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Institute Working Paper 27
October 2019

DOI: https://doi.org/10.21034/iwp.27
Keywords: Household bargaining; Education; Child welfare; Divorce; Family law; Difference-in-difference estimation
JEL Codes: D12, D13, J1, J12, I21, H0
The views expressed herein are those of the author and not those of the Federal Reserve Bank of Minneapolis, the Federal Reserve System, or the U.S. Census Bureau. The author has no competing interests to declare.

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Improving Child Welfare in Middle Income Countries: The Unintended Consequence of a Pro-Homemaker Divorce Law and Wait Time to Divorce

By MISTY L. HEGGENESS [1]

This study identifies the impact of access to and the speed of divorce on the welfare of children in a middle income largely Catholic country. Using difference-in-difference estimation techniques, I compare school enrollment for children of married and cohabiting parent households before and after the legalization of divorce. Implementing pro-homemaker divorce laws increased school enrollment anywhere from 3.4 to 5.5 percentage points, and the effect was particularly salient on secondary school students. I provide evidence that administrative processes influencing the speed of divorce affect household bargaining and investments in schooling. With every additional six months wait to the finalization of divorce, school enrollment decreased by approximately one percentage point. The impact almost doubles for secondary schooling. When contemplating development policies, advocates, policymakers, and leaders should not overlook the impact changes in family policies and administrative processes can have on advancements in child welfare and, ultimately, economic development.

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"…any divorce-law change that alters the financial well-being of divorcing women and their children will also impact the welfare of individuals in families that do not dissolve…these indirect effects should not be ignored when designing effective social and economic policies (p. 639)."

– Jeffrey S. Gray, AER 1998

I. Introduction

Implementing development policies focused on improving child welfare can be expensive, time consuming, and, sometimes, ineffective (Filmer 2003, Glewwe and Kremer 2006, Ingram and Kessides 1994). This is particularly true if the intervention is country-wide and focused on major activities like building infrastructure, purchasing supplies, or developing advocacy groups (Ingram and Kessides 1994). Can alternative policy paths like changes to family law lead to major advancements in child welfare and, eventually, economic development? To study this, one needs a rarefied environment where a national policy is implemented with enough leverage to induce a redistribution of household resources in certain households, restrict the household's ability to manipulate the magnitude or speed of the redistribution, and data overtime of those exposed and not exposed to the policy to estimate outcomes. I find a natural experiment environment close to this in Chile where gendered family norms are relatively rigid, geographic immobility is common, family legal procedures are tied to the local geography where one lives and local family court districts are independent, and a major policy shock – the legalization of divorce – happened in 2004.

After almost a decade of intense national debate, the Chilean Congress passed a revised Civil Marriage Act in 2004. For the first time in the country's history, Chileans could divorce. The Act, progressive in nature, included a requirement that breadwinners (mostly men) provide an *economic compensation*[2] to homemakers (almost always women) equivalent to lost wages incurred while engaging in home production during the marriage. Prior research shows that shifting resources to homemakers increases their bargaining power within marriage (Lundberg, Pollak, and Wales 1996, Voena 2015, Wong 2016), and evidence exists that, on average, women invest more in household goods, such as children's education and clothing, than men (Lundberg, Pollak, and Wales 1996, Quisumbing and Maluccio 1999, Rubalcava et al. 2004, Rangel 2006, Schady and Rosero 2007, Nunley and Seals 2011). Additionally, Chilean social networks and family structures are such that "…men [within the household] exercise overt and subtle forms of control over family monetary allocation, spending choices and earnings strategies (Stillerman 2004)." A shift in resources away from men into the hands of women could change household consumption and investment patterns towards woman's preferences.

In this paper, I study the impact of two changes: the introduction of divorce with *economic compensation* and exogenous variation in geographically local wait times to finalize a divorce[3]. To shed light on the question of whether family policies in middle income countries can accelerate development, I explore school enrollment under the new family law regime. I build on an already existing trove of literature on the effects of divorce and expand it by demonstrating that both the advancement of policies that shift property rights to family members who invest in household goods like children's education and the bureaucratic idiosyncrasies involved in implementing said policy, influence household bargaining and can accelerate or deter economic development. I use a natural experiment in exogenous variation of family court's average length of time to divorce to study this phenomenon.

[2] The economic compensation was paid as a cash lump sum or in regular installments until paid in full. Some breadwinners converted the lump sum cash payment into a property transfer (e.g. rights to full ownership of the family house) (Cox 2011).

[3] Individuals could not manipulate their environment to invoke shorter wait times, described in more detail later in the paper.

II. Background

A. Household Bargaining Models, Family Law, and Child Outcomes

Unitary models assume a household maximizes a single, well defined utility function subject to a household budget constraint where preferences are decided by consensus among household members or determined by a benevolent dictator (Samuelson 1956, Becker 1981). In the late 20th century, economists dissatisfied with the unitary model's simplicity began developing models in which household members bargained over consumption decisions based on the resources they controlled or the separate spheres they occupied (Manser and Brown 1980, McElroy and Horney 1981, Lundberg & Pollak 1993) and this brought about a groundswell of literature advancing the notion that the unitary model of household utility does not accurately describe household economic behavior (Lundberg & Pollak 1994, Lundberg & Pollak 1995, Alderman et al. 1995, Behrman 1997, Bergstrom 1997, Gray 1998, Chiappori et al. 2002, Ermisch 2003). Empirical studies have, for the most part, provided evidence in support of cooperative bargaining models (Schultz 1990, Thomas 1990, Lundberg, Pollak, and Wales 1996, Fortin and Lacroix 1997, Browning and Chiappori 1998, Rangel 2006), and, by now, enough evidence has accumulated demonstrating that household consumption differs depending on who is making the decision and who controls the resources.

The impact of divorce laws on divorce rates and marital instability has been a topic studied extensively within the U.S. Gray (1998) found the expansion to unilateral divorce in the 1970s and related reassignment of property rights within marriage did not affect divorce rates, arguing that instead within household transfers may have taken place to prevent increased marital dissolution. Gruber (2004), on the other hand, used four decades of decennial data to show that unilateral divorce increased the incident of divorce, and Wolfers (2006) showed that while divorce rates rose in the short run in response to unilateral divorce laws, the rise dissipated overtime. He found that changes in family law explain little of the rise in divorce in the U.S. during the late 20th century. In a slightly different study, Wong (2016) used state variation in U.S. "homemaking" provisions of the 1980s to show that reinforcement of wives' post-divorce property rights increased her bargaining power and, subsequently increased marriages. While the overall impact of divorce law on marital instability is relevant, in this study I focus instead on household bargaining among married couples within the context of a shift in outside options due to the legalization of divorce in Chile and not on changes to the incidence of divorce.[4]

Most research on household bargaining models within the context of divorce tests the changing responsiveness of female labor supply to changes in family law, showing mixed results. All else equal, Fernández and Wong (2014) found increasing divorce risk associated with married couples investing less in joint household savings and wives working more. Bargain et al. (2012) examined the effect of legalizing divorce in Ireland in 1996 and found, compared to U.S. studies, a larger magnitude increase in female labor supply attributable to divorce. They argue that the larger magnitude is driven by differences between expanding already existing laws compared to creating a divorce law where none previously existed.

While Gray (1998) found no evidence on divorce rates, he did find evidence consistent with a standard household bargaining model where women's labor force participation and leisure time increased as a function of increasing bargaining power. Gray's study is particularly relevant because he examined the expansion of

[4] The reasons for this are twofold: not enough time had passed post the introduction of divorce in my data to analyze an impact on divorce rates, and, since there was no divorce prior to the legalization of divorce, during the first years of after the implementation of divorce not many families divorced. The sample of divorced adults is small in the immediate years following its legalization.

divorce within the context of varying degrees of property laws that reallocated or redistributed resources upon divorce. He highlighted an important policy implication for this study, which is that "…any divorce-law change that alters the financial well-being of divorcing women and their children will also impact the welfare of individuals in families that do not dissolve…these indirect effects should not be ignored when designing effective social and economic policies (p. 639)."

Stevenson (2007) also found that unilateral divorce induced spouses to invest less in marriage-specific capital like home production and showed an increase in wives' labor force participation and decrease in fertility. Findings on fertility, however, can be difficult to interpret. Divorce reform in China reduced the likelihood of having a son after a firstborn child, attributable to an increase in women's empowerment within marriage due to improved outside options (Sun and Zhao 2016). The introduction of Chile's divorce law decreased the age at first birth for highly educated women (Gallegos and Ondrich 2017). While this might seem contradictory, the authors argued that highly education women had the most to gain from *economic compensation* within the new divorce law. High-skilled married women post-divorce law had less to lose from leaving the labor market since their home production could be compensated for upon divorce.

The responsiveness of labor supply has been shown particularly sensitive for mothers (Nunley and Seals 2011, Genadek et al. 2007). Nunley and Seals found that the expansion of joint custody laws in the U.S. transferring bargaining power away from married mothers increased female labor supply. Genadek et al. found the labor supply of mothers more responsive to no-fault divorce and property division rules in the U.S. than for married women without children.

While there is a rich literature on divorce rates and the effect of divorce on female labor supply and fertility, less is known about what other impacts family law has on intrahousehold bargaining of other household outcomes important to economic development, such as health, education, and child welfare. While there is evidence that easier access to divorce decreases female suicide, domestic violence, and possibly female homicides (Stevenson and Wolfers 2006), easier access to divorce has also been shown to increase the odds of adult suicide of the children of divorced parents (Gruber 2004).

Child health and education are future household investments in the form of informal social security for both parents in old age. Gruber (2004) found that in the long run unilateral divorce decreased children's schooling. Nunley and Seals (2011) found a decrease in investments in private school education for children in married parent families in states favoring joint custody instead of sole custody to mothers upon divorce.

While U.S. studies provide conflicting evidence on the impact of women's increased bargaining power on children's education, there is evidence that, at least in the short run, health and education investments in other countries increase when women gain more bargaining power within the household (Quisumbing and Maluccio 1999, Rubalcava et al. 2004, Rangel 2006, Schady and Rosero 2007, Martínez 2013). Two studies in South America take advantage of alimony and child support law changes to evaluate the impact on education. Rangel (2006) studied an expanded alimony law to cohabitating women in Brazil and found that the expansion increased schooling of first-born girls. Martínez (2013) showed how changes in child support for out-of-wedlock children in Chile increased school attendance for high school aged children by 3.6 percentage points, for primary school children by 1.0 percentage point, and for preschool children by 2.4 percentage points. She also found a decrease in father's employment and hours worked.

Family policies that transfer resources to women have generally been found to positively influence marriage rates, married women's labor supply and health, and investments in children. While studies in South America have shown a positive benefit to children's education of expanding alimony rights and child support, there have been no studies to date in middle income countries specifically testing whether access to progressive divorce options can influence development indicators like education. To my knowledge, this is the first study examining whether reforms that introduce access to divorce with substantial *economic compensation* affects investments in children's education and whether the administrative process to finalize a divorce matters.

B. The Case of Chile

Marital Instability.—Married couples wishing to dissolve their relationship before 2004 had two options: separation (while remaining legally married) or a legal annulment (Haas 2010, Cox 2011).[5] Informal separation could leave custodial parents economically vulnerable. A partner could request a legal separation and child support via the court, but this rarely occurred. There were no more than 70 legal separation cases annually between 2005 and 2008.[6]

The Chilean process for legal annulment was such that, "[according to law] …couples must marry at the civil registrar in the home district of the man or woman. The easiest way to [then] void the marriage contract is to have witnesses testify that at the time of the marriage neither party actually lived in the district where they registered (p. 127, Haas 2010)." Spouses could only annul if they agreed to cooperate with each other to report or manufacture an inaccuracy in their marriage certificate and they had the necessary financial resources to pay for the annulment (Haas 2010, Cox 2011). Before divorce became legal, spouses wanting to end their relationship but choosing not to cooperate with each other or not having

[5] See: https://www.nytimes.com/2005/01/30/weekinreview/divorce-ties-chile-in-knots.html (accessed on November 18, 2018).

[6] Data on legal separations: http://www.registrocivil.cl/f_estadisticas.html (accessed on May 01, 2009).

the necessary finances could informally separate, but they remained legally married.

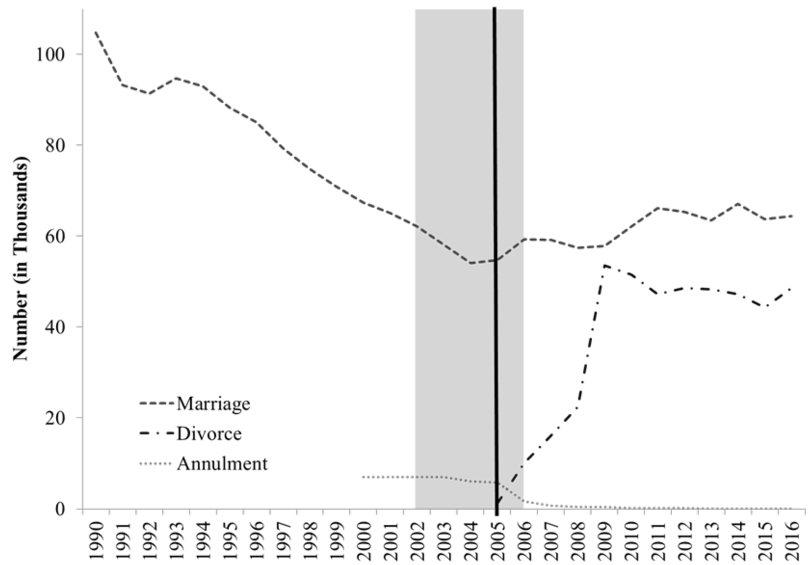


FIGURE 1. NUMBER OF MARRIAGES, ANNULMENTS, AND DIVORCES IN CHILE

Sources: Ministerio de Justicia, Servicio de Registro Civil E Identificación

Notes: The grey vertical area represents the time period of this study. The black vertical line is the first year divorce became available through the judicial system.

Creating a Pro-Homemaker Divorce Law.—Legalizing divorce in Chile, a conservative, middle-income country in South America, had been a goal of women's advocates and left-to-center-leaning politicians throughout the 1990s and into the early 2000s (Haas 2010). By the turn of the 21st century, Chile was only one of three countries with no formal legal process of divorce, forcing unstable married couples to maintain legal marriage arrangements long after physical separations.[7] In May 2004, after years of continuous decline in the number of new

[7] For a more detailed description, see https://foreignpolicy.com/2015/01/19/the-last-country-in-the-world-where-divorce-is-illegal-philippines-catholic-church/ (accessed on November 18, 2018).

marriages (see Figure 1) and numerous failed attempts to legalize divorce, the Chilean Congress passed the country's first divorce law (Rohter 2005, Haas 2010).[8] Because of repeated failures over the course of a decade and, due to national media attention, Chileans were aware of the efforts but had no guarantee that in 2004 the law would actually pass (Cox 2011).[9]

With divorce, came the right to a civil law second marriage, child support for children under age 22 (or under age 29 if in college), and financial compensation for household production during the marriage (Cox 2011). The financial compensation, called *economic compensation*, was a lump sum of money to be paid all at once[10] or in monthly installments to the homemaker from the breadwinner until the entire amount was paid in full. Cox notes that "…strictly speaking, according to the law, nothing indicates that only women can receive [*economic compensation*] from their ex-[spouses], but in practice men qualify for it only under very peculiar situations and the cases are very rare (p. 156)" Judges calculated the payment based on the assumed lost wages of the homemaker spouse. A reservation wage based on the homemaker's education, family background, socioeconomic factors, and market rates for various occupations was estimated. It was then multiplied by the number of years married during which the homemaker stayed at home taking care of the family.[11]

The Establishment of Family Courts.—The Chilean family court system was established the same year under Chilean Law No. 19.968 (Haas 2010).[12] The startup

[8] See also: https://www.bcn.cl/leyfacil/recurso/divorcio (accessed on October 17, 2018).

[9] For additional information, see: http://www.nbcnews.com/id/6523667/ns/world_news/t/chile-enacts-first-divorce-law-its-history/#.W_iijeKQzIV (accessed on November 18, 2018).

[10] In some instances, the lump sum payment would transform into the wife getting the house or other big assets upon divorce.

[11] Over time, the technique used to calculate economic compensation changed. 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.

[12] For more information, see: https://www.oas.org/dil/esp/Ley_19968_Tribunales_familia_Chile.pdf (accessed on December 9, 2018).

of family courts was seen as somewhat chaotic and the judicial system eventually reorganized family courts into a more coherent system in later years. But, in 2004, each court established itself autonomously (Haas 2010). Administrators and judges were charged with developing efficient procedures and protocols for their court. Each family court was composed of a small group of *comunas* [cities or townships]. Generally ranging from one to nine *comunas* per family court district; the average was three or four. An exception was the capital city of Santiago, in which one family court district encompassed 19 *comunas* and where 10 diverse family court systems existed.

All cases related to family legal issues, including cases of inheritance, domestic violence, child custody, adoptions, paternity cases, abandonment, child neglect and abuse, juvenile delinquency, and adjudication cases were transferred to the family courts upon creation. Divorce cases were submitted and finalized within the new family court system; they were a minority of all cases. The independent creation of each court and their administrative management procedures and protocols for handling cases, in addition to the diversity and magnitude of caseloads within each court, independence and working speed of judges, and general management methods employed by the court administrator and his/her staff all influenced wait time for finalizing a divorce.

Geographic Constraints.—Chileans rarely move far from their nuclear family and usually marry and start a family in the same township or neighborhood where they themselves were raised. Araos-Bralic (2015) found through ethnographic studies in Chile that both members from poor and well-off families live in close geographic proximity to their descendant groups (of at least three-generations), a phenomenon locally called “allegamiento.” The Chilean divorce law stipulated that when couples decided to divorce, they could only divorce in the family court corresponding to the county or address where they lived.[13] This stipulation is critical to this study because it restricts the ability of individuals to manipulate the length of time to divorce.

TABLE 1—DESCRIPTIVE STATISTICS OF PARENTS BY MARITAL STATUS, 2002

	Married Parents	Cohabiting Parents	diff.	t-stat
Demographics				
Average age (all parents)	41.7	38.4	-3.27 ***	-14.86
	(0.004)	(0.009)		
Household composition				
Female homemaker (age 25 to 54 and not in labor force) (%)	54.5	48.0	-6.45 ***	-3.98
	(0.031)	(0.069)		
Male homemaker (age 25 to 54 and not in labor force) (%)	2.5	2.8	0.00	0.30
	(0.010)	(0.023)		
Average number of children	1.9	1.8	-0.06 **	-3.23
	(0.000)	(0.001)		
Characteristics influencing bargaining power				
Educational attainment of female homemakers (age 25 plus and not in labor force)				
No schooling	1.3	3.0	1.65 **	2.90
	(0.009)	(0.033)		
Some schooling, no high school diploma	41.1	53.1	11.80 ***	5.34
	(0.040)	(0.097)		
High school, diploma	47.2	39.6	-8.43 ***	-3.80
	(0.040)	(0.095)		
Some technical college, university, or more	10.0	3.7	-5.17 ***	-4.25
	(0.024)	(0.037)		
Average hours worked (for those age 25 to 54 in the labor force)				
Female working parent	43.0	41.3	-1.68 *	-2.26
	(0.014)	(0.033)		
Male working parent	50.4	50.0	-0.31	-0.71
	(0.008)	(0.019)		
Median Monthly Earnings $USD (for those working 35 hours or more per week)				
Female working parent	$296.52	$237.22		
Male working parent	$355.82	$296.52		
Total Sample	12,275	2,639		

Source: Author calculations. Encuesta de Protección Social (EPS), Chile.

Breadwinning, Homemaking, and Household Relationships.—Prior to the Civil Marriage Act of 2004, over half (54.5 percent) of all married mothers and 48.0 percent of cohabiting mothers were not in the labor force (Table 1). While cohabiting couples may not be a perfect counterfactual to married couples, on

[13] See: https://www.leychile.cl/Navegar?idNorma=225128&idParte=0 (accessed on December 9, 2018).

average, the magnitude of differences between then on most indicators is modest. Cohabiting couples are slightly younger. Cohabiting women have a somewhat closer attachment to the labor market (on the extensive margin) but work less hours per week than married mothers (intensive margin). Cohabiters appear to earn less and have lower education. That said, both family types have around the same number of children (two). Fathers are rarely homemaker, almost always breadwinners, and strongly attached to the labor market – were they appear to work a lot (50 hours per week).

III. Data

A. Encuesta de Protección Social (EPS)

I use panel data from the Chilean Encuesta de Protección Social[14] (EPS) combined with administrative records on divorce from the Chilean Supreme Court. The EPS data used in this study come from three waves of the survey (2002, 2004, and 2006) and follow the same individuals and their representative households over time. Since the original intent of the survey was to collect labor and social security pension fund data, the first wave is nationally representative of all individuals who contributed to a public pension fund. The 2004 and 2006 waves are nationally representative samples of the entire population (Centro de Microdatos 2011). The panel is unbalanced. The survey includes complete marital, fertility, and labor histories, as well as detailed information on the family in which the interviewee was raised. I add county-level identifiers to the public use EPS dataset in order to merge the family court administrative records on divorce.

[14] The EPS 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. For more information, see: https://www.previsionsocial.gob.cl/sps/biblioteca/encuesta-de-proteccion-social/ (accessed on September 30, 3019).

TABLE 2—DESCRIPTIVE CHARACTERISTICS BY WAVE, AGED 6 TO 18

	2002 (Wave 1)		2004 (Wave 2)		2006 (Wave 3)	
	Children with:		Children with:		Children with:	
	Married Parents	Cohabiting Parents	Married Parents	Cohabiting Parents	Married Parents	Cohabiting Parents
Full Sample						
School (%)	93.11	92.94	91.96	92.29	93.75	90.58
	(0.249)	(0.602)	(0.302)	(0.749)	(0.275)	(0.008)
Primary School Age (%)	96.48	96.30	97.38	97.78	97.43	95.48
	(0.269)	(0.623)	(0.278)	(0.609)	(0.289)	(0.794)
(N)	4,693	919	3,283	586	2,996	686
Secondary School Age (%)	90.29	89.47	88.26	87.59	91.43	85.73
	(0.395)	(1.028)	(0.465)	(1.261)	(0.407)	(1.328)
(N)	5,624	893	4,794	685	4,737	694
Average Age	11.98	11.54	12.41	11.96	12.57	11.70
	(0.036)	(0.084)	(0.041)	(0.105)	(0.042)	(0.103)
Female (%)	47.84	47.13	49.26	47.21	49.28	46.67
	(0.492)	(0.012)	(0.556)	(0.014)	(0.569)	(0.013)
Total	10,317	1,812	8,077	1,271	7,733	1,380
Urban Sample						
School (%)	93.19	92.58	91.90	91.87	94.06	90.80
	(0.301)	(0.774)	(0.357)	(0.932)	(0.300)	(0.900)
Primary School Age (%)	96.43	95.88	97.35	97.71	97.66	95.59
	(0.327)	(0.824)	(0.330)	(0.756)	(0.307)	(0.899)
(N)	3,219	583	2,374	393	2,432	522
Secondary School Age (%)	90.43	89.17	88.16	86.97	91.75	85.91
	(0.479)	(1.311)	(0.550)	(1.558)	(0.447)	(1.541)
(N)	3,781	563	3,454	468	3,784	511
Average Age	11.94	11.55	12.41	12.00	12.54	11.62
	(0.043)	(0.107)	(0.048)	(0.125)	(0.046)	(0.118)
Female (%)	48.36	46.42	49.16	46.92	49.13	46.66
	(0.597)	(0.015)	(0.655)	(0.017)	(0.634)	(0.016)
Total	7,000	1,146	5,828	861	6,216	1,033
Unique Individuals	TOTAL	URBAN				
Married Parents	14,362	9,977				
Cohabiting Parents	2,744	1,815				

Source: Author calculations. Encuesta de Protección Social (EPS), Chile.

A subset of school age children (ages 6 to 18) whose parents were married or cohabitating over the sample time period is constructed for the main analysis.[15] Since the estimates reported are calculated using a method that differences over time and across groups, any unobserved heterogeneity stable over time and between

[15] I follow a similar methodology to Martínez (2013) in selecting age of children as those under age 19, which draws from legal laws governing coverage of child support and alimony in Chile. I restrict to age six as a lower bound due to the general start of school enrollment.

children from married and cohabiting parents is differenced out and will not bias the observed estimates. The final sample includes 14,362 children from married parent families and 2,744 children from cohabiting parents (Table 2).

B. Administrative Records on Divorce

In partnership with the University of Chile's Center for Microdata, we acquired administrative records on divorce through a special request to the Chilean Supreme Court. Electronic records on divorce cases were only available for urban areas. Therefore, any analysis using length of time to finalizing a divorce was limited to children living in urban areas, around 70 percent in the sample (Table 2). The court administrative records included divorce cases from October 1, 2005[16] to December 30, 2006 and contain basic information about the type of resolution, start date and end date of each divorce case within each respective family court district. The data include 38,870 divorce cases. Cases beginning after November 1, 2006 were excluded from this analysis because data collection of the 2006 wave began in November 2006. Of the remaining cases (33,475), 95.9 percent had been finalized: 71.4 percent ended in a successful divorce, and 24.5 percent of cases were closed for other reasons, such as the couple decided to stay married. Around 4.1 percent of cases were still pending in 2009. The data on divorce cases were merged with the EPS panel data by county code for this analysis.

IV. Methodology

A. Household Bargaining

This paper assumes a standard household bargaining model, such as those developed by Manser and Brown (1980), McElroy and Horney (1981), McElroy

[16] This was the earliest date electronic records were available.

(1990), and McElroy (1997), in which individuals within the household bargain based on the power or resources they hold. A change in individual resource control shifts household investments towards the preferences of the individual controlling the resource. The rest of this section will focus on the application of this standard model, related extensions to the case of Chile, and describe the difference-in-difference estimation methods in context.

Opportunity Costs, Credible Threats, and Intertemporal Choices.—The legalization of divorce and its requirements of economic compensation for homemakers caused the opportunity cost of staying married to decrease for breadwinners because the law transferred potentially large sums of money to homemakers upon divorce. For the same reason, the opportunity cost of staying married for homemakers' increased. The introduction of *economic compensation* shifted the threat point within the household bargaining model, and intrahousehold bargaining power subsequently increased for homemakers in married-parent families. Given that prior research shows women invest more in household public goods (Quisumbing and Maluccio 1999, Rubalcava et al. 2004, Rangel 2006, Schady and Rosero 2007, Martínez 2013), I expect to see an increase in consumption of related items like children's education. The same shift would not occur in families not directly eligible for divorce.

If divorce shifts the opportunity cost of remaining married, it does so only in the sense that the threat of divorce or costs associated with divorce are truly credible. Shorter wait times make the threat of divorce more imminent and, thereby, more credible. If true, lengthy wait times should decrease the opportunity costs for homemakers and increase them for breadwinners resulting in less bargaining power for the homemaker in married couple households. Credibility, in this case, is measured by the time distance between when a homemaker threatens divorce and when the divorce can be actualized. The shorter the distance is, the more credible the threat. Take two extreme examples. In one, a homemaker threatens a divorce from his/her spouse and the divorce can be finalized in court the following day and economic compensation will be due immediately. The breadwinner will see the threat as very credible because a financial fine would be imposed the next day and will adjust accordingly by yielding to the demands of the homemaker's stated preferences.

If, however, the homemaker threatens divorce but finalizing the divorce will take more than one year, the breadwinner will see the immediate threat as less credible. The breadwinner could, for example, ask the homemaker to leave the home the next day, but the economic compensation would be due far into the future – once the divorce was finalized. In the meantime, the homemaker would need to find alternative living arrangements and, even if he/she is able to find resources for housing or alternative living arrangements with family or friends, a lot can happen in one year. Perhaps the breadwinner believes the short term separation and lack of resources will bring the homemaker back to the home before the divorce is finalized, giving in to the preferences of the breadwinner renegotiating resource allocation from a weakened bargaining position.

Spouses engaged in balancing time differing behaviors in this way are exercising behavior similar to hyperbolic time discounting. The time dimension of the credible threat problem identified here is an intertemporal choice decision where spouses are making trade-offs between the costs and benefits of the reality of a divorce occurring now versus later (Loewenstein et al. 2003).

Complete Information.—Complete information implies common knowledge. This is a critical component of the analysis because if spouses have no knowledge of wait times until they exercise their option for divorce, then variations in the length of time to divorce cannot influence behavior of intact households. Given the high propensity of Chilean families to live near each other among multiple generations (Araos-Bralic 2015), and the strong social and communal networks present in Chilean society, I argue the following.

Housewives have strong local social networks. When they get together for family or social events, generally husbands tag along. Wives tend to socialize in one group (e.g. around food preparation in the kitchen), while husbands tend to socialize in a separate group (e.g. around a grill or outside smoking). Gossip travels and, if a family member or close friend in the local *comuna* is getting a divorce, people will hear about it in social gatherings or while meeting one-on-one with friends and family. They will also hear details related to the divorce. Was it messy? Was it quick? Did it take a long time? Who got the house? What about child support?

All of this information travels through the social network, and, since Chileans are not commonly mobile, information about the process, ease, and length of time to divorce will travel within their *comuna*. Those in unstable or unhappy marriages are likely to hear this information or search it out and use it to update their priors on the opportunity costs associated with staying in their marriage. Those in unhappy marriages may also seek out advice from lawyers, family court staff, public servants, and others who would be knowledgeable about wait times and would share this information along with information about the requirements and process for divorce.

B. Difference-in-Difference Estimation

A difference-in-differences (DD) approach with panel data can generate unbiased estimates of the impact of a policy by comparing over time the group that experienced the policy change (treatment) to a similar group that did not (control). In this case, children from married parent families are the treatment group since they are directly exposed to and potentially affected by the legalization of divorce after 2004. Children from cohabiting parent families are the control group because these intact two-parent family households are not eligible for economic compensation upon separation.[17] In the next section, I test the validity of the control group under the parallel trends assumption.

Assuming the control group is valid for now, the basic individual-level equation in the DD analysis is the following.

$$(1)\ S_{igt} = \beta_0 + \beta_1 M_g + \beta_2 T_2 + \beta_3 T_3 + \beta_4 (M_g T_2) + \beta_5 (M_g T_3) + v_{gt} + \varepsilon_{igt} \ \ \forall\, i = 1 \dots I_{gt}$$

where S_{igt} is the binary dependent variable indicating whether child i from group g at time t is in school, M_g is a dichotomous variable that equals one for children from married parent families and zero for children from cohabitating parent families, T_t is a set of year dichotomous variables (time fixed effects), υ_{gt} is unobserved group effects at time t, ε_{igt} is the individual-specific error term, and E[υ_{gt}] = E[ε_{igt}] = 0.

The three time periods in the estimation are $T_1 = 2002$, $T_2 = 2004$, and $T_3 = 2006$. The reference year variable, T_1, is omitted from the equation above. To obtain consistent estimates, I assume E[ε_{igt} |M_g,T_t] = 0. While the treatment (exposure to divorce) became an option in November 2004, a time lag exists in the administrative process creating family courts and the information transfer of wait times. Because of this, I assume 2004 to be a pre-intervention year or a placebo year in that β_4 should not be significant if the theoretical predications about the impact of divorce are correct. The estimate of β_5 is the average treatment effect (ATE) of exposure to divorce on children's schooling.

In general, this basic DD equation is sufficient to produce unbiased estimates of the coefficient of interest, β_5. However, if decisions regarding education are made

[17] A similar method of selecting treatment and control groups to understand the impact of changes in family law has also been employed by Rangel (2006) and Martínez (2013). Although in both cases, Rangel and Martínez use children from cohabiting parent families as the treated because they are interested in the impact of alimony rights and child support expansions to cohabiting couples with children.

differently for subgroups within the sample and the decision-making process is correlated with explanatory variables not included in the regression equation, omitted variable bias can occur. In the case of school enrollment, the parental decision-making process may be different based on the gender and age of the child. For this reason, two approaches are considered. First, I add controls for gender and age of the child to Equation (1). Adding these variables is expected to improve the estimation since, for example, parental decisions to enroll their children in primary school are different from decisions to enroll them in secondary or tertiary school. Parents might also have different preferences in terms of schooling for daughters compared to sons (Gertler and Glewwe 1992, Rubalcava and Contreras 2000). Rubalcava and Contreras find evidence in Chile of gender preferences in children's education. Adding these controls gives the following equation.

$$(2)\ S_{igt} = \beta_0 + \beta_1 M_g + \beta_2 T_2 + \beta_3 T_3 + \beta_4 (M_g T_2) + \beta_5 (M_g T_3) + \boldsymbol{\gamma_1 Z_{igt}} + v_{gt} + \varepsilon_{igt} \ \ \forall\ i = 1 \ldots I_{gt}$$

where $\boldsymbol{Z_{igt}}$ are the gender and age specific control variables. All other variables are the same as in Equation (1).

C. Modeling Wait Times

For wait times to be a valid exogenous source of variation, individuals should not be able to manipulate wait times with their own behavior. Aside from the fact that married individuals cannot manipulate or change the *comuna* they divorce in[18], I also motivate this argument by examining other potential factors that influence wait times, dividing them into three major categories: overall family court case volume,

[18] Recall the law requires individuals to divorce in the family court district corresponding to the *comuna* in which they reside.

propensity for divorce cases, and environmental characteristics. The model is shown in Equation (3).

$$(3) \qquad WaitTime_j = \boldsymbol{X_j^1} + \boldsymbol{X_j^2} + \boldsymbol{Z_j} + \varepsilon_j$$

where $\boldsymbol{X_j^1}$ includes overall case volume and related characteristics of family court district *j* like the rate of all cases in family court and the percent of children under age 6, which serves as an overall indicator of families (and potential workload) within the district. $\boldsymbol{X_j^2}$ includes propensity to divorce indicators for each family court *j*, which are the percent married age 18 plus in the district and the rate of all divorce cases. $\boldsymbol{Z_j}$ are general community indicators for each *j* such as the percent of individuals aged 25 plus with a university degree and the percent of individuals aged 18 plus in labor force.

If divorce indicators influence wait times, then factors like the divorce case load and the percent married age 18 plus will drive wait times. If true, wait times may be endogenous to divorce. If overall court case volume drives wait times, then variables associated with other types of family court cases like the percent of kids under age 6 should be significant, providing some evidence of the wait time exogeneity.

D. Difference-in-Difference-in-Difference (DDD) Estimation

To capture the effect of bureaucratic processes on intrahousehold bargaining, I use the average wait time to divorce by *comuna*. Adding wait time with the appropriate interaction terms to Equation (2) generates a difference-in-difference-in-difference (DDD) estimation as shown in Equation (4).

$$(4) \quad S_{igt} = \beta_0 + \beta_1 M_g + \beta_2 T_2 + \beta_3 T_3 + \beta_4 (M_g T_2) + \beta_5 (M_g T_3) + \delta_0 W_3^c + \delta_1 (W_3^c M_g) + \delta_2 (W_3^c T_2) + \delta_3 (W_3^c T_3) + \delta_4 (W_3^c M_g T_2) + \delta_5 (W_3^c M_g T_3) + \boldsymbol{\gamma_1 Z_{igt}} + v_{gt} + \varepsilon_{igt} \;\; \forall\, i = 1 \ldots I_{gt}$$

where W_3^c is the wait time for divorce by court district in the last time period. All other variables are labeled as in Equations (1) and (2). Because wait times are only observed in 2006 (or T_3), multiple components drop out of Equation (4). Specifically, $\delta_0, \delta_1, \delta_2, \delta_3, and\ \delta_4$ for the following reasons.

$\delta_0 = 0$; $\delta_3 = 0$	This holds because wait times are not generalizable over the entire population, but specific only to a certain subgroup (married parents).
$\delta_2 = 0$; $\delta_4 = 0$	This holds because wait time equals zero for all *comunas* in 2004.
$\delta_1 = \delta_5$	This is true by construct of the data since wait times (W_3^c) are only available in year 2006 (T_3). To avoid collinearity, I drop δ_1 from the specification.

Given the above, the final reduced form DDD equation I estimate is:

$$(5) \; S_{igt} = \beta_0 + \beta_1 M_g + \beta_2 T_2 + \beta_3 T_3 + \beta_4 (M_g T_2) + \beta_5 (M_g T_3) + \delta_5 (W_3^c M_g T_3) + \boldsymbol{\gamma_1 Z_{igt}} + v_{gt} + \varepsilon_{igt} \;\; \forall\, i = 1 \ldots I_{gt}$$

The coefficients of interest in Equation (5) are β_5 (the pure effect of offering divorce where no option previously existed) and δ_5 (the added effect of waiting an additional 6 months for the divorce to finalize). In the case of the DDD variable, δ_5:

$$(6) \qquad \hat{\delta}_5 = \left(\bar{S}_{W,M,3} - \bar{S}_{W,M,1}\right) - \left(\bar{S}_{0,M,3} - \bar{S}_{0,M,1}\right) - \left(\bar{S}_{W,C,3} - \bar{S}_{W,C,1}\right) - \left(\bar{S}_{0,C,3} - \bar{S}_{0,C,1}\right)$$

where *W* is average wait time in 2006, 0 is the average wait time in 2002, *M* is married parent child, *C* is cohabiting parent child, 3 is the last time period (2006), and 1 is the first time period (2002). In this estimation, $\left(\bar{S}_{W,M,3} - \bar{S}_{W,M,1}\right)$ is the time change in schooling for married parent children by wait time, $\left(\bar{S}_{0,M,3} - \bar{S}_{0,M,1}\right)$ is the time change in schooling for married parent children when wait time is zero, $\left(\bar{S}_{W,C,3} - \bar{S}_{W,C,1}\right)$ is the time change in schooling for cohabiting parent children by wait time, and $\left(\bar{S}_{0,C,3} - \bar{S}_{0,C,1}\right)$ is the time change in schooling for married cohabiting parent children when wait time is zero. The treatment effect, $\hat{\delta}_5$, is the effect of administrative processes on married children's schooling relative to cohabiting children's schooling for each six month wait time.

I interpret β_4 in Equation (5) as a placebo (or test) year. If the theory of the effect of divorce law and wait times hold, then I expect $\beta_4 = 0$ because the divorce law was implemented at the end of 2004 around the same time the survey was in field. If β_4 is significant, it could be an anticipation effect related to the law change.

E. Regression Models

Since the binary dependent variable, children's school enrollment, is a variable indicating one if the child is in school and zero otherwise, all estimation equations are estimated using a linear probability model (LPM), logit, and probit. For ease of interpretation, the LPM model results are discussed. Logit and probit models are shown to produce similar results in both direction and significance. I conducted a Bruesch-Pagan Lagrange multiplier (LM) test for random effects (results not shown) and found that random effects are appropriate.[19] Using random effects in my analysis allows for individual effects, and I use Huber-White robust standard errors for heteroscedasticity in all regressions.

V. Results

A. Validating the Parallel Trend Assumption

A key non-trivial identifying assumption with DD estimation is that the *trend* in outcomes of interest must be similar for both treatment and control groups pre-intervention (Angrist and Pischke 2009). To test this assumption, I report schooling trends for two time periods before the intervention. If the parallel assumption holds prior to the treatment, then the two groups can be compared using difference-in-differences estimation.

Figure 2 shows school attendance for children from married and cohabiting parent families. The rate of enrollment is similar for both groups before treatment; the differences observed are not significant. However, after the legalization of divorce, cohabiting parent children continue to decrease school enrollment while children from married parent families do not, providing evidence that supports the use of a DD method for this analysis. Further comparisons of married parent families to single and divorced parent families are in the Appendix.

[19] Further down in Table 3 (column (0)), I include results using fixed effects for comparative purposes only.

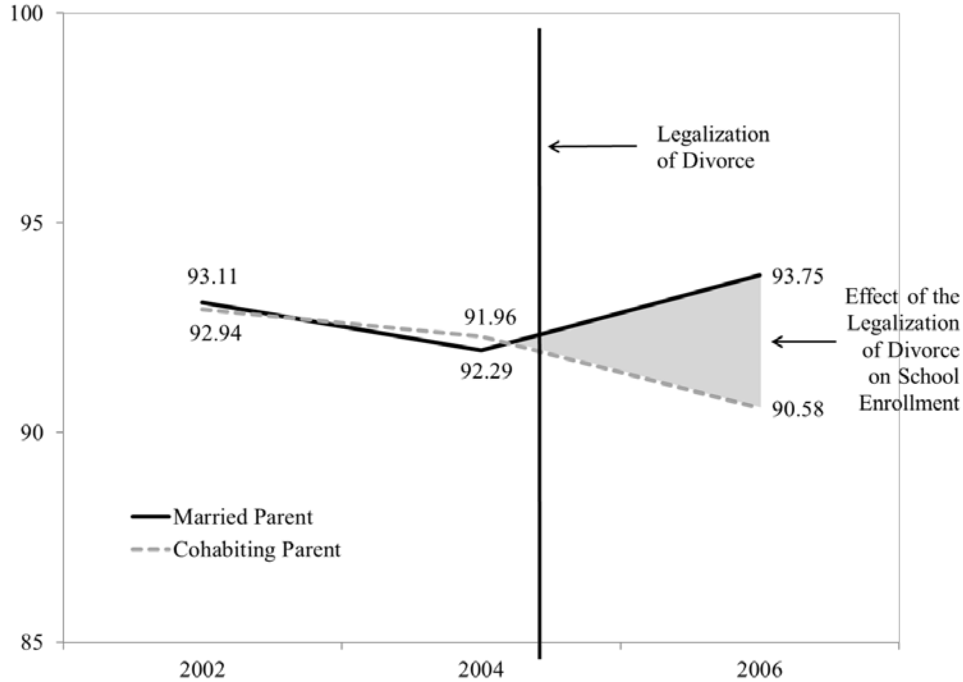


FIGURE 2. PERCENT OF CHILDREN (AGE 6 TO 18) ATTENDING SCHOOL BY PARENTAL MARITAL STATUS

Source: Authors calculations. Encuesta de Protección Social (EPS), Chile.

Notes: The differences between children from married parent families and children from cohabiting parent families in 2002 and 2004 is not statistically significant. The difference between these children in 2006 is statistically significant at the 90% level.

B. The Effect of Access to Divorce Law on Schooling

Table 3 shows the results from difference-in-difference (DD) estimations of linear probability model, logit, and probit equations.[20] Column (1) represents Equation (1) above and includes marital status of the parent, panel year, and interaction terms interacting parental marital status with panel year. The interaction term (*married parent*2006*) estimates $\hat{\beta}_5$, the effect of legalizing divorce on school enrollment. As expected, the coefficient is positive and significant; legalizing divorce increased school enrollment by 2.4 percentage points (p = 0.021). The

[20] As previously mentioned, column (0) reports results using fixed effects for comparative purposes.

dummy variable for 2006, $\hat{\beta}_3$, is negative and statistically significant at p = 0.000 implying that, all else equal, enrollment rates decreased, on average, for everyone in 2006 compared to 2002. This regression, however, does not control for age or gender. Adding these variables improves the estimation (Column (2) in Table 3). Once age[21] and gender are controlled for, the impact of exposure to divorce, $\hat{\beta}_5$, is larger, 3.3 percentage points, and significant at p = 0.001. All else equal, school enrollment for girls increased by almost one percentage point (0.8) compared to boys.

TABLE 3—EFFECT OF ACCESS TO DIVORCE ON CHILDREN'S SCHOOLING

	LPM			Logit		Probit	
	(0)	(1)	(2)	(3)	(4)	(5)	(6)
Marital Status of Parent (Reference=Cohabiting Parent)							
Married Parent	-0.029 **	-0.003	0.004	0.004	0.184	0.004 *	0.091
	(0.012)	(0.066)	(0.006)	(0.121)	(0.132)	(0.063)	(0.067)
Year (Reference=2002)							
2004	-0.034 ***	-0.014	0.005	-0.171	0.159	-0.086	0.093
	(0.011)	(0.009)	(0.009)	(0.165)	(0.177)	(0.087)	(0.092)
2006	-0.086 ***	-0.041 ***	-0.016 *	-0.515 **	-0.206	-0.268 ***	-0.098
	(0.011)	(0.009)	(0.009)	(0.159)	(0.167)	(0.084)	(0.086)
Interaction Terms (Reference=Married Parents in 2002)							
Married parent * 2004	-0.006	-0.007	-0.006	-0.101	-0.155	-0.055	-0.078
	(0.012)	(0.009)	(0.009)	(0.176)	(0.191)	(0.093)	(0.099)
Married parent * 2006	0.025 **	0.024 **	0.033 ***	0.440 *	0.591 **	0.228 **	0.305 ***
	(0.012)	(0.009)	(0.010)	(0.172)	(0.185)	(0.228)	(0.096)
Demographics							
Girl	No	No	0.008 ***	No	0.149 *	No	0.077 **
			(0.003)		(0.062)		(0.032)
Age	No	No	Yes	No	Yes	No	Yes
Other Model Considerations							
Fixed effects	Yes	No	No	No	No	No	No
Random effects	No	Yes	Yes	Yes	Yes	Yes	Yes
Constant	0.9839 ***	0.936 ***	0.943 ***	3.530 ***	3.685 ***	1.944 ***	1.999 ***
	(0.011)	(0.006)	(0.007)	(0.148)	(0.197)	(0.078)	(0.101)
N observations		30,590	30,590	30,590	30,590	30,590	30,590
N individuals		17,280	17,280	17,280	17,280	17,280	17,280

Source: Authors calculations. Encuesta de Protección Social (EPS), Chile.

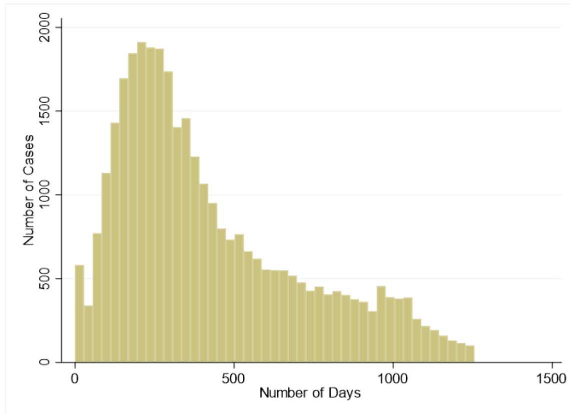
Notes: Linear Probability Model (LPM) using a national sample. Random effects and Huber-White robust standard errors; * p<0.05, ** p<0.01, *** p<0.001.

[21] Dummy variables are used for age since the sample is not representative of the total population with its related underlying functional form.

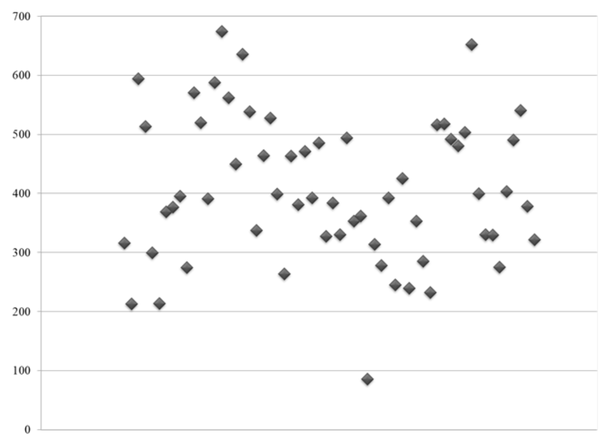
C. The Effect of Wait Times on Schooling

Variation in Wait Times.—Aside from the geographic constraints imposed on divorcing couples, is there enough variation in wait times to warrant its use? Panel A of Figure 3 highlights variation in the length of divorce wait times for cases ending between October 1, 2005 and October 31, 2006. Wait times range anywhere from zero days (same day divorce, N=2) to 1,225 days (N=2). Panel B of Figure 3 shows variation in the average wait time to divorce by family court district. The fastest *comuna* average wait time to finalize a divorce is 85.9 days or around 3 months. The longest average wait time is 674.4 days, almost two years. I will use this variability in wait time, which the individual household member cannot control but can acquire information on by talking with friends, neighbors, family, lawyers, or court administrators, as an exogenous shock defining the credibility or opportunity cost of the homemaker's threat of divorce.

In 2006, divorce cases made up between 3.6 to 13.7 percent of all family court cases (Table 4, Panel A). Panel B of Table 4 shows OLS regression results of wait time as described in Equation (3). The results show that the overall family court case rates and the percent of children under age 6 are weakly significant and in the expected direction. Factors associated with divorce filings like the rate of divorce cases and the percent of married individuals aged 18 plus (who are the only individuals at risk of divorce) are not significant and the coefficients are in the opposite direction. These results provide some evidence that wait times are driven by a broad set of factors outside of the control of the married couple household wanting to divorce. That, in combination with the legal constructs driving the geographic process of divorce, are used to assume sufficient independence of an average wait time variable by family court district.



PANEL A. NUMBER OF DAYS TO DIVORCE BY DIVORCE CASES (N=33.475)



PANEL B. MEAN NUMBER OF DAYS TO DIVORCE BY FAMILY COURT TRIBUNAL (N=60)

FIGURE 3. VARIATION IN THE TIMING OF DIVORCE

Source: Authors calculations. Chilean Supreme Court, 2005-2006.

TABLE 4—FAMILY COURT CHARACTERISTICS AND WAIT TIME REGRESSIONS, 2006

Panel A. Statistics on Cases by Family Court	N	(%)
All Family Court Cases		
Min (N)	807	
Max (N)	21,235	
Divorce Cases		
Min (N/(%) of all cases)	65	3.62
Max (N/(%) of all cases)	2,667	13.74
Percent of Divorce Cases in All Family Court Cases (Mean)		8.08
Percent of Divorce Cases in All Family Court Cases (Mode)		8.05
Panel B. Family Court Wait Time Regressions	**Coef.**	**Coef.**
Overall Volume of Family Court Cases and Related Characteristics		
Family Court Case Rate (per Population Aged 18 Plus)	0.044 *	0.045 *
	(0.023)	(0.023)
Percent of Population Under Age 6	13.729 *	14.346 *
	(7.765)	(7.802)
Propensity to Divorce and Related Characteristics		
Divorce Case Rate (per Population Aged 18 Plus)	-0.252	-0.288
	(0.220)	(0.215)
Percent of Population Aged 18 and Older Who Are Married	-2.967	-3.553
	(2.350)	(2.424)
Community Characteristics		
Percent of Population Aged 25 and Older with a University Degree		-3.290
		-3.031
Percent of Population Aged 18 and Older in Labor Force	1.798	
	(1.799)	
Constant	1.168	2.546 *
	(1.504)	(1.317)
N	56	56

Source: Authors calculations. Encuesta de Protección Social (EPS), Chile.

Notes: Random effects and Huber-White robust standard errors; * p<0.05, ** p<0.01, *** p<0.001.

The Effect of Administrative Processes of Divorce.—The sample used to analyze the influence of the speed of divorce on schooling is children from married-parent and cohabiting-parent families living in urban areas.[22] To provide an accurate comparison of the results with and without wait time, prior analysis from Table 3

[22] Administrative records on divorce cases were not available electronically for rural areas.

are run using the subset of urban children. Table 5 presents the comparison results for urban children in Columns (1) and (2). Notice the results for the urban sample are similar to the national sample showing a positive (3.4 percentage points) and significant effect of access to divorce on children's education in urban areas.

TABLE 5—EFFECT OF ADMINISTRATIVE PROCESSES OF DIVORCE ON CHILDREN'S SCHOOLING

	LPM			Logit	Probit
	(1)	(2)	(3)	(4)	(5)
Marital Status of Parent (Reference=Cohabiting Parent)					
Married Parent	0.001	0.005	0.005	0.206	0.105
	(0.008)	(0.008)	(0.008)	(0.165)	(0.084)
Year (Reference=2002)					
2004	-0.013	0.003	0.003	0.106	0.071
	(0.011)	(0.011)	(0.011)	(0.215)	(0.111)
2006	-0.037 ***	-0.015	-0.015	-0.179	-0.081
	(0.012)	(0.011)	(0.011)	(0.204)	(0.105)
Interaction Terms (Reference=Married Parents in 2002)					
Married parent*2004	-0.009	-0.004	-0.004	-0.100	-0.054
	(0.012)	(0.011)	(0.011)	(0.231)	(0.119)
Married parent*2006	0.022 *	0.034 ***	0.055 ***	1.170 **	0.594 ***
	(0.013)	(0.012)	(0.017)	(0.368)	(0.189)
Triple Interaction Term					
Married parent*2006*Wait Time (Six Months)			-0.009 *	-0.230 **	-0.115 *
			(0.005)	(0.123)	(0.063)
Demographics					
Girl		0.008 **	0.008 **	0.142 *	0.074 *
		(0.004)	(0.004)	(0.074)	(0.038)
Age	No	Yes	Yes	Yes	Yes
Constant	0.933 ***	0.943 ***	0.943 ***	3.668 ***	1.982 ***
	(0.008)	(0.009)	(0.009)	(0.236)	(0.120)
N observations	22,084	22,084	22,084	22,084	22,084
N individuals	11,922	11,922	11,922	11,922	11,922

Source: Authors calculations. Encuesta de Protección Social (EPS), Chile.

Notes: Linear Probability Model (LPM) using an urban sample. Random effects and Huber-White robust standard errors; * p<0.05, ** p<0.01, *** p<0.001.

The DDD analysis of Equation (4) in its reduced form (Equation (5)) is presented in Table 5 Column (3). Administrative wait time, $\hat{\delta}_5$, is negative and significant (p = 0.070). Every six-month increase in wait time to finalize a divorce results in an approximate one percentage point (0.9) decrease in school enrollment. The coefficient measuring the overall effect of access to divorce, $\hat{\beta}_5$, increases to 5.5

percentage points and is significant (p = 0.001). Gender remains a significant factor in determining whether children attend school.

Logit and probit models for the urban sample provide similar results; however, the coefficients themselves do not directly explain the estimated effect of each independent variable. For that reason, the marginal effects of the national sample and urban sample logit regressions are presented in Table 6, comparing children of married parents in each year to the base comparative group (children of cohabiting parents). The results show that the only significant difference occurs between the two groups in 2006, post the lagged implementation of administrative divorce processes. The coefficients are similar in magnitude to the LPM results and significant.

TABLE 6—LOGIT MARGINAL EFFECTS ON CHILDREN'S SCHOOLING

	National Sample		Urban Sample	
	(dy/dx)		(dy/dx)	
Marital Status of Parent				
Cohabiting	(base outcome)		(base outcome)	
Married				
2002	0.009		0.011	
	(0.007)		(0.009)	
2004	0.001		0.006	
	(0.007)		(0.009)	
2006	0.038	***	0.063	***
	(0.007)		(0.015)	
N observations	30,590		22,084	
N individuals	17,280		11,922	

Source: Authors calculations. Encuesta de Protección Social (EPS), Chile.

Notes: Regressions include age, gender, and wait time (urban sample only) variables. dy/dx for factor levels is the discrete change from the base level; random effects and Huber-White robust standard errors; * p<0.05, ** p<0.01, *** p<0.001.

D. Who Benefits the Most?

School enrollment varies. Table 2 shows that enrollment rates of primary school children from cohabitating parent families fluctuate over the panel. The same is not true of their secondary school counterparts, whose school enrollment rate continually decreased between 2002 and 2006. Married parent children in secondary school experienced a decreased between 2002 and 2004 followed by an increase in school enrollment rates from 2004 to 2006, while their primary school counterparts' school enrollment continually increased over time. Little variation is observed in primary school enrollment for either group. Increasing variation over time in the percent of children in school is observed for secondary school aged children; of note, rates for married parent children went up in 2006 while rates for cohabitating parent children decreased. Figure 2 masks deviations by school age, but Table 2 gives clear indication that including dummy variables for age and running separate regressions by school type are appropriate steps in the estimation process.

Although access to divorce for parents had a positive effect on children's education, interpreted as increasing women's bargaining power within married couple families, age was also significant. All else equal, older children were less likely to attend school compared to six year olds (results not shown). A question still remains as to which school age children benefitted the most from the implementation and speed of divorce. For that reason, and because it is possible that enrollment decisions are made differently depending on the level of school (primary versus secondary), the LPM regressions are replicated separately for elementary school-aged and secondary school-aged children for both the national and urban samples (Table 7).

TABLE 7—EFFECT OF DIVORCE ON CHILDREN'S SCHOOLING BY SCHOOL LEVEL

	Primary School		Secondary School	
	National Sample	Urban Sample	National Sample	Urban Sample
Marital Status of Parent (Reference=Cohabiting Parent)				
Married Parent	0.003	0.007	0.001	0.004
	(0.007)	(0.009)	(0.011)	(0.014)
Year (Reference=2002)				
2004	0.013	0.016	-0.031 *	-0.033 *
	(0.009)	(0.011)	(0.016)	(0.020)
2006	-0.004	-0.001	-0.068 ***	-0.065 ***
	(0.010)	(0.012)	(0.017)	(0.021)
Interaction Terms (Reference=Married Parents in 2002)				
Married parent * 2004	-0.007	-0.010	-0.002	-0.002
	(0.009)	(0.012)	(0.017)	(0.021)
Married parent * 2006	0.014	0.008	0.051 ***	0.090 ***
	(0.011)	(0.018)	(0.018)	(0.029)
Triple Interaction Term				
Wait Time (Six Months)		0.002		-0.017 **
		(0.006)		(0.008)
Gender (Reference=Boy)				
Girl	0.006 *	0.005	0.014 ***	0.014 **
	(0.003)	(0.004)	(0.005)	(0.006)
Constant	0.958 ***	0.955 ***	0.905 ***	0.904 ***
	(0.007)	(0.008)	(0.905)	(0.014)
N observations	13,163	9,523	17,427	12,561
N individuals	9,133	6,425	11,205	7,863

Source: Authors calculations. Encuesta de Protección Social (EPS), Chile.

Notes: Random effects and Huber-White robust standard errors; * p<0.05, ** p<0.01, *** p<0.001.

These DDD results show that legalizing divorce had a significant impact on school enrollment for children in secondary school and no impact on children in primary school. This makes sense given that schooling is compulsory in Chile and attendance rates are particularly high for primary school-aged children. Additionally, in a middle income country like Chile, there is little reason to keep primary school aged children out of school. However, there may still be families struggling to make ends meet who are faced with financially constraining decisions of encouraging elder children to work for pay in lieu of continued schooling. This theory is somewhat supported by the slightly higher impact of the law in urban areas, as the magnitude of the effect is larger for the urban sample compared to the national sample as a whole (Table 7). Financial constraints and decisions between school and working in, for example, the service sector as a teen, might be more salient in higher population density urban areas where paid or informal work to support the family business is more easily available to teenagers in middle income countries.

VI. Conclusion

Attempts to analyze the effects of divorce on child and family wellbeing are challenging due to selection bias and endogeneity issues. I take advantage of a unique natural experiment to advance our understanding of the role family policies play in economic development. In Chile, a society comprised of strong gendered family norms (Stillerman 2004), access to divorce with *economic compensation* shifted household bargaining power into the hands of married women. Legal constraints regarding where individuals could divorce, along with the expansion of new independent family court districts, provided a ripe context for a natural experiment where I use exogenous wait times for divorce to study not only the effect of access to divorce with strong property right redistribution, but also relevance of administrative processes in hindering or facilitating economic development.

In a middle income, largely Catholic country, introducing a path to legal divorce induced changes in household bargaining that lead to improvements in child welfare, advancing economic development, by inducing higher investments in children's education. In particular, children in secondary school increased school enrollment between 5.1 and 9.0 percentage points. Second, when the threat of divorce was less imminent (defined by lagged processing times), investments in children's education decreased by 1.7 percentage points every six months for secondary school-aged children. These results are larger in magnitude than those found by Rangel (2006) and Martínez (2013). However, given that the policy change was (1) more drastic (from no divorce to a progressive divorce law requiring the back pay of wages to homemakers), (2) influential for a larger group of individuals (married parents) instead of an extension of an already existing policy focused on cohabiters, these results are not surprising. All results together provide evidence that households are sensitive to the magnitude and breadth of policy shocks. Future studies should examine differences in short run versus long run effects given Gruber's (2004) finding of the long run effects on child outcomes and Wolfer's (2006) work showing the long run decreasing impact of unilateral divorce on divorce rates.

More generally, I analyze the effect of access to divorce on *household behavior*. I test whether a law mandating resource reallocation to homemakers gives more bargaining power to wives in married couple households. Using the bargaining household model framework, I provide evidence that pure access to divorce options that increase wives' opportunity cost of remaining married transfers power to homemakers. Additionally, I show that burdensome administrative processes influence household bargaining power and resource allocation by altering opportunity costs and the "*credibility*" of the threat. Said another way, *bureaucracy* can alter transaction costs associated with intrahousehold bargaining separately from the policy design and intention itself. If we do not recognize this and pay close attention to how administrative processes influence the role out of laws and policies, we could undermine good public policies that have the potential to advance development.

Family policies not traditionally seen as development tools have the power to advance economic development and are an undervalued resource. Legal changes to marriage and divorce laws shifting power towards family members more likely to invest in household public goods can have a positive effect on outcomes like educational attainment. This, in turn, can advance economic development. Policies that reduce burden and facilitate quick and easy reallocation of resources also have real life implications on family wellbeing, household investments, and economic development. Development economists, planners, and policy makers alike should consider these types of public policy mechanisms as a potential option to accelerate children’s education and foster economic development.

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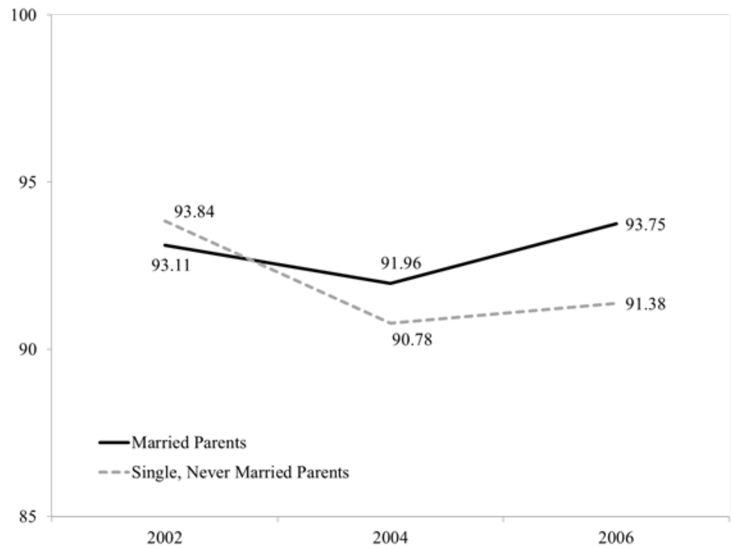
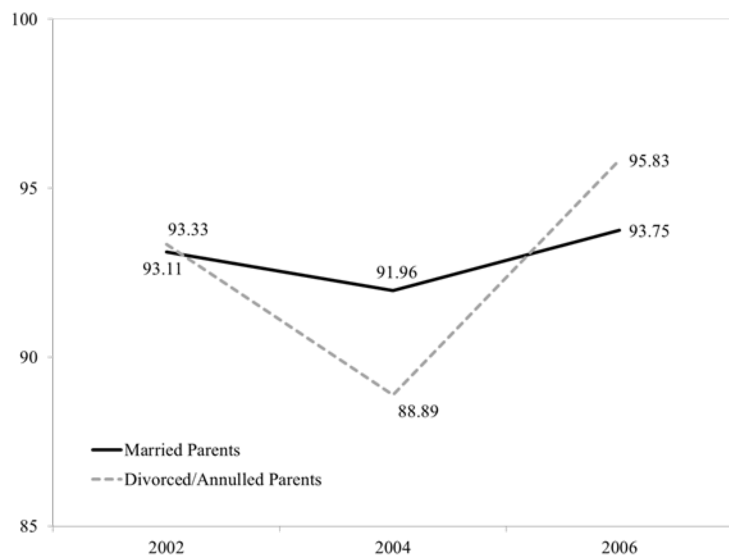
APPENDIX – ALTERNATIVE CONTROL GROUPS

There are two other potential comparison groups as counterfactuals, or controls, for this analysis: single, never-married and annulled/divorced parent families. Neither of these groups would be eligible for divorce and its economic compensation package. These two groups are less similar to married-parent families than cohabiting families for many reasons including the number of decision-making adults in the household, the propensity to have an adult staying home as a homemaker, and the number of potential income sources entering the household. Appendix Figure A1a and A1b demonstrate violations of the parallel trends assumption for both these alternative potential control groups.

There are additional reasons to believe that these groups are not good comparison candidates for married-parent families. First, household bargaining is assumed nonexistent in these households. There is only one adult in the household responsible for decision making, domestic chores, and guiding child development. In addition, one parent logistically has less time to devote to helping with homework or attending school events so the environments are different. Second, these parents have a higher propensity to be living in poverty or economically stressed situations. And, finally, deviating from the social norm to live as a single parent head of household in a middle income mostly Catholic country directly implies unobserved characteristics of grit, strength, and resolve that most likely deviate from the general case of the average married parent. For all these reasons and because the parallel trend assumption holds, in addition to the fact that Rangel (2006) and Martínez (2013) have used similar comparisons, the primary comparison I use in this analysis is married parent compared to cohabiting parent families.



PANEL A. MARRIED PARENTS COMPARED TO SINGLE, NEVER-MARRIED PARENTS



PANEL B. MARRIED PARENTS COMPARED TO DIVORCED/ANNULLED PARENTS

FIGURE A1. SCHOOL ENROLLMENT OF CHILDREN (AGE 6 TO 18) BY MARITAL STATUS

Source: Authors calculations. Encuesta de Protección Social (EPS), Chile.