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Subnational Birth Squeezes

Male-Female TFR Differences Across Eight High- and Middle-Income Countries over Time

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Abstract

Men and women can have different levels of fertility indicators; for instance, the total fertility rate of women can be higher than the total fertility rate of men. A large difference between the fertility of men and women is called a birth squeeze. Evidence on the occurrence of birth squeezes has so far been limited. In this study, we conduct a subnational analysis of differences in the total fertility rate between men and women, utilizing high-quality administrative registers from the United States, Mexico, Australia, Colombia, Finland, France, Germany, Mexico, and Spain (n = 5,278), and we suggest several approaches to measuring birth squeezes. Our results show that birth squeezes regularly occur at the subnational level, with the notable exception of France. We discuss how birth squeezes are related to population structures and age-patterns of fertility, which are, in turn, influenced by development progress and urban-rural settlement structures. Overall, our findings demonstrate the importance of considering both male and female fertility indicators to gain a more nuanced understanding of population dynamics. (170 words)

1 Introduction

In high-income countries, the quantum of fertility at the national level is often very similar for women and men. This insight comes from a recent strand of research built on methodological innovations and new data sources (e.g., Dudel & Klüsener, 2019a, 2021; Schoumaker, 2017). Previously, fertility research was predominantly focused on women due to the lack or limitation of data on male fertility. This new research shows that differences in indicators of the quantum of fertility between men and women, such as differences in the total fertility rate (TFR), are often modest with a few exceptions (Bratsberg et al., 2021; Dudel & Klüsener, 2016, 2019a, 2021; Schoumaker, 2019; Zhang, 2011). These gender differences in the level of fertility can be explained by differences in the population sizes of men and women, gender differences in the timing of fertility, and the interaction of fertility timing and differences in the cohort sizes of men and women (Dudel & Klüsener, 2021; Schoumaker, 2019). Theoretically, gender differences in fertility quantum could be much more pronounced given specific conditions; however, at the national level, it seems that the factors impacting fertility differentials cancel each other out to some extent (Canudas-Romo et al., 2023).

Less is known about gender differences in the quantum of fertility at the subnational level. It can be speculated, however, that differentials could be larger at the subnational level. This is because of gender-selective subnational migration (Gulczynski, 2023), subnational variation in gender differences in mortality (Sauerberg et al., 2023), and subnational variation in the causes of gender differences in fertility behavior (Campisi et al., 2020; Fox & Myrskylä, 2015). A recent study showed considerable heterogeneity in subnational adult sex ratios worldwide, which tend to be more skewed than those at the national level (Gulczynski, 2023). The few studies looking at fertility differentials at the subnational level have confirmed this expectation. For instance, Dudel and Klüsener (2016, 2019a) found that the difference in the TFR between men and women is bigger in Eastern Germany than in Western Germany, likely due to gender-differential internal migration.

This paper examines subnational differences in the quantum of fertility, as reflected in the TFR, with a particular emphasis on potential birth squeezes (Schoen, 1985). A birth squeeze is a situation in which the fertility of one gender is considerably lower than of the other gender. We discuss which indicators and thresholds to use for assessing whether there is a birth squeeze. We harmonize and analyze high-quality register data from Australia, Colombia, Finland, France, Germany, Mexico, Spain and the United States, covering a wide range of regions and institutional settings. We find birth squeezes at the subnational level in several of these countries. For Finland, we are able to conduct additional analyses by parity and for childlessness, which shed some light on demographic explanations for the observed patterns.

The term "*birth squeeze*" was introduced by Schoen (1985) in analogy to the concept of a "marriage squeeze" (e.g., Akers, 1967; Muhsam, 1974), although the basic idea was recognized earlier (Brouard, 1977). Essentially, a marriage squeeze occurs if there is an imbalance in the numbers of women and men leading to differences in marriage

rates between women and men (Schoen, 1983). Thus, a birth squeeze is a difference in fertility rates between women and men caused by an imbalance in adult sex ratios (Schoen, 1985). In contrast to the marriage squeeze, the birth squeeze concept has received very little attention in the literature, likely due to the above-mentioned fact that differentials in the quantum of fertility between women and men are mostly small at the national level. Moreover, there is little guidance on what level of difference can be really considered a birth squeeze.

Birth squeezes, as well as skewed adult sex ratios, have been hypothesized to have a significant impact on various domains, including social, family, and economic outcomes (for a review, see Dyson, 2012). The primary concern is the effect on fertility and childlessness, as cohorts exposed to birth squeezes face a structural constraint to childbearing, potentially leading to increased childlessness among men and women (Klein, 2003; Kravdal, 2021; Schubert et al., 2024). Birth squeezes may also indicate imbalanced mating markets in which the more abundant gender faces structural constraints on partnership formation, and may affect union compositions in terms of age gaps between partners and bargaining power (Abramitzky, 2009; Akers, 1967; Albrecht, 2001; Filser & Willführ, 2022; Muhsam, 1974). Furthermore, birth squeezes may have downstream implications for social and health outcomes, including increased violence, the spread of sexually transmitted diseases, and economic vulnerability, particularly among unpartnered and childless individuals, who may suffer from loneliness and have fewer kin to care for them at older ages (Angrist, 2002; Angrist, 2000; Edlund, 2005; Gupta, 2010; Hesketh, 2009; Hesketh & Xing, 2006; Tucker et al., 2005).

The selection of countries in this study represents a variety of contexts and exhibits substantial subnational heterogeneity. For instance, Colombia and Mexico are middle-income countries, while Germany and the United States are among the wealthiest nations. Furthermore, the countries reflect different world regions, as evidenced by the inclusion of the continents of Australia, South America, North America, and Europe. The countries are at very different stages of the demographic transition, which is reflected in the broad range of fertility levels. Spain has consistently reported lowest low fertility for several decades, while Mexico just recently experienced fertility levels below the replacement level. Finally, the selected countries exhibit considerable subnational heterogeneity with regard to historical, economic, and developmental characteristics. In the United States, for instance, Alaska is noteworthy as a highly industrial and population-sparse state, in sharp contrast to the state of New York. In Mexico and Colombia, certain regions have been impacted by internal armed conflicts. Germany has historically been divided into Western and Eastern Germany. Additionally, the countries under investigation provide high-quality micro-level birth registration data.

Overall, this paper makes several contributions to the literature. At the conceptual and methodological level, we provide a detailed discussion of the birth squeeze concept and how it can be measured. In particular, we provide several thresholds that allow us to assess what level of gender difference in the TFR can be considered high. In our empirical analysis, we provide comparative evidence of the relevance of the birth squeeze, and of fertility differentials more broadly, at the subnational level, covering not only high-income countries, but also middle-income countries. While this paper focuses on the subnational

level, the indicators and thresholds we discuss are generally applicable at the national level. Finally, we present a new high-quality data source for the study of fertility differentials at the subnational level that draws on vital statistics and register data from eight countries, some of which have received little attention in the previous literature.

The remainder of this paper is structured as follows. In the next section, we discuss male-female differentials in the quantum of fertility at the national and the subnational level and their causes. In the third section, we provide a detailed discussion of the birth squeeze, indicators to measure male-female differentials, and thresholds to decide whether a value of an indicator shows a birth squeeze. The data we use for our analysis is described in Section 4. Results are presented in Section 5 and discussed in Section 6.

2 Gender differences in the quantum of fertility

2.1 Differences at the national level: reasons and results

The average number of children born to women can be quite different from the average number of children born to men (Dudel & Klüsener, 2021; Schoumaker, 2019), which has been known for some time (Karmel, 1947). These differences can be explained by four main demographic factors: first, differences in the size of the male and female populations; second, gender differences in the timing of fertility; third, the interaction of differences in the timing of fertility and cohort sizes (Dudel & Klüsener, 2021; Schoumaker, 2019); and fourth, gender differences in data quality, which may underlie and exacerbate gender differences in fertility. We explain these factors below and illustrate them using Alaska and Oaxaca as examples in Figure 1.

Figure 1 about here

First, differences in the number of women and men impact fertility differentials via exposures. In a very simple scenario, in a population of 100 women and 80 men, with 10 births, the average number of children per woman will be 0.10, while the average number of children per man will be 0.125. For example, Panel A) in Figure 1 shows that there are more men than women in every age group in Alaska. Differences in the population size can be caused by three factors: the sex ratio at birth (SRB), mortality differentials, and differences in migration. The SRB is influenced by biological and behavioral aspects, and averages around 105 male births per 100 female births (Chao et al., 2019; Hesketh, 2009). Mortality differentials exhibit a consistent pattern of advantage for women, with men usually displaying higher mortality rates across all age groups (Luy & Gast, 2014; Oksuzyan et al., 2008; Remund et al., 2018; Sauerberg et al., 2023). Thus, the impact of mortality differentials on fertility differentials will usually counterbalance the impact of the SRB, at least to some extent. Finally, differentials in migration have a less clear impact. Historically, men have marginally dominated in numbers in international migration (Castles, 2003; Donato & Gabaccia, 2016), likely due to gender norms and economic roles, but international migration among women has been catching up, and has surpassed the migration among men in some places (Donato & Gabaccia, 2016). Moreover, immigration

to one country generally implies emigration from another country; thus, for instance, the inflow of men to one country will mean a reduction in the number of men in one or several other countries.

Second, gender differences in fertility timing also influence male-female fertility differences. On average, men have their children later than women (Dudel et al., 2021; Schoumaker, 2019). In many high-income countries, the father tends to be around 3 to 4 years older than the mother at childbirth (Dudel et al., 2023). Globally, there is considerable variation, and much larger age differences can be found (Schoumaker, 2019). For example, the male fertility schedule is shifted to older ages compared to the female schedule in Panel B) in Figure 1. Age differences can have an impact even if the total population size and the age structure are the same for men and women. To illustrate, consider a population comprising three age groups. In the first age group, there are 10 men and 10 women; in the second age group, there are 20 men and 20 women; and in the third age group, there are 5 men and 5 women. Let us further assume that there are 5 births in this population, all to mothers from the first age group. However, these women have their children exclusively with men from the second age group. The total fertility rate (TFR) will be 0.5 for women and 0.25 for men, using the population counts as exposures.

Third, the interaction between fertility timing and cohort size provides an additional explanation for gender differences in the quantum of fertility. This interaction has been well established in the literature on marriage squeezes (Akers, 1967; Muhsam, 1974; Schoen, 1985). As an alternative to cohort size, this point can also be viewed as an interaction of fertility timing and age structure. As stated above, men tend to have children with younger women. If, for instance, cohort size declines, there will be fewer women relative to men. In a simplified example, assume that in each cohort there is the same number of men and women, but cohort size declines. If men from any cohort c exclusively have children with women from the cohort $c-3$ and vice versa, and the relative decline in cohort size between cohorts c and $c-3$ is a factor d, then the ratio of the TFR of men relative to the TFR of women will be d (for simplicity using cohort size as exposure).

Fourth, data quality is also potentially relevant for gender differentials in fertility. Generally, information on the fertility of men is often of inferior quality and has more issues compared to information on the fertility of women. For instance, even in high-quality birth register data, it is quite common that the age of the father is not known for a large proportion of births (Dudel & Klüsener, 2019b). Such issues can be expected to lead to an underestimation of the quantum of male fertility. It has been demonstrated, for instance, that data quality may contribute to gender differences in childlessness in register data (Kravdal, 2021). Moreover, there is strong evidence that surveys may underestimate male fertility because men are more likely than women to underreport the number of their children when participating in such studies (Joyner et al., 2012).

Empirical evidence shows that the level of male fertility usually follows the level of female fertility very closely (Dudel & Klüsener, 2021). For instance, in the United States in 2015 the male TFR was 1.824 and the TFR of women was 1.845. However, differences may occur. Gender differences in fertility quantum, expressed in the ratio of the male TFR to the female TFR (TFR ratio), range from 0.54 for Bahrain in 2011 to 2.5 for Gambia in 2011 (Schoumaker, 2019). In general, higher values of the TFR ratio are reported for

countries in sub-Saharan Africa, while lower values are reported for high-income countries. This pattern is produced by the stage of the demographic transition, which affects the factors mentioned above (Schoen, 1985; Schoumaker, 2019): the gender gap in mortality is narrowing as a result of improved access to health care and declines in lifestyle-related deaths; the gender gap in fertility timing is narrowing as a result of progress in gender equality and significant postponement of chilbearing among women; and the population growth rates approaching equilibrium. Some African countries diverge from this pattern, as in these countries age differences are stagnating due to the high prevalence of polygamy and the unique fertility transition associated with this practice (Bongaarts, 2017; Garenne & Van De Walle, 1989; Schoumaker, 2019). However, once the demographic transition is completed in a country and it has reached below replacement fertility, the trends are less clear and can fluctuate (Dudel & Klüsener, 2021).

2.2 Differences at the subnational level

Both the level of fertility and the differentials in the quantum of fertility between men and women can be heterogeneous at the subnational level. Geographic heterogeneity in the level of fertility at the subnational level is mainly due to the characteristics of the subnational units (e.g., Campisi et al., 2020; Nisén et al., 2021). For instance, Fox and Myrskylä (2015) demonstrated that subnational variations in income play an important role in fertility. Moreover, rural-urban disparities have been well documented, with fertility rates generally being lower in urban areas and higher in rural areas (Campisi et al., 2020; Kulu & González-Ferrer, 2014). The characteristics of subnational areas not only increase or reduce the individual risk of having children and impact the timing of births through various channels, but they also lead to differences in population size and age structure between men and women due to internal migration. These difference, in turn, affect local partner markets, which have been demonstrated to be highly relevant for fertility outcomes (Chudnovskaya & Ueda, 2021; Kravdal, 2021; Schubert et al., 2024). Finally, general socioeconomic developments such as changes in gender relations might arrive at different times in different subnational units, and might spread spatially over time (e.g., Blanc, 2024; Goldstein & Cassidy, 2014).

There may be subnational variations in the aforementioned factors that lead to gender differences in fertility. First, Gulczynski (2023) found that the sizes of the female and male populations of reproductive ages differ considerably at the subnational level in many countries. Gender-selective subnational migration was found to be the main driver of this gap, but differences in cohort size also play a role. Second, beyond migration, gender differences in mortality also vary considerably at the subnational level, although the gender gap in mortality has narrowed recently in some places (Sauerberg et al., 2023). Third, the age pattern of the fertility differences between men and women is heavily influenced by gender inequalities and economic development (Dudel et al., 2023; Presser, 1975; Schoumaker, 2019). Given the sometimes considerable variation in economic development within countries (Scherbov & Gietel-Basten, 2020), the timing of fertility may also differ between men and women.

3 The birth squeeze

3.1 Concept and previous evidence

The term "*birth squeeze*" was introduced by Schoen (1985) to describe discrepancies in fertility rates between men and women in analogy to the "*marriage squeeze*". A "*marriage squeeze*" occurs if there are differences in the marriage rates of men and of women; i.e., the marriage rates of one gender are lower than those of the other gender. This is usually assumed to be caused by imbalances in the population size, the age structure, and the age preferences of men and of women (e.g., Akers, 1967; Muhsam, 1974). These imbalances can arise from the same factors as those driving the differentials in the quantum of fertility described in the previous section. The assumption is often implicit that the "marriage squeeze" leaves some people who actually want to marry unmarried (Muhsam, 1974); that is, if the imbalances leading to the squeeze were reduced, the differences in rates would also be reduced, and the involuntarily unmarried would marry, at least to some extent.

A birth squeeze is defined as a substantial difference in the fertility of men and women. For instance, a birth squeeze occurs if the TFR of women is substantially higher than the TFR of men, or *vice versa*. This difference is implicitly assumed to have implications for childlessness and birth counts. The present study will test this assumption for birth squeezes and examine the relationship between birth squeezes and childlessness in Finland (see Section 5.4).

There is little empirical research on the occurrence of birth squeezes, likely because birth squeezes at the national level seem to be rare, at least for high-income countries (Dudel & Klüsener, 2021). There are, however, some notable exceptions. Although the term birth squeeze was not used, Brouard (1977) reported substantial differences between male and female fertility in France after World War I, likely due to pronounced differences in war-related mortality. Schoen (1985) was the first to use the term when he found that abruptly changing cohort sizes led to a birth squeeze in England and Wales as well as the United States in the 1970s. Kravdal (2021) attributed the growing sex differences in childlessness in the last 15 years in Norway to fluctuating cohort sizes between the mid-1950s and the mid-1970s, although no reference was made to birth squeezes. Finally, sex-selective migration after reunification caused a birth squeeze in East Germany (Dudel & Klüsener, 2016; Klein, 2003).

Another potential reason why there is little empirical research on birth squeezes is that most of the literature only vaguely defines what level of difference in the quantum of fertility can be considered a birth squeeze, if at all. Generally, it is likely that differences in quantum will always occur, as the mechanisms causing them discussed in the previous section will be present in most cases, and they will not necessarily cancel each other out. This means that some threshold is required to determine whether there is a birth squeeze or just a minor difference in quantum.

In what follows, we will approach and operationalize the birth squeeze in two steps. First, we will discuss indicators that can be used to measure differences in the quantum of fertility between women and men. Second, we will provide a set of thresholds that allow us to determine if a quantum differential can be considered as indicating a birth squeeze. This set of thresholds will be developed from several ideas, including stable population

theory and its relevance for downstream outcomes such as childlessness. While the focus of this paper is on subnational birth squeezes, the indicators and thresholds we discuss are generally applicable at the national level as well.

3.2 Indicators: measuring quantum differentials

We assess fertility quantum through the total fertility rate (TFR). The TFR is a widely used demographic measure that describes the average number of children a person would conceive under the fertility regime of a given year (period), if the person survives the entire reproductive period (Preston et al., 2008). It is calculated as the sum of age-specific fertility rates. Usually, it is preferred over crude birth rates, because it accounts for the age distribution. However, in our discussion below it could easily be replaced with any indicator that provides a measure of the average number of children, be it a period indicator or a cohort indicator.

We quantify gender differences in fertility quantum with the ratio of the TFR of men (TFR_m) relative to the TFR of women (TFR_w) (Dudel & Klüsener, 2019a): $R = \frac{TFR_m}{TFR_w}$. Values higher than 1 indicate that the TFR is higher for men than for women, while values below 1 indicate a higher TFR for women. The measure is comparable to other sex ratios in demography, such as the population sex ratio or the sex ratio at birth, and it can take values from 0 to ∞. The TFR ratio is easy to calculate, to understand, and to interpret, and it is closely related to more elaborate measures; as discussed below. For instance, it can be applied as a scaling factor to the female TFR, enabling the calculation of the male TFR and providing a clear and interpretable measure of the gap between the two. Furthermore, its data demands are relatively modest.

An alternative measure is the birth squeeze index U, as suggested by Schoen (1985). U is defined as $\frac{TFR_m - TFR_w}{0.5 \cdot TFR_m + 0.5 \cdot TFR_w}$; i.e., the difference in TFRs relative to the average TFR of men and women. It differs from the TFR ratio R by a multiplicative factor M and the constant 1. M is given by $M = 0.5 \cdot (R+1)$. Then $U = \frac{(R-1)}{M}$ and $R = U \cdot M + 1$. This means that U and R will differ more if M is further away from 1. Given the approximation, the thresholds we provide below to determine a birth squeeze can be easily translated to apply to U, and we provide corresponding tables in the supplementary materials.

Several other approaches have been used in the literature to assess gender differences in fertility, for instance for childlessness. Kravdal (2021) estimated and decomposed the differences in the level of childlessness between 45-year-old women and men who are two years older. This approach requires data on the parity of the resident population, which are not available for other countries. Klein (2003) took a cohort perspective and found a higher probability of remaining childless among men than among women using German panel data. Therefore, we analyze gender differences in parity-specific fertility in Finland (see Section 5.4), which is, unfortunately, not possible for the other countries as parity-specific exposures are missing.

3.3 Thresholds: When is it a birth squeeze?

Before we describe different approaches to derive thresholds for high or low values of the TFR ratio, we first briefly discuss what value of the TFR ratio can be considered to be "balanced". While a TFR ratio of 1 indicates that the TFRs of women and men are the same, and are in this sense balanced, some deviation from this value is to be expected. For example, assume that both the male and the female population have a rectangular age structure, but their population sizes differ because of the sex ratio at birth. If the sex ratio at birth is 105 boys to 100 girls, then the TFR ratio in this population will be 1/1.05, or around 0.95. Allowing the age structures to differ due to mortality, mortality differentials by gender will push the TFR ratio somewhat toward 1, as men usually experience higher mortality than women (Rogers et al., 2010; Sauerberg et al., 2023). More complexity could be introduced, but these examples show that typical demographic patterns will lead to TFR ratios deviating at least slightly from 1. We will nevertheless use a TFR ratio of 1 as a reference point, as it is straightforward to interpret; however, not all deviations from this reference are of substantive relevance, highlighting the need for thresholds.

Thresholds that specify what level of quantum differential can be considered a birth squeeze can be derived in several ways. We explore four different approaches: first, an expert-based approach, in which the threshold is derived from experts' assessments in the literature; second, a data-based approach, which determines empirical thresholds using outliers in observed TFR ratios; third, a stable population approach, which uses mathematical relationships and certain combinations of input parameters to determine the threshold; and, fourth, an outcome-based approach, which first empirically assesses to what extent TFR ratios are predictive of childlessness using regression modelling, and then uses regression parameters to find thresholds. A summary of the results for the different thresholds is provided in Table 1 at the end of this subsection.

Expert-based approach: First, the expert-based approach obtains thresholds from the existing literature. This assumes that the authors of previous studies had good reasons to choose the thresholds they used, regardless of whether they made the reason explicit. Dudel and Klüsener (2021) took the lowest TFR ratio reported by Schoen (1985) as a lower threshold: i.e., 0.89, which occurred in 1973 in England and Wales. Schoen (1985), in addition, described the TFR ratio of 1.12 in the United States in 1970 as relatively high, at least among high-income countries. Taken together, these observations would suggest that thresholds of 0.89 and 1.12 may be used. Coleman (2000) reported sex ratios between 0.91 (Denmark in 1988) and 1.1 (German birth cohorts from 1902 to 1928). Gulczynski (2023) did not study TFR ratios, but instead examined adult sex ratios, and suggested that deviations from 1 greater than ± 0.085 are rare and indicate gender-skewed populations. This leads to slightly narrower thresholds of 0.915 and 1.085. In what follows, we will use 0.9 and 1.1 as the expert-based thresholds, because these values fall in between the three variants, differing only slightly from each. Moreover, they are equivalent to rounding the TFR ratio thresholds from the literature.

Data-based approach: Second, the data-based approach builds on the observation made by Schoumaker (2019) that TFR ratios tend to be high if the TFR of women is high, and they tend to decline if the TFR of women declines. We utilize this pattern and integrate existing country-level data with a regression approach to identify thresholds for outliers. We

estimate a Bayesian linear regression model of the TFR ratio on female TFR using country-level data from the register-based Human Fertility Collection (Dudel & Klüsener, 2021) and the survey-based as well as the UN-based estimates by Schoumaker (2019), which cover a wide range of low-, middle-, and high-income countries. The thresholds for birth squeezes are obtained from the 90% posterior prediction intervals; that is, TFR ratios that lie outside the range of likely values are considered to indicate a birth squeeze, conditional on the level of fertility of women. Using this approach, we find a lower threshold of $exp(-0.161+0.304\cdot log(TFR_w))-0.12$ and an upper threshold of $exp(-0.161+0.304\cdot log(TFR_w))+0.12$. Technical details are provided in the supplementary materials. This method relies on the assumption that the relationship at the national level holds at the subnational level, or at least is a good benchmark. Plugging in a value of the TFR of 1.5, which is typical for high-income countries, gives bounds of 0.843 to 1.082.

Stable population approach: Third, the stable population approach builds on the stable population approximation of the TFR ratio developed by Schoumaker (2019) and Tucker and Van Hook (2013). In this approximation, the TFR ratio depends on the sex ratio at birth, the survival differential between men and women until the mean age at childbearing, the mean generation length for men and women, and the growth rate of the population. For each of these parameters, we use a set of input values that represent common demographic scenarios. For instance, we use values between 1.04 and 1.06 for the sex ratio at birth (Chao et al., 2019). We calculate the result of the equation for all possible combinations of input values. Then, we take the minimum and the maximum value resulting from this exercise as the lower and the upper threshold, which gives thresholds of 0.78 and 1.15, respectively. Technical details, including information on the input parameters and their derivation, are provided in the supplementary materials. This approach essentially derives what can be considered common TFR ratios based on common demographic scenarios, and it replaces the direct choice of thresholds for the TFR ratio with the choice of input parameters for well-known demographic characteristics.

Outcome-based approach: Fourth, the outcome-based approach identifies thresholds for TFR ratios based on their impact on individual-level outcomes. Here, we use childlessness as an outcome in the comprehensive Finnish register data. Specifically, we assess the impact of the TFR ratio at age 30 on the probability of men being childless at age 40 using a logistic regression model. Based on the coefficient estimates, we can calculate the difference from a TFR ratio of 1 needed to increase or decrease childlessness by 1 percentage point. Applying this difference to get both lower and upper thresholds yields values of 0.86 and 1.13, respectively. More details are again provided in the supplementary materials. Unfortunately, we can do the analysis of the impact on individual-level childlessness only for Finland and not for the other countries, and we have to assume that the results for Finland can be carried over to the other countries.

Table 1 about here

An overview of all thresholds is provided in Table 1. Overall, the upper and lower thresholds are close to each other across approaches, with some minor variation. This indicates that there is a small gray area in which the result will depend on the choice of

threshold. Similar thresholds for the birth squeeze index of Schoen (1985) are provided in the supplementary materials.

4 Data

We calculate TFR ratios using birth counts and population exposure by age and by gender at the subnational level. In our final data, we have 5,278 region-year observations. For birth counts, we use vital registration data from Australia, Colombia, Finland, France, Germany, Mexico, and the United States. The coverage of years differs across countries, as do the size and the population density of the spatial units that make up the subnational level. For Finland, we can estimate fertility indicators for small spatial units (*fin.* Maakunta), while Australia and the United States provide only data at the state level with relatively large units in terms of area and/or population. The information for maternal age was complete for all births in all countries, while the information for paternal age was missing for 1-10% of the births. Missing information was imputed following Dudel and Klüsener (2019a). The population exposures are based on the two-year average of population counts for each age, gender, and regional unit (Wachter, 2014). Table 2 summarizes the data sources used in the study.

Table 2 about here

Generally, the data we use can be considered to provide high-quality information. First, the information is based on birth certificates, which are both mandatory and official documents. Legal enforcement minimizes case and item non-response. Second, birth certificates are issued shortly after birth, which reduces the risk of measurement error from recall bias. It also minimizes the risk of missing information on the father, as it is unlikely that the father has left the country, and the father is less likely to be neglected due to couple separation. The data quality may differ across countries due to different administrative procedures and reporting behaviors. Regardless, the male-female comparisons should be unaffected, because we include the exact same number of births in the male TFR as in the female TFR due to imputation of missing ages for men.

5 Results

5.1 Descriptive results

Figure 2 shows the male TFR and the female TFR for each region-year observation in our data; points above the diagonal line indicate that the TFR of men is higher than the TFR of women, and points below the line indicate a higher TFR of women. For instance, the point in the top-right of the figure is the region Chiapas in Mexico in the year 1990, in which men had a TFR of 9.01 and women had a TFR of 7.91. The TFR of men exceeds the TFR of women in about half of the region-year observations, as the majority of points lie below the diagonal line (50%, 2,633 observations). The geometric mean of the TFR ratios

across N regions is 1.005, indicating a higher TFR for men than for women. A second pattern emerges, as the gender differences are larger at higher levels of TFR and diminish as fertility declines. Larger gaps between the point and the diagonal line appear more toward the right part of the figure, while the points get closer to the line or even cross it at lower fertility levels. For instance, the average TFR ratio (geometric mean) for regions with a high fertility regime, i.e., a female TFR higher than 3, is at 1.164. In a substantial share of regions the TFR among women is higher than the TFR among men, as some points land below the diagonal line, especially those with low fertility. The average TFR ratio in regions with low fertility, i.e., the female TFR is below 2, is 0.975. Figure S.3 in the supplementary materials compares our subnational results with the existing country-level data. We find that the spread of male-female TFR ratios is larger at lower levels of fertility, while data from Schoumaker (2019) show a larger spread at higher fertility levels. While the data from Schoen (1985) and Dudel et al. (2021) and our own data land on a line, the data from Schoumaker (2019) deviate substantially from the line, which may be related to the practice of polygamy in some African countries or data quality issues with male fertility information in surveys (Joyner et al., 2012).

Figure 2 about here

Moreover, male-to-female TFR ratios also depend on the characteristics of the subnational unit. Metropolitan and capital regions exhibit comparatively high male-female TFR ratios for the overall level of fertility. A simple regression analysis indicates that capital regions have a TFR ratio that is 0.04 higher than the TFR ratio in non-capital regions, holding the level of female TFR constant. For instance, theTFR ratio was 1.01 (2009) in the Australian capital Region, 1.06 in Washington D.C (in 2004), 0.95 in Berlin (in 2018), 1.05 in Île-de-France/Paris (in 2013), and 1.01 in Mexico City in 2020. In metropolitan regions, the TFR ratios are also higher relative to the fertility level, such as in Massachusetts, where it was 1 in 2004 or in Hamburg, where it was at 0.99 in 2018. This pattern may be related to particularly delayed childbearing of women and female skewed population structures. The presence of universities and strong service sectors may lead to fertility postponement among women (Blossfeld, 2009), and may also specifically attracts women, because women are more likely than men to enroll in tertiary education (Gulczynski, 2023).

Furthermore, some regions in Germany, Australia and Spain exhibit idiosyncrasies. Eastern German states show particularly low TFR ratios (TFR ratio < 0.85), which may be related to the highly female skewed outmigration after reunification (see Figure 3). As a consequence, the population structures are male skewed in East German states, which increases the size of the exposures for men relative to women, and therefore depresses the TFR of men. In more recent years, the gender difference in fertility may be sustained by the decline of cohort sizes resulting from the small cohorts born in the 1990s after reunification, e.g., the interaction of cohort changes and age patterns discussed in Section 2.1. The second case is the Northern Territory in Australia, which exhibits a similar pattern in terms of TFR ratios. Overall, the TFR ratio is very low, which may point to similarly

skewed population structures. For instance, in this region, there are 11% more men than women in the 20-24 age group, which reduces the male TFR.

Figure 3 about here

5.2 Temporal trends in the TFR ratios

Figure 4 shows TFR ratios over time for different subnational units across countries. Note that higher values indicate higher TFRs among men, while values below 1 indicate higher TFRs among women. The male-female TFR ratios follow different trends over time across countries. Australia, France, and Finland have stable levels of TFR ratios over time with some minor fluctuations up and down, although the levels differ markedly. The United States, Colombia, Germany, Mexico, and Spain exhibit noticeable temporal trends over the observation period. For instance, the male-female TFR ratios for West Germany have risen in recent years, particulary in city-states like Hamburg and Berlin, which suggests that the male TFR increased relative to the female TFR. Over the same period, the TFR ratios of the East German states stagnated at low levels. Therefore, instead of an East-West convergence, the polarization between East and West-Germany has been reinforced three decades after reunification. By contrast, in the United States, Colombia, Mexico and Spain, the TFR ratios have declined sharply over time. The TFR ratio decreased first in the United States in the 1980s, when the TFR ratios declined from above 1.1 to below 1. In Mexico, the TFR ratio declined from 1.18 in the 1990s to 1.06 in 2021. In Colombia, the regional TFR ratios converged to a level of 1.04 between 1998 and 2020.

However, some regions in Mexico and in Colombia have deviated from the larger country-level trend. Figure S.4 in the supplementary materials shows that the TFR ratios in some Mexican states actually rebounded between 2000 and 2010. These increases occurred in states with amplified cross-border migration to the United States and Canada, and mortality related to the war on drugs. Previous research has highlighted the demographic impact of the war on drugs on mortality (Aburto et al., 2016), but fertility seems to be affected as well. The impact of violence on TFR ratios is clearly observed through the impact on population exposures: first, by increasing mortality predominantly among men; and, second, by increasing migration. This underscores the impact of external factors, such as exposure to conflict and its consequences, on the TFR and gender differences in the measure (Castro Torres & Urdinola, 2019; Keilman et al., 2014). In Colombia, some departments show TFR ratios while others have declining TFR ratios over time. As in the case for Mexico, this may be related to the internal armed conflict and migration. These findings persisted in our sensitivity analysis when we accounted for potential under-coverage of men in areas affected by armed conflict (see section D in the supplementary materials).

Figure 4 about here

5.3 Birth squeezes

Figure 5 shows the existence of birth squeezes; the region-year observations are represented as dots and the thresholds are represented as lines. The x-axis shows the TFR of women on a logarithmic scale and the y-axis shows the logarithm of the male-female TFR ratio. The various approaches to defining thresholds are distinguished by the colors of the lines, which yield disparate insights into the data. Furthermore, Figure 6 shows the country share of birth squeezes among region-year observations across measurement approaches. We discuss below which countries may be classified as experiencing a birth squeeze, and why.

Figure 5 about here

The derivation of cut-offs for birth squeezes from previous research underpins the expert-based approach (orange line in Figure 5), which yields 1,714 observations (32.47 %) pertaining to the birth squeeze phenomenon, with a high concentration of birth squeezes in Mexico (n = 586, 57.28%), Germany (n = 254, 66.15%), and Colombia (n = 358, 48.44%). In Germany, for instance, the East German states of Brandenburg, Mecklenburg-Vorpommern, Saxony, Saxony-Anhalt, and Thuringia are consistently identified as experiencing birth squeezes, whereas a combination of birth squeeze and non-birth squeeze observations are found for the West German states. The occurrence of birth squeezes in Germany is likely attributable to two key factors: the migration of women from the East to the West German states and the sudden decline of birth cohorts in the 1990s. In Mexico and Colombia, the regions with high emigration and those that are affected by internal conflict are likely to be categorized as experiencing birth squeezes.

The data-based approach (gray line in Figure 5) accounts for the robust relationship between the fertility level and the male-female fertility differences, which classifies about the same number of region-year observations as birth squeezes as the expert-based approach (n = 1179, 22%). However, the pattern of birth squeezes is different from that observed in the expert-based approach, because it is relatively equally spread across countries (see Figure 6). Moreover, in a few cases the birth squeezes are shifted to other regions. For instance, Finland has a high prevalence of birth squeezes (n = 214, 34.13%). Finland exhibits a discernible central-periphery pattern, with the Helsinki region diverging from the other regions by not having any birth squeeze during this period. The eastern and northern parts of Finland exhibit a notable prevalence of birth squeezes, a finding that aligns with the conclusions presented in Schubert et al. (2024). Female-skewed birth squeezes appear to be concentrated at higher fertility levels, defined as a TFR for females exceeding 5. This is because most of the previous research analyzed country-level data, and found a steeper decline in TFR ratios with fertility levels than it is the case in our subnational data (see Figure S.3). However, some caution is needed here as the country-level data may be biased, particularly the data on male fertility based on survey estimates (Joyner et al., 2012).

The stable population approach (purple line in Figure 5) is more conservative regarding the identification of birth squeezes, which translates into few birth squeeze observations. The overwhelming majority of region-year observations are classified as exhibiting normal deviations (n = 4780, 90.6%), while only a negligible proportion demonstrate TFR ratios indicative of a birth squeeze (n = 498, 9.4%). The stable population approach is most

effective in detecting birth squeezes in regions with rapid population change, such as those observed in Mexico and Colombia, where up to 35% and 14% of region-year observations, respectively, are classified as such. In contrast, the occurrence of birth squeezes is infrequent in Australia (except for the Northern Territory), the United States (District of Columbia), Spain (Teruel), and Germany (Brandenburg, Saxony-Anhalt, Mecklenburg-Vorpommern). Therefore, it can be concluded that the approach detects significant outliers with regard to age-sex structure. All observations in the United States, Spain, Finland, and France fall within the normal range.

The outcome-based approach (red line in Figure 5) identifies a significant proportion of birth squeezes (n = 910, 17.24%), although the majority of region-year observations fall within the normal range (n = 4383, 82.75%). According to this approach, the most birth squeezes are found in Mexico (445, 43.50%), Germany (158, 41.15%), and Colombia (183, 24.76%). In Mexico, as much as 44% of the region-year observations are classified as birth squeezes, and the proportions are similarly elevated in Colombia (25%) and Germany (41%). In contrast, only a few birth squeezes are observed in the United States (43, 2.55%) and none are observed in France.

Figure 6 about here

France is notable for the small number of subnational birth squeezes across time, across regions, and across measurement approaches. These findings are consistent with research on the population structure in France, which has been very balanced between the sexes in recent decades (Breton et al., 2024). The balance of men to women may be attributed to the demographic history of France, which has influenced the population structure. Due to stable fertility rates, cohort sizes have changed only marginally and gradually. The TFR exhibited only small fluctuations between 1.8 and 2 over the past four decades. Moreover, the very low female mortality in combination with a noticeable mortality hump among males compensates for the sex ratio at birth (Breton et al., 2024). Finally, the population is more female than the EU average due to migration patterns (Breton et al., 2024). Migration from former colonies attracts both genders.

5.4 Parity analysis for Finland

We exploit the richness of the Finnish register data by studying the gender differences in childlessness and parity-specific fertility. Our estimates, e.g., the parity-specific hazard-ratio between men and women across time ($hazardratio(t) = \frac{\gamma_m(t)}{\gamma_f(t)}$,) region and age, show that the main gender differences pertain to the transition to first birth (more details in C). We thus proceed to discuss the results pertaining to gender differences in childlessness in greater detail, which are presented in Figure 7.

Figure 7 about here

Childlessness among men is consistently and substantially higher than childlessness among women, as presented in Figure 7. The large majority of region-year observations

are above the diagonal line, which indicates that childlessness is more frequent among men than women. Note that this observation even accounts for the wider fertility time window among men, as we compare the level of childlessness among men at ages 50-55 to their female counterparts at ages 45-50. The central-periphery pattern emerges again, which is highlighted by color, showing that the gender differences are larger in rural areas and smaller in urban regions. Helsinki is the only region that had higher levels of female childlessness in some years, and thus falls below the diagonal line. The transparency highlights the temporal dimensions, showing more recent observations in darker colors. The temporal trend indicates that region-year observations have become less spread out and converged to a difference of about 8 percentage points.

6 Discussion

6.1 Main findings

In this paper, we examined gender differences in fertility at the subnational level over time and across countries. We systematically assessed whether these differences can be considered a birth squeeze; that is, a gender differential in the quantum of fertility that can be considered high. To measure differentials and birth squeezes, we employed the TFR ratio, i.e., the ratio of the TFR of men relative to the TFR of women. We revealed substantial differences between male and female fertility, and found that birth squeezes occurred in several countries. Intriguingly, the differences at the subnational level often exceed the national-level results. Capital and metropolitan regions often have relatively high TFR ratios, which may be related to female-skewed populations. The results for Australia, Colombia, Mexico and East Germany show that historical events, which skew population age-sex structures, can have strong impacts on gender differences in fertility. France stands out with balanced TFRs for men and women.

We argue for the use of the TFR ratio as an indicator of quantum differentials by gender, and developed a series of thresholds for the identification of birth squeezes. The TFR ratio is the ratio of the TFR of men relative to the TFR of women. Moreover, we have suggested several approaches to identify strong deviations in TFR ratios from unity, e.g., birth squeezes. These thresholds are derived from the literature (*expert-based approach*), from empirical patterns in country-level data (*data-based approach*), from the demographic mechanisms driving TFR ratios (*stable population approach*), and from the potential impact on childlessness (*outcome-based approach*). While there is some variation in the thresholds, and thus in what levels of the TFR ratio can be considered to indicate a birth squeeze, the thresholds are mostly close to 0.9 as a lower bound and 1.1 as an upper bound, respectively. This is largely consistent with the thresholds used previously in the literature. Moreover, our findings on the occurrence of birth squeezes also match those from the previous literature; e.g., our findings for Eastern Germany match those of Dudel and Klüsener (2016) and our results for the United States are in line with those of Dudel and Klüsener (2021) and Schoen (1985).

More broadly, we find that fertility differentials at the subnational level, whether birth squeezes or not, are larger than those observed at the national level. This observation

is in line with Dudel and Klüsener (2016), who found significantly larger differentials in Eastern Germany than in total Germany. Moreover, our subnational results on male-female fertility differentials are consistent with those from cross-country comparisons by Schoen (1985), Schoumaker (2019), and Dudel and Klüsener (2021). In line with Schoen (1985) and Schoumaker (2019), we find declining fertility differentials in all countries except for Australia and France. At low fertility levels, TFR ratio fluctuations support the country-level findings of Dudel and Klüsener (2021).

The results of our study indicate that population structures exert a considerable influence on gender differences in fertility rates. The most significant factor influencing gender differences in fertility appears to be the population growth rate, which impacts the change in subsequent cohort sizes due to the decline in TFR ratios associated with the female TFR. Another noteworthy observation is the tendency for capital regions to exhibit relatively high TFR ratios compared to the level of female TFR, which may be attributed to the presence of a large tertiary sector and a well-developed education system. Additionally, regions with distinct cultural or economic characteristics often have higher TFR ratios, as evidenced by examples such as Massachusetts and the Northern territory in Australia.

Beyond these common patterns, the demographic history of countries and regions is reflected in their TFR ratios. For instance, regions in Mexico and Colombia, which were affected by conflict related-mortality and substantial cross-border emigration, show high TFR ratios and even reversals of the narrowing of the gender difference in fertility. Moreover, Eastern German states exhibit considerably lower fertility rates for men than for women, which may be related to the gender-selective migration in conjunction with declining cohort sizes. This has meaningful implications, as the record low TFR of 0.703 was observed for men in Brandenburg in 1995, which is substantially lower than the value for women in Brandenburg in that year (0.772) and the record low TFR at the country level in South Korea of 0.72 (2023).

Furthermore, parity-specific analysis for Finland revealed that the most-robust gender differences exist for the transition to first birth, with women having a higher risk relative to men (Chudnovskaya & Ueda, 2021; Kravdal, 2021). This also translates into a robust gender difference in ultimate childlessness, as childlessness is more widespread among men across almost all region-year observations. However, at higher ages and higher parities, men also show higher relative birth risks, which may be related to the more gradual age-related decline in fecundity among men (Hassan & Killick, 2003).

6.2 Strengths and limitations

The data we use in this paper come with challenges. First, the subnational units vary strongly in terms of population size and surface area, making cross-country comparisons difficult. Nonetheless, the variation in spatial units also allows us to examine how the level of analysis may impact the results. Second, data quality may vary across countries and over time. This problem is particularly relevant for regions affected by internal conflicts, such as regions in Mexico and Colombia. Finally, variations in subnational age-sex population structures might reflect a delay in household registration or job-related migration. Moving from one region to another within the same country may not result

in immediate registration of a new residence, which could distort demographic data. Consequently, some age-sex population structures might be data artifacts rather than accurate demographic representations.

Our key indicator is the TFR ratio. It is straightforward to calculate and easy to interpret. However, a potential drawback is that changes in the numerator (male TFR) and the denominator (female TFR) affect it differently. That is, an increase in the TFR of men by a certain amount δ will not have the same effect as the same increase in the TFR of women. Relatedly, inverting the TFR ratio and using the TFR of men as the reference can lead to inconsistent results when thresholds symmetric around 1 are used. If, for instance, the thresholds are 0.9 and 1.1, then a TFR ratio of 1.11 will be considered a birth squeeze; while its inverted value, $1/1.11$, is slightly above 0.9 and thus is not a birth squeeze. Finally, TFR ratios are period measures, using the concept of a synthetic cohort. This means that the findings do not readily translate to any lived experience by real cohorts. If the fertility of one sex is substantially lower than the fertility of the other in one year, this gap could be compensated for over the span of their reproductive life course. Nevertheless, Klein (2003), Kravdal (2021), and Schubert and Dudel (2024) have demonstrated that regional population imbalances affected the number of children born in the more abundant sex, and in this paper we also linked the TFR ratio to childlessness.

We combine the TFR ratio with a series of thresholds to determine birth squeezes. These thresholds are not without issues. The expert-based thresholds partly pick up at least seemingly arbitrary choices from the previous literature. The data-based approach might be problematic if many observations from the input data are actually affected by birth squeezes, as this approach would classify these as not affected by a birth squeeze. The assumptions of the stable population approach are very simplifying, and in particular ignore migration, which seems to be a key driver of fertility differentials. Finally, the outcome-based approach only builds on Finnish data, and requires the assumption that findings for Finland also apply in other, sometimes very different contexts. Despite these limitations, however, these heterogeneous approaches lead to relatively consistent thresholds.

6.3 Outlook

Our findings underscore that there is not just one number summarizing the state of fertility, but one for men and one for women, which can diverge substantially at times. Therefore, we encourage publishing the TFR of men along with the TFR of women, and argue that empirical analysis should use both the TFR of men and the TFR of women, as showcased in Schubert et al. (2024). An avenue for future research is the development and application of two-sex fertility measures, in line with work by Keilman et al. (2014).

We detected the existence of subnational birth squeezes across eight middle and high-income countries, but there is reason to expect that similar imbalances would be observed in other countries around the world. For instance, gender differences in fertility might be even more widespread and even larger in countries simultaneously experiencing sudden fertility declines and high SRBs, such as China (SRB in 2017 = 1.1143) and India (SRB in 2017=1.098) (Dyson, 2012). Subnational migration related to education, marriage prospects, and employment may reinforce these gender imbalances (Edlund, 2005; Gupta,

2010). Moreover, while we have linked birth squeezes to increased levels of childlessness, further studies on the demographic and socio-economic consequences of a birth squeeze would be valuable.

Acknowledgements

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7 Figures

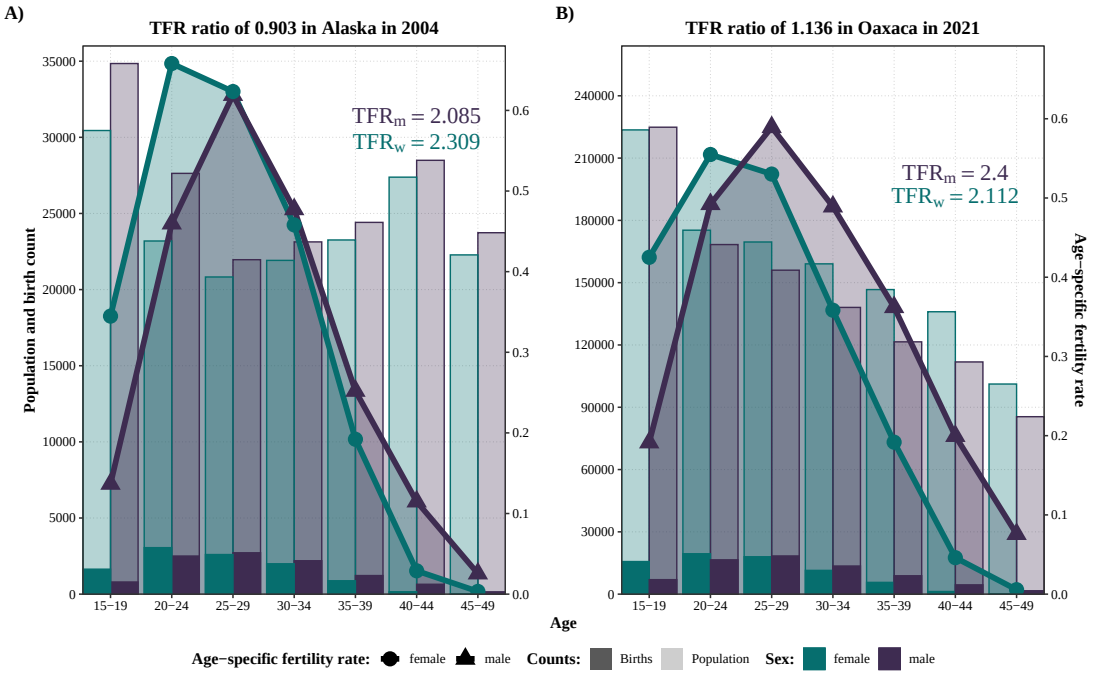

Figure 1: This figure illustrates the concept of birth squeezes in Alaska (United States of America) and Oaxcaca (Mexico) in 2004. The count data refers to the left y-axis, population exposures are illustrated in the shaded bars and the birth counts are illustrated in the solid bars. The age-specific fertility rates shown in lines refer to the right y-axis. The TFR ratio is the ratio of the area under the purple line to the area size under the green line. Panel B) illustrates a male birth squeeze in Alaska, where the TFR ratio was at 0.9 as the male TFR was lower relative to the female TFR. Panel B) shows a female birth squeeze for Oaxaca where the TFR ratio was at 1.14, indicating a higher male TFR relative to female TFR.

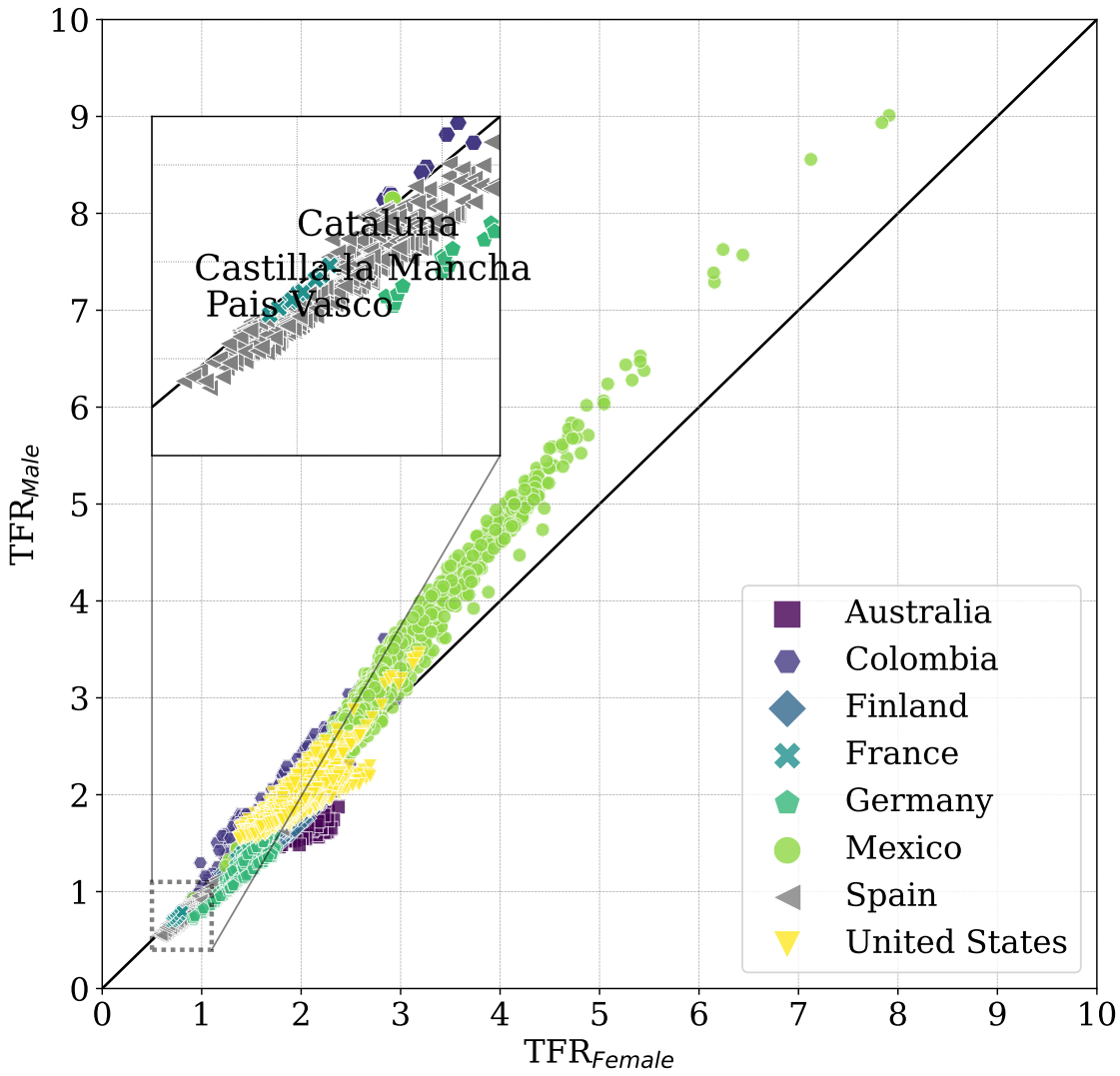

Figure 2: Gender differences in fertility quantum at the regional level. The figure shows the relationship between the female TFR displayed on the x-axis and the corresponding value for the male TFR in the same region in the same year. The lower the overall fertility level the smaller the gender difference.

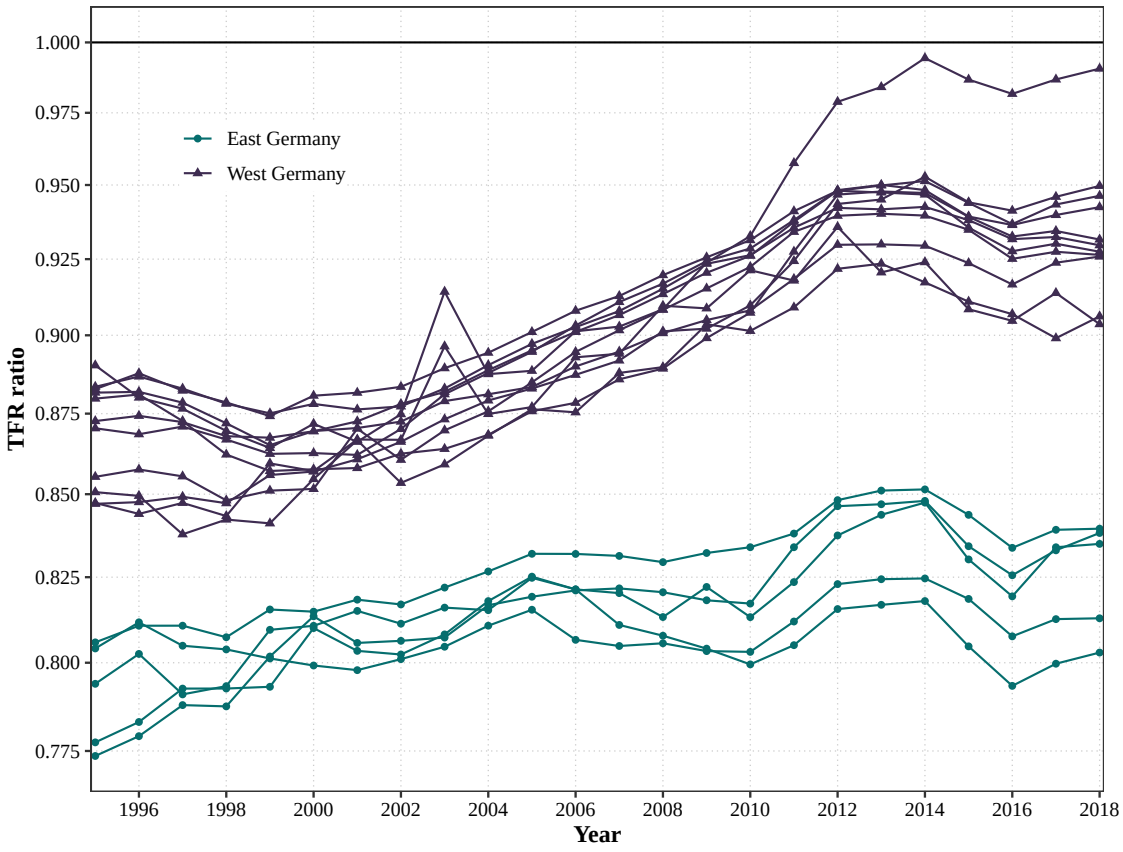

Figure 3: Male TFR / Female TFR ratio over time and across states (*German* Bundesländer) in Germany.

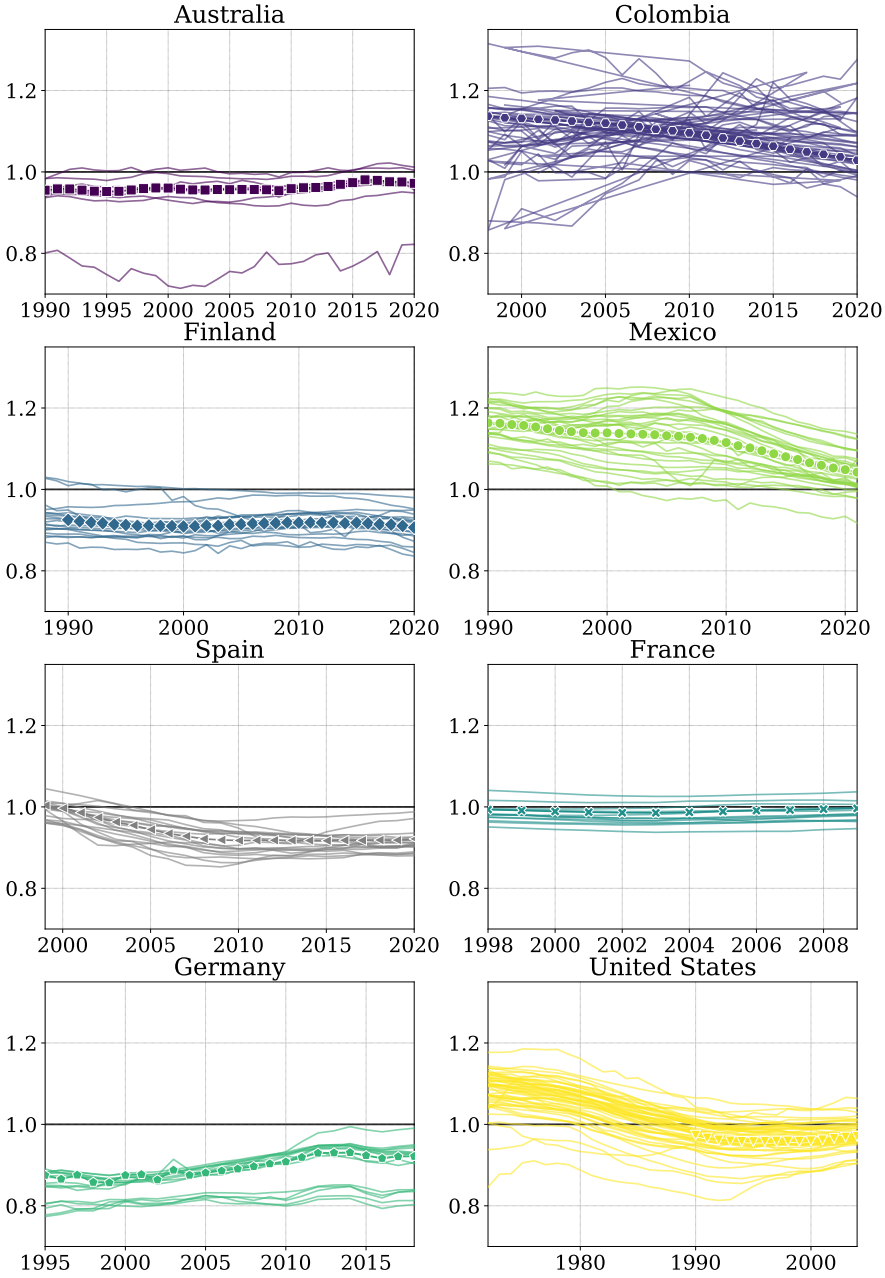

Figure 4: This figure displays the temporal trend of the male female TFR ratio over time across countries. The panels display the values for different countries. Note the different scales on the x-axis.

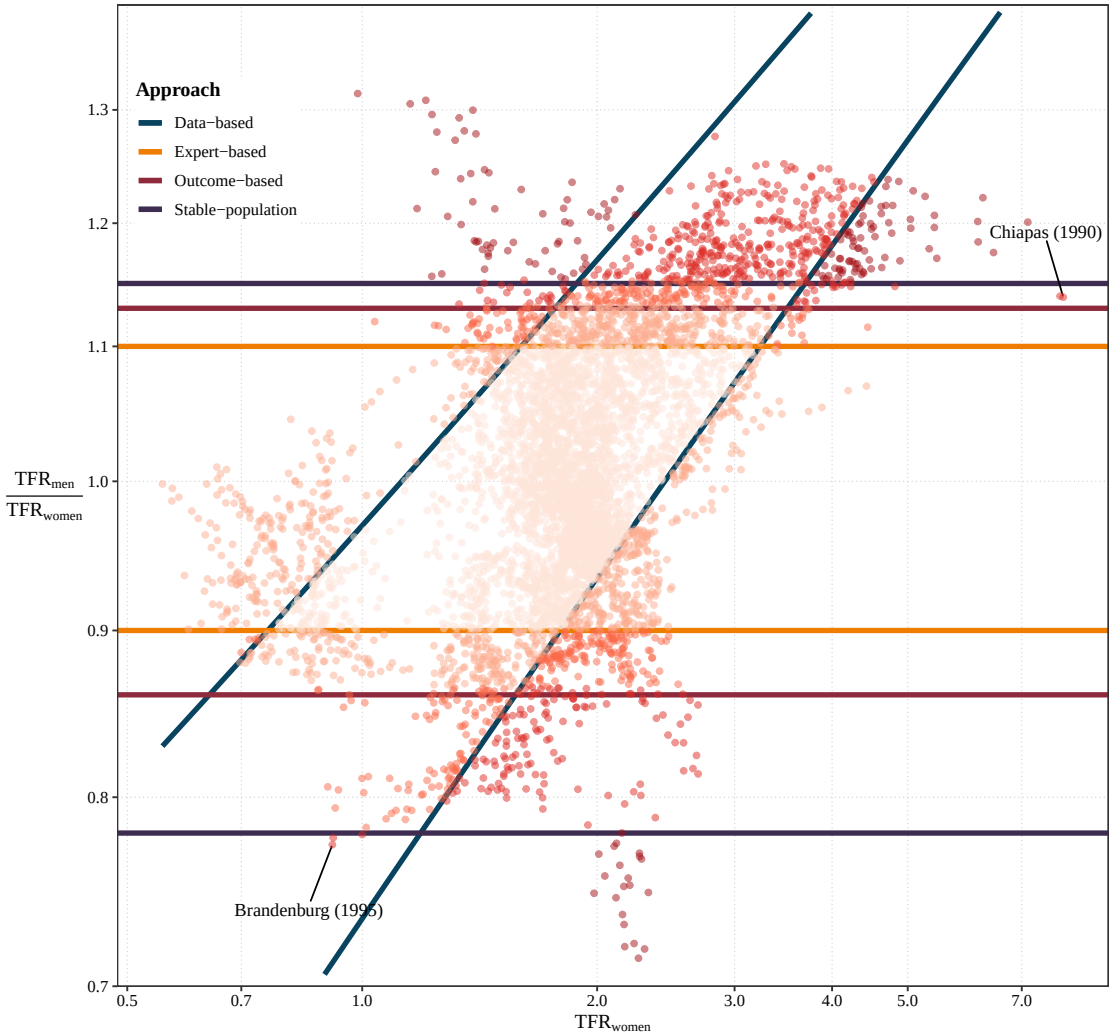

Figure 5: This figure displays the thresholds for detecting birth squeezes in a cartesian coordinate system showing TFR ratios (y-axis, logarithm) in relation to the TFR of women (x-axis, logarithm). The lines indicate the thresholds for the different threshold approaches of birth squeezes. The points are the subnational region-year observations and the colour coding of the points indicate how many threshholds classify the points as a birhth squeeze, if zero it is not classified by any approach as a birth squeeze, and if it is four, all threshold approaches classify the observation as birth squeeze. For example, the point in the bottom left corner represents Brandenburg in the year 1995, where the TFR of women was 0.916 and the male-female TFR ratio was 0.774. In the top right corner lies the point for Chiapas in 1990, where the TFR of women was 7.91 and the male-female TFR ratio was 1.14. For a different representation of the data, see Figure S.5 in the supplementary materials.

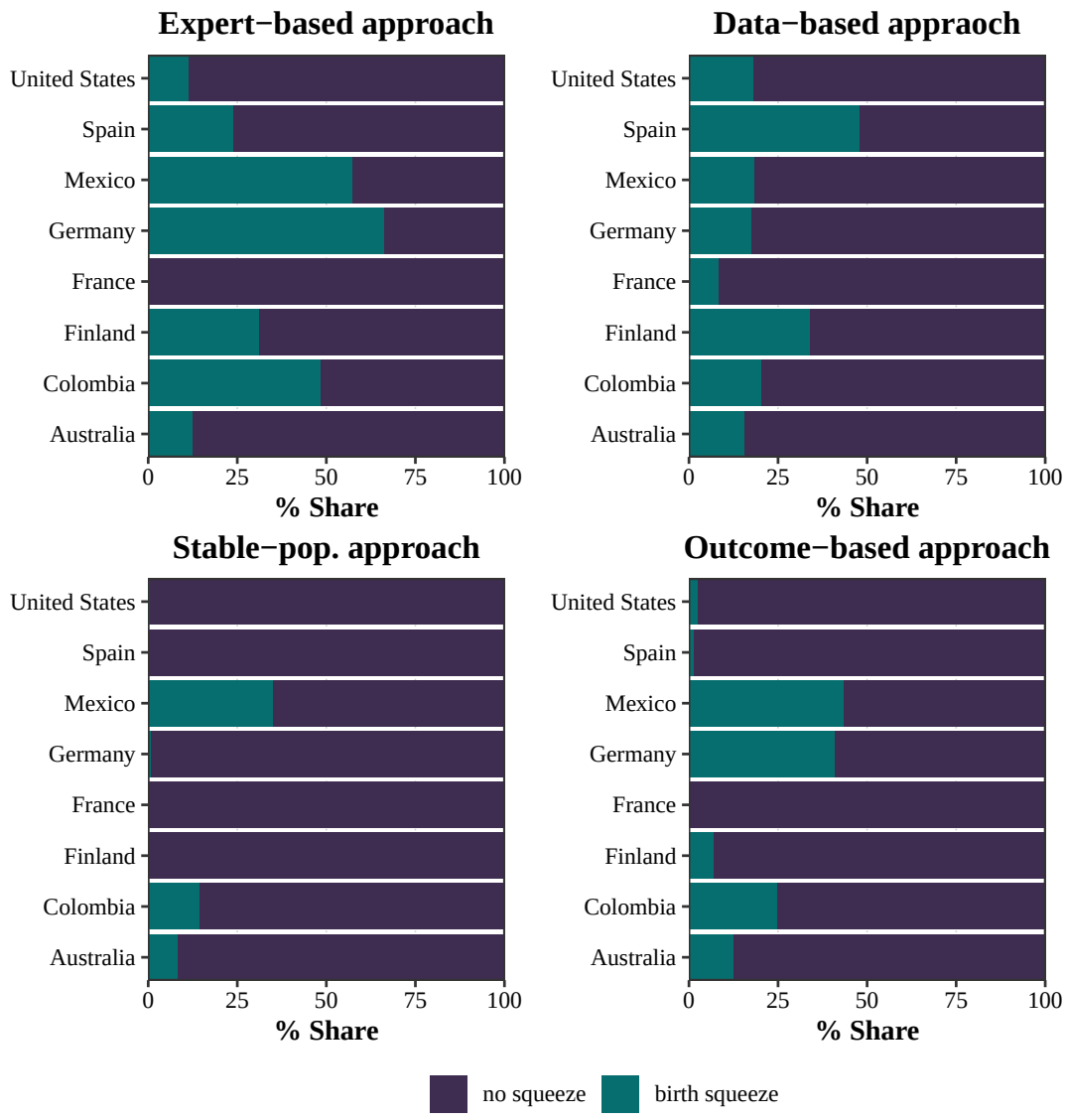

Figure 6: This figure displays the share of birth squeezes across countries. The yellow bar indicates the share of birth squeezes, whereas the purple bar illustrates the share of normal fluctuations.

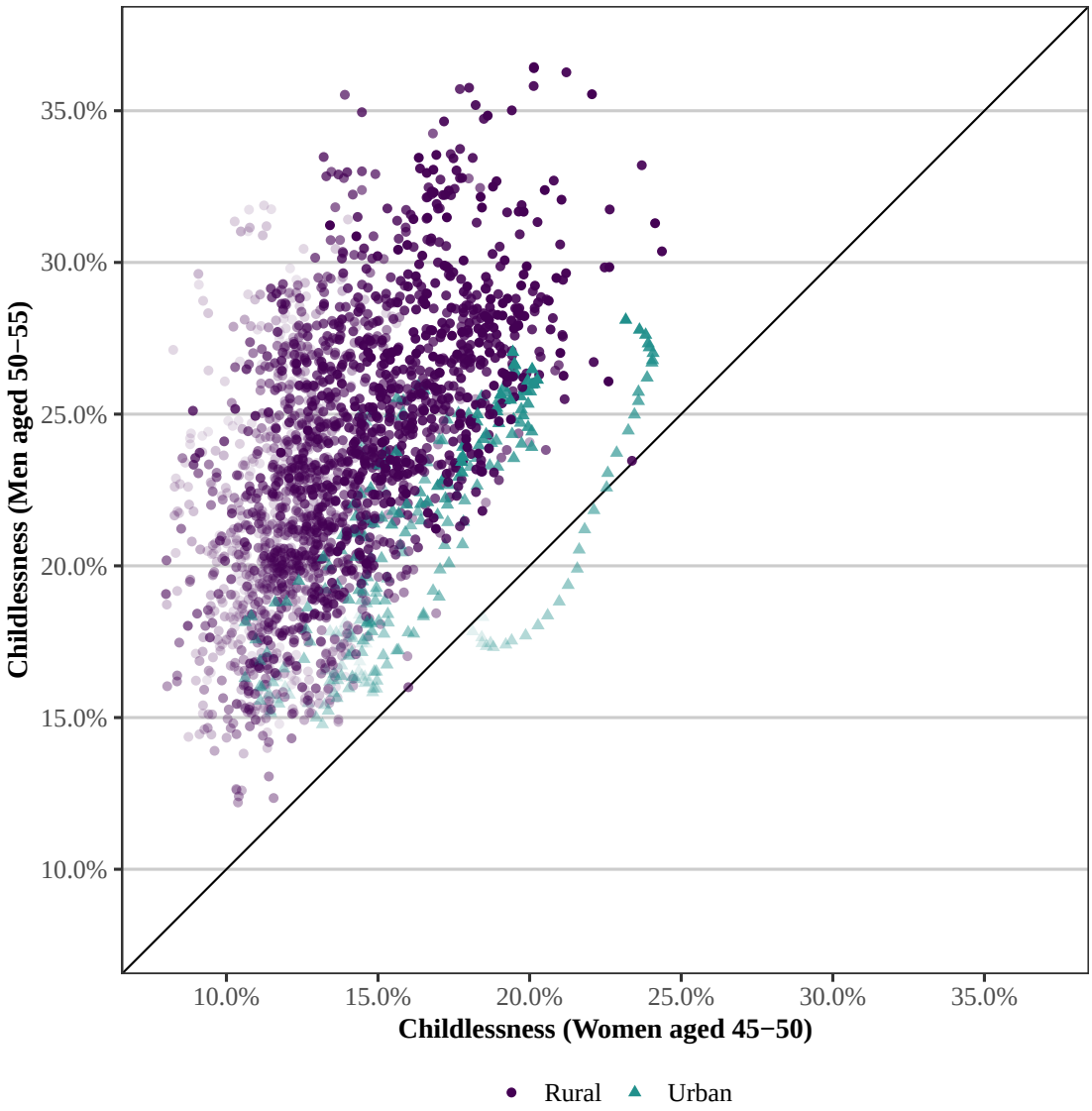

Figure 7: This figure illustrates the relationship of male and female childlessness across regions in the period between 1987 and 2020. The purple points represent rural regions and the turquoise urban regions. The transparency of the points indicates the time of the observation, showing more recent observations in stronger colours.

8 Tables

Table 1: This table summarizes the different approaches and their respective lower and upper cut-offs for birth squeezes.

Approach	Reasoning	Lower	Upper
Expert-based (E)	TFR ratios in previous literature.	0.9	1.1
Data-based (D)	TFR ratios conditional on fertility level.	$exp(-0.161 + 0.304 \cdot log(TFR_w)) \pm 0.12$	
Stable population (S)	Sources of abnormal TFR Ratios.	0.78	1.15
Outcome-based (O)	Impact of TFR ratios on childlessness.	0.86	1.13

Table 2: The table summarises the data used in the study, providing information on the country, the observation period, the spatial unit, the number of spatial units, and a link to the source.

Country	Period	Level	Units	Source
Australia	1990-2020	States, Territories	8	explore.data.abs.gov.au
Colombia	1998-2020	Departments	32	https://microdatos.dane.gov.co
Finland	1990-2020	Regions	19	https://www.stat.fi/
France	1989-2013	Departments	81	insee.fr/fr/statistiques
Germany	1990-2018	States	16	https://www.destatis.de
Mexico [a]	1990-2021	Regions	32	inegi.org.mx/programas/natalidad
USA	1969-2004	States	51	https://data.nber.org/natality/
Spain	1998-2020	Provinces	32	https://www.ine.es/

[a] The time-series is not complete for all states. There is no information for Aguascalientes, Baja California, Baja California Sur, Campeche, Chiapas, Chihuahua, Ciudad de México, Coahuila de Zaragoza and Colima after 2015.

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Appendix

Table of Contents

A Birth squeezes thresholds

A.1 Expert-based approach

The expert-based approach relies on thresholds established in the existing literature, assuming that the authors of previous studies had valid reasons for selecting these thresholds, even if they were not explicitly stated. For instance, Dudel and Klüsener (2021) adopted the lowest total fertility rate (TFR) ratio reported by Schoen (1985) as a lower threshold, which was 0.89 in England and Wales in 1973. Additionally, Schoen (1985) noted that the TFR ratio of 1.12 in the United States in 1970 was relatively high compared to that in other high-income countries. This suggests a range of thresholds between 0.89 and 1.12. Coleman (2000) reported sex ratios ranging from 0.91 (Denmark in 1988) to 1.1 (German birth cohorts from 1902 to 1928). Gulczynski (2023) studied sex ratios and found that deviations from 1 greater than ± 0.085 were rare, indicating gender-skewed populations. This led to slightly narrower thresholds of 0.915 and 1.085. For the purposes of this analysis, we will use 0.9 and 1.1 as thresholds, as they fall within the range of the three variants with minimal differences. Furthermore, these values are equivalent to rounding the TFR ratio thresholds from the literature.

A.2 Data-based approach

We estimate a Bayesian log-log linear regression model to assess the relationship between TFR_{women} and the TFR ratio. We opted for a Bayesian model and not for a frequentist model, because our objective is to express uncertainty in parameter estimates, which can be more consistently achieved within the Bayesian framework (Gelman et al., 2020). The model estimation follows,

$$Likelihood: \ R = Normal(log(\mu_i), \sigma), \tag{1}$$
$$\text{Linear model}: \ log(\mu_i) = \alpha + \beta \ log(x_i), \tag{2}$$
$$Prior: \ \alpha = Normal(\mu = 0, \sigma = 2), \tag{3}$$
$$Prior: \ \beta = Gamma(\kappa = 1, \theta = 1), \tag{4}$$
$$Prior: \ \sigma = HalfNormal(\mu = 0, \sigma = 1), \tag{5}$$

where μ_i is the mean prediction, which depends on the transformation of x_i (TFR of women) using α_i and β_i, see equation 2. We also obtain an estimate of the standard deviation σ.

We selected a normal distribution as Likelihood-Function, but we take the log of the outcome to obtain positive predictions.Thus, our predictions fulfill that TFR ratios are a 1) continuous variable, and 2) range between 0 and ∞.

The priors were selected based on both mathematical and substantial considerations. The prior for α is neutral, as it is centered on 0 and has a standard deviation of 2. The prior for β is set to a gamma distribution with a shape parameter $\kappa = 1$ and a rate parameter of $\theta = 1$, because the gamma distribution ranges only over positive values, so that it captures the observation that the TFR ratio increases with the TFR of women (Schoumaker, 2019). The prior for σ follows a half-normal distribution with a $\mu = 0$ and $\sigma = 1$, which yields only positive values.

In order to obtain confidence intervals (c_α), we draw on the normal distribution and the estimated μ_i and σ:

$$c_\alpha = exp(\alpha + \beta \cdot log(TFR_w)) + 1.96 \cdot \sigma \tag{6}$$

Table S.1: This table displays the results from the regression model of the country-level relationship between TFR_{women} and the male-female TFR ratio using data from Dudel et al. (2021), Schoumaker (2019) and Schoen (1985).

	variable	median	5%-Credible Interval	95%-Credible Interval
1	α	-0.161	-0.171	-0.15
2	β	0.304	0.29	0.319
3	σ	0.091	0.088	0.096

A.3 Stable population approach

Stable population theory describes the asymptotic behaviour of closed populations with constant fertility and mortality rates. A closed population assumes no migration. The stable population approximation of the male female TFR ratio by Schoumaker (2019) and Tucker and Van Hook (2013) looks as follows:

$$\frac{TFR_m}{TFR_f} \approx \frac{1}{SRB} \cdot \frac{p(MAC)}{p(MAF)} \cdot exp(r \cdot (T^m - T^f)), \tag{7}$$

where SRB is the sex ratio at birth, $p(MAC)$ and $p(MAF)$ are the probability of surviving until the mean age at childbearing, r is the growth rate, and T is the generation length.

A.3.1 Data sources for the stable population approach

We can study the behavior of male-female TFR ratios through the lens of the stable population approximation using realistic input values. For that exercise, we derive input values using data from different sources. Female and male life tables (1-year) were obtained the Human Mortality Database (Wilmoth et al., 2017). Male and female age-specific fertility rates were obtained from the Human Fertility Collection and the Human Fertility Database. Estimates of the sex ratio at births are drawn from (Chao et al., 2019) and the World Population Prospects 2024 (United Nations Department of Economic and Social Affairs, 2024).

A.3.2 Estimation for the stable population approach

Some input values require further transformation steps, which we are going to explain below. We estimate the probability of surviving to the mean age of childbearing in a sequential way. First, we estimate the mean age of childbearing based on age-specific fertility rates: $\frac{\sum_{x=15}^{59} f_x \times x}{\sum_{x=15}^{59} f_x}$. Second, once we have the mean age of childbearing, we look up the survival probability until that age from the female life table, which is given by ℓ_x/ℓ_0. The estimation of the survival probability until the mean age of fatherhood follows an analogous procedure, but uses data for men.

The growth rate in a stable population is the first *eigenvalue* of the Leslie matrix ($\boldsymbol{A}$) based on the life table and age-specific fertility rates. The Leslie matrix is a matrix expression of the stable population. The matrix contains only values in the first row (fertility) and subdiagonal (survival) and we follow the notation in Wachter (p. 107, 2003). We estimate the elements in the first row by $A_{1,j(x)} = \frac{{}_nL_0}{2\ell_0}\left({}_nF_x + {}_nF_{x+n}\frac{{}_nL_{x+n}}{{}_nL_x}\right)f_{fab}$. The values on the subdiagonal follow $A_{j+1,j} = \frac{{}_nL_{x+n}}{{}_nL_x}$. The growth rate r is equal to the harmonic mean of the male and female *eigenvalues*: $r = \frac{2}{1/\lambda_1^{Am}+1/\lambda_1^{Am}}$.

The estimation of the generation length follows Keyfitz and Caswell (p. 169, 2005) and looks as follows:

$$T^f = \frac{\sum_{x=0}^{59} x \cdot p_x \cdot f_x}{\sum_{x=0}^{59} p_x \cdot f_x} \tag{8}$$

A.4 Parameter estimates for the stable population approximation

The estimated parameter values are shown in Table S.2 in the column input range. The sex ratio at birth is set to range between 1.04 and 1.06. The population survival ratio to the mean age at childbearing is set to 0.915 and 1.085. The difference in generation length ranges between 2 and 4. The population growth rate is set to -0.01 and 0.01.

Table S.2: This table summarizes the simulation experiment using the stable population approximation by Schoumaker (2019) and realistic input values for the components.

Component	Estimation of components	Input range	TFR ratio	U (Schoen, 1985)
Sex ratio at birth	$SRB = \frac{B_m}{B_f}$	min = 1.04	0.85 to 1.09	-0.17 to 0.08
		max = 1.06	0.83 to 1.07	-0.19 to 0.06
Population ratio	$p(MAC) = \frac{\ell_{MAC}}{\ell_0}$	min = 0.915	0.83 to 0.92	-0.19 to -0.09
		max = 1.085	0.98 to 1.09	-0.02 to 0.08
Generation length	$T = \frac{\sum_{x=15}^{55} p_x \cdot m_x \cdot x}{\sum_{x=15}^{55} p_x \cdot m_x}$	min = 2	0.84 to 1.08	-0.18 to 0.07
		max = 4	0.83 to 1.09	-0.19 to 0.08
Growth rate	$r = \lambda_1^A$	min = -0.01	0.83 to 1.01	-0.19 to 0.01
		max = 0.01	0.89 to 1.09	-0.12 to 0.08
Overall	$\frac{1}{SRB} \cdot \frac{p(MAC)}{p(MAF)} \cdot exp(r \cdot (T^m - T^f))$	See above	0.83 to 1.09	-0.19 to 0.08

Note: The columns show from left to right the components how it is commonly estimated, shows the range of the input values,the range of resulting TFR ratios and U values. The rows are ordered following the appearance in the equation 7.

Source: Authors' analysis of data from Dudel and Klüsener (2021), Chao et al. (2019), and the Human Mortality Database (Wilmoth et al., 2017).

A.5 Outcome-based approach

The outcome-based approach uses regression analysis on individual-level data from Finnish administrative registers to study the strength of the relationship between TFR ratios and childlessness. In this section, we describe the data processing, the model estimation, and the algebra to derive the cut-offs.

First, we link the region-year TFR ratio data with individual data by year and residence in a region. Thus, we obtain data on the TFR ratio at the place of residence in a specific year. Second, we obtain information on the transition to first birth for all individuals. If a person does not have a child, the person is considered childless. Third, we filter respondents at age 30. Moreover, we filter persons who complete age 40 before the last year of observation, e.g., 2020. We then have the following data structure.

Table S.3: Example for the data structure used in the impact approach. Note that the data is fictive.

ID	Year	Age	Region	TFR ratio	Childless
1	1990	30	Uusimaa	0.99	1
⋮	…	⋱	⋱	⋱	⋮
2	2005	30	Lappland	0.75	0

Source: Authors' analysis of population register data from Statistics Finland.

We then estimate a logistic regression to estimate the association between the TFR ratio at age 30 and being childless at age 40. We model childlessness using a logistic regression model, such that

$$\Pr(C|TFR_m/TFR_f) = \frac{\exp(\alpha + \beta \frac{TFR_m}{TFR_f})}{1 + \exp(\alpha + \beta \frac{TFR_m}{TFR_f})}, \tag{9}$$

where C is childlessness, TFR_m/TFR_f is the TFR ratio, and α and β are the coefficients.

What we are interested in: how much deviation δ from a balanced ratio, $TFR_m/TFR_f = 1$, is needed to change $\Pr(C)$ by some value μ. That is, what is δ such that $\Pr(C|1+\delta) - \Pr(C|1) = \mu$. We are solving this equation below.

Plugging in definitions

$$\mu = \frac{\exp(\alpha + \beta(1+\delta)))}{1 + \exp(\alpha + \beta(1+\delta))} - \frac{\exp(\alpha + \beta))}{1 + \exp(\alpha + \beta)} \tag{10}$$

For simplification we write $l_\delta = \exp(\alpha + \beta(1+\delta))$ and $l = \exp(\alpha + \beta)$ such that we have

$$\mu = \frac{l_\delta}{1 + l_\delta} - \frac{l}{1 + l} \tag{11}$$

After rearranging we obtain

$$\delta = \frac{\log\left(\frac{-l-\mu l-\mu}{\mu-1+\mu l}\right)-\alpha-\beta}{\beta} \tag{12}$$

If the point of reference is not $TFR_m/TFR_f = 1$ but some ratio $TFR_m/TFR_f = r$, the last equation becomes

$$\delta = \frac{\log\left(\frac{-l_r-\mu l_r-\mu}{\mu-1+\mu l_r}\right)-\alpha-\beta r}{\beta} \tag{13}$$

where $l_r = \exp(\alpha+\beta r)$.

B Birth squeeze thresholds using Schoen's U

Table S.4: This table summarizes the different approaches and their respective lower and upper cut-offs for birth squeezes.

Approach	Reasoning	Lower	Upper
Expert-based (E)	TFR ratios in previous literature.	-0.105	0.095
Data-based (D)	TFR ratios conditional on fertility level.	$0.66 + \delta^{a}$	$0.85 + \delta^{a}$
Stable population (S)	Sources of abnormal TFR Ratios.	-0.247	0.140
Outcome-based (O)	Impact of TFR ratios on childlessness.	-0.151	0.122

$\delta = 0.136 \cdot TFR_w$.

C Parity-specific analysis

This section presents the results on gender differences in age-specific transition risks at different parities. The results are based on Poisson regression models with p-splines. The risk ratio is the ratio of the predicted male age-specific transition hazard to the predicted female age-specific transition hazard for the same parity. The results in Figure S.1 indicate gender differences in the risk of first birth, which are particularly pronounced at young ages, and no systematic pattern for higher parities.

The top panels of Figure S.1 show values below 1, which indicates a higher risk of first birth for women across all ages, across all regions, and over time. However, the gender difference in the risk of first birth diminishes with age, as the lines approach 1. For instance, after 5 years of process time (age 20) men have a 50% lower risk of transitioning to a first birth. However, the figure approaches unity after 20 years of process time (age 35). Geographic heterogeneities in gender differences result from differences in settlement type, as rural areas have large gender differences. For instance, at age 35, there is almost no difference in the risk of first birth between the sexes in Uusimaa, which is the Helsinki region, while in more rural areas, the risk remains lower for men.

Regarding higher parities, we see erratic fluctuations around equilibrium, which relate to small case numbers. However, it seems that at higher values of process time, which occur at later ages, we observe higher transition risks for men relative to women. For instance, we find a relative risks of up to 3 for men for higher parities at the end of the process time.

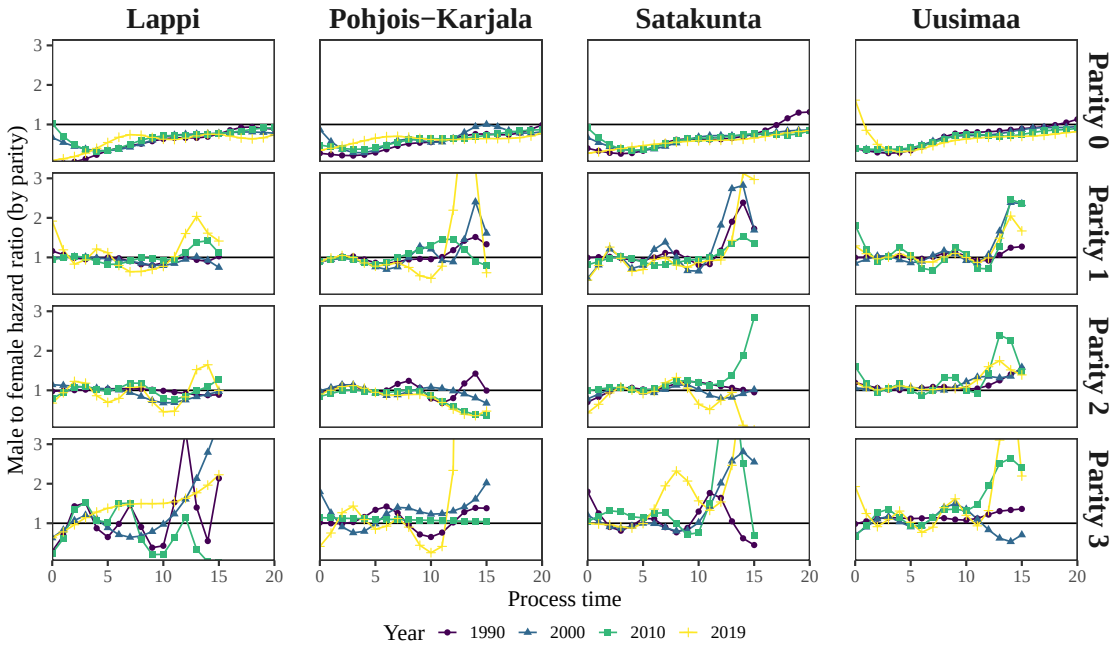

Figure S.1: Male-female hazard ratio of transitioning to the next birth. Values higher than 1 indicate a higher risk of men. For first births (Parity 0) the process time refers to time since age 15. For higher parity births, the process time is measuring the time since the last birth.

D Misreporting in war zones

Mexico's war on drugs and Colombia's internal armed conflict may affect the data quality, leading to a potential bias of our estimates on gender differences in fertility (Castro Torres & Urdinola, 2019). Our results would be biased, if the resulting undercoverage is different for men and for women. Gender-symmetric biases cancel out in our measure, given that we analyse TFR ratios. We use sensitivity analysis to account for the possibility of such a bias. We found increasing TFR ratios in war zones (see Section 5), which could be explained by two patterns. First, excess male mortality reduces male exposures stronger than female exposures during conflicts. Second, conflicts may lead to administrative delays and malfunctionings, which may contribute to undercounting men, which diminishes male exposures relative to female exposures. We want to make sure that the conclusions are robust to the latter.

We pursue two approaches to account for potential undercoverage. Both approaches draw on weights that account for the bias in the male exposures (w_m). We simply multiply the male population exposures by the weight that accounts for the undercoverage (${}^mL_i(x) = {}^mL_i(x) \cdot w_m$). The approaches differ in the way the weights are estimated and applied.

On one side, we assume that men are constantly undercount (constant bias across regions), so that we multiply all male exposures by a weight. We assume that the undercount is 10%, so that the weight w becomes 1.1 (see equation 14).

$$L_i^m(x) = {}^mL_i^m(x) \cdot w^m; w^m = 1.1 \tag{14}$$

This rather crude approach, however, does not distinguish between war zones and the rest, which may over-adjust in less affected areas.

Therefore, we use a more nuanced approach that derives the weight from a measure of conflict intensity, e.g. excess homicide rates (see equation 15). Drawing on mortality statistics, we estimate the proportion of homicide deaths for all regions, which is the ratio of deaths caused by homicide to all deaths. The higher the number, the more people die from homicides. In the next step, we estimate an expected value of proportion of homicide deaths, which we say is the country mean increased by two standard deviation. This rather complicated measure is inspired by the fact that everything that is one standard deviation higher from the mean homicide rate is an extreme value, similar to box plots. Finally, we derive the weight w^m from the excess homicide rate, which is the percent difference between the expected and the observed homicide rate.

$$w_i(x) = \frac{(m_i^h(x) - E[m^h(x)])}{E[m^h(x)]}; E[m^h(x)] = \sum_{i=1}^{n} m_i^h(x) \cdot \frac{1}{n}, \tag{15}$$

where m is the homicide rate, x indicates the year, the subscripts ^{m} and ${}_i$ indicate men and region respectively.

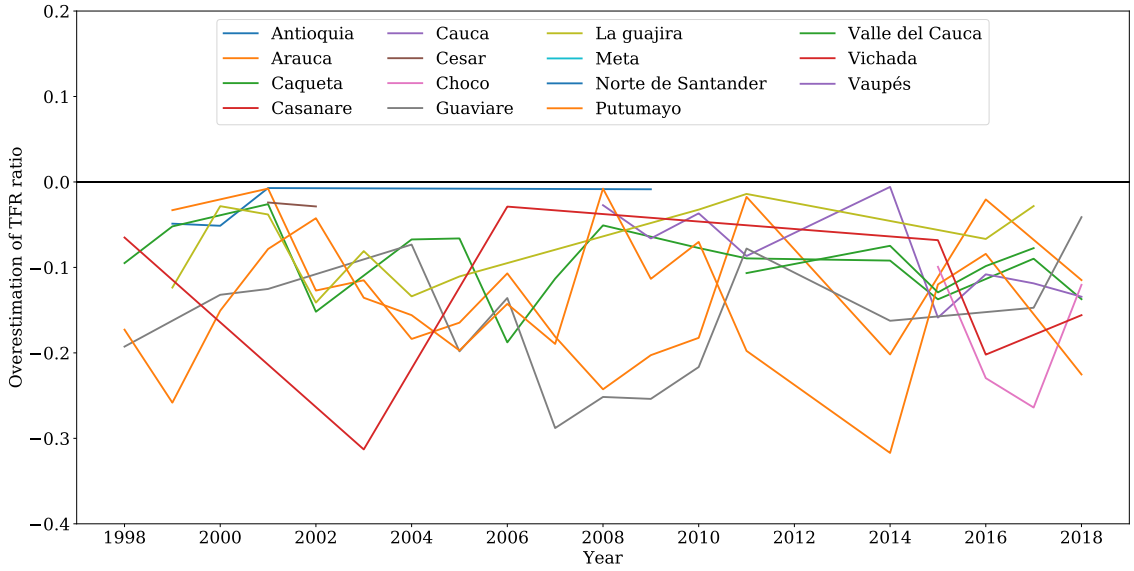

Figure S.2: This figure illustrates the potential impact of conflict-related undercoverage of men in the population. The estimate is the percentage-difference between the observed TFR ratio and the TFR ratio, that has re-weighted male exposures in conflict-affected areas. The weights are estimated following equation 15.

E Additional figures

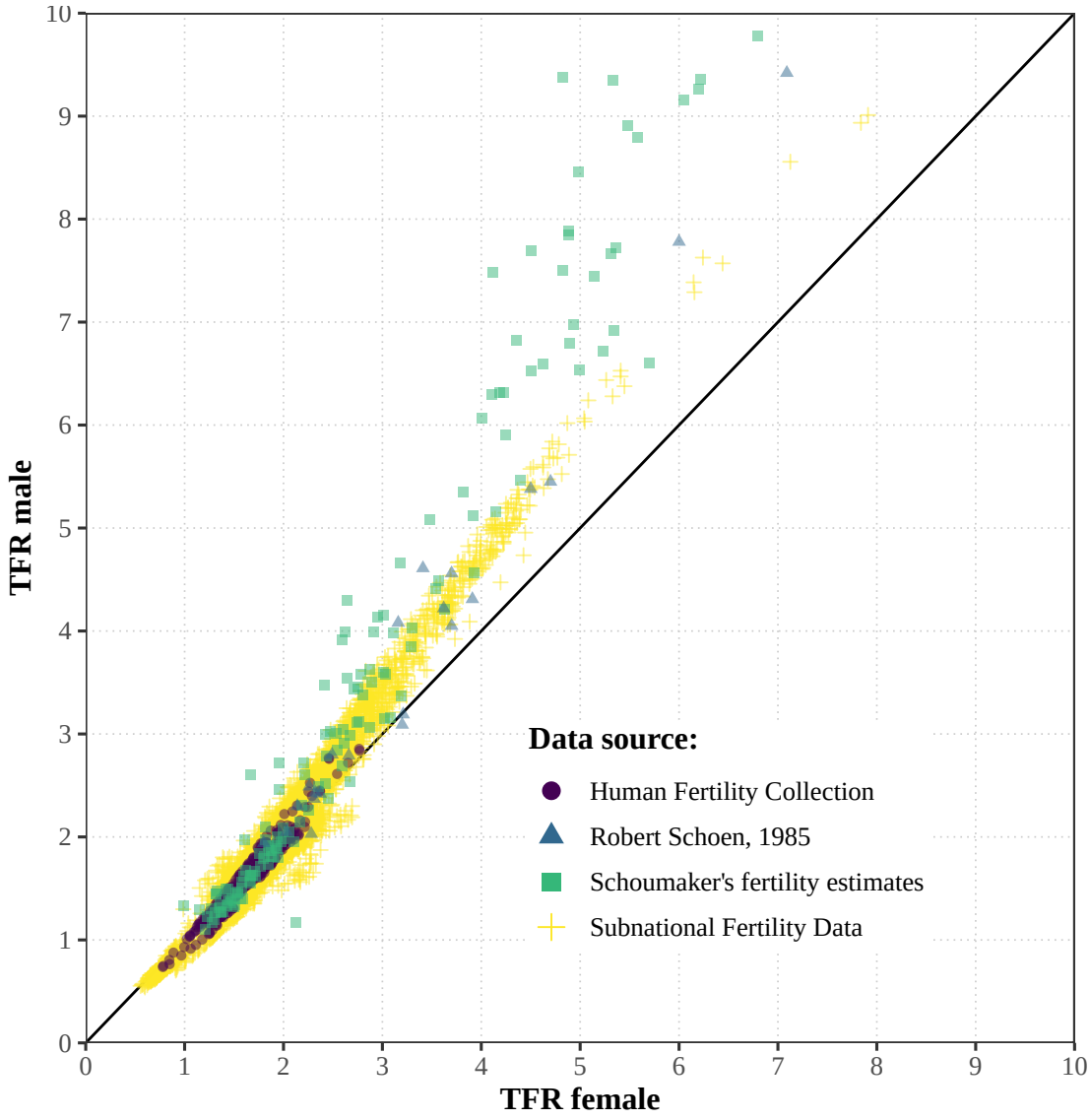

Figure S.3: This figure displays the distribution of region-year observations for the country and subnational data sources.

Source: The data on male fertility in the human fertility collection was created by Dudel and Klüsener (2021). Schoumaker's fertility estimates are obtained from Schoumaker (2019). Data was also obtained from Schoen (1985). The subnational fertility data was created by the author's.

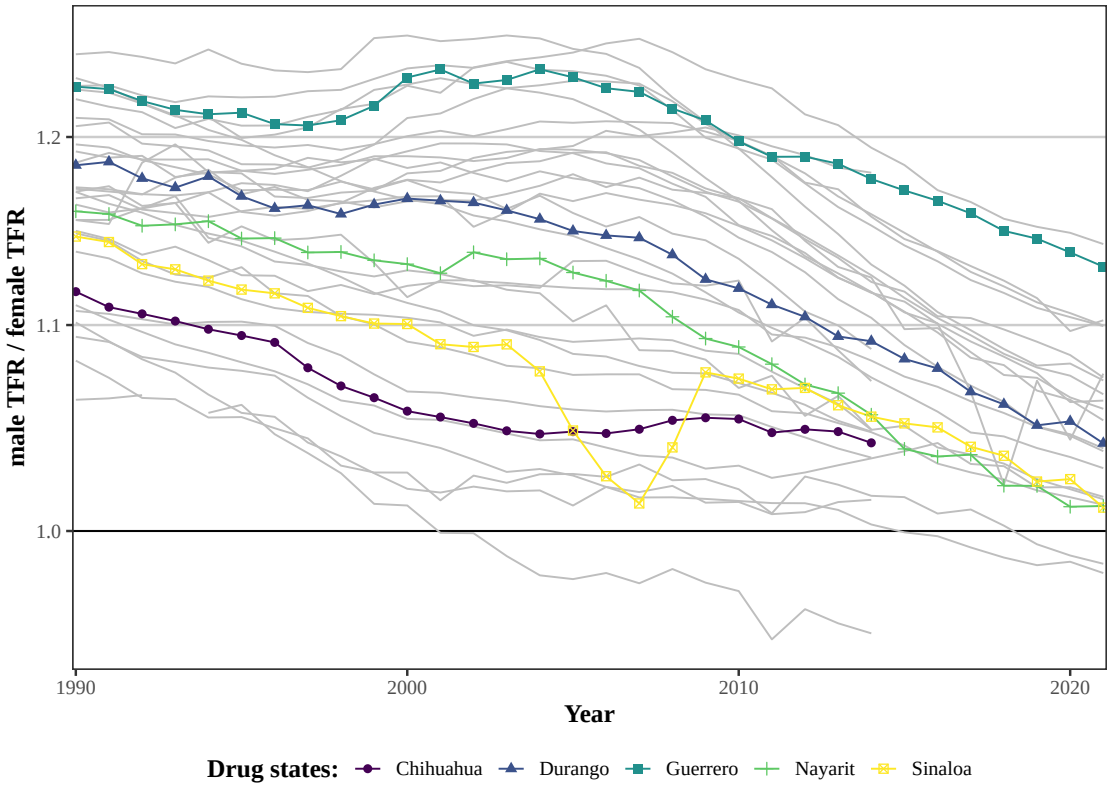

Figure S.4: This figure displays the time-trend in the male-female TFR ratios in the Mexican states. The coloured lines are the states most affected by the war on drugs according to Aburto et al. (2016), while the grey lines are the less and non-affected states.

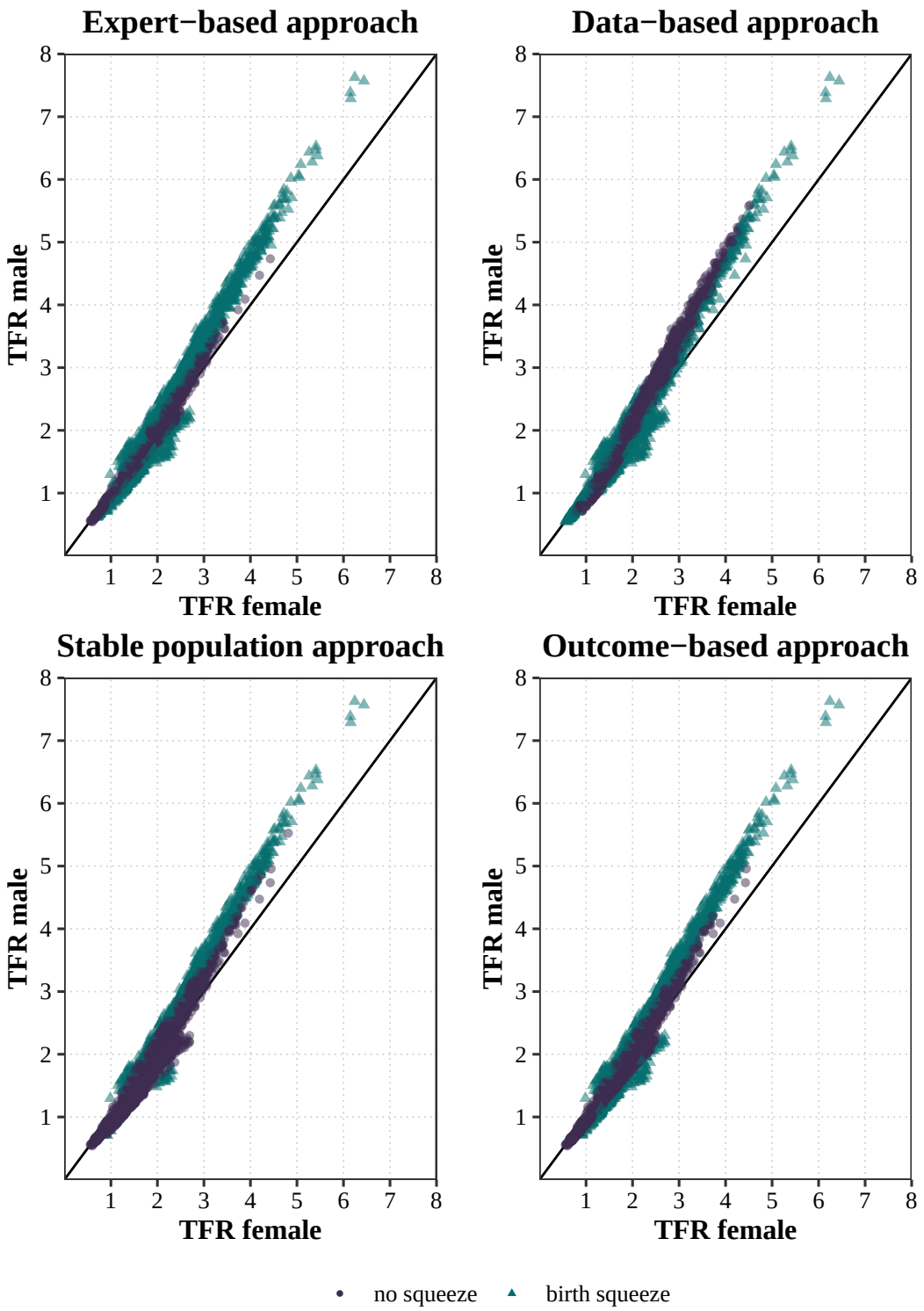

Figure S.5: The panels illustrate the occurrence of birth squeezes at the subnational level dependent on the measurement approach. Every point represents a region-year observation. The purple colour indicates male-female differences in the normal range, whereas yellow points indicate birth squeezes.

F Additional tables

Table S.5: Quantiles of the male to female TFR ratio across different data sets.

	Source	N	10%	50%	90%
1	Human Fertility Collection	516	0.920	0.968	1.037
2	Robert Schoen, 1985	22	0.995	1.098	1.296
3	Schoumaker's fertility estimates	163	0.938	1.159	1.623
4	Subnational Fertility Data	5,278	0.882	0.994	1.146

Source: Dudel and Klüsener (2021), Schoen (1985), and Schoumaker (2019).

Table S.6: Regression model underlying the impact approach. Logit-binomial regression of TFR ratio at age 30 on the probability remaining childless at age 40. The results are displayed as logit-coefficients.

	Males	Females
α	-1.136^{***}	-5.167^{***}
	(0.048)	(0.056)
β	0.347^{***}	4.089^{***}
	(0.0005)	(0.059)
Observations	782,021	747,063

*p<0.1; **p<0.05; ***p<0.01

Source: Authors' analysis of population register data from Statistics Finland.