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Essays on the economics of migration: an empirical perspective

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Ai miei nonni,
che furono emigranti,
per il piú bel dono:
le mie radici,
da custodire e preservare con cura

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Contents

Introduction

This thesis comprises three essays on determinants and consequences of international labor migration at individual, family and country level.

The first chapter explores the effect of parents' absence due to migration on the educational outcomes of children remaining in the home country. The reduced form model estimated in this paper relies on an educational production function where the total effect of migration is the combination of a positive effect, coming from remittances, and a negative effect, coming from a parent's absence. The results suggest that parental absence offsets the positive impact of remittances on children's schooling achievements and that a mother's absence is more detrimental than a father's absence, especially for girls experiencing a mother's migration during adolescence.

The second chapter examines the effects of family size and demographic structure on offspring's international migration. Exploiting a data set containing detailed information on fertility histories, the potential endogeneity of parental fertility choices is addressed by using infertility shocks and miscarriage before first birth as exogenous variation in family size. There is 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, a deeper analysis on the birth order and the gender composition of siblings reveals that the chances to migrate are not equally distributed across children within the same family. Older siblings, especially firstborn males, are more likely to migrate, while having more sisters than brothers may increase the chances of migration, particularly among girls.

The third chapter studies the effect of granting the right to vote in local elections to non citizens of non Eu origin[1]. The introduction of a new law is exploited as a natural experiment and two sources of variation are used in order to identify the effect of the new law: the first is time variation coming from the introduction of the reform; the second

[1] This is a work in progress and results discussed below must be intended as preliminary.

source of variation is cross-sectional and arises from the differences in municipalities' share of new voters. In the spirit of a differences-in-differences strategy, election results of municipalities with more immigrants entitled to vote are compared with municipalities with a smaller share of immigrant voters (intensity of the treatment), before and after the reform. Preliminary results indicate that, depending on the economic and demographic context, and on the immigration policies, the effect of enfranchising non citizens from a different ethnic group can vary substantially: the pro immigrants parties can either gain on lose votes, with relevant consequences for public good expenditures.

Chapter 1

All you need is love... The effect of a mother's or father's migration on the education of children left behind

1.1 Introduction

During the last thirty years, an increasing share of women has started to migrate alone in order to pursue better economic conditions rather than to join other family members, thus leaving part of the family, including children, back at home (Oishi, 2002; Pedraza, 1991). This paper aims at providing causal estimates of the effects of parental migration on the educational outcomes of children left behind and addresses the heterogeneity of the effects driven by the gender of the migrant spouse.

The economic literature emphasizes that a mother's and father's inputs affect differently the production function of children's outcomes. On the one hand, it is well established that income and assets managed by women have a higher impact on children's health with respect to income managed by men (Duflo, 2003; Thomas, 1990, 1994). The evaluation of aid programs targeting women as income recipients shows significant improvements in children's health and education (e.g. Gertler et al., 2001; Schultz, 2000, on the evaluation of Progresa in Mexico). On the other hand, studies on the determinants of children's education stress that a mother's human capital is more closely related to children's attainment than the father's, and that maternal child care time significantly increases a child's completed years of education, especially if mothers are highly educated (Datcher-Loury, 1988).

Only recently, the economic literature has started to include gender in migration research (Cortes, 2015; Docquier et al., 2012; Morrison et al., 2007, among others). However, the so called 'feminization of migration' is not a new phenomenon (Jolly et al., 2005), and according to UN figures[1] women represented half of the migrant population already back in the 90s. Studying female international migration and assessing the differences to the male migration is essential to deeply understand the consequences for economic development of the sending countries.

This paper relies on Mexican households data, taken from the 1992 and 1997 cross sections of the *Encuesta Nacional de la Dinámica Demográfica* (ENADID). The ratio of women moving from Mexico to the United States increases over time, reaching its maximum in the middle of 90s, when women represented the 57% of Mexican legal migrants Cerrutti and Massey (2001b). Even if women are less likely than men to migrate to work in the United States, Cerrutti and Massey (2001b) find that the determinants of female migration change through the generations, suggesting common labor force motivations for the international migration of both younger men and women.

Following Antman (2012b)'s identification strategy, the problem of households selection into migration is overcome by exploiting a family fixed effect strategy and comparing children suffering mother's or father's absence at different stages of their educational career. Such a strategy addresses the 'double' selection which arises in the framework of this paper: on the one hand, migrant families differ fundamentally from non-migrant families; on the other, households in which the mother migrates can also systematically differ from those in which the father migrates.

The main results show that the gender of the migrant spouse and the timing of migration, relative to the child's educational progress, are both relevant. The absence of a parent due to migration has a negative effect on the children's years of completed schooling; moreover, a mother's absence is more detrimental than is a father's absence, especially for girls. These results help to shed light on the sign of the causal effect of parental migration on education, given that the total effect is theoretically ambiguous and that also the available empirical evidence is mixed. There are, in fact, at least two main channels, working in opposite directions, that determine the total effect of migration on education: remittances from abroad, which should positively affect children's schooling; and parental absence from home, the 'love' component, which is instead expected to have a negative impact on children's attainment.

The remainder of the paper is structured as follows: Section 2 reviews the relevant

[1] *Trends in International Migrant Stock: the 2008 Revision*, United Nation - Population Division.

literature; Section 3 describes the details of the empirical strategy; Section 4 presents data and descriptive statistics; Section 5 illustrated the results, Section 6 provides the robustness checks and Section 7 concludes.

1.2 Literature Review

This paper relates to two different strands of the literature: the first one studies the effect of a household member's migration on the outcomes of children left behind; while the second focuses on the determinants of education, in terms of parental investment in time and income.

Migration is an investment often made to improve families' welfare (Kennan and Walker, 2011a; Chen et al., 2003a). The individual who migrates is expected to earn more in the destination country than at home, and the remittances from abroad should improve children's left behind outcomes by relaxing the household budget constraint. However, the total effect on children is controversial, given that migration may impose a psychological cost to children, due to the absence of the migrant parent. A parent's migration can also have a negative impact on children outcomes, balancing the positive effect of remittances.

The total effect of parental migration on children education is then ambiguous and it remains an empirical question. The existing evidence is mixed and provides statistically significant estimates of both a positive and negative sign.

Hanson and Woodruff (2003) find a positive effect of living in a household with external migrants by focusing on Mexico. The effect is significant only for girls whose parents have a low level of education. However, in the same context, McKenzie and Rapoport (2011) find a negative effect of living in a migrant household on schooling attendance and attainments for boys aged 12-18 years and girls aged 16-18 years, as a consequence of increased housework for girls and of increased migration to the United States for boys.

By specifically focusing on paternal migration, Antman (2011a) finds that a father's migration from Mexico to the United States has a negative impact on educational outcomes of children staying at home. More specifically, the results on children's time use data suggest a negative short-term effect on hours devoted to study in the period following paternal migration and an increase in work hours of boys aged 12-15 years. A different results is reached by Antman (2012b) who estimates a positive effect of a father's migration to the USA on daughter's completed schooling in Mexico.

There are also few works studying the effect coming from remittances only, without accounting for the effect driven by the absence of the migrant household member. Bansak and Chezum (2009) and Edwards and Ureta (2003) find, respectively, that remittances from abroad have a positive effect on school attendance for school-age children in Nepal, and a significant impact on school retention in El Salvador. Given the lack of data, there is not much evidence on remitting behavior by gender. In general, since women are more likely to earn less in the host country, one may expect they remit less in absolute value. With regard to Mexico, Massey and Basem (1992) find evidence that men remit relatively more than women, that is a larger portion of their wage. The same results hold also in a larger area, i.e. Latin America and the Carribbean (Orozco, 2006).

Concerning the literature on the determinants of education, there is evidence that gender matters for explaining children attainments as a result of parental investment in time and income (Haveman and Wolfe, 1995, is a good reference for a complete review of the methods and findings on children's attainment). This literature shows that children growing up in single-parent households (not only as a result of migration) suffer from disadvantages, including poor school performance. Lyle (2006) analyzes the impact of parental absence and household relocation on children's academic achievements, by exploiting military deployment in the United States. The main results about parental absence show that the absence of the mother has a more adverse effect then the absence of the father. Chen et al. (2009) study the impact of parental death on children's education by using data on Taiwan. They find that losing a father has a very small and insignificant effect on children's college attainment, irrespective of the cause of death. Losing a mother causes instead a significant decrease in college enrollment, whatever the reason of the unexpected death. The authors interpret these findings as evidence of the fact that maternal roles, in terms of non financial support, are no less important than the provision of financial support in shaping children's cognitive ability.

In addition to the direct impact on children's outcomes, the absence of a spouse can also determine the change of spousal control over the intra-household allocation of resources. By exploiting changes in the household environment as a consequence of migration, Antman (2011b) finds evidence of increased women's decision making power and of increased gender discrimination among children while the husband is away. The fraction of clothing expenditure spent on boys decreases while the Mexican father is in the United States and increases in families in which the father is already back.

The results of Antman (2011b) suggest that the gender of the migrant and the resulting change in the decision making power over household allocations play crucial roles in determining the investment in the children remaining at home with the non-migrant spouse. However, the heterogeneity driven by the gender of the migrant is an issue not addressed by most of the previous works. One exception is the recent paper by

Cortes (2015), who explores the effect of a mother's migration on children's wellbeing in the Philippines. She compares children of migrating mothers with children of migrating fathers, by using an instrumental variable approach. Specifically, she uses economic shocks and changes in immigration laws in the destination countries as a random source of variation of migration rates. The probability of lagging behind (taking a value of one when the child has dropped out of school or is enrolled in a lower level than expected at her age) is used as the dependent variable. The results show that a mother's migration increases by 35% the probability that a child is lagging behind in school. The author interprets the results as evidence that the effect is mainly driven by lower parental inputs, rather than by fewer remittances (given that the same results persist when controlling for remittances).

There are two main aspects that differentiate this paper with respect to Cortes (2015). The first difference is in the outcome variable: she looks at the probability of lagging behind, which is a contemporaneous outcome with respect to a parent's migration; the focus of this paper is instead on completed education, as measured even many year after the experience of a parent's migration, i.e. a long run measure of children's achievements. The second different concerns with the identification strategy: while she relies on an IV strategy, in this paper a family fixed effects is adopted. More specifically, the time at which children experience a parent's absence is exploited in order to assess the impact of migration on their educational outcomes. This study tries to assess how decisions on schooling investments, at any given level, respond to changes in income, due to remittances coming from abroad, and to changes in time parents spend with their children.

1.3 Empirical Strategy

The empirical specification used throughout this paper exploits the different times of the educational career during which each sibling experiences a parental absence because of migration to the United States. Three crucial moments in the children's educational career are defined, each corresponding to the entrance to a new level of schooling. The empirical model captures these crucial moments through the creation of a series of dummies taking a value of one if the child experiences parental absence when she is 6-7 years old (corresponding to the moment in which decisions on entering primary education must be taken), 12-13 years old (the decision on entering secondary education) or 15-16 years old (the decision on entering upper secondary education).

The empirical strategy is inspired by Antman (2012b), however the main hypothesis on which the two papers rely is different. In Antman (2012b), the basic assumption is that schooling of children suffering migration after turning 20 years old is not affected by a father's being in the United States. The results she finds can be affected by a birth order effect, given that, by construction, only later born siblings are affected by a father's migration. In contrast, the identification strategy used in this paper does not generate any systematic relations between a parent's migration and the birth order of the children, because it exploits migration episodes at different moments of each child's educational career.

The underlying assumption that household decisions on investment in education, made at the beginning of each level, will affect the attainment of the entire level is confirmed by the statistics reported in Table 1.1 about the schooling attainment in Mexico for the schooling years 2000-2001 and 2001-2002, computed at a national level.

The takeover index is defined as the percentage of pupils of a cohort enrolls in a specific level who also enrolls in the next schooling level. More in details, this indicator shows that there is high continuity between primary and secondary school and that more than 90% of pupils completing primary school enroll also in secondary school.

The drop-out index indicates the percentage of pupils who abandon school before completing the specific educational level: only 2% of pupils abandon primary school before completion, while about 7% and 16% of pupils abandon secondary and upper secondary school before completion. The second index, *eficiencia terminal*, is defined as the percentage of pupils who manage to complete each educational level during the n years devoted to that specify level, that is, with no delay or repetition. This value is high especially for primary and secondary schools (almost 90 and 80% respectively). The two indicators show that the abandon rate is low (especially for primary and secondary schools) and a high percentage of pupils manage to complete each school level on time.

Given the above evidence, it is reasonable to assume that the schooling attainment, as measured by the years of completed education, of children who suffered parental absence outside of the *critical* ages defined above, is not affected by a parental migration experience. However, this assumption does not imply that parental absence does not affect other educational outcomes, for example, grades.

The following equation should be estimated in order to assess the impact of parental migration on children's education:

$$\begin{aligned} Years\_of\_schooling_i = \alpha + & \beta_1 primary_i + \beta_2 secondary_i + \\ & + \beta_3 upper\_secondary_i + \beta_4 X_i + v_i \end{aligned} \tag{1.1}$$

The dependent variable is the number of years of schooling that child i has completed by the time of the survey. The coefficients of interest are the ones for a mother's or father's migration at any stage (primary, secondary and upper secondary education), namely β_1, β_2 and β_3. The list of controls includes age, age squared and gender of each child, birth order dummies and a dummy for school attendance at the time of the survey. In 1992, the Secretariat of Public Education officially increased compulsory education from completion of primary school (grade six) to completion of lower secondary school (grade nine). A dummy for the reform is included in the controls. v_i is the idiosyncratic component.

One issue with the specification in equation (1.1) concerns the selection into migration, which can bias OLS estimates of the parameters β_1, β_2 and β_3, given the heterogeneity between migrant and non-migrant households. In this paper, the use of a family fixed effect strategy overcomes the problem of selection into migration, by comparing children's educational outcomes within the same family.

A family fixed effect is added to equation (1) and the following equation is estimated:

$$\begin{aligned} Years\_of\_schooling_i = \alpha + \beta_1 primary_i + \beta_2 secondary_i + \\ \beta_3 upper\_secondary_i + \beta_4 X_i + u_f + v_i \end{aligned} \tag{1.2}$$

In equation (1.2) the coefficients β_1, β_2 and β_3 refer to the effects of a mother *or* father's absence. In order to shed light on the heterogeneity of the effects driven by the gender of the migrant spouse, the regressors are split according to the gender of the migrating parent and the following equation is estimated:

$$\begin{aligned} Years\_of\_schooling_{if} = \alpha + \sum_p \beta_{1p} primary_{ifp} + \sum_p \beta_{2p} secondary_{ifp} + \\ + \sum_p \beta_{3p} upper\_secondary_{ifp} + \beta_4 X_{if} + u_f + v_{if}, \qquad p = father, mother \end{aligned} \tag{1.3}$$

Previous literature mainly relies on the IV approach to overcome the problem of the endogeneity of migration. Most studies use as an instrument either the historical migration rate (McKenzie and Rapoport, 2011) or the labor market conditions in the country of destination (Antman, 2011a; Cortes, 2015). However, neither set of instruments is free from criticism concerning the validity of the exclusion restrictions.

Employment conditions in the country of destination can have a direct impact on decisions about children's schooling in the country of origin, especially if the economic indicators of both countries are highly correlated. Moreover, there can be channelsother than parental migration, through which shocks in the labor market conditions of the destination country can affect schooling at the origin (such as the employment and wage conditions of other family members who have already migrated). When historical state-level migration rates are used as the instrument, high migration rates can be a proxy for higher development in the region and it can have a direct impact on current schooling attainment. Moreover, such an instrument can address the selectivity problem at a household level, but is not helpful to address the issue of who migrates and for how long.

Instead, the family fixed effect strategy controls for maternal and paternal selection into migration, given that the comparison is not between children of migrant fathers and children of migrant mothers, but *between* siblings *within* the same family. Also the observed and unobserved heterogeneity at a family level is controlled for, to the extent that the heterogeneity is assumed to be constant over time.

An underlying assumption of this identification strategy is that there are no other shocks at the family level, except for migration, that potentially affect children's education. The identifying assumption would be violated only if the unobserved shocks at a family level affect children depending on their ages, exactly in the same way as migration does (which means that only children at the beginning of a new level of schooling should suffer or enjoy the shock).

Another critical issue relates to the measure of schooling and to the inclusion of censored observations: the observed completed years of schooling for children who are still in school does not represent the final attainment and not taking it into account can result in biased estimates.

Therefore, the problem of censored observations of the dependent variable is addressed in two ways. First, the analysis focuses on children older than 16 years, because at this age they are not expected to be enrolled in school; indeed, this is a sample for which censoring is not a major problem (a robustness check is performed by using only children older than 19 years). Second, an age and gender-specific school index is build. The index is defined as the difference between the schooling of the individual i of gender g of cohort c, S_{igc}, and the mean schooling of her (or his) age and gender cohort, $\bar{S}_{gc}$. By using this approach, individual schooling attainments are rescaled relative to the performance of students of the same gender and cohort. This has the advantage of using all the observations. But, this method does not distinguish between individuals attending school and individuals outside of the educational process, since the information on current attendance is ignored.

1.4 Sample and Descriptive Statistics

The data used in this paper come from the 1992 and 1997 cross sections of the Encuesta Nacional de la Dinámica Demográfica (ENADID, National Survey of Demographic Dynamics) conducted by Mexico's National Statistical Agency (INEGI). The ENADID is a large survey, representative of the Mexican population, that collects personal information on household characteristics (such as the occupants' age, level of education, religion, job and income), details on pregnancy of women aged 15 to 54 years as well as information on migration. The ENADID also contains detailed information on the migration of both thefather and mother, on the country of destination and on the time spent away.

In order to identify parental migration episodes three data sources are used: the questionnaire section on international migration, in which migration episodes up to five years prior to the survey are registered and two questions from the section of general information on individuals living in the household. The first question collects data on an individual's prior residence, including details on time and place, and the second one asks about labor migration to the United States, again gathering details allowing reconstruction of the timing of such episodes.

In Table 1.2, summary statistics on the original and the final sample are reported. More than fifty thousand households were surveyed each year, and information on more than one hundred and fifty thousands children[2] within the whole country were collected. More than half of fathers in the original sample experienced migration to the United States, while only between four and six percent of mothers moved to the United Sates.

The final sample includes only children of migrants, in families where there are at least two siblings, and excludes children ever migrated to the United States. The sample of estimation is then made of more than 20,000 individuals, with 99% of them aged from six to thirty-five years. More than 90% of children experienced paternal migration at some point of their lives, while only about 6% suffered maternal migration.

In Panel A of Table 1.3 households with migrant mothers are compared with households with migrant fathers. As expected, the two types of families differ significantly in many dimensions (with the exception of the fraction of female children). In families in which the father leaves to the United States, it is more likely that the woman declares herself to be the head of the household; both mothers and fathers in such families are less educated and have more children. These differences do not invalidate the results:

[2]All daughters or sons of the head of household and her/his spouse are considered as children here.

the strategy used here exploits variation within the family, such that all observed and unobserved heterogeneity at the family level is cleaned away.

Table 1.3 (Panel B and C) compares migrant families with at least two children and migrant families with an only child. Only child families are very few, but also very different with respect to larger families. Families with only one child are, on average, more educated and experience longer migration. One major consequence of these tests is that the results of this study cannot be generalized to one child families, but they are valid only for families with two or more children.

The estimates presented below use two outcome variables: one is the years of schooling in formal education and the other is a measure of the years of schooling also including the years attended in training or in vocational courses (training courses in what follows). Figure 1.1 plots the two distributions separately by gender. It clearly emerges that females are more likely than males of the same age to attend training courses. Vocational courses are very important for girls: controlling for girls' age, attendance at such courses results in a higher probability of working out of the household and in a smaller probability of doing household works.

1.5 Results

The results unambiguously suggest a negative effect of parental absence on children's schooling due to migration. The magnitude of the effect depends on the time and the gender of the migrant, with mothers having a larger effect (in absolute value)[3] than do fathers. Moreover, there is evidence of heterogeneity driven by the gender of the child: daughters are more affected than sons, both in terms of the magnitude of the effect and because they are affected throughout their entire educational career. The effect is not monotonic with respect to the age of children at the time of parental migration; a mother's migration has a U-shaped effect, reaching its highest effect during secondary school, for both daughters and sons; a father's migration instead has a decreasing pattern for boys, reaching its highest effect when sons suffer his absence at the beginning of upper secondary school and a less clear pattern for daughters.

In the tables below, Panel A reports results when only years of schooling in formal education are used as the dependent variable, while in Panel B information on the years of schooling acquired by attending training courses are also used. Column 1 shows the

[3] All coefficients of interest have a negative sign.

results for the whole sample of children, then only girls (column 2) and only boys (column 3) are considered. Looking at the results by gender of the child is more informative about the parental decision process underlying the results, given that parents invest differently in the schooling of boys and girls.

Table 1.4 shows the results for the sample of all children aged 16 years and older, with dummies for the absence of at least one parent, without distinguishing by gender. The results in Panel A show that one parent's absence while starting secondary school makes girls attending half of a year of schooling less with respect to their sisters enjoying parents at home at the same age. The effect is lower and less significant when one parent is away at the beginning of upper secondary school. In Panel B, previous results are confirmed, with slightly higher coefficients.

In Table 1.5 the gender of the migrant spouse is accounted for. Estimation in Panel A column (2) suggests that both maternal and paternal migration has a detrimental impact on years of completed schooling. Suffering maternal absence when secondary school should start, namely at 12-13 years old, significantly reduces the schooling for girls by more than one year with respect to their sisters enjoying the presence of the mother at the same age. The magnitude of the effect on girls is even larger when one also considers the attendance of training courses. In order to enroll in most of the training courses, a student must complete secondary school. Thus, losing the opportunity to attend (and eventually complete) secondary school (as shown in Panel A) also prevents girls from attending a training course later on. No significant results are found for boys. At the bottom of Panel A and B of Table 1.5, p-values of one-sided tests for a mother's coefficients being larger than a father's coefficients at any stage are reported. For most specifications, a mother's absence has a significantly more detrimental effect on children's education with respect to a father's absence. This holds for both daughters and sons and especially at the beginning of secondary education.

The results of the family fixed effects strategy applied to the index of schooling (Table 1.6) support previous evidence that daughters are more affected than are sons, and suggest that girls suffer more from maternal migration and sons from paternal migration. The effect of parental migration on girls is never significant in Panel A (column 2), when only years of formal education are included. A one-sided test on the coefficient of maternal absence at 12-13 years old, however, results in the non-rejection of the null hypothesis that such a coefficient is negative. As for previous results, a maternal absence at 12-13 years old has a highly significant and detrimental effect on girls' years of schooling, when training courses are also considered. The p-values at the bottom of Panel B, relative to the test on a mother's coefficients being larger than a father's coefficients, suggest again that, at least for girls, a mother's absence at 12-13 years old is significantly more detrimental than a father's absence at the same age.

The presence of parents at home, when the decision on entering a new level of schooling must be taken, can be seen as the *extensive* margin effect. By exploiting information on the duration of parental absence, it is possible to shed light on the *intensive* margin effect. Duration is measured here as the years of a parent's absence *before* the child turns, respectively, 6, 12 and 15 years old. Table 1.7 and 1.8 show the results of the duration of parental absence on the years of schooling for the sample of children older than 16 years and on the index of schooling, respectively. The results for the sample of girls suggest that the length of maternal absence significantly decreases education during the child's adolescence (in line with results on the *extensive* margin). While the absence of the mother *per se* does not significantly affect boys' education, the length of the absence matters also for boys' schooling attainment. Surprisingly, the duration of maternal absence has a very negative effect at the beginning of boys' education (one more year of absence results in almost one year of schooling less for boys), while a longer absence has a positive but smaller effect when the child has to decide on the attendance of upper secondary school.

1.5.1 Discussion of the Results

There are two main channels through which migration can affect household investment on children's education: remittances and time, each affecting education in an opposite direction. Suppose that the education production function depends on two inputs: economic resources, represented in this framework by remittances, R; and parental time, T:

$$E(R_p, T_p), \qquad p = parental \tag{1.4}$$

The effect of parental migration, M_k, on E is then given by the following:

$$\frac{\partial E}{\partial M_k} = \frac{\partial E}{\partial R_p}\frac{\partial R_p}{\partial M_k} + \frac{\partial E}{\partial T_p}\frac{\partial T_p}{\partial M_k}, \quad k = father, mother \tag{1.5}$$

The sign of the derivative in (1.2) is ambiguous and depends on which effect dominates; the positive one coming from remittances or the negative one coming from absence.

Under the assumption that a dollar coming from maternal or paternal resources has the same effect on children's educational achievements ($\frac{\partial E}{\partial R_f} = \frac{\partial E}{\partial R_m}$) and one hour invested in children's education by the mother or the father has the same effect ($\frac{\partial E}{\partial T_f} = \frac{\partial E}{\partial T_m}$), it seems reasonable to expect that children experiencing a mother's migration will suffer more with respect to children experiencing migration of the father. This last expectation is justified by the evidence from previous studies on the effect of remittances and time by gender of the migrant parent, which suggests that the positive effect from remittances should be higher when a father migrates, while the negative effect deriving from a parent's absence is bigger when a mother migrates[4]. In short:

$$\frac{\partial E}{\partial R_p}\frac{\partial R_p}{\partial M_m} < \frac{\partial E}{\partial R_p}\frac{\partial R_p}{\partial M_f} \quad \text{and} \quad \frac{\partial E}{\partial T_p}\frac{\partial T_p}{\partial M_m} > \frac{\partial E}{\partial T_p}\frac{\partial T_p}{\partial M_f}, \quad m = mother, f = father.$$

The empirical exercise of this paper does not disentangle the two effects of remittance and time. However, looking at the sign of the total effect, $\frac{\partial E}{\partial M_k}$, allows us to infer that the negative effect coming from the lack of time inputs dominates the positive effect from remittances. Moreover, the results suggest that the negative effect is stronger when mother migrates. Descriptive evidence shows that migrant mothers are, on average, more educated than migrant fathers. One can then imagines that the detected effect operates through the reduction of time spent with the most skilled parent. This conjecture is in line with the results on the duration of a parent's absence: the longer is a mother's migration, the less she can exert influence on children's behavior and, consequently, the children will attend less schooling.

If the simple framework above helps to explain the negative sign of the coefficients and why, in the data, mothers' migration is more detrimental, it provides no intuitions on why mother's negative coefficients are mainly significant for girls. Migration of a parent implies that the spouse remaining at home has more decision making power on children's investments: if a father has higher preferences for sons, while the mother is away, he decides to devote more resources to them. The opposite is true for the mother, who has higher preferences for girls. Such an interpretation is consistent with other studies showing that an increase of the woman's bargaining power generates an increase of girls' outcomes but not of boys' (Duflo, 2003; Thomas, 1994).

Consistently across most of the specifications, no significant results are found for children experiencing parental migration early in their educational career (namely when they are 6-7 years old). Reasonably the opportunity cost of sending children to school is lower at this stage, while as children grow up it is more valuable to make them work instead of

[4]Refer to Section 1.2.

study, even to sustain the cost of a parent's migration.

The literature identifies an additional channel through which parental migration can influence the education of children remaining in the home country: the prospect of future migration. According to this explanation, parents decide not to educate children since they expect children to migrate later in life and return to education is higher in Mexico than for Mexican individuals moving to the United States (Chiquiar and Hanson, 2005). This mechanism does not play a major role in explaining previous results since the main effect is found for girls and descriptive statistics show that migrant women positively select with respect to education. Moreover, McKenzie and Rapoport (2011) find that living in a migrant household increases the probability of migration for younger males, while 'female youth in migrant households are not that likely to migrate themselves'.

1.6 Robustness Checks

To test the robustness of the results, it is necessary to show that they are not affected by the sample composition driven by the cut at 16 years old and not driven by the difference in timing of the mother's or father's migration with respect to their children's births.

First, in Table 1.9, the results for an alternative sample including only children aged 19 years or older are shown. This sample shows a lower average attendance rate (relaxing the potential bias coming from comparing children still in school with children already out) at the cost of being smaller in size. The coefficients on daughters (column 2) are not significant and the change with respect to the coefficients for girls older than 16 is small. This indicates that between 17 and 19 years old, girls do not acquire additional schooling, independently of any parental migration episode they experience. A one-sided test on the coefficient of a mother's absence at 12-13 years old suggests we cannot reject the hypothesis that the effect of a mother's absence at 12-13 years old is negative. The results on informal education, in Panel B, are fully robust to the previous results: the absence of the mother when a daughter is aged 12-13 years results in a significant reduction in the schooling acquired of almost three years.

Looking at the sample of boys (column 3), almost all coefficients are higher in magnitude with respect to the coefficients in Table 1.5. The coefficient for a mother's absence at 12-13 years old is now also highly significant. It suggests that, differently from girls, boys are more likely to attend a higher level of schooling after they turned 16 years old, and that experiencing a mother's absence when decision on entering secondary school must be taken is crucial for enrollment in higher grades (sons experiencing a mother's absence at 12-13 years old have almost five years of schooling less than their brothers enjoying the presence of the mother at the same age).

Column 1 of Table 1.10 shows the results for a sample of children all aged 16 years or older whose siblings are all in the household[5] . Such a sample ensures that the comparison is between *all* siblings and none are missing, being already out of the household for any reason. In both Panel A and B, the result that a mother's absence at 12-13 years old is highly detrimental is confirmed, with coefficients slightly higher with respect to previous specifications.

A possible threat to the identification strategy comes from the eventual correlation between a mother's and father's migration and their decision on fertility: mothers, for instance, may decide to migrate only after the births of all children, while fathers may be more likely to migrate in between the births. This discrepancy might generate correlations between the timing of mothers' and fathers' migrations and their fertility decisions. If this was true for any household size, only first borns would have suffered a mother's migration later in their educational career. If a correlation between the timing and spacing of births and migration exists, a correlation between birth order and children's age at the time of parental migration must also exist. First, descriptive evidence is provided in order to show that it does not happen in this empirical study. After that, robustness checks are run on a selected sample for which such a problem does not exist.

Figure 1.3(a) plots the average number of children born before a mother's or father's migration for any given household size (on the x-axis). For families with four children or fewer (40% of the sample), a very small difference is observed in the patterns of a mother's and father's migrations. As the number of children increases, it appears more likely that mothers migrate before giving births to all children, while fathers migrate later. However, data clearly show that, for any household size, both parents are likely to migrate before the births of all kids, with mothers moving earlier, especially in a large family.

Figure 1.3(b) (1.3(c)) shows the probability of experiencing a mother's (father's) migration at each critical threshold, that is when child is aged 6-7, 12-13 or 15-16 years. Figure 1.3(b) suggests that, at almost any birth order (x-axis), a mother's migration is more likely to happen when the child is very young (6-7 years). A father's migration at any stage (Figure 1.3(c)) is distributed more uniformly with respect to birth order. A mother's and father's time of migration do not show a systematically different pattern with respect to the birth order of children. Then, even if there are small differences in the timing of migration by the gender of the migrant, they do not represent a threat to

[5]The sample is not split by the gender of the chid because of a lack of sufficient observations.

the validity of the proposed identification strategy.

The sample in column 2 of Table 1.10 uses all children aged 16 or older who have no more than three siblings; that is, a sample for which there are no concerns about the different timing of a mother's and father's migration. Again, in both Panel A and B, a mother's absence at 12-13 years old has the highest effect, while also a father's absence at 15-16 years old has a negative effect on schooling when the attendance of training courses is also accounted for.

1.7 Conclusions

This paper studies the impact of parental migration on the educational outcomes of children left behind in Mexico and controls for the heterogeneity of the effect driven by the gender of the migrant spouse.

The problem of selection into migration is overcome by exploiting a family fixed effect strategy, which ensures that the selection bias arising from the differences between female and male migrant households is properly taken into account.

The results unambiguously suggest a negative effect of parental absence due to migration on children schooling. The reduced form model estimated here relies on a simple educational production function where the effect of migration is the combination of a positive effect coming from remittances and a negative effect coming from a parent's absence. Then, the empirical evidence suggests that a parental absence offsets the positive impact of remittances on children's schooling achievements. The results on the intensive margin, i.e. the duration of a parent's migration, are also in line with the above framework, i.e. a longer absence of the parent who is mainly devoted to caring for their children and helping them with their homework and who is more educated, namely the mother, negatively affects children's educational achievement, for both girls and boys.

The effect of migration on the educational attainment of the children left behind is heterogeneous: on the one hand, maternal absence is more detrimental than paternal absence during child adolescence; on the other hand, the timing at which the child experiences parental absence plays a crucial role. Experiencing migration early in the child's educational career does not have a significant effect. A possible explanation for this relies on the opportunity cost of sending the child to school: it is lower when the child is very young and it increases as the child grows up, when she can contribute to a family's income by working outside the household or by participating at household occupations. In line with this, McKenzie and Rapoport (2011) find that girls in migrant households are significantly more likely to do housework than girls in non-migrant households; while Antman

(2011a) finds that boys aged 12-15 years increase working hours and work participation outside the home in response to a father's migration.

The increase of migrant women in response to global changes in the labor market raises relevant policy questions. Women are now more likely to migrate independently and to work abroad in order to fulfil the demand for cheap female labor supply, especially as caregivers or to perform other *women's work*. This path can promote the reproduction and exploitation of gender inequalities by reinforcing existing patterns that oppress women and which tend to perpetuate a women's role in the family and the society. Moreover, as this study points out, maternal absence during child adolescence can be very detrimental for children's educational attainments, especially for daughters. Women's migration can then have a negative impact on women themselves through the perpetuation of gender roles and also through the effect on the next generation of women.

Gender is a key element in migration and it can influence any aspect of migration: causes, patterns and consequences. This paper shows that the gender of the migrant spouse is an important dimension when assessing the impact of migration on the family left behind. Incorporating gender into migration research is essential to fully understand this phenomenon.

1.8 Figures and Tables

Figure 1.1: Distribution of years of completed schooling by age and gender

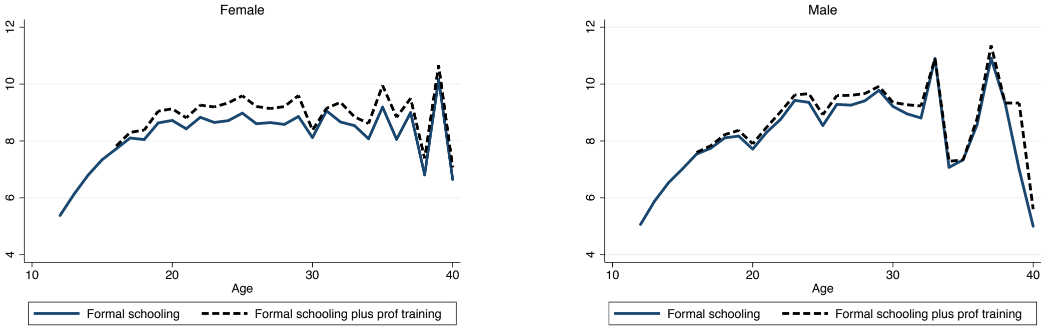
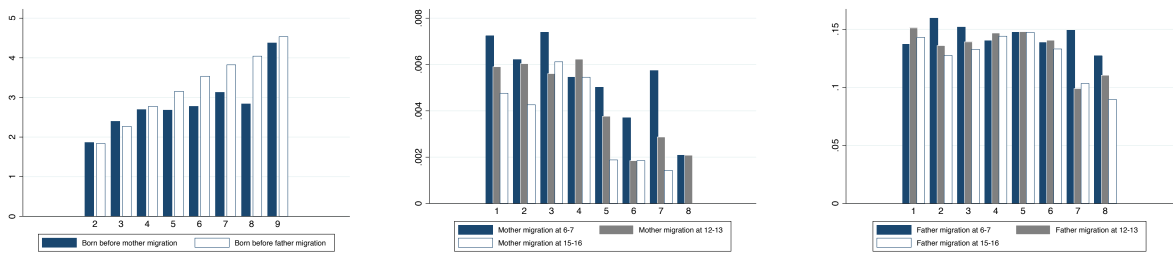


Figure 1.2: Fertility and migration



(a) For each family size, on the x-axis, the number of children born before a mother's and father's migration is plotted

(b) For each birth order, on the x-axis, the percentage of children experiencing a mother's migration at each *critical age* is plotted

(c) For each birth order, on the x-axis, the percentage of children experiencing a father's migration at each *critical age* is plotted

Table 1.1: Mexican Schooling Indicators

	Drop-out Rate		***Eficiencia Terminal***	
Schooling year	2000/2001	2001/2002	2000/2001	2001/2002
Primary School	2%	2%	86%	88%
Secondary School	8%	7%	75%	78%
Upper Secondary School	17%	16%	59%	59%

Note. Source: Instituto Nacional de Estadística y Geografía (INEGI).

Table 1.2: ENADID surveys

A.	**Original Sample**		
	1992	1997	Total
Observations	284,980	362,905	647,885
Number of households	53,874	72,724	126,598
Number of children	160,365	191,033	351,398
Migrant mothers	0.040	0.063	0.053
Migrant fathers	0.547	0.709	0.639

B.	**Final Sample**		
	1992	1997	Total
Number of households	2,652	4,152	6,804
Number of children	8,689	12,957	21,646
Migrant mothers	0.064	0.066	0.065
Migrant fathers	0.936	0.934	0.935

Note. *Panel A*: Original sample after merging survey sections; *Panel B*: Only households for which either the mother or the father migrated to the United States and with at least two children.

Table 1.3: Tests on household characteristics

A. Migrant Mothers vs Migrant Fathers

	Mean migrant mothers	Mean migrant fathers	P-values
Male household head	0.454	0.953	0.00
Mother's education	7.458	6.004	0.00
Father's education	9.124	5.933	0.00
Kid's education	8.375	7.146	0.00
Duration of migration	5.027	3.353	0.00
Number of kids	2.380	3.230	0.00
Fraction of female	0.486	0.500	0.16
Kid's age	13.968	12.378	0.00
Number of households	1,457	12,838	

B. Migrant Mother with 2 or more Kids vs Migrant Mother with Only Child

	Mean 2 or more kids	Mean only 1 kid	P-values
Male household head	0.559	0.493	0.06
Mother's education	7.287	8.237	0.00
Father's education	8.896	9.821	0.06
Kid's education	8.211	9.114	0.02
Duration of migration	3.796	5.213	0.00
Fraction of female	0.487	0.472	0.64
Kid's age	13.968	14.690	0.19
Number of households	706	276	

C. Migrant Father with 2 or more Kids vs Migrant Father with Only Child

	Mean 2 or more kids	Mean only 1 kid	P-values
Male household head	0.948	0.961	0.02
Mother's education	5.924	6.939	0.00
Father's education	5.873	6.490	0.00
Kid's education	7.100	7.664	0.00
Duration of migration	2.675	3.507	0.00
Fraction of female	0.500	0.494	0.57
Kid's age	12.323	13.348	0.00
Number of households	8,679	1,843	

Note. *Panel A*: The characteristics of the households in which the mother migrates are compared with the characteristics of the households in which the father migrates; the p-values in the last column derive from a t-test on the equality of the mean between the two groups; *Panel B (C)*: The characteristics of migrating mother (father) families with two or more children are compared with the characteristics of migrating mother (father) families with an only child. Statistics on children's education are computed by using only children with completed education. The duration of migration is expressed in years.

Table 1.4: The effect of parental migration on the years of schooling of children left behind

A. Dependent Variable:	**Years of Schooling in Formal Education Only**		
	All children (1)	Females only (2)	Males only (3)
One parent's absence when child is 6-7 years old	0.0417	0.0478	-0.0827
	(0.154)	(0.264)	(0.227)
One parent's absence when child is 12-13 years old	-0.1812	-0.5726***	-0.1930
	(0.146)	(0.217)	(0.250)
One parent's absence when child is 15-16 years old	-0.2540**	-0.3459*	-0.3349
	(0.119)	(0.196)	(0.204)
Female	0.0677		
	(0.076)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	7,494	2,196	2,069
Number of Families	3,683	938	889
Average Schooling	8.286	8.359	8.300
Average Attendance Rate	0.227	0.196	0.193
R-squared	0.0527	0.0566	0.0634
B. Dependent Variable:	**Years of Schooling including Training Courses**		
One parent's absence when child is 6-7 years old	0.0229	-0.1000	-0.0176
	(0.158)	(0.246)	(0.234)
One parent's absence when child is 12-13 years old	-0.2134	-0.8690***	-0.0963
	(0.155)	(0.230)	(0.253)
One parent's absence when child is 15-16 years old	-0.3531***	-0.5445**	-0.3403
	(0.125)	(0.216)	(0.209)
Female	0.2474***		
	(0.080)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	7,494	2,196	2,069
Number of Families	3,683	938	889
Average Schooling	8.554	8.767	8.473
Average Attendance Rate	0.227	0.196	0.193
R-squared	0.0552	0.0659	0.0649

Note. Only children older than 16 years old. Dependent variable: *Panel A*: Years of completed schooling in formal education only; *Panel B*: Years of completed schooling including the attendance of training courses. Controls include: age, age squared, birth order dummies, a dummy for attendance, a dummy for the 1992 schooling reform. Standard errors clustered at the family level are in parentheses. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.

Table 1.5: The effect of parental migration on the years of schooling of children left behind, by the gender of the migrant

A. Dependent Variable:	**Years of Schooling in Formal Education Only**		
	All children (1)	Females only (2)	Males only (3)
Paternal absence when child is 6-7 years old	0.0336	0.0823	-0.0701
	(0.159)	(0.270)	(0.231)
Maternal absence when child is 6-7 years old	0.2290	-0.6968	-1.4212
	(0.535)	(0.703)	(1.002)
Paternal absence when child is 12-13 years old	-0.1040	-0.5019**	-0.1044
	(0.148)	(0.222)	(0.246)
Maternal absence when child is 12-13 years old	-1.5192**	-1.4793**	-2.7097
	(0.686)	(0.675)	(1.876)
Paternal absence when child is 15-16 years old	-0.2458**	-0.2913	-0.3360
	(0.122)	(0.194)	(0.209)
Maternal absence when child is 15-16 years old	-0.2534	-0.9570	-0.3967
	(0.490)	(1.114)	(0.768)
Female	0.0684		
	(0.076)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	7,494	2,196	2,069
Number of Families	3,683	938	889
Average Schooling	8.286	8.359	8.300
Average Attendance Rate	0.227	0.196	0.193
R-squared	0.0540	0.0582	0.0663
Mother > Father at 6-7 years old	0.637	0.149	0.0944
Mother > Father at 12-13 years old	0.0220	0.0842	0.0842
Mother > Father at 15-16 years old	0.494	0.277	0.470
B. Dependent Variable:	**Years of Schooling including Training Courses**		
Paternal absence when child is 6-7 years old	0.0358	-0.0524	0.0183
	(0.163)	(0.252)	(0.236)
Maternal absence when child is 6-7 years old	-0.3366	-1.1896	-2.5422
	(0.600)	(0.930)	(1.746)
Paternal absence when child is 12-13 years old	-0.1122	-0.7588***	-0.0107
	(0.157)	(0.233)	(0.251)
Maternal absence when child is 12-13 years old	-1.9724***	-2.5631***	-2.6750
	(0.659)	(0.713)	(1.761)
Paternal absence when child is 15-16 years old	-0.3351***	-0.5095**	-0.3410
	(0.128)	(0.215)	(0.215)
Maternal absence when child is 15-16 years old	-0.4789	-0.7007	-0.1368
	(0.532)	(1.087)	(0.814)
Female	0.2463***		
	(0.080)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	7,494	2,196	2,069
Number of Families	3,683	938	889
Average Schooling	8.554	8.767	8.473
Average Attendance Rate	0.227	0.196	0.193
R-squared	0.0572	0.0687	0.0683
Mother > Father at 6-7 years old	0.275	0.118	0.0733
Mother > Father at 12-13 years old	0.00301	0.00809	0.0671
Mother > Father at 15-16 years old	0.396	0.431	0.596

Note. Only children older than 16 years old. Dependent variable: *Panel A*: Years of completed schooling in formal education only; *Panel B*: Years of completed schooling including the attendance of training courses. Controls include: age, age squared, birth order dummies, a dummy for attendance, a dummy for the 1992 schooling reform. At the bottom of Panel A and B, p-values associated to one-sided tests on the mother 's coefficients being bigger than the father's coefficients at each stage are reported. Standard errors clustered at the family level are in parentheses. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.

Table 1.6: The effect of parental migration on the index of schooling of the children left behind

A. Dependent Variable: **Index of Schooling computed on Formal Education Only**			
	All children (1)	Females only (2)	Males only (3)
Paternal absence when child is 6-7 years old	0.0255	-0.0340	0.0285
	(0.036)	(0.069)	(0.060)
Maternal absence when child is 6-7 years old	0.2491	0.2413	-0.1564
	(0.212)	(0.272)	(0.536)
Paternal absence when child is 12-13 years old	-0.0238	-0.0022	-0.0255
	(0.048)	(0.077)	(0.080)
Maternal absence when child is 12-13 years old	-0.3884	-0.3549	-0.5237
	(0.241)	(0.354)	(0.539)
Paternal absence when child is 15-16 years old	-0.0623	0.0906	-0.2873**
	(0.070)	(0.099)	(0.120)
Maternal absence when child is 15-16 years old	-0.0058	-0.2098	0.4572
	(0.249)	(0.611)	(0.344)
Female	-0.0069		
	(0.026)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	21,638	8,263	8,412
Number of amilies	6,804	3,254	3,334
R-squared	0.0800	0.0877	0.0741
Mother > Father at 6-7 years old	0.851	0.837	0.366
Mother > Father at 12-13 years old	0.0690	0.165	0.180
Mother > Father at 15-16 years old	0.586	0.314	0.980
B. Dependent Variable: **Index of Schooling including Training Courses**			
Paternal absence when child is 6-7 years old	0.0225	-0.0587	0.0507
	(0.038)	(0.070)	(0.062)
Maternal absence when child is 6-7 years old	0.1655	0.0933	-0.2961
	(0.220)	(0.310)	(0.565)
Paternal absence when child is 12-13 years old	-0.0224	-0.0289	-0.0103
	(0.050)	(0.080)	(0.082)
Maternal absence when child is 12-13 years old	-0.5236**	-0.6863*	-0.4966
	(0.257)	(0.409)	(0.528)
Paternal absence when child is 15-16 years old	-0.0764	0.0643	-0.2834**
	(0.073)	(0.107)	(0.123)
Maternal absence when child is 15-16 years old	-0.0625	-0.0644	0.4884
	(0.274)	(0.615)	(0.340)
Female	-0.0117		
	(0.027)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	21,638	8,263	8,412
Number of Families	6,804	3,254	3,334
R-squared	0.0825	0.0961	0.0770
Mother > Father at 6-7 years old	0.739	0.684	0.271
Mother > Father at 12-13 years old	0.0277	0.0572	0.181
Mother > Father at 15-16 years old	0.519	0.418	0.984

Note. Only children aged 6 years and older. Dependent variable: *Panel A*: Index of schooling computed on formal education; *Panel B*: Index of schooling computed on formal education including the attendance of training courses. Controls include: age, age squared, birth order dummies, a dummy for attendance, a dummy for the 1992 schooling reform. At the bottom of Panel A and B, p-values associated to one-sided tests on the mother's coefficients being bigger than the father's coefficients at each stage are reported. Standard errors clustered at the family level are in parentheses. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.

Table 1.7: The effect of parental migration duration on the years of schooling of children left behind

A. Dependent Variable: **Years of Schooling in Formal Education Only**			
	All children (1)	Females only (2)	Males only (3)
Duration of paternal absence before turning 6 years old	-0.0198	0.0136	0.0269
	(0.052)	(0.101)	(0.070)
Duration of maternal absence before turning 6 years old	0.1348	-0.3700	-0.7955**
	(0.172)	(0.789)	(0.348)
Duration of paternal absence before turning 12 years old	0.0660	-0.0672	0.0936
	(0.055)	(0.076)	(0.097)
Duration of maternal absence before turning 12 years old	-0.4030	-0.3213*	0.2723
	(0.366)	(0.164)	(0.554)
Duration of paternal absence before turning 15	-0.1156	-0.0571	-0.3553**
	(0.090)	(0.136)	(0.152)
Duration of maternal absence before turning 15 years old	-0.0545	-0.5108	0.2750**
	(0.284)	(0.598)	(0.123)
Female	0.0648		
	(0.076)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	7,494	2,196	2,069
Number of family	3,683	938	889
R-squared	0.0527	0.0522	0.0669
Mother > Father at 6-7 years old	0.805	0.315	0.0104
Mother > Father at 12-13 years old	0.103	0.0809	0.624
Mother > Father at 15-16 years old	0.581	0.229	0.999
B. Dependent Variable: **Years of Schooling including Training Courses**			
Duration of paternal absence before turning 6	-0.0113	-0.0180	0.0739
	(0.056)	(0.105)	(0.070)
Duration of maternal absence before turning 6	0.0400	-1.0958	-1.2498*
	(0.239)	(1.171)	(0.645)
Duration of paternal absence before turning 12	0.0780	-0.1388	0.1366
	(0.061)	(0.095)	(0.094)
Duration of maternal absence before turning 12	-0.4895	-0.4894*	0.2661
	(0.370)	(0.278)	(0.546)
Duration of paternal absence before turning 15	-0.1696*	-0.1612	-0.3585**
	(0.096)	(0.158)	(0.154)
Duration of maternal absence before turning 15	-0.2333	-0.6273	0.2756**
	(0.347)	(0.774)	(0.122)
Female	0.2450***		
	(0.080)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	7,494	2,196	2,069
Number of Families	3,683	938	889
R-squared	0.0552	0.0577	0.0706
Mother > Father at 6-7 years old	0.583	0.180	0.0208
Mother > Father at 12-13 years old	0.0654	0.118	0.592
Mother > Father at 15-16 years old	0.430	0.277	0.999

Note. Only children aged 16 years and older. Dependent variable: *Panel A*: Years of completed schooling in formal education only; *Panel B*: Years of completed schooling including the attendance of training courses. Controls include: age, age squared, birth order dummies, a dummy for attendance, a dummy for the 1992 schooling reform. At the bottom of Panel A and B, p-values associated to one-sided tests on the mother's coefficients being bigger than the father's coefficients at each stage are reported. Standard errors clustered at the family level are in parentheses. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.

Table 1.8: The effect of parental migration duration on the index of schooling of children left behind

A. Dependent Variable: **Index of Schooling computed on Formal Education Only**			
	All children (1)	Females only (2)	Males only (3)
Duration of paternal absence before turning 6 years old	0.0071	0.0045	0.0144
	(0.018)	(0.033)	(0.027)
Duration of maternal absence before turning 6 years old	-0.0187	0.0404	-0.0826
	(0.065)	(0.152)	(0.175)
Duration of paternal absence before turning 12 years old	0.0259	0.0037	0.0424
	(0.018)	(0.031)	(0.029)
Duration of maternal absence before turning 12 years old	-0.0755	-0.1626*	0.1812
	(0.118)	(0.087)	(0.267)
Duration of paternal absence before turning 15 years old	-0.0449	0.0413	-0.1719***
	(0.038)	(0.052)	(0.065)
Duration of maternal absence before turning 15 years old	-0.0570	-0.0953	-0.0218
	(0.133)	(0.227)	(0.184)
female	-0.0081		
	(0.026)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	21,638	8,263	8,412
Number of Families	6,804	3,254	3,334
R-squared	0.0798	0.0878	0.0742
Mother > Father at 6-7 years old	0.350	0.592	0.292
Mother > Father at 12-13 years old	0.198	0.0361	0.697
Mother > Father at 15-16 years old	0.465	0.279	0.779
B. Dependent Variable: **Index of Schooling including Training Courses**			
Duration of paternal absence before turning 6 years old	0.0107	-0.0064	0.0315
	(0.019)	(0.035)	(0.029)
Duration of maternal absence before turning 6 years old	-0.0387	-0.0145	-0.1332
	(0.066)	(0.144)	(0.182)
Duration of paternal absence before turning 12 years old	0.0277	-0.0040	0.0452
	(0.019)	(0.034)	(0.030)
Duration of maternal absence before turning 12 years old	-0.1048	-0.1796**	0.1523
	(0.118)	(0.086)	(0.273)
Duration of paternal absence before turning 15 years old	-0.0482	0.0351	-0.1701**
	(0.040)	(0.058)	(0.066)
Duration of maternal absence before turning 15 years old	-0.0756	-0.0877	0.0040
	(0.142)	(0.272)	(0.188)
female	-0.0130		
	(0.027)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	21,638	8,263	8,412
Number of Families	6,804	3,254	3,334
R-squared	0.0823	0.0958	0.0773
Mother > Father at 6-7 years old	0.235	0.478	0.186
Mother > Father at 12-13 years old	0.134	0.0294	0.652
Mother > Father at 15-16 years old	0.426	0.329	0.810

Note. Only children aged 6 years and older. Dependent variable: *Panel A*: Index of schooling computed on formal education; *Panel B*: Index of schooling computed on formal education including the attendance of training courses. Controls include: age, age squared, birth order dummies, a dummy for attendance, a dummy for the 1992 schooling reform. At the bottom of Panel A and B, p-values associated to one-sided tests on the mother's coefficients being bigger than the father's coefficients at each stage are reported. Standard errors clustered at the family level are in parentheses. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.

Table 1.9: The effect of parental migration on the schooling of children left behind - Children aged 19 years or older

A. Dependent Variable: **Years of Schooling in Formal Education Only**			
	All children (1)	Females only (2)	Males only (3)
Paternal absence when child is 6-7 years old	0.0625	0.5308	0.0381
	(0.231)	(0.407)	(0.350)
Maternal absence when child is 6-7 years old	0.0940	-1.0239	-2.2806
	(0.738)	(0.890)	(1.800)
Paternal absence when child is 12-13 years old	0.0433	-0.4465	0.2594
	(0.254)	(0.354)	(0.407)
Maternal absence when child is 12-13 years old	-1.4120*	-1.3814	-5.6969***
	(0.832)	(0.967)	(1.795)
Paternal absence when child is 15-16 years old	-0.1306	0.1483	-0.5112
	(0.276)	(0.376)	(0.462)
Maternal absence when child is 15-16 years old	0.0488	-0.4104	-0.8297
	(0.913)	(1.653)	(1.767)
Female	-0.1246		
	(0.128)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	4,270	1,459	1,378
Number of Families	2,343	812	789
Average Schooling	8.607	8.624	8.567
Average Attendance Rate	0.137	0.129	0.134
R-squared	0.0601	0.0767	0.0803
Mother > Father at 6-7 years old	0.516	0.0535	0.104
Mother > Father at 12-13 years old	0.0472	0.180	0.000624
Mother > Father at 15-16 years old	0.575	0.371	0.431
B. Dependent Variable: **Years of Schooling including Training Courses**			
Paternal absence when child is 6-7 years old	0.0510	0.3282	0.0986
	(0.237)	(0.387)	(0.341)
Maternal absence when child is 6-7 years old	-0.2759	-0.9118	-2.0135
	(0.630)	(0.843)	(1.643)
Paternal absence when child is 12-13 years old	0.0061	-0.8158**	0.4234
	(0.274)	(0.391)	(0.401)
Maternal absence when child is 12-13 years old	-2.0295***	-2.8820***	-4.8145***
	(0.715)	(1.077)	(1.640)
Paternal absence when child is 15-16 years old	-0.3397	-0.1551	-0.4627
	(0.284)	(0.435)	(0.468)
Maternal absence when child is 15-16 years old	-0.1511	-0.1232	-0.5773
	(0.970)	(1.658)	(1.607)
Female	0.1315		
	(0.133)		
Family Fixed Effect	√	√	√
Birth Order Dummies	√	√	√
Observations	4,270	1,459	1,378
Number of Families	2,343	812	789
R-squared	0.0570	0.0846	0.0731
Average Schooling	8.970	9.132	8.790
Average Attendance Rate	0.137	0.129	0.134
Mother > Father at 6-7 years old	0.314	0.0873	0.106
Mother > Father at 12-13 years old	0.00381	0.0351	0.000973
Mother > Father at 15-16 years old	0.574	0.507	0.473

Note. Only children older than 19 years. Dependent variable: *Panel A*: Years of completed schooling in formal education only; *Panel B*: Years of completed schooling including the attendance of training courses. Controls include: age, age squared, birth order dummies, a dummy for attendance, a dummy for the 1992 schooling reform. At the bottom of Panel A and B, p-values associated to one-sided tests on the mother's coefficients being bigger than the father's coefficients at each stage are reported. Standard errors clustered at the family level are in parentheses. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.

Table 1.10: The effect of parental migration on the schooling of children left behind - Selected samples

A. Dependent Variable: **Years of Schooling in Formal Education Only**	All children HH (1)	Less than 4 kids HH (2)
Paternal absence when child is 6-7 years old	0.1904	0.1834
	(0.306)	(0.332)
Maternal absence when child is 6-7 years old	0.7405	0.6784
	(0.962)	(0.697)
Paternal absence when child is 12-13 years old	0.0544	0.1564
	(0.303)	(0.332)
Maternal absence when child is 12-13 years old	-1.6392**	-2.3363***
	(0.718)	(0.850)
Paternal absence when child is 15-16 years old	-0.7135***	-0.3243
	(0.247)	(0.258)
Maternal absence when child is 15-16 years old	-0.5765	-0.1019
	(0.761)	(0.687)
Female	-0.0169	0.0423
	(0.140)	(0.146)
Family Fixed Effect	√	√
Birth Order Dummies	√	√
Observations	2,653	2,993
Number of Families	1,210	1,636
R-squared	0.0756	0.0806
Average Schooling	9.506	9.474
Average Attendance Rate	0.277	0.296
Mother > Father at 6-7 years old	0.708	0.739
Mother > Father at 12-13 years old	0.0151	0.00319
Mother > Father at 15-16 years old	0.568	0.619
B. Dependent Variable: **Years of Schooling including Training Courses**		
Paternal absence when child is 6-7 years old	0.0302	0.0060
	(0.310)	(0.335)
Maternal absence when child is 6-7 years old	0.4566	0.4156
	(0.827)	(0.612)
Paternal absence when child is 12-13 years old	0.1463	0.1948
	(0.318)	(0.350)
Maternal absence when child is 12-13 years old	-3.4825***	-3.0357***
	(0.550)	(0.689)
Paternal absence when child is 15-16 years old	-0.8972***	-0.5456**
	(0.258)	(0.266)
Maternal absence when child is 15-16 years old	-0.7878	-0.2605
	(0.800)	(0.736)
Female	0.2407*	0.2685*
	(0.145)	(0.152)
Family Fixed Effect	√	√
Birth Order Dummies	√	√
Observations	2,653	2,993
Number of Families	1,210	1,636
R-squared	0.0819	0.0897
Average Schooling	9.506	9.474
Average Attendance Rate	0.277	0.296
Mother > Father at 6-7 years old	0.686	0.721
Mother > Father at 12-13 years old	1.20e-08	1.51e-05
Mother > Father at 15-16 years old	0.552	0.642

Note. Column (1): the sample includes children all aged 16 years or older and whose siblings are all in the household. Column (2): the sample includes all children aged 16 years or older whose family size if four or less. Dependent variable: *Panel A*: Years of completed schooling in formal education only; *Panel B*: Years of completed schooling including the attendance of training courses. Controls include: age, age squared, birth order dummies, a dummy for attendance, a dummy for the 1992 schooling reform. At the bottom of Panel A and B, p-values associated to one-sided tests on the mother's coefficients being bigger than the father's coefficients at each stage are reported. Standard errors clustered at the family level are in parentheses. Significance at the 10 % level is represented by *, at the 5% level by **, and at the 1% level by ***.

Chapter 2

Family Size, Sibling Rivalry and Migration: Evidence from Mexico[1]

2.1 Introduction

Migration from poor to rich countries is one of the most important ways through which workers can increase their income opportunities as well as their families' welfare back home (Chen et al., 2003b; Kennan and Walker, 2011b; Clemens, 2011). A key feature of migration is that it mainly involves young adults who are more likely to have a positive net expected return to 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).

Given the profitable nature of labor mobility, which involves both the (young) 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). Indeed, in many developing countries labor migration is a family strategy to diversify income sources, improve earning potentials and increase household security through remittances (e.g. Stark and Bloom, 1985; Yang, 2008; Antman, 2012a).

[1]This is a joint work with M. Bratti and M. Mendola

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, and 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 across their children, due to either limited household resources or (perceived) different returns to the migration investment (e.g., pro-male bias).[2] This may generate resource diluition effects in large families or competition (rivalry) among siblings from the same household (Garg and Morduch, 1998; Black et al., 2005). Although the determinants of migration have already been studied extensively, far less is known about the role of the size and the structure of the origin household – in particular the role of siblings – 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 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.[3] We address this question in the context of the the Mexico-U.S. mass migration in the 1990s. Mexico is one of the largest migrant-sending and remittance-recipient countries worldwide, with a migration wave that swelled in the 1970s and kept growing in the 1980s and 1990s, accounting from 5.2 percent of Mexico's national population in 1990 to a peak of 10.2 per cent in 2005 (Hanson and McIntosh, 2010). According to the Mexico's Population Census, during the 1990s alone, 9 percent of Mexicans 16 to 25 years old (based on age in 1990) migrated to the United States. A distinguishing feature of last century Mexico-U.S. migration is that most migrants have usually low levels of education and many of them have their first U.S. jobs in the seasonal agriculture (Martin, 1993).[4] According to U.S. Census data, in 1990 70.4 percent of Mexican immigrant men were high-school dropouts, against 12.9 percent of male native-born working population and 21 percent of non-Mexican immigrant working men (Borjas and Katz, 2005). Yet,

[2]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; Schultz, 1990; Thomas, 1990)

[3]Several studies document sibship size and birth order effects in outcomes as varied as schooling, height, IQ (see Black et al., 2005; Angrist et al., 2010; Pande, 2003; Jayachandran and Pande, 2015, among others).

[4]U.S. policy supported the recruitment of rural Mexicans under bilateral agreement between 1940s and 1960s (e.g., the Bracero Program) but most of the 20th century Mexican migrants arrived and were employed outside guestworker programs (Martin, 1993).

the American Dream creates opportunities for upward mobility such that Mexican immigrants enjoy income gains with respect to their counterparts living in Mexico, and family members at home share in these gains through remittances (Hanson, 2004; Ozden and Schiff, 2006; Rosenzweig, 2007; Clemens et al., 2010). Importantly, emigration rates differ by age and gender. As reported by Hanson and McIntosh (2010) by using Mexico's population censuses, a significant fraction of males migrates by age 16 with emigration increasing sharply until around age 30 and decreasing thereafter, presumably as a result of return migration. For Mexican women instead, there is less emigration by age 16, with subsequent rates being relatively stable over the course of their lives.[5]

Moreover, interestingly enough the 1990s wave of Mexican migration has been crossing over a demographic boom that petered out years later. Mexico's birth rate stood at about seven children per mother in 1970. The gradual spread of family planning practices contributed to impel the fertility transition in the country where, by 2005, the number of children per woman declined to slightly more than two (Cabrera, 1994).[6] Yet, despite the abundant evidence on the potentially significant implications of high fertility rates for child investments and economic outcomes, the existing literature provides scant rigourous analysis of the link between family size[7] and offspring's international migration.

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 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 may be the case as, in a context such as Mexico with weak 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).[8] Thus, the lure of international migration from Mexico to the U.S. may

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

[6]In 1974 a new population policy was designed in Mexico with the aim to reduce population growth and to promote development. The new institutional structure set up to ensure the policy implementation (the National Population Council- CONAPO) has expandend geographically and socially over time (Zuniga Herrera, 2008).

[7]We use 'family size' (i.e. the number of children) and 'sibship size' (i.e. the number of siblings) as synonymous: the first takes the point of view of parents, the second the one of children.

[8]We use data on young adults in Mexico in the mid 1990's whereby fertility decisions of their mothers were taken across the '70s and '80s. At that time the country was classified as a developing poor economy and the lack of markets or institutions were more likely to be mitigated by the family than it is the case

increase the likelihood of upstream transfers from children to parents, and hence raise the economic returns of high fertility for parents (Stark, 1981). We address this endogeneity issue by exploiting exogenous variation in family size induced by either infertility shocks or miscarriage at first pregnancy (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 between siblings variation only. This is important in order to shed light on the intra-household selection process into migration, which has important implications for child welfare, gender disparities and the ultimate impact on origin families (Chen, 2006; Mourard, 2015).

We find no evidence that high fertility drives migration choices at the household level. 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 firstborn males, 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.

These findings have relevant implications. First, our analysis can contribute to explaining the impact of 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 circle of high-fertility, poverty and economic stagnation (Miller and Babiarz, 2014; Schultz, 2008). Some of these programs have been implemented in high fertility societies with significant out-migration rates, such as Mexico, but little is known on the (intended and unintended) consequences of the former on the latter. By observing a positive association between fertility and economic migration, implications may be drawn that smaller families could lead to lower rates of mobility. 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 dynamic fertility decisions in contexts of poor resources and high emigration opportunities. The reason is that, in developing settings, 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 2.2 describes the link between household structure and migration as considered by the related literature on human capital investment.

in nowadays Mexico.

Section 2.3 presents the data and sample selection. The methodology and empirical strategy is described in Section 2.4. Section 2.5 presents our main results on birth order and sibship size estimates, plus robustness checks. Section 2.6 reports results on the siblings' compostion effect. Finally, Section 2.7 summarizes our main findings and concludes.

2.2 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, which are higher for young people, are conceived in terms of both migrants' earnings and remittances sent back home (Stark and Bloom, 1985; Yang, 2008; Amuedo-Dorantes and Pozo, 2011). Indee, people decide to migrate because they expect their own or their family's payoff to be higher in terms of a different and higher profile of earnings, quality of life, health or security or because migration mitigates risks and household portfolios at origin (Chen et al., 2003b; Kennan and Walker, 2011b; Clemens, 2011). Recent evidence shows that – after controlling for self-selection– 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).[9] Beyond income gains for migrants, cross-border migration typically brings additional liquidity to family members left behind through remittances, which significantly support consumption and investment decisions, as well as the management of risk and credit constraints in the household of origin.

Given the key economic role of migrants' remittances, especially in developing contexts, several contributions in the migration literature point at the household as the main unit for migration choices (Rosenzweig, 1988; Stark, 1991; Ghatak and Price, 1996). The core feature of this collective decision-making framework is that the family aims at maximizing household income and therefore can take the costly decision to dispatch one (or

[9]By combining household data in Mexico with U.S. and Mexico population censuses, and controlling for self-selection on observables and unobservables in migration, recent works estimate that yearly income gains to Mexico-U.S. migration are around 6,700 to 8,000 U.S. dollars (Hanson, 2004; Clemens et al., 2010). Moreover, according to the 2009 poll by the Pew Research Center, a third of all Mexicans would move to the U.S. if they could, and half of these potential Mexican migrants report to be prepared to move illegally to the U.S. According to more than 55 percent of those polled in Mexico, Mexicans who move to the U.S. have a better life despite well-known hardships, while less than 15 percent report life is worse in the U.S.

more) young member 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 investment strategy whereby one or more members are assigned to work in the local economy while others are sent abroad to act as a source of insurance or financial enhancement. 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 (2012a) shows that children migrated to the U.S. (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 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 (e.g. Nugent, 1985; Cigno, 1993; Ray, 1998).[10] 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. Children embody income-earning possibilities both for themselves and for their parents, and this may be the reason why in poor contexts (i.e. with weak formal markets and social safety-nets) people generally choose to invest in their future in the form of children (Duflo and Banerjee, 2011). The traditional system of family arrangement, though, may have important consequences on economic choices and

[10]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).

offspring's outcomes (Platteau, 1991).

While an extensive empirical literature provides evidence on the role of household size and composition on parental investments in other forms of children's human capital, such as education or health, (Garg and Morduch, 1998; Black et al., 2005), within-family considerations have been less analyzed in the context of migration decisions.[11] 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 (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, larger families may increase the pressure of the family hierarchy, higher dependency ratio and the amount of disposable resources, to support the family members. Hence, a reallocation of resources from children to parents may become necessary 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.

2.3 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

[11]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).

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 in 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.[12] 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 offspring's 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.[13] Figure 2.1 reports the incidence of non-tied migration by age and gender in Mexico and shows that, overall, migrants are massively 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 of the origin family.[14]

[12]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 under represent 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.

[13]Yet, we investigated the robustness of our findings to the inclusion of tied-migrants as well (9 percent of the sample), including parents' migration status among the controls (results available upon request). In their study of Mexican migration to the U.S., Cerrutti and Massey (2001a) report that nearly half of all male migrants leave to the U.S. before or without a wife or a parent.

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

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 Section 2.4.2 for more details). In line with the medical definition of infertility [15] and with the literature (e.g., 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 80 per cent of the original sample).

Our final estimation sample is made of 26,743 children in the age range 15-25, whose mother's fertility history is likely to be completed such that full siblings' information is observed.[16] In our sample of individuals, 5.2 percent are migrants with male and female migration rates equal to 7.07 and 2.92 percent, respectively. In Figure 2.2 we plot the average migration rate of boys and girls in our sample by sibship size. A positive association between sibship size and migration for sons clearly emerges. Individual sample characteristics are reported in Table 2.1 according to the 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.

In Figure 2.3 we plot the ratio of migrant children in the household by family size, in the sample of households with at least one migrant child (against the distribution of migrant households in the population. The plot shows a negative association between the child migrant ratio and sibship size, which means that all households, of any size, hardly have more than one young member migrated abroad (the average number of young migrant members per household is 1.14 in the sample of households with migrants). This is suggestive of an intra-household selection process into offspring's migration which we

[15] The medical literature defines infertility as the failure to conceive after a year of regular intercourse without contraception.

[16] Mothers of individuals in our estimation sample are on average 45 years old. The average birth spacing between the first and the last child in our estimation sample is 13 years, that is below the minimum age of individuals we consider (15). Our sample of children does not include those with mothers older than 54 years of age (9 percent of the total population aged 15-25) since fertility information was not collected from them.

explore further through inferential analysis in the following sections.

2.4 Empirical strategy and identification

2.4.1 Sibship size and birth order effects

In our analysis we are interested in the effects of sibship size and composition on an individual's 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). Indeed, if parents have a preference for the first children they have, and invest comparatively more resources in them, 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 birth order and sibship size are highly correlated. In particular, while one can assess the effect of family size on firstborns by looking at firstborns' outcomes in families of different size, it is not possible to look, for instance, at the outcome of a fourth-born child when sibship size changes from two to three, as fourth born children are only found in larger families.

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

$$M_{ij} = \alpha_0 + \sum_{k=1}^{K} \boldsymbol{\alpha}_{1k} bo_{ijk} + \boldsymbol{\alpha}_2 \mathbf{X}_{ij} + u_j + \epsilon_{ij} \tag{2.1}$$

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_{ijk} is a dichotomous indicator for the child being of birth order $k = 1, ..K$ where K is the maximum number of children (i.e. family size) 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).[17] u_j is a family fixed effect,[18] and ϵ_{ij} an idiosyncratic error.

The effect of sibship size is captured in equation (2.1) by the family fixed effects, which

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

[18]Since our sample includes siblings born from the same mother, we are *de facto* controlling for mother fixed effects.

control for any (observed and unobserved) difference between families. The inclusion of family fixed effects also helps address the birth order endogeneity due, for example, to birth-order selective child fostering (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 subtract the birth order effects from the dependent variable, i.e. we compute the difference $\hat{NM}_{ij} = M_{ij} - \sum_{k=1}^{K} \hat{\boldsymbol{\alpha}}_{1k} bo_{ijk}$ where NM stands for 'netted migration', and use it as the dependent variable in a second step.[19] Hence, the following equation is estimated:

$$\hat{NM}_{ij} = \beta_0 + \beta_1 S_{ij} + \boldsymbol{\beta}_2 \mathbf{X}_{ij} + \boldsymbol{\beta}_3 \mathbf{W}_j + v_{ij} \tag{2.2}$$

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 that 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. In some specifications, we also control for maternal health, the father not being 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.). Since the dependent variable has been generated by a regression, standard errors are corrected by weighting the estimation with the inverse of the standard error of $\hat{NM}_{ij}$.[20] We estimate equation (2.2) by using either WLS or 2SLS (see the next Section). Throughout, standard errors are clustered at the household level to account for potential error correlation across siblings.

2.4.2 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's effect estimate in equation (2.2) will provide spurious evidence. In other words, parental fertility may be endogenous

[19] Coefficients of all birth order indicators are recovered using the method described in Suits (1984).

[20] See, for instance Lewis and Linzer (2005). We also run estimates using OLS and White robust standard errors, and the results on the effect of sibship size did not change.

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), for instance, postulate 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 association 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 contraceptive methods or who are not currently using them. More specifically, we construct an indicator variable for infertility shocks taking value one if a woman declares she never used contraception or she has stopped using the previous method because of infertility ('infertility shock') and zero otherwise (Agüero and Marks, 2008).[21] 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, relevance and the exclusion restriction assumption — which we discuss below.

Infertility or subfertility conditions have been already used in the economic literature to estimate the effect of the number of children and fertility timing on mothers' labor market outcomes (see, for instance, Agüero and Marks, 2008, 2011; Schultz, 2008). There is evidence that infertility is virtually random, i.e. it is independent of the background

[21]Such shocks may only be temporary, and have emerged relatively recently. This means that even subfertile women may have large families. Unfortunately, we do not have any information about the age when these problems first showed up.

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 U.S. data (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 will be controlled for in our models).

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). Their exogeneity is generally supported by the medical literature. A few papers using administrative data in which rich labor market and health data are merged show, for example, that miscarrying is not generally significantly associated with worse labor market outcomes (e.g., work absences) before miscarriage (Karimi, 2014; Markussen and Strøm, 2015). 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.[22] Only 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's 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). There is a potential issue of measurement error with this instrument, since women may be unaware of miscarriages or, especially older women, may fail to recall them. Misreporting would generally affect the power of the instrument, but we do not expect

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

any specific pattern of correlation between it and parents' attitudes towards child out-migration conditional on the observables (including a quadratic in mother's age). 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 and Mexico had the stricter anti-abortion legislation in Latin America during the period under consideration.[23]

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 have an 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, chronic illness/disability, marital status and husband's characteristics (age, education and absence).

[24]

In Figures 2.4 and 2.5 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 2.5). In particular, we use ENADID data to plot the life-cycle profile of the total number of live births by women's age and infertility shock and miscarriage status. Figure 2.4 shows that women who ever experienced an infertility condition generally have a lower number of children, and that differences in fertility tend to become evident after

[23] For women who voluntary abort, the instrument would be endogenous. However, there is no evident sign in our data that a relevant share of the recorded abortions could be voluntary. For instance, Catholic women in our sample do not tend to abort significantly less than other women (this check can be done only for the 1997 wave, which includes information on religion): for the first group the incidence of abortion is 4.6 percent and for the second group is 4.8 percent. In case the instrument is substantially contaminated by voluntary abortions, we would expect IV estimates to be biased in the same direction as OLS. Indeed, omitting subscripts and in the models without controls, if we define as $M = \beta_0 + \beta_1 S + v$ the migration equation, where M and S are child migration status and sibship size, respectively, and $S = \gamma_0 + \gamma_1 Z + u$ the sibship size equation (the first stage) and Z the instrument (abortion), $\beta_{1,OLS} = \beta_1 + Cov(S, v)/Var(S)$ while $\beta_{1,IV} = \beta_1 + Cov(Z, v)/Cov(Z, S)$, where $Cov(Z, S) < 0$ and $sign(Cov(S, v)) = -sign(Cov(Z, v))$. In case, for instance, unobserved mother's total desired fertility is positively correlated with children's migration and a substantial share of abortions are voluntary, both OLS and IV will be similarly upward biased.

[24] In Table A1 in Appendix we report women's characteristics by fertility and infertility status. Descriptive statistics suggest that women who experienced infertility differ from those who did not in some aspects, such as age, schooling, single-status and husband's welfare (Panel A). A similar 'positive' selection is found for women who had an abortion at their first pregnancy (Panel B). Hence, we do include these characteristics as controls in our regression models (with the exception of husband's income that is only available for the 1997 wave).

the age of 30. Similarly, Figure 2.5 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 family size, overall Mexican women were able to achieve a generally high fertility rate by the end of their fecund life span. This is due to the fact that exogenous infertility shocks, as defined in this and related papers, clearly affect the number of children a woman can have but they may also be temporary (i.e. secondary infertility) or treatable so that fertility may eventually be restored.

2.5 Results

2.5.1 Birth order effects

We start by estimating the impact of birth order on individual migration, as specified in equation (2.1), 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 or for other omitted family-specific unobservable factors. The first column of Table 2.2 reports 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).[25] Indeed, child age is correlated with birth order and it is also likely to have a (non-linear) relationship with migration (this is why we include the age quadratic term).

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 3.6 percentage points (p.p. hereafter). Moreover, 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 2.1 p.p. less likely to migrate than firstborns (column 2). 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 migration rate). The coefficients for the following birth orders are larger in absolute value

[25]By including child age and cohort dummies, with household fixed effects, we are also de facto controlling for birth spacing between siblings.

and peak for birth orders 9 and 10 or more (-16.6 and -20 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.[26] We observe a negative birth order gradient for boys (the coefficients on third and higher parities are negative and significant), consistently with average results above. The interactions of being female with birth order dummies are not statistically significant, suggesting that the birth-order gradient in child migration is not statistically different between boys and girls. Yet, the latter holds for all parities but for firstborns: in column (4) the female main effect shows that female firstborns are significantly less likely to migrate than male firstborns. Overall, these estimates suggest that the chances of migration are not equally distributed across children within the same family. Low-parity children are in general more likely to migrate and this may be explained by the fact that, if migration is also a household-level investment strategy, the family will have more time to reap the benefits of migration. Yet, from our fidings a firstborn daughter is significantly less likely to migrate than a firstborn son by 3 p.p., which means a reduction in the probability of migration of roughly 60 percent at the sample average migration rate. We further explore these gendered effects in light of potential parental preferences later on in Section 2.6.

2.5.2 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 described above, we start by reporting OLS estimates as a benchmark model, where the dependent variable is 'netted migration' (see Section 2.4.1). The number of siblings is tallied as the number of currently living biological brothers and sisters of each child.[27] The first column of Table 2.3 reports OLS results for a linear specification including sibship size. The highly significant coefficient implies that, on average and

[26]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 a sample selection, we rather adopt a pooled estimation including interaction effects with gender.

[27]We drop from our definition of siblings those currently deceased. This is so because of two reasons: (i) 70 per cent of deceased children in our sample died before the first year of life, 90 per cent of them before the second one; (ii) the focus of our analysis is not on young children so that we need to take into account siblings who actually 'had enough time' to both receive and compete over household resources, so that can exclude infant deaths. In Appendix, we report robustness checks related to concerns about the endogeneity of our definition of sibship size and birth order based on ever-born children, i.e. currently alive or deceased.

after controlling for birth order effects, adding one sibling is associated with a 1.1 p.p. higher likelihood to migrate of young adults (+17 percent at the sample mean). The same effect holds once we include individual level controls, namely child gender, age, age squared and years of birth indicators (column 2). When we allow for differential effects by child gender (column 3), 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 extra sibling raises the migration probability more for sons than for daughters by 0.8 p.p. In columns (4) to (7), we run the same regressions above while adding further parental, household and aggregate-level controls in order to account for potential confounding factors of the relationship between family size and offspring's migration. Specifically, in column (3) and (4) we include parental covariates, which may predict completed fertility and affect child migration, namely mother's years of birth indicators, age at first pregnancy, chronic illness, single status (i.e. widow, divorced, single de facto), father's decade of birth indicators, mothers' and father's (quadratic) age and years of schooling.[28] In column (5) and (6) we further add municipality fixed effects that, conditional on family size, control for rural vs. urban residence along with many other local factors related to different cultural or economic conditions, which may have an effect on fertility and migration (e.g. employment rates, migration intensity, access to contraception, social services etc). All in all, the sibiship size effect is essentially unchanged when we control for all of the aformentioned factors, and the same holds for the differential effect by gender.

Yet, as mentioned in the methodological section, the coefficients on sibship size reported in Table 2.3 are still likely to be biased, even when including a rich set of demographic and economic controls. This is so as fertility may be endogenous with respect to child out-migration. Thus 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 2.4 we present two-stage least squares (2SLS) estimates using a linear version of our 'saturated' specification (with controls) and the two-step methodology, as outlined above, to estimate equation (2.2). In column (1) we instrument sibship size with an indicator variable for infertility shocks taking value one if the woman declares she never used or

[28] We are de facto also controlling for mother's age at birth, which is a linear combination of child's age and mother's age.

she stopped using contraception because of infertility. In column (2), instead, we report results using a woman's experience of miscarriage on her first pregnancy as an instrument. Eventually, in column (3) we present results using both instruments in an over-identified equation model. Throughout all models, the first stage results point to a strong and highly significant relationship between infertility /fertility shocks and completed fertility. In particular, women who experienced an infertility shock have a reduction in their number of children of nearly 0.5 ($t = -5.2$) with an $F-$statistic of 26.9 (column 1). The negative impact of miscarriage on completed fertility is similar in magnitude (-0.437) with an $F-$statistic of 19.13 (column 2). Also the $F-$statistic of the joint significance of the instruments in the over-identified model is as high as 23.37 (column 3). The sibship size effects estimated using 2SLS are always small and statistically insignificant at standard confidence levels. The Anderson-Rubin $F-$statistic cannot reject in none of the models that the coefficient of the instrument is zero in the reduce form, and the Hansen $J-$statistics confirms the validity of the instruments in the overidentified model. Interestingly enough, the point estimate of the effect of sibship size on child migration obtained with the abortion instrument (which might include voluntary abortions) is lower than that obtained with the infertility instrument, which we consider much less (or not) affected by endogeneity issues, and much lower than the OLS estimate, a fact that is inconsistent with induced abortions being a substantial share of total abortions (cf. footnote 23).

In Table 2.5 we report results of the same 2SLS regressions as above while testing the sibship size differential effect by gender in the pooled sample with interaction terms.[29] Results do not point to any significant difference in the impact of sibiship size bewteen boys and girls, as it turns out to be insignificant for both (columns 1-3). When using miscarriage as an instrument, though, we cannot draw strong conclusions as the $F-$statistic for the interacted endogenous variable is rather low (4.27, column 2). However, even in this case the Anderson-Rubin $F-$statistic confirms that we cannot reject the hypothesis of sibship size not affecting child migration.

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 abroad.

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

2.5.3 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.[30] In so doing we are able to check the robustness of our baseline family size effect 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 consequence, household-level regressions allow us to include also one-child households in the sample.[31] Thus, we estimate a specification as follows:

$$m_j = \gamma_0 + \gamma_1 n_j + \boldsymbol{\gamma}_2 \mathbf{W}_j + v_j \tag{2.3}$$

where the dependent variable is the number of migrants in household j and the independent variable of interest is n_j, i.e. the total number of children 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).

Results are reported in Table 2.6. 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 0.012 ($t = 10.6$). Computed at the average number of child migrants per household in the sample (0.077), this corresponds to a 16 percent increase. Column (2) reports the IV estimate using the infertility instrument. The first stage shows a reduction of -0.64 ($t = -10.2$) in the total number of children per woman who experienced an infertiliy shock, with an $F-$statistic of 104. The first-stage coefficient is a bit higher in magnitude than that obtained in the child-level estimates (-0.5). 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

[30] More precisely, our unit of analysis is the biological family.

[31] 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.

and statistically insignificant. Column (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 of about 40. 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. -0.44 ($t = -6.3$). Like for the previous instrument, also in this case no significant effect is detected in the second stage. The same happens in the overidentified model in column (4).

The household-level estimates in this section confirm the results of Section 2.5.2 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.

2.6 Sibling gender composition

Our estimates so far show that gender is a robust predictor of migration and, *ceteris paribus*, boys – especially firstborns – are systematically more likely to migrate to the U.S. than girls in Mexican families. This points to a migration male-dominated phenomenon (e.g., Cerrutti and Massey, 2001a) 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 parental 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). Thus, in order to explore the scope of sibling rivalry by gender, we test how sibling composition influences child migration investment by running two sets of regressions as reported in Table 2.7. First, we estimate migration equations with family and birth order fixed effects (i.e., conditioning on both family size and birth order) on the full sample of children as a function of the number of their older brothers, while controlling for the number of older siblings, and child's age (linear and squared). Results in column (1) show that, ceteris paribus, 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 gender and age composition of siblings in children's migration outcomes, which does not differ significantly by the child's gender (column 2).

We further exploit the gendered migration pattern and the fact that siblings are likely to migrate in order of birth (with higher parities being less likely to migrate, as shown by our former estimates in section 5.1) to test the hypothesis of parental son preference. 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.[32] Results in columns (3) and (4) in Table 2.7 show that, conditional on older siblings' composition, 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 1.2 p.p. ($t = -2$). 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 daughter with a next-born brother may be less likely to migrate than a girl with a next-born sister. In other words, when parents face the decision whether to send a daughter abroad, they seem to prefer to invest in the migration of her 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) and typically girls turns out to be disadvantaged when they compete with boys. (Dunn and Plomin, 1990; Kristin and Anne, 1994; Morduch, 2000). Indeed, our findings are suggestive that children, especially girls, with relatively more brothers than sisters are less likely to migrate abroad than their peers.

These results, combined with the birth order effects reported above, are consistent with the argument that a low-parity Mexican boy may be more valuable to send as a migrant abroad than a girl. Indeed, labor market returns for Mexican boys in the U.S. were relatively higher in the 1990s (e.g., in the farm sector). In addition, the opportunity cost of sending girls abroad may be higher because they usually take care of chores and family

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

duties at home or are in charge of being close to parents in their elderly age. Hence, social norms or practices combined with market returns on the migration investment may explain the pro-male biased pattern of mass Mexico-U.S. migration and document, similarly to other developing contexts, that young females tend to have less access to human capital investment and enhancing economic opportunities than it is the case for males.

2.7 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 rates 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 money to old-age parents' living arrangements.

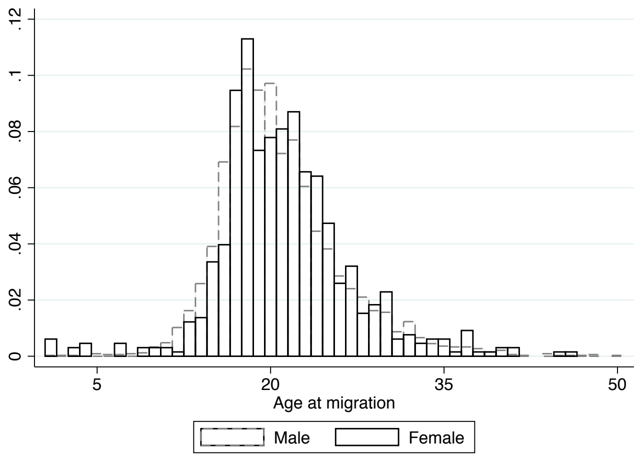
We use a rich household-survey dataset on teenagers and young adults to examine the causal effects of sibship size, birth order and sibling composition on migration outcomes in Mexico. Mexican migration, mainly to the U.S., is an enduring flow accounting for one third of total U.S. immigration and one-tenth 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 firstborn males, are more likely to migrate, while having relatively more sisters than brothers systematically increases the likelihood to migrate. Moreover, girls are less likely to migrate when their next parity is a male. This is consistent with the argument that, in scarce-resource contexts, girls can be seen as more economically and socially costly to parents, so that boys end up having more economic opportunities than girls even through migration.

Our findings adds to the migration literature by shedding new light on the role of the family in determining international migration choices. Labor mobility, especially from poor to rich settings, is one of the most important ways through which young adults can expand their human capital and earning potentials. The type of family-based migration from Mexico to the U.S. during the 1990s is of substantial and growing importance for many other developing countries (e.g. in Asia and Africa) currently affected by both high fertility and international migration (e.g., Hatton and Williamson, 2003). Despite the easy to see association between fertility rates and migration, we provide evidence that large families are unlikely to be a systematic driver of migration. This is in line with other recent findings that show high fertility in developing contexts is not necessarily bad for children's economic outcomes (e.g., Qian, 2009). In terms of policy, understanding the link between fertility and migration is especially relevant today as many governments in developing countries have attempted to curb population growth as a way of increasing average human capital investment and possibly reduce migration (e.g., China and India, the world's two most populous countries, have experimented with different family planning policy first to control family size). Yet, while our empirical findings do not point to a causal link between fertility and migration, they hint to the fact that parental investment in offspring's migration may matter for lifetime fertility choices. This is so as in a context of poor resources and weak safety net institutions, children may be a key social security valve for parents such that high migration opportunities to rich countries increase the value of having children. Hence, effective social safety nets (such as old age pensions) or even the development of credit and insurance markets could lead to a reduction in both migration and fertility, and perhaps also to less bias against girls.

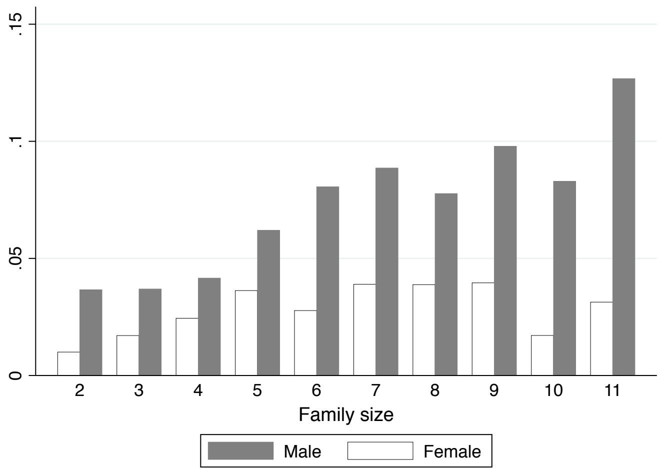
2.8 Figures and Tables

Figure 2.1: Mexican individual (non-tied) migration by age and gender



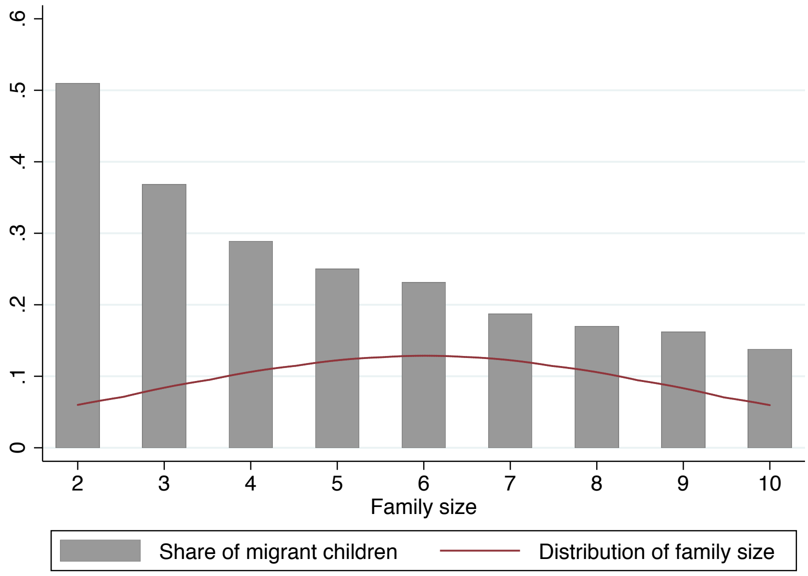
Source: Our computations on ENADID, 1992 and 1997.

Figure 2.2: Individual migration rate by sibship size



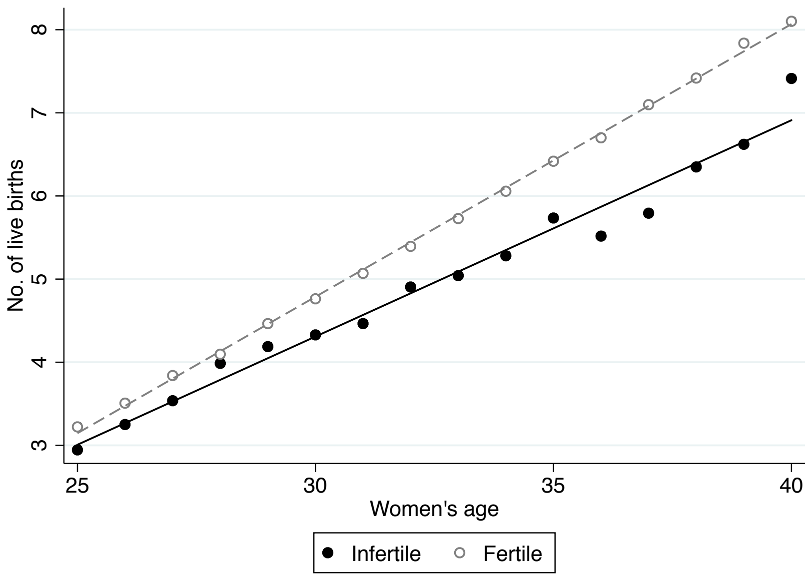
Source: Our computations on ENADID, 1992 and 1997.

Figure 2.3: Ratio of migrant children by sibship size



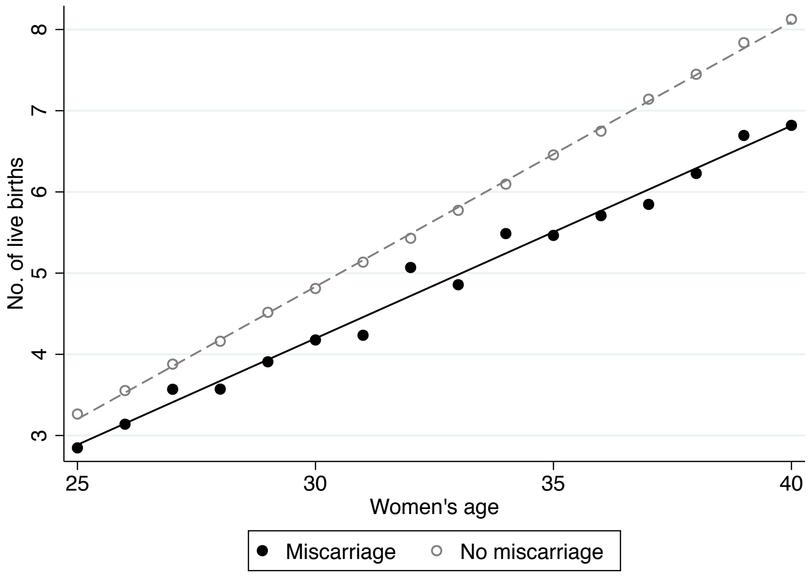
Source: Our computations on ENADID, 1992 and 1997 (subsample of households with migrants).

Figure 2.4: Cumulative number of children by women's infertility shock status



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

Figure 2.5: 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 2.1: Individual sample characteristics by migration status

	Non-migrants	Migrants	P-values
Age	18.878	20.982	0.000
Female	0.458	0.250	0.000
N. of siblings	5.071	5.869	0.000
Birth order 1	0.181	0.192	0.300
Birth order 2	0.231	0.225	0.555
Birth order 3	0.178	0.178	0.978
Birth order 4	0.137	0.154	0.077
Birth order 5	0.102	0.102	0.993
Birth order 6	0.071	0.073	0.781
Birth order 7	0.046	0.041	0.343
Birth order 8	0.028	0.021	0.100
Birth order 9	0.014	0.009	0.121
Birth order 10+	0.011	0.006	0.107
Mother's age	44.847	46.065	0.000
Mother's age at first pregnancy	20.025	19.554	0.000
Mother's years of schooling	3.953	3.286	0.000
Mother chronic illness	0.022	0.017	0.203
Single mother	0.176	0.220	0.000
Mother's labor income	552.094	775.354	0.000
Father's age	38.745	19.605	0.000
Father's years of schooling	3.683	1.328	0.000
Father's labor income	2,011.533	3,135.883	0.000
Observations	25,349	1,394	

Note. Source: ENADID, 1992 and 1997. The estimation sample includes individuals aged 15–25 whose mothers are not using contraceptive methods or never did.

Table 2.2: Birth order effects

Variables	(1)	(2)	(3)	(4)
female	-0.036***	-0.035***	-0.032***	-0.031***
	(0.003)	(0.003)	(0.006)	(0.007)
birth order	-0.019***		-0.019***	
	(0.003)		(0.003)	
birth order × female			-0.001	
			(0.001)	
birth order 2		-0.002		0.002
		(0.005)		(0.006)
birth order 3		-0.021***		-0.023***
		(0.007)		(0.008)
birth order 4		-0.038***		-0.034***
		(0.010)		(0.011)
birth order 5		-0.068***		-0.070***
		(0.013)		(0.014)
birth order 6		-0.086***		-0.077***
		(0.016)		(0.017)
birth order 7		-0.112***		-0.103***
		(0.019)		(0.020)
birth order 8		-0.136***		-0.140***
		(0.022)		(0.023)
birth order 9		-0.161***		-0.166***
		(0.026)		(0.028)
birth order 10+		-0.199***		-0.188***
		(0.030)		(0.033)
birth order 2, female				-0.011
				(0.009)
birth order 3, female				0.005
				(0.010)
birth order 4, female				-0.010
				(0.010)
birth order 5, female				0.006
				(0.011)
birth order 6, female				-0.018
				(0.012)
birth order 7, female				-0.017
				(0.015)
birth order 8, female				0.010
				(0.018)
birth order 9, female				0.012
				(0.024)
birth order 10+, female				-0.022
				(0.027)
age	0.020**	0.021**	0.020**	0.021**
	(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	YES	YES	YES	YES
Family fixed effects	YES	YES	YES	YES
Observations	26,743	26,743	26,743	26,743
Number of households	10,139	10,139	10,139	10,139
R-squared	0.050	0.052	0.050	0.053

Note. The dependent variable is a dichotomous indicator of the child's migration status. The model is estimated using OLS. Sibship size is absorbed by family fixed effects. Standard errors clustered at the household level in parentheses. *,** and *** denote statistical significance at 10, 5 and 1 percent level, respectively.

Table 2.3: Sibship size effect: WLS estimates

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)
N. siblings	0.011***	0.011***	0.014***	0.010***	0.013***	0.010***	0.013***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
N. siblings $\times$ female$^{(a)}$			-0.008***		-0.007***		-0.006***
			(0.001)		(0.001)		(0.001)
female		-0.038***	-0.036***	-0.033***	-0.031***	-0.033***	-0.031***
		(0.003)	(0.006)	(0.003)	(0.006)	(0.003)	(0.003)
Individual's controls	NO	YES	YES	YES	YES	YES	YES
Mother's controls	NO	NO	NO	YES	YES	YES	YES
Father's controls	NO	NO	NO	YES	YES	YES	YES
Municipality indicators	NO	NO	NO	NO	NO	YES	YES
Observations	26,743	26,743	26,743	26,743	26,743	26,743	26,743
R-squared	0.013	0.054	0.055	0.177	0.178	0.202	0.203

Note. The dependent variable is *netted migration* (see Section 2.4). The model is estimated using Weighted Least Squares (WLS), in which weights are the inverse of the standard errors of *netted migration.* Individual's controls include year of birth indicators, age, age squared; mother's controls include year of birth indicators, age and age squared, age at first pregnancy, years of schooling, indicators for mother's chronic illness and being single; father's controls include decade of birth indicators, age and age squared, years of schooling. Standard errors clustered at the household level in parentheses. $^{(a)}$ The number of siblings is demeaned before taking the interaction. *,** and *** denote statistical significance at 10, 5 and 1 percent level, respectively.

Table 2.4: Sibliship size effect: 2SLS estimates

Variables	(1)	(2)	(3)
Second stage			
N. siblings	0.004	-0.018	-0.005
	(0.014)	(0.023)	(0.012)
female	-0.033***	-0.033***	-0.033***
	(0.003)	(0.003)	(0.003)
IV:	infertility	miscarriage	overidentified
Anderson-Rubin $F-$statistic	0.0734	0.686	0.389
	[0.787]	[0.407]	[0.678]
Hansen $J-$statistic			0.737
			[0.391]
First stage — N. siblings			
infertility	-0.494***		-0.491***
	(0.095)		(0.095)
miscarriage		-0.437***	-0.433***
		(0.10)	(0.10)
Angrist-Pischke $F-$statistic instrument(s)	26.90	19.13	23.37
Individual's controls	YES	YES	YES
Mother's controls	YES	YES	YES
Father's controls	YES	YES	YES
Municipality indicators	YES	YES	YES
Observations	26,743	26,743	26,743

Note. The dependent variable is *netted migration* (see Section 2.4). Observations are weighted by the inverse of the standard error of *netted migration*. Individual's controls include year of birth indicators, age, age squared; mother's controls include year of birth indicators, age and age squared, age at first pregnancy, years of schooling, indicators for mother's chronic illness and being single; father's controls include decade of birth indicators, age and age squared, years of schooling. Standard errors clustered at the household level in parentheses. $P-$values are reported in brackets. *,** and *** denote statistical significance at 10, 5 and 1 percent level, respectively.

Table 2.5: Child gender and sibliship size effect: 2SLS estimates

Variables	(1)	(2)	(3)
Second stage			
N. siblings	0.005	-0.065	-0.007
	(0.016)	(0.048)	(0.015)
N. siblings × female[(a)]	-0.005	0.112	0.005
	(0.013)	(0.079)	(0.013)
female	-0.032***	-0.064***	-0.034***
	(0.004)	(0.022)	(0.005)
IV:	infertility	miscarriage	overidentified
Anderson-Rubin $F-$statistic	0.0744	2.210	1.150
	[0.928]	[0.110]	[0.331]
Hansen $J-$statistic			4.399
			[0.111]
First stage — N. siblings			
infertility	-0.567***		-0.564***
	(0.109)		(0.108)
infertility × female	0.168		0.169
	(0.115)		(0.115)
miscarriage		-0.453***	-0.450***
		(0.117)	(0.117)
miscarriage × female		0.037	0.038
		(0.106)	(0.105)
Angrist-Pischke $F-$statistic instrument(s)	28.62	11.98	15.68
First stage — N. siblings × female			
infertility	0.125***		0.125***
	(0.038)		(0.038)
infertility × female	-0.694***		-0.691***
	(0.131)		(0.131)
miscarriage		-0.067	-0.068
		(0.044)	(0.043)
miscarriage × female		-0.261**	-0.254*
		(0.131)	(0.130)
Angrist-Pischke $F-$statistic instrument(s)	26.93	4.27	13.83
Individual's controls	YES	YES	YES
Mother's controls	YES	YES	YES
Father's controls	YES	YES	YES
Municipality indicators	YES	YES	YES
Observations	26,743	26,743	26,743

Note. The dependent variable is *netted migration* (see Section 2.4). Observations are weighted by the inverse of the standard error of *netted migration.* Individual's controls include year of birth indicators, age, age squared; mother's controls include year of birth indicators, age and age squared, age at first pregnancy, years of schooling, indicators for mother's chronic illness and being single; father's controls include decade of birth indicators, age and age squared, years of schooling. Standard errors clustered at the household level in parentheses. $P-$values are reported in brackets. [(a)] The number of siblings is demeaned before taking the interaction. *,** and *** denote statistical significance at 10, 5 and 1 percent level, respectively.

Table 2.6: Sibship size effect: Household-level estimates

Variables	(1) OLS	(2) 2SLS	(3) 2SLS	(4) 2SLS
Second stage				
N. children	0.012***	-0.001	-0.014	-0.005
	(0.001)	(0.013)	(0.022)	(0.011)
IV:	—	infertility	miscarriage	overidentified
Anderson-Rubin $F-$statistic		0.006	0.400	0.203
		[0.940]	[0.527]	[0.816]
Hansen $J-$statistic				0.265
				[0.607]
First stage — N. children				
infertility		-0.643***		-0.640***
		(0.063)		(0.063)
miscarriage			-0.442***	-0.438***
			(0.070)	(0.070)
Angrist-Pischke $F-$statistic instrument(s)		103.87	39.81	72.40
Mother's controls	YES	YES	YES	YES
Father's controls	YES	YES	YES	YES
Municipality indicators	YES	YES	YES	YES
Observations	17,544	17,544	17,544	17,544

Note. The dependent variable is a dummy for the child's migration status. Mother's controls include year of birth indicators, age and age squared, age at first pregnancy, years of schooling, indicators for mother's chronic illness and being single; father's controls include decade of birth indicators, age and age squared, years of schooling. $P-$values are reported in brackets. *,** and *** denote statistical significance at 10, 5 and 1 percent level, respectively.

Table 2.7: Siblings' composition effect: OLS estimates

Variables	(1)	(2)	(3)	(4)
N. older brothers	-0.014***	-0.016***	-0.017***	-0.018***
	(0.004)	(0.004)	(0.004)	(0.005)
N. older siblings	-0.015***	-0.013***	-0.013***	-0.012***
	(0.004)	(0.004)	(0.004)	(0.004)
female	-0.028***	-0.023***	-0.025***	-0.014**
	(0.003)	(0.005)	(0.004)	(0.006)
N. older brothers × female		0.001		-0.001
		(0.003)		(0.003)
N. older siblings × female		-0.002		-0.002
		(0.002)		(0.002)
next brother			-0.005	0.000
			(0.003)	(0.004)
next brother × female				-0.012**
				(0.006)
Age, age squared	YES	YES	YES	YES
Birth order fixed effects	YES	YES	YES	YES
Year of birth indicators	YES	YES	YES	YES
Family fixed effects	YES	YES	YES	YES
Observations	26,743	26,743	26,743	26,743
Number of hid	10,139	10,139	10,139	10,139
R-squared	0.053	0.053	0.053	0.053

Note. The dependent variable is a dichotomous indicator of the child's migration status. The model is estimated using OLS. Sibship size is absorbed by family fixed effects. Standard errors clustered at the household level in parentheses. *,** and *** denote statistical significance at 10, 5 and 1 percent level, respectively.

Chapter 3

The effect of immigrants' voting right: evidence from a natural experiment

3.1 Introduction

During the last decades, immigration to European countries has increased exponentially, contributing to the creation of more ethnically heterogenous societies. Together with the increase in immigrants, politics in Europe has taken a rightward turn, driven by the success of nationalist and anti immigrants parties. Among scholars, it has risen questions about the relationship between ethnic diversity and political and economic outcomes. The main hyphotesis addressed by the existing literature concerns with the effect of the *mere presence* of immigrants and the increasing ethnic diversity on policy outcomes, through the changes in native voters' political attitudes and voting behavior. The contribution of this paper is to show how political and economic outcomes change as a consequence of an *increase in political power of ethnic minorities*, given by the extension of voting rights to immigrants of non EU origin.

Belgium recently extended the right to vote in local elections to immigrants from outside the European Union with at least five years of legal residence in the country. The act providing for non citizen voting was adopted by the Belgian Parliament as the Law of 19 March 2004, and went into effect in 2006. This natural experiment is exploited here to study how the new group of enfranchised affects political outcomes at municipal level, by looking at electoral results for 589 municipalities during the period 1988-2012.

Since the 70s, extending the right to vote to immigrants in local elections has been a sensitive issue of the political debate in Europe. Non-citizens' active political participation is widely recognized as a stimulating factor for immigrants' integration in the host society. Supporters of the expansion of voting rights claim that all residents who pay taxes, contribute to the social security and take part at the community life should have the right to contribute to the selection of representatives. On the other hand, there is a common fear that immigrants' votes could tilt the political balance with potentially unfavorable outcomes for natives, e.g. with respect to the level of redistribution. In Belgium, Flemish politicians were the most resistant to the enfranchisement of immigrants, arguing that the introduction of a new group of voters would shatter the delicate balance of power between Dutch and French-speaking communities and possibly create a disproportionate benefit for Francophone political parties. Behind anecdotal evidence, there is very little systematic evidence on the socio-economic consequences of enfranchisement of immigrants. The aim of this paper is to fill this gap by exploiting the 2004 Belgian law.

This paper relates to two strands of literature. The first one takes a political economy approach and examines the effects of the extension of voting rights on the size of government (Husted and Kenny, 1997; Lott, 1999; Cascio and Washington, 2014). The main message from this literature is that the enfranchisement of minorities (e.g. black, poorer or women) determines an increase in welfare or government expenditures, mainly driven by the change of the median voter. The second strand of literature examines the relationship between the mere presence of immigrants (and immigration-driven increases in ethnic diversity) and natives' voting behavior or political attitudes (Razin et al., 2002; Böheim and Mayr, 2005; Dahlberg et al., 2012; Harmon, 2013; Barone et al., 2014; Mayda et al., 2015). The negative relationship between ethnic diversity and left-wing political preferences, as well as the level of public spending at local level, is confirmed in most studies.

Belgium is an interesting setting to study the impact of enfranchising non citizens of non Eu origin. First, Belgium has historically attracted significant immigration flows and the number of immigrants has steadily increased since the 80s. Second, according to the Belgian Constitution, local governments can take any initiatives that is beneficial to local interests and no other government has legal responsibility for the concerned field of action. Local governments have great autonomy and responsibility on a number of important issues, including expenditures on education, security and social policies, culture, urbanism and environment. It suggests that local politics is the relevant decision making level for many aspects of life. Moreover, Belgium is a country of different linguistic and political realities. The three regions of the Flanders, Wallonia and Brussels differ in many dimensions, including political preferences and immigration policy (see Section 3.2.1). These differences offer the opportunity to explore heterogeneous effects of the 2004 Law and can provide key insights on the interpretation of the results.

The paper is structured as follows: in Section 2, the institutional setting of Belgium is discussed; Section 3 presents data and the empirical strategy; in Section 4 results of the empirical analysis are presented; Section 5 concludes.

3.2 Institutional setting of Belgium

Specific features of the Belgian immigration policy and political system are particularly relevant for the empirical analysis. This section gives a brief overview of these two aspects of the Belgian context.

3.2.1 The Belgian immigration policy

In Belgium, competencies over immigration issues are split between the federal state and the regions. The federal state is responsible for admissions, removals, residence rights and laws on citizenship of immigrants; while regional authorities have jurisdiction over 'the reception and integration of immigrants', by promoting their participation in the host society (as established by the law of 8 August 1980).

The regions of Flanders and Wallonia adopted very different immigrants' integration policies, inspired by diverse values and pursuing contrasting objectives. It results in immigrants having different rights and duties, according to their place of residence. Flanders' policy highlights the importance of cultural diversity and ethnic identities and combines an assimilationist policy targeting new comers with a multicultural policy addressed to established ethic minority groups. A large network of implementing actors, coordinated by a centralized organization, administers language courses and a civic integration course (focusing on rules and regulations, common norms and values), as part of the compulsory integration trajectories. Importantly, non attendance of the courses is punished by the means of an administrative fine.

The Francophone Belgium, instead, puts emphasis on economic and social inclusion and promotes it through color-blind policies, as a good strategy to promote assimilation, opposite to targeted or multiculturalist policies, responsible for stressing differences among natives and immigrants. The Walloon integration process is very decentralized, and most of the initiatives promoting integration are taken at municipality level and implemented by local organizations. Immigrants are not obliged to attend integration courses and the the total budget devoted to integration policy is much lower in Wallonia with respect to Flanders (Adam and Jacobs, 2014).

The region of Brussels combines together Flemish and Francophone integration policies, led by local organizations, mainly in poor and deprived neighborhoods.

Another relevant difference at regional level concern parties' position regarding immigration and immigrants' rigths. On the one hand, Francophone political parties are willing to promote naturalization as a natural mean for foreigners' integration; on the other, Flemish parties claim the importance of integration before naturalization and ask for a stricter immigration policy.

3.2.2 The Belgian political system

The Belgian electoral system is divided according to linguistic lines: Flemish parties compete for votes in the region of Flanders, while Francophone parties run in the region of Wallonia. Only in the region of Brussels there are both parties. During the 70s the major statewide parties split along regional lines, and all parties created after that are organized at community level and only represent the interest of part of the population. The Belgian political system can then be classified as a two party systems (Dandoy, 2014).

Immigration is a relevant issue in the electoral platforms of Belgian parties, and it has been highly politicized. It allows labeling and ranking parties according to their position on immigration issues. The attitudes of local parties toward immigration is recovered by looking at the manifestos of their national counterparts; more specifically, by looking at the analysis conducted by Dandoy (2014), who studied the electoral platforms of the main Belgian national parties between 1977 and 2007. Belgian local politics is highly nationalized, with many national parties taking part at local elections. It happens more in Flanders than the Walloon and Brussels regions. As Figure 3.1 reveal, for the local elections in the period 1988-2012, in Flanders, more than 80% of the votes was gained by parties with a national counterparts (or parties that can be labeled as belonging to one major party family with a national counterpart); the share is much lower (about 40%) in the other two regions, where municipal elections still remain more 'localized', i.e. there are many local party with no national counterparts running for elections.

Many significant aspects come out from the analysis of national parties' political manifestos and allow constructing a rank of parties over immigration attitudes. First, it emerges that, overall, Flemish parties allocate more attention to immigration than Francophone parties. Second, as expected, extreme right parties' manifestos dedicate more space to immigration issues than other parties, with an exception. When looking at each single election in the period 1977-2007, it emerges that during the 90s the Flemish liberal party focused on the issue of immigration more than any other parties. Third, by looking at a specific policy sector of immigration, namely *immigration integration*, it appears that extreme-right and liberal parties dedicate more attention to the issue with respect to socialist and Christian Democrat parties. Last, when looking at the specific issues of the migration debate a clear pattern emerges in both regions: while Christian Democrat and socialist parties give more space to the discussion of democracy and rights, including the debate on voting rights; extreme-right and liberal parties focus on law and order, with special emphasis on the relation between immigrants and crime and the implications for public order. Given these elements of the parties' manifestos, main parties taking part at local election are ranked from *very-pro* to *very-anti* immigrants parties.

3.3 Data and methods

Data come from multiple sources. As for the outcomes variables, the focus is on both election results and expenditures at municipal level. Data on five election rounds results (1988,1994 2000 before the reform and 2006, 2012 after the reform) have been collected from the web, using Python. The regional authorities of Flanders provided data on municipal expenditures and revenues at local level (for the period 2003-2013).

The Ministry of Interior released data on potential and registered non Eu voters for local elections in 2006 and 2012. Data on the characteristics of the Belgian municipalities over time are (mainly) available online through the Statistics Belgium.

Table 3.1 shows the share of potential non Eu non citizens voters measured in 2006 and 2012 elections. Interestingly, this share has increased over time, above all in Flanders. Moreover, it is important to notice that potential non Eu voters do not spread equally across regions: the municipalities of the Brussels region have a much higher share, while Flanders and Wallonia have a smaller and similar share of non Eu potential voters.

3.3.1 Empirical strategy

The empirical strategy exploits two sources of variation: the first is time variation coming from the introduction of the reform; the second source of variation is cross-sectional and arises from the differences in municipalities' share of non Eu immigrants allowed to vote. In the spirit of a differences-in-differences strategy, election results of municipalities with more immigrants entitled to vote are compared with municipalities with a smaller share of immigrant voters (intensity of the treatment), before and after the reform.

The introduction of municipality and time-period fixed effects controls for all time-invariant differences across municipalities and secular changes over time. The strategy relies on the absence of any other shocks occurred around the same time the reform was introduced and correlated with the share of potential immigrant voters. The latter identification concern is addressed by controlling for time and municipality-varying factors that may bias the estimates, such as population density, population size (in logarithmic form) and the share of foreign population. In addition, controls for turnout and a dummy for winning incumbent are added.

The paper focuses on the identification of the parameter δ in equation 3.1 below:

$$y_{mt} = \alpha + \eta_m + \gamma_t + \delta T_m * post_t + \beta X_{mt} + \epsilon_{mt} \tag{3.1}$$

where y_{mt} is an electoral (or economic) outcomes, η_m is a municipality fixed effect, γ_t an election (or year, depending on having a political or economic outcome) fixed effect, T_m is the fraction of non Eu non citizens allowed to vote, as it is measured in 2006 (the variable that captures the treatment intensity) and X_{mt} includes a set of time-varying municipal characteristics, as listed above.

The availability of two or more pre- and post-treatment period allows to estimate a flexible model that includes leads and lags of the treatment. This allows to assess the presence of anticipatory effects or other violations of the common trend assumption. Formally, the model in equation 3.2 below is estimated:

$$y_{mt} = \alpha + \eta_m + \gamma_t + \sum_{j=t+1}^{T} \delta_j T_m * I_t^j + \beta X_{mt} + \epsilon_{mt} \tag{3.2}$$

where everything is defined as above, with the exception that the effect of the treatment is identified in each election going from t+1 to T (with t being the reference category).

Equation 3.2 imposes no parametric assumptions on the pre-treatment dynamics and allows for a the test of the null hypothesis of no common pre-treatment trends ($H_0 : \delta_j = 0$ for all pre-treatment periods). Moreover, it also allows the implementation of tests on the dynamics of the treatment effect, i.e, it is possible to test whether the effect is constant in the post-treatment period.

3.4 Results

This section presents the results of the effect of the reform on political and economic outcomes, by region. For what concerns political outcomes, first, all ranked parties are considered, then the analysis will focus on outcomes for which the reform was effective.

3.4.1 Flanders

Table 3.2 show results on the share of votes taken by each family of parties in Flanders. The reform significantly affects only the pro-immigrant party in the region, which is losing power in the post reform period. The magnitude of the coefficient indicates that an increase of 1 p.p. (percentage point) in the treatment determines a decrease of more than 3 p.p. in the share of votes to the pro-immigrant party. There are no significant results on the outcomes of the other parties.

From now on, the focus will be on the share of votes to the pro-immigrant party.

In order to give a causal interpretation to the effect of the reform, municipalities with different intensity of the treatment must have similar pre reform trend in the outcome variable. This hypothesis can be tested by estimating a fully flexible model, as expressed in equation 3.2. Results of the flexible estimates for Flanders are reported in Table 3.3. In all columns, the coefficients associated to the pre-treatment periods are small and non significantly different from zero; while there is a negative and increasing (in absolute value) effect in the post treatment election rounds. As shown in column (6), this result is also robust to the inclusion of district specific time trend. Figure 3.2 plots the coefficients of the interactions between the intensity of the treatment and each election's dummies, as reported in column (6). There is a zero and flat effect before the 2004 reform and a negative and significant effect starting from 2006 election.

In Table 3.4 equation 3.1 is estimated by using a restricted sample, accounting only for 1988 to 2000 elections. In this subsample a placebo effect of the reform is estimated, by assuming that the reform took place in 1994 or 2000. Given that the reform was effective only for the 2006 election onwards, finding significant effects in 1994 or 2000 elections would suggest violations of the identifying assumptions, since they could not be attributed to the real reform. Conversely, finding no effects on this subsample suggest that the baseline estimates can be interpreted as causal. In column (1) of Table 3.4 the sample is restricted to three rounds of election, and the treatment is imposed in 1994 and 2000. In column (2) only elections in 1988 and 1994 are used and 1994 is used as a fake post treatment period. In column (3) only elections in 1994 and 2000 are considered and 2000 is the post reform period. The coefficients of the interaction between the post dummy and the treatment intensity are always much smaller than in Table 3.2 and they are never significantly different from zero. The placebo experiments suggest no evidence of a differential relationship between share of votes to the pro-immigration party and the treatment intensity in the pre reform elections.

If voting behavior reflects voters' demand for public goods, changes in municipalities' expenditures and revenues should be expected as a consequence of the changes in parties' share of votes. In the case of Flanders, given that pro immigrants parties also favor more redistribution, less public expenditures and revenues are expected as a result of the reducing power of the pro-immigration parties. Results in Table 3.5 confirm this expectations and show that municipalities with a higher treatment experience a higher decrease in total and welfare per capita expenditures and per capita revenues in the post reform period.

3.4.2 Brussels

Table 3.6 reports results of the estimation of equation 3.1 for the 19 municipalities of the Brussels region. The enfranchisement of non citizens of non Eu origin has a negative effect on the share of votes of the pro immigration parties, that loses about 1 p.p. of votes as the treatment intensity increases by 1 p.p.

However, a deeper analysis of this effect, as shown by the fully flexible estimates in Table 3.7 points to no evidence of a casual effect of the reform. Specifically, even if the post reform coefficients are negative and bigger (in absolute value) than the pre reform coefficients, they are not statistically significant.

3.4.3 Wallonia

When looking at the results for the region of Wallonia, an opposite effect of the reform with respect to Flanders emerges: the very pro-immigrants's parties gain votes in municipalities where the share of potential non Eu voters is higher. More precisely, an increase of 1 p.p. in the treatment determines an increase of almost 3 p.p in the share of votes to the very pro-immigrant parties.

In order to interpret causally the previous result, there should be no violation of the common trend assumption. The fully flexible estimates, reported in Table 3.9 point to similar pre reform trends in the outcome variable: the interacted coefficients are never significantly different from zero in the pre-reform elections. The main effect for Wallonia comes from the second election after the reform: in 2012 an increases of 1 p.p. in the share of immigrant potential voters determines an increase in the share of votes to the very pro-immigrants parties of more than 5 p.p.. Figure 3.3 reports the interacted coefficients from column (6) in table 3.9, for a better visual inspection of the effect.

The placebo experiments for Wallonia are reported in Table 3.10. As for Flanders, only a subsample of elections is considered and the treatment is imposed in the pre reform period. The coefficients of the interaction between the dummy post and the treatment intensity are never significantly different from zero, pointing to a non differential relationship between share of votes to the very pro-immigrants party and the treatment intensity in the pre reform elections.

3.5 Conclusions

This paper studies the effect of granting non citizens of non Eu origin with the right to vote in local elections on political and economic outcomes. The introduction of a reform in Belgium in 2004 is exploited as an exogenous variation, allowing for a causal interpretation of the results.

Results point to an opposite effect of the reform in the region of Flanders and Wallonia: while in the former the main effect is a decrease in the votes for the pro-immigrants parties and, consequently, a reduction in welfare expenditures and total revenues; in the latter the reform determines an increase of votes to the left and pro immigrants parties [1]. The magnitude of the effects, if compared with the small group of new voters, suggests that the law generates a reaction in natives' voting behavior.

As discussed in section 3.2.1, Flanders and Wallonia pursue very different immigration policies and have an opposite attitude toward immigration. Moreover, the two regions are characterized by a very different economic and demographic context (Dandoy, 2014): compared to Wallonia, Flanders are a more prosperous economy, with no need of immigrant workers. These differences might explain the results: the rich Flanders have stronger (negative) ethnic preferences and care less about public good, then natives vote less for the pro immigrants and pro public good parties.

Most of the existing literature studying the political effect of migration in Europe focuses on the impact of the mere presence of immigrants on natives' local behavior. The main contribution of this paper is to show that a big natives' changes in voting behavior comes from the enfranchisement of an ethnically different group. The opposite results on the two regions suggest that enfranchisement *per se* has not necessarily positive effects for immigrants: if they are a small and non-integrated group, as in the case of Flanders, the reaction of natives, who support less the pro immigrants parties, result in a decrease in the demand public goods, that mostly hurts the low income population, including immigrants.

[1] Results in the region of Brussels do not allow for a causal interpretation of the reform effect.

3.6 Figures and Tables

Figure 3.1: Share of local parties labelled according to their national counterparts

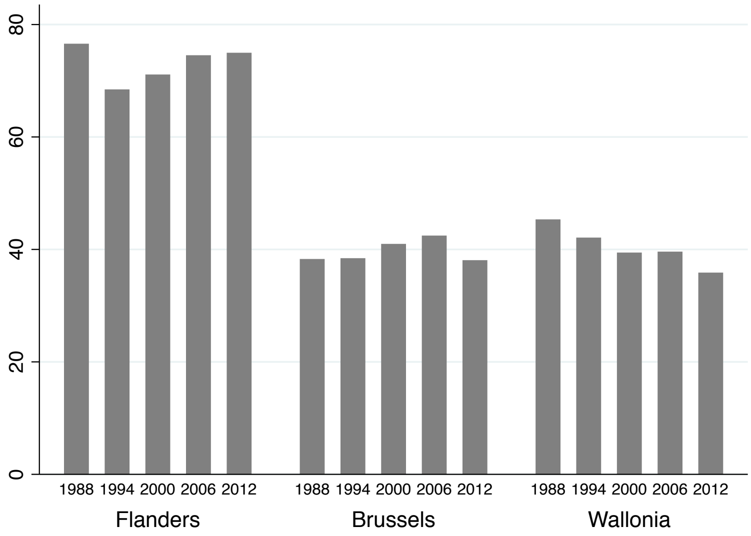


Figure 3.2: Interacted coefficients from fully flexible estimates: Flanders

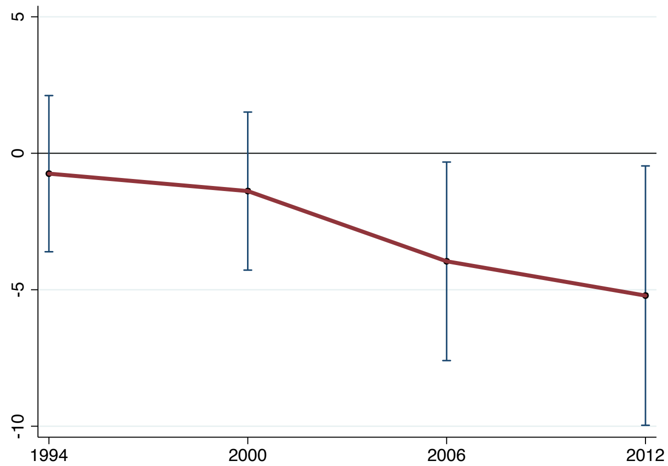


Figure 3.3: Interacted coefficients from fully flexible estimates: Wallonia

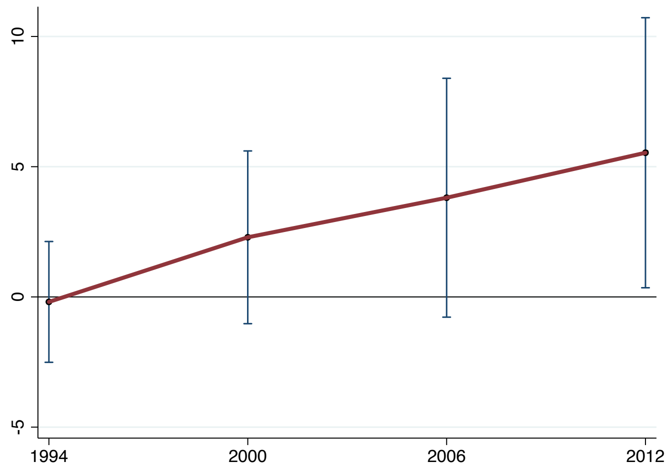


Table 3.1: Potential non citizens non Eu voters as a percentage of all other voters

	2006	2012	Difference
Belgium	0.6%	0.79%	32%
Flanders	0.41%	0.62%	51%
Wallonia	0.46%	0.59%	28%
Brussels	5.42%	6.06%	11%

Table 3.2: Effect of immigrants' enfranchisement: Flanders

	(1) Very pro-imm	(2) Pro-imm	(3) Anti-imm	(4) Very anti-imm
Treatment×Post	0.264	-3.225**	0.999	-0.246
	(1.372)	(1.364)	(1.349)	(1.074)
Share of Foreigners	0.096	-0.143	-0.437	0.782**
	(0.216)	(0.424)	(0.280)	(0.327)
Log of population	6.586	2.251	44.309***	-3.929
	(11.754)	(17.266)	(16.270)	(13.599)
Population Density	-0.000	0.001	-0.003**	0.001
	(0.002)	(0.001)	(0.001)	(0.001)
Turnout	0.074	0.645**	0.343	0.089
	(0.163)	(0.291)	(0.298)	(0.219)
Winning Incumbent	0.997*	1.988**	0.605	-0.835
	(0.530)	(0.822)	(0.730)	(0.625)
Observations	1,536	1,536	1,536	1,536
R-squared	0.815	0.809	0.743	0.740
Municipalities' indicators	YES	YES	YES	YES
Years of elections' indicators	YES	YES	YES	YES
District Time Trend	YES	YES	YES	YES

Standard errors clustered at municipality level in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 3.3: Fully Flexible Estimates Results: Flanders

	(1) Pro-imm	(2) Pro-imm	(3) Pro-imm	(4) Pro-imm	(5) Pro-imm	(6) Pro-imm
Share of foreigners		-0.327	-0.341	-0.298	-0.296	-0.141
		(0.336)	(0.332)	(0.336)	(0.328)	(0.434)
Log of population			9.660		10.785	-1.861
			(15.654)		(15.674)	(18.574)
Population density			-0.001	-0.000	-0.001	0.001
			(0.001)	(0.001)	(0.001)	(0.001)
Turnout				0.720**	0.688**	0.576*
				(0.290)	(0.295)	(0.309)
Winning Incumbent					2.361***	1.965**
					(0.778)	(0.818)
Treatment×1994	-0.718	-0.646	-0.530	-0.391	0.004	-0.747
	(1.347)	(1.343)	(1.368)	(1.354)	(1.344)	(1.460)
Treatment×2000	-0.779	-0.813	-0.632	-0.506	-0.117	-1.385
	(1.111)	(1.104)	(1.152)	(1.117)	(1.174)	(1.477)
Treatment×2006	-1.955*	-2.229*	-1.932	-2.569**	-2.032*	-3.958**
	(1.113)	(1.168)	(1.199)	(1.176)	(1.225)	(1.855)
Treatment×2012	-3.188***	-3.429***	-2.923**	-2.797**	-2.107	-5.215**
	(1.204)	(1.243)	(1.368)	(1.290)	(1.384)	(2.424)
Observations	1,536	1,536	1,536	1,536	1,536	1,536
R-squared	0.773	0.773	0.773	0.774	0.777	0.809
Municipalities' indicators	YES	YES	YES	YES	YES	YES
Years of elections' indicators	YES	YES	YES	YES	YES	YES
District Time Trend	NO	NO	NO	NO	NO	YES

Standard errors clustered at municipality level in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 3.4: Placebo test: Flanders

	1988-2000 Post=1994,2000	1988-1994 Post=1994	1994-2000 Post=2000
	(1)	(2)	(3)
	Pro-imm	Pro-imm	Pro-imm
Treatment×Post	-0.428	-0.780	-0.314
	(1.799)	(3.440)	(2.663)
Observations	921	614	614
R-squared	0.878	0.930	0.936
Controls	YES	YES	YES
Municipalities' indicators	YES	YES	YES
Years of elections' indicators	YES	YES	YES
District Time Trend	YES	YES	YES

Standard errors clustered at municipality level in parentheses
*** p<0.01, ** p<0.05, * p<0.1

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Table 3.5: Expenditures and revenues: Flanders

	(1)	(2)	(3)	(4)	(5)	(6)
	Per capita Log-Expenditures		Per capita Log-Welfare Expenditures		Per capita Log-Revenues	
Treatment×Post	-0.032**		-0.419**		-0.030**	
	(0.013)		(0.169)		(0.015)	
Treatment×I round after the law		-0.031**		-0.421**		-0.028*
		(0.013)		(0.168)		(0.015)
Treatment×II round after the law		-0.011		-0.443		0.001
		(0.022)		(0.272)		(0.028)
Observations	3,076	3,076	2,748	2,748	3,076	3,076
R-squared	0.773	0.773	0.382	0.382	0.674	0.675
Municipalities' indicators	YES	YES	YES	YES	YES	YES
Years of elections' indicators	YES	YES	YES	YES	YES	YES
District Time Trend	YES	YES	YES	YES	YES	YES
Share of foreigners×year	YES	YES	YES	YES	YES	YES
District Time Trend	YES	YES	YES	YES	YES	YES

Standard errors clustered at municipality level in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 3.6: Effect of immigrants' enfranchisement: Brussels

	(1) Very pro-imm	(2) Pro-imm	(3) Anti-imm	(4) Very anti-imm
Treatment×Post	-0.825	-1.168**	2.049	1.448
	(0.704)	(0.562)	(1.416)	(1.277)
Share of Foreigners	0.217	-0.739***	0.809	0.475
	(0.342)	(0.201)	(0.607)	(0.480)
Log of population	-48.931	26.035	-7.084	-46.905
	(61.682)	(29.345)	(85.156)	(49.550)
Population Density	0.001	0.000	-0.000	-0.000
	(0.000)	(0.000)	(0.000)	(0.000)
Turnout	2.137	0.450	-0.660	0.150
	(1.307)	(0.434)	(1.364)	(0.656)
Winning Incumbent	0.777	-2.066	1.351	4.099*
	(2.610)	(1.418)	(4.473)	(2.178)
Observations	95	95	95	95
R-squared	0.684	0.696	0.516	0.769
Municipalities' indicators	YES	YES	YES	YES
Years of elections' indicators	YES	YES	YES	YES
District Time Trend	YES	YES	YES	YES

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 3.7: Fully Flexible Estimates Results: Brussels

	(1) Pro-imm	(2) Pro-imm	(3) Pro-imm	(4) Pro-imm	(5) Pro-imm	(6) Pro-imm
Share of foreigners		-0.311	-0.338	-0.380	-0.436*	-0.656**
		(0.241)	(0.246)	(0.254)	(0.259)	(0.300)
Log of population			-11.342		-1.988	24.680
			(22.477)		(23.945)	(30.567)
Population density			-0.000	-0.000	-0.000	-0.000
			(0.000)	(0.000)	(0.000)	(0.000)
Turnout				0.297	0.431	0.373
				(0.418)	(0.502)	(0.521)
Winning incumbent					-1.618	-2.148
					(1.449)	(1.428)
Treatment×1994	0.575	0.492	0.510	0.471	0.451	0.331
	(0.458)	(0.446)	(0.427)	(0.436)	(0.430)	(0.376)
Treatment×2000	1.117***	0.780	0.852*	0.747	0.726	0.325
	(0.395)	(0.476)	(0.488)	(0.507)	(0.502)	(0.500)
Treatment×2006	1.051**	0.352	0.543	0.299	0.133	-0.736
	(0.502)	(0.762)	(0.765)	(0.850)	(0.858)	(0.968)
Treatment×2012	1.645***	0.932	1.395*	1.187	1.074	-0.137
	(0.431)	(0.713)	(0.808)	(0.925)	(0.937)	(1.087)
Observations	95	95	95	95	95	95
R-squared	0.573	0.579	0.588	0.590	0.598	0.705
Municipalities' indicators	YES	YES	YES	YES	YES	YES
Years of elections' indicators	YES	YES	YES	YES	YES	YES
District Time Trend	NO	NO	NO	NO	NO	YES

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 3.8: Effect of immigrants' enfranchisement: Wallonia

	(1) Very pro-imm	(2) Pro-imm	(3) Anti-imm	(4) Very anti-imm
Treatment×Post	2.751**	-1.537	0.604	-0.463
	(1.328)	(1.206)	(1.420)	(0.498)
Share of Foreigners	0.026	0.127	0.263	0.090
	(0.516)	(0.420)	(0.424)	(0.157)
Log of population	-2.208	-1.530	-10.682	-3.317
	(17.279)	(8.936)	(12.162)	(4.805)
Population Density	0.001	0.002	0.002	0.000
	(0.003)	(0.002)	(0.003)	(0.001)
Turnout	-0.155	-0.036	0.157	0.023
	(0.186)	(0.119)	(0.183)	(0.056)
Winning Incumbent	-0.874	2.679**	-0.645	-0.117
	(1.163)	(1.104)	(1.057)	(0.439)
Observations	1,262	1,262	1,262	1,262
R-squared	0.882	0.768	0.788	0.387
Municipalities' indicators	YES	YES	YES	YES
Years of elections' indicators	YES	YES	YES	YES
District Time Trend	YES	YES	YES	YES

Standard errors clustered at municipality level in parentheses

*** p<0.01, ** p<0.05, * p<0.1

Table 3.9: Fully Flexible Estimates Results: Wallonia

	(1) Very pro-imm	(2) Very pro-imm	(3) Very pro-imm	(4) Very pro-imm	(5) Very pro-imm	(6) Very pro-imm
Share of foreigners		0.108	0.274	0.213	0.261	0.203
		(0.380)	(0.423)	(0.403)	(0.421)	(0.583)
Log of population			-6.873		-6.813	-0.024
			(10.920)		(10.930)	(17.423)
Population density			-0.002	-0.003	-0.002	0.001
			(0.003)	(0.003)	(0.003)	(0.003)
Turnout				0.042	0.052	-0.132
				(0.173)	(0.177)	(0.183)
Winning incumbent					-0.840	-0.764
					(1.120)	(1.171)
Treatment×1994	-0.446	-0.348	-0.390	-0.336	-0.393	-0.191
	(0.947)	(1.023)	(1.071)	(1.081)	(1.067)	(1.183)
Treatment×2000	1.718	1.913	1.610	1.682	1.549	2.289
	(1.057)	(1.288)	(1.391)	(1.421)	(1.380)	(1.692)
Treatment×2006	2.841**	3.177*	3.061	3.157	3.007	3.808
	(1.416)	(1.769)	(1.905)	(1.923)	(1.882)	(2.339)
Treatment×2012	4.742***	5.092***	5.182***	5.464***	5.130***	5.536**
	(1.357)	(1.760)	(1.870)	(1.889)	(1.854)	(2.645)
Observations	1,262	1,262	1,262	1,262	1,262	1,262
R-squared	0.851	0.851	0.851	0.851	0.852	0.882
Municipalities' indicators	YES	YES	YES	YES	YES	YES
Years of elections' indicators	YES	YES	YES	YES	YES	YES
District Time Trend	NO	NO	NO	NO	NO	YES

Standard errors clustered at municipality level in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Table 3.10: Placebo experiments: Wallonia

	1988-2000 Post=1994,2000	1988-1994 Post=1994	1994-2000 Post=2000
	(1) Very pro-imm	(2) Very pro-imm	(3) Very pro-imm
Treatment×Post	-0.080	0.598	2.759
	(1.572)	(2.532)	(2.854)
Observations	756	504	504
R-squared	0.937	0.973	0.969
Controls	YES	YES	YES
Municipalities' indicators	YES	YES	YES
Years of elections' indicators	YES	YES	YES
District Time Trend	YES	YES	YES

Standard errors clustered at municipality level in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Bibliography

Adam, I. and D. Jacobs (2014). Divided on immigration, two models for integration. the multilevel governance of immigration and integration in belgium. In *The Politics of Immigration in Multi-Level States*, pp. 65–85. Springer.

Agüero, J. M. and M. S. Marks (2008). Motherhood and female labor force participation: Evidence from infertility shocks. *American Economic Review 98*(2), 500–504.

Agüero, J. M. and M. S. Marks (2011). Motherhood and female labor supply in the developing world: Evidence from infertility shocks. *Journal of Human Resources 46*(4), 800–826.

Amuedo-Dorantes, C. and S. Pozo (2011). Remittances and income smoothing. *The American Economic Review, Papers and Proceedings 102*(3), 582–587.

Angrist, J., V. Lavy, and A. Schlosser (2010). Multiple experiments for the causal link between the quantity and quality of children. *Journal of Labor Economics 28*(4), 773–824.

Antman, F. (2012a). Elderly care and intrafamily resource allocation when children migrate. *Journal of Human Resources 47*(2), 331–63.

Antman, F. M. (2011a). The intergenerational effects of paternal migration on schooling and work: What can we learn from children's time allocations? *Journal of Development Economics 96*(2), 200–208.

Antman, F. M. (2011b). International migration and gender discrimination among children left behind. *The American economic review 101*(3), 645.

Antman, F. M. (2012b). Gender, educational attainment, and the impact of parental migration on children left behind. *Journal of Population Economics*, 1–28.

Ashenfelter, O. (2012). Comparing real wage rates: Presidential address. *American Economic Review 102*(2), 617–42.

Bagger, J., J. A. Birchenall, H. Mansour, and S. Urzúa (2013). Education, birth order, and family size. Technical report, National Bureau of Economic Research.

Bansak, C. and B. Chezum (2009). How do remittances affect human capital formation of school-age boys and girls? *The American economic review*, 145–148.

Barone, G., G. De Blasio, and P. Naticchioni (2014). Mr. rossi, mr. hu and politics: The role of immigration in shaping natives' political preferences.

Becker, G. S. (1960). An economic analysis of fertility. In G. S. Becker (Ed.), *Demographic and Economic Change in Developed Countries,*. Princeton, NJ: Princeton University Press.

Becker, G. S. and H. Lewis (1973). On the interaction between the quantity and quality of children. *Journal of Political Economy 81*(2), S279–S288.

Becker, G. S. and N. Tomes (1976). Child endowments and the quantity and quality of children. *Journal of Political Economy 84*(4), S143–62.

Billari, F. C. and V. Galasso (2014, April). Fertility decisions and pension reforms. Evidence from natural experiments in Italy. IdEP Economic Papers 1403, USI Università della Svizzera italiana.

Black, S. E., P. J. Devereux, and K. G. Salvanes (2005). The more the merrier? The effect of family size and birth order on children's education. *The Quarterly Journal of Economics*, 669–700.

Böheim, R. and K. Mayr (2005). Immigration and public spending. *IZA Discussion Paper*.

Bongaarts, J. and R. Potter (1983). *Fertility, Biology, and Behavior: An Analysis of the Proximate Determinants.* New York: Academy Press.

Borjas, G. J. and L. F. Katz (2005). The evolution of the Mexican-born workforce in the United States. NBER Working Papers 11281, National Bureau of Economic Research, Inc.

Bratti, M. and L. Cavalli (2014). Delayed first birth and new mothers' labor market outcomes: Evidence from biological fertility shocks. *European Journal of Population 30*(1), 35–63.

Brezis, E. S. and R. D. S. Ferreira (2014). Endogenous fertility with a sibship size effect. Working Papers of BETA 2014-03.

Buck, G. M., L. E. Sever, R. E. Batt, and P. Mendola (1997). Life-style factors and female infertility. *Epidemiology 8*(4), 435–41.

Cabrera, G. (1994). Demographic dynamics and development: The role of population policy in mexico. *Population and Development Review 20*, 105–120.

Cascio, E. U. and E. Washington (2014). Valuing the vote: The redistribution of voting rights and state funds following the voting rights act of 1965*. *Quarterly Journal of Economics 129*(1).

Casterline, J. (1989). Collecting data on pregnancy loss: A review of evidence from the World Fertility Survey. *Studies in Family Planning 20*(2), 81–95.

Cerrutti, M. and D. Massey (2001a). On the auspices of female migration from Mexico to the United States. *Demography 38*(2), 187–200.

Cerrutti, M. and D. S. Massey (2001b). On the auspices of female migration from mexico to the united states. *Demography 38*(2), 187–200.

Chen, J. J. (2006). Migration and imperfect monitoring: Implications for intra-household allocation. *American Economic Review 96*(2), 227–231.

Chen, K.-P., S.-H. Chiang, and S. F. Leung (2003a). Migration, family, and risk diversification. *Journal of Labor Economics 21*(2), 353–380.

Chen, K.-P., S.-H. Chiang, and S. F. Leung (2003b). Migration, family, and risk diversification. *Journal of Labor Economics 21*(2), 323–352.

Chen, S. H., Y.-C. Chen, and J.-T. Liu (2009). The impact of unexpected maternal death on education: First evidence from three national administrative data links. *The American Economic Review*, 149–153.

Chiquiar, D. and G. H. Hanson (2005). International migration, self-selection, and the distribution of wages: Evidence from mexico and the united states. *Journal of Political Economy 113*(2).

Cigno, A. (1993). Intergenerational transfers without altruism: Family, market and state. *European Journal of Political Economy 9*(4), 505–518.

Clemens, M. (2011). Economics and emigration: Trillion-dollar bills on the sidewalk? *Journal of Economic Perspectives 25*(3), 83–106.

Clemens, M., C. E. Montenegro, and L. Pritchett (2010). The place premium: Wage differences for identical workers across the us border. Working papers, University of Chile, Department of Economics.

Clemens, M. A., Ç. Özden, and H. Rapoport (2014). Migration and development research is moving far beyond remittances. *World Development 64*, 121–124.

Cortes, P. (2015). The feminization of international migration and its effects on the children left behind: Evidence from the philippines. *World Development 65*, 62–78.

Dahlberg, M., K. Edmark, and H. Lundqvist (2012). Ethnic diversity and preferences for redistribution. *Journal of Political Economy 120*(1), 41–76.

Dandoy, R. (2014). Regionalist parties and immigration in belgium. In *The Politics of Immigration in Multi-Level States*, pp. 200–222. Springer.

Datcher-Loury, L. (1988). Effects of mother's home time on children's schooling. *The Review of Economics and Statistics*, 367–373.

Docquier, F., A. Marfouk, S. Salomone, and K. Sekkat (2012). Are skilled women more migratory than skilled men? *World development 40*(2), 251–265.

Duflo, E. (2003). Grandmothers and granddaughters: Old-age pensions and intrahousehold allocation in south africa. *The World Bank Economic Review 17*(1), 1–25.

Duflo, E. and A. Banerjee (2011). *Poor Economics: A Radical Rethinking of the Way to Fight Global Poverty*. PublicAffairs.

Dunn, J. and R. Plomin (1990). *Separate Lives: Why Siblings are so Different.* Basic Books, New York.

Dustmann, C. and A. Glitz (2011). Migration and education. In E. Hanushek, S. Machin, and L. Woessmann (Eds.), *Handbook of the Economics of Education*, Volume 4, Chapter Chapter 4, pp. 327–439. Amsterdam: Elsevier.

Edwards, A. C. and M. Ureta (2003). International migration, remittances, and schooling: evidence from el salvador. *Journal of development economics 72*(2), 429–461.

Fitzsimons, E. and B. Malde (2014, January). Empirically probing the quantity-quality model. *Journal of Population Economics 27*(1), 33–68.

Gábos, A., R. I. Gál, and G. Kézdi (2009). The effects of child-related benefits and pensions on fertility by birth order: A test on hungarian data. *Population studies 63*(3), 215–231.

Garg, A. and J. Morduch (1998). Sibling rivalry and the gender gap: Evidence from child health outcomes in Ghana. *Journal of Population Economics 11*(4), 471–493.

Gertler, P., S. Boyce, et al. (2001). An experiment in incentive-based welfare: The impact of progresa on health in mexico. *University of California, Berkeley*, 30–37.

Ghatak, S., P. L. and S. W. Price (1996). Migration theories and evidence: An assessment. *Journal of Economic Surveys 10*(2), 159–197.

Gibson, J. and D. McKenzie (2012). The economic consequences of 'brain drain'of the best and brightest: Microeconomic evidence from five countries*. *The Economic Journal 122*(560), 339–375.

Giles, J. and R. Mu (2007). Elderly parent health and the migration decisions of adult children: Evidence from rural China. *Demography 44*(2), 265–288.

Hanson, G. (2004). Illegal migration from Mexico to the United States. *Journal of Economic Literature Vol. 44*(4), 869–924.

Hanson, G. H. and C. McIntosh (2010). The Great Mexican migration. *The Review of Economics and Statistics 92*(4).

Hanson, G. H. and C. Woodruff (2003). Emigration and educational attainment in mexico. *Documento de Trabajo del IR/PS. Disponible en http://irpshome. ucsd. edu/faculty/gohanson/working_papers. htm*, 1–39.

Hanushek, E. A. (1992). The trade-off between child quantity and quality. *Journal of Political Economy 100*(1), 84–117.

Harmon, N. A. (2013). Immigration ethnic diversity and political outcomes.

Hatton, T. J. and J. G. Williamson (1998). *The Age of Mass Migration: Causes and Economic Impact*. New York: Oxford University Press.

Hatton, T. J. and J. G. Williamson (2003). Demographic and Economic Pressure on Emigration out of Africa. *Scandinavian Journal of Economics 105*(3), 465–486.

Haveman, R. and B. Wolfe (1995). A review of methods and findings. *Journal of economic literature 33*(4), 1829–1878.

Hotz, V. J., S. W. McElroy, and S. G. Sanders (2005). Teenage childbearing and its life cycle consequences: Exploiting a natural experiment. *Journal of Human Resources 40*(3), 683–715.

Husted, T. A. and L. W. Kenny (1997). The effect of the expansion of the voting franchise on the size of government. *Journal of Political Economy*, 54–82.

Jayachandran, S. and R. Pande (2015). Why are indian children so short? *NBER Working Papers 21036, National Bureau of Economic Research, Inc.*.

Jensen, R. and N. Miller (2011). Keepin' 'em down on the farm: Old age security and strategic underinvestment in children. *mimeo UCLA* (4).

Joffe, M. and I. Barnes (2000). Do parental factors affect male and female fertility? *Epidemiology 11*(6), 700–705.

Jolly, S., H. Reeves, and N. Piper (2005). *Gender and migration: Overview report*. Bridge.

Karimi, A. (2014, August). Effects of the timing of births on women's earnings – Evidence from a natural experiment. Working Paper Series 2014:17, IFAU – Institute for Evaluation of Labour Market and Education Policy.

Kennan, J. and J. R. Walker (2011a). The effect of expected income on individual migration decisions. *Econometrica 79*(1), 211–251.

Kennan, J. and J. R. Walker (2011b). The Effect of Expected Income on Individual Migration Decisions. *Econometrica 79*(1), 211–251.

Kristin, B. and C. Anne (1994). The effect of sibling sex composition on women's education and earnings. *The Quarterly Journal of Economics 109*(3).

Lewis, J. B. and D. A. Linzer (2005). Estimating regression models in which the dependent variable is based on estimates. *Political Analysis 13*(4), 345–364.

Lott, J. R. (1999). How dramatically did women's suffrage change the size and scope of government? *Journal of Political Economy 107*(6 Part 1), 1163–1198.

Lucas, R. E. B. and O. Stark (1985). Motivations to remit: Evidence from Botswana. *Journal of Political Economy 93*(5), 901–18.

Lyle, D. S. (2006). Using military deployments and job assignments to estimate the effect of parental absences and household relocations on children's academic achievement. *Journal of Labor Economics 24*(2), 319–350.

Markussen, S. and M. Strøm (2015). The effects of motherhood. Technical report, MEMORANDUM No 19/2015, Department of Economics, University of Oslo.

Martin, P. L. (1993). *Trade and Migration: NAFTA and Agriculture*. Washington: Institute for International Economics.

Massey, D. S. and L. C. Basem (1992). Determinants of savings, remittances, and spending patterns among us migrants in four mexican communities. *Sociological Inquiry 62*(2), 185–207.

Mayda, A. M., G. Peri, and W. Steingress (2015). Immigration to the us: A problem for the republicans or the democrats?

Mckenzie, D. and H. Rapoport (2007). Network effects and the dynamics of migration and inequality: Theory and evidence from Mexico. *Journal of Development Economics 84*(1), 1–24.

McKenzie, D. and H. Rapoport (2011). Can migration reduce educational attainment? evidence from mexico. *Journal of Population Economics 24*(4), 1331–1358.

Miller, A. (2011). The effect of motherhood timing on career path. *Journal of Population Economics 24*(3), 1071–1100.

Miller, G. and K. S. Babiarz (2014). Family planning program effects. *NBER Working Paper No. 20586*.

Morduch, J. (2000). Sibling rivarly in Africa. *American Economic Review 90 (2)*, 405–409.

Morrison, A. R., M. W. Schiff, and M. Sjöblom (2007). *The international migration of women*. World Bank Publications.

Mourard, E. (2015). The impact of migration on the family left behind: Estimation in presence of intra-household selection of the migrants. *Paris School of Economics. Mimeo.*.

Nugent, J. (1985). The old-age security motive for fertility. *Population and Development Review 11*(1), 75–97.

Oishi, N. (2002). Gender and migration: an integrative approach. *Center for Comparative Immigration Studies*.

Orozco, M. (2006). Gender remittances: Preliminary notes about senders and recipients in latin america and the carribean. *United Nations, Nations Unies*.

Ozden, C. and M. Schiff (2006). *International Migration, Remittances, and the Brain Drain*. Washington, DC: The World Bank and Palgrave McMillan.

Pande, R. (2003). Selective gender differences in childhood nutrition and immunization in rural India: The role of siblings. *Demography 40*(3), 395–418.

Parish, W. L. and R. J. Willis (1993). Daughters, education, and family budgets Taiwan experiences. *Journal of Human Resources 28*(4), 863–898.

Pedraza, S. (1991). Women and migration: The social consequences of gender. *Annual review of sociology*, 303–325.

Platteau, J. (1991). *Traditional systems of social security and hunger insurance: Past achievements and modern challenges*. Clarendon Press, Oxford.

Plouffe, L. J., T. S. White E. W., C. Sweet, L. Layman, G. Whitman, and P. McDonough (1992). Etiologic factors of recurrent abortion and subsequent reproductive performance of couples: Have we made any progress in the past 10 years? *American Journal of Obstetrics and Gynecology 167*(2), 313–321.

Qian, N. (2009). Quantity-Quality and the One Child Policy:The Only-Child Disadvantage in School Enrollment in Rural China. NBER Working Papers 14973, National Bureau of Economic Research, Inc.

Rapoport, H. and F. Docquier (2006, 00). *The Economics of Migrants' Remittances*, Volume 1 of *Handbook on the Economics of Giving, Reciprocity and Altruism*, Chapter 17, pp. 1135–1198. Elsevier.

Ratha, D., S. Mohapatra, and E. Scheja (2011). Impact of migration on economic and social development: A review of evidence and emerging issues. *Policy Research Working Paper Series 5558, The World Bank 61*(3), 380–89.

Ray, D. (1998). *Development Economics*. New York: Princeton, NJ: Princeton University Press.

Razin, A., E. Sadka, and P. Swagel (2002). Tax burden and migration: a political economy theory and evidence. *Journal of Public Economics 85*(2), 167–190.

Rosenzweig, M. (2007). Education and migration: A global perspective. *Yale University. Mimeo*.

Rosenzweig, M. R. (1988). Risk, implicit contracts and the family in rural areas of low-income countries. *Economic Journal 98*(393), 1148–70.

Rosenzweig, M. R. and O. Stark (1989). Consumption smoothing, migration, and marriage: Evidence from rural India. *Journal of Political Economy 97*(4), 905–26.

Rosenzweig, M. R. and K. I. Wolpin (1980). Testing the quantity-quality fertility model: The use of twins as a natural experiment. *Econometrica 48*(1), 227–240.

Schultz, P. (2000). The impact of progresa on school enrollments: Final report. *Washington: International Food Policy Research Institute*.

Schultz, T. (2008). Population policies, fertility, women's human capital, and child quality. Volume 4, Chapter 52, pp. 3249–3303. Amsterdam: Elsevier.

Schultz, T. P. (1990). Testing the neoclassical model of family labor supply and fertility. *Journal of Human Resources 25*(4), 599–634.

Schultz, T. W. (1972). Human capital: Policy issues and research opportunities. In *Economic Research: Retrospect and Prospect Vol 6: Human Resources*, pp. 1–84. UMI.

Sjaastad, L. A. (1962). The costs and returns of human migration. *The journal of political economy*, 80–93.

Stark, O. (1981). The asset demand for children during agricultural moderninzation. *Population and Development Review 4*(3), 671–675.

Stark, O. (1991). *The migration of labour.* Basil Blackwell, Cambridge.

Stark, O. and D. E. Bloom (1985). The New Economics of Labor Migration. *American Economic Review 75*(2), 173–78.

Stöhr, T. (2015). Siblings' interaction in migration decisions: who provides for the elderly left behind? *Journal of Population Economics 28*(3), 593–629.

Suits, D. B. (1984). Dummy variables: Mechanics v. interpretation. *The Review of Economics and Statistics 66*(1), 177–80.

Thomas, D. (1990). Intra-household resource allocation an inferential approach. *Journal of human resources 25*(4).

Thomas, D. (1994). Like father, like son; like mother, like daughter: Parental resources and child height. *Journal of Human Resources*, 950–988.

UN (2013). World youth report. Technical report, United Nations.

UNDESA (2011). The age and sex of migrants. Technical report, United Nation, Department of Economics and Social Affairs, Population Division.

Vogl, T. S. (2013). Marriage institutions and sibling competition: Evidence from south Asia. *The Quarterly Journal of Economics 128*(3), 1017–1072.

Wilcox, L. S. and W. D. Mosher (1993). Use of infertility services in the United States. *Obstetrics & Gynecology 82*(1), 122–127.

Williamson, J. (1990). *Coping with City Growth during the British Industrial Revolution.* Cambridge, MA: Harvard University Press.

Yang, D. (2008). International migration, remittances and household investment: Evidence from Philippine migrants' exchange rate shocks. *Economic Journal 118*(528), 591–630.

Zuniga Herrera, E. (2008). Population policy in Mexico and its relationship to development. presented at Contribution of Population Policy to the Achievement of the Internationally Development Goals including the MDGs. Panel discussion, New York, 15 December 2008.