Rev Econ Household (2017) 15:857–877
DOI 10.1007/s11150-015-9307-8

The effects of the Chilean divorce law on women's first birth decisions

Jose V. Gallegos[1] **· Jan I. Ondrich**[2]

Received: 2 July 2014 / Accepted: 17 August 2015 / Published online: 29 August 2015

Abstract In 2004 a new law in Chile allowed couples to divorce for the first time. The law also established compensation in case of divorce where one spouse sacrifices professional development or advancement for the good of the household. Using birth histories constructed from the Chilean Social Security Survey (Encuesta de Prevision Social–EPS) Panel 2002–2009, we investigate the effect of the divorce law on a woman's decision of when to have a first child. We find that the divorce law increases the hazard rate of having the first child by 61 % for more educated women, controlling for socioeconomic characteristics, length of marriage and the negative trend in fertility rates observed in Chile since the mid-1960s. We also find that a given percentage increase in a woman's potential income will increase the hazard rate by a greater percentage increase after the passage of the law.

Keywords Divorce law · Time to first birth · Chile

JEL Classification J12 · J13

1 Introduction

In this paper we evaluate the effect of the introduction of divorce in Chile in 2004 on a woman's decision regarding when to have a first child. The evolution of Chilean society and the historical trends of the regulation of marriage in Chile

✉ Jose V. Gallegos
jvgalleg@maxwell.syr.edu

Jan I. Ondrich
jondrich@maxwell.syr.edu

[1] Universidad de Piura, Lima, Peru

[2] Department of Economics, Syracuse University, Syracuse, NY, USA

provide a rich context to study a topic where the economic literature has yet to form a consensus. In 1884 the regulation and registration of marriages passed from the Catholic Church to the State. However, the influence of the Church was not eliminated, especially with regard to the termination of marriages. Couples were not allowed to divorce until 2004, when new legislation introduced the concept of "divorce a vincula matrimonii" or complete divorce, under which either spouse can request a complete termination of the marriage. The new legislation also introduced monetary compensation for a spouse who gives up professional development or advancement for the good of the household.

The Chilean divorce law (henceforth, DL) standardizes the process through which a marriage is terminated and thereby removes the uncertainty and potential negotiation costs the couple would incur if they decide to obtain the divorce. Becker (1974) suggests that the outcome of a divorce is related to specific investments made during the marriage, and lists children as one such type of investment. A married couple receives utility from children, but if the time cost of raising a child falls mainly on the woman, her career prospects are reduced. If this cost outweighs the benefit to the woman herself, having children reduces her net utility, and she would opt not to do so. If children are potentially Pareto-improving, the man might offer to transfer resources to the woman in order to compensate her for the costs she bears, so that having the child would become individually rational for both parties. While such a transfer may be fairly easy to implement for couples who anticipate being able to cooperate indefinitely, others may be unable to make such an arrangement unless they could count on an external enforcement mechanism. The system in place in Chile before the divorce law was introduced could not provide that enforcement, and many separated men were able to avoid making spousal support payments. Thus, the new divorce law in effect raises fertility because it rectifies a market failure by making some desirable implicit contracts feasible.

The increase in fertility may actually have little to do with the ability to divorce itself, but instead is driven by the fact that the divorce law now enforces implicit contracts concerning alimony. It may also account for an increase in non-marital fertility, assuming that the courts now make a greater effort to enforce child support payments more generally, which suggests that the divorce reform is more closely associated with decreases in the mean age of first-time mothers than with decreases in the mean age of first-time brides. Moreover, the effect on the fertility may be greater for women with higher potential incomes, because they have more to lose if their career is derailed.[1]

The goal of this paper is to determine the degree to which the DL affects a woman's decision regarding when to have her first child. Using the Social Security Survey (Encuesta de Previsión Social–EPS) Panel 2002–2009, we find that the DL has a positive effect on the hazard rate of having a first child for more educated women, controlling for the woman's age and length of marriage, the negative trend observed on fertility rates in Chile since the mid-1960s, and other socioeconomic

[1] The amount of compensation is determined by a court, which takes into account the beneficiary's age, level of education, length of marriage, and number of children, as well as other socio-demographic characteristics.

characteristics. We also find that a higher potential income level is associated with an increase in the hazard rate for first birth.

This paper is structured as follows. Section 2 presents the context in which the Chilean DL was introduced, and our motivation to focus on first birth. The third section consists of a review of the literature which focuses on divorce laws and fertility decisions. In the fourth section we present the data, describe the empirical strategy and discuss the results of the estimation. Finally, in the fifth section we present our conclusions.

2 Background and motivation

Historically religion has played an important role in marital legislation in Latin America, where, in most countries, divorce was prohibited throughout the twentieth century (Rossetti 1993). Attempts to implement divorce in Chile can be traced back more than a century.[2] The most recent was in November 1995 and was finally approved by the Chilean Congress in March 2004 and implemented in November of the same year.

Before 2004 Chilean marital law was based on the 1884 civil code, which permitted dissolution of marriage through annulment only. To obtain an annulment, spouses must argue before a judge that some aspect of the marriage was improper, for example, that one of the witnesses was under 18 years of age, or that the name of the spouse was misspelled on the marriage certificate. The 1884 civil code also allowed for the separation of the couple but this separation did not dissolve the marriage and marital obligations continued.[3] The DL suspended the succession rights, introduced a monetary compensation given to the spouse who sacrifices personal development for the good of the household, and most importantly, introduced a standard procedure to terminate a marriage, thereby removing the uncertainty and potential negotiation costs that seeking the annulment implied.[4] Moreover, the DL allows unilateral and no-fault divorce, requiring in both cases a 1-year separation period. However, if one of the spouses files for divorce on the grounds of a fault, no separation period is required.

The DL has contributed to the observed trends of first births and first marriages in Chile in the last decade. The left panel in Fig. 1 shows a clear change in the trend of first births after the implementation of the DL in late 2004. Between 1997 and 2004 the proportion of first births decreased from 1.45 to 1.16 %, but increased from 1.21 % in 2005 to 1.31 % in 2008, before decreasing to 1.22 % in 2012. On the other hand, the right panel in Fig. 1 shows a persistent decline in the proportion of

[2] Chilean Library of Congress. http://www.bcn.cl/carpeta_temas/temas_portada.2005-10-27.7388460505.

[3] Separation provided a woman with a food pension, but there is evidence that these pensions are frequently not paid: administrative data shows that 80 % of petitions to the courts concerning children's issues are related to nonpayment of a food pension. See Senate of Chile, Records of Session 12 (15 July 2003).

[4] We would like to thank an anonymous referee for this argument.

first marriages before 2005.[5] Between 1997 and 2005 marriage records show that the proportion of single women who got married with respect to the total female population decreased from one to 0.63 %. In 2006 this figure increased to 0.67 %, and decreased to 0.61 % in 2012.

The decreasing trend in first marriages is in line with the increasing preference towards cohabitating relationships: according to Census information, the share of the population (15 or over) in a cohabitating relationship increased from 5.7 % in 1992 to 16.2 % in 2012. The decline in first marriages significantly decreases the contribution to first births by married couples between 1999 and 2006 (from 36.5 to about 22 %), according to administrative records (Larrañaga 2006). Palma (2006) suggests that 50 % of out-of-wedlock births are to women in a cohabitating relationship; separated women and partnerless women contribute 10 and 40 %, respectively. Therefore, given that the DL clarifies the procedure to terminate a marriage, cohabitating couples may decide to get married and have a first child, which may explain in part the relative stability of the proportion of first marriages and the positive trend in first births after 2005.

Another important factor in understanding women's first birth decisions and the contribution of married and cohabiting women is the trend in average age at first marriage for women and the average age at which they have their first child. The trends for both are positive, suggesting that Chilean women have become less willing to sacrifice certain personal development for marrying and starting a family (see Fig. 2 below). This trend is not surprising given the increased participation of women in the labor market and significant increases in female educational attainment (see Larrañaga 2004). Moreover, the overall fertility rate has declined since the mid-1960s (Economic Commission for Latin America and Caribbean, ECLAC, and INE 2002). However, after the DL was approved, administrative data from INE suggests a decrease in the average age at which women give birth to their first child: in the period 1997–2005 this average increased from 22.73 to 23.39, but fell to 23.03 in 2006 and 23.14 in 2007. Although it is still too early to conclude that there is a change in the trend, it appears that the average age at first pregnancy has declined after the DL.

Moreover, a simple comparison of both trends in Fig. 2 suggests that more Chilean women have the first child before getting married for the first time.[6] This trend is consistent with the increasing preference towards cohabitation mentioned before. Given the DL, and considering that 45 % of Chileans agree with the idea that cohabitating couples should get married when they have a child (35 % disagree and the rest are indifferent), according to the 2008 Bicentennial Survey (see Salinas

[5] These figures are provided by the National Institute of Statistics of Chile (INE). The INE produces the Vital Statistics Report (*Reporte de Estadísticas Vitales*) annually, using administrative information on births, marriages and other variables collected from birth registry offices. In Fig. 1 the number of marriages for the years 2001, 2002 and 2004 is estimated using the annualized growth rate of first marriages for the periods 2000–2003 and 2003–2005, respectively. The number of first births for the years 2001 and 2002 is also estimated using the annualized growth rate of the number of first births for the period 2000–2003.

[6] The administrative data do not distinguish the average ages at first marriage of women who have children and women who do not.

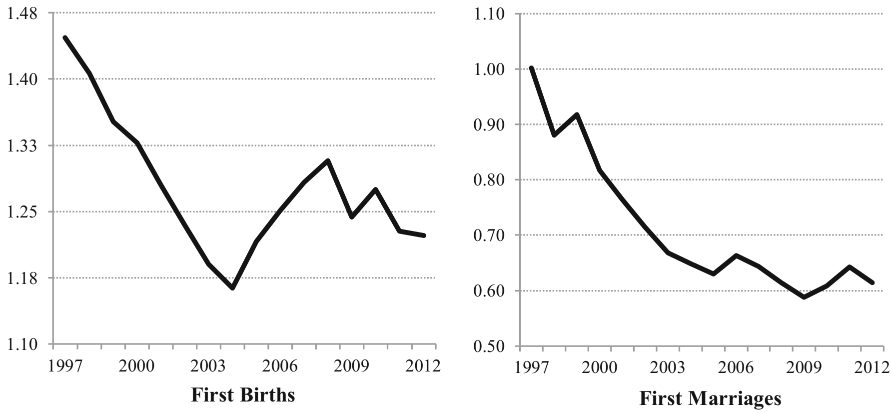

Fig. 1 Proportion of first births and first marriages for women in Chile 1997–2012, with respect to total female population (%). *Source*: National Institute of Statistics, Chile

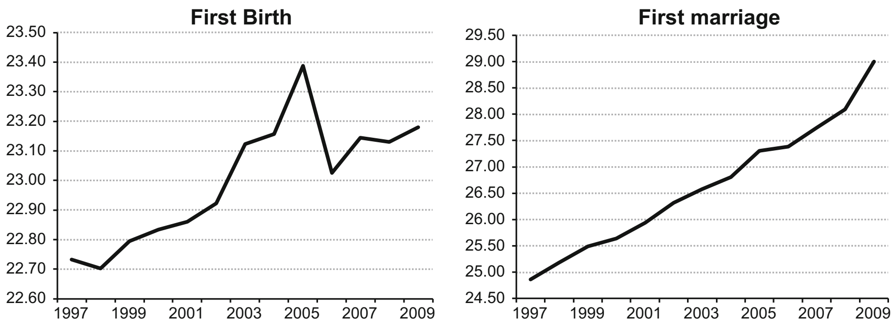

Fig. 2 Average age at first birth and first marriage for women in Chile 1997–2008. *Source*: National Institute of Statistics, Chile

2008), it is possible that more cohabitating women have the first child and marry afterwards.

The evidence presented in Figs. 1 and 2 suggests that the DL had an effect on first births for both married women and cohabitating women. Married women may delay the first birth to see if the relationship succeeds, while in the case of cohabitating women the effect of the DL on first birth may be related to the decision of whether to get married. However, the available data do not suggest a substantial change in the average age of women at first marriage after the passage of the DL.

The trends in first marriages and first births may seem surprising considering that Chilean society is very traditional—the Chilean Marital Law of 1884 aimed to protect this tradition. Considering that Chile has reached a high level of development and economic growth in the last 20 years, the negative trends shown in Fig. 1 are less surprising. In fact, decreasing marriage and fertility rates are also observed in many western European countries. For instance, between 1998 and 2002 the largest declines in the birth rate are observed in Portugal (19.7 %), Malta (18.4 %), Germany (16 %), Luxemburg (13.5 %), and the Netherlands (11 %).

Other Eurozone countries, which receive larger flows of migration or have child-benefit programs and a relatively more flexible post-natal legislation, show a positive or relatively lower decreasing trend in birth rates in the same period: for instance, Spain (28.9 %), Greece (15.6 %), France (0 %) and Italy (−2.6 %).[7] Since Chile does not have a child benefit program as many European countries do and post-natal leave was extended to 3 months only in 2011,[8] we would not expect the effect of the DL to be confounded by any other child-related legislation change.[9]

It is also important to consider unplanned pregnancies in Chile (see Diaz, year unknown). Unfortunately, administrative records on pregnancies do not distinguish between planned and unplanned pregnancies. As far as under-age pregnancies are concerned, Molina et al. (2004) reports that most of the under-age pregnancies are to eventual high-school dropouts. The conclusions on first births presented in this section do not change when women with first conceptions at age 19 or younger are included in the analysis.

3 Literature review

The US literature on divorce law has examined the introduction of no-fault and unilateral divorce in the 1970s.[10] No-fault divorce means a divorce granted without a "fault" by one of the spouses (for instance, adultery, physical or mental abuse, and abandonment), while unilateral divorce means that the consent of both spouses is not required. These new laws passed at different times in different states,[11] and much of the literature has taken advantage of this variation to identify the effects of these laws. The literature on divorce laws and fertility decisions is more scarce.

Alesina and Giuliano (2007) use unilateral divorce to analyze the relationship between divorce and fertility. They claim that more flexible divorce laws have two possible effects. The first effect is that the number of children born to married couples decreases, and people have more children out of wedlock. The second effect is that the cost of entering the 'wrong' marriage is lower, because it is easier to dissolve the relationship. Alesina and Giuliano say that the literature on divorce law has focused on the first effect and has not given sufficient attention to the second effect.

[7] In general, between 1998 and 2009, the 15 countries of the Eurozone decreased by 0.8 %. http://ec.europa.eu/eurostat.

[8] In 2013, a benefit scheme was introduced to provide financial support to couples with more than two children.

[9] In 2005, new domestic violence legislation was introduced. However, there is no clear evidence of the effect of domestic violence on fertility decisions. Moreover, to our understanding, there are no exclusive social programs targeting single mothers.

[10] Peters (1986, 1992), Allen (1992), Zelder (1993), Weiss and Willis (1997), Friedberg (1998) and Wolfers (2006) analyze the effects of divorce laws on divorce rates. Regarding the relationship between divorce laws and women's labor supply, the literature focuses on the response of women's labor supply after divorce: Johnson and Skinner (1986), Gray (1998) and Stevenson (2007). Regarding human capital accumulation, King (1982) and Stevenson (2007) analyze the effect of divorce laws on couples' investments in the education of spouses and children, as well as household specialization.

[11] Gruber (2004) provides a summary of when these laws were implemented in each state.

To analyze both effects, Alesina and Giuliano use birth certificates from the National Vital Statistics of the US to construct different measures of fertility at the state level. Using a panel of states for the period 1968–1999, they recover the impact of the introduction of unilateral divorce on state fertility rates. Their results show that the introduction of no-fault divorce is associated with a decline of 3 % in the fertility rate for those states that introduce the new legislation. Alesina and Giuliano also construct a panel of state-age records based on Census data from 1960 to 1990. Their dependent variable is the number of children ever born to women residents aged 15–44 in states that adopt unilateral divorce. Their results corroborate the negative effect of no-fault divorce on fertility rates. They conclude that as divorce becomes easier, couples feel less locked in, and women who are considering having children or are already pregnant are more willing to "try" marriage. Therefore out-of-wedlock fertility declines and marriage rates increase.

Drewianka (2008) tests the effect of no-fault and unilateral divorce law in the United States on US marriage divorce, fertility and legitimacy rates. As far as the latter two rates are concerned, he finds that no-fault divorce lowers birth rates by about 2–3 %, with the effect concentrated on married couples, whereas unilateral divorce increases marital fertility but decreases non-marital birth rates. Moreover, the negative effect of unilateral divorce on non-marital fertility becomes stronger as time passes.

A study by Ekert-Jaffe and Grossbard (2008) relates directly to the monetary compensation introduced in the DL. Ekert-Jaffe and Grossbard develop a model of fertility in which it is assumed that a woman makes decisions regarding whether to have marital fertility, non-marital in-couple fertility, or out-of-couple fertility. The less to be gained from marital or in-couple fertility, respectively, the more likely a woman is to choose non-marital or out-of-couple fertility. The authors demonstrate that laws regulating divorce settlements influence the benefits of in-couple fertility relative to out-of-couple fertility. Using a sample of close to 20 different legal regimes, Ekert-Jaffe and Grossbard find that the likelihood of an unpartnered birth is lower when a country or province has community property laws guiding division of assets in case of partnership dissolution that are more favorable to women than the alternatives. The most convincing case is that of New Zealand, where there were large decreases in out-of-couple fertility after community property law favorable to women was introduced.

In a recent paper Bellido and Marcen (2014) examine the effect of no-fault unilateral divorce law reforms on fertility for the case of Europe. They find that the negative effects of divorce law reform on marital fertility rates are transitory. On the other hand, the negative effects on non-marital fertility are permanent. Thus, in both studies by Drewianka and by Bellido and Marcen there are strong negative effects on non-marital birth rates, and mixed or transitory effects on marital birth rates.

We conclude that the literature on divorce law and fertility decisions is scarce and focuses on the response of fertility rates. While rates are important, there are other important aspects of fertility to consider. The overall fertility rate in Chile has decreased from 2.09 to 1.9 births per women between 2003 and 2010,[12] and it is

[12] CIA World Factbook. https://www.cia.gov/library/publications/the-world-factbook/geos/ci.html.

possible that the introduction of the DL in 2004 may partially explain this trend. However, the monetary compensation provisions of the DL may have substantially different effects for different people, calling for a micro-level analysis.[13] In the next section we present the data and the empirical strategy, and discuss our results.

4 Data and estimation

According to the Vital Statistics Report 2001–2008 provided by INE, Chilean fertility rates are declining. At the same time, the data suggest that the DL has shortened the time to first births for women with more than 12 years of schooling in the period 2005–2008.[14] To quantify the conditional probability of first births controlling for the socioeconomic characteristics of the woman, we estimate a hazard model for first births.

The estimation of the hazard of first births raises several issues. In duration data it is not always possible to observe the occurrence of the event of interest. In the case of first births, some women have a first child during their sample period, while others do not. Biases are introduced if we discard the childless women from the analysis. At the same time, these women cannot be treated as women who give birth to a child. Relevant work on applying hazard analysis to fertility has been developed by Newman and McCulloch (1984). The authors explain that hazard analysis corrects for censoring bias without generating selectivity bias. Kravdal (2001) studies the duration of second and third birth intervals in Norway. Li and Choe (1997) study second birth intervals in China, where the one-child policy creates disincentives for couples to have a second child. The present study uses a semi-parametric Cox proportional hazard model and includes the DL information to examine whether this policy has an effect on women's hazard of having a first child.

4.1 Data: the social security survey (EPS)

The 2002–2009 waves of the Social Security Survey (Encuesta de Previsión Social–EPS), Social Security Subsecretary (2002), provide 8126 birth histories.[15] For the main analysis we use all women who turned 18 in 1980 or later so that the birth histories comprise no more than 27 years starting at age 18, because we believe that the decisions of younger women are less likely to be determined by potential

[13] To our knowledge, the only study that examines the Chilean DL is Heggeness (2009), which analyzes the law's effect on educational enrollment using a difference-in-difference approach.

[14] To see this, we compute the difference between the share of first births for more educated women (those with more than 12 years of education) and the share for women with 10–12 years of education for each year after the DL was passed in 2004. We compare these differences with the corresponding difference for the year 2003, which is one year before the DL was passed. The formula is $DD^{y}_{13} = \left(f^{y}_{13} - f^{y}_{10\,to\,12}\right) - \left(f^{03}_{13} - f^{03}_{10\,to\,12}\right)$ for y from 2005 to 2008. The difference-in-differences increases from 0.77 % in 2005 to 3.06 % in 2008 (see Fig. 2). The result is unchanged if women 19 or younger are excluded from the analysis.

[15] The original sample was collected in 2002. The 2004 wave added around 3000 new individuals to the sample, while the 2006 and 2009 waves update the individuals' information.

education or income.[16] Accordingly, our results should be interpreted as not having conceived their first child at age 18 or before. A woman's birth history is terminated when she conceives her first child (2820 cases), when she turns 45 (1046 women), or when she is no longer observed in the sample.

The socioeconomic information for the woman is time-invariant and sample means are presented in Table 1. However, the indicator for whether the DL has been passed is a time-varying indicator variable. The birth history of a woman conceiving her first child before November 2004 (when the DL was approved), is unaffected by the DL. The DL only affects the birth histories of women conceiving their first child after this date. Therefore, the DL divides the duration of the spell into two intervals.

4.2 The DL and women's expected level of education

As mentioned before, the DL standardizes the procedure to terminate a marriage and the possible monetary compensation may create incentives for a woman to have her first child earlier, as she would be entitled to compensation in case of a future divorce that would make it easier to live alone as a single parent. These incentives are likely to be greater for women with higher levels of education. To recover the effect of the DL we estimate the following equation:

$$\ln h(t) = \gamma_0 T(t) + \gamma_1 T80(t) + \gamma_2 M(t) + \beta_w W + \beta_E E + \delta DL(t) + \theta E \cdot DL(t), \quad (1)$$

where ln $h(t)$ is the log-hazard of first birth at time t, $T(t)$ is a piecewise-linear spline representing the woman's age, $T80(t)$ is a piecewise-linear spline that provides a time trend with 1980 as the origin, $M(t)$ is a piecewise-linear spline representing the number of years since the woman first married or started cohabitating (t_m),[17] Ws a vector of time-invariant socioeconomic characteristics,[18] Es a vector of potential levels of education, and $DL(t)$ is a time-varying indicator for whether the DL is in effect.

The vector E describes the woman's potential education at age 18, and not her actual education level. The purpose of this time-invariant covariate is to recover the final level of education that an 18-year-old woman expects to eventually attain.[19]

[16] The results are conditional on not having conceived before the age of 18. In the administrative data, the percentage of women conceiving their first child at age 18 or before is very small. The calculations presented in the text are based on administrative data. According to the Ministry of Health of Chile (2013) the percentage of conceptions for women under the age of 19 has remained relatively flat from 2004 to 2011 at 15.5 % (there are no figures for first conceptions alone). In the administrative data there were substantially fewer first conceptions for this age group, and their exclusion did not significantly affect the difference-in-difference analysis, and the average age at first birth in Fig. 2.

[17] Less than 3 % of women in the sample are married or cohabitating before reaching 18 years of age. In such cases, we assume that their relationship started at age 18. Before 2009 the survey does not collect information on the timing of the current marriage or cohabitation. Therefore the exact duration of a marriage or cohabitating relationship may be mismeasured when more than one occurs. Less than 18 % of the women in the sample report having terminated a first marriage or cohabitating relationship, and only 11.5 % of non-single women report more than one marital or cohabitating relationship.

[18] The variables in W are the level of education level of the woman's father and mother, an indicator variable for whether the woman married before 2004, and a set of indicator variables to control for the region of residence of the woman when first interviewed.

[19] The actual final level of education is not available in the EPS.

Table 1 Table of means–time—invariant covariates

Time—invariant covariate	Mean	SD	Min.	Max.
Woman's expected education				
Basic education				
Incomplete	0.36	0.479	0	1
Complete	0.22	0.416	0	1
Post-secondary education				
Technical/professional				
Incomplete	0.11	0.314	0	1
Complete	0.10	0.296	0	1
University				
Incomplete	0.09	0.288	0	1
Complete	0.10	0.301	0	1
Graduate education	0.02	0.141	0	1
Woman's place of residence				
Woman lives in same household where born	0.41	0.455	0	1
Missing—woman lives in same household where born	0.17	0.377	0	1
Education of the woman's parents				
Mother's education				
None	0.05	0.217	0	1
Primary				
Incomplete	0.58	0.494	0	1
Complete	0.06	0.246	0	1
Secondary				
Incomplete	0.14	0.347	0	1
Complete	0.13	0.337	0	1
Post-secondary	0.04	0.186	0	1
Doesn't know	0.04	0.200	0	1
Missing	0.24	0.425	0	1
Father's education				
None	0.04	0.206	0	1
Primary				
Incomplete	0.58	0.493	0	1
Complete	0.04	0.199	0	1
Secondary				
Incomplete	0.14	0.343	0	1
Complete	0.13	0.341	0	1
Post-secondary	0.06	0.237	0	1
Doesn't know	0.06	0.240	0	1
Missing	0.27	0.444	0	1
Woman's potential income at age 23[a]				
Woman's potential income at age 23 (mother's education)	8.65	0.43	8.07	9.59

Table 1 continued

Time—invariant covariate	Mean	SD	Min.	Max.
Woman's potential income at age 23 (father's education)	8.52	0.46	7.79	9.53
Woman's potential income at age 23 (mother's and father's education)	8.57	0.45	7.89	9.55
Woman's potential income at age 23 (partner's estimated income)	8.17	0.55	6.98	9.34

[a] Variable included in the selection equation of the Heckman 2-step procedure

The education system in Chile allows the student to choose a professional track or an academic track once they complete primary school (8 years) and enter secondary school (4 years). The professional track is completed in a professional institute, while the academic track typically ends in a university degree. At age 18, an individual has completed primary and secondary school and already started post-secondary education in a professional institute or university.

For the 92 % of women in our sample who are no longer attending school when they are interviewed by the EPS, their potential education is the highest education level they report. The remaining women report attending school in at least one of the EPS interviews. The potential education of these women will differ from their level reported in the survey if they obtain a new degree at the last institution or in the last educational track reported.[20]

The estimates are obtained using the *aML* software developed by Lillard and Panis (2003). Their software estimates a Cox proportional hazard model in which the duration baseline is estimated as a piecewise-linear spline in the log hazard, rather than being given a specific functional form such as Weibull. Table 2 presents the coefficients and marginal effects for the estimation of Eq. (1). In column (A), E is a dummy variable for all women with some university experience, i.e., the baseline category is for women with no university experience, while in column (B) the education indicator is for women with a university degree, i.e., the baseline category is for women with no university degree. Using the estimates in Table 2, 3 and 4 present the effect of DL on the hazard of first birth.

The covariate of interest in Eq. (1) is $DL(t)$, which is included both by itself and interacted with E. Therefore, the coefficient δ gives the effect of the DL for less educated women, while the sum of δ and θ gives the effect of the DL for more educated women.

The results in Table 3 shows that the hazard of first birth for more educated women increases by 61 % with the DL. The effect of the DL for women with lower education is insignificantly different from zero. When we add women who turn 18 between 1970 and 1980 to the sample, the magnitude of the effect of the DL decreases to 47 % and when we further include women who turn 18 between 1960 and 1970 the effect doesn't change.[21] The magnitude of the estimates increases in

[20] Specifically we assume the following: (a) women who last report primary or secondary education will complete that education level; and (b) women who last report enrollment in the final year of secondary education or who report attending post-secondary education in every wave of the EPS will complete the corresponding track of post-secondary studies.

[21] We would like to thank an anonymous referee for suggesting these robustness checks.

Table 2 Hazard of first birth—Eq. (1)—woman's attained education

	(A) Some university experience		(B) University degree	
	Coefficient	Exp (coef.)	Coefficient	Exp (coef.)
Splines				
Woman's age				
18–19	0.2331*	1.2625	0.2317	1.2607
	(0.1415)		(0.1414)	
20–24	−0.012	0.9881	−0.0096	0.9904
	(0.0149)		(0.0149)	
24–35	−0.0477	0.9534	−0.0457***	0.9553
	(0.0105)		(0.0105)	
35 or more	−0.1860***	0.8303	−0.1831***	0.8327
	(0.0664)		(0.0664)	
Time trend				
1980–1990	0.0036	1.0036	0.0025	1.0025
	(0.012)		(0.0119)	
1990–2000	−0.0317***	0.9688	−0.0354***	0.9652
	(0.007)		(0.007)	
2000 or After	−0.1997***	0.8190	−0.2061***	0.8138
	(0.0245)		(0.0245)	
Years since first marriage or cohabitating relationship				
1 or less	0.2250***	1.2523	0.2240***	1.2511
	(0.0084)		(0.0084)	
1–5	0.1830***	1.2008	0.1831***	1.2009
	(0.0208)		(0.0207)	
5–10	−0.1965***	0.8216	−0.1955***	0.8224
	(0.0474)		(0.0477)	
10 or more	0.0225	1.0228	0.0279	1.0283
	(0.0909)		(0.0909)	
Covariates				
Constant	−0.9893***	0.3718	−1.1077***	0.3303
	(0.2006)		(0.2011)	
DL	0.0198	1.0200	0.051	1.0523
	(0.1814)		(0.1739)	
Education level attained[1]	−0.6313***	0.5319	−0.6468***	0.5237
	(0.0587)		(0.0768)	
DL × education level attained	0.4580**	1.5809	0.6900***	1.9937
	(0.1892)		(0.2278)	
Married or started cohabitating before 2004	−0.9490***	0.3871	−0.9088***	0.4030
	(0.1254)		(0.1267)	
Mother's education complete primary or less	0.1030**	1.1085	0.1248***	1.1329
	(0.0467)		(0.0466)	

Table 2 continued

	(A) Some university experience		(B) University degree	
	Coefficient	Exp (coef.)	Coefficient	Exp (coef.)
Mother's education unknown	−0.0603	0.9415	−0.0422	0.9587
	(0.089)		(0.0883)	
Mother's education missing	0.2500***	1.2840	0.2600***	1.2969
	(0.0922)		(0.0936)	
Father's education—incomplete secondary or less	0.0330	1.0336	0.0738	1.0766
	(0.0551)		(0.0545)	
Father's education unknown	0.1507*	1.1626	0.1360*	1.1457
	(0.0783)		(0.0779)	
Father's education missing	−0.0910	0.9130	−0.1131	0.8931
	(0.0859)		(0.0872)	
Ln–L	−15,677.87		−15,692.41	

Sample: women who turned 18 in 1980 or later: 3866 observations

Specifications include a set of 12 dummy variables for region of residence (base category: Santiago de Chile) of the woman when first interviewed

Asymptotic SEs in parentheses; significance: '*' = 10 %; '**' = 5 %; '***' = 1 %

Table 3 Effect of DL on first birth hazard—specification (1A)

Sample	Effect of DL on women			
	Some university experience ($\delta + \theta$)		No university experience (δ)	
	Estimate	Exp (est.)	Estimate	Exp (est.)
1980	0.4778	1.6125	0.0198	1.0200
(n = 3866)	(0.2051)		(0.1814)	
1970	0.3855	1.4703	0.0986	1.1036
(n = 5522)	(0.2023)		(0.18)	
1960	0.3736	1.4530	0.114	1.1208
(n = 6653)	(0.2013)		(0.1798)	

Asymptotic SEs in parenthesis

Table 4. However, the conclusion remains unchanged: the DL increases the hazard for more educated women and has no significant effect for less educated women.[22]

[22] To test for anticipation effects, we substitute the time-varying covariate that goes from 0 to 1 to indicate the implementation of the DL in November 2004, for a duration spline. This duration spline enters the hazard in 1999, when the DL was presented to the congress. We also include nodes in May and November of 2004 for the approval and implementation of the DL. The results (not shown in the paper) suggest that the DL has no anticipation effect on the hazard of having a first child. Indeed, this result is not surprising, given that the DL is the last of multiple attempts to introduce divorce in Chile. Moreover, none of the results in this study change substantively when the indicators for region of residence are replaced with the woman's region of birth.

Table 4 Effect of DL on First birth hazard—specification (1B)

Sample	Effect of DL on women			
	University degree ($\delta + \theta$)		No university degree (δ)	
	Estimate	Exp (est.)	Estimate	Exp (est.)
1980	0.7410	2.0980	0.0510	1.0523
(n = 3866)	(0.2452)		(0.1739)	
1970	0.6142	1.8482	0.1077	1.1137
(n = 5522)	(0.2405)		(0.1727)	
1960	0.6109	1.8421	0.1158	1.1228

Asymptotic SEs in parenthesis

The coefficient β_E recovers the effect of potential education on the hazard of having a first child for the period before the DL passes. The effect after the DL passes is given by $\beta_E + \theta$. The top panel in Table 5 shows these estimates. Before the DL passes, women with more education have a lower hazard of having a first child. However, after the DL passes, the effect of education is insignificant. When we add women who turn 18 before 1980, the conclusion remains unchanged.

In specification (A), the coefficient for the baseline hazard spline is significant between the woman's 18th and 19th birthdays—the hazard rate increases by 26 %. Between the 24th and 35th birthdays, the baseline hazard spline is insignificantly different from zero. After the 35th birthday, the hazard decreases about 16.97 % every year. The estimates are relatively similar in both specifications. The second term, $T80(t)$, captures the effect of the decrease in fertility rates observed in the administrative data. The estimates suggest a negative slope after 1990 in both specifications with a decrease of 3.2 % a year between 1990 and 2000. After 2000, the hazard decreases by about 19.97 % per year. When women who turn 18 in 1970 (1960) or later are included in the sample, we set the origin of the spline in 1970 (1960). The results for $M(t)$ suggest that the hazard of having a first child increases by 23 % in the first year of the relationship. The increase is not as large for each year between the 1st and the fifth year of the relationship. The hazard starts decreasing after the fifth year, with no significant change after the tenth year. Similar results are observed in both specifications. Finally, the education attained by the woman's mother increases the first birth hazard. This result is not surprising since we would expect a relationship between the woman's education and the education of her parents.[23,24] In the next subsection we perform a robustness check on our results by replacing expected education level with potential income.

[23] Ermisch and Pronzato (2010) show that an additional year of parent's education increases their children's education by at least one tenth of a year.

[24] We have also included a piece-wise linear time trends with knots at every 10 years after the start of the sample. The knot at 1990 is critical since it represents the start of the post Pinochet period. The knot at 2000 coincides with a presidential election. Moving the knot from the year 2000 to non-election years increase the variability of the estimate associated with the DL, and hurts our significance levels in the regression where the highest education level is some university experience, but not in the case where the highest education level is university degree.

Table 5 Effect of potential attained education on first birth hazard

Sample	Some university experience (specification 1A)				University degree (specification 1B)			
	Before DL (β_E)		After DL ($\beta_E + \theta$)		Before DL (β_E)		After DL ($\beta_E + \theta$)	
	Estimate	Exp (est.)	Estimate	Exp (est.)	Estimate	Exp (est.)	Estimate	Exp (est.)
1980	−0.6313	0.5319	−0.1733	0.8409	−0.6468	0.5237	0.0432	1.0441
(n = 3866)	(0.0587)		(0.1719)		(0.0768)		(0.1562)	
1970	−0.4805	0.6185	−0.1936	0.8240	−0.4782	0.6199	0.0283	1.0287
(n = 5522)	(0.0481)		(0.1737)		(0.0608)		(0.1621)	
1960	−0.4566	0.6334	−0.1970	0.8212	−0.4574	0.6329	0.0377	1.0384
(n = 6653)	(0.0431)		(0.1747)		(0.0531)		(0.1646)	

Asymptotic SEs in parenthesis

4.3 The DL and women's potential income

In this section we assume that an 18-year-old woman is able to foresee her potential income given her expected education level and her future labor experience and further assume that her decision of when to have her first child will be influenced by the income level she would expect to have in the future.

We predict the natural logarithm of the woman's potential income level at age 25 based on the estimates from a Heckman two-step procedure using data from the EPS 2004. In the wage equation we regress the natural logarithm of hourly income in 2004 on the woman's education, labor force experience and its square. In the selection equation we also include the education level of the woman's mother.[25] Estimates from the wage equation are used to predict log hourly income at age 25 according to the woman's potential education. The prediction does not include the estimate for the inverse of the Mill's ratio, since the goal is to obtain a measure of the woman's expected income, which does not take into account whether the woman chooses to work or not at that particular age.

We estimate a hazard model which includes the interaction of the DL indicator and the woman's potential income:

$$\ln h(t) = \gamma_0 T(t) + \gamma_1 T80(t) + \gamma_2 M(t) + \beta_w W + \beta_I I + \delta DL(t) + \theta I \cdot DL(t), \quad (2)$$

where I is the woman's potential income at age 25. Based on the estimates of column (2) in Table 6, column (2) in Table 7 shows the estimated hazard elasticity of the woman's potential income before (β_I) and after ($\beta_I + \theta$) the DL. When the sample includes women who turned 18 in 1980 or later (first panel), the estimates suggest that a 1 % increase of the woman's potential income is associated to a reduction of 0.82 % in the hazard of first birth before the DL, while for the period

[25] See Heckman (1979).

Table 6 Hazard of first birth—introducing the woman's potential income at age 25

	(2)	(3)	
		A	B
Splines			
Woman's age			
18–19	0.2375*	0.2366*	0.2475*
	(0.1417)	(0.1417)	(0.1418)
20–24	−0.0062	−0.0065	−0.0065
	(0.0149)	(0.0149)	(0.0149)
24–35	−0.0480***	−0.0482***	−0.0481***
	(0.0104)	(0.0104)	(0.0104)
35 or more	−0.1877***	−0.1879***	−0.1880***
	(0.0665)	(0.0665)	(0.0665)
Time trend			
1980–1990	0.0054	0.0054	0.0056
	(0.012)	(0.012)	(0.012)
1990–2000	−0.0278***	−0.0277***	−0.0274***
	(0.007)	(0.007)	(0.007)
2000 or After	−0.1962***	−0.1957***	−0.1961***
	(0.0244)	(0.0244)	(0.0244)
Years since first marriage or cohabitating relationship			
1 or less	0.2187***	0.2190***	0.2195***
	(0.0085)	(0.0085)	(0.0085)
1–5	0.1734***	0.1736***	0.1722***
	(0.0204)	(0.0204)	(0.0204)
5–10	−0.1996***	−0.1996***	−0.1994***
	(0.0469)	(0.047)	(0.0469)
10 or more	0.0211	0.0207	0.0211
	(0.0889)	(0.0891)	(0.0886)
Covariates			
Constant	5.9792***	5.5704***	6.7780***
	−(0.5113)	(0.6619)	(0.653)
DL	−5.1550***	−5.2775***	−4.8625***
	(1.8134)	(1.8192)	(1.8189)
Potential log-hourly income at age 25	−0.8179***	−0.7684***	−0.9128***
	(0.055)	(0.075)	(0.0733)
Educational attainment	−0.0702	0.1851*	
	(0.0766)	(0.0992)	
DL × potential log-hourly income at age 25	0.6079***	0.6219***	0.5743***
	(0.2056)	(0.2063)	(0.2063)
Married or started cohabitating before 2004	−0.9182***	−0.9222***	−0.9236***
	(0.126)	(0.126)	(0.1259)

Table 6 continued

	(2)	(3)	
		A	B
Father's education—incomplete secondary or less	0.0107	0.0086	0.0089
	(0.0497)	(0.0498)	(0.0498)
Father's education unknown	0.1040	0.1058	0.1025
	(0.0678)	(0.0681)	(0.0675)
Father's education missing	0.0923**	0.0966**	0.0833*
	(0.0431)	(0.0432)	(0.0432)
Ln–L	−15,744.04	−15,743.66	−15,742.29

Sample: women who turned 18 in 1980 or later (n = 3866)

Specifications include a set of 12 dummy variables for region of residence (base category: Santiago de Chile) of the woman when first interviewed

Asymptotic SEs in parentheses; Significance: '*' = 10 %; '**' = 5 %; '***' = 1 %

Table 7 Hazard of first births—interacting DL with woman's potential income at age 25

		(2)	(3)	
			A	B
1980	Before DL (β_I)	−0.8179	−0.7684	−0.9128
(n = 3866)	SE	(0.055)	(0.075)	(0.0733)
	After DL ($\beta_I + \theta$)	−0.21	−0.1465	−0.3385
	SE	(0.1989)	(0.21)	(0.2099)
1970	Before DL (β_I)	−0.6438	−0.6630	−0.7793
(n = 5522)	SE	(0.0437)	(0.061)	(0.0602)
	After DL ($\beta_I + \theta$)	−0.2149	−0.2397	−0.391
	SE	(0.1979)	(0.206)	(0.2053)
1960	Before DL (β_I)	−0.5863	−0.6078	−0.7195
(n = 6653)	SE	(0.0389)	(0.0569)	(0.056)
	After DL ($\beta_I + \theta$)	−0.2007	−0.2283	−0.3688
	SE	(0.1974)	(0.2048)	(0.2037)

Sample: women who turned 18 on or after

Asymptotic standard errors in parenthesis

after the DL passed, the hazard elasticity is insignificantly different from zero.[26] These results do not change significantly when the woman's potential income is predicted based on a Heckman 2-step procedure that uses the father's education instead of the mother's or uses both father's and mother's education.

[26] The standard errors for the estimates of the interaction of the DL variable with the woman's potential income have not been adjusted to take into account the fact that the woman's potential income is a generated regressor.

We conclude that the hazard of having a first child increases for women with higher potential income level after the DL was passed. This result is consistent with the positive effect of the DL on the hazard for more educated women from Eq. (1), and does not change when the sample includes women who turn 18 in 1970 or later (second panel), or who turn 18 in 1960 (third panel) or later: before the DL the reduction in the hazard is 0.64 and 0.58 %, respectively, while the effect after the DL is insignificantly different from zero in both cases.

To further investigate these results, we estimate Eq. (2) adding the woman's expected education level (E):

$$\ln h(t) = \gamma_0 T(t) + \gamma_1 T80(t) + \gamma_2 M(t) + \beta_w W + \beta_E E + \beta_I I + \delta DL(t) + \theta I \cdot DL(t). \tag{3}$$

Column (3A) of Table 7 shows the hazard elasticity before and after the DL for Eq. (3) when the education variable reflects university experience.[27] Column (3B) considers the specification when the education indicator reflects university degree. The hazard elasticity in column (3A) for both before and after the DL do not differ significantly from the results shown in column (2) for any of the three samples. However, in column (3B), the decrease in the hazard is significantly larger than in column (2), both before and after the DL. These results suggest that the DL has increased the hazard of women with higher potential income given her education level. Moreover, women with higher levels of education are likely to have higher potential income, corroborating the results in Sect. 4.2.[28,29]

The results described so far are based on women's characteristics at age 18, and do not take into consideration the information of a potential partner. We now explore the possibility that an 18-year-old woman considers the expected income of the future partner in her decision of when to have the first child.[30] The partner's expected income affects the woman's labor force participation decision. Therefore, we predict a new measure of the woman's potential income at age 25 using the estimates from a Heckman two-step procedure that includes the partner's expected income in the selection equation. When we estimate Eqs. (2) and (3) replacing I with the new income measure, the results corroborate the conclusions obtained in

[27] Column (3) of Table 6 shows the full set of results for Eq. (3).

[28] As an alternative specification, we include in Eq. (3) the interaction of DL with the woman's expected education ($E \cdot DL(t)$). The estimates from this specification corroborate the negative effect of the woman's potential income on the hazard of having a first child.

[29] Our regressions do not control for general economic conditions. A sustained economic growth in Chile starting in 1982 has resulted in low levels of unemployment and improvement in women's educational attainment, which we do control for. Therefore it is unlikely that fertility decisions are negatively affected by deteriorating economic conditions.

[30] We use a Mincerian equation to estimate the partner's expected income. Using the EPS 2004 male sample, we regress the natural logarithm of hourly income in 2004 on education, labor force experience, and its square. The equation is used to predict log hourly income at age 25. The EPS provides information about the partner's education only in the marriage histories updated after the EPS 2006 (in earlier waves this information is not collected). However, for a marriage or cohabiting relationship starting before 2004, it is still possible to recover the partner's education from the household members' demographic characteristics if a married woman reports only one marriage (58 % of the sample). We mean-fill by woman's education in the case of missing values.

the previous subsection: before the DL the hazard elasticity is negative, and becomes insignificantly different from zero under the DL. This conclusion does not change when we include women who turned 18 before 1980 in the sample.

4.4 Summary of findings

The main result of our study is that the divorce law that passed in Chile in 2004 decreases the age at first birth for women in Chile with higher expected education levels, whether measured as university experience or university degree. The results are similar when potential income replaces expected education level in the estimation. There are no significant changes in age at first birth for women with lower levels of education. We feel that the important factor underlying these results is the monetary compensation introduced in the DL for divorced women who have sacrificed professional development or advancement for the good of the household. These sacrifices are likely to be greater for women with higher expected education levels and higher potential income.

Our results are consistent with those of Ekert-Jaffe and Grossbard (2008), Drewianka (2008) and Bellido and Marcen (2014) on the assumption that a lower age at first birth is likely to be correlated with higher fertility. Eckert-Jaffe and Grossbard find that unpartnered fertility decreases with the introduction of community property laws, while the latter two studies find that unilateral divorce law has strong negative effects on non-marital birth rates.

A woman who sacrifices professional development for the good of the family may subsequently find herself in a bad marriage that she cannot leave because it would be difficult to raise the child alone. Highly educated women and women with high potential incomes have the most to lose with this outcome. The monetary compensation in the Chilean DL that rewards this sacrifice in case of dissolution of the relationship makes investment in children less risky for these women and they have children earlier. The monetary compensation is an enforcement mechanism for the division of assets that is directly comparable to the community property laws favorable to women in the study by Eckert-Jaffe and Grossbard. Partnered fertility becomes more attractive and unpartnered fertility declines. More generally, the dissolution of a bad marriage is made easier, as with the cases of unilateral divorce studied by Drewianka and by Bellido and Marcen, marriage rates increase and non-marital fertility declines.

5 Conclusions

This study analyzes the effect of the new Chilean divorce law (DL) of 2004 on a woman's decision of when to have her first child. The new DL introduces divorce in Chile and implements a monetary compensation regime, which benefits the spouse who gives up personal and professional development for the good of the household. The effect of the DL may depend critically on this compensation, since it creates an external enforcement mechanism for the transfer of marital resources to the woman in the case where the couple separates. Educated women and women with higher potential income stand to gain the most from this compensation.

Using the 2002–2009 EPS panel, we construct birth histories for women who turn 18 in 1980 or later (3688 women). The hazard model estimates show that, controlling for age, marital duration, education, declining fertility rates, region of residence when the woman was first interviewed and other socio-demographic characteristics, the DL increases the hazard of first births for women with post-secondary education by 61 % (110 % for women with a post-secondary degree). Moreover, the estimates show that before the DL the hazard elasticity with respect to potential income at age 25 is significantly negative but insignificantly different from zero after the DL. These results are in line with trends observed in the administrative data. Birth records provided by the National Institute of Statistics show that the contribution to first births of women who complete secondary education increased from 26 to 30 % in the period from 2004 to 2008. The trend is the same when births from women who are 19 years-old or younger are excluded from the analysis.

While there is a large literature on the effects of no-fault divorce in the United States in the 1970s, there is no consensus on the effects of introducing divorce for the first time on fertility decisions. In this sense, research on the effect of the DL on women's fertility decisions is needed. The estimates we obtain suggest that the DL increased the hazard of having a first child for more educated women because their higher potential income loss while caring for the child is now offset by alimony if the marriage fails. It did not significantly change the hazard for less educated women.

Acknowledgments We would like to thank Susan Gensemer, Alfonso Flores-Lagunes, Pieter Serneels, two anonymous referees for their useful comments and the Subsecretary of Social Security of Chile for providing the data. We are also thankful to the participants at the Gender Economics session at the 2012 Meeting of the Latin American and Caribbean Economic Association. All errors are our own.

References

Alesina, A., & Giuliano, P. (2007). *Divorce, fertility and the value of marriage*. Harvard Institute of Economic Research, discussion paper no. 2136.

Allen, D. (1992). Marriage and divorce: Comment. *The American Economic Review, 82*(3), 679–685.

Becker, G. (1974). A theory of marriage: Part II. *Journal of Political Economy, 82*(2), S11–S26.

Bellido, H., & Marcen, M. (2014). Divorce laws and fertility. *Labour Economics, 27*, 56–70.

Drewianka, S. (2008). Divorce law and family formation. *Journal of Population Economics, 21*, 485–503.

ECLAC, INE. (2002). Chile: Proyección y Estimaciones de Poblacion. Total País. 1950–2050 (Chile: Population Forecast and Estimates. Total Country. 1950–2050).

Ekert-Jaffe, O., & Grossbard, S. (2008). Does community property discourage unpartnered births? *European Journal of Political Economy, 24*(1), 25–40.

Ermisch, J., & Pronzato, C. (2010). Causal effects of parent's education on children's education. Institute for Social and Economic Research. ISER working papers series no. 2010-16.

Friedberg, L. (1998). Did unilateral divorce raise divorce rates? Evidence from panel data. *The American Economic Review, 88*(3), 608–627.

Gray, J. (1998). Divorce-law changes, household bargaining, and married women's labor supply. *The American Economic Review, 88*(3), 628–642.

Gruber, J. (2004). Is making divorce easier bad for children? The long-run implications of unilateral divorce. *Journal of Labor Economics, 22*(4), 799–833.

Heckman, J. (1979). Sample selection bias as a specification error. *Econometrica, 47*(1), 153–161.
Heggeness, M. (2009). *Evidence of shifts in intra-household allocation under exogenous changes in family policy and administrative procedures: The case of school enrollment in Chile*. Agricultural and Applied Economics Association, Annual Meeting, July 26–28.
Johnson, W. R., & Skinner, J. (1986). Labor supply and marital separation. *American Economic Review, 76*(3), 455–469.
King, A. (1982). Human capital and the risk of divorce: An asset in search of a property right. *Southern Economic Journal, 49*(2), 536–541.
Kravdal, Ø. (2001). The high fertility of college-educated women in Norway. *Demographic Research, 5*(6), 187–216.
Larrañaga, O. (2004). *Participación Laboral de la Mujer en Chile: 1958–2003 (Women's Labor Force Participation in Chile: 1958–2003)*. Chile: Department of Economics, Universidad de Chile.
Larrañaga, O. (2006). Comportamientos Reproductivos y Natalidad: 1960–2003 (Reproductive behavior and fertility: 1960–2003). In J. S. Valenzuela, E. Tironi, & T. Scully (Eds.), *El Eslabón Perdido: Familia, modernización y bienestar en Chile*. Santiago de chile, Editorial Tauros.
Li, L., & Choe, M. K. (1997). A mixture model for duration data: Analysis of second births in China. *Demography, 34*(2), 189–197.
Lillard, L., & C. Panis (2003) aML multilevel multiprocess statistical software, version 2.0. Los Angeles, CA: EconWare.
Ministry of Health, Government of Chile. (2013). Situación Actual del Embarazo Adolescente en Chile (Current state of adolescent pregnancy in Chile).
Molina, M., et al. (2004). Embarazo en la Adolescencia y su Relación con la Deserción Escolar (Adolescent pregnancy and its relationship to schooling dropout). *Revista Médica de Chile, 132*, 65–70.
Newman, J., & McCulloch, C. (1984). A hazard rate approach to the timing of births. *Econometrica, 52*(4), 939–961.
Palma, I. (2006). Transformaciones de la fecundidad en la sociedad chilena. http://www.pasa.cl/index.php?option=com_docman&task=doc_download&gid=153.
Peters, H. E. (1986). Marriage and divorce: Informational constraints and private contracting. *The American Economic Review, 76*(3), 437–454.
Peters, H. E. (1992). Marriage and divorce: Reply. *The American Economic Review, 82*(3), 686–693.
Rossetti, J. (1993). Hacia un perfil de la familia actual en Latinoamérica y el Caribe (Towards a profile of the current family in Latin America and the Caribbean). In Cambios en el perfil de las familias: la experiencia regional, Libros de la CEPAL, No. 36 (LC/G.1761-P), Santiago de Chile, Economic Commission for Latin America and the Caribbean (Comisión Económica para América Latina y el Caribe–CEPAL).
Salinas, V. (2008). *Matrimonio y Convivencia a la Luz de la Encuesta Bicentenario (Marriage and cohabitation in the bicentennial survey)*. Institute of Sociology, Bicentennial Survey Seminar 2008 (Instituto de Sociología, Seminario Encuesta Bicentenario 2008).
Social Security Subsecretary (Subsecretaría de Previsión Social). Chile. (2002). Primera Encuesta de Protección Social (First survey of social security). www.proteccionsocial.cl.
Stevenson, B. (2007). The impact of divorce laws on marriage-specific capital. *Journal of Labor Economics, 25*(1), 75–94.
Weiss, Y., & Willis, R. J. (1997). Match quality, new information, and marital dissolution, Part 2: Essays in Honor of Yoram Ben-Porath. *Journal of Labor Economics, 15*(1), S293–S329.
Wolfers, J. (2006). Did unilateral divorce laws raise divorce rates? A reconciliation and new results. *American Economic Review, 96*(5), 1802–1820.
Diaz, S. (Year unknown). *Información Sobre la Situación de la Planificación Familiar en Chile (Information on the Situation of Family Planning in Chile)*. Instituto Chileno de Salud Reproductiva. http://www.icmer.org/documentos/anticoncepcion_de_emergencia/inf_situac_planific_familiar_chile.pdf.
Zelder, M. (1993). Inefficient dissolutions as a consequence of public goods: The case of no-fault divorce. *The Journal of Legal Studies, 22*(2), 503–520.