observed by potential employers to belong to the 20 % or so longest spells in those groups. Then, the duration at which IUT/STS males start to get a stigma is shorter than the corresponding duration for university males. This may explain the fact that there is more genuine negative duration dependence for university males than for IUT/STS males, before the fourth month of unemployment duration. If this explanation is correct, than one should expect duration dependence for IUT/STS males to become more negative after 4 months of unemployment duration.

Note that in the microeconomic literature on unemployment duration it is always assumed that the duration dependence parameters do not depend on individual characteristics like gender.

Equilibrium models of stigma (Berkovitch, 1990) seem to predict that the stigmatization of long-term unemployed individuals is stronger if there is more unobserved heterogeneity in the quality of potential employer-employee matches. Now suppose that that a large variance of this type of heterogeneity is associated with a large variance of the unobserved heterogeneity in unemployment durations. Then one may expect that, over the six groups we distinguish, a large variance of the unobserved heterogeneity distribution is associated with relatively strong negative duration dependence. The results discussed below do not empirically confirm this. Presumably, other differences between these groups are more important.

The fact that the  $g_2$  estimates significantly exceed one means that the data confirm that observed duration dependence (when going from  $t=0$  to  $t=1$ ) is more negative in the top of the cycle than in a recession. Blanchard and Diamond (1994) so-called ranking model of unemployment predicts the opposite result. Apparently, in these data, the dynamic selection due to unobserved heterogeneity is empirically more important than ranking phenomena.

In general, the interaction between  $t$  and  $\mathbf{t}$  in the observed exit probabilities that is caused by unobserved heterogeneity is such that the observed degree of duration dependence is less negative in a recession than at the top of the business cycle. For example, eq.[7] (see appendix 1) implies that, if  $g_2 > 1$  and  $h_1 < 1$ , the decrease of  $q(t/\mathbf{t})$  when going from  $t=0$  to  $t=1$  is smaller in a recession ( $q(0/\mathbf{t} - 1)$  small) than at the top of the cycle ( $q(0/\mathbf{t} - 1)$  large). This is because in a recession the weeding out of individuals with a high quality (i.e., a large  $n$ ) cannot occur as

fast as in the other case. Note that the model implication on the interaction sign is testable. In particular, if  $g_2 < 1$ . Then the interaction sign in (7) is opposite to above.

We finish the discussion of our results by comparing them to the results on duration dependence and unobserved heterogeneity in the literature on parametric empirical analysis of unemployment durations within the MPH framework. Butler and McDonald (1986) estimate these phenomena in a parametric setting using CSP data. They take a weibull specification for  $y_1(t)$  and assume that  $G(\mathbf{n})$  is generalized gamma distribution, and they find evidence for the presence of unobserved heterogeneity and positive duration dependence (so  $y_1(t)$  is increasing). However their model does not account for dependence of the unemployment duration hazard on individual characteristics  $x$  on calendar time  $t$ . Consequently, the model is non parametrically unidentified, and the results on duration dependence by changing the exit probabilities for all unemployed individuals simultaneously.

The results in Flinn and Heckman (1983) are not statistically significant. Heckman and Singer (1984) find evidence for the presence of unobserved heterogeneity. They take a weibull specification for  $y_1(t)$  and find that that the results on the sign of the duration dependence are very sensitive to the assumed family of distribution of  $G(\mathbf{n})$ .

Meyer (1990) uses a flexible functional form for  $y_1(t)$  and assumes  $G(\mathbf{n})$  belongs to the gamma family. He find evidence for unobserved heterogeneity. Further, in general  $y_1(t)$  does not display spikes near durations at which benefits entitlement ends, but these durations are well beyond the 4-months period we examine in our analysis.

## CONCLUSION

In this article, we show both theoretically and empirically that it is possible to distinguish unobserved heterogeneity from genuine duration dependence in unemployment durations by using aggregate time series on exit probabilities out of unemployment.

We analyse French unemployment data for the period 09-1994|03-1997 distinguishing six groups of workers with different sex and the type of institution. We find that unobserved heterogeneity is relevant for all six groups, causing the observed exit probability out of unemployment to decline over the duration of unemployment. Furthermore, we find diverging

duration dependence effects between university, IUT/STS, commerce/engineers workers. For university males, the genuine duration dependence is most negative. The effect for university female individual is smaller, but significant. For IUT/STS, we do not find significant negative duration dependence. From this, we conclude that, in the French labour market segment, duration dependence stigma effects related to unemployment durations are dominant for commerce/engineers workers, but not for IUT/STS workers. Except for commerce/engineers males, though, the effect of unobserved heterogeneity dominates the duration dependence.

Several topics for future research emerge. First, it seems worthwhile to combine the aggregate data with micro data containing information on explanatory variables. This might make it possible to estimate the quantities of interest under weaker assumptions. Another topic for further research would be to improve the foundation of the stochastic specification of the equations of the empirical model. However, it can be shown that incorporating this would lead to a model (i) with a complicate error covariance structure and (ii) that cannot be estimated with the method of this article.

## Appendix 1: A preliminary application

To explain the gist of our model, we present a preliminary application using exit probabilities for the first 2 months. Consider an expression (4) for  $t = 0$  and  $t = 1$  :

$$\mathbf{q}(0/t) = \mathbf{y}_1(0) \cdot \mathbf{y}_2(t) \mathbf{m}_1 \text{ et } \mathbf{q}(1/t) = \frac{\mathbf{y}_1(1) \mathbf{y}_2(t) [\mathbf{m}_1 - \mathbf{y}_1(0) \mathbf{y}_2(t-1) \mathbf{m}_2]}{1 - \mathbf{y}_1(0) \mathbf{y}_2(t-1) \mathbf{m}_1} \quad (5)$$

and

$$\mathbf{q}(1/t) = \mathbf{y}_1(1) \mathbf{y}_2(t) \mathbf{m}_1 \frac{1 - \left( \frac{\mathbf{m}_2}{\mathbf{m}_1} \right) \mathbf{q}(0/t-1)}{1 - \mathbf{q}(0/t-1)} \quad (6)$$

Substituting a 1-period lagged version of (4) into (5), we obtain

$$\mathbf{q}(1/t) = \mathbf{y}_1(1) \mathbf{y}_2(t) \mathbf{m}_1 \frac{1 - \left( \frac{\mathbf{m}_2}{\mathbf{m}_1^2} \right) \mathbf{q}(0/t-1)}{1 - \mathbf{q}(0/t-1)}$$

We can eliminate  $\mathbf{y}_2(t)$  by taking the ratio of (6) and (5). Denoting  $\frac{\mathbf{m}_2}{\mathbf{m}_1^2}$  by  $\mathbf{g}_2$  and  $\frac{\mathbf{y}_1(1)}{\mathbf{y}_1(0)}$  by  $\mathbf{h}_1$ , we get

$$\frac{\mathbf{q}(1/t)}{\mathbf{q}(0/t)} = \mathbf{h}_1 \cdot \frac{1 - \mathbf{g}_2 \mathbf{q}(0/t-1)}{1 - \mathbf{q}(0/t-1)} \quad (7)$$

It holds that  $\frac{\mathbf{q}(1/t)}{\mathbf{q}(0/t)}$  depends on  $\mathbf{q}(0/t-1)$  if and only if  $\mathbf{g}_2 \neq 1$ , which in turn holds if and only if there is unobserved heterogeneity (note that  $\mathbf{g}_2 = 1 + \text{var}(\mathbf{n})/\mathbf{m}_1^2$  and that always  $\mathbf{g}_2 \geq 1$ ). The exit probability  $\mathbf{q}(0/t-1)$  varies with  $t$  by virtue of assumption 3, so in sum,  $\frac{\mathbf{q}(1/t)}{\mathbf{q}(0/t)}$  varies with  $t$  if and only if there is unobserved heterogeneity. Clearly  $\mathbf{h}_1$  and  $\mathbf{g}_2$  are identified.

In this preliminary empirical analysis we adopt the following regression-type stochastic specification:  $\text{Log} \frac{\mathbf{q}(1/t)}{\mathbf{q}(0/t)}$  equals the log of the r.h.s. of (7), plus an error term. This error term is

assumed to represent an independently and identically distributed (i.i.d.) specification error. We then estimate the regression using nonlinear regression « *Seemingly Unrelated Nonlinear Regression* » (SUNR) (Jensen,-Mark-J, 1995 et Smith, M., R. Kohn, 2000).

First, if the  $g_2$  is significantly larger than one, which implies  $\text{var}(\mathbf{n}) > 0$ , so unobserved heterogeneity is present. Second, if the duration dependence parameter  $h_1$  is significantly smaller than one, this implies that the average conditional exit probability decreases from the 1<sup>st</sup> to the 2<sup>nd</sup> month of unemployment, so there is negative duration dependence in the first two months.

## Appendix 2: The duration dependence and heterogeneity for the period 09-1994\03-1997

	University workers		Commerce/Engineers workers		IUT/STS	
	Male	Female	Male	Female	Male	Female
<b>Unobserved heterogeneity :</b>						
$g_2$	1.28 (0.02)	1.18 (0.01)	1.10 (0.01)	1.09 (0.01)	1.24 (0.01)	1.20 (0.01)
$g_3$	1.78 (0.06)	1.67 (0.06)	1.29 (0.04)	1.59 (0.05)	1.75 (0.05)	1.62 (0.05)
$g_4$	2.93 (0.15)	2.12 (0.13)	1.83 (0.11)	1.81 (0.11)	2.20 (0.13)	2.05 (0.12)
<b>Duration dependence</b>						
$h_1$	0.97 (0.04)	0.95 (0.03)	1.18 (0.03)	1.14 (0.02)	1.14 (0.04)	1.09 (0.03)
$h_2$	0.83 (0.05)	0.97 (0.04)	1.12 (0.05)	1.10 (0.04)	1.08 (0.05)	1.12 (0.05)
$h_3$	0.79 (0.05)	0.82 (0.06)	1.08 (0.06)	1.03 (0.05)	0.98 (0.05)	1.01 (0.06)
<b>SUNR residual covariance matrix</b> $\Sigma \equiv (s_{ij})$						
$s_{11}$	0.17	0.08	0.24	0.19	0.17	0.07
$s_{12}$	-0.09	-0.07	-0.15	-0.11	-0.92	-0.63
$s_{13}$	-0.03	-0.01	-0.01	-0.00	-0.22	-0.00
$s_{22}$	0.27	0.26	0.10	0.08	0.28	0.27
$s_{23}$	-0.18	-0.20	-0.06	-0.04	-0.16	-0.2
$s_{33}$	0.45	0.54	0.28	0.22	0.41	0.64

NOTE.-Standard errors are in parentheses.

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