A Pooled Time-Series Analysis on the Relation Between Fertility and Female Employment

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May 25, 2005

Abstract

Various authors find that in OECD countries the cross-country correlation between the total fertility rate and the female labour force participation rate turned from a negative value before the 1980s to a positive value thereafter. Based on pooled time series analysis the literature seems to agree that this change is due to unmeasured country and time heterogeneity with respect to female employment. However, the role of female employment for time and country heterogeneity remains unclear. Using data of 22 OECD countries from 1960-2000 we estimate pooled time series models of fertility and female labour force participation by applying random effects and fixed effects panel models as well as Prais-Winsten regressions with panel-corrected standard errors and autoregressive errors. Proceeding with Prais-Winsten regressions our empirical findings reveal substantial differences across countries and time periods in the effects of female employment on fertility. Initial increases in female employment strongly lowers fertility, but continued increases have a progressively less negative effect. The country heterogeneity in the effect of female employment can also be confirmed for different regions as well as for varying welfare and gender regimes.

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1 Introduction

The common wisdom of an inverse relation between fertility and female employment has been challenged in previous years. Several authors in a variety of disciplines have recently noted an aggregate reversal in the cross-country correlation between the the total fertility rate (TFR) and the female labour market participation rate (FLP) among OECD countries. Figure 1 illustrates this change for 22 countries.[1] The countries that now have the lowest levels of fertility are those with relatively low levels of female labour force participation and the countries with higher fertility levels tend to have relatively high female labour force participation rates.

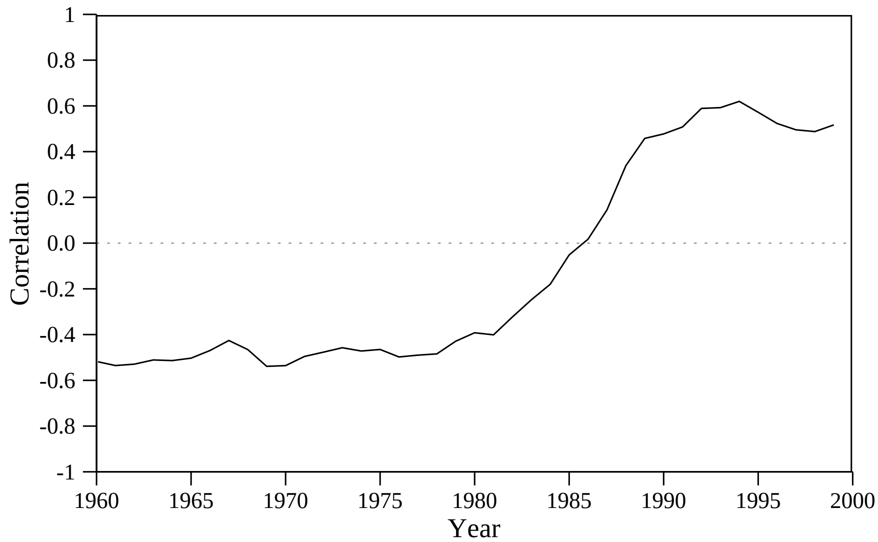
This change in the sign of the cross-country correlation between TFR and FLP has often been mistakenly associated with a change in the time series association between TFR and FLP (Benjamin 2001; Brewster and Rindfuss 2000; Esping-Andersen 1999; Rindfuss et al. 2003). A recent study by Engelhardt et al. (2004a) shows that neither the causality nor the time series association between TFR and FLP has changed over time. By applying error-correction models to six industrialised countries the paper finds Granger causality in both directions, which is consistent with simultaneous movements of both variables brought about by common exogenous factors.

Though this study provides econometric evidence that the time series association of single countries has not changed its sign, it does not investigate the factors that may actually explain the change in the cross-country correlation coefficient. Adsera (2004), Ahn and Mira (2002), Benjamin (2001), Pampel (2001), Castles (2003), de Laat and Sevilla-Sanz (2003), Kögel (2004), and Engelhardt and Prskawetz (2004b) offer some theories and data that may explain why the sign of the cross-country correlation changed. The empirical analyses by Engelhardt and Prskawetz, Ahn and Mira, Castles, and de Laat and Sevilla-Sanz remain on a descriptive bivariate level. The studies by Benjamin, Pampel, Adsera and Kögel include multivariate analysis based on pooled cross-section time series. Table 1 provides a summary of the data, variables, and methods used in these multivariate studies, as well as of the respective results.

Benjamin (2001) presents an extensive discussion on factors that may

[1] The countries included are Australia, Austria, Belgium, Canada, Denmark, Finland, France, West Germany, Greece, Ireland, Italy, Japan, Luxembourg, Norway, the Netherlands, New Zealand, Portugal, Spain, Sweden, Switzerland, United Kingdom, and the United States.

Figure 1: Cross-country correlation between the total fertility rate and female labour force participation rate, 1960-2000



cause the reversal of the cross-country correlation coefficient. To account for cross-sectional heteroscedasticity and time-wise autocorrelation Benjamin applied Prais-Winsten regressions to pooled time series, thereby neglecting possible non-stationarity of the variables. She finds that the relationship between female labour participation and fertility becomes positive over time, although the timing of this shift depends on the country group (broadly reflecting role incompatibility).

Pampel (2001) estimates the effect of female employment and country dummies by applying a fixed effects GLS model which also adjusts for autocorrelation and heteroscedasticity. Like Benjamin, Pampel uses all variables in levels, thereby disregarding the possibility of non-stationarity of the data. Averaged across all nations he finds negative effects of female employment on fertility depending on class and gender equality of the respective country group. Moreover, Pampel finds that initial increases in female labour participation strongly lower fertility, but continued increases have a progressively less negative influence on fertility.

Adsera (2004) explicitly tests for non-stationarity and estimated the effects of labour market arrangements on pooled fertility rates using levels,

Table 1: Summary of pooled time series studies

Author	Method	Data	Variables	Results
Benjamin (2001)	Prais-Winsten OLS	21 countries, 1970-95, annual	all in levels	change in relationship, timing dependent on context
Pampel (2001)	fixed effects GLS	18 countries, 1951-94, annual	all in levels	negative relationship weakened over time, dependent on context
Adsera (2004)	random effects GLS	23 countries, 1960-97, annual	levels, logs, first differences	indirect evidence that change in corr. is due to labour market contexts
Kögel (2004)	fixed effects OLS, random effects GLS	21 countries, 1960-2000, quinquennial	all in logs	no change in relationship, unmeasured country-specificity, heterogeneity in association

logs, and first differences of the variables depending on the test results. By applying a random effects model Adsera ignores any country heterogeneity. By regressing different labour market indicators on fertility, Adsera does not explicitly address the question of the factors behind the change in the correlation. She finds indirect evidence that labour market institutions shaped the changing correlation.

To avoid serial correlation of the data Kögel (2004) uses quinquennial data (i.e., only the data points 1960, 65, 70, 75, 80, 85, 90, 95 and 99). Since the data are not difference stationary, Kögel applies all variables in logarithms. Kögel shows that the time series association between TFR and FLP has not changed and offers two convincing elements that may explain the change in the cross-country correlation. These are the presence of unmeasured country-specific factors and country heterogeneity in the magnitude of the negative time-series association between fertility and female employment.

As this short review demonstrates, the literature is quite heterogenous in terms of data, variables, and methods applied to the same research question. Concerning the results, most of the empirical evidence points in the direction that the change in correlation is due to unmeasured country-specific factors and country and time heterogeneity with respect to female employment. The exact role of female employment, however, remains unclear.

The aim of our paper is to study the role of female employment for time and country heterogeneity that explains the change in the sign of the cross-country correlation coefficient. We begin similar to Pampel (2001, chapter 5) by describing the relationship between female labour force participation and fertility across all time periods and nations and then continue to describe differences in the relationship across time and across nations. Unlike Pampel, however, we apply Prais-Winsten estimations to take care of the high serial correlation in our data. We also model time heterogeneity in more detail as compared to the setting in Pampel. To test for country heterogeneity in the relation between female labour force participation and fertility we apply alternative country groupings that classify countries according to their welfare state regimes (Esping-Andersen 1990), family policies (Gauthier 1996) and their regional clustering. The latter grouping takes care of spatial autocorrelation and most importantly also singles out Southern Europe as a distinct group. As discussed in Dalla Zuanna and Micheli (2004) and Bettio and Villa (1998) lowest low fertility in Southern Europe took place in a context where there is low compatibility between childbearing and labour market participation owing to less flexible working hours and difficulties in re-entering the labour market after child-birth and at the same time where strong values and social norms towards the good family are still prevalent. The relevance of Mediterranean countries to explain the change in the cross-country correlation between fertility and female labour force participation is thereby taken care off (cf. Ahn and Mira 2002). The grouping by Esping-Andersen and Gauthier rather refer to the explanation of Brewster and Rindfuss (2000) who contribute the reversal in the sign of the cross-country correlation between female labour force participation and fertility to temporal and spatial heterogeneity in institutional arrangements such as differences in family policies, childcare-systems, welfare-state systems and norms towards the combination of childrearing and labour force behaviour of females. Since many of those institutional factors, and in particular representations of norms and values, are not available as cross-country time series, our approach to group countries to reflect common trends in these variables may indirectly help to discern the determinants of the temporal and spatial heterogeneity that led to the reversal in sign of the cross-country correlation of fertility and female labour force participation.

The paper is organised as follows: In the ensuing section we discuss the variables and econometric methods. Section three compiles the pooled time series analysis for the basic model selection and for the analysis of time and country heterogeneity with respect to the female employment. We close with a short discussion and an outlook for future research.

2 Data and Methodology

In the empirical analysis we assembled annual time series of the total fertility rate and women's labour force participation rate from 1960 to 2000 for 22 OECD countries. The countries included are Austria, Australia, Belgium, Canada, Denmark, Finland, France, West Germany, Greece, Italy, Ireland, Japan, Luxembourg, Norway, the Netherlands, New Zealand, Portugal, Spain, Sweden, Switzerland, United Kingdom, and the United States.

The total fertility rate is defined as the average number of children that would be born alive to a woman during her lifetime if she experiences a given set of age-specific fertility rate observed in a population during a given year. The data are compiled from United Nations Demographic Yearbook, New Cronos (Eurostat Database), and the German Federal Statistical Office. The female labour participation rate is defined as the number of females working part- or full-time or actively seeking employment at ages 15-64 divided by the total female population aged 15-64. The source of our data is the Comparative Welfare Data Set assembled by Huber et al. (1997) and the OECD Labor Force Statistics.

Our methodological approach is to pool cross-sectional time series. This technique incorporates both the cross-sectional effect of the independent variables on fertility as well as the time-series effects within nations. The critical assumption of pooled cross-sectional times series models is that of "pooling". That is, all units are characterised by the same regression equation at all points in time:

$$y_{it} = x'_{it}\beta + \varepsilon_{it}, \quad i = 1, \ldots, N; t = 1, \ldots, T \tag{1}$$

where y_{it} and x_{it} are observations for the ith unit at time t and β is a vector of coefficients. ε_{it} is the residual with the usual properties (mean 0, uncorrelated with itself, uncorrelated with x, and homoscedastic).

Pooled time series can be difficult to estimate. As Hicks (1994, p. 172) notes, "errors for regressions equations estimated from pooled data using OLS [ordinary least squares regression] procedures tend to be (1) temporally autoregressive, (2) cross-sectionally heteroscedastic, and (3) cross-sectionally correlated as well as (4) conceal unit and period effects and (5) reflect some causal heterogeneity across space, time, or both".

2.1 Temporal and Spatial Heterogeneity

To deal with causal heterogeneity across space, often fixed country effects are assumed. Formally, the fixed effects model is given by:

$$y_{it} = x'_{it}\beta + \nu_i + \varepsilon_{it}, \tag{2}$$

where ν_i are assumed to be fixed parameters which may be correlated with x_{it}. Such a model focuses on the within-country variation, and the coefficients represent a cross-country average of the longitudinal effect. Time effects γ_t, in contrast, capture developments over time that are common to all countries. Combining both country and time intercepts in a single specification results in a model from which all unobserved country- and time-specific effects are removed:

$$y_{it} = x'_{it}\beta + \nu_i + \gamma_t + \varepsilon_{it}, \tag{3}$$

If the unobserved country- or time-specific heterogeneity, however, can be assumed to be realisations of a random process and uncorrelated with the included variables, then the model is a random effects model. Thus, the crucial distinction between the fixed and the random effects model is whether the unobserved country- and time-specific effect embodies elements that are correlated with the regressors in the model (Greene 2003).

Whether the fixed or random effects model should be used is both a substantial and statistical question. If there is no substantial reason to assume a significant correlation between the unobserved country-specific random effects and the regressors, then the random effects model may be more powerful and parsimonious. If there is such a correlation, the random effects model would be inconsistently estimated and the fixed effects model would be the model of choice. The Hausman specification test is the classical test for statistical model selection.

2.2 Autocorrelation

Both random and fixed-effects panel models do not deal explicitly with temporally and spatially correlated errors often contained in pooled time series models. If there is autocorrelation in the model, it is necessary to deal with it because autocorrelation in the residuals causes seriously inefficient estimates. One can apply one or more of the several tests for residual autocorrelation, for instance the modified Durbin-Watson test for first-order autocorrelation in the residuals by Baltagi and Wu (1999) to handle unbalanced panel and equally spaced data. An autoregression on lags of the residuals may indicate the presence or absence of autocorrelation and the need for dynamic panel analysis (Greene 2003).

In principle, there are three ways to deal with autocorrelation. On the one hand, autocorrelation is regarded as a nuisance in the residuals that has to be corrected. On the other hand, autocorrelation may indicate persistency in the dependent variable that can be captured by modelling an autoregressive process including a lagged dependent variable. The literature refers to the latter approach as the "dynamic model" and to the former as the "static model" (Beck and Katz 1996). Finally, autocorrelation in pooled time series can be seen as the result of unit roots in the single series which can be corrected for by differencing the series.

Most commonly the static approach is used where the nuisance in the residuals is modelled as a first-order autoregression or AR(1) process:

$$\varepsilon_{it} = \rho\varepsilon_{i,t-1} + \eta_{it}, \tag{4}$$

where η_{it} are independent and identically distributed with mean 0 and ρ is the so-called autocorrelation parameter, which is less than one in absolute values.[2]

The static model is usually estimated by "feasible generalised least squares" (FGLS). This method proceeds by first estimating equation (1) by OLS and then using the residuals from this estimation to estimate ρ in equation (4). This estimate of ρ is used to transform the data and the transformed model can be estimated by OLS. The estimator by Prais and Winsten (1954) transforms the data as follows:

$$y^*_{it} = \begin{bmatrix} \sqrt{1-\hat{\rho}^2}y_{i1} \\ y_{i2} - \hat{\rho}y_{i2} \\ y_{i3} - \hat{\rho}y_{i3} \\ \vdots \\ y_{iT} - \hat{\rho}y_{iT} \end{bmatrix}, \quad x^*_{it} = \begin{bmatrix} \sqrt{1-\hat{\rho}^2}x_{i1} \\ x_{i2} - \hat{\rho}x_{i2} \\ x_{i3} - \hat{\rho}x_{i3} \\ \vdots \\ x_{iT} - \hat{\rho}x_{iT} \end{bmatrix}. \tag{5}$$

In contrast, the Cochrane and Orcutt (1949) estimator omits the first observation. In terms of the transformed data, the model is now only het-

[2] Higher-order models are often constructed as a refinement of AR(1) processes. The first-order autoregression is a reasonable model for impenetrably complex underlying processes (Greene 2003: 257). Some analysts allow for unit specific ρ_i. Beck and Katz (1995: 121) make a case against assuming a unit-specific autoregressive process.

eroscedastic; the transformation has removed the autocorrelation (Greene 2003: 325).

In the empirical approach we shall only consider the static approach and we therefore refer the reader to Arellano and Bond (1991) and Arellano and Bover (1995) for the dynamic approach to deal with autocorrelation.

3 Empirical Results

In the first part (section 3.1) of our empirical results we consider the relation between TFR and FLP across all time periods and nations. We use this first step of our analysis for model selection. In the second part of our analysis we allow for heterogeneity in the relationship between TFR and FLP across time (section 3.2) and across countries (section 3.3).

3.1 Basic Model Selection

Table 2 summarises the estimation results if we apply the level of the variables and use alternative panel data estimations. The first column in this table shows the results of between-group estimation. In this case all data are converted into country-specific time averages and OLS is applied to these transformed data. In case of between-group estimation one does not use time-series information to account for country effects. Hence, the between estimator only uses the cross-sectional information in the data. The second column gives the results for the pooled least -squares estimation with fixed country effects (approximated with a dummy variable for each country).[3] Fixed country effects estimation is identical to within-group estimation, i.e., a pooled least-squares regression on the deviation of each variable from its time series average. In contrast to the between estimator the fixed effects estimator represents the time series association between the TFR and FLP within each country. The third column shows the results of generalised least-squares estimation with random country effects.[4]

[3]Fixed effect models and random effects models, respectively, control for country effects and assume that the time series association between TFR and FLP is the same across countries.

[4]As is well known from the econometrics literature, the random effects estimator is a matrix-weighted average of the between and fixed effects estimator and therefore contains some cross-country information in addition to the time series association between TFR and FLP.

Table 2: Comparison of estimated coefficients from between effects, random effects, and fixed country effects model; t-values in paranthesis

$\mathrm{TFR}_t = \alpha + \beta \mathrm{FLP}_t$	BE	FE	RE
β	-0.007	-0.044***	-0.042***
	(-1.085)	(-28.856)	(-27.995)
α	2.404***	4.287***	4.187***
	(6.629)	(53.982)	(40.938)
R^2	0.242	0.242	0.242
BIC	11.249	616.663	734.869
F(df, n)	1.178	832.696***	
Wald			783.702***
Chow		42.23***	
Breusch Pagan			2365.98***
Hausman	26.25***	72.02***	
ρ		0.980	0.980
Baltagi-Wu LBI		0.143	0.143
Wooldridge		37.582***	

Notes: BE requests the OLS-estimator for the between-effects model; RE requests the GLS estimator of the random-effects model; FE requests the within OLS-estimator of the fixed country effects model. BIC = –2 log-likelihood + log(N) p, where p is the number of parameters of the model, and N is the number of observations (for RE estimated by MLE). F(df, n)-test on poolability of the data, $H_0 : \beta_i = \beta$. Wald performs a χ^2-test for $H_0 : \beta = 0$. Chow test on absence of fixed country effects, $H_0 : \nu_i = 0$. Breusch Pagan lagrange multiplier test on absence of random effects: $H_0 : var(\nu_i) = 0$. Hausman test: H_0: difference in coefficients between BE or FE and RE model is not systematic. ρ is the estimated autocorrelation coefficient from the AR(1) model. LBI is the Baltagi-Wu (1999) locally best invariant test statistic from the AR(1) model, $H_0 : \rho = 0$. If LBI is far below 1.5 we have positive serial correlation. Wooldridge test for serial correlation from the regression of the first-differences variables, H_0 : no first-order autocorrelation. *** $p \leq .001$, ** $p \leq .01$, * $p \leq .05$, † $p \leq .10$.

The results in Table 2 indicate that independent of the specific estimation procedure the association between TFR and FLP was negative. Only in case of the between-group estimation the negative association is not significant. Since the fixed effects estimator gives the expected change in the TFR within each country if FLP changes by one unit this result demonstrates that the time series association between TFR and FLP does not change its sign. On the other hand, the between estimator indicates the expected effect of a unit change in the independent variable FLP on the value of TFR between two countries. The latter estimate which is not significant may therefore explain why the sign in the cross-country association between TFR and FLP is reversed (see Kögel 2004).

Moreover, Table 2 shows results of the Hausman test with H_0: the differences in the estimated coefficients between the fixed effects model and the random effects model are not systematic. Since the test rejects the null hypothesis and the model fit of the fixed effects model (BIC) is better than the fit of the random effects model, this suggests that fixed effect estimations are more appropriate than random effect estimations. In addition the test for poolability of the data indicates that it is not appropriate to assume a common constant coefficient for FLP in case of the fixed effects model (in section 3.3 we will therefore allow for country effects in the time series association between TFR and FLP). Moreover, the null hypothesis of absence of country effects can be rejected in case of the fixed effects estimations (Chow test) as well as in case of the random effects model (Breusch-Pagan test).

Independent of the estimation procedure, additional calculations shown in the lower panel of Table 2 indicate high serial correlation of our data when estimating a model with first-order autocorrelation or performing the Wooldridge test from the regression of the first-differences variables (Wooldridge 2002). In a next step we therefore apply the Prais-Winsten estimator with AR(1) disturbance terms as described in section 2.2.

In Table 3 we summarise the estimation results if we apply the Prais-Winsten model. In addition we also present results if we apply the FLP lagged by one year as proposed in Pampel (2001). The argument to use the lagged independent variable is based on micro foundations arguing that a lagged FLP variable may prevent reverse causality.

Including the lagged value of the independent variable (column (2)) does not change the results compared to applying our estimations to contemporaneous values of TFR and FLP (column (1) in Table 3).

The most important message of Table 3 is that by applying the Prais-

Table 3: Fixed country effects Prais-Winsten estimations with panel-corrected standard errors and AR(1) disturbances for different model specifications; t-values in parenthesis

Model	(1)	(2)
β	-0.026***	-0.025***
	(-6.526)	(-5.934)
α	3.742***	3.682***
	(14.877)	(14.120)
R^2	0.737	0.742
Wald	257.266***	344.177***

Notes: Model (1) $\text{TFR}_t = \alpha + \beta\text{FLP}_t$; (2) $\text{TFR}_t = \alpha + \beta\text{FLP}_{t-1}$. Wald performs a χ^2-test for $H_0 : \beta = 0$. ρ is the estimated autocorrelation coefficient. Wooldridge test for serial correlation from the regression of the first-differences variables, H_0 : no first-order autocorrelation. *** $p \leq .001$, ** $p \leq .01$, * $p \leq .05$, † $p \leq .10$.

Winsten estimation with AR(1) disturbance terms increases the fit of the model. In addition, the coefficient on FLP slightly decreases from -0.044 to -0.028. We have also transformed the variables (e.g., logarithmic transformation and first differences of FLP and TFR) and applied the Prais-Winsten estimation with AR(1) disturbance terms. However these transformations have not resulted in further improvements of our model criteria.[5]

From the comparison of alternative models we may conclude that the most appropriate among the models considered is the Prais-Winsten estimation with AR(1) disturbance terms applied to the levels of the TFR and FLP (first column in Table 3). This model will be the starting framework for the next section where we allow for time and country heterogeneity in the time series association between TFR and FLP.

3.2 Time Heterogeneity

As indicated by Kögel (2004) and Pampel (2001) the coefficient on FLP as presented in Table 3 represents an average across all nations and does not account for country heterogeneity in the magnitude of the negative time series association between fertility and female employment. However, as noted by Kögel (2004, p. 11) the latter argument may explain the reversal in

[5]These results are available from the authors on request.

the cross-country correlation which cannot be explained by the models that only account for the presence of unmeasured country-specific factors. In the following we shall test for country- and time-specific effects of the female participation rate.

In a first step we investigate the appropriate representation of the time effect, i.e., we successively choose more detailed split-ups of the time period which are presented by effect-coded dummy variables (results not shown here).[6] Independent of the specific time dummy applied we find that compared to the average value of the FLP effect over all nations and time periods there exists an independent time effect which was positive during the 60s and early 70s and negative during the mid 70s up to the end of the 20th century (Table 4, column 1 and 2). Not only does the sign of the coefficients on the time dummies switches around 1974 from positive to negative, but also the quantitative effects are reduced. As our results indicate, a split-up of the time dimension in the period before 1985 and thereafter (as suggested by the change in the cross-country correlation coefficient in Figure 1) would not be correct.[7]

We next study the effect of the female participation on fertility for the different time periods, i.e., we include interaction effects between the FLP variable and each time dummy (Table 4, column 3). The fact that the coefficients on the time dummies are reduced between model (1) and model (2) already indicates an association between time and FLP trends. A comparison of the estimated slope coefficient of FLP between the Prais-Winsten estimation in Table 3, column (1) and model (2) in Table 4 indicates that part of the negative effect of the female labour force participation works through the time specific effect.[8] Moreover, as the value of R^2 indicates, the inclusion of time dummies slightly improves the fit of the model.

The results in model (3), which includes the interaction between the time dummies and the FLP variable, confirm the hypothesis that the negative

[6]The coefficients for all effect-coded periods are obtained by two separate estimation procedures with changing omitted categories.

[7]Kögel (2004) for instance uses such a split-up and applied separate estimations for the data from 1960-85 and 1985-2000.

[8]Pampel (2001, p. 105, Table 13) finds a different result. Controlling for time increases the effect of the female participation rate compared to not including any time effects because Pampel's set of countries excluded Luxembourg, Greece, Portugal, and Spain. Especially the three latter countries have shown the fastest decline in TFR accompanied by rather modest increase in FLP.

Table 4: Effects of female labour participation and its time interaction on fertility; Prais-Winsten regressions with panel-corrected standard errors and AR(1) disturbances

Model	(1)	(2)	(3)	
			Main effect	FLP_t*time
FLP_t		-0.015***	-0.016***	
1960-1964	0.435***	0.342***	1.108***	-0.014***
1965-1969	0.294***	0.209***	0.850***	-0.012***
1970-1974	0.134**	0.085†	0.542***	-0.008***
1975-1979	0.042	-0.063	0.132	-0.003
1980-1984	-0.122**	-0.107**	-0.172	0.002
1985-1989	-0.204***	-0.156***	-0.593***	0.006***
1990-1994	-0.216***	-0.133**	-0.846***	0.013***
1995-2000	-0.280***	-0.178**	-1.021***	0.015***
Constant	2.083***	2.879***	2.805***	
R^2	0.747	0.765	0.798	
χ^2	44.955***	83.902***	182.545***	

Notes: *** $p \leq .001$, ** $p \leq .01$, * $p \leq .05$, † $p \leq .10$.

effect of FLP on fertility became smaller over time. By summing up the main effect of FLP of -0.016 and the estimates on the interaction term $FLP_t * \text{time}$ we may summarise the effect of female labour force participation rate on total fertility for selected time periods. For instance, the negative effect of FLP on TFR declined from -0.030 in the time period 1960-1964 to -0.019 in the time period 1975-1979 and to -0.001 in 1995-2000. The coefficient in the last period reaches only about 1/30 of the size of the coefficient in the first time period.[9]

3.3 Country Heterogeneity

In order to find common patterns of fertility and female employment among groups of countries, we classify the countries according to Esping-Andersen's welfare state regimes, Gauthier's family policy types, and according to geographical criteria.

Esping-Andersen (1990) outlines three types of Western, industrialised welfare state regimes, based on their provision of social rights, contribution

[9] Pampel (2001, p. 106) found a reduction in the negative effect of FLP on TFR between 1951-61 and 1984-94 of 50 per cent only.

to social stratification, and nexus of state-market-family relations. He classifies welfare states into liberal, conservative, and social-democratic regimes. Liberal regimes are characterised by heavy dependence on the market for economic security and means-tested welfare benefits for those who are unsuccessful in the market. Conservative regimes provide social provision for all citizens, but their social policies enforce status and class distinctions. The social-democratic regimes provide all citizens with state provisions and with a minimum income. These are the most egalitarian welfare states. The welfare state grouping according to Esping-Andersen (1990) of our countries is as follows: conservative regimes (Australia, Austria, Belgium, France, Germany, Greece, Italy, Japan, Luxembourg, Portugal, Spain, Switzerland), social-democratic regimes (Denmark, Finland, Netherlands, Norway, Sweden), liberal regimes (Canada, Ireland, United Kingdom, USA, New Zealand).

In a study on family policy Gauthier (1996: 203ff) presented a historical review of the development of family policy in OECD countries. She clustered countries into four different groups. First, in countries belonging to the pro-family/pro-natalist model the major concern is low fertility and because of this the main task of family policy is to encourage families to have children. This is done by helping mothers reconcile work and family life. In this model, relatively high levels of support are provided for maternity leave and child-care facilities. Great emphasis is placed on cash benefits and more particularly, towards the third child. In the second, pro-traditional model the preservation of the family is the main concern. Government takes some responsibility for supporting families, but the most important sources of support are seen as the families themselves and voluntary organisations. Under this model, a medium level of state support for families is provided. The low provision of childcare does not give women the opportunity to combine employment and family responsibility easily. The third, pro-egalitarian model seeks to promote gender equality. Men and women are treated as equal breadwinners and equal carers and policy aims to support dual parent/worker roles. Liberal policies on marriage, divorce and abortion mean that there are few restrictions on how people can choose their family life. Fourth, in the countries belonging to pro-family but non-interventionist model the main concern is the families in need. The participation of women in the labour force is not discouraged, but limited benefits are provided by the state to support them. Families are viewed as basically self-sufficient and able to meet their own needs through the private market with only limited help from the state. It is believed that the market will meet any needs that emerge, as long as it is not hindered by government regulation. In our study, we applied Gauthier's classification system as follows: Pro-family/pro-natalist countries (Belgium, France, Luxembourg), pro-traditional countries (Australia, Austria, Germany, Greece, Italy, Japan, Portugal, Spain, Switzerland), pro-egalitarian countries (Denmark, Finland, Netherlands, Norway, Sweden), and non-interventionist countries (Canada, Ireland, United Kingdom, USA, New Zealand).

A comparison of the country groupings by Esping-Andersen and Gauthier reveals that the latter grouping singles out Belgium, France and Luxembourg as separate countries representative of a Pro-family/pro-natalist model. The other groups coincide among these two settings as follows: countries of the liberal model in Esping-Andersen classification conform to countries of the pro-family but non-interventionist model in the classification by Gauthier, while countries of the conservative and the social-democratic models in Esping-Andersen grouping correspond to countries of the pro-traditional and pro-egalitarian model, respectively, in Gauthier's classification.

Neither in Esping-Andersen nor in Gauthier are Southern European countries singled out as a distinct model, although Gauthier points out that in her analysis of benefit levels Southern European countries were placed in a separate category. To capture also regional differences and in particular to account for the distinct role of Mediterranean countries in explaining the change in the sign of the cross-country correlation we additionally clustered the countries into six groups according to their geographical location: north European countries (Finland, Norway and Sweden), south European countries (Greece, Italy, Portugal, Spain), west European countries (Belgium, France, Luxembourg, Netherlands), central European countries (Austria, Germany, Switzerland), other European countries (United Kingdom, Ireland, Denmark), and non-European countries (Australia, Canada, USA, New Zealand, Japan).

The regional country grouping coincides closely with the groupings by Gauthier. The groups that cover central and respectively south European countries consist of countries that have been classified as countries where the pro-traditional models are prevalent. The group of western European countries coincides with the pro-family/pro-natalist type of countries and additionally also includes the Netherlands which has been classified rather as a pro-egalitarian type of model. The group of northern European countries covers countries of the pro-egalitarian type. The group of other European countries includes one country of the pro-egalitarian model (Denmark) while the remaining countries are of the pro-family/non-interventionist type. The group of non-European countries covers countries of pro-traditional and pro-family/non-interventionist type. As compared to Gauthier's classification the country setting allows for a more detailed specification of country heterogeneity. We are particularly interested in the separate group of southern European countries that is merged with the central European countries in Gauthier's classification.

In the following we test for country-specific effects in the slope of FLP applying each of the three country typologies.[10] The results of our estimations with regional country groups are summarised in Table 5. As model (1) shows, the northern and other European countries have a fertility rate above average while the southern, western and central European countries have a below average fertility rate. The interaction between regional country groups and FLP can be seen in model (2). In the southern and other European countries the negative effect of FLP becomes stronger while the effect is not significantly reduced in northern, western and central European countries. These latter effects become significant under additional consideration of time (model 3). Moreover, under control of time heterogeneity in the effect of FLP, we find again a significant negative effect of FLP during the whole period of time.

The logic of separate time and region interactions might further suggest more complex interactions: for instance, the modifying effect of regions on the female participation rate might change over time. This implies interaction among region, time, and female participation in determining fertility and inclusion of three-way interaction terms. Due to the inflation of parameters, we estimated the model separately for each country group, thereby evading three-way interaction terms. The coefficients of the separately estimated models are identical with a single model including three-way interaction terms.[11] The results of the models with dummy trend terms are depicted in Figure 2. The figure shows that the widely varying effect of FLP in the early 1960s narrowed at the end of the 1990s. The results for the single country groups differ, however. For central European countries and western European countries, the effect of FLP on TFR was less negative

[10]Note that fixed country effects are excluded in all regressions that include dummy variables for country groupings. The decision was made based on the fact that inclusion of many time-invariant variables reduces the significance of these variables.

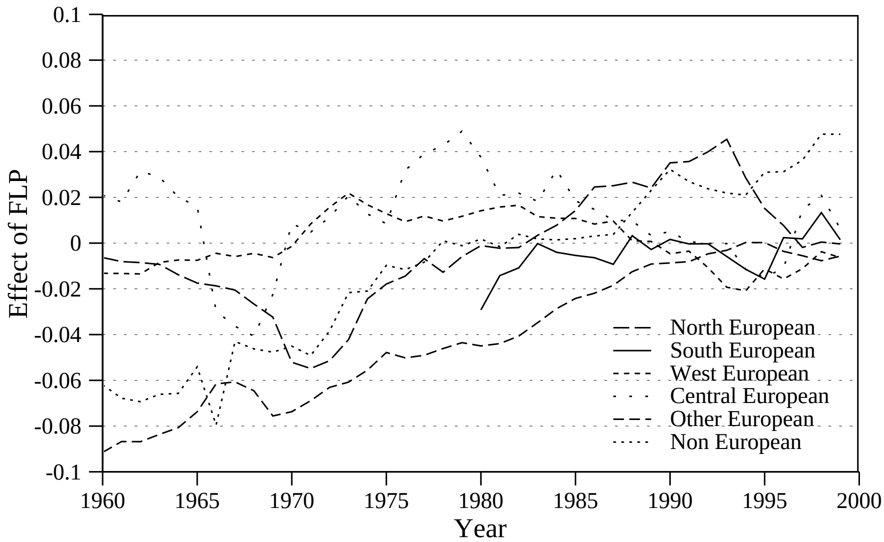
[11]Our results are based on the following equation: $\mathrm{TFR}_t = \alpha + \beta \mathrm{FLP}_t + \gamma \mathrm{region} + \delta \mathrm{FLP}_t * \mathrm{region}$.

Table 5: Regional country grouping; Prais-Winsten regressions with panel-corrected standard errors and AR(1) disturbances

Model	(1)	(2)	(3)
FLP_t	-0.026***	-0.027***	-0.025***
North	0.268***	-0.198	-0.574***
South	-0.521***	0.025	0.279
West	-0.217***	-0.537**	-1.103***
Central	-0.142**	-0.711*	-0.108
Other European	0.307***	0.794***	0.985***
Non-European	0.306***	0.627*	0.520**
FLP_t*North		0.008*	0.013***
FLP_t*South		-0.013*	-0.016**
FLP_t*West		0.008*	0.019***
FLP_t*Central		0.011†	0.000
FLP_t*Other European		-0.008*	-0.012***
FLP_t*Non-European		-0.005	-0.005†
1960-1964			1.094***
1965-1969			0.870***
1970-1974			0.565***
1975-1979			0.112
1980-1984			-0.229*
1985-1989			-0.673***
1990-1994			-0.863***
1995-2000			-0.875***
FLP_t*1960-1964			-0.014***
FLP_t*1965-1969			-0.013***
FLP_t*1970-1974			-0.009***
FLP_t*1975-1979			-0.003†
FLP_t*1980-1984			0.003
FLP_t*1985-1989			0.010***
FLP_t*1990-1994			0.014***
FLP_t*1995-2000			0.013***
Constant	3.226***	3.003***	3.050***
R^2	0.724	0.747	0.853
Wald χ^2	118.325***	208.585***	786.568***

Notes: *** $p \leq .001$, ** $p \leq .01$, * $p \leq .05$, † $p \leq .10$.

Figure 2: Effects of female labour participation over time for different regional country groups



which may be explained by the fact that female labour force participation in those countries did not change much until the end of the 1970s (for western European countries) or until the end of the 1980s (for central European countries), respectively, and the FLP in those countries is still one of the lowest among European countries. Moreover, among the western European countries we observe the highest TFR among countries that also experienced the highest FLP (e.g., France). For the set of other European countries and non-European countries the negative effect of FLP has been more pronounced which may partly be explained by the fact that those countries experienced the strongest increase in FLP from sometimes a value of only slightly above 30 per cent in 1960 up to more than 70 per cent in 2002 (e.g., Canada). At the same time these groups also include countries that had one of the highest TFR in the 1960s (close to 4 for Canada in 1960). Among the countries included in the group of other European countries and non-European countries, the rank order is similar for TFR and FLP. Put differently, those countries that experienced the strongest increase in FLP also experienced the strongest drop in TFR. For northern European countries the effect of FLP on TFR is rather non-monotonic mirroring the "roller-coaster" fertility pattern observed in those countries (this holds particularly for Sweden). The first decline in TFR during the 60s and early 70s was followed by an increase in FLP that lagged the TFR fall by about 5 years. The continued increase in FLP during the late 1980s and early 1990s, however, was accompanied by a "baby boom" in the early 1990s which was particularly pronounced in Sweden. For southern European countries the effect of FLP on TFR is among the most negative ones. In those countries a modest increase of FLP went together with the most pronounced decrease in TFR among all European countries reaching values close to one during the late 1990s while the FLP is close to 50 per cent in 2002 (except for Portugal where it has always been above the Southern European average and peaked at close to 68 per cent in the early years of this century).

In Table 6 we apply the country grouping suggested by Esping-Andersen. Our results indicate a significantly below average fertility for conservative welfare state regimes and an above average fertility level for social-democratic and liberal regimes (model 1). These results coincide with those obtained in Table 2 since social-democratic and liberal regimes cover northern European and non-European countries for which we found a similar result. However we find that the country heterogeneity in the slope coefficient of FLP is insignificant for all three considered country groups (model 2). These results remain when we additionally control for heterogeneity across time periods in model (3). However, the results change when we introduce a three-way interaction between female employment, time, and welfare regime.

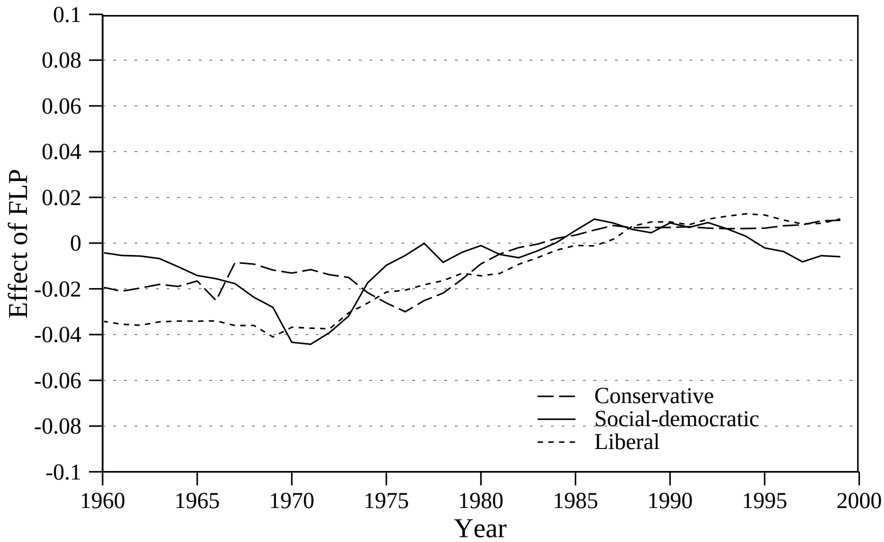
Figure 3 displays the effects of FLP over time for the different welfare regimes by applying piecewise (dummy) trend terms in separate estimations for each welfare state regime. While we find negative effects of FLP for liberal countries which seems to have been stalling since the beginning of the 1980s, the results for social-democratic countries differ greatly. The separate models with dummy trend terms reveal a roller-coaster course of FLP on TFR mainly caused by the roller-coaster course of TFR in northern European countries that was accompanied by a steady increase in FLP (cf. Figure 2). The relatively small negative effect of FLP on TFR for the group of conservative countries reflects the experience of the central European countries as already discussed in Figure 2. A comparison between Figure 2 and Figure 3 clearly reveals that a more heterogenous grouping of countries will produce a more pronounced variance in the difference of FLP on TFR across countries. The widely varying effect of FLP on TFR is nevertheless also clear in the more heterogenous country grouping of Esping-Andersen.

Table 6: Country grouping according to Esping-Andersen; Prais-Winsten regressions with panel-corrected standard errors with AR(1) disturbances

Model	(1)	(2)	(3)
FLP_t	-0.016***	-0.016***	-0.016***
Conservative	-0.314***	-0.342*	-0.299**
Social-democratic	0.014	-0.026	-0.303
Liberal	0.291*	0.307	0.398*
FLP_t*Conservative		0.001	0.002
FLP_t*Social-democratic		0.001	0.005†
FLP_t*Liberal		-0.000	-0.003
1960-1964			0.963***
1965-1969			0.523***
1970-1974			0.321***
1975-1979			0.018
1980-1984			-0.141*
1985-1989			-0.416***
1990-1994			-0.494***
1995-2000			-0.526***
FLP_t*1960-1964			-0.011***
FLP_t*1965-1969			-0.008***
FLP_t*1970-1974			-0.005***
FLP_t*1975-1979			-0.001
FLP_t*1980-1984			0.002
FLP_t*1985-1989			0.006***
FLP_t*1990-1994			0.008***
FLP_t*1995-2000			0.008***
Constant	3.093***	3.088***	2.918***
R^2	0.712	0.715	0.824
Wald χ^2	65.994***	66.081***	233.486***

Notes: (1) $\text{TFR}_t = \alpha + \beta\text{FLP}_t + \gamma\text{region}$; (2) $\text{TFR}_t = \alpha + \beta\text{FLP}_t + \gamma\text{region} + \delta\text{FLP}_t * \text{region}$; (3) $\text{TFR}_t = \alpha + \beta\text{FLP}_t + \gamma\text{region} + \delta\text{FLP}_t * \text{region} + \epsilon\text{time} + \lambda\text{FLP}_t * \text{time}$. *** $p \leq .001$, ** $p \leq .01$, * $p \leq .05$, † $p \leq .10$.

Figure 3: Effects of female labour participation over time for different welfare regimes



In Table 7 we apply the country grouping by Gauthier. Our results in Model (1) indicate that pro-natalist and pro-traditional countries have below average fertility rates while pro-egalitarian and non-interventionist regimes have above average fertility rates. We find a significant country heterogeneity for the slope coefficient of FLP once we include time heterogeneity in addition to country heterogeneity (model 3). However, these effects are rather small and would indicate that the negative effect of FLP is reduced only for pro-natalist countries while it would be higher for the other country groups.

The results of separate estimations for the single gender regimes with piecewise trend terms are shown in Figure 4. Most interestingly, we find for pro-natalist countries a positive effect of FLP till the beginning of the 1990s. Thereafter, the effect turned negative and is decreasing over time. This result mirrors the experience among countries included in the group of western European countries that coincides with the pro-natalist regime (cf. Figure 2). In those countries we observe the highest TFR for countries which experienced the highest levels of FLP as well.[12] Like in the case of

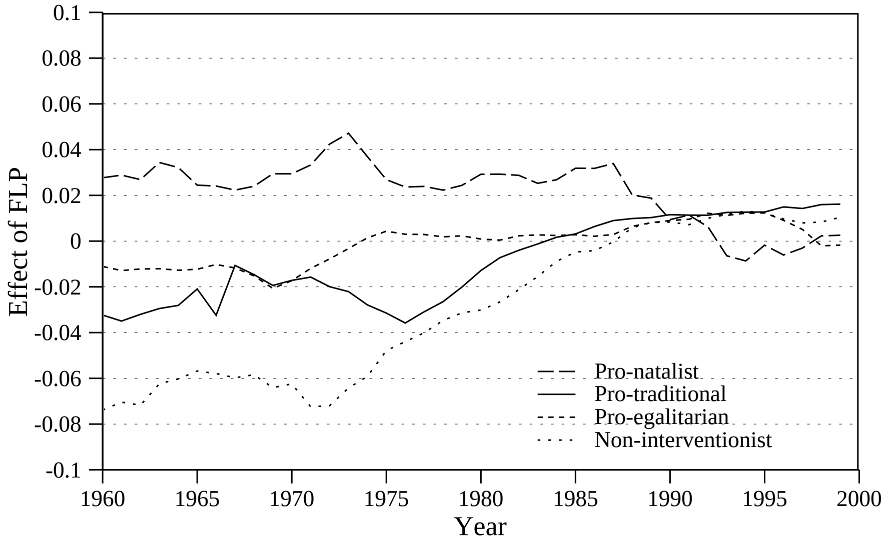
[12]Recall that by ignoring fixed country effects our results are identical to between-group estimations. The positive coefficient on FLP for pro-natalist countries therefore gives the

Table 7: Country grouping according to Gauthier; Prais-Winsten regressions with panel-corrected standard errors and AR(1) disturbances

Model	(1)	(2)	(3)
FLP_t	-0.022***	-0.022***	-0.014***
Pro-natalist	-0.239***	-0.380†	-0.981***
Pro-traditional	-0.243***	-0.230	0.086
Pro-egalitarian	0.093†	-0.188	-0.331***
Non-interventionist	0.389***	0.798***	1.227***
FLP_t*Pro-natalist		0.003	0.017***
FLP_t*Pro-traditional		-0.000	-0.005*
FLP_t*Pro-egalitarian		0.005†	0.005**
FLP_t*Non-interventionist		-0.008*	-0.018***
1960-1964			1.071***
1965-1969			0.861***
1970-1974			0.574***
1975-1979			0.151
1980-1984			-0.176†
1985-1989			-0.632***
1990-1994			-0.870***
1995-2000			-0.980***
FLP_t*1960-1964			-0.013***
FLP_t*1965-1969			-0.012***
FLP_t*1970-1974			-0.009***
FLP_t*1975-1979			-0.004*
FLP_t*1980-1984			0.002
FLP_t*1985-1989			0.009***
FLP_t*1990-1994			0.013***
FLP_t*1995-2000			0.014***
Constant	3.281***	3.254***	2.719***
R^2	0.712	0.713	0.842
Wald χ^2	90.867***	96.516***	581.883***

Notes: *** $p \leq .001$, ** $p \leq .01$, * $p \leq .05$, † $p \leq .10$.

Figure 4: Effects of female labour participation over time for different gender regimes



social-democratic welfare state regimes in Figure 3, the effect of FLP for pro-egalitarian regimes is moving up and down over time and is even positive since the mid-1980s. But also for pro-traditional regimes the effect of FLP on fertility turned from a negative to a positive value since the mid-1980s.

Summing up our findings we may argue that a regional country grouping seems to be more appropriate compared to a welfare state or a gender regime grouping. Our results also indicate that the regional country grouping seems to be valid across time while the welfare and gender regime grouping may not apply for the single countries for the whole time period 1960 to 2000 as indicated by the fact that the interaction of FLP and country groups becomes significant only if we include time heterogeneity.

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effect of a unit change in FLP on the value of TFR between two countries included in the group of pro-natalist countries.

4 Discussion

In this paper we analyse how country specificities, and changes over time, affect the relationship between female labour force participation and fertility. By moderating the conflict between work and family roles experienced by women, family-friendly country-specific institutions may reduce the negative influence of female labour force participation on fertility. Our empirical analysis is based on pooled time series data from 22 OECD countries. The basic model selection and specification follows both methodological and substantial reasons.

Averaged across all countries and years, female labour participation has a negative effect on fertility that persists with controls for between-country and between-time heterogeneity. Based on Prais-Winsten regressions with panel-corrected standard errors and autoregressive disturbances our empirical findings reveal substantial differences across countries and time periods in the effects of female labour force participation. Initial increases in female labour force participation strongly lowers fertility, but continued increases in female labour force participation have a progressively less negative effect on fertility.

Across regions, the inhibiting effects of female labour force participation are smaller in Scandinavian countries, in West European countries as well as in non-European countries than in South European countries with weak family/-work-friendly institutions. In particular, policies in Scandinavian countries aid women in combining work and family and, therefore, reduce the negative effects of female labour force participation on fertility more effectively than in other OECD countries.

Taking explicitly into account the different sociopolitical contexts of the countries by applying the country grouping suggested by Esping-Andersen, we find for conservative welfare state regimes a significantly below average fertility and for social-democratic and liberal regimes above average fertility levels. Under additional consideration of gender equality in the country grouping by Gauthier, our results indicate that pro-egalitarian and non-interventionist regimes have fertility rates above average, while pro-traditional and pro-natalist countries have below average fertility.

As we have shown in the paper, the effects of female labour force participation for the different regional country groups as well as for country groups with varying sociopolitical contexts vary widely over time. Under explicit consideration of a three-way interaction of female employment, time and country group we found a narrowing effect in the effect of FLP on fertility over time, which is even positive for north European countries. For the other European and non-European regions under consideration, the effect of FLP oscillates around zero.

Applying the welfare state grouping suggested by Esping-Andersen we find positive effects of FLP on fertility also for social-democratic countries since the mid-1980s. Taking additionally into account gender issues by gender regime grouping as suggested by Gauthier, we positive effects of female employment on fertility for pro-egalitarian and even for pro-traditional countries since the same period of time. Countries following a pro-natalist family policy seem to lose the positive effect of female employment on fertility.

However, given that sociopolitical contexts are not stable over time (as assumed in our analysis), the results might look different when accounting additionally for changing special institutional settings. Therefore, we regard our results on regional country groupings as more valid than the findings based on welfare and gender regimes. Overall, the results emphasise the importance of country heterogeneity in institutions not only for the levels of fertility but also for processes that determine levels of fertility. Varying contexts across nations and time periods affect the decisions either for working or childbearing, or for doing both.

A further contribution of our paper is the attempt to reconcile the micro and macro studies on the relationship between fertility and female labour force participation. The positive cross-country correlation coefficient used to be, and often still is, regarded to be in conflict with the micro level evidence of a negative relationship between fertility and female labour force participation. As already evidenced in earlier studies, e.g., Kögel (2004), we also found that country and time heterogeneity with respect to female labour force participation may explain the change in the cross-country correlation between TFR and FLP. In addition, we also present clear evidence that the modifying effect of regions on the female participation rate has changed over time. Hence, any analysis on the macrolevel requires to control for country heterogeneity but also for time-varying country heterogeneity.

As noted earlier, a daunting problem in our analysis comes from the crude measures of female labour force participation. Measures that distinguish between rates by age and hours worked would allow to take a closer look into the components of fertility changes. Moreover, to better understand cross-national variation in the effect of female employment on fertility it is necessary to consider a broader spectrum of confounding indicators such as those related to male and female economic status, institutional arrangements and the role of proximate determinants of fertility across countries.

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