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Essays on Economic Development in Costa Rica

by

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Abstract

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Costa Rica had two major demographic changes in the last 20 years. First, Costa Ricans have fewer children. Second, many Nicaraguans have migrated into Costa Rica. Estimating the effects of migration and fertility changes is difficult. Both migration and fertility decisions are endogenous to labor market and social conditions. To address this concern, in this thesis I use two natural experiments. First, I use a regression discontinuity approach that takes advantage of the exact timing of a 2001 change in child support legislation. I find that birthrates fell significantly after the change. Single mothers now find it easier to receive child support payments. This affects marital status decisions. Indeed, births within marriage also fell, suggesting a decline in shotgun marriage. Maximilian Kasy and I analyze the impact of changing family structures on the female income distribution in Costa Rica between 1993 and 2009, using decomposition methods based on reweighting, as in [DiNardo *et al.*, 1996], and influence function regression, as in [Firpo *et al.*, 2009]. We find that family structures changes before the law change had a general negative effect on incomes. After the law passed, the trend reverses for higher income women, while lower incomes continue to be negatively affected by changing family structures. The second natural experiment is Hurricane Mitch in 1998. The damage it caused to the Nicaraguan countryside generated a surge in Nicaraguan immigrants. This surge located disproportionately in the regions and economic sectors that already had large Nicaraguan participation. I use difference-in-differences to find that Nicaraguan immigration after Mitch had no impact on native labor market outcomes, except for a small negative income effect on low-education men in the regions that share a border with Nicaragua.

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To my parents and to Cristie, with all my love

Contents

Chapter 1

Introduction

In 1986, a Costa Rican mother was going to have between three and four children over her lifetime. Today, she is going to have two. In 1986, virtually all newborns had Costa Rican mothers. Today, nearly a fifth have a foreign mother. Those demographic changes reflect economic opportunity in Costa Rica. Opportunities for women who decide to enter the labor force and delay childbearing. Opportunities for foreigners, a majority Nicaraguan, who seek jobs in Costa Rica and start families there.

Migration affects labor market outcomes. Lower fertility and higher female labor participation affects the female income distribution. *Essays on Economic Development in Costa Rica* is an attempt to estimate the impact of those changes. This is a challenge because fertility and migration decisions are endogenous labor market and income outcomes. The literature suggests addressing these difficulties with natural experiments that provide an observable exogenous shock (e.g. [Card, 1990] on migration). I found two natural experiments in Costa Rica: the Ley de Paternidad Responsable (Responsible Paternity Law) for fertility outcomes and Hurricane Mitch for migration. The rich literature on natural experiments tends to use developed country examples. Adding a developing country example such as Costa Rica is one of the contributions of this thesis. I am able to do this thanks to the availability of good Costa Rican databases of births and household surveys.

In Chapter 2 I estimate the effects of the Ley de Paternidad Responsable (Responsible Paternity Law) on fertility and marital status decisions. This Law made DNA tests compulsory for presumptive fathers who refused to acknowledge their children. Payment of child support is well enforced in Costa Rica, but only to a legally affiliated father. The relative costs of out-of-wedlock sex have changed after the law; in the event of pregnancy, men have to pay (expensive) child support. Such a reallocation of resources from fathers to mothers may affect childbearing and marital status decisions. The literature frequently finds weak or inconclusive results of the effect of family policies on fertility decisions. I use a regression discontinuity approach that takes advantage of the knowledge of the exact day of the Law passage and the availability of daily data on births in Costa Rica. The main finding is that the Responsible Paternity Law caused the birthrate and the total fertility rate to fall an additional 5% in the most conservative specifications and possibly as much as 10% on top of already rapidly-falling Costa Rican fertility. Also, this drop was not concentrated only among unmarried mothers, but it is also observed in married mothers, which suggests a decline in shotgun marriages.

In Chapter 3, written in collaboration with Maximilian Kasy, we seek to isolate the effect of changing family structures in Costa Rica, from 1993 to 2009, on the female income distribution. We assume that the family status of individuals is independent of the unobserved determinants of their income, conditional on a set of demographic covariates. We consider counterfactual changes in the income distribution that would have taken place as a consequence of changing family structures if the distribution of demographic covariates, as well as the distribution of equivalent income given family status and covariates, had remained constant during the period of interest. We find that the changes in family structure had an inequality-increasing effect for the entire working-age female population over the 1993/94-2008-09 period, but with quite different patterns of change before and after 2000/01. In the earlier period, the relative changes of incomes are roughly constant and negative across income levels; in the later period, higher incomes grew due to changing family structures, while lower incomes decreased further. In most subgroups we find similar effects, with the exception of urban women, where both high and low incomes were decreased more strongly than intermediate incomes by changing family structures.

In Chapter 4 I estimate the impact of a surge in Nicaraguan migration on native labor market outcomes. In 1998, Hurricane Mitch caused crop losses and forced many Nicaraguans to migrate to Costa Rica, which had been spared that kind of damage. The Hurricane Mitch immigrants chose to disproportionately locate in those Costa Rican regions near the border with Nicaragua and attempted to find work in agriculture and construction, encouraged by the presence of previous Nicaraguan immigrants in those regions and economic sectors. I use differences-in-differences and find that immigration after Hurricane Mitch had no immediate impact on native Costa Rican incomes and employment, except for a negative impact on incomes among men with only primary education in the border regions, but even this finding is small relative to the variation in incomes observed before and after the hurricane.

Chapter 2

Responsible Fathers: The impact of child support legislation in Costa Rica

2.1 Introduction

In 2000, thirty-one percent of Costa Rican children were born without a registered father. In early 2001, Costa Rican lawmakers passed the Ley de Paternidad Responsable (Responsible Paternity Law). This Law made DNA tests compulsory for presumptive fathers who refused to acknowledge their children. Child support enforcement in Costa Rica is very credible but applicable only to a legally affiliated father. Hence, the primary concern of the Law was to ensure children their access to the economic resources of the father[1]. In this, it was a success: the number of children without a registered father at birth fell 73% in 2002[2].

The relative costs of out-of-wedlock sex have changed after the law; in the event of pregnancy men face a higher likelihood of having to pay child support. These payments are a large fraction of parents' incomes in Costa Rica. Such a reallocation of resources from fathers to mothers may affect childbearing and marital status decisions. The literature frequently finds weak or inconclusive results of the effect of family policies on fertility decisions[3]. Evaluating family policies is difficult because they may reinforce or counteract other policies or social trends. In this chapter, I use a regression discontinuity approach to estimate the impact of the Responsible Paternity Law on childbearing and marital status decisions. This approach takes advantage of the knowledge of the exact day of the Law passage and the availability of daily data on births in Costa Rica.

To study the relationship between policy and family formation, the seminal rational choice model of marriage and childbearing was developed by [Becker, 1974]. In it, the

[1]The President at the time, Miguel Ángel Rodríguez, has stated that his goal was to "ensure the intergenerational wealth transfer from fathers to their children".

[2]See Table 2.1 and Figure 2.2a

[3]For overviews, see [Gauthier, 2007], [Lundberg and Pollak, 2007] and [Neyer and Andersson, 2008].

Table 2.1: Registered Births in Costa Rica, 1998-2005

	Total births		Unknown father		Unmarried mothers	
	N	Change	% Births	Change	% Births	Change
1998	76,749	-1.6%	28%	0.4%	50%	1.3%
1999	78,885	2.8%	30%	11.4%	52%	6.6%
2000	78,271	-0.8%	31%	1.7%	53%	2.0%
2001	76,041	-2.8%	29%	-9.4%	54%	-1.6%
2002	70,683	-7.0%	8%	-73.4%	57%	-2.7%
2003	72,356	2.4%	8%	-5.4%	58%	4.7%
2004	71,408	-1.3%	8%	5.2%	60%	2.1%
2005	71,002	-0.6%	7%	-16.6%	61%	1.4%

Notes: Estimates from Database of Births, INEC. Unknown father means one that is not registered at birth. Unmarried mothers includes single, cohabiting, widowed or separated. Change is in number of births.

family behaves as if maximizing the preferences of one parent who has altruistic preferences. Altruistic preferences are those in which the parent defers to the consumption patterns of other family members, who are egoistic. Marriage is a mechanism for the couple to maximize their production through specialization, where one partner specializes in producing public goods (e.g. cooking, children) for the family and the other partner specializes in obtaining income for consumption. Marriage market behavior is explained as a rational search for a best match[4]. [Lundberg and Pollak, 1994] extend the model to account for the role of initial resource control within the marriage as a threat point that affects the expenditure pattern of the couple. This encouraged the redesign of family policies such as giving aid to women instead of men, because women were shown to transfer more of this aid to their children.

Empirical attempts to estimate the impact of family policy changes has turned out to be controversial. Consider abortion. [Ananat *et al.*, 2009] make a distinction between the effect on pregnancies and the effect on whether the pregnancy is carried to term. If more couples engage in unprotected sex because abortion is available, pregnancies would increase. This increase should be offset by abortions, but women might be unable or unwilling to follow through with an abortion after they become pregnant. If they are unable to get an abortion due to cost or distant clinics, then birth outcomes are expected to worsen. If they are unwilling to abort, the pregnancy is desired and the set of birth outcomes should be better than before abortion.

Empirically, there is little agreement on what happened after abortion in the United States became legal in 1973. Identifying the effects of abortion is difficult because the Pill

[4] While he didn't analyze out-of-wedlock births in depth, he suggested that the difficulty of obtaining a divorce in Latin America (in 1974) encouraged cohabiting as part of the search for a mate before marriage

became available at roughly the same time and unilateral divorce laws were becoming more common[5]. Nevertheless, [Akerlof *et al.*, 1996] found that out-of-wedlock births increased because women who want children lost the ability to guilt men into shotgun-marrying them. This conflicts with [Gruber *et al.*, 1999] who found a positive selection effect of abortion. Positive selection means that terminated pregnancies are those where the women believe the outcome will be adverse[6]. [Ananat *et al.*, 2009] offer a potential solution to this conflict: even if there are more out-of-wedlock mothers after abortion, as long those mothers strongly desire their children and are willing to invest in them to ensure good outcomes, there will be positive selection.

[Donohue and Levitt, 2001] found a negative relationship between abortion in the seventies and crime in the eighties and nineties. Crime should decrease both because the cohort is smaller (young people commit a disproportionate fraction of crimes) and because positive selection implies a smaller fraction of children born whose mothers believed they would have bad outcomes. This line of research has been criticized as flawed. For example, [DiNardo, 2007] shows that the abortion ratio measure was poorly constructed, [Foote and Goetz, 2008] found coding errors that weaken or reverse the results when corrected[7], and the recent summary by [Joyce, 2009] suggests that the results are simply not robust.

Another example of the difficulties in estimating the determinants of fertility behavior is the "missing women" literature started by [Sen, 1990]. Some countries have a skewed sex ratio, where the proportion of men to women is much higher than normal. Sen's hypothesis is that families prefer boys and neglect girls (who then die of malnutrition and poor health). In more extreme cases, families engage in selective abortion of baby girls[8]. [Oster, 2005] explored the possibility that Hepatitis B infections increased the likelihood of a son in pregnant women. This could have explained an important fraction of the missing women. However, later research by [Lin and Luoh, 2008] in Taiwan and [Oster *et al.*, 2010] in China failed to replicate the male birth-Hepatitis B correlation.

Both [Akerlof *et al.*, 1996] and [Levine, 2001] point out that a family policy change that made it harder for men to avoid their responsibility towards the child could have very different results from the ones described so far in the literature which tend to have ambiguous gender effects. This chapter seeks to add to this line of research by estimating the impact of the Responsible Paternity Law on fertility and marital outcomes.

The main finding is that the Responsible Paternity Law caused the birthrate and the total

[5]see [Goldin and Katz, 2000] on the Pill and [Alesina and Giuliano, 2006] on divorces

[6][Pop-Eleches, 2006] finds positive selection when he estimates the effect of outlawing abortion in Romania. Cohorts born after the ban have better average education and labor outcomes. However, highly educated Romanian women used abortions the most. After controlling for mother background, the ban is estimated to have had a negative effect on outcomes

[7]The rise of out-of-wedlock births may play a role, see [Lott and Whitley, 2007].

[8]Extreme gender bias occurs mostly in Asia, but [Dahl and Moretti, 2008] find evidence of gender bias in the United States. For example, divorce is more likely if the firstborn is a daughter

fertility rate to fall an additional 5% in the most conservative specifications and possibly as much as 10% in the context of already rapidly-falling Costa Rican fertility. Also, this drop was not concentrated only among unmarried mothers, but it is also observed in married mothers, which suggests that at least some marriages were shotgun marriages. The effect is stronger in urban areas and among younger mothers and first time mothers. Controlling for nationality yields robust results. There is some evidence of a latent preference for sons because there was a spike in the sex ratio towards men after the law.

The remainder of this chapter is organized as follows: a discussion of the Costa Rican family law environment and changes introduced by the Responsible Paternity Law in section 2.2. Then section 2.3 describes the construction of birthrate time-series from data in the database of registered births in Costa Rica during the period 1986-2009, a set of population estimates and the Costa Rican Household Surveys for 1993-2008. Section 2.4 specifies a model of sexual participation, based on [Akerlof *et al.*, 1996] and modified to account for changes in the level of child support. The regression discontinuity methodology used to estimate the impact of the Responsible Paternity Law can be found in 2.5. Results are discussed in section 2.6, but tables and figures are found in section 2.8. Section 2.7 concludes and discusses potential new avenues of research.

2.2 Institutional Background

Costa Rica is a middle-income country with a political and legal system that has long made it a priority to protect the well-being of children[9]. In 2000, concerned Costa Rican lawmakers and policymakers began drafting the Responsible Paternity Law in response to the rising number of children born without a registered father [Solano, 2000]. They were spurred into action by [Budowski and Rosero-Bixby, 2000] and other studies that had linked fatherless children to childhood poverty in Costa Rica.

2.2.1 The Responsible Paternity Law and previous legislation

The Responsible Paternity Bill was signed into Law by President Rodríguez on April 16th, 2001. It was then published in the official newspaper on April 27th, 2001[10]. However, the most important date for analyzing its impact is February 26th, 2001. That was when the law bill attracted major media discussion due to its near-unanimous passage in the first debate of the Costa Rican Congress. This level of political support virtually guaranteed that the

[9]Infant mortality, which is closely linked to fertility outcomes, has been low the last three decades: in 1988 it was 14.7 deaths per thousand newborns and declined to 8.9 in 200. See [INEC, 2008]. This is similar to the levels of developed countries.

[10]Laws in Costa Rica are enforceable only after publication in *La Gaceta*, the official newspaper. For the full text of the Responsible Paternity Law, refer to [Congress, 2001]

bill would become law[11].

After the Law, a written and signed statement of the mother suffices to request the DNA test. This is a big shift in the burden of proof of paternity towards the presumptive father. Before the Law, compulsory DNA tests were available only with a court order[12]. To obtain this court order, mothers had to find witnesses and other proof of their relationship with the presumptive father. [Camacho, 2006] explains that this costly process could take up to three years and was subject to the discretion of the judge, so most mothers gave up if they were liquidity constrained or if they lacked credible witnesses to the relationship.

The Law introduced four additional major changes. First, mothers now get retroactive child support for the costs of pregnancy. Second, fathers lose any custody rights if they refuse to acknowledge the child. They still have to pay child support. Third, the DNA tests are performed in neutral government-funded labs. The cost of the test is paid by the man if he turns out to be the father; by the woman if not. Fourth, the presumptive father of the newborn can be registered in the hospital even in his absence. However, final legal affiliation of out-of-wedlock paternity still has to be decided by a court, subject to the voluntary acknowledgement of the father or to the DNA test results.

The base of Costa Rican family law is the Código de Familia (Family Code) [Congress, 1973]. For a child to receive support, the father must be legally affiliated. Registering affiliation within marriage is straightforward given the high quality of the marriage registry: the husband of the mother is automatically the father. To affiliate a child outside marriage, there are three major venues: a)Direct acknowledgement[13]. b) The administrative procedure described in the Responsible Paternity Law. c) A court declaration of paternity. In the past, having an out-of-wedlock affiliation meant the child had fewer rights, but this is not true anymore [Camacho, 2006]. Furthermore the [Congress, 1995] amendment to the Family Code considers a 3-year proven cohabitation equivalent to marriage for asset and child support purposes. No-fault divorce exists since 1970 which is important because it could be a confounding factor when analyzing fertility decisions if its implementation had been more recent [Alesina and Giuliano, 2006].

2.2.2 Enforcement of Child Support Legislation

The Responsible Paternity Law by itself only means that more children are affiliated with their biological father, but parental affiliation in Costa Rica guarantees access to child support payments. Judges have discretion in the size of the child support payments, but awards

[11] Law bills in Costa Rica have to be debated and passed twice in Congress. Their constitutionality may be challenged in the Supreme Court and they can be vetoed by the President.

[12] Even this option was introduced in a 1997 reform to the Family Code [Congress, 1997]. Paternity affiliation is automatic if the presumptive father refuses the test.

[13] Direct acknowledgement can be done even against the wishes of the mother if the father does it immediately after birth [Aguirre *et al.*, 2006].

for one child are usually 25-35% of the father's income and up to 50% for multiple children [Mata and Vargas, 2010]. Payments continue until the child reaches legal adulthood (18 years old) or up to 25 years old if the child is still a student.

The Law of Child Support Payments [Congress, 1996][14] also has strong enforcement mechanisms: The obligor (the parent that has to pay child support) cannot travel outside Costa Rica unless 13 months of payments are deposited in advance. Salaried workers can have their wages embargoed. If the obligor is not salaried and refuses to pay, he can be sent to jail for up to six months[15]. Trying to hide assets to reduce or avoid payment can result in fines up to 20 times the monthly payment.

Costa Rican courts have very experienced family law judges and lawyers thanks to a long period of development and specialization and heavy caseloads (23 thousand child support cases in 2004). There are six specialized child support courts and sixty-six small-issue courts that can handle child support problems. Children are protected by this system only if they are affiliated to a father. This explains the strong political and social support for the Responsible Paternity Law[16].

2.2.3 The direct effects of the Paternity Law

The Law intended to reduce the number of children without a registered father at birth. In this, it was very successful. Births without a registered father fell from roughly 30% of all births in 2001 to 8% in 2002. Unregistered fathers also fell from 55% to 18% of all births from unmarried mothers see Figure 2.2a. This is partly an artifact of changes in the way births are registered, but it also implies that the majority of unmarried mothers knew who was the father but were unable to get him to register as such.

From the Law publication until December 2008, mothers made 33,830 formal paternity requests to the Civil Registry. Of those, 44% of the presumptive fathers accepted or were given paternity outright and 25% were problematic because the information provided by the mother was incomplete or insufficient to find the father. Hence, as of December 2008 only 10,332 DNA tests were requested (31% of paternity requests) and one third had not been performed yet (or results were pending). Of the two-thirds of the DNA tests for which results were known (a total of 6595 completed DNA tests), 81% were positive [TSE, 2008].

No DNA test was needed for the great majority of the new registered fathers. The threat sufficed. Consider the following: every year there were 20,000 new births with a registered father. From 2002 to December 2008, this adds to 140,000 births. Hence, less than 7,000 DNA tests were sufficient to get 140,000 children to have a registered father at birth.

[14] Early versions of Child Support laws in Costa Rica date to 1916. See [Benavides, 2007]

[15] This is a credible threat: in June 19th, 2010 there were 187 men in prison due to child support nonpayment [Mata, 2010]

[16] Polls found 90% of the population supporting the bill while it was discussed in Congress [Vega, 2001a]

In September 2000, [Vega, 2001b] interviewed 100 single mothers that gave birth in two of the major hospitals. According to her research, 82% of the women claimed to have a good relationship with the father. Additionally, a majority of those fathers had helped during the pregnancy. In the end, her sample of 100 single mothers between August 25th and September 20th, 2000 had only 42% of registered fathers[17]. Even though 82% of women had such a good relationship to their unmarried partners, half of them had difficulties getting the father to register. This would partly explain why after the law so few DNA tests were needed relative to the number of additional children that had a registered father.

I could find only one previous analysis of the direct effects of the Law itself. [Vega, 2006] estimated how likely was a mother to request a DNA test, conditional on a set of individual characteristics of the mothers for the years 2002 to 2005. His main findings are that young adult urban women are more likely to request the DNA test.

A natural question would be the effect of the Law on abortion, but non-therapeutic abortions are illegal in Costa Rica. Therefore, the exact number of induced abortions is unknown. [Gómez, 2008] estimates that roughly 25% of pregnancies in 2006 were terminated with an induced abortion (Gómez 2008). This is double the 1991 estimate (the only other estimate available). The fifteen year gap between the two estimates makes them unhelpful to determine the direct effect of the Law on abortions.

Within stable partnerships, the rate of self-reported contraceptive use in Costa Rica is high, at around 80%. It has remained at this level in the last two reproductive health surveys (1999 and 2009)[Gómez, 2009]. This would suggest that the Law didn't affect contraceptive use. However, the specific contraceptive method has changed significantly in this decade. For example, condom use has fallen substantially, while male sterilization has risen. Health statistics on vasectomies performed at public hospitals confirm this result[18].

In 2008, 16% of children born in Costa Rica had a Nicaraguan mother. In 1986, only 2.8% of newborns in Costa Rica had a Nicaraguan mother. Nicaraguans are more likely to be unmarried than Costa Ricans. They also face more trouble getting the father to register at birth. The fast growth of the Nicaraguan immigrant population and its distinct marital status makeup suggest that both the entire population of mothers and the subpopulation of Costa Rican mothers should be tested to ensure that results are not driven by Nicaraguan behavior.

[17] The Census and Statistics database shows 561 births from single mothers in those two hospitals over this period, of which 40% had a registered father. This would seem to imply that the interviewed sample is close to the true population.

[18] The increase in male sterilization is more likely to be linked to the removal of certain restrictions on it in 1999

2.3 Data description

The three main data sources to estimate the impact of the Law are:

1. The Database of Registered Births, compiled by the Costa Rican National Institute of Census and Statistics(INEC). It has basic information on 1.8 million births from 1986 to 2008, which is the entire population of registered births (registration coverage is estimated by INEC to be above 98% of births). For example: It has day and location at the hospital-level of the birth. Of the child it has gender and whether the child is a singleton or part of a multiple birth. Of the mother we have age, civil status and nationality. If known, of the father we have age and civil status. There is also an affiliation variable that has a value of 1 if the father of the child acknowledges his affiliation to the hospital register This is important as a measure of true stability of cohabitation because many fathers do not show up to register even if the mother claims to be cohabiting with him. Since 2002, the database collects additional information on the mother and the father, such as education level and occupation. This information isn't as useful because there is no pre-law baseline and [Vega, 2006] finds it less reliable than the more basic information. Table 2.2 shows some summary statistics of registered births in Costa Rica, before and after the Law.

Any child born in Costa Rican soil is a citizen of Costa Rica. Registration of births is very good because as a prerequisite of citizenship, the child has to be inscribed in the Civil Registry. Hence, in every Costa Rican hospital there is an office of the Civil Registry, with at least one official in charge of birth inscriptions (see [Vega, 2001b]).

2. The Costa Rican Household Surveys, also compiled by INEC. I use the Surveys from 1993 to 2009. Each survey covers about 10,000 households and 40,000 individuals. The variables change from year to year, but in all years it is possible to derive the structure of the household, its income, the ages and gender of its members, whether they are employed and the sector of employment, education levels and location, plus an expansion factor (how many people are represented by a given observation in the Survey sample) consistent to the level of region.

3. To be able to estimate the general birthrates and total fertility rates, I use the population estimates from [INEC, 2008]. This document shows the yearly estimated population by age and location. The most important source of the estimates is the 2000 Census. The Law change occurs in 2001 when the accuracy of the estimates is high.

The INEC database on births has very few noticeable problems. The only data cleaning issues are: First, the birth data has both year of registration and year of birth. I always use year of birth. Due to this distinction, some registration years have people born before 1986. These people are dropped. Also, there are a few observations with extreme age values, but I restrict my attention to the range from 13 to 49 years. All others are dropped. I then group the ages in 5 year groups except for the oldest mothers who have a group of 12 years.

To estimate the effects of the law I employ time series data constructed as follows: first the individual birth is collapsed by date. There are 8401 days from January 1st, 1986 to

December 31st, 2008. This collapse can be done for all births or by the subgroups of interest (e.g. age or urban/rural subgroups). This yields only basic birth data. February 29th is problematic and in the original dataset it has missing data. I just assign the mean of the day before and the day after. The same fix is applied to November 15th, 1993 which shows three times the normal number of births.

To obtain birthrates, take the yearly population estimates for either the entire population or the subgroup of interest. These estimates are centered on June 30th of each year. Take the difference between yearly estimates and divide by the number of days to get the daily change in the size of the population. This daily change can then be added to the estimate of the previous year, day by day, to interpolate daily total population estimates. Divide the daily birth observations by the daily total population estimates and the daily birthrate is obtained. For estimations I use the log of this birthrate, so that the estimates are approximates of percentage changes.

The total fertility rate is constructed with the daily birthrate for each age group, then multiplying it by 365 to get annualized birthrate. Then multiply each age-group annualized birthrate by the number of years within the age groups (as noted above, most have 5 years, but the 38 to 49 group has 12 years) and add up the results: this yields the total fertility rate, which is interpreted as the number of children a woman would have over her lifetime if her birthrates at each stage of her life mimic the current birthrates of the respective age.

To estimate changes in the ages of parents, I collapse the individual ages of mothers and fathers into a daily mean, again by population and subgroups. To estimate changes in the percentage of married people, I simply define as born within marriage all those births where both the father and mother are known, declared to be married and where the affiliation is positive. All other births are assigned as unmarried. Collapse the population by this marital status dummy and then divide the resulting daily number of births of married mothers by the total of births. This yields the daily percentage of married people.

2.4 Child support and sexual participation: A model

[Akerlof *et al.*, 1996] model premarital sexual activity, childbearing and marriage as a sequence of decisions in a game tree where a man marries if his own cost of getting married is lower than his perception of the cost to his partner of becoming a single mother. I make an extension of this model, where both the possibility of cohabitation and of child support payments are added to the payoffs. This model helps explain two of the main findings of this paper: the large drop in the birthrates and the drop in births within marriage.

Figure 2.1 is a tree diagram showing the sequence of decisions and their payoffs for a couple deciding whether or not to initiate a sexual relationship. In the beginning, the man decides whether or not to initiate a sexual relationship with her partner. If she accepts to have sex, there are potential future consequences. With probability p the woman becomes pregnant. As in [Akerlof *et al.*, 1996], for simplicity ignore contraception and take p as fixed. If the woman becomes pregnant, she next chooses whether or not to have an abortion. If she chooses not to have an abortion, her partner must then decide whether to marry her, start cohabiting with her, or leave her. She has to decide whether or not to accept his offer.

For notational convenience, [Akerlof *et al.*, 1996] normalize the payoffs so that the couple's payoff if they engage in sex and the woman does not become pregnant is 0 for both of them. If they both enjoy sex, forgoing it yields a loss of s_f for the woman and s_m for the man. If a pregnancy results, the woman may decide to abort at a cost of $a_f - c_a$ which is the cost of the abortion itself with a potential offset of c_a which is any transfer the man makes towards defraying the costs of abortion. Abortion is also costly for men, they pay both their abortion cost and any transfer to the partner, for a payoff of $-a_f - c_a$.

If a pregnancy results and the woman decides to carry it to term, there are three possibilities for the man. 1. He may marry the woman, in which case the payoff for the woman is the cost bf of childbearing within marriage and for the man it is. Marriage is also costly for men (e.g. they may forego other parts of their social life) and this is represented by the $-d_{mm}$ payoff.

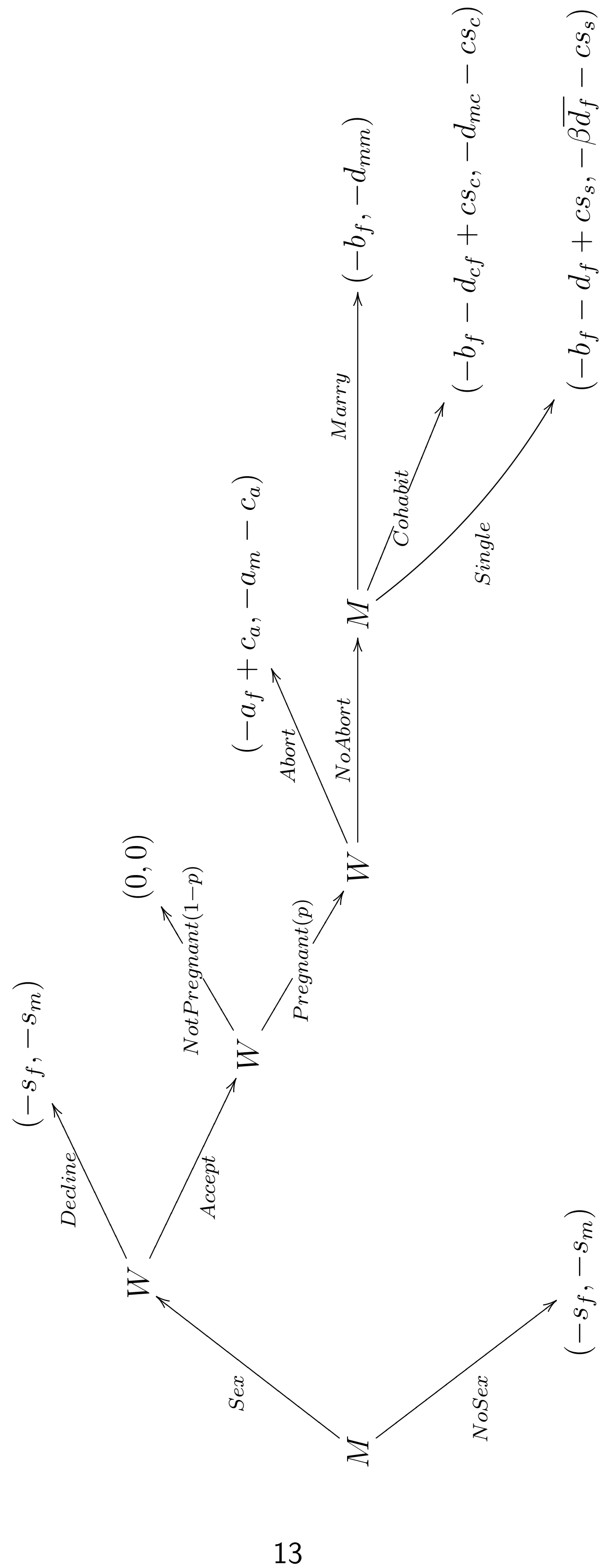
2. He may offer to cohabit with the woman, in which case the payoff for the woman is $-bf - d_{cf} + cs_c$. This addresses two costs: that of childbearing as in marriage, minus a cost d_{cf} due to the added relationship uncertainty of cohabiting (e.g. under Costa Rican law, if the man leaves within three years, she may receive no assets after separation), plus a potential child support transfer cs_c. For the man, the payoff is $-d_{mc} - cs_c$ where d_{mc} is the social cost of cohabiting with the woman and cs_c is the additional cost of child support transfers.

3. He may leave the woman as a single mother, in which case the mother has a payoff of $-b_f - d_f + cs_s$ which is the cost of childbearing minus the cost d_f (emotional and financial) of being a single mother plus potential child support transfers of cs_s. For men, leaving the woman and the child may make them guilty, reflected in the parameter β and in the expected cost $\overline{d_f}$ of being a single mother for the woman, defined by [Akerlof *et al.*, 1996] as the mean value of d_f in the population of women who decide not to have abortions following unplanned pregnancies (this value is used under the assumption that the true cost of childbearing d_f is unobservable for the man). Finally, there is an additional cost to men in potential transfers of child support cs_s.

Before the Responsible Paternity Law, all potential child support transfers are assumed to have been zero. That is, $cs_s = cs_c = 0$ and fathers wouldn't pay anything towards abortion, so $c_a = 0$. For tractability, assume that $d_f > d_{cf} > 0$. This condition implies that in the absence of child support transfers, women will always prefer marriage to cohabitation and cohabitation to being single. Also for tractability, assume that the cost of (illegal) abortions is so high that they are not a choice.

Now consider the introduction of the Law so that $cs_s > cs_c > 0$. The additional child support transfers to a single mother are assumed to be larger than child support transfers

Figure 2.1: Sexual participation and child support



Women's payoffs are on left side, men's on right side.

to a cohabiting partner because a cohabiting man probably was more likely to pay some child support even before the Law. There is a group of men for whom $s_m < \beta\overline{d_f}$ but $s_m > \beta\overline{d_f} + cs_s$. In other words, there is a group of men that was willing to abandon a woman and child instead of foregoing sex, but is also willing to forego sex if a pregnancy will result in paying child support to the mother. Another group of men has preferences such that $s_m < d_{mc}$ but $s_m > d_{mc} + cs_c$. This second group will no longer cohabit after the law, but will rather forego sex.

If those two groups of men are large enough, a large drop in the birthrate should be observed after the passage of a law that increased the expected child support payments, because so many men are foregoing sex. If the assumption of high abortion costs is relaxed, men will also be willing to pay $c_a > 0$ to avoid the much more expensive lifetime costs of child support. This should increase the number of terminated births and yields a similar result: a reduction in the birthrate, except that now the terminated births might reflect any biases men have, such as a latent preference for sons.

However, men are not acting in isolation. The condition of $d_f > d_{cf} > 0$ doesn't preclude the possibility that for some groups of women, $d_{cf} < cs_c$ and even that $d_f < cs_s$. After the law, child support payments may outweigh the relative gain from marrying or cohabiting. Some women will shift from marriage to cohabitation or even to single motherhood. Depending on the relative size of the payments, the direction of the shift from single to cohabiting is unclear, but it is clear that shotgun marriage will decline.

After the law, this model predicts that we should observe a decline in the birthrate and a decline of births within marriage relative to births from cohabiting mothers and single motherhood. Without stronger assumptions, it is not possible to predict the relative changes between cohabiting mothers and single motherhood. Another prediction of the model is that the law may encourage abortions, but reflecting the preferences of men willing to pay for them, which could partly explain the brief appearance of a sex ratio bias towards boys after the Law passed.

2.5 Identification and estimation

This chapter takes advantage of our exact knowledge of each step of the Law's passage. As explained in section 4.2, the most important date for our purposes is February 26th, 2001. On this day, the Responsible Paternity Bill passed its first Congressional debate and became virtually certain to pass. It was widely discussed in the media and a large fraction of the population became aware of it. For example, [Ortega *et al.*, 2005] found that 77% of Costa Rican men knew about the Law in 2002.

Even if the impact of the law on fertility decisions were immediate, the natural length of a pregnancy is between 37 and 42 weeks. Therefore, I chose November 26th, 2001 (nine months after the February first-debate passage), as the date where the impact of the Law on births would be at its peak, because births after this date were very likely to be conceived after the Law became well-known. This date was chosen under the assumption that pre-law pregnancies were not terminated through increased abortion. Otherwise, birthrates could have shifted downwards after February, 2001 but well before late November, 2001.

Section 2.3 describes how the daily time series on variables of interest Y_t such as log birthrates and total fertility rates were constructed for the period between January 1st, 1986 and December 31st, 2008. To estimate the impact of the Law, I construct a dummy D_t where t is the daily trend variable[19]:

$$D_t = \begin{cases} 0 & \text{if } t \leq \text{November 26th, 2001} \\ 1 & \text{if } t > \text{November 26th, 2001} \end{cases} \tag{2.1}$$

The trend in births may have changed because of the law, so both a time trend and a time trend interacted with the Law dummy are added to the regression. The baseline is the following:

$$Y_t = \alpha + \rho D_t + \beta_1 t + \beta_2 D_t t + \eta_t \tag{2.2}$$

A simple linear regression implies strong assumptions, so I also use a more general version of the model by adding squared and cubed time trends and interacted with the dummy. Births in Costa Rica have a strong seasonal pattern, where many more children are born in October than in March. To account for this, take a variable s equal to the day and construct the seasonal adjustment mechanism with up to the sixth higher order. So the most general regression with seasonality adjustment is:

$$Y_t = \alpha + \rho D_t + \beta_1 t + \beta_2 D_t t + \beta_3 t^2 + \beta_4 D_t t^2 + \beta_5 t^3 + \beta_6 D_t t^3 + \Gamma S + \eta_t \tag{2.3}$$

Where ρ is the coefficient of interest that the impact of the Law on the dependent variable. The seasonal adjustment ΓS is:

$$\Gamma S = \gamma_1 s + \gamma_1 s^2 + \gamma_1 s^3 + \gamma_1 s^4 + \gamma_1 s^5 + \gamma_1 s^6 \tag{2.4}$$

Inertia in demographic variables suggests there may be autocorrelation. To correct for this, all estimations are done with Newey-West standard errors (see [Newey and West, 1987]) with 365 lags (to account for a year lag in daily data). To estimate a regression discontinuity, Hahn et al (2001) advise to use linear functions in a short period to ensure identification of the treatment effect[Hahn *et al.*, 2001]. Hence I test robustness with regression specifications where the time period is much shorter. Instead of the full 1986-2008 set, I use only 4 years before and 4 years after the law, which is approximately the period of late November 1997 to late November 2005.

19 [Angrist and Pischke, 2009] notation

The regression above works for log birthrates and total fertility rates, but estimating the impact on marital status is a bit more complicated. The "correct" model would be to consider the law impact as affecting the joint conditional distribution of $P(B, M|X, L)$ where B is births, M is marital status, X is other covariates and L is the Responsible Paternity Law. To be able to test for the impact of the law, we make the following decomposition:

$$P(B, M|X, L) = P(M|B, X, L) * P(B|X, L) \tag{2.5}$$

Here $P(M|B, X, L)$ can be estimated with the proportion of births of a given marital status and $P(B|X, L)$ is estimated with total births. As long as the dependent variable is the proportions of births that come from married mothers then we can estimate $P(M|B, X, L)$ with:

$$M_t = \alpha + \rho D_t + \beta_1 t + \beta_2 D_t t + \beta_1 t^2 + \beta_2 D_t t^2 + \eta_t \tag{2.6}$$

Where ρ is the coefficient of interest that estimates the impact of the Law on the percentage of births that come from married mothers. This makes sense as long as the story is one of couples deciding to have premarital sex and then if a pregnancy results, deciding their marital status. This fits with the model described in section 2.4 in which the sequence of decisions is precisely that.

2.6 Results

The number of children without a registered father fell because the Responsible Paternity Law made it easier to register the father. This could have happened without any further changes in parents' behavior. However, in Figure 2.2b, based on yearly data, there is some evidence that both the total number of births and the number of births within marriage had a large negative change from 2001 to 2002.

Daily data on births permits a more precise estimation of the effect of the change in births. As described in section 2.5, November 26th, 2001 is a good choice to estimate the discontinuity. For example, in Figure 2.3a, the number of births for every date is displayed. To ease visual identification of the discontinuity, I superimposed the OLS predicted values of daily births yielded by regression 2.3 (without seasonality adjustment). Figure 2.3a suggests the Law caused a drop of roughly 25 births off the daily average.

The fertile-age female population in Costa Rica has been growing very quickly over the past two decades[20] and the age structure within this population has also changed. To account for both phenomena I use the total fertility rate (TFR), for which Figure 2.3b shows the discontinuity in November 26th, 2001, using the same superimposition of predicted values

[20][INEC, 2008] estimated this population to have grown from less than 700,000 in early 1986, to more than 1.2 million women in 2008.

of regression 2.3. Because of the Law, women went from expecting to have 2.4 children over their reproductive lifetimes to expecting to have 2.2 children.

From institutional circumstances, I have argued that November 26th, 2001 is the appropriate date to estimate the impact of the Law. For further evidence that this is a good choice, consider Figure 2.4a. It shows the pointwise estimate and 95% confidence bands[21] of the discontinuity in births for every day between May 31st, 1996 and May 14th, 2007 (2000 days before and after my date of interest). The largest estimated discontinuities on the birth data are precisely around our date of interest. As a robustness check, Figure 2.4b shows the highest R^2 values for each of those daily regressions are at or near the date of interest.

Another robustness check is to compare Figure 2.5a and Figure 2.5b. In both, the estimates of the discontinuity in November 26th are significant (5% fall in birthrates, 1.5% fall in percentage of mothers who are married). However, after running the same regression using other dates as the discontinuity in the percentage of married mothers, it is easy to find later or earlier dates with larger estimated discontinuities. This is not true for the birthrates graph.

The main findings of this chapter are:

1. The Responsible Paternity Law had a large immediate impact on both birthrates and the Total Fertility Rate. Even the most conservative specifications yield at least a 4% fall and most results show a fall in the 6%-10% range. This is striking because the time trends show that Costa Rican fertility reached a plateau soon after the Law at a total fertility rate around the replacement rate (2.1 children per woman). Hence, there was little margin for birthrates to fall even further. See Table 2.3.

2. As discussed in section 2.2, Nicaraguan mothers were a rising fraction of all births in Costa Rica during the 1990s. To address concerns that a decrease in immigration in the early 2000s might be driving the results, Table 2.4 shows that results are robust to restricting the population to only Costa Rican mothers.

3. First-time mothers had a much larger reaction to the Law. From Table 2.4, their fertility outcomes fell between 7% and 12%.

4. Women older than 33 years old were little affected by the Law. The largest fall in fertility outcomes occurred in the 23 to 27 years old group[22]. See 2.5. Those results hold if we consider only Costa Rican women and become stronger if we consider only first-time mothers (Table 2.6)

5. Any differences in the effects of the Law in urban and rural areas are not very robust, but this may be due to less precision in the population estimates which weaken the quality of the birthrate estimates. See Table 2.7.

6. Births within marriage fell by as much or even more than total births, so that the

[21] The point estimates were obtained using a simple linear regression specification with the Law dummy, linear trends and the interaction. Robust standard errors were estimated for the confidence bands

[22] Risky sex decisions might be different across age groups, especially if youths have higher self-control problems. See [O'Donoghue and Rabin, 2001]

percentage of children born within marriage fell between 0% and 2% but this result is not robust (See Table 2.8 and Figure 2.5b). Given that people who marry before a pregnancy results are not affected by the Law, this suggests that premarital sex and shotgun marriages are common in Costa Rica and that they may have declined after the Law as women realized they could get child support without marrying. Unfortunately, I cannot find evidence of a sharp drop in the marriage rate in Figure 2.6a.

7.Table 2.9 is surprising because it shows that the Law generated a statistically significant increase in the percentage of children that are boys. Table 2.10 confirms the effect, especially among urban first-time mothers. Section 2.1 discusses the "missing women" phenomenon It is unexpected because health outcomes of girls in Costa Rica are the same as boys and similar to those of developed countries. The size of the effect suggests that 1.5 to 2% of girls in 2002 were "missing". It may be that the sudden shock of the Law made latent preferences for boys more salient.

8. After the law, mothers became older by about two and a half months. The effect on fathers is less clear. Some subgroups of fathers such as first-timers became younger; those who got minors pregnant became older(see Tables 2.11 and 2.12).

9. Impact of the law on marital status was restricted to whether the mother was married or not. A more fine-grained breakdown of marital status is not available until after the Law. Table 2.13 shows that cohabitation has been growing fast enough that it more than compensates the decline of marriage. The upshot is that single mothers are declining. Household Survey Data has cohabitation information before 2001. I use this to build two time series of the percentage of mothers in the survey that claim to be cohabiting. One of mothers in nuclear families (which would be the most direct substitute to traditional marriage) and one of all cohabiting mothers. The results are in Figure 2.6b. The levels of cohabitation in the Surveys suggest that the initial values of cohabiting mothers in the Birth Database were underestimating the true values but they converge afterwards. There seems to be no evidence of a positive shock on cohabitation decisions after the Law.

2.7 Conclusion

Single mothers and their children represent a large fraction of the poor in Costa Rica. They are especially likely to be poor if the father of the child refuses to pay child support. During the 1990s, the percentage of children born without a registered father rose to more than 30% of all births. In 2001, concerned Costa Rican lawmakers decided to address this problem. The Responsible Paternity Law shifted the burden of proof of paternity from mothers to fathers. Presumptive fathers now have to take DNA tests to prove they are not the father.

The Law was very successful in decreasing the number of children without a registered father. In Costa Rica child support is paid only by a legally affiliated father. Child support is expensive for fathers and strongly enforced by the courts. Therefore, the Law should affect childbearing and marital status decisions.

I take advantage of the knowledge of the exact date the Law was passed and publicized in the media to estimate its impact on fertility and marital status. For this, I construct daily time-series of the number of births and other outcomes of interest, depending on the subgroup of interest. Then I employ a simple regression discontinuity with a dummy that takes a value of 1 if the date is after the Law passage and 0 otherwise[23]. I also control for time trends, time trends interacted with the Law dummy and seasonality.

I find that the Responsible Paternity Law affected childbearing and marital status decisions. Birthrates fell and it fell the most in those women who had not taken the decision to start their childbearing. This discontinuity in fertility decisions induced by the Law is striking because it happened in the context of a rapidly falling total fertility rate that was approaching the replacement rate of 2.1 children per woman.

The Law doesn't change the childbearing incentives for married couples. Married couples should be a "control group" that shows no change in fertility. However, I find that births within marriage fell as much or more than total births. A possible explanation is that premarital sex is common and shotgun marriage has declined as women realize they don't need to marry to receive child support.

The most surprising result after the Law is the immediate increase in the percentage of boys, which is usually very stable. This could be evidence for some latent "preference for sons". For example, fathers might be willing to pay for selective abortion of girls if they do not want to pay child support for them.

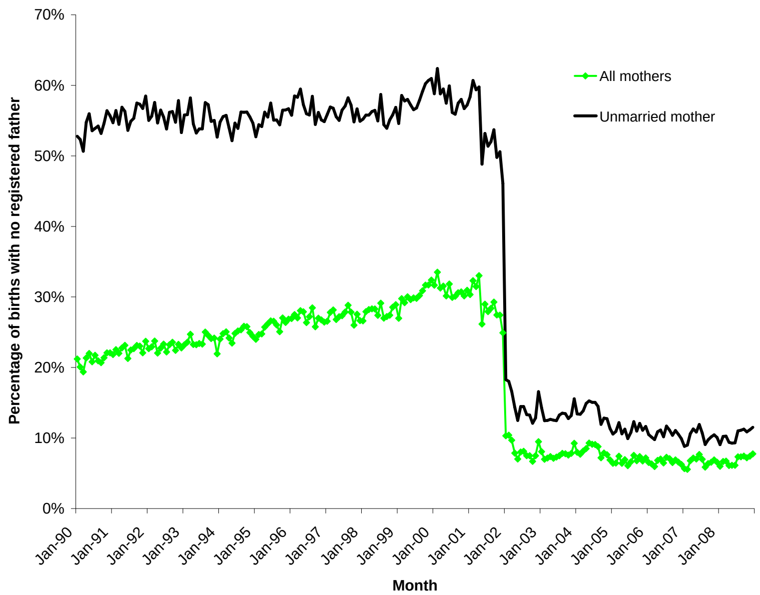
For future research, this permanent and dramatic drop in the birthrate and especially in total births implies potential discontinuities across cohorts in educational outcomes, crime, access to health services and so on. Eventually, with adult outcomes, it will be possible to determine if there was a positive selection effect due to the Responsible Paternity Law. Some suggestive evidence on the economic impact of the law is discussed in Chapter 3.

2.8 Tables and Figures

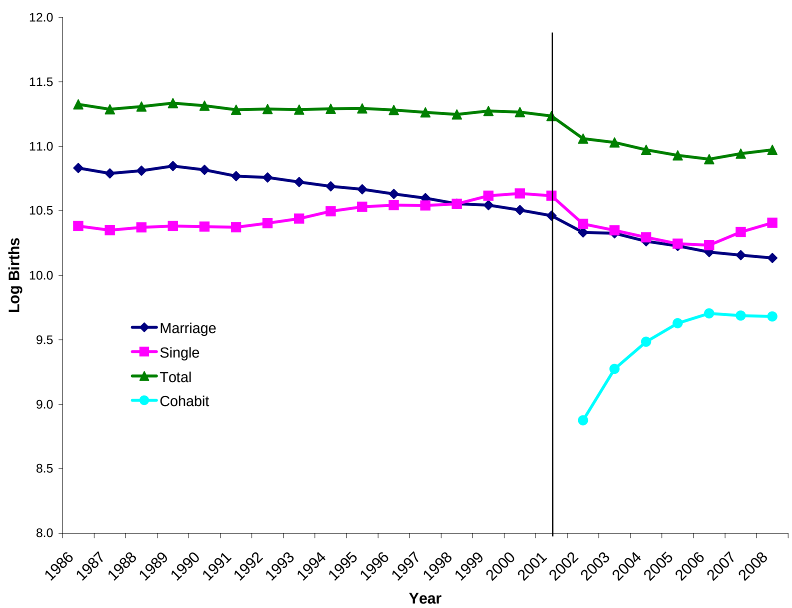
[23] More precisely, I use the date of the law passage plus nine months. Women who were already pregnant cannot be affected by the Law unless they decide to have an abortion.

Figure 2.2: Births with no registered father and Log Births

(a) Percentage of children without a registered father, 1990-2008



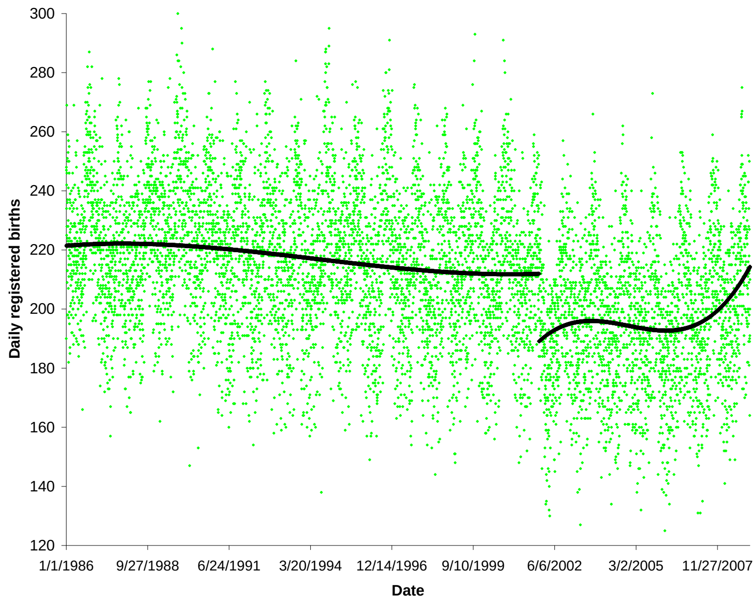
(b) Log births from 1986 to 2008, total and by marital status



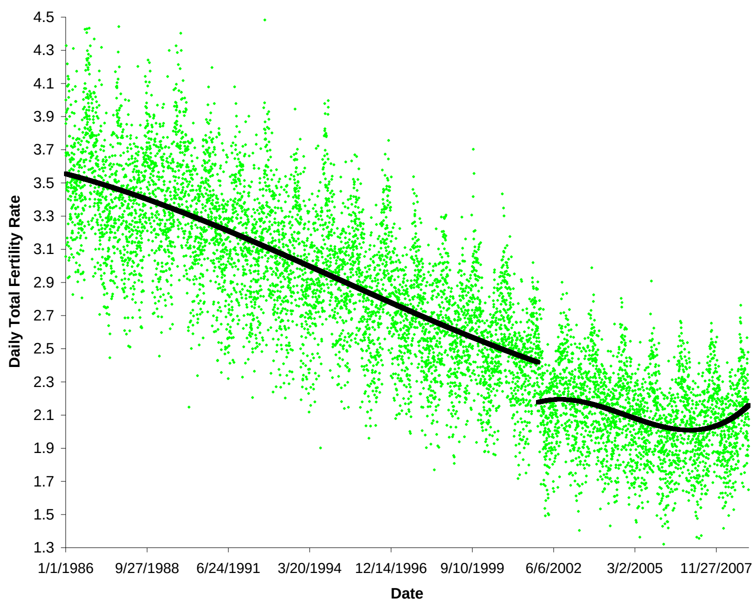
Notes: Black vertical line emphasizes the year 2001, when the law was passed. Source of data is the Birth Database of the Institute of Census and Statistics

Figure 2.3: Daily observed values and regression predicted values, 1986-2008

(a) Total Births



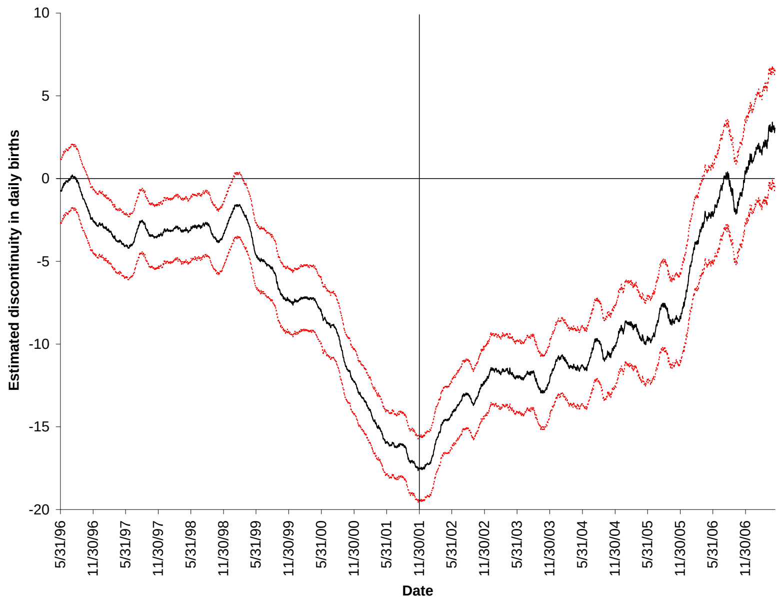
(b) Total Fertility Rate



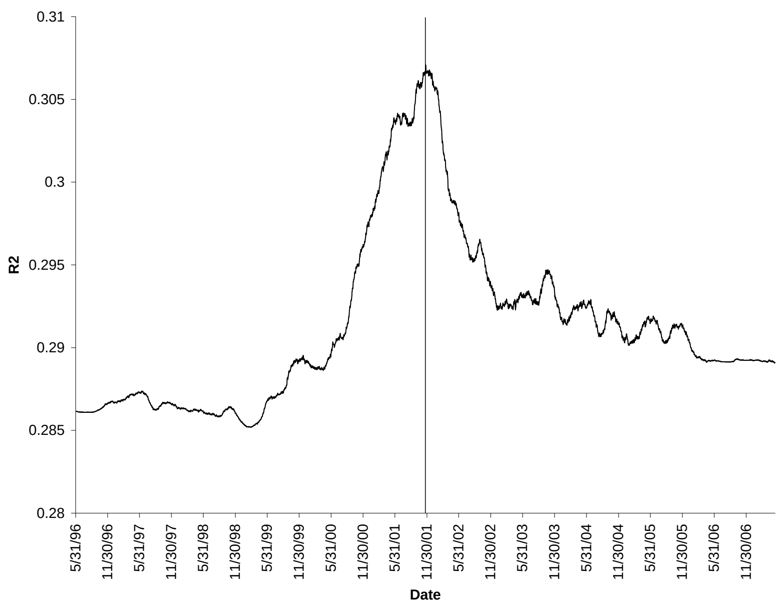
Notes: These are registered births from the Births Database and estimated Total Fertility Rate using births and population estimates from INEC. Black lines are the predicted values from an OLS regression with a time trend, its square and cube, a law dummy that starts on Nov 26th, 2001 and the interactions.

Figure 2.4: Daily births discontinuity, May 1996- May 2007

(a) Estimated discontinuity in daily births



(b) Values of the coefficient of determination, R^2

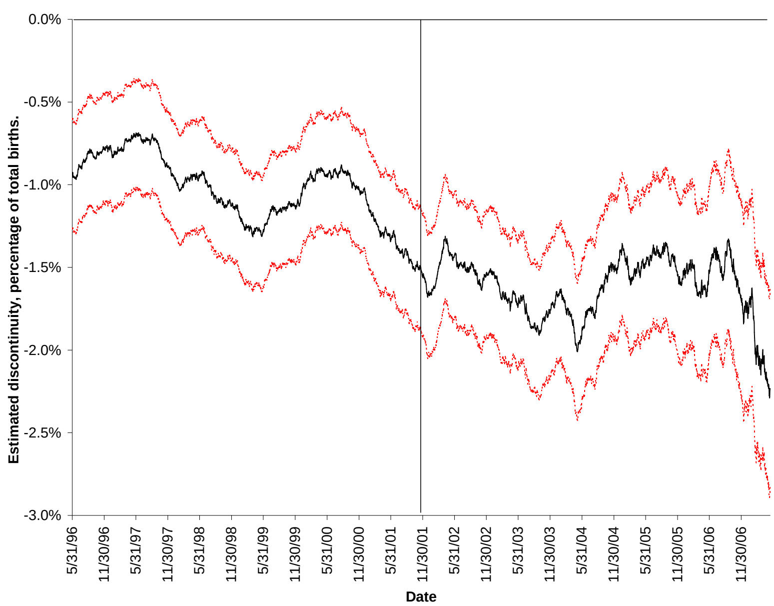


Notes: For every day between May 31st, 1996 and May 14th, 2007, a regression was run to estimate the discontinuity in daily births, controlling for long-term trends and seasonality. The discontinuity estimated for each day and the R2 of the fit are displayed. Vertical lines represent Nov 26th, 2001

Figure 2.5: A comparison of estimated discontinuities, May 1996- May 2007

(a) In log daily birthrate

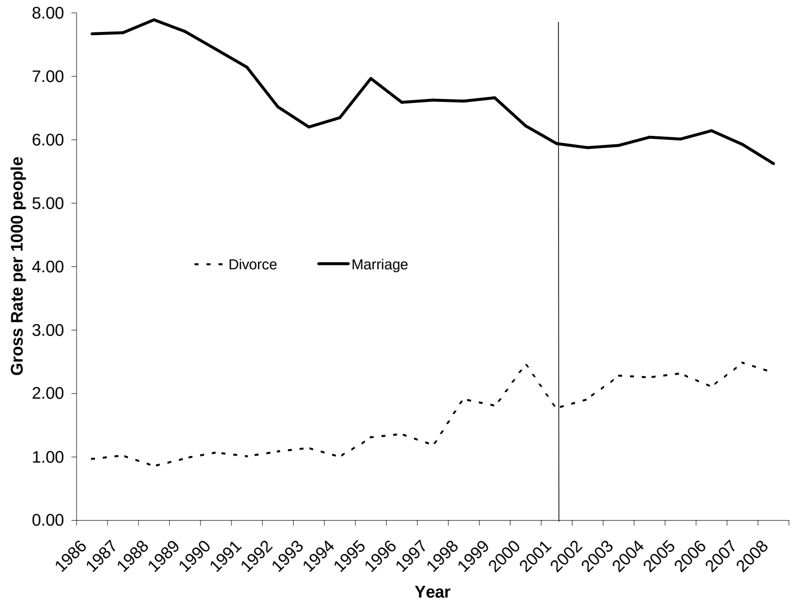
(b) In percentage of children born within marriage



Notes: For every day between May 31st, 1996 and May 14th, 2007, the discontinuity in the log birthrate and the percentage of children born to a married mother was estimated, controlling for long-term trends and seasonality. Vertical lines represent Nov 26th, 2001

Figure 2.6: Marriage, Cohabitation and Divorce

(a) Marriage and Divorce Rates, 1986-2008



(b) Cohabitation, 1993-2008

Notes: Data sources are the Civil Registry for Marriage and Divorce Rates. The Household surveys and Database of Births of INEC for Cohabitation trends

Table 2.2: Summary statistics, before and after the Responsible Paternity Law

	Number of Births		
	Before	After	
Years	16	7	
All Births	1,269,215	510,511	
Births per year	79,326	72,930	
	Percentage of births		
	Before	After	Change
Age of the mother			
Unknown	0.6%	0.4%	-33%
13 to 17	8.0%	9.0%	4%
18 to 22	27.4%	29.9%	0%
23 to 27	27.6%	27.0%	-10%
28 to 32	20.4%	18.8%	-15%
33 to 37	11.4%	10.6%	-14%
37 to 49	4.7%	4.3%	-16%
Age of the father			
Unknown	25.4%	30.2%	9%
13 to 17	0.2%	0.4%	92%
18 to 22	9.3%	10.5%	3%
23 to 27	20.2%	18.2%	-17%
28 to 32	19.8%	16.8%	-22%
33 to 37	13.3%	12.2%	-16%
37 to 49	11.8%	11.8%	-8%
Other characteristics			
Urban	42.4%	42.8%	-7%
Costa Rican	91.8%	82.3%	-18%
Unmarried mother	44.7%	61.2%	26%
No registered father	19.5%	7.6%	-64%

Notes : Source Data from the Database of Births, INEC. Change is the change in the average number of births within a given category before and after the Law.

Table 2.3: Impact of the Law on the Log Birthrate and Log Total Fertility Rate of all mothers

	Log Birthrate			Log Total Fertility Rate		
	(1)	(2)	(3)	(1)	(2)	(3)
Jan 1986- Dec 2008						
Law	$-.104^{***}$	$-.0464^{***}$	$-.0785^{***}$	$-.113^{***}$	$-.0502^{***}$	$-.0731^{***}$
	(.0153)	(.0169)	(.0169)	(.0157)	(.0167)	(.0165)
Trend(year)	$-.033^{***}$	$-.0385^{***}$	$-.0308^{***}$	$-.0252^{***}$	$-.0327^{***}$	$-.0329^{***}$
	(.00096)	(.00254)	(.00717)	(.00096)	(.00266)	(.00709)
Law*Trend(year)	$.0223^{***}$	$-.00215$	$.027^{*}$	$.0135^{***}$	$-.00893$	$.0305^{**}$
	(.00375)	(.00939)	(.0151)	(.00381)	(.00984)	(.0152)
Constant	-8.55^{***}	-8.52^{***}	-8.51^{***}	$.887^{***}$	$.91^{***}$	$.911^{***}$
	(.00698)	(.0118)	(.0139)	(.00744)	(.0118)	(.0137)
R^2	0.794	0.832	0.833	0.729	0.778	0.779
N	8401	8401	8401	8401	8401	8401
Nov 1997- Nov 2005						
Law	$-.102^{***}$	$-.058^{***}$	$-.0424^{**}$	$-.1^{***}$	$-.0573^{***}$	$-.0418^{**}$
	(.0223)	(.0161)	(.0212)	(.0225)	(.0162)	(.0213)
Trend(year)	$-.023^{***}$	$-.0657^{***}$	$-.128^{***}$	$-.0189^{***}$	$-.062^{***}$	$-.125^{***}$
	(.0061)	(.0131)	(.0288)	(.00615)	(.0134)	(.0294)
Law*Trend(year)	$.00752$	$.0573^{***}$	$.138^{***}$	$.00262$	$.0539^{***}$	$.135^{***}$
	(.00854)	(.0163)	(.0468)	(.0086)	(.0167)	(.047)
Constant	-8.54^{***}	-8.55^{***}	-8.57^{***}	$.889^{***}$	$.874^{***}$	$.851^{***}$
	(.0141)	(.0139)	(.017)	(.0142)	(.0144)	(.0172)
R^2	0.382	0.494	0.495	0.354	0.468	0.470
N	2921	2921	2921	2921	2921	2921

Notes : $^{*}p < 0.10,^{**}p < 0.05,^{***}p < 0.01$. Newey-West standard errors in parentheses. (1) Only dependent variables shown (2) Additional controls are seasonality adjustment and squared trends interacted with Law dummy. (3) Additional controls are seasonality adjustment, squared and cubed trends interacted with Law dummy. Birth data from the Database of Births, INEC. Population estimates from INEC.

Table 2.4: Impact of the Law on the Log Birthrate and Log Total Fertility Rate, 1986-2008

	Log Birthrate			Log Total Fertility Rate		
	(1)	(2)	(3)	(1)	(2)	(3)
Costa Rican Mothers						
Law	−.125***	−.0428**	−.076***	−.133***	−.0455***	−.0701***
	(.0161)	(.0171)	(.017)	(.0171)	(.0166)	(.0169)
Trend(year)	−.0427***	−.0586***	−.0439***	−.0347***	−.0526***	−.046***
	(.00144)	(.00298)	(.00692)	(.00146)	(.00284)	(.00709)
Law*Trend(year)	.0292***	.0175**	.0256	.0196***	.00987	.0298*
	(.00344)	(.00785)	(.0158)	(.00354)	(.00818)	(.0159)
Constant	−8.71***	−8.72***	−8.7***	.724***	.714***	.723***
	(.0104)	(.0154)	(.0141)	(.0115)	(.0146)	(.0145)
R^2	0.848	0.875	0.875	0.807	0.841	0.841
N	8401	8401	8401	8401	8401	8401
First-time Mothers						
Law	−.0775***	−.0822***	−.116***	−.121***	−.0727***	−.109***
	(.0192)	(.0191)	(.0215)	(.0172)	(.0201)	(.0207)
Trend(year)	−.0168***	−.0101***	−.00563	−.0066***	−.0173***	−.0135
	(.00109)	(.0034)	(.00973)	(.00113)	(.0034)	(.00949)
Law*Trend(year)	.0284***	.0171	.0598***	.0192***	.0192	.0673***
	(.00326)	(.0123)	(.0188)	(.00357)	(.0126)	(.0184)
Constant	−9.55***	−9.5***	−9.49***	−.0934***	−.0887***	−.0825***
	(.011)	(.016)	(.0214)	(.0102)	(.0158)	(.0201)
R^2	0.340	0.428	0.429	0.172	0.279	0.281
N	8401	8401	8401	8401	8401	8401

Notes : $^{*}p < 0.10, ^{**}p < 0.05, ^{***}p < 0.01$. Newey-West standard errors in parentheses. (1) Only dependent variables shown (2) Additional controls are seasonality adjustment and squared trends interacted with Law dummy. (3) Additional controls are seasonality adjustment, squared and cubed trends interacted with Law dummy. Birth data from the Database of Births, INEC. Population estimates from INEC.

Table 2.5: Impact of the Law on the Log Birthrate of all mothers by age groups

Age Group	13 to 18	18 to 22	23 to 27	28 to 32	33 to 37	38 to 49
Jan 1986- Dec 2008						
Law	−.0862***	−.0395**	−.0791***	−.0531***	−.00328	−.0194
	(.0273)	(.0171)	(.0193)	(.0203)	(.0235)	(.025)
Trend(year)	−.0134*	−.032***	−.0344***	−.0215***	−.0431***	−.0809***
	(.00727)	(.00271)	(.00381)	(.00378)	(.0042)	(.00509)
Law*Trend(year)	−.0654***	−.00816	.0106	−.016	.00733	.00583
	(.016)	(.00926)	(.0125)	(.00983)	(.0135)	(.0136)
Constant	4.15***	4.81***	4.83***	4.57***	4***	2.42***
	(.0314)	(.0139)	(.0141)	(.0166)	(.0344)	(.0254)
R^2	0.197	0.600	0.605	0.493	0.484	0.519
N	8401	8401	8401	8401	8401	8401
Nov 1997- Nov 2005						
Law	−.0468**	−.0425***	−.0784***	−.0724***	−.0389	−.0637***
	(.0229)	(.0163)	(.0149)	(.0269)	(.0243)	(.024)
Trend(year)	−.11***	−.077***	−.0613***	−.0199	−.0379*	−.0823***
	(.023)	(.0136)	(.0107)	(.0187)	(.0218)	(.0316)
Law*Trend(year)	.0915***	.0741***	.0786***	−.00942	.0184	.0318
	(.0257)	(.0164)	(.018)	(.0255)	(.0265)	(.038)
Constant	4.06***	4.76***	4.79***	4.59***	3.99***	2.42***
	(.0266)	(.0156)	(.0183)	(.0248)	(.0498)	(.0346)
R^2	0.196	0.315	0.281	0.181	0.105	0.153
N	2921	2921	2921	2921	2921	2921

Notes : $^{*}p < 0.10,^{**} p < 0.05,^{***} p < 0.01$. Newey-West standard errors in parentheses. Besides the dependent variables shown, all regressions have a seasonality adjustment and squared trends interacted with Law dummy. Birth data from the Database of Births, INEC. Population estimates from INEC.

Table 2.6: Impact of the Law on the Log Birthrate by age groups of mothers, 1986-2008

Age Group	13 to 18	18 to 22	23 to 27	28 to 32	33 to 37	38 to 49
Costa Rican Mothers						
Law	−.0574**	−.0448***	−.0646***	−.0547**	−.00673	−.0266
	(.0234)	(.0165)	(.0197)	(.0223)	(.0287)	(.0271)
Trend(year)	−.0385***	−.0524***	−.0585***	−.0404***	−.0568***	−.0916***
	(.00616)	(.00338)	(.00307)	(.00457)	(.00481)	(.00558)
Law*Trend(year)	−.0308**	.0275***	.0325***	−.0151	.00723	.00865
	(.0136)	(.00852)	(.0116)	(.00941)	(.0152)	(.0142)
Constant	3.91***	4.6***	4.61***	4.39***	3.86***	2.27***
	(.0288)	(.0154)	(.0142)	(.0206)	(.0373)	(.0286)
R^2	0.295	0.699	0.709	0.606	0.557	0.544
N	8401	8401	8401	8401	8401	8401
First-time Mothers						
Law	−.0851***	−.0505**	−.0936***	−.12***	−.0458	−.00637
	(.0288)	(.0238)	(.0276)	(.0304)	(.0384)	(.0151)
Trend(year)	−.0115	−.0251***	−.0252***	−.00044	.00901	−.00752
	(.00699)	(.00422)	(.00306)	(.00657)	(.00788)	(.00488)
Law*Trend(year)	−.0454***	.0271**	.0642***	.0884***	.0452*	−.00077
	(.017)	(.0132)	(.0163)	(.0104)	(.0242)	(.011)
Constant	3.98***	4.23***	3.53***	2.73***	1.47***	.235***
	(.0293)	(.0215)	(.0206)	(.0426)	(.0685)	(.029)
R^2	0.116	0.274	0.097	0.054	0.041	0.009
N	8401	8401	8401	8401	8401	8401

Notes : $^{*}p < 0.10,^{**}p < 0.05,^{***}p < 0.01$. Newey-West standard errors in parentheses. Besides the dependent variables shown, all regressions have a seasonality adjustment and squared trends interacted with Law dummy. Birth data from the Database of Births, INEC. Population estimates from INEC.

Table 2.7: Impact of the Law on Log Birthrates in urban and rural areas

	All mothers		Costa Rican mothers		First-time mothers	
	Rural	Urban	Rural	Urban	Rural	Urban
Jan 1986- Dec 2008						
Law	$-.0597^{*}$	$-.024$	$-.0479$	$-.0308$	$-.0875^{**}$	$-.0716^{*}$
	(.0333)	(.0456)	(.0317)	(.0472)	(.0393)	(.0373)
Trend(year)	$-.0528^{***}$	$-.0306^{***}$	$-.0722^{***}$	$-.0516^{***}$	$-.0218^{**}$	$-.00486$
	(.00949)	(.00995)	(.00884)	(.0118)	(.0111)	(.00767)
Law*Trend(year)	$.0314^{**}$	$-.0352^{**}$	$.0485^{***}$	$-.013$	$.0546^{***}$	$-.0155$
	(.0156)	(.014)	(.0152)	(.0154)	(.0198)	(.0127)
Constant	-2.57^{***}	-2.82^{***}	-2.76^{***}	-3.03^{***}	-3.62^{***}	-3.74^{***}
	(.0348)	(.0359)	(.0329)	(.043)	(.0432)	(.0275)
R^2	0.842	0.657	0.872	0.745	0.454	0.217
N	8401	8401	8401	8401	8401	8401
Nov 1997- Nov 2005						
Law	$.0337$	$-.15^{***}$	$.0294$	$-.168^{***}$	$.0194$	$-.16^{***}$
	(.0308)	(.0243)	(.0349)	(.0252)	(.0325)	(.0208)
Trend(year)	$-.195^{***}$	$.0688^{***}$	$-.203^{***}$	$.0762^{***}$	$-.203^{***}$	$.0311^{*}$
	(.0304)	(.0193)	(.0332)	(.0218)	(.0348)	(.0167)
Law*Trend(year)	$.204^{***}$	$-.105^{***}$	$.214^{***}$	$-.132^{***}$	$.287^{***}$	$-.00978$
	(.0323)	(.0361)	(.0345)	(.0381)	(.0382)	(.0295)
Constant	-2.71^{***}	-2.73^{***}	-2.88^{***}	-2.91^{***}	-3.77^{***}	-3.71^{***}
	(.0288)	(.0183)	(.0306)	(.0235)	(.0315)	(.0148)
R^2	0.450	0.430	0.468	0.466	0.168	0.177
N	2921	2921	2921	2921	2921	2921

Notes : $^{*}p < 0.10, ^{**}p < 0.05, ^{***}p < 0.01$. Newey-West standard errors in parentheses. All regressions have the dependent variables shown, seasonality adjustment and squared time trends interacted with the Law dummy. Birth data from the Database of Births, INEC. Population estimates from INEC.

Table 2.8: Impact of the Law on the percentage of children born within marriage

	All mothers		Costa Rican mothers		First-time mothers	
	(1)	(2)	(1)	(2)	(1)	(2)
Jan 1986- Dec 2008						
Law	$-.0153^{***}$	$.00644$	$-.0169^{***}$	$-.00015$	$-.0206^{***}$	$.00113$
	(.00487)	(.00611)	(.00414)	(.00546)	(.00597)	(.00527)
Trend(year)	$-.0114^{***}$	$-.0198^{***}$	$-.00838^{***}$	$-.0144^{***}$	$-.0119^{***}$	$-.021^{***}$
	(.00076)	(.00176)	(.0006)	(.00128)	(.00087)	(.00169)
Law*Trend(year)	$-.00481^{***}$	$.00412^{*}$	$-.00963^{***}$	$-.00424^{*}$	$-.0025^{**}$	$.00861^{***}$
	(.00084)	(.00216)	(.00082)	(.00233)	(.00098)	(.00227)
Constant	$.461^{***}$	$.439^{***}$	$.511^{***}$	$.495^{***}$	$.401^{***}$	$.377^{***}$
	(.00472)	(.00552)	(.00356)	(.00335)	(.00579)	(.00466)
R^2	0.840	0.847	0.790	0.795	0.695	0.702
N	8401	8401	8401	8401	8401	8401
Squared terms	*No*	*Yes*	*No*	*Yes*	*No*	*Yes*
Nov 1997- Nov 2005						
Law	$-.00365$	$-.0087^{**}$	$-.00632^{*}$	$-.009^{**}$	$-.00372$	$-.00744$
	(.00334)	(.00352)	(.00354)	(.00444)	(.00447)	(.00605)
Trend(year)	$-.0141^{***}$	$-.00571$	$-.0115^{***}$	$-.00899^{***}$	$-.018^{***}$	$-.00915^{*}$
	(.00093)	(.0036)	(.00054)	(.00293)	(.00153)	(.00553)
Law*Trend(year)	$-.00212^{*}$	$-.0114^{**}$	$-.00664^{***}$	$-.00763$	$.00407^{**}$	$-.00799$
	(.00125)	(.00557)	(.00139)	(.00643)	(.00167)	(.00652)
Constant	$.449^{***}$	$.455^{***}$	$.501^{***}$	$.502^{***}$	$.383^{***}$	$.389^{***}$
	(.00258)	(.0026)	(.00145)	(.00206)	(.00359)	(.00349)
R^2	0.468	0.469	0.440	0.440	0.305	0.306
N	2921	2921	2921	2921	2921	2921
Squared terms	*No*	*Yes*	*No*	*Yes*	*No*	*Yes*

Notes : $^{*}p < 0.10, ^{**}p < 0.05, ^{***}p < 0.01$. Newey-West standard errors in parentheses. (1) Regression has only the dependent variables shown (2) Regression has additional squared time trends interacted with the Law dummy. Birth data from the Database of Births, INEC.

Table 2.9: Impact of the Law on the percentage of children that are boys

	All mothers		Costa Rican mothers		First-time mothers	
	(1)	(2)	(1)	(2)	(1)	(2)
Jan 1986- Dec 2008						
Law	.00095	.00617***	.00135	.00646***	.00293	.00616**
	(.00208)	(.00173)	(.00206)	(.00175)	(.00279)	(.00275)
Trend(year)	−.00014	−.00037	$-9.0e-05$	−.00035	−.00011	.00042
	$(9.9e-05)$	(.00043)	(.0001)	(.00046)	(.00021)	(.00062)
Law*Trend(year)	−.00022	−.00388***	−.0004	−.00386***	−.00064	−.00507***
	(.00045)	(.00092)	(.00045)	(.00112)	(.00054)	(.00176)
Constant	.512***	.512***	.513***	.512***	.514***	.515***
	(.00095)	(.00137)	(.00102)	(.00152)	(.00159)	(.00173)
R^2	0.001	0.002	0.000	0.001	0.000	0.001
N	8401	8401	8401	8401	8401	8401
Squared Terms	*No*	*Yes*	*No*	*Yes*	*No*	*Yes*
Nov 1997- Nov 2005						
Law	.00357*	.00726**	.00432**	.00831***	.00431*	.00877*
	(.00202)	(.00306)	(.00211)	(.00282)	(.00258)	(.00499)
Trend(year)	$-3.7e-05$	−.00074	$-2.4e-05$	−.00115	.00064	−.00233
	(.00077)	(.0029)	(.00088)	(.00336)	(.00103)	(.00377)
Law*Trend(year)	−.002*	−.00614*	−.00231*	−.00603	−.00377**	−.00451
	(.00109)	(.00316)	(.00123)	(.00415)	(.00149)	(.00461)
Constant	.513***	.512***	.513***	.512***	.516***	.514***
	(.00161)	(.00243)	(.00185)	(.00284)	(.00171)	(.0029)
R^2	0.002	0.003	0.003	0.003	0.002	0.002
N	2921	2921	2921	2921	2921	2921
Squared Terms	*No*	*Yes*	*No*	*Yes*	*No*	*Yes*

Notes : $^{*}p < 0.10,^{**}p < 0.05,^{***}p < 0.01$. Newey-West standard errors in parentheses.(1) Regression has only the dependent variables shown (2) Regression has additional squared time trends interacted with the Law dummy. Birth data from the Database of Births, INEC.

Table 2.10: Impact of the Law on the percentage of births that are boys, by rural or urban area

	All mothers		Costa Rican mothers		First-time mothers	
	Rural	Urban	Rural	Urban	Rural	Urban
Jan 1986- Dec 2008						
Law	.00306	.0104***	.00273	.012***	−.00095	.0132**
	(.00264)	(.00302)	(.00352)	(.00345)	(.00375)	(.00531)
Trend(year)	−.00034	−.00039	−.00042	−.00029	.00042	.00051
	(.00054)	(.00065)	(.00052)	(.0007)	(.00097)	(.00096)
Law*Trend(year)	−.00256	−.00592***	−.00201	−.00689***	−.00159	−.00831***
	(.00174)	(.00133)	(.00222)	(.0016)	(.00262)	(.0031)
Constant	.512***	.511***	.512***	.511***	.516***	.515***
	(.00178)	(.00205)	(.00161)	(.00237)	(.00292)	(.0028)
R^2	0.001	0.002	0.001	0.002	0.000	0.001
N	8401	8401	8401	8401	8401	8401
Nov 1997- Nov 2005						
Law	.00667	.0078***	.00897**	.00781**	−.00081	.0177***
	(.00454)	(.00292)	(.00384)	(.00352)	(.00559)	(.00686)
Trend(year)	.00073	−.00166	−.00081	−.00065	.00482	−.00937**
	(.00441)	(.00302)	(.00414)	(.00385)	(.00598)	(.00391)
Law*Trend(year)	−.0109**	−.00166	−.0108**	−.0019	−.0144*	.0061
	(.00495)	(.00395)	(.00536)	(.00511)	(.00756)	(.00657)
Constant	.513***	.512***	.512***	.513***	.521***	.508***
	(.00395)	(.00229)	(.00353)	(.00303)	(.00507)	(.00269)
R^2	0.002	0.002	0.002	0.003	0.001	0.003
N	2921	2921	2921	2921	2921	2921

Notes : $^{*}p < 0.10,^{**} p < 0.05,^{***} p < 0.01$. Newey-West standard errors in parentheses. All regressions have the dependent variables shown and squared time trends interacted with the Law dummy. Birth data from the Database of Births, INEC.

Table 2.11: Impact of the Law on average mother and father age, 1986-2008

	Parents		Costa Rican First-time parents		
	All	Costa Rican	All	Married	Unmarried
Age of Mothers					
Law	.197***	.174**	.0671	.0944	.0924*
	(.0672)	(.0695)	(.057)	(.0732)	(.0507)
Trend(year)	−.151***	−.128***	−.0515***	.0344**	−.00841
	(.0158)	(.0139)	(.0135)	(.0149)	(.0154)
Law*Trend(year)	.0992***	.00654	.153***	.197***	.122***
	(.0191)	(.0191)	(.0191)	(.0313)	(.0228)
Constant	25.1***	25.3***	21.4***	23.5***	19.9***
	(.0638)	(.0565)	(.0485)	(.0509)	(.0509)
R^2	0.189	0.178	0.071	0.423	0.062
N	8401	8401	8401	8401	8401
Age of Fathers					
Law	.00028	−.00272	−.166***	−.113	.00561
	(.0459)	(.0379)	(.0413)	(.0942)	(.0883)
Trend(year)	−.0302***	−.0268***	.0365***	.112***	−.00315
	(.00822)	(.00916)	(.0106)	(.0191)	(.0178)
Law*Trend(year)	−.0338	−.0875***	.106***	.238***	.118***
	(.0328)	(.0229)	(.0207)	(.0329)	(.0413)
Constant	30.4***	30.4***	26.6***	27.4***	25.1***
	(.0291)	(.0328)	(.0364)	(.0735)	(.0502)
R^2	0.031	0.055	0.196	0.451	0.008
N	8401	8401	8401	8401	8401

Notes : $^{*}p < 0.10,^{**} p < 0.05,^{***} p < 0.01$. Newey-West standard errors in parentheses. All regressions have the dependent variables shown and squared time trends interacted with the Law dummy. Average age of parents data from the Database of Births, INEC.

Table 2.12: Impact of the Law on average age of the father by mother age groups, 1986-2008

Age Group	13 to 18	18 to 22	23 to 27	28 to 32	33 to 37	38 to 49
Law	.209**	−.144	−.183**	−.0202	.00093	.192
	(.103)	(.0996)	(.0733)	(.0573)	(.0978)	(.138)
Trend(year)	−.0612***	.0225	.0722***	.0914***	.0159	−.133***
	(.0225)	(.0182)	(.0156)	(.0108)	(.0186)	(.0339)
Law*Trend(year)	.0475	.0226	−.00554	−.0403	.0762	.176**
	(.0601)	(.0337)	(.0279)	(.0389)	(.0568)	(.0866)
Constant	23.0***	25.7***	29.4***	33.2***	36.8***	40.9***
	(.0719)	(.0761)	(.058)	(.0405)	(.0591)	(.103)
R^2	0.006	0.012	0.046	0.016	0.022	0.033
N	8401	8401	8401	8401	8401	8401

Notes : $^{*}p < 0.10,^{**} p < 0.05,^{***} p < 0.01$. Newey-West standard errors in parentheses. All regressions have the dependent variables shown and squared time trends interacted with the Law dummy. Average age of parents data from the Database of Births, INEC.

Table 2.13: Simple linear trends in the percentage of mothers within a given marital status after the Law

	(1) Marriage	(2) Cohabit	(3) Stable	(4) Claim Cohabit	(5) All Cohabit	(6) Single
All mothers						
Trend(year)	-0.016^{***}	0.019^{***}	0.003	0.015^{***}	0.034^{***}	-0.018^{***}
	(0.000)	(0.004)	(0.004)	(0.002)	(0.004)	(0.004)
Constant	0.443^{***}	0.121^{***}	0.564^{***}	0.089^{***}	0.209^{***}	0.348^{***}
	(0.002)	(0.015)	(0.016)	(0.006)	(0.018)	(0.019)
R^2	0.413	0.496	0.020	0.446	0.687	0.384
N	2557	2557	2557	2557	2557	2557
Costa Rican mother						
Trend(year)	-0.018^{***}	0.020^{***}	0.002	0.012^{***}	0.033^{***}	-0.015^{***}
	(0.001)	(0.004)	(0.004)	(0.002)	(0.004)	(0.004)
Constant	0.491^{***}	0.105^{***}	0.597^{***}	0.072^{***}	0.177^{***}	0.332^{***}
	(0.003)	(0.014)	(0.014)	(0.006)	(0.017)	(0.017)
R^2	0.421	0.515	0.010	0.391	0.678	0.305
N	2557	2557	2557	2557	2557	2557
First-time mother						
Trend(year)	-0.014^{***}	0.021^{***}	0.006	0.011^{***}	0.032^{***}	-0.018^{***}
	(0.000)	(0.004)	(0.004)	(0.002)	(0.004)	(0.004)
Constant	0.379^{***}	0.123^{***}	0.502^{***}	0.079^{***}	0.202^{***}	0.420^{***}
	(0.002)	(0.014)	(0.014)	(0.006)	(0.016)	(0.016)
R^2	0.225	0.395	0.041	0.227	0.540	0.252
N	2557	2557	2557	2557	2557	2557

Notes : $^{*}p < 0.10,^{**} p < 0.05,^{***} p < 0.01$. Newey-West standard errors in parentheses. All regressions have the dependent variables shown. (1) Married mothers (2) Cohabiting mothers whose partner registered as the father. (3) Both married and cohabiting with registered father. (4) Cohabiting mothers but father did not register. (5) All cohabiting mothers. (6) Single mothers. Birth data from the Database of Births, INEC.

Chapter 3

Family structures and the female income distribution in Costa Rica

3.1 Introduction

Single motherhood is highly correlated with female poverty in many countries. For instance, table 3.1 shows that single mothers in Costa Rica in 2008/09 were 8.5% more likely to be poor than women who live in households without dependents.[1] There are several reasons for this correlation: Having more adults in a household means having more potential wage earners. Having children or other dependents such as elderly parents means that there are more people to share the income with. Having (young) children also implies that more work has to be done within the household. This in turn makes it harder for single parents to earn a living by working outside the household. Finally, the correlation might also be due to causal effects from income levels to the likelihood of single motherhood, or due to confounding factors.

The goal of this chapter is to isolate the effect of changing family structures in Costa Rica, from 1993 to 2009, on the female income distribution. We assume that the family status of individuals is independent of the unobserved determinants of their income, conditional on a set of demographic covariates including age, geographic location, and education. We consider counterfactual changes in the income distribution that would have taken place as a consequence of changing family structures if the distribution of demographic covariates, as well as the distribution of equivalent income given family status and covariates, had remained constant during the period of interest.

[0] Chapter written in collaboration with Maximilian Kasy

[1] More precisely, column 2 of table 3.1 shows the poverty rate of women living in households headed by a single parent. A person here is defined to be poor if the equivalent income of her household lies below the poverty line defined by the National Institute of Census and Statistics of Costa Rica for that year. Equivalent income is defined as total income of a household divided by the OECD equivalent scale, see section 3.2.

Table 3.1: Population shares and poverty rates for women in different types of households

	Population share			Population Poverty rate	
	Couples w/dependents	Single w/children	All others	Actual	Counterfactual
1993-94	47.8%	20.4%	31.8%	27.32%	19.41%
2000-01	46.2%	22.7%	31.1%	22.99%	19.52%
2008-09	35.7%	24.8%	39.5%	19.12%	19.12%
	Poverty rate				
2008-09	20.6%	23.5%	15.0%		

Notes: This table shows population shares and poverty rates for working-age women in different types of households, as defined in section 3.2. Counterfactual poverty rates are calculated based on current population shares and 2008-09 poverty rates, see text. All estimates are based on the Costa Rican Household Surveys, INEC.

Table 3.1 illustrates the basic idea, without controlling for covariates. The counterfactual poverty rates in the last column of this table are calculated by holding constant poverty rates for each of the three groups shown, at the level of 2008/09, while letting the population shares of the groups vary over time. Counterfactual poverty is constructed by taking the average of poverty rates in 2008/09, weighted by the changing population shares of each group. According to this decomposition, table 3.1 would imply that changes in family structures led to a 0.11 % increase in the female poverty rate from 1993/94 to 2000/01, and a subsequent decline of 0.4 % from 2000/01 to 2008/09. This adds up to an overall decline of 0.29 % during the entire 1993/94 to 2008/09 period, contributing to the sizable (7.2%) drop in actual poverty rates among working-age women. The identification and estimation approach described in section 3.3 below is essentially a generalization of this construction of counterfactual poverty rates. In this generalization, we (i) control for covariates, (ii) consider a finer classification of family status, as discussed in section 3.2, (iii) use nonparametric estimation methods, and (iv) consider the entire distribution of female equivalent incomes instead of just the poverty rate.

We find that the changes in family structure had an inequality-increasing effect for the entire working-age female population over the 1993/94-2008-09 period, but with quite different patterns of change before and after 2000/01. In the earlier period, the relative changes of incomes are roughly constant and negative across income levels; in the later period, higher incomes grew due to changing family structures, while lower incomes decreased further. In most subgroups we find similar effects, with the exception of urban women, where both high and low incomes were decreased more strongly than intermediate incomes by changing family structures.

This chapter draws on the literature on income decomposition techniques. Our methods build on the reweighting proposed in [DiNardo *et al.*, 1996], as well as the influence function regression developed in [Firpo *et al.*, 2009]. A general discussion of issues arising in the analysis of income inequality, such as the choice of the unit of observation and income measurement can be found in [Atkinson and Bourguignon, 2000]. Decompositions of the household income distribution are discussed in [Bourguignon *et al.*, 2008].

This chapter contributes to the literature in several ways. First, we make a point about the determinants of the income distribution. Market factors, such as demand shifts, trade and technological change, as well as institutional factors, such as (de)unionization and the minimum wage, have received a lot of attention in the literature. In this chapter, we document how family structures play a role in determining the income distribution. Second, we provide a comparison of two estimation methods for the impact of changes in the distribution of a determinant of incomes on the unconditional distribution of incomes: reweighting and influence function regression. Third, we provide suggestive evidence on the impact of the Ley de Paternidad Responsable (Responsible Paternity Law), which was passed in 2001. This law made it much easier for single mothers to get child support payments from the fathers, and seems to have had a considerable impact on family structures.[2] Dividing our period of observations, 1993/94 to 2008/09, into the two subperiods 1993/94 to 2000/01 and 2000/01 to 2008/09, we find that the initial trend of changing family structures leading to lower incomes was reversed in the latter period, at least for higher income levels. This reversion might be due at least in part to the introduction of the Ley de Paternidad Responsable.

There are some empirical papers in the literature emphasizing the impact of family structures on inequality and poverty, in the context of other countries. [Peichl *et al.*, 2010], in particular, have recently discussed the impact of changes in household structure on the income distribution in Germany. Papers that consider the role of family structures in determining the income distribution in the United States include [Lerman, 1996], [Iceland, 2003] and [Martin, 2006]. More recently, [Le and Booth, 2010] used the method of [Firpo *et al.*, 2009] in an analysis of Vietnamese living standards.

Our approach has a number of limitations if our object of interest is the effect of family structures on the distribution of welfare. We only analyze the distribution of income across households, not within households. This is because data on the distribution within households are not available, and we therefore study the distribution of equivalent household income across women. We do not consider nonmonetary consequences of family structures on welfare, such as the division of labor in raising children or taking care of elderly family members. And we do not study the effects of family structures on adult men or children. The reason is that this chapter is motivated by the correlation between single parenthood and the risk of poverty, and the fact that single parenthood is much more common among women then men.

The rest of the chapter is structured as follows: In section 3.2 we describe the institutional

[2]See Chapter 2

background and the dataset that we use, the Costa Rican Household Survey by the Costa Rican National Institute of Census and Statistics. We also discuss how we define variables such as family status and equivalent income. Section 3.3 provides a formal discussion of our identification approach and of nonparametric decompositions of distributional change. In this section we review two estimation methods of the impact of changes in covariate distributions on the unconditional outcome distribution of interest: reweighting and influence function regression. Section 3.4 discusses the main empirical results of this chapter, the estimated impact of the historical changes in family structures on the female income distribution. This analysis is repeated for various subpopulations defined by age, location, and education. Section 3.5 concludes. Appendix 3.6 contains further empirical results.

3.2 Background and data description

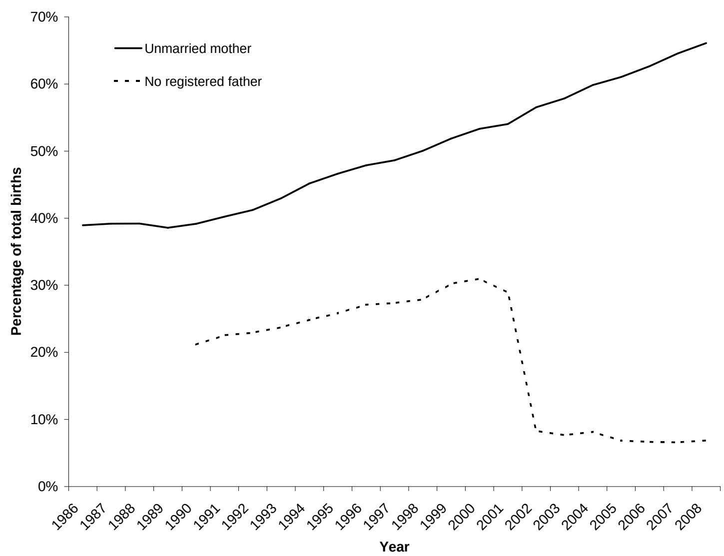
Latin America in general and Costa Rica in particular have historically had a high fraction of births from unmarried mothers. During the 1990s, birth data show that this fraction increased significantly, and especially the fraction of children without a registered father at birth, see figure 3.1.

In response to this development, the "Ley de Paternidad Responsable" (Responsible Paternity Law) was passed in April 2001. The Law shifted the burden of proof of paternity from the mother, who used to have to prove she had a relationship with the presumptive father, to the presumptive father, who now has to submit to a DNA test to prove he is not the father. This makes it much easier for single mothers to obtain child support, and has dramatically decreased the fraction of births without a registered father, as can be seen in figure 3.1.

In Costa Rica, birthrates fell very quickly during the 1990s. Immediately following the introduction of the law, birthrates dropped by an additional 5%. By 2003, the birthrates were below the replacement rate (less than 2.1 children per woman over her reproductive lifetime), and the 2000s have been a period of relative stability in fertility rates, albeit at much lower levels than in the 1990s, see [Ins, 2008]. Such large changes in fertility also lead to large changes in family structures. Since the Costa Rican population is comparably young, the rapidly decreasing number of children was not compensated by an increase in the number of dependent elderly parents, which means that the overall dependency ratio has gone down. Also, the drop in the number of young children lead to a larger percentage of childless households and of households with only adult children, who potentially contribute to household income. We explore the effects of these changes in family structures on the female income distribution, in contrast to other studies of the income distribution in Costa Rica which have emphasized the role of education and occupational choices or social class, see [EstadodelaNación, 2009].

The data used in this chapter come from the yearly Costa Rican Multiple-Purpose House-

Figure 3.1: Fraction of births from unmarried mothers and of births without a registered father



Notes: This figure shows the fraction of births from unmarried mothers and of births without a registered father, calculated using the Birth Database of the Costa Rican Institute of Census and Statistics

hold Surveys (Encuesta de Hogares de Propósitos Múltiples). These surveys were conducted from 1993 to 2009 by the Costa Rican National Institute of Census and Statistics (Instituto Nacional de Estadísticas y Censos - INEC). Each yearly survey comprises observations between July of the previous year and July of the year it is published. Since the Fatherhood Responsibility Law was passed in April 2001, the Surveys of 2000 and 2001 (with its July 2001 cutoff) are the last pre-Law surveys.

These income surveys distinguish three sources of income for each individual: the main job, a potential secondary job and other sources of income. Based on these individual incomes, a household income is calculated. While the dataset also reports whether a household receives non-monetary income such as housing or food, we do not include such non-monetary income in our definition of household income. We impute a zero income for all individuals with missing income observations. Income observations are missing in roughly 5% of households in the sample. We calculate household equivalent income by dividing total household income by the "Oxford equivalence scale", as defined in [OECD, 2008]. This scale assigns a value of one to the first adult of the household, a value of 0.7 to all further adults, and a value of 0.5 to all children (household members aged less than 15 years). Costa Rica had very high inflation rates over the period under consideration (more than 10% yearly on average). We discount equivalent incomes using the Central Bank of Costa Rica official deflators to obtain a "real equivalent income" variable for estimations. Every individual in the survey is assigned their households real equivalent income as our main outcome variable of interest.

A major practical challenge in our analysis was the categorization of households according to their composition. In order to be useful for our purposes, such a categorization has to distinguish households according to the main determinants of equivalent income - the number of (potential) wage earners, the number of dependents, and the number of potential caretakers of dependents. The categories have to be mutually exclusive and cover all households. Given the astonishing diversity in living arrangements, this turned out to be a larger difficulty than might be expected. Finally, in order to avoid estimation problems (non-overlapping support), the number of categories must not be too large. Based on these considerations, we constructed the following categorization of households:

1. Couple with at least one dependent (mostly own children), no other working-age adults or adult children.

2. Couple with at least one dependent and at least one additional working-age adult (mostly own adult children).

3. Couple with at least one adult child and no dependents, sometimes other working-age adults.

4. Single parent as head of household with at least one extra working-age adult (mostly own adult children), possibly small children or dependents.

5. Single or Couple with no children and no dependents, sometimes there are other working-age adults.

6. Single parent, with at least one dependent, no other working-age adults or adult children

Our population of interest are working-age women (15 to 65 years old). This includes the population of reproductive-age women (15 to 49 years old), and in particular women who are less than 28 years old, who presumably were most affected by the Responsible Paternity Law, and who show the largest drop in births following the introduction of the law. The dataset includes variables on rural/urban area, region of Costa Rica (there are six major regions), age and level of education. These are the controls we use in this chapter. The survey data also include expansion factors, depending on the local district of the household, to adjust for sampling rates and response rates. We use these expansion factors throughout.

3.3 Identification and estimation

Suppose we observe repeated cross-sections with i.i.d. draws from the time t distributions P^t of the variables (Y, S, X), where X denotes covariates such as age, education, and location and S is family status (household composition). The variable Y denotes equivalent real income, that is, household income adjusted for household size and for inflation. We are interested in isolating the effect of historical changes in family structures on the distribution of female (equivalent) incomes Y, $P(Y)$, or statistics thereof, $\nu(P(Y))$. Possible choices for ν include the mean, the variance, the share below the poverty line, quantiles or the Gini coefficient.

Let $P^1(Y, S, X)$ in particular denote the joint distribution of (Y, S, X) in period 1 (e.g., 2008/09), and $P^0(Y, S, X)$ the corresponding distribution in period 0 (e.g., 1993/94). Our goal is to identify the counterfactual distribution P^* of Y in which the effect of changing family structures is "undone", while holding constant the current (period 1) distribution of covariates X as well as the distribution of income Y given X and family status S. The change from P^* to P^1 will be interpreted as the causal effect of changing family structures on the female income distribution. Formally, define P^* as

$$P^*(Y) := \int_{X,S} P^1(Y|X,S)P^0(S|X)P^1(X)dSdX. \tag{3.1}$$

This counterfactual distribution is constructed similarly to the counterfactual changes in the wage distribution of the united states, ascribed to changes in unionization and the minimum wage etc., which were analyzed in [DiNardo *et al.*, 1996]. This counterfactual distribution can be interpreted causally under an assumption of conditional independence. Denote $Y(S, X)$ the potential equivalent income of a woman with family status S and exogenous covariates X. Stability of $P(Y|X, S)$, under changes of family structures $P(S|X)$, is implied by

$$Y(s, X) \perp S|X \; \forall \; s, \tag{3.2}$$

where this conditional independence is assumed to hold in time periods 0 and 1. This assumption states that there is no self selection into family status correlated with potential income, conditional on the covariates X. This assumption is not unproblematic, but reasonably credible with a rich set of covariates, as we have at our disposition. We can rewrite the distribution P^* as

$$P^*(Y \leq y) = E^1\left[\mathbf{1}(Y \leq y)\theta^*\right], \tag{3.3}$$

where

$$\theta^* := \frac{P^0(S|X)}{P^1(S|X)}. \tag{3.4}$$

Equation 3.3 states that P^* is a reweighted version of the current distribution, P^1. Any counterfactual distributional characteristic ν of P^* can be estimated based on estimates of P^*, as in [DiNardo *et al.*, 1996]. This requires estimation of the ratio (3.4).

Alternatively, assume for a moment that ν can be written as the expectation of a function

f of Y, $\nu = E[f(Y)]$. Then the counterfactual ν^* can be obtained from

$$\nu^1 - \nu^* = \int\int E^1[f(Y)|X,S][P^1(S|X) - P^0(S|X)]P^1(X)dSdX. \tag{3.5}$$

In general, ν will not have this linear form but can be approximated by a linear first order expansion around P^1. This idea underlies the influence-function regression approach proposed in [Firpo *et al.*, 2009]. It requires estimation of the difference $P^1(S|X) - P^0(S|X)$ and of $E^1[f(Y)|X, S = 1]$.

Corresponding to these two representations of the counterfactual ν^*, we consider two estimation approaches; reweighting observations and influence-function regression. The reweighting approach estimates the weight θ^* and calculates counterfactual ν from the reweighted distribution P^*.

The influence-function regression approach is based on the first order approximation of ν, as a function of P, around P^1:

$$\nu(P) = \nu(P^1) + \int IF(y;\nu,P^1)d(P - P^1)(y) + R, \tag{3.6}$$

where IF is the influence function of the parameter ν at P^1 and R is a second order remainder term. Ignoring the remainder, this representation of ν has the linear form required for the use of the representation (3.5), i.e.,

$$\nu(P) \approx E[\nu(P^1) + IF(Y;\nu,P^1)].$$

We can hence calculate first order approximations to the counterfactual ν based on estimates of $P^1(S|X) - P^0(S|X)$ and of $E^1[IF|X, S = 1]$. For details, the reader is referred to [Firpo *et al.*, 2009].

For either approach, we need to estimate the ratio or difference between $P^1(S|X)$ and $P^0(S|X)$, corresponding to the change in family structures within demographic groups defined by X. We use a multinomial logit model for the distribution of S given X, with parameters changing over time:

$$P^t(S = s|X) = \frac{\exp(X\cdot\beta^{s,t})}{\sum_{s'}\exp(X\cdot\beta^{s',t})}. \tag{3.7}$$

Based on estimates of the parameters $\beta^{s,t}$, we can calculate the weights θ^*, as

$$\theta^* = \frac{P^0(S|X)}{P^1(S|X)} = \frac{\exp(X\cdot\beta^{S,0})}{\sum_{s'}\exp(X\cdot\beta^{s',0})}\cdot\frac{\sum_{s'}\exp(X\cdot\beta^{s',1})}{\exp(X\cdot\beta^{S,1})}. \tag{3.8}$$

Similarly

$$P^1(S|X) - P^0(S|X) = \frac{\exp(X \cdot \beta^{S,1})}{\sum_{s'} \exp(X \cdot \beta^{s',1})} - \frac{\exp(X \cdot \beta^{S,0})}{\sum_{s'} \exp(X \cdot \beta^{s',0})}. \tag{3.9}$$

For the influence-function regression approach, we also need estimates of $E^1[IF|X, S = 1]$. We run the following regression, with full interactions between X and S, for the sample of household observed in 2008/09,

$$IF = (X \times S) \cdot \beta^{IF,1} + \epsilon, \tag{3.10}$$

and assume $E^1[IF|X, S] = (X \times S) \cdot \beta^{IF,1}$.

Note that, while the above approach is parametric, identification does not rely on the parametric choices: both the logit specification for $P(S|X)$ and the linear specification for $E[IF|X, S]$ are in fact "nonparametric" if we allow for sufficiently rich interactions and powers between the components of X and S. Furthermore, following the arguments of [Newey, 1994], the choice of nonparametric estimator for this "first stage", if it is consistent, does not affect the asymptotic variance of root-n estimable parameters. This covers all our examples for ν, except for the counterfactual densities. Confidence sets for all estimators are obtained by bootstrapping the entire procedure.

3.4 Results

To give a preliminary idea about the relationship between family categories and poverty (income), table 3.2 shows regressions of poverty and of log real equivalent income on the family categories, with and without controls. The omitted category in these regressions is the category of couples with dependents but no adult children. The signs and magnitudes of the coefficients are as we would expect. In particular, households consisting of a single head and dependents have significantly higher poverty levels and lower income relative to the baseline, while couples with adult children and no dependents are less likely to be poor and are richer on average. The coefficients on the controls are also as expected; e.g., having tertiary education significantly reduces poverty and increases income relative to other education levels.

Let us now turn to a discussion of our main empirical results. Figure 3.2 and all following figures are constructed as follows. The left column shows counterfactual changes calculated using the reweighting method described in section 3.3, the right column shows the analogous estimates calculated using influence function regression. The top row shows counterfactual changes over the entire period 1993/94 -2008/09, the middle row over the period 1993/94-2000/01, and the bottom row over the period 2000/01-2008/09. Every graph in these figures

Table 3.2: The impact of family structures on poverty rates and real equivalent income

	(1) Poverty	(2) Poverty	(3) Income	(4) Income
Couple w/adult children, dependents	0.037	0.032	-0.105	-0.074
	(0.008)	(0.007)	(0.018)	(0.015)
Couple w/adult children, no dependents	-0.043	-0.012	0.242	0.081
	(0.007)	(0.007)	(0.018)	(0.015)
Single w/adult children	0.039	0.056	-0.101	-0.186
	(0.008)	(0.008)	(0.018)	(0.016)
Childless households	-0.019	0.004	0.399	0.218
	(0.009)	(0.009)	(0.024)	(0.020)
Single w/dependents	0.119	0.131	-0.223	-0.305
	(0.017)	(0.016)	(0.042)	(0.033)
Age		-0.004		0.010
		(0.001)		(0.002)
Age^2		4.4×10^{-5}		4.3×10^{-5}
		(1.4×10^{-5})		(2.9×10^{-5})
Primary Education		-0.145		0.253
		(0.017)		(0.031)
Secondary Education		-0.256		0.650
		(0.017)		(0.031)
Tertiary Education		-0.326		1.340
		(0.018)		(0.033)
Rural		0.065		-0.223
		(0.005)		(0.010)
Constant	0.187	0.440	11.521	10.836
	(0.005)	(0.026)	(0.014)	(0.051)
N	32171	32171	31300	31300
Region controls	No	Yes	No	Yes

Notes: This table shows regressions of poverty dummies and of real equivalent income on dummies for different types of household composition. The omitted household composition category are couples with dependents and no other adult household members. The regressions in columns (2) and (4) control additionally for covariates. Standard errors are in parentheses, the regressions are calculated for the 2008/09 waves of the Costa Rican Household Survey.

plots the function

$$\frac{Q^1(.) - Q^*(.)}{Q^1(.)},$$

where $Q^1(q)$ is the q th quantile of the income distribution at the end of the period under consideration, and $Q^*(q)$ is the corresponding quantile of the counterfactual distribution which constructed as discussed in section 3.3. This function gives the percentage change in income levels at different quantiles of the distribution induced by changing family structures. For instance, for q=0.5, the graphs show the percentage change in the median income of working age women over the period under consideration due to changing family structures. The figures also plot pointwise 95% confidence bands obtained by bootstrapping. Panels 3.2 and 3.3 show these graphs for all working age women and for rural women only, the panels in appendix 3.6 show similar graphs for urban women, for women below 23 and women below 28, as well as for women with and without more than primary education.

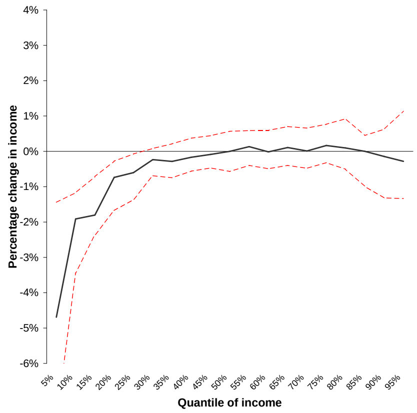
Several messages emerge from these graphs: (i) Changing family structures had an inequality-increasing effect for working age Costa Rican women over the period 1993-2009, leaving median incomes unchanged. (ii) This inequality increasing effect came in particular from changes in the period after 2000/01, i.e., after the Responsible Paternity Law was introduced. In contrast, the changes in the period prior to 2000/01 had a negative effect on equivalent incomes across the entire income distribution. (iii) The inequality increasing effect was strongest among rural women. (iv) The changing family structures after 2000/01 had a particularly positive effect for younger women, except for the lowest income groups.

One possible interpretation of these results is that the general trend towards single parenthood adversely affected female equivalent incomes across the distribution prior to 2000/01, but that the Responsible Paternity Law was effective in stopping this trend. We would expect the law to affect, in particular, younger women, who are the most likely to bear children - consistently with the results in figure 3.5. It appears, however, that the law was not effective in reversing trends at the bottom end of the income distribution.

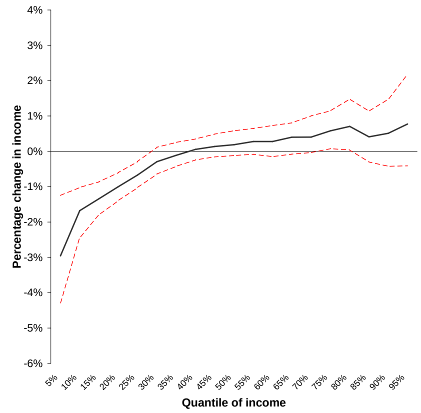
Methodologically, another interesting feature of our results lies in the comparison of the estimates obtained using reweighting with those obtained using influence function regression. Both methods yield very similar point estimates, which increases our confidence in the robustness of the results. It appears that the method based on influence function regression generally leads to tighter confidence bands and produces smoother estimates across quantiles. This difference in finite sample performance stands in contrast to the asymptotic equivalence, in the case of linear ν, suggested by the arguments of [Newey, 1994]. On the other hand, the use of a first order approximation in the influence function regression approach leads to a bias since higher order terms are ignored. In our case, given the small size of the counterfactual changes in the distribution, this bias is likely to be negligible, though. In combination, these observations suggest a possible bias-variance trade-off in the use of reweighting versus influence function regression.

Figure 3.2: Counterfactual changes of the income distribution of all Costa Rican women

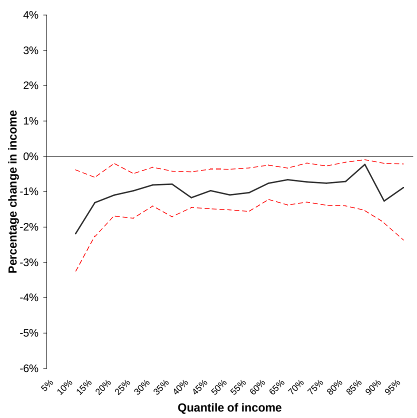
(a) Reweighting, 1993/94-2008/09



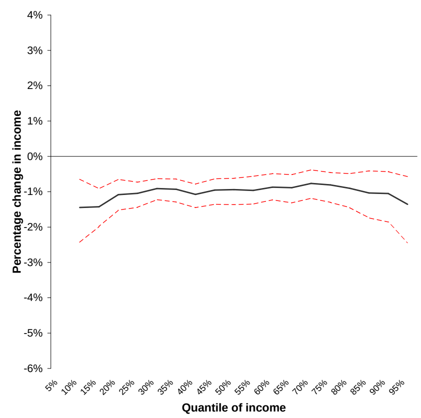
(b) Regression, 1993/94-2008/09



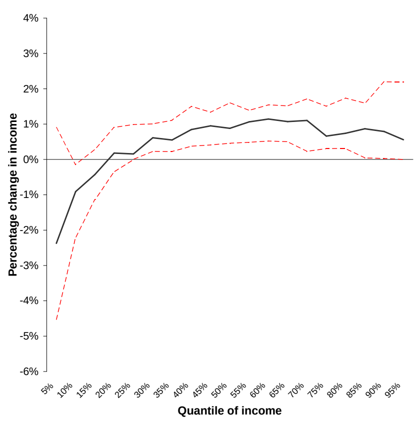
(c) Reweighting, 1993/94-2000/01



(d) Regression, 1993/94-2000/01



(e) Reweighting, 2000/01-2008/09



(f) Regression, 2000/01-2008/09

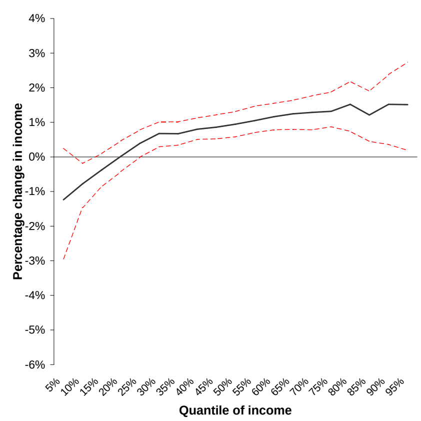
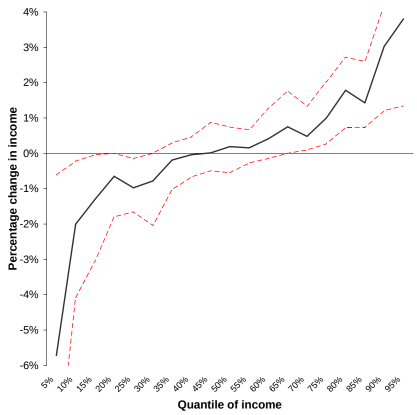
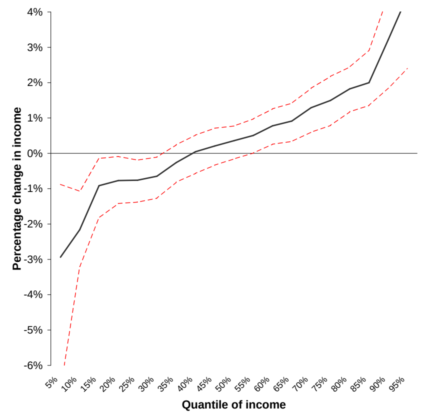


Figure 3.3: Counterfactual changes of the income distribution of rural women

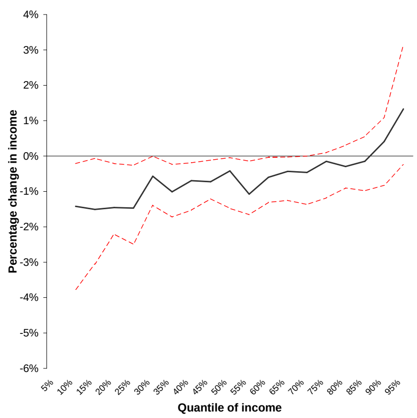
(a) Reweighting, 1993/94-2008/09



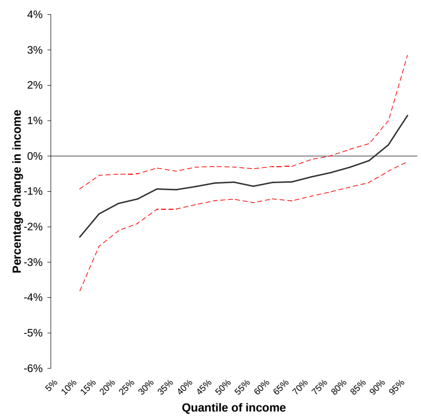
(b) Regression, 1993/94-2008/09



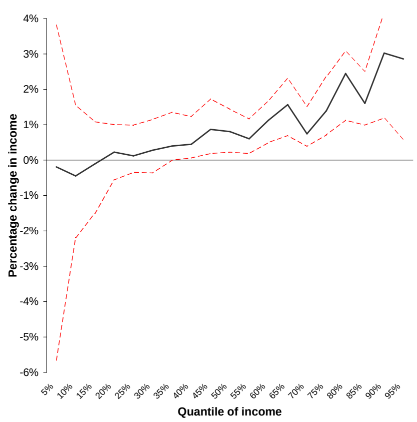
(c) Reweighting, 1993/94-2000/01



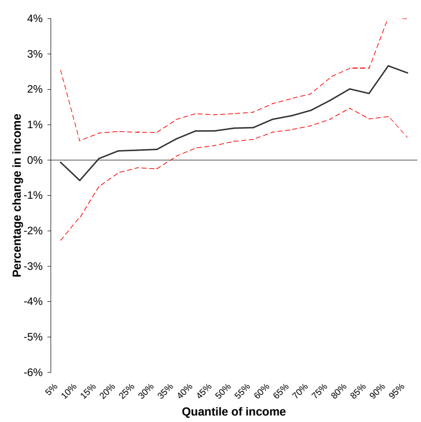
(d) Regression, 1993/94-2000/01



(e) Reweighting, 2000/01-2008/09



(f) Regression, 2000/01-2008/09



3.5 Conclusion

In this chapter, we have studied the effect of changing family structures on the female income distribution in Costa Rica between 1993 and 2009. We find that changing family structures had an inequality-increasing effect for working age Costa Rican women over this period, in particular due to changes in the period after 2000/01, i.e., after the Responsible Paternity Law was introduced. Changes in the period prior to 2000/01 had a negative effect on equivalent incomes across the entire income distribution. The inequality increasing effect was particularly strong among rural women, while the change family structures after 2000/01 had a particularly positive effect for younger women, except for the lowest income groups.

Our results illustrate the role of family structures in the determination of the distribution and level of incomes. These results imply, in particular, that various policies affecting family structures also have an important distributional effect. Such policies include policies related to family planning, the public provision of child care, retirement systems for the elderly, and marriage, divorce, and child support legislation. It is an interesting area for future research to explore the effects of such policies on the income distribution.

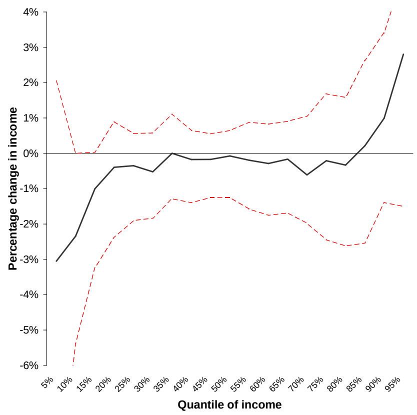
In a comparison of the reweighting and influence function regression methods proposed in [DiNardo *et al.*, 1996] and [Firpo *et al.*, 2009], we find that the influence-function technique gives smoother results with narrower confidence bands. This is contrasted by a potential bias in influence function regression, since it is based upon a first-order approximation, suggesting a potential bias-variance trade-off between the two methods.

A valuable extension of the research presented in this chapter might be a combination of the distributional decomposition methods discussed with credible estimates of the structural relationship between child support legislation and household composition. Such a combination would allow to assess the distributional impact of the Responsible Paternity law, and open interesting methodological perspectives for distributional policy evaluation.

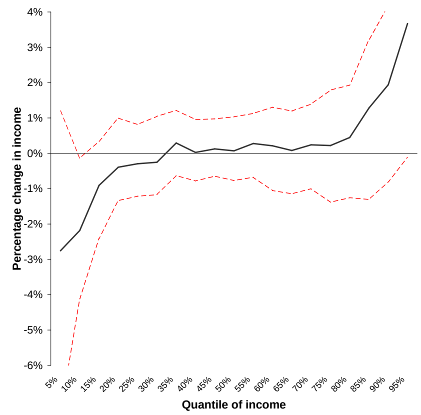
3.6 Further results

Figure 3.4: Counterfactual changes of the income distribution of women under 23 years old

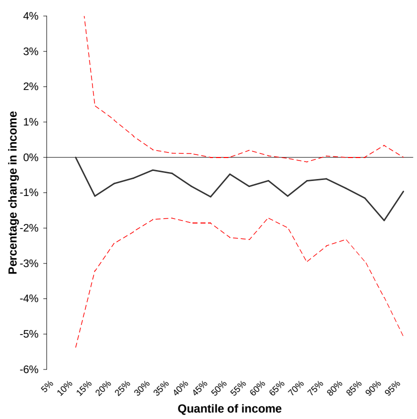
(a) Reweighting, 1993/94-2008/09



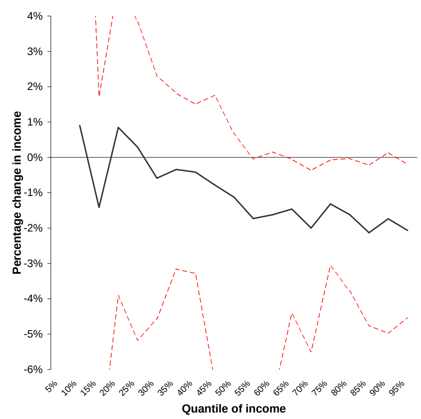
(b) Regression, 1993/94-2008/09



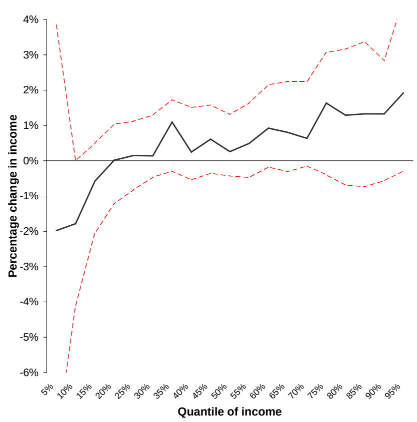
(c) Reweighting, 1993/94-2000/01



(d) Regression, 1993/94-2000/01

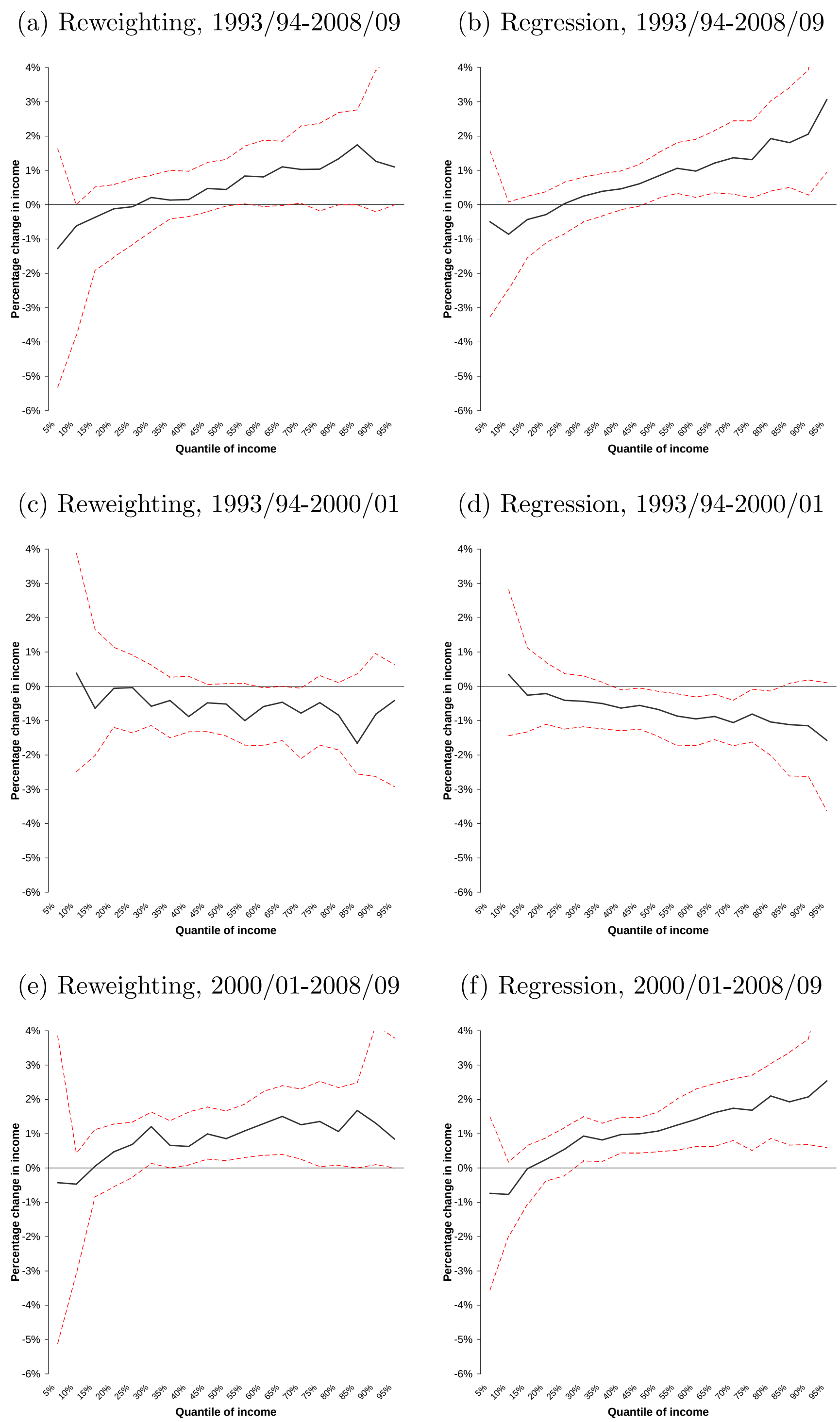


(e) Reweighting, 2000/01-2008/09



(f) Regression, 2000/01-2008/09

Figure 3.5: Counterfactual changes of the income distribution of women under 28 years old



(a) Reweighting, 1993/94-2008/09

(b) Regression, 1993/94-2008/09

(c) Reweighting, 1993/94-2000/01

(d) Regression, 1993/94-2000/01

(e) Reweighting, 2000/01-2008/09

(f) Regression, 2000/01-2008/09

Figure 3.6: Counterfactual changes of the income distribution of women with no secondary education

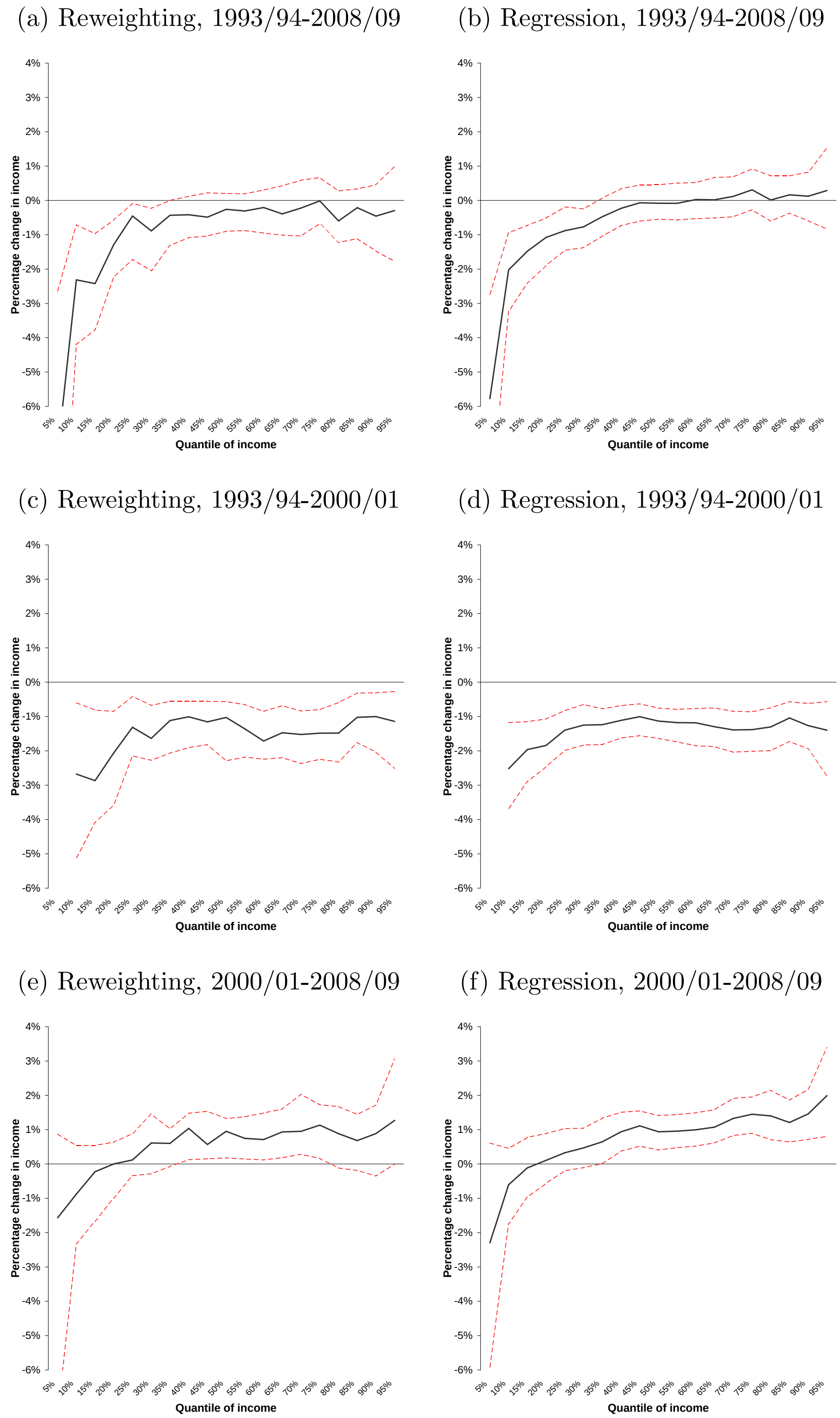
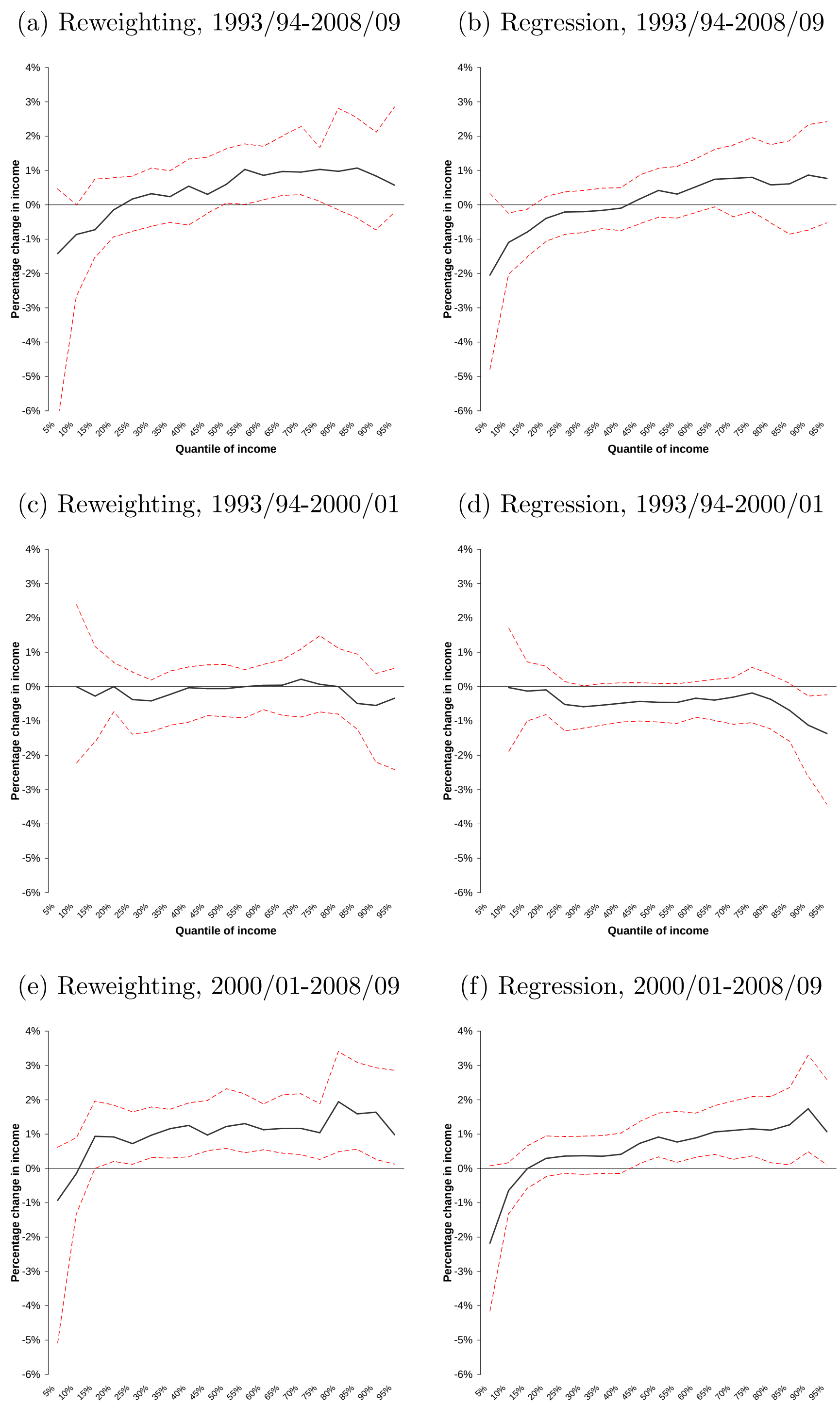


Figure 3.7: Counterfactual changes of the income distribution of women with at least secondary education



(a) Reweighting, 1993/94-2008/09

(b) Regression, 1993/94-2008/09

(c) Reweighting, 1993/94-2000/01

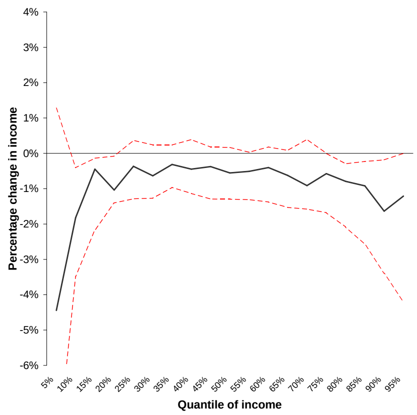
(d) Regression, 1993/94-2000/01

(e) Reweighting, 2000/01-2008/09

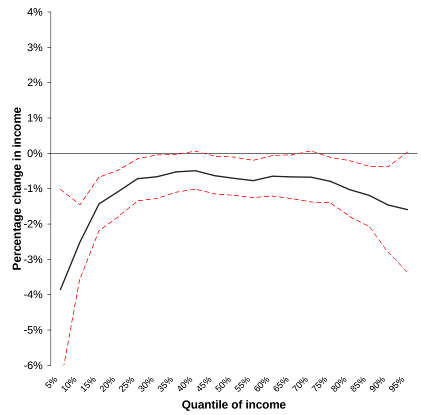
(f) Regression, 2000/01-2008/09

Figure 3.8: Counterfactual changes of the income distribution of urban women

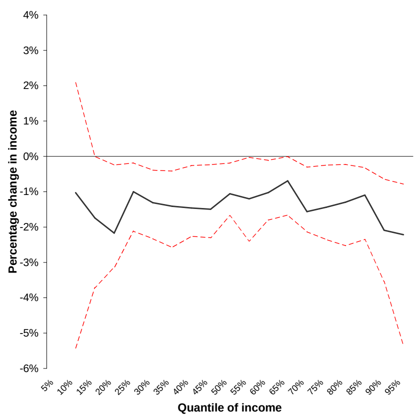
(a) Reweighting, 1993/94-2008/09



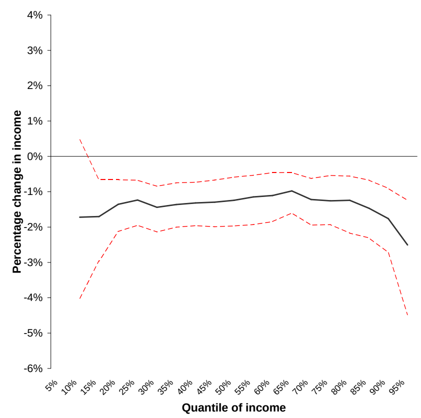
(b) Regression, 1993/94-2008/09



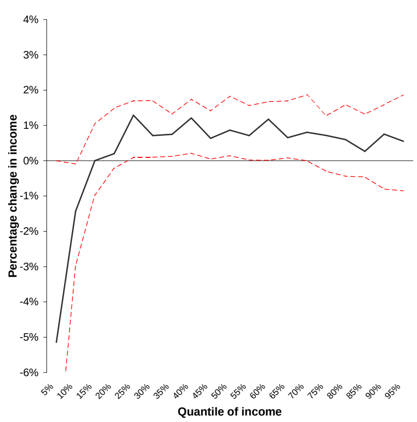
(c) Reweighting, 1993/94-2000/01



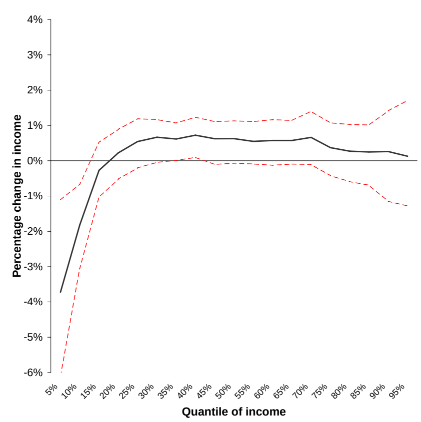
(d) Regression, 1993/94-2000/01



(e) Reweighting, 2000/01-2008/09



(f) Regression, 2000/01-2008/09



Chapter 4

After Hurricane Mitch: the effects of Nicaraguan immigration in Costa Rica

4.1 Introduction

Today, nearly half a million Nicaraguans live in Costa Rica, in addition to about four million Costa Ricans. Most Nicaraguans arrived in the past 20 years. The size of the working-age immigrant population makes their impact on native Costa Rican labor market outcomes a salient question. Evaluating the effect of immigration is difficult due to the endogeneity of migration decisions. If Nicaraguans migrate into Costa Rica only when jobs are easily available, estimates of their impact will be biased towards a positive correlation between immigration and native labor market outcomes. In the literature, various approaches have been proposed to address this difficulty; [Card, 1990] suggests finding a natural experiment that corresponds to an exogenous increase in the supply of immigrants into the labor force of interest.

In October 1998, Hurricane Mitch caused floods and mudslides in Honduras and Nicaragua, killing thousands and leaving hundreds of thousands homeless. After the hurricane, crop losses forced many Nicaraguans to migrate to Costa Rica, which had been spared that kind of damage. The Hurricane Mitch immigrants chose to disproportionately locate in those Costa Rican regions near the border with Nicaragua and attempted to find work in agriculture and construction, encouraged by the presence of previous Nicaraguan immigrants in those regions and economic sectors.

Empirical estimation of the effects of immigration is necessary because economic theory doesn't provide us with an unique prediction. Neoclassical competitive labor market models predict immigration, which is essentially a labor supply increase, to have a smooth negative impact on native wages. Other models predict small to no changes in native wages. such as the [Harris and Todaro, 1970] model where rigid labor contracts imply that immigrants are either unemployed or they get a formal contract at the same wages as natives. Similarly, the [Bencivenga and Smith, 1997] model of information asymmetry suggests that migration flows continue until the expected[1]) wage in the destination country is equalized to the certain wage in the source country[2].

The recent literature on estimating the effects of immigration and addressing its endogeneity has two major strands. First, [Card, 1990] emphasizes finding a natural experiment, such as the Mariel Boatlift Cuban refugees in Miami, which is an exogenous labor supply shock to the destination labor market and then examining the effects on labor outcomes. Second, [Borjas, 2003] suggests looking at nationwide education-experience groups who are unevenly affected by immigration, to capture potential extra-regional effects. An example of Borjas' critique is that Miami residents might have left that city due to competition from Mariel Cubans, which would depress wages elsewhere and estimates of Miami wage effects would be biased upwards[3].

This research seeks to contribute to Card's line of research by exploiting an exogenous event (Hurricane Mitch) to identify the effects of migration from one developing country to another (South-South migration). Most immigration studies are either focused on the United States or on European natural experiments, such as [Hunt, 1992] who finds little impact of 1960s Algerian repatriates in France. Or [Carrington and de Lima, 1996], who look at Portugal repatriates from Angola and Mozambique in the 1970s and find a negative effect in Portugal but are unable to isolate it from European-wide downturn effects. Yet another example is immigration into Israel from the Soviet Union in the early 1990s, discussed by [Friedberg, 2001] who finds that native Israelis were not affected by it. However, in the European examples, the repatriates had a similar set of skills to the natives, whereas in both immigration into the United States and Nicaraguan immigration into Costa Rica the immigrants have lower skills, on average, than the natives.

The main finding of this chapter is that the hypothesis of immigration after Hurricane Mitch having no immediate impact on native Costa Rican incomes and employment cannot be rejected, except for a negative impact on incomes among men with only primary education in the border regions, but even this finding is small relative to the variation in incomes observed before and after the hurricane. This suggests that a middle-income country like Costa Rica can absorb an exogenous labor supply shock without much of an impact on incomes or employment among natives. These results also fit with the small effects found in the rest of the natural-experiment immigration literature. A complication is that, partly due to the Mitch emergency in the rest of Central America, in early 1999 the Costa Rican government granted an amnesty to all Central American migrants that had arrived before November 1998. However, under the assumption that the amnesty made it easier for em-

[1]Destination country wages are a lottery of a probability p of getting a native-wage job and a probability (1-p) of being unemployed

[2]For a more complete description of these models, the reader is referred to [Bardhan and Udry, 1999].

[3][Card, 2000] addresses Borjas' critique by accounting for intercity mobility and still finds small to no effects of immigration on wages.

ployers to hire Nicaraguans, this would bias downwards estimates of the impact of Mitch immigrants in native wages and employment and strengthen the finding that there was no effect.

The rest of the chapter is structured as follows: In section 4.2 the patterns of Nicaraguan migration into Costa Rica and the impact of Hurricane Mitch are discussed. Then section 4.4 describes the differences-in-differences estimation method which follows [Card, 1990]. The data sources and database construction from the Household Surveys can be found in section 4.3. The simple differences between the labor market outcomes of the groups of interest and the differences-in-differences estimation results are in section 4.5. Section 4.6 concludes.

4.2 Institutional background

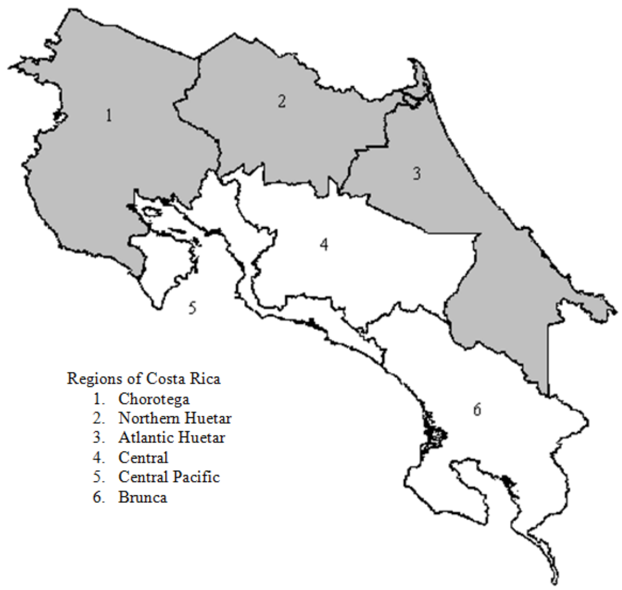
Costa Rica is a middle income country, while Nicaragua, its northern neighbor, is the second-poorest country in the Americas. According to the Human Development Report by [UNDP, 2010], in 2008 the average Costa Rican had an income 4.23 times larger than the average Nicaraguan. For comparison, the average citizen of the United States had an income 3.37 times larger than the average Mexican[4]. The border between Costa Rica and Nicaragua is unpopulated forest, porous and hard to police. Nicaraguans speak Spanish and share many cultural patterns with Costa Ricans, easing their integration and making it attractive to migrate to Costa Rica even if other options such as the United States offer much higher potential incomes. The 1980s civil war in Nicaragua weakened its economy, so when peace returned in the 1990s, many Nicaraguans continued migrating towards the United States as they had during the civil war, but a new migration flow started towards Costa Rica.

The Costa Rican Census of 2000 asked the year of entry of foreign residents and confirms that the most Nicaraguans entered during the 1990s, at a rate of more than 15 thousand per year, with a 1998 peak, when more than 20 thousand Nicaraguans entered the country, see [Marquette, 2006]. The regions of Costa Rica that share a border with Nicaragua had 34% of Nicaraguan immigrants, compared with 22% of the native Costa Rican population in the Census. These border regions are the Chorotega, Northern Huetar and Atlantic Huetar, highlighted in Figure 4.1. The Census also showed that Nicaraguans are concentrated in agriculture, construction and domestic services, because Nicaraguans represented roughly 6% of the total population, but 15% of the population working in those sectors. Not coincidentally, the border regions also happen to have strong agriculture and construction sectors.

This Census question is problematic because more recent Nicaraguan immigrants were less likely to be counted in the Census since earlier arrivals had benefited from the 1999 amnesty and would have felt safer answering the Census interviewers. Fortunately, there are

[4] The income measure is GNI (Gross National Income) per capita, adjusted for PPP (Purchasing Power Parity). 2008 PPP GNI per capita for each country was: United States, $47,094; Mexico, $13,971; Costa Rica, $10,870 and Nicaragua, $2,567.

Figure 4.1: Northern Border Regions



other measures that show a Mitch-related immigrant spike in late 1998 and early 1999. The first measure is the number of registered births in Costa Rica that had a Nicaraguan mother. As can be seen in Table 4.1, births from Nicaraguan mothers grew 17% in 1999[5]. The second measure is self-identified Nicaraguans observed in Household Surveys (adjusted by the survey expansion factor). Since the Household Survey is completed every July, any Nicaraguan immigrant spike should appear in the July 1999 survey. Table 4.1 shows that observations of Nicaraguans in the Survey grew 70% in 1999. This last number could overstate the true increase in immigration, but from both proxy measures and the Census question, it is possible to say with some confidence that Mitch an increase in the size of the flow of migrants.

Costa Rican labor demand is unlikely to have been affected by Hurricane Mitch because Costa Rica was not hit directly by it. It is also unlikely that Costa Rica suffered indirect price or wage shocks even when taking into account reduced production in Nicaragua, because in most Costa Rican export products such as bananas and coffee, Nicaragua is too small an exporter for its reduction in production to affect world markets. Also, there is little competition between staple goods in Nicaragua and Costa Rica.

Table 4.1 shows that the immigrant population established in Costa Rica was large before Mitch (e.g. in 1997 almost 10% of births were from Nicaraguan mothers), so most of the Mitch migrants tried to use the existing social networks and located in similar jobs and

[5]It is the largest yearly increase in births from Nicaraguan mothers in the entire 1986-2009 period of the births database. Also, monthly data shows an increase of the percentage of births from Nicaraguan mothers in November and December 1998 relative to the rest of 1998 (Hurricane Mitch hit in October 1998)

Table 4.1: Proxies for measuring the growth of Nicaraguan immigration

	Nicaraguan Births			Nicaraguans Surveyed		
	Total	% of all	Growth	Total	% of all	Growth
1997	7654	9.8%	11%	75,490	2.3%	NA
1998	8319	10.8%	9%	78,487	2.3%	4%
1999	9757	12.4%	17%	133,548	3.9%	70%
2000	10351	13.2%	6%	157,401	4.1%	18%
2001	10391	13.7%	0%	165,617	4.2%	5%

Notes: Source of data is Births Database and Costa Rican Household Surveys, INEC. Nicaraguan Births means Costa Rican children born from a Nicaraguan mother. Nicaraguans Surveyed means the number of self-identified Nicaraguans surveyed in the Costa Rican Household Surveys, adjusted for the expansion factor provided by the survey data.

regions as previous migrants, independently of the labor demand at the time. ”Treatment” regions and economic sectors can be defined as those that received disproportionate numbers of Nicaraguan immigrants. In the case of the regions, these are those that share a border with Nicaragua (Chorotega, Northern Huetar and Atlantic Huetar) while all other regions are controls. The treatment economic sectors are agriculture and construction and all other sectors are controls.

To see why the native Costa Rican population in those regions and sectors is the one affected or ”treated” by Hurricane Mitch, consider the yearly Household Survey population proportions in Table 4.2. Household Surveys understate the number of Nicaraguans relative to the Census. However, the treatment/control regional ratio in the Household Surveys was 165% and the same ratio was 154% in the Census. This implies that even if the absolute numbers are biased downwards, relative ratios in the Household Surveys can be trusted. Both the regional and sectoral treatment/control ratios are over 1.5 for all years and they both increased in 1999. Hence, both the treatment regions and treatment sectors already had a higher initial proportion of Nicaraguans relative to the control regions and sectors, but this imbalance became even larger immediately after Mitch.

Previous research on Hurricane Mitch-related immigration was done by [Kugler and Yuksel, 2008], but focused on Honduran and Nicaraguan migration into the United States, especially into southern U.S. states and California. Due to Mitch, the State Department gave these migrants a special Temporary Protected Status in late 1998 for 18 months, which made it easier for these migrant to find jobs. However, the researchers find no important wage effects and no out-migration effects (to address Borjas' critique). The only effect they find is on employment of other Latin migrants.

Nicaraguan migration into Costa Rica has been explored in several sociological and psy-

Table 4.2: Sectoral and regional differences of the Nicaraguan presence in Costa Rica

Panel A: Nicaraguans by Regions					
	1997	1998	1999	2000	2001
Control					
Population	51,809	50,634	81,501	106,124	106,666
% of total pop.	2.0%	1.9%	3.1%	3.6%	3.5%
Treatment					
Population	23,681	27,853	52,047	51,277	58,951
% of total pop.	3.3%	3.8%	7.0%	5.9%	6.6%
Ratio Treat/Control	1.65	1.97	2.28	1.65	1.88

Panel B: Nicaraguans by Economic Sectors				
	1997	1998	1999	2000
Control				
Population	69,794	72,612	121,989	137,668
% of total pop.	2.2%	2.3%	3.8%	3.8%
Treatment				
Population	5,696	5,875	11,559	19,733
% of total pop.	3.6%	3.6%	7.0%	11.0%
Ratio Treat/Control	1.59	1.56	1.87	2.89

Notes Source of data is Costa Rican Household Surveys, INEC. Treatment regions share the Nicaraguan Border (Chorotega, Northern and Atlantic Huetar), Controls are all other regions. Treatment economic sectors are agriculture and construction; Controls are all other economic sectors. The Ratio Treat/Control shows how much more likely is it to find a Nicaraguan in a random population sample of the treatment group relative to a same-sized sample of the control group.

chological studies[6], but to my knowledge, there are only two quantitative economic impact analysis. First, [Castillo *et al.*, 2010] emphasize the role of substitutability of immigrant and native skills. As long as immigrants' skills are imperfect substitutes for natives' skills (for example, because of language barriers), most additional immigration will impact previous immigrants, not natives. They find that skills of immigrants in Costa Rica are much closer to those of natives than in the United States. Even so, they find little evidence of an impact of immigration on native labor market outcomes in both countries. Second, [Gindling, 2008] finds that Nicaraguan immigration in selected industries didn't seem to impact the wages in those industries negatively. However, he is concerned that Nicaraguans are being drawn into fast-growing industries, so he employs the [Borjas, 2003] skill-cells method. He finds that only low-skilled females show skill-substitutability with Nicaraguans, while high-skilled females have skill-complementarity.

4.3 Data description

For this chapter, there were three sources of data, all of them developed by the National Institute of Census and Statistics (Instituto Nacional de Estadísticas y Censos - INEC). One is the Costa Rican Census of the year 2000, which has gender, age, nationality, year of arrival (if a foreigner), education level, economic activity sector and geographical information. Since the Census attempted to count the entire population residing in Costa Rica, it is one of the most reliable sources to find out which regions and economic sectors have a disproportionate number of Nicaraguans relative to the rest of the country, as described in Section 4.2. The 2000 Census counted 226,374 Nicaraguan-born residents in Costa Rica, that is, 5.9% of a total population of 3,810,179.

However, the second source of information, the birth database also collected by INEC reveals that the 225,000 Nicaraguans found by the Census is likely to be an understatement of the true number. [Chen *et al.*, 2001] uses a Reproductive Data Survey to estimate the average fertility of Nicaraguan women that reside in Costa Rica. Then it is possible to extrapolate the total number of Nicaraguans in Costa Rica from the number of Costa Rican children born to Nicaraguan mothers[7]. With this method, [Chen *et al.*, 2001] estimated 315,000 Nicaraguans in mid-1998, which is just before Hurricane Mitch. Two years later, the growth in births from Nicaraguan mothers suggests a Nicaraguan population easily exceeding 350,000. The striking difference between this number and the Census one suggests a high percentage of Nicaraguan migrants in irregular migratory condition, especially considering that [Mora, 2004] finds that the 1999 amnesty covered only 150,000 Nicaraguans.

The third source of information used is the annual Household Surveys from 1995 to 2003, which have the same variables as the Census plus average monthly income data, except for

[6] [Marquette, 2006] is a good survey of those studies

[7] Any child born in Costa Rican soil receives citizenship automatically upon registration of the birth

the year of arrival question. The surveys have exact country of origin only for the period 1997-2001, other years it is possible to distinguish whether the individual surveyed is Costa Rican or a foreigner. A disadvantage of the Surveys is that they have even more under-sampling of Nicaraguans than the Census. This is clear from the fact that in 2000, the sample of Nicaraguans, when appropriately expanded using the expansion factor, is just 157,000, well below the 225,000 of the Census. As discussed in section 4.2, there is reason to believe that the under-sampling is proportionally constant across regions or economic activity sectors so that the samples are still useful to measure the impact of variations.

To estimate the impact of Hurricane Mitch, a multi-year database that covers the period 1995-2003 was built from the yearly Household Surveys, focusing only on working-age (15-64) native Costa Rican adults (all foreigners are excluded). Education levels were grouped into Primary (up to 6th grade completed), Secondary (up to 11th grade completed) and Tertiary (at least some education beyond Secondary). The treatment region dummy had a value of 1 if the individual lived in one of the Northern Border regions. The treatment economic sector dummy had a value of 1 if the individual worked in any of the sectors coded between 1000 and 1199 (Agriculture) or in the sector coded 5000 (Construction). The employment variable is a dummy of whether income was observed. Incomes well below the minimum wage (a sign of very sporadic or unstable employment) were coded as missing. Finally, the expansion factor[8] of the Survey was recoded into a probability weight of the observation through dividing the individual factor expansion by the sum of all expansion factors in the database.

An important limitation of the Costa Rican Household Surveys is that the definition of some covariates is not necessarily consistent over time (although changes are properly documented). For example, the economic sector classification was modified in 2001 and the new classification doesn't map cleanly into the old one. Also, while it is possible to separate Costa Ricans from foreigners in all years from 1995-2003, the question of specific foreign nationality was asked only for the years 1997-2001, which is the question that allows us to identify Nicaraguans. The main years of interest are 1998 and 1999, so these issues are not problematic for the purposes of measuring the immediate impact of migration due to Mitch.

4.4 Identification and estimation

The estimation approach requires the assumption that Nicaraguan immigrants after Mitch disproportionately located in specific economic sectors such as Construction and Agriculture, or specific regions such as those close to the border, during the year immediately after Mitch, that is, 1999. Let us call these the Treatment Groups (T) and they will be our treated group in 1999. The key assumption is that the rest of economic sectors had similar (common)

[8] This is the variable that summarizes how many people each survey observation represents in the entire population.

wage trends but suffered a much smaller immigration impact and so can be considered control or comparison groups (C). This will let us follow the natural experiment differences-in-differences approach of [Card, 1990].

Using [Angrist and Krueger, 1999] notation, think of a person f in a T (treated group) as having a pair of income outcomes: $Y_{0f} = X\beta + \epsilon_f$ if there had been no migration and $Y_{1f} = Y_{0f} + \delta$ which we observe. Then the outcome of interest is $E[Y_{0f}|g = T, t = 1999]$ that is, the counterfactual average income for that group if Mitch-caused immigration had never happened. To identify the causal effects, the differences-in-differences method requires a restriction on the conditional mean function. Concretely:

$$E[Y_{0f}|g, t] = \beta_t + \gamma_g \tag{4.1}$$

Where β_t is time fixed effects and γ_g is group fixed effects. The effect of Mitch-caused immigration is assumed to be:

$$E[Y_{1f}|g, t] = E[Y_{0f}|g, t] + \delta \tag{4.2}$$

Therefore, the wage outcome of individuals working on T or control groups can be written as:

$$Y_f = \beta_t + \gamma_g + \delta T_f + \epsilon_f, E[\epsilon_f|g, t] = 0 \tag{4.3}$$

Where T_f is a dummy that equals 1 if the individual was exposed to the Mitch immigrant impact in 1999. Then estimate:

$$E[Y_f|g = T, t = 99] - E[Y_f|g = C, t = 99] - E[Y_f|g = T, t = 98] - E[Y_f|g = C, t = 98] = \delta \tag{4.4}$$

[Angrist and Krueger, 1999] point out that one of the key identifying assumptions is that the interaction terms would have been zero in the absence of Hurricane Mitch. To test this assumption, income trends should be the same across economic sectors or geographical regions whether there are immigration shocks or not. Since the Household Surveys are cross sections, not panel data, the trends are tested using a vector of year dummies Z_f. α and δ are vectors of coefficients and the coefficient of interest is the δ for the year 1999.

$$Y_f = Z_f\alpha + (T_f * Z_f)\delta + \epsilon_f \tag{4.5}$$

It is possible to control for other individual characteristics X_f if it is believed that they might be correlated to T_f, conditional on group and year, since this would change our estimate of δ. For example, the groups are defined on geographical and economic activity criteria which are likely to be correlated to individual characteristics of age, level of education and sex.

$$Y_f = Z_f\alpha + (T_f * Z_f)\delta + X'_f\beta_0 + \epsilon_f \tag{4.6}$$

4.5 Results

As part of the difference-in-differences methodology, the time trends in both control and treatment groups should be the same. For this to hold,there should be very little variation over time in the simple-difference in employment and income between the groups. Specifically, the simple-difference should be steady until 1998, then there should be a large change between 1998 and 1999 (the effect of Mitch) and then a steady difference again. Figures 4.2 and 4.3 do not show this pattern. The simple differences sometimes show larger changes in the non-Mitch years, before and after, than in the change between 1998 and 1999 (after the black vertical line). These simple-difference figures are shown only for the total working-age native population(sex-education subgroups are not shown), but these figures already suggest very weak effects of Hurricane Mitch migration. It seems that any impact of Mitch is swamped by the natural volatility in labor market outcomes in Costa Rica.

To see why, consider Figure 4.2a, which shows the simple difference in log monthly incomes of all working-age natives between the treatment northern border regions and the rest. The simple-difference jumps more than 7% from July 1997 to July 1998, which breaks the ”same trend” condition. And then, from 1998 to 1999, which is when a large change is expected due to Mitch, the small change that occurs is well within the confidence intervals. A second violation of the ”same trend” occurs when the simple-difference jumps down again from 1999 to 2000. In other words, 1998 was such a positive outlier in the relative income of the treatment regions that the differences-in-differences estimations beyond 1999 (e.g. the 2000 and 2001 estimations) will show a strong negative effect that is not necessarily related to Mitch. The other simple labor market outcome differences present similar problems.

In Tables 4.3 and 4.4, the treatment groups are the Costa Rican natives who lived in the northern border regions and the Costa Rican natives that were identified as working in the construction and agriculture economic sectors. The results were obtained as follows: First, estimate the baseline difference-in-differences treatment coefficient for all individuals with control variables (age, sex, education). Then find the diff-in-diffs coefficient for each of six subgroups, broken up by sex (men and women) and level of education, controlling for age.

The 1999 results do not allow us to reject the null hypothesis that the migration spike after Hurricane Mitch had no impact on labor market outcomes. This is true for almost all sex-education subgroups. The exceptions are employment outcomes for the subgroup of highly educated (tertiary education) women in the treatment economic sectors (which is likely to have small sample problems) and, more importantly, a negative log income effect in low (primary only) education men in the northern border regions. This effect is harder to dismiss. It is a relative -6% change in income in response to a 10% increase in the relative size of the entire labor force for that subgroup (due to immigration), which yields a 0.6 income-labor elasticity, larger than most findings in the U.S., but not in Europe (see [Hunt, 1992]).

In later years, such as 2000 and 2001, measured with respect to the base year of 1998, it is

Figure 4.2: Simple difference in labor market outcomes of working-age natives, between the Northern Border Regions and the rest

(a) In log income

(b) In employment (observed incomes)

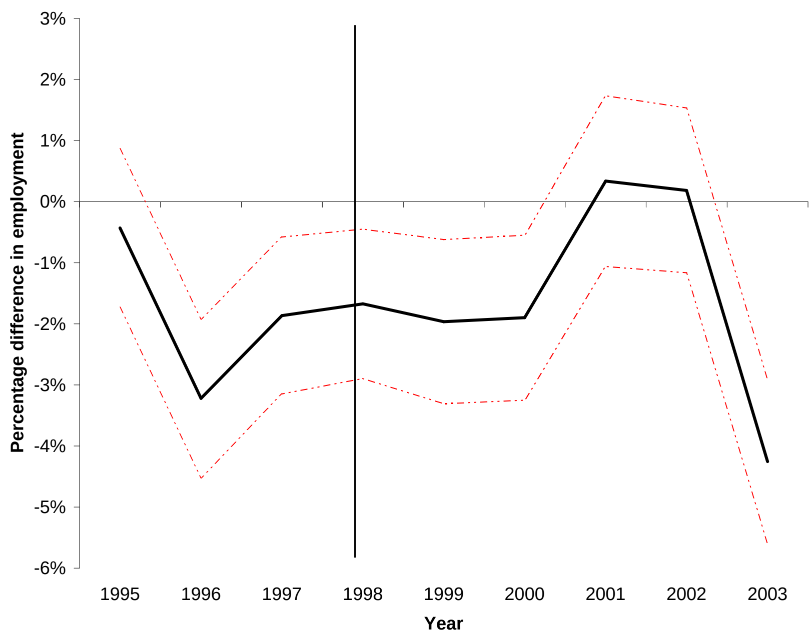
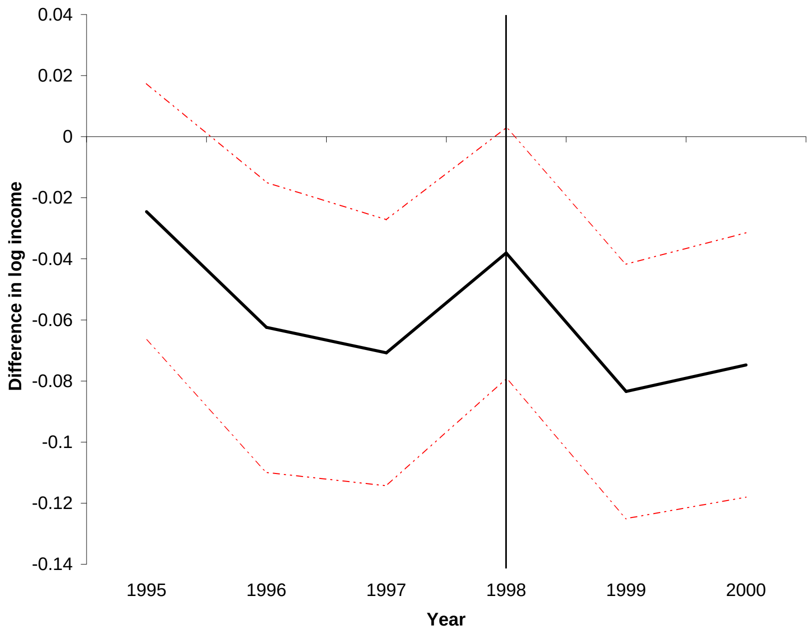


Figure 4.3: Simple difference in labor market outcomes of working-age natives, between the Agriculture-Construction sectors and the rest

(a) In log income



(b) In employment(observed income)

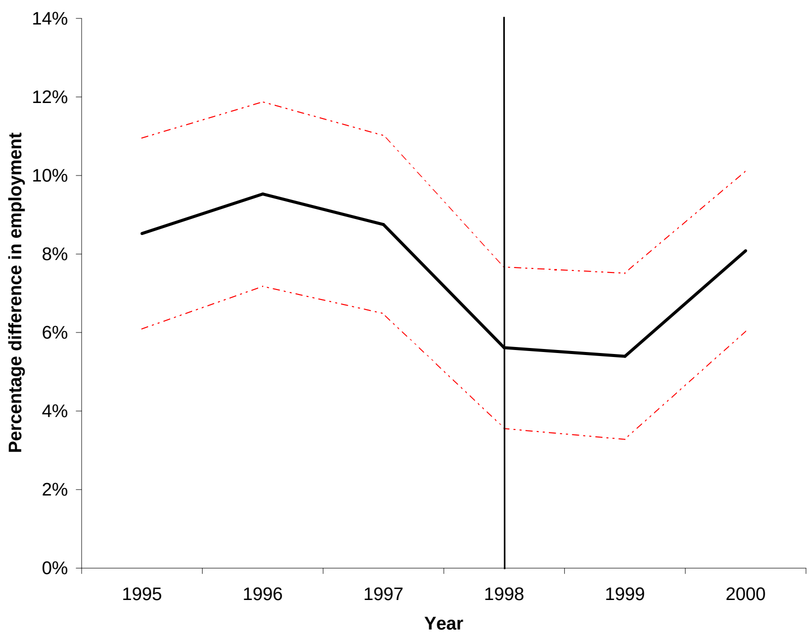


Table 4.3: Differences-in-Differences between the Agriculture/Construction sectors and the rest, using 1998 as base year

	Total	Men by Education Level			Women by Education Level		
		Primary	Secondary	Tertiary	Primary	Secondary	Tertiary
Year 1999							
Employment %	-0.002	0.003	-0.017	0.086	0.088	-0.031	0.469
	(0.015)	(0.018)	(0.039)	(0.090)	(0.094)	(0.174)	(0.237)**
Log Income	-0.045	-0.035	-0.028	-0.159	-0.183	-0.256	-0.395
	(0.030)	(0.035)	(0.075)	(0.219)	(0.277)	(0.621)	(0.437)
Log Labor	-0.031	-0.015	0.089	-0.042	-0.028	-0.374	0.366
Income Elasticity	1.478	2.320	-0.309	3.779	6.590	0.683	-1.077
Year 2000							
Employment %	0.025	0.026	-0.008	-0.039	0.105	0.129	0.403
	(0.015)*	(0.018)	(0.038)	(0.096)	(0.088)	(0.181)	(0.237)*
Log Income	-0.037	-0.006	-0.037	-0.089	-0.184	0.287	-0.135
	(0.030)	(0.037)	(0.075)	(0.184)	(0.273)	(0.339)	(0.236)
Log Labor	-0.048	-0.050	0.055	0.043	0.119	-0.108	0.477
Income Elasticity	0.767	0.122	-0.679	-2.075	-1.546	-2.660	-0.283

Notes: Source of data is Costa Rican Household Surveys, INEC. Employment % reflects the change in the percentage of non-zero incomes among working-age adults of the subgroup. Log income is the logarithm of self-reported individual income in the survey. Log Labor is the logarithm of the size of the labor force, including Nicaraguans. Income Elasticity is with respect to the change in the size of the labor force.

Table 4.4: Differences-in-Differences between the Northern Regions and the rest, using 1998 as base year

	Total	Men by Education Level			Women by Education Level		
		Primary	Secondary	Tertiary	Primary	Secondary	Tertiary
Year 1999							
Employment %	-0.003	-0.005	-0.001	-0.004	-0.019	-0.011	0.005
	(0.009)	(0.015)	(0.024)	(0.043)	(0.018)	(0.026)	(0.042)
Log Income	-0.021	-0.063	0.006	0.025	-0.011	0.026	-0.078
	(0.020)	(0.030)**	(0.045)	(0.089)	(0.057)	(0.074)	(0.085)
Log Labor	0.028	0.099	-0.035	0.021	0.019	-0.082	0.104
Income Elasticity	-0.733	-0.637	-0.177	1.152	-0.610	-0.323	-0.752
Year 2000							
Employment %	-0.002	-0.006	-0.012	0.112	-0.015	0.008	0.062
	(0.009)	(0.015)	(0.024)	(0.040)***	(0.018)	(0.026)	(0.042)
Log Income	-0.092	-0.129	-0.052	0.016	-0.062	-0.139	-0.022
	(0.021)***	(0.031)***	(0.047)	(0.090)	(0.058)	(0.079)*	(0.083)
Log Labor	0.067	0.126	-0.002	0.055	0.110	-0.063	0.122
Income Elasticity	-1.365	-1.023	33.090	0.293	-0.569	2.196	-0.185
Year 2001							
Employment %	0.020	0.028	0.026	0.047	-0.016	0.055	0.004
	(0.009)**	(0.016)*	(0.024)	(0.043)	(0.019)	(0.026)**	(0.042)
Log Income	-0.094	-0.123	-0.143	0.081	-0.074	0.004	-0.170
	(0.021)***	(0.032)***	(0.049)***	(0.081)	(0.058)	(0.076)	(0.080)**
Log Labor	0.068	0.163	0.073	0.108	0.101	-0.050	0.158
Income Elasticity	-1.383	-0.756	-1.960	0.752	-0.736	-0.087	-1.080

Notes: Source of data is Costa Rican Household Surveys, INEC. Employment % reflects the change in the percentage of non-zero incomes among working-age adults of the subgroup. Log income is the logarithm of self-reported individual income in the survey. Log Labor is the logarithm of the size of the labor force, including Nicaraguans. Income Elasticity is with respect to the change in the size of the labor force.

possible to find larger labor market impacts (e.g. -9% change in incomes of all working-age natives in the border regions, in response to a 7% increase in the labor force). However, as pointed out from the simple-difference figures, it is possible to explain these results as a return to the true levels of income from the outlier that is the base year of 1998. A less likely possibility is to think of the latter year results as the effect of the combined increase in Nicaraguan immigrants due to Mitch and the increased ability of Nicaraguans to compete with Costa Ricans in the labor market thanks to the normalization of the migratory status of 150,000 Nicaraguans during the amnesty of the first half of 1999. However, these years show both an increase in the size of the labor force and of the percentage employed in the treated labor force at the same time that incomes in the treatment regions decline 9% simply suggests that volatility in labor outcomes is large.

4.6 Conclusion

There is a rich literature on South-South migration and a separate literature that uses natural experiments to determine the impact of migration on labor market outcomes. The purpose of this chapter is to contribute to both, by presenting new evidence of immigration impact on labor market using an example of a natural experiment in a South-South migration context.

After two decades of continuous migration from Nicaragua to Costa Rica it is likely that Nicaraguan migration responds to labor demand increases in Costa Rica. This endogeneity makes it harder to identify the effects of these migratory flows on native labor market outcomes. I employ a natural experiment to solve this problem. In October 1998, Hurricane Mitch severely affected Honduras and Nicaragua and generated a large exogenous labor supply shock of Nicaraguan immigrants in Costa Rica, a shock noticeable in Costa Rican Birth data, Census data and Household Survey data. The supply push is shown, using Census and Household Survey data, to have been especially strong in those regions that share the border with Nicaragua and in the agriculture and construction economic sectors. This permits the use of a difference-in-differences methodology where the treated native population is that affected by disproportionate migratory flows relative to the rest of Costa Rica.

The main finding of this chapter is that immigration had no labor market outcome effects for nearly all working-age native subgroups in 1999, the year immediately after Mitch. The exception is the subgroup of low-education men in the border regions, who had a -6% income decline with a 10% labor force size change, implying an income-labor elasticity of 0.6, below what has been found in other work such as [Hunt, 1992]. Even then, it can be argued that this result is downwards biased due to a special amnesty that increased the legal labor force with 150,000 Nicaraguans in the first half of 1999[9] and that it is swamped by the natural volatility of labor outcomes in Costa Rica. For example, in later years, it is possible to find larger negative changes in labor market outcomes relative to the pre-Mitch values. However,

[9]For a more complete description of this amnesty, see [Mora, 2004]

these results are distorted because 1998, the base year used for the difference-in-differences methodology, is a positive outlier for the treatment groups, a fact found by inspecting the time trends of the control and treatment groups before and after Mitch. Therefore, this chapter adds to the growing evidence that immigration doesn't have a large impact on native labor market outcomes, either in developed countries such as the United States or France, or in developing countries such as Costa Rica.

Chapter 5

Conclusion

In *Essays on Economic Development in Costa Rica* I attempt to address two kinds of decisions that are usually a challenge to evaluate: fertility decisions and the decision to migrate. With the caveat that the results might apply only to the Costa Rican environment, I believe this thesis contributes to the economic literature on both types of decision.

The literature rarely finds an effect of a family policy on fertility decisions. This is partly due to the small size of most family policies relative to the overall cost of childbearing. In chapter 2 I estimate the impact of the Responsible Paternity Law on childbearing and marital status decisions. This law makes it easier for single mothers to get child support payments. Child support payments are large in lifetime income terms. I find that birthrates fell and they fell the most in those women who had not taken the decision to start their childbearing. I also find that births within marriage fell as much or more than total births. A possible explanation is that premarital sex is common and shotgun marriage has declined as women realize they don't need to marry to receive child support.

In chapter 2 I look at the impact of a policy on family structures and fertility. It is then logical to ask about the impact of family structures on other outcomes. In chapter 3, I studied, in collaboration with Maximilian Kasy, the effect of changing family structures on the female income distribution in Costa Rica. We find that changing family structures had an inequality-increasing effect for working-age Costa Rican women over this period, in particular due to changes in the period after 2000/01, i.e., after the Responsible Paternity Law was introduced. Changes in the period prior to 2000/01 had a negative effect on equivalent incomes across the entire income distribution. The inequality increasing effect was particularly strong among rural women, while the change family structures after 2000/01 had a particularly positive effect for younger women, except for the lowest income groups. Our results illustrate the role of family structures in the determination of the distribution and level of incomes. These results imply, in particular, that various policies affecting family structures also have an important distributional effect.

There is a rich literature on South-South migration and a separate literature that uses natural experiments to determine the impact of migration on labor market outcomes. The purpose of Chapter 4 is to contribute to both. I use the surge in Nicaraguan immigrants after Hurricane Mitch in 1998 as a natural experiment to solve the endogeneity problem of migration. I use differences-in-differences, taking advantage of the uneven location decisions of this migrant surge. The main finding is that immigration had no labor market outcome effects on working-age native subgroups in 1999, the year immediately after Mitch. The exception is the decrease in incomes for the subgroup of low-education men in the regions that share a border with Nicaragua, but even then, this result is small relative to the high volatility of incomes in Costa Rica.

For future research, the permanent and dramatic drop in the birthrate and especially in total births found in chapter 2 implies potential discontinuities across cohorts in educational outcomes, crime, access to health services and so on. Eventually, with adult outcomes, it will be possible to determine if there was a positive selection effect due to the Responsible Paternity Law. A valuable extension of the research presented in chapter 3 might be a combination of the distributional decomposition methods discussed with credible estimates of the structural relationship between child support legislation and household composition. Such a combination would allow to assess more precisely the distributional impact of the Responsible Paternity law, and open interesting methodological perspectives for distributional policy evaluation.

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