Max-Planck-Institut für demografische Forschung
Max Planck Institute for Demographic Research
Doberaner Strasse 114 · D-18057 Rostock · GERMANY
Tel +49 (0) 3 81 20 81 - 0; Fax +49 (0) 3 81 20 81 - 202;
http://www.demogr.mpg.de

MPIDR WORKING PAPER WP 2001-006
MARCH 2001

The Age at First Birth and Completed Fertility Reconsidered: Findings from a Sample of Identical Twins

Hans-Peter Kohler (kohler@demogr.mpg.de)
Axel Skytthe (askytthe@health.sdu.de)
Kaare Christensen (kchristensen@health.sdu.de)

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Hans-Peter Kohler* Axel Skytthe† Kaare Christensen‡

March 6, 2001

Abstract

In this paper we use new methods and data to reassess the relationship between the age at first birth and completed fertility. In particular we attempt to properly estimate the *postponement effect*, i.e., the reduction in fertility associated with a delay in childbearing, using a sample of Danish monozygotic twins born 1945–60 to control for unobserved heterogeneity. Within-MZ twin pair estimates of the postponement effect indicate that a one year delay in the first birth reduces completed fertility by about 3% for both males and females. The effect is significantly stronger for older cohorts, and it is stronger for females with a late desired entry into parenthood. Analyses that fail to control for unobservables underestimate this postponement effect between 10–25%, and they underestimate the annual decline of this effect by up to 50%. Moreover, our estimates indicate important changes across cohorts in the relevance of child-preferences and ability characteristics for the age at first birth and the pace and level of subsequent fertility.

JEL Classification: J13 (this version: twfirstb02c.tex)

1 Introduction

In the last two decades the mean age at first birth has substantially increased in many countries, and this rise has been associated with a substantial decline in the observed total

*Head of Research Group on Social Dynamics and Fertility, Max Planck Institute for Demographic Research, Doberaner Str. 114, 18057 Rostock, Germany. *Tel:* +49-381-2081-123, *Fax:* +49-381-2081-423, *Email:* kohler@demogr.mpg.de, *www:* http://user.demogr.mpg.de/kohler.

†Research Scientist, Instiute of Public Health and Danish Center for Demographic Research, SDU-Odense, Sdr. Boulevard 23A, 5000 Odense C, Denmark, *Tel:* +45-6550-3034, *email:* askytthe@health.sdu.dk

‡Research Professor, Institute of Public Health and Danish Center for Demographic Research, SDU-Odense, Sdr. Boulevard 23A, 5000 Odense C, Denmark, *Tel:* +45-6550-3049, *email:* kchristensen@health.sdu.dk

fertility rates. For instance, the mean age at first birth in Italy has risen between 1980 and 1996 from 25.0 to 28.4 years, and the total fertility rate (TFR) has declined from 1.64 to 1.19. A similarly striking development occurred in Spain, where the mean age at first birth has risen from 25.0 to 28.9 years during 1980 to 1998, while the TFR declined from 2.2 to 1.16 (Council of Europe 1999, 2000). These delays in childbearing in recent decades are frequently seen as a consequence of increased returns to (female) education and labor-market participation, and the difficulties to combine childbearing with either labor force participation or higher education (Brewster and Rindfuss 2000; Cigno 1994; Gustafsson 1999; Lesthaeghe and Willems 1999; Rindfuss et al. 2000).

In view of these unprecedented increases in the mean age at birth and reductions in fertility, it is essential to revisit an old question in demography and the economics of fertility: the relation between the age at first birth and completed fertility. In particular, we attempt in this paper to properly estimate the *postponement effect*, i.e., the reduction in fertility that is causally associated with a delay in childbearing. At least four reasons render this postponement effect of considerable theoretical and empirical interest to demographers and economists: (a) The onset and completed level of childbearing constitute two central aspects of contemporary fertility behavior that need to be incorporated in any theoretical framework that includes the timing as an important component of fertility decisions (Hotz et al. 1997). (b) Understanding the causal relation between the age at first birth and completed fertility helps to assess the long-term implications of a broad range of socioeconomic changes that transform the human-capital accumulation, labor market participation, partnership formation and related behaviors during early adulthood (e.g., see Buchmann 1989 for a discussion of such changes in early adulthood). (c) The ability to forecast the completed fertility of cohorts, who have not finished childbearing, is greatly facilitated by knowledge that relates early fertility indicators—such as the age at first birth—to completed fertility (Lee 1981). (d) The relation between the postponement of fertility and the completed level of fertility is essential for evaluating recent demographic methods that adjust the commonly used total fertility rate for the distortions caused by a delay in childbearing (Bongaarts and Feeney 1998; Kohler and Ortega 2001; Kohler and Philipov 2001).

The negative association between the age at first birth and completed fertility has been known for quite some time and constitutes an important empirical regularity in the fertility patterns of individuals and cohorts (e.g., Billari and Kohler 2000; Bumpass and Mburugu 1977; Bumpass et al. 1978; Frejka and Calot 2001; Marini and Hodsdon 1981; Presser 1971;

Trussell and Menken 1978). In addition to the rapid change in the timing of fertility in developed countries, the renewed interest in this issue is also stimulated by advances in economic life-cycle models of fertility (e.g., Gustafsson 1999; Happel et al. 1984; Heckman et al. 1985; Ribar 1996; Yamaguchi and Ferguson 1995). In particular, these models suggest several causal mechanisms that can potentially explain a negative relation between the onset of childbearing and the level of completed fertility: an early age at first birth is associated with a longer risk of pregnancy due to imperfect contraception (e.g., Rosenzweig and Schultz 1985); human-capital accumulation and labor market experience prior to the first birth affect the incentives for the timing and number of subsequent children (e.g., Cigno and Ermisch 1989; Gustafsson 1999; Happel et al. 1984; Hotz and Miller 1988; Moffitt 1984); and fecundity tends to decline with age, although recent research suggests that this aspect may be quite modest until age 35 (e.g., Christensen et al. 1998; Frank et al. 1994; Menken et al. 1986).

Despite the apparently ample evidence on the relation between the onset and subsequent level of fertility, the interpretation of the empirical findings is subject to a heated debate. In particular, several authors have pointed out that only a part of the observed negative association is due to causal mechanisms, while another part is due spurious effects (Heckman et al. 1985; Marini and Hodsdon 1981). The distinction between the causal and spurious effects poses a substantial challenge to analysts. In particular, standard estimations of the relation between the age at first birth and completed fertility are hampered by unobserved differences among individuals that affect both the age of entry into parenthood and the pace or level of subsequent fertility. Three broad classes of such factors, which cause biased or inconsistent estimates of the postponement effect in conventional analyses, seem relevant: (a) heterogeneity in the preferences or desires for children; (b) heterogeneity in biological fecundity of women or couples; (c) heterogeneity in 'ability' that affects the incentives to invest in education or labor market skills. Even comprehensive socioeconomic survey data on individual characteristics do not allow to fully control for these unobserved factors, and the resulting estimates thus remain potentially biased even in sophisticated multivariate analyses.

In this paper we therefore choose an alternative strategy and use monozygotic (identical) Danish twins born between 1945–60 in order to overcome the estimation problems caused by unobserved characteristics. Under certain assumptions within-MZ twin estimates allow the identification of the true postponement effect even when individuals differ with respect to their child-preferences, fecundity and ability. The results obtained from these within twin estimates reemphasize some known facts about the interrelation between the onset and level of fertility, and they add some new insights about its changes over time.

Our analyses confirm the existence of a relevant postponement effect for both males and females. On average, an additional year of delay in childbearing reduces completed fertility by 3% for females and 3.4% for males. If interactions with birth years are included, a clear trend towards a reduced relevance of this postponement effect in younger cohorts emerges for both males and females. The failure to account for unobserved factors such as 'preferences' for children and 'economic ability' can substantially distort these estimates of the postponement effect and its change over time. On one hand, ordinary least square regression (OLS) substantially underestimates the relevance of first-birth timing for completed fertility for cohorts born around 1945. In addition, standard OLS estimations also underestimate the pace at which this effect is reduced in younger birth cohorts: the decline in the magnitude of the postponement effect is up to twice as large in the within-MZ estimation as in the OLS results.

The models in this paper also allow an evaluation of how unobserved factors affect the level and timing of fertility. In particular, standard OLS results underestimate the postponement effect for females in older cohorts and overestimate it in younger cohorts. For males, the underestimation is substantially reduced over time. Combined with our theoretical analyses, this pattern of distortions suggests that variation in unobserved characteristics related to the economic costs and returns of a fertility postponement is more important for differences in the timing and level of fertility in older cohorts, while differences in fertility seem to be more related to unobserved variation in the preferences for children in younger cohorts. Across merely 15 birth cohorts born during 1945–60 our analyses suggest important transformations in the determinants of early fertility behavior. Differences in preferences have apparently gained in importance relative to characteristics that determine the economic costs and returns of a delay in childbearing. Such a shift is to be expected in societies that provide increasing compatibility between female labor market careers and fertility and that provide increasingly egalitarian life-course options for individuals. Moreover, this shift is also consistent with related findings that show an increasing relevance of variations in genetic dispositions for differences in fertility, especially for females, and the argument that these emerging genetic influences pertain, at least in part, to genetically mediated differences in motivations and preferences for children (Kohler et al. 1999, 2000; Rodgers et al. 2001).

Finally, our analyses reveal an important sex-difference in the general trend towards a reduced relevance of the postponement effects. For females, but not for males, our analyses show a dependence of the postponement effect on the desired age at first birth. For instance, the fertility of women who have high returns to investing in human-capital and labor market experience depends more strongly on the timing of their first child than the fertility of women with lower returns. This effect is not only relevant within birth cohorts, but has also implications for the changes in the postponement effect over time. In particular, socioeconomic developments that lead to an increasing delay of childbearing partially compensate the general trend towards a reduced relevance postponement effect because completed fertility reacts more sensitive to the timing of first birth at later ages of entering parenthood. Depending on the future increases in the mean age at first birth, this effect can substantially reduce a further weakening of the postponement effect.

2 Theoretical Framework

The empirical identification of the postponement effect from observational data requires assumptions about the influences of unobserved factors on the age at first birth (henceforth AFB) and completed fertility. In order to provide a structure for our empirical analysis, we set out a simple optimizing model for the timing of first births and the choice of subsequent fertility. The three key facts this theory must capture are: (a) the relationship between the AFB and the (logarithm) of completed fertility is approximately linear across a substantial age range; (b) the timing and level or fertility are strongly influenced by the preferences for children, biological fecundity, the incentives to invest in education and labor market experience; and (c) some unobserved factors determine both the AFB and subsequent fertility given the AFB.

Since our empirical estimation is based on data on identical twins, we consider in the following model the fertility behavior of twin i within twin pair j. We allow for heterogeneity in preferences and economic incentives to postpone fertility by introducing the *preference parameter* ϕ_{ij}, the *ability parameter* λ_{ij}, and a parameter reflecting the *age-related costs of postponement* μ_{ij}. These parameters are specific for twin i in pair j and determine systematic differences in the desired timing and level of fertility. In our model we assume that variation with respect to these determinants of fertility behavior is due to either genetic dispositions or socialization in the parental household.

For simplicity we abstract in the subsequent model from intertemporal optimization and joint household decision-making about fertility (an extension of the model that includes assortative mating and joint fertility decisions is presented in the Appendix A.1). We assume that parents maximize a utility function $U(c_{ij}, n_{ij}) = (1-\phi_{ij}) \log c_{ij} + \phi_{ij} \log n_{ij}$ that depends on completed fertility n_{ij}, which is a continuous variable in our model, and (lifetime) consumption c_{ij}. The parameter ϕ_{ij}, with $0 < \phi_{ij} < 1$, reflects that individuals may differ with respect to their child-preferences. Individuals with high levels of ϕ_{ij} tend to gain relatively more utility from children, while individuals with low levels of ϕ_{ij} obtain relatively more utility from the consumption of goods. The above utility function also implies that intended childlessness does not occur (childlessness is associated with a utility level of minus infinity) and all individuals will desire to have at least some children so that $n_{ij} > 0$.

Because important determinants of fertility outcomes, such as partnership formation, contraception and conception, are subject to random influences beyond the control of individuals, parents can only imperfectly control their age at first birth T_{ij} and their completed fertility n_{ij}. In particular, we assume that $\log n_{ij} = \log n_{ij}^d + \varepsilon_{ij}$ and $T_{ij} = T_{ij}^d + \nu_{ij}$, where n_{ij}^d and T_{ij}^d are respectively the desired number of children and the desired AFB, and ε_{ij} and ν_{ij} represent random influences with mean zero. Parents can influence the level and timing of their fertility by choosing a desired number of children n_{ij}^d and a desired age at first birth T_{ij}^d. Their actual timing and level of fertility is then determined by the combination of these desired values and the random influences ε_{ij} and ν_{ij}.

In addition to the preferences given above, the incentives for the timing and level of fertility depend on the following aspects of our theoretical framework. *Human capital formation:* human capital increases the later an individual enters parenthood as $h_{ij} = \lambda_{ij} T_{ij} + \zeta_{ij}$. The term ζ_{ij} is a random influence on this human capital level, and λ_{ij} is the ability parameter that determines the incentives for delaying childbearing in order to facilitate the accumulation of human capital. The dependence of human capital on the AFB captures the fact that children are often incompatible with either formal education or careers with a high accumulation of human-capital on the job. Important differences among individuals, however, exist with respect to their returns of delaying births. On one hand, 'high ability' individuals have large values of λ_{ij} and thus have a large incentive to postpone fertility in order to increase their human capital; on the other hand, 'low ability' individuals have small values of λ_{ij} and therefore expect only small human capital gains from delaying parenthood (e.g., see Behrman and Taubman 1989 for estimates of ability differences among individuals and the relative contributions of environmental and genetic factors to these differences). *Annual wages:* the annual wages w_{ij} earned during each year an individual participates in the labor market is given by $\log w_{ij} = h_{ij}$. *Labor market participation:* women's total years in the labor market depends on the AFB as $L_{ij} = \gamma_{ij} \exp(-\delta_{ij}/2 \cdot (\tilde{T} - T_{ij})^2)$, where $\tilde{T}$ is the timing of the first child that yields the highest labor market participation and is thus optimal in terms of combining childbearing and labor market participation. Earlier and later childbearing is associated with lower labor market participation, and the parameter δ_{ij} reflects the—potentially individual-specific—sensitivity of labor market participation with respect to the timing of fertility. The above relation reflects the fact that women with early childbearing have lower human capital and wages and are therefore more likely not to work after childbirth. Women with a late entry in parenthood, on the other hand, may experience more medical problems and incompatibilities of work and childbearing that tend to reduce their labor market participation after the birth of the first child. *Costs of children:* The specific number of hours of labor, which are supplied by a woman during her time in the labor market, depends on the level of fertility, and we capture this aspect in terms of foregone wages (for analyses of the labor force participation of women after childbirth see for instance Brewster and Rindfuss 2000; Gustafsson et al. 1996; Joshi et al. 1996; Klerman and Leibowitz 1999; Leibowitz and Klerman 1994). The overall costs of children therefore depend on the number of years in the labor market, L_{ij}, and the annual costs of children $p_{ij} = \theta h_{ij} + \mu_{ij} T + \pi_{ij}$. In this latter expression, the term π_{ij} captures costs that are independent of the timing of fertility, the term θ reflects the increase in the costs of children with human capital due to higher foregone wages (we assume $\theta < 1$), and the term μ_{ij} reflects increases in the costs of children with age. The latter, for instance, are associated with higher pre- and postnatal medical costs, higher investments in child-quality, potentially higher health related expenses per child, costs associated with declining fecundity, or similar factors that tend to increase the costs of children as the age of entering parenthood increases (for instance, a well-documented case is the increased risk of Down Syndrome in children that is associated with higher maternal age; see Newberger 2000). We assume for simplicity that θ is constant across individuals, but we allow for the fact that individuals can differ in the age-related costs of postponement μ_{ij}. For instance, these differences can be due variations in fecundity. Moreover, the assumption that $\theta < 1$ reflects a situation where women with higher earning ability or human capital can find more effective child-care arrangements so that the opportunity costs of children in terms of foregone wages increase less rapidly with human capital h_{ij} as does the earning ability w_{ij}.

The above simple specification captures an important aspect in the decision about the timing of fertility: there is a central trade-off between a desire for late fertility in order to accumulate human capital prior to childrearing, either via formal education or through labor market experience, and an incentive to have children early because the costs associated with childbearing tend to increase with human capital and the age at birth. The desired timing T_{ij}^d and level n_{ij}^d of fertility then follow by maximizing the expected utility associated with a choice of T_{ij}^d and n_{ij}^d. In particular, this maximization is possible in two steps. In the first step parents optimally plan the age of entry into parenthood, taking into account their fertility behavior afterwards. In the second step, parents choose consumption c_{ij} and the fertility level n_{ij}^d given their age at first birth T_{ij}. In this utility maximization, individuals take into account their knowledge about their own ability λ_{ij}, the determinants of child costs θ, μ_{ij} and π_{ij}, and the preference parameter ϕ_{ij}. In addition, we assume that the random human capital influence ζ_{ij} is revealed relatively early in life prior to the first birth. The second stage of the decision process can thus include the knowledge of ζ_{ij}.

We first consider the choice of c_{ij} and n_{ij}^d conditional on the AFB. The maximization of expected utility $EU(c_{ij|T}, n_{ij|T}^d)$ is subject to the budget constraint $c_{ij|T} + L_{ij|T} \cdot p_{ij|T} \cdot En_{ij|T} = L_{ij|T} w_{ij|T}$, where the notation '$|T$' in the subscript indicates that the respective terms are conditional on the age at first birth T_{ij}. The first order condition for the number of children yields $\log n_{ij|T}^{d*} = \log[w_{ij|T}/p_{ij|T}] + \log \phi_{ij}$, where $n_{ij|T}^{d*}$ is the optimal fertility choice given the AFB. This desired level of fertility conditional on the AFB follows in terms of the model parameters as

$$\log n_{ij|T}^{d*} = -\beta_{ij} T_{ij} + \zeta_{ij} - \sigma_\varepsilon^2/2 - \pi_{ij} + \log \phi_{ij}, \tag{1}$$

where $\beta_{ij} = (\mu_{ij} - (1-\theta)\lambda_{ij})$. The term β_{ij} in the above relation represents the *postponement effect* that is to be estimated below. The sign and magnitude of this term determines the relation between the desired level of fertility $n_{ij|T}^d$ and the AFB. Since this relation is empirically negative, i.e., later fertility is associated with lower fertility, the parameters in Eq. (1) must satisfy $\beta_{ij} = (\mu_{ij} - (1-\theta)\lambda_{ij}) > 0$. This condition is equivalent to the statement that children must become relatively more costly if parenthood is entered late. If this condition holds, then the postponement effect β_{ij} has the interpretation as the *relative reduction* in fertility that is associated with an increase in the AFB by one year. This reduction is large if: (a) the age-related increases in the costs of children are large (i.e., μ_{ij} is large); (b) the costs of children, particularly in terms of foregone wages, increase rapidly with human capital because fertility and labor market participation are difficult to combine (i.e., θ is large); (c) a delay of fertility has small returns in terms of human capital and wages (i.e., λ_{ij} is small). This latter effect may initially seem surprising. However, if individuals are able to substantially increase their earning ability by delaying childbirth (i.e., λ_{ij} is large), then the income gained by postponing fertility partially compensates for the age-related increases in the costs of children. Hence, a delay in fertility has a smaller effect on the fertility level for women who can off-set for the age-related increase in child costs via higher wages, and vice versa. In addition to the influences operating through the age at first birth, the desired fertility level in Eq. (1) increases the stronger are the preferences for children (i.e., the larger is ϕ_{ij}), the more fertility can be controlled (i.e., the smaller is σ_ε^2), the lower are costs of children that are independent of the AFB and human-capital level (i.e., the smaller is π_{ij}), and the larger is the human-capital shock ζ_{ij}.

The completed fertility of individuals conditional on the AFB, denoted as $n_{ij|T}$, equals the desired fertility level $n^d_{ij|T}$ in Eq. (1) plus the random influences ε_{ij} that result from the imperfect control about the number of children. Based on the knowledge of the relation between completed fertility and the AFB, we can now investigate the first step of the decision process about the optimal choice of the desired AFB. This optimal age follows by maximizing $EU(c_{ij|T}, n_{ij|T})$ with respect to the age at first birth T^d_{ij} and is given by

$$T^{d*}_{ij} = \tilde{T} + \frac{(1-\phi_{ij}\theta)\lambda_{ij} - \phi_{ij}\mu_{ij}}{(1-\phi_{ij})\delta_{ij}}. \tag{2}$$

The observed AFB will differ from this desired age T^{d*}_{ij} due to random influences on conception, partnership formation, etc., and is given by $T_{ij} = T^{d*}_{ij} + \nu_{ij}$.

Despite its simplicity, the above model captures many stylized facts characterizing the relation between the AFB and completed fertility (see Table 1): (a) An increasing preference for children, ϕ_{ij}, leads to an earlier entry into parenthood, to a higher fertility given the AFB, and to a higher completed fertility level (in part these implications depend on the assumption that $\beta_{ij} > 0$). (b) Higher ability in human capital accumulation leads to a later onset of fertility, a smaller postponement effect, and a higher fertility given the AFB. The effect on the overall fertility n^{d*}_{ij} is ambiguous. (c) Higher age-related or human-capital related costs of children lead to an earlier onset of fertility, a larger postponement effect, and a lower fertility given the AFB. (d) In addition, changes in the human-capital shock ζ_{ij} and the parameters σ_ε^2, π_{ij} affect the level of fertility given the AFB, but they do not lead to a different timing of the first child. Similarly, differences in the parameter

Table 1: Effect of parameter changes in our theoretical model on the desired AFB, the postponement effect, the desired fertility level conditional on the AFB, and the desired fertility without conditioning on the age of entry into parenthood.

Outcome variable		**Effect of parameter change on** Desired AFB	Postponement effect	Desired fertility given AFB	Desired fertility
Notation		T_{ij}^{d*}	β_{ij}	$n_{ij\|T}^{d*}$	n_{ij}^{d*}
Increase in parameter					
Preference for children	ϕ_{ij}	$-$	no effect	$+$	$+$
Ability in HC accumulation	λ_{ij}	$+$	$-$	$+$	$+/-$
Age-related costs of postponement	μ_{ij}	$-$	$+$	$-$	$+/-$
HC-related costs of children	θ	$-$	$+$	$-$	$+/-$

Notes: The relations in Table 1 rely in part on the assumptions that $\theta < 1$ and $\beta_{ij} = (\mu_{ij} - (1-\theta)\lambda_{ij}) > 0$

δ_{ij} and the random influences ν_{ij} imply changes in the desired AFB without affecting the desired level of fertility conditional on the AFB.

An important aspect of Table 1 are the two distinct patterns of how parameter changes affect the age at first birth T_{ij}^{d*} and the desired fertility given the age at first birth $n_{ij|T}^{d*}$. On one hand, changes in the economic parameters λ_{ij}, θ or μ_{ij} affect the timing T_{ij}^{d*} and level of fertility $n_{ij|T}^{d*}$ in the same direction; on the other hand, changes in the preference parameter ϕ_{ij} lead to opposite adjustments in the timing T_{ij}^{d*} and level of fertility $n_{ij|T}^{d*}$. The former aspects therefore cause a positive correlation between unobserved determinants of the age at first birth and the level of fertility conditional on the AFB, while the latter aspect causes a negative correlation.

These different patterns have important implications regarding the inference of the postponement effect β_{ij} from observational data. In particular, if the variation in the observed AFB is due to random effects, or more generally, factors that only affect the AFB but not the fertility conditional on the AFB, then the observed relation between the AFB and completed fertility reveals the true postponement effect β_{ij}. However, if the variation in T_{ij} is due to either differences in preferences or ability, the inference of the postponement effect is biased. In particular, the postponement effect is *overestimated* if differences in T_{ij} are due to differences in preferences ϕ_{ij}, while it is *underestimated* if differences in

T_{ij} is due to differences in the ability parameters λ_{ij} and or the age-related postponement costs μ_{ij}. Standard analyses that ignore the existence of such unobserved determinants of fertility behavior, such as preferences for children, ability in human-capital accumulation, and fecundity, can therefore yield a—possibly substantially—distorted inference of the postponement effect. Moreover, neither magnitude nor direction of the bias can be determined from theoretical reasoning and we do not know whether standard analyses over- or underestimate the postponement effect.

3 Empirical Implementation

Our theoretical model implies individual level heterogeneity in the postponement effect β_{ij} that is systematically related to individual characteristics. It is clear, that the individual level variation in the postponement effect β_{ij} cannot be revealed by our—or most other—data on the timing and level of fertility. Our study, however, will identify the average postponement effect $\bar{\beta} = E\beta_{ij}$ in male and female cohorts, and we can investigate systematic changes of this average postponement effect over time as well as with systematic influences on the desired AFB.

Basic model: The analyses of the average postponement effect is based on the estimation of the following relation, which is the empirical equivalent of Eq. (1) in our theoretical model:

$$\log n_{ij|T} = -\bar{\beta} T_{ij} + \eta_{ij}, \tag{3}$$

where $\eta_{ij} = -(\beta_{ij} - \bar{\beta})T_{ij} + \zeta_{ij} - \sigma_{\varepsilon}^2/2 - \pi_{ij} + \log \phi_{ij} + \varepsilon_{ij}$. The residual term η_{ij} in this expression represents the unobserved determinants of the level of fertility conditional on the AFB. This term includes the deviation between the individual and average postponement effect, $(\beta_{ij} - \bar{\beta})$, multiplied with the age at first birth T_{ij}, and it also includes the random variation in the number of children ε_{ij}.

Unfortunately, the OLS estimate of Eq. (3) is biased because the age at first birth T_{ij} is correlated with the residual η_{ij}. In particular, our theoretical analyses have shown that variations in the preference for children ϕ_{ij} induce a negative correlation, while variations in the economic parameters pertaining to the returns to postponing childbearing λ_{ij} and the age-related increase in child-costs μ_{ij} cause a positive correlation (see Table 1). The former correlation will therefore tend to bias the estimate of the average postponement effect upward, while the latter correlation will tend to bias it downward. The net bias resulting from unobserved heterogeneity in both preferences and economic aspect is theoretically ambiguous, but will be revealed by our empirical estimation.[1]

Two primary approaches exist to overcome the estimation problems caused by unobserved preference and economic aspects that affect both the level and timing of fertility. Fixed effect estimations difference out common determinants of behavior over time (in the case of multiple observations over time) or across individuals (in the case when individuals share common prices and other determinants of behavior). Instrumental variable estimations purge the right-hand-side variables of their correlation with the residual by using suitable instrumental variables that are correlated with the endogenous right-hand-side variables but not with the residual.

In our analyses we pursue the former strategy. Since multiple observations of the AFB are impossible for one individual, we use fixed effect analyses of Eq. (3) within monozygotic (identical) twins in order to control for a wide range of unobserved factors that affect both the level and timing of fertility. A similar approach has been used extensively in the analysis of the returns to education (e.g., Ashenfelter and Krueger 1994; Ashenfelter and Rouse 1998; Behrman and Rosenzweig 1999; Behrman et al. 1996; Behrman and Taubman 1976) or household allocations and marriage market effects (Behrman and Rosenzweig 2001; Behrman et al. 1994).

The key to identifying the true postponement effect $\bar{\beta}$ in Eq. (3) with data on monozygotic twins is the assumption that identical twins share the same genetic and shared environmental factors that affect their preferences for children and their costs and returns to postponing childbearing. We therefore assume that the relations $\phi_{1j} = \phi_{2j}$, $\lambda_{1j} = \lambda_{2j}$, $\mu_{1j} = \mu_{2j}$ and $\delta_{1j} = \delta_{2j}$ hold within identical twin pairs. This assumption implies an identical postponement effect and an identical desired AFB for twins within a MZ pair.[2]

The above assumption about identical preferences for children and equal returns and costs of fertility postponement within a MZ twin pair is central for the validity of our estimation technique, but it is not as restrictive as it may initially seem. In terms of our theoretical framework, MZ twins can differ with respect to their attained level of human

[1] The results about the distortions in the OLS estimates of the postponement effect $\bar{\beta}$ are identical with analogous results in the econometric literature about omitted variable bias. Denote as $\Theta_{ij} = -(\beta_{ij} - \bar{\beta})T_{ij} + \zeta_{ij} - \sigma_\varepsilon^2/2 - \pi_{ij} + \log \phi_{ij}$ the total effect of all unobserved characteristics on the fertility level conditional on the age at first birth in (1). The coefficient $\hat{\bar{\beta}}$ that is obtained from a regression of $\log n_{ij|T}$ on the age at first birth then satisfies $\operatorname{plim} \hat{\bar{\beta}} = \bar{\beta} - Cov(\Theta_{ij}, T_{ij}) / \operatorname{var}(T_{ij})$ (e.g., see Griliches 1979).

[2] The second implications that MZ twins share an identical age at first birth can potentially be tested in future surveys that ask relatively young MZ twins about their desired age at first birth.

capital h_{ij}, the fixed component of child-costs π_{ij}, and the labor supply as a function of the AFB (determined by the parameter γ_{ij}). Moreover, the AFB can and will differ among twins due to random influences ν_{ij} on conception, partnership formation, etc. Our assumption about MZ twins is therefore consistent with a broad number of influences on the timing (and also level) of fertility that render the fertility behavior of MZ twins different. The assumption, however, requires that twins are identical with respect to two key influences that pertain on one side to the preferences for children (represented by ϕ_{ij}) and on the other side to the marginal human-capital returns and age-related costs of postponing childbearing (represented by λ_{ij}, μ_{ij}, and δ_{ij}). For MZ twins, the individual-specific terms ϕ_{ij}, λ_{ij}, μ_{ij} and δ_{ij} can therefore be replaced with the twin-pair specific terms ϕ_j, λ_j, μ_j and δ_j that reflect the common genetic dispositions and socialization of identical twins. Moreover, we can denote the common postponement effect and the common desired AFB within MZ twin pairs respectively as β_j and T_j^{d*}.

Since the decisions about the accumulation of human capital and the timing of the first birth are based on expectations about the preferences for children and the economic costs and returns to a fertility postponement, the assumption that MZ twins share these systematic determinants of fertility timing is very plausible. Preferences for children have been shown to depend strongly on socialization in the parental household (e.g., Axinn and Thorton 1996; Barber 2000) and genetic dispositions (Kohler et al. 1999, 2000; Rodgers et al. 2001). Moreover, parental influences, shared parental household conditions, and genetic dispositions constitute important determinants of the returns to human capital and the productivity in human-capital accumulation, and MZ twins are therefore likely to share the same expectations about their human capital returns to delaying fertility. Similarly, the expectations about age-related increases the costs of children are also likely to be shared by MZ twins due to their similar socialized attitudes towards investments in child-quality and their similar health status and functional abilities, especially since this latter similarity persists until relatively high ages (Christensen et al. 2000).

The within-MZ twin estimator for $\bar{\beta}$ is obtained by subtracting Eq. (3) for individuals $i = 1, 2$ within a twin pair j. A slight complication in this method, however, arises because not all individuals experience a first birth. In addition, one may hesitate to extend the above linear relation (3) to relatively late first births, say, above age 32. In order to estimate the postponement effect $\bar{\beta}$ in Eq. (3) we therefore include in our primary analyses only twin pairs in which both twins experience a first birth before some maximum age $T^{\max}$, which we set to age 32. Fortunately, this selection of the sample preserves the properties of the within-MZ twin estimator. In order to see this, denote with $\check{T}_{ij}$ the observed AFB and specify that this AFB is only observed when the latent age at first birth $T_{ij} = T_j^{d*} + \nu_{ij}$ is smaller or equal to $T^{\max}$. The observed age at first birth $\check{T}_{ij}$ is therefore given by

$$\check{T}_{ij} = T_j^{d*} + \mathrm{E}[\nu_{ij} | \nu_{ij} \leq T^{\max} - T_j^{d*}] + \zeta_{ij}, \tag{4}$$

where $\mathrm{E}[\nu_{ij} | \nu_{ij} \leq T^{\max} - T_j^{d*}]$ is the distortion in the observed AFB arising from the fact that we observe the AFB only for the subsample with $T_{ij} \leq T^{\max}$ (e.g., Amemiya 1985). The final term ζ_{ij} in Eq. (4) is a random influence on the AFB that is independent across individuals with $\mathrm{E}\,\zeta_{ij} = 0$. Since the expectation in Eq. (4) depends only on factors that are common to twins in a MZ pair, differencing this relation within twin pairs sweeps out all common unobserved characteristics that affect both the timing of the first birth and the level of fertility conditional on the AFB, including the distortions caused by the fact that both twins in a pair need to experience a first birth prior to age $T^{\max}$ in order to be part of the sample.

An unbiased estimator of the postponement effect $\bar{\beta}$ is thus obtained from the within-MZ twin pair regression

$$\log n_{1j|T} - \log n_{2j|T} = -\bar{\beta} \Delta \check{T}_j + \Delta \eta_j, \tag{5}$$

where $\Delta \check{T}_j = \check{T}_{1j} - \check{T}_{2j}$ and $\Delta \eta_j = \eta_{1j} - \eta_{2j}$. The residual $\Delta \eta_j$ in this within-twin pair relation has an expectation of zero, and most importantly, it is not correlated with the right-hand-side variable $\Delta \check{T}_j$.

Some caveats and robustness tests: The results obtained from the within-MZ twin regression are our *a priori* preferred estimates because we have strong priors that unobserved influences affecting both the timing and level of fertility are relevant. In addition, we can implement Hausman (1978) specification tests that compare the within-MZ twin results with standard OLS analyses of the undifferenced relation in Eq. (3). Unfortunately, several other methods that have been developed to test 'ability bias' in the returns to schooling literature (e.g., Behrman and Rosenzweig 1999) cannot be applied here due to differences in the model structure (the existing tests do not allow for influences of unobserved parameters on the slope coefficients, i.e., in our case the postponement effect $\bar{\beta}$). It is possible, however, to investigate the presence of a twin-specific component in the AFB, which constitutes a violation of our assumptions about identical child-preferences ϕ_{ij}, returns λ_{ij} and costs μ_{ij} of fertility postponement within MZ twin pairs, by using ancillary data on other observable characteristics that differ across and within families (Ashenfelter and Rouse 1998; Griliches 1979). We will present the corresponding correlation analyses in Appendix A.2.

In order to verify the robustness of our analyses, we also report in the appendix an alternative specification that is based on the number of children instead of on the logarithm of fertility as in Eq. (5). We prefer the analysis of log fertility because of its theoretical motivation in Section 2 and its interpretation in terms of relative fertility reductions caused by delays in the age of entering parenthood. Moreover, the logarithm also reduces the skewness of the fertility distribution (see also Section 4 below) and this specification is therefore preferable also for statistical reasons. In addition, we also present some estimates that are based on a measurement of fertility at a different age as a further robustness test of our results.

Extensions of the basic model: The postponement effect $\bar{\beta}$ obtained from the within-MZ twin regression in Eq. (5) pertains to all individuals in our data independent of their birth cohort and of unobserved personal characteristics. In order to relax this assumption, we explore two extensions of our basic model. First, we allow for systematic changes in the average postponement effect across cohorts, which can be due to secular changes in socioeconomic determinants in the average values of the parameters λ_{ij}, π_{ij} and θ that affect the postponement effect. For instance, such changes can be due to improvements in child-care provision that reduce the foregone wages associated with childbearing, or due to progress in health technologies that reduce the age-related increase in child costs. Time trends in the postponement effect that are caused by these secular changes can be identified in our model by including an interaction with the birth-year of cohorts as $\bar{\beta} = \bar{\beta}_0 + \bar{\beta}_1 \cdot (\text{birth-year})$. If the coefficient $\bar{\beta}_1$ is negative, the average postponement effect declines in more recent cohorts.

Second, it is possible that the fertility of 'late starters', i.e., parents whose preference, ability and costs characteristics ϕ_{ij}, λ_{ij}, μ_{ij} imply a relatively late desired AFB in Eq. (2), is more sensitive to variations in the onset of childbearing than the fertility of parents who tend to be 'early starters' according to their unobserved characteristics. We can investigate this hypothesis by testing for a dependence of the postponement effect $\bar{\beta}$ on the average AFB within a twin pair. This average AFB in a twin pair is an estimator of the expected AFB defined as $T_{ij}^{e*} = T_j^{d*} + \mathrm{E}[\nu_{ij}|\nu_{ij} \leq T^{\max} - T_j^{d*}]$, where T_j^{d*} is the common desired AFB for twin pair j, and $\mathrm{E}[\nu_{ij}|T_{ij} \leq T^{\max}]$ is the distortion caused by censoring the observations at a maximum age at first birth $T^{\max}$. Moreover, if $T^{\max}$ is chosen sufficiently high so that censoring does not occur, then the average AFB within a twin pair is an estimator of the desired age at first birth T_j^{*d} in Eq. (2).

A dependence of the postponement effect on the average AFB in a twin pair can be incorporated in the above theoretical and empirical model by specifying

$$\bar{\beta}_j = \bar{\beta}_0 + \bar{\beta}_2 \bar{T}_j, \tag{6}$$

where $\bar{T}_j$ is the average AFB within twin pair j. If $\bar{\beta}_2$ in this relation is positive, then the fertility of 'late starters' with a late desired onset of childbearing is more strongly affected by variations in the AFB than is the fertility of 'early starters'. This specification yields a dependence of completed fertility on the AFB as $\log n_{ij|T} = -\bar{\beta}_0 \check{T}_{ij} - \bar{\beta}_2 \bar{T}_j \check{T}_{ij} + \eta_{ij}$. While an OLS estimation of this relation is biased, the parameters $\bar{\beta}_0$ and $\bar{\beta}_2$ can be estimated consistently by a within-MZ twin regression that is obtained by differencing the above relation within twin pairs. The respective regression then contains the difference between the AFB within a twin pair, $\Delta\check{T}_j$, and an interaction of this difference with the average AFB in a twin pair, $\bar{T}_j$, as

$$\Delta \log n_j = -\bar{\beta}_0 \Delta\check{T}_j - \bar{\beta}_2 \bar{T}_j \Delta\check{T}_j + \Delta\eta_j. \tag{7}$$

This estimation can also be combined with a dependence of the postponement effect on cohort (or birth year), which we discussed above.

From a theoretical perspective, a slightly different specification for the postponement effect may be more appealing than a dependence on the average AFB in a twin pair. In particular, we would like to investigate a dependence of the postponement effect on the desired AFB as $\bar{\beta}_j = \tilde{\beta}_0 + \tilde{\beta}_2 \bar{T}_j^{*d}$. Unfortunately, the desired age at first birth in a twin pair, $\bar{T}_j^{*d}$, is unobserved. Its only available measurement is the average AFB within a twin pair, but this measurement is subject to two caveats: First, it is only unbiased if there is no relevant censoring of the age at first birth. Second, it is subject to random measurement error. This measurement error equals $\frac{1}{2}(\nu_{1j} + \nu_{2j})$, where ν_{ij} is the random deviation of the observed AFB of twin i in pair j from the desired AFB in twin pair j. Due to these caveats, the estimate of the coefficient $\bar{\beta}_1$ on the basis of Eq. (7) is a biased estimator of the coefficient $\tilde{\beta}_1$ that reflects changes in the postponement effect due to increases in the desired AFB. However, in the specific case when there is no censoring of the age at first birth, the standard results for measurement error in regression analyses suggests that $\bar{\beta}_1$ is biased towards zero as compared to $\tilde{\beta}_1$. The estimates of $\bar{\beta}_1$ obtained from the within-MZ twin regression in Eq. (7) are therefore a conservative estimate for the dependence of the postponement effect on the desired age at first birth.[3]

4 Data

Our analyses are based on the *Danish Twin-Fertility Database* (*DTFD*) and include the fertility of all same-sex twin pairs born after 1945 whose twin status and zygosity could be identified. The verification of the zygosity of same-sexed twin is based on a survey including four questions about the similarity of the twins, and this method been proved to determine the zygosity correctly in approximately 95% of the twin pairs (Hauge 1981). The Danish Twin-Fertility Data are generated by merging the Twin Register with the fertility information in the Danish Civil Registration System (*CRS*), which encompasses all persons who have lived in Denmark since 2 April 1968 and have registered with the national registration offices. This link between the *CRS* and the Twin Register provides a complete fertility history of all twins, and the linkage includes all births until December 31, 1998. The linkage was performed if the information in the *CRS* contained at least one parental reference to a twin (father or mother).[4] The information obtained in this manner about each child born to twins in the Danish twin registry includes the year of birth, age of parent at birth, sex of child, and the year of death if the child has died. Infant deaths before April 2, 1968 are not included in the data set since these events are not registered in the *CRS*. However, only relatively few births for the cohorts 1945 and later have occurred prior to 1968, and with an infant mortality rate about 20 per 1000 live births in the 1960's the number of missing children due to infant deaths is very low. Hence, for twins born

[3]Some progress regarding the distortion caused by measurement error in the desired AFB is possible by using the specific structure of the twin data. These analyses, however, are beyond the scope of the present paper. Nevertheless, if there is no censoring in the age at first birth, one can show that the estimate of $\bar{\beta}_1$ obtained from the within-MZ twin regression in Eq. (7) is biased towards zero as compared to the correct parameter $\tilde{\beta}_1$ if $E\nu_{ij}^3 = 0$ and $E[\nu_{ij}^3 T_j^{d*}] < .5 \cdot E[\nu_{ij}^4] + .5 \cdot [E[\nu_{ij}^2]^2]$, where T_j^{d*} is the desired AFB and ν_{ij} is the random deviation of the observed AFB of twin i in pair j from the desired AFB. Therefore, if the overall distribution of ν_{ij} is not skewed, and if the distribution of ν_{ij} conditional on the desired AFB does not shift substantially from very right-skewed to at low levels of T_j^{d*} to very left-skewed at high levels of T_j^{d*}, then the estimate of $\bar{\beta}_1$ obtained from the within-MZ twin regression in Eq. (7) is a conservative estimate of the $\tilde{\beta}_1$.

[4]The links in the *CRS* between children and parents represent the legal parenthood, and the register contains no information about the biological parents of adopted children. Therefore it is not possible to distinguish between biological and adoptive parents in the data set. However only about 1.2% of the children born in the study period are adopted according to the official statistics, and this proportion is likely to be much lower for early fertility. Moreover, still born children are not included in the data set, since no Personal Number is assigned to them.

Table 2: Summary statistics for all twins in MZ pairs born 1945–60, and for twins included in the estimation sample consisting of all MZ twin pairs in which both twins experienced a first birth by age 32

	Females			**Males**		
	Mean	Std. Dev.	within MZ pair corr.	Mean	Std. Dev.	within MZ pair corr.[b]
All twins in MZ pairs born 1945-60						
No. of twins	1712			2286		
Prop. childless at age 32	0.196	0.397	0.28	0.315	0.465	0.27
# of childless at age 38	0.137	0.344	0.26	0.225	0.418	0.28
# of children at age 38	1.800	1.035	0.31	1.608	1.145	0.33
Age at first birth[a]	25.10	4.666	0.41	27.11	4.979	0.31
MZ twins included in estimation						
No. of twins	1182			1208		
# of children at age 38	2.163	0.795	0.25	2.224	0.860	0.15
Log of fertility at age 38	0.703	0.378	0.15	0.724	0.398	0.11
Age at first birth	24.00	3.620	0.36	25.38	3.372	0.32

Notes: (*a*) only for twins who experienced at least one birth; (*b*) Pearson correlation coefficient.

after 1945 the link with the *CRS* provides an almost comprehensive coverage of fertility, and the quality of the fertility information in *DTFD* can be considered as very high.

Our analyses are based on all male and female MZ twin pairs in the *DTFD* who are born during 1945–60. In addition, we restrict the analyses to twin pairs in which both twins experienced their first birth prior to age 32 (in Section 3 we show that this restriction does not bias the estimates of the postponement effect). Moreover, we exclude a few twins (10 females and 7 males) who experienced a first birth at age 16 or earlier (teenage fertility is quite rare in Denmark, especially when compared to the US). Fertility at age 38 is taken as a proxy for completed fertility. This choice is a trade-off between, on one hand, choosing a relatively late cut-off age in order to capture completed fertility as much as possible, and on the other hand, including a relatively broad range of cohorts in the analyses in order to investigate changes in the postponement effect over time. Table 2 summarizes the relevant fertility information about all MZ twins in our data and for those twins who are included in the subsequent analyses.

The summary statistics for the fertility of twins in Table 2 shows that about 80% (68%) of all female (male) twins experience a first birth by age 32, and 85% (78%) by age 38. Hence, about 93% (88%) of females (males), who have at least one child at age 38, experience their first birth up to age 32. The censoring caused by the selection of our sample is therefore quite modest, especially for females. The average fertility level at age 38 is between 1.6 and 1.8 children for all twins, and around 2.2 for those twins who are included in our estimation sample. The investigation of the logarithm of completed fertility in our subsequent analyses is suggested by our theoretical model, but this transformation is also appropriate from a statistical viewpoint because it reduces the skewness of the fertility distribution. On average, the logarithm of fertility for males and females in our estimation sample is slightly above 0.7. The age at first birth is on average between 25–27 years for all twins, and between 24 and 25.3 years for twins included in the subsequent analyses.

Since we observe fertility behavior for the subset of twins born 1945–55 in our estimation sample at least until age 43, we can also verify that fertility at age 38 is a good approximation of completed fertility in our sample of twins who have a first birth prior to age 32. Females have on average of an additional 0.018 children after age 38 and males have an additional 0.026 children. This additional fertility after age 38 does not represent a relevant truncation of fertility, and fertility at age 38 captures almost completed fertility. Moreover, the additional number of children after age 38 is virtually uncorrelated with the age at first birth in our estimation sample (the correlation is less than 0.01 for females, and 0.026 for males). There is thus no indication that fertility at age 38 is a less suitable measure of completed fertility for those with a late onset of childbearing.

The combination of a shared parental environment and identical genetic dispositions accounts—perhaps not surprisingly—for a substantial part of the variation in fertility outcomes across individuals. The within-MZ twin pair correlations for females range from .25 to .31 for the number of children and from .36 to .41 for the age at first birth. For males, the correlation in the number of children varies from .15 to .33 and equals about .3 for the age at first birth. These correlations also indicate for the twins included in our estimations that genetic and shared environmental influences tend to be stronger for outcomes that occur relatively early in life, such a the age at first birth, as compared to outcomes that are determined latter in life, such as the completed fertility (see also Kohler et al. 1999, 2000; Rodgers et al. 2001). Moreover, the negative and approximately linear association between the onset of fertility and completed fertility also applies to the Danish twin data and is shown in Figure 1 along with a fitted regression line. For both males and females, fertility decreases systematically and importantly with the age at first birth, and the average number of children declines by about 2.6% for each additional year by which the first birth is postponed. An interesting aspect of Figure 1 is the absence of significant

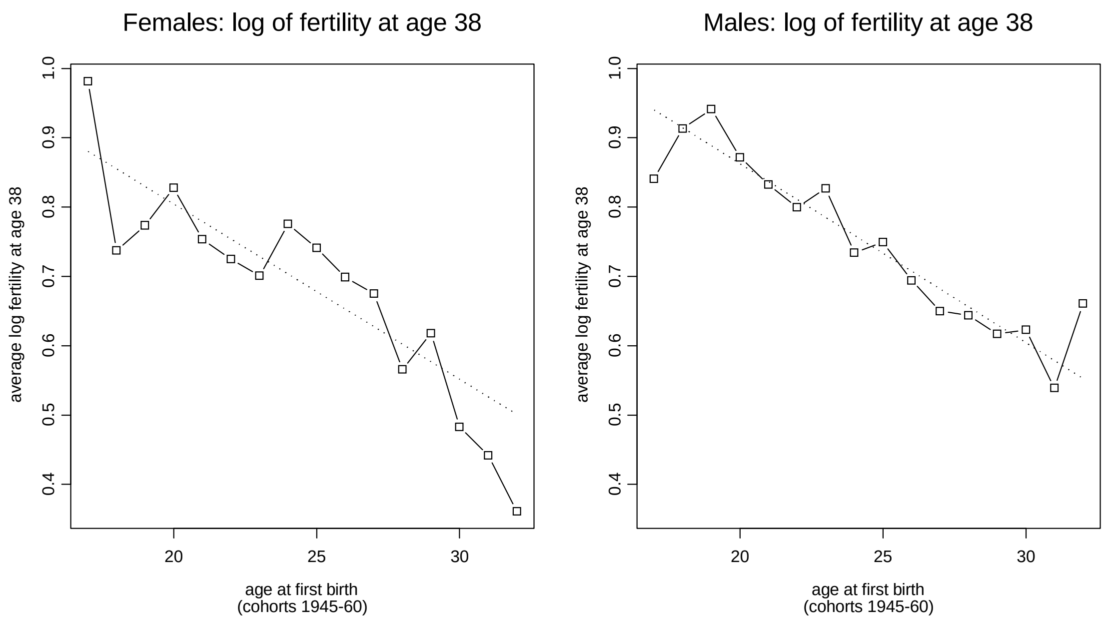

Figure 1: Age at first birth and fertility at age 38 for female and male MZ twins born 1945–60

male-female differences in this relation between the age at first birth and fertility at age 38.

An important concern in using monozygotic twins in the estimation of the returns to human-capital are potential errors in the measurement of schooling because fixed effect estimations tend to exacerbate problems of measurement error (e.g., Ashenfelter and Krueger 1994; Ashenfelter and Rouse 1998; Behrman and Rosenzweig 1999; Behrman et al. 1994, 1996). The analogous issue in this paper is measurement error in the age at first birth. Since the information on the age at first birth and completed fertility is obtained from the Danish civil registration system—which is considered to be of very high quality for the period when the twin cohorts used in the estimation experience their fertility—problems of measurement error do not seem to be relevant for our estimations. Moreover, our data do not include the information that would allow us to implement the various instrumental variable techniques that are used in the corresponding literature on the returns to human-capital.

5 Results

5.1 Estimate of postponement effect and its change across cohorts

Main findings: Table 3 reports the postponement effect $\bar{\beta}$ obtained from Danish twins born 1945–60 using the logarithm of fertility at age 38 as a measure of completed fertility. We report the consistent estimates of the postponement effect obtained from the within-MZ pair regression in Eq. (5), and we include for comparison the corresponding results obtained from an OLS estimation of Eq. (3). We report robust standard errors for the within-MZ and OLS estimates in order to accommodate the potential heteroscedasticity of the residual in these regressions. For the OLS analyses, these robust standard errors also account for the within twin pair correlation of the residual term. The percentages in squared parentheses below the standard errors in Table 3 report the relative difference between the OLS results and the consistent within-MZ estimates, an these percentages assess the relative bias of OLS as compared to the within-MZ estimates. In addition to the separate analyses for females and males, which do not show statistically different results in Table 3, we also report the results of a joint estimation that includes male and female MZ twins.

The first set of estimates in Panel A of Table 3 considers a constant postponement effect for all cohorts. According to these estimates, delaying the onset of fertility by one additional year reduces completed fertility by 3% for females and 3.3% for males. A standard OLS regression of the relation between the age at first birth and completed fertility underestimates this postponement effect by 11.4% for females and 20% for males. If both sexes are combined, the consistent within-MZ pair estimates suggests that postponing the first birth by one year reduced completed fertility 3.24%, while OLS yields a postponement effect of 2.68%.

The direction of the distortion in the OLS results may initially seem surprising. If variations in preferences are the most important determinant in the timing and level of fertility, that is if 'early starters' have a particularly strong desire for children and thus choose to have relatively many children and a relatively early onset of fertility, the postponement effect obtained from OLS analyses should be biased upwards. The distortions in Panel A, however, are in the opposite direction. The fact that the OLS estimates are biased downward as compared to the within-MZ estimates implies that the sum of the unobserved characteristics included in the random influences η_{ij} on the level of fertility in Eq. (3) are positively correlated with the age at first birth. This suggests that the

Table 3: Postponement effect $\bar{\beta}$ for Danish twins born 1945–60. Dependent variable is logarithm of fertility at age 38. Twin pairs are included in the sample if both twins in a pair have experienced their first birth by age 32 (for summary statistics see Table 2). The percentages in square parentheses below the standard errors report the relative differences between the OLS and within-MZ pair estimates.

	Females		**Males**		**Females and Males**	
Method	**Within MZ pair**^{a}	**OLS**^{b}	**Within MZ pair**^{a}	**OLS**^{b}	**Within MZ pair**^{a}	**OLS**^{b}
Panel A: Constant postponement effect for all cohorts						
	Model 1	*Model 2*	*Model 3*	*Model 4*	*Model 5*	*Model 6*
$\bar{\beta}$	0.0299	0.0265	0.0336	0.0267	0.0317	0.0264
	(0.0049)**	(0.0032)**	(0.0059)**	(0.0036)**	(0.0038)**	(0.0024)**
		[-11.4%]		[-20.5%]		[-16.7%]
Panel B: Interaction with birth year as $\bar{\beta} = \bar{\beta}_0 + \bar{\beta}_1 \cdot (\text{birth-year}-1945)$						
	Model 7	*Model 8*	*Model 9*	*Model 10*	*Model 11*	*Model 12*
$\bar{\beta}_0$	0.0434	0.0337	0.0487	0.0391	0.0460	0.0347
	(0.0082)**	(0.0054)**	(0.0102)**	(0.0060)**	(0.0065)**	(0.0040)**
		[-22.4%]		[-19.7%]		[-24.6%]
$\bar{\beta}_1$	-0.0020	-0.0010	-0.0022	-0.0018	-0.0021	-0.0012
	(0.0010)$^+$	(0.0007)	(0.0011)$^+$	(0.0007)**	(0.0008)**	(0.0005)*
		[+50.0%]		[+18.2%]		[+42.9%]
Panel C: Cohort-specific estimation						
	Model 13	*Model 14*	*Model 15*	*Model 16*	*Model 17*	*Model 18*
$\bar{\beta}$ for cohorts 1945–52	0.0381	0.0310	0.0487	0.0360	0.0431	0.0328
	(0.0062)**	(0.0042)**	(0.0080)**	(0.0048)**	(0.0050)**	(0.0032)**
		[-18.6%]		[-26.1%]		[-23.9%]
$\bar{\beta}$ for cohorts 1953–60	0.0185	0.0198	0.0161	0.0140	0.0173	0.0175
	(0.0077)*	(0.0048)**	(0.0084)$^+$	(0.0052)**	(0.0057)**	(0.0035)**
		[+7.0%]		[-13.0%]		[+1.2%]
χ^2 test for equal $\bar{\beta}$	3.95*	3.1$^+$	7.88**	9.81**	11.6**	10.8**
	(df = 1)	(df = 1)	(df = 1)	(df = 1)	(df = 1)	(df = 1)
N	591	1182	604	1208	1195	2390

Notes: Robust standard errors in parentheses. *p-values:* $^+$ $p < 0.10$; * $p < 0.05$; ** $p < 0.01$. Standard errors are calculated using White's (1980) heteroscedasticity consistent variance estimator. The OLS standard errors are additionally adjusted for within twin-pair correlation of the error term. (a) The fixed effect regressions do not include a constant term. (b) The OLS models 2, 4, 6, 8, 10, 12 include a third-order polynomial in (birth-year – 1945) in order to account for cohort trends in completed fertility, and OLS models 14, 15, 18 include a dummy for birth cohorts 1953–60. The combined OLS models 6, 12, 18 additionally include a dummy for females.

characteristics related to the economic returns of a fertility postponement and the age-related increases in the costs of childbearing, which are captured by the parameters λ_{ij} and μ_{ij} in our theoretical model, are more important for variations in the age at first and completed fertility than unobserved preference parameters. This inference is possible since only the former imply a positive correlation of η_{ij} with the observed age at first birth, while variations in preferences tend to imply a negative correlation.

In a recent analysis, Morgan and Rindfuss (1999) have argued on the basis of CPS data in the United States that the link between the age at first birth and completed fertility has been weakening over time. We provide two tests for this hypothesis in Table 3 based on (a) an interaction of the postponement effect with birth-year as $\bar{\beta} = \bar{\beta}_0 + \bar{\beta}_1 \cdot$ (birth-year$-$1945) in Panel B, and (b) a cohort-specific estimation of the postponement effect $\bar{\beta}$ in Panel C. Both results indicate that the magnitude of $\bar{\beta}$ has been declining over time in a statistically significant pattern, and the postponement effect has become less pronounced in more recent cohorts. In particular, the interaction of $\bar{\beta}$ with birth year in the joint male-female model implies that 1-year increases in the age at first birth reduces completed fertility by about 4.6% for cohorts born around 1945, and that this effect has been decreasing by approximately .21 percentage points per birth year. If this linear trend continues, the postponement effect would vanish for cohorts born in the late 1960s. A similar substantial reduction of the postponement effect with time is also suggested by the cohort-specific estimation in the bottom part of Table 3. While the postponement effect is around 3.8–4.9% for cohorts 1945–52, it decreases to about 1.7–1.85% for cohorts 1953–60.

A second interesting aspect of Table 3 is the comparison between the OLS and within-MZ twin estimates. The former underestimate the postponement effect for cohorts born 1945 in Panel B, but they also underestimate the extent to which this postponement effect has reduced across birth cohorts. For instance, according to the OLS results of model 12, the postponement effect is reduced by .10 percentage points per birth year while the fixed effect estimates suggest an annual reduction by more than .21 percentage points (the p-value of this coefficient is .056). The decline may therefore be almost twice as fast as is suggested by standard OLS analyses that fail to account for the implications of unobserved characteristics. Our analyses therefore support the conclusions of Morgan and Rindfuss (1999) about a declining relevance of the age at first birth completed fertility. More importantly, this decline may even be occurring at a faster pace than revealed by existing analyses that do not account for potential biases due to unobserved characteristics.

The underestimation of the time trend is also noteworthy because it additionally suggests that the direction and/or magnitude of the OLS distortion is changing across cohorts. This presupposition is confirmed in the cohort-specific estimation: OLS *underestimates* the postponement effect in Panel C for both males and females born during 1945–52, while it *overestimates* the postponement effect for females and the combined male and female analyses for the cohorts 1953–60. The direction of the distortion is not reversed for males, but it is nevertheless substantially reduced (in relative terms) by a factor of .5.

Specification and robustness tests: Hausman tests provide a general possibility to test for the presence of unobserved factors that lead to a correlation between the residual in Eq. (3) and the right-hand-side variables. In the separate analyses for males and females, however, this test does not reject the null-hypothesis, which is due to the relatively small number of observations in these sex-specific analyses. In the joint male-female analyses Table 3, on the other hand, the null-hypothesis that OLS provides consistent estimates of the postponement effect is rejected in all models with a p-value of .075 or lower.

In order to verify the robustness of our results, we also provide two alternative specifications of our analyses in the Appendix. First, we use the number of children at age 38, instead of the logarithm of fertility at age 38, as dependent variable. In this case the postponement effect measures the absolute—instead of relative—reduction of in completed fertility for each year the entry into parenthood is delayed. The corresponding estimates in Table A1 agree with the results discussed above. The overall postponement effect in the combined analyses is .067, and the results also reveal a substantial reduction in this postponement effect in more recent cohorts. Moreover, similar to our analyses in Table 3, the negative bias of the OLS estimates in older cohorts turns into a positive bias in younger cohorts in the female analyses, while the negative bias is substantially reduces in the male and joint male-female regressions.

Second, we re-estimate our analyses using only the subset of MZ twins who are born during 1945–55 in order investigate whether our results are sensitive to the cutoff point of age 32 for the age at first birth and the measurement of fertility at age 38. The results in Table A2 are based on all twins pairs born during 1945–55 where both twins experience a first birth prior to age 35, and completed fertility is measured at age 43. This specification of our model eliminates any relevant censoring in the age at first birth because only 5% of all MZ twins born during 1945–55 experience a first birth after age 35. Moreover, the only very few women experience have additional children after age 43, and our dependent variable is very close to completed fertility. Fortunately, the results in Table A2 agree highly with the corresponding results discussed in the main text on the basis of Table 3. Interestingly, the secular change in the postponement effect across cohorts in Table A2 is even stronger than in Table 3 (-.005 instead of -.0021), which indicates that the trend towards a smaller postponement effect may have been particularly fast across the early cohorts in our study.

Interpretation: Our primary results in Tables 3 and A1–A2 show that (*a*) a relevant postponement effect exists in all cohorts, and the relevance of this postponement effect has been declining in younger cohorts; (*b*) there are no systematic male-female differences in this postponement effect and its change over time, despite the fact that the costs and benefits of children usually have very sex-specific patterns; (*c*) standard analyses tend to underestimate the magnitude of the postponement effect and the extent of decline in more recent cohorts; and (*d*) important structural changes occur in the correlation between unobserved determinants of the AFB and the level of fertility conditional on the AFB. In particular, an initial positive correlation vanishes across cohorts in all specifications, and for females it even reverses into a negative correlation. The reduction of the negative OLS bias caused by this changing correlation structure, or its reversal into a positive bias as for females, must be due to the fact that variation in the level of fertility, conditional on the age at first birth, is increasingly related to variation in the preferences for children (i.e., the parameter ϕ_{ij} in our theoretical model), and/or decreasingly related to variation in the returns and costs of fertility postponement (i.e., the parameters λ_{ij} and μ_{ij} in our theoretical model).

Although this evidence is indirect, the interpretation of these results in terms of our theoretical model suggests that the relative importance of unobserved characteristics leading to variations in the age at first birth and completed fertility has been shifting. In earlier cohorts, characteristics pertaining to the costs and returns of fertility postponement seem to be most important, and these aspects seem to have lost some of their relevance for variation in fertility behavior while characteristics pertaining to preferences potentially have gained in their importance.

This interpretation is very plausible in terms of the socioeconomic conditions and changes in Denmark. A shift towards an increased role of preferences is to be expected in societies that provide increasing compatibility between female labor market careers and fertility, and that provide increasingly egalitarian life-course options for individuals (for related discussions, see Gauthier 1996; Gustafsson et al. 1996). In particular, improved compatibility of childbearing and labor market participation reduces the postponement effect via a reduction of the opportunity costs of children in terms of foregone wages. A further important mechanism leading to a reduced postponement effect is a higher return to a fertility postponement in terms of wages and human capital, and this higher return can be caused by changes in the educational system, technological progress, and changes in the occupational structure. Moreover, the shift towards an increased relevance of preferences for variation in the onset and level of fertility is also consistent with findings that show an increasing importance of genetic dispositions for differences in fertility, and the argument that these emerging genetic influences pertain, at least in part, to genetically mediated differences in motivations and preferences for children (Kohler et al. 1999, 2000; Rodgers et al. 2001).

5.2 Dependence of postponement effect on average (desired) age at first birth

The above analyses assumed that the average postponement effect $\bar{\beta}$ applies equally to all individuals independent of the characteristics that determine the desired age at first birth. In this Section we relax this assumption and investigate whether important differences in the marginal impact of a fertility postponement exist between twin pairs with a different average AFB. In Section 3 we have shown that the average AFB within a twin pair is an estimator of the expected age at first birth (in the presence of censoring) or the desired age at first birth (if there is no relevant censoring). The results can therefore be interpreted as systematic differences in the postponement effect for individuals with different desired (expected) timing of entry into parenthood.

Main findings: Table 4 reports the results obtained from the within-MZ twin regression that estimates a dependence of the postponement effect on the average AFB within a twin pair (see Eq. 7). The corresponding OLS results are included for comparisons. Moreover, because males and females differ in important aspects, we do not include a combined model for males and females.

The results in Panel A of Table 4 show that the postponement effect for females depends significantly on the average age at first birth within a twin pair. The completed fertility of women in twin pairs with a high average AFB, i.e., women who in terms of our theoretical model tend to have a relatively low preference for children and/or large economic incentives to delay childbearing, is more sensitive to variations in the age at first birth than is the completed fertility of women with a relatively low desired age at first birth. For males such a dependence does not seem to be present. Moreover, OLS analyses

Table 4: Dependence of the postponement effect on the average AFB a within twin pair, cohorts 1945–60. The dependent variable and the sample of twins are identical to our earlier analyses in Table 3.

	Females		**Males**	
Method	**Within MZ pair**^{a,c}	**OLS**^{b,c}	**Within MZ pair**^{a,c}	**OLS**^{b,c}
Panel A: Postponement effect depends on desired age at first birth as $\bar{\beta} = \bar{\beta}_0 + \bar{\beta}_2 \bar{T}_j$				
$\bar{\beta}_0$	0.02674	0.02873	0.03351	0.03574
	(0.00519)**	(0.00496)**	(0.00593)**	(0.00583)**
$\bar{\beta}_2$	0.00476	-0.00014	-0.00069	-0.00055
	(0.00197)*	(0.00025)	(0.00240)	(0.00028)*
Panel B: Additional interaction with birth year as $\bar{\beta} = \bar{\beta}_0 + \bar{\beta}_1(\text{birth-year} - 1945) + \bar{\beta}_2 \bar{T}_j$				
$\bar{\beta}_0$	0.04243	0.03515	0.04869	0.04789
	(0.00815)**	(0.00646)**	(0.01024)**	(0.00762)**
$\bar{\beta}_1$	-0.00235	-0.00100	-0.00218	-0.00182
	(0.00103)*	(0.00066)	(0.00114)$^+$	(0.00070)**
$\bar{\beta}_2$	0.00549	-0.00010	-0.00003	-0.00054
	(0.00199)**	(0.00025)	(0.00240)	(0.00028)$^+$
N	591	1182	604	1208

Notes: Robust standard errors in parentheses. *p-values:* $^+$ $p < 0.10$; * $p < 0.05$; ** $p < 0.01$. Standard errors are calculated using White's (1980) heteroscedasticity consistent variance estimator. The OLS standard errors are additionally adjusted for within twin-pair correlation of the error term. (a) Fixed effect regressions do not include a constant term. (b) The OLS models include a third-order polynomial in (birth-year − 1945) in order to account for cohort trends in completed fertility. (c) The overall male and female mean age at first birth is subtraced from $\bar{T}_j$ in the estimation of Eq. (7) in order to make the coefficient β_0 interpretable as the postponement effect for individuals with an average desired age at first birth.

that do not account for potential unobserved characteristics are not able to detect this dependence of the postponement effect on the desired age at first birth.

Panel B in Table 4 additionally includes an interaction of the postponement effect with birth-year in order to account for the trend towards a reduced relevance of the age at first birth for completed fertility. The results of these analyses show the presence of two parallel processes that affect the relation between the age at first birth and fertility. On one hand, the within-MZ twin estimates confirm the earlier-noted trend in younger birth cohorts towards a reduced postponement effect. This time-trend towards an overall smaller postponement effect is approximately equal for males and females. On the other hand, the postponement effect for females—but not for males—seems to increase significantly with the average age at first birth in a twin pair. Women in twin pairs with a higher average AFB are subject to a larger postponement effect. For instance, a difference of two years in the desired onset of fertility increases the postponement effect by about one percentage point, or equivalently, about 25% as compared to the average postponement effect reflected by the coefficient β_0. The completed fertility of women who plan to time their first child relatively late can therefore be substantially more sensitive to variations in the age at first birth than the fertility of women with a younger age at first birth. Additional analyses, not reported in Table 4, did not indicate that this aspect has been subject to any relevant change across time.

In Table 5 we reestimate the above analyses for the cohorts born 1945–55. In these analyses we include all twin pairs where both twins experience a first birth prior to age 35—as compared to age 32 in our previous analyses—and we measure completed fertility at age 43. This restriction to older cohorts and the extension of the age range is relevant for at least three reasons. First, the dependence of the postponement effect may be sensitive on the censoring of our earlier sample at an maximum AFB of age 32. Second, the average AFB within a twin pair is only an unbiased estimator of the desired age at first birth if there is no relevant censoring. Third, in the absence of censoring we can use the special structure of our data to assess the extend to which the measurement error in the desired AFB distorts the estimated postponement effect.

The results obtained from this alternative specification confirm the presence of a relevant dependence of the postponement effect on the desired AFB for females. Moreover, despite the fact that completed fertility is measured at a later age and the limit for the maximum age at first birth has been extended to age 35, the estimates for $\bar{\beta}_2$ indicate again an increase in the postponement effect by approximately .47 percentage points due

Table 5: Dependence of the postponement effect on the average AFB within a twin pair, cohorts 1945–55. Dependent variable is logarithm of fertility at age 43, instead of at age 38 as in Tables 4. Twin pairs are included in the sample if both twins in a pair have experienced their first birth by age 35.

	Females		**Males**	
Method	**Within MZ pair**^{a,c}	**OLS**^{b,c}	**Within MZ pair**^{a,c}	**OLS**^{b,c}
Panel A: Postponement effect depends on desired age at first birth as $\bar{\beta} = \bar{\beta}_0 + \bar{\beta}_2 \bar{T}_j$				
β_0	0.02766	0.02999	0.03922	0.03933
	(0.00575)**	(0.00539)**	(0.00631)**	(0.00608)**
β_2	0.00412	-0.00008	-0.00109	-0.00047
	(0.00181)*	(0.00026)	(0.00203)	(0.00029)
Panel B: Additional interaction with birth year as $\bar{\beta} = \bar{\beta}_0 + \bar{\beta}_1(\text{birth-year} - 1945) + \bar{\beta}_2 \bar{T}_j$				
β_0	0.05131	0.04379	0.05196	0.04805
	(0.00980)**	(0.00690)**	(0.00929)**	(0.00742)**
β_1	-0.00548	-0.00327	-0.00308	-0.00214
	(0.00169)**	(0.00096)**	(0.00172)$^+$	(0.00107)*
β_2	0.00473	0.00002	-0.00054	-0.00039
	(0.00179)**	(0.00025)	(0.00202)	(0.00028)
N	433	866	468	936

Notes: Robust standard errors in parentheses. *p-values:* $^+$ $p < 0.10$; * $p < 0.05$; ** $p < 0.01$. Standard errors are calculated using White's (1980) heteroscedasticity consistent variance estimator. The OLS standard errors are additionally adjusted for within twin-pair correlation of the error term. (a) Fixed effect regressions do not include a constant term. (b) The OLS models include a third-order polynomial in (birth-year − 1945) in order to account for cohort trends in completed fertility. (c) The overall male and female mean age at first birth is subtraced from ($\bar{T}_j$ in the estimation of Eq. (7) in order to make the coefficient β_0 interpretable as the postponement effect for individuals with an average desired age at first birth.

to a one-year increase in the average AFB within a twin pair. Moreover, because the average AFB within a twin pair is a noisy estimate of the desired AFB of twins in a MZ pair, the coefficients in Table 5 are likely to be an underestimate of the extent to which the postponement effect depends on the desired age at first birth. It is therefore likely that the dependence of the postponement effect on the desired AFB is even stronger than revealed by the coefficient $\bar{\beta}_2$ in Table 5 (see also footnote 3).

Interpretation: In summary, the analysis in Tables 4 and 5 suggest important differences in the relevance of the first-birth timing for completed fertility across women with different desired timing of their fertility. In the context of relatively low fertility such an interaction is very plausible: on one hand, completed fertility is less sensitive to the timing of the first child for those women who plan to start childbearing relatively early due to their personal abilities, characteristics and preferences. On the other hand, completed fertility can be considerably more sensitive to variations in the timing of their first child for women who prefer to have a relatively late start of childbearing on the basis of their personal characteristics.

This dependence of the postponement effect on the average AFB in a twin pair is depicted in Figure 2(a), which is based on the within-MZ regression for females in Panel B of Table 4. The full line shows how the postponement effects for females in cohorts born during 1945–48 increases as the average AFB in a twin pair rises. For instance, around age 20 the postponement effect is close to zero and it rises to above 8% as the average AFB approaches age 32. The youngest cohort in our sample born during 1957–60 are subject to a lower postponement effect at any given age as compared to the oldest cohorts in our sample. This effect is due to the secular decline in the postponement effect across cohorts. For instance, women in twin pairs with a very low average AFB are subject to a slightly negative postponement effect in the youngest cohorts, and even twin pairs with very high average AFB are subject to a postponement effect that does not exceed 6%.

The dependence of the postponement effect on the average—or desired—age at first birth is particularly relevant because it can offset and potentially limit the trend towards a reduced relevance of first-birth timing on completed fertility. Younger cohorts increasingly contain women with a relatively late desired onset of fertility, caused by variety of socioeconomic changes that create incentives for a delay of childbearing (e.g., see our discussion in Section 2). The fertility of these women with a late desired onset of fertility is likely to react substantially more sensitive to variations in the age at first birth than the fertility of women with an early desired onset. For instance, Figure 2(b) depicts the distribution of

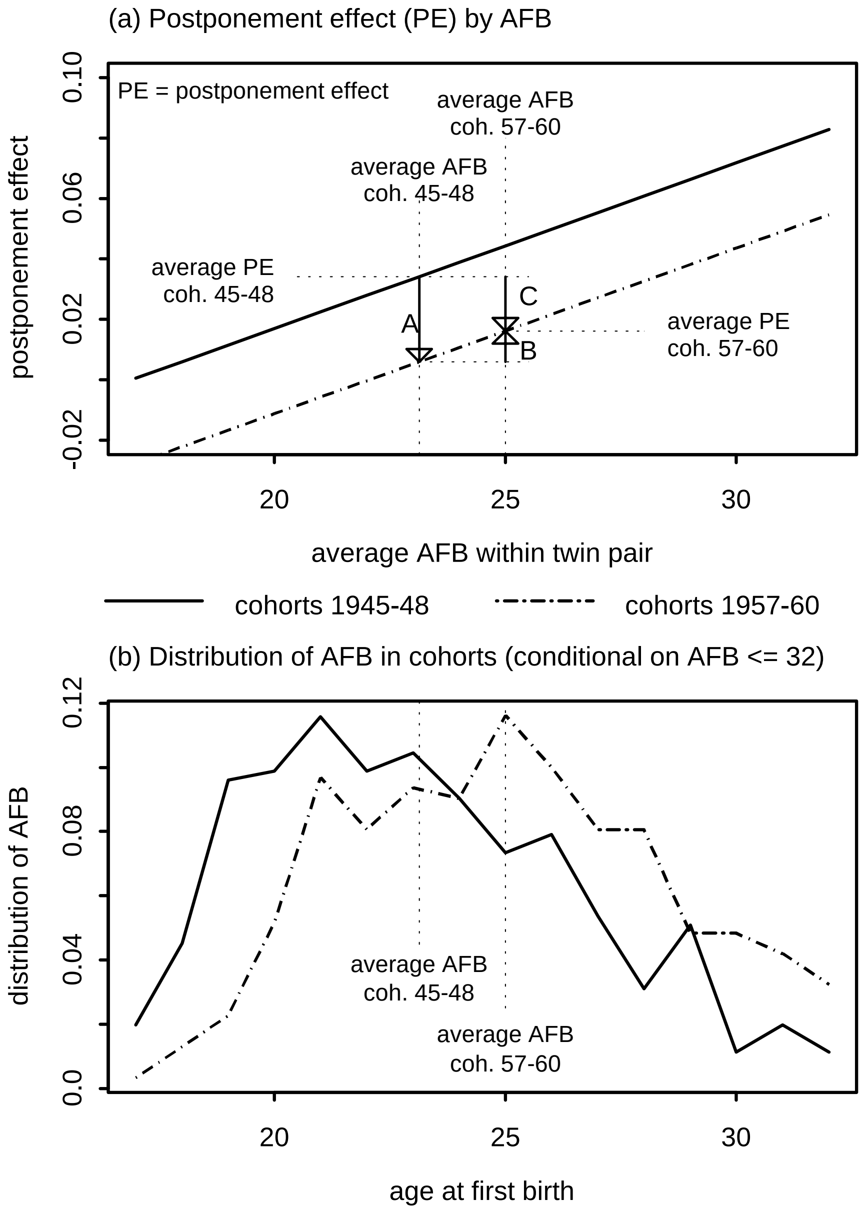

Figure 2: Postponement effect and AFB in the oldest and youngest cohorts in our sample. (a) Dependence of the postponement effect on the average AFB within twin pairs for females. (b) Distribution of AFB in the oldest and youngest cohorts included in our analyses (only for twins with AFB ≤ 32).

the AFB in the oldest and youngest cohorts in our sample. The younger cohorts clearly exhibit a later entry into parenthood, and on average the AFB differs by 1.8 years between these two cohorts.

The overall postponement effect in these two cohorts is indicated in Figure 2(a) and equals 3.4% (cohorts 1945–48) and 1.6% (cohorts 1957–60). This overall decline of the postponement effect is marked by the arrow **C**. Across the cohorts 1945–60, therefore, the general time trend towards a reduced postponement effect dominates the effect caused by the dependence of the postponement on the desired AFB and the increasing delay of childbearing across cohorts. This decline in the postponement effect can be decomposed into two parts: On one hand, arrow **A** shows the decline at a constant average AFB that is due to various secular trends that affect the costs and returns of a fertility postponement. Arrow **B**, on the other hand, indicates the increase in the postponement effect that is due to fact that the overall entry into parenthood occurs at a later age in younger cohorts. The increasing delay of childbearing in younger cohorts therefore decreases the impact of the secular time trend towards a reduced postponement effect for females. Depending on the average age at first birth in more recent cohorts, this may reduce or even limit further declines in the postponement effect for females.

The dependence of the postponement effect occurs only for females, and it is absent for males. The postponement effect across all ages would therefore be indicated by horizontal lines in Figure 2(a), and the secular trend towards a reduced relevance of the postponement effect is not offset by the overall increase in the mean age at first birth in younger cohorts.

Since all twins in our sample experienced a first birth until age 32, this male-female difference in the postponement effect is *not* due to the fact that women in our sample face relevant declines in fecundity with age or reach the 'limits' of fertility postponement (e.g., see Menken 1985). The male-female differences are more likely to be related to important differences in the socioeconomic determinants of the timing and level of fertility. While the detailed analyses of these determinants is beyond the scope of the paper, we can nevertheless consider the potential mechanisms causing this pattern in terms of our theoretical model. The postponement effect in our model is due to an increase in the relative costs of children associated with a later timing of childbearing. The model therefore suggests that these relative costs of children are more sensitive with respect to the timing of fertility for 'high-ability women' with high economic returns to a delay of childbearing.

6 Conclusions

The relation between the age at first birth and completed fertility is a central aspect for understanding, modelling and predicting fertility behavior because it relates the timing of entering parenthood to the completed level of fertility. The *postponement effect* estimated in this paper measures the relative decrease in completed fertility caused by an additional year of delay in the onset of childbearing. Standard estimates of this postponement effect, however, are potentially distorted by the presence of unobserved preference, ability and fecundity characteristics of individuals that affect both the timing of the first child and the level and pace of subsequent childbearing. Moreover, neither the magnitude nor the direction of the distortion can be specified from theoretical reasoning.

In this paper we use within-MZ twin estimations based on all identical twins with verified zygosity born in Denmark during 1945–60 in order to overcome the problems caused by unobserved characteristics. In a theoretical model we establish the effect of these characteristics on the timing and level of fertility, and we show that unobserved variations in preferences for children tend to bias the conventional OLS estimates of the postponement effect upward as compared to the consistent within-MZ twin estimates. Variations in characteristics that determine the economic costs and returns of a delay in childbearing cause a downward bias in OLS analyses. This pattern derived from our theoretical model therefore implies that the distortions of OLS results are informative about unobserved factors that cause variations in the timing of the first birth and the level and pace of subsequent childbearing.

Our analyses confirm the existence of a relevant postponement effect for both males and females. On average, an additional year of delay in childbearing reduces completed fertility by 3% for females and 3.4% for males. Similar to a recent US study by Morgan and Rindfuss (1999), our analyses reveal a clear trend towards a reduced relevance of this postponement effect in younger cohorts emerges for both males and females. The failure to account for unobserved factors like child-preferences for children and economic ability in the estimation can substantially distort these estimates. For instance, OLS underestimates the relevance of first-birth timing for completed fertility substantially for cohorts born around 1945 by about 20-25%, and it also underestimate the pace of change in this postponement effect by up to 50%. Moreover, the direction of the OLS distortions reverses for females over time. This reversal reveals that variations in unobserved ability- or cost characteristics are more important for variations in the timing and level of fertility in older female cohorts, while variation in the timing and level of fertility seem to be more related to unobserved differences in the preferences for children in younger cohorts. For males a similar but less pronounced shift is present.

In addition, our analyses show a dependence of the postponement effect on the desired age at first birth for females. In particular, the fertility of women with a late desired onset of fertility is substantially more sensitive to variations in the age at first birth than the fertility of women with a relatively early desired age at first childbirth. This effect operates in parallel to the general trend towards a reduced postponement effect. Socioeconomic developments that lead to an increasing delay of childbirth, therefore, partially compensate the reduction of the postponement effect across cohorts. Depending on the future increases in the mean age at first birth, this effect can substantially reduce a further weakening of the postponement effect.

Acknowledgments: We grateful to Jere Behrman, Francesco Billari, Jan Hoem, Karsten Hank, Michaela Kreyenfeld and participants at various seminars and conferences for many helpful comments and suggestions. We are also gratefully acknowledge the financial support provided by the Max Planck Institute for Demographic Research in Rostock, the Danish Interdisciplinary Research Council and the Helsefonden grant (#11/2 12-97). The activities of the Danish Center for Demographic Research are funded by a grant from the Danish National Research Foundation.

A Appendix

A.1 Fertility Decisions in the Presence of Household Bargaining

The theoretical and empirical model in Sections 2 and 3 does not explicitly consider the implications of joint household decisions about the timing and level of fertility. In this Appendix we therefore augment our model to reflect joint decision processes and assortative mating with respect to child preferences and labor-market ability.

In this appendix the superscript M denotes variables pertaining to a twin in a monozygotic twin pair, while the superscript S denotes variables pertaining to his/her spouse. We assume that the desired age at first birth of the household, T_{ij}^{d*}, is the weighed average of the desired age at first birth T_{ij}^{M} of twin i in pair j, which is specified in Eq. (2), and the desired age at first birth of his/her spouse T_{ij}^S; that is, we assume that

$$T_{ij}^{d*} = \psi T_{ij}^{d*M} + (1-\psi) T_{ij}^{d*S}, \tag{8}$$

where ψ determines the relative influence of the twin and his/her spouse on the timing of first child. Similarly, we assume that the number of children conditional on the age at first birth is given by

$$\log n_{ij|T}^{d*} = \varphi \log n_{ij|T}^{d*M} + (1-\varphi) \log n_{ij|T}^{d*S}, \tag{9}$$

where $n_{ij|T}^{d*M}$, $n_{ij|T}^{d*S}$ and φ are respectively the desired fertility levels of twin i in pair j conditional on the AFB (see Eq. 1), the desired fertility of his/her spouse, and the relative influences of the twin on the joint household decision.

In addition to joint decision-making about fertility, we also assume assortative mating with respect to preferences for children and ability characteristics. In particular, we assume that the spouse's desired age at first birth and completed fertility are given by a transformation of the corresponding desires and underlying characteristics of twin i in pair j plus additional random influences as

$$\begin{aligned} T_{ij}^{d*S} &= T_{ij}^{d*M} + \tau_1 \lambda_{ij}^M + \tau_2 \pi_{ij}^M + \tau_3 \delta_{ij}^M + \tau_4 \log \phi_{ij}^M + \xi_{ij}^S & (10) \\ \log n^{*S}(T_{ij}) &= \log n_{ij|T}^{d*M} + \rho_1 \lambda_{ij}^M + \rho_2 \pi_{ij}^M + \rho_3 \delta_{ij}^M + \rho_4 \log \phi_{ij}^M + \varsigma_{ij}^S, & (11) \end{aligned}$$

where ξ_{ij}^S and ς_{ij}^S are random influences that are independent of λ_{ij}^M, π_{ij}^M, δ_{ij}^M and ϕ_{ij}^M. Substituting Eqs. (10–11) into Eqs. (8–9) yields the desired timing and level of fertility in households as a function of the preference and ability characteristics of twin i in pair j and additional influences resulting from the spouse. The derivation of the consistent within-MZ twin estimator in Section 3 continues to hold if the timing and level of fertility is determined in joint household decisions, given the above assumptions about assortative with respect to preference and ability characteristics. In particular, differencing within MZ pairs removes the unobserved characteristics that affect both the timing of the first child and the level of fertility conditional on the age at first birth and the within-MZ estimates yield consistent estimates of the postponement effect even when fertility results from joint household decisions.

A.2 Investigation of common and individual-specific influences on the AFB in MZ twin pairs

A central element of our within-MZ twin pair analysis of the postponement effect is the existence of shared characteristics that determine the desired AFB within a twin pair. In particular, our assumptions imply that twins in MZ twin pairs share an identical desired AFB, and differences in the observed AFB within a twin pair represent 'true' random variation. We can provide support for this assumption in our analyses by investigating the between- and within-twin pair correlations of observable characteristics that are likely to be correlated with the age at first birth (the below approach closely follows the analyses in Ashenfelter and Rouse 1998).

We have specified in Eq. (4) that the observed AFB of twin i in pair j consists of random variation around the expected AFB, and the expected AFB is shared by both twins in twin pair. In particular, the observed AFB is given as $\check{T}_{ij} = T_j^{e*} + \zeta_{ij}$, where ζ_{ij} are random influences on the AFB and $T_j^{e*} = T_j^{d*} + \mathrm{E}[\nu_{ij}|\nu_{ij} \leq T^{\max} - T_j^{d*}]$ is the expected (or in the absence of censoring, the desired) AFB in pair j.

An alternative specification of the above relation between the expected and observed AFB, which violates the assumptions of our within-MZ estimation, includes additional individual-specific influences, denoted as $\tilde{T}_{ij}^{e*}$, that are specific to twin i in pair j. In this alternative specification, the observed AFB of a twin is determined by

$$\check{T}_{ij} = T_j^{e*} + \alpha_1 \tilde{T}_{ij}^{e*} + \zeta_{ij}, \tag{12}$$

where the parameter α_1 measures the relevance of the individual-specific influences on the observed AFB. Since the expectation of these individual-specific influences is zero, i.e., $E\tilde{T}_{ij}^{e*} = 0$, the average AFB in a twin pair is $\bar{T}_j = T_j^{e*} + \bar{\zeta}_j$. Now suppose that other observable characteristics X, which vary within and across twin pairs, are correlated with the desired or expected AFB so that

$$\begin{aligned} X_{ij} &= \alpha_2 T_j^{e*} + \alpha_3 \tilde{T}_{ij}^{e*} + \xi_{ij}, \text{ and} \\ \bar{X}_j &= \alpha_2 T_j^{e*} + \bar{\xi}_j. \end{aligned} \tag{13}$$

The correlation across twin pairs between the average AFB and other characteristics is therefore $\mathrm{corr}(\bar{T}_j, \bar{X}_j) = \alpha_2 \,\mathrm{var}(T_j^{e*})/[(\mathrm{var}(T_j^{e*}) + \mathrm{var}(\bar{\zeta}_j)) \cdot \mathrm{var}(\bar{X}_j)]^{1/2}$. This correlation is significant only if (a) the variance of the expected AFB in twin pair j is large relative to the variance of the random influences $\bar{\zeta}_j$, and (b) the expected AFB is systematically related to the other characteristics $\bar{X}_j$, i.e., $\alpha_2 \neq 0$. Differencing Eqs. (12) and (13) shows that the within-twin pair correlation between the AFB and other characteristics is $\mathrm{corr}(\Delta\bar{T}_j, \Delta\bar{X}_j) = \alpha_1\alpha_3 \, \mathrm{var}(\Delta\tilde{T}_j^{e*})/[(\mathrm{var}(\Delta\tilde{T}_j^{e*}) + \mathrm{var}(\Delta\zeta_j)) \cdot \mathrm{var}(\Delta\bar{X}_j)]^{1/2}$.

If the other characteristics X are systematically related to the expected or desired AFB of a twin, then it is reasonable to assume that they are related to both the twin-pair specific component (so that $\alpha_2 \neq 0$), and the individual specific component (so that $\alpha_3 \neq 0$). We can therefore test for the presence of within-twin pair differences in their expected AFB by comparing across-twin pair correlations of the AFB and other characteristics with the corresponding within-twin pair correlations. If the former are significant for some characteristics, we can conclude that these characteristics are systematically related to the unobserved determinants of the expected AFB. If additionally the within-twin pair correlations are significant, the results provide evidence that within-twin pair differences in the expected AFB exist. In this case, our within-MZ regressions may therefore be biased due to individual-specific components that are not controlled for by our within-MZ analyses. On the other hand, if the within-twin pair correlations are negligible and substantially reduced as compared to the across-twin pair correlations, the results support our assumptions that unobserved differences in individual-specific determinants of the expected AFB are absent and do not provide a potential source of bias in our within-MZ analyses.

Unfortunately, the analyses of these paper are based on pure register-based data and do not include any additional information about the twins except the timing and level of fertility. The above test, therefore, cannot be implemented for all twins pairs that underlie our analyses. A subset of these twins born during 1953–60, however, has participated in a survey conducted in 1994 on health related issues that provides information on several personal characteristics that are potentially related to the desired AFB (Christensen et al. 1998). This survey provides personal characteristics for twins in 404 female and 322 male MZ twin pairs born during 1953–60 who are also included in our analyses of the postponement effect. The characteristics include married or cohabiting at time of survey (MARR), years of primary and secondary education (PS-ED), total years of schooling (TOT-ED), body weight (WGHT; we use pre-pregnancy weight if a woman is pregnant at time of the survey), the extent of smoking (SMOKE; variable has a value of zero for not smoking, one for smoking casually, and two for smoking daily), and the frequency of contact between twins (CONT; variable ranges from one for rare contacts to five for daily contacts). All variables are first regressed on a polynomial in age in order to remove systematic age patterns, and the correlation analyses are preformed with the residuals.

The correlations matrices for the average twin-pair characteristics and the within-twin pair differences are shown in Table A3 separately for males and females. The former correlations reveal the existence of shared characteristics that affect both the AFB and related characteristics, and the latter provide an indirect test for the presence of individual-specific differences in the AFB within twin pairs. In the top panel for females and males, many characteristics are significantly correlated with the average AFB, and with each other, indicating that the correlation between unobserved determinants in the AFB and related personal characteristics is strong. On the other hand, the correlation of within-twin pair differences in the AFB and other characteristics are not significant and of very small magnitude. For females, the only exception is the persistent correlation between the years of primary/secondary education and the total years of education, which is likely to be caused by the fact that more years of primary/secondary education is a prerequisite for higher education. For males, additional modest correlations between within-twin pair differences in the AFB and the years of schooling remain significant, but these correlations are substantially reduced as compared to the correlations observed for the average twin pair characteristics. Moreover, the emerging correlation between ΔSMOKE and ΔWGHT is likely to be due to the causal effect of smoking on body weight and not to aspects related to the desired AFB.

Especially for females, Table A3 therefore provides strong evidence that within-twin pair differences in the desired AFB are small and negligible compared to the across twin pair variability. Particularly compelling is the strong correlation between the level of education and the AFB in Panel A of Table A3, which indicates that systematic ability differences between twin pairs lead to a higher investment in human capital as well as an later onset of childbearing. This strong positive correlation vanishes in Panel B after the systematic twin-pair components of ability are removed, which indicates that individual-specific ability differences that affect the AFB are likely to be absent.

For females, Table A3 therefore provides strong support for our argument that the within-MZ twin regressions implemented in this paper provide consistent estimates of the postponement effect that are preferable to the biased results obtained via standard OLS regressions.

Although some within-MZ twin pair correlations remain statistically significant for males, the correlations for all characteristics are substantially reduced and its significance is weakened. The effect of individual-specific influences on the desired AFB is therefore likely to be very small in the within-MZ analyses, and the respective results are likely to be considerably more accurate than standard OLS estimates of the postponement effect.

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Table A 1: Postponement effect $\bar{\beta}$ for Danish twins born 1945–60. Dependent variable is number of children at age 38 instead of logarithm of fertility as age 38 as in Tables 3.

	Females		Males		Females and males	
Method	**Within MZ pair**[a]	**OLS**[b]	**Within MZ pair**[a]	**OLS**[b]	**Within MZ pair**[a]	**OLS**[b]
Panel A: Constant postponement effect for all cohorts						
$\bar{\beta}$	0.0571	0.0549	0.0782	0.0622	0.0673	0.0580
	(0.0097)**	(0.0065)**	(0.0125)**	(0.0078)**	(0.0078)**	(0.0051)**
		[-3.9%]		[-20.5%]		[-13.8%]
Panel B: Interaction with birth year as $\bar{\beta} = \bar{\beta}_0 + \bar{\beta}_1 \cdot$ (birth-year$-$1945)						
$\bar{\beta}_0$	0.0840	0.0696	0.1048	0.0901	0.0941	0.0762
	(0.0160)**	(0.0107)**	(0.0222)**	(0.0139)**	(0.0136)**	(0.0085)**
		[-17.1%]		[-14.0%]		[-19.0%]
$\bar{\beta}_1$	-0.0039	-0.0021	-0.0039	-0.0041	-0.0039	-0.0026
	(0.0020)*	(0.0013)	(0.0025)	(0.0016)*	(0.0016)*	(0.0010)**
		[+46.2%]		[-5.1%]		[+33.3%]
Panel C: Cohort-specific estimation						
$\bar{\beta}$ for cohorts	0.0734	0.0646	0.1039	0.0813	0.0876	0.0713
1945–52	(0.0125)**	(0.0086)**	(0.0170)**	(0.0110)**	(0.0104)**	(0.0068)**
		[-12.0%]		[-21.8%]		[-18.6%]
$\bar{\beta}$ for cohorts	0.0344	0.0404	0.0484	0.0358	0.0416	0.0392
1953–60	(0.0149)*	(0.0098)**	(0.0177)**	(0.0107)**	(0.0116)**	(0.0071)**
		[+17.4%]		[-26.0%]		[-5.8%]
χ^2 test for	4.00*	3.99*	5.10*	8.78**	8.74**	10.85**
equal $\bar{\beta}$	(df = 1)	(df = 1)	(df = 1)	(df = 1)	(df = 1)	(df = 1)
N	591	1182	604	1208	1195	2390

Notes: Robust standard errors in parentheses. *p-values:* $^+$ $p < 0.10$; * $p < 0.05$; ** $p < 0.01$. Standard errors are calculated using White's (1980) heteroscedasticity consistent variance estimator. The OLS standard errors are additionally adjusted for within twin-pair correlation of the error term. (a) The fixed effect regressions do not include a constant term. (b) The OLS models in Panel A and B include a third-order polynomial in (birth-year – 1945) in order to account for cohort trends in completed fertility, and OLS models Panel C include a dummy for birth cohorts 1953–60. The combined OLS models for males and females additionally include a dummy for females.

Table A 2: Postponement effect $\bar{\beta}$ for Danish twins born 1945–55. Dependent variable is logarithm of fertility at age 43, instead of at age 38 as in Tables 3. Twin pairs are included in the sample if both twins in a pair have experienced their first birth by age 35. The percentages in square parentheses below the standard errors report the relative differences between the OLS and within MZ pair estimates.

	Females		**Males**		**Females and males**	
Method	**Within MZ pair**[a]	**OLS**[b]	**Within MZ pair**[a]	**OLS**[b]	**Within MZ pair**[a]	**OLS**[b]
Panel A: Constant postponement effect for all cohorts						
$\bar{\beta}$	0.0319	0.0286	0.0385	0.0313	0.0355	0.0300
	(0.0053)**	(0.0033)**	(0.0059)**	(0.0037)**	(0.0040)**	(0.0025)**
		[-10.3%]		[-18.7%]		[-15.5%]
Panel B: Interaction with birth year as $\bar{\beta} = \bar{\beta}_0 + \bar{\beta}_1 \cdot (\text{birth-year}-1945)$						
$\bar{\beta}_0$	0.0541	0.0440	0.0519	0.0422	0.0525	0.0412
	(0.0098)**	(0.0056)**	(0.0092)**	(0.0061)**	(0.0068)**	(0.0042)**
		[-18.7%]		[-18.7%]		[-21.5%]
$\bar{\beta}_1$	-0.0050	-0.0033	-0.0031	-0.0023	-0.0039	-0.0024
	(0.0017)**	(0.0010)**	(0.0017)$^+$	(0.0011)*	(0.0012)**	(0.0007)**
		[+34.0%]		[+25.8%]		[+38.5%]
Panel C: Cohort-specific estimation						
$\bar{\beta}$ for cohorts	0.0413	0.0368	0.0466	0.0377	0.0441	0.0366
1945–52	(0.0065)**	(0.0041)**	(0.0078)**	(0.0048)**	(0.0052)**	(0.0032)**
		[-10.9%]		[-19.1%]		[-17.0%]
$\bar{\beta}$ for cohorts	0.0141	0.0154	0.0258	0.0219	0.0208	0.0195
1953–60	(0.0083)$^+$	(0.0053)**	(0.0087)**	(0.0057)**	(0.0061)**	(0.0039)**
		[+9.2%]		[-15.1%]		[-6.3%]
χ^2 test for	6.64**	10.1**	3.18$^+$	4.52*	8.49**	11.8**
equal $\bar{\beta}$	(df = 1)	(df = 1)	(df = 1)	(df = 1)	(df = 1)	(df = 1)
N 433	866	468	936	901	1802	

Notes: Robust standard errors in parentheses. *p-values:* $^+$ $p < 0.10$; * $p < 0.05$; ** $p < 0.01$. Standard errors are calculated using White's (1980) heteroscedasticity consistent variance estimator. The OLS standard errors are additionally adjusted for within twin-pair correlation of the error term. (a) The fixed effect regressions do not include a constant term. (b) The OLS models in Panel A and B include a third-order polynomial in (birth-year – 1945) in order to account for cohort trends in completed fertility, and OLS models Panel C include a dummy for birth cohorts 1953–60. The combined OLS models for males and females additionally include a dummy for females.

Table A 3: Correlation analysis for the investigation of common and individual-specific influences on the AFB and other personal characteristics in MZ twin pairs

Females (N = 404 twin pairs)						
Panel A: *Correlation matrix of average twin pair characteristics*						
	AFB	MARR	PS-ED	TOT-ED	WGHT	SMOKE
AFB	1					
MARR	0.041	1				
PS-ED	0.426**	0.068	1			
TOT-ED	0.421**	0.029	0.884**	1		
WGHT	-0.163*	0.115	-0.015	-0.031	1	
SMOKE	-0.023	-0.239**	-0.136^{+}	-0.155*	-0.101	1
CONTACT	0.048	-0.056	-0.018	-0.082	-0.035	0.170*
Panel B: *Correlation matrix of within-twin pair differences in characteristics*						
	Δ AFB	Δ MARR	Δ PS-ED	Δ TOT-ED	Δ WGHT	Δ SMOKE
Δ AFB	1					
Δ MARR	-0.067	1				
Δ PS-ED	0.037	-0.071	1			
Δ TOT-ED	0.099	0.001	0.769**	1		
Δ WGHT	-0.035	-0.090	0.015	0.022	1	
Δ SMOKE	-0.083	-0.046	-0.071	-0.061	-0.010	1
Δ CONT	-0.005	0.059	-0.117	-0.013	0.102	0.086
Males (N = 322 twin pairs)						
Panel C: *Correlation matrix of average twin pair characteristics*						
	AFB	MARR	PS-ED	TOT-ED	WGHT	SMOKE
AFB	1					
MARR	0.051	1				
PS-ED	0.309**	0.060	1			
TOT-ED	0.359**	0.007	0.883**	1		
WGHT	-0.179*	-0.025	0.059	0.015	1	
SMOKE	-0.274**	-0.133^{+}	-0.312**	-0.271**	-0.122^{+}	1
CONT	-0.133^{+}	-0.083	-0.250**	-0.284**	0.049	0.022
Panel D: *Correlation matrix of within-twin pair differences in characteristics*						
	Δ AFB	Δ MARR	Δ PS-ED	Δ TOT-ED	Δ WGHT	Δ SMOKE
Δ AFB	1					
Δ MARR	0.017	1				
Δ PS-ED	0.153*	-0.096	1			
Δ TOT-ED	0.169*	-0.117	0.745**	1		
Δ WGHT	-0.047	0.122^{+}	-0.019	-0.070	1	
Δ SMOKE	0.025	-0.090	-0.064	0.015	-0.265**	1
Δ CONT	0.083	0.012	-0.031	-0.005	0.011	-0.037

Notes: $^{+}$ $p < 0.10$ * $p < 0.05$. ** $p < 0.01$