The Impact of Economic Variables on Timing of Births and Parity Progression Ratio. A Question Re-examined for a Panel of French Women

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1. Introduction

The question of the impact of women's wages on the timing of births and parity progression is attracting renewed interest. With increasing unemployment spells for men, the traditionally dominant substitution effect, which caused fertility to decline as women's wages increase, since a child becomes much more costly, may be cancelled out for two reasons:

- the income effect: a high wage also has the consequence of relaxing the budget constraints which prevent the number of children per family from increasing. This positive relationship has been found in Europe, for wages (Ekert, 1986), and university degrees (which are correlated with wages) (Hoem and Hoem, 1989; Kravdal 1992, 2001) Similarly, we have found that a job is a prerequisite for partnership and first child. (Ekert-Jaffe- Solaz, 2002;).
- family-friendly policy can partly offset the negative relationship by helping to reconcile childbearing and career: for instance, using longitudinal data from linked census and vital statistics (data1), Ekert-Jaffe and al (2002) compare the progression and timing of births between France and Britain. They point out the role of French family policy that allows more prevalent family formation in upper-level occupations and offsets their traditional low fertility.

Recently these findings have been discussed and attributed to the model's mechanical timing effect and selection bias rather than actual behaviour. When time since previous birth is estimated separately for second and third births, the model fails to recognise (i) that the shorter interval before second conception for upper occupations is due to higher age at first birth. (ii) that it may be due to unobserved heterogeneity and selection bias: with higher child cost, women with higher potential wage who give birth to a first child have a higher preference for children and greater transition to subsequent births. When all three parity transitions are modelled jointly, with a common unobserved factor included, negative effects of educational level appear. This was first shown in Norway (Kravdal 2001). We find the same selection effect in France using data1 (Mougin 2004): when the selection effect linked to possible unobserved characteristics among the most highly qualified women is taken into account, the positive effect of higher degrees on the likelihood of bearing a second child diminishes, and disappears for the third birth. Consequently, there are indeed unobserved personal dimensions — such as preference for children — that contribute to the decision to have a second or third child. Similarly the impact of wages on the timing of births must be re-examined with individual panel data to take account of unobserved heterogeneity.

Besides, some models do successfully predict family formation with a purely demographic approach in which. the timing and occurrence of birth order are explained by previous demographic behaviour (Rodriguez and al.). The question is: Are wages really important?

This paper seeks to assess the actual impact of women's and men's wages on the timing of births, and to test the neo-classical model against a purely demographic model.

We investigate these relations in detail, using a multi-state dynamic duration model for flexible modelling of the timing of births. We implement the model using a long-term longitudinal individual dataset (Birth registration data were linked to three census extracts carried out from 1975 to 1991 and a microwage dataset) and we used instrumental microwages. The model includes three sets of explanatory factors. The first are economic variables such as wages and/or maternity leave. The second are demographic variables such as age at marriage and/or at first birth and the intervals between previous births. The third set represents unobserved exogenous variables specific to individuals, where values may vary over time. Following Heckman and Walker (1987, 1990) in their study of Swedish women, we take unobserved heterogeneity into account in both ways: we introduce an heterogeneity term that is specific to women, and a mover-stayer structure specific to parity j, to take account of women who are no longer likely to give birth. Our best fit model is slightly different from the one these authors found for Sweden, showing that French women's behaviour is more heterogeneous than Swedish women's. In accordance with Heckman and Walker's results, for our sample of French women from 3 birth cohorts (1946-50, 1952-56 and 1958-62), followed from 1975 to 1995, we show that the economic – versus demographic – model produces the better results according to the tests. There is a strong robust negative impact of women's wages on the timing of births, whatever the cohort. This impact increases with parity. The steady increase in women's wages over more than thirty years explains the decrease and delay in fertility. Nevertheless generous maternity leave partly offsets this negative impact.

The influence of men's wages is less clear. Their effect is statistically significant for the first birth, but often absent for second and third births, notably for the earlier birth cohorts. In particular, when non-observable factors and the intervals between births are taken into account, their influence is considerably reduced. Moreover, when the observation start date is defined as the date of marriage, the effect of men's wages becomes insignificant for first births and can thus be interpreted as a "partner" or "spouse" effect.

The plan of the paper is as follows: Section 2 briefly recapitulates the economic background of the possible influence of wages on births; Section 3 presents the model and the data used in his study. In Section 4, we select the best fit model and present the results. Section 5 discusses our results compared with previous studies.

2. Wages and fertility according to economic theory

Economic theory suggests three mechanisms to explain the effect of income on fertility (Becker, 1981; Cigno and Ermisch, 1988).

1) Time cost or budget constraint?

When women's wages increase, the price of time increases correspondingly. While a woman is looking after her child, she is not able to use that time to earn money in the labour market, and the opportunity cost is a function of her wage (the time cost of children being borne principally by women). By a *substitution effect*, then, the greater the cost of the child, the larger the decline of fertility. This effect is therefore stronger for women in higher socio-occupational groups. Moreover, career women may choose to limit their fertility in order to devote themselves more fully to their working life. In helping women to combine family life and working life, family policy can tend to minimise these two effects. The state (partially) compensates women for the loss of time and minimises the reduction of fertility induced by women's employment for all women, and especially for those in managerial positions (Ermisch, 1989).

But a high wage also has the effect of relaxing the budget constraints that limit the number of children in a family: this is the *income effect*. For women, this effect is generally offset by the substitution effect; for men, the income effect dominates and fertility rises with the level reached on the social ladder. By helping to reduce the monetary cost of children by a greater proportion for the lower and middle classes, French family policy diminishes fertility differences linked to men's income.

2) The “quantity” versus “quality” effect

When social class rises, couples seek to extend the schooling of their children in order to obtain for them of the best possible social position, resulting in a limitation of the number of children (Becker and Lewis, 1973). This explains the lower fertility of the middle class compared with the working class. At the same time, the upper class has sufficient resources not to limit the number of children.

These three effects produce the classical inverted J shape of fertility according to social class.

Prediction of the effect of a rise in wages is less clear. Theory would predict a negative influence of women’s wages, which might decline at the highest wage levels, and could be reduced by family policy. Men’s wages will have a positive effect on the first birth, which might be moderated for later births if men decide rather to devote their disposable income to their child or children’s education.

3. Data and methods

3.1. Data

Our data are drawn from a combination of census records, birth records (EDP), and annual company data declarations (DADS) from 1976 to 1996. The permanent demographic sample (*Échantillon Démographique Permanent*, EDP) is a longitudinal panel linking data from the census records of 1968, 1975, 1982, and 1990 to data from civil registration records collected between 1968 and 1995, for such major demographic events as births, marriages, deaths and recognitions of paternity. The date of marital separations is not known precisely. For greater accuracy, we decided to take only women born in France, present at every census, married and still cohabiting in 1990, all of whose children were born of the same partner. The selection bias due to the decline in marriages ending in divorce is only slight since it has been shown that divorced couples do not exhibit different fertility behaviour before the divorce, except perhaps in the last two years (Festy 1994, Toulemon 1994).

The annual company data declarations (Déclarations Annuelles des Données Sociales, DADS) are administrative declarations filed by employers of the gross income of each of their employees who is subject to income tax. Statistically it covers all employees in the private sector and major state-owned enterprises aged from 18 to 65, who have worked full- or part-time at least once in the year. We then restricted our sample to individuals whose professional careers were in the private sector or major state-owned enterprises, or who never worked.

3.2. Model

The general form of duration dependence and unobserved heterogeneity.

In duration models, unobserved heterogeneity and duration dependence play an identical or concurrent role, which may lead to many inference errors if heterogeneity is omitted. However, it is difficult to distinguish heterogeneity effects from those of duration dependence (Hoem 1990). Furthermore, the parametric approach adopted in many studies presupposes defining a priori, often with no information, a functional form for the mixing distribution of the unobserved heterogeneity. The work of Heckman and Singer (1982, 1984b) shows that this hypothesis as to the functional mixing form is not essential if one adopts a non-parametric representation of unobserved heterogeneity. They propose replacing the unobserved heterogeneity distribution by discrete approximation, and show that this method gives results that are often more accurate when estimating the various models used.

The birth process is defined as a continuous multi-state process, where the state of possible attained birth states (parity $j=1,2,3,..$) is finite. The first birth can occur after $\tau = 0$; and every birth constitutes a transition that occurs after τ =0. We need to estimate conditional hazard. We define $H(\tau)$ as the relevant conditioning set at time τ: the history of the women's birth process at time τ and anticipations about the future formed at time τ may be a part of $H(\tau)$. The birth process is characterised by T1,T2 T3 conditional on H, and each random variable can be described by a conditional density that integrates to a conditional distribution function. If a women has her j-1 birth at time τ(j-1) and therefore becomes at risk for the jth birth at this time, we have to estimate the conditional hazard of having a birth of parity j. at duration t_{j},; it depends on the conditioning history H at time of birth τ(j-1)+ t_j We approximate this jth conditional hazard by using the following functional form:

(Eq.1) $$h_j(t_j | \mathrm{H}(\tau(j-1)+t_j), \theta) = \exp\{ \gamma_{oj} + Z(\tau(j-1)+t_j)\beta_j + \sum_{k=1}^{K_j} (\frac{t_j^{\lambda_{kj}} - 1}{\lambda_{kj}})\gamma_{kj} + \varepsilon_j \theta \}$$

where γ_{oj}, γ_{kj}, β_j, and ε_j are parameters to estimate, $Z(\tau(j-1)+t_j)$ includes all observed regressors exogenous or endogenous to the birth process — such as wages — depending on time or not, possibly including durations from previous spells and spline functions of current duration.. The rest of the right member represents the baseline duration dependence conditioning on unobserved heterogeneity $\varepsilon_{\{j\}}\theta$. θ can be correlated over spells and ε_j enables the same θ to have a varied impact according to parity. Distribution of θ does not depend on the parameters of the model or the process.

Hazard specification encompasses different commonly used base lines. Setting β=0, $\varepsilon_{j=0}$,, the conditions K_j =1and λ_{1j} =0 specialise to a Weibull model, K_j =1and λ_{1j} =1 to a Gumpertz, and K_j =2, λ_{1j} =1 and λ_{2j} =2 to a quadratic model. etc; As many of these models are nested, they can be compared with likelihood ratio. Nevertheless, some models – such as quadratic model versus Weibull model — are non-nested and we will use other tests to test competing specifications. Then, setting $\beta\neq0$ et $\varepsilon_j \neq0$, we allow for general time-varying covariates and unobserved heterogeneity component as in Heckman et Walker (1990a, b).

Since θ is unobserved, we have to estimate its distribution $M(\theta)$. We estimate it by the non parametric maximum likelihood (NPMLE) procedure described in Heckman and Walker (1987, 1990) and in Heckman and Singer (1984). This procedure approximates any distribution of unobservable function by a finite mixture of distributions that maximise the sample likelihood. The cumulative function $M(\theta)$ is then approximated by a piecewise linear function, in I pieces, θ being estimated as a discrete variable that converges to the true distribution when the sample size increases. We estimate (pi,θi, 0<i<I), where pi, the probability that θ= θi, is a weight put on the point-mass θi, the θi being ordered from lowest to highest. I is estimated with other parameters. It is as if the population were divided into I homogeneous groups whose ε_j θ= θi(j). The distribution of the population among the θi can be independent, or correlated across spells.

The structure of the unobservable is as follows. The first hypothesis is that the unobservable determinants can be decomposed in the sum of a parity-specific unobservable term and a purely women-specific unobservable perfectly correlated across spells.

$\theta_j = \varepsilon^*_j \theta^* + \nu_j$,

θ^* here is the individual-specific unobserved heterogeneity, namely the unobservables non-exogenous to the birth process, and ν_j is the parity-specific unobserved heterogeneity (unobservables randomly distributed between parity and individuals). Vector $(\nu_1,\ldots,\nu_n\},\theta^*)$

defines mutually independent random variables ν_n of density $k(\nu_j)$ and θ of density $m(\theta)$. Then $\varepsilon_j \theta$ is the result of the integration of θ over ν_j.

Under the second hypothesis, the density of ν_j is estimated by allocating a point-mass at $\nu_j = -\infty$ that corresponds to a point mass $\theta = -\infty$ in the estimation of equation 1. This value of θ sets the hazard of leaving parity j-1, defined in equation 1, to zero. So the survivor function utilised in this empirical work is based on hazard (equ1) augmented to allow for parity-specific stopping behaviour. By dividing the definition domain of variable ν_j in two, such that $\nu_j = -\infty$ and $\nu_j = (-\infty, \infty]$, a mover-stayer structure specific to parity j is introduced into the model with latent variable ν_j (Goodman 1961). By integrating over ν_j, the survivor function for the j birth is:

$$Sj(tj \mid H, \varepsilon_j \theta) = P^{(j-1)} + (1 - P^{(j-1)})\, S(tj \mid H, \varepsilon_j \theta)$$

where $P^{(j-1)}$ is the likelihood that a women with j-1 children will never be exposed to the hazard of a jth birth ($\nu_{j-1} = -\infty$) in the population).

In this way it is possible to model a mover-stayer structure as a particular case of Heckman and Singer's non-parametric maximum likelihood estimator.

3.3. Instrumental wages

Women with low wages may more often have a third child because of the many career interruptions they have had by placing their family life ahead of their professional careers. The correlation between women's microwages and their fertility gives no indication of the direction of causality: whether it is wages that affect fertility or fertility that depresses wages. Individual preference for children can influence both past career and past fertility; using observed wages leads to spurious regressions that appear to express the impact of fertility on wages together with the weaker impact of wages on births.

This is one justification for Heckman and Walker's use (1990) of aggregate wages, which are less sensitive to women's individual careers. However, they have been criticised for this option (Tasiran, 1995): when using aggregate wage measures, women are assumed to respond to an aggregate wage when presumably they respond to their own wage profile. By using microwage data linked with fertility history-like Tasiran-, we can avoid this drawback; but we need to purge individual data of any endogenous variation due to either simultaneity or selection: we must isolate and use only the exogenous variation of wages.

Any unbiased estimate of the effect of these variables consequently requires considerable care in order to avoid the endogeneity bias in the measurement of the effects of wages on fertility, and the heterogeneity bias.

Here, instrumental variables were used to determine microwages uncorrelated with the unobserved determinants of past decisions concerning time allocation, career and fertility (Mougin and Ekert-Jaffé, 2002). The specification of these instrumental wage equations, based on the human capital earnings functions introduced by Mincer (1963), presupposes the use of measurements of the current labour supply, such as actual professional experience — measurements that are themselves exogenous to the wages received in the past or past labour supply decisions. Furthermore, these wage equations were corrected for the endogenous selection biases in the sample used for the estimates (Heckman 1979, see also Verbeek and Nijman 1992a, Powell 1994 and Vella 1998 for a comprehensive review of the literature in this field). The reason is that when earnings functions are adjusted, individual characteristics relating to individuals' decision to enter the labour market may also influence the process determining their salary. When these characteristics cannot be observed and correlate with the explanatory variables of the earnings function, use of usual methods of estimation leads to biases in the estimates made.

The task, therefore, was to allow for the effects of correlation between the explanatory variables of the wage equation, such as education and professional experience, and the unobserved characteristics of individuals, such as their aptitudes and personal preferences, which are also determining factors for wages (Card 1999). From this point of view, some correlation between the explanatory variables of the function and the individual component of the random disturbance in the model does not make it possible to apply usual methods of estimation. In this case, the use of methods designed for panel data overcomes this difficulty (see Chamberlain 1985, Hsiao 1986, Baltagi 1995 and Arellano and Honoré 2000 for details of all these methods).

In practice, the variables used in estimating women's earnings functions are age on completing education, professional experience and professional experience squared (instrumental), seniority in job and time indicators. Characteristics of the post occupied, such as sector of activity or size of company, are not considered in order to avoid assimilating the situation of women outside the sample with those of the reference categories. We also used these variables in new estimates. The results obtained show that women's wage coefficient vary little whatever the measurements used. For men, we estimated a wage equation as a function of age on completing education, potential experience in the labour market, experience squared, spouse's educational qualifications, and sector of activity. We also added an individual permanent wage term and a temporary residue whose parameters were calculated from DADS data.

These wage estimates were corrected by the reduced selection equation with a lagged dependent variable to allow for dynamic factors attributed to working behaviour. The estimate method is proposed by Vella and Verbeek (1999).

We also use current values of women and men's wages and do not include expected future wages despite the importance of such variables in life-cycle theory. In models in which agents have stationary expectations, are short-sighted, or in which wages are uncertain and are first order Markov, current wages are sufficient statistics for future wages (Heckman and McCurdy 1980). Furthermore, current wages at the time of a child's birth may be a good predictor of the wages expected by the couple when they decided to have the child, and our results support this hypothesis since the models estimated with wages measured at the time of conception are dominant on all criteria and not goodness-of-fit tests. The third reason for excluding future wages is a practical one. Future and current wages are highly correlated and models that include current and future values are numerically unstable and highly correlated (Heckman and Walker 1990 a, b, 1992).

Heckman and Walker's CTM package is used to estimate the models**.**

3.4 Variables

To analyse the duration from the age of 16 to the first birth, and the spacing of the two subsequent births, the dependent variables are the first three transition rates.

The main variables taken into account in the main neo-classical model are women's and spouse's estimated wages and a dummy indicating whether the women had grown up in a rural zone till the age of 16. The demographic model uses the past timing and spacing of births and the rural dummy. Following Heckman and Walker, we first selected among these two models and we will present the best fitting model.

We also tested other specifications of the best model; we added variables as regressors and analysed how they modify the impact of the wages. We controlled successively for age effect, time period effect, previous birth spacings, lagged wages - measured at the time of conception, the effect of interacting male wages with a dummy indicating women's labour force participation status as suggested by Willis, a term interacting between men's wages and women's working, and a model using the date of marriage as the origin of duration.

3.5. Tests

To select among alternative models (some of the alternatives being unnested), we use two kinds of tests.

(i) Chi-squared tests for predicting attained fertility by ages 20, 25, 30 and 35 for each cohort, for each spell, then across spells for each age and ten across ages: because tests within a cohort are not independent, a Bonferroni test is used to evaluate the joint hypothesis that predicted parity distributions fit at each of the selected ages. This test is based on the maximum chi-squared statistic over all age groups. It is performed for each cohort.

(ii) the Akaike information criterion, and the Leamer–Schwarz–test (Bayesian Information Criterion, BIC) are based on maximum likelihood metric corrected in order to penalise complexity (Table 1).

4. Results

4.1. Best fitting model

Heckman and Walker tested 148 models on Swedish data, differing in their specification of duration dependence, including models with time trend dummy variables. In all of them, the negative effect of women's wages was significant. Their best fitting model was a Weibull model, with a mover-stayer heterogeneity control. and, as in all the models where they introduce the mover-stayer structure, the nonparametric maximum likelihood converges to one point distribution: thus there is no unobservable correlated across spell and the authors emphasise the homogeneity of Swedish women's behaviour.

We also find a robust negative effect of women's wages, as in Sweden, for each parity, for each cohort and for each specification, but our results are slightly different.

- We find that the quadratic model fits the data better than the Weibull. In our quadratic model, the base line hazard duration increases and then decreases in accordance with the influence of age on the first birth. For further birth order transitions, the maximum is reached earlier and the decrease is steeper, in accordance with the general patterns of observed timing shown in graphs 1. Both models pass the chi-squared test but the fit of the quadratic model predicts more accurately the proportion of women in Cohort 1 who are childless at 35, and the parity reached at 20. Its likelihood is also much higher and this is not due to the larger number of parameters estimated in the model, as it dominates the Weibull model for the AIC and BIC tests that penalise complexity (Table1).

As in the Swedish case, we compare the performance of the neo-classical model with the purely demographic model described by Rodriguez et al. (1984). In this model, only the length of interval between births is used as an explanatory factor. In the demographic literature, the study of fertility behaviour shows that the timing of fertility is largely determined by the length of intervals between previous births (Rodriguez et al. 1984).

This length of birth intervals can be used to represent persistent unobserved factors between births. In the fertility process, when these factors represent biological differences in fecundity between women, the women with lower fecundity must have longer than average birth intervals, which translates a negative effect of birth intervals on each transition. This result is confirmed in our estimates. A longer interval between the first two births increases the length of interval between subsequent births (negative coefficient values for the various transitions), which may be interpreted either as the influence of unobservable biological factors or as behaviour dependence in planning family size (Yamaguchi and Ferguson 1995).

To compare the models, tests were made for a quadratic birth process ($K_j = 2, \lambda_{1j} = 1,$ and $\lambda_{2j} = 2$), with a mover stayer heterogeneity control, but the results obtained with other specifications, the Weibull ($K_j = 1, \lambda_{1j} = 0$) for instance, were the same. We find that both of them pass the chi-squared test at 5% significance level . Nevertheless, the BIC ad AIC tests, based on modified likelihood ratios show that the neo-classical model fits the data better than the purely demographic model (Table 1).

Can we improve the specification of the neo-classical model? What is the best heterogeneity structure? With this two-parameter quadratic base-line model, using the test in Table 1, we find that the best fitting model is the one that combines both sorts of heterogeneity: the parity-specific term associated with the mover-stayer structure and a women-specific heterogeneity term, correlated across spells.

This model passes the chi-squared tests (Table 2) for predicting attained fertility by ages 20, 25, 30 and 35 for each cohort, at a 1% level of significance for all parities, ages and cohorts, except for predicting childless women of Cohort 1 at age 35.

Since all the specifications used (under the various hypotheses made as to unobserved heterogeneity) are nested, we can also compare them with the test of maximum likelihood ratio (Table 3). It clearly shows that this more general specification of the model is the best. This initial result differs from that obtained by Heckman and Walker (1990a, b), who are unable to obtain a convergent model in the presence of women-specific heterogeneity (NPMLE), once they have allowed for parity-specific heterogeneity. Consequently, our study shows that the fertility behaviour of the women studied is less homogeneous that that of the Swedish population analysed by Heckman and Walker (1990a, b). It also shows that, unlike in Sweden, the conventional approach of estimating each birth parity independently may be erroneous for France. The reason is that the presence of unobservables correlated over spells can produce inconsistent estimates of the model parameters because it tends to maximise erroneous likelihood.

We now analyse the results of this model.

4.2. Effects of model variables on the timing and the spacing of births

Table 4 presents the estimates of the quadratic model with a parity-correlated women-specific unobservable and a mover-stayer heterogeneity control. Estimated female wage coefficients are negative at 1% significance level for all transitions and for all cohorts. Their magnitude is larger for the third birth. Men's wages have a positive effect at 1% significance on the first birth, but no significant effect on the second birth. This positive effect kicks in again for the third birth with women of Cohort 2 and at the 8% threshold limit for women of Cohort 3. The influence of a rural upbringing does not appear to be significant except for a positive effect —

growing up in a rural environment advances the first birth of Cohort 3 — and negative effects on the second birth for Cohort 1 and the third birth of Cohort 3.[1]

Base line duration coefficients are all significant. The base line hazard duration increases and then decreases, according to the influence of age on the first birth. For later birth order transitions, the maximum is reached earlier and the decrease is steeper, especially for the second birth. All cohorts show the same pattern but the details are different: compared to the earlier cohort, the women of the youngest cohort postpone their first birth and Cohorts 2 and 3 have their third birth sooner after the second.[2]

To assess the robustness of our findings, we estimated about 200 models, changing the specification of the birth process — Weibull, Exponential, Gompertz, Cubic and Spline — for every base line specification, we estimated models with no unobserved heterogeneity, with heterogeneity introduced as parity specific stopping probability controlled with and without a parity-correlated women-specific unobservable (NPMLE). We attempted to enhance our best model by adding various control variables (table 5); we controlled for women education level by adding three educational levels; we controlled for age by adding age as a regressor; we controlled for period effect by adding a linear time trend. Women's wage effect is robust for all these specifications, although their magnitude is smaller with period or age control. Men's wage significantly advances the first birth for all cohorts, and with most of the specifications; it advances the third birth of Cohorts 2 and 3. When no *women-specific heterogeneity term* is taken into account; its positive effect becomes significant for the second birth of Cohorts 2 and 3.

We estimated further models including other demographic variables. In the first model, we combined both demographic and economic approaches: we added previous birth intervals and time from marriage till first birth as regressors[3] In this enhanced model**,** the introduction of demographic variables increases the influence of women's wages and reduces that of men's wages, which become non-significant where they had been significant, except for the first birth. This fact may indicate a correlation between men's wages and a general advance in timing of births for each birth order. The negative effect of women's wages is robust and even higher than in the pure neo-classical model. At all events, this augmented model with both economic and demographic variables does not pass the chi-squared goodness-of-fit test for attained parity at age 30 and 35.

Our results show that higher women's wages generally postpone the timing of births and lengthen the spacing especially between the second and third births. This delaying effect holds for each transition, whatever the generation and generally increases with birth rank.

Adding the spacing between previous births as regressor in the model or taking account of a women-specific heterogeneity term increases this effect.

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[1] Place of upbringing has little significant influence in most models when account is taken of all the other explanatory variables. There is probably an indirect effect on fertility of this variable, which may be a determining factor for educational level and consequently wages.

[2] The difference in parameter values between cohorts may, however, be linked to the fact that the observation periods are not comparable between cohorts. For example, the fertility of women born between 1958 and 1962 (Cohort 3) is analysed until the age they achieved at the end of the observation period in 1995, namely 33-37, when many of these women have not yet completed their families. To examine this possibility, we adjusted our models to observation periods cut off at ages 25, 30 and 35. The estimated parameters in the various models are in every case consistent with the comments we have made.

[3] The model with a mover-stayer heterogeneity control does not admit women-specific heterogeneity.

The influence of men's wages is more problematic.

The effect here is robust for the first birth, but the results for other parities are much more fragile. An increase in men's wages advances the first birth but when all results are compared, this effect is not so clear for the second and third births, particularly for the earlier generation. It greatly depends on the specification of the model for the younger generations, and indeed it becomes non significant when demographic variables are included. For more accurate information on this role of men's wages, we estimated another model where the starting date for observation is taken as the date of marriage. In this model, the effect of men's wages on the first birth is no longer significant and may therefore be interpreted as a positive "spouse" effect. Our results here are comparable with those obtained by Heckman and Walker (1990, 1992) when they take account of the effect of marital status in their analyses. Nevertheless, men's wages do advance the second and third births of the younger generations in this model.

One value of the method use is the possibility of identifying the proportion of individuals who are never likely to change parity.[4] The parity-specific heterogeneity parameters for each birth rank — 1, 2 and 3 — are all negative and statistically significant. This procedure can therefore be used to identify biological sterility and permanent, efficient use of contraception. The probability of never being exposed to the risk of having a child increases with parity, but its value is never high. The probability of never going beyond parity 2 is highest for Cohort 3 at 26%, and lowest for Cohort 2 at 10%.

Overall, the inclusion of a women-specific heterogeneity term correlated across does not radically alter the coefficients of the other variables. It does however increase in absolute terms the negative effect of women's wages on the second and third births. It considerably reduces the effect of men's wages: without taking into account a women-specific heterogeneity term correlated across spells, higher men's wages speed the arrival of second and third births for couples in the younger generations. When these unobserved factors are taken into account, the influence of men's wages is noticeably lower than previously observed.

In the quadratic model, the women-specific unobserved heterogeneity terms ε_1, ε_2 and ε_3 take positive values (significant except for rank 1 for Cohorts 2 and 3). Their probability terms p_i are also significant for each cohort. From these results one may see that some of women's unobserved characteristics have positive effects on the likelihood of having a second and third child. These new hypotheses based on unobserved heterogeneity also reduce the effect of men's wages and increase in absolute terms the effect of women's wages on the second birth. So the effect of the unobserved characteristics compensates for the negative effect of higher wages on fertility.

What could these characteristics be? They have a very strong effect on the second birth but less on the third. Similarly the effect on the second birth is greater for Cohort 1 and gradually declines for the other two cohorts. Might these characteristics be related to women's attitudes towards the family? If women have a definite preference for a particular number of children (say, two) and these children provide them with positive "utility" for the future, the women will certainly intend to have them. We have seen that for the early cohorts, the decreased likelihood of a third order parity progression was part of a trend towards smaller variety in family size, mainly for two children. For more recent generations the stabilisation and then rise in the likelihood of a third order parity progression coincides with the end of the focus on

[4] For example, the existence of a strong preference for a particular number of children will always correspond to a non-zero likelihood for parities below the desired number and, in most cases, zero for the likelihood of subsequent births.

two children. These two observations may therefore be consistent with the greater effect of unobserved characteristics on the second birth for the early generations than for the last two. Furthermore, these unobserved characteristics cannot correspond to such factors as fertility, or the sign observed would have been negative. We may suppose that they correspond to the effect of women having a strong preference for children, who in other words accord them considerable significance. This interpretation is in line with the influence of unobserved heterogeneity on the observed effect of women's wages on the second birth. It is compatible with the lower influence of men's wages if we assume that a women with a greater preference for children will attempt to choose a husband earning higher wages.

4.3. Influence of maternity leave

Maternity leave policy in France has two aspects, length of leave and partial remuneration, which depends directly on the wages women receive. Table 8 (32) gives a description of these measures in France from 1962 to 1995. The legal duration of maternity leave is expressed in weeks and remuneration during that period as a percentage of usual women's wages. Before 1980, maternity leave only lasted 14 weeks. From 1980 this rose to 16 weeks for the first two children and 26 weeks for the third and subsequent children. Compensation was 50% of wages from 1962 to 1969 for the first two births and 66.6% for the third and subsequent births. The figure rose to 90% of wages in 1970 and then fell to 84% in 1986, whatever the number of children.

We computed the total amount of remuneration received during this period (based on current women's wages and legal duration of maternity leave).

This measurement was combined with a variable that indicates women's participation in the labour market; in other words, this measurement only enters the model for working women. The variable then enters independently into the regressions (quadratic model with unobserved heterogeneity controlled by parity-specific stopping and women-specific heterogeneity.

For the first birth, the influence of various variables measuring maternity leave was non-significant for whatever cohort (Table 6). For the second and third births, the estimated coefficients of the total amount of remuneration received during that period are clearly positive and significant for Cohorts 2 and 3. The magnitude of the coefficients is roughly one-quarter of the strong negative impact of women's wages.

The influence of maternity leave arrangements is never significant for Cohort 1 (Table 6). This result most likely reflects the fact that the proportion of women who had already had a second or third child before the years when the legal duration and compensation of maternity leave changed was already considerable. For example, in 1980, the proportion of women in Cohort 1 with two children was close to 43% by or near the age of 30, and varied little afterwards, reaching 45% by or near the age of 35 (Table 4).

The direct effect of the total sum received during maternity leave is positive for the second and third births and partly compensates (roughly one-quarter) for the negative effect of women's wages. Taken as a single measure, the extension of the legal duration of maternity leave also had a positive effect when it affected generations of women during their main child-bearing period.

<u>5. Conclusion</u>

This article gives one answer to the question of the economic determinants of the timing and spacing of births in France. The economic literature has presented the relationship between economic variables, such as prices and wages, and fertility behaviour, and has advanced two hypotheses: (i) increases in women's wages, as a measurement of the price of women's time, cause a reduction or postponement in fertility, whereas (ii) men's wages or income have a positive effect on fertility. Furthermore, the endogeneity of wage variations — which themselves depend on fertility — is a major problem at individual level that few authors have emphasised and which we have considered in this article. To that end we have amended the method described by Heckman and Walker for Sweden and have analysed the fertility, career and wage history of 7,500 women between 1975 and 1995.

Our results first show the major effect of women's wages on the timing and spacing of births, supporting Heckman and Walker's conclusions. In line with the neo-classical fertility model, an increase in women's wages causes a general delay in the timing of births. Furthermore, this result holds whatever the specification of the model and the definition of wages adopted — current wages or past wages — and the extra explanatory variables introduced into the analysis. The substitution effect linked to the increase in women's wages considerably outweighs the income effect. This compensation of the income effect by the substitution effect is generally greater for the birth of the third child than for the first two.

The influence of men's wages is less clear. We have observed that when the observation starting date is taken to be the date of the marriage, the positive effect of men's wages on the first birth is no longer significant; this effect must therefore be interpreted as a "spouse" effect. Our results here are identical to those obtained by Heckman and Walker (1990a, b, 1992), when they take account of marital status in their analyses. However, when unobservable factors correlating with births and the length of interval between births are taken into account, this noticeably reduces the influence of men's wages on second and third births. The effects of men's wages on fertility depend on the birth rank of the child and the various measurements of wages used, and differ from one cohort to another. One may not, therefore, conclude that there is a generally positive effect of men's wages, as presented in Heckman and Walker's work (1990a, b, 1992), based on aggregate wage series. Our results agree in this way with those obtained by Tasiran (1995), who also uses microwage data and demonstrates a limited effect of men's wages.

The extent of the effects of women's and men's wages (where we have found them to be significant) is lower than that found by Heckman and Walker (1990a, b, 1992) for Sweden. This corresponds to Tasiran's findings (1995), when, using the same biographical data for Sweden, he proposes replacing some of the aggregate series used by Heckman and Walker (1990a, b) — smoothed wages where the only individual variations are annual variations corresponding to average wage increases by age — by microwage data for part of the observation period. Before considering this result, it must be borne in mind that the use of aggregate or microwages in these two studies does not make it possible to entirely escape biases due to endogeneity and wage selection on fertility.

The gross microwages used by Tasiran are the product of past decisions to invest in human capital, allocate one's time, career choices and fertility; they correlate highly with the unobserved determinants of participation in the labour market. Consequently, some of the correlation observed between fertility and wages may be the result of the correlation between wages and couples' preferences for children. Estimates of instrumental earnings functions we have made (Ekert-Jaffé and Mougin 2002) can be used to determine microwages uncorrelated with individuals' past decisions and avoid endogeneity biases in measuring wage effects. Our analysis points out the considerable caution that is required when individual data are used.

The most controversial aspect of Heckman and Walker's analysis (1990a, b, 1992) is the use they make of aggregate rather than microwages. Although using annual wage series reduces the likelihood of simultaneous bias between wages and fertility, variations in aggregate wages may also be linked to unobserved labour supply factors that may be jointly determined with fertility — women's wages rise at times of low fertility and intensive investment in the labour market. In that case, market prices are potentially affected by the same simultaneous equation biases as those found at individual level (Rosenzweig and Schultz 1985 and Schultz 1985).

Since both Tasiran and Heckman and Walker (1990a, b, 1992) define the wage formation process as exogenous to the fertility process, they do not use educational level or professional experience in their analyses. So their results may well reflect the major negative effect on fertility of age at end of education and professional experience.

This debate illustrates the complexity of analysing the dynamic aspects of fertility. The two greatest problems are the lack of microwage data linked over time to family biographies, and the modelling of individual data. These difficulties may explain the divergences in the findings of these empirical studies. We show that when both aspects are taken into account, the negative effect of women's wages on the timing and spacing of each birth rank is smaller but does still exist. In this case, our results confirm those initially obtained by Heckman and Walker (1990a, b, 1992) from aggregate wage series.

In addition, we estimated the impact of maternity leave with this model. The direct effect of the total sum received during maternity leave is positive for the second and third births and partly compensates (roughly one-quarter) for the negative effect of women's wages. Taken as a single measure, the extension of the legal duration of maternity leave also had a positive effect when it affected generations of women during their main child-bearing period.

REFERENCES

Akaike, H., (1973), 'Information theory and the extension of the maximum likelihood principle', dans *Second International Symposium on Information Theory*}, B.N. Petrov et F. Csaki (éd.), Akailseiniai-Kiudo, Budapest.

Arellano M. and B. Honoré (2001) 'Panel Data Models: Some Recent Developments', in *Handbook of Econometrics,* Volume 5, Heckman, J.. and E. Leamer (eds.), I, North-Holland, Amsterdam.

Becker, G., (1981), *A Treatise on the Family*, Cambridge, Harvard University Press.

Becker, G., et Lewis, H., G., (1973),'On the interaction between quantity and quality of children', *Journal of Political Economy*, 82, 279-288.

Baltagi (1995) B. H., *Econometric Analysis of Panel Data,* John Wiley &. Sons, 1995

Blanchet, D., and Ekert-Jaffé, O., (1994), 'The Demographic Impact of Family Benefits: Evidence From a Micro-model and From Macro Data', in J. Ermish et N. Ogawa (ed.), *Family, the Market and the Sate in Ageing Societies*, Clarenton Press, Oxford.

Card, D., (1999), 'The Causal Effect of Education on Earnings', in *Handbook of Labor Economics*, Vol.3, O.Ashenfelter et R. Layard (éd.), I, North-Holland, Amsterdam.

Chamberlain, G., (1985), 'Heterogeneity, omitted variables bias and duration dependence', in *Longitudinal Analysis of Labour Market Data*, J.J. Heckman et B. Singer (éd.), Cambridge

University Press, Cambridge.

Cigno, A., and J. Ermisch, (1989), 'A microeconomic analysis of the timing of births', European Economic Review, 33, 737-760.

Ekert-Jaffé, O., (1986), 'Effets et limites des aides financières aux familles : une expérience et un modèle', *Population*, 2, 327-346.

Ekert-Jaffé, O., Solaz, A., (2001), 'Unemployment, marriage, and cohabitation in France', *Journal of Socio-Economics*, 30, 75-98.

Ermisch, J., (1989), 'Purchased childcare, optimal family size and mother's employment. Theory and econometric analysis', *Journal of Population Economics*, 2 (2), 79-102.

Festy P., (1994), 'L'enfant dans la famille : vingt ans de changement dans l'envionnement familial des enfants', *Population*, 6, 1245-1296.

Goodman, L.A., (1961), 'Statistical methods for the mover-stayer model', *Journal of American Statistical Association*, 56 (296), 841-868.

Heckman, J.J., (1979), 'Sample Selection Bias as a Specification Error', *Econometrica*, 47, 153-161.

Heckman, J.J., et Singer, B., (1982), 'Population heterogeneity in demographic models', in *Multidimensional Mathematical Demography*, K. Land and A. Rogers (éd.), Academic Press, New-York.

Heckman, J.J., et Singer, B., (1984), 'The identifiability of proportional hazard model', *Review of Economic Studies*, 51, 231-241.

Heckman, J.J., et Singer, B., (1984b), 'A Method for minimizing the impact of distributional assumptions in econometric models for duration data', *Econometrica*, 52, 271-320.

Heckman, J.J., and Walker, J.R. (1987), 'Using goodness of fit and other criteria to choose among competing duration models: a case studies of Hutterite data', in *Sociological Methodology*, C. Clogg (éd.), American Sociological Association, Washington, DC.

Heckman, J.J.,and Walker, J.R. (1990a), 'The relationship between wages and income and the timing and spacing of births : evidence from Swedish longitudinal data', *Econometrica*, 58 (6) 1411-1441.

Heckman, J.J., and Walker, J.R. (1990b), 'Economic models of fertility dynamics: a case study of Swedish fertility', in *Research of Population Economics*, ed. P. Schultz, VII, (JAI Press, Greenwich, conn.).

Heckman, J.J., and Walker, J.R. (1992), 'Understanding the Third Births in Sweden', in *Demographic Applications of Event History Analysis*, eds. J. Trussell, R. Hankinson, et J. Tilton, (Clarenton Press, Oxford).

Hoem, B., et Hoem, J., (1989), 'The impact of women's employment on second and third births in modern Sweden', *Population Studies*, 43, 47-67.

Hoem, J., (1990), Hoem, J.M., (1990a), 'Limitations oh heterogeneity technique: selectivity issues in conjugal union disruption at parity zero in contemporary Sweden', in *Convergent Issues in Genetics and Demography*, J. Adams, D.A. Lam, A.I. Hermalin, et P.E. Smouse(ed.), Oxford University Press, Oxford.

Hsiao, C., (1986), *Analysis of Panel Data*, Cambridge University Press.

Kravdal, O., (1992), 'The emergence of a positive relation between education and third birth rates in Norway with supportive evidence from United States', *Population Studies*, 46, 459-475.

Kravdal, O., (2001), 'The high fertility of college educated women in Norway. An artefact of a separate modelling of each parity transition', *Demographic Research*, 5(6), 187-214.

Mincer, J., (1963), 'Market Prices, Opportunity Cost and Income Effects', in *Measurement in Economics*, C. Christ and al. eds, (Stanford University Press, Stanford, CA)

Montgomery, M., et Trussel, J., (1986), 'Models of marital status and childbearing', in *Handbook of Labor Economics*, O.Ashenfelter et R. Layard (éd.), I, North-Holland, Amsterdam.

Mougin, R. 2004. *Fécondité et salaires, une analyse des déterminants économiques du calendrier des naissances en France*, Economics thesis, 15 March, IEP, Paris. 471 p.

Mougin R.and O. Ekert-Jaffé 2002. "Estimating wage equations with sample selection bias on a French panel", Conference on Econometrics of Wages, *Applied Econometrics Association*, Brussels, 28-29 May.

Nijman, T., and M. Verbeek, (1992), 'Nonresponse in Panel Data: The Impact on Estimates of a Life Cycle Consumption Function', J*ournal of Applied Econometrics*, 7, 243-257.

Powell, J.L., (1994), 'Estimation of semiparametric models', in *Handbook of Econometrics*, Vol. IV, 2444-2521, Amsterdam : North-Holland.

Ridder, G., (1987), 'The sensitivity of duration models to misspecified unobserved heterogeneity and duration dependence', document de travail, Groningen University, Groningen.

Rodriguez, G., Hobcraft, J., McDonald J., Menken, J., et Trussell,J., (1984), 'A Comparative Analysis of Determinants of Birth Intervals', Comparative Studies}, 30 (May), World Fertility Survey (London).

Rosenzweig, M., and Schultz, T.P., (1985), ' The supply of and demand for births : Fertility and its life-cycle consequences', *American Economic Review*, 75, 992-1015.

Schultz, T.P., (1985), 'Changing world prices, women's wages, and the fertility transition : Sweden, 1860-1910', *Journal of Political Economy*, 93, 1126-1154.

Schwartz, G., (1978), 'Estimating the dimension of a model', *Annals of Statistics,* 6, 461-464.

Taniguchi, H., (1999), 'The timing of childbearing and women's wage', *Journal of Marriage and the Family*, 61(4), 1008-1019.

Tasiran, A.C., (1995), *Fertility Dynamics: Spacing and Timing of Births in Sweden and the United States*, Contribution to Economic Analysis, (North-Holland, Amsterdam).

Toulemon, L.,(1994), Laplace des enfants dans l'histoire des couples', *Population*, 6, 1321-1346.

Trussel, J., et Rodriguez, G., (1990), 'Heterogeneity in demographic Research', in *Convergent Issues in Genetics and Demography*, J. Adams, D.A. Lam, A.I. Hermalin, and P.E. Smouse(eds.), Oxford University Press, Oxford.

Vella, F., (1998), 'Estimating Models with Sample Selection Bias : A Survey', *Journal of Human Ressources*, 33, 127-169.

Vella, F., et Verbeek, M., (1999), 'Two-Step Estimation of Panel Data Models with Censored Endogenous Variables and Selection Bias', *Journal of Econometrics*, 90, 239-263.

Yamaguchi, K., et Ferguson, L., (1995), 'The stopping and spacing if childbirths and their birth-history predictors: rational-choice theory and event-history analysis', *American Sociological Review*}, 60, 272-298.

Table1. Model Selection : Criterion of model Fit

Model and specification	Cohort 1 1946-1950		Cohort 2 1952-1956		Cohort 3 1958-1962	
	AIC[a]	BIC[b]	AIC[a]	BIC[b]	AIC[a]	BIC[b]
Neoclassic						
Weibull	10,70	-8613,6	10,6	-7853,1	10,21	-7714,5
Quadratic	10,51	-8467,0	10,5	-7815,1	10,21	-7727,3
Quadratic with a women specific heterogeneity term	10,41	-8390,5	10,4	-7750,7	10,09	-7643,9
Pure demographic						
Quadratic	11,68	-9394,7	11,9	-8812,1	11,2	-8481,0

[a] Akaike Information Criterion (1973): (2/N) log likelihood. + (2/N) (# estimated parameters)

[b] Leamer-Schwarz Criterion (1978) (Bayesian Information Criterion) : log likelihood – [InN/2] (# estimated parameters).

All models have a mover-stayer structure.

Table 2. Chi-square tests for goodness of fit of Table 4 quadratic models.

	Cohort 1 1946-1950			Cohort 2 1952-1956			Cohort 3 1958-1962		
Number of births	act.[a]	pred.[b]	test	act.[a]	pred.[b]	test	act.[a]	pred.[b]	test
by age 20									
n = 0	0.861	0.849	0.31	0.800	0.784	0.53	0.852	0.841	0.36
n = 1	0.118	0.114	0.05	0.166	0.165	0.50	0.128	0.122	0.59
n = 2	0.020	0.033	1.56	0.028	0.025	0.02	0.019	0.015	0.16
n = 3+	0.001	0.002	0.39	0.007	0.005	0.82	0.001	0.000	1.34
joint			2.03			1.29			2.47
by age 25									
n = 0	0.333	0.325	0.24	0.339	0.331	0.24	0.343	0.325	1.25
n = 1	0.381	0.378	0.02	0.372	0.363	0.18	0.381	0.375	0.10
n = 2	0.227	0.229	0.05	0.224	0.227	0.02	0.226	0.229	0.08
n = 3+	0.059	0.062	0.11	0.065	0.063	0.02	0.050	0.042	0.95
joint			1.02			1.09			2.58
by age 30									
n = 0	0.110	0.105	0.30	0.112	0.104	0.41	0.081	0.075	0.96
n = 1	0.298	0.292	0.02	0.299	0.289	0.22	0.268	0.259	0.20
n = 2	0.428	0.430	0.09	0.404	0.406	0.01	0.457	0.437	1.28
n = 3+	0.165	0.169	0.08	0.185	0.190	0.09	0.194	0.190	0.03
joint			1.49			1.25			1.54
by age 35									
n = 0	0.068	0.059	1.65	0.061	0.059	0.04			
n = 1	0.226	0.222	0.04	0.228	0.222	0.11			
n = 2	0.453	0.46	0.15	0.443	0.441	0.01			
n = 3+	0.254	0.263	0.79	0.268	0.265	0.02			
joint			2.09			1.01			

[b] predicted probability

X^2critical values : Bonferroni test critical values :

d.f.	10%	5%	1 %	d.f.	1.67%	1.25%
3	6.25	7.81	11.35	3	10.25	10.88
1	1.64	2.74	5.41			

Table 3. Chi-square tests for likelihood ratio of various Specifications of Unobserved Heterogeneity[a] of the quadratic model.

	Cohorte 1 1946-1950		Cohorte 2 1952-1956		Cohorte 3 1958-1962	
	MSH*	UM**	MSH*	UM**	MSH*	UM**
WH***	*36,38 (3)*	*210,92 (3)*	*97,1 (3)*	*262,63 (3)*	*66,39 (3)*	*261,61 (3)*
MSH*		*174,54 (3)*		*165,53 (3)*		*195,21 (3)*

Note : degrees of freedom in parentheses.

[a] Twice the difference between the log-likelihood of the model (in column)and that of the reference model (in line).

[*] Heterogeneity control by only Parity Specific Stopping (proportion of women that never leave parity j).

[**] Heterogeneity control by Both Parity Specific Stopping and women specific heterogeneity term, correlated across spells (NPMLE).

[***] No control of the Unobserved Heterogeneity.

Table 4. Quadratic Birth Process Model (Kj=2, λ_{1j} =1, λ_{2j}=2, j = 1, 2, 3) with instrumented wages variables and Unobserved Heterogeneity control by Both Parity Specific Stopping and a women specific heterogeneity term, correlated across spells **(NPMLE)**

	Cohort 1 1946-1950		Cohort 2 1952-1956		Cohort 3 1958-1962	
	Estimate	Std.-Err.	Estimate	Std.-Err.	Estimate	Std.-Err.
Variable/transition	First Birth		First Birth		First Birth	
intercept	-4.2368	.4119	-3.9074	.7196	-5.1663	.7174
γ_1	0.3469	.0330	0.2202	.0400	0.25	.0422
γ_2	-0.0322	.0026	-0.0193	.0032	-0.0251	.0042
rural	0.0471	.0639	0.1078	.0737	0.1625	.0677
women wage	-0.377	.0504	-0.3692	.0832	-0.2077	.0684
male wage	0.4467	.0228	0.4719	.0235	0.4158	.0296
ε_1	0.1691	.0883	0.0764	.1045	0.1146	.1377
Variable/transition	Second Birth		Second Birth		Second Birth	
intercept	3.7674	.5754	3.2742	.9357	-1.1018	1.013
γ_1	1.6746	.0702	1.3424	.0629	1.5294	.0711
γ_2	-0.2587	.0098	-0.1824	.0085	-0.2046	.0185
rural	-0.183	.0856	-0.1683	.0956	-0.0374	.0877
women wage	-0.9097	.0601	-1.1096	.0965	-0.5247	.0988
male wage	-0.0143	.0486	0.1807	.1037	0.0336	.0918
ε_2	3.142	.1370	2.9512	.1401	2.7313	.1389
Variable/transition	Third Bird		Third Birth		Third Birth	
intercept	4.9806	1.042	5.4552	1.289	7.3286	1.886
γ_1	0.2969	.0544	0.4866	.0656	0.768	.0766
γ_2	-0.0586	.0069	-0.0986	.0105	-0.1532	.0154
rural	-0.1036	.1096	-0.1558	.1086	-0.2828	.1208
women wage	-0.8793	.0934	-1.4805	.1307	-1.411	.1379
male wage	-0.0242	.1105	0.4797	.1298	0.2971	.1727

ε_3	0.8921	.1502	1.1315	.1900	1.0857	.2078
Estimates[a] :						
	Parity 0		Parity 0		Parity 0	
μ_0	-3.7909	.2817	-3.2073	.171	-4.2081	.5302
Probabilities	0.0221		0.0389		0.0147	
	Parity 1		Parity 1		Parity 1	
μ_1	-2.8519	.3402	-2.3527	.9449	-2.0360	.1949
Probabilities	0.0546		0.0869		0.1155	
	Parity 2		Parity 2		Parity 2	
μ_2	-1.5941	.461	-3.5181	2.889	-1.217	.3414
Probabilities	0.1688		0.0288		0.2285	
p_i[b]	0.5207	.0206	0.4183	.0213	0.2970	.025
- log-L	8302.01		7663.15		7556.17	
N	1600		1472		1502	
K	24		24		24	

[a] The probability of never leaving parity j is : $P_j=(1 + \exp-\mu_j)^{-1}$,

[b] With two points mass

AIC	10.408	10.111	10.09
BIC	-8390.543	-7750.68	-7643.94

$$h_j(t_j|\mathrm{H}(\tau(j-1)+t_j),\theta)=\exp\{\ \gamma_{oj}+Z(\ \tau(j-1)+t_j)\beta_j+\sum_{k=1}^{K_j}(\frac{t_j^{\lambda_{kj}}-1}{\lambda_{kj}})\gamma_{kj}+\varepsilon_j\theta\ \}$$

Table 5. Robustness of Estimated Men's Wage and Women's Wage Coefficients to Different Variables added as a Covariate to baseline regressors, and to Different Control of Unobserved Heterogeneity.

	First transition		Second transition		Third transition			
			Panel A : Cohorte 1 -1946-1950				Model : variables added as covariate	AIC[d]
	Estimate	Std. Error	Estimate	Std. Error	Estimate	Std. Error		
No unobserved heterogeneity (UH) control :								
Women's wage	-0.3769	.045	-0.5703	.041	-0.7061	.075		
Men's wage	0.04440	.021	-0.0461	.051	-0.1847	.100	Basic model	10,526
UH Control by parity specific stopping[e] :								
Women's wage	-0.3800	.0498	-0.6241	.047	-0.7841	.089	(1,2,3)[a]	
Men's wage	0.4480	.0219	-0.0265	.0534	-0.1657	.107	Basic model	10,513
Women's wage	-0.3784	.050	-0.6595	.046	-0.7865	.088	(1,2,3)	
Men's wage	0.4475	.022	-0.0440	.051	-0.2921	.109	With Education	10,486
Women's wage	-0.3802	.050	-0.630	.049	-0.813	.091	Bac, Tertiary	
Men's wage	0.4483	.022	-0.043	.052	-0.119	.120	Durmar, second spacing	10,427
Women's wage	-0.6276	.043	-0.6808	.046	-0.9416	.090	(1,2,3)	
Men's wage	0.3506	.021	-0.0524	.047	-0.2299	.109	Age at marriage	10,381
Women's wage	-0.3802	.050	-0.6559	.050	-0.8140	.091	(1,2,3)	
Men's wage	0.4481	.022	-0.0474	.0575	-0.1172	.133	Intervalle1, intervalle 2	10,486
Women's wage	-03810	.0498	-0.5126	.0500	-0.6650	.0885	Age	10,49
Men's wage	0.4478	.0221	0.1676	.0597	0.2417	.1330		
Women's wage	-0.2589	.0503	-0.5168	.0515	-0.7059	.0912	Time Period	10,42
Men's wage	0.4639	.0222	0.1119	.0627	0.1324	.1366		
UH Control by parity specific stopping and by women specific heterogeneity[f] :								
Women's wage	-0.377	.050	-0.9097	.060	-0.8993	.093	(I =2)[b](1,2,3)	
Men's wage	0,4467	.022	-0,0143	.0486	-0,0242	.110	Basic model	10,814
Women's wage	-0.6258	.044	-0.9874	.064	-1.0174	.095	(I = 2) (1,2,3)	
Men's wage	0.3490	.022	-0.1371	.050	-0.1033	.113	Age at marriage	10,286
Women's wage	-0. 3770	0.51	-0.9655	.071	-0.8576	.104	(I=2) (1,2,3)	
Men's wage	0.4466	.023	-0.1716	.055	-0.1128	.134	Previous Intervalles	10,403

(I=2) (1,2,3)

	First transition		Second transition		Third transition			
			Panel B : Cohorte 2 -1952-1956				Model : variables added as covariate	AIC[d]
	Estimate	Std. Error	Estimate	Std. Error	Estimate	Std. Error		
No unobserved heterogeneity (UH) control :								
Women's wage	-0.3892	.077	-0.9876	.072	-1.5208	.118		
Men's wage	0.4831	.023	0.4063	.080	0.4839	.127	Basic model	10,568
UH Control by parity specific stopping[e] :								
Women's wage	-0.3700	.082	-0.9919	.074	-1.5228	.129	(1,2,3)[a]	
Men's wage	0.4723	.023	0.4028	.081	0.4826	.127	Basic model	10,543
Women's wage	-0.3737	.085	-1.0223	.073	-1.5590	.130	(1,2,3)	
Men's wage	0.4753	.023	0.2572	.086	0.3558	.135	With Education	10,50
Women's wage	-0,226	.076	-1.1487	.072	-1.7793	.164	(1,2,3)	
Men's wage	0.4793	.024	0.1646	.081	0.1085	.133	Durmar, second spacing	10,291
Women's wage	-1.1189	.060	-1.0301	.067	-1.7846	.138	(1,2,3)	
Men's wage	0.3370	.022	0.2171	.077	0.1266	.125	Age at marriage	10,231
Women's wage	-0.1260	.0753	-1.6535	.086	-2/2037	.114	(1,2,3)	
Men's wage	0.4793	.0240	0.0484	.072	0.0080	.127	Intervalle1, intervalle 2	10,724
Women's wage	-03700	.0828	-0.8672	.0956	-1.3771	.1647	Age	10,6
Men's wage	0.4724	.0233	0.4668	.0883	0.5416	.1346	8.42*	
Women's wage	-0.3247	.0818	-1.0357	.0924	-1.6376	.1656	Time Period	10,57
Men's wage	0.4864	.0233	0.3765	.0892	0.4302	.1378	227.12**	
UH Control by parity specific stopping and by women specific heterogeneity[f] :								
Women's wage	-0,3692	.083	-1,1096	.096	-1,4805	.130	(I =2)[b] (1,2,3)	
Men's wage	0,4719	.023	0,187	.103	0,4797	.129	Basic model	10,444
Women's wage	-1.3828	.079	-1.0284	.068	-1.7847	.138	(I = 2)[b] (1,2,3)	

Men's wage	0.3027	.023	0.2630	.081	0.1460	.128	Age at marriage	10,182

	First transition		Second transition		Third transition			
	Panel C : Cohorte 3 -1958-1962						Model : variables added as covariate	AIC[d]
	Estimate	Std. Error	Estimate	Std. Error	Estimate	Std. Error		
No unobserved heterogeneity (UH) control :								
Women's wage	-0.2117	.067	-0.5082	.071	-1.3338	.117		
Men's wage	0.4171	.030	0.2118	.066	0.2536	.152	Basic model	10,235
UH Control by parity specific stopping[e] :								
Women's wage	-0.2089	.068	-0.4716	.075	-1.4710	.138	(1,2,3)[a]	
Men's wage	0.4163	.030	0.2368	.068	0.3647	.171	Basic model	10,215
Women's wage	-0.2097	.069	-0.5231	.077	-1.5221	.141	(1,2,3)[a]	
Men's wage	0.4190	.030	0.1374	.073	0.1345	.184	With Education	10,183
Women's wage	-0,2087	.068	-0.7600	.083	-1.4718	.162	Bac, Tertiary (1,2,3)	
Men's wage	0.4161	.0300	0.1276	.067	0.1576	.145	Durmar, second spacing	10,059
Women's wage	-0.6357	.066	-0.7221	.075	-1.6341	.143	(1,2,3)	
Men's wage	0.3575	.029	0.1818	.065	0.2510	.146	Age at marriage	10,032
Women's wage	-0.2085	.068	-0.5921	.092	-1.3653	.156	(1,2,3)	
Men's wage	0.4163	.029	0.0933	.072	0.2702	.171	Intervalle1, intervalle 2	10,176
Women's wage	-0.2089	.0683	-0.2681	.0902	-0.9216	.1517	Age	10,172
Men's wage	0.4163	.0303	0.3872	.0696	0.5212	.1721		
Women's wage	-0.1947	.0673	-0.3054	.0886	-1.1497	.1526	Time Period	10,168
Men's wage	0.4397	.0305	0.3985	.0706	0.4520	.1753		
UH Control by parity specific stopping and by women specific heterogeneity[f] :								
Women's wage	-0,2077	.068	-0,5247	.098	-1,411	.137	(I =2)[b] (1,2,3)	
Men's wage	0,4158	.029	0,0336	.091	0,2971	.172	Basic model	10,093
Women's wage	-0.2073	.069	-0.5492	.100	-1.4613	.140	(I = 2)[b] (1,2,3)	
Men's wage	0.4187	.029	-0.0154	.109	0.1153	.185	Bac, Tertiary (1,2,3)	10,066
Women's wage	-0,2079	.068	-1.0924	.101	-1.4836	.167	(I =2 [b] (1,2,3)	
Men's wage	0,4157	.029	0.0906	.087	0.1602	.145	Durmar, second spacing	9,952

Basic model – added regressor : rural.

[a] probability of stopping before transitions 1, 2, 3.

[b] number of points used to estimate women specific heterogeneity.

[c] Singular estimated Hessian matrix.

[d] Akaike Information Criterion.

[e] Heterogeneity control by only Parity Specific Stopping (proportion of women that never leave parity j).

[f] Heterogeneity control by Both Parity Specific Stopping and women specific heterogeneity term, correlated across spells (NPMLE-Non Parametric Procedure for Likelihood Maximisation) – Heckman et Singer (1982, 1984b).

Table 6. Quadratic Birth Process Model (Kj=2, λ_{1j} =1, λ_{2j}=2, j = 1, 2, 3), Unobserved Heterogeneity control by Both Parity Specific Stopping and a women specific heterogeneity term, correlated across spells (NPMLE). Impact of the total amount received for maternity leave.

	Cohort 1		Cohort 2		Cohort 3	
	1946-1950		1952-1956		1958-1962	
	Estimate	Std.-Err.	Estimate	Std.-Err.	Estimate	Std.-Err.
Variable/transition	First Birth		First Birth		First Birth	
intercept	0.7144	.0321	-0.2276	.0362	-0.25025	1.039
γ_1	0.341	.0322	0.2381	.0365	0.2626	.0437
γ_2	-0.0312	.0021	-0.0211	.0032	-0.0247	.0049
rural	0.0265	.0644	0.0747	.0694	0.1593	.0685
women wage	-0.3149	.0586	-0.3214	.091	-0.2366	.1051
male wage	0.4238	.0222	0.4604	.0231	0.4410	.0311
maternity benefits x LFP	0.0346	.0811	-0.0403	.0952	-0.0098	.1191
LFP[b]	-0.5569	.6530	-0.1088	.8114	-0.3480	1.070
ε_1	0.1571	.0851	0.1127	.1071	0.1409	.1234
Variable/transition	Second Birth		Second Birth		Second Birth	
intercept	3.7945	.6961	4.6528	1.155	-0.1480	1.291
γ_1	1.6876	.0711	1.2099	.0591	1.4999	.0650
γ_2	-0.2637	.0092	-0.1671	.0081	-0.2004	.0143
rural	-0.1247	.0864	-0.2057	.0992	-0.0702	.0850
women wage	-0.8883	.0777	-1.2459	.1285	-0.9106	.1448
male wage	-0.1878	.1426	0.2146	.0985	0.2985	.0720
maternity benefits x LFP	-0.0017	.1226	0.3186	.1346	0.3168	.1572
LFP[b]	-0.0360	1.020	-1.2443	.5694	-1.3473	.5221
ε_2	3.1401	.1401	2.6192	.1401	2.7054	.1374
Variable/transition	Third Bird		Third Birth		Third Birth	
intercept	5.5316	1.181	7.9137	1.452	9.0149	1.993
γ_1	0.2887	.0551	0.5508	.0621	0.798	.0753
γ_2	-0.0577	.0071	-0.0983	.0101	-0.144	.0153
rural	-0.0686	.1111	-0.1876	.1262	-0.2855	.1213
women wage	-0.8911	.1242	-1.6489	.1626	-1.7136	.2114
male wage	-0.0872	.1072	0.3721	.1355	0.3735	.1456
maternity benefits x LFP	0.2352	.1412	0.4393	.2182	0.4863	.2204

LFP[b]	-2.2439	1.233	-1.7882	.8645	-1.2358	.5028
ε_3	0.9280	.1502	1.4469	.2091	1.3015	.2141
Estimates[a] :						
	Parity 0		Parity 0		Parity 0	
μ_0	-3.6667	.2292	-3.3496	0.1884	-3.7563	.2865
Probabilities	0.0249		0.0339		0.0228	
	Parity 1		Parity 1		Parity 1	
μ_1	-2.7669	0.3442	-2.3681	.9574	-2.1961	.1785
Probabilities	0.0591		0.0856		0.1001	
	Parity 2		Parity 2		Parity 2	
μ_2	-1.5894	.5000	-1.5252	.6857	-0.9997	.2032
Probabilities	0.1695		0.1787		0.2690	
- log-L	8301.81		7588.39		7451.61	
N	1600		1472		1502	
K	27		27		27	

[a] The probability of never leaving parity j is ：$P_j=(1 + \exp -\mu_j)^{-1}$,

[b] With two points mass

AIC	10.51494	10.5232065	10.2107457
BIC	-8484.55975	-7816.55409	-7740.01646

Figure1.A- Distribution of age at first birth, cohorts 1946-1950, 1952-1956and 1958-1962.

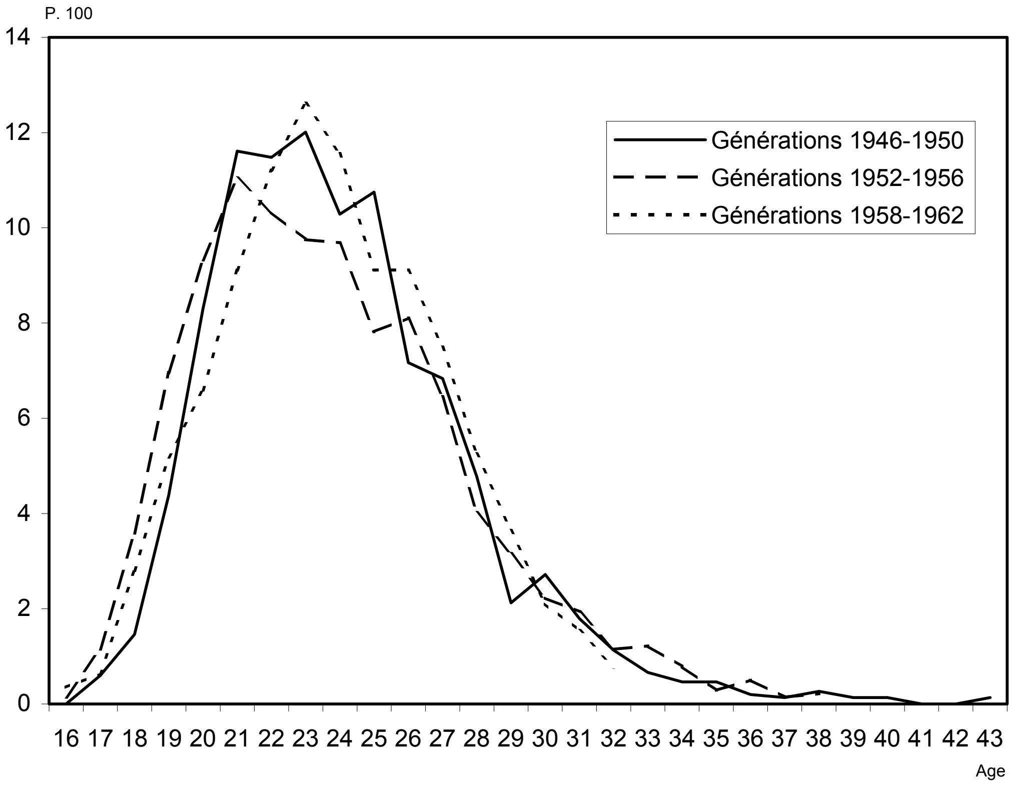

Figure 1B. - Distribution of second birth intervals, générations 1946-1950, 1952-1956 and 1958-1962,

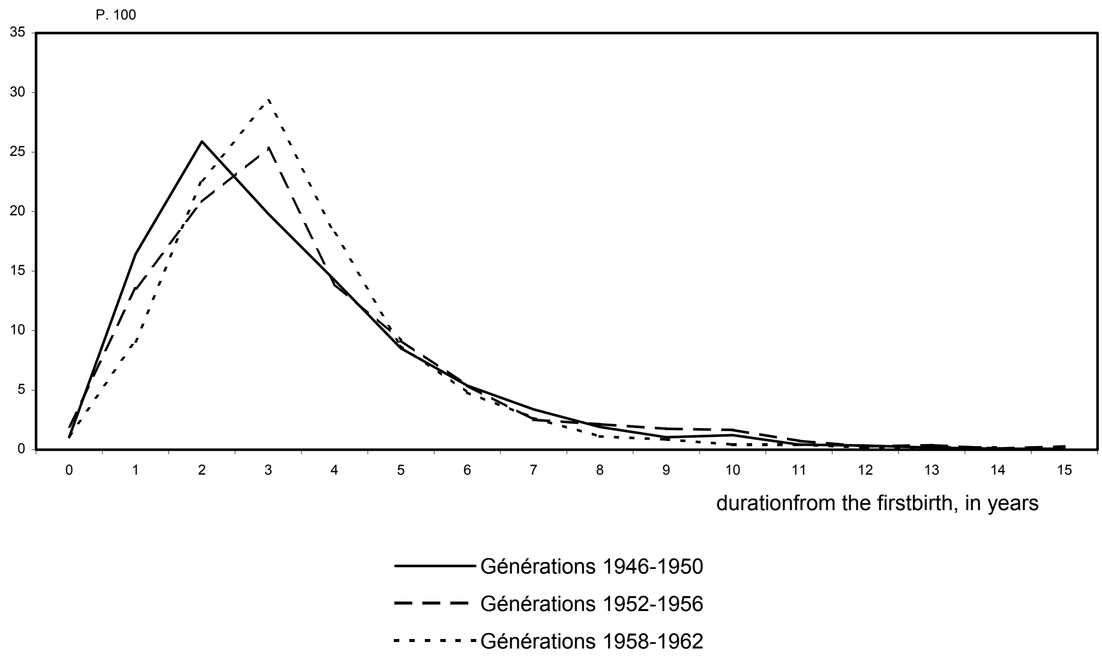

Figure 4. - Distribution of third birth intervals, générations 1946-1950, 1952-1956 et 1958-1962,