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Does Economic Advancement 'Cause' a Re-increase in Fertility? An Empirical Analysis for OECD Countries (1960–2007)

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Abstract In the light of the recent reversal of fertility trends in several highly developed countries, we investigate the impact of economic development and its components on fertility in OECD countries from 1960 to 2007. We find that the strong negative correlation between GDP per capita does no longer hold for high levels of per capita economic output; the relation and fertility instead seems to turn into positive from a certain threshold level of economic development on. Survival of an inverse J-shaped association between GDP per capita and fertility is found when controlling for birth postponement, omitted variable bias, non-stationarity and endogeneity. However, gaps between actual and predicted fertility rates show implicitly the importance of factors influencing fertility above and over per capita income. By decomposing GDP per capita into several components, we identify female employment as co-varying factor for the fertility rebound that can be observed in several highly developed countries. Pointing out to important differences with regard to the compatibility between childbearing and female employment, our results suggest that fertility increases are likely to be small if economic development is not accompanied by institutional changes that improve parents' opportunities to combine work and family life.

Keywords Fertility · Economic development · Female employment

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1 Introduction

Over the last decades, in many OECD countries fertility rates dropped drastically while at the same time average income levels continued to increase. However, in a limited number of highly developed countries, fertility rates have been somewhat re-increasing since the early 2000s, simultaneously with continuing economic development.

The reversal of the fertility trend along the process of economic development in many, but not all highly developed countries suggests that the impact of economic development on fertility is ambiguous. Knowing whether further economic advancement is likely to sustain a 'rebound' of fertility in highly developed countries is of major political, social and economic interest. As fertility affects population growth and the age structure of the population, changes in fertility in the immediate future have far-reaching consequences on economic development, productivity growth and aspects of welfare systems.

Most recently, Myrskylä et al. (2009) found a so-called 'inverse J-shaped' relation between the human development index (HDI) and total fertility rates (TFR) for over 100 countries, suggesting an increase in fertility rates from a certain level of human development on. However, the use of a composite measurement of human development masks the particular contributions of each of the indicator's components (GDP per capita, life expectancy and school enrolment) and thus does not reveal *why* in some, especially highly developed countries, a rise in fertility comes along with increases in human development.

We want to find out more about how the reversal of fertility trends is related to economic advancement in highly developed countries. Therefore, we first empirically estimate the impact of GDP per capita on fertility, using data for 30 OECD countries that spans the years 1960–2007. In order to identify the driving factors behind the re-increase in aggregated fertility rates observed in several highly developed countries, we decompose GDP per capita in a second step and estimate the impact of labour productivity, working hours and employment on fertility while taking into account the gender composition of each of these possible determinants.

Our empirical analysis confirms a change in the relationship between GDP per capita and fertility, implying that an increase in GDP per capita 'causes' a sharp decrease in fertility rates at lower levels of per capita GDP, while this association disappears at high levels of economic output. Our results even suggest a slight positive association between fertility and GDP per capita levels, in line with the inverse J-shaped relation suggested by earlier studies. Nevertheless, this 'regime switch' happens for the few most economically advanced countries only once they achieve relatively high levels of GDP per capita. Moreover, actual fertility is often at much higher (resp. lower) levels than the value predicted by GDP per capita, which suggest that other institutional factors matter besides economic development.

We further investigate this issue by decomposing GDP per capita in labour productivity, working hours and employment components in the regression analysis. We find that fertility rates are positively correlated with increases in female employment rates, which suggests that increases in fertility are likely to be observed in those highly developed countries where economic development goes hand in hand with increasing opportunities for women to combine work with family life. This finding explains why countries with similarly high levels GDP per capita can nevertheless have very different fertility rates, as the association between GDP and fertility seems to depend on the institutional setup helping parents (and especially women) to combine childbearing and work. Thus, increases in GDP per capita may not be sufficient to lift fertility to a significant higher level if not accompanied by changes in institutions which facilitate this combination.

Our article is organised as follows. Section 2 presents an overview of the existing theoretical literature on the impact of economic development and fertility, while Sect. 3 discusses hitherto existing empirical findings. Section 4 presents our data, Sect. 5 the empirical strategy and Sect. 6 the estimation results. Section 7 concludes by summarising the main findings and identifying directions for future research.

2 The Impact of Economic Development on Fertility in Theory

The impact of economic development on fertility is found to be rather ambiguous not only on the empirical side but also in theory. An increase in income per capita can either bring an increase in the demand for children because the explicit costs are more easily borne ('income effect') or a decrease in the demand for children.

To explain the negative impact of income on fertility, the main arguments are provided by the so-called 'new home economic theory'. Becker (1960, 1981) interprets fertility reduction as a rational behaviour of households by explaining that the impact of an increase in income on fertility is subject to a *quality–quantity* trade-off. A household income increase raises not only the indirect but also the direct costs of children, because in modern societies parents place more focus on children's 'quality' to raise the chances of their children, which induces a substitution effect against the number of children in favour of the 'quality' per child (education) and the living standards of the household (Becker and Lewis 1973; Willis 1973). Jones et al. (2008) suggest furthermore that advances in 'education technology' and the accompanied rise in education costs are other important factors explaining the decrease in fertility that comes along with economic development. Becker et al. (1990) examine formally a negative effect of increasing returns of investments in individual human capital on fertility. Under these assumptions, families find it optimal to have fewer children, and to provide each child with a higher level of human capital. This high level of human capital leads, at the aggregate level, to economic growth, which explains why economic development goes hand in hand fertility decline. This latter might also be heightened by limits in intergenerational altruism which imply that an individual's future utility of consumption and savings is reduced by the number of descendants. In such circumstances, technological progress is likely to induce both a higher growth rate of consumption and a lower rate of fertility (Barro and Becker 1989; Doepke 2004).

Another argument in favour of a negative impact of economic development on fertility stresses an increase in the 'opportunity costs' of having children which is essentially borned by women. These costs are derived from the increase in women's educational achievement and thus the increase in women's earnings potentials which come along with economic development from a certain level of development on (Boserup 1970; Goldin 1994; Cagatay and Özler 1995; Galor and Weil 1996; Mammon and Paxson 2000; Luci 2009). Increasing earning potentials incite women to participate in the labour market. In the absence of possibilities to combine work with family life, women are likely to 'substitute' work for children, as staying at home to care for children implies an implicit wage loss for women. This wage loss represents indirect costs of having children. As economic development is likely to increase these indirect costs, economic development is expected to decrease fertility due to a 'substitution effect' between fertility and female employment (Becker 1965; Willis 1973; Hotz et al. 1997). Women postpone childbirth until a period of life when raising children is less damaging to the career opportunities of women and/or reduce their completed number of children (Blossfeld 1995).

However, this decrease in fertility happens only if women (or parents in general) have to choose between work and family life. If parents are given the possibilities to combine both, that is for example by substituting parental care with purchased services, the negative effect of economic development and of female employment on fertility might be weakened (Day 2004).

Martinez and Iza (2004) argue that this negative effect weakens all the more in the case of decreasing *relative* costs of child care services. This decrease in relative costs, which has been initiated by skill biased technological change, has been observed in several highly developed countries over the last two decades. Increasing relative wages for female skilled labour make child care costs more bearable and might therefore lead to a turn from negative to positive in the association between economic development and female employment on the one side and fertility on the other side. The professionalization of the sector of child-minders in some advanced countries is likely to increase the cost of childcare services and therefore alters the relative benefit that households may get from externalising childcare. However, salaries of childcare workers are still significantly lower than the average wage and are combined with childcare subsidies. This makes the use of formal childcare a profitable option for a large fraction of working parents (OECD 2007).

In this context, the connection between economic development, technological and societal change and female employment is crucial for explaining fertility trends. The changing context of institutions and norms regarding childbearing, gender relations and the division of work is also one dimension to be considered in order to understand the impact of economic development on fertility (Lesthaeghe and Surkyn 1988; Jones et al. 2008; Philipov et al. 2009). Thus, the increasing use of contraceptives and changes in the norms concerning childbearing age are parameters that, one the one hand, enable households choose more freely in terms of timing and number of births. On the other hand, changing attitudes toward female employment and the care of young children also facilitate the adaptation of childbearing behaviours (Lesthaeghe 2010; Goldstein et al. 2009). These changes have also been accompanied in many economically advanced countries with the development of policies supporting families with children and working parents who get now more opportunities to combine work and family life than few decades ago (OECD 2007, 2011; Thévenon 2011).

To sum up, the income effect produced by an increase in income per capita is expected to gain in relative importance for fertility after a certain stage of development is attained, when institutions are developed that alleviate direct and indirect costs of having children. These institutions that come along with economic development reflect not only new political or economic but also social dimensions such as modern attitudes and norms toward the family and gender roles, which allow women to combine work, childbearing and child-raising (Philipov et al. 2009).

In addition, parts of the re-increase in fertility may also be explained by an end of the process of birth postponement (Goldstein et al. 2009). Birth postponement does not always reduce parents 'demand' for the total number of children. Consequently, increased education and later transition to employment for women leads to a postponement of childbirth (tempo effect), but does not necessarily affect the total number of children a woman has (quantum effect) (Rindfuss et al. 1980; Lesthaeghe 2001; Bongaarts 2002). Period measures like TFR thus tend to decrease with birth postponement and re-increase once this process has come to an end (Sobotka 2004). Estimations of tempo (and/or parity)-adjusted fertility rates help at limiting the variations of fertility rates due to the changes in the timing of births. They show decrease and recent upswing of fertility rates which are weaker than the trends given by the traditional measure of period fertility (Bongaarts and Sobotka 2012).

3 Previous Empirical Findings on the Impact of Economic Development on Fertility

The existence of divergent relations between economic growth and fertility rates are also assessed empirically. Butz and Ward (1979) observe that fertility rates in the US were pro-cyclical until the 1960s, but started to decline in a period of persistent economic growth from the 1960s until the late 1970s. The study by Butz and Ward (1979) has been challenged, however, for several reasons. While some studies such as Mocan (1990) still provide figures of persistent counter-cyclical fertility patterns, other studies raise objections to the empirical strategy pursued by Butz and Ward (1979) and propose different estimates that do not confirm the negative impact of real wages and income on fertility rates at higher levels of income (McDonald 1983; Krämer and Neusser 1984; Macunovich 1995). Moreover, Butz and Ward's (1979) prediction of continuous fertility decline with further economic advancement only applies to a limited number of countries. In many OECD countries, the negative correlation between fertility and economic advancement has weakened within the last decade and in some highly developed countries, a reversal of fertility trends and a rebound of fertility rates back to replacement levels can be observed simultaneously with continuous economic growth.

Most recently, Myrskylä et al. (2009) argue that a fundamental change occurred during the last quarter of the last century in the relation between fertility and human development. On the basis of both cross-sectional and longitudinal data covering more than 100 countries for the years 1975–2005, Myrskylä et al. (2009) estimate the impact of human development (measured by the United Nations HDI) on TFR. They use a graphical analysis to identify the potential level of HDI that turns the correlation between human development and fertility from negative to positive

(HDI = 0.85–0.9). This critical level is then tested by including it as a parameter in a maximum likelihood function. For the year 1975 Myrskylä et al. (2009) find a strictly negative correlation between HDI and fertility for all countries. Yet, for the year 2005, they find a negative correlation between HDI and TFR only for countries with a HDI level below that minimum. For countries with a HDI level above that minimum, Myrskylä et al. (2009) find that the two variables are positively correlated. This suggests that in highly developed countries like the USA, Norway and Ireland, human development implies a rebound of fertility, whereas at low and medium development levels, human development continues to decrease fertility.

Furuoka (2009) provides a further empirical test of the critical level of HDI that leads to a turn in the correlation. The test for the threshold effect of HDI on fertility constructs asymptotic confidence intervals for the threshold parameter. Like Myrskylä et al. (2009), Furuoka (2009) splits the sample in two regimes in order to test linear correlations. Furuoka (2009) contests the study by Myrskylä et al. (2009) by finding that in countries with a high HDI, higher levels of HDI still tend, albeit weakly, to be associated with lower fertility rates. Moreover, Harttgen and Vollmer (2012) revisit this topic with revised data and find that the reversal in the HDI-TFR relationship is neither robust to UNDP's recent revision in the HDI calculation method nor to the decomposition of HDI into education, standard of living and health sub-indices.

Besides ambiguous findings, both Myrskylä et al. (2009) and Furuoka (2009) assume a linear relation between economic advancement and fertility trends, but allow for a change in slope or even in the sign of the association. However, both studies use a composite measure of human development, containing GDP per capita, life expectancy and school enrolment. The combination of the three components makes it difficult to interpret the estimated coefficients for two reasons. Firstly, due to limited HDI-data availability, in both studies the analysis of the fertility rebound is focused on cross-country variations only. Secondly, it is unclear which of the HDI components initiates the fertility rebound. In addition, as life expectancy and school enrolment are correlated with GDP per capita, interpretation problems arise because of multi-collinearity. Consequently, it is unclear what elements behind human development drive the fertility rebound in highly developed countries.

The most recent empirical studies jointly suggest that the impact of development on fertility turns from negative to positive from a certain development stage on in highly developed countries. What is still unclear is at which level of development one can expect the correlation to turn and which components of development do exactly drive the fertility rebound.

We therefore empirically investigate to what extent fertility variations are connected with trends in GDP per capita and which specific components of GDP are most likely to cause an increase in fertility levels, such as the slight upswing of fertility rates observed recently in the most economically advanced countries.

4 Data Discussion

In order to identify the driving factors of the fertility rebound, we consider it appropriate to focus our analysis on OECD countries only, as the rebound is mainly observable in highly developed countries. A closer look at the separate HDI components for OECD countries shows that for this limited group of countries, the variation is highest for GDP per capita in comparison with life expectancy and school enrolment. We therefore find it appropriate to focus our measure of development on GDP per capita for this particular group of highly developed countries that differ quite weakly in terms of life expectancy and school enrolment. In addition, using GDP per capita as determinant of fertility instead of HDI allows focussing on within-country variations, as observations of GDP per capita are available on a yearly basis.

When estimating the impact GDP per capita on fertility, we use a large macroeconomic panel data set from OECD databases that includes observations from 30 OECD countries[1] over four decades (1960–2007). We use TFR as standard measure for fertility, but also try to distinguish in how far economic development influences the tempo and the quantum effect of fertility. We are aware of the fact that TFR as a period measure only gives an accurate estimation of completed fertility levels if there is no change in the timing of births across cohorts. In the opposite case, such as when there is an increase in the mean age of mothers at childbirth, the number of births in a given period is reduced. Consequently, the postponement of birth to older ages reduces TFR and the end of postponement increases TFR (Bongaarts and Feeney 1998; Bongaarts 2001, 2002; Kohler et al. 2002; Goldstein et al. 2009). As *TFR* is sensitive to changes in the timing of childbirth, we also use tempo-adjusted fertility rates (*adjTFR*), which come from the Human Fertility Database and cover the years 1961–2005, but are only available for a subset of 18 OECD countries.[2] Taking tempo changes into account, tempo-adjusted fertility rates are usually higher than TFR. Tempo-adjusted fertility rates are available as 3-year moving averages. By weighting *TFR* by changes in women's mean age at childbirth, the *tempo-adjusted TFR* focuses on the quantum-component of fertility changes (Bongaarts and Feeney 1998; Sobotka 2004). However, *adjTFR* only corresponds to a pure quantum measure of fertility on the assumption of uniform postponement of all stages, i.e. an absence of cohort effects (Kohler and Philipov 2001). Consequently, *adjTFR* implies only an imperfect control for tempo effects. To intensify the control for birth postponement, we integrate two different measures of women's age at childbirth as control variables.

Finally, we analyse the role of different GDP components (labour productivity, working hours, employment and its gender composition) for re-increases in fertility. This allows answering the question why in certain countries, GDP increases go hand in hand with increases in fertility. For this purpose, we focus on the time period 1995–2007, therefore ending our time window before the on-going economic recession, for which the effects on fertility are not our concern here.

Table 4 in the appendix provides an overview of all data used in this study.

[1] Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Japan, South Korea, Luxembourg, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Spain, Sweden, Switzerland, Turkey, United Kingdom, USA.

[2] OECD countries without: Australia, Belgium, Canada, France, Germany, Greece, Korea, Luxembourg, Mexico, New Zealand, Switzerland, Turkey.

4.1 Trends in Total Fertility Rates in OECD Countries

The dominant feature regarding fertility trends is the sharp decline in TFR in OECD countries over the last four decades. Looking back to the early 1970s, the fall appears substantial with an average *TFR* that fell from 3.23 children per woman in 1970 to 1.71 in 2008, e.g. a level well below the 2.1 threshold required to replace the population with no contribution from immigration (Fig. 1 panel 1). In 2008, only a few countries had a fertility rate around or above the so-called replacement rate level (United States, Ireland, New Zealand, Iceland, and Mexico and Turkey).

As a result of fertility decline, 'lowest-low' fertility countries (with *TFR* below or around 1.3 on average since 2000) include Austria, Czech Republic, Germany, Greece, Hungary, Korea, Japan, Poland, Portugal, Slovak Republic, Spain and Switzerland.

Despite this overall decline in fertility, many countries have recently experienced a reversal of trends, with an increase in fertility rates (Fig. 1 panel 2). The 'rebound' has been especially high (above 0.3 children per women, comparing TFR in 2008 with the minimum since 1970) in Denmark, Sweden, Czech Republic, United States, Finland, France, United Kingdom, Belgium, Netherlands, Spain, Norway and New Zealand. The timing and pace of this change varies from country to country. Only a few countries experienced such a reversal in trends in the mid-1990s (Belgium, France, Ireland, Italy, Netherlands, Spain and the US), while a significant increase (by above 0.2 children per woman) has occurred since 2000 in Sweden, Czech Republic, United Kingdom, Greece, Spain, New Zealand and Ireland). Nevertheless, most OECD countries have seen such an increase since 2000, though often very slight, the only exceptions being Germany, Korea, Mexico, Portugal, Switzerland, and Turkey. Fertility rates continue to decline in this latter set of countries, but the pace of decrease slowed down.

4.2 Trends in GDP per Capita in OECD Countries

GDP per capita is measured at purchasing power parity (PPP) in constant 2005 US $. On average in all 30 OECD countries, GDP per capita at PPP increased from $11,915 in 1970 to $28,134 in 2007. Constant-price measures of GDP are considered here in order to filter out the increase in GDP per capita that is due to price inflation without relating to any increase in the consumption basket.

In all countries, the increase is more or less continuous with common breaks around 1975, 1980, 1990 and 2000 due to economic shocks that affected all countries at about the same time. Countries with high GDP levels are Luxembourg, and somewhat closer to the average level Norway, the United States and Sweden, with highest levels in the decade after 2000. The lowest levels of GDP per capita can be observed in Korea, Turkey and Mexico in the 1970s, followed at some distance by Poland in the 1990s and Portugal in the 1970s.

The descriptive analysis suggests that whereas until the late 1980s in all observed countries economic advancement went hand in hand with fertility decline, since the early 1990s the picture is threefold: generally speaking, countries with the lowest

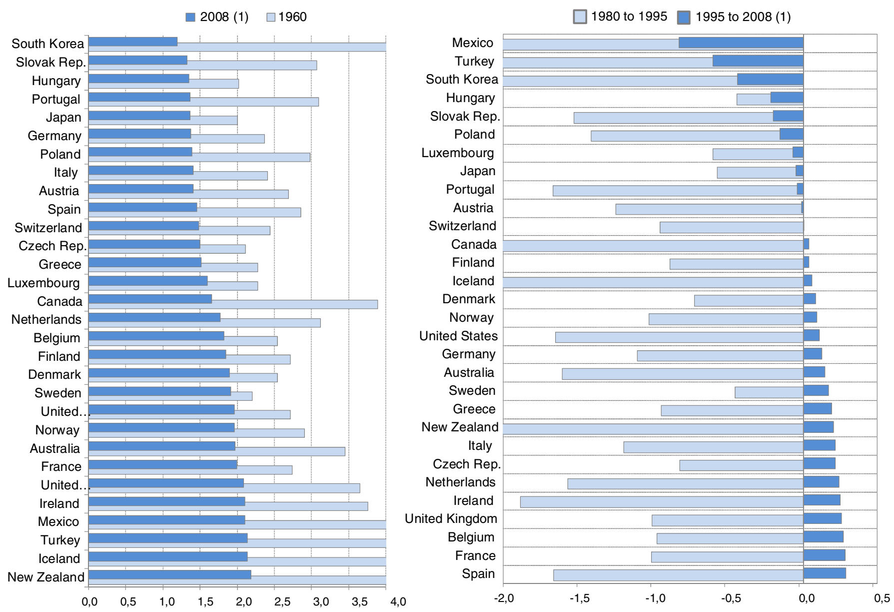

Fig. 1 Fertility trends in OECD countries *Panel 1* TFR in 1960 and 2008 *Panel 2* relative change 1980–1995, 1995–2008 source OECD family database (2010) (1) Year 2007 for Canada, Czech Republic, Estonia and Slovenia

income levels record continuously declining fertility rates. Countries with medium income levels record stagnant fertility levels below replacement levels and countries with the highest income levels record a fertility rebound. This observation supports the hypothesis of a reversal of fertility trends along the process of economic development in OECD countries and suggests a convex impact of economic advancement on fertility.

Figure 2 plots the observations of GDP per capita against those of total fertility and shows an inverse J-shaped pattern between the two variables. This suggests that at low-income levels, economic growth lowers fertility whereas form a certain higher level of income on, income growth increases fertility. In this data plot, countries that risk over-accentuating the inverse J-shaped pattern are dropped. This concerns Luxembourg, which has an outstandingly high level of GDP per capita among OECD countries, especially in the 2000s. This also concerns Korea, Mexico and Turkey, as these emerging countries have outstandingly low levels of GDP per capita and high levels of fertility, especially in the 1960s and 1970s.

5 Empirical Strategy

Our empirical analysis looks at the extent to which fertility levels are a predictable function of per capita GDP. Hereby, we especially test whether a change in the sign

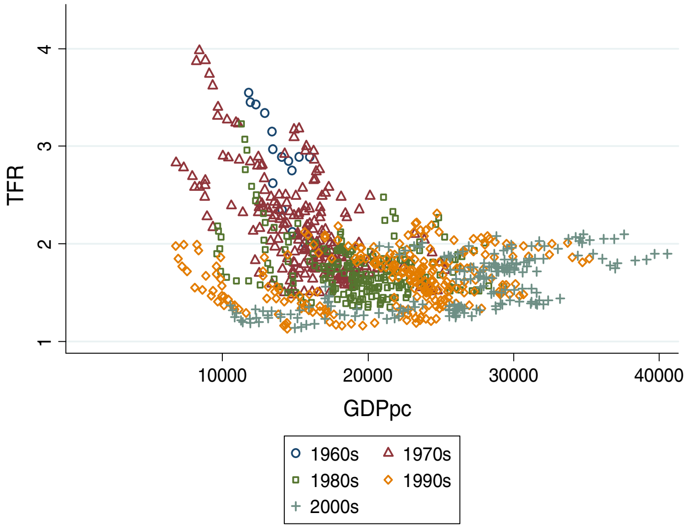

Fig. 2 GDP per capita against TFR for 26 OECD countries, 1960–2007 data source: OECD Family Data Base (2010)

of the association between economic advancement—as measured by levels of GDP per capita—and fertility trends can be identified and be part of the explanation of the slight fertility rebound observed in some of the most highly developed countries. In addition, we analyse the role of different GDP components for re-increases in fertility.

In this perspective, we first test the assumption that, from some level of per capita income upward, TFRs may switch from a decreasing to an increasing function of per capita GDP.[3] Testing this assumption requires to disentangle as much as possible the influence of GDP from those of the other determinants of fertility and GDP that may be linked over time. To do so, the estimation takes into account possible unobserved determinants of fertility by including country-specific time trends in the model. There are limits, however, in the extent to which time trends can accurately account for unobservable changes, for example if GDP per capita follows the same trend, possibly confounding the relation between GDP and fertility. This can occur if countries are exposed to the same processes, albeit not necessarily with the same strength, influencing both GDP per capita and fertility. One may think, for example, of the important development of policies reconciling work and family in OECD countries (Thévenon 2011) or of the evolution of cultural norms concerning childbearing (Lesthaeghe 2010). Such circumstances would

[3] Other assumptions regarding this relationship are of course possible. For instance, one can anticipate the relationship between GDP per capita and fertility to change from a certain point in time onward, which may correspond to a certain situation of institutional background. Hence, we aim at controlling as much as possible for time-varying unobserved characteristics. However, this control is not perfect due to the correlation between trends of these characteristics and those of GDP per capita (thanks to anonymous reviewers for having drawn our attention on this issue).

create 'cross-section dependence' between countries. This dependence cannot be properly wiped out by controlling for time trends, and thus there remains a risk of obtaining biased estimation coefficients measuring the effect of GDP per capita on fertility. Testing of the cross-section independence of the residuals obtained by the model is then a mean to check whether results are affected by this issue (Pesaran 2004). In this latter case, a strategy to account for these common unobserved factors is the Common Correlated Estimator proposed by Pesaran (2006), which includes cross-section averages of the dependent and independent variations in the regression equation.[4] However, conditions for a satisfactory application are not fully met here, which will lead us to interpret results with caution.[5] For this reason, these results will not be shown in the tables included in the core text, but are included in the Annexes.

Against this Backdrop, Our Empirical Procedure Consists of Four Steps (I–IV).

I. Alternative specifications are tested to measure the effect of an increase in GDP per capita on fertility, including models with linear as well as nonlinear specifications. This allows testing for a change in the magnitude or sign of the relation between per capita GDP and fertility levels. By applying pooled OLS with robust standard errors, a linear specification is first estimated, with TFR as endogenous variable and the log of GDP per capita (*lnGDPpc*) as exogenous variable. Then, we test an 'exponential' specification where the log of total fertility rates (*lnTFR*) is modelled as a function of GDP per capita (*GDPpc*). A third specification with TFR expressed as a quadratic function of the log of GDP per capita is estimated as follows:[6]

$$\mathrm{TFR}_{i,t} = \alpha + \beta_1 * \ln GDPpc_{i,t} + \beta_2 * \ln \left(GDPpc_{i,t}\right)^2 + \alpha_i T + \varepsilon_{i,t}. \qquad (1)$$

This model allows for a change in the sign of the effect of an increase in GDP per capita on fertility levels, which is compatible with a reversal of fertility trends. A positive estimated coefficient β_2 would suggest that the correlation between TFR and GDP per capita is first negative up to certain threshold level of GDP per capita and then turns into positive for higher levels of GDP per capita. As it will be shown, the quadratic model turns out to be better suited than the linear and the exponential specification to represent the existing nonlinear relation between fertility and economic development.

The robustness of the quadratic model is then checked by applying a fixed effects estimation with robust standard errors, which allows capturing unobserved time-

[4] Country-specific influences of these averages are then estimated to approximate the incidence of unobserved factors which may vary across countries. This allows for more flexibility as the impact of the unobserved common factors can differ across country while the evolution of these factors may be nonlinear or even non-stationary.

[5] In particular, to be unbiased, the CCE estimator requires that the number of unobserved factors is not larger than $K + 1$ (K being the number of independent variables and equal to one here); or it requires that the factors loadings of independent and dependent variables (i.e. TFR and *GDPpc*) to be uncorrelated, which is not likely to be the case (Sarafidis and Wansbeek 2012).

[6] For the linear and the quadratic model, we use the natural logarithm of GDP per capita (*lnGDPpc*) which is standard in most macro-econometric works, as the logarithmic form reduces absolute increases in the levels of GDP per capita and therefore captures proportional rather than absolute differences in the distribution of GDP per capita levels.

constant variables which may affect fertility (i.e. country-specific characteristics linked to historical geography, population build-up or certain norms/attitudes, etc.). Controlling for these country-specific factors also moves the focus on within-country variations, so as to assess the impact of GDP per capita increases on fertility over time. As mentioned before, a key question is also to separate the influence of GDP on fertility levels from those of the period which comes along with an increase in per capita GDP. This period effects might, for instance, capture changes in attitudes, life-style or institutions that may also affect how GDP per capita influence fertility rates. The inclusion of 10 year-period effects in the quadratic fixed effects model helps us disentangle the pure effect of GDP from those of other time-varying unobserved factors. Period effects ($\alpha_i T$) are, in addition, assumed to be country-specific as the influence of time can vary across countries. As shown in the next section on results, this approach, however, does not properly account for the correlation between time trends and GDP per capita, which ends in cross-section dependence of error terms. A common correlated estimator is then applied to sort out this issue, but there are limits in the extent to which this procedure (or other procedure to deal with unobserved common factors) can be applied here.

Subsequently, to further test the existence of a nonlinear relation and—more precisely—to identify a structural break in the correlation between the two variables, we express TFR as a piecewise linear function of *lnGDPpc*. By applying fixed effects with robust standard errors, we estimate the impact of GDP per capita on fertility rates as follows:

$$\mathrm{TFR}_{i,t} = \alpha + \alpha_i + \beta_{\mathrm{pre}} * \ln GDPpc_{i,t}^{\mathrm{pre}} + \beta_{\mathrm{post}} * \ln GDPpc_{i,t}^{\mathrm{post}} + \varepsilon_{i,t} \qquad (2)$$

(α_i = country fixed effects), where the coefficients β_{pre} and β_{post} measure the effects of economic development on the total fertility rate at GDP per capita levels below and at or above a critical value $\mathrm{GDPpc}^{\mathrm{crit}}$. The hypothesis of a reversal of the relation implies that $\beta_{\mathrm{pre}} < 0$ and $\beta_{\mathrm{post}} > 0$.

For testing this hypothesis, we estimate $\mathrm{GDPpc}^{\mathrm{crit}}$ via an iterative search process, using all countries that attained a log value of GDPpc above 10 (which is the lower bound for the critical point suggested by figure A, as well as by the quadratic OLS estimation). The statistical estimate of $\mathrm{GDPpc}^{\mathrm{crit}}$ is obtained using maximum likelihood, by including $\mathrm{GDPpc}^{\mathrm{crit}}$ as a parameter in the likelihood function of Eq. (2).[7]

II. In a second step, we apply several robustness checks for the quadratic model in order to capture possible biases caused by unobserved variables, endogeneity and non-stationarity (2 stage least squares, random effects, between effects (BE), first difference estimator, system GMM).

Endogeneity is controlled for by applying 2SLS. In the absence of more accurate instruments to handle with possible endogeneity, lagged values of GDP per capita

[7] The log-likelihood function is actually maximised by using a two-stage grid-search algorithm that in the first stage varies the value of $\mathrm{lnGDPpc}^{\mathrm{crit}}$ from 9.5, 9.6, 10.0, 10.8, …, and in a second stage refine the search with a step size of 0.01 in the neighbourhood of the best-fitting first stage of $\mathrm{lnGDPpc}^{\mathrm{crit}}$. The likelihood profile is available on request.

serve as instrumental variables as a second best option.[8] Granger Causality (Granger 1969) has also been tested (see Table 5 in the appendix) and the test suggest that *lnGDPpc* 'Granger causes' *TFR*, whereas *TFR* does not 'Granger cause' *lnGDPp*c: i.e. lagged values of the *lnGDPpc* provide statistically significant information about actual values of *TFR* but at the same time, lagged values of *TFR* do not provide statistically significant information about actual values of *lnGDPpc*.

We compare the FE model to a BE model, which is based on time averages of each variable for each country. We also apply a random effects model (RE) which also controls for unobserved country heterogeneity but captures both within and between-country variation. The FE model turns out to be superior to the BE- and the RE-model.

For most countries, neither *TFR* nor GDP per capita follow clear time trends for the observed period. Graphical tests (correlogram, partial correlogram), an augmented Dickey Fuller (1979) and a Phillips and Perron (1988) test for unit root in time series and a Levin et al. (2002) test for unit root in panel data suggest the existence of an autocorrelation in some, but not all of the time series of TFR and lnGDPpc (results available on request). As the tests suggest that all series are difference stationary, we apply a First Difference Estimator. The differencing process implicitly controls for fixed effects and removes the unit root from the residual autocorrelation that can come from non-stationarity of data series.

We finally use a one-step system generalised method of moments estimator, which not only considers unobserved heterogeneity and non-stationarity, but at the same time also endogeneity. In addition, system GMM allows controlling for the dynamics of adjustment (see Box in the appendix for more details about System GMM estimation).

III. We then test the robustness of the quadratic model and the piecewise regression by controlling for birth postponement. For this purpose, first add two measures of women's age at childbirth as control variables as exogenous variables and then use tempo-adjusted fertility rates as endogenous variable, again by applying fixed effects with robust standard errors.

IV. Finally, we aim at getting a deeper insight in the economic mechanisms behind fertility increase. For this purpose, we decompose *GDP per capita* into its three standard components, which are labour productivity, average working hours per worker and the employment ratio.[9] Because we are now particularly interested in the specific determinants of the fertility *rebound*, and also because of limitation in data availability,

[8] More precisely, we use lagged variables of *lnGDPpc* as instruments for *lnGDPpc* and lagged variables of $lnGDPpc^2$ as instruments for $lnGDPpc^2$ We perform the IV-regression in two steps (two-stage least squares estimator) by using 1-year lags as well as 5-year lags. The use of lagged exogenous variables lessens the risk of obtaining biased and inconsistent estimators due to inverse causality between the endogenous and the exogenous variables. It is for example likely that variations of fertility that lead back to changes in the economic environment appear time-lagged. At the same time, it is less likely that *TFR* observed in 1984 impacts *GDP per capita* levels of the year 1980. However, the use of lagged exogenous variables does not completely rule out the problem of inverse causality between fertility and *GDP per capita*. *TFR* of 1984 may affect *GDP per capita* of 1980 due to birth postponement, for example. Women who delay childbirth from 1980 to 1984 in favour of labour market participation actually do influence GDP levels measured in 1980 by the fact that they have a child in 1984 only.

[9] Labour productivity = GDP/sum of working hours; avrg. working hrs. per worker = sum of working hours/active population; employment ratio = active population/total population.

we focus on linear impacts of the decomposition variables on fertility and consider the time period 1995–2007 by applying the FE model. The decomposition we propose is done with a sequence of different steps. The first step is to estimate the impact of our three decomposition variables on *TFR* (with country fixed effects):

$$\begin{aligned} \mathrm{TFR}_{i,t} = {} & \alpha + \alpha_i + \beta_1 * \ln(\text{labour productivity})_{i,t} \\ & + \quad \beta_2 * \ln(\text{avrg.hrs.per worker}) + \beta_3 * \ln(\text{employment ratio}) + \varepsilon_{i,t} \end{aligned} \quad . \qquad (3)$$

The second step is to split the employment ratio into two variables, which are the employment rate (ages 25–54) and the ratio of the active population.[10] We limit the observed age group in order to better capture the impact of the employment variables on fertility. We estimate the impact of our four decomposition variables on TFR as follows:

$$\begin{aligned} \mathrm{TFR}_{i,t} = {} & \alpha + \alpha_i + \beta_3 * \ln(\text{labour productivity})_{i,t} + \beta_4 * \ln(\text{avrg.hrs.per worker}) \\ & + \beta_5 * \ln(\text{employment rate}) + \beta_6 * \ln(\text{ratio active population}) + \varepsilon_{i,t} \end{aligned} \quad . \qquad (4)$$

The third step is to use our decomposition variables disaggregated by gender and estimate our model as follows:

$$\begin{aligned} \mathrm{TFR}_{i,t} = {} & \alpha + \alpha_i + \beta_1 * \ln(\text{labour productivity})_{i,t} \\ & + \beta_2 * \ln(\text{avrg.hoursper worker\_men}) + \beta_3 * \ln(\text{avrg.hours.per worker}_{\text{women}}) \\ & + \beta_4 * \ln(\text{employment rate\_men}) + \beta_5 * \ln(\text{employment rate\_women}) \\ & + \beta_6 * \ln(\text{ratio active population\_men}) \\ & + \beta_7 * \ln(\text{ratio active population\_women}) + \varepsilon_{i,t} \end{aligned} \qquad (5)$$

6 Estimation Results

6.1 The Impact of GDP per Capita on Fertility

I. We start with testing alternative models capturing the relation between GDP per capita and fertility, i.e. we test a linear against several nonlinear specifications. Regression results are shown in Table 1.

In comparison to the linear and the exponential specification (column 1 et 2), the goodness of fit *(R^2)* is highest for the quadratic model (column 3). The quadratic model allows a change in the sign of the association between GDP per capita and TFR.[11] The fact that the estimated coefficient for $lnGDPpc^2$ is significantly positive

[10] Ratio active population = active population (ages 25–54)/total population (ages 25–54).

[11] However, the goodness of fit of the exponential model cannot be directly compared to the goodness of fit of the linear and quadratic model, as the form of the endogenous variable of the exponential model (*lnTFR*) differs from the other two models (*TFR*). We therefore also test the linear and the quadratic model using *lnTFR* as endogenous variable while keeping GDP in its logarithmic form. For this specification (not presented here), R^2 of the linear model is 0.33, while R^2 of the quadratic model is 0.41. The goodness of fit of the quadratic model, using *lnTFR* as endogenous variable and GDP per capita and its squared form as exogenous variable is 0.34. The form of the quadratic model presented in column 3 of Table 1 is thus confirmed in having the highest goodness of fit.

Table 1 Specification tests

Type of regression	(1) Pooled OLS (robust SE)	(2) Pooled OLS (robust SE)	(3) Pooled OLS (robust SE)	(4) Fixed effects[1] (robust SE)	(5) Fixed effects (robust SE)
Endogenous variable	TFR	lnTFR	TFR	TFR	TFR
Specification	Linear model	Exponential model	Quadratic model	Quadratic model	Piecewise regression
Regressors					
GDPpc		0.0000166*** (−11.30)			
lnGDPpc	−1.013*** (−14.38)		−15.63*** (−9.45)	−16.95*** (−12.80)	
lnGDPpc^2			0.760*** (9.09)	0.818***£ (12.19)	
lnGDPpc (pre 10)					−2.02*** (−24.30)
lnGDPpc (post 10)					0.13** (3.08)
Constant	11.87*** (16.89)	0.943*** (28.57)	81.92*** (10.02)	89.58*** (13.69)	19.9*** (−0.73)
N	1050	1050	1050	1050	1050
nb. of countries	30	30	30	30	30
Time period	1960−2007	1960−2007	1960−2007	1960−2007	1960−2007
R^2	0.359	0.200	0.460 (overall)	0.542 (within)	0.544 (within)
R^2 adj.	0.359	0.200	0.459	0.542	0.544
Estim. minimum lnGDPpc			10.28	10.39	10
Estim. minimum GDPpc US$ (PPP)			29.228	31.746	22.000
Estim. minimum TFR			1.56	1.69	
Test of cross-section independence of residuals abs. (*p* value)[2]	0.45 (0.00)	0.52 (0.00)	0.47 (0.00)	0.45 (0.00)	0.41 (0.00)

t statistics in parentheses, * $p < 0.05$; ** $p < 0.01$; *** $p < 0.001$

(1) The model includes country-specific time dummies for each period of 10 years

(2) Pesaran (2004) CD test, the null hypothesis assuming that all residuals are cross-section independent. Absolute correlation and *p* value of the test are reported; a *p* value below 0.05 leads to the rejection of cross-section independence

reveals the existence of a minimum point in the relation between *TFR* and *lnGDPpc*. This suggests that the correlation between TFR and GDP per capita is first negative up to certain threshold level of GDP per capita and then turns into positive for higher levels of GDP per capita.

The quadratic specification is then tested by a Fixed Effects estimation (column 4), which confirms a convex impact of GDP per capita on TFR. The estimation controls for the influence of time with country-specific dummies for each period of 10 years. Hence we estimate how fertility rates evolve with GDP per capita within countries, irrespectively of the effect of time.[12] The fact that the FE regression results are significant indicates that the hypothesis of a convex impact of *lnGDPpc* on *TFR* is confirmed when focusing only on within-country variation over time (and not caused by cross-country distortions). The higher goodness of fit as well as the higher significance of the fixed effects estimation compared to a BE estimation (see BE-results in Table 6, column 2, in the appendix) suggests that the convex impact of economic development on fertility is actually dominated by within-country variation. Yet, potential bias in these estimates cannot be ignored since a Pesaran (2004) test applied to regression residuals suggests that the assumption of cross-section dependence cannot be rejected. For this reason, the model is re-estimated with cross-section averages of dependent and independent variables that are expected to wipe out the incidental common correlated factors (Table 6, column 5). The influence of GDP per capital on fertility rates is still convex but yields prediction of a turning point at a much lower level of per capita income.

Thus, for the quadratic model, the FE estimation results (Table 1, column 4) indicate that the critical GDP per capita level is located at US$ 31 746 (PPP) and a fertility level of 1.69 children per woman in the absence of country-specific characteristics (column 4). Slightly lower estimates of these critical values are given by the pooled OLS estimation, with a minimum TFR estimated at 1.56 and a corresponding GDP per capita at US$ 29 230 (PPP) for the quadratic model (column 3). This suggests that economic development decreases fertility until a relatively high income level, but then, economic growth is associated with a re-increase in fertility rates. Nevertheless, fertility rates are also found to depend on unobserved time-varying factors which are approximated here by country-specific time trends, but the rejection of cross-section independence of the residuals suggests that they do not completely account for unobserved variables that are correlated with GDP per capita.[13]

[12] Fertility and economic outcomes may also follow a non-stationary evolution path, in which case there is some risk of getting spurious regression results. To deal with this issue and to test the robustness of our results, we also carry out First Difference and System GMM estimations, which take into account non-stationarity. The results are presented in Table 6 in the appendix.

[13] By contrast, the Common Correlated estimator shown in the Annex (Table 6) seems to provide a better control for these unobserved variables with a slightly lower correlation of residuals for which the assumption cross-section independence is no longer rejected (column 5). Here the model also foresees a switch to a positive impact of GDPpc on fertility rates, which is however predicted to happen at much lower level of income per capita (US$ 8929) once correlated factors are controlled for. In practice, for all countries which experienced an upturn in fertility rates, this upturn happened at much higher levels of GDP per head which makes this estimate very implausible (see Figs. 3, 4).

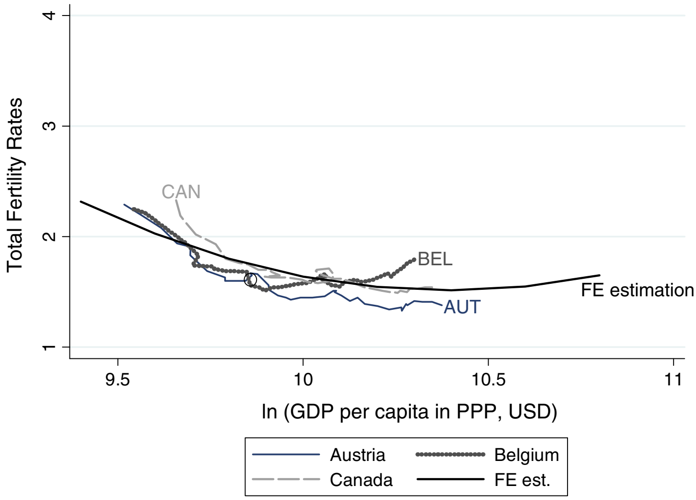

Fig. 3 Fixed effect estimation (based on 30 countries, 1960–2007) against observed within-country variation in Austria, Canada and Belgium (1960–2007). Data source: OECD Family Database (2010) and authors' estimates

Then, results of the piecewise linear regression (with fixed effects) also confirm the change from a negative association between economic advancement and aggregated fertility levels to a positive relations from a certain level of GDP per capita onward (column 5).[14] The estimated GDP-breakpoint is lower here (US$ 22 000 PPP) than for the quadratic function, since the log-likelihood function is actually maximised for an estimated *lnGDPpc* at 10.[15]

We illustrate the FE results of the quadratic model (column 4), as the quadratic model gives information about the estimated breakpoint levels of GDP per capita and TFR. Figures 3 and 4 compare our estimated pattern between GDP per capita and TFR with true within-country variations of selected OECD countries.

The curved line presents the FE results graphically. The line confirms a flattening relation between economic development and fertility and shows that the estimated pattern between *TFR* and *lnGDPpc* is presumably '*inverse J-shaped*'. The declining branch on the left-hand side is longer than the rising branch at the right-hand side, i.e. increases in GDP per capita lead to increases in fertility only from a relatively high level of income on. In addition, Fig. 3 compares the FE estimation results with real within-country variations in countries which are close to the estimated path: Austria, Canada and Belgium. We can observe that in Belgium, the fertility rebound

[14] Similar results are obtained when regressions are run separately on two sub- samples, one for *lnGDPpc* higher than 10 and one for *lnGDPpc* lower than 10 (results available on request).

[15] Note that the estimated GDP per capita-breakpoint varies largely between the different applied estimation models (see last rows of Tables 1, 6). The predicted minimum never falls outside the observed range of GDP per capita-values (max. GDP pc observed: 65 000 USD) and therefore the trend reversal is not a statistical artefact. However, for most of the models, the location of the predicted minimum is on a relatively high level of GDP per capita, especially for the FE estimation (GDP pc levels above 31 746 US$ are observed for 7 out of 30 countries), and the predicted fertility upswing is relatively slight (see Figs. 3, 4: inverse J-shaped pattern).

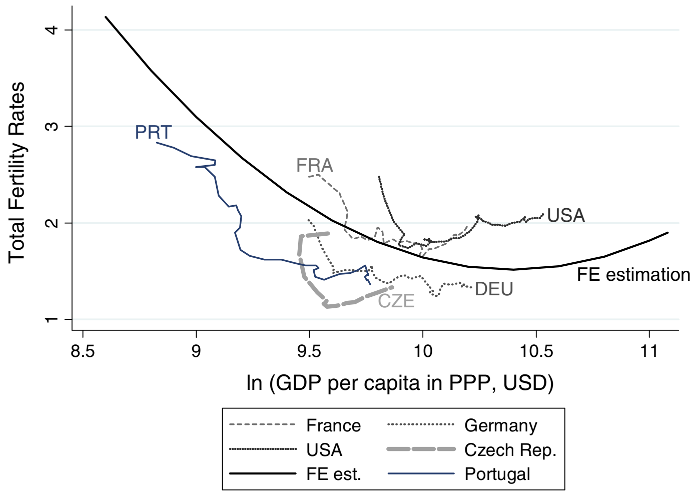

Fig. 4 FE estimation (based on 30 countries, 1960–2007) against observed within-country variation (1960–2007) in France, Germany, Portugal, the Czech Republic and the USA. Data source: OECD Family Database (2010) and authors' estimates

is larger than suggested by the FE results and happened at a quite low level of GDP per capita. In Austria, the impact of immediate further economic growth on fertility is quite inconclusive and the pattern as a whole is situated on a lower fertility level.

Figure 4 illustrates the cases of countries which mostly deviate from the expected path concerning the level of fertility. However, irrespective of periodical fluctuations, the pattern between fertility and income is rather inverse J-shaped in all these countries, which confirms that economic growth decreases fertility up to a certain relatively high level of income, and then increases it. The critical level of GDP per capita actually varies from country to country, these differences being smoothed by the FE estimation. The fertility rebound coming along with a certain level of economic development is particularly observable in France, the United States and the Czech Republic, whereas in Germany and Portugal, the impact of immediate further economic growth on fertility is quite inconclusive.

Figures 3 and 4 lead to the following conclusion: in Eastern and Southern European countries and Germany, economic development comes along with a lower level of fertility than suggested by our empirical results, whereas in countries like France, for example, the regression analysis suggests a lower level of fertility given the country's increase and level of GDP per capita. It is striking that the German pattern is almost parallel to the French one. This means that in these two countries, changes in fertility are almost identically related to changes in income. Yet, the German pattern as a whole is situated on a much lower fertility level than the French one. Moreover, recent economic growth (on highest GDP per capita levels) has induced a much more significant fertility rebound in France than in Germany.

We conclude from Figs. 3 and 4 that in general, our empirical results prove a change in fertility trends going with the process of economic development in OECD countries. Hence, we identify fertility to have a strong negative association with first stages of economic development, while it fades up to a quite high level of GDP per capita from which the relationship seems to reverse. This result is thus consistent with the inverse J-shaped profile suggested by former analysis by Myrskylä et al. (2009), but the switch to a positive relation is found here to happen at quite high GDP per capital level as so far experienced by only few countries (Sweden, Netherlands, United States, Ireland, Luxembourg, Iceland, Norway).

This implies that further economic development is likely to increase fertility in many OECD countries in the future, even though very large increases in per capita economic output would be necessary to raise significantly fertility levels in the current low fertility countries. Moreover, the empirical model estimated so far does not succeed in explaining why in some OECD countries, the inverse J-shaped pattern is situated at quite different fertility levels. The issue remains unclear why in some countries, economic growth increases fertility more significantly than in other countries. The actual GDP per capita level from which fertility rates started to re-increase also varies significantly among countries.

In countries like France, Belgium and New Zealand, it seems that other factors beyond economic advancement are responsible for the relatively high fertility levels and the significant fertility rebound that occurred already at relatively low GDP per capita levels. At the same time, in Japan, Germany, Austria and Eastern and Southern European countries, low fertility levels cannot, or not only, be explained by insufficient economic advancement. Even though our analysis suggests that in these countries too further economic growth increases fertility, it seems likely that fertility increases at a much lower level.

This leads to the question which elements *above and beyond* GDP per capita could make the difference between those two groups of countries. A country typology by Thévenon (2011) shows that the first group provides comparatively high assistance to working parents with young children, whereas the second group is characterised by a relatively limited assistance to families and rather low support for a combination of work and family life. This suggests the benefit we may get by looking further to the relations between female employment and other GDP components with fertility trends.

II. However, before investigating the impact of certain labour market factors on fertility, which are captured by GDP per capita, we now apply some further robustness checks for the quadratic specification. Table 6 in the appendix presents regression results for 2SLS, BE, RE, FDE, CCE and System GMM. Most estimations confirm a convex impact of economic development on fertility with a clear shift in the correlation between the two variables from negative to positive. Only the BE estimator does not confirm a significant breakpoint, implying that the convex impact of *lnGDPpc* on *TFR* is clearly dominated by within-country variation (besides the higher R^2 for the FE model). The estimated break point varies with the applied estimation method. Nevertheless, the important result from the robustness checks is that the inverse J-shaped relation between per capita output and fertility rates is confirmed when running procedures that are designed to best control for potential endogeneity of GDP per capita and for non-stationarity of time series data.

III. We now control whether the convex impact of GDP per capita on TFR still holds when taking into account tempo effects of fertility. This is necessary as the

Table 2 Control for birth postponement

Type of regression	Fixed effects (robust SE)						
Endogenous variable	Total fertility rate (TFR)				Tempo-adjusted total fertility rate (adjTFR)		
Specification	(1) Quadratic model	(2) Quadratic model	(3) Quadratic model	(4) Piecewise regression	(5) Quadratic model	(6) CCE	(7) Piecewise regression
Regressors							
lnGDPpc	−19.42*** (−17.36)	−15.81*** (−8.74)	−14.19*** (−7.51)		−13.95*** (−8.45)	−14.71* (−2.36)	
lnGDPpc^2	0.933*** (16.70)	0.779*** (8.58)	0.701*** (7.45)		0.678*** (8.12)	0.713* (2.27)	
lnGDPpc (pre 10)				−1.10*** (0.137)			−1.15*** (−16.52)
lnGDPpc (post 10)				0.27*** (0.107)			−0.01 (−0.29)
MAB	0.0323* (2.27)						
MA1B		−0.0580*** (−3.35)	−0.577*** (−4.26)	−0.59*** (0.132)			
MA1B^2			0.00965*** (4.04)	0.102*** (0.002)			
Constant	102.0*** (18.11)	83.51*** (9.30)	82.11*** (9.09)	21.3*** (1.73)	72.63*** (8.94)	−23.44 (−0.39)	12.30*** (0.63)
N	845	582	582	582	406	406	406
nb. of countries	30	30	30	30	18+	18	18+
Time period	1960–2007	1960–2007	1960–2007	1960–2007	1961–2006	1961–2006	1961–2006
Test of cross-section independence of residuals abs. (*p* value)					0.43 (0.13)	0.34 (0.02)	0.44 (0.00)
R (within)	0.538	0.493	0.508	0.546	0.433		0.475
R^2 adj.	0.522	0.464	0.479	0.607	0.405	0.88	0.476

t statistics in parentheses, * $p < 0.05$; ** $p < 0.01$; *** $p < 0.001$

+ OECD countries without: Australia, Belgium, Canada, France, Germany, Greece, Korea, Luxembourg, Mexico, New Zealand, Switzerland, Turkey delay in childbirth can be a main determinant of decreases in TFR and the end of birth postponement can be a main determinant for a re-increase of TFR. For this purpose, we first keep *TFR* as endogenous variable and add the mean age of mothers at childbirth (*MAB*) as well as the age of mothers at first childbirth (*MA1B*) as control variables to the FE model. We introduce these two control variables in their linear as well as in their quadratic form because changes in the timing of births can have a nonlinear impact on TFR. In a second step, we use tempo-adjusted TFR (*adjTFR*) as endogenous variable. The tempo-adjusted fertility rate is intended to measure fertility levels within a given period in the absence of postponement. Data on *adjTFR* is available as 3-year moving averages, which smoothes out short-term fluctuations, but covers only 18 OECD countries. We test the quadratic model and the piecewise regression by applying fixed effects. Regression results are shown in Table 2.

All specifications with TFR estimated as a quadratic function of *lnGDPpc* confirm a significantly convex impact of economic development on fertility when indicators of mothers' mean age at childbirth are included (column 1–3).[16] Column 1 suggests that an increase in mothers' mean age at childbirth slightly increases TFR when taking into account the effect of economic development on fertility.[17] This implies that the possibility of birth postponement can contribute to explain why there is a re-increase in fertility at higher stages of economic development. However, as the GDP variables are still significant, it seems that the tempo effect alone is not sufficient to explain the fertility rebound. Column 3 suggests that the average age of mothers' at *first* childbirth has a convex impact on TFR, just like *lnGDPpc*. This suggests that when women start delaying childbirth, fertility rates decrease, but once the birth postponement process stagnates at a relatively high average age of mothers at first childbirth (around 30), TFR re-increase. Once again, we see here that birth postponement actually plays a role in explaining the fertility re-increase that is observed along the process of economic development, but other factors captured by GDP per capita also contribute to the rebound.

The reversal of the impact of economic development, from negative to non-negative or positive, is also confirmed when controlling for the mean age of women at their first birth for the piecewise regression (column 4).

The convex impact of GDP per capita on fertility is also confirmed once we use tempo-adjusted fertility rates as endogenous variable, for which residuals can also be reasonably assumed as cross-sectionally independent (column 5 and 6). The estimation by piecewise linear modelling (column 7) fails to corroborate this finding, however, in a context where data cover a sub-sample of countries only, and

[16] Because data on mean age at birth are available for only a limited time period which also varies across countries, the panel becomes highly unbalanced, which makes it impossible to run tests of cross-section independence for the residuals.

[17] The rate of change of *MAB* is found to have a significantly negative impact on *TFR*, however, while the convex impact of *lnGDPpc* on *TFR* stays unchanged (results available on request). This indicates that the mean age of mothers at childbirth might have an ambiguous impact on *TFR* for our covered time period. Introducing *MAB* and its square as exogenous variables yields insignificant results for both *MAB*-coefficients (results available on request). However, column 3 shows that the impact of *MA1B* on *TFR* is actually convex.

Table 3 GDP decomposition

Type of regression	Fixed effects (robust SE)			
Endogenous variable	Total fertility rate (TFR)			
Specification	Linear model			
Regressors				
InGDPpc	−9.226***			
	(−6.41)			
InGDPpc^2	0.456***			
	(6.32)			
ln(labour productivity)		−0.219***	−0.252***	−0.160***
		(−3.48)	(−4.74)	(−3.29)
ln(avrg. hours per worker)		−0.380	−0.524	
		(−0.72)	(−1.21)	
ln(avrg. hours per worker men)				0.366
				(1.40)
ln(avrg. hours per worker women)				−0.951**
				(−2.91)
ln(employment ratio)		0.632***		
		(4.02)		
ln(employment rate 25–54)			1.392***	
			(7.79)	
ln(employment rate 25–54 men)				0.553
				(1.72)
ln(employment rate 25–54 women)				0.520***
				(3.61)

Table 3 continued

Type of regression	Fixed effects (robust SE)			
Endogenous variable	Total fertility rate (TFR)			
Specification	Linear model			
ln(ratio active population)			−1.785***	
			(−7.78)	
ln(ratio active population men)				3.150***
				(2.57)
ln(ratio active population women)				−4.986***
				(−3.79)
Constant	48.43***	2.945	7.987**	9.594**
	(6.74)	(0.67)	(1.87)	(2.64)
N	368	356	356	315
No. of countries	30	30	30	27*
Time period	1995–2007	1995–2007	1995–2007	1995–2007
R^2 (within)	0.232	0.091	0.327	0.326
R^2 adj.	0.161	0.001	0.259	0.247

t statistics in parentheses, * $p < 0.05$; ** $p < 0.01$; *** $p < 0.001$

+ OECD countries without: Japan, Turkey, USA where the reversal trends regarding the adjusted tempo fertility rates are much smoother than those of the TFR.

We conclude that economic advancement does not only affect the timing of fertility but seems also to affect the 'quantum'-component of fertility. In this case, the reversal of the relation between economic development and TFR from negative to positive is not only a mechanical consequence of the process of birth postponement coming to its end. The results above suggest rather that inherent dimensions of economic development are at play to enable an upturn of fertility trends.

IV. Finally, to investigate *w*hat exactly *behind* economic advancement increases fertility from a certain level of development on, we now decompose GDP per capita into a number of more specific variables. Due to the limited observable time period and due to our particular interest in the fertility re-increase, we analyse *linear* impacts of the decomposition variables on TFR by focussing on the time period 1995–2007. Estimation results are again presented for the FE model with robust standard errors.

Column 1 of Table 3 shows that economic development measured by GDP per capita continues to have a convex impact on fertility when limiting the observed time period to the years 1995–2007. This may be due to the fact that the different components of GDP have ambiguous linear impacts on TFR. In fact, decomposing GDP per capita shows that labour productivity and the ratio of the active population is significantly negatively correlated with fertility, while the employment variables are positively correlated with fertility (columns 2–4). Among the employment variables, a particularly strong association is found between TFRs and women's employment (observed for women aged 25–54—column 4), pointing to the fact that female employment is a key factor for the fertility rebound in OECD countries.[18] Countries with increasing female employment rates are thus likely to experience a fertility rebound. However, unobserved factors may play an important role for the positive within-country association between female employment and fertility, as for example increasing investments in policies supporting the compatibility of family and career for women. This consideration is supported by our finding that women's average working hours are negatively associated with fertility: long average working hours negatively affect the aggregated rate of fertility, in spite of a positive association between this latter and female employment rates.

This again suggests that institutions increasing the compatibility between women's labour market participation and childbearing play an important role for the positive association between female employment and fertility.

[18] Granger Causality tests (Granger 1969) suggest that female employment Granger causes TFR, whereas TFR does not Granger cause female employment. However, Granger causality is not sufficient to imply true causality when the true relation involves three or more variables (Granger 1969). Nevertheless, GMM results taking into account endogeneity issues (and capturing both within- and between-country variation) confirm a significantly positive impact of female employment on TFR (Granger tests and GMM results available on request).

7 Discussion

This study shows that the influence of economic development on fertility trends has changed radically in OECD countries. Our empirical findings confirm a convex impact of economic advancement on fertility rates in OECD countries over the last decades, while there is, in a first stage, a strong negative association between fertility rates and the increase in GDP per capita, the relation weakens and even seems to turn into positive at high level of per capita GDP. We find that this inverse J-shaped pattern of fertility along the process of economic development is actually dominated by within-country variation. This implies that recent economic advancement has been coming along with a slight re-increase in fertility rates in some of the most economically advanced countries. This finding is robust when controlling for postponement of birth. However, while our results unambiguously show that the negative association between GDP per capita and fertility rates weakens with economic advancement, the extent economic development actually produces (or will produce in the near future) a re-increase in fertility is more uncertain. Unobserved factors which co-vary with GDP and influence fertility are important to account for, but their control in the estimation presented here remains non-perfect.

Our finding suggests that further economic development is likely to induce a fertility re-increase in the richest societies, but this increase will be small if driven by increase in GDP per capita only. The Fixed Effects estimation illustrated above suggests that GDP per capita has to reach US$ 66,000 for fertility to increase back to replacement level (2.1 children per women) if we disregard country-specific factors and trends that affect fertility besides economic development (for comparison, the GDP per capita level reached on average US$ 28,100 in the OECD in 2007).

Besides, we also find that several OECD countries do not follow the estimated path of fertility along the process of economic development due to country-specific factors. Some countries demonstrate significantly lower actual fertility rates than the one predicted from GDP trends. Eastern and Southern European countries as well as Germany, Japan and Korea are clearly in that situation. By contrast, Northern European and English-speaking countries and France exhibit higher fertility rates than their expected values. We conclude that economic development is likely, but not sufficient to lift fertility to a higher level in all OECD countries without additional institutional changes.

To gain a deeper insight in the factors that 'cause' the recent increase in fertility rates, we decompose GDP per capita into a number of more specific variables (labour productivity, working hours, and employment) and estimate their impact on fertility. Hereby, we find a positive association between female employment and fertility for within-country variations. This implies that a change in the impact of economic development on fertility from negative to positive is only likely to happen in those countries where economic development has come along with increases in female employment.

The growing participation of women in the labour market is one of the big development changes of the past decades that concern most OECD countries since the 1960s. Its correlation with fertility has changed over time, however. While higher fertility rates were clearly observed in countries with lower rates of female employment in the early 1980s, the opposite seems now to operate with higher fertility rates observed in countries where female employment rates are also higher (OECD 2011). Our finding of a positive association between female employment and fertility is in line with Ahn and Mira's (2002) results, but different from those by Engelhardt et al. (2004) and Engelhardt and Prskawetz (2004), for example, who find for six OECD countries and the years 1960–2000 that the correlation between female labour market participation and fertility is significantly negative only up to the year 1975 and gets insignificant afterward. Kögel (2004) even find a persistent but weakened negative association between female employment and fertility rates. Based on a larger database that includes more countries and more recent time periods, we find a significantly positive association between female employment and fertility even when focussing on within-country variations only. It should be emphasised that this association observed at the 'macro' country level does not always hold at the 'micro' individual level. Matysiak and Vignoli (2008) do a meta-analysis of existing studies of the relationship between women's employment and fertility at the micro-level and show that most micro studies find that women with a continuous career have lower completed fertility than those with interrupted employment spells. The strength of the association is stronger where the male breadwinner model prevails.

Making this coincidence possible implicitly points out the role of unobserved but time-variant institutional factors such as labour market institutions and policy support that enable parents to combine work and family life.

The role of these factors role is accentuated by our observation that countries which combine high fertility and female employment rates generally facilitate a combination of work and family life. Differences in institutional settings among countries with high fertility remain quite large in the OECD (Thévenon 2011). To date, high female employment rates (ages 25–54) over 80 % along with high fertility rates can especially be observed in Finland, Norway, Sweden, Denmark and Iceland. These are countries with high income levels and high public assistance to working parents with young children at the same time. Parental leave schemes are comparatively generous and child care services are also provided area-wide. English-speaking countries support a combination of work and child rearing mainly by in-work benefits, flexible working hours, and both in-cash and in-kind support which target primarily low-income families and preschool children. In contrast, those countries with low fertility and female employment levels, like Eastern and Southern European countries or Germany, are characterised by a relatively low support for work and family reconciliation. France contrasts with these countries with higher female full-time employment rates and at the same time higher fertility rates than Germany, even though Germany has somewhat higher GDP levels. A key difference stands also in the support granted to households with children under preschool to combine work and family. Thus, our results suggest that changes in the impact of economic development on fertility reflect changes in institutional patterns helping parents to balance work and family life.

Further investigation of the relations between economic growth, labour market institutions, the design of work-life balance policies, societal norms and fertility trends is now required to better understand the variety of cross-national patterns. Our estimation results suggest that economic advancement increases fertility in countries that enable female employment, but they do not allow any statements concerning the role of public or private reconciliation instruments, as these are only part of our GDP measures but are not modelled explicitly in this study. An in-depth analysis of the linkages between fertility, institutional settings like norms and family policies, and women's labour market participation has proved to be a fruitful research area (Luci-Greulich and Thévenon 2013).

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Appendix

The IV-estimation results are based on 5-year lags as instruments for the exogenous variables. The estimated coefficients based on 1- to 4-year lags do not differ much and thus are not presented in particular.

The Hausman (1978) test comparing the fixed effects to the RE model suggests that the difference of the estimation results of the fixed and the RE models is systematic. The fact that the *p* value is below 0.05 (0.0371) implies that the hypothesis that the unobserved country effects are not correlated with the error term in the RE model must be rejected. Hence, for our data the fixed effect specification is superior to a RE specification in controlling for unobserved country heterogeneity.

The estimated critical GDP per capital level which leads to a turn in the correlation between fertility and *lnGDPpc* is strikingly low for the First Difference Estimator. This is due to the fact that the first difference of the natural logarithm of GDP per capita approximates the year-to-year relative changes of GDP per capita. Hence, the First Difference Estimator estimates the impact of GDP per capita growth on fertility variations and therefore risks obtaining biased estimates due to an 'underdevelopment' effect. Low levels of GDP per capita are likely to go hand in hand with steeper increases (due to convergence mechanism) and thereby might be rather associated with fertility declines than with fertility increases, referred as a period of demographic transition. This is likely to bias the estimated critical level of GDP per capita. Furthermore, as the First Difference Estimator is not based on level variations, the estimated constant differ largely from the constants estimated by the other estimation methods presented in Tables 2 and 6 and makes it impossible to calculate the minimum level of TFR. Consequently, the First Difference Estimator confirms a convex impact of GDP per capita on TFR while controlling for non-stationarity, but does not permit clear statements about the exact turning point in the correlation between the two variables.

See Tables 4, 5, 6 and Fig. 5.

Table 4 Summary statistics

Variable	Definition	nb. of obs.	nb. of countries	Time period	Mean	Std. dev.	Min.	Max.	Source
TFR	TFR (average number of births per woman)	1418	30	1960–2007	2.19	0.96	1.08	7.26	OECD Family Data Base
adjTFR	Tempo–adjusted TFR, 3 years MA	519	18	1961–2006	1.97	0.32	1.34	3.43	Bongaarts & Feeney Human Fertility Database
GDPpc	Gross domestic product per capita in purchasing power parities (in constant 2005 USD)	1072	30	1960–2007	19812.53	8234.63	2859.90	65001.25	OECD FDB
lnGDPpc	Natural logarithm of *GDPpc*	1072	30	1960–2007	9.80	0.46	7.96	11.08	own calculation
MAB	Mean age of mothers at childbirth	1097	29	1960–2007	27.79	1.40	24.55	31.20	OECD FDB
MA1B	Age of mothers at first childbirth	702	26	1960–2007	25.79	2.09	20.70	30.70	OECD FDB
Labour productivity	GDP/sum of working hours	693	30	1980–2007	26.28	12.29	2.66	78.29	OECD FDB
Avrg hours per worker	Average working hours per worker = sum of working hours/active population	711	30	1980–2007	1800.89	247.41	1334.00	2922.73	OECD FDB
Avrg hours per worker men	Average working hours per male worker = sum of working hours men/active population men	508	27	1980–2007	2198.20	160.94	1871.34	2891.76	OECD FDB
Avrg hours per worker women	Average working hours per female worker = sum of working hours women/active population women	508	27	1980–2007	1814.64	238.13	1244.73	2653.42	OECD FDB
Employment ratio	Active population/total population	787	30	1980–2007	44.18	6.49	27.87	70.08	OECD FDB
Employment rate 25–54	Number of employed persons/working age population (ages 25–54)	710	30	1980–2007	75.72	7.73	53.21	91.60	OECD FDB
Employment rate 25–54 men	Number of employed men/working age population men (ages 25–54)	710	30	1980–2007	87.98	4.32	73,01	97.30	OECD FDB

Table 4 continued

Variable	Definition	nb. of obs.	nb. of countries	Time period	Mean	Std. dev.	Min.	Max.	Source
Employment rate 25–54 women	Number of employed women/working age population women (ages 25–54)	710	30	1980–2007	63.48	14.21	25.59	89.60	OECD FDB
Ratio active population	Active population (ages 25–54)/total population (ages 25–54)	710	30	1980–2007	63.24	2.85	54.65	69.63	OECD FDB
Ratio active population men	Active population men (ages 25–54)/total population men (ages 25–54)	710	30	1980–2007	63.50	2.93	54.05	70.69	OECD FDB
Ratio active population women	Active population women (ages 25–54)/total population women (ages 25–54)	710	30	1980–2007	62.99	2.85	55.20	69.61	OECD FDB

Table 5 Granger causality

Endogenous variable Type of regression	TFR Pooled OLS	InGDPpc Pooled OLS
Regressors		
$(TFR)_{t-1}$	1.261*** (43.03)	−0.0139 (−1.32)
(TFR)	−0.285*** (−9.91)	0.00963 (0.93)
$In(GDPpc)_{t-1}$	0.306*** (3.88)	1.342*** (47.42)
$In(GDPpc)_{t-2}$	−0.268*** (−3.43)	−0.353*** (−12.60)
Constant	−0.357*** (−5.62)	0.131*** (5.80)
N	990	1012
	0.992	0.997

t statistics in parentheses, * $p < 0.05$; ** $p < 0.01$; *** $p < 0.001$

Table 6 Robustness checks

Type of regression	(1) V (2SLS)	(2) Between effects	(3) Random effects	(4) First difference	(5) CCE[1]	(6) System GMM[2]	(7) System GMM with limited lags	(8) System GMM with limited and increased lags
Regressors								
InGDPpc	−14.49*** (−12.29)	−19.14* (−2.05)	−16.89*** (−20.86)	−13.75*** (−11.18)	8.09*** (−4.25)	−15.98*** (−25.97)	−2.369** (−3.00)	−13.62*** (−4.56)
InGDPpc2	0.708*** (11.63)	0.960 (1.98)	0.813*** (19.45)	0.716*** (11.10)	0.44*** (−4.51)	0.788*** (24.58)	0.110** (2.81)	0.711*** (4.62)
Lagged TFR								0.537*** (7.15)
Constant	75.81*** (13.28)	97.10* (2.18)	89.14*** (22.72)	0.036*** (−11.12)	62.63* (2.42)	82.62*** (28.03)	14.93*** (3.55)	65.81*** (4.54)
nb of observations	900	1050	1050	1020	1050	224	164	164
nb. of countries	30	30	30	30	30	30	30	30
Time period	1960–2007	1960–2007	1960–2007	1960–2007	1960–2007	1960–2007	1960–2007	1960–2007
R^2	0.424	0.327 (between)	0.4580 (overall)	0.110				
R^2 adj.	0.422	0.327		0.108	0.977			
Test of cross-section independence of residuals abs. correlation (*p* value)[3]					0.37 (0.30)			
	1 (5 years lags)					81	8	16
nb. of estim. param.	3	3	3	3		3	3	4
Hausman (*p* value)			0.0371					
Sargan (*p* value)						0	0.447	0.172
Sargan-diff. (*p* value)						0	0.251	0.189

Table 6 continued

Type of regression	(1) V (2SLS)	(2) Between effects	(3) Random effects	(4) First difference	(5) CCE[1]	(6) System GMM[2]	(7) System GMM with limited lags	(8) System GMM with limited and increased lags
Instruments for first differences equation						L.(lnGDFpc lnGDFpc2)	L7.(lnGDFpc lnGDFpc2)	L5.(L.TFR L2.lnGDFpc L2.lnGDFpc2)
Instruments for levels equation						D.(lnGDFpc lnGDFpc2)	DL6.(lnGDFpc lnGDFpc2)	DL4.(L.TFR L2.lnGDFpc L2.lnGDFpc2)
Estim. minimum lnGDFpc	9.97	9.96	10.39	9.6	9.1	10.14	10.77	9.58
Estim. min. GDFpc US$	25,964	21,349	32,450	14,794	8,929	25,500	47,500	14,444
Estim. minimum TFR	1.57	1.7	1.42	(−66.05)		1.6	2.17	0.58

t statistics in parentheses, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

(1) Common correlation estimator with country dummies and common correlated effects

(2) System GMM models are estimated on 5-years observations

The GMM method goes back to Arellano and Bond (1991), who propose a difference GMM estimator that transforms the regressors by first differencing, which removes the fixed country-specific effect. Moreover, the use of lagged levels of the regressors as instruments for the first-differenced regressors controls for endogeneity. However, lagged levels of the regressors are likely to be poor instruments for the first-differences equation. We therefore use an augmented version, which implies an efficiency gain over the basic first-difference GMM: a one-step System GMM estimator that goes back to Arellano and Bover (1995) and Blundell and Bond (1998). The System GMM estimator combines a set of first-differenced equations with equations in levels as a "system", using different instruments for each estimated equation simultaneously. This involves the use of lagged levels of the exogenous variables as instruments for the difference equation and the use of lagged first-differences of the exogenous variables as instruments for the levels equation. In addition, System GMM is a dynamic panel estimator that makes it possible to control for the dynamics of adjustment by including a lagged endogenous variable among the exogenous variables.

However, even though System GMM implies an efficiency gain difference GMM by using additional instruments, the System GMM does not completely resolve the problem of weak instruments, as not only lagged levels are likely to be poor instruments for differences, but differences are also likely to be weak instruments for levels (Roodman 2009; Stock and Yogo 2002). Hence, even though the System GMM model proposes the most comprehensive control for a variety of econometric pitfalls, it does not offer a complete control for endogeneity.

Moreover, the fact that the System GMM method uses more instruments than the difference GMM increases the risk that the estimation model is over-identified (Bowsher 2002; Roodman 2009). In order to reduce the number of instruments, we apply the System GMM estimator to edited data. We obtain quinquennial data by dividing the measured time period into five-year sections as follows: we use five-year means for the observations of the endogenous variable and observations of the beginning year of the respective mean for the exogenous variables for every country. This data transformation reduces the number of periods from over 40 to 10 and therefore implies a significant reduction in the number of instruments (from over 800 to around 100 depending on the number of exogenous variables). Moreover, the transformation of the data into quinquennial data allows us to limit time trends, because five-year intervals are less likely to be serially correlated than annual data. In addition, the transformed data makes it possible to intensify the control for endogeneity: for example, if a country's observation of *TFR* is the mean of the years 1980-1984, the corresponding observation of *lnGDPpc* is from 1980, which limits capturing impacts of fertility on GDP per capita.

However, the use of around 100 instruments still implies a significant risk of obtaining a severe overfitting bias (Bond 2002) and reduces the power of the Sargan test to detect invalid instruments (Bowsher 2002). In order to further reduce the number of instruments, we limit the number of lags of the instruments for the first difference and for the levels equation instead of using all available moment conditions. Moreover, we increase the length of the lag of the instruments. By doing so, we obtain a limited number of instruments that does not outnumber the degrees of freedom.

We report the number of instruments and the statistics of the Sargan test of over-identifying restrictions. The Sargan test tests the validity of the instruments and has a null hypothesis of "the instruments are exogenous as a group". A p-value above 0.05 makes it possible to accept this hypothesis. The Sargan difference statistics validate the extra moment restrictions imposed by the level equations in the System-GMM specification in comparison to the Difference-GMM specification.

Box Generalised method of moments applied to the analysis of fertility trends

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