How does the Right to Divorce Affect Resource Allocation within Households? The Case of Chile

Misty L. Heggeness[1,2]

Abstract: In recent years, many economists have argued the unitary household model, which assumes households maximize a single utility function given an overall household budget constraint, does not accurately describe the economic behavior of households. Instead, they argue, models should acknowledge that household members' individual bargaining power influences the allocation of household resources. This study examines the effects of exogenous changes in family policy and administrative processes in Chile on the allocation of resources towards children's education. Specifically, the legalization of divorce and family court wait times for divorce are analyzed. Using panel data and a difference-in-differences approach, I show that implementing a pro-female divorce law shifts the bargaining power within married couple households towards the wife, as does the speed with which family courts process divorce cases. By increasing women's bargaining power, both family policy and administrative processes have had an influence on household consumption decisions relating to children's education.

Keywords: Bargaining Power, Household Behavior, Family Economics, Policy Analysis, Education, Difference-in-Differences Estimation

JEL Classification: D12, D13, I00, J12, J18

---

[1] Misty L. Heggeness is a Ph.D. candidate in the Department of Applied Economics and a research assistant at the Minnesota Population Center, University of Minnesota, 50 Willey Hall, 225 19th Avenue South, Minneapolis, Minnesota, 55455. Comments can be sent via email to hegg0018@umn.edu.
[2] The author wishes to thank David Bravo and the Centro Microdatos, Facultad de Economía y Negocios, Universidad de Chile for the invaluable support and guidance they provided for this project, as well as the Center for International Food and Agricultural Policy, University of Minnesota for financial support.

Introduction

In recent years, many economists have argued that the unitary model of household utility, which assumes that households maximize a single utility function given an overall household budget constraint, does not accurately describe the economic behavior of households (Alderman et al. 1995, Behrman 1997, Bergstrom 1997, Gray 1998, Ermisch 2003). Some have found empirical evidence rejecting this model (Schultz 1990, Thomas 1990, Fortin and Lacroix 1997, Browing and Chiappori 1998, Rangel 2006). Instead, they argue, models should acknowledge the bargaining power of individuals to influence the allocation of household resources. This study examines the effects of exogenous changes in family policy and administrative processes on one household decision, investments in children's education.

Until November 2004, divorce did not exist in Chile. Instead, married couples wishing to dissolve their relationship had two options: separation (while remaining legally married) or annulment. This study analyzes the effect of legalizing divorce in Chile on married family households. This new law created an option for dissolving marriage. While in most cases it is very difficult to study the effects of divorce on children and families because of sample selection issues, this study takes advantage of the new law and panel data that follow the same individuals before and after the law change in order to tease out the effect of legalizing divorce on intrahousehold allocation decisions regarding children's education. It also estimates the effects of exogenous variation in wait time for access to divorce via family courts on the same variable.

While testing the accuracy of the traditional unitary household model by capturing these types of effects using a difference-in-differences methodology with cross-sectional data has been accomplished in previous research (Rangel 2006, Martínez 2007), this study is the first to use panel data to analyze the specific effects of having, versus not having, a divorce option, as well as the unintended consequences of administrative processes, on household bargaining and resource allocation. It argues that pro-female divorce legislation increases the bargaining power of women within marriage. In game theoretic models, each player has alternative options for game play. The pro-female divorce legislation transfers resources to wives upon divorce, which increases the threat point, or the point at which the alternative option of leaving the marriage is preferred to staying in the marriage, making the opportunity cost of marriage higher for wives. Additionally, shorter wait times to divorce increase the credibility of the threat to divorce, and thus the bargaining power of the wife.

Previous studies have shown that women invest more in some types of household goods, such as children's education and clothing, than men (Quisumbing and Maluccio 1999, Rubalcava et al. 2004, Schady and Rosero 2007). If, via increased bargaining power of married women through the legalization of divorce and shorter wait times to divorce, there are significantly more investments in one type of household good, children's education, then this study reinforces the recent finding that collective bargaining household models are a more accurate depiction of household behavior. More surprisingly, however, it provides evidence of unexpected impacts of family policies and administrative processes relating to marital instability on the behavior of stable families. The most interesting results of this study are that policies created for unstable families directly influence the intrahousehold allocation decisions of stable families, as do administrative processes at the local level. In other words, the creation of divorce as an option for unstable families and the speed with which family court districts can process a divorce have significant and positive effects on stable families' investments in children's education.

Literature review

There is limited research on the effects of government policies pertaining to marital instability on household resource allocation and even less discussion of the implications in developing countries. This section reviews the economic literature on household behavior and intrahousehold allocation changes associated with laws or programs.

As explained in the introduction, a common practice in economics, until the 1980s, was to model a household as maximizing a single, well defined utility function subject to a household budget constraint, which is now known as the unitary household model. Many economists have criticized this model (Alderman et al. 1995, Behrman1997, Bergstrom 1997, Gray 1998, Ermisch 2003), and others have found empirical evidence that rejects the unitary model and income pooling (Schultz 1990, Thomas 1990, Fortin and Lacroix 1997, Browning and Chiappori 1998). Game theory models have been developed in which household members bargain over decisions related to household consumption based on the bargaining power they hold within the household or based on the separate spheres they occupy within the household (Manser and Brown 1980, McElroy and Horney 1981, Lundberg and Pollak 1993, Lundberg and Pollak 1994, Lundberg and Pollak 1996).

Both unitary and bargaining household models in their most general form are classified by Haddad et al. as collective models (1997). Models under this collective model format include Becker’s (1973, 1974, 1981) altruism model, where an altruistic parent or partner cares about the preferences of their child or spouse/partner and, therefore, transfers income to that person, Chiappori’s (1988) income-sharing rule model where sharing rules are developed based on individual incomes, and the Manser and Brown (1980) and McElroy and Horney (1981) models of a specific bargaining process using game theory. McElroy (1990) defines her model as a Nash bargaining model that allows both non-wage income and external factors called “extra-household environmental parameters” (EEPs), such as policy changes to marriage or divorce law, to influence bargaining power within the household. EEPs shift the opportunity cost of being married, and, therefore, have the potential to increase or decrease the gains to being married for men and women. Lundberg and Pollak create a separate spheres bargaining model (1993) where they show that shifts in intrahousehold allocation can be caused by simply making cash payments (i.e. for child allowances) to a mother instead of a father, which can imply different equilibrium distributions.

Several studies have examined the effects of changes in divorce law and alimony rights on families and intrahousehold allocation (Gray 1998, Chiappori et al. 2002, Rangel 2006). Gray examines divorce-law changes, household bargaining, and married women’s labor supply in the U.S. Using a bargaining model, he takes advantage of an exogenous change in state divorce laws to analyze the response of women’s labor supply to unilateral divorce laws. He finds evidence to reject the neoclassical unitary model assumption of income pooling and accepts the bargaining model of household behavior as a plausible interpretation of household time allocation and decision-making.

Chiappori, Fortin, and Lacroix (2002) also analyze marriage markets, divorce legislation, and household labor supply. They find a causal relationship between marriage markets (sex ratios), divorce laws, and labor supply in that both sex ratios and pro-female divorce laws affect women’s labor supply behavior and decision processes in the ways that one would expect, and those effects are sizeable. Passing divorce laws that are favorable to women increases the amount of money transferred from the husband to the wife after divorce. In addition, an increase in the proportion of males in the population increases the transfer of money to their wives because more men relative to women implies a better marriage and remarriage market for women, which increases the available options outside the marriage.

While the above studies analyze the effect of changes in divorce laws in the United States, family policies towards alimony and child support have also been shown to affect household allocation decisions in Latin American countries (Rangel 2006, Martínez 2007). Rangel finds that an exogenous policy change extending alimony rights and obligations to cohabitating couples in Brazil increased the bargaining power of cohabitating women, as shown by a decrease in their total hours worked (in formal labor as well as household labor) and increased investments in the education of their children. His study provides evidence of gender-specific intrahousehold allocation preferences. Martínez finds that extending child support enforcement laws to out-of-wedlock children in Chile decreases the probability that men work, while increasing the probability that children attend school, again providing evidence that family policies have the potential to increasing women’s bargaining power within the household. Both of these studies use an exogenous policy shock to analyze changes in women's bargaining power within the household, and both find that when mothers have more resources after union dissolution, increased investments are made in their children's education.

Several recent papers argue that women are more likely than men to invest in household goods, like children’s education, clothing, or food for household meals (Quisumbing and Maluccio 1999, Rubalcava et al. 2004, Schady and Rosero 2007). While child health and education are future household investments in the form of informal social security for the both parents in old age, investments in children’s health and education have been shown to increase when women gain more bargaining power within the household. Quisumbing and Maluccio show that having more assets controlled by women is associated with increased investments in children's education and clothing in four countries. Rubalcava et al. find that money put in the hands of women via a cash transfer program is more likely to be spent on children's goods, better nutrition, and investments in small livestock, all of which are investments back into the household. Schady and Rosero find that unconditional cash transfers to women in Ecuador increase income shares spent on food expenditures in households with both men and women compared to female-only households. This is evidence that gender-specific bargaining occurs in Ecuadorian households and, when more resources are put into the hands of women, increased investments in household items such as food expenditures, can be observed.

These studies provide evidence that bargaining exists in households and that women allocate resources differently than men. While they show that government policies giving more power to women and cash transfer programs that transfer money to women shift the bargaining power from men to women and, thereby, influence intrahousehold allocation, more research is needed to understand specifically the effects of divorce on intrahousehold allocation decisions in married couple families, particularly in developing countries. The contribution of this study is that it uses rigorous econometric techniques to tease out the effect of divorce on intrahousehold allocation decisions regarding children's education in Chile. It is also one of the first studies taking advantage of a natural experiment setting of random variations in the administrative length of time to finalize a divorce to show the effects of unintended governmental processes on household behavior.

Background

> As an interesting exceptional case, Chile has evolved a widely understood body of procedures for *annulment*, remarkably akin in their ingeniousness to the elaborate grounds for annulment in Church courts in Europe over the several centuries after the indissolubility of marriage was finally imposed (in 1563). They were then, as they are now in Chile, most easily utilized by families with adequate means to pursue their goals with the aid of lawyers.
>
> Since a legal marriage in Chile can go forward only after a number of official facts are filed, it follows that any proof that the official record contains errors could become the grounds for annulment. This can be as trivial as the claim that the addresses of the prospective spouses were not correct. Needless to say, this possibility is not written explicitly into the law. On the other hand, it can only be done with the collusion of the couples as well as the court judges. Because an annulment does permit remarriage, it is, then, the Chilean "substitute" for a real divorce. (Annulment does not apply to consensual unions, which legally are not marriages.) (Goode 1993, p. 189)

Prior to November 2004, no formal mechanism existed with which to divorce in Chile.[3] Disputing spouses either informally separated but remained legally married, meaning they were unable to marry anyone else, or legally annulled their marriage. Informal separation left the custodial parent vulnerable because limited formal mechanisms existed for transferring resources from the noncustodial parent to the custodial parent. While either partner can request a legal separation via the family court system and the custodial parent can formally request child support this rarely occurred.[4]

Legal annulment in Chile requires both spouses to cooperate with each other because they must agree to report inaccuracies in their marriage license application (such as an inaccurate

[3] This overview of the creation and existence of the Chilean divorce law comes from interviews with Gabriel A. Hernandez Paulsen, Professor of Family Law, University of Chile, in May 2009 and Luis Perez, Chilean Family Court Aide and Lawyer, in May 2009.

[4] The total number of legal separations in the entire country was less than 70 cases each year between 2005 and 2008 (http://www.registrocivil.cl/f_estadisticas.html). Number of legal separations prior to 2005 is unavailable but via interviews I have learned there is not much difference before and after the legalization of divorce (Interview with Gabriel A. Hernandez Paulsen, May 2009).

living address) to the judge who married them in order to annul their marriage.[5] In addition, legal annulment usually requires financial resources to pay legal fees. Therefore, spouses can only annul if 1) they agree to cooperate with each other and 2) they have the necessary financial resources to pay for the annulment. Before divorce became legal, spouses wanting to end their relationship but choosing not to cooperate with each other or not having the necessary finances were able to be separated, but had to remain legally married.

In November 2004, divorce became part of the Chilean family law. With the implementation of the divorce law, disputing spouses had the option of formally divorcing their partner, thereby acquiring 1) the right to remarry and 2) the right to receive an economic compensation if they stayed in the household to take care of children or the home during the marriage. According to this new law, upon divorce, the partner who set aside his or her career to take care of the family home or children is entitled to a payment from their partner, called an economic compensation. The economic compensation is a lump sum of money to be paid all at once or in monthly installments until the entire amount is paid. Judges calculate the payment based on the assumed lost wages of the homemaker spouse. An average wage per year is calculated based on the homemaker's education, family background, and other socioeconomic factors. This wage is then multiplied by the number of years married where the homemaker was staying at home taking care of the family. Over time, the technique used to calculate the economic compensation has changed.[6] However, during the time period for which the study data

[5] Any discrepancies about name, address, or other standard information given by the couple to the courts at the time of marriage, is justification to claim the marrying judge "incompetent," which provides a case for annulling the marriage.

[6] Today, the goal of the economic compensation is to give the homemaker spouse enough money after divorce so that she does not become impoverished, but instead maintains a more or less equal status as she had during marriage, at least for the first few years after the divorce. It is assumed by the courts that providing this resource the first couple of years will allow the homemaker spouse enough time to be able to be independent after she has used up all of the economic compensation money. (Interview with Gabriel A. Hernandez Paulsen, May 2009; interview with Luis Perez, May 2009).

were collected, calculating economic compensation was calculated in the fashion described above.

When couples decide to divorce, they can divorce only in the family court district corresponding to the county in which they live. Couples living in a county corresponding to a family court district with a very long wait time have no choice but to wait for their divorce to become finalized. They cannot go to a neighboring county/family court district with a shorter wait time to expedite the process.

While divorce is now legal in Chile, it is still relatively uncommon. In 2008 there were approximately 22,000 divorces in a country of more than 10 million adults.[7] Divorce was even less common in the years immediately following the legalization of divorce; in 2005 and 2006 together, there were less than 12,000 total divorces (Chart 1). If one makes relatively harmless assumptions that all divorces involve two adults and that no one individual divorced more than once between 2005 and 2008, there were approximately 50,000 divorce cases in that time period, resulting in 100,000 individuals divorced. This upper bound constitutes less than 1 percent of the adult population. While there is not enough transition to divorce to study divorce rates or the implications of divorce on divorced parent households in the early years after the legalization of divorce, I will show that the legalization of divorce had significant effects on the bargaining power and intrahousehold allocation decisions of married couple households shortly after the law went into effect. In other words, the divorce legislation appears to have had a larger impact on intact families than on unstable family units.

[7] According to the 2002 Census of Chile, 74.3 percent of the population (11.2 million persons) where age 15 or older (http://www.ine.cl/cd2002/sintesiscensal.pdf).

Data

This paper uses panel data from the Encuesta de Protección Social (EPS),[8] as well as data from the Chilean court system. The EPS currently consists of three waves or rounds (2002, 2004, and 2006) that follow the same individuals over time. Since the original purpose of the survey was to collect labor and social security pension fund data, the first wave (2002) is nationally representative of all individuals who contribute to a public pension fund. The 2004 and 2006 waves, however, are nationally representative samples of the entire population.[9] The survey includes detailed information on complete marital, fertility, and labor histories, as well as detailed information on the family in which the interviewee was raised. For the purposes of this study, county-level identifiers are included in the dataset so that family court administrative data may be appended.

For the purposes of this study, I analyze school attendance data of the children of the interviewees. A sample of school age children (ages 4 to 21) whose parents were married or cohabitating with the same person over the entire sample time period (2002 to 2006) is constructed. The sample includes approximately 900 children from cohabiting parents and approximately 4,200 children from married parent families (Table 1). Constructing the sample this way implies that children from parents whose legal marital status changed over time are excluded. Excluding this group is beneficial because it eliminates any confusion regarding whether those who change marital status are somehow confounding the results. However, approximately five percent of the interviewee sample (and, hence, their children) are lost by limiting the sample to stable relationships.

---

[8] The Encuesta de Protección Social, or Social Protection Survey [title translation by author], is a survey administered by the University of Chile and the Chilean Ministry of Work and Social Prevention, in partnership with the University of Pennsylvania and the University of Michigan.

[9] A new subsample of individuals was added to the 2004 wave to make the panel representative of the entire population.

Since complete marital histories exist for the interviewees, the actual marital history of the parent is used to construct parental marital status, instead of a variable for marital status or civil status at the time of the interview. A concern with using a marital status variable in household survey and census data is whether one's marital status from one survey to the next refers to the same partner. The definition of marriage and cohabitation used in this study ensures that a child identified as having married or cohabiting parents has parents who have been married to or cohabiting with the same partner in 2002, 2004, and 2006.

A second source of data was collected by the author together with the director of the Microdata Center at the University of Chile by making a special request to the administrative offices of the Chilean Supreme Court.[10] These data contain basic information about the date each divorce case started within each respective local family court district and the date the divorce case was finally settled by the family court from the beginning of 2005 to the end of 2006. A divorce case is started when all paperwork is turned into the court, which means that all forms have been filled out completely and all requested information has been received. The data create a natural experiment environment because the wait time between submitting one's paperwork to the court and receiving a court date to finalize the divorce is driven solely by each court's individual backlog and administrative procedures.[11] This additional dataset is merged with the panel data from the Encuesta de Protección Social (EPS) by county code and used to examine the effects of variation in local court administrative procedures on household consumption decisions (described in detail below).

[10] Electronic data on the dates of divorce cases in the Chilean family court system exists only for urban areas. Therefore, for the regressions that include average wait time for a divorce are limited to couples living in urban areas. According to the 2002 Chilean Census, 86.6 percent of the population lived in urban areas.

[11] Any cases related to the family, including cases not associated with divorce, such as the distribution of inheritances from wills, adoption, or domestic violence cases, are also processed by the family court system. The backlog is driven by the combination of these cases and the way in which each family court administrator manages his or her court.

Methodology

The household bargaining model used in this paper is based on those of Manser and Brown (1980) and McElroy and Horney (1981), which, in turn, are based on the Nash (1953) two-person cooperative game model. An application of the model for this paper is described below.[12]

In the model there are two individuals, *m* and *f*, in a married or cohabitating couple household, and they jointly allocate resources via a solution to a two-person, Nash cooperative game.[13] Each player in this game has a threat point, or a point at which some alternative situation becomes preferred to their original play in the game. The threat point is the utility received from dissolving the marriage. If the utility, or benefit, from remaining married falls below the threat point for one (or both) player(s), and that player's partner cannot transfer enough resources to him or her without the partner's own utility from marriage falling below the utility he or she would receive from leaving the marriage, then the first individual will choose to leave the marriage, and it will dissolve.

Each individual has the following utility function:

$$U^i(\boldsymbol{x_0}, \boldsymbol{x_i}, \ell_i) \; \forall \; i = m, f \tag{1}$$

s.t. $\boldsymbol{p_0 x_0} + \boldsymbol{p_{xi} x_i} + w_i \ell_i = I_i + w_i T + \alpha_i \;\; \forall \; i = m, f$ (full income constraint)

where $\boldsymbol{x_0}$ are household public goods including children's education

$\boldsymbol{x_i}$ are private goods consumed by i

ℓ_i is the leisure consumed by i

α_i is the income transferred to or from partner *j* to partner *i* upon divorce.

[12] Also see McElroy (1990) and McElroy (1997).
[13] In this model, *m* and *f* can be thought of as *male* and *female* or *mother* and *father*, etc.

U^i is assumed nonnegative. Let T be the total time endowment for both m and f and I_i be the nonwage income for $i = m, f$. If not married or cohabitating, each person would maximize his or her own utility subject to a full income constraint, leading to their respective indirect utility functions $V^i(\boldsymbol{p_0}, \boldsymbol{p_{xi}}, w_i; I_i, \alpha_i)\ \forall\ i = m, f.$

Assuming m and f are married, V^i is the threat point for leaving the marriage for $i = m, f$ in a Nash bargaining model. The α_i affects only in the indirect utility function because its influence is on the individual's outside option or what the individual can gain from choosing to divorce. If the individual stays married, they do not receive the benefits (or costs) of the divorce law or the wait time to divorce as they would should they end the marriage. In other words, $\alpha_i = 0$ in equation (1). Therefore, divorce legislation and administrative wait times affect household behavior of married couple families by increasing or decreasing the value of one's utility outside of the marriage and, in this way, directly influence the V^i and bargaining power of each individual and do not affect U^i. Since divorce laws and wait times for a divorce only affect married couples, $\alpha_i = 0$ always for cohabitating couples.

An individual considering marriage dissolution has multiple threat points. For the case of Chile, there are three threat points: V^i_d = the threat point under divorce, V^i_s = the threat point under de facto separation, and V^i_a = the threat point under annulment. Whichever threat point is the highest is the true threat point used by the individual in considering whether to stay in the marriage or dissolve it. If $V^i_d > V^i_s \geq V^i_a$ or $V^i_d > V^i_a \geq V^i_s$ the legalization of divorce will increase the opportunity cost of staying married for mothers and decrease the opportunity cost for non-custodial fathers primarily because of the economic compensation clause tied to the divorce law. For cohabitating couples, V^i_s is the threat point before and after the legalization of divorce since the only outside option is the utility gained from separating from their partner.

For the couples in this model, the Nash-bargained solution to the joint maximization of the product of their gains from marriage or cohabitation is:

$$\text{Max}_{\{\mathbf{x}\}}\ [U^m(\boldsymbol{x_0}, \boldsymbol{x_m}, \ell_m) - V^m(\boldsymbol{p_0}, \boldsymbol{p_{xm}}, w_m; I_m, \alpha_m)][\, U^f(\boldsymbol{x_0}, \boldsymbol{x_f}, \ell_f) - V^f(\boldsymbol{p_0}, \boldsymbol{p_{xf}}, w_f; I_f, \alpha_f)] \qquad (2)$$

s.t.

(i) $\boldsymbol{p_0x_0} + \boldsymbol{p_{xm}x_m} + \boldsymbol{p_{xf}x_f} + w_m\ell_m + w_f\ell_f = (w_m + w_f)T + I_m + I_f \equiv$ full income constraint

(ii) $\ell_i \leq T_i$

(iii) $\alpha_f = - \alpha_m$

Under this problem, m and f will choose to dissolve the marriage if the gains to dissolving ($g^i{}_d$) outweigh the gains to remaining married ($g^i{}_m$). In other words, for this household maximization problem to be solved, $g^i{}_m > 0$, where $g^i{}_m = U^i - V^i\ \forall\ i = m, f$. If $g^i{}_m < 0$ for partner i and $g^j{}_m > 0$ for partner j where $j = m, f$ and $j \neq i$, then partner j may choose to transfer resources to partner i to keep the marriage together if the transfer of resources still leaves partner j with some gain to marriage ($\breve{g}^j{}_m > 0$ where $g^j{}_m > \breve{g}^j{}_m$). If the transfer of resources is enough so that partner i's gains to marriage become positive ($\breve{g}^i{}_m > 0$ after the transfer, where $\breve{g}^i{}_m > g^i{}_m$), then the marriage will not dissolve.

Also note that by constraint (iii) $\alpha_f = - \alpha_m$. This constraint implies that as α_f increases, α_m will decrease, V^f will increase, and V^m will decrease. In other words, an increase in pro-female divorce legislation and shorter wait times to divorce increases the value of married women's threat point under divorce ($V^f{}_d$). This increases their bargaining power within the marriage, and, because of this increased bargaining power, increases investments in goods that women value, such as children's education. At the same time, the value of married men's threat point under divorce ($V^m{}_d$) decreases, as does their bargaining power within the marriage.

The solution to the above maximization problem yields a system of demand equations.

$$x_j^* = h_j(\boldsymbol{p};\, I_m, I_f, \alpha_f) \qquad \forall\, j = \boldsymbol{x_0}, \boldsymbol{x_m}, \boldsymbol{x_f}, \ell_m, \ell_f$$

$$\text{where } \boldsymbol{p} = (\boldsymbol{p_0}, \boldsymbol{p_{xm}}, \boldsymbol{p_{xf}}, w_m, w_f)$$

Notice that the demand for each good is a function of a price vector, non-wage income, the external divorce law, and the administrative wait time for a divorce. With this model, one can analyze the effects of shifts in the threat point, or opportunity costs of remaining together, from exogenous shocks, in this case the legalization of divorce and wait time to finalize a divorce. This paper analyzes changes in the demand for children's education, S_{igt}, driven by shifts in the threat point to divorce because of the new divorce law and variation in the wait time to divorce.

This paper assumes the mother's preferences imply that she will invest more in her children's education than would the father (Quisumbing and Maluccio 1999, Rubalcava et al. 2004, Schady and Rosero 2007). This model implies two hypotheses, the last of which is tested in this paper. If hypothesis two is found to be true, then hypothesis one must also be true.

I. <u>Hypothesis 1:</u> The legalization of divorce, which includes requirements for the economic compensation of homemakers, will cause the following changes to the opportunity cost of staying married:
    a. The opportunity cost of staying married for men decreases (because of the implied transfer of money from husbands to wives via economic compensation).
    b. The opportunity cost of staying married for women increases.

II. <u>Hypothesis 2:</u> Hypothesis 1 implies that married women's bargaining power must have increased, so investments in their children's education among those families who stay married will increase after the legalization of divorce.

Does a divorce threat point, which includes an economic compensation, shift more intrahousehold bargaining power into the hands of women in those families that stay married? If so, given that prior research shows that women invest more in certain types of household goods and resources like children's education, one expects to see an increase in investments to child education in those families that remain married. Therefore, an increase in the bargaining power of married women in Chile, via the threat of divorce and its associated economic compensation, should increase investments in children's education in married couple households. It's also reasonable to expect the exogenous variation in divorce wait time to influence intrahousehold allocation of resources. The shorter the wait time, the more of a credible threat is the divorce. Therefore, shorter wait times are expected to translate into an increase in investments in household goods valued by the mother, such as children's education.

Estimation Methodology

There are two difficulties in analyzing the effect of divorce on household bargaining: sample selection bias and endogeneity bias. Comparisons of households experiencing versus not experiencing divorce have sample selection problems in countries where divorce has existed for many years, which is the case for most countries around the world, because couples are self-selecting into a divorced status. These couples could have similar unobserved characteristics or traits that confound any estimation results.

Problems of endogeneity are also common in these types of studies. While it may seem straightforward, for example, to analyze shares of income in the household per individual as a proxy for bargaining power within the household, it is unclear whether income creates more bargaining power for that individual within the household or whether one's individual characteristics (including the ability to persuade and other favorable characteristics associated with both increased income and household bargaining power) are increasing one's income as well as one's bargaining power. The method used in this study to minimize issues of endogeneity is to introduce two exogenous factors, the legalization of divorce and variation in the wait time to divorce, as proxies for analyzing shifts in household bargaining structures. This method has been applied in other studies as well (Chiappori et al. 2002, Rangel 2006, Martinez 2007), but, to the author's knowledge, this is the first time having (versus not having) a divorce option is analyzed using panel data.

A difference-in-differences (DID) approach is used to identify the effects of the legalization of divorce on child education. This approach uses panel data to estimate the impact of a program or policy change on a variable of interest by comparing the change in that variable for the group that experienced the program or policy to the same change in a group that did not experience the program or policy. This estimation technique essentially uses the former as a treatment group and the latter as a control group. Children from married parent families are the treatment group as they are direct recipients of the treatment, in this case, the legalization of divorce. Children from cohabiting parent families are used as a control group because their households are not influenced by the legalization of divorce, since their parents are not married. In other words, the legalization of divorce is not expected to change women's bargaining power in cohabitating households. Upon separation, cohabitating women are not eligible for the economic compensation that married women are via the new divorce law.

The following is the basic individual-level equation for this analysis.

$$S_{igt} = \beta_0 M_g + \beta_1 T_1 + \beta_2 T_2 + \delta_1 M_g * T_1 + \delta_2 M_g * T_2 + \upsilon_{gt} + \varepsilon_{igt} \qquad \forall\ i = 1,\ldots,I_{gt} \qquad (3)$$

where S_{igt} is a dummy variable indicating whether child i from group g at time t is in school, M_g is a dummy variable that equals one for the treatment group and zero for the control group (married parent versus cohabitating parent families, where the variable is equal to one if the child lives with married parents and zero if living with cohabitating parents), T_t is a set of year dummy variables or the time effects, υ_{gt} is unobserved group effects at time t, ε_{igt} is the individual-specific error term, and $E[\upsilon_{gt}] = E[\varepsilon_{igt}] = 0$. Since the treatment, the ability to divorce, became an option in early 2005, one assumes that $\delta_1 = 0$. There are three time periods in the estimation, T_0, as the reference year, is omitted from the equation above. To obtain consistent estimates of this equation one needs to assume that $E[\varepsilon_{igt} | M_g, T_j] = 0$. The estimate of δ_2 is the average treatment effect of divorce on children's school enrollment.

An equation describing the difference-in-differences coefficient is: $\delta_2 = [(S_m^2 - S_m^0) - (S_c^2 - S_c^0)]$, where δ_2 is the estimated average effect of divorce on children's school enrollment, S_g^j is the sample average of the variable of interest for the group g in time period j.[14] When estimating the effect of the legalization of divorce on children's education using a difference-in-differences estimation, an unbiased estimate of the coefficient of interest, δ_2, can be obtained by regression methods.

In general, adding other explanatory variables provides more explanatory power in the estimation. Thus, age group dummy variables, grouped by school age of the children (primary school, etc.), are added. Since parental decisions to enroll their children in primary school are different from decisions to enroll them in secondary or tertiary school, adding these dummy variables will improve the estimation. Adding these dummy variables gives the following two equations.

---

[14] Note S_m^2 = E[S_{igt}|g = *Married parent, t = 2006*] = $\beta_0 + \beta_2 + \delta_2$; S_m^0 = E[S_{igt}|g = *Married parent, t = 2002*] = β_0; S_c^2 = E[S_{igt}|g = *Cohabitating parent, t = 2006*] = β_2; and S_c^2 = E[S_{igt}|g = *Cohabitating parent, t = 2002*] = *0*.

$$S_{igt} = \beta_0 M_g + \beta_1 T_1 + \beta_2 T_2 + \delta_1 M_g*T_1 + \delta_2 M_g*T_2 + \boldsymbol{\gamma_1 Z_{igt}} + \upsilon_{gt} + \varepsilon_{igt} \qquad \forall\ i = 1,\ldots,I_{gt} \qquad (4)$$

where $\boldsymbol{Z_{igt}}$ are the individual-specific variables indicating whether the child is in a particular school age group. All other variables are the same as in equation (3).

To account for the additional exogenous factor of variation in wait time for divorce, wait time and the appropriate interaction terms can be added to the equation.

$$S_{igt} = \beta_0 M_g + \beta_1 T_1 + \beta_2 T_2 + \delta_1 M_g*T_1 + \delta_2 M_g*T_2 + \boldsymbol{\gamma_1 Z_{igt}} + \mu_1 W_{c2} + \mu_2 W_{c2}*M_g + \qquad (5)$$
$$\mu_3 W_{c2}*T_1 + \mu_4 W_{c2}*T_2 + \alpha_1 M_g*T_1*W_{c2} + \alpha_2 M_g*T_2*W_{c2} + \upsilon_{gt} + \varepsilon_{igt} \qquad \forall\ i = 1,\ldots, I_{gt}$$

where W_{c2} is the average wait time for divorce by court district in the last time period. All other variables are labeled as in equation (3). The variables of interest in equation (5) are δ_2 (the effect of legalizing divorce) and α_2 (the effect of one additional month wait time to finalize a divorce). One expects $\mu_1 = \mu_2 = \mu_3 = \mu_4 = \alpha_1 = 0$.

A key non-trivial identifying assumption with a difference-in-differences estimation is that the *trends* in school enrollment would have been the same for both groups in the absence of the legalization of divorce (Angrist and Pischke 2009). This implies that the variable of interest, in this case children's enrollment in school, is affected similarly by any other environmental changes over time for both the treatment and control groups. One way to test this assumption is to observe the variable for each group before the actual treatment. While the means or percent of individuals affected do not need to have the same outcome, the trend from one time period to the next must be parallel. If the parallel assumption holds prior to the treatment, then the two groups can be compared using difference-in-differences estimation.

Chart 2 shows the rates of school attendance for children from married parent families compared to children from cohabiting parent families. The rate of school enrollment is parallel for both groups before treatment. However, after the legalization of divorce, cohabiting parent family children continue to experience a decrease in school enrollment while children from married parent families experience an increase.

All three estimation equations described above are estimated by a logit regression using household fixed effects. Fixed effects controls for any household-specific time-invariant omitted variables that are the same for all children in a given household but vary across households. In the difference-in-difference estimation, using household fixed effects will eliminate any time invariant variables in the model, as well as any other variables that are the same for households in each time period. For this reason, variables like M_g and W_{c2} are dropped from the estimated equations. Finally, since the dependent variable, children's school enrollment, is a dummy variable indicating one if the child is in school and zero otherwise, a logit model is used to estimate the effect of legalizing divorce on children's education.

Results

While Chart 2 shows an overall decrease in school enrollment for cohabitating parent children ages 4 to 21 and an overall increase after the legalization of divorce for married parent children of the same age, Table 1 reports more variation when school enrollment rates are separated by age group of school type. While primary school and university aged children of cohabitating parent families experience a continual decrease in percent enrollment, the same is not true of their secondary school counterparts, whose school enrollment increases between 2002 and 2004 but decreases in 2006. Married parent children in secondary and tertiary school experienced continual increases in their school enrollment rates, while their primary school counterparts did not. While not much variation is observed in primary school rates for both groups, increasing variation over time in the percent of children in school can be observed between secondary and tertiary school aged children; rates for married parent children are tending up while rates for cohabitating parent children tend down. Chart 1 masks the deviations based on type of schooling, but Table 1 gives clear indication that including dummy variables for school age categories and running separate regressions by school age type are appropriate steps in the estimation process.

Table 2 shows the results from two DID estimations (equation (3) and equation (4) from above) for a nationally representative sample. *Model One* is a standard DID using time dummy variables and interaction terms for time and marital status of the parent (equation (3)). We expect the married parents in 2006 variable (which is an interaction term of a married parent dummy variable with the dummy variable for 2006 and estimates the effect of legalizing divorce on school enrollment) to have a positive coefficient. While the coefficient is positive, this variable is insignificant in this regression ($p = 0.112$). The dummy variable for 2006 is negative and statistically significant at $p = 0.020$.

This regression, however, does not control for ages of the children analyzed and, as shown above, there is variation in school enrollment by school age group. For more explanatory power, dummy variables for school age groups are added to the regression (see equation (4) and *Model Two* in Table 2). Once school age groups are controlled for, the legalization of divorce variable has a larger, positive coefficient and is significant at $p = 0.038$. All age groups are strongly significant at $p = 0.000$.

In order to accurately capture the reality of divorce in the Chilean context, an additional component is added to the regression. With the legalization of divorce came the creation of family courts. Each family court is composed of a small group of *comunas*, ranging from one to nine *comunas* in each group where the average is three or four *comunas* per family court district.[15] As mentioned previously, individuals are required to process their divorce in the family court corresponding to the *comuna* in which they live. The exogenous difference in administrative wait times should influence bargaining power within the household. If divorce shifts the opportunity cost of remaining married, it does so only in the sense that the threat of divorce is truly credible. Specifically, the shorter the wait time, the more credible the threat of divorce becomes. If true, by adding a variable that identifies the average wait time for married couples to divorce in 2006 by family court district, one would expect to see a negative and significant coefficient. In other words, the longer the wait time, the less credible a threat the divorce is, and the less bargaining power the woman will have in married couple households.

Administrative wait time data is available only for family court districts in urban areas. Therefore, the sample used to analyze administrative wait times is individuals living in urban *comunas*. In order to provide an accurate comparison of the results with and without wait time, *Model One* and *Model Two* from Table 2 are run again using the subset of individuals living in urban *comunas*. Notice that *Model One* and *Model Two* results for the urban sample show a positive and significant effect of legalizing divorce on children's education.

When administrative wait time added into the equation (see equation (5) and *Model Three* in Table 3) for an urban sample, it is negative and weakly significant. The coefficient on the variable measuring the effect of the legalization of divorce remains strongly significant. The age group categories are still significant factors in predicting whether children attend school. Finally, the coefficient on the dummy for 2006 is negative and significant. There was some change between 2002 and 2006 that had a weakly significant negative effect on all school children's enrollment compared to previous years, possibly an educational policy change or

---

[15] An exception is parts of the capital city Santiago, in which one family court district encompasses 19 comunas. As an urban area, comunas in Santiago are geographically very small but densely populated. The metropolitan area of Santiago has a total of 10 different family court districts.

shifts in macroeconomic trends driving a need for younger adults to work, or at least not be in school.

While the sign and significance of the coefficients in a logit regression provide relevant information, the coefficients themselves do not explain the estimated effect of each independent variable. For that reason, the marginal effects of the full sample regressions and the urban sample regressions are shown in Table 4 and Table 5, respectively. Marginal effects show the effect on y, the dependent variable, from a one unit change in x, the explanatory variable, holding all else constant. There are multiple ways to calculate marginal effects for a logit model. Two examples of how marginal effects can be calculated include calculating them at the average or for a representative agent (Cameron and Trivedi 2009). While in practice these methods tend to give similar results, if the independent variables are dummy variables, calculating marginal effects for a representative agent is more meaningful because calculating marginal effects at the average for a dummy variable will not refer to any particular category (neither the 0 nor the 1 case). Table 4 and Table 5 report these results for a representative agent who is a primary school aged child in 2006 living with married parents in a family court district with an average wait time of 4 months, which is the average wait time for the entire sample.

The results shown for *Model Three* in Table 5 show an estimated marginal effect of the divorce law on primary school aged children in 2006 living with married parents in a family court district with an average wait time of four months to be 29.7 percent.[16] In other words, holding all else constant, legalizing divorce increased school enrollment by 29.7 percent for children of married parents compared to children of cohabitating parents. Each additional month

[16] The sample in *Model Three* is smaller because data on wait times to divorce is only available in urban areas. Therefore, the results also represent the effect of legalizing divorce on education and families in urban areas, where as *Model One* and *Model Two* are nationally representative. This could explain the large variation in magnitude of the marginal effects of divorce between *Model Two* and *Model Three*.

of wait time to finalize a divorce is estimated to decrease school enrollment by 1.2 percent for this representative agent group. These marginal effects explain the magnitude of the effect of legalizing divorce and administrative wait times. The effect of legalizing divorce is strong, positive, and large and was responsible for a significant increase of children from married parent families to be enrolled in school, holding all else constant.

Although the legalization of divorce clearly had a positive effect on children's education, interpreted as increasing women's bargaining power within married couple families, a question still remains as to which school age children benefitted the most. For that reason, Tables 6 and 7 replicate *Model Two* and *Model Three* regressions in Table 2 and Table 4, respectively, but report separated regression results by school age group. Table 6 shows the estimation equation (4) from above. Table 7 shows the full estimation equation (5), which includes average wait time to finalize a divorce. Table 7 results show that legalizing divorce had a significant impact on schooling of youth ages 12 to 17, or secondary education children. It had no effect on primary school children, nor did it have any effect on tertiary, or university, aged children.

These estimates imply that the legalization of divorce had an effect of raising school enrollment for secondary or high school aged children. This makes sense given that these children might still be too young to be independently working, as is the case with those of university age, but are old enough to where their parents might consider having their children work informally to earn additional income for the household than to have them in school. Marginal effects (Table 8) show that for a representative agent, the effects are large. For a child in 2006 who is living with married parents in a family court district with an average wait time of four months, the legalization of divorce increased school enrollment for youth ages 12 to 17 by

51.7 percent. An additional month added to the wait time for a divorce decreased school enrollment by 2.2 percent for this same group.

Conclusion

Studies analyzing the effects of divorce on child and family wellbeing perpetually face selection bias issues because individuals who divorce can have systemically different characteristics than those who remain married. This study takes advantage of national household survey panel data from 2002, 2004, and 2006 and a 2004 external shock to households in Chile in the form of family policy, the legalization of divorce, to analyze the effects of divorce on child education using a difference-in-differences (DID) approach and, thereby, minimizing selection bias and endogeneity issues. Using panel data before and after the legalization of divorce, and a difference-in-differences methodological approach, this paper investigates the effect of the legalization of divorce on household resource allocation decisions regarding children's education. Specifically, child education is analyzed in cohabitating parent families, who are not affected by the legalization of divorce, and married parent families, who are affected by the new law.

More generally, this paper analyzes the effect of divorce on household behavior. It tests whether a divorce law that mandates an economic compensation be transferred to a homemaker upon divorce gives more bargaining power to wives in married couple households by identifying the effects of the law on children's education. Based on previous literature on gender and intrahousehold allocation, it assumes that women invest in household public goods, like children's education, at higher rates than men. Using the bargaining household model framework developed by Manser and Brown (1980) and McElroy and Horney (1981), this paper provides evidence that the legalization of divorce, via an increase in the opportunity cost of remaining married for wives, actually increased school enrollment for children within married couple households in Chile, specifically for children of secondary school age. Additionally, it shows that exogenous administrative processes to obtaining a divorce also influence household bargaining power and resource allocation by altering the credible threat of divorce.

Legalizing divorce has had a large and significant effect on school enrollment for high school students from married parent families in Chile. While other macro-level factors decreased school enrollment in 2006, legalizing divorce caused an approximately 52 percent increase in school enrollment for married parent children on average (holding all else constant). These results show that family policies and laws favoring women can have positive, unintended consequences on families and investments in households.

Family policies created for one specific group can have unintended or unexpected effects on other groups. In this case, divorce legislation was created for unstable families, but this paper has shown that it influences resource allocation decisions in stable family households. It has also shown that family policies providing more bargaining power to women have the potential to increase investments in household goods that women value. Although this study analyzes the effect of legalizing divorce, it can also be argued that changes to divorce laws and family policies that empower women by increasing their bargaining power within the marriage could have similar effects. In this sense, my results are not just specific to the case of Chile but have implications for many countries.

References

Alderman, H., Chiappori, P., Haddad, L., Hoddinott, J., & Kanbur, R. 1995. "Unitary Versus Collective Models of the Household: Is It Time to Shift the Burden of Proof?" *World Bank Research Observer*, 10(1): 1-19.

Angrist, J., D., & Pischke, J. S. 2009. *Mostly Harmless Econometrics: An Empiricist's Companion.* Princeton, New Jersey: Princeton University Press.

Becker, G. S. 1973. "A Theory of Marriage: Part I." *Journal of Political Economy*, 81(4): 813-846.

Becker, G. S. 1974. "A Theory of Marriage: Part II." *Journal of Political Economy*, 82(2): S11-S26.

Becker, G. S. 1981. *A Treatise on the Family*. London: Harvard University Press.

Behrman, J. R. 1997. "Intrahousehold Distribution and the Family". In M. Rosenzweig and O. Stark (Eds.), *Handbook of Population and Family Economics* (pp. 125-187). Netherlands: Elsevier.

Bergstrom, T. C. 1997. "A Survey of Theories of the Family." In M. Rosenzweig and O. Stark (Eds.), *Handbook of Population and Family Economics* (pp. 21-79). Netherlands: Elsevier.

Bravo, D. 2009. *Encuesta de Proteccion Social: Panel 2002, 2004, 2006*. Santiago, Chile: Universidad de Chile, Centro de Microdatos.

Browning, M., & Chiappori, P. A. 1998. "Efficient Intrahousehold Allocations: A General Characterization and Empirical Tests." *Econometrica*, 66(6), 1241-1278.

Bumpass, L., & Rindfuss, R. R. 1979. "Children's Experience of Marital Disruption." *American Journal of Sociology*, 85(1): 49-65.

Burstein, N. R. 2007. "Economic Influences on Marriage and Divorce." *Journal of Policy Analysis and Management*, 26(2): 387-429.

Cameron, A. C., & Trivedi, P. K. 2009. *Microeconomics using Stata.* College Station, Texas: Stata Press.

Chiappori, P. 1988. "Rational Household Labor Supply." *Econometrica*, 56(1): 63-90.

Chiappori, P., Fortin, B., & Lacroix, G. 2002. "Marriage Market, Divorce Legislation, and Household Labor Supply." *Journal of Political Economy*, 110(1): 37-72.

Duncan, G. J., & Rodgers, W. 1987. "Single-Parent Families: Are Their Economic Problems Transitory or Persistent." *Family Planning Perspectives*, 19(4): 171-178.

Ermisch, J. 2003. *An Economic Analysis of the Family*. Princeton, New Jersey: Princeton University Press.

Fortin, B., & Lacroix, G. 1997. “A Test of the Unitary and Collective Models of Household Labour Supply.” *Economic Journal*, 107(443): 933-955.

Goode, W. J. 1993. *World Changes in Divorce Patterns*. New Haven: Yale University Press.

Gray, J. S. 1998. “Divorce-Law Changes, Household Bargaining, and Married Women's Labor Supply.” *American Economic Review*, 88(3): 628-642.

Haddad, L., Hoddinott, J., & Alderman, H. 1997. “Introduction: The Scope of Intrahousehold Resource Allocation Issues.” In L. Haddad, J. Hoddinott & H. Alderman (Eds.), *Intrahousehold Resource Allocation in Developing Countries: Models, Methods, and Policy* (pp. 1-16). Baltimore, Washington, D.C.: International Food Policy Research Institute; distributed by Johns Hopkins University Press.

Lundberg, S., & Pollak, R. A. 1993. “Separate Spheres Bargaining and the Marriage Market.” *Journal of Political Economy*, 101(6): 988-1010.

Lundberg, S., & Pollak, R. A. 1994. “Noncooperative Bargaining Models of Marriage.” *American Economic Review*, 84(2): 132-137.

Lundberg, S., & Pollak, R. A. 1996. “Bargaining and Distribution in Marriage.” *Journal of Economic Perspectives*, 10(4), 139-158.

Manser, M., & Brown, M. 1980. “Marriage and Household Decision-Making: A Bargaining Analysis.” *International Economic Review*, 21(1): 31-44.

Martinez, C. A. 2007. *Intrahousehold Allocation and Bargaining Power: Evidence from Chile.* Unpublished Ph.D., University of Michigan, Ann Arbor.

McElroy, M. B. 1990. “The Empirical Content of Nash-Bargained Household Behavior.” *Journal of Human Resources*, 25(4): 559-583.

McElroy, M. B. 1997. “The Policy Implications of Family Bargaining and Marriage Markets.” In L. Haddad, J. Hoddinott & H. Alderman eds (Eds.), *Intrahousehold Resource Allocation in Developing Countries: Models, Methods, and Policy* (pp. 53-74). Baltimore and London: Johns Hopkins University Press for the International Food Policy Research Institute.

McElroy, M. B., & Horney, M. J. 1981. “Nash-Bargained Household Decisions: Toward a Generalization of the Theory of Demand.” *International Economic Review*, 22(2): 333-349.

Nash, J. 1953. “Two-Person Cooperative Games.” *Econometrica*, 21(1): 128-140.

Quisumbing, A. R., & Maluccio, J. A. 1999. "Intrahousehold Allocation and Gender Relations: New Empirical Evidence." *Policy Research Report on Gender and Development, Working Paper Series no. 2*. Washington, DC: International Food Policy Research Institute.

Rangel, M. A. 2006. "Alimony Rights and Intrahousehold Allocation of Resources: Evidence from Brazil." *Economic Journal*, 116(513): 627-658.

Rubalcava, L., Teruel, G., & Thomas, D. 2004. "Spending, Saving and Public Transfers Paid to Women." Working Paper Series CCPR-024-04. Los Angeles: California Center for Population Research, University of California.

Schultz, T. P. 1990. "Testing the Neoclassical Model of Family Labor Supply and Fertility." *Journal of Human Resources*, 25(4): 599-634.

Shady, N., & Rosero, J. 2007. "Are Cash Transfers Made to Women Spent Like Other Sources of Income?" Policy Research Working Paper no. 4282. Washington, DC: World Bank.

Thomas, D. 1990. "Intrahousehold Resource Allocation an Inferential Approach." *Journal of Human Resources*, 25(4): 635-664.

Wooldridge, J. M. 2002. *Econometric Analysis of Cross Section and Panel Data*. Cambridge, Massachusetts: MIT Press.

Table 1. Descriptives of School Age Children (5 to 21) by Legal Civil Status of Parent, 2002 to 2006

	Parents are:	
	Married	Cohabitating
2002		
Total Number	4,203	897
Percent in School by Age Category:		
Age 5 to 11	94.0	93.9
Age 12 to 17	95.2	93.7
Age 18 to 21	44.5	43.8
Average Age	11.8	11.9
2004		
Total Number	4,274	894
Percent in School by Age Category:		
Age 5 to 11	92.6	92.2
Age 12 to 17	95.6	96.1
Age 18 to 21	45.6	40.0
Average Age	13.0	12.9
2006		
Total Number	4,182	907
Percent in School by Age Category:		
Age 5 to 11	93.1	91.0
Age 12 to 17	96.0	92.5
Age 18 to 21	50.1	36.0
Average Age	13.9	13.7

Source: Encuesta de Protección Social (EPS), 2009

Table 2. Logit Regression of School Attendance, National Sample, 2002 to 2006†

	Model One		Model Two	
	β		β	
Year dummies				
2002	reference		reference	
2004	-0.2485		-0.1022	
	(0.1602)		(0.1960)	
2006	-0.3480	**	-0.3174	*
	(0.1491)		(0.1854)	
Interaction terms				
Married parents in 2002	reference		reference	
Married parents in 2004	-0.0621		0.0271	
	(0.1765)		(0.2158)	
Married parents in 2006	0.2640		0.4274	**
	(0.1662)		0.2064	
Age groups				
Ages 4 to 10	–		reference	
Ages 11 to 17	–		0.5047	***
			0.0927	
Ages 18 to 21	–		-3.1355	***
			0.1205	
Log likelihood	-2571.52		-1633.53	
N observations	7365		7365	
N groups	1053		1053	

Standard errors are in parenthesis [* = significant at p<0.10, ** = significant at p<0.05, and *** = significant at p<0.01].

† All models include household-level fixed effects.

Source: Encuesta de Protección Social (EPS), 2009

Table 3. Logit Regression of School Attendance, Urban Sample, 2002 to 2006†

	Model One	Model Two	Model Three
	β	β	β
Year dummies			
2002	reference	reference	reference
2004	-0.0948	0.1994	0.1991
	(0.1894)	(0.2340)	(0.2340)
2006	-0.4825 ***	-0.5531 **	-0.5535 **
	(0.1774)	(0.2207)	(0.2208)
Interaction terms			
Married parents in 2002	reference	reference	reference
Married parents in 2004	-0.2124	-0.2601	-0.2604
	(0.2074)	(0.2548)	(0.2549)
Married parents in 2006	0.4083 **	0.6765 ***	1.4451 ***
	(0.1950)	(0.2435)	(0.4623)
Age groups			
Ages 4 to 10	–	reference	reference
Ages 11 to 17	–	0.5539 ***	0.5577 ***
		(0.1065)	(0.1065)
Ages 18 to 21	–	-3.0995 ***	-3.1000 ***
		(0.1354)	(0.1354)
Administrative changes			
Average wait time for married couples in 2006	–	–	-0.0579 *
			(0.0294)
Log likelihood	-1977.58	-1264.47	-1262.54
N observations	5618	5618	5618
N groups	808	808	808

Standard errors are in parenthesis [* = significant at p<0.10, ** = significant at p<0.05, and *** = significant at p<0.01].

† All models include household-level fixed effects.

Source: Encuesta de Protección Social (EPS), 2009

Table 4. Marginal Effects for Logit Regression of School Attendance, National Sample, 2002 to 2006^

	Model One	Model Two
	dy/dx	dy/dx
Year dummies		
2002	reference	reference
2004	-0.0614	-0.0255
	(0.0390)	(0.0490)
2006	-0.0866 **	-0.0778 *
	(0.0368)	(0.0443)
Interaction terms		
Married parents in 2002	reference	reference
Married parents in 2004	-0.0155	0.0067
	(0.0440)	0.0537
Married parents in 2006	0.0651	0.1062 **
	(0.0405)	0.0505
Age groups		
Ages 4 to 10	–	reference
Ages 11 to 17	–	0.1215 ***
		0.0214
Ages 18 to 21	–	-0.4812 ***
		0.0201

Standard errors are in parenthesis [* = significant at p<0.10, ** = significant at p<0.05, and *** = significant at p<0.01].

^ Marginal effects are calculated for a primary school aged child in 2006 living with married parents in a family court district with an average wait time of 4 months.

Source: Encuesta de Protección Social (EPS), 2009

Table 5. Marginal Effects for Logit Regression of School Attendance, Urban Sample, 2002 to 2006^

	Model One	Model Two	Model Three
	dy/dx	dy/dx	dy/dx
Year dummies			
2002	reference	reference	reference
2004	-0.0236	0.0492	0.0403
	(0.0470)	(0.0570)	(0.0458)
2006	-0.1192 **	-0.1321 ***	-0.1030 **
	(0.0426)	(0.0494)	(0.0403)
Interaction terms			
Married parents in 2002	reference	reference	reference
Married parents in 2004	-0.0526	-0.0649	-0.0576
	(0.0509)	(0.0634)	(0.0589)
Married parents in 2006	0.0998 **	0.1657 ***	0.3454 ***
	(0.0465)	(0.0572)	(0.0999)
Age groups			
Ages 4 to 10	–	reference	reference
Ages 11 to 17	–	0.1323 ***	0.1037 ***
		(0.0242)	(0.0259)
Ages 18 to 21	–	-0.4823 ***	-0.6032 ***
		(0.0227)	(0.0496)
Administrative changes			
Average wait time for married couples in 2006	–	–	-0.0122 *
			(0.0062)

Standard errors are in parenthesis [* = significant at p<0.10, ** = significant at p<0.05, and *** = significant at p<0.01].

^ Marginal effects are calculated for a primary school aged child in 2006 living with married parents in a family court district with an average wait time of 4 months.

Source: Encuesta de Protección Social (EPS), 2009

Table 6. Logit Regression of School Attendance (*Model One* from Table 2) by School Age Group and without Wait Times, National Sample, 2002 to 2006†

		Primary school	Secondary school	Tertiary school
		β	β	β
Year dummies				
	2002	reference	reference	reference
	2004	-0.1933	0.0890	-2.7539 **
		(0.3113)	(0.4038)	(1.0628)
	2006	0.3680	-1.0352 ***	-0.7817 *
		(0.3473)	(0.3482)	(0.4615)
Interaction terms				
	Married parents in 2002	reference	reference	reference
	Married parents in 2004	0.1027	-0.3444	2.6813 **
		(0.3417)	(0.4465)	(1.0837)
	Married parents in 2006	0.0865	0.9583 **	0.5398
		(0.3849)	(0.3864)	(0.4950)
Log likelihood		-463.55	-335.88	-209.86
N observations		1253	983	666
N groups		299	248	238

Standard errors are in parenthesis [* = significant at p<0.10, ** = significant at p<0.05, and *** = significant at p<0.01].

† All models include household-level fixed effects.

Source: Encuesta de Protección Social (EPS), 2009

Table 7. Logit Regression of School Attendance (*Model Three* from Table 3) by School Age Group and with Wait Times, Urban Sample, 2002 to 2006†

	Primary school	Secondary school	Tertiary school
	β	β	β
Year dummies			
2002	reference	reference	reference
2004	0.0738	0.4546	-1.5855
	(0.3976)	(0.4842)	(1.1822)
2006	0.2668	-1.2463 ***	-1.0424 **
	(0.4226)	(0.4351)	(0.5268)
Interaction terms			
Married parents in 2002	reference	reference	reference
Married parents in 2004	-0.0142	-0.4635	1.5809
	(0.4317)	(0.5285)	(1.2192)
Married parents in 2006	0.7854	2.3683 ***	1.1640
	(0.8656)	(0.8603)	(1.0550)
Administrative changes			
Average wait time for married couples in 2006	-0.0493	-0.1008 *	-0.0229
	(0.0546)	(0.0546)	(0.0657)
Log likelihood	-338.39	-253.3397	-141.33
N observations	899	756	426
N groups	205	192	149

Standard errors are in parenthesis [* = significant at p<0.10, ** = significant at p<0.05, and *** = significant at p<0.01].

† All models include household-level fixed effects.

Source: Encuesta de Protección Social (EPS), 2009

Table 8. Marginal Effects for Logit Regression of School Attendance by School Age Group and with Wait Times, Urban Sample, 2002 to 2006^

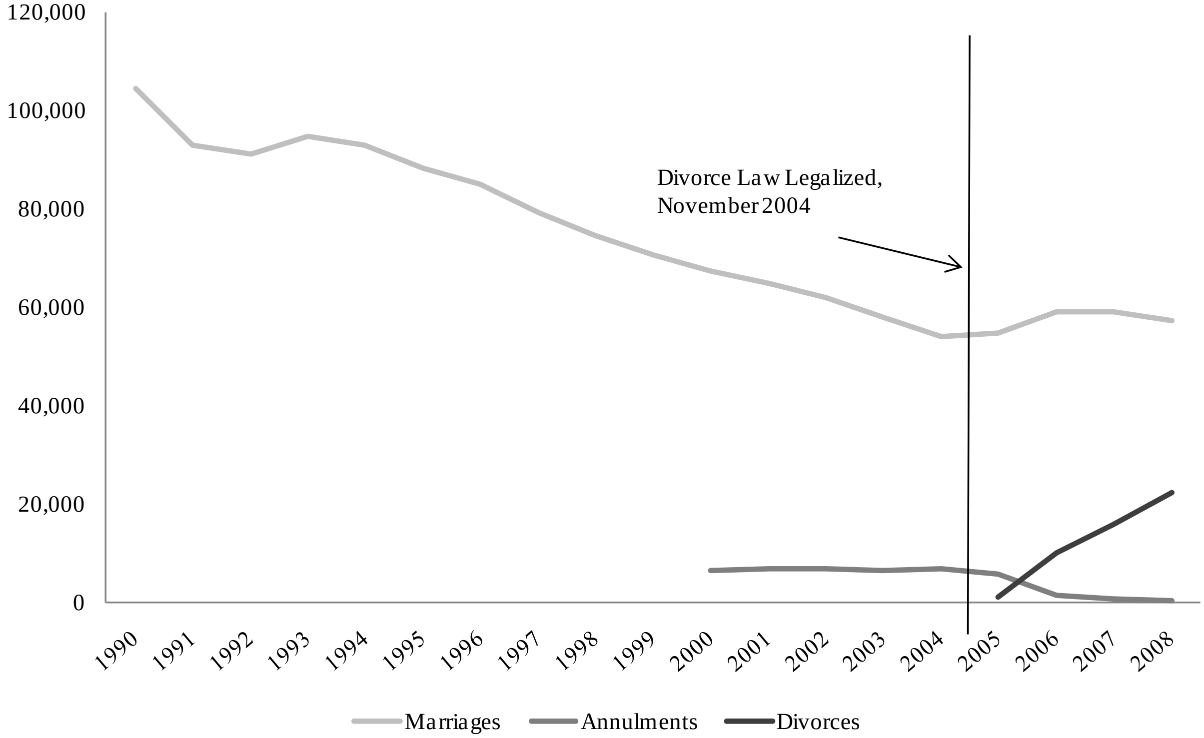
	Primary school	Secondary school	Tertiary school
	dy/dx	dy/dx	dy/dx
Year dummies			
2002	reference	reference	reference
2004	0.0152	0.0867	-0.3332 *
	(0.0808)	(0.0904)	(0.1840)
2006	0.0587	-0.2096 ***	-0.2376 **
	(0.0976)	(0.0768)	(0.1076)
Interaction terms			
Married parents in 2002	reference	reference	reference
Married parents in 2004	-0.0030	-0.1024	0.3260 *
	(0.0908)	(0.1289)	(0.1870)
Married parents in 2006	0.1842	0.5172 ***	0.2641
	(0.1973)	(0.1556)	(0.2328)
Administrative changes			
Average wait time for married couples in 2006	-0.0103	-0.0222 ***	-0.0057
	(0.0114)	(0.0083)	(0.0164)

Standard errors are in parenthesis [* = significant at p<0.10, ** = significant at p<0.05, and *** = significant at p<0.01].

^ Marginal effects are calculated for a child in 2006 living with married parents in a family court district with an average wait time of 4 months.

Source: Encuesta de Protección Social (EPS), 2009

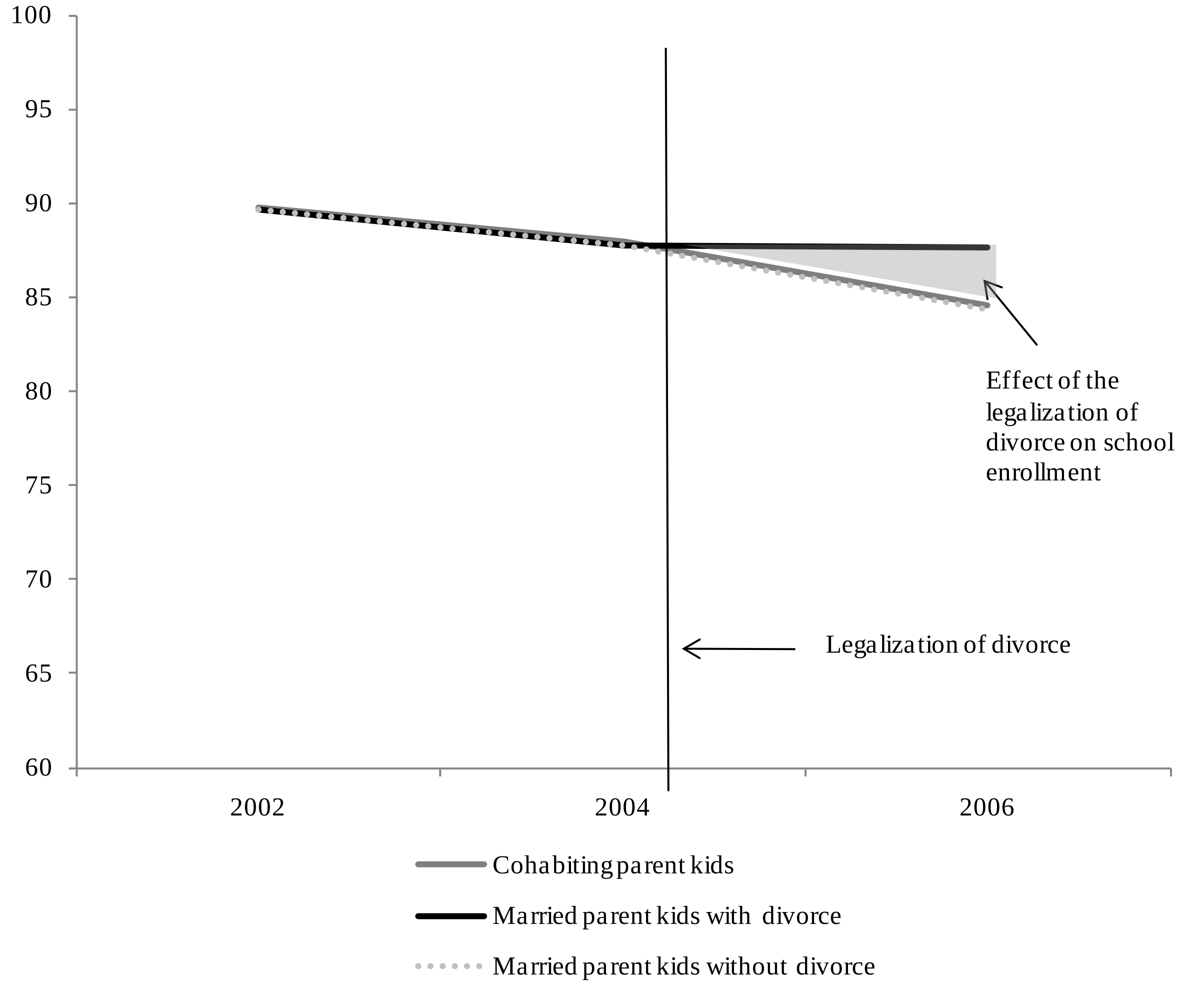
Chart 1. Number of marriages, annulments, and divorces in Chile, 1990 – 2008[17]



Source: Ministerio de Justicia, Servicio de Registro Civil E Identificacion (http://www.srcei.cl/f_estadisticas.html)

[17] Data on annulments is not publicly available before the year 2000. Although divorce was legalized in 2004, the implementation of the law began in 2005.

Chart 2. Percent of school age children attending school by parental legal civil status and year, 2002 to 2006



Source: Encuesta de Protección Social (EPS), 2009