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An explanation of the positive correlation between fertility and female employment across Western European countries*

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Abstract: *Recent literature shows the puzzling result of a positive and significant cross-country correlation between the total fertility rate and the female labour force participation rate across Western European countries. The present paper shows that this cross-country correlation becomes negative and significant, once one corrects the total fertility rate for a distortion, caused by an increasing age of childbearing, and controls in cross-country regressions for purchased child care use and female long-term unemployment. This result survives an empirical analysis in which the female labour force participation rate is treated as an endogenous variable.*

Keywords: Total fertility rate, female labour force participation rate, purchased child care, female unemployment.

JEL classification: J10, J11, J13.

1. Introduction

The issue of demographic change is today on the top of the policy agenda in many OECD countries. Most OECD countries experienced considerable declining fertility rates since the 1960s. They also experienced increasing female labour force participation rates during these decades. On the one hand, considerable decreasing fertility rates strongly reduce the potential sustainability of public pension systems, as it implies considerable declining future contributors to public pension systems. On the other hand, increasing female labour force participation rates improve the prospects for sustaining public pension systems, as it implies increasing contributors to public pension systems. Declining aggregate fertility and increasing aggregate female employment over time is consistent with microeconomic models of fertility behaviour (see, e.g., De Cooman et al., 1987, and Ermisch, 2000). According to these models, time and budget constraints imply that the number of births and female labour market activity are determined by common external variables; in particular by the relative wage of women. In Galor and Weil (1996) such a microeconomic model of fertility behaviour is built into a general equilibrium model with an endogenously rising relative wage of women. Their model implies declining fertility rates and rising female labour force participation rates.

*I benefited from insightful comments from participants of the British Society for Population Studies Annual Conference 2005 at the University of Kent at Canterbury.

While time series data since the 1960s are consistent with standard economic models of fertility, this is not true for cross-country data. As I will be discussing in section 2, a recent analysis has shown a reversal of the cross-country correlation between the fertility rate (TFR) and the female labour force participation rate (FLP) from a negative value until the begin of the 1980s to a positive value today. This finding is a challenge for economics, as standard economic models of fertility imply a negative association between the TFR and the FLP over time, as well as, across countries. However, Kögel (2004) showed that the association between the TFR and the FLP remains negative and significant, once one includes country dummy variables in pooled time series and cross-country data. This implies that the puzzling positive cross-country correlation between the TFR and the FLP is due to omitted variables that can be approximated with country dummy variables. Since the values of the coefficients of the country dummy variables are correlated with the FLP, failure to control for these country dummy variables leads to an omission bias of the coefficient of the FLP. Due to the fact that the country dummy variables explain more of the cross-country variation of the TFR today than the FLP, the bias of the coefficient of the FLP even causes the sign of this coefficient to change to a positive value today.

Given the fact that these country dummy variables are so important for the sign of the association between the TFR and the FLP today, it seems important to know what the omitted variables are that are behind the country dummy variables. Section 3 shows that a distortion of the TFR, caused by an increasing age of childbearing, purchased child care use and the female long-term unemployment rate are behind the country dummy variables. It is shown that the cross-country correlation between the TFR and the FLP across Western European countries today changes from a positive and significant value to a negative and significant value, once one corrects the TFR for the aforementioned distortion and controls in cross-country regressions for the aforementioned two omitted variables. The underlying mechanisms are identical to those that caused the cross-country correlation between the TFR and the FLP to change its sign, once one controls for country dummy variables.¹

According to the logic of economic models of fertility, fertility and female employment are determined by common external variables. Hence, in a regression of the TFR on the FLP the coefficient of the FLP should contain an endogeneity bias. In addition, the FLP most likely contains measurement errors that lead to a measurement error bias of the coefficient of the FLP. Both problems can be solved by use of instrument variable estimation with instrument variables that are correlated with the FLP, but that are uncorrelated with the TFR (after controlling for the FLP) and the measurement errors of the FLP. Section 4 applies instrument variable estimation to the

¹ See also Engelhardt and Prskawetz (2004) for an attempt to find omitted variables that might be behind the country dummy variables in Kögel (2004). Castles (2003) also controlled for various family policy variables and labour market indicators in a regression of the TFR on the FLP or female education. In the paper he only shows multivariate regression results for his best-fit model, which is a regression of the TFR on female tertiary education and female unemployment. In this model female tertiary education remains positively and significantly correlated with the TFR. However, Castles neither corrected for the distortion of the TFR nor controlled for purchased child care use.

regression equation of section 3 and shows that the results of section 3 survive and become even slightly stronger with instrument variable estimation. This is most likely due to a reduction of measurement errors.

Cross-country differences in purchased child care use could reflect differences in the availability of purchased child care. However, alternatively it could also reflect differences in attitudes towards purchased child care. For this reason, section 5 briefly investigates possible determinants of purchased child care.

2. A Review of Recent Literature

Recently Ahn and Mira (2002) and Rindfuss et al. (2003) challenged economic models of fertility by showing that in OECD countries the cross-country correlation coefficient between the TFR and the FLP had changed from a negative value until around 1985 to a positive value thereafter. Figure 1 replicates their result for twenty-one OECD countries for 1960-1999 (see Kögel, 2004, Appendix D for all countries and data sources). Similarly, Esping-Andersen (1999) and Brewster and Rindfuss (2000) showed that in cross-country regressions of the TFR on the FLP the coefficient of the FLP was significant and negative in the 1970s, but significant and positive by the 1990s. The puzzle of a positive cross-country correlation since the 1990s has been confirmed for Western European countries in Del Boca et al. (2003).

However, after introducing purchased child care in the basic model of Galor and Weil (1996), Apps and Rees (2004) show, among other results, that an increase in child care subsidies simultaneously increases the TFR and the FLP. In addition, after also introducing purchased child care in the basic model of Galor and Weil and assuming that purchased child care is produced with unskilled labour, Martinez and Iza (2004) show that in an increasing

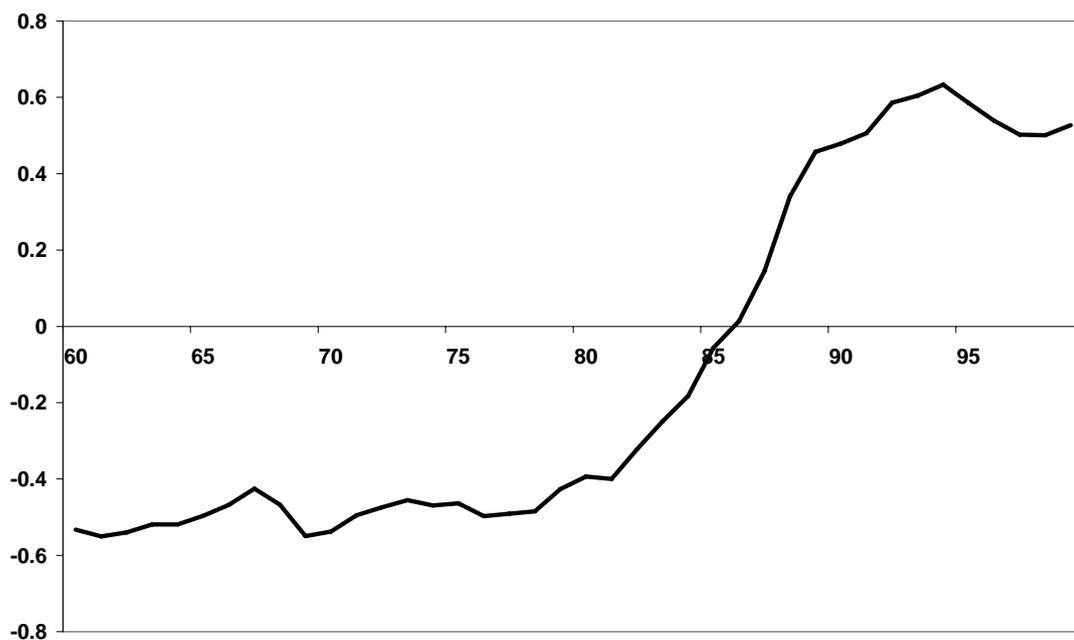


Figure 1: *Annual cross-country correlation coefficient between fertility and female labour force participation.*
Data source: Kögel (2004).

relative wage of skilled labour (relative to the wage of unskilled labour) also simultaneously increases the TFR and the FLP. Therefore, cross-country differences or changes over time of child care subsidies or the relative wage of skilled labour can produce a positive association between the TFR and the FLP across countries or over time. In addition, a model of Da Rocha and Fuster (2006) implies that high female unemployment in an economy increases the risk that a woman does not find a job after interrupting a career for childrearing. The authors show in a calibrated general equilibrium model that for this reason high female unemployment leads to postponement of childbearing and reduced female labour force participation and can produce a positive association between the TFR and the FLP across countries and over time. Hence, macroeconomic models of fertility extended with purchased child care or female unemployment are consistent with Figure 1.

The aforementioned literature still maintains the common assumption in economics that the female wage has a negative effect on fertility. Recently, demographers went further. Various recent studies by demographers show in Swedish micro-level data that surprisingly female wage income or female education have a positive effect on fertility, in particular on second and third order births (see, among others, Andersson, 2000, Berinde, 1999, and Olah, 2003). The demographic literature explains this finding with Swedish family policies of high subsidies for purchased child care and generous parental leave benefits that are calculated on the basis of a woman's prior wage income. It is argued, both policies would cause the substitution effect from an increase in female wages on fertility to be dominated by its income effect. The logic behind this explanation follows a model of Ermisch (1989), who shows that the possibility of purchased child care reduces the effect of female wages on fertility. He shows that women with high wage income can afford more purchased child care and that this reduces the magnitude of the substitution effect for these women.² The demographic literature argues that Swedish family policy would, similarly to the presence of purchased child care in Ermisch's model, reduce the magnitude of the substitution effect in Sweden. However, Kögel (2006) shows within a model similarly to the model of Ermisch and the aforementioned model of Apps and Rees (2004) that there are offsetting effects from Swedish family policy that cause the reduction in the magnitude of the substitution effect of female wages to be most likely rather small. This is due to the fact that the effect of child care subsidies on the substitution effect and the effect of parental leave benefits on the substitution effect do more or less offset each other (see Kögel, 2006, for details).

As mentioned before, the present paper finds that a distortion of the TFR, caused by an increasing age of childbearing, purchased child care use, and female long-term unemployment are behind the country dummy variables. This finding is consistent with the cross-country implication of the aforementioned models of Apps and Rees (2004), Martinez and Iza (2004) and Da Rocha and Fuster (2006). In the former two models cross-country differences in child care subsidies or the relative wage of skilled labour lead to

² See also Ermisch (1989) for empirical support for this result in British micro-level data.

cross-country differences in purchased child care use and in turn these difference in purchased child care use are positively correlated with the TFR and the FLP. In the model of Da Rocha and Fuster the female unemployment rate is positively correlated with the TFR and the FLP. For this reason, these three models predict the present paper's result that, apart from failure to correct for a distortion of the TFR, omitting purchased child care use and female unemployment from the a cross-country regression leads to a spuriously positive cross-country correlation between the TFR and the FLP.³

The result of Figure 1 of a reversal of the sign of the cross-country correlation between the TFR and the FLP is in the literature (explicitly or implicitly) often interpreted as evidence for a changing sign in the time series association between the TFR and the FLP. However, by showing that the association between the TFR and the FLP remains negative and significant, once one includes country dummy variables in pooled time series and cross-country data, the aforementioned study by Kögel (2004) showed that the time series association between the TFR and the FLP remained negative and significant today. The reason for this is the fact that, as is well-know in econometrics literature, a regression of the TFR on the FLP in pooled time series and cross-country data with country dummy variables is equivalent to a regression of $(TFR_{it} - \overline{TFR}_{it})$ on $(FLP_{it} - \overline{FLP}_{it})$, where the indexes *i* and *t* denote country *i* and time, while \overline{TFR}_{it} and \overline{FLP}_{it} denote the average values of the TFR and the FLP over time for country *i*. This implies that in the regressions in Kögel the negative coefficient of the FLP represents a negative time series association between the TFR and the FLP.

Hence, the finding in Kögel is in contrast to the time series implications of the aforementioned models of Apps and Rees (2004), Martinez and Iza (2004) and Da Rocha and Fuster (2006). However, probably in reality the contrast between the time series and the cross-country association between the TFR and the FLP is not as extreme as the analysis in Kögel suggest when taken at face value. Instead, in reality the aforementioned control variables behind the country dummy variables in Kögel seem to change over time. Because of this, if changes of the control variables are omitted, then the magnitude of the negative time series association between the TFR and the FLP changes over time. There is evidence for this view in Engelhardt et al. (2004), who omitted the control variables and found in time series analysis for six representative OECD countries (France, Italy, Sweden, the United Kingdom, the United States and Western Germany) that the magnitude of the negative time series association between the TFR and the FLP fell over time for all countries except Italy. In addition, Engelhardt et al. found country-heterogeneity in the time series association between the TFR and the FLP. This could imply that the control variables changed differently for different countries.⁴

³ Note that, if purchased child care use and female long-term unemployment were the only determinants of the FLP, then after controlling for these two variables in cross-country regressions of the TFR on the FLP, the coefficient of the FLP should be insignificant. The fact that the coefficient of the FLP is negative and significant after controlling for these two variables implies that there exists other variables that are important determinants of the FLP.

⁴ Kögel (2004, section 3), respectively, Engelhardt and Prskawetz (2005) confirmed the results of Engelhardt et al. in pooled time series and cross-country data, using quinquennial time series data (i.e. data for 1960, 1965,...2000), respectively, annual data. Furthermore,

Figure 2 shows a graph that plots for Italy and Sweden in 1975 and 2000 the FLP on the horizontal axis and the TFR on the vertical axis. Due to country dummy variables the value of the intercept (i.e. the value of the TFR, if the FLP equals zero) is larger for Sweden than for Italy. Further, the negative time series association between the TFR and the FLP is stronger for Italy than for Sweden. The figure shows that both elements, significance of country dummy variables and country-heterogeneity in the magnitude of the negative time series association between the TFR and the FLP, together imply a reversal of the sign of the cross-country association between the TFR and the FLP from a negative value in 1975 to a positive value in 2000. It should be clear that a weakening time series association between the TFR and the FLP over time can make a reversal of the sign of the cross-country association between the TFR and the FLP even more likely. Unfortunately, there is lack of data to calculate time series variation of the distortion of the TFR for various countries and lack of time series data of purchased child care use. This prevents a confirmation of the aforementioned interpretation that changes in the distortion of the TFR, changes in purchased child care use and changes in female long-time unemployment over time caused country-heterogeneity in the time series association between the TFR and the FLP and a weakening time series association between the TFR and the FLP; features which can according to Figure 2 contribute to a reversal of the sign of the cross-country association between the TFR and the FLP.

The aforementioned literature measured fertility with the total fertility rate (TFR). The total fertility rate is the sum of all age specific fertility rates of women age of 15 to 49. This measure is independent of the age structure women age of 15 to 49. However, the fertility decline in OECD countries came

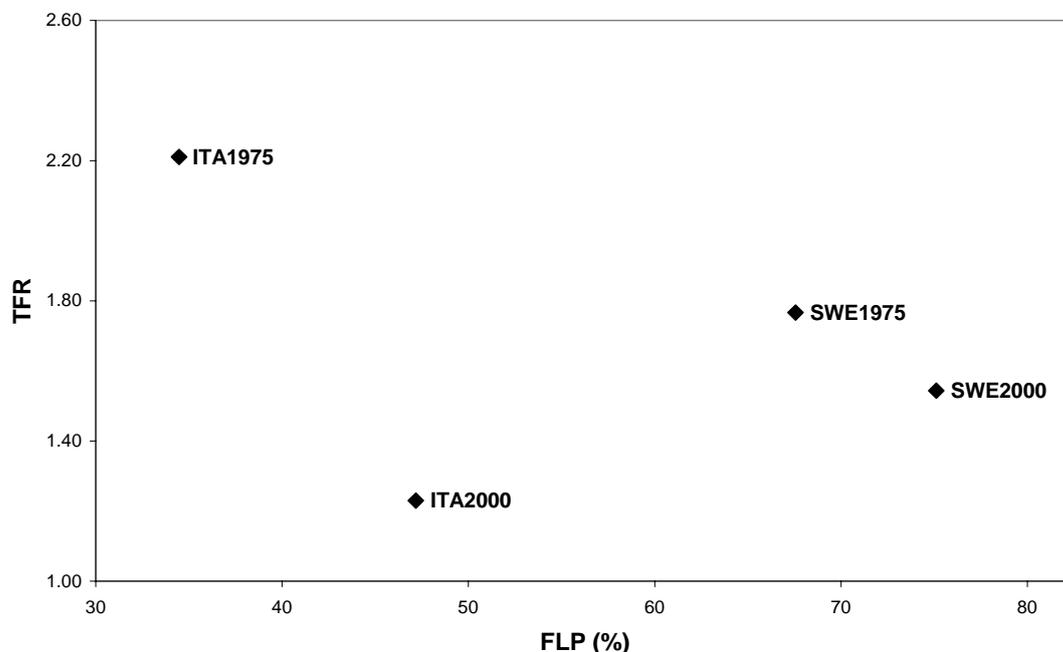


Figure 2: *TFR and FLP in Italy and Sweden in 1975 and 2000.*
Data source: Kögel (2004).

also Pampel (2001) found in pooled time series and cross-country data with country dummy variables a weakening negative association between the TFR and the FLP over time.

along with increasing age of childbearing. Since path breaking work of Bongaarts and Feeney (1998), recent demographic literature emphasises that an increasing age of childbearing leads to a temporary “tempo effect” that reduces the TFR temporarily. The fact that this reduction of the TFR is only temporarily implies that it is a distortion that makes the comparison of the TFR over time and across countries more difficult. Bongaarts and Feeney (1988) developed a method to calculate an adjusted TFR (henceforth TFRadj) that is supposed to be free of the distortion from the tempo effect. Demographers label a TFR that is free of a tempo effect “fertility quantum”. Sobotka (2004) has recently shown that the tempo effect is the largest in Mediterranean and Central European and Eastern European countries. Most of these countries have what Kohler et al. (2002) labelled “lowest-low fertility” and which they defined as a TFR below 1.3. Sobotka applied the Bongaarts and Feeney adjustment method to the average TFR of 1995 to 2000 and found that the TFRadj is for none of the lowest-low fertility countries below the threshold of 1.3. Hence, while the TFRadj in these countries is still low, it is not as dramatically low as data of the TFR suggest. For the present study most importantly, the Mediterranean countries, with the largest tempo effect across the Western European countries, also have the lowest FLP across the Western European countries. For this reason, in a regression of the TFR on the FLP, the FLP captures the tempo effect of the TFR because it is correlated with the tempo effect. However, there is no causal effect from a low FLP to a high tempo effect. While an increase in the FLP might lead to postponement of childbearing, a low level of the FLP obviously does not lead to such postponement of childbearing. Clearly, this implies a necessity to use the data of Sobotka (2004) with the TFRadj, which, for this reason, the present study uses.⁵

3. The Basic Empirical Model

Given the fact that according to Sobotka (2004) Mediterranean countries are the countries in Western Europe with the highest magnitude of the tempo effect and given the fact that Mediterranean countries have also the lowest FLP in Western Europe, it is possible that the positive cross-country correlation today is simply due to the distortion of the TFR from the tempo effect. For this reason, Table 1 examines with an OLS regression applied to cross-country data whether the result of a positive coefficient of the FLP survives in case one regresses the TFRadj on the FLP. Later in this paper I will show results were I controlled in addition for purchased child care use and the female long-term unemployment rate. Unfortunately, the complete set of all variables of the present study is only available for 12 Western European

⁵ See Sobotka (2004) for a brief explanation of the adjustment method of the TFR of Bongaarts and Feeney and his data sources. Kohler and Ortega (2002) developed recently a more sophisticated method to calculate an adjusted TFR. Several demographers argue that with this method one can calculate an adjusted TFR that comes closer to the “true” fertility quantum than with the adjustment method of Bongaarts and Feeney. Unfortunately, lack of data prevents the use of the adjustment method of Kohler and Ortega for many countries.

Table 1: OLS regression without control variables, cross-country data, average 1995-2000.

Dependent variable	TFRadj
Constant	1.20
FLP	0.91 (0.11)
R ²	0.24

Note: The p-value is in parentheses behind the value of the coefficient.

countries.⁶ For comparability with later results, the regression of Table 1 contained only data for these 12 twelve countries.⁷ As mentioned before, data of the TFRadj are taken from Sobotka (2004). Since according to Sobotka the TFRadj contains some irregularities over time, he used the mean over the five year period 1995-2000 (due to data availability problems, he used for a few countries the mean of data in the neighbourhood of 1995-2000). The FLP is the labour force participation rate of women of age 25-49, i.e. of childbearing age (see the appendix for exact definitions of all variables of this paper and all data sources used in this paper). In Table 1 the FLP is the average of the FLP of 1995 and 2000. A glance at Table 1 reveals that the puzzling result of a positive coefficient of the FLP since the 1990s survives when one regresses the TFRadj on the FLP. For 12 data points a p-value of 0.11 indicates also moderate significance of the positive coefficient of the FLP. Hence, other factors than a tempo effect must also play a role in explaining the positive cross-country correlation between fertility and female employment today.

Figure 3 compares purchased child care use with the FLPs for three broad country groups within Western Europe. The black bars show the percent rate of children below age three for which purchased child care (publicly or market provided) is used (this variable is abbreviation with child care). The age group of children below three is shown, because this is the most care intensive age group and hence, according to demographers, purchased child care use is in this age group most relevant for the incompatibility between childrearing and female employment. The year of reporting of purchased child care use varies between 1995 and 2000. The grey bars show the FLPs for 1995-2000. The abbreviation Nordic and Med denote the group of Nordic countries (Denmark, Finland, Norway and Sweden) and the group of Mediterranean countries (Greece, Italy, Portugal and Spain). Rest denotes the remaining Western European countries (Austria, France, Ireland and The Netherlands). The figure shows that in 1995-2000 purchased child care uses, as well as, the FLPs in the Nordic countries were the highest in Western Europe, while the values of both variables for the Mediterranean countries were the lowest in Western Europe. Hence, purchased child care use and the FLP are positively correlated. This implies that, if purchased child care use increases the TFRadj, then the FLP captures the positive effect from purchased child care use, if this variable is omitted from the regression. Hence, in this case one needs to control for purchased child care use to avoid a biased coefficient of the FLP.

⁶ The twelve countries are Austria, Denmark, Finland, France, Greece, Ireland, Italy, The Netherlands, Norway, Portugal, Spain and Sweden.

⁷ The reader should note that, while use of data for only 12 countries is unusual, the effect of use of so few countries is mainly to reduce the efficiency of the estimates; making it only more difficult to find significant effects.

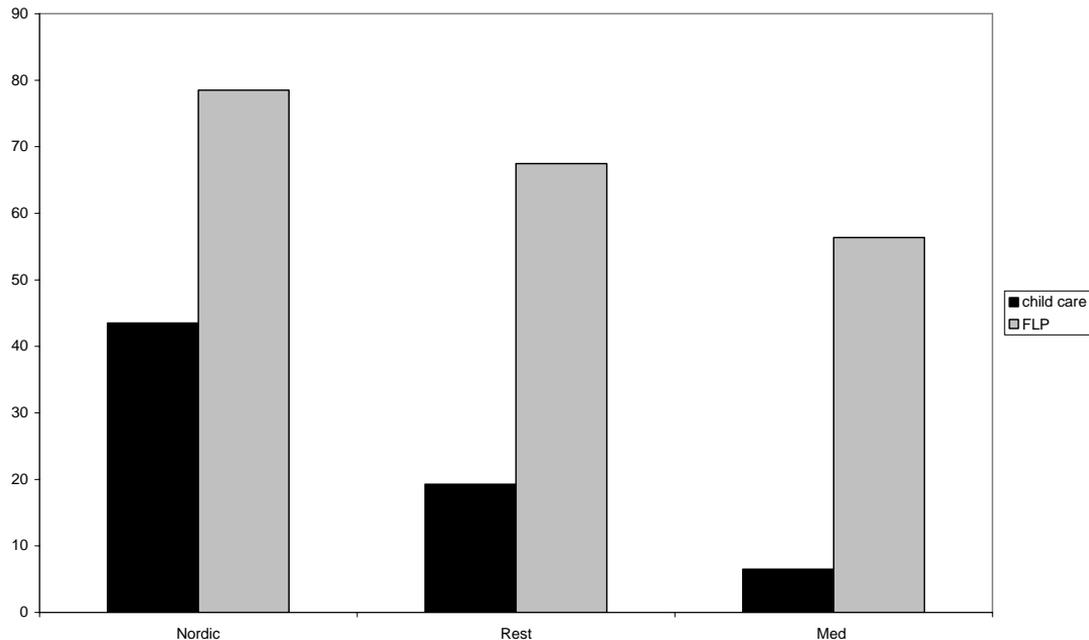


Figure 3: *Percent rate of children below three for which purchased child care is used (abbreviated with child care) and FLP (%), average 1995-2000. Sources: OECD Employment Outlook (2001, Table 4.7) (for purchased child care use) and Eurostat (2005a)(for the FLP).*

Figure 4 compares, for the same three broad country groups as in Figure 3, the female long-term unemployment rate (abbreviated with *ufemale_long*) in black bars and the FLP (again in grey bars). The female long-term unemployment rate is the proportion of women of age above 15 in the female labour force who are unemployed since twelve months or longer. Recently, Adsera, (2005) and, as mentioned before, Da Rocha and Fuster (2006) argued that female unemployment leads due to increased uncertainty to postponement of childbearing (possibly because of increased risk that a women does not find a job after interrupting a career for childrearing). Adsera found that in particular the long-term unemployment rate of women reduces fertility significantly in micro-level data. Empirical exercises with my data (not shown) confirmed also that the negative effect on the TFR_{adj} from female long-term unemployment is stronger than the negative effect from the sum of short-term and long-term female unemployment or the effect from male unemployment.⁸ Postponement of childbearing reduces the fertility quantum, if there is no sizable "recuperation" with increased fertility at higher age for fertility forgone at earlier age. Lesthaeghe and Moors (2000) do not find evidence for sizable recuperation in Italy and Spain, implying that postponement of childbearing reduces the fertility quantum in these countries. Figure 4 shows for 1995-2000 that the Mediterranean countries had the highest female long-term unemployment rates in Western Europe, while the Nordic countries had the lowest female long-term unemployment rates. The fact that those countries that lack sizable recuperation also had the highest female long-term unemployment rates implies that female long-term

⁸ See also Castles (2003) for evidence for a significant and negative effect of female unemployment on the TFR in OECD countries.

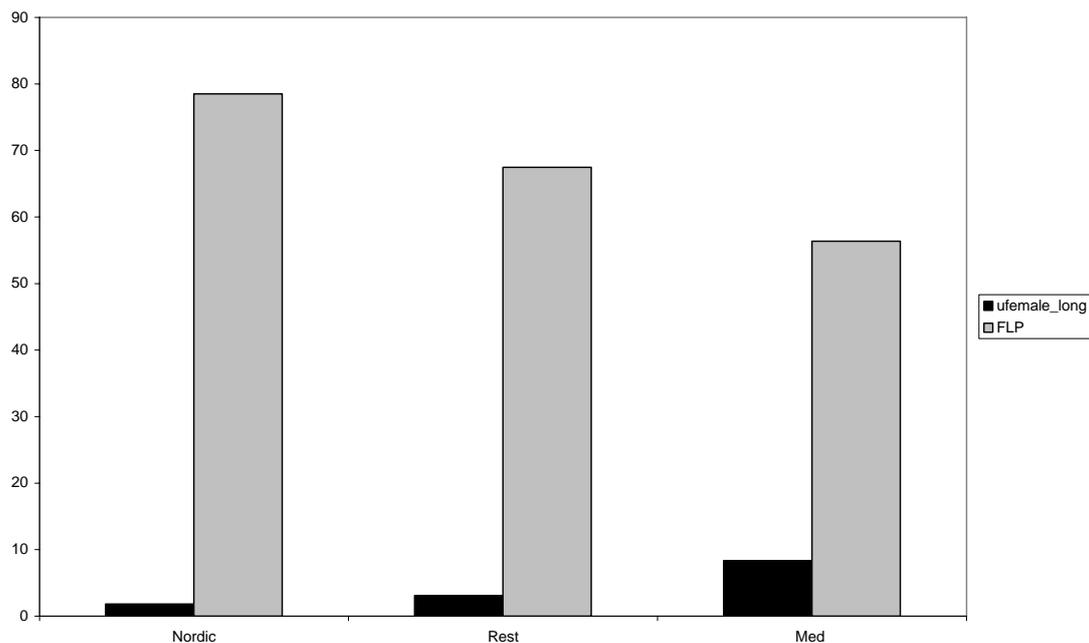


Figure 4: Female long-term unemployment rate (%) (abbreviated with *ufemale_long*) and FLP (%), average 1995-2000.

Sources: Eurostat (2005b) (for Female long-term unemployment rate) and Eurostat (2005a) (for the FLP).

unemployment contributed to the low TFRadj in these countries, if female long-term unemployment caused postponement of childbearing. Further, Figure 4 shows that the female long-term unemployment rate and the FLP are negatively correlated. As a consequence, a negative effect from female long-term unemployment on the TFRadj implies that the FLP captures the negative effect from female long-term female unemployment on the TFRadj, if this variable is omitted from the regression. Obviously, this implies that one also needs to control for female long-term unemployment to avoid a biased coefficient of the FLP.

Table 2 shows the result of an OLS regression applied to cross-country data of the twelve aforementioned Western European countries. In this regression the TFRadj is regressed on the FLP and the two aforementioned control variables purchased child care use and the female long-term unemployment rate. The data are the same as in Figure 4 and 5.⁹ The table shows that the coefficient of the FLP becomes negative and significant after controlling for the aforementioned control variables. In addition, the coefficients of the control variables are both significant with the expected sign (i.e. positive for the natural logarithm of purchased child care use and negative for female long-term unemployment).¹⁰ In addition the regression in Table 2 gives rise to an extremely high R^2 of 0.87. As explained before, the

⁹ Throughout the present study, the decision whether the natural logarithm of a variable or its level was included in the regression was based on which specification gives the lowest p-value of that variable. A lower p-value for a specification in natural logarithm implies a better statistical fit for an approximation of a non-linear relationship with the natural logarithm.

¹⁰ Note, that the significant effect from purchased child care use on the TFRadj cannot reflect reverse causality, as purchased child care use does not measure the absolute amount of

Table 2: OLS regression with control variables, cross-country data, average 1995-2000.

Dependent variable	TFRadj
Constant	3.42
FLP	-1.53 (0.04)
Natural logarithm of purchased child care use	0.19 (0.00)
Female long-term unemployment rate	-4.55 (0.04)
R ²	0.87

Note: The p-values are in parentheses behind the values of the coefficients.

differences in the value of the coefficient of the FLP between the regression in Table 1 without control variables and the regression in Table 2 with control variables is due to the fact that failure to control for the control variables of Table 2 leads to an omission bias of the coefficient of the FLP, because the FLP picks up some of the effects of the omitted variables. As also explained before, the fact that the coefficient of the FLP has even different signs in Table 1 and 2, is due to the fact that the control variables of Table 2 explain more of the cross-country variation of the TFRadj than the FLP and hence omitting these control variables even changes the sign of the coefficient of the FLP.¹¹

The significant and positive effect of cross-country differences in purchased child care use for fertility differences between Western European countries is consistent with recent findings in Hilgeman and Butts (2004). These authors applied to data from OECD countries a multi-level regression, i.e. a regression of micro-level fertility on micro-level characteristics and macro-level variables that vary across countries, and found a significant and positive effect of cross-country differences in purchased child care use on micro-level fertility. Hilgeman and Butts controlled also in their regression, among other mainly micro-level variables, for micro-level female labour market activity and macro-level female labour force participation rates. Both variables were found to have a negative and significant effect on micro-level fertility. The negative and significant effect of the two variables remained, when the authors dropped from the regression the macro-level variable purchased child care use. However, this finding is not surprising in light of the present study's result, which argues that the positive correlation between macro-level fertility and macro-level female employment across OECD countries is only caused by the fact that the macro-level female labour force participation rate is correlated with omitted macro-level variables and becomes negative, if one controls for these omitted macro-level variables. Obviously, differences in fertility between individuals within countries are not correlated with these omitted macro-level variables, causing no such omission

purchased child care use per women of age 15-49. Instead, it measures, as mentioned before, the proportion of children for which purchased child care is used.

¹¹ I added various further control variables to the regression of Table 2 (results not shown), such as, the percentage of women in employment working part-time on a voluntary basis, parental leave benefits as a percentage of average wages (data source: OECD, 2001) and public expenditures on family cash benefits as share of GDP (data source: Bertelsmann Foundation, 2006). None of these variables turned out to be significant.

bias. Apparently, in the authors' data set, data variation in differences in fertility between individuals within countries dominates data variation in differences in fertility between countries.

In contrast to Hilgeman and Butts, Hank and Kreyenfeld (2003) applied to Western German data multi-level regressions of micro-level fertility on micro-level variables and public child care use in Western Germany districts and found an insignificant effect from public child care use on micro-level fertility. Using the same statistical approach, Andersson et al. (2004) also found barely an effect from public child care use in Swedish municipalities on micro-level fertility. However, using also multi-level regressions, Del Boca (2002) found a positive and significant effect of purchased child care use in regions of Italy on micro-level fertility. A possible explanation of these conflicting findings in the multi-level literature might be that one only finds evidence for an effect of purchased child care use on fertility, if there is enough data variation in purchased child care use across the macro-level units. Apparently between industrialised countries this seems to be the case. However, possibly between some regions within countries this might not be the case.

4. Instrument-Variable Estimation

As mentioned before, according to the logic of economic models of fertility, fertility and female employment are determined by common external variables. As a consequence, according to these models the FLP in Table 2 of the last section is not an exogenous variable, but an endogenous variable instead. This leads to an endogeneity bias. Because of the endogeneity problem, the standard economic approach is to regress the TFRadj and the FLP in separate equations on their common external variables. After estimation, one could then compare the signs of the coefficients of the common external variables in both equations to see whether variations of the common external variables lead indirectly to a positive or a negative association between the TFRadj and the FLP.¹² However, to make a comparison with earlier literature and the result of the last section easier, this section contains instead the results of instrument variable estimation or more precisely two-stage least squares (2SLS) regressions. As is well-known to applied econometricians, in 2SLS regressions, in a first stage regression the FLP is regressed on its determinants and in a second stage regression the TFRadj is regressed on the predicted values of the FLP from the first stage regression and, in addition, on purchased child care use and on female long-term unemployment.¹³ Apart from avoiding an endogeneity bias, this approach has also the advantage that it reduces possible measurement errors that the variable FLP surely contains. It will turn out that measurement errors were actually a more important source of bias in the results of Table 2 in the last section than an endogeneity bias.

¹² Applying such an empirical exercise, with the control variables of Table 2 included in the list of common external variables, showed evidence for an indirectly negative association between the TFRadj and the FLP (results are not shown).

¹³ To receive the correct standard errors, the first stage regression and the second stage regression are estimated simultaneously.

Table 3: Gender wage gap (i.e. ratio of the male wage to the female wage), 1997.

	Gender wage gap
Nordic countries	1.15
Mediterranean countries	1.29
Remaining countries in sample	1.31

Source: Gauthier (2002).

The model of Galor and Weil (1996) implies the ratio of the male wage to the female wage (henceforth the gender wage gap) to be a natural candidate for an instrument for the FLP. However, in 2SLS regressions (not shown) the gender wage gap turned out to be insignificant for the FLP.¹⁴ Table 3 gives an explanation of this result. The table shows that Nordic countries have the lowest gender wage gap in Western Europe, consistent with their highest FLPs in Western Europe (because the model of Galor and Weil implies a negative effect of the gender wage gap on the FLP). However, surprisingly the gender wage gap in Mediterranean countries is not higher than in the remaining four countries in the sample (i.e. the group of countries that was labelled Rest in Figure 3 and 4). However, the gender wage gap could only explain the lower FLP in Mediterranean countries than in the remaining countries in the sample (as seen in Figure 3 and 4), if the gender wage gap were higher in Mediterranean countries than in the remaining countries in the sample. Using micro-level data, Olivetti and Petrongolo (2006) explain the surprisingly low gender wage gap in Mediterranean countries with non-random selection of women into work across countries. In particular, if women who are employed tend to have high-wage characteristics, then a low FLP in Mediterranean countries implies that fewer women with low-wage characteristics are part of the labour force, implying a higher weight of women with high-wage characteristics in the calculation of the average female wage.

Because of the lack of significant of the gender wage gap for the FLP in 2SLS regressions, the present study uses two alternative instruments for the FLP. The first instrument is the natural logarithm of the proportion of public employment in total employment. This is motivated by the fact that women seem to have a preference for clerical occupations and similar occupations, for which the government has particular demand (see, e.g., also Cavalcanti and Tavares, 2003, and Hakim, 2000). For this reason, a large proportion of public employment in total employment should have a positive effect on the FLP. In the framework of Galor and Weil (1996) large labour demand of the government increase the FLP by increasing the relative wage of women and therefore encouraging women to devote more time to labour market participation. Since there might be a feedback effect from a high FLP to high labour demand of the government, the present study uses data from 1994 for the proportion of public employment in total employment. This variable is a lagged variable, because average values for 1995-2000 are used for the FLP. By definition a variable measured in 1995-2000 cannot influence a variable

¹⁴ Using the real male wage and the real female wage separately as two instruments showed also in 2SLS regressions insignificance of each of the two variables for the FLP.

Table 4: Second stage regression results of 2SLS with control variables, cross-country data, average 1995-2000.

Dependent variable	TFRadj
Constant	3.56
FLP	-1.70 (0.03)
Natural logarithm of purchased child care use	0.20 (0.00)
Female long-term unemployment rate	-5.01 (0.04)

Notes: ¹The p-values are in parentheses behind the values of the coefficients.

² The FLP is the instrumented variable and instruments are the proportion of government employment in total employment and the proportion of women who strongly agreed or agreed to the statement “it is just as important for a women to have a job as it is for a man” (according to Eurobarometer 1997)

that is measured in 1994 (although, admittedly, possible autocorrelation of the FLP and the TFR implies that one can never get completely rid of all endogeneity). As a second instrument for the FLP the present paper uses an index of the proportion of women who strongly agreed or agreed to the statement “it is just as important for a women to have a job as it is for a man” from Eurobarometer (1997). This variable is in the following table abbreviated with “attitudes towards female employment”. As this abbreviation already tells, the variable is supposed to reflect attitudes in the society towards female employment and should influence female labour force participation. Again, a feedback effect from the FLP towards attitudes is possible. However, since the attitudes have been reported in Eurobarometer in 1997, they should also measure lagged attitudes, which reduces the endogeneity problem.¹⁵

Table 4 shows the second stage regression results of the 2SLS regression with the control variables purchased child care use and female long-term unemployment included. It is evident that the results are almost identical to the results in Table 2. The only difference is the fact that the significance of the FLP increased slightly. This is most likely due to a reduction of measurement errors, which increases the value of the coefficient of the FLP slightly. Hence, the results are even slightly stronger with 2SLS regression than with OLS.

Table 5 shows the results of the first stage regression of the 2SLS regression of Table 5. Most importantly, the two instrument variables for the FLP are very significant with the expected positive signs of the coefficients. As is well-known to applied econometricians, in 2SLS regressions, the first stage regression has to include the aforementioned two control variables of the second stage regression. Of these control variables, the female long-term unemployment rate has a negative and significant effect, indicating that female long-term unemployment discourages female labour force participation (as is an implication in the model of Da Rocha and Fuster, 2006). The other control variable, purchased child care use has an almost significant and negative effect on the FLP. This result is surprising, because the aforementioned models of Apps and Rees (2004) and Martinez and Iza

¹⁵ There exists alternative indexes for attitudes towards female employment in the literature (see, e.g., Algan and Cahuc, 2005, and Fortin, 2005). However, I tried several such indexes and the index that I present in this study was the most significant index for the FLP in 2SLS regressions. In addition, once I controlled for this index, no other index for attitudes towards female employment was significant for the FLP in empirical exercises that are not shown.

Table 5: First stage regression results of 2SLS with control variables in second stage regression, cross-country data, average 1995-2000.

Dependent variable	FLP
Constant	0.03
Natural logarithm of purchased child care use	-0.02 (0.06)
Female long-term unemployment rate	-2.65 (0.00)
Natural logarithm of the proportion of public employment in total employment	0.11 (0.00)
Attitudes towards female employment	0.44 (0.00)

Note: The p-values are in parentheses behind the values of the coefficients.

(2004) imply that a low price of purchased child care simultaneously increases both, the TFR and the FLP. Regressing the FLP with OLS on the independent variables of Table 5 and, in addition, the TFRadj as a further independent variable increases the p-value of purchased child care use to 0.33 (results not shown). Hence, the almost significance of purchased child care use for the FLP in Table 5 is due to the necessary omission of the TFRadj in the first stage regression of the 2SLS regression. Nevertheless, even controlling for the TFRadj does not give the expected significant and positive effect of purchased child care use on the FLP. However, a closer look at the data of the present study explains the insignificant effect of purchased child care use on the FLP. Table 6 reveals that, on the one hand, Ireland's purchased child care use of 38% has a higher value than the average value within Western European countries of 23%, while its FLP of 58% has a lower value than the average value in Western European countries of 67%. On the other hand, the purchased child care uses of Austria and Portugal have with 4%, respectively, 12% lower values than the average value within Western European countries, while their FLPs of 74% have a higher value than the average value in Western European countries. Since the data variation of the other independent variables of Table 5 can explain the FLP of Ireland, Austria and Portugal quite well (data not shown), it is not surprisingly that for these three countries purchased child care use does not have a positive effects on the FLP. Apparently, the cross-country variations of purchased child care uses and of the FLPs of the other countries in the sample cannot offset this result that strongly that purchased child care use has a positive and significant effect on the FLP in the overall sample.

Table 6: Purchased child care use and the FLP in Ireland, Austria and Portugal in comparison to their average values within the sample.

	Purchased child care use (%)	FLP (%)
Ireland	38	58
Austria	4	74
Portugal	12	74
Average within sample	23	67

Sources: OECD Employment Outlook (2001, Table 4.7) (for purchased child care use) and Eurostat (2005a)(for the FLP).

5. Determinants of Purchased Child Care

Cross-country differences in purchased child care use for children below age three are usually interpreted as reflecting cross-country differences in availability of purchased child care. Indeed, data not shown reveal that purchased child care use for children of this age group is very high in Nordic countries and France, countries which are known for strong state support for purchased child care. However, purchased child care use for children of this age group is also very high in the UK and the US, two countries which are known for weak state support for purchased child care. Availability of purchased child care could be high in the UK and the US because of better working markets for purchased child care in these countries. However, Fagnani (2002) argues that differences in purchased child care use would not only reflect availability of purchased child care. Instead, it would also reflect differences in attitudes towards purchased child care. In addition, in case of Western Germany and France state support of purchased child care and attitudes towards purchased would mutually influence each other.

To check whether or not purchased child care use reflects both, state support for purchased child care and attitudes towards purchased child care, this section shows the results of an OLS regression of purchased child care use on these two variables. For comparability, the regression was applied to the same Western European countries as in the last section (however, with Ireland, due to lack of data, dropped from the regression). In addition, data for purchased child care use were, again as in the last section, data for 1995-2000. State support for purchased child is measured with public expenditures on family services as a share of GDP (with years of reporting varying between 1995 and 1998), using data from Bertelsmann Foundation (2006). Attitudes towards purchased child care are measured by an index of strong agreement or agreement to the statement that “a pre-school child is likely to suffer if his or her mother works” from the International Social Survey Programme (2002). Interviewed persons were women and men of all ages. The fact that for this index also the respond of interviewees above childbearing age was used reduces possible reverse causality from purchased child care to attitudes (as people above childbearing age probably partly built their attitudes before 1995).

Table 7 shows the results of the aforementioned regression of purchased child care use on a constant, on “public expenditures on family services as a share of GDP” and on the natural logarithm of attitudes towards purchased child care. For comparability, the coefficients in Table 7 are standardised coefficients. Since the standardised coefficient of the constant is undefined, the constant is not shown in the table. The standardised coefficients of both independent variables have about the same magnitude (with the expected signs of the coefficients). Hence, state support and attitudes towards purchased child care explain about the same proportion of cross-country differences in purchased child care use. The R^2 in Table 7 is with 0.87 extremely high. Both variables are only significant at the 10 percent level. However, this is without any doubt due to high multicollinearity between the two variables. Indeed, an OLS regression of “public expenditures on family services as a share of GDP” on the natural logarithm of attitudes towards purchased child care or the other way around gives the extremely

Table 7: OLS regression, cross-country data, average 1995-2000.

Dependent variable	Purchased child care use
Public expenditures on family services as a share of GDP	0.45 (0.09)
Natural logarithm of attitudes towards purchased child care	-0.53 (0.06)
R ²	0.87

Note: ¹ Coefficients are standardised coefficients.

² The regression contains a constant (not shown).

³ The p-values are in parentheses behind the value of the coefficient.

high R² of 0.72. This is consistent with the argument of Fagnani (2002) that state support of purchased child care and attitudes towards purchased child care would mutually influence each other. Alternatively or additionally, state support and attitudes towards purchased child care could be the result of common cultural roots or advancement of modern values.

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Appendix: Definitions und data sources

Adjusted total fertility rate (TFRadj)

Definition: Sum of age specific birth rates of women, in general of age 15-49, adjusted for tempo effects according to the method of Bongaarts and Feeney (average of 1995 and 2000 or for a few countries in the neighbourhood of this interval – see Sobotka, 2004).

Data source: Sobotka (2004)

Female labour force participation rate (FLP)

Definition: Sum of full or part time employed women of age 25-49 and as all unemployed registered women of age 25-49 divided by the total number of women of age 25-49 (average of 1995 and 2000).

Data source: Eurostat (2000a).

Natural logarithm of purchased child care use

Definition: Natural logarithm of the percent of children below three for which purchased child care is used (years of reporting varies between 1995 and 2000).

Data source: OECD (2001, Table 4.7).

Female long-term unemployment rate

Definition: Number of women of age above age 15 in the labour force who are unemployed since twelve months or longer divided by the sum of full or part time employed women above age 15 and all unemployed registered women of age above 15 (average of 1995-2000).

Data source: Eurostat (2005b)

Gender wage gap

Definition: Ratio of hourly male wage to hourly female wage in manufacturing.
Year for which measured: 1997.

Data source: Gauthier (2002), the data point for Finland is the average of the values of the other Nordic countries (Denmark, Norway and Sweden).

Natural logarithm of proportion of public employment in total employment

Year for which measured: 1994.

Data source: Adsera (2005).

Attitudes towards female employment

Definition: Proportion of women who strongly agreed or agreed to the statement "it is just as important for a women to have a job as it is for a man".

Data source: Eurobarometer (1997), the data point for Norway is the average of the values of the other Nordic countries (Denmark, Finland and Sweden).

Natural logarithm of attitudes towards purchased child care

Definition: Index of agreement to the statement that "a pre-school child is

likely to suffer if his or her mother works”.

Calculation of Index: Index was calculated by weighting the proportion of interviewees commenting with “strongly agree” with 2, weighting the proportion of interviewees commenting with “agree” with 1, weighting the proportion of interviewees commenting with “neither agree nor disagree” with zero, weighting the proportion of interviewees commenting with “disagree” with -1 and weighting the proportion of interviewees commenting with “strongly disagree” with -2.

Data source: International Social Survey Programme (2002), data for Greek and Italy are the average values of the other Mediterranean countries (Spain and Portugal).

Public expenditures on family services as a share of GDP

Definition: Public expenditures as a share of GDP on formal day care, personal services, household services and other benefits in kind.

Year for which measured: Years of reporting varies between 1995 and 1998.

Data source: Bertelsmann Foundation (2006), the data point for Norway is the average of the values of the other Nordic countries (Denmark, Finland and Sweden), while the data points for Greece and Portugal are the average of the values of the other Mediterranean countries (Italy and Spain).