The Role of Conflict in Sex Discrimination

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The Role of Conflict in Sex Discrimination: The Case of Missing Girls*

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ABSTRACT: Conflicts between ethnic groups can threaten group survival and exacerbate son preference in conformity with the traditional role of men as group defenders. We study the impact of the Armenia-Azerbaijan conflict over the Nagorno-Karabakh region on the subnational variation of sex ratios among children in Armenia that has one of the world's highest sex ratios at birth. Difference-in-differences analysis show that communities closer to the conflict region have higher sex ratio among children relative to the communities further away. The findings from household surveys show that fear of war is associated with a stronger son preference at the individual level.

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1 Introduction

A decades-long smoldering conflict between Armenia and Azerbaijan over the enclave of Nagorno-Karabakh (NK) erupted into a 44-day war in late September 2020, killing thousands in combat (Kramer, 2020). This was the worst fighting since the vicious ethnic war in the 1990s, which ended with a ceasefire agreement in 1994. The escalation of the unresolved conflict into a fierce war was not unexpected due to regular ceasefire breaches of various intensity on the front lines of the NK (De Waal, 2016a,b). Yet, the social consequences of the long-standing fear of war in the region are hardly understood.

In the ceasefire period of the NK conflict, Armenia and Azerbaijan experienced the world's highest skewed sex ratios at birth (World Bank, 2015), making the South Caucasus region one of the hot spots for "missing girls". Amartya Sen coined the term "missing girls" to describe females who would have been alive if their birth or survival had not been intentionally interrupted. Skewed sex ratios reflected in an increase of the natural ratio of 105 males for every 100 females imply missing girls in the population and a revealed preference for sons (Sen, 1990; Klasen, 1994; Klasen and Wink, 2003).

In this study, we postulate a link between the threat of war and preference for sons. Conceptually, the risk of war can revive group survival values, namely, preference for sons due to the traditional role of men as defenders. Within the group, boys will be valued more than girls, influencing son preference at the individual level. In particular, parents may respond to the risk of war by changing their reproductive behavior and have more sons than daughters. This choice will ensure the continuation of family lineage and influence the number of group defenders (males) in the future. At the population level, highly skewed sex ratios at birth will reflect this behavior, given low fertility levels.[1] We empirically investigate the role of the decades-long, simmering NK

[1] Duthé et al. (2012) associate highly skewed sex ratios in Armenia with low levels of fertility and access to modern (ultrasound) technology. Similarly, Jayachandran (2017) provides robust evidence for the relationship between low fertility levels and higher sex ratios at birth in India's case.

conflict in son bias and, subsequently, skewed sex ratios among children in Armenia.

The NK region lies within the internationally recognized borders of Azerbaijan, yet, historically, it has also been home to ethnic Armenians. During the post-1994 ceasefire period, ethnic Armenians populated and governed the unrecognized republic of NK (or the Republic of Artsakh), where armed forces of Armenia and NK jointly defended its de facto borders with Azerbaijan. Ceasefire clashes at the front lines were not uncommon, but in 2008 the first round of intense clashes took place. These skirmishes involved heavy artillery and arsenal, killing dozens in combat and increasing the likelihood of a revival of war (International Crisis Group, 2011). This increased threat of future war may have influenced the preference for sons in Armenia. Recent studies (Michael et al., 2013; Dudwick, 2015) and media articles (Jack, 2017; Moore, 2018) on Armenia's highly skewed sex ratios among children denote the role of the NK conflict in preference for sons. The below quote provides a perspective for such an explanation:

> *The headteacher of a school in Gavar [close-border town in Armenia], where the classes have more boys than girls, Araxia Vardanyan, says the impact of the war hangs over its people. Armenia is at war with Azerbaijan over Nagorno-Karabakh. "Our soldiers are killed on a daily basis. We need girls to reproduce. We need boys to defend the border." (Moore, 2018).*[2]

We study the social consequences of the first episodes of post-war ceasefire breaches in the front lines. Anecdotal evidence shows that such ceasefire breaches and skirmishes usually followed failed diplomatic negotiations between Armenia and Azerbaijan, threatening the status quo (Giragosian, 2014; Broers, 2015). The timing of these skirmishes was thus exogenous for the ordinary people in Armenia. Meanwhile,

[2]What the statement effectively implies is that the primary role of the males in Armenia is defending the homeland, whereas the primary role of the females is reproduction. If the perception is that soldiers are killed daily because the country is "at war", it could increase the preference for boys over girls, given that each woman can give birth more than once.

clashes along the NK front lines likely transmitted stronger fear of future war in Armenian communities closer to the NK region due to the spillover risks.

The case of Armenia provides an ideal context for ”all else equal” analysis. In particular, the country has a homogeneous population in terms of ethnicity (98.1 percent Armenians), language (97.9 percent Armenian) and religion (92.1 percent Apostolic Christians), and is small in size (population of 3,045,191) (The World Factbook, 2017). The ethnolinguistic and religious homogeneity among Armenians presumes similar group-level preferences and collective memory on critical historical events, such as conflicts and violence. Yet, homogeneity in ethnicity, religion and language does not presume within-group homogeneity in individual-level values (cultural homogeneity). For example, Desmet et al. (2017) use cross-country data from the World Value Surveys and show that within-group variation in values trumps between-group variation, which is also the case for Armenia. However, the authors also show that ethnic and cultural identity tend to overlap when it comes to values linked to national identity, which is essential in our context. One commonly shared view among the ethnic Armenians is that NK, which they call Artsakh, is part of their historical homeland and national identity (De Waal, 2003).

This study is based on several complementary datasets from Armenia. First, we use the three latest population censuses (1989, 2001 and 2011) of Armenia that include records of permanent residents by age group, gender, and municipality (settlements with more than 5000 people). The population census takes place every 10 years; therefore, we compute sex ratios for under 5-years-olds (under-5) using 0-4 and 5-9 age groups in each census, which results in municipality-level panel data of under-5 sex ratios, covering the period from 1984 to 2011. We then divide our sample of municipalities into treated and control groups. The treated group includes municipalities within less than the mean distance to the capital of the NK region, Stepanakert. All other municipalities belong to the control group.

We use a difference-in-difference model augmented with a full set of fixed effects and time trends to estimate the changes in under-5 sex ratios in the treated group, relative to the control group, from pre- to post-ceasefire periods. The municipality-level analysis shows that under-5 sex ratios increased by 6-10 points (depending on the estimation) in the treated group, relative to the control group, in the last period of our study. This period coincides with the first intense clashes that flared up along the NK front line. We test the robustness of this finding by altering the cut-off point for the treatment and using both travel and geographic distances as well as providing a placebo test based on the municipality distance to the 1988 earthquake epicenter in Armenia. Our main finding remains robust to these alterations.

Nonetheless, our census-based empirical setup does not eliminate the potential effect of some confounding factors which do not enter the regression due to data unavailability. Most significantly perhaps, there is scope for preference-based differential selection into migration, as well as for gender differences in under-5 child mortality rates between treated and control regions. To investigate if these are in fact at play, we employ a second dataset, the Demographic and Health Surveys (DHS) collected in Armenia in 2005 and 2010 that correspond to our study period. The DHS data includes respondent information on migration and preference for offspring sex, as well as birth records of children, including the incidence of mortality, sex of the child and the birth order. This information also helps us construct an alternative measure of revealed preference for sons based on the sex ratio of last births (SRLB), which picks up not only the sex-selective abortions but also continued childbearing until a son is born (Jayachandran, 2017). We define treated and control groups based on the geo-referenced information on DHS survey clusters.

The findings from the DHS-based analysis do not provide evidence for differential selection into migration between treated and control regions depending on the respondents' preference for sons, controlling for the desired fertility and other individual characteristics. Neither do we find evidence for gender differences in under-5 child mortality rates between treated and control groups. It is, therefore, unlikely that these factors confound our baseline inference. Furthermore, the results based on the use of SLRB as an alternative dependent variable for son preference, corroborate our finding based on under-5 sex ratios showing that in 2010 the treated group had a higher SRLB relative to the control group.

We then assess the validity of our hypothesized mechanism by exploring the link between individuals' fear of war and expressed son bias using a third dataset, the Caucasus Barometer, a nationally representative household survey, which has been conducted in Armenia since 2004. We use the 2010 wave as it features a question on son preference. We construct a survey measure for fear of war based on the respondents' ranking of peace or territorial integrity as the most pressing issue in the country. The estimation results from the most extensive model suggest that having a fear of war is associated with 10 percentage points higher likelihood of having a son preference. This finding is robust to applying matching and partial identification approaches to mitigate the unobserved heterogeneity problems.

Our study provides insights into the literature that links armed conflict with changes in sex ratios among children based on biological mechanisms. One of the well-known biological explanations for changes in the sex ratios at birth goes back to the study by Trivers and Willard (1973), which postulates that through the evolutionary process, women have developed the ability to *unconsciously* adjust their offspring's sex based on the maximum number of potential grandchildren. It suggests that stress hormones, due to poor conditions such as wars, lead to the termination of male fetuses differentially from female fetuses, whereby males are less likely to be born (James, 2009; James and Valentine, 2014). Once born, they are more likely to survive due to in-utero selection (scarring and culling effects). These mechanisms suggest a decrease in the male-to-female sex ratios at birth and higher child mortality rates for girls.[3] Valente (2015) finds evidence for culling and scarring effects resulting from the in-utero exposure to the civil war in Nepal from 1996 to 2006. Similarly, Dagnelie et al. (2018) find that

[3] Kemper (1994) and James (1997) find a higher likelihood of male births after World War II, which they argue is due to a 'returned soldier' effect, where coital frequency and the phase of menstrual cycle play a role. Empirical evidence from the civil war in Tajikistan also shows that during the civil war the sex ratios at birth increased from 104 to 106 but returned to the baseline values shortly after (Hohmann et al., 2010).

in-utero exposure to the civil war in the Democratic Republic of Congo significantly reduced the number of live male births, supporting the Trivers-Willard hypothesis.

We find some descriptive support for the Trivers-Willard hypothesis within the long period of our study. In particular, in the period of all-out war and shortly after (1992 to 1996) we observe a short-lived, small decrease in the under-5 sex ratios in the treated region in 1996. However, in the post-1996 ceasefire period, we observe a large increase in the under-5 sex ratios from 106 in 1996 to 115 in 2011 in the same region. Coupled with qualitative evidence on sex-selective abortions, our findings on under-5 sex ratios and child mortality suggest that biological factors are not the driving forces behind the increased sex ratios among children in Armenia.[4]

Our study contributes to the growing literature on the behavioral consequences of conflict on fertility (Buvinic et al., 2013), revealed son preference and child mortality. Sng and Zhong (2018) show that regions with past historical conflict within China have higher juvenile sex ratios contemporaneously.[5] A recent study by Tapsoba (2020) suggests that the fear of conflict can trigger behavioral responses and affect child mortality rates before (or without) actual exposure to violent events. Our study provides comprehensive evidence on the relationship between a threat of war and revealed preference for sons, triggered by the NK conflict.

This study is also related to the broader literature on the causes and consequences of son preference and missing girls. This literature has established that the reason behind missing girls is the stronger preference for sons, evidenced by attitudes (Westoff and Rindfuss, 1974; Haughton and Haughton, 1998; Kureishi and Wakabayashi, 2011), cultural norms (Li et al., 2000; Das Gupta et al., 2003; Fogarty and Feldman, 2011) and behavior (Park, 1983; Sen, 1990; Coale, 1991; Klasen, 1994; Ebenstein, 2010; Li et al., 2011). Patrilineal kinship systems, patrilocality, old-age support as well as lineage

[4] A 2012 study on Armenia provides qualitative evidence on the link between highly skewed sex ratios at birth and sex-selective abortions in the country (Abrahamyan et al., 2012).

[5] A large share of India's missing girls comes from northern part of the country (Das Gupta and Shuzhuo, 1999), which has been home to several conflicts (i.e., Sikh-Hindu clashes in Punjab and Haryana as well as Kashmir insurgency).

and inheritance through males play an important role in explaining the son preference (Rose, 1999; Clark, 2000; Das Gupta et al., 2003; Ebenstein, 2014). This preference can also be impacted by resource constraints, such as additional costs associated with daughters and benefits associated with sons (Ben-Porath and Welch, 1976; Rosenzweig and Schultz, 1982), and a perception that males are the most productive sex (Ahn, 1995). Moreover, son preference could be stronger among upper-class parents because well-off males have the highest probability of successful mating in the marriage market (Edlund, 1999). Son bias can also become self-perpetuating and endanger female survival due to women's relatively low earnings potential (Qian, 2008) and deteriorating bargaining power in the household (Klasen, 1998). Broadly, son preference correlates with deep-rooted gender inequalities in the society (Branisa et al., 2014), which can lead to unequal economic and social opportunities for half of the population (Sen, 1989). In the long run, the gender bias at the country level can translate into loss in economic growth and development (Lagerlöf, 2003). We contribute to the literature on missing girls by offering an additional explanation for son preference that is deeply embedded in the traditional gender role of males as group defenders.

2 Conceptual framework

Fundamental choices and preferences related to the survival of offspring, such as life and death, reproduction, and group survival can be altered under certain conditions such as famines, combats and other disasters as modeled by McDermott et al. (2008). These survival (loss domain) choices are regarded as risk-seeking and can be explained using the prospect theory (Kahneman and Tversky, 1984). Empirically, Voors et al. (2012) show that individuals who have been exposed to violent conflicts are more risk-seeking and altruistic, confirming the predictions made by McDermott et al. (2008). Beyond the physical exposure to violent events, the mere expectation of being exposed to such events can also alter the behavior of individuals (Tapsoba, 2020).

Evidence from experimental economics shows that in the presence of an external threat to a group's survival, individual preferences derive from in-group preferences, which lead to a higher probability of group-conforming behavior among the in-group individuals (Weisel and Zultan, 2016). In contrast, when the threat is at an individual level, the individual preferences are not overwritten by in-group preferences. Hence, one can deduce that a perceived external threat to group survival, such as war, can lead to more group-conforming and traditional choices, such as having a preference for sons.

Depending on the context, preference for sons can be seen both as risk-seeking and as a risk-averse preference. If the group survival is not perceived to be under threat, but parents prefer sons for old-age support and family line continuation, it is a risk-averse preference. However, suppose there is a perceived threat to group survival. In that case, preference for sons can be a risky choice. The sons are more likely not to survive a possible future war as defenders of the group (as has historically been the case), and the family line may discontinue. But since not only family but also group survival is at stake, individual choices are likely to surrender to group pressures for sons.[6]

We illustrate our argument linking the threat of war to son preference and skewed sex ratios among children (skewed SRC) in Figure 1. First, the threat of war can enhance individual-level survival values behind son preference, such as family lineage and inheritance via males. Second, the threat of war can trigger group-level survival values such as a defense of the homeland, leading to a purposeful increase in male births. In this process, the group-specific survival values may actuate son preference for their traditional role as group defenders and, concurrently, enhance individual-level survival values congruent with other traditional gender roles, such as males as breadwinners and guarantors of income. As a result, the expectation of war can exacerbate the overall preference for sons. Coupled with fertility decline, which adds con-

[6]One might argue that the choice over a male offspring is not rational from an individual (parent's) perspective when they expect a war. However, an increased number of males is more likely to ensure the group's survival - a perception shared among the group members due to past conflicts. This shared perception or norm can put pressure on individuals to opt for male offspring.

straint over individual choices, the increased preference for male offspring can manifest in the skewed sex ratios among children or live births. Although parents cannot predict when the active war will erupt, they can change their behavior to influence the number of males in the future by consciously manipulating their offspring's sex composition. This conscious manipulation can take the form of sex-selective abortions or the continuation of childbearing until a son is born.

[Figure 1 about here.]

In our empirical investigation of the relationship between the risk of war and (revealed) preference for sons, we draw on the argument on group survival values whereby males have traditionally taken the role of defenders. However, our argument is compatible with other potential mechanisms at play. For example, the risk of war may decrease the labor supply of males in the non-military sector, increasing their demand and, hence, their wages in the labor market. This possible increase in male wages relative to female wages can lead to a revival of other traditional motives for son preference, such as financial or old-age support.

3 Context

Under-5 sex ratios

Several countries in the world have experienced significant increases in the under-5 sex ratios since the 1980s. Figure 2 depicts these trends for selected countries with 'unnatural' sex ratios among children, based on the World Bank (2015) data. In the mid-1990s, three countries in the South Caucasus (Armenia, Azerbaijan and Georgia) and a few countries in southeast Europe exhibited sharp increases in under-5 sex ratios. This timing coincided with the period when ultrasound technologies became widely available in the region (Duthé et al., 2012; UNFPA, 2015). The under-5 sex ratios became exceptionally high in Armenia and Azerbaijan, reaching China's and surpassing India's levels. Estimates suggest that Armenia, with roughly 3 million population, may lose about 93,000 women by 2060 if the trend continues (Moore, 2018).

[Figure 2 about here.]

In addition to the wide availability of ultrasound technology, the region experienced inter-ethnic conflicts, including the Bosnian war (1992-1995), the Albanian civil war (1997) and the Kosovo war (1998-1999) in southeast Europe. Conflicts in the South Caucasus include the war in Abkhazia (1998), the Adjaria crisis (2004), the Kodori crisis (2006), the Russo-Georgian war (2008), and lastly, the NK war between Armenia and Azerbaijan (1992-1994). With many inter-ethnic conflicts remaining unresolved, the region has seen increases in militarization in anticipation of a revival of war. In particular, the NK conflict gave rise to the most militarized zone in wider Europe - the de facto *line of contact* between ethnic Armenian and Azerbaijani armed forces (De Waal, 2016a).

Abrahamyan et al. (2012), Michael et al. (2013) and Dudwick (2015) show that sex-selective abortions in Armenia and in the wider South Caucasus have become a common method for controlling fertility levels and achieving the desired offspring sex. Since the late 1980s, the fertility levels have been declining in Armenia (World Bank, 2015). According to the data from the DHS Armenia, since the 2000s, the total fertility rate has been below the replacement level (1.7). At the same time, child mortality rates have decreased significantly in Armenia. In 2010, the infant and child mortality rates were 13 and 16 deaths per 1000 live births, computed for the preceding 5-year period (DHS, 2012).

Background on conflict

The NK conflict dates back to 1988, although the dispute between Armenians and Azerbaijanis over this region has its roots in the early 20th century. In 1920, Armenia and Azerbaijan became part of the Soviet Union, which resulted in the allocation of the Armenian-majority NK to Soviet Azerbaijan (De Waal, 2003).[7] In 1988, in the con-

[7]In 1989, before the outbreak of the ethnic expulsions and the NK war, about 72 percent of the population in NK were Armenians and around 22 percent were Azerbaijanis (USSR, 1991).

text of liberalization reforms in the Soviet Union such as *Glastnost* and *Perestroika*, the Armenians of NK campaigned to join Armenia in 1988 and voted to leave Azerbaijan in an independence referendum in 1991. Neither the union with Armenia nor the independence referendum was officially recognized by Baku or Moscow.

Mutual fear led Armenians and Azerbaijanis into inter-ethnic violence: the Armenians of NK feared to be "swallowed up or driven from their homeland by Azerbaijan; the Azerbaijanis [feared] that Armenian-led Karabakh constituted a fifth column threatening the integrity of their republic" (De Waal, 2016b). The (Soviet) Armenian government supported the Karabakh movement while the inter-ethnic violence spilled over into Armenian regions with Azerbaijani minorities and Azerbaijani cities with Armenian minorities (De Waal, 2003; Freizer, 2014). In the fall of 1991, the Soviet Union collapsed and Armenia and Azerbaijan became independent states. The inter-ethnic violence escalated into a full-scale war by the winter of 1992.

The United Nations High Commissioner for Refugees (UNHCR) reports that before and during the all-out war in 1992, around 300,000 Armenians living in Soviet Azerbaijan fled to Armenia, while around 150,000 Azerbaijanis living in Soviet Armenia fled to Azerbaijan (UNHCR, 1994). Additionally, the massive displacement of Azerbaijanis from NK to other parts of Azerbaijan led to one of the worst refugee and internally displaced person (IDP) crises in the world. Various sources estimate the total number of casualties from civil and combat violence to be around 17 to 25 thousand people (UNHCR, 1994; De Waal, 2003, 2010).

Figure 3 presents the map of the NK conflict region, illustrating the de facto Armenian-controlled territories and the *line of contact*, e.g., the front lines, as of 2010. As the map shows, as a result of the war (1992-1994), ethnic Armenians won control of the former autonomous NK region (majority Armenian) and other nearby Azerbaijani regions (majority Azerbaijani), connecting the NK to Armenia by land.[8]

[Figure 3 about here.]

[8] Figure A1 in the Online Appendix shows the time-line of the main events related to the NK conflict until 2011.

Thus, while NK and the surrounding regions de jure remained part of Azerbaijan, de facto, these regions were populated by ethnic Armenians and controlled by the joint Armenian and NK armed forces until 2020. The prolonging of this status-quo led to the de-facto union of NK and surrounding territories with the Armenian state.

Ceasefire period

The ceasefire agreement in May 1994 put a hold on the full-scale war and set the stage for diplomatic negotiations between Armenia and Azerbaijan. The Minsk Group of the Organization for Security and Cooperation in Europe (OSCE), co-chaired by Russia, France and the United States, became the official mediator of the NK conflict. However, the two parties of the conflict, Armenia and Azerbaijan, would not agree on conflict resolution protocols for many years to come. The ceasefire remained fragile. The clashes between ethnic Armenian and Azerbaijani armed forces at the front lines increased in frequency and intensity, rendering it likely that war would break out again (International Crisis Group, 2011).

Anecdotal evidence suggests that inter-state diplomatic debacles were the main drivers of the ceasefire breaches as intense violations followed the important but failed diplomatic talks between Armenian and Azerbaijani leaders, conducted confidentially. The collapsed negotiations triggered war rhetoric and ceasefire breaches at the NK front lines, threatening the status quo (Giragosian, 2014; Broers, 2015).[9]

In Figure 4, we depict the number of battle-related deaths from 1991 to 2011 based on the Uppsala Conflict Data Program (UCDP) (Pettersson and Eck, 2018; Sundberg and Melander, 2013). The figure shows that the violence truly scaled down after the ceasefire agreement in 1994. During the full-blown war from 1992 to 1994, the number of battle-related deaths exceeded 9000. The periodic ceasefire breaches in the post-1994 period led to dozens of battle-related deaths. Ceasefire violations took place in

[9] Up to 2011, there were three significant diplomatic efforts for the NK conflict resolution, where the OSCE Minsk Group developed conflict resolution protocols (Hopmann, 2015). The governments of both Armenia and Azerbaijan rejected these conflict resolution protocols at different times.

the Post1 (1997 to 2001), Post2 (2002 to 2006) and Post3 (2007-2011) periods, signaling increasing threat of war over time.

[Figure 4 about here.]

There is a lack of reliable data for overtime comparison of fear of war before the last period of our study (Post3). Nonetheless, it is not unlikely that the unprecedented use of heavy military weapons during Post3 clashes may have triggered an increased fear of war. In particular, the 2008 Martakert skirmishes in Post3 were the first intense ceasefire violations since 1994, which, unlike the previous episodes, involved heavy military artillery and arsenal (International Crisis Group, 2009; BBC, 2008), alongside the similar number of battle-related deaths.[10] The timing of Post3 skirmishes followed the failure to reach mutual acceptance on the implementation of the latest resolution protocol, the Madrid Principles, which the OSCE Minsk Group presented to both sides in Madrid, Spain in 2007. In 2009, the International Crisis Group warned on the difficulty of sustaining the NK status quo (International Crisis Group, 2009). Their 2011 report called for an urgent action to prevent another war between Armenia and Azerbaijan (International Crisis Group, 2011, p.1): "An arms race, escalating front-line clashes, vitriolic war rhetoric and a virtual breakdown in peace talks are increasing the chance Armenia and Azerbaijan will go back to war over Nagorno-Karabakh. Preventing this is urgent." [11]

Males as defenders

In 2011, the United Nations Population Fund (UNFPA) in Armenia sponsored a collection of qualitative surveys concerning the highly skewed sex ratios, son preference

[10]Martakert is a town in NK, located within 50 kilometers northeast of Stepanakert, close to the front lines.

[11]The later ceasefire violations such as the four-day war in April 2016 and the 44-day war in 2020 were more violent, involving modern military equipment and artillery. The latter led to unprecedented human, infrastructure, and military losses since the first NK war and changed the status quo of the conflict.

and sex-selective abortions in the country (Abrahamyan et al., 2012). The survey interviewed over 2760 women across all regions (provinces) in Armenia.

The survey results reveal that in the social circle of women (e.g., friends, relatives) the share of people who have son preference (60 percent) is about ten times higher than the share of people who have a daughter preference (6 percent). Regarding own family preferences, the survey results show that the share of families with a son preference (44 percent) is about six times greater than that of families with a daughter preference (7 percent) (Abrahamyan et al., 2012, p.28).

Using the data from the UNFPA qualitative surveys, in Figure 5, we present the cited reasons for son preference, comparing the survey results of the Armenian provinces that share a border with NK and Azerbaijan (*border*) to the other provinces that do not (*no border*). The results show that the top three cited reasons for son preference are the continuation of family lineage through sons, that sons are seen as inheritors of property (though by law, women and men have equal inheritance rights in Armenia) as well as guarantors of income or financial support.

[Figure 5 about here.]

The literature has long established these particular reasons for son preference in East and Southeast Asian countries, and Armenia is not particularly different in that sense. Yet, the fourth most cited reason for son preference in Armenia is that "boys are defenders of homeland", and a slightly higher share of women in border regions cite this reason relative to those in the non-border regions. As Dudwick (2015) points out, the NK conflict has indirectly created an environment in Armenia where males are valued as the defenders of the country. This qualitative evidence lends support to this statement and to our conceptual framework linking the threat of war to a stronger preference for sons allowing us to focus on the role of males as defenders. Nonetheless, other mechanisms, such as expectations of financial support from sons or male-biased inheritance practices, are also important. We test for the potential role of these mechanisms in the individual-level analysis to the extent possible.

4 Threat of war and under-5 sex ratios: Municipality-level analysis

Data and descriptive analysis

We use the last three population censuses of Armenia conducted in 1989, 2001 and 2011, which provide information on population numbers disaggregated by gender, age group, and municipality.[12] The 1989 census is the last one collected in the Soviet Socialist Republic of Armenia, while the 2001 and 2011 ones are the first (and currently only) two censuses collected in the independent Republic of Armenia.

We compute under-5 sex ratios by municipality and period (year) using the 0 to 4 and 5 to 9 age cohorts in each population census (1989, 2001 and 2011), given that the census runs every 10 years. For example, in the case of the 2011 census, we draw on the 5 to 9 age cohort to compute the under-5 sex ratios as of 2006, where the corresponding period of live births is the period from 2002 to 2006. Similarly, we calculate the under-5 sex ratios as of 2011 using the 0 to 4 age group, where the corresponding period of live births is the period from 2007 to 2011. Table A1 in the Online Appendix shows the concordance of periods, censuses and cohorts that we used. We based our computations on the permanent residence (*de jure*) records to avoid concerns related to temporary migration.[13] The sample includes 76 geo-referenced municipalities observed over six points in time.

During the NK war, the battle locations were not limited to the NK region but spilled over into the Armenian regions that share a state border with Azerbaijan (see Figure A2 in the Online Appendix). A nationwide military mobilization ensured broad recruitment of soldiers across all Armenian provinces. But those provinces bordering

[12] This data was obtained specifically for this project from Armenia's National Statistical Service. The publicly available census data and the Integrated Public Use Microdata Series (IPUMS-International) census records for Armenia are disaggregated by 11 provinces but not at the municipality level.

[13] In the post-war period, Armenia has been characterized by high rates of emigration, largely for temporary purposes, and biased towards adult males with seasonal/temporary jobs in Russia (Collyer, 2015).

the conflict region had the highest per capita number of soldiers killed in combat (see Figure A3 the Online Appendix).[14] Therefore, we expect that the perceived threat of future war varies with the distance to NK.

Therefore, we use municipality distance to the capital of NK, Stepanakert (shown on the map in Figure 3), as a proxy measure for the perceived threat of war.[15] Stepanakert is important as an epicenter because the war in NK started with heavy missile attacks over its capital. Most intense battles in the NK war (1992-1994) and the recent locations of ceasefire breaches are in close proximity of Stepanakert (see maps in Figures A2 and A4 in the Online Appendix). A possible adversary advancement beyond Stepanakert would move the front lines closer to the Armenian state borders.

We treat the timing of ceasefire breaches as municipality-level shocks relative to the threat of war and divide our sample of municipalities into treated and control groups, taking the mean distance to Stepanakert as the cut-off point. We assign municipalities to the treated group if their distance to the epicenter is less than the cut-off; otherwise, we assign them to the control group. In the robustness tests, we vary the cut-off point using the median and 25th percentile of distance to Stepanakert.

The map in Figure 6 shows that the treated municipalities in our sample are mostly located in the south of Armenia and are much closer to the NK region and Soviet Azerbaijan borders. The proximity of these municipalities to the conflict region also implies that the refugees would first arrive in these regions. The annual data from the UNHCR show that the stock of refugees in Armenia peaked in 1993 and 2000 and declined sharply after that (see Figure A5 in the Online Appendix). If the norms and traditions of refugees in the treated regions have been driving the skewed sex ratios under-5, we should expect to observe these differences by 2005, given that refugee flows do not increase thereafter. We check below whether that is the case.

[14] As a result of the nation-wide military mobilization, one-third of ethnic Armenians killed in combat (KIC) were from Armenia (Arstamyan, 2009). We geo-referenced the birthplaces (e.g., village, town or city) of the KIC from Armenia, shown on the map in Figure A2 in the Online Appendix.

[15] The conflict literature (Voors et al., 2012; Verwimp and Van Bavel, 2013) commonly uses the distance measure as a proxy for exposure to violence.

[Figure 6 about here.]

We plot the trends in under-5 sex ratios in treated and control groups without any controls in Figure 7. The under-5 sex ratios were at the natural levels (around 104) in the first period of our study (Pre2, 1980-1984) in both treated and control groups. In the second period of our study (Pre1, 1985-1989), the under-5 sex ratios increased by 5 points in the treated group but they hardly changed in the control group. It is plausible that the fear of inter-ethnic violence in the pre-war period affected the slight increase in the under-5 sex ratios in the treated group before the war. During the all-out war and shortly after (1992-1996), the under-5 sex ratios decreased in the treated group, possibly due to biological factors, such as an increase in stress hormones. Yet, in the control group, the under-5 sex ratios slightly increased during the all-out war. This could have led to the contraction of the gap in under-5 sex ratios between the treated and control groups. In the period from 1996 to 2006, we observe parallel trends in under-5 sex ratios for treated and control groups, rendering the differential impact of refugees implausible. In 2011, the under-5 sex ratios increased by more than 10 points in the treated group relative to its pre-war levels, which may be due to the stronger fear of war in the region as the corresponding period includes the 2008 Martakert skirmishes. In contrast, the 2011 under-5 sex ratios decreased in the control group.[16]

[Figure 7 about here.]

Next, we check the difference in pre-trends between treatment and control groups by running a regression with a fully flexible specification, where the treatment variable interacts with each period dummy. This specification does not include any controls and excludes outliers - observations with less than 100 births. Figure 8 presents the results at the 95 percent confidence interval. In the upper graph, the cut-off point for the treated group is the mean distance, while in the lower graph, the cut-off point is the

[16] Biological factors, such as stress hormones, during the 2008 mass civil unrests in the capital city, Yerevan, may have played a role in the decline of the under-5 sex ratios in the control group, which includes Yerevan as well as the second and third largest cities of Armenia.

25th percentile distance to Stepanakert. First, the upper graph shows no statistically significant difference in pre-trends in the under-5 sex ratios between the treated and control groups before Post3. In the Post3 period, the difference between the treated and control groups is about 11 points, statistically significant at the 5 percent level. Second, the lower graph in Figure 8 shows a similar pattern but with larger deviations in magnitude. In particular, we observe largely a statistically insignificant and small difference in trends before Post3. In the Post3 period, the difference between the treated and control groups becomes quite large (14 points), statistically significant at the 5 percent level.

[Figure 8 about here.]

The differences observed in under-5 sex ratios between treated and control groups in the census-based analysis may be due to selective migration, where son preference may differ between those who migrate and those who stay behind. Also, gender differences in under-5 mortality rates could be a concern if these differ between treated and control groups. Moreover, there might be some concerns over the use of under-5 sex ratios computed retrospectively based on the census data as a measure for revealed son preference.

Therefore, we also use geo-referenced data from the DHS Program collected in 2005 and 2010 in Armenia, which helps us mitigate the limitations of the census data and provide a series of robustness checks featuring the use of alternative measures of son preference and addressing potential issues associated with selective migration and under-5 mortality rates. The limitation of the DHS is that the data does not cover pre-war periods.

The 2005 DHS wave provides information on migration, namely the number of years the respondent has lived in the current place of residence (village, town or city).[17] We exploit this information to check whether migration rates between treated and control regions vary for men and women based on their preferences for sons as

[17] Information on migration is not available in the 2010 DHS wave.

well as the (ideal) number of children, education, wealth group and employment status. We use both 2005 and 2010 DHS waves to explore the scope for concerns related to gender differences in under-5 mortality rates and to construct an alternative measure for son preference, sex ratios at last birth, using the DHS birth records in the 5 years preceding the survey year.

Based on around 300 geo-referenced DHS clusters in Armenia, we construct treated and control groups by calculating the distance from each DHS cluster to the conflict epicenter, Stepanakert, and taking the median as the cut-off point, similar to the approach in the census data analysis.[18] Clusters that are closer than the median to the conflict epicenter are in the treated group; otherwise, they are in the control group.

We present descriptive statistics based on the 2005 DHS Armenia in Table A2 in the Online Appendix by treated and control groups and the gender of respondents. The sample means to show that the share of respondents who changed the place of residence in the last five years (migration) is very low among men (0.8 percent) and slightly higher in the control group (1 percent) relative to the treatment group (0.6 percent). Women are more likely to have changed their place of residence in the past 5 years (on average 6.5 percent). Son preference (expressed desire for more boys than girls) is around 3 percentage points higher in the treated group for both men and women, reaching 40 percent for men and 22 percent for women. Actual fertility is slightly higher in the treated group while the desired fertility hardly differs between the treated and control groups (2.6 for women and 2.8 for men). Slightly more men are employed in the control group (50 percent) relative to the treated group (45 percent). In the treated group marginally more women are employed (31 percent) relative to the control group (25 percent) but they have fewer years of education (8.5) compared to women in the control group (9.4). In terms of wealth, on average, both groups across genders sit around the middle-income category. The average age for women and men in the sample is 31 years old.

[18] The results are similar when we use the mean as the cut-off point. We prefer the median in this case as the number of clusters in the treated group is much smaller when we take the mean distance.

Estimation strategy

In our main estimations, we use the census-based panel data and explore the effects of the threat of war on under-5 sex ratios in Armenia by employing a difference-in-difference (DiD) technique with a full set of fixed effects and time trends. The DiD estimations include three binary time variables for the post-war periods (Post1, Post2 and Post3), as shown in the fully flexible specification (Figure 8). We compute the under-5 sex ratio in the reference period by taking the average value for the years from 1984 to 1996.[19] Qualitative evidence shows that sex-selective abortions are important in the context of Armenia (Abrahamyan et al., 2012; DHS, 2012), and our analysis of the under-5 sex ratios would ideally control for access to ultrasound technology. In the absence of a direct measure of such access, we proxy for it using the distance to the capital of Armenia, Yerevan.[20]

We estimate a fixed-effects panel data regression, which takes the following form:

$$SRC_{ct} = \alpha + \sum_{t=1}^{3} \gamma_t Post_t + \sum_{t=1}^{3} \theta_t (Treat_c \times Post_t) + \sum_{t=1}^{3} \zeta_t (CloseCapital_c \times Post_t) + \\ + \mu_c + T_c + \eta_{rt} + \epsilon_{ct} \quad (1)$$

where SRC_{ct} is a continuous variable denoting male over female under-5 sex ratios in municipality c in time t. α is the constant term that denotes the average SRC_{ct} when $t = 0$. $Treat_c$ equals 1 if the municipality c is in the treated group; otherwise it is in the control group and equals 0. $\sum_{t=1}^{3} \gamma_t Post_t$ denote the three binary variables that correspond to the post-periods shown in Figure 8 (Post1, Post2 and Post3). The interaction

[19] Taking the average helps normalize the distribution of under-5 sex ratios in the pre-period and address possible serial correlation issues at the municipality level (Bertrand et al., 2004).

[20] Anecdotal evidence suggests that ultrasound technology first became available in Yerevan and then spread towards the adjacent regions over time. For example, a newsletter from the largest Diaspora Armenian aid organization (AGBU) shows that in the period from 1996 to 2010, the top medical and research institutions in Yerevan were the first recipients of the donated ultrasound technologies and accessories (AGBU, 2010).

terms $\sum_{t=1}^{3} \theta_t (Treat_c \times Post_t)$ estimate the average effects of ceasefire breaches on the municipalities in the treated group relative to those in the control group for each $Post_t$ period relative to the reference period.

$\sum_{t=1}^{3} \zeta_t (CloseCapital_c \times Post_t)$ are interaction terms between the time variables, $Post_t$ and $CloseCapital_c$. *CloseCapital* equals 1 if the distance between the municipality and Yerevan is less than the sample mean and 0 otherwise. These terms control for municipality-specific time-varying factors that correlate with the degree of remoteness of the municipality, such as access to labor and product markets, better education and health facilities, internal migration to the capital city, etc. Besides controlling for a number of unobserved municipality-level confounders over time, these variables also tease out the effect of municipality-level differences in access to ultrasound technology. The inclusion of these terms, therefore, strengthens the identification in our model.

The term μ_c denotes municipality fixed effects and controls for constant municipality-specific unobservable factors such altitude, mode of development (agricultural or industrial), type of place (urban or rural), culture, social norms, and the like. The term T_c denotes the municipality-level linear time trend, which is implemented as an interaction term between the municipality and a time trend. This term controls for the time-varying, linear, municipality-specific effects such as migration, fertility and employment trends for each municipality. The inclusion of this term also relaxes the common trend assumption of the DiD setting. Nonetheless, this term does not capture time-varying and non-linear effects, such as other confounding shocks that may have coincided with the ceasefire breaches.[21] Therefore, we also include the term η_{rt} to capture other within-country regional shocks (non-linear effects), which is an interaction term between three macro-regions of Armenia and the binary time, $Post_t$, variables.[22] The map in Figure A3 in the Online Appendix indicates these macro-regions. ϵ_{ct} is the

[21] For example, in 2008, Armenia held national elections, followed by mass protests and violence against civilians in Yerevan. State violence against its population could have affected the under-5 sex ratios.

[22] We omit $Treat_c$, $CloseCapital_c$ and the macro-regions as separate terms in equation 1 as these are perfectly collinear with the municipality fixed effects.

error term clustered at the municipality level.

We make use of travel distances between two points since they absorb altitude differences along the path. However, changes in the roads in the post-war period may not be exogenous. Therefore, as an alternative approach, we also use geographic distances, which measure the length of the shortest path between two points along the surface of a mathematical model of the earth, based on the Vincenty (1975) formula in geodesy. It is independent of changes in roads and infrastructure.

We use the mean distance as the cut-off point, which is arbitrary. Therefore, we also vary the cut-off for treatment assignment, using the median and 25th percentile of the distance. Additionally, we run a placebo test measuring the municipality distance to the 1988 earthquake epicenter, located 30 kilometers from the second-largest Armenian city, Gyumri. The map in Figure 6 indicates the location of the earthquake's epicenter.[23] As a supplementary analysis, we also test the validity of distance as a continuous variable in a panel fixed effects regression.

The model in equation 1 does not directly include municipality-level covariates due to data limitations, but the model's high saturation with fixed effects and time trends addresses the problem of confounders to a large extent. Nonetheless, concerns related to the gender-specific within-group and time-varying unobserved factors remain.

Therefore, we estimate linear probability models to explore the most plausible confounders using the DHS data. First, we estimate the likelihood of selective migration for men and women by treated and control groups in a model where the dependent variable is the *migration* status of the individual, which takes the value of 1 if the person has lived in the current place of residence for less than 5 years, and 0 otherwise.[24] The variable of our interest *son preference* takes the value of 1 if the reported ideal number of boys is larger than the reported ideal number of girls. We measure desired

[23] On December 7, 1988, a 6.8-magnitude strong earthquake hit the north of Armenia. It destroyed Gyumri and other towns and villages within 50-60 kilometers. The death toll of the earthquake is comparable to the death toll of the NK war. More than 25,000 people died and 150,000 were injured.

[24] The results hardly change when we use a longer reference period, e.g., 10 years.

fertility using the data on the expressed ideal number of children.[25] Additionally, we include controls for education, employment and wealth index in these regressions.

Second, we estimate the gender differences in under-5 mortality rates in the treated and control groups in 2010 relative to 2005. The dependent variable, *child mortality* (u5cm), takes a value of 1 if the child died before the age of 5, otherwise, it is 0. We use the birth records in the 5 years before the survey year to arrive at this measure. Additionally, we construct an alternative measure for revealed preference for sons (SLRB) based on the sex ratio of last-born children, which captures both sex-selective abortions and continuation of fertility until a son is born (Jayachandran, 2017). We estimate the SRLB for treated and control groups in 2010 relative to 2005, where the dependent variable takes a value of 1 if the last born child is male, otherwise, it is 0. In the regression models of u5cm and SRLB we also control for mother's characteristics such as education, employment status, migration, number of children, rurality and wealth index.

Results

Table 1 presents the main results for the census-based municipality-level analysis. In column 1 of Table 1, we estimate the differential effects of the post (ceasefire) periods on the treated and control municipalities without the CloseCapital terms. The coefficient on the interaction term Treat×Post3 shows that in the third period when the ceasefire breaches intensified, the under-5 sex ratios increased by 9 points in the treated municipalities relative to the control municipalities, statistically significant at the 5 percent level.

[Table 1 about here.]

In column 2 of Table 1, we exclude the *Treat* terms and include the CloseCapital terms to separately test the possible effects of access to ultrasound technology and other distance-related factors, which are not related to the NK conflict. The results

[25]The results hold when we instead use the actual number of children born.

show that municipalities in the CloseCapital group had higher under-5 sex ratios in the Post1, the first period after the ceasefire, statistically significant at the 10 percent level. The interaction terms with Post2 and Post3 are not statistically significant. In column 3 of Table 1, we add the interaction terms $\sum_{t=1}^{3} \theta_t (Treat_c \times Post_t)$ and relax the common trends assumption by including municipality-specific linear time trends. As a result, the coefficients on Post1 and Post2 become statistically insignificant. The statistically significant interaction effect in the Post3 period for the treated group remains but the effect size becomes slightly smaller.[26] In column 4 of Table 1 we add the macro-region-specific time fixed effects, which corresponds to the full model in equation 1. The coefficient size of Treat × Post3 further increases by 3 points. The results in column 4 show that the under-5 sex ratios in the treated group increased by 10.5, compared to the change in the control group from the reference period to the Post3 period.

We run two additional robustness tests for the result in column 4 of Table 1. First, we include a term for municipality-specific quadratic time trend instead of the macro-region-specific time effects to capture the non-linear municipality-specific changes over time. Second, we exclude the capital Yerevan from the sample to test for the possible overestimation of the treatment effect where the civil unrest in Yerevan coincided with the Martakert skirmishes in the Post3 period. The main effect on Treat×Post3 remains robust to these changes with slightly reduced effect size from 10.5 to 7.6, see Table A3 in the Online Appendix.

Furthermore, the main these results might be affected by outliers and small sample bias as in a few observations, the under-5 population size is less than 100. Therefore, in column 1 of Table 2 we re-estimate the model in equation 1 limiting the sample to observations with over 100 under-5 population size, which reduces the sample size slightly. This hardly changes the precision of the estimates but the size of the coefficient on Treat× Post3 decreases slightly.

Next, we test whether and how the results change as we alter the distance cut-off

[26]The overall correlation coefficient between the two distance variables, CloseCapital and Treat, is 0.46 and negative, which relegates concerns over possible multicollinearity issues. See Figure A6 in the Online Appendix for details.

point. In Table 2, we use median distance (column 2) and the 25th percentile distance (column 3) as cut-off points to define both Treat and CloseCapital variables. These changes hardly affect the statistical significance of the estimates. They alter the size of the coefficients slightly. In the case of the 25th percentile cut-off, the effect size for Treat $\times$ Post3 becomes larger. This implies that the closer the municipality is to the conflict region, the higher is the under-5 sex ratio, which validates the distance measure as a proxy for the threat of war.

[Table 2 about here.]

In Table 3 we use geographic distances to address the endogeneity concerns associated with the use of travel distances. In column 1 we estimate our full model in equation 1, where the main result on Treat$\times$Post3 remains statistically significant at the 1 percent level but has a smaller coefficient size (a drop from 10.5 to 5.9). Besides, we also observe effects for Treat$\times$Post2, which are statistically significant only at the 10 percent level. In column 2, we do not take Stepanakert as the distance point for conflict but use distance to the closest location of the front line clashes in the post-ceasefire period (see the locations on the map in Figure A4 in the Online Appendix). The results are similar to those in column 1, where the result for Treat$\times$Post3 hardly changes but the effect for Treat$\times$Post2 increases in size and becomes statistically significant at the 5 percent level. These results likely suggest that increases in the under-5 sex ratios first took place only in a few municipalities of the treated region in Post2 (weaker statistical significance) and with the increase in the intensity of clashes, these spread over more municipalities of the treated region in Post3 (stronger statistical significance). At the same time, it is not implausible that the use of ultrasound technology for sex-selective abortions became a more widespread practice within the treated group as ultrasound scanners became widely available over time and space. This development could potentially explain why we do not observe effects in Post 1 but rather in the later periods.

[Table 3 about here.]

In column 3 of Table 3, we run a placebo test by defining the treated group based on the distance to the 1988 earthquake epicenter, 30 kilometers from Armenia's second-largest city, Gyumri. We do not find any statistically significant results in this case, which confirms that our results are unlikely to be driven by unobserved heterogeneity correlated with distance itself. We also test the validity of the distance measure as a continuous variable and present the results in Table A4 in the Online Appendix, which show that in the Post3 period, there is a negative correlation between distance to Stepanakert and under-5 sex ratios, statistically significant at the 5 percent level - a result that is consistent with our main finding.

To assess whether the differences in under-5 sex ratios between treated and control groups in the census-based analysis could be due to selective migration, we undertake an analysis based on DHS data presented in Table 4. The estimated coefficients on *Treated* in columns 1 to 4 of Table 4 suggest that there is no statistically significant difference in migration rates between the treated and control regions for men and women. Furthermore, there is no statistically significant difference between the treated and control regions in selection into migration based on son preference (*Treated*×*Son preference*), after controlling for the desired fertility levels and other individual characteristics (columns 2 and 4). These results suggest that concerns related to differential selection into migration are unlikely to confound our main findings based on the census data.

[Table 4 about here.]

Gender differences in under-5 mortality rates could be another source of concern in our analysis if these differ between treated and control groups. Concerns over the use of under-5 sex ratios computed retrospectively as a measure for revealed son preference could be there too. We engage with these concerns through the analysis presented in Table 5 where child mortality rates and SRLB are employed as dependent variables. The estimated coefficients on the binary variable *Male* in columns 1 and 2 show that the under-5 child mortality rate is slightly higher for male relative to female births, statistically significant at the five percent level. The results do not show any statistically significant gender difference in the child mortality rates between the treated and control regions as all the estimates on the *Treated* variable as well as its interaction terms with the *Male* and survey year dummies are statistically insignificant. The SLRB results in column 3 show that in the treated group the likelihood of the last born child being a male increased by almost 9 percentage points in the last 5-year period prior to the survey year, statistically significant at the 5 percent level (Treated×2010). This increase in the revealed preference for sons in the period from 2006 to 2010 in the treated group corroborates our main finding based on the census data.

[Table 5 about here.]

5 Fear of war and son preference: Individual-level analysis

Data and descriptive analysis

We use data from the Caucasus Barometer (CB) of Armenia to study the relationship between the fear of war and son preference at the individual level. The CB is an annual nationally representative household survey, which includes a wide range of demographic, social, economic and political variables collected by the Caucasus Research Resource Centers (CRRC) since 2004. It is one of the few high-quality datasets in the countries of the South Caucasus and other published studies on the region have also used it (e.g., Habibov and Afandi, 2011; Antinyan, 2016; Mavisakalyan and Meinecke, 2016; Mavisakalyan, 2018).

We use the 2010 wave of the CB survey since it includes a question on son preference. The year 2010 also falls into the last period of the census data analysis, Post3. Our sample consists of adults aged 18 to 80 years old, comprising 97 percent of the raw sample.[27] The baseline analysis includes 1,676 non-missing observations.[28]

At the individual level, we define the fear of war based on the response ranking of the two (out of eighteen given) most important issues facing the country. In particular, the variable FEAR OF WAR takes a value of 1 if a respondent reported either Insurance of Peace or Territorial Integrity as one of their top two concerns and 0 otherwise. Descriptive statistics show that 21 percent of people in our sample have a fear of war (see Table B1 in the Online Appendix). The measure of son preference, SON BIAS, equals 1 if the respondent's preferred sex for a single-child family is a boy and 0 otherwise. While 54 percent of all respondents have a preference for sons, it is significantly more prevalent among those who have a fear of war compared to those without a fear of war (a difference of 14 percentage points), as shown in Table B1.

We also observe that the fear of war appears to increase with educational attainment. Namely, university education is more common among individuals with a fear of war (27 percent) than among those without such fear (22 percent). Similarly, the descriptive statistics in Table B1 show that among individuals who have a fear of war, 43 percent are employed; this is considerably higher than the employment rate of 38 percent among those with no fear of war. Moreover, the share of individuals with self-reported good economic standing is slightly higher among those who have a fear of war compared to those who do not have it.

Estimation strategy

We evaluate the baseline relationship between the fear of war and son bias by constructing an estimation model in which the propensity for son bias, $Bias_i^*$ for an individual i is assumed to depend on the fear of war, $Fear_i$, together with series of controls X_i for demographic, socio-economic and location characteristics. Unobserved factors

[27]The remaining 3 percent are those over the age of 80. The results are largely insensitive to the presence of these individuals.

[28]The number of missing observations for the key variables of interest is small; e.g. data on the variables used in the construction of measures for son bias and war as a primary concern are missing for 0.64 percent of individuals only.

ε_i further contribute to the propensity for son bias, leading to an equation of the form:

$$Bias_i^* = X_i\kappa + \beta Fear_i + \varepsilon_i \text{ for all } i = 1, ..., N. \tag{2}$$

We assume that observed son bias $Bias_i$ relates to latent propensity through the criterion $Bias_i = 1(Bias_i^* \geq 0)$, so that the probability of having a son bias under an assumption of normality for ε_i becomes

$$Pr(Bias_i = 1|X_i, Fear_i) = \Phi(X_i\kappa + \beta Fear_i), \tag{3}$$

with the marginal effect of fear of war derived from the estimated model thus:

$$\frac{\partial Pr(Bias_i = 1|X_i, Fear_i)}{\partial Fear_i} = \beta\phi(X_i\kappa + \beta Fear_i). \tag{4}$$

The control variables in the estimation model include those shown in Table B1 in the Online Appendix, which are gender, age and family status, education, employment and income standing as well as urban/rural status of an individual's place of residence. The inclusion of observable characteristics of individuals as controls potentially mitigates the problem of omitted variable bias to a certain extent; however, it does not eliminate it.[29] A conventional approach to directly address such endogeneity concerns would be to exploit an instrumental variable and estimate a bivariate probit model. However, we do not have persuasive instrumental variables in the dataset. Therefore, we take two alternative approaches to mitigate the concerns around endogeneity further. First, we employ a matching approach to account for differences in individuals' observable characteristics and the likelihood of having a fear of war. Second, we use a partial identification approach proposed by Oster (2019) to evaluate how large the amount of selection on unobservable variables would need to be relative

[29]Note that we extend the list of controls beyond the baseline set as presented in Table 6 in the results sub-section.

to the amount of selection on observables to explain away the entire effect of FEAR OF WAR on SON BIAS.[30]

The general idea of the matching approach is straightforward: we examine the impact of FEAR OF WAR on SON BIAS for individuals who have a fear of war (treatment group), compared to those who do not have such fear but are as similar as possible in terms of observable characteristics that affect the outcome variable of interest (control group). Formally, we estimate the following model:

$$\tau_{ATT}(x) = E[Bias(1)|T = 1, X = x] - E[Bias(1)|T = 0, X = x] \tag{5}$$

where *Bias* is our outcome variable, T indicates whether an individual is exposed to treatment, i.e., has a fear of war $(T = 1)$ or not $(T = 0)$, and x is a vector of relevant characteristics that affect the outcome variable. First, we employ entropy balancing to select matches for the units exposed to treatment (Hainmueller, 2012). Entropy balancing is in a way a generalization of conventional matching approaches since it employs a synthetic control group that represents "a virtually perfect image of the treatment group" (Neuenkirch and Neumeier, 2016, p. 113). Second, we follow a more traditional approach employing propensity score matching (Rosenbaum and Rubin, 1983).

The matching approach makes our identification strategy robust against selection on observables. However, despite the fact that we employ a rich set of conditioning variables, there may still be concerns over unobserved heterogeneity. We apply a partial identification approach to assess the extent of the associated bias. To do that, we evaluate the bias-adjusted coefficient derived by Oster (2019):

$$\beta^* \approx \tilde{\beta} - \delta[\dot{\beta} - \tilde{\beta}]\frac{R_{max} - \tilde{R}}{\tilde{R} - \dot{R}} \tag{6}$$

[30]See Oster (2019) for detailed description of the test; for recent applications please see Baranov et al. (2015); Birthal et al. (2015); Freier et al. (2015); Mavisakalyan et al. (2018).

where $\dot{\beta}$ and $\dot{R}$ are the coefficient and the R-squared from a regression with the treatment only ($\kappa = 0$) and $\tilde{\beta}$ and $\tilde{R}$ are the coefficients and R-squared from the regression with the treatment and the observed controls ($\kappa \neq 0$). δ denotes the relative importance of observable relative to unobservable variables in generating bias; R_{max} is the R-squared from a hypothetical regression of SON BIAS on all observable and unobservable variables. Both δ and R_{max} are unknown. Oster (2019) proposes a bounding approach whereby the estimated effect of FEAR OF WAR should range from $\tilde{\beta}$ to β^* estimated under an assumption of $\delta = 1$, and given values of $R_{max} \in [\tilde{R}, 1]$. We follow Oster's proposal, based on evidence from randomized control studies published in reputable economics journals, to set $R_{max} = min\{1.3\tilde{R}, 1\}$ and estimate the coefficient bounds accordingly.

Results

Column 1 of Table 6 reports the marginal effects described in equation 4 evaluated at the sample means based on a regression that includes the set of baseline controls (demographic, socio-economic and location characteristics of individuals); see Table B2 in the Online Appendix for the expanded set of baseline estimations. Consistent with the descriptive statistics in Table B1 in the Online Appendix, the estimated marginal effect confirms a statistically significant and positive relationship between FEAR OF WAR and SON BIAS. This baseline estimate suggests that relative to individuals who have no fear of war, experiencing fear of war is associated with an increase in the probability of expressing son preference of 12.2 percentage points.

[Table 6 about here.]

Furthermore, we aim to mitigate the omitted variable bias to the extent possible by adding proxy variables for unobserved factors that could be correlated with the unexplained components of SON BIAS. The results of this analysis are presented in columns 2 to 6 of Table 6. Specifically, one of the most plausible omitted variables in the analysis could be nationalist attitudes among individuals, which may be correlated with stronger concerns for peace and territorial integrity as well as SON BIAS. In fact, as the results in Table 6 show the variable RACIAL TOLERANCE, which is our proxy measure for non-conservative and open (not nationalistic) attitudes, has a statistically significant (and negative) association with SON BIAS.[31] Yet, the inclusion of this variable does not undermine the statistically significant association between SON BIAS and FEAR OF WAR.

The estimated marginal effects on other controls, such as the number of overseas trips in the past 5 years and religiosity (NO OVERSEAS TRIPS, RELIGIOUS MAJORITY, VERY RELIGIOUS), which are alternative proxy measures for a conservative background; proficiency of Russian language (RUSSIAN FLUENT, RUSSIAN NATIVE), which captures broader dimensions of human capital in the post-Soviet context (Duncan and Mavisakalyan, 2015; Mavisakalyan, 2017); and the measures of individual vulnerability (HAS CLOSE PEOPLE, FEELS EMPTINESS) and institutional trust (DISTRUST IN ARMY) are not statistically significant at the conventional levels. As a result, we estimate a positive association between FEAR OF WAR and SON BIAS statistically significant at the 1 percent level. The size of the association drops from 12.2 to 10.8 percentage points after including the additional controls. Nevertheless, this is an economically significant size, implying that the probability of having a son bias is 20 percent higher for those who have a fear of war.[32]

One may argue that the relationship between fear of war and son bias could be capturing both the preference for sons because of their role as defenders, and because

[31] RACIAL TOLERANCE equals 1 if the individual approves marriage with Chinese people, otherwise 0. The choice of this population group is driven by the data availability.

[32] In the Online Appendix we demonstrate that this result largely persists when we employ alternative measures of FEAR OF WAR (see Table B3, columns 1-3) and SON BIAS (see Table B4). In Table B4 we additionally show that FEAR OF WAR is also positively and significantly associated with the share of sons that an individual has. We also estimate the relationship between FEAR OF WAR and SON BIAS by limiting the analysis to specific groups of individuals based on education, gender and age groups (see Table B5 in the Online Appendix). While in all samples we confirm the baseline findings, we estimate particularly large marginal effects of FEAR OF WAR on SON BIAS among males, individuals over the age of 45 and those with post-secondary education.

of their role as breadwinners. This is a plausible interpretation given the qualitative evidence presented in Figure 5. Yet, empirically, we do not find any suggestive evidence in support of this interpretation. Namely, as shown in columns 4-6 of Table B3 in the Online Appendix, we substitute fear of war with proxy variables for economic and institutional concerns such as concerns over the labor market, including unemployment or wages (column 4); concerns over social security issues such as affordable healthcare, quality education, pensions or poverty (column 5); concerns over various dimensions of institutions including corruption, the court system, fairness of elections, human rights, freedom of speech or property rights (column 6). None of these variables are statistically significant. Moreover, the inclusion of these variables in the same specification with fear of war does not affect our central finding: the estimated marginal effect on FEAR OF WAR (columns 7-9). Assuming these variables at least partially capture the economic considerations behind son preference, we do not find any evidence that they explain away the established association between fear of war and son preference.

Further, we employ a matching approach where we match and compare individuals who have a fear of war to individuals who do not have such fear but are as similar as possible with regards to characteristics that affect son bias. The results are presented in Table 7, where the full set of matching covariates are included as control variables. Column 1 reports the estimates obtained from weighted regressions where observations in the treatment group have a weight of 1 and in the control group they have a positive weight obtained from matching using entropy balancing. In columns 2 and 3 we report the estimates from kernel and radius matching estimators. The estimated coefficients in all three cases are statistically significant and similar in magnitude.

[Table 7 about here.]

By employing a matching approach, we address the issue of selection on observables at least partially. To assess the extent of bias from unobservables, we apply the partial identification method proposed by Oster (2019). The results of this exercise are summarized in Table 8. In column (2) we report the coefficient bounds $[\tilde{\beta}, \beta^*]$ for models with baseline and comprehensive lists of controls. $\tilde{\beta}$ comes from the specifications controlling for all baseline/comprehensive observables. β^* is evaluated using equation (6) by setting $\delta = 1$ and assuming $R_{max} = min\{1.3\tilde{R}, 1\}$ - the rule of thumb proposed by Oster (2019). Reassuringly, the identified sets $[\tilde{\beta}, \beta^*]$ exclude zeros. Moreover, as we see in column (1), in both cases $\delta > 1$, i.e. the unobservables would have to be more important than the observables in explaining SON BIAS. Thus, the method developed by Oster (2019) corroborates our findings.

Overall, the individual-level analysis sheds light on the likely behavioral mechanism underlying the link between the threat of war and skewed sex ratios at birth at the aggregate level. It confirms that individuals facing the fear of war are more likely to express son preference.

[Table 8 about here.]

6 Conclusions

This study shows that unresolved conflict between ethnic groups can contribute to the increased preference for sons and give rise to missing girls. We argue that the threat of war may lead to a revival of survival values and son bias, reflected in the juvenile sex ratios at the population level. We provide empirical support for our argument by exploiting the period of ceasefire breaches observed during the ceasefire period (1996-2011) of the Nagorno Karabakh conflict between Armenia and Azerbaijan.

Our findings from the census data, based on difference-in-difference estimations, show that in the post-war period, Armenian communities closer to the conflict region (treated group) experienced 6-10 points higher under-5 sex ratios in 2011. The corresponding period of live births (2007-2011) coincides with the period of intense ceasefire breaches in Nagorno-Karabakh. The results based on an alternative measure for son preference, sex ratios of last-born children, using DHS data show that the likelihood of the last birth to be a male increased by 9 percentage points in the treated group in the period from 2006 to 2010, corroborating our main finding based on the census data. We test the validity of our hypothesized mechanism on the link between fear of war and son preference using data from a nationally representative household survey. We find that having a fear of war is associated with 10 percentage points increase in the likelihood of having a son bias at the individual level.

Previous literature has established the importance of economic and cultural factors (e.g., financial support and inheritance practices) behind the son preference. We show that in a simmering conflict, the role of males as potential defenders of homeland is one of the important mechanisms for increased son preference along with the existing economic motives.

While our findings are country-specific, they provide some relevant insights for the research on missing girls in other regions. For example, in India, the Sikh-Hindu clashes in Punjab and Haryana could have led to group survival concerns in those provinces, contributing to son bias and high sex ratios at birth. Similarly, the intense militarization in South Korea up to the 1990s and the recent conflicts in Eastern Europe and the South Caucasus are likely to have increased survival concerns within the affected ethnic groups and contributed to the increased preference for sons in those regions. At this stage, these are mere speculations to be explored in future research.

It is paramount to relate our argument to countries with inter-ethnic conflicts where the fertility rates are above the replacement level (several countries in Central Asia, Africa and the Middle East). In this case, one would not necessarily observe highly skewed sex ratios for all live births, hence no missing girls, but one might expect skewed sex ratios among the last births - another measure for son preference.

That conflict may matter for exacerbated preference for sons and skewed sex ratios among children is a novel hypothesis in the literature. This article provides empirical evidence supporting this hypothesis for the case of Armenia, which has one of the highest sex ratios at birth in the world. Future research can help identify whether such effects are also observed across other regions, given similar conditions. In broader terms, our study points towards potentially long-lasting consequences of conflict and highlights the deep historical sources of gender inequalities.

References

Abrahamyan, R., Avagyan, G., Gyulkhasyan, V., Kuyumjyan, K., Saghatelyan, A., Aghajanyan, A., and Arushanyan, A. (2012). Prevalence of and reasons for sex-selectve abortions in Armenia. *Report. United Nations Population Fund Armenia. Yerevan*.

AGBU (2010). Chercherian family foundation donates additional probes and printers for ultrasound machines to the ultrasound training and research center. *AGBU (Armenian General Benevolent Union) Armenia Newsletter. Issue 5, March*.

Ahn, N. (1995). Measuring the value of children by sex and age using a dynamic programming model. *The Review of Economic Studies*, 62(3):361–379.

Antinyan, A. (2016). Reference group income and subjective well-being: Empirical evidence from low-income transition economies. *Social Indicators Research*, 127(3):1333–1348.

Arstamyan, G. (2009). *Chronicle of Fame: NKR National Liberation Struggle 1988-2009 [textitin Armenian]*. Stepanakert: Dizak Plus, ISBN: 978-9939-800-68-4.

Baranov, V., Bennett, D., and Kohler, H.-P. (2015). The indirect impact of antiretroviral therapy: Mortality risk, mental health, and HIV-negative labor supply. *Journal of Health Economics*, 44:195 – 211.

BBC (2008). Karabakh casualty toll disputed. *BBC News, Retrieved: http://news.bbc.co.uk/2/hi/europe/7278871.stm*.

Ben-Porath, Y. and Welch, F. (1976). Do sex preferences really matter? *The Quarterly Journal of Economics*, 90(2):285–307.

Bertrand, M., Duflo, E., and Mullainathan, S. (2004). How much should we trust differences-in-differences estimates? *The Quarterly journal of economics*, 119(1):249–275.

Birthal, P. S., Roy, D., and Negi, D. S. (2015). Assessing the impact of crop diversification on farm poverty in india. *World Development*, 72:70 – 92.

Branisa, B., Klasen, S., Ziegler, M., Drechsler, D., and Jütting, J. (2014). The institutional basis of gender inequality: The social institutions and gender index (SIGI). *Feminist Economics*, 20(2):29–64.

Broers, L. (2015). From “frozen conflict” to enduring rivalry: reassessing the Nagorny Karabakh conflict. *Nationalities Papers*, 43(4):556–576.

Buvinic, M., Gupta, M. D., Casabonne, U., and Verwimp, P. (2013). *Violent conflict and gender inequality: An overview*. The World Bank.

Clark, S. (2000). Son preference and sex composition of children: Evidence from India. *Demography*, 37(1):95–108.

Coale, A. J. (1991). Excess female mortality and the balance of the sexes in the population: An estimate of the number of ”missing females”. *The Population and Development Review*, pages 517–523.

Collyer, M. e. a. (2015). Migration and Skills in Armenia and Georgia. Comparative Report. *European Training Foundation*.

Dagnelie, O., De Luca, G. D., and Maystadt, J.-F. (2018). Violence, selection and infant mortality in Congo. *Journal of Health Economics*, 59:153–177.

Das Gupta, M. and Shuzhuo, L. (1999). Gender bias in China, South Korea and India 1920–1990: Effects of war, famine and fertility decline. *Development and Change*, 30(3):619–652.

Das Gupta, M., Zhenghua, J., Bohua, L., Zhenming, X., Chung, W., and Hwa-Ok, B. (2003). Why is son preference so persistent in East and South Asia? A cross-country study of China, India and the Republic of Korea. *The Journal of Development Studies*, 40(2):153–187.

De Waal, T. (2003). *Black Garden: Armenia and Azerbaijan through Peace and War*. New York University Press.

De Waal, T. (2010). Remaking the Nagorno-Karabakh peace process. *Survival*, 52(4):159–176.

De Waal, T. (2016a). Nagorno-Karabakh's cocktail of conflict explodes again. *BBC News, Retrieved: https://www.bbc.com/news/world-europe-35954969, April 3*.

De Waal, T. (2016b). Prisoners of the Caucasus: Resolving the Karabakh security dilemma. *Carnegie Europe*, 16.

Desmet, K., Ortuño-Ortín, I., and Wacziarg, R. (2017). Culture, ethnicity, and diversity. *American Economic Review*, 107(9):2479–2513.

DHS (2012). Armenia Demographic and Health Survey 2010. Technical report, National Statistical Service/Armenia, Ministry of Health/Armenia, and ICF International.

Dudwick, N. (2015). *'Missing women' in the South Caucasus*. World Bank, Washington, DC.

Duncan, A. and Mavisakalyan, A. (2015). Russian language skills and employment in the Former Soviet Union. *Economics of Transition*, 23(3):625–656.

Duthé, G., Meslé, F., Vallin, J., Badurashvili, I., and Kuyumjyan, K. (2012). High sex ratios at birth in the Caucasus: Modern technology to satisfy old desires. *Population and Development Review*, 38(3):487–501.

Ebenstein, A. (2010). The 'missing girls' of China and the unintended consequences of the one child policy. *Journal of Human Resources*, 45(1):87–115.

Ebenstein, A. (2014). Patrilocality and missing women. *SSRN*.

Edlund, L. (1999). Son preference, sex ratios, and marriage patterns. *Journal of Political Economy*, 107(6):1275–1304.

Fogarty, L. and Feldman, M. W. (2011). The cultural and demographic evolution of son preference and marriage type in contemporary China. *Biological Theory*, 6(3):272–282.

Freier, R., Schumann, M., and Siedler, T. (2015). The earnings returns to graduating with honors — evidence from law graduates. *Labour Economics*, 34:39 – 50.

Freizer, S. (2014). Twenty years after the Nagorny Karabakh ceasefire: An opportunity to move towards more inclusive conflict resolution. *Caucasus Survey*, 1(2):109–122.

Giragosian, R. (2014). Nagorno Karabakh situational assessment. *Yerevan: Regional Studies Center*.

Habibov, N. N. and Afandi, E. N. (2011). Self-rated health and social capital in transitional countries: Multilevel analysis of comparative surveys in Armenia, Azerbaijan, and Georgia. *Social Science & Medicine*, 72(7):1193 – 1204.

Hainmueller, J. (2012). Entropy balancing for causal effects: A multivariate reweighting method to produce balanced samples in observational studies. *Political Analysis*, 20(1):25–46.

Haughton, J. and Haughton, D. (1998). Are simple tests of son preference useful? An evaluation using data from Vietnam. *Journal of Population Economics*, 11(4):495–516.

Hohmann, S., Roche, S., and Garenne, M. (2010). The changing sex ratios at birth during the civil war in Tajikistan: 1992–1997. *Journal of Biosocial Science*, 42(6):773–786.

Hopmann, P. T. (2015). Minsk Group Mediation of the Nagorno-Karabakh Conflict: Confronting an "Intractable Conflict". In *OSCE Yearbook 2014*, pages 163–180. Nomos Verlagsgesellschaft mbH & Co. KG.

International Crisis Group (2009). Nagorno-Karabakh: Getting to a Breakthrough. *Crises Group Europe Briefing 55. International Crisis Group. Tbilisi/Baku/Yerevan/Istanbul/Brussels*.

International Crisis Group (2011). Armenian and Azerbaijan: Preventing War. *Crises Group Europe Briefing 60. International Crisis Group. Tbilisi/Baku/Yerevan/Istanbul/Brussels*.

Jack, A. (2017). Our community loves boys more. Armenia's missing girls. *Financial Times*.

James, W. H. (1997). Coital rates and sex ratios. *Human reproduction*, 12(9):2083–2085.

James, W. H. (2009). The variations of human sex ratio at birth during and after wars, and their potential explanations. *Journal of Theoretical Biology*, 257(1):116–123.

James, W. H. and Valentine, J. (2014). A further note on the rises in sex ratio at birth during and just after the two world wars. *Journal of Theoretical Biology*, 363:404–411.

Jayachandran, S. (2017). Fertility decline and missing women. *American Economic Journal: Applied Economics*, 9(1):118–139.

Kahneman, D. and Tversky, A. (1984). Choices, values, and frames. *American psychologist*, 39(4):341.

Kemper, T. D. (1994). Social stratification, testosterone, and male sexuality. *In Social Stratification and Socioeconomic Inequality, Reproductiveand Interpersonal Aspects of Dominance and Status, edited by Lee Ellis. Westport, Conn.*, 2:47–62.

Klasen, S. (1994). 'Missing women' reconsidered. *World Development*, 22(7):1061–1071.

Klasen, S. (1998). Marriage, bargaining, and intrahousehold resource allocation: Excess female mortality among adults during early German development, 1740–1860. *The Journal of Economic History*, 58(2):432–467.

Klasen, S. and Wink, C. (2003). "Missing women": Revisiting the debate. *Feminist Economics*, 9(2-3):263–299.

Kramer, A. E. (2020). Armenia and Azerbaijan: What sparked war and will peace prevail? *The New York Times. Dec. 1. Url: https://nyti.ms/2IjbSNt*.

Kureishi, W. and Wakabayashi, M. (2011). Son preference in Japan. *Journal of Population Economics*, 24(3):873–893.

Lagerlöf, N.-P. (2003). Gender equality and long-run growth. *Journal of Economic Growth*, 8(4):403–426.

Li, H., Yi, J., and Zhang, J. (2011). Estimating the effect of the one-child policy on the sex ratio imbalance in China: Identification based on the difference-in-differences. *Demography*, 48(4):1535–1557.

Li, N., Feldman, M. W., and Li, S. (2000). Cultural transmission in a demographic study of sex ratio at birth in China's future. *Theoretical Population Biology*, 58(2):161–172.

Mavisakalyan, A. (2017). Returns to language skills in transition economies. *IZA World of Labor*, 416.

Mavisakalyan, A. (2018). Do employers reward physical attractiveness in transition countries? *Economics & Human Biology*, 28:38 – 52.

Mavisakalyan, A. and Meinecke, J. (2016). The labor market return to academic fraud. *European Economic Review*, 82:212 – 230.

Mavisakalyan, A., Tarverdi, Y., and Weber, C. (2018). Talking in the present, caring for the future: Language and environment. *Journal of Comparative Economics*, 46(4):1370 – 1387.

McDermott, R., Fowler, J. H., and Smirnov, O. (2008). On the evolutionary origin of prospect theory preferences. *The Journal of Politics*, 70(2):335–350.

Michael, M., King, L., Guo, L., McKee, M., Richardson, E., and Stuckler, D. (2013). The mystery of missing female children in the Caucasus: An analysis of sex ratios by birth order. *International perspectives on sexual and reproductive health*, pages 97–102.

Moore, S. (2018). "We lose 1,400 girls a year. who will our boys marry?":Armenia's quandary. *The Guardian*.

Neuenkirch, M. and Neumeier, F. (2016). The impact of US sanctions on poverty. *Journal of Development Economics*, 121:110 – 119.

Oster, E. (2019). Unobservable selection and coefficient stability: Theory and evidence. *Journal of Business & Economic Statistics*, 37(2):187–204.

Park, C. B. (1983). Preference for sons, family size, and sex ratio: An empirical study in Korea. *Demography*, 20(3):333–352.

Pettersson, T. and Eck, K. (2018). Organized violence, 1989–2017. *Journal of Peace Research*, 55(4):535–547.

Qian, N. (2008). Missing women and the price of tea in China: The effect of sex-specific earnings on sex imbalance. *The Quarterly Journal of Economics*, 123(3):1251–1285.

Rose, E. (1999). Consumption smoothing and excess female mortality in rural India. *Review of Economics and Statistics*, 81(1):41–49.

Rosenbaum, P. R. and Rubin, D. B. (1983). The central role of the propensity score in observational studies for causal effects. *Biometrika*, 70(1):41–55.

Rosenzweig, M. R. and Schultz, T. P. (1982). Market opportunities, genetic endowments, and intrafamily resource distribution: Child survival in rural India. *The American Economic Review*, 72(4):803–815.

Sen, A. (1989). Women's survival as a development problem. *Bulletin of the American Academy of Arts and Sciences*, 43(2):14–29.

Sen, A. (1990). More than 100 million women are missing. *The New York Review of Books*, 37(20):61–66.

Sng, T.-H. and Zhong, S. (2018). Historical violence and China's missing women. *Working Paper. Economic History Association*.

Sundberg, R. and Melander, E. (2013). Introducing the UCDP georeferenced event dataset. *Journal of Peace Research*, 50(4):523–532.

Tapsoba, A. (2020). The cost of fear: Impact of violence risk on child health during conflict. *TSE Working Paper. No. 1127*.

The World Factbook (2017). *Central Intellegance Agency, Washington DC*.

Trivers, R. L. and Willard, D. E. (1973). Natural selection of parental ability to vary the sex ratio of offspring. *Science*, 179(4068):90–92.

UNFPA (2015). Preventing gender-biased sex selection in Eastern Europe and Central Asia. *UNFPA EECARO*.

UNHCR (1994). Report of the United Nations High Commissioner for Refugees,1994, a/49/12.

USSR (1991). The National Composition of the Population of the USSR, according to All-Union Population Census of 1989. Finance and Statistics, Moscow.

Valente, C. (2015). Civil conflict, gender-specific fetal loss, and selection: A new test of the Trivers–Willard hypothesis. *Journal of Health Economics*, 39:31–50.

Verwimp, P. and Van Bavel, J. (2013). Schooling, violent conflict, and gender in Burundi. *The World Bank Economic Review*, 28(2):384–411.

Vincenty, T. (1975). Direct and inverse solutions of geodesics on the ellipsoid with application of nested equations. *Survey review*, 23(176):88–93.

Voors, M. J., Nillesen, E. E., Verwimp, P., Bulte, E. H., Lensink, R., and Van Soest, D. P. (2012). Violent conflict and behavior: a field experiment in Burundi. *American Economic Review*, 102(2):941–64.

Weisel, O. and Zultan, R. (2016). Social motives in intergroup conflict: Group identity and perceived target of threat. *European Economic Review*, 90:122–133.

Westoff, C. F. and Rindfuss, R. R. (1974). Sex preselection in the United States: Some implications. *Science*, 184(4137):633–636.

World Bank (2015). World Development Indicators. *The World Bank. Washington DC*.

Table 1: Under-5 sex ratios in Armenia: Baseline results

	(1)	(2)	(3)	(4)
Post1	9.637***	5.890**	-0.895	-8.659
	(2.017)	(2.753)	(4.226)	(6.469)
Post2	10.563***	9.484***	-0.162	0.841
	(1.846)	(1.977)	(3.215)	(5.209)
Post3	8.147***	14.521***	-4.037**	-6.959**
	(1.606)	(3.674)	(1.878)	(2.942)
Treat × Post1	-0.507		3.582	10.747
	(3.883)		(5.799)	(7.569)
Treat × Post2	-0.025		0.745	-0.195
	(3.196)		(3.936)	(5.192)
Treat × Post3	9.295**		7.807***	10.509***
	(4.526)		(2.219)	(3.271)
CloseCapital × Post1		6.752*	11.027**	1.729
		(3.505)	(5.026)	(4.565)
CloseCapital × Post2		1.987	9.916***	0.800
		(3.033)	(3.559)	(4.014)
CloseCapital × Post3		-5.411	9.653***	-6.525***
		(4.022)	(2.125)	(2.085)
Constant	105.102***	105.132***	105.305***	105.305***
	(0.772)	(0.782)	(0.935)	(0.932)
Within R-squared	0.176	0.181	0.483	0.498
Municipality FE	✓	✓	✓	✓
Municipality time trend	-	-	✓	✓
Macro-region time FE	-	-	-	✓
Number of municipalities	76	76	76	76
Observations	298	298	298	298

Note. The dependent variable is male over female sex ratios for children under-5. *Treat* is a binary variable equal to 1 for municipalities closer than the average distance to Stepanakert. *CloseCapilal* is a binary variable equal to 1 for municipalities closer than the average distance to Yerevan. Standard errors are clustered at the municipality level. * denotes significance at 10 percent; ** at 5 percent; *** at 1 percent levels.

Table 2: Under-5 sex ratios in Armenia: Robustness to sample and cut-off choice

	(1) Sample: births≥100	(2) Cut-off: median	(3) Cut-off: p25
Post1	-2.378	-4.810	-0.085
	(4.311)	(7.342)	(4.919)
Post2	-0.959	2.140	-1.074
	(5.054)	(5.247)	(4.346)
Post3	-4.266**	-6.733***	-10.026***
	(1.792)	(2.531)	(1.521)
Treat × Post1	7.277	5.770	-0.393
	(4.663)	(6.920)	(8.375)
Treat × Post2	3.591	-1.107	0.220
	(4.935)	(4.449)	(6.246)
Treat × Post3	8.090***	9.778***	13.221***
	(1.738)	(2.551)	(2.101)
CloseCapital × Post1	4.213	8.324*	5.965
	(3.552)	(4.427)	(4.960)
CloseCapital × Post2	13.731***	14.413***	15.019***
	(4.957)	(5.246)	(4.946)
CloseCapital × Post3	9.512***	11.785***	15.628***
	(1.760)	(2.546)	(2.022)
Constant	105.301***	105.305***	105.305***
	(0.722)	(0.935)	(0.936)
Within R-squared	0.520	0.492	0.488
Municipality time trend	✓	✓	✓
Municipality FE	✓	✓	✓
Macro-region time FE	✓	✓	✓
Number of municipalities	75	76	76
Observations	287	298	298

Note. The dependent variable is male over female sex ratios for children under-5. *Treat* is a binary variable equal 1 for municipalities closer than the average (median/25th percentile) to Stepanakert. *CloseCapilal* is a binary variable equal to 1 for municipalities closer than average (median/25th percentile) to Yerevan. Standard errors are clustered at the municipality level. Standard errors are clustered at the municipality level. * denotes significance at 10 percent; ** at 5 percent; *** at 1 percent levels.

Table 3: Under-5 sex ratios in Armenia: Robustness to alternative distance measures

	(1) Conflict distance	(2) Battle distance	(3) Earthquake distance
Post1	-2.617	-7.360	-0.105
	(6.482)	(6.194)	(5.521)
Post2	-5.573	-8.838**	2.697
	(4.384)	(3.967)	(3.500)
Post3	-3.431**	-3.300***	2.135
	(1.301)	(1.176)	(1.376)
Treat × Post1	3.183	8.614	3.175
	(6.797)	(7.046)	(6.806)
Treat × Post2	7.822*	11.566**	-6.443
	(4.505)	(4.575)	(5.411)
Treat × Post3	5.938***	5.877***	2.062
	(1.420)	(1.416)	(2.651)
CloseCapital × Post1	5.594	5.068	4.695
	(4.810)	(4.724)	(6.400)
CloseCapital × Post2	6.537	6.177	-0.917
	(4.994)	(5.012)	(5.910)
CloseCapital × Post3	9.944***	9.707***	3.385*
	(1.372)	(1.456)	(1.787)
Constant	105.305***	105.305***	105.305***
	(0.941)	(0.935)	(0.942)
Within R-squared	0.479	0.483	0.489
Municipality time trend	✓	✓	✓
Municipality FE	✓	✓	✓
Macro-region time FE	✓	✓	✓
Number of municipalities	76	76	76
Observations	298	298	298

Note. The dependent variable is male over female sex ratios for children under-5. *Treat* is a binary variable equal 1 for municipalities closer than the average to Stepanakert (column 1)/battle location (column 2)/earthquake epicenter (column 3). *CloseCapilal* is a binary variable equal to 1 for municipalities closer than average to Yerevan. Standard errors are clustered at the municipality level. * denotes significance at 10 percent; ** at 5 percent; *** at 1 percent levels.

Table 4: Selection into migration

	(1) Men	(2) Men	(3) Women	(4) Women
Treated	-0.011	-0.010	-0.013	-0.010
	(0.013)	(0.013)	(0.010)	(0.010)
Son preference	0.038**	0.036*	-0.013	-0.011
	(0.018)	(0.018)	(0.011)	(0.011)
Treated× Son preference	-0.031	-0.031	0.010	0.007
	(0.024)	(0.024)	(0.015)	(0.015)
Ideal number of children	0.000	0.001	-0.017***	-0.016***
	(0.004)	(0.004)	(0.003)	(0.003)
Currently employed		-0.007		-0.042***
		(0.010)		(0.006)
Educ: secondary		0.017		0.028
		(0.013)		(0.041)
Educ: tertiary		0.047**		0.038
		(0.022)		(0.042)
Wealth: poorer		0.028		0.007
		(0.017)		(0.012)
Wealth: middle		0.017		0.014
		(0.016)		(0.013)
Wealth: richer		0.004		-0.006
		(0.017)		(0.013)
Wealth: richest		0.033		-0.003
		(0.022)		(0.014)
Constant	0.039***	0.002	0.131***	0.104**
	(0.014)	(0.016)	(0.011)	(0.043)
R-squared	0.007	0.016	0.005	0.010
Number of dhs clusters	294	294	304	304
Observations	1358	1356	6408	6405

Note. Analysis based on DHS Armenia 2005. The dependent variable, *migration status*, equals 1 if respondents have spent less than 5 years in their current place of residence, and 0 otherwise. *Son preference* equals 1 if the respondents' preferred ideal number of boys is larger than the ideal number of girls, otherwise it equals 0. Coefficients from a linear probability model; standard errors clustered by DHS clusters. * denotes significance at 10 percent; ** at 5 percent; *** at 1 percent levels.

Table 5: Child mortality rates and sex ratios at last birth

	(1) u5cm	(2) u5cm	(3) SRLB
Treated	0.007	0.005	-0.046
	(0.009)	(0.008)	(0.030)
Year 2010	0.007	0.005	-0.032
	(0.008)	(0.009)	(0.033)
Treated × 2010	-0.005	-0.003	0.087**
	(0.013)	(0.013)	(0.044)
Male	0.023**	0.020**	
	(0.010)	(0.010)	
Treated × Male	-0.009	-0.008	
	(0.016)	(0.015)	
Year 2010 × Male	-0.019	-0.018	
	(0.013)	(0.013)	
Treated × 2010 × Male	-0.005	-0.006	
	(0.021)	(0.020)	
Constant	0.009*	0.001	0.476***
	(0.005)	(0.014)	(0.055)
Controls	No	Yes	Yes
Number of dhs clusters	308	308	308
Observations	2861	2861	2252

Note. Analysis based on DHS Armenia 2005 and 2010 (birth records). In columns 1 and 2, the dependent variable takes the value of 1 if the child died before age 5. Column 3 estimates the probability of a last born to be a male (SRLB). Controls include mother's education, mother's employment status, change of residence in the last 5 years, the total number of births, rurality and wealth index. Coefficients from a linear probability model; standard errors are clustered at the DHS cluster level. * denotes significance at 10 percent; ** at 5 percent; *** at 1 percent levels.

Table 6: Individual model with comprehensive controls

Control variables	(1)	(2)	(3)	(4)	(5)	(6)
	A: Univariate probit estimates					
FEAR OF WAR	0.122***	0.121***	0.119***	0.107***	0.104***	0.108***
	(0.030)	(0.030)	(0.030)	(0.032)	(0.033)	(0.033)
RUSSIAN FLUENT		-0.013	-0.011	-0.027	-0.030	-0.036
		(0.039)	(0.039)	(0.042)	(0.043)	(0.043)
RUSSIAN NATIVE		-0.084	-0.102	-0.108	-0.111	-0.126
		(0.076)	(0.078)	(0.084)	(0.084)	(0.086)
NO OVERSEAS TRIPS			0.010	0.026	0.025	0.028
			(0.029)	(0.032)	(0.032)	(0.032)
RACIAL TOLERANCE			-0.084***	-0.071**	-0.075**	-0.074**
			(0.030)	(0.032)	(0.032)	(0.033)
RELIGIOUS MAJORITY				0.006	0.007	-0.008
				(0.072)	(0.072)	(0.074)
VERY RELIGIOUS				-0.027	-0.024	-0.018
				(0.029)	(0.029)	(0.030)
HAS CLOSE PEOPLE					-0.015	-0.001
					(0.040)	(0.041)
FEELS EMPTINESS					-0.008	-0.002
					(0.029)	(0.029)
DISTRUST IN ARMY						0.038
						(0.034)
Baseline controls included	Yes	Yes	Yes	Yes	Yes	Yes
Pseudo R^2	0.069	0.070	0.074	0.077	0.074	0.074
N	1676	1671	1649	1430	1409	1387

Note. Dependent variable is SON BIAS. Marginal effects for a discrete change of a variable from 0 to 1 for a person with FEAR OF WAR=0; controls are fixed at sample means. Standard errors in parentheses. *Denotes significance at the 10 percent; ** 5 percent; *** 1 percent levels.

Table 7: Matching estimations

	(1)	(2)	(3)
Fear of war	0.107***	0.104**	0.095**
	(0.025)	(0.046)	(0.041)
N	1387	1387	1387

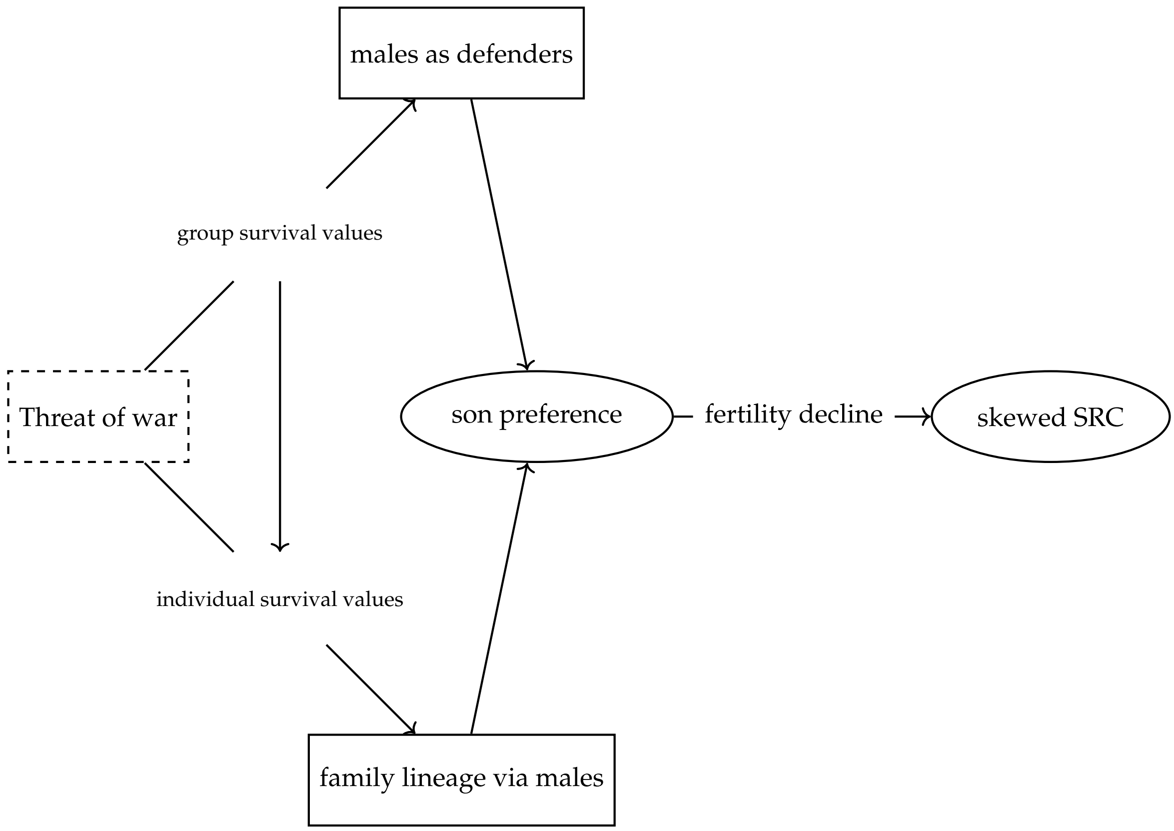
Note. Column 1 reports average treatment effects on the treated based on matching using entropy balancing. Column 2 and column 3 report the average treatment effects from kernel and radius matching estimators with bandwidth = 0.0009 and with standard errors calculated from bootstrapping with 50 replications; the propensity scores are calculated using the entire set of comprehensive controls from Table 6; * denotes significance at the 10 percent level; ** 5 percent level; *** 1 percent level.

Table 8: Test of omitted variable bias

	(1)	(2)
Specification	$\delta_{R_{max}=min\{1.3\tilde{R},1\}}$	$\left[\tilde{\beta},\beta^*_{\left(R_{max}=min\{1.3\tilde{R},1\},\delta=1\right)}\right]$
BASELINE CONTROLS	9.341	$[0.114, 0.105]$
COMPREHENSIVE CONTROLS	9.730	$[0.101, 0.093]$

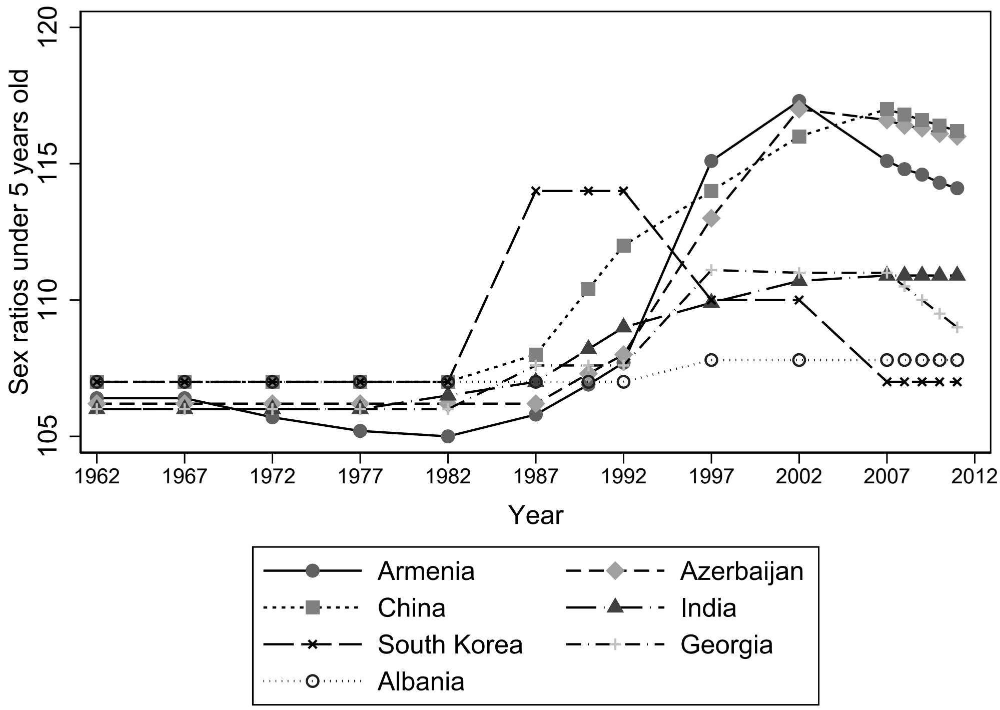
Note. δ indicates the value of a proportional selection of unobservables to observable assuming the maximum value of theoretical R^2 is R_{max}. The coefficient bounds are calculated assuming the unobservables are as important as the observable in explaining the outcome variable (i.e. $\delta = 1$). Since the test can only be performed with linear models, a linear probability model is used to estimate the probability of having a son bias instead of a probit model.

Figure 1: Threat of war, son preference and sex ratios at birth



Note. Conceptualization on the linkages between the threat of war, exacerbation of son preference, and the behavioral response reflected in skewed sex ratios at birth.

Figure 2: Trends in under-5 sex ratios, selected countries

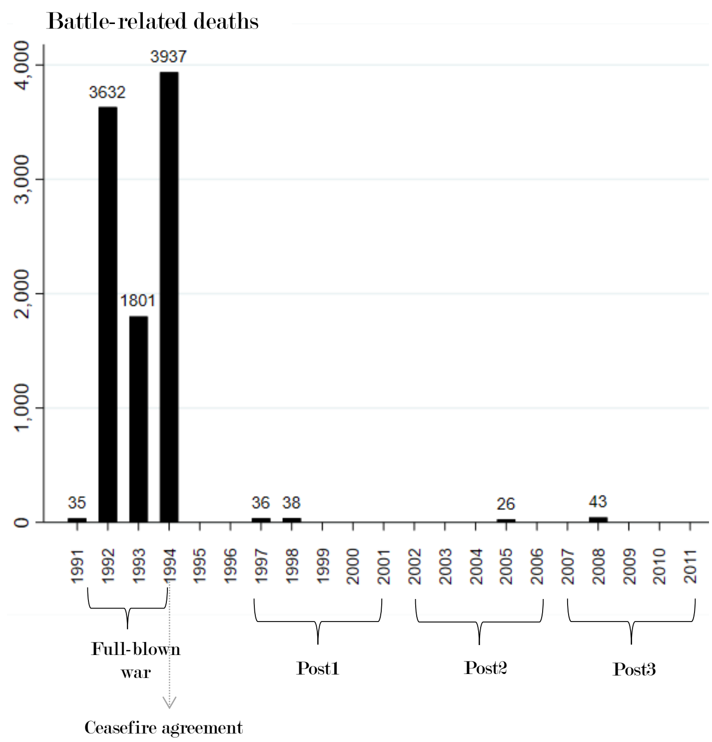


Source: World Development Indicators (World Bank, 2015). Sex ratios at birth (male over female births) from 1962 to 2012 for countries with relatively high values, 5-year averages.

Figure 3: Nagorno Karabakh conflict: Map of unresolved territorial disputes

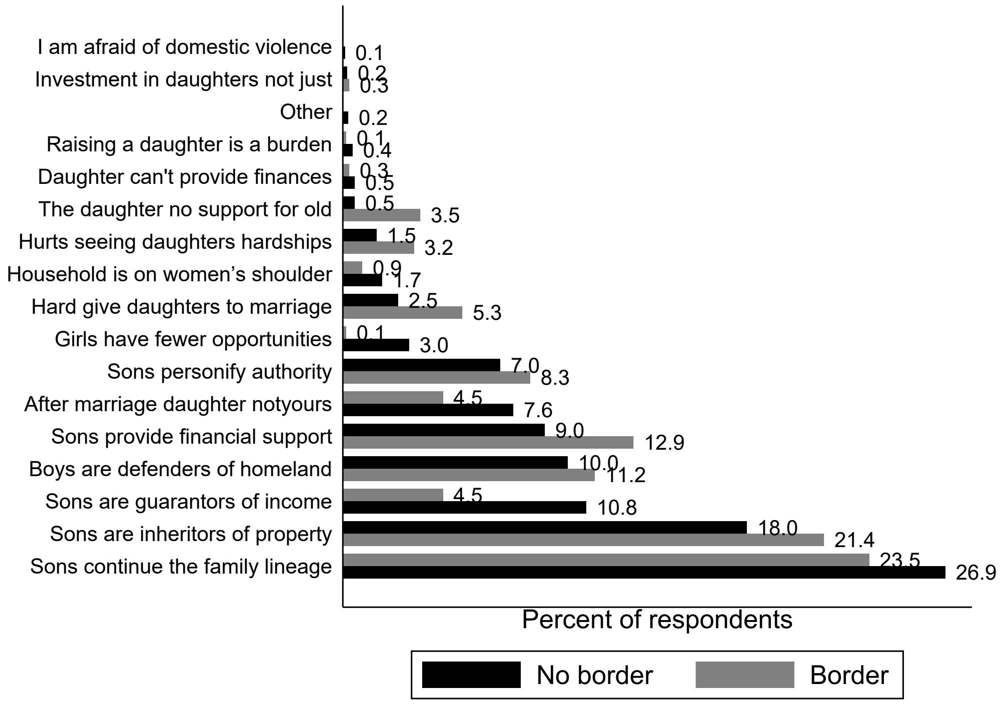
Source: De Waal (2010), p.161. Note. The map depicts the internationally recognized borders of Republic of Armenia, the disputed Nagorno Karabakh region and adjasent territories controlled by Armenian troops from 1994 to 2020. Stepanakert is the capital of the disputed Nagorno Karabakh region. The line of contact denotes the de-facto borders between Armenian and Azerbaijani sides.

Figure 4: Nagorno Karabakh conflict: Battle-related deaths, 1987-2011



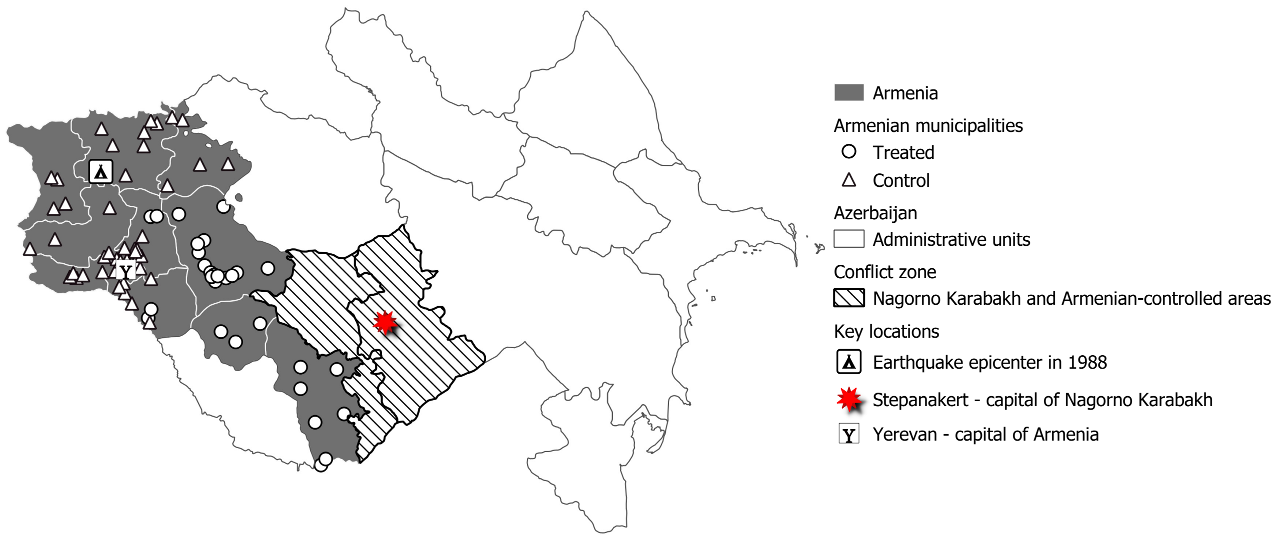
Source: Uppsala Conflict Data Program (UCDP) (Pettersson and Eck, 2018; Sundberg and Melander, 2013). Note: Post1, Post2 and Post3 indicate first, second and third (5-year) period after the ceasefire agreement in 1994.

Figure 5: Reasons for son preference: Armenian households, 2011



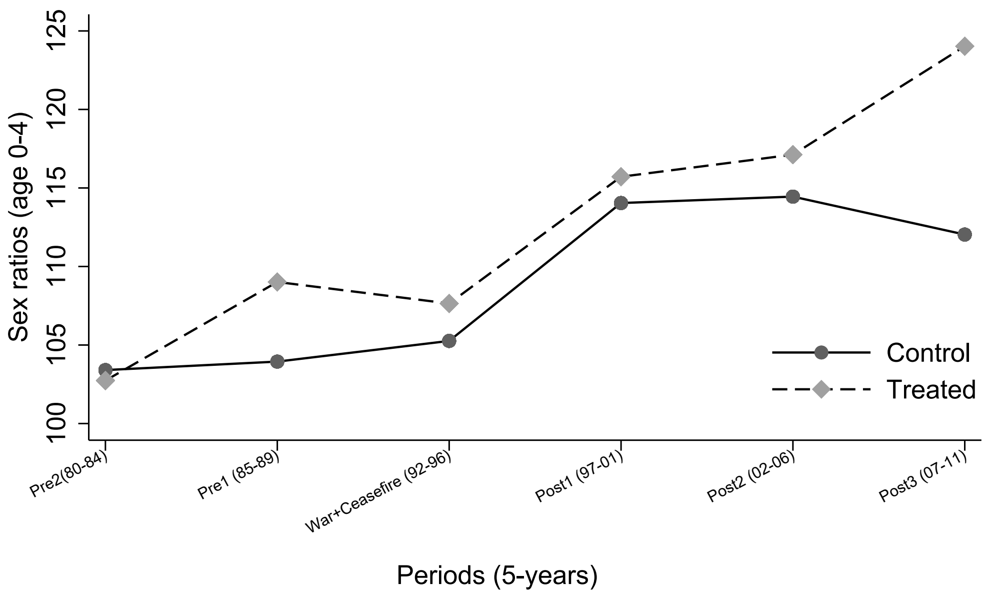
Source: UNFPA project on Strengthening Sexual and Reproductive Health Services (Abrahamyan et al., 2012). Note. Responses are based on a survey question asked to women who express son preference: "Why does your family give preference to sons rather than to daughters?" More than one answer could be given. The responses are compared among the administrative units (marz) that have a border with Nagorno-Karabakh/Azerbaijan (gray bars) and those that do not (black bars). The total number of respondents is 2762 (women).

Figure 6: Treated and control municipalities



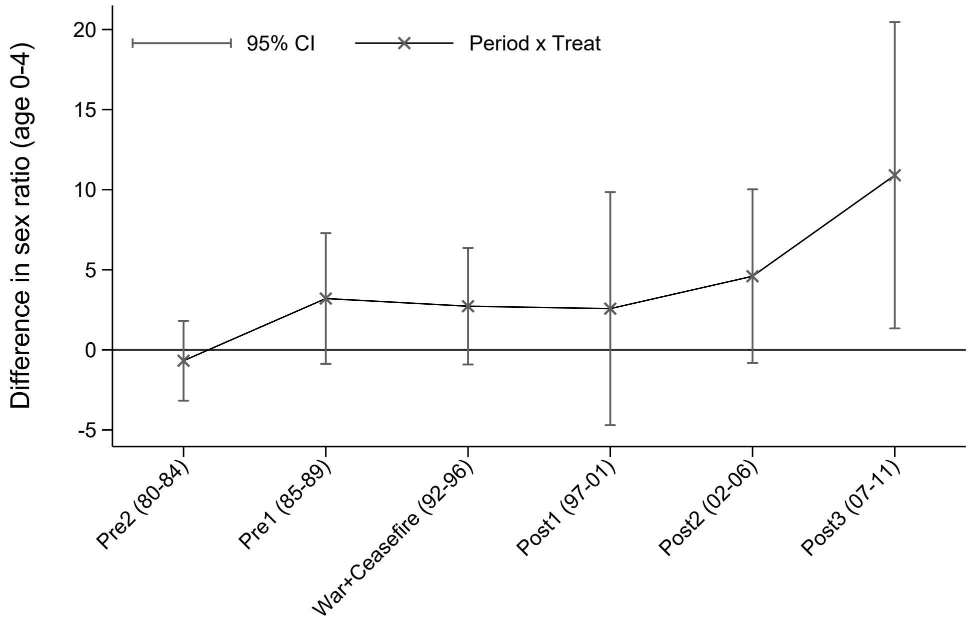
Note. The map displays the location of treated and control municipalities in the sample. All the municipalities are within the internationally recognized borders in Armenia. The location of three important cities Yerevan - the capital of Armenia, Stepanakert - the capital of Nagorno Karabakh region and Gyumri - the placebo location, are depicted on the map.

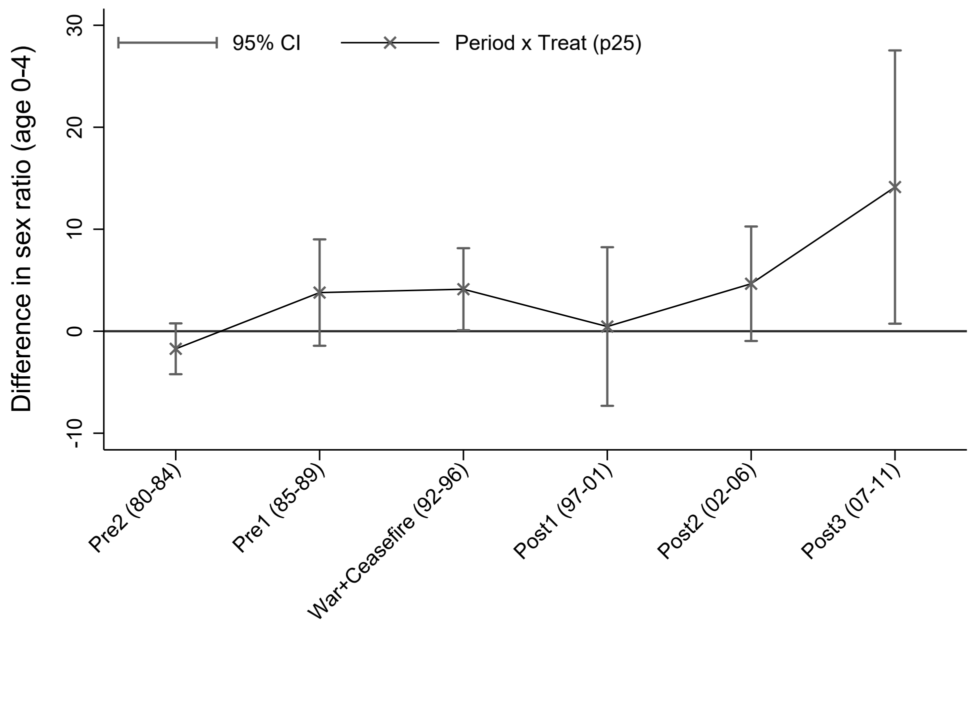
Figure 7: Trends in under-5 sex ratios in treated and control municipalities



Note. The trends show under-5 sex ratios in the municipalities in our sample that have less than average distance (Treat) and larger than average distance (Control) to Stepanakert - the capital of Nagorno Karabakh region.

Figure 8: Difference in pre-trends of treated and control municipalities





Note. Fully flexible specification. Interaction of treatment variable (less than average municipality distance to Stepanakert) with each period dummy. Confidence intervals are at the 95 percent level. The sample size includes 76 municipalities and 363 observations.