# The Effects of Childbearing Regional Contexts on Ideal Family Size in Europe: A Multilevel Analysis

Maria Rita Testa
Leonardo Grilli

Maria Rita Testa is research scientist at the Vienna Institute of Demography of the Austrian Academy of Sciences. Leonardo Grilli is research scientist at the Department of Statistics, University of Florence.

**Abstract**

The persistence of the two-child family ideal is no longer universally widespread in Europe (Goldstein et al. 2003), but the reasons why people prefer a given number of children have not yet been systematically investigated.

We examine the individual and regional factors of ideal family size by taking into account the unobservable similarities of people sharing the same demographic and socio-economic environment.

Multilevel binary and ordered logistic regression models are implemented, by using 2001 Eurobarometer data. The hierarchical structure is defined by respondents embedded within regions of the countries of the European Union.

The main result is that the context of actual fertility of the older generations influences the preferences of the younger cohorts: the lower, on average, the past actual childbearing in the region, the higher is the individual probability of people in reproductive ages to prefer smaller families.

## 1. INTRODUCTION: THE EMERGENCE OF BELOW-REPLACEMENT FAMILY SIZE IDEALS IN EUROPE

After several decades of a large predominance of the two-child norm in fertility preferences, sub-replacement family size ideals have emerged in the German-speaking countries where the younger women aged 20 to 34 reported an average ideal family size of 1.7 children in 2001 (Goldstein et al. 2003). Although Germany has always shown the lowest family size ideals among the European Union countries (European Commission 1979 and 1990), so far such lowest-low levels had not been registered.

The decline of ideal family size for young German and Austrian women has been explained by the drop of period fertility after the baby boom, which in Germany and Austria occurred earlier than in the other European countries. The German-speaking children born during the baby bust of the 1970s would have been socialised in smaller families and would have perceived as normal a family with only one child, or maximum two children. These cohorts would have benefited from all the advantages of being grown up in small families without any pressure to have more than one child being exerted by either their close neighbours or society at large. An interesting hypothesis made for the mechanism behind this process is that through social learning the young generations would have taken the actual childbearing behaviour of the previous generation as a standard for their own ideal fertility (Goldstein et al. 2003).

The starting point of our analysis is that fertility ideals are, on average, lower in Germany and Austria than in the rest of Europe, but these cross-national differences may hide considerable within-country heterogeneity.

Our research hypothesis is twofold. We aim to:

(1) verify whether the lower fertility preferences of Germans and Austrians still hold when we control for several demographic and socio-economic factors;

(2) test the hypothesis that ideal fertility of the young cohorts is influenced by the actual childbearing of the previous generations.

We adopt a multi-level framework in order to capture the role of different levels of determinants on individual preferences. The approach is considered as a promising avenue for comparative research, but so far there have been only few examples of cross-national multilevel analysis (Gauthier 2002; Festy 2003).

In the existing literature there is no broad theoretical framework on the process of forming ideal family size. The only consolidated finding is that childlessness is mostly involuntary[1]: women – who eventually remain childless – reach that state through a series of decisions to postpone childbearing, rather than definitively deciding at a young age against having children (Rindfuss et al. 1988; Morgan and Chen 1992; Toulemon 1996). While the low prevalence of voluntary childlessness is largely documented, the reasons why individuals may want a given number of children are still not clear. Nevertheless, future fertility preferences are a crucial topic for the post-transitional fertility (Bongaarts 2001) and a phenomenon that needs to be thoroughly investigated.

We intend to contribute answers to the following questions: Are smaller families preferred by a particularly set of people who can – perhaps – be considered as trendsetters? Who are these people? Are family size preferences influenced by context factors? Is the decline in German fertility ideals due to the higher proportion of German childlessness? The answer to these questions may shed light on the possible spread of sub-replacement family size ideals in other European countries.

The rest of the paper proceeds as follows: the next section discusses the measurement problems linked to the use of different questionnaire items on fertility preferences. We then describe the data and the models used in the

[1] In Italy some studies have shown that part of the childlessness may be due to the explicit choice not to have a baby, in order to remain free from the commitments required by children (Tanturri and Mencarini, 2004).

multi-level analysis, and finally we give the results and discuss some of their implications.

## 2. Is the 'ideal family size' a useful concept? Answers and criticisms from the literature

The tradition to ask about future childbearing preferences in fertility surveys has often been debated in the literature. The discussion has mainly focused on the measurement problems related to the items on ideal family size, such as validity and reliability of different measures – ideal, desired, intended or expected fertility –, the most appropriate question wording, the possibility of comparing them across countries, the treatment of non-response cases and of the "don't know" answers, and the usefulness itself of such questions for measuring future actual childbearing.

Many of these objections proceed from the consideration that ideal family sizes, showing little variation around the two-child norm across countries and generations, are not really responsive to changes of actual fertility levels. Indeed, the vast majority of interviewed people have been consistently responding that they would ideally like to have at least two children, despite the relevant declines in period fertility – which has gone well below replacement level in many European countries during the last thirty years.

The significance of fertility preferences is compromised even further by the ongoing trends towards postponing childbearing, especially as stated by young women at the beginning of their reproductive career.

Several concepts can be used to measure fertility preferences: ideal number of children, desired number of children, expected number of children (in addition to those already born), intended number of children.

The main objections relate to the real concept of "ideal family size": the abstract character (Toulemon 2001) of the indicator aimed at measuring not well-defined concepts such as reproductive goals, the nature of not very demography specific/oriented questions that do not make any reference to the childbearing already experienced (van de Kaa 2001), the lack of sensitivity to the fluctuations of actual fertility (Livi Bacci and De Santis 2001).

The fertility ideal has been interpreted as reflecting the normative context, while the desires for children have been considered as a personal norm. The most accurate measure to predict fertility has been found in the ultimately expected number of children, that is the sum of the children already born plus the children expected for the future (usual item: do you expect to have a(nother) child? If yes, how many?). This variable is more accurate because it takes into account constraints on childbearing that may be encountered in implementing desires. Although more valuable in the perspective of forecasting demographic behaviour (van de Kaa 2001), the number of children ultimately expected has proved insufficient to overcome measurement difficulties. This indicator, has indeed shown only little variation across countries and generations in the Fertility and Family Surveys data, turning out to be not very responsive to cross-national and generational differences.

The search for more suitable concepts of measuring preferred family size is still open among demographers.

New question items on desired and/or wanted fertility now often refer to a given time window as suggested by Miller and Pasta (1995), such kinds of question wordings can be found, for example, in the Bulgarian and French surveys[2]. As the normal attitude of the people is 'to treat the future as open', there would be a natural tendency to overestimate one's own future fertility by answering, "Yes, I want a child" if the questionnaire items do not explicitly state a specific reference time. Another option is to ask people about their evaluation (or perceived likelihood) of having a child, as, for example, in the General Household Survey conducted in Great Britain or, again, in the French survey.

[2] In Bulgaria a survey on "Young people, partnership, marriage, children" has been carried out in 2002. In France a longitudinal survey ("Enquête permanente sur les conditions de vie des ménages") has been conducted in the period 1998-2003.

Ideal fertility has also been codified with more specific questions, such as the “situated” ideal number of children, which makes explicit reference to the family size ideal of people in the same social group and with the same living standard of the respondents (Toulemon 2001). The question has been asked together with other items concerning ideal childbearing ages (van Nimwegen et al. 2002), since the timing of fertility is as important as the quantum.

The persistence of a discrepancy between actual reproductive behaviour and fertility preferences, whatever indicator being used, has given great relevance to studies looking into explanations of such a gap (Van Peer 2002).

Longitudinal studies have been carried out in order to monitor changing preferences over an individual’s life cycle. Some of these follow-up surveys, e.g., those carried out in Greece and Italy, have revealed high consistency between expected and achieved fertility, once controlled for several biographical variables (Symeonidou 2000; Menniti 2001). The consistency becomes particularly high as women move through their reproductive years. This evidence is also documented in the Family and Fertility data (van de Kaa 2001). The reasons for this trend are still not clear: it may be due to the fact that older women are more familiar with the obstacles to childbearing, such as the costs of rearing children and competing demands from jobs, divorce, medical problems and so forth; or simply that they are making an ex-post rationalisation of their expected number of children, an adjustment typical for couples beyond the fertile age.

A broad framework reconciling the inconsistency between intended and actual fertility has been provided by the sequential-conditional nature of the decision process to have a child (Namboodiri 1983). As stated in the rational choice theories approach (Yamaguchi and Ferguson 1995), fertility intentions may change after each new birth and are not taken only once for the whole reproductive life (Monnier 1987; Miller and Pasta 1995). The number of children a person wants may be constantly under reconsideration in response to changes in economic prospects and other important factors, such as the marital relationship (Ruokolainen and Notkola 2001) and partners' preferences combination process (Thomson 1997; Voas 2003). But even without any external event, people might simply change their mind and revise their preferences upwards or downwards. In this perspective, a discrepancy does not have to be necessarily viewed as an unmet demand for children (Smallwoods and Jefferies 2003), as other authors have argued (Chesanis 2000). The mechanism of translating intentions into behaviour becomes more and more complex in the presence of an increasing individualism that may erode all normative criteria in the decision-making process in favour of individual initiatives (Liefbroer 1999). In this case, a higher consistency between intended and achieved fertility would be reached if subjects were able to reach a high level of accuracy in their rational choices.

The relationship between attitudes and behaviours works reciprocally as in an interaction process: opinions, preferences and intentions influence the actual number of children and actual childbearing in turn influences preferences and intentions (Van Peer 2002). This is particularly true for people who do not yet have children and cannot easily anticipate the experience linked to the status of 'being a parent' (McMahon 1995; Wolf 2002).

Answering to the question posed at the beginning of the section we can assume that personal ideal family size is not a useful concept in predicting fertility or measuring the demand for children that is highly related to constraints and trade-offs[3]. It merely reflects personal values and attitudes toward childbearing. These attitudes are related to actual fertility without taking into account other factors, such as educational level, labour market and family experiences (Barber, Axinn and Thornton 2002). But the features relevant for childbearing ideals may also affect actual fertility (Kohler 2001; National Research Council 2001; van de Kaa 2001). Family size ideals have been considered as an upper bound of fertility, ideals being

[3] But even the predictive power of desire or intention remains to be proved.

usually larger than desires, and desires being usually larger than actual fertility (Van Peer 2002). In this sense, they can be indicative of future fertility as long as they are lower than the achieved parity. If the normative ideal of the "two-child family" corresponds to the stated fertility intentions, the consistency between fertility ideals and actual reproductive behaviour is more likely (Quesnel-Vallée and Morgan. 2003). Although more stable over different ages as compared to fertility intentions, ideals may also be subject to changes due to life course events and experiences and they are not decided simply once in the early stage of the reproductive career.

## 3. DATA

We use data from the Eurobarometer sample 2001, designed for comparative analysis among national populations. The stratified sampling procedure assures nearly equal probability samples of about 1,000 respondents in each of the 15 EU nations before the recent expansion. The sample size allows equally precise estimates for small and large countries, as well as comparisons between sub-groups broken down by sex, age, educational attainment, marital status and so on. The survey used a single uniform questionnaire design, with particular attention being paid to equivalent question wording across languages.

We restrict the analysis to individuals aged 20-39, as they are the group most involved in the reproductive process. This broad age group includes people who presumably have largely finished childbearing and those who have not yet begun or are still in the middle of their childbearing career, but it ensures a sufficient sample size[4]. The selected sample includes

[4] A rough comparison between the Eurobarometer data and various national statistical sources may raise doubts on the presence of possible (downward) biases in the ideal fertility levels of the first source of data. However, there are no reasons to believe that the possible biases vary from region to region. Moreover, in the other two international statistical sources containing information on ideal family size – i.e., the World Value Survey and the Family and Fertility Survey – the ranking of the countries is very close to that of Eurobarometer 2001. Thus, since our analysis is mainly aimed at a comparison among countries, the possible biases in the

5302 individuals nested in 72 regions belonging to 16 countries (as Germany is kept divided in West and East here). The hierarchical structure of the data is described in Table 1.

As discussed later, the models adopted here are formally based on two levels, namely individuals and regions (referred to as "clusters"). To summarise, the hierarchical structure consists of 72 clusters, with minimum and maximum clusters sizes equal to 10 and 220, respectively (Table 1). The unbalanced structure is not a problem, as it is efficiently handled by maximum likelihood methods. The number of clusters and their sizes are sufficient to achieve high power and good accuracy of the asymptotic distributions of the estimators (Snijders and Bosker 1999; Maas and Hox, 2004).

**Table 1** Respondents, regions and countries. Ages 20 -- 39. EU-15.

| COUNTRIES | N. REGIONS | N. RESPONDENTS | RESPONDENTS IN REGIONS | |
|---|---|---|---|---|
| | | | Minimum | Maximum |
| Belgium | 3 | 319 | 39 | 181 |
| Denmark | 4 | 357 | 22 | 170 |
| W. Germany | 4 | 278 | 27 | 111 |
| Greece | 3 | 380 | 45 | 220 |
| Italy | 5 | 331 | 34 | 94 |
| Spain | 7 | 340 | 12 | 85 |
| France | 5 | 419 | 48 | 105 |
| Ireland | 3 | 308 | 38 | 192 |
| Luxembourg | 2 | 215 | 53 | 162 |
| Netherlands | 4 | 343 | 30 | 160 |
| Portugal | 4 | 310 | 28 | 125 |
| U. Kingdom | 7 | 475 | 15 | 166 |
| E. Germany | 4 | 295 | 26 | 124 |
| Finland | 5 | 321 | 34 | 116 |
| Sweden | 6 | 305 | 10 | 99 |
| Austria | 6 | 306 | 36 | 71 |
| TOTAL | 72 | 5302 | 10 | 220 |

---

Eurobarometer data should be of little consequence. A deep comparison of the Eurobarometer data with data coming from several national statistical sources would be on our agenda, once the question wording adopted in the different national sources would be the same, allowing for a European comparative analysis.

### 3.1 Measure of ideal fertility

The response variable used in the multivariate analysis is the ideal number of children, which is surveyed through the following item: "And for you personally, what would be the ideal number of children you would like to have or would have liked to have had?" [5]. It has been stressed that such a wording does not include the term "family", and this might be crucial in determining the stated childbearing ideals, at least in the most traditional countries. This objection may be countered by pointing out that the question on personal ideals comes immediately after a previous item – to which it is strictly linked with an "and" – worded as follows: "Generally speaking, what do you think is the ideal number of children for a family?" The two questions allow us to distinguish between a societal and a personal ideal family size. We consider the second variable, as it is the most direct measure of the respondent's attitude towards fertility. To simplify the analysis, values greater than or equal to 3, in the light of their low frequency, are collapsed into a single category. The distribution of the response variable used in the multilevel analysis is shown in Table 2.

The possibility of comparing the ideals throughout countries relies on the similar question wordings adopted in the Eurobarometer items on personal ideal family size. This comparability is strengthened by the similar definition and connotation of the term 'ideal' in the German, English, as well as in all the Roman languages.

Some individuals in the sample (8%) did not report the ideal family size, so their response is coded as "don't know". A missing answer may be symptomatic of certain family ideals. However, given the absence of reliable auxiliary information on this point and in order to avoid relevant complications, all individuals who did not report any ideal family size are simply excluded from the analysis. In this way the results obtained from the

[5] Henceforth, we refer to this concept of ideal family size whatever term is used to indicate the fertility preferences (desires, wishes, etc.).

respondents are valid under the standard missing at random (MAR) assumption (Little and Rubin 2002).

**Table 2** Personal ideal number of children for people aged 20-39

| IDEAL FAMILY SIZE | FREQUENCY | PERCENTAGE | CUMULATE |
|---|---|---|---|
| No child | 370 | 6.98 | 6.98 |
| One child | 621 | 11.71 | 18.69 |
| Two children | 2764 | 52.13 | 70.82 |
| Three or more children | 1547 | 29.18 | 100.0 |
| TOTAL | 5302 | 100.0 | |

## 4. METHODS

### 4.1 Models

The multilevel analysis described in the following relies on the random intercept version of two well-known models, namely the logit model for binary responses and the proportional odds model for ordinal responses (e.g. Agresti 2002). In the models presented below $Y_{ij}$ denotes the response variable of individual $i$ of cluster (i.e., region) $j$ ($i=1,\dots,n_j$, $j=1,\dots,J$) and $\mathbf{x}_{ij}$ is the corresponding vector of covariates, including both individual-level and cluster-level variables. Moreover, $u_j$ denotes the cluster-level error term, also called random effect. Throughout the analysis we make the standard assumptions on random effects, namely: (i) the random effects are independent and identically distributed following a normal distribution with zero mean and an unknown, estimable variance $\sigma_u^2$; (ii) the random effects are independent of the covariates[6].

[6] The assumption that the random effects are independent of the covariates is analogous to the independence assumption on the error terms usually made in standard linear regression. However it should be noted that the independence assumption concerning the random effects is not as stringent as it may appear, since Snijders and Bosker (1999) show that if the random effects are correlated with an individual-level variable, such correlation is removed as soon as the cluster mean of such variable is introduced as a further covariate.

When the response variable is binary, taking the values 0 or 1, one can define $\pi_{ij} = P(Y_{ij} = 1 \mid u_j)$ and adopt the random intercept *logit* model:

$$\log\left(\frac{\pi_{ij}}{1-\pi_{ij}}\right) = \alpha + \boldsymbol{\beta}'\mathbf{x}_{ij} + u_j\,,$$

where $\alpha$ is the intercept and $\boldsymbol{\beta}$ is the vector of regression coefficients.

When the response variable is ordinal, taking the values $1, 2, \ldots, M$, one can define $\gamma_{ij}^{(m)} = P(Y_{ij} \leq m \mid u_j)$ and adopt the random intercept *proportional odds* model, which can be viewed as a set of linear models for the *M*-1 cumulative logits:

$$\log\left(\frac{\gamma_{ij}^{(m)}}{1-\gamma_{ij}^{(m)}}\right) = \tau^{(m)} - \left(\alpha + \boldsymbol{\beta}'\mathbf{x}_{ij} + u_j\right) \qquad m = 1, \ldots, M-1\,,$$

where $\alpha$ is the intercept, $\boldsymbol{\beta}$ is the vector of regression coefficients and $\tau^{(m)}$ are the cutpoint parameters. The cutpoints must be ordered, $\tau^{(1)} \leq \tau^{(2)} \ldots \leq \tau^{(M-1)}$, and the first cutpoint, $\tau^{(1)}$, is fixed to zero for identifiability reasons. The minus sign preceding the linear predictor is necessary in order to interpret the effects of the covariates in the more natural way (i.e., a positive regression coefficient means that higher values of the covariate tend to yield higher values of the response variable).

The assumption that the vector of regression coefficients $\boldsymbol{\beta}$ is constant for all the *M*-1 cumulative logits, sometimes called the *parallel regression assumption*, leads to the *proportional odds* property, i.e., the ratio of the odds of two individuals does not depend on the category. The parallel regression assumption is very convenient as for parsimony and interpretation, and can be checked using, for instance, the test developed by Brant (1990).

Since the individual-level variance implied by the logit link is $\pi^2/3$, the intraclass correlation coefficient is $\sigma_u^2 / \left(\sigma_u^2 + \pi^2/3\right)$ for both the logit model and the proportional odds model (Snijders and Bosker 1999).

The present analysis concerns the ideal number of children, which is an ordinal variable with values 0, 1, 2 and 3 or more (as values over 3 are collapsed for convenience). Therefore a natural choice would be to study such a variable by means of a proportional odds model. However, the no-child ideal is somewhat different from the other ideals since the desire to have no children may be determined by specific factors, and consequently, may be particularly informative of the attitude towards childbearing (Westoff 1990). In many countries it is simply the result of repeated postponements, rather than a deliberate desire to remain childless. Moreover, the social pressure may have a great impact on the extreme alternative between families with children and families without children. The huge cross-country variation in the proportion of those preferring a 'zero-child' family may be linked to different levels of this social pressure in the countries under analysis.

From a statistical point of view the peculiarity of the no-child ideal leads to a violation of the parallel regression assumption. In fact, the Brant test on the proportional odds model (without random effects) for the ideal number of children including the "zero" rejects the parallel regression assumption, implying that the violation is mainly due to the "zero" category. The model could be extended to handle partial proportional odds (Terza 1985; Peterson and Harrel 1990), but then the interpretation becomes somewhat tortuous.

The theoretical and statistical issues just mentioned induced us to perform two separate analyses: first, an analysis of the at-least-one-child ideal versus the no-child ideal, fitting on the whole sample a logit model based on the following response variable:

$$Y_{ij} = \begin{cases} 0 & \text{if } INC_{ij} = 0 \\ 1 & \text{if } INC_{ij} > 0 \end{cases}$$

where $INC_{ij}$ is the ideal number of children reported by individual $i$ of cluster $j$.

The second analysis concerns the ideal number of children, but only for those desiring at least one child. To this end we fit a proportional odds model on the sub-sample of those desiring at least one child ($INC_{ij}$>0) using the following response variable:

$$Y_{ij} = \begin{cases} 1 & \text{if } INC_{ij} = 1 \\ 2 & \text{if } INC_{ij} = 2 \\ 3 & \text{if } INC_{ij} \geq 3 \end{cases}$$

In this case the parallel regression assumption passes the Brant test.

In the application the two models are developed independently, since the determinants for having the ideal of a family with children may be rather different from the factors driving the preference for a given number of children.

## 4.2 Multilevel analysis

The clustering of the individuals in regions is considered as a phenomenon of interest rather than a mere disturbance. Therefore, we adopt a multilevel approach (Snijders and Bosker 1999), explicitly considering regional-level variables and regional means of individual variables and adding the regional random effects in the predictor of both models. Random coefficients and cross-level interactions are also considered in the model selection process.

In the present investigation the data allow us to set up a multilevel model where individuals are nested in regions. The regional context has been defined as a ‘meso-level’, a link between macro social structures and micro-demographic behaviors. In principle, a third level represented by the countries could be added, though in the present case the limited number of countries also limits reliable inferences (Maas and Hox 2004). Anyway, the residual variability at country level turns out to be modest and not significant, so the country level can be ignored in the model specification.

There is an extensive literature on the advantages of explicitly considering the influence of broader social contexts on the attitudes and behaviours of individuals by making use of multilevel models (Di Prete and Forristal 1994; Snijders and Bosker 1999; Guo and Zhao 2000; Teachman and Crowder 2002; Gauthier 2002; Courgeau 2003; Festy 2003). In general, a multilevel model attempts to represent the complex causal process underlying the behaviour of individuals living in a social context, allowing to draw valid inferences on the relationships at the relevant hierarchical levels.

There are also many challenges and concerns linked to multilevel modelling, one of the most relevant being the risk of “contextual fallacy”: the statistically significant effects of aggregate-level variables may merely be the result of a poor specification of the individual-level relationships. Particular caution is required when choosing the variables and it is recommended to think well about the meaning of contextual effects (Hauser 1974).

Regional and national differences are analysed by also looking at the predicted random effects (Empirical Bayes residuals) of the two models. In fact the random effects convey all the unobserved factors at regional level, so regions with high positive or negative residuals have fertility ideals, which are somewhat unexpected in the light of the covariates. In order to identify possible outliers, residuals at regional level are standardised with the sampling standard deviation, so that values outside +/-2 should be regarded as anomalous. Then the country means of the predicted random effects are computed in order to look at cross-national differences. Obviously, when taking the country means, high and low residuals tend to counterbalance each other, so it is unlikely to find country-level residuals exceeding the +/-2 interval.

### 4.3 Individual-level covariates

As for the individual covariates, several background characteristics, aspects of life history and structural constraints are considered in the analysis. All covariates are referred to the time of interview. Unfortunately, the data do not carry any retrospective information concerning the previous history of the respondents, which could allow us to estimate the role of biographical trajectories on the process of forming family size ideals in a dynamic framework.

The same covariates are tried in the two separate models formerly outlined: the logit model for the at-least-one-child ideal versus the no-child ideal, and the proportional odds model for the ideal number of children for those desiring at least one child. Tables 3A and 3B report a brief description of the covariates that turned out to be significant in either model. Other covariates are generated in the model selection process as interaction terms, or as quadratic and cubic terms of continuous covariates.

Individual covariates having a significant effect are: age, sex, marital status, school enrolment, educational attainment (in years of schooling), employment status, household size and household income. Urban level (rural area, middle town, large town) and political opinion (left, centre, right) are not significant.

The *age* of the respondents and the *educational level* are the only two continuous covariates. Both are centered on the rounded mean value, 30 and 14 years, respectively. The educational level is defined as the total number of years of education. For those still enrolled in school, which are marked by a specific covariate, the value refers to the *current* number of years of education.

All other covariates are categorical, so they are transformed into suitable dummy variables. Some collapsing of categories is often needed: in such cases several alternative collapsing schemes are tried in the model selection process. In the following the covariates are described with the categorisation used in the final models.

Individuals with any missing values on the covariates are not excluded from the sample: rather, the missing value is first treated as a distinct category and then, as long as no relevant difference emerges, it is included in the baseline category.

The *marital status* is codified using four categories: single, married, cohabiting and separated. The category 'separated' includes also divorced persons, while the 'married' respondents are grouped together with the remarried and the widowed ones.

The *employment status* has two categories: employed persons and people not in the labour market or unemployed. A more refined breakdown of this variable (e.g., unemployed versus non active, or manual worker versus non manual worker) is not supported by the data.

The *household size* refers to the number of household members aged 15 or over, with categories 1, 2, 3 and 4 or more.

The *household income*, which is harmonised at the international level, is dichotomised: households with an income over the average are contrasted with households with an income below the average or not declared.

A special individual covariate is the *actual number of children*, which is obviously a good predictor of the *ideal* number of children. However, its inclusion in the models raises a severe problem of endogeneity: in fact, even though it is known that actual fertility influences ideal fertility, the stronger effect presumably works in the opposite direction, i.e., ideal fertility affects actual fertility. Moreover, the meaning of actual fertility varies from person to person, being dependent on the period of the childbearing career at which it is measured. Therefore the interpretation of the regression coefficient of actual fertility would be quite difficult. Moreover, since actual fertility is a very strong predictor, its inclusion in the models would drown the effects of the other covariates, as shown by our trials. For these reasons we do not include the actual number of children in the models. However, to take into account its key role in the analysis of fertility preferences, we fit the models also on the sub-sample of individuals without any children, reporting the resulting estimates along with the estimates springing from the whole sample (Tables 4 and 5)[7].

It is clear that a satisfactory analysis of the relationship between ideal and actual fertility would require good quality longitudinal data and the implementation of adequate event-history models.

### 4.4 Regional-level covariates

In order to verify the hypothesis of declining preferences in the contexts of lower actual childbearing, three regional-level covariates are added: the mean actual number of children ever born, the percentage of childless people, and the mean age at the birth of the first child. The first two covariates are calculated considering both males and females, while the third one is computed considering only the women. The three regional-level covariates are referred to the older generations, as they are computed from the Eurobarometer sample by taking the means for people aged 40-60. In this way the covariates reflect the cultural context in which the studied individuals, which are now aged 20-40, have grown up and have been socialised.

Other regional-level covariates are the compositional variables obtained by computing the regional means of the individual covariates. The inclusion of such covariates allows us to disentangle between and within

[7] We also tried the models on the complementary sub-sample of individuals with at least one child, but the results are not worth showing since this sub-sample is nonetheless quite heterogeneous. Moreover, the logit model shows some problems with this sub-sample since among those with children only very few respondents report a zero-child ideal. Presumably, there is a problem of measurement: once people become parents they change their ideal family size or they simply have difficulties to admit that they had a no-child ideal. A similar measurement problem probably affects the results of the corresponding proportional odds model as well.

regression coefficients, which in principle could even have an opposite sign. For easier interpretation the regional means of the binary covariates are multiplied by 100, i.e., they are expressed as percentage.

All the regional-level covariates are centered on the value of the Western region of West Germany, which had the greatest number of respondents.

**Table 3A** Description of the individual-level and the regional-level covariates in the sample (5032 individuals)

| VARIABLES | MEAN | STD. DEV. | MIN | MAX | DESCRIPTION |
|---|---|---|---|---|---|
| *Individual-level covariates* | | | | | |
| AGE (in years) | 29.80 | 5.66 | 20 | 39 | Age of respondents |
| SEX | | | | | |
| Female | 0.54 | 0.50 | 0 | 1 | 1=female; 0 otherwise |
| MARITAL STATUS | | | | | |
| Married | 0.43 | 0.49 | 0 | 1 | 1=married; 0=otherwise |
| Cohabiting | 0.18 | 0.39 | 0 | 1 | 1= cohabiting; 0=otherwise |
| Single | 0.34 | 0.47 | 0 | 1 | 1=unmarried; 0=otherwise |
| Divorced | 0.05 | 0.22 | 0 | 1 | 1=divorced; 0=otherwise |
| EDUCATION | | | | | |
| School enrolment | 0.12 | 0.32 | 0 | 1 | 1=attending school; 0=otherwise |
| Educational level | 13.70 | 4.19 | 2 | 33 | Number of years of education |
| EMPLOYMENT | | | | | |
| Employed | 0.69 | 0.46 | 0 | 1 | 1=employed; 0=otherwise |
| Unemployed or no active | 0.20 | 0.40 | 0 | 1 | 1=unemployed or no active; 0=otherwise |
| HOUSEHOLD SIZE | | | | | |
| One-adult | 0.21 | 0.41 | 0 | 1 | 1=1 adult; 0=otherwise |
| Two-adult | 0.54 | 0.50 | 0 | 1 | 1=2 adults; 0=otherwise |
| Three-adult | 0.13 | 0.34 | 0 | 1 | 1=3 adults; 0=otherwise |
| Four-adult (or more) | 0.12 | 0.33 | 0 | 1 | 1=4 or more adults; 0=otherwise |
| HOUSEHOLD INCOME | | | | | |
| Above median | 0.28 | 0.45 | 0 | 1 | 1=above median; 0=otherwise |
| Below median | 0.72 | 0.45 | 0 | 1 | 1=below m. & "DK"; 0=otherwise |
| *Regional-level covariates* | | | | | |
| MEAN NUMBER OF CHILDREN EVER BORN | 1.97 | 0.36 | 1 | 3.19 | Children born among cohorts aged 40 to 60 |
| PROPORTION CHILDLESS (%) | 14.30 | 5.66 | 0 | 33.34 | Childlessness among cohorts aged 40 to 60 |
| MEAN AGE AT THE BIRTH OF FIRST CHILD (years) | 24.15 | 1.23 | 21.67 | 29.67 | Mean age at the first birth among female cohorts aged 40 to 60 |

**Table 3B** Description of the individual-level and the regional-level covariates in the sub-sample of those without children (2563 individuals)

| VARIABLE | MEAN | STD. DEV. | MIN | MAX | DESCRIPTION |
|---|---|---|---|---|---|
| *Individual-level covariates* | | | | | |
| AGE (in years) | 26.98 | 5.23 | 20 | 39 | Age of respondents |
| SEX | | | | | |
| Female | 0.45 | 0.50 | 0 | 1 | 1=female; 0 otherwise |
| MARITAL STATUS | | | | | |
| Married | 0.14 | 0.34 | 0 | 1 | 1=married; 0=otherwise |
| Cohabiting | 0.22 | 0.41 | 0 | 1 | 1= cohabiting; 0=otherwise |
| Single | 0.63 | 0.48 | 0 | 1 | 1=unmarried; 0=otherwise |
| Divorced | 0.02 | 0.13 | 0 | 1 | 1=divorced; 0=otherwise |
| EDUCATION | | | | | |
| School enrolment | 0.22 | 0.41 | 0 | 1 | 1=attending school; otherwise |
| Educational level | 14.40 | 3.81 | 3 | 33 | Number of years of education |
| EMPLOYMENT | | | | | |
| Employed | 0.67 | 0.47 | 0 | 1 | 1=employed; 0=otherwise |
| Unemployed or no active | 0.11 | 0.31 | 0 | 1 | 1=unemployed or no active; 0=otherwise |
| HOUSEHOLD SIZE | | | | | |
| One-adult | 0.31 | 0.46 | 0 | 1 | 1=1 adult; 0=otherwise |
| Two-adult | 0.36 | 0.48 | 0 | 1 | 1=2 adults; 0=otherwise |
| Three-adult | 0.15 | 0.36 | 0 | 1 | 1=3 adults; 0=otherwise |
| Four-adult (or more) | 0.18 | 0.38 | 0 | 1 | 1=4 or more adults; 0=otherwise |
| HOUSEHOLD INCOME | | | | | |
| Above median | 0.33 | 0.47 | 0 | 1 | 1=above median; 0=otherwise |
| Below median | 0.67 | 0.47 | 0 | 1 | 1=below median and "DK"; 0=otherwise |
| *Regional-level covariates* | | | | | |
| MEAN NUMBER OF CHILDREN EVER BORN | 1.93 | 0.39 | 1 | 2.97 | Children born among cohorts aged 40 to 60 |
| PROPORTION CHILDLESS (%) | 15.03 | 5.87 | 4 | 30.0 | Childlessness among cohorts aged 40 to 60 |
| MEAN AGE AT THE BIRTH OF FIRST CHILD (years) | 24.17 | 1.05 | 22.57 | 26.62 | Mean age at the first birth among female cohorts aged 40 to 60 |

### 4.5 Model selection and estimation

All models are fitted by maximum likelihood using suitable Stata commands, namely logit for the single-level logit model, ologit for the single-level proportional odds model and gllamm (Rabe-Hesketh et al. 2001) for the corresponding multilevel, random effects versions. Fitting logit or proportional odds models with normally distributed random effects is computationally demanding, as it requires the numerical approximation of some intractable integrals. The gllamm command approximates the integrals through adaptive Gaussian quadrature, leading to accurate results. However, in our application the computational time for a random intercept model is about ten minutes, so in order to select the best model we rely on the single-level, standard versions, using the fast logit and ologit commands. Such a trick is quite common and yields reliable results (Agresti 2002). In the provisional single-level models the covariates, along with interactions and quadratic terms, are retained according to prudential Wald tests at a 10% level. Then a random effect on the intercept is added and the model is refined using Wald tests at a 5% level. Random effects on the regression coefficients are also tried. The significance of the variances and covariances of the random effects are not evaluated with the Wald test, but with the more reliable likelihood ratio test with corrected *p*-value (Snijders and Bosker 1999).

## 5. RESULTS

Our multilevel analysis of fertility preferences relies on two distinct models: first, a random effects logit model explaining the desire to have or not to have children; second, a random effects proportional odds model for the ideal number of children for those desiring at least one child. The model selection process begins with a common set of covariates, but then it proceeds independently for the two models, so the final models have different covariates: in fact, only the school attendance, the marital status and the household size play a role in both models.

### 5.1 Model for a family ideal with at least one child

The results from the first model (Table 4 and Figure 1) show an interesting interaction between gender and school attendance: in fact, among people who have terminated the studies, females are more likely to desire children, but among people who are still studying, females are less likely to do so. Alternatively, we can say that the effect of school attendance on the ideal of a family with at least one child is positive for males and negative for females.

The other classical demographic variable, namely age, has a significant negative effect only for singles.

As for the other covariates, the likelihood to have a family ideal with children increases for persons living in a household of four or more adults or in regions with a high percentage of people still attending school or cohabiting. A preference for families with children is less likely for people who are single, cohabiting or separated, or for individuals living in single-person households, or in regions with a high percentage of childlessness among the older generations or a high percentage of one-adult households.

It is worth to note that, among the regional-level variables concerning the older generations, actual fertility has no significant effect, while the proportion of childless people has a strong negative effect.

Figure 1 shows the predicted probabilities of desiring a family with children for some hypothetical individuals. The base individual has all the individual covariates set to the base category (age is set to 30), while the regional-level covariates are set to the value of the Western region of West Germany and the random effect is set to zero.

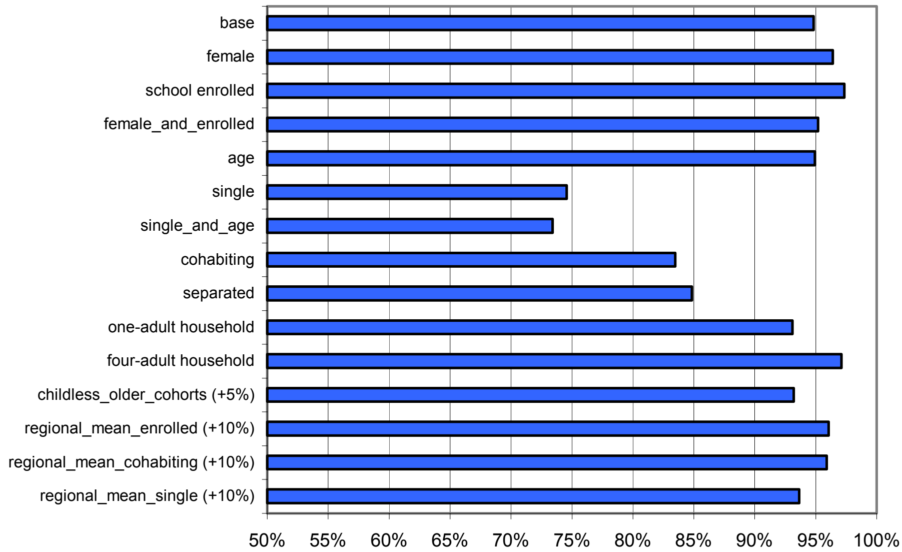
In the model fitted on those without children most of the covariates keep their magnitude and are still significant. Interestingly, the variables related to the marital status show smaller effects and are no longer significant. Moreover, for those not attending school the difference between males and females vanishes and age seems to have a negative effect in general, not only for singles.

**Table 4** Random intercept logistic models for an ideal of at least one-child family. EU-15, Year 2001.

| | All respondents | | | Respondents without children | | |
|---|---|---|---|---|---|---|
| | Beta | S.E. | P>z | Beta | S.E. | P>z |
| *Individual covariates* | | | | | | |
| Sex | | | | | | |
| (base: male) | | | | | | |
| Female | 0.38 | 0.12 | 0.00 | 0.04 | 0.14 | 0.78 |
| School enrolment | | | | | | |
| (base: not enrolled) | | | | | | |
| Enrolled | 0.69 | 0.30 | 0.02 | 0.63 | 0.30 | 0.04 |
| Enrolled*female | -0.99 | 0.36 | 0.01 | -0.71 | 0.37 | 0.05 |
| Age | | | | | | |
| (Age-30) | 0.02 | 0.02 | 0.32 | -0.06 | 0.02 | 0.01 |
| (Age-30)*single | -0.08 | 0.02 | 0.00 | -0.02 | 0.03 | 0.44 |
| Marital status | | | | | | |
| (base: married) | | | | | | |
| Cohabiting | -1.29 | 0.21 | 0.00 | -0.10 | 0.23 | 0.67 |
| Single | -1.83 | 0.20 | 0.00 | -0.34 | 0.22 | 0.12 |
| Separated | -1.19 | 0.29 | 0.00 | -0.56 | 0.42 | 0.18 |
| Household size | | | | | | |
| (base: Two or three adults) | | | | | | |
| One adult | -0.31 | 0.16 | 0.06 | -0.22 | 0.18 | 0.21 |
| Four or more adults | 0.61 | 0.24 | 0.01 | 0.64 | 0.26 | 0.01 |
| *Regional-level covariates (a)* | | | | | | |
| Percentage of childless | -0.06 | 0.02 | 0.00 | -0.06 | 0.02 | 0.00 |
| *Regional means (b)* | | | | | | |
| 'Enrolled' | 0.03 | 0.01 | 0.01 | 0.02 | 0.01 | 0.01 |
| 'Cohabiting' | 0.02 | 0.01 | 0.01 | 0.00 | 0.01 | 0.99 |
| 'One-adult household' | -0.02 | 0.01 | 0.01 | -0.01 | 0.01 | 0.01 |
| Constant | 2.91 | 0.31 | 0.00 | 1.63 | 0.34 | 0.00 |
| Regional-level variance | 0.19 | 0.09 | | 0.22 | 0.10 | |
| Y==1 | 4932 | 93.0% | | 2223 | 86.7% | |
| level-1 units | 5302 | | | 2563 | | |
| level-2 units | 72 | | | 72 | | |
| Log-Likelihood | -1154.52 | | | -898.98 | | |

Note. (a) Computed on the individuals aged 40 to 60 and centered on the value of the Western region of West Germany; (b) The regional mean covariates are multiplied by 100.

**Figure 1** Predicted probability for an ideal family size with at least one child. European Union. All respondents aged 20 to 39. EU-15. Year 2001.




### 5.2 Model for the ideal number of children for those desiring at least one child

The results from the second model (Table 5 and Figure 2) show that the propensity toward larger families for those desiring children decreases for people who are single, cohabiting or separated, or for individuals with an income above the median. In contrast, the ideal family size is higher for individuals who are still attending school or not employed, or live in households of four or more adults. The effect of the educational level, as measured by years of schooling, is not at individual-level but entirely at cluster level: living in a region with more educated persons increases the probability of desiring more children, though the personal educational level is not relevant in this respect. Finally, in contrast to the previous model, the percentage of childless people of the older generations has no significant effect, while the actual fertility of the older generations has a strong positive effect. Also the average age at birth of the first child has a positive effect, which is somewhat unexpected and difficult to interpret.

Figure 2 shows the predicted probabilities of desiring a family with one child, two children and more than two children for some hypothetical individuals. The base individual has all the individual covariates set to the base category, while the regional-level covariates are set to the value of the Western region of West Germany and the random effect is set to zero.

In the model estimated on the sub-sample of those without children the regional-level variables remain fairly stable in magnitude and they are still highly significant, while the individual-level variables, with the only exception of the school enrolment, show weaker effects and lose their significance. We are particularly puzzled by the fact that the variable marital status becomes irrelevant in the group without children.

Considering again the model fitted on all the people desiring at least one child without distinguishing the actual number of children, the effect of older generations' actual fertility is quadratic and thus difficult to interpret. Therefore it is useful to draw some graphs (Figure 3 and Figure 4) showing how the predicted probabilities for the base individual depend on the mean actual number of children ever born among the older generations. Since in the last decades the total fertility has shown a decreasing trend, the graphs are better understood if read from right to left. At high levels of actual fertility the probability to prefer one-child families is rather stable, while the preference for large families (3 or more children) declines and the likelihood of an ideal of exactly 2 children is first stable and then increasing, capturing those leaving the ideal of large families. At below replacement fertility levels the likelihood of two-child families continues to increase, while preferences for large families decline more steeply and the probability of a one-child ideal also starts to increase. Only at very low levels of actual fertility, marked in the tables with the vertical line at 1.5 children, the two-child norm starts to decrease and the ideal of the one-child family becomes more and more likely. The cumulated probabilities (Figure 4) tell us a similar story: the ideal of two or more children is strongly constant and declines only when the regional actual fertility reaches low (<2) and very low (<1.5) levels. The picture described does not change much when a different base individual is chosen, in this case a shift of the curves up or down is observed, but their shape and their general pattern do not change. Although the trends have to be interpreted in a cross-section framework (as they reflect the effects of different regional fertility contexts, provided by the reproductive experience of the older generations, and affecting the individual fertility preferences of younger people), they can shed some light on the reasons why the two-child norm has been so stable in the last few decades.

For each model, the standardised residuals at regional level are useful for pointing out the atypical regions, i.e., those where the fertility ideals are rather different from what is predicted by the covariates. Figure 5 reports the country means of such residuals, showing that West Germany, East Germany and Austria are the only countries whose average residual is negative for both models, revealing the presence of unobserved factors that reduce the family ideals. However, considering the residuals of the proportional odds model for the desired number of children, it appears that the average residual of West Germany is not as extreme as that of East Germany or Austria. Therefore, the regional-level covariates introduced in the model to a considerable extent account for the exceptionally low ideal fertility level of West Germany. The propensity for lower fertility preferences in the German speaking countries does not change if the models are run on the group of those without children: Austria and Germany (both East and West) are still the only countries showing negative residuals in both models and now Austria becomes the champion in the unobserved factors encouraging the preference for families without children.

The Netherlands and Spain also takes atypical positions. The Netherlands has strong unobserved factors which increase the number of those not desiring children, while, given the model covariates, it is in line with other countries as for the desired offspring of those having a family ideal with children. Spain on the other hand has strong unobserved factors increasing the number of those desiring smaller families while, given the model covariates, it is in line with other countries as for the desire of a family ideal with children. An inspection of the regional residuals shows that these results are due to the Eastern and Southern parts for the Netherlands and to the Central-Southern regions for Spain (see Appendix).

**Table 5** Random intercept proportional odds model for the ideal number of children for those desiring at least one child. EU-15, Year 2001.

| | All respondents | | | Respondents without children | | |
|---|---|---|---|---|---|---|
| | Beta | S.E. | P>z | Beta | S.E. | P>z |
| *Individual covariates* | | | | | | |
| School enrolment | | | | | | |
| (base: not enrolled) | | | | | | |
| Enrolled | 0. 57 | 0.10 | 0.00 | 0.65 | 0.11 | 0.00 |
| Marital status | | | | | | |
| (base: married) | | | | | | |
| Cohabiting | -0.42 | 0.08 | 0.00 | -0.04 | 0.15 | 0.77 |
| Single | -0.42 | 0.07 | 0.00 | -0.01 | 0.13 | 0.93 |
| Separated | -0.24 | 0.13 | 0.07 | -0.21 | 0.37 | 0.58 |
| Employment status | | | | | | |
| (base: employed) | | | | | | |
| Unemployed or no active | 0.34 | 0.07 | 0.00 | 0.11 | 0.14 | 0.44 |
| Household size | | | | | | |
| (base: less than 4 adults) | | | | | | |
| Four-adult household | 0.28 | 0.09 | 0.00 | 0.09 | 0.12 | 0.44 |
| Household income | | | | | | |
| (base: below median) | | | | | | |
| Above median | -0.14 | 0.07 | 0.04 | -0.03 | 0.10 | 0.77 |
| *Regional-level covariates (a)* | | | | | | |
| Mean number of children | 1.89 | 0.34 | 0.00 | 1.49 | 0.44 | 0.00 |
| Mean number of children ^2 | -0.63 | 0.27 | 0.02 | -0.27 | 0.35 | 0.44 |
| Average age at birth of the first child | 0.15 | 0.05 | 0.00 | 0.13 | 0.03 | 0.04 |
| *Regional means* | 0.15 | 0.03 | 0.00 | 0.15 | 0.05 | 0.01 |
| Educational level | | | | | | |
| First cutpoint | -1.49 | 0.11 | 0.00 | -1.13 | 0.17 | 0.00 |
| Second cutpoint | 1.46 | 0.11 | 0.00 | 1.89 | 0.18 | 0.00 |
| Regional-level variance | 0.14 | 0.04 | | 0.18 | 0.06 | |
| Y==1 | 621 | 12.6% | | 310 | 14.0% | |
| Y==2 | 2764 | 56.0% | | 1290 | 58.0% | |
| Y==3 | 1547 | 31.4% | | 623 | 28.0% | |
| level-1 units | 4932 | | | 2223 | | |
| level-2 units | 72 | | | 72 | | |
| Log-likelihood | -4474 | | | -2015 | | |

Note. (a) Computed on the individuals aged 40 to 60, and centered on the value of the Western region of West Germany.

**Figure 2** Predicted probability for a given ideal family size for those desiring at least one child. All respondents aged 20 to 39 and desiring at least one child. EU-15. Year 2001.

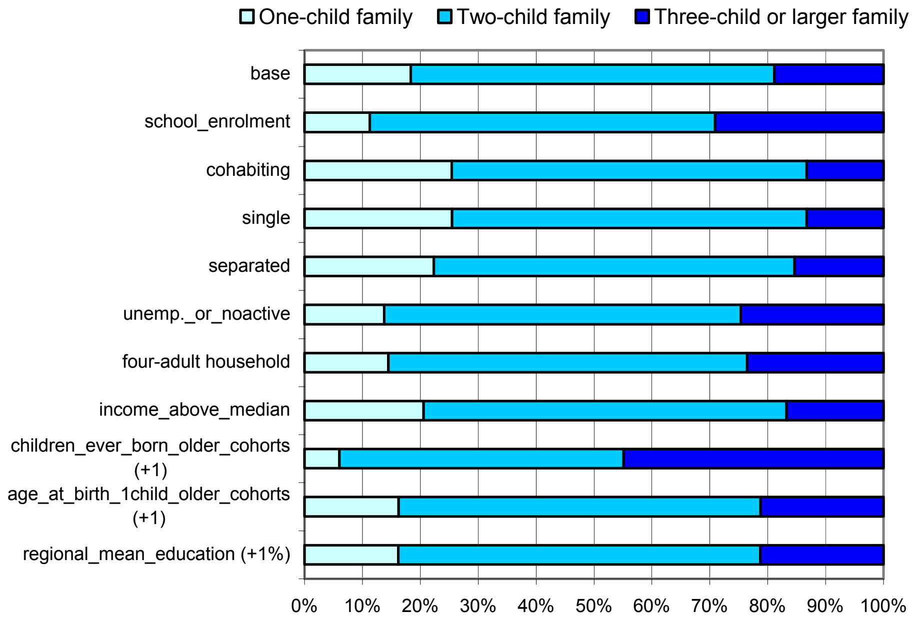
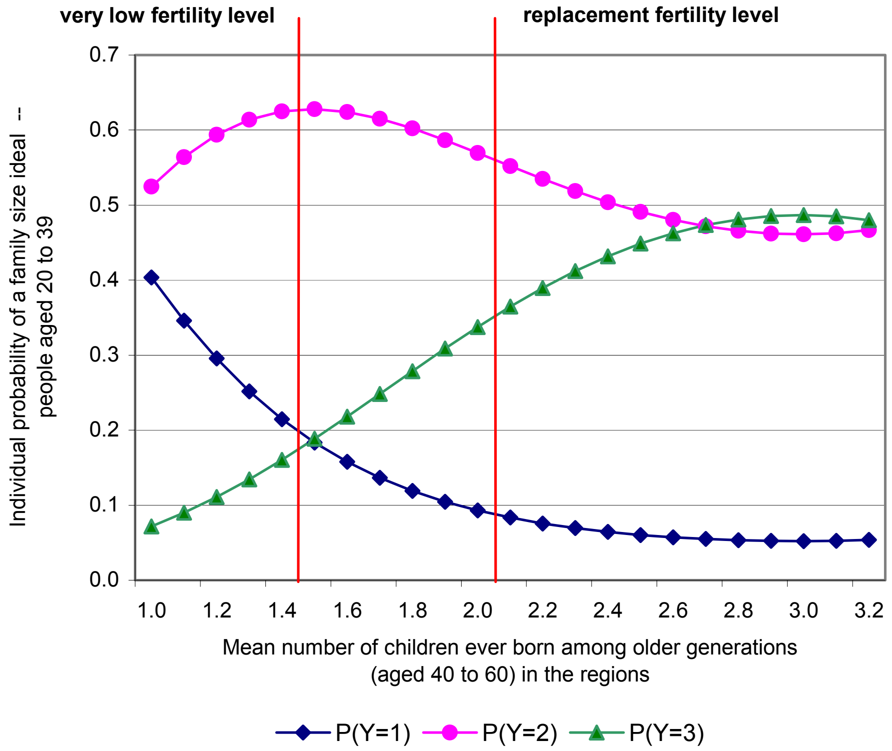
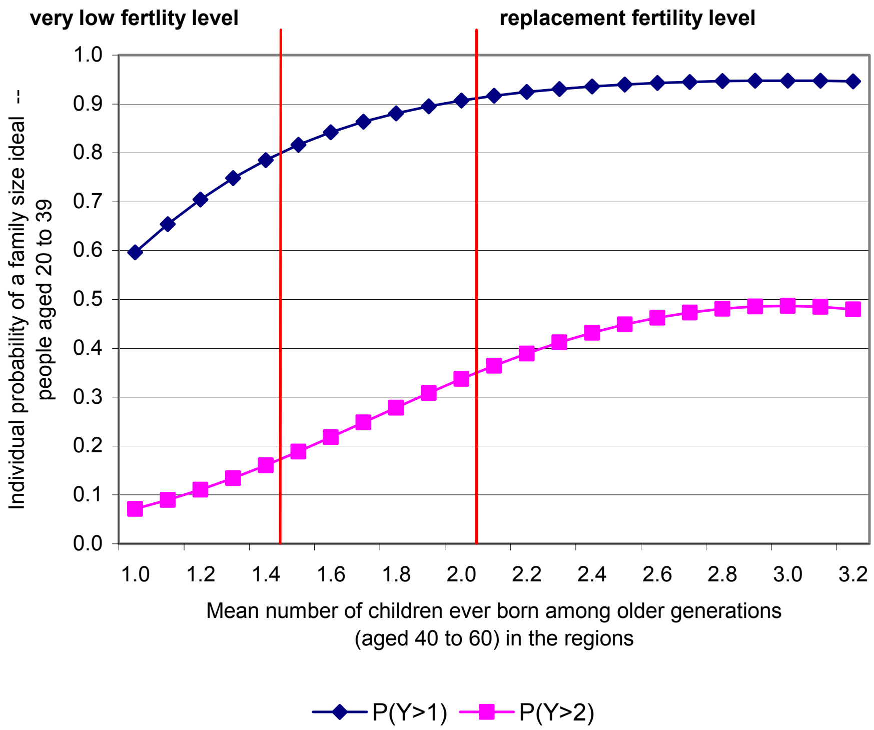


**Figure 3** Effect of the mean number of children ever born among older generations on the younger generations' individual probability of a given ideal family size. All respondents aged 20 to 39 and desiring at least one child. EU-15. Year 2001.




Note. Probabilities computed for the base individual (all the individual covariates are set to the base category, while the regional-level covariates are set to the value of the Western region of West Germany and the random effect is set to zero). The two vertical lines denote the replacement fertility level (2.1 children) and the very low fertility level fixed at 1.5 children.

**Figure 4** Effect of the mean number of children ever born among older generations on the younger generations' individual cumulated probability of a given ideal family size. All respondents aged 20 to 39 and desiring at least one child. EU-15. Year 2001.




Note. Probabilities computed for the base individual (all the individual covariates are set to the base category, while the regional-level covariates are set to the value of the Western region of West Germany and the random effect is set to zero). The two vertical lines denote the replacement fertility level (2.1 children) and the very low fertility level fixed at 1.5 children.

## 6. DISCUSSION

As evidenced in our analysis, Germany and Austria have the lowest fertility preferences in Europe (EU-15) even after controlling for different individual and contextual features. For both countries the driving force of low fertility ideals is the propensity for a smaller number of children rather than the preference for families without children. However, there are differences between the countries: while in West Germany the context of actual childbearing to a good extent explains individual family size ideals, in Austria and East Germany there are some unobserved factors that cause the ideals for families with children to be so low (Figure 5, A). Generally, the no-child ideal is not more common in the German speaking countries than in the other European countries, there is much more positive attitude towards childlessness in the Netherlands, for example, and Austrians become the champion in the preferences for no-child families only if we restrict the analysis to persons without children (Figure 5, B). The results are consistent with those coming from a recent survey carried out among young Austrians aged 16-24 among whom only 27% answered that having children is "very important" for a couple (Friesl 2001).

The hypothesis of a German bipolarisation between the family and the no-family sector, due to difficulties in combining parenthood and work that would make individuals decide to either not have any child at all, or in favour of a two-child family (Dorbritz and Hoehn 2000), is not verified for the fertility preferences.

Interestingly, the regional context of actual childbearing, as measured by the mean actual family size and by the proportion of childless women among the older generations living in the region, exerts a strong impact on fertility preferences of the young people, which largely explain the between-regions differences. The hypothesis that changing fertility ideals lag behind the changes in the actual reproductive behaviours (Lee 1980; Goldstein et al. 2003) is thus supported and interpreted in an integrated micro-macro framework, where the social context plays a major role.

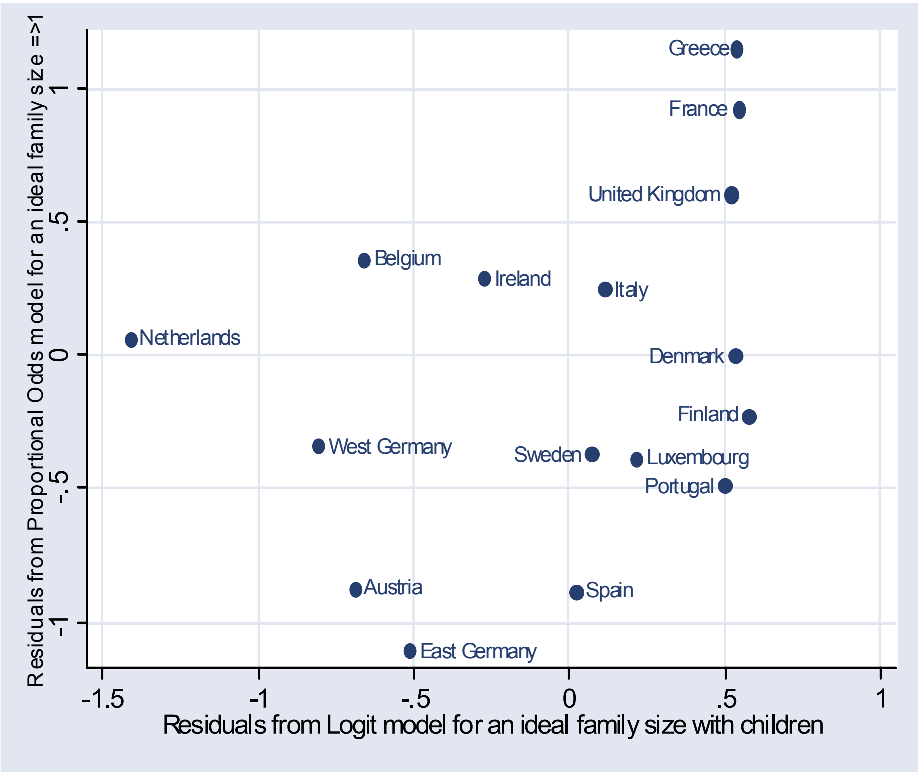
Several mechanisms may affect the relationship between contextual childbearing features and individual behaviour, one of these being collective socialisation, or social control: adults in a given neighbourhood influence young people who are not their own children (Jencks and Mayer 1990). This mechanism supports our hypothesis of a social learning process between young and old people that does not necessarily have to go through the children-parents relationship. Of course, the family context in which people have grown up may take a very important role in this process, as more prolific parents normally correspond to more prolific children (Murphy 1999).

There are other possible causal mechanisms in which social contexts may affect individual behaviour: diffusion and reproduction of cultural patterns, opportunity structure, and destination of selective migration (Nauck 1995). This last cause raises the topic of potential endogeneity of the contextual effects: unobserved characteristics of individuals might influence the phenomenon of interest, e.g., fertility preference, as well as the choice of the environment to live in, e.g., the region. In this case people with high preference for children should tend to move into regions with opportunity structures that they perceive to be favourable to family formation and socialisation of children and vice versa. However, we can assume that in our analysis the endogeneity is not that serious, as the relationship between contextual fertility and personal fertility preferences works through a generational lag, and it would be hard to assume individuals being consciously and deliberately attracted towards contexts where the actual childbearing of the older generation is closer to their own family size ideals.

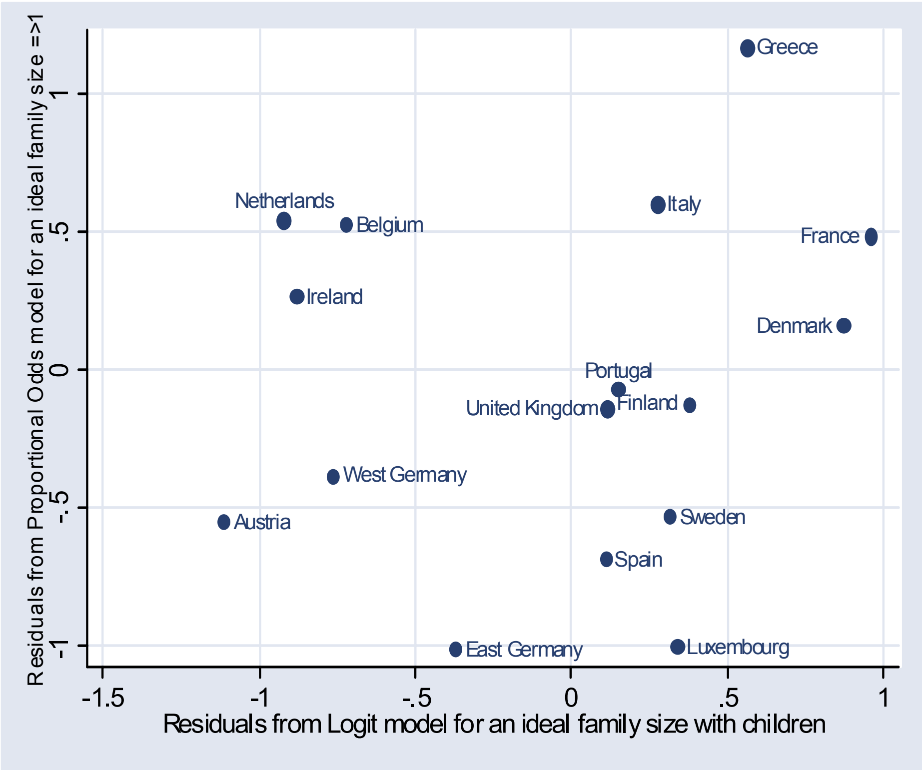
Of course, the influence of the context is not limited to the childbearing experience of the older generations, and, it would be interesting to enlarge the analysis including other non-demographic context factors as well as the influence of the childbearing experience or ideal exerted by the *peers*, that can counterbalance the pressure coming from the older cohorts and might be very relevant in the diffusion of new demographic behaviours (Nazio and Blossfeld 2003). Unfortunately, we do not have enough information in our data set for this, but the topic is extremely interesting and would certainly deserve further investigation once the data become available.

**Figure 5** Country means of standardized predicted random effects (Empirical Bayes residuals). EU-15. Year 2001.

A. All respondents



B. Only respondents without children




## 7. CONCLUSION

We propose two different models in order to investigate the factors affecting family size ideals in Europe. First, we implement a multilevel logistic model for a family ideal with children. Second, we perform a proportional odds model for a given number of children for those desiring at least one child. The individuals are assumed to be part of a complex system where the relations are defined in a contextual framework, and therefore personal individual preferences are explained by both micro level variables and actual childbearing context. In particular, we use the mean number of children ever born to the older generations in order to identify the neighbourhood's influence on the young generations' ideals. We argue that the mechanism of social interaction is responsible for the transmission/diffusion of fertility ideals from the older to the younger cohorts. The two multi-level models provide significant and consistent results: given the model covariates, the proportion of childless people among the older generations in the region has a positive significant effect on the no-child family ideal; while the mean actual fertility of the older generations in the regions is positively and significantly correlated with the ideal number of children for those desiring at least one child.

However, there are still other important observed and unobserved contextual factors determining the lowest-low fertility preferences of the German speaking countries and the cross-regional differences in Europe. The study of such factors will be the topic of further research.

The past history of fertility is thus confirmed as a key context indicator explaining cross-regional differences in the fertility ideals. These findings are rich in implications for policy makers: if the preferences for smaller families spread out in Europe as a results of the persistent experience of low actual fertility, any further recovery of fertility may be compromised and the introduction of family-friendly policies may become more and more challenging in the future.

**Acknowledgments**

We would like to thank Francesco Billari and Dimiter Philipov for their valuable comments and suggestions.

**References**

Agresti, Alan. 2002. *Categorical Data Analysis.* Second Edition, New York, Wiley.

Barber, Jennifer S., William G. Axinn, and Arland Thornton. 2002. The influence of attitudes on family formation processes, in: Lesthaeghe, Ron. *Meaning and choice: value orientations and life course decisions*. The Hague, Netherlands Interdisciplinary Demographic Institute, pp.45-95.

Bongaarts, John. 2001. Fertility and reproductive preferences in post-transitional societies, Global fertility transition, Supplement to *Population and Development Review*, Vol.27, pp.260-281.

Brant, Rollin. 1990. Assessing proportionality in the proportional odds model for ordinal logistic regression. *Biometrics*, No.46, pp.1171-1178.

Chesnais, Jean-Claude. 2000. Determinants of below replacement fertility, in: Below replacement fertility, *Population Bulletin of the United Nations*, Special Issue 1999, No.40/41, pp.126-136.

Courgeau, Daniel. 2003 (ed.). *Methodology and Epistemology of Multilevel Analysis*, Boston/Dordrecht/London, Kluwer Academic Publishers.

Di Prete, Thomas A. and Jerry D. Forristal. 1994. Multilevel Models: Methods and Substance, *Annual Review of Sociology*, Vol.20, pp.331-357.

Dorbritz, Jürgen, and Charlotte Hoehn. 2000. The future of the family and future fertility trends in Germany. *Population Bulletin of the United Nations*, Special Issue 1999, No.40/41, pp.218-234.

European Commission. 1979. *The Europeans and their children*, Eurobarometer 11, organised and supervised by DG Press and Communication, Brussels.

European Commission. 1990. *Family and the desire for children*, Eurobarometer 32, organised and supervised by DG Press and Communication, Brussels.

Festy, Patrick. 2003. Can we build comparative analysis on micro-data? FFS in the debate. Paper presented at the European Comparisons Seminar, INED, 3 November.

Friesl, Christian. 2001. *Experiment Jung-Sein. Die Wertewelt Österreichischer Jugendlicher*, Vienna, Czernin Verlag.

Gauthier, Anne. 2002. The Promises of Comparative Research, Schmollers Jahrbuch, *Journal of Applied Social Sciences Studies*, 122, pp.5-30.

Grilli, Leonardo and Carla Rampichini. 2002. Specification issues in stratified variance component ordinal response models. *Statistical Modelling*, Vol.2, pp. 251-264.

Goldstein, Joshua, Lutz Wolfgang, and Maria Rita Testa. 2003. The emergence of sub-replacement family size ideals in Europe. *Population Research and Policy Review,* No.5-6, pp.479-496.

Guo, Guang, and Hongxin Zhao. 2000. Multilevel Modeling for Binary Data, *Annual Sociological Review*. Vol.26, pp.441-462.

Hauser, Robert M. 1974. Contextual Analysis Revisited, *Sociological Methods and Research*, 2, pp.365-375.

Jencks, Christopher, and Susan Mayer. 1990. The social consequences of growing up in a poor neighbourhood, in L.E. Lynn and M.G.H. McGeary (eds.) *Inner-city poverty in the United States*. Washington, National Academic Press, pp.111-186.

Kohler, Hans-Peter. 2001. *Fertility and social interaction. An economic perspective*. Oxford, Oxford University Press.

Lee, Ronald. 1980. Aiming at moving target: period fertility and changing reproductive goals, *Population Studies*, 34, pp.205-226.

Liefbroer, Aart C. 1999. Life-planning during young adulthood: do intentions about timing of family life events come true? Paper presented at the European Population Conference. The Hague, The Netherlands, 30 August - 3 September.

Little, Roderick J. A., and Donald B. Rubin. 2002. *Statistical Analysis with Missing Data*, II Edition, New York, Wiley.

Livi Bacci, Massimo, and Gustavo De Santis. 2001. Reflections on the economics of the fertility decline in Europe, Paper presented at the ESF/EURESCO Conference on The Second Demographic Transition in Europe, Bad Herrenhalb, Germany, June 23-28.

Maas, Cora J.M., and Joop J. Hox. 2004. Robustness issues in multilevel regression analysis. *Statistica Neerlandica*, Vol.58, No.2, pp.127–137.

McMahon, Martha. 1995. *Engendering motherhood*. New York, Guilford Press.

Menniti, Adele. 2001. Fertility intentions and subsequent behaviour: first results of a panel study, paper presented to the EAPS, European Population Conference, Helsinki, June 7-9.

Miller, Warren B., and David J. Pasta. 1995. How does childbearing affect fertility motivations and desires? *Social Biology*, Vol.42, No.3-4, pp.185-198.

Monnier, Alain. 1987. Projects de fécondité et fécondité effective. Une enquête longitudinale: 1974, 1976, 1979, *Population*, No.6, pp.819-842.

Morgan, Philip S., and Renbao Chen. 1992. Predicting childlessness for recent cohorts of America women. *International Journal of Forecasting*, Vol.8, pp.477-493.

Murphy, Mike 1999. Is the relationship between parents and children really weak? *Social Biology.* Vol. 46, No.1-2, pp.122-145.

Namboodiri, Krishnan. 1983. Sequential fertility decision making and the life course. Rodolfo A. Bulatao, and Ronald D. Lee (eds.). *Determinants of fertility in developing countries*, Vol.2, New York, Academic Press.

National Research Council. 2001. *Diffusion Processes and fertility transition: selected perspectives,* Committee on Population. John B. Casterline, (ed.). Division of Behavioural and Social Sciences and Education, Washington, National Academy Press.

Nauck, Bernhard. 1995. Regionale Millieus von Familien in Deutschland nach der politischen Vereiningung. In: Nauck, Bernhard, and Corinna Onnen-Isemann (eds.) *Familie im Brennpunkt*. Neuwied, Luchterhand, pp.91-121.

Nazio, Tiziana, and Hans-Peter Blossfeld. 2003. The diffusion of cohabitation among young women in West Germany, East Germany and Italy. *European Journal of Population*, Vol.19, No.1, pp.47-82.

Peterson, Bercedis and Frank E. Harrel jr. 1990. Partial proportional odds models for ordinal response variables. *Applied Statistics*, No.2, Vol.39, pp.205-217.

Quesnel-Vallée, Amélie, and Philip S. Morgan. 2003. Missing the target? Correspondence of fertility intentions and behaviour in the U.S. *Population Research and Policy Review*. Vol.22, No.5-6, pp.497-525.

Rabe-Hesketh, Sophia, Andrew Pickles and Anders Skrondal. 2001. *GLLAMM Manual. Technical Report 2001/04*. Department of Biostatistics and Computing. Institute of Psychiatry, King's College, University of London.

Rindfuss, Ronald R., Philip S. Morgan, and Gray Swicegood. 1988. *First births in America: changes in the timing of parenthood*, Berkeley and Los Angeles, University of California Press.

Ruokolainen, Anne and Irma-Leena Notkola. 2001. Third-child intentions and uncertainty: the effects of family characteristics and attitudinal factors. *Journal of Marriage and the Family*.

Smallwood, Steve, and Julie Jefferies. 2003. Family Building intentions in England and Wales: Trend, Outcomes, and Interpretations. *Population Trends*. 112, pp.15-28.

Snijders, Tom A.B. and Roel Bosker J. 1999. *Multilevel Analysis. An Introduction to Basic and Advanced Multilevel Modeling*. London, Sage Publications.

Symeonidou, Haris. 2000. Expected and Actual Family Size in Greece: 1983-1997, *European Journal of Population*, Vol.16, No.4, pp.335-352.

Tanturri, Maria Letizia, and Letizia Mencarini. 2004. Childless or childfree? A qualitative insight into childlessness in Italy, paper presented at the Annual Meeting of the Population Association of America, Boston.

Teachman, Jay and Kyle Crowder. 2002. Multilevel models in family research: Some conceptual and methodological issues. *Journal of Marriage and Family*. Vol.64, No.2, pp.280-294.

Terza, Joseph V. 1985. Ordinal probit: a generalization. *Communications in Statistics, Theory and Methods*, 14, pp.1-11.

Thomson, Elisabeth. 1997. Couple Childbearing Desires, Intentions and Births. *Demography*. 34, pp.343-354.

Toulemon, Laurent. 1996. Very few couples remain voluntary childless, *Population: An English Selection*, 8, pp.1-28.

Toulemon, Laurent. 2001. Why fertility is not so low in France, paper presented at the IUSSP – International Union for Scientific Study of Population - Seminar on International perspectives on Low fertility: trends, theories and policies, Tokyo, March 21-23.

Van de Kaa, Dirk J. 2001. Postmodern fertility preferences: from changing value orientation to new behaviour, Global fertility transition, Supplement to *Population and Development Review*, Vol.27, pp.290-331.

Van Nimwegen, Nico, Marieke Blommesteijn, Hein Moors, and Gijs Beets. 2002. Late motherhood in the Netherlands: current trends, attitudes and policies. *GENUS*, Vol. LVIII, No.2, pp.9-34.

Van Peer, Christine. 2002. Desired and achieved fertility. In: Klijzing Erik and Corijn Martine (eds.): *Dynamics of fertility and partnership in Europe: insights and lessons from comparative research*. Vol.2, New York and Geneva, United Nations, pp.117-141.

Voas, David. 2003. Conflicting preferences: a reason fertility tends to bee too high or too low. *Population and Development Review*. Vol.29, No.4, pp.627-646.

Westoff, Charles F. 1990. Reproductive Intentions and Fertility Rates. *International Family Planning Perspectives*, Vol.16, No.3, pp.84-89.

Wolf, Naomi 2002. *Misconceptions: Truth, lies, and the unexpected on the journey to motherhood*. New York, Doubleday.

Wong, George Y., and William M. Mason. 1985. The hierarchical logistic regression model for multilevel analysis, *Journal of American Statistical Association*, 80, 391, pp.513-524.

Yamaguchi, Kazuo, and Linda R. Ferguson. 1985. The stopping and spacing of childbirths and their birth-history predictors: rational-choice theory and event-history analysis, *American Sociological Review*, Vol.60, No.2, pp.272-298.

## APPENDIX

### List of regions by country:

| Countries | Regions | Frequencies | Standardized Empirical Bayes Residuals | |
|---|---|---|---|---|
| | | | Logit model | Prop.odds model |
| BELGIUM | South | 99 | -1.00 | -0.15 |
| | Brussels | 39 | 1.76 | 1.11 |
| | North | 181 | -1.00 | 0.47 |
| DENMARK | Copenhagen | 124 | -0.25 | 0.94 |
| | Sjaelland | 41 | 1.28 | 0.75 |
| | Fyn | 22 | 0.58 | 0.62 |
| | Jutland | 170 | 0.92 | -0.92 |
| WEST GERMANY | North | 48 | -1.66 | -1.09 |
| | West | 111 | -0.63 | -0.33 |
| | South | 92 | -1.01 | 0.08 |
| | Berlin area | 27 | 0.67 | -0.45 |
| GREECE | North | 115 | -0.66 | -0.10 |
| | Centre | 220 | 0.88 | 1.63 |
| | South | 45 | 1.93 | 2.00 |
| ITALY | North-West | 49 | 0.33 | 0.72 |
| | North-East | 65 | 0.22 | 0.71 |
| | Center | 65 | -0.33 | 0.14 |
| | South | 73 | -0.26 | 0.49 |
| | Islands | 34 | 0.98 | -2.11 |
| SPAIN | North-West | 37 | 0.92 | 1.06 |
| | North-East | 38 | 2.03 | 0.38 |
| | Madrid area | 43 | -0.47 | -1.04 |
| | Centre | 50 | -1.16 | -2.04 |
| | East | 85 | 0.60 | -0.81 |
| | South | 75 | -1.16 | -1.92 |
| | Canary Islands | 12 | 0.91 | 0.19 |
| FRANCE | Center | 98 | 0.04 | 0.71 |
| | North | 105 | 0.35 | 0.27 |
| | East | 86 | 0.10 | 1.65 |
| | West | 82 | 0.78 | 1.51 |
| | South | 48 | 2.40 | 0.50 |
| IRELAND | North | 38 | 0.03 | 1.31 |
| | Centre | 192 | -1.22 | 0.04 |
| | South | 78 | 1.90 | 0.37 |

| *(Continued):* Countries | Regions | Frequencies | Standardized Empirical Bayes Residuals | |
|---|---|---|---|---|
| | | | Logit model | Prop. odds model |
| LUXEMBOURG | Luxembourg | 162 | 0.20 | -0.83 |
| | Rest of country | 53 | 0.26 | 0.98 |
| NETHERLANDS | North | 30 | -1.58 | 0.84 |
| | East | 72 | -2.17 | 0.86 |
| | West | 160 | -0.77 | -0.90 |
| | South | 81 | -1.93 | 0.92 |
| PORTUGAL | North | 99 | 0.79 | 0.03 |
| | Centre | 58 | 1.47 | 0.88 |
| | Lisboan area | 125 | 0.17 | -1.63 |
| | South | 28 | -1.07 | -0.06 |
| GREAT BRITAIN | North | 60 | -0.97 | 0.62 |
| | Centre | 31 | 1.52 | 1.86 |
| | South | 166 | 0.95 | 0.64 |
| | West | 79 | 0.25 | 0.78 |
| | Wales | 15 | 0.93 | -0.42 |
| | Scotland | 22 | -0.14 | -0.29 |
| | North Ireland | 102 | 0.69 | 0.35 |
| EAST GERMANY | North | 26 | 0.40 | -0.42 |
| | Centre | 93 | -0.71 | -1.27 |
| | South | 124 | -0.67 | -0.96 |
| | Berlin area | 52 | -0.23 | -1.61 |
| FINLAND | South | 91 | 0.14 | -0.06 |
| | Centre | 116 | 0.65 | -1.25 |
| | East | 44 | 0.52 | 0.33 |
| | West | 36 | 1.47 | 0.42 |
| | North | 34 | 0.61 | 1.45 |
| SWEDEN | South | 99 | 0.10 | -0.47 |
| | Centre | 56 | -0.68 | 0.09 |
| | North | 47 | 0.80 | 0.28 |
| | Göteborg | 56 | 0.89 | -0.43 |
| | Stockholm | 37 | -1.06 | -1.50 |
| | Malmö area | 10 | 0.25 | -0.64 |
| AUSTRIA | Tirol | 43 | -0.06 | -2.66 |
| | Carynthia | 36 | -1.50 | -1.13 |
| | Upper Austria | 48 | -0.30 | -0.40 |
| | Styria | 51 | -0.99 | -1.32 |
| | Lower Austria | 57 | 0.23 | 0.33 |
| | Vienna | 71 | -1.44 | -0.70 |