Sibling Rivalry and Migration*

Massimiliano Bratti† Simona Fiore‡ Mariapia Mendola§

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Abstract

This paper examines the effects of family size and demographic structure on offspring's international migration. We use rich survey data from Mexico to estimate the impact of sibship size, birth order and sibling composition on teenagers' and young adults' migration outcomes. We find little evidence that high fertility drives migration. The positive correlation between sibship size and migration disappears when endogeneity of family size is addressed using biological fertility (miscarriages) and infertility shocks. Yet, the chances to migrate are not equally distributed across children within the family. Older siblings, especially firstborns, are more likely to migrate, while having more sisters than brothers may increase the chances of migration, especially among females. [*JEL codes:* J13 F22 O15.]

Keywords: International Migration, Mexico, Family Size, Birth Order, Sibling Composition.

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†Università degli Studi di Milano and IZA, email:massimiliano.bratti@unimi.it

‡Università degli Studi di Bologna, email:simona.fiore3@unibo.it

§Università degli Studi di Milano–Bicocca and IZA, email:mariapia.mendola@unimib.it

1 Introduction

Migration is one of the most important ways through which workers can increase their income opportunities as well as improve their families' welfare, especially in developing countries (Clemens et al. 2009, Gibson and McKenzie 2012). Given the profitable nature of labor mobility, which involves both the migrant and her origin family, an extensive literature on the determinants of migration has emphasized the important role of household (along with individual) factors in the migration decision (e.g. Rosenzweig and Stark 1989, Stark 1991). In many developing countries labor migration is a family strategy to diversify income sources, improve earning potentials and increase household security through remittances (Clemens 2012, Yang 2008, Antman 2012, among others).

A key feature of migration is that it mainly involves young adults who are more likely to have a positive net expected return on migration due to their longer remaining life expectancy (Sjaastad 1962). According to recent UN figures, international migrants aged 15 to 24 in the world account for 12.5 per cent of total migrants worldwide, and when migrants between the ages of 25 and 34 are added, young migrants represent over 30 per cent of the total (UNDESA 2011). The proportion of youth migrants is much higher in developing countries than in developed ones and it more than doubles if we consider internal migrants as well (UN 2013).

As a result, family migration strategies in developing countries may involve the costly parental decision to dispatch one of their children to work in a different city or abroad, or to invest in a potentially remitting child (Lucas and Stark 1985, Jensen and Miller 2011). Yet, parents face a number of trade-offs when allocating resources among their children, due to either limited household resources or (perceived) different returns to the migration investment (e.g., pro-male bias) (Garg and Morduch 1998, Black et al. 2005).[1] Although the determinants of migration have already been studied extensively, far less is known about the role of the origin household's structure and especially sibship size on migration investment decisions. This is a surprising gap given the popular view that migrants come from high-fertility countries and typically leave behind several household members who oftentimes are

[1] A well-established theoretical literature in economics rationalizes a causal link running from children's economic resources to their lifetime opportunities and their adult outcomes (Becker and Tomes 1976, Shultz 1990, Thomas 1990).

siblings (Hatton and Williamson 1998).

To the best of our knowledge, this is the first paper to assess the causal effect of demographic characteristics of one's childhood household, i.e. sibship size, birth order and composition of siblings (by gender and age), on the likelihood to migrate abroad. We address this question in the context of Mexico, one of the largest migrant-sending and remittance-recipient countries worldwide. Mexican emigrants, mostly leaving to the US, have surged in the last two decades accounting from 5.2 percent of Mexico's national population in 1990 to 10.2 per cent in 2005 (Hanson and McIntosh 2010). Importantly, Mexican migration patterns differ by age and gender. As shown by Hanson and McIntosh (2010) by using Mexico's population censuses, a significant fraction of males migrate by age 16 with emigration increasing sharply until around age 30 and decreasing thereafter, presumably as as result of return migration. For women instead, there is less emigration by age 16, with subsequent rates being relatively stable over the course of their lives. [2]

By using two waves of a large and nationally-representative demographic household survey, we focus on the determinants of migration of Mexican adolescents and young adults in the age range 15-25. Our large dataset allows us to overcome limitations of small samples of children and includes detailed information on fertility histories, infant and general mortality. Importantly, it further allows us to address the potential endogeneity of parental fertility choices which arises from the fact that families who choose to have more children may also be those who value child out-migration more. This is so as, in a context such as Mexico with limited institutions and imperfect credit or insurance markets, children may be viewed as a way to acquire old-age security and support (Becker 1960, Cigno 1993, Ray 1998).[3] Thus, the lure of international migration to the US may increase the likelihood of acquiring such a social security and hence increase fertility (Stark 1981). We address this endogeneity issue by exploiting exogenous variation in family size induced by either infertility shocks or miscarriage before first birth (Agüero and Marks 2008, Miller 2011). We further investigate birth order, sibling-sex and sibling-age composition effects on migration by using family fixed effects, i.e. by exploiting within-family variation.

[2]See Figure 2 reported in Hanson and McIntosh (2010).

[3]We use data from Mexico in the mid 1990's when the country was classified as a developing economy. During that decade the country was hit by a severe financial crisis so that all indicators for wage, income, consumption and wealth inequality fared even worse.

We find no evidence that high fertility drives migration. The positive correlation between fertility and migration disappears when the potential endogeneity of sibship size is addressed. On the other hand, the chances to migrate are not equally distributed across children within the same family. Older siblings, especially firstborns, are more likely to migrate, while having more sisters than brothers may increase the chances of migration, especially among females. Results are robust to several changes in both the estimation sample and the estimation strategy.

Our findings have relevant implications. First, by easily observing a positive association between fertility and economic migration, implications may be drawn that smaller families will lead to lower rates of mobility. As a consequence, one may expect a negative impact on migration from fertility-reducing programs – such as investments in family planning, sex and reproductive health — which have been endorsed in many developing countries as a policy response to the apparent vicious cycles of high-fertility, poverty and economic stagnation (Miller and Babiarz 2014, Schultz 2008b). Some of these programs have been implemented in high fertility societies with significant out-migration rates, such as Mexico (Cabrera 1994). Yet, we provide little evidence that the causal relationship goes in this direction. Second, our empirical findings hint to the fact that parental investment in offspring's migration may matter for fertility decisions (i.e., fertility is endogenous) in a context of poor resources and high emigration opportunities. The reason is that, in developing countries, offspring are the primary caretakers of parents and they may do so by providing support to their origin family through emigration and spatial diversification in residential location.

The paper unfolds as follows. Section 2 describes the link between household structure and migration as considered by the related literature on human capital investment. Section 3 presents the data and sample selection. The methodology and empirical strategy is described in Section 4. Section 5 presents our main results on birth order and sibship size estimates, and Section 6 some robustness checks. Finally, Section 7 summarizes our main findings and concludes.

2 Related literature

Standard economic theory conceives labor migration as an investment in human capital whereby relocation requires up-front resources followed by a positive payout in the future (Sjaastad 1962, Schultz 1972, Dustmann and Glitz 2011). Positive returns on migration are coinceived in terms of both migrants' earnings and remittances sent back home (Clemens et al. 2009, Yang 2008, Amuedo-Dorantes and Pozo 2011, Stark and Bloom 1985). Recent evidence shows that workers moving from a poor to a rich country can experience immediate, lasting, and very likely increases in earnings, even for exactly the same tasks (Gibson and McKenzie 2012, Ashenfelter 2012). Even though the original economic model of migration does not distinguish between personal and family decisions about migration, early research focused on the individual investment decision (Borjas 1987; 1991). Accordingly, migration choices depend on individual-level characteristics — e.g. age, gender, skills — which influence personal discounted net-gain from moving (Chiquiar and Hanson 2005). Yet, more recent contributions in the migration literature, especially from developing countries, point at the household as the main unit for migration choices (Stark and Levhari 1991, Katz and Stark 1986). The core feature of this collective decision-making framework is that the family, unlike the individual, aims at maximizing household income and therefore can take the costly decision to dispatch one (or more) young members to work in foreign labor markets in order to receive remittances (Stark and Bloom 1985). In the absence of well-functioning credit or insurance markets then, migration can be a household diversification strategy whereby one or more members are sent abroad while others are assigned to work in the local economy (Rosenzweig 1988). By the same token, migrant members can be considered as financial intermediaries providing the origin household with a potential source of insurance as well as liquidity through remittances (Stark 1991). Empirical evidence on the implications of migration as a family security strategy in developing countries is abundant (see Ratha et al. 2011, for a review). Rapoport and Docquier (2006), for instance, survey the different motives for remittances sent by migrants, which are found to be used also as a form of support to the elderly (see also Clemens et al. 2014). By using data from Mexico, Antman (2012) shows that children migrated to the US (strategically) provide financial contributions to health cares for their parents (see also Stöhr 2015). At the same time though, little evidence exists on the degree to which family environment — in particular family size[4] and composition — affects children's out-migration decisions.

The link between the household structure and parents' investments on the human capital of their children has received substantive attention in the Household Economics literature. Theoretical models of fertility choices have been widely influenced by the argument of the 'quantity-quality (Q-Q) trade-off'. The Q-Q model treats the quantity and quality of children in a similar fashion as other 'consumption goods' in the household so that, in the absence of parental discrimination between children, there is a trade-off between child 'quality' or outcomes, and the number of children within a family (Becker 1960, Becker and Lewis 1973, Becker and Tomes 1976). However, in many of today's developing countries (as well as in rich countries around the time of their industrial revolution) parents have often used their children as a substitute for missing institutions and markets, notably social security in old age (Cigno 1993, Ray 1998).[5] According to this framework — known as the old-age security hypothesis — on top of the 'consumption-good' aspect of children, fertility choices are influenced by the child role of 'investment-good' or household 'asset' (Neher 1971, Caldwell 1976, Cigno 1993). Chidren embody income-earning possibilities both for themselves and for their parents, and this may be the reason why in weak institutional contexts (i.e. with poor formal markets and social safety-nets) people generally choose to invest in their future in the form of children (Duflo and Banerjee 2011).

Empirical evidence on the role of household size and composition on parental investments in other forms of human capital (such as education or health) has received substantive attention (Garg and Morduch 1998, Manacorda 2006, Black et al. 2005), whereas within-family considerations have been less analyzed in the context of migration decisions.[6] Yet, if migration is costly and migrants move at a relatively young age, it is plausible that migration is the result of family decision-making in which parents decide on their children's relocation

[4] We use family size and sibship size as synonymous, the first takes the point of view of parents, the second of children.

[5] Recent contributions on contemporary developed societies show that when pensions and income from retirement decrease, the old-age security motive matters for fertility decisions even in these settings (see Gábos et al. 2009, Billari and Galasso 2014).

[6] Findings on the impact of family size on child outcomes are mixed. Early results tended to predominantly show that children from larger families have worse outcomes, especially in terms of human capital investment and earnings (Rosenzweig and Wolpin 1980, Hanushek 1992, Parish and Willis 1993). Yet, after controlling for the endogeneity of fertility, in more recent papers family size turns out not to adversely affect child outcomes (see Black et al. 2005, Angrist et al. 2010, Fitzsimons and Malde 2014, among others).

(potentially retaining some control over their children's earnings as well), or children are influenced by their family background (e.g. household characteristics, number of siblings) while deciding to move. Thus, in families with limited resources and more than one child to raise, greater sibship size may *negatively* affect child out-migration through a resource dilution effect (i.e. a smaller share of resources per child) or because more family-work is needed at home, e.g., care for younger children (Becker and Lewis 1973, Giles and Mu 2007). On the other hand, more children may increase the average maturity level of the household and allow a reallocation of resources from children to parents so that young household members are dispatched abroad in order to send remittances or offer potential support back home. In particular, if children contribute to family income either through child-labor, economic diversification or parental-care, a larger number of siblings may have a *positive* effect on the out-migration of one (or more) of them (Brezis and Ferreira 2014, Stöhr 2015). The relative strength of these competing forces is ultimately an empirical question. This is what we turn to in the following sections.

3 Data and sample selection

This study uses data from the 1992 and 1997 waves of the *Encuesta Nacional de la Dinámica Demográfica* (ENADID), conducted by the National Institute of Statistics and Geography (INEGI) in Mexico. Each ENADID's wave surveys more than 50,000 households from all over the country and is representative of the Mexican population. The dataset is very reach and unique in collecting comprehensive information on women's fertility as well as migration history of all household members, along with standard socio-economic characteristics. Importantly, by using detailed demographic information on age (month and year of birth) and gender of individuals in the same household with the same mother, we are able to identify all biological families within the sample and recover complete information on the number and gender of all siblings (also those not currently living in the household of origin).

The ENADID allows us to define household members' international migration experience from three separate questions, i.e. (i) whether there is any household member (even temporarily absent) migrated abroad during the five years prior to the survey; (ii) whether any household member has ever worked in or looked for work in the United States (and the year in which this occurred); (iii) whether the respondent reports a period of residence abroad at any point in time prior to the survey. The use of these three different sources of information for migration episodes ensures that we are able to capture a relevant part of the phenomenon.[7] Overall, almost 18 percent of households in Mexico reports having a member migrated abroad in 1997, and up from 15 percent in 1992.

Since we are interested in the effect of family size on parental investment in offsprings' migration, we define individual migration episodes as *non-tied* migration, i.e. we exclude from the sample children who experienced episodes of migration joint with their parents and those whose parents have an international migration experience.[8] Figure 1 reports the incidence of non-tied migration by age and gender in Mexico and shows that, overall, migrants are clearly concentrated (more that 70%) in the age range 15-25. Hence, throughout our analysis we restrict the sample to individuals aged 15 to 25. This is also consistent with the argument that Mexican youngsters finish compulsory schooling and potentially enter the labor market at the age of 15, whereas beyond the age of 25 they are more likely to make their own life out the origin family.[9]

[Figure 1 about here]

The ENADID further collects detailed information on fertility for all women aged 15 to 54 at the time of the survey. Women answered specific questions on the number of the children ever born, their gender and birth order, current and past contraceptive use, fertility preferences, and their socioeconomic and marital status. Such information allow us to construct our key explanatory variable, that is the total number of biological siblings of each individual in the sample. Moreover, it enables us to identify parental exogenous shocks to fertility induced by self-reported infertility episodes and miscarriage at first pregnancy (see

[7] By containing information on migrants who have either returned to Mexico, or who have at least one household member remaining in Mexico, excluding households which have migrated abroad in their entirety, the ENADID tends to underrepresent permanent tied migrants (see also Hanson 2004, Mckenzie and Rapoport 2007). Yet, the latter form of potential selection is of little concern to us since our main outcome of interest is the effect of family size on parental investment in children's migration, so that we do need to exclude 'family migration' and focus on households left behind by one or more migrant member.

[8] Yet, we investigated the robustness of our findings to the inclusion of tied-migrants, including parents' migration status among the controls (results available upon request). In their study of Mexican migration towards the US, Cerrutti and Massey (2001) found that nearly half of all male migrants left for the US before or without a wife or a parent.

[9] Yet, our findings are also robust to the sample cut on individuals aged 15 to 35. Results are available upon request.

Section 4.2 for more details on this). In line with the medical definition of infertility and with the previous literature (Agüero and Marks 2011), we restrict our sample to children of non-sterilized women who are not currently using contraceptives or who never did (about the 80% of the sample).

Our final estimation sample is made of 26,723 children in the age range 15-25, among whom 5.2 percent are migrants. Main individual characteristics are reported in Table 1, according to their migration status. Migrants are mostly males (75 percent) and they report significantly more brothers and sisters than non migrants. Moreover, migrant children appear to be slightly older and to live in less educated but richer (in terms of income) households with respect to non migrant youngsters.

[Table 1 about here]

4 Empirical strategy and identification

4.1 The sibship size and birth order effects

In our analysis we are interested in the effects of sibship size and composition on the individual likelihood to migrate. In order to estimate the effect of sibship size, though, we need to control for children's birth order (see, for instance Black et al. 2005). Let us assume that parents have a preference for the first children they have and invest comparatively more resources in them. In this case, a spurious negative correlation between sibship size and human capital investments may emerge just because in larger families we also find children with higher birth orders. In other words the two variables do not have independent variation, e.g. while we can assess the effect of family size on firstborns by looking at the outcomes of the firstborns in families of different sizes, we cannot look at the outcome of a fourth-born child when sibship size changes from two to three as fourth born children are found in larger families.

Recently, Bagger et al. (2013) have proposed a theoretically-grounded methodology to disentangle the two effects. We employ a two-step estimation strategy which draws on their idea. In a first step we estimate the following regression:

$$M_{ij} = \alpha_0 + \sum_{k=1}^{K} \alpha_{1k} bo_{ijk} + \alpha_2 \mathbf{X}_{ij} + u_j + \varepsilon_{ij} \tag{1}$$

where the outcome variable M_{ij} pertains to the migration status of child i in household j and is a dichotomous indicator for either current or past migration experiences abroad. bo_{ij} is a dichotomous indicator for the child being of birth order $k = 1, ..K$ where K is the maximum number of children in the families in our sample (so as the maximum sibship size is $K-1$); $\mathbf{X}_{ij}$ is a vector of individual covariates including child gender, age, age squared and cohort indicators (one for each year of birth)[10]. u_j is a family fixed effect, and ε_{ij} an idyonsincratic error.

The effect of sibship size is captured in equation (1) by the family fixed effects. The inclusion of family fixed effects also helps address the birth order endogeneity (cf. Bagger et al. 2013). The birth order fixed effects capture the differences in the probability of migration between children of different order within the same family. Only within-family variation is exploited in these estimates, and birth order effects are not contaminated by between-family variation in family sizes, i.e. the fact that children in larger families also have higher average birth orders.

In a second step, we substract the birth order effects from the dependent variable, i.e. we compute the difference $\hat{NM}_{ij} = M_{ij} - \sum_{k=1}^{K} \hat{\alpha}_{2k} bo_{ijk}$ where NM stands for 'netted migration', and use it as the dependent variable in a second step in which the following equation is estimated:

$$\hat{NM}_{ij} = \beta_0 + \beta_1 S_{ij} + \beta_2 \mathbf{X}_{ij} + \beta_3 \mathbf{W}_j + v_{ij} \tag{2}$$

where S_{ij} is sibship size. The coefficient β_1 captures the effect on migration of being grown in a family with sibship size S_{ij} for the 'average child' in this family, i.e. irrespective of his/her birth order. $\mathbf{X}_{ij}$ is a vector of individual covariates defined as above and $\mathbf{W}_j$ includes family background characteristics such as mother's and father's age and age squared, and mother's and father's years of completed education. We also control for the father not being

[10]We can include both a control for age and birth cohort indicators since we are using two cross-section surveys.

in the household (i.e. widowed and divorced single-mother families) and municipality fixed effects, which also capture the rural vs. urban residence along many other factors related to different local cultural or economic conditions, access to contraception, etc. We estimate equation (2) by using a linear probability model (LPM) and employing both a linear specification and flexible dummy indicators for sibship size. Since the dependent variable has been generated by a regression, we correct standard errors by weighting the estimation with the inverse of the variance of $\hat{NM}_{ij}$. We also cluster standard errors at the household level as to account for potential error correlation across siblings.

4.2 The sibship size effect: Identification strategy

If the number of children and investment in child out-migration are both outcomes over which parents exercise some choice, the sibship size estimation as in equation (2) will generate spurious evidence. In other words, parental fertility may be endogenous with respect to children's migration. It is plausible, for instance, that the opportunity to send some children abroad modifies parents' fertility choices. In developing countries, children are a valuable asset for parents and a source of old-age support. If offspring's migration opportunities are not equally distributed across families, it may happen that households with lower migration costs or higher benefits for their members will also decide to have more children. Alternatively, unobservable parental preferences for children and old-age support through migration may co-vary positively. Stark (1981) and Williamson (1990) postulate, for instance, that heterogeneity in parents' preferences for childbearing and for migration are systematically related, and in a context such as Mexico where migration cum remittances is an essential lifeline to households of origin, they are generally positively related. In both these cases the positive association observed between fertility and child out-migration is likely to overstate the true causal relationship. This pattern of preferences' or migration costs' heterogeneity would lead to a larger positive correlation between fertility and child out-migration than it would be observed if fertility changes due to exogenous shocks.

Hence, to clearly identify the relationship between sibship size and migration, a presumably exogenous source of variation in family size is needed. ENADID allows us to identify self-reported infertility from specific questions asked to non-sterilized women who never used contraceptives or who are not currently using them. More specifically, we construct an indicator variable for infertility shocks taking value one if the woman declares she is not currently using contraception or she has stopped using the previous method because of infertility ('infertility shock') and zero otherwise. ENADID also enables us to build a second indicator variable which equals one if a woman experienced a miscarriage at first pregnancy ('fertility shock') and zero otherwise.

In order for our identification strategy to be valid, the two instruments must satisfy three conditions — i.e. exogeneity, exclusion restriction assumption and relevance— which we discuss below.

Infertility or subfecundity conditions have been already used in the economic literature to estimate the effect of family size and fertility timing on mothers' labor market outcomes (see, for instance, Agüero and Marks 2008; 2011, Schultz 2008a). There is evidence that infertility is independent of the background characteristics of infertile women. For example, variables such as father's social class and parity have been shown to be unrelated to observed heterogeneity in fertility (Joffe and Barnes 2000). In an article summarizing the epidemiological literature regarding the role of lifestyle factors (cigarette smoking, alcohol and caffeine consumption, exercise, BMI, and drug use) on female infertility, Buck et al. (1997) conclude that few risk factors have been assessed or identified for secondary infertility. Also, education, occupation, and race have been shown to be unrelated with impaired fecundity using US data from the National Survey of Family Growth (Wilcox and Mosher 1993) . By using data on a large set of developing countries, Agüero and Marks (2011) present evidence that infertility is generally uncorrelated with background characteristics of women, with a few exceptions such as women's education and rural residence, which therefore become important controls in the regression model.

Also miscarriages and stillbirths have been used to identify fertility *tempo* and *quantum* effects on women's labor market outcomes, mainly in advanced countries (Hotz et al. 2005, Miller 2011, Bratti and Cavalli 2014, Karimi 2014). Their exogeneity is generally supported by the medical literature. Miscarriage or spontaneous abortion usually refers to any pregnancy loss that takes place before the 20th week of pregnancy. For their nature, miscarriages should have a negative effect on total fertility, and in our context on sibship size.[11] Only

[11] According to Bongaarts and Potter (1983) overall spontaneous loss rates are about 20 percent of recognized

two etiological factors for miscarriage are recognized by different authors in the obstetric literature, i.e. uterine malformations and the presence of balanced chromosomal rearrangements in parents (Plouffe et al. 1992). The latter though, are unlikely to be correlated with women's attitudes towards offspring' migration. The number of miscarriages and stillbirths will generally increase with the number of pregnancies, which depend in turn on desired fertility, and this could potentially generate a spurious positive correlation between the number of miscarriages and observed fertility. For this reason, we only consider miscarriages occurred at the first pregnancy (Miller 2011). Moreover, there is a potential issue of measurement error with this instrument, since women may be unaware of miscarriages or may fail to recall them. Misreporting would generally affect the power of the instrument, but we do not expect any specific pattern of correlation between it and parents' attitudes towards child out-migration. Finally, the question, as it was formulated in ENADID, does not distinguish between voluntary and involuntary abortions. Thus, it may be the case that some of the reported abortions were actually voluntary, even though induced abortion was illegal in Mexico during the period under consideration.

In general, it is worth noting that compared with the literature that used fertility and infertility shocks to identify the causal effect of fertility on *mothers'* outcomes (e.g. in the labor market), we expect endogeneity issues to be less severe in our case as we require these biological shocks to be exogenous with respect to out-migration of *children*.

For our instruments to be valid, in addition to exogeneity, they have to satisfy the exclusion restriction assumption, i.e. fertility and infertility shocks must impact on children's migration only through sibship size. For this reason, in the child migration equation we control for many variables that may act as a confounding factor and those which may be affected by the shocks while having a direct effect on children's migration. Among these variables, we include mother's age, age at first pregnancy, education, marital status and husband's characteristics (age, education and absence).

Finally, in Figures 2 and 3 we report a preliminary visual representation of the relevance of our instruments (more compelling evidence is given by the first-stage of the IVs reported in Section 5). In particular, we use ENADID data to plot the total number of live births

pregnancies (i.e. one out of five). Casterline (1989) stresses how in most societies pregnancies losses produce a reduction of fertility of 5-10% from levels expected in the absence of miscarriages and stillbirths.

by women's age and infertility shock and miscarriage status. Figure 2 shows that women who experience an infertility condition generally have a lower number of children, and that differences in fertility tend to become evident after the age of 30. Similarly, Figure 3 displays a negative association between miscarriage at first birth and the total number of live births. Both figures suggest that our instruments are relevant. They also suggest that, even though the shocks we consider have a negative impact on fertility, overall Mexican women are able to achieve a generally high fertility rate by the end of their fecund life span.

[Figure 2 and 3 about here]

5 Results

5.1 Birth order effects

We start by estimating the impact of birth order on individual migration, as specified in expression (1), controlling for family fixed effects. The within-family estimator sweeps out all parental- and family-level heterogeneity, including completed family size. Moreover, family fixed effects account for potential endogeneity of birth order effects due, for example, to parental preferences for specific birth orders of children (birth-order selective child fostering), or for other omitted family-specific unobservable factors. The first column of Table 2 report estimates with a linear specification of birth order on the full sample, while in column (2) we allow for a more flexible specification by adding birth-order-specific dichotomous indicators. Regressions control for individual age and gender plus child cohort dummies (one for each year of birth).

First, in column (1) we observe that, after controlling for family fixed effects, birth order and individual characteristics, females are significantly less likely to migrate than males by 3.5 percentage points (p.p. hereafter). Moreover, from column (2) the birth order point estimate is negative and statistically significant. The effect starts to be economically significant from children of birth order 3, which are about 2.3 p.p. less likely to migrate than firstborns. Although this appears to be a small effect in absolute value, it roughly represents a 40 percent decrease in migration at the sample average (5.2 percent of migration rate). The coefficients for the following birth orders are larger in absolute value and peak for birth orders 9 and 10 or more (-16.6 and -20.4 p.p. respectively).

In columns (3) and (4) we estimate the same regressions as above by adding interaction effects between birth order and gender to the models.[12] Results show that birth-order interacted effects for females are generally not statistically significant. In other words, the birth order impact on the probability of migration does not significantly differ according to the child's gender. Overall, these estimates suggest that the chances of migration are not equally distributed across children within the same family. Older siblings are indeed more likely to migrate.

[Table 2 about here]

5.2 Sibship size effect: OLS and 2SLS results at the individual-level

In this Section, we turn to the estimation of the sibship size effects. By applying the two-step procedure outlined above, we start by reporting OLS estimates as a benchmark model, where the dependent variable is 'netted migration' (see Section 4.1). The number of siblings is tallied as the number of biological brothers and sisters of each child living in the household. The first column of Table 3 reports OLS results for a linear specification including sibship size. The highly significant coefficient implies that, on average and after controlling for birth order effects, adding one sibling is associated with a 0.9 p.p. higher likelihood to migrate of young adults (+17 percent at the sample mean). When we allow for differential effects by child gender (column 2), the significant negative coefficient for the interaction term indicates that the female likelihood to migrate increases less due to sibship size than for males. Specifically, one more sibling raises the migration probability more for sons than for daughters by 0.2 p.p.

[Table 3 about here]

Yet, as mentioned in the methodological section, the coefficients on sibship size reported in Table 3 are still likely to be biased, even when including a rich set of demographic controls. This is so as fertility may be endogenous with respect to child out-migration. Thus

[12] As our two-step procedure relies on family fixed effects, when estimating separate regressions by gender only families with at least two sons and at least two daughters can be included in the estimates for males and females, respectively. In order to avoid such sample selection, we rather adopt a pooled estimation including interaction effects with gender.

we employ an IV approach and exploit the arguably exogenous fertility variation generated by episodes of infertility and miscarriage. Since these events can vary the actual family size from the desired one, we use infertility shocks and miscarriage at first pregnancy to identify the effect of sibship size on child out-migration. In Table 4 we present two-stage least squares (2SLS) estimates using a linear specification with and without gender interaction effects. Throughout all of the models we employ the two-step methodology as outlined above to estimate equation (2). In columns (1) and (2) we instrument sibship size with an indicator variable for infertility shocks taking value one if the woman declares she never used or she stopped using contraception because of infertility.[13] In columns (3) and (4) we report results using a woman's experience of miscarriage on her first pregnancy as an instrument. Throughout all models, the first stage results point to a strong and highly significant relationship between infertility and fertility shocks and completed fertility. In particular, women who experienced an infertility shock have a reduction in their number of children of more than 0.7 (t=8.4) with an F-statistic of 70.14 (column 1). The negative impact of miscarriage on completed fertility is smaller (nearly 0.5) with an F-statistic of 20.10 (column 3).

The sibship size effects estimated using 2SLS are always small and statistically insignificant at standard confidence levels. The same holds for sons when we include interaction effects by gender (column 2 and 4). For daughters though, in the IV estimates using miscarriage as an instrument, we find that higher sibship size has a significantly positive effect on emigration (column 4). Yet, we cannot draw strong conclusions since the F-statistic for the interacted instrument is very low (5.41)[14]. Overall, findings in this section point to the little role of family size on children's migration outcomes. This evidence is not in line with the popular view that high-fertility in developing countries is a major cause of international emigration: according to our estimates this correlation is driven by unobservable variables which make some families more prone to both have more children and send some of them

[13] The interaction effect sibship size*female is instrumented using the interaction instrument*female, where the instrument is infertility or miscarriage depending on the specification.

[14] This specific finding for females differs depending on the instrument used. Thus, it seems to suggest that, unlike infertility, miscarriage is not a very good predictor of fertility of daughters' mothers. We interpret this result in terms of the different nature of fertility shocks the two instruments represent. Indeed, while infertility is a permanent shock to fecundity, miscarriage at first pregnancy is more unexpected and potentially temporary as a shock. Thus, women with a first-born daughter(s) may have some son preference and so are more likely to have an addition child (Fitzsimons and Malde 2014). The preference for sons has not such a deep social and cultural roots in Mexico as it is the case in some East and South Asian societies, for example, but there is some evidence that Mexico, especially at the time we consider, was a traditional and pro-male biased economy (WDR 2012)

abroad.

[Table 4 about here]

5.3 Sibling gender composition

Our estimates so far show that gender is a robust predictor of migration and, *ceteris paribus*, Mexican boys are systematically more likely to migrate to the US than girls. This points to a migration male-dominated phenomenon (e.g., Cerrutti and Massey 2001) that may be explained by (perceived) higher migration returns for boys (due to either higher expected wages abroad than at home or by lower moving costs for males with respect to females) or by a pure parents' preference for sons. In practice, if migration is costly and not all children are in the position to migrate, a pro-son migration bias may lead to a situation in which children compete for household resources in order to migrate and such 'rivalry' can yield gains to having relatively more sisters than brothers (Garg and Morduch 1998). Moreover, we also find that higher sibship parity decreases the likelihood to migrate. Thus, in order to explore the scope of sibling rivalry by gender, we test the effect of the sibling's gender by running two sets of regressions as reported in Table 5. First, we estimate migration equations with family fixed effects (i.e., conditioning on family size as given) on the full sample of children as a function of the number of their older brothers, while controlling for the number of older siblings, child's age (linear and squared) and birth order. Results in column (1) show that, while the number of older siblings does not significantly affect the likelihood to migrate, having an older brother (sister) instead of an older sister (brother) decreases (increases) the migration probability by 1.4 p.p. ($t = 3.6$). This result points to a significant role of the genders and ages of siblings in children's migration outcomes which does not differ significantly by the child's gender (column 2).

[Table 5 about here]

We further exploit the pro-male bias and the fact that siblings are likely to migrate in order of birth (with later born being less likely to migrate, as shown by our evidence above) to test whether boys and girls are treated equally within the household with respect to migration investments. We do so by including a control for having a next-born brother in the family fixed effects regressions on the pooled-sample (with and without interactive effects), as above. If a child has at least one younger sibling, the gender of his/her next-born sibling is random and a comparison of children with next-born brothers with children with next-born sisters, while controlling for older siblings composition, can identify the effect of the sibling's gender. [15] Results in columns (3) and (4) in Table 5 show that having a next-born brother does not play any role for sons, but reduces the likelihood to migrate for girls with respect to boys by 1 p.p. ($t = 1.7$). This result suggests that when parents decide the level of investment in their children's out-migration, the siblings' composition by gender and age matters. More specifically, from our results it seems that a first-born daughter with a next-born brother may be less likely to migrate than a first-born girl with a next-born sister. A possible reason is that when parents face the decision whether to send a daughter abroad, they may prefer to invest in the migration of the next-born brother. Consistent with this explanation, when looking at sample raw statistics, the average migration rate of daughters' next-born brothers and next-born sisters are 7 percent and 3 percent respectively. These results are in line with other evidence from developing countries that, when there are high returns to investing in the human capital of children but resources are limited, children may become rivals, even in the absence of any explicit strategic behavior on the part of any family member (Dunn and Plomin 1990, Morduch 2000). In particular, our findings are suggestive that children, especially girls, with relatively more sisters than brothers are more likely to migrate abroad than their peers.

6 Robustness checks: household-level estimates

In this section, we estimate the migration equation while using the household instead of the individual as the unit of analysis.[16] In doing so we are able to check the robustness of our baseline results to changing the estimation sample and strategy. Indeed, the two-step procedure reported above is based on household fixed effects and therefore can only be applied to households with more than one child in the full sample. By contrast, while focusing on the total number of migrants in the household as a function of total fertility, we do not need to control for birth order effects and we can use a standard IV procedure. As a

[15] A similar empirical strategy has been used in Vogl (2013) to study sibling rivalry over arranged marriages in South Asia.

[16] More precisely, our unit if analysis is the biological family.

consequence, household-level regressions allow us to include also one-child households in the sample.[17] Thus, we estimate a specification as follows:

$$n_j = \gamma_0 + \gamma_1 f_j + \gamma_2 \mathbf{W}_j + v_j \tag{3}$$

where the dependent variable is the number of migrants in household j and the independent variable of interest is f_j, i.e. the total fertility in household j. The coefficient γ_1 captures the increase in the number of migrants associated with a unitary increase in the number of children. Like in the child-level estimates, $\mathbf{W}_j$ includes family background characteristics such as the mother's and the father's age, age squared, and years of completed education, mother's age at first pregnancy, an indicator for the father not being in the household and municipality fixed effects; v_j is an household-level error term. This specification is estimated both with OLS and with IVs (namely two-stage least squares).

[Table 6 about here]

Results are reported in Table 6. Column (1) shows that a unit increase in the number of children is associated with an average increase in the number of migrants in the household of 0.018 (t=9). Computed at the average number of child migrants in the sample, this corresponds to a 23.2 percent increase of child migrants per household. Column (2) reports the IV estimate using the infertility instrument. The first stage shows a reduction of -0.77 (t=12.8) in the total number of children for women who experience an infertiliy shock, with an $F-$statistic of 163. The first-stage coefficient is very similar to that obtained in the child-level estimates (-0.73). In spite of the strength of the instrument, the second stage does not show any evidence of a positive effect of fertility on migration: the coefficient on the number of children turns out to be negative and statistically insignificant. Columns (3) reports the IV results using the variation in the number of children generated by miscarriage. Also in this case the first-stage coefficient is highly statistically significant and negative, with an $F-$statistic above 62. The negative impact of miscarriage on total fertility is smaller than the one exerted by infertility, yet it is quite large and precisely estimated, i.e. -0.52 (t=7.95). Like for the previous instrument, also in this case no significant effect is detected in the

[17]Thus, in these estimates we also exploit individuals who do not have siblings, and look at whether they are more (less) likely to migrate than individuals with siblings.

second stage.

The household-level estimates in this section confirm the results of Section 5.2 of a positive correlation between family size and migration, but of no causal effect of the former on the latter. Also in this case, as with individual-level estimates, the larger magnitude of OLS estimates relative to the IV ones points to an upward biased estimation because of endogeneity, i.e. families more likely to send young migrants abroad tend to have more children.

7 Conclusions

In this paper we provide novel and rigorous evidence on the extent to which international labor mobility is affected by the demographic conditions of the migrant's household. Migration is largely a youth phenomenon occurring in households which ever dispatch all of their children to work abroad. The 'resource dilution' hypothesis predicts that with larger sibship size, children's migration outcomes fall. Yet, in poor contexts, parents are likely to depend on their grown up children for the provision of care and income, and high rates of migration can significantly contribute to old-age living arrangements.

We use a rich household-survey dataset on teenagers and young adults to examine the causal effects of sibship size, birth order and sibling composition on migration outcomes in Mexico. Mexican migration, mainly to the US, is an enduring flow accounting for one third of total US immigration and one-tenth of the entire population born in Mexico. Importantly, migration patterns differ by age and gender, with a significant fraction of Mexican males migrating in the age between 15 and 30.

We focus on the determinants of migration of adolescents and young adults in Mexico. Our large dataset allows us to overcome limitations of small samples of children and includes detailed information on both women's fertility and household members' migration histories. We find little evidence that fertility has a causal impact on migration. The positive correlation between fertility and migration disappears when the potential endogeneity of sibship size is addressed using biological fertility and infertility shocks. On the other hand, we find differences in the chance to migrate between siblings within the same family (sibling rivalry). Older siblings, especially firstborns, are more likely to migrate, while having relatively more sisters than brothers systematically increases the likelihood to migrate, especially among girls.

Our findings adds to the migration literature by shedding new light on the household-level determinants of migration, which is one of the most important ways through which young adults can increase their human capital, especially in developing countries. Morevoer, we complement previous work in both advanced and developing areas that investigates the relationship between family environment and child achievements in terms of both human capital accumulation and labor supply.

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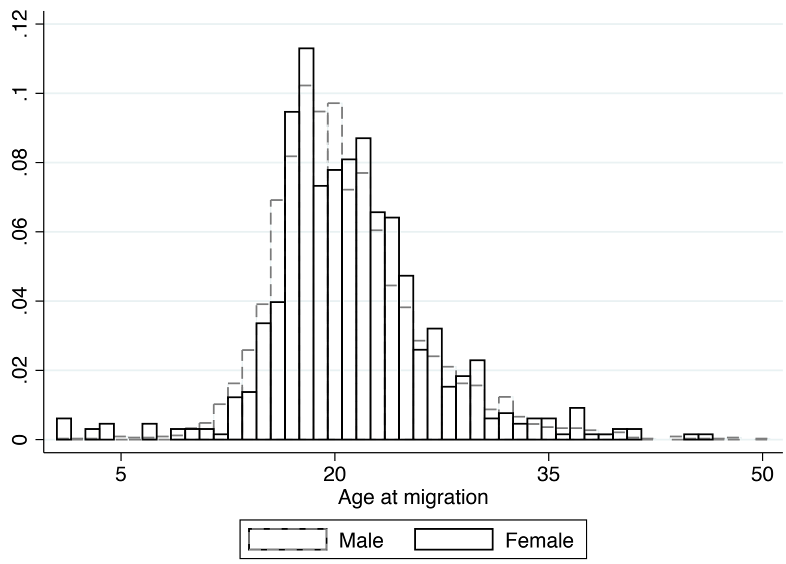
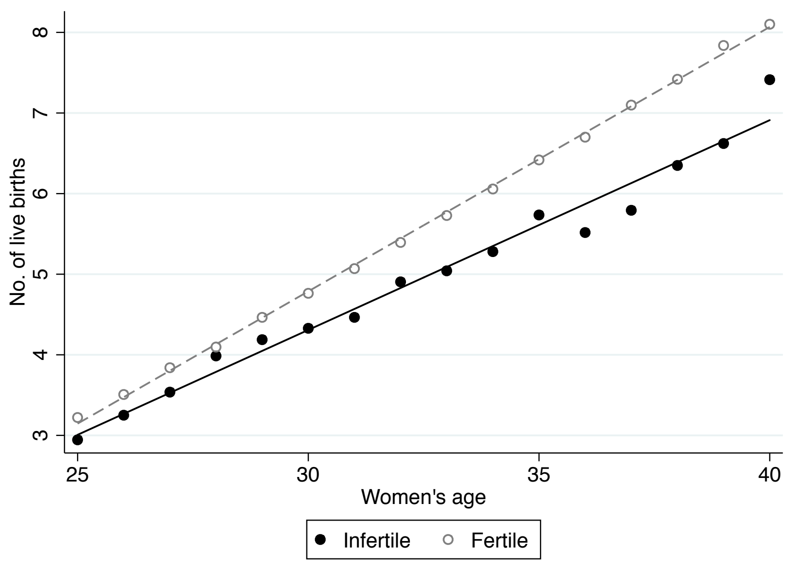
Figure 1: Mexican individual migration by age and gender



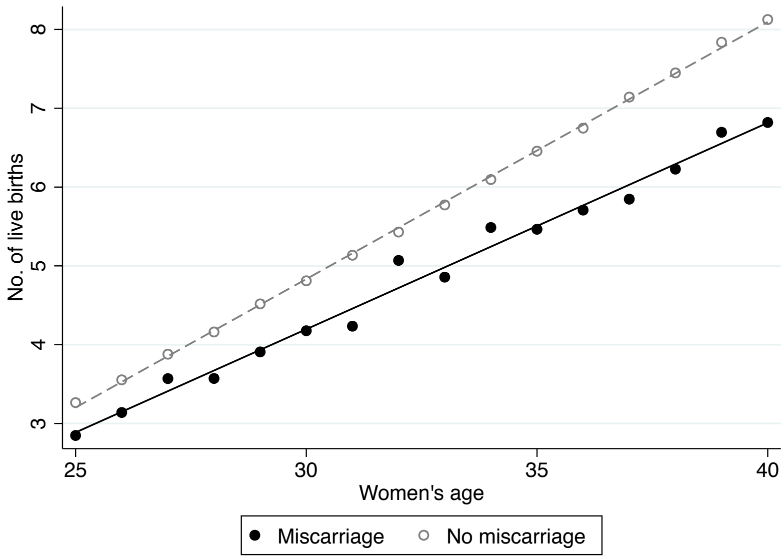
Note. Source: Our computations on ENADID, 1992 and 1997.

Figure 2: Cumulative number of children by women's fertile shocl status



Note. Source: ENADID, 1992 and 1997. This figure reports the total (cumulative) number of live births by women's fertile shock status and age (it refers to women belonging to different birth cohorts). Regression lines are super-imposed to the cross-plot.

Figure 3: Cumulative number of children by women's miscarriage at first pregnancy



Note. Source: ENADID, 1992 and 1997. This figure reports the total (cumulative) number of live births by miscarriage at first pregnancy and age (it refers to women belonging to different birth cohorts). Regression lines are super-imposed to the cross-plot.

Table 1: Individual characteristics by migration status

	Non-migrants	Migrants	P-values
Age	18.878	20.991	0.000
Female	0.458	0.251	0.000
Number of sisters	2.524	2.926	0.000
Number of brothers	2.600	3.060	0.000
Birth order 1	0.181	0.192	0.289
Birth order 2	0.231	0.225	0.595
Birth order 3	0.178	0.177	0.928
Birth order 4	0.138	0.154	0.063
Birth order 5	0.102	0.102	0.996
Birth order 6	0.071	0.073	0.776
Birth order 7	0.046	0.040	0.233
Birth order 8	0.028	0.021	0.103
Birth order 9	0.014	0.009	0.124
Birth order 10+	0.007	0.004	0.124
Mother's age	44.847	46.077	0.000
Mother's years of schooling	3.951	3.287	0.000
Mother's income	551.693	773.946	0.000
Father's age	49.026	51.431	0.000
Father's years of schooling	4.768	3.560	0.000
Father's income	2,047.259	3,260.356	0.000
Observations	25,331	1,392	

Note. The sample includes individuals aged 15 to 25 from 1992 and 1997 ENADID surveys. Monthly income is expressed in local currency and is only available in the 1997 ENADID wave.

Table 2: Birth order effects

	(1)	(2)	(3)	(4)
Birth order	-0.020***		-0.020***	
	(0.003)		(0.003)	
Birth order*Female			-0.001	
			(0.001)	
Birth order 2		-0.003		0.002
		(0.005)		(0.006)
Birth order 3		-0.023***		-0.025***
		(0.007)		(0.008)
Birth order 4		-0.040***		-0.036***
		(0.010)		(0.011)
Birth order 5		-0.070***		-0.072***
		(0.013)		(0.014)
Birth order 6		-0.089***		-0.080***
		(0.016)		(0.017)
Birth order 7		-0.116***		-0.109***
		(0.019)		(0.020)
Birth order 8		-0.140***		-0.145***
		(0.022)		(0.023)
Birth order 9		-0.166***		-0.172***
		(0.026)		(0.028)
Birth order 10		-0.204***		-0.193***
		(0.030)		(0.033)
Female	-0.035***	-0.035***	-0.033***	-0.032***
	(0.003)	(0.003)	(0.006)	(0.007)
Birth order 2*Female				-0.010
				(0.009)
Birth order 3*Female				0.006
				(0.010)
Birth order 4*Female				-0.009
				(0.010)
Birth order 5*Female				0.007
				(0.011)
Birth order 7*Female				-0.015
				(0.015)
Birth order 8*Female				0.010
				(0.018)
Birth order 9*Female				0.013
				(0.024)
Birth order 10*Female				-0.021
				(0.027)
Age	0.019**	0.019**	0.019**	0.019**
	(0.009)	(0.009)	(0.009)	(0.009)
Age squared	0.000	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Year of birth indicators	√	√	√	√
Family Fixed Effects	√	√	√	√
Observations	26,723	26,723	26,723	26,723

Note. The dependent variable is child migration. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.

Table 3: Sibship size effect: OLS estimates

	(1)	(2)
Number of siblings	0.009***	0.009***
	(0.001)	(0.001)
Number of siblings*Female		-0.002*
		(0.001)
Female	-0.029***	-0.022***
	(0.002)	(0.005)
Age	0.016**	0.016**
	(0.007)	(0.007)
Age suqared	-0.000	-0.000
	(0.000)	(0.000)
Mother's age	0.006	0.006
	(0.007)	(0.007)
Mother's age squared	-0.000	-0.000
	(0.000)	(0.000)
Mother's years of schooling	-0.000	-0.000
	(0.001)	(0.001)
Mother's age at first pregnancy	-0.005***	-0.005***
	(0.001)	(0.001)
Father's age	-0.002	-0.002
	(0.003)	(0.003)
Father's age _s	-0.000	-0.000
	(0.000)	(0.000)
Father's years of schooling	0.000	0.000
	(0.000)	(0.000)
Single mother	-0.202***	-0.202***
	(0.017)	(0.017)
Controls	√	√
Municipality Fixed Effects	√	√
Observations	26,723	26,723

Note. The dependent variable is a child's *netted migration* (see Section 4). Controls include year of birth indicators, mother's year of birth indicators and father's decade of birth indicators. Standard errors clustered at the household level are in parentheses. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.

Table 4: Sibship size effect: IV estimates

	(1)	(2)	(3)	(4)
Number of siblings	-0.010	-0.008	-0.003	-0.049
	(0.011)	(0.013)	(0.020)	(0.038)
Number of siblings*Female		-0.006		0.105**
		(0.013)		(0.052)
Female	-0.029***	-0.004	-0.029***	-0.460**
	(0.002)	(0.055)	(0.002)	(0.215)
Age	0.018**	0.018**	0.017**	0.025**
	(0.007)	(0.008)	(0.008)	(0.011)
Age squared	-0.000	-0.000	-0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Mother's age	0.003	0.003	0.004	0.005
	(0.008)	(0.008)	(0.008)	(0.009)
Mother's age squared	-0.000	-0.000	-0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Mother's years of schooling	-0.003*	-0.003*	-0.002	-0.002
	(0.002)	(0.002)	(0.003)	(0.003)
Mother's age at first pregnancy	-0.010***	-0.010***	-0.008	-0.009
	(0.003)	(0.003)	(0.006)	(0.006)
Father's age	-0.003	-0.003	-0.003	0.002
	(0.003)	(0.003)	(0.003)	(0.004)
Father's age squared	0.000	0.000	-0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Father's years of schooling	-0.001	-0.001	-0.001	-0.001
	(0.001)	(0.001)	(0.002)	(0.002)
Single mother	-0.218***	-0.219***	-0.212***	-0.211***
	(0.020)	(0.020)	(0.025)	(0.025)
First Stage				
Infertility	-0.728***	-0.805***		
	(0.0870)	(0.010)		
Miscarriage			-0.457***	-0.455***
			(0.102)	(0.123)
Angrist-Pischke multivariate F-stat	70.14	71.06	20.10	10.84
Infertility*Female		-0.822***		
		(0.125)		
Miscarriage*Female				-0.337**
				(0.144)
Angrist-Pischke multivariate F-stat		43.32		5.41
Controls	√	√	√	√
Municipality Fixed Effects	√	√	√	√
Observations	26,723	26,723	26,723	26,723

Note. The dependent variable is a child's *netted migration* (see Section 4). Controls include year of birth indicators, mother's year of birth indicators and father's decade of birth indicators. Standard errors clustered at the household level are in parentheses. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.

Table 5: Sibling composition effects

	(1)	(2)	(3)	(4)
Number of older brothers	-0.014***	-0.016***	-0.018***	-0.017***
	(0.004)	(0.004)	(0.005)	(0.005)
Number of older brothers*Female		0.001		
		(0.003)		
Number of older siblings	-0.003	-0.000	-0.001	-0.002
	(0.014)	(0.014)	(0.014)	(0.014)
Number of older siblings*Female		-0.002		
		(0.002)		
Female	-0.028***	-0.022***	-0.025***	-0.021***
	(0.003)	(0.005)	(0.004)	(0.005)
Next brother			-0.006	-0.001
			(0.003)	(0.004)
Next brother*Female				-0.010*
				(0.006)
Controls	√	√	√	√
Family Fixed Effects	√	√	√	√
Observations	26,723	26,723	26,723	26,723

Note. The dependent variable is child migration. Controls include year of birth indicators, age, age squared and birth order. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.

Table 6: Sibship size effects: household level estimates

	(1)	(2)	(3)
	OLS	IV	
Number of siblings	0.018***	-0.003	-0.031
	(0.002)	(0.015)	(0.033)
First Stage			
Infertility		-0.768***	
		(0.060)	
Miscarriage			-0.517***
			(0.065)
F-test		162.5	62.28
Controls	√	√	√
Municipality Fixed Effects	√	√	√
Observations	18,215	18,215	18,215

Note. The dependent variable is the number of migrants in the household. Controls include mother's year of birth indicators, mother's age and age squared, mother's age at first pregnancy, mother's years of schooling, indicator for single mother, father's decade of birth indicators, father's age and age squared and father's years of schooling, Standard errors clustered at the municipality level are in parentheses. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.