# THE IMPACT OF FAMILY POLICY EXPENDITURE ON FERTILITY IN WESTERN EUROPE*

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*This article analyzes the impact on fertility of changes in national expenditure for family allowances, maternity- and parental-leave benefits, and childcare subsidies. To do so, I estimate a model for the timing of births using individual-level data from 16 western European countries, supplemented with data on national social expenditure for different family policy programs. The latter allow approximation of the subsidies that households with children receive from such programs. The results show that increased expenditure on family policy programs that help women to combine family and employment—and thus reduce the opportunity cost of children—generates positive fertility responses.*

Decreasing fertility and increasing life expectancy has aged Europe's population in what is perhaps the final stage of a demographic transition that begun when early nineteenth century medical innovations and rising incomes augmented the expected life span (Lee 2003). The fertility decline in Europe, which had already begun by the late nineteenth century (Yule 1906), may have been a response not only to decreasing mortality rates but also to the increased cost of raising children owing to economic development. Increasing returns to education, particularly, have improved investment opportunities in the human capital of children and, correspondingly, have increased the incentive for parents to so invest. Such increases, however, have in turn raised the cost of child rearing and may have caused a parental shift toward child quality and away from quantity (Becker 1981; Willis 1973).[1] Whatever its cause, this fertility decline has resulted since the mid-1980s in a below-replacement fertility rate across western Europe[2] and has contributed to the aging of western European populations. These factors have raised concerns about the future labor supply of young skilled workers and about the social cohesion and sustainability of the welfare state (Lee 2003; McDonald 2006; Neyer 2006).

Interestingly, national fertility rates appear to be stabilizing at quite different levels across western Europe; for instance, Mediterranean countries have total fertility rates below 1.4, while the Nordic countries have total fertility rates around 1.8 (Organisation for Economic Co-operation and Development [OECD] 2007a). This cross-national variation in western European fertility rates has often been largely explained in terms of family policy: in particular, labor market policy aimed at creating opportunities for women to combine family and employment (Adserà 2005; Ahn and Mira 2002; Chesnais 1996; d'Addio and d'Ercole 2005; McDonald 2006; Neyer 2006). Theoretical support for this explanation that conforms to the economic theory of fertility (Becker 1981) focuses on the increasing importance of the opportunity cost of children over recent decades as changing gender roles across western Europe have increased the demand for policies that facilitate women's economic empowerment. Such policies often aim to keep women attached to the labor market while

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1. I refer to Galor and Weil (1996) for an alternative economic explanation, and to Folbre (1994) for a discussion on the changing consequences of raising children.

2. Ireland has had a below-replacement fertility rate since the early 1990s (OECD 2007a).

enabling them to have the desired number of children (Adserà 2006; Chesnais 1998).[3] An important issue for public policy, therefore, is the extent to which these family-friendly policies that help women combine labor market participation with child rearing, and thereby reduce the opportunity cost of children, result in higher fertility.

Nonetheless, the empirical evidence on the extent to which fertility is affected by family-friendly labor market policies, such as maternity leave and childcare, is scarce and inconclusive. It ranges from the insignificant effects of maternity leave for Canada (Zhang, Quan, and van Meerbergen 1994), to a significantly negative effect of parental leave duration but a significantly positive effect of the replacement rate for a panel of OECD countries (d'Addio and d'Ercole 2005), and to a significantly positive effect of both maternity-leave availability for the United States (Averett and Whittington 2001) and maternity-leave duration for a panel of OECD countries (Adserà 2004). In addition, Lalive and Zweimuller (2009) reported that the 1990 extension of parental leave in Austria had a significantly positive effect on fertility but that the 1996 reduction had no such effect. Likewise, Blau and Robins (1989) reported insignificant effects on fertility of a childcare tax credit, and Castles (2003) found that fertility outcomes are positively associated with formal childcare provisions but negatively associated with publicly funded childcare provisions. One possible reason for this inconclusive empirical evidence is that these labor market policies have many dimensions—including duration, benefit, entitlement, cash or in kind, and public or mandatory private (see Neyer 2006)—that limit the comparability of a single policy indicator (e.g., length of maternity leave) over time or across countries. Such limitations affect the empirical evaluation of a policy change when not all policy indicators are controlled for. In contrast, the empirical findings on the relationship between family (or child) allowances and fertility are unambiguous, showing that transfers to families with children (cash and tax exemption) have a positive and significant effect on fertility (see d'Addio and d'Ercole 2005; Ermisch 1988; Gauthier and Hatzius 1997; Milligan 2005; Whittington 1992; Whittington, Alm, and Peters 1990; Zhang et al. 1994).

This article aims to contribute substantially to the empirical literature on the effect that changing national expenditure on family policy programs has on fertility in western Europe. Most specifically, it focuses on how the timing of births and completed fertility are affected by changes in public and mandatory private expenditure on family allowances, maternity- and parental-leave benefits, and childcare subsidies. The empirical framework has several methodological advantages over most of the aforementioned empirical studies. First, following Adserà (2005), the empirical model analyzes the timing of births, which is of particular importance given that policy reform may cause women to postpone births, or to have children earlier, without actually affecting completed fertility. Also, by analyzing the timing of births, this model explicitly allows for the fact that fertility decisions are irreversible; that is, the number of a woman's children, in principle, cannot be reduced by a family policy reform. The empirical model extends on the model of Adserà (2005) by including random individual-specific effects to avoid dynamic selection bias resulting from unobserved time-constant individual characteristics (see Cameron and Heckman 1998).

Second, the empirical model takes into account the complexity of family policy programs (discussed earlier) by analyzing fertility responses to a change in the mean subsidy that a woman receives for a child, which is essentially a combination of all policy indicators. For example, the model evaluates the fertility response to a change in the maternity-leave benefit (whether in replacement rate or leave duration) that an employed woman would receive, on average, during the year that she has a child. This approach is

3. Beyond the scope of this article are family planning programs designed to reduce fertility (see Angeles, Guilkey, and Mroz 2005; Brackett, Ravenholt, and Chao 1978; and Tsui 2001). I refer to Van de Kaa (2006) for a critical discussion on pronatal policies, and to Goldstein, Lutz, and Rita (2003) for a discussion and recent evidence on family-size ideals in Europe.

persuasively supported by Zhang et al.'s (1994) empirical finding that three very distinct tax-transfer programs (family allowances) in Canada have no differential effect on fertility, implying that individuals are concerned primarily with the programs' cumulative value (generosity). Examining the generosity of a family policy rather than the policy rules allows comparisons over time and across countries.

Finally, the empirical model aims to avoid the estimation of spurious relationships between fertility and family policy expenditure that may have occurred in earlier within- or cross-country studies when confounding variables at a country level influence the estimated fertility responses (Adserà 2004; d'Addio and d'Ercole 2005; Gauthier and Hatzius 1997; Whittington et al. 1990; Zhang et al. 1994). The model includes country-specific effects to control for time-invariant institutional differences across countries, thereby exploiting the variation over time in family policy expenditure to estimate its impact on birth timing. The model also controls for several time-variant macroeconomic variables (e.g., social expenditure per capita and the national unemployment rate) that the literature identifies as possible confounding variables. Moreover, policymakers may have responded to declining aggregate fertility, or to the (unobserved) socioeconomic changes causing it, by introducing pronatal reforms to family policy, which may result in confounded policy effects (Milligan 2005). Hence, aggregate fertility, being determined by these socioeconomic changes and possibly correlated with birth timings at an individual level and family policy expenditure at a country level, is potentially a confounding variable. To control for such potentially confounded effects, the modeling of birth timings at an individual level includes countries' crude birth rate and total fertility rate as additional country-level covariates.

The remainder of the article is organized as follows. The next section describes the individual-level data on women's fertility history for 16 European countries, which are supplemented by OECD data on national family policy expenditure during 1980–2003. The third and fourth sections, respectively, outline the empirical model and present the estimation results that form the basis of the simulations to assess the quantitative impact of changes in family policy program expenditure on birth timing and completed fertility. The last section summarizes the findings and concludes the article.

## DATA

The fertility histories of women in 16 western European countries are taken from the 2004 European Social Survey (ESS 2004), which is a biennial survey of more than 20 countries designed as a representative sample of all people aged 15 and older living in private households in each country. The data set derived from the ESS 2004 survey comprises second-round data collected in 2004/2005 in face-to-face interviews. The ESS 2004 survey questions probed a wide variety of topics, including social and public trust, social exclusion, and well-being and health. This current analysis uses variables derived from the core questionnaire on household demographic composition and the educational attainment of all household members, which the ESS made comparable for women across countries, using the 1997 International Standard Classification of Education (ISCED).[4]

Because modeling a woman's complete fertility history up to 2004 requires that she be observed from age 15 (assumedly the start of her fertile period) through the years 1980–2003, the empirical analysis includes women born between 1965 and 1984. Thus, the oldest women enter the sample at age 15 in 1980 and are followed to age 38 in 2003; the youngest women enter at age 15 in 1999 and are followed to age 19 in 2003. Whereas the original sample included 5,337 women, 30 observations were removed because of missing information on educational attainment. An additional 51 women were excluded because

4. Available online at http://www.unesco.org/education/information/nfsunesco/doc/isced_1997.htm. See also Appendix Table A1 for the definitions of the levels of education used in this article.

**Table 1. Birth Cohorts 1965–1984: From Age 15 Through 1980–2003**

| Country | Number of Women | Gave Birth to One Child (%) | Gave Birth to Two Children (%) | Gave Birth to Three or More Children (%) | Conceived While Enrolled in School[a] (%) |
|---|---|---|---|---|---|
| Sweden | 283 | 48 | 31 | 11 | 15 |
| Norway | 275 | 58 | 34 | 15 | 17 |
| Finland | 305 | 45 | 27 | 9 | 14 |
| Denmark | 228 | 51 | 33 | 12 | 9 |
| Germany | 388 | 54 | 30 | 9 | 3 |
| Austria | 371 | 54 | 36 | 10 | 6 |
| Belgium | 272 | 52 | 29 | 9 | 6 |
| Netherlands | 308 | 56 | 33 | 11 | 6 |
| Switzerland | 398 | 43 | 27 | 8 | 3 |
| United Kingdom | 345 | 63 | 38 | 15 | 9 |
| Ireland | 406 | 50 | 35 | 25 | 6 |
| France | 304 | 56 | 36 | 12 | 3 |
| Portugal | 376 | 48 | 22 | 4 | 2 |
| Spain | 300 | 38 | 18 | 2 | 1 |
| Italy | 252 | 37 | 16 | 2 | 3 |
| Greece | 445 | 51 | 32 | 8 | 3 |
| All Countries | 5,256 | 50 | 30 | 10 | 7 |

*Notes:* All women were childless at age 15. All statistics are weighted means.

*Source:* 2004 European Social Survey.

[a]Based on the assumption that the child was conceived during the year before his or her birth.

they reported having given birth before the age of 15, most (33) even before the age of 10. Therefore, the final sample includes a total of 5,256 observations.

Table 1 shows the number of women sampled for each country (first column). As mentioned, observations for all the women began when they were 15 years old and childless. For all countries combined, over the observation period, about 50% of the women in the sample gave birth to one child (second column), 30% gave birth to two children (third column), and only 10% gave birth to three or more children (fourth column). In addition, as shown in the last column of Table 1, 7% of the women conceived a child while enrolled in school.

Table 2 shows that women aged 35–38 (cohorts 1965–1968) had, on average, 1.76 children. The mean number of children for most countries is somewhat higher than the national statistics reported by the OECD (2007a), possibly because of definitional differences between cohort and total fertility rates.[5] In line with the statistics reported by Gustafsson (2001), Table 2 shows that a considerable percentage of women aged 35–38 were childless (18%) and that, on average, women had births in their late twenties. Appendix Table A1 shows the age and educational attainment distribution for each sample country.

As in the studies referenced, the data set contains no information on the actual cost of having and raising a child nor on the subsidy that a woman receives for having children or would receive if she had children. However, the child subsidy (i.e., the reduction in the cost of having a child) is related to the national average family allowance per child, maternity- and parental-leave benefits per infant for an employed woman, and the

5. In these countries, cohort fertility rates are somewhat higher than the total fertility rates. See Bongaarts (1999, 2002) for a discussion of the tempo and quantum of fertility.

**Table 2. Mean Number of Children, Percentage Childlessness, and the Average Age at First and Second Birth of Women Aged 35–38 (in 2003)**

| Country | Number of Children (mean) | Childless Women (%) | Average Age at First Birth | Average Age at Second Birth |
|---|---|---|---|---|
| Sweden | 1.99 | 13.9 | 26.9 | 30.1 |
| Norway | 1.92 | 13.5 | 27.0 | 29.2 |
| Finland | 1.78 | 23.2 | 27.2 | 28.8 |
| Denmark | 2.00 | 11.7 | 27.2 | 30.5 |
| Germany | 1.55 | 19.9 | 27.1 | 29.2 |
| Austria | 1.76 | 17.2 | 25.6 | 26.8 |
| Belgium | 1.78 | 13.3 | 26.9 | 28.4 |
| Netherlands | 1.69 | 15.9 | 28.0 | 29.2 |
| Switzerland | 1.53 | 27.9 | 28.2 | 28.8 |
| United Kingdom | 1.75 | 15.8 | 26.1 | 27.7 |
| Ireland | 2.39 | 12.0 | 26.6 | 28.6 |
| France | 1.97 | 15.1 | 27.3 | 27.8 |
| Portugal | 1.53 | 21.5 | 26.9 | 28.6 |
| Spain | 1.41 | 23.9 | 28.8 | 28.2 |
| Italy | 1.47 | 19.4 | 27.7 | 29.3 |
| Greece | 1.79 | 14.7 | 24.6 | 25.0 |
| All Countries | 1.76 | 17.7 | 26.8 | 28.4 |

*Note:* All statistics are weighted means.

*Source:* 2004 European Social Survey.

childcare subsidy per young child for an employed woman. Therefore, in an approach similar to Whittington et al.'s (1990) use of the average tax value of a personal exemption to analyze the exemption's effect on the national birth rate in the United States, I model these child subsidies using the expected or mean subsidy. For this article, national expenditures on family allowances, maternity- and parental-leave benefits, and childcare subsidies are taken from the 2007 OECD Social Expenditure Database, SOCX 2007 (OECD 2007b), which contains information on national social expenditure on public and mandatory private programs for the years 1980–2003 for 16 western European countries that are also included in the 2004 ESS. One primary category of SOCX 2007 is social family expenditure, which includes cash benefits (e.g., child or family allowances, maternity payments, or childcare support), social services (e.g., childcare), tax breaks (e.g., a tax exemption for families with children), and mandatory private expenditure (e.g., through legislation). These public and mandatory private social expenditures are generally referred to collectively as "family expenditures," which are classified for this study as follows: (1) family allowances, (2) maternity- and parental-leave benefits, and (3) childcare subsidies (day care or home help services). However, even though the OECD refers to these programs as "family policy programs," from a different perspective, family allowance is an income-transfer program, whereas maternity- and parental-leave benefits and childcare subsidies are labor-market programs. The (mean) family allowance per child, maternity- and parental-leave benefits per infant for employed women, and childcare subsidy per young child for employed women in selected years are reported in Tables 3, 4, and 5, respectively. In the construction of these statistics, a child is defined as younger than 16; a young child, as younger than 5; and an infant, as up to age 1. To facilitate

**Table 3. Mean Family Allowance Per Child in Selected Years**

| Country | Year 1980 | 1985 | 1990 | 1995 | 2000 | 2003 | Change per Year 1980–2003 (%) |
|---|---|---|---|---|---|---|---|
| Sweden | 1,189 | 1,255 | 1,112 | 1,115 | 1,201 | 1,254 | 0.2 |
| Norway | 927 | 1,030 | 1,498 | 1,923 | 1,485 | 1,638 | 2.5 |
| Finland | 762 | 783 | 852 | 1,587 | 1,418 | 1,376 | 2.6 |
| Denmark | 834 | 559 | 1,101 | 1,351 | 1,486 | 1,524 | 2.7 |
| Germany | 984 | 919 | 905 | 991 | 1,209 | 1,382 | 1.5 |
| Austria | 1,887 | 2,013 | 2,215 | 2,541 | 3,181 | 3,741 | 3.0 |
| Belgium | 2,402 | 2,275 | 2,285 | 2,383 | 2,290 | 2,465 | 0.1 |
| Netherlands | 1,228 | 1,379 | 1,292 | 1,221 | 1,039 | 990 | –0.9 |
| Switzerland | 1,078 | 1,292 | 1,406 | 1,519 | 1,700 | 1,932 | 2.6 |
| United Kingdom | 1,256 | 1,399 | 925 | 1,012 | 1,197 | 1,179 | –0.3 |
| Ireland | 433 | 443 | 518 | 656 | 937 | 1,860 | 6.5 |
| France | 1,523 | 1,598 | 1,120 | 1,315 | 1,523 | 1,578 | 0.2 |
| Portugal | 965 | 490 | 364 | 367 | 413 | 468 | –3.1 |
| Spain | 339 | 141 | 86 | 146 | 185 | 310 | –0.4 |
| Italy | 1,331 | 834 | 834 | 487 | 755 | 725 | –2.6 |
| Greece | 1,347 | 653 | 409 | 647 | 507 | 657 | –3.1 |
| All Countries | 1,155 | 1,066 | 1,058 | 1,204 | 1,283 | 1,443 | 1.0 |

*Note:* All financial statistics are in 2000 euro and are corrected for differences in purchasing power across countries.
*Source:* 2007 OECD Social Expenditure Database (OECD 2007b).

comparisons over time and across countries, all amounts are real (in 2000 euro) and corrected for purchasing power parities.

Tables 3–5 show little convergence despite European Union efforts to influence national family policy.[6] Rather, they show considerable variation in (the changes in) expenditure on the three family policy programs. For instance, looking at change per year in 1980–2003, Table 3 shows that family allowances have increased in most countries, most strongly in Ireland, but have decreased in the Netherlands, the United Kingdom, and Mediterranean countries. In 2003, the mean yearly family allowance per child ranged from 310 euro in Spain to 3,741 euro in Austria. In the same year (see Table 4), maternity- and parental-leave benefits per infant for employed women ranged from 2,842 euro in Ireland to 34,575 euro in Norway. Nonetheless, over the observation period, maternity- and parental-leave benefits remained relatively high for Sweden, Austria, France, and Italy; increased strongly in Norway and Switzerland; but decreased in Portugal and Greece. As Table 5 illustrates, the mean yearly childcare subsidy per young child for employed women has increased in all 16 countries—particularly in Belgium and Ireland, and in Mediterranean countries—resulting in 2003 in a range from 1,432 euro in Greece to 15,544 euro in Denmark.

6. For instance, the European Union sets minimum standards, such as the Parental Leave Directive (Council Directive 96/34/EC), which introduced the individual right to a three-month parental leave for fathers and mothers. See Hantrais (1997) for a discussion of European family policy. See also Gauthier (2002) for an analysis of trends in family policies in 22 industrialized countries.

**Table 4. Mean Maternity- and Parental-Leave Benefits per Infant for Employed Women in Selected Years**

| | Year | | | | | | |
|---|---|---|---|---|---|---|---|
| Country | 1980 | 1985 | 1990 | 1995 | 2000 | 2003 | Change per Year 1980–2003 (%) |
| Sweden | 23,272 | 21,990 | 25,245 | 25,251 | 22,520 | 25,210 | 0.3 |
| Norway | 6,265 | 6,324 | 10,629 | 24,379 | 31,898 | 34,575 | 7.7 |
| Finland | 7,279 | 17,674 | 26,280 | 29,009 | 22,344 | 24,096 | 5.3 |
| Denmark | 8,272 | 11,975 | 12,625 | 25,751 | 17,414 | 21,614 | 4.3 |
| Germany | 4,270 | 3,873 | 9,894 | 14,986 | 14,359 | 14,521 | 5.5 |
| Austria | 13,561 | 13,637 | 15,560 | 25,402 | 19,010 | 16,257 | 0.8 |
| Belgium | 3,133 | 3,405 | 6,629 | 7,459 | 8,287 | 9,936 | 5.1 |
| Netherlands | 1,815 | 1,796 | 3,957 | 3,957 | 3,559 | 3,240 | 2.6 |
| Switzerland | 491 | 441 | 3,012 | 3,789 | 3,613 | 4,546 | 10.2 |
| United Kingdom | 2,787 | 2,281 | 2,679 | 2,312 | 3,031 | 4,087 | 1.7 |
| Ireland | 1,637 | 2,444 | 2,026 | 2,198 | 2,278 | 2,842 | 2.4 |
| France | 9,066 | 7,069 | 9,816 | 11,993 | 13,066 | 13,699 | 1.8 |
| Portugal | 5,910 | 3,037 | 2,279 | 2,676 | 3,009 | 3,805 | –1.9 |
| Spain | 3,400 | 3,880 | 4,570 | 5,704 | 6,832 | 6,712 | 3.0 |
| Italy | 12,548 | 10,864 | 7,226 | 8,388 | 8,982 | 11,984 | –0.2 |
| Greece | 9,192 | 6,692 | 7,446 | 4,062 | 3,259 | 4,202 | –3.3 |
| All Countries | 7,056 | 7,336 | 9,243 | 12,189 | 11,319 | 12,458 | 2.5 |

*Note:* All financial statistics are in 2000 euro and are corrected for differences in purchasing power across countries.

*Source:* 2007 OECD Social Expenditure Database (OECD 2007b).

## EMPIRICAL MODEL

The first part of the empirical analysis estimates the effects of the covariates on the timing of births using a discrete-time proportional hazard model (Cameron and Trivedi 2005: chap. 17). The timing of births is essentially modeled as a sequence of yearly birth decisions. I define the probability of a woman (indexed by $i$) giving birth after $d_{ik}$ years at birth parity $k$ at calendar time $t$ as follows:

$$\Pr(B_{ic}(t) = 1 | \, d_{ik}(t), \mathbf{X}_{ic}(t), \mathbf{Z}_c(t), \eta_i; \boldsymbol{\theta}) = F(\lambda(d_{ik}(t); \boldsymbol{\alpha}_k) + \mathbf{X}_{ic}(t)'\boldsymbol{\beta}_k + \mathbf{Z}_c(t)'\boldsymbol{\gamma}_k + \eta_i), \qquad (1)$$

where $B_{ic}(t)$ is a random variable equal to 1 if woman $i$ gives birth at time $t$, and 0 otherwise. $F(.)$ is the logistic cumulative distribution function, and $k$ denotes the birth parity, $k \in \{0,1,2,\ldots,K\}$. $K$ denotes the maximum number of births, $\lambda(.)$ is a nonparametric baseline function for modeling duration dependence, $c$ is a country index, $\mathbf{X}_{ic}(t)$ and $\mathbf{Z}_c(t)$ are vectors of covariates that are allowed to depend on calendar time $t$, and $i \in \{1,\ldots,n\}$. The model explicitly accounts for unobserved individual-specific effects, or random effects, denoted by $\eta_i$, which in turn allows for dynamic selection on unobservable time-constant individual characteristics (see Cameron and Heckman 1998). In the context of this article, $\eta_i$ can be thought of as a preference for having children. I denote the parameters of interest by $\boldsymbol{\theta} = (\theta_0,\ldots,\theta_K)$, with $\theta_k = (\alpha_k, \beta_k, \gamma_k)$. I address the possible importance for the birth timing/policy relationship of conception occurring nine months before birth, by using a conception date rather than a birth date but designate it as 1 year

**Table 5. Mean Childcare Subsidy per Young Child for Employed Women in Selected Years**

| | Year | | | | | | |
|---|---|---|---|---|---|---|---|
| Country | 1980 | 1985 | 1990 | 1995 | 2000 | 2003 | Change per Year 1980–2003 (%) |
| Sweden | 11,666 | 11,732 | 10,390 | 9,209 | 12,778 | 14,190 | 0.9 |
| Norway | 2,325 | 2,660 | 3,958 | 6,254 | 11,591 | 8,678 | 5.9 |
| Finland | 3,539 | 4,672 | 6,166 | 7,158 | 8,169 | 8,304 | 3.8 |
| Denmark | 9,458 | 9,742 | 10,188 | 11,755 | 14,143 | 15,544 | 2.2 |
| Germany | 1,413 | 1,535 | 1,781 | 3,467 | 3,921 | 4,286 | 4.9 |
| Austria | 2,321 | 2,424 | 2,808 | 3,355 | 5,320 | 5,934 | 4.2 |
| Belgium | 972 | 938 | 701 | 595 | 6,181 | 8,460 | 9.9 |
| Netherlands | 4,170 | 3,599 | 4,029 | 2,744 | 5,792 | 6,695 | 2.1 |
| Switzerland | 993 | 1,041 | 982 | 959 | 2,920 | 3,692 | 5.9 |
| United Kingdom | 1,473 | 1,265 | 942 | 1,262 | 4,051 | 4,667 | 5.1 |
| Ireland | 247 | 267 | 216 | 749 | 1,521 | 1,982 | 9.5 |
| France | 1,077 | 1,346 | 3,833 | 4,990 | 9,692 | 9,579 | 10.0 |
| Portugal | 40 | 23 | 46 | 75 | 1,606 | 1,923 | 18.3 |
| Spain | 278 | 300 | 261 | 210 | 5,061 | 5,182 | 13.6 |
| Italy | 2,104 | 1,776 | 1,343 | 1,370 | 8,349 | 7,891 | 5.9 |
| Greece | 160 | 137 | 1,961 | 1,290 | 1,218 | 1,432 | 10.0 |
| All Countries | 2,640 | 2,716 | 3,100 | 3,465 | 6,395 | 6,777 | 4.2 |

*Note:* All financial statistics are in 2000 euro and are corrected for differences in purchasing power across countries.
*Source:* 2007 OECD Social Expenditure Database (OECD 2007b).

(rather than 9 months) previously because the data are yearly. Nonetheless, because still births are excluded, the term "timing of births" still seems the most appropriate.

The covariates included in $\mathbf{X}_{ic}(t)$ are a woman's age, to control for such factors as age-related decline in fecundity (Van Noord-Zaadstra et al. 1991), and the woman's educational attainment, whose effect on fertility is assumed to be the same across countries. This latter assumption is empirically supported by Björklund's (2006) finding that western European household data typically reveal a negative association between women's educational attainment and fertility (see Gustafsson and Kalwij 2006). However, because this present analysis cannot take into account a woman's marital and employment history, the model does not control for either a woman's employment status or a partner's educational attainment. Thus, given that educational attainment may, for instance, affect fertility through differences in employment opportunities (see Kalwij 2000), this model may be considered a reduced-form model.

The country-level covariates are included in $\mathbf{Z}_c(t)$. Specifically, the family policy expenditure variables are the logarithms of the mean child allowance per child, the mean maternity- and parental-leave benefit per infant for employed women, and the mean childcare subsidy per young child for employed women. These statistics, reported in Tables 3–5, are country- and time-specific. However, as previously pointed out, because confounding country-level variables may yield spurious relationships between fertility and family policy expenditure, the empirical model controls for country-specific effects and for several time-variant macroeconomic variables identified as possible confounding variables. I include GDP per capita and the unemployment rate to control for cyclical fertility patterns (Butz and Ward 1979) not captured by possible cyclical variation in family policy expenditure. In addition, Boldrin, De Nardi, and Jones (2005) reported a strong negative relationship between Social Security spending (and, in particular, government-provided pensions) and fertility. Social Security spending also may be positively related to family policy expenditure, and the model therefore controls for (total) social expenditure per capita. The model also includes the female employment rate as an indicator for labor market institutions that facilitate combining employment and child rearing and that can be related to expenditure on the three family policy programs (see d'Addio and d'Ercole 2005). Finally, I include the crude birth rate and the total fertility rate to control for the possibly confounded effects of introducing pronatal reforms of family policy in response to declining aggregate fertility or the (unobserved) socioeconomic changes causing it, as argued by Milligan (2005).[7]

As mentioned, the analysis includes neither the actual subsidy a family receives when, for instance, making use of childcare nor the childcare subsidy that a childless woman would have received if she had a child. Surveys do not include such counterfactual information, but even if available, it might be subject to self-selection because women desiring children may move into jobs that offer such benefits as maternity leave (see Averett and Whittington 2001).[8] Hence, this research assumes that women's fertility decisions are influenced by the subsidy that they expect to receive from each of the three family policies based on the mean subsidy received by eligible women, the route through which national expenditure on family policy programs is assumed to affect fertility outcomes. Identification of this relationship is thus the focus of the empirical analysis.

## ESTIMATION

A woman enters the panel at age 15 (at time $t_i$) and remains until 2003 (time $T$). I model the timing of births as a sequence of birth decisions. Hence, the likelihood function of individual $i$ with an observed sequence of births denoted by $b_{ic} = (b_{ic}(t_i),\ldots,b_{ic}(T))$ is as follows:

$$\Pr(B_{ic} = b_{ic} \mid \mathbf{X}_{ic}, \mathbf{Z}_c, \eta_i; \theta) = \prod_{t=t_i}^{T} \left(\Pr(B_{ic}(t) = 0 \mid d_{ik}(t), \mathbf{X}_{ic}(t), \mathbf{Z}_c(t), \eta_i; \theta)\right)^{1-b_{ic}(t)} \times \left(\Pr(B_{ic}(t) = 1 \mid d_{ik}(t), \mathbf{X}_{ic}(t), \mathbf{Z}_c(t), \eta_i; \theta)\right)^{b_{ic}(t)}, \tag{2}$$

where $B_{ic} = (B_{ic}(t_i),\ldots,B_{ic}(T))$, $\mathbf{X}_{ic} = (X_{ic}(t_i),\ldots,X_{ic}(T))$ and $\mathbf{Z}_c = (Z_c(t_i),\ldots,Z_c(T))$. The probabilities on the right side of Eq. (2) are defined in Eq. (1). I treat unobserved individual characteristics (the $\eta_i$'s) as random effects and assume them to be normally distributed with mean 0 and variance $\sigma^2$. I estimate the model by maximizing the logarithm of the likelihood function in which the unobserved individual-specific effects are integrated out:

$$(\hat{\theta}, \hat{\sigma}) = \underset{(\theta,\sigma)}{\arg\max} \sum_{i=1}^{n} \log\left( \int_{-\infty}^{+\infty} \Pr(b_{ic} \mid \mathbf{X}_{ic}, \mathbf{Z}_c, \eta_i; \theta)\, d\Phi\left(\frac{\eta_i}{\sigma}\right)\right), \tag{3}$$

where $\Phi$ denotes the standard normal cumulative distribution function. As Table 1 shows, third births are relatively few; therefore, the parameter vectors are restricted to being the same for all births after the first up to a constant that depends on whether the woman has already conceived two or three times: that is, $\boldsymbol{\theta} = (\theta_0, \theta_1)$. I calculate the standard errors by taking into account serial correlation and clustering on an individual and a country level (see Cameron and Trivedi 2005: chap. 24; and Moulton 1990).

7. As discussed in relation to Eq. (1), a conception rather than a birth date is used when modeling the timing of births. Including the total fertility and therefore the crude birth rates does not exclude the possibility of an immediate impact of a particular family policy expenditure on current period conception decisions (and next period births).

8. Averett and Whittington (2001) argued this point but found no empirical evidence to support it.

## MONTE CARLO SIMULATIONS

Whereas the parameter estimates of the model provide insights into the direction and relative size of the effects of the covariates on the birth probabilities, they offer no clear insights into the quantitative effects on birth timing and completed fertility. Therefore, the second part of the empirical analysis simulates life cycle fertility patterns and examines how they are affected by the family policy expenditure variables included in the model. To be more specific, an increase in family policy expenditure may have only a small effect on the probability of giving birth in a specific year but may have a substantial effect on completed fertility for young women who start planning childbearing based on this increased family policy expenditure. The simulation exercise throws light on such long-run effects.

I perform the Monte Carlo simulations as follows. For a group of reference women (the baseline), arbitrarily chosen as being from Sweden and having an ISCED educational level of 3 or 4 (secondary education or postsecondary/nontertiary education), the simulation starts at age 15. The parameter estimates, Eq. (3), enable calculation of the probability that each woman will give birth, following which, based on a random drawing from the uniform distribution, it is possible to simulate whether each does give birth (see Law and Kelton 1982). Next, using the estimated birth probabilities, and given the (simulated) past fertility outcomes, I simulate year-by-year births for each woman up to age 40. I present these statistics as the means for this homogenous reference group: for example, the average number of children these women have up to age 25. I then rerun these simulations with a change in expenditure on one of the three family policy programs so that the differences between these and the baseline simulation outcomes can reveal the effect of this change on the timing of births and completed fertility. I perform these Monte Carlo simulations for 10,000 (identical) women. The standard errors for the differences from the baseline situation are based on 1,000 drawings from the asymptotic distribution around the parameter estimates.

## EMPIRICAL RESULTS

Table 6 reports the estimation results of the model, and Tables 7 and 8 report the simulation results. In the following discussion, I use a 5% level of significance. As Table 6 shows, the higher the woman's level of education, the lower the probabilities of first and subsequent births, which results in a negative relationship between fertility and educational attainment that is well established in the literature (for an overview, see Björklund 2006; and Gustafsson and Kalwij 2006). The economic rationale for this finding, in line with Becker (1981), is that if price effects dominate income effects, higher educational attainment is associated with a higher wage rate and thus higher opportunity costs of having children. This dynamic lowers the demand for the quantity of children, and women can postpone childbirth and increase human capital investments in their children (i.e., a higher demand for quality). Even though the analysis does not identify the age effects on the probability of first birth separately from duration dependence, the age effects for subsequent births show a significant increase in the birth probability up to age 31, after which the birth probability declines with age. This latter finding echoes the empirical evidence on the age-related decline in fecundity for women over age 31 (Van Noord-Zaadstra et al. 1991), although, of course, other factors may also explain it.

As discussed earlier, the model includes countries' crude birth rate and total fertility rate to control for the possibly confounded effects of introducing pronatal reforms in family policy in response to declining aggregate fertility or the socioeconomic changes causing it. As Table 6 shows, the effects of these two variables are (jointly) significant (Wald-test statistic = 52.1; degrees of freedom = 4; $p$ value = 0). More important, however, is the effect that including these two variables has on the parameter estimates of the other country-level variables. To illustrate this point, Table 6 reports the estimation results when the crude birth rate and total fertility rate are not controlled for. A comparison of the four leftmost columns

**Table 6. Estimation Results: Probability of Giving Birth (based on date of conception), From Eq. (1),[a] With and Without Controlling for Crude Birth Rate and Total Fertility Rate ($N$ = 5,256)**

| | With Controls | | | | Without Controls | | | |
|---|---|---|---|---|---|---|---|---|
| | First Birth | | Subsequent Births | | First Birth | | Subsequent Births | |
| Covariates | Estimate | $z$ Value | Estimate | $z$ Value | Estimate | $z$ Value | Estimate | $z$ Value |
| Spline Function for Age Effects | | | | | | | | |
| Slope for age ≤ 31 | | | 0.142 | 12.58 | | | 0.139 | 12.38 |
| Slope for age >31 and age ≤ 35 | | | –0.014 | –0.575 | | | –0.016 | –0.690 |
| Slope for age > 35 | | | –0.274 | –3.014 | | | –0.277 | –3.067 |
| Two or More Children | | | –1.881 | –33.78 | | | –1.868 | –33.70 |
| Three or More Children | | | –0.863 | –9.305 | | | –0.850 | –9.168 |
| During Education | –0.876 | –10.51 | –0.193 | –1.663 | –0.884 | -10.65 | –0.189 | –1.636 |
| Lower Level of Education, ISCED 0, 1, or 2 | 0.957 | 12.45 | 0.363 | 5.229 | 0.953 | 12.44 | 0.355 | 5.138 |
| Higher Level of Education, ISCED 5 or 6 | –0.581 | –9.542 | –0.074 | –1.048 | –0.573 | -9.471 | -0.069 | -0.983 |
| Time Trend (in years) | –0.013 | –2.326 | –0.018 | –1.902 | –0.021 | -3.735 | –0.035 | –4.147 |
| Log(GDP per Capita) | 0.009 | 0.037 | 0.151 | 0.478 | 0.841 | 3.770 | 1.295 | 4.990 |
| Log(Social Expenditure per Capita) | –0.278 | –1.186 | –0.051 | –0.167 | –0.949 | -4.435 | –1.161 | –4.252 |
| Female Employment Rate (in %) | –0.001 | –0.165 | –0.007 | –1.211 | 0.011 | 3.446 | 0.009 | 1.868 |
| National Unemployment Rate (in %) | –0.030 | –3.432 | –0.009 | –0.703 | 0.002 | 0.242 | 0.030 | 2.740 |
| Crude Birth Rate (births per 1,000 people) | –0.117 | –2.895 | 0.020 | 0.443 | — | | — | |
| Total Fertility Rate | 1.690 | 6.020 | 1.215 | 3.498 | — | | — | |
| Log(Family Allowance per Child) | 0.019 | 0.328 | 0.089 | 1.186 | 0.180 | 3.535 | 0.320 | 4.438 |
| Log(Maternity- and Parental-Leave Benefits per Infant for an Employed Woman) | 0.159 | 3.570 | –0.102 | –1.572 | 0.204 | 4.705 | -0.062 | -0.954 |
| Log(Childcare Subsidy per Young Child for an Employed Woman) | –0.014 | –0.486 | 0.166 | 3.851 | 0.018 | 0.619 | 0.223 | 5.354 |
| $p$ Value: Joint Significance of Country-Specific Effects | .000 | | .000 | | .000 | | .000 | |
| SD of the Distribution of Unobserved Individual Heterogeneity (σ) | 1.093 | | 13.42 | | 1.080 | | 13.34 | |
| Logarithm of the Likelihood Function | –16,943 | | | | –16,948 | | | |

[a]Duration dependence is controlled for by including dummy variables for each year-duration (for both probabilities).

with the four rightmost columns reveals that when these two variables are controlled for, the effects of GDP per capita, social expenditure per capita, and the female employment rate on birth probabilities become (jointly) insignificant (Wald-test statistic = 8.9; degrees of freedom = 6; *p* value = .18); the effects of child allowances become insignificant; and the effects of maternity- and parental-leave benefits and child care subsidies are reduced (but not significantly).

This finding that the effects of GDP per capita, social expenditure per capita, and the female employment rate on the probabilities of giving birth are individually and jointly insignificant are not in line with the findings in some earlier studies. For instance, Adserà (2004) reported a positive effect of GDP per capita on the fertility of women aged 30–34, d'Addio and d'Ercole (2005) found a positive effect of the female employment rate on the fertility rate, and Boldrin et al. (2005) showed a negative association between Social Security expenditure and fertility. These results from previous cross-country studies are, however, in line with those in the four rightmost columns of Table 6, suggesting that the relationships may be spurious. In regards to the cyclical sensitivity of fertility, as shown in the four leftmost columns of Table 6, an increase in the national unemployment rate has a small and negative effect on the probability of first birth but no significant effect on the probability of subsequent births. Overall, this finding may explain the often-reported countercyclical movements of fertility (Butz and Ward 1979).[9]

The primary focus of this analysis, however, is fertility responses to expenditure changes in family allowances (for children), maternity- and parental-leave benefits, and childcare subsidies. I find that in contrast to previous empirical findings (Ermisch 1988; Gauthier and Hatzius 1997; Whittington et al. 1990; Zhang et al. 1994), family allowance has no significant effect on birth probabilities (see Table 6). Without controlling for the crude birth rate and total fertility rate, however, the results are in line with previous empirical findings that family allowance has a positive and significant effect on fertility (see Table 6). This discrepancy suggests that the positive relationship identified in previous empirical work may be spurious. As also shown in Table 6, maternity- and parental-leave benefits have a significant positive effect on the probability of first birth but no significant effect on the probability of subsequent births. Conversely, childcare subsidy has no significant effect on the probability of first birth but a significant and positive effect on the probability of subsequent births.

The results given in Table 7, which are based on Monte Carlo simulation, quantify the effects on birth timing and completed fertility of changes in family policy expenditure on child allowances, maternity- and parental-leave benefits, and childcare subsidies. The simulations are based on the estimation results presented in the four leftmost columns of Table 6. As mentioned earlier, these changes in family policy expenditure are effective from the age of 15, when women are assumed to start planning childbearing. In the life cycle fertility simulation outcomes for a baseline reference group (first column), 88% of the women have children before age 40, with the average being 1.7 children (last row). In line with the estimation results given in Table 6, the simulation results in the second column show that a 10% increase in family allowance has no significant effects on the timing of maternity or on completed fertility. The third column illustrates that although a 10% increase in maternity- and parental-leave benefits has no significant effect on the number of children borne by maternal women, it significantly affects the probability of having children, reducing childlessness at age 36–40 by about 3.2%. Hence, in contrast to the finding of Averett and Whittington (2001), the effect on completed fertility is insignificant. This absence of any effect of increased maternity- and parental-leave benefits on subsequent births may result from a concomitant increase in the career prospects of women who have had their first child and are thus free to continue on their career paths. The result is an increase

9. Without controls (Table 6), the results for the effects of unemployment are in line with Kravdal (2002).

**Table 7. Simulated Effects on Life Cycle Fertility of Changes in Family Policy Program Expenditures**

| Age Group | Baseline Prediction | 10% Increase in Family Allowance, Difference From Baseline | | 10% Increase in Maternity- and Parental-Leave Benefits, Difference From Baseline | | 10% Increase in Childcare Subsidies, Difference From Baseline | |
|---|---|---|---|---|---|---|---|
| | | Estimate | *z* Value | Estimate | *z* Value | Estimate | *z* Value |
| Probability of Having Children | | | | | | | |
| 16–20 | 0.019 | 0.000 | 0.240 | 0.000 | 2.160 | –0.0000 | –0.380 |
| 21–25 | 0.162 | 0.000 | 0.290 | 0.002 | 3.070 | –0.0002 | –0.440 |
| 26–30 | 0.489 | 0.001 | 0.320 | 0.005 | 3.420 | –0.0004 | –0.470 |
| 31–35 | 0.791 | 0.001 | 0.340 | 0.005 | 3.220 | –0.0004 | –0.490 |
| 36–40 | 0.883 | 0.000 | 0.340 | 0.004 | 2.780 | –0.0004 | –0.500 |
| Average Number of Children Conditional on Having Children | | | | | | | |
| 16–20 | 1.037 | 0.000 | 0.260 | 0.000 | –0.180 | 0.001 | 0.520 |
| 21–25 | 1.164 | 0.001 | 0.900 | –0.001 | –0.920 | 0.002 | 2.020 |
| 26–30 | 1.379 | 0.002 | 1.150 | –0.002 | –0.860 | 0.004 | 2.720 |
| 31–35 | 1.721 | 0.004 | 1.210 | –0.002 | –0.930 | 0.006 | 3.090 |
| 36–40 | 1.944 | 0.004 | 1.200 | –0.003 | –1.070 | 0.008 | 3.130 |
| Average Number of Children | | | | | | | |
| 16–20 | 0.020 | 0.000 | 0.270 | 0.000 | 2.080 | 0.000 | 0.240 |
| 21–25 | 0.189 | 0.000 | 0.460 | 0.003 | 2.740 | 0.000 | 0.160 |
| 26–30 | 0.675 | 0.002 | 0.650 | 0.006 | 2.560 | 0.001 | 0.760 |
| 31–35 | 1.363 | 0.004 | 0.910 | 0.006 | 1.640 | 0.004 | 1.610 |
| 36–40 | 1.718 | 0.005 | 1.040 | 0.004 | 0.950 | 0.006 | 2.080 |

in the opportunity cost of the second child, which may explain the small negative but insignificant effects on subsequent births. Conversely, a 10% increase in childcare subsidy has no significant effect on childlessness, but it significantly increases the number of children borne by maternal women (fourth column), yielding a 0.4% increase in completed fertility. I am not aware of any comparable numbers in the empirical literature.

Although these quantitative fertility responses are small (see Table 7), the relatively large changes in expenditure on family policy in western Europe over the period 1980–2003 (see Tables 3–5) may have generated considerable fertility responses. Table 8 provides insights into this using a Monte Carlo simulation and thereby keeping all other factors affecting fertility constant. The observed changes in the average family policy expenditures are taken from Tables 3–5. Tables 3–5 show that over this period, on average, expenditure on child allowances per child increased by 26%, expenditure on maternity- and parental-leave benefits per infant for employed women increased by 76%, and expenditure on childcare subsidy per young child for employed women increased by 158%. The baseline predictions are identical to those in Table 7. Table 8 shows that the increased family policy expenditure in western Europe during 1980–2003 had significant effects on the timing of maternity and completed fertility. Childlessness at age 36–40 declined by about 19%, and completed fertility increased by about 8.3%.

**Table 8. Simulated Effects on Life Cycle Fertility of Observed Changes in Family Policy Program Expenditure, 1980–2003: A 26% Increase in Family Allowance per Child, a 76% Increase in Maternity- and Parental-Leave Benefits per Infant for Employed Women, and a 158% Increase in Childcare Subsidies per Young Child for Employed Women (see Tables 3–5)**

| Age Group | Baseline Prediction | Difference From Baseline | |
|---|---|---|---|
| | | Estimate | *z* Value |
| Probability of Having Children | | | |
| 16–20 | 0.019 | 0.002 | 1.810 |
| 21–25 | 0.162 | 0.017 | 1.910 |
| 26–30 | 0.489 | 0.034 | 2.020 |
| 31–35 | 0.791 | 0.030 | 2.060 |
| 36–40 | 0.883 | 0.022 | 1.990 |
| Average Number of Children, Conditional on Having Children | | | |
| 16–20 | 1.037 | 0.006 | 1.570 |
| 21–25 | 1.164 | 0.027 | 2.400 |
| 26–30 | 1.379 | 0.059 | 2.690 |
| 31–35 | 1.721 | 0.092 | 2.850 |
| 36–40 | 1.944 | 0.109 | 2.860 |
| Average Number of Children | | | |
| 16–20 | 0.020 | 0.002 | 1.890 |
| 21–25 | 0.189 | 0.025 | 2.140 |
| 26–30 | 0.675 | 0.078 | 2.430 |
| 31–35 | 1.363 | 0.127 | 2.830 |
| 36–40 | 1.718 | 0.143 | 2.990 |

## SUMMARY AND CONCLUSIONS

This analysis of the impact on fertility of changes in national expenditure on family policy programs focuses on family allowances, maternity- and parental-leave benefits, and childcare subsidies. For this purpose, it estimates a model for the timing of births using individual-level data from 16 western European countries, supplemented with data on national social expenditure for different family policy programs to approximate the subsidies that households with children receive from these programs. Analysis of the three family policy programs studied here yield different fertility responses.

The first program, family allowance, is an income policy program, targeted at households with children, that aims to alleviate the financial burden of having children and consequently increase the quality of children. The results show that an increase in child subsidy through a family allowance program's increased generosity has no significant impact on the timing of births or on completed fertility (see Table 7). This finding can be explained by the fact that a family-allowance policy subsidizes the direct costs of children and not the opportunity costs of children, which have arguably become much more important for fertility decisions in recent decades during which changing gender roles have increased the demand for policies that facilitate women's economic empowerment.

In contrast, maternity and parental leave and childcare provisions, which are both labor-market policy programs, aim to reduce the opportunity costs of children and facilitate the combination of employment and child rearing. According to the results of this analysis, these family-friendly labor-market policy programs significantly affect life cycle fertility. A 10% increase in maternity- and parental-leave benefits results in about a 3.2% reduction in childlessness at ages 36–40 but has no significant effect on completed fertility. Conversely, a 10% increase in childcare subsidies has no significant effect on the proportion of women who have children but results in about a 0.4% increase in completed fertility (see Table 7).

Overall, these empirical findings suggest that increased expenditure on family policy programs aimed at empowering women through opportunities to combine family and employment—thereby reducing the opportunity cost of children—generate positive fertility responses. More specifically, extending maternity and parental leave and childcare provisions causes women to have children earlier in life and to have more children. Even though the quantitative fertility responses to changes in family-friendly labor-market policy expenditure identified here are small, the relatively large changes in expenditure on family policy programs over recent decades in western Europe have generated considerable fertility responses (see Table 8).

**Appendix Table A1. Age Distribution and Educational Attainment in 2003 (in %)**

| | Age Distribution | | | | | Educational Attainment | | |
|---|---|---|---|---|---|---|---|---|
| Country | 15–19 | 20–24 | 25–29 | 30–34 | 35–38 | ISCED[a] 0–2 | ISCED[a] 3–4 | ISCED[a] 5–6 |
| Sweden | 5 | 25 | 22 | 23 | 25 | 18 | 33 | 49 |
| Norway | 4 | 17 | 22 | 30 | 27 | 5 | 46 | 49 |
| Finland | 5 | 21 | 26 | 26 | 23 | 7 | 48 | 46 |
| Denmark | 3 | 16 | 26 | 28 | 26 | 12 | 54 | 34 |
| Germany | 4 | 19 | 20 | 25 | 32 | 12 | 75 | 14 |
| Austria | 7 | 21 | 13 | 25 | 33 | 59 | 35 | 6 |
| Belgium | 3 | 26 | 25 | 25 | 22 | 12 | 45 | 43 |
| Netherlands | 3 | 14 | 23 | 31 | 29 | 22 | 50 | 28 |
| Switzerland | 3 | 15 | 22 | 30 | 30 | 11 | 72 | 17 |
| United Kingdom | 5 | 21 | 24 | 27 | 24 | 11 | 57 | 32 |
| Ireland | 7 | 28 | 23 | 21 | 22 | 18 | 56 | 26 |
| France | 5 | 21 | 22 | 25 | 27 | 31 | 21 | 47 |
| Portugal | 4 | 25 | 23 | 28 | 20 | 57 | 27 | 16 |
| Spain | 6 | 19 | 28 | 26 | 22 | 42 | 25 | 33 |
| Italy | 3 | 20 | 27 | 26 | 23 | 23 | 61 | 16 |
| Greece | 3 | 19 | 26 | 27 | 24 | 27 | 55 | 18 |
| All Countries | 4 | 21 | 23 | 26 | 26 | 24 | 48 | 28 |

*Source:* 2004 European Social Survey.

[a]The 1997 International Standard Classification of Education (ISCED) categorizes as follows: Level 0 = preprimary education; Level 1 = primary education or first stage of basic education; Level 2 = lower secondary or second stage of basic education; Level 3 = (upper) secondary education; Level 4 = postsecondary nontertiary education; Level 5 = first stage of tertiary education; and Level 6 = second stage of tertiary education.

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