Are Intra-Household Allocations Policy Neutral? A Theory and Some Evidence*

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Abstract

We develop a collective household model with spousal matching in which there exists marital gains to assortative matching and marriage quality for each couple is revealed ex post. Marriages, intra-marital allocations and divorce are determined endogenously. Changes in alimony laws affect existing couples and couples-to-be differently. For existing couples, legislative changes that favor (wo)men benefit them especially if the marriage match quality is low. For couples not yet married, however, they generate offsetting intra-household transfers and lower intra-marital allocations for the spouses who are the intended beneficiary. Thus, changes in alimony rights produce policy neutrality in the marriage markets. We then estimate the effect of granting alimony rights to cohabiting couples in Canada. Canadian provinces extended these marital rights to cohabiting couples in different years and using different eligibility rules in terms of cohabitation length, allowing precise estimates of their causal effects in a triple-difference framework. We find that obtaining the right to petition for alimony led women to lower their labor force participation as these laws increased their bargaining power within the household. These results, however, do not hold — and, in some cases, are reversed — for cohabiting couples whose unions formed after the legislative changes occurred.

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1 Introduction

Since the seminal work of Becker (1973), economists have greatly progressed in their understanding of household behavior. There is now a fair level of agreement on the fact that spousal bargaining determinants, such as divorce laws and marriage-market sex ratios, affect the allocation of household resources between partners (Chiappori et al., 2002, and Angrist, 2002). However, if individuals are forward-looking and there are changes in the determinants of spousal bargaining power, individuals might alter their behavior before entering unions (as discussed in Lafortune, 2010). Moreover, altered expectations might also impinge upon who matches with whom. If either of these factors are at play, one would expect that individuals who are "caught" by changes in one of these determinants would behave differently than those who are able to react to the policy before entering into a union.

In this paper, we present some theory and empirics that highlight such marriage-market and intra-household allocation effects. To that end, we present an *integrated collective household* model where the matching process as well as the prospect of divorce or separation are embedded into the collective analysis.[1]

The main ingredients of our model are as follows: There is a continuum of men and women who live for two periods. Each agent is characterized by a single attribute, income (or human capital), with continuous distributions of incomes on both sides of the marriage market, so that each agent has a close substitute. The economic gains from marriage arise from joint consumption of a public good and a non-monetary common factor that is match specific. This match quality for each couple is revealed ex post and those with poor matches may divorce. Finally, we rely on a 'Becker-Coase' framework, in the sense that utility is transferable both between spouses and after divorce.

We employ this model to investigate the impact of changes in the policy environment. We specifically consider a reform that increases the wives' share of incomes *after divorce*, such as a universal increase in the mandatory (share of) alimony payments. Such a change in post-divorce property rights cannot affect divorce probabilities in our Becker-Coase world. However, it can influence the allocation of resources within a household, both before and

[1] The collective models encompass the early- and late-generation marital bargaining theories, whereby spousal preferences affect household choices via an intra-household sharing mechanism. However, the collective models also cover more generalized versions that only assume that household decisions are Pareto efficient, while remaining agnostic with respect to the precise bargaining mechanism with which intra-household allocations are made. Among the earliest examples of the collective model are Becker (1981), Chiappori (1988, 1992), and Bourguignon and Chiappori (1994), and those of marital bargaining are Manser and Brown (1980), McElroy and Horney (1981) and Sen (1983).

after divorce — even among couples who do not eventually divorce.

We show that the short- and long-term consequences of the reform are different and generally opposite of one another. For *partnerships already in existence* at the time of the legislative change, an increase in mandatory alimony payments can only improve the wives' welfare at the husbands' expense. While the exact scope of the reform depends on assumptions regarding commitment, either some or all women will strictly gain from the reform and no woman can lose (equivalently, no man can gain). Regarding couples who marry *after* the reform, the logic is quite different, because the new divorce settlement is taken into account at the matching stage, resulting in a different inter-temporal allocation of resources and welfare between spouses. Specifically, a change in alimony settlement laws aimed at favoring women typically generate offsetting intra-household transfers, eventually resulting in *lower* intra-marital allocations for all married women.

To the extent that there are non-transferabilities in spousal utility, the adjustments we described above would not be complete. Consequently, one would see an increase in the dissolution of partnerships that were formed *before* the legislative changes as well. But, due to the fact that adjustments in spousal utility payoffs would have to be made for partnerships to be viable given the new outside options, one ought to detect no impact on the dissolution rate of partnerships formed *after* the legislative changes.

In the second part of our paper, we turn to an empirical exploration of our model by exploiting the legislative changes which granted the right to petition for alimony upon separation for cohabiting couples in Canada over the last 35 years. The fact that these new laws were implemented at different times in different provinces with different eligibility rules enables us to convincingly estimate the causal effect of these rules. Furthermore, one can easily distinguish between couples who started their relationships before and after the legislative changes. Our empirics thus compare the causal estimates of granting alimony rights to partnerships already in existence, when the new rules were implemented, with those that potentially reflect how individuals respond to these changes before entering into a union.

Empirically, we estimate the impact of granting cohabiting couples the capacity to petition for alimony upon separation. Finding the impact of such a legislation is not easy as there is an obvious endogeneity problem: regions that implement such a rule may be distinct from those that do not. Similarly, comparing couples who "register" their union with those who do not is not likely to lead to a causal effect, due to the obvious selection bias. Furthermore, in the case of cohabiting couples, few countries have implemented rules with variations which allow the construction of a credible "control group" for the estimation of a causal impact (see Rangel, 2006, for a notable exception).

The context studied here is particularly interesting because not only were "common-law spouses" — as cohabiting couples are called in Canada — granted alimony rights at different time periods in different provinces, but also each province defined the length of cohabitation required to qualify for such rights differently. This provides a very rich source of variation for our current analysis in which we employ a triple-difference strategy (based on province, time and relationship duration) in order to identify the causal impact of the legal change. Furthermore, many of these legal changes were implemented not in response to a demand from cohabiting couples but as a way to offer homosexual couples — who, at that time, were unable to legally marry — the same legal protection as their heterosexual counterparts, thus diminishing the potential problem of endogenous adoption of the laws.

Using labor supply as a proxy for one's share of household resources (something that has been employed previously but mostly in the context of married couples), we directly search for evidence that alimony rights influence spousal bargaining power. Alimony rights are likely to benefit women as they are rarely granted to men. Moreover, men are still more likely to be earning more than their companions, rendering it more likely that transfers are made from men to women even when laws operate on the basis of 'equitable distribution' principles.

The results obtained here suggest that as cohabiting relationships were suddenly granted alimony rights, women were more likely to attend school and stop working and less likely to work full-time whereas men became more likely to work and less likely to study or have work interruptions. These results hold within a given relationship over time, but they do not apply to individuals who were *married*, as those already benefited from these rights and thus were unaffected by the new laws. More importantly, however, we find contrasting outcomes for the new alimony rights' impact on the behavior of cohabiting couples who entered a union *after* the alimony rights were granted: among such couples, the impact of the law is limited and when observed, it is women — and not men — who were less likely to study and have fewer work interruptions, whereas they were more likely to work or work full time.

Our results suggest that the institution of alimony rights for cohabiting couples led to longer periods of cohabitation but also that fewer of these unions eventually lead to marriage. This appears to be only economically and statistically significant for couples who were matched before the legislative changes occurred, as we see no such effects among couples who entered a cohabiting union after they were introduced.

These results contribute to our understanding of the dynamics between cohabitation and marriage, a topic that has been mostly the focus of sociologists and demographers (see Smock, 2000, for a review). Couples who marry after cohabitation have lower marriage quality in terms of length of relationship, propensity to divorce, etc., although this does not seem to be the case in most recent years (Schoen, 1992). And the only Canadian study on the subject (White, 1987) reaches the opposite conclusion that cohabitation reduces the probability of divorce upon marriage. Also, there are studies which document that children who live in cohabiting households perform worse in most measures (see Manning, 1995, and Manning, 2001). Amador and Bernal (2008) attempt to correct for the obvious endogeneity problem in these comparisons, but still find that children with cohabiting parents have worse outcomes than those with married parents in Colombia (despite the fact that both types of households have the same rights).

An empirical application of our main 'marriage-market induced policy-neutrality' idea in an economic development context is provided by Ambrus et al. (2010). They document that *mehr,* a form of Islamic bride-price which functions as a prenuptial agreement in Bangladesh due to the practice of it being only payable upon divorce, influences *dowries* positively in the marriage markets.

To the best of our knowledge, the only paper that has explored the impact of alimony rights on cohabiting households is Rangel (2006), who also finds that such a rule decreases female labor supply. He obtains a causal estimate of granting alimony rights to cohabiting women in Brazil by using the fact that couples with children obtained such a right, but not those without. The identification assumption we use here has the advantage of relying on a much more similar control group through the use of a tripe-difference estimator. Nevertheless, our key contribution lies in our empirics' capacity to estimate the effects of changing spousal bargaining power in existing unions and comparing them with those in relationships yet to be formed.

Finally, our results also mirror those obtained in the case of divorce laws in the United States (Peters, 1986, Friedberg, 1998, Chiappori et al., 2002, and Wolfers, 2006).

The rest of the paper is organized as follows: Section 2 presents our theoretical framework; Section 3 summarizes the legal setting; Section 4 covers our estimation methodology and the data, while the subsequent section presents the empirical results. Our final section then concludes.

2 The Model

2.1 Preferences

The economy is made up of individuals who live for two periods. They are characterized by their income, y for men and z for women. In each period, they derive utility from consumption of n private goods, $q^1, ...q^n$ and N public goods $Q^1, ..., Q^N$.[2] Let $p^1, ...p^n$ and $P^1, ..., P^N$ denote the corresponding prices, with the normalization $p^1 = 1$. Married people also derive satisfaction from the quality of their match, θ. The husbands' and wives' individual utilities take the form

$$U_i = u_i(q_i, Q) + \theta, \qquad i = h, w, \tag{1}$$

where $q_i = (q_i^1, ...q_i^n)$ is the vector of private consumption of member i, $Q = \left(Q^1, ..., Q^N\right)$ is the vector of public consumption by the couple, and θ is the quality of the couple-specific match.[3]

In order to remain as close as possible to the standard 'Becker-Coase' framework, which relies on transferable utilities, we assume that preferences of married individuals are of the *generalized quasi-linear* (GQL) form (see Bergstrom, 1989).

$$u_i\left(q_i, Q\right) = A\left(Q\right) q_i^1 + B_i^m\left(Q, q_i^{-1}\right) + \theta, \tag{2}$$

where $Q = \left(Q^1, ..., Q^N\right)$ and $q_i^{-1} = (q_i^2, ...q_i^n)$. Here, A and B_i^m, $i = h, w$, are positive, increasing, concave functions such that $A(0) = 1$ and $B_i^m(0) = 0$, and good 1 is the 'numeraire' that can be used to transfer utility between spouses at a constant 'exchange rate'.

Similarly, when single or after divorce, preferences take the *strictly quasi-linear* form:[4]

$$u_i^s\left(q_i, Q\right) = q_i^1 + B_i^s\left(Q, q_i^{-1}\right), \tag{3}$$

[2]The number of public and private goods need not be strictly greater than one. Our main conclusions go through intact in a more specific version of the model in which $n = N = 1$.

[3]In our empirical investigation, we shall study the impact of the legal granting of alimony rights to partners in a cohabitation relationship in Canada. In this sense, we use the terms 'husband' and 'wife' loosely to refer to the man and the woman in a cohabitation relationship, respectively.

[4]Both GQL preferences when married *and* quasi linear utilities when single are necessary to generate the Becker-Coase benchmark in which, in a static context, divorce laws do not affect divorce probabilities; see Clark (1999) and Chiappori, Iyigun, Weiss (2007).

Since one of our primary objectives is to explore if and when alimony divorce laws affect divorce rates, we adopt these preference specifications as our stringent benchmark. See Chiappori, Iyigun and Weiss (2008) for an example of a household model which generates a strictly linear Pareto frontier after divorce.

where, again, the B_i^s, $i = h, w$, are increasing concave functions, with $B_i^s(0) = 0$. This utility is quasi-linear; in particular, the optimal consumptions of public and private goods other than good 1 are given by the conditions:

$$\frac{\partial B_i^s\left(Q, q_i^{-1}\right)}{\partial Q^j} = P^j, \;\; 1 \leq j \leq N \qquad \text{and} \qquad \frac{\partial B_i^s\left(Q, q_i^{-1}\right)}{\partial q_i^k} = p^k, \;\; 2 \leq k \leq n \; .$$

Neither these conditions nor the optimal levels of all private and public consumptions (except for good 1) depend on income. Let the latter be denoted $\left(\bar{Q}, \bar{q}_i^{-1}\right) = \left(\bar{Q}^1, ..., \bar{Q}^N, \bar{q}_i^2, ...\bar{q}_i^n\right)$. To simplify notation, we choose units such that $B_i^s\left(\bar{Q}, \bar{q}_i^{-1}\right) = \sum_{j=1}^N P^j \bar{Q}^j + \sum_{k=2}^n p^k \bar{q}_i^k$, $i = h, w$. Then, the indirect utility of a single person equals his or her income.

If a man with income y is matched with a woman with income z, they can pool their incomes. Given GQL preferences, utility is *transferable* between spouses. There is a unique efficient level for the consumption of each of the public goods and each of the private goods 2 to n. Moreover, these levels depend only on the total income of the couple. The Pareto frontier is linear and given by

$$\begin{aligned} u_h + u_w &= \max_{\left(Q, q_h^{-1}, q_w^{-1}\right)} \left\{ \begin{array}{c} A(Q)\left[t - \sum_{j=1}^N P^j Q^j - \sum_{k=2}^n p^k \left(q_h^k + q_w^k\right)\right] \\ + B_h\left(Q, q_h^{-1}\right) + B_w\left(Q, q_w^{-1}\right) \end{array} \right\} + 2\theta \\ &\equiv \eta(t) + 2\theta, \end{aligned} \tag{4}$$

where $t \equiv y + z$ is the *total* family income while u_h and u_w are the attainable utility levels that can be implemented by the allocations of the private good q^1 between the two spouses, given the efficient consumption levels of all other goods. Assuming, as is standard, that the optimal public consumptions are such that $A(Q)$ is increasing in Q, we see that $\eta(t)$ is *increasing and convex* in t.[5]

Due to the consumption of public goods, the two individual traits, y and z of a couple, are *complements* within the household. This generates positive economic gains from marriage in the sense that the material output $\eta(t)$ the partners generate together exceeds the sum of the outputs that the partners can obtain separately. Specifically, the marital surplus $\eta(t) - t$ rises with the total income of the partners, t, and it vanishes only when both partners have

[5] By the envelope theorem, the derivative $\eta'(t)$ is equal to $A(Q)$. Therefore, η is increasing in t and, if $A(Q)$ is increasing in t as well, then η is convex. Note that a sufficient (but, by no means, necessary) condition is that public consumptions are all normal.

no income.

For any couple, match quality θ is drawn from a fixed distribution Φ with a mean $\bar{\theta} \geq 0$. Upon union, both spouses expect to derive the same non-monetary utility from marriage, $\bar{\theta}$. At the end of the first period, the match quality is revealed; a realized value of θ that is below the expected level $\bar{\theta}$ constitutes a negative surprise that may trigger legal separation.[6]

2.2 Family Decisions and Commitment

An important modeling issue is how families make decisions and, in particular, whether or not they can commit to agreed-upon individual allocation schemes. In our model, the decision variables for a couple in each period are the amounts of the public and private goods that the couple purchases and the division of the 'numeraire' private good between them. At the beginning of each period, partners agree to buy the unique efficient levels of all goods but the numeraire, namely the quantities which shift the linear utility Pareto frontier outward as much as possible. There are no commitment issues involved here because (by construction) the levels of consumption within a period cannot be changed and each spouse can predict that, in the second period, consumption will be chosen at the unique and efficient level.[7]

Concerning the division of the numeraire good, however, there is a conflict between the two partners and the question is how it is resolved. As we shall show, competition at the time of marriage fully determines the expected lifetime utility shares of the partners. The marriage market is cleared by a set reservation values of the expected lifetime utility that each agent requires to match with anyone of the opposite sex. Each agent 'marries' the spouse that provides him or her with the highest surplus given these requirements (see Browning et al., in progress, Chapter 8). This outcome requires the possibility of bidding away potential spouses by offering them a larger amount of the private good within the union. In the two period context discussed here, the share of the private good that each married spouse receives can vary across time and the second-period division is anticipated when partners choose to unite. We shall consider here two cases: Either the next period allocation is determined by some known mechanism such as Nash Bargaining and the marriage market clears based only on the flexibility in the first-period allocations. Or alternatively, partners can sign binding

[6] One could incorporate random income shocks into our model. In that case, such shocks could trigger divorce as well, but our qualitative conclusions would not be altered as long as the shocks in question are transitory.

[7] One could also imagine that partners play a non-cooperative contribution game that ends up with lower utility for both spouses. But given that the efficient level of Q can be easily implemented simply by buying and consuming that quantity of the public good, such an assumption is hard to justify here.

contracts which determine allocations in both periods, with the agreed-upon, second-period allocations being relevant only in marriage.

2.3 Endowments

There exists a continuum of men and a continuum of women. The measure of men is normalized to unity and the measure of women is denoted by r, where $r \gtreqless 1$. Each man receives an idiosyncratic income at the beginning of each period; their incomes, denoted y, are distributed over the support $[y_m, y_M]$, $0 < y_m < y_M$, according to some distribution F. Similarly, each woman gets an income z at the beginning of each period, and the z's are distributed over the support $[z_m, z_M]$, $0 < z_m < z_M$ according to the distribution G.

Following divorce, there can be income transfers (i.e., alimony payments) between the ex-spouses. We assume here that these transfers are fully determined by law and no further voluntary transfers are made. Specifically, if a man with income y marries a woman with income z, her income following divorce is $z' = \beta(y+z)$ and his income is $y' = (1-\beta)(y+z)$.[8] Note that the net income of a divorced person is generally different from what his or her income would have been had he or she not married.

Income in our model can be interpreted as either labor or property income.[9] Redistribution corresponds to a legal approach where property incomes or spousal earnings are treated as a common resource and each spouse has some claim on the income of the other. The special case in which all incomes are considered private, implying no redistribution via alimony payments, is represented by a β that is couple-specific, namely $\beta \equiv \frac{z}{y+z}$.

[8]In essence, this means that the wife is granted an alimony payment equal to $\beta y - (1-\beta)z$ upon legal separation (or, in the case of marriage, divorce). And, analogously, the husband gets $(1-\beta)z - \beta y$ following legal separation. Thus, net transfers flow from the husband to the wife if and only if $\frac{y}{z} > \frac{1-\beta}{\beta}$.

[9]For simplicity, we do not allow savings or human capital investments during marriage so that both property and human capital are constant. Given that we abstract from savings and the accumulation of wealth or human capital, the distinction between the post-divorce division of property and alimony payments is mostly semantic here. But one can intepret the variables y' and z' as the stream of incomes generated from the (underlying) assets of the couple which were redistributed according to the alimony laws that apply in legal separation (or divorce).

Labor supply is also assumed fixed. Endogenizing the labor supply decisions could complicate our analysis although, in the context of the assortative model we employ here, it would not alter the patterns of matching in the marriage market. Nevertheless, as shown by Mazzocco and Yamaguchi (2007) and Stevenson (2008), spousal labor supply decisions could be influenced by the property division laws upon divorce.

2.4 The Marriage Market

In the first period, all men and women wish to 'marry' because the expected economic and non-monetary gains from marriage are positive. For heuristic purposes, assume that $r > 1$, so that some women remain single. As usual, we solve the model backwards, starting with the legal separation (or, alternatively, the divorce) decision.

2.4.1 Stable Matches and Lifetime Utilities

Divorce At the end of the first period, the true value of match quality is revealed and each partner of a couple (y, z) can decide whether or not to stay in the marriage, based on the realization of θ. Because utility is transferable within marriage and upon divorce, the Becker-Coase theorem applies and divorce occurs whenever the total surplus generated outside the relationship is larger than what can be achieved within it.[10] Denoting total income of the partners by $t = y + z$, divorce occurs whenever

$$\eta(t) + 2\theta < t \ , \tag{5}$$

or, equivalently,

$$\eta(t) + 2\theta < t \Leftrightarrow \theta < \hat{\theta}(t) = -\frac{1}{2}[\eta(t) - t] \ . \tag{6}$$

In words, a union dissolves if the sum of the outside options, here t, exceeds $\eta(t) + 2\theta$, implying that reservation utilities are outside the Pareto frontier if the partnership continues.

On this basis, the ex-ante probability of divorce for a couple with endowments of y and z is

$$\alpha(t) \equiv \Phi[\hat{\theta}(t)] \ . \tag{7}$$

Note that the threshold $\hat{\theta}(t)$ rises with the income of the couple, t, and consequently the probability of divorce $\alpha(t)$ declines. Because of the complementarity of individual incomes in the household production process, the economic loss generated by divorce is higher for wealthier couples.

The expected marital output (i.e. sum of utilities) generated over the two periods of a couple with incomes y for the husband and z for the wife is

[10] See Clark (1999) and Chiappori, Iyigun and Weiss (2007) for detailed investigations of the transferability in the presence of public goods.

$$S(t) = \eta(t) + 2\bar{\theta} + [1 - \alpha(t)] \left\{ \eta(t) + 2E\left[\theta \mid \theta \geq \hat{\theta}(t)\right] \right\} + \alpha(t)\, t\ .$$

Note, first, that $S(t) > 2t$, because $\eta(t) \geq t$ and $E\left[\theta \mid \theta \geq \hat{\theta}(t)\right] > \bar{\theta} \geq 0$. Thus, all individuals prefer to get married rather than stay single. Secondly, $S(t)$ is increasing in t, hence in each partner's income. In particular, whenever women strictly outnumber men so that $r > 1$, women belonging to the bottom part of the female income distribution remain single. Finally, individuals sort positively into unions. Indeed, since the 'marriage' surplus only depends on total income t, the cross partial $\partial^2 S/\partial y \partial z$ is equal to $S''(t)$. One can readily prove that $S(t)$ is convex and therefore that the traits of the two partners are *complements* even after the risk of divorce is taken into account.[11]

Matching: Who Marries Whom? Given the results of transferable utility and the complementarity of individual incomes in generating marital surplus, a stable assignment must be characterized by *positive assortative matching*. That is, if a man with an endowment y is married to a woman with an endowment z, then the mass of men with endowments above y must exactly equal the mass of women with endowments above z. This implies the following marriage market clearing condition:

$$1 - F(y) = r[1 - G(z)]. \tag{8}$$

As a result, we have the following, spousal matching functions:

$$y = F^{-1}[1 - r(1 - G(z))] \equiv \phi(z) \tag{9}$$

or equivalently:

$$z = G^{-1}\left[1 - \frac{1}{r}(1 - F(y))\right] \equiv \psi(y)\,. \tag{10}$$

For $r > 1$, all men are married and women with incomes below $z_0 = G^{-1}(1 - 1/r)$ remain single. Women with incomes exceeding z_0 are then assigned to men according to $\psi(y)$ which indicates positive assortative matching.

Positive assortative matching has immediate implications for the analysis of divorce. Because divorce is less likely when a couple has higher total income and individuals sort into

[11] See Chiappori, Iyigun and Weiss (2008) for the complete proof.

marriage based on income, individuals with higher income are less likely to divorce.[12]

Stability Conditions The allocations which support a stable assignment must be such that the implied expected lifetime utilities of the partners satisfy

$$U_h(y) + U_w(z) \ \geq S(t) \ ; \qquad\qquad \forall \ y, \ z \ , \tag{11}$$

where $U_h(y)$ and $U_w(z)$ respectively represent the expected lifetime utilities of the husband and the wife over the two periods. For any stable marriage, equation (11) is satisfied as an equality, whereas for a pair that is not married, (11) would be satisfied as an inequality. In particular, we have

$$\begin{aligned} U_h\left(y\right) &= \max_z \left[S\left(t\right) - U_w\left(z\right)\right] \ , \\ U_w(z) &= \max_y \left[S\left(t\right) - U_h\left(y\right)\right] \ . \end{aligned} \tag{12}$$

It is important to note that the stability conditions above constrain the total (two-period) expected utilities U_h and U_w, but have no implication for the *intertemporal distribution* of utility over the two periods.

2.4.2 Determination of Expected Lifetime Utilities

General Characterization Conditions in subsection (2.4.1) lead to an explicit characterization of the intra-household allocations. The envelope theorem applied to these conditions yields the differential equations :

$$U_h'\left(y\right) \ = S'[y + \psi\left(y\right)], \tag{13}$$

and

$$U_w{}'\left(z\right) \ = S'[\phi\left(z\right) + z] \ . \tag{14}$$

[12]Such a result is consistent with empirical findings on marriage and divorce patterns by schooling: individuals sort positively into marriage based on schooling and individuals with more schooling are less likely to divorce. See Browning, Chiappori, Weiss (in progress, ch. 1).

To derive the expected spousal allocations over the two periods and along the assortative marital order, we integrate the expressions in (13) and (14). Hence, surplus share of a *married man* with income y is

$$U_h(y) \ = \ k^h + \int_{y_m}^{y} U_h'(x)\,dx \ , \tag{15}$$

and the surplus share for a *married woman* with income z is

$$U_w(z) = k^w + \int_{z_m}^{z} U_w'(x)\,dx \ , \tag{16}$$

for some constants k^h and k^w which we determine below.

Pinning Down the Constants The constants k^h and k^w are pinned down by two conditions. First, for all married couples, the total output is known as expressed by equations (15) and (16). Hence,

$$k^h \ + \ k^w \ = S[y + \psi(y)] - \int_{y_m}^{y} U_h'(x)\,dx - \int_{z_m}^{\psi(y)} U_w'(x)\,dx \ , \tag{17}$$

where the left-hand side, by construction, does not depend on y. Secondly, it must be the case that 'the last married person' is just indifferent between marriage and singlehood. In the case with more women than men, $r > 1$, we have

$$U_w(z_0) \ = 2z_0 \qquad \Leftrightarrow \qquad k^w = 2z_0 - \int_{z_m}^{z_0} U_w'(x)\,dx \ , \tag{18}$$

with $z_0 \equiv \Phi(1-r)$. Hence,

$$k^h = S[\phi(z_0) + z_0] - 2z_0 \ ,$$

$$U_w(z) = 2z_0 + \int_{z_0}^{z} U_w'(x)\,dx \ , \tag{19}$$

$$U_h(y) \ = S[y + \psi(y)] - U_w[\psi(y)] = S[y + \psi(y)] - \left(2z_0 + \int_{z_0}^{\psi(y)} U_w'(x)\,dx\right).$$

It is important to stress that the stability conditions apply without any assumption about the level of commitment attainable by the spouses or the option of remarriage. The insight is that the conditions on the marriage market determine the allocation of lifetime utilities between spouses: because of competition, a wife would not agree to marry a husband who would provide less than the equilibrium utility — since many perfect substitutes exist — and neither would the husband.

2.5 The Intertemporal Allocations of Utility

2.5.1 The Commitment Issue

We continue our analysis with a consideration of the allocation of lifetime utilities U_h and U_w between the two periods. At this point, commitment issues become crucial. While some degree of commitment is clearly achievable, there may be limits on the extent to which couples are able to commit — after all, couples could not and would not commit not to divorce. Two broad views emerge from the existing literature. Some contributors argue that only short-term commitment is attainable and that long-term decisions are generally open to renegotiation at a further stage. Others authors point out that a set of instruments, including prenuptial agreements, are available to sustain commitment. They, therefore, claim that divorce is the only limitation on commitment. Technically, marriage contracts should be seen as long-term efficient agreements under one constraint — namely that a person who wants to divorce can always choose to do so.[13]

In our framework, these two alternative views about commitment have a natural translation. Specifically, we can entertain two scenarios: In the first case ('commitment'), couples can commit to their spousal allocations in both periods conditional on the continuation of their marriage; the corresponding contingent allocations are ex-ante efficient under the sole constraint that divorce is unilateral. Therefore, the only constraint on intra-temporal allocation of resources is that second-period utility should exceed singles' utility, at least insofar as divorce is not an efficient outcome. Finally, should an unexpected event occur between the two periods, such as a reform of the alimony-payment laws (an example we consider below), this would not trigger a renegotiation of the initial agreement, unless the new individual rationality constraint is violated for one spouse. In the latter case, such a spouse

[13]As in standard contract theory, we assume in all cases that a minimal level of commitment, whereby agents are able to at least commit to *first-period* allocations when they get married, is attainable. See Lundberg and Pollak (1993) for alternative assumptions. Also see Lundberg and Pollak (1993) and Mazzocco (2007) for further discussions of commitment issues within marriage.

would receive an additional share of household resources so that she becomes just indifferent between marriage and singlehood under the new law.[14]

In the alternative, polar case ('no commitment'), serious limits exist on the spouses' ability to commit. In this case, couples may be able to commit to the immediate (i.e. first period) allocation of resources; but future allocations cannot be contracted upon and will therefore be determined by a bargaining mechanism at the beginning of the second period. Of course, this feature is known ex ante by the agents and it influences the decisions regarding first-period allocations. Finally, if a reform occurs between the two periods, the new situation is taken into account during second-period bargaining; i.e., bargaining always take place 'in the shadow of the law'.

2.5.2 Second-period Utilities

For illustrative purposes, consider the case in which couples can commit to their spousal allocations in marriage ex ante. No renegotiation can therefore take place unless divorce is credible. Moreover, if renegotiation does occur, it results in the minimal change needed for a union to continue, if that is indeed optimal.

Let $u_h^2(y)$ and $u_w^2(z)$ denote the *monetary* components of utility derived from the intramarital allocations respectively of husband with endowment y and wife with endowment z in the second period should they continue with their partnership. Hence, the husband's (wife's) total second-period utility is $u_h^2(y) + \theta$ (resp. $u_w^2(z) + \theta$) if the union continues. Feasibility constraints require that

$$u_h^2(y) + u_w^2(z) = \eta(t) . \tag{20}$$

Under unilateral divorce, each spouse can walk away with the share of family income determined by law, βt for the wife and $(1-\beta)t$ for the husband, where $t = (y+z)$ is family income. Individual rationality implies that these outside options cannot exceed the utility payoffs if the marriage continues. Therefore, it must be the case that

$$u_h^2(y) + \theta \geq (1-\beta)\,t \quad \text{and} \quad u_w^2(z) + \theta \geq \beta t, \tag{21}$$

[14] Such contracts are actually (second best) efficient under the constraint that agents cannot commit not to divorce. Similar ideas are used in different contexts, in particular risk sharing agreements under limited commitment. See Ligon et al. (2002) and Kocherlakota and Pistaferri (2008).

which we shall hereafter refer as the *individual rationality constraints* (IR). Note that these conditions jointly imply that

$$u_h^2(y) + \quad u_w^2(z) + 2\theta = \eta(t) + 2\theta \geq t \ , \tag{22}$$

or equivalently that $\theta \geq \hat{\theta}(t)$, so that divorce is not the efficient outcome.

Any allocation such that (21) is satisfied can be implemented as part of a feasible marital contract:[15]

Proposition 1 *With commitment and unilateral divorce, there exists exactly one allocation that is not θ-contingent and guarantees that all the constraints are satisfied for any realization of θ.*

Proof. *The key remark is that the individual rationality constraints (21) must be binding when $\theta = \hat{\theta}(t)$ since, for that value, the couple is indifferent between marriage and divorce. Hence,*

$$u_h^2(y) = (1-\beta)t - \hat{\theta}(t) = \frac{1}{2}(\eta(t) + (1-2\beta)t), \tag{23}$$

$$u_w^2(z) = \beta t - \hat{\theta}(t) = \frac{1}{2}(\eta(t) - (1-2\beta)t). \tag{24}$$

Note that, for any realization of θ, either $\theta < \hat{\theta}(t)$ and divorce takes place or $\theta \geq \hat{\theta}(t)$ and utilities are equal to $(1-\beta)t + \theta - \hat{\theta}(t)$ and $\beta t + \theta - \hat{\theta}(t)$ for the husband and the wife respectively, so that the time-consistency constraints are fulfilled for both spouses. ■

Interestingly, the second-period utilities in union exactly reflect the utilities if separated, with the addition of the difference between the actual match quality θ and the threshold $\hat{\theta}$.[16] In particular, we have

[15] A natural question, however, is whether the material allocation (u_h^2, u_w^2) can be contingent upon the realization of θ. Contingent allocations raise specific problems. For instance, depending on the enforcement mechanism, they may require that the quality of the match be verifiable by a third party. Whether such verifiability is an acceptable assumption is not clear. It turns out, however, that under our assumption of common θ, verifiability is not an issue because there exists (exactly) one allocation allocation that satisfies the incentive compatibility constraints for *all* θ.

[16] If, instead, one entertains the case in which couples cannot make pre-marital allocative commitments, renegotiation would systematically take place at the beginning of the second period. If such couples reach a Nash-bargaining solution, with the utility of the husband and the wife in case of divorce as the relevant threat points, then the allocations will be such that they correspond exactly to the non θ-contingent allocations under commitment. In other words, the unique second-period allocation that is not θ-contingent and

Corollary 2 *Any increase of, say, the wife's utility in divorce is exactly reflected in her second-period utility even if divorce does not take place.*

2.5.3 First-period Utilities

For each choice of k, we can now recover the first-period allocations. The expected two-period utilities equal

$$U_h(y) = u_h^1(y) + \bar{\theta} + (1-\alpha(t))\left\{u_h^2(y) + E\left[\theta \mid \theta \geq \hat{\theta}(t)\right]\right\} + \alpha(t)(1-\beta)t\ , \tag{25}$$

$$U_w(z) = u_w^1(z) + \bar{\theta} + (1-\alpha(t))\left\{u_w^2(z) + E\left[\theta \mid \theta \geq \hat{\theta}(t)\right]\right\} + \alpha(t)\beta t\ , \tag{26}$$

where $\alpha(t) = \Pr\left(\theta < \hat{\theta}\right)$ is the separation (or divorce) probability. These utilities must coincide with the equilibrium values derived above. Therefore, for $r > 1$,

$$\begin{aligned} u_w^1(z) &= z_0 + \int_{z_0}^{z} S'\left[\phi(x) + x\right]dx \\ &\quad - (1-\alpha(t))\left\{u_w^2(z) + E\left[\theta \mid \theta \geq \hat{\theta}(t)\right]\right\} - \alpha(t)\beta t\ , \\ u_h^1(y) &= S[y + \psi(y)] - z_0 - \int_{z_0}^{\psi(y)} S'\left[\phi(x) + x\right]dx \\ &\quad + (1-\alpha(t))\left\{u_h^2(y) + E\left[\theta \mid \theta \geq \hat{\theta}(t)\right]\right\} - \alpha(t)(1-\beta)t\ . \end{aligned} \tag{27}$$

---

guarantees that the individual rationality constraints are satisfied for any realization of θ is also the Nash solution to a second-period bargaining.

Finally, if one were to allow different valuations of the match quality by the husband and wife, then spousal allocations would be contingent on the realization of (θ_h, θ_w) reflecting the fact that, should the marriage continue, a spouse whose evaluation is poor must be compensated by an adequate monetary transfer. Again, this allocation can be implemented in the commitment case as well, although it is now θ-contingent (that is, it requires transfers that depend on the *difference* in the valuations of the husband and wife). For further details as well as the complete proof of the overlap between second-period allocations with and without commitment, see Chiappori, Iyigun and Weiss (2008).

2.6 Reforming Alimony Laws

Consider now a change in alimony payment laws such that the wives' share of household income is increased from β to $\hat{\beta}$. This, of course, does not rule out the possibility that β's may be couple-specific (as it would be in a private-property regime).

As long as utility is transferable, the Becker-Coase theorem applies and such a change does not affect divorce probabilities. In particular, the threshold $\hat{\theta}(t)$ only depends on the surplus generated by a union, not on its post-divorce division between (ex-) spouses; a couple splits if and only if its realized θ lies below the threshold, irrespective of the β in place. But, under unilateral divorce laws, changes in β typically result in a redistribution of the surplus between spouses during marriage. Whether a wife would benefit from the new property division rules would depend on her income, her marriage match quality and the level of commitment achieved between the spouses.

Concerning the impact on the division of marital gains, it is crucial to distinguish between *existing couples*, who are together when the change becomes effective, and *future couples,* who are not. For the former, unexpected legislative changes may trigger a renegotiation within the household and alter the original contract implemented. For the latter, the new legislation would be taken into account at the matching stage and reflected in the expected allocations entering marriage. We now consider these two cases successively.

2.6.1 Existing Couples

Consider a couple with endowments y and z for the husband and wife, respectively, whose match quality θ strictly exceeds the threshold $\hat{\theta}(t)$. Since the intra-household spousal allocations, as determined in the marriage market, were individually rational, it must have been the case that neither spouse had an incentive to get divorced with the original β in place.

Commitment Assume, first, that the spouses feel committed by the contract they initially chose, although they do not feel obligated to remain together. If θ is large enough, the wife's individual rationality requirements given by (21) are satisfied for both β *and* $\hat{\beta}$. This occurs if

$$\theta \geq \hat{\beta}t - u_w^2(z), \tag{28}$$

where $u_w^2(z)$ denotes the continuation utility of the wife under the current agreement. Then, due to the commitment assumption, the change in divorce laws has no impact on intra-household allocations. If, on the contrary, θ is such that

$$\hat{\beta}t - u_w^2(z) > \theta \geq \beta t - u_w^2(z) \ , \tag{29}$$

then the initial agreement is no longer enforceable, since it would violate the wife's individual rationality. Hence, her second-period allocation must be adjusted upward to $\hat{u}_w^2(z) = \hat{\beta}t - \theta$, which requires an additional transfer equal to

$$T = \left(\hat{\beta} - \beta\right) t - \theta - \frac{\eta(t) - t}{2} \geq 0 \ . \tag{30}$$

From a comparative perspective, the probability of a renegotiation taking place depends on the distribution of θ. In the benchmark case where θ is more or less uniform over a 'large enough' support, the probability is proportional to $\left(\hat{\beta} - \beta\right) t$. When both β and $\hat{\beta}$ are identical across couples, the reform affects a larger proportion of higher-income couples. Regarding the size of the transfer, one can readily check that if β and $\hat{\beta}$ are identical across couples, the transfer T given by (30) is concave in total wealth t. It increases in t for small t but if the surplus function $\eta(t)$ is convex enough, it decreases in t when t is large enough. Then, the magnitude of the transfer is non-monotonic in income; it is smaller for the poorest and the highest-income couples and maximal for intermediate income levels. In the special case of a move from private to common property, then $\beta = z/(y+z)$, and the reform, not surprisingly, is more likely to affect those couples for whom the initial distribution of incomes was biased in favor of the husband. The transfer can be written as

$$T = \hat{\beta}t - z - \theta - \frac{\eta(t) - t}{2}. \tag{31}$$

It is still concave in t. Moreover, for any given t, it decreases in z, implying that it is larger for initially unequal couples.

We conclude that the reform will affect intra-household allocations of some — but not all — couples. For couples with a low realized match quality, the second-period marital allocation of the wife may no longer be sustainable. As a result, there will be more recontracting in favor of women among such couples. And since first-period spousal allocations would have already been sunk for all of the existing unions at the time of the legislative change, a more generous settlement rule for the wives would imply higher allocations for them in the second period *and* over their lifetimes.[17]

[17] In the absence of commitment, renegotiation takes place between all spouses. The reform directly impacts

2.6.2 Future Couples

Now consider a couple who is not yet together at the time of change in the alimony laws. The expected lifetime allocations of such a couple, as given by equations (25) and (26), can be decomposed into three parts: the first-period utility, the second-period utility if the union is continued, and the second-period utility in case of legal separation. Unlike existing unions, however, this effect is fully anticipated by the agents in the matching phase and reflected in the equilibrium allocations. This has two consequences. First, the reform influences intra-household allocation in *both* periods. This is because the allocation of *lifetime* utility, which involves first- and second-period welfare, is decided during the matching process, taking into account the new law. A second and more subtle implication is that the impact of the reform on a future union is the same whether or not agents are able to commit to specific intra-household allocations ex ante. Indeed, we have seen in subsection 2.4.2 that the (non-θ-contingent) allocation decided ex ante is the same in both contexts.

Using (23) and (24), we can compute the impact of a change in post-divorce allocations on individual utilities. If β is identical across couples before and after the reform — one may think of a redefinition of 'equitable distribution' in a sense more favorable to women — the variations in individual utilities are given by:

$$\Delta u_h^1 = \Delta u_w^2 = \left(\hat{\beta} - \beta\right) t, \qquad \Delta u_h^2 = \Delta u_w^1 = -\left(\hat{\beta} - \beta\right) t,$$

while if the switch is from private to common property, then,

$$\Delta u_h^1 = \Delta u_w^2 = \hat{\beta} t - z, \qquad \Delta u_h^2 = \Delta u_w^1 = -\left(\hat{\beta} t - z\right).$$

In both cases, a divorce law that mandates more generous divorce settlements for women increases their utility in the second period whether or not the couple separates. However, the reform also *lowers* their first-period allocations by the same amount. Implicit in the above argument is what we have already established in (19): in unions not yet formed, a

the respective threat points. Therefore, it affects all couples. Assuming, as above, a Nash bargaining solution with utilities in case of divorce as threat points, we see that the wife's gain from the reform is

$$\hat{v}_w^2(z) - v_w^2(z) = \left(\hat{\beta} - \beta\right) t\ , \tag{32}$$

while the husband loses the same amount.

We conclude that when a reform of divorce laws is favorable to women and there is no commitment to ex-ante spousal allocations between spouses, all wives will benefit and all husbands will lose. This exemplifies the case of 'bargaining in the shadow of the law'.

legislative change has no effect on the expected *lifetime* allocations of each spouse, $U_h(y)$ and $U_w(z)$. But given that equilibrium spousal allocations need to be individually rational, more favorable divorce rules may lead to a more rapidly rising allocation path for the wives-to-be in order to ensure that their marital commitments are time consistent; in practice, they get more at the end, therefore less at the beginning of the union. In particular, all wives' expected intra-marital allocations *conditional on remaining married* are reduced and the reduction exactly offsets their gain in case of divorce.

We conclude with the following general proposition:

Proposition 3 *A change in the rules governing property rights over the distribution of family assets has* no *impact on welfare as measured by expected lifetime utilities at the time of marriage. To the extent that the policy raises the utility of women following divorce, it must* reduce *their total utility while married.*

These neutrality results are related to the literature on Ricardian equivalence (see Barro, 1974) in that an attempt by the government to redistribute income among agents is completely undone by a redistribution over time *within* family units. The neutrality of mandated divorce settlements is also similar to Lazear's (1990) result on the neutrality of mandated severance payments in the context of worker-firm relationships. In both cases, an attempt by the government to redistribute income among agents is completely undone by a redistribution over time *within* families or firms and does not affect the competitive outcome.[18] Indeed, this point is made by Lundberg and Pollak (1993) regarding child allowances.

While extremely stylized, this framework is meant as an illustrative example of how an estimate of the impact of granting alimony rights to couples may be extremely different for couples already in a union at the time of the legislative change than for couples who unite after the law is enacted. In what lies ahead, we explore whether there is any empirical evidence of this kind of policy neutrality in the context of cohabiting partners in Canada.

3 Alimony Rights of Cohabiting Partners in Canada

The rights of cohabiting individuals in the case of separation in Canada have changed dramatically over the last 35 years, mirroring the pattern of other nations. However, what

[18]Note, however, that our result relies on market forces rather than altruism to endogenize redistribution between spouses.

sets the Canadian provinces apart from most European nations is that no "registration" of unions is required. Cohabitation, in itself, is the sign required by law for demonstrating one's commitment to the relationship.

Family law in Canada is mostly governed by provincial authorities who are allowed to legislate these matters independently, as long as the rights guaranteed in the constitution are respected. Table 1 presents a summary of the legislative changes studied in this paper. These laws only granted spouses the right to petition for alimony upon separation. They did not grant rights to an equal division of assets, which is still, in most jurisdictions, granted to married individuals.[19]

The legislative shifts analyzed here occurred between 1978 (in the province of Ontario) and 2000 (in Newfoundland). There appears to be no general trend for provinces close to one another to have coordinated their legislation. It also does not appear that provinces that were more liberal or with a higher proportion of common-law spouses adopted these legislations earlier than others. Actually, the province with the most common-law relationships (Quebec) is the only province that has continued not to offer any protection to partners in that form of unions.[20] Furthermore, most of the shifts were brought upon by cases in provincial and national courts. The majority of recent changes in legislation was due to cases involving homosexual couples rather than heterosexual couples, who were then granted these benefits on the grounds of equality. This should reduce the potential for the passing of these laws endogenously according to changes in cohabiting couples' behavior.

What provides yet more source of variation for identification is that, as shown in Table 1, each geographical entity differed greatly on how they defined a common-law relationship. The duration of cohabitation differs between provinces and ranges from one year in Nova Scotia to five years in Manitoba. Also, six provinces reduced the requirement in terms of cohabitation length for couples with children.

How are these laws enforced? It appears that it is left to the petitioner to prove that the relationship lasted the required amount of time. Evidence such as common leases, common bank accounts are useful in this matter. However, since 1993, this is facilitated by the existence of tax records that specify the length of the relationship. In that year, common-

[19] Asset division was granted to common-law spouses in 2001 in Saskatchewan. We have excluded from our analysis the territories where, in general, spousal benefits and asset division rights were granted simultaneously.

[20] In July 2009, the Quebec Superior Court ruled that such a law was constitutional and allowed Quebec to continue with this policy in a case involving a famous tycoon and his cohabiting partner of many years who was seeking monthly alimony payments of CAN$56,000 in addition to a payout of CAN$50 million. In December 2009, the case was sent to the Supreme Court.

law spouses were granted the same tax-related privileges and obligations as married couples. Actually, as of 1993, common-law partners having lived together for more than 12 months (or less but with a child) *must* file their income tax jointly. This shift affected all couples in all provinces at the same time.

One also needs to mention that in all provinces, cohabitation agreements are legal and could be signed upon entry into cohabitation and stipulate the financial exchanges that would be accepted if separation ever occurs. These were not invalidated by the change in the law. However, the courts have a record of refusing to enforce agreements that are judged to be "unfair". Furthermore, such cohabitation agreements are actually rarely signed by partners.

Finally, cohabitation in Canada is not uncommon and rising in popularity. According to Statistics Canada (2001), 14 percent of all families were living in common-law partnerships (compared to 70 percent of married and 16 percent of single parents). Among all couples, 16 percent of all couples were cohabiting. This is driven by the very large number of common-law unions in Quebec (where 30 percent of all unions are cohabitations) but the proportion of common-law relationships in the rest of Canada in 2001 (11.7) is still larger than that in the United States (8.2). Common-law relationships differ observationally from legal unions in many ways: they are shorter-lived and have lower fertility rates. Cohabitation is also correlated with being younger, French-speaking, Catholic or no religious beliefs. Individuals who are cohabiting are also less likely to have attended religious services as kids and they are slightly more educated. For further details, see Statistics Canada (2001).

4 Estimation Framework and Data

The structure of the law discussed above seems to suggest the use of a Difference-in-Difference-in-Differences (DDD) estimator since whether a relationship was subject to the law depended on 3 distinct components: the year in which the relationship started, the duration of the relationship and the province where the relationship was occurring. The special rules for parents also imply a 3-component variation: the year when the relationship started, when the child was born and in which province determined whether the relationship was under the influence of the new law. This section describes in detail the empirical strategy pursued and the data used in order to operationalize this identification strategy.

4.1 The Estimation Equation

Assume that one wants to use the legislative framework presented above to estimate the impact of granting cohabiting partners the right to petition for alimony on an outcome y_{iptdc} for an individual i, in province p, whose relationship began in year t and lasted for at least d years and whose first child in the relationship was born in year c (let $c = 0$ when no child was born). For each province, define the year in which the new law is implemented as T_p, the required duration as D_p and the duration required when a child is present as C_p.

Let us define two variables indicating whether the household would be eligible to petition for alimony in case of separation. Define the variable A_{ptdc} as an indicator equal to 1 if the relationship was, at any point in time, subject to the new rules regarding alimony. Define the variable A^*_{ptdc} if the relationship became eligible as the new law entered into place but was formed before the rules were changed.

Formally, we can then define these indicator functions as:

$$A_{ptdc} = \min\left[\sum_{k=1}^{10} 1(p=k) * 1(t+d > T_k) * (1(d > D_k) + 1(d > C_d) * 1(c > 0)), 1\right] \tag{33}$$

$$A^*_{pdtc} \quad = \quad \min\left[\sum_{k=1}^{10} 1(p=k) * 1(t+d > T_k) * \Big(1(\tilde{d}_k > D_k) + 1(\tilde{d}_k > C_k) * 1(c < T_k)\Big), 1\right] \tag{34}$$

where 1(.) represents the indicator function and $\tilde{d}_k = T_k - t$. In each case, a couple is eligible if the relationship is still active at the time of the legal change and if it lasted (or has lasted at the moment of the change in the case of equation (34)) more than the required amount of time. The second term in the summation refers to the special rules linked to the presence of a child and, given that a couple could qualify through both the duration test and the presence of a child, the minimum function ensures that our indicator is never larger than one. Note, however, that since all provinces that have special rules for cohabitations involving children also impose a relationship duration of at most one year, $1(d > C_d)$ and $1(\tilde{d}_k > C_k)$ will simply equal one when the province has such a provision and zero otherwise.

It is obvious that in both equations and for both criteria of eligibility, the indicator function is equal to one when three distinct characteristics are satisfied simultaneously. This intuitively suggest the following identification assumption: controlling for the interactions between each pair of these characteristics should allow us to properly isolate the effect of the new rule. Intuitively, this strategy corresponds to the following argument: Compare two couples that were formed in the same year; one whose duration was long enough to qualify and one who wasn't. Obviously, the difference in the outcome of these two couples would depend on a large number of factors that would obscure the impact of being eligible for the right to petition alimony. We could then compare the difference in the outcomes of these couples to that of two couples with the same difference in duration, but who were formed before the laws were passed and thus were both ineligible. While this would ensure that the estimate is not contaminated by the effect of duration, there is still the possibility that a general time trend in the outcomes would bias our estimate. To remove this possibility, the double-difference described above is repeated using couples in a province not yet affected by a change in the law and then the two are differenced from each other. This difference should purge out any time trends common to all provinces and lead us to an estimate of the effect of the legislation.

Formally, provided that we ignore the issue of endogenous relationship formation, this triple difference strategy translates into a regression equation which is given by:

$$\begin{aligned} y_{ipdtc} &= \alpha A_{pdtc} + \beta X_i + \sum_{k=1}^{10} \delta_k 1(p=k) * 1(t+d > T_k) + \sum_{k=1}^{10} \lambda_k 1(p=k) * 1(d > D_k) \\ &+ \sum_{k=1}^{10} \phi_k 1(t+d > T_k) * 1(d > D_k) + \sum_{k=1}^{10} \theta_k 1(p=k) * 1(d > C_d) * 1(c > 0) \\ &+ \sum_{k=1}^{10} \tau_k 1(t+d > T_k) * 1(d > C_d) * 1(c > 0) + \sum_{k=1}^{10} \theta_k 1(t+d > T_k) \\ &+ \sum_{k=1}^{10} \rho_k 1(c > 0) + \sum_{k=1}^{10} \gamma_k 1(d > D_k) + \nu_p + \mu_t + \varepsilon_{ipdtc} \end{aligned} \tag{35}$$

We will estimate this equation to measure the impact of granting alimony rights to new couples by restricting the sample to new couples and those couples formed before the law that never became eligible.

If one worries that endogenous relationship formation may be at play, the solution is to first restrict the sample to couples whose relationship began before they could modify their behavior to avoid becoming subject to the new ruling (that is where $t < T_p - D_p$ for each province) and, second, to use the following equation instead:

$$
\begin{aligned}
y_{ipdtc} &= \alpha A^{*}_{pdtc} + \beta X_i + \sum_{k=1}^{10} \phi_k 1(t + \tilde{d}_k > T_k) * 1(\tilde{d}_k > D_k) \\
&+ \sum_{k=1}^{10} \lambda_k 1(p = k) * 1(\tilde{d}_k > D_k) + \sum_{k=1}^{10} \theta_k 1(p = k) * 1(\tilde{d}_k > C_k) * 1(c < T_k) \\
&+ \sum_{k=1}^{10} \tau_k 1(t + d > T_k) * 1(\tilde{d}_k > C_k) * 1(c < T_k) + \sum_{k=1}^{10} \theta_k 1(t + \tilde{d}_k > T_k) \\
&+ \sum_{k=1}^{10} \rho_k 1(c < T_k) + \sum_{k=1}^{10} \gamma_k 1(\tilde{d}_k > D_k) + \nu_p + \mu_t + \varepsilon_{ipdtc}
\end{aligned}
\tag{36}
$$

In the two equations above, α represents the causal estimate of granting cohabiting couples the right to petition for alimony. To control for other individual-specific characteristics that could influence y, a set of individual controls X_i will be added (including gender, age, religion and education). In order to allow for serial correlation, standard errors are clustered at the province level. Furthermore, in order to alleviate the problem of selective migration by which individuals may be moving to a particular province in response to the changes in the legal framework, we shall assume that individuals were subject to the legislative changes in their province of birth. And while a more complete set of interactions may have been ideal, the sample size at hand for this study was too small to allow for more detailed interactions between province, year of formation and duration.

The required identifying assumption here is that there was no other contemporaneous shock affecting cohabiting couples who were living together for more than a certain period in provinces where the legislation was changed. This is robust to shocks occurring in a province at a given time or to couples with longer durations in a particular province being different than those in another one. Furthermore, as we specified above, in most of the recent legislative changes, the impetus for modifying the law was not a desire to modify the legal rights of cohabiting partners but more of a need to offer homosexual couples (who, at that time, were not allowed to marry legally) the same type of legal protection married heterosexual couples were afforded in case of separation.

In many instances, the data presented below allow further analyses because outcomes are observed in every single year of the relationship, not only once per relationship. In this case, one can modify equations (35) and (36) to allow for relationship-specific fixed effects and simply compare the behavior of individuals before and after their relationship becomes subject to the new rules. In this case, the estimation equations for an outcome y_{ipd_jtcj} for relationship i, in province p, which began in year t and lasted d_j years in year j, had a child in year c and where the outcome is observed for year j become:

$$\begin{aligned}
y_{ipdtcj} &= \alpha A_{pd_jtc_j} + \beta X_{ij} + \sum_{k=1}^{10} \delta_k 1(p=k) * 1(j > T_k) + \sum_{k=1}^{10} \lambda_k 1(p=k) * 1(d_j > D_k) \\
&+ \sum_{k=1}^{10} \phi_k 1(j > T_k) * 1(d_j > D_k) + \sum_{k=1}^{10} \theta_k 1(p=k) * 1(d_j > C_d) * 1(c_j > 0) \\
&+ \sum_{k=1}^{10} \tau_k 1(j > T_k) * 1(d_j > C_d) * 1(c_j > 0) + \sum_{k=1}^{10} \theta_k 1(j > T_k) \\
&+ \sum_{k=1}^{10} \rho_k 1(c_j > 0) + \sum_{k=1}^{10} \gamma_k 1(d_j > D_k) + \mu_j + \nu_i + \varepsilon_{ipdtcj}
\end{aligned} \tag{37}$$

and

$$\begin{aligned}
y_{ipdtcj} &= \alpha A^*_{pd_jtc_j} + \beta X_{ij} + \sum_{k=1}^{10} \lambda_k 1(p=k) * (1(\hat{d}_j > D_k) + \sum_{k=1}^{10} \phi_k 1(j > T_k) * 1(\hat{d}_j > D_k) \\
&+ \sum_{k=1}^{10} \theta_k 1(p=k) * 1(T_k - t > C_k) * 1(c_j < T_k) + \sum_{k=1}^{10} \theta_k 1(t + \tilde{d}_k > T_k) \\
&+ \sum_{k=1}^{10} \tau_k 1(j > T_k) * 1(T_k - t > C_k) * 1(c_j < T_k) \\
&+ \sum_{k=1}^{10} \rho_k 1(c_j < T_k) + \sum_{k=1}^{10} \gamma_k 1(\hat{d}_j > D_k) + \mu_j + \nu_i + \varepsilon_{ipdtcj}
\end{aligned} \tag{38}$$

where $c_j = 1(c < j) * c$ and $\hat{d}_j = 1(j < T_k) * d_j + 1(j > T_k) * \tilde{d}_k$. In this case, the only year-specific controls included in the regression are the age and the square of the age of the individual in year j. Standard errors are once more clustered by province.

Finally, one can also only restrict the sample only to new relationships and use a difference-in-difference model using the province and the duration of the relationship as the only variables determining treatment. In this case, the identifying assumption would be that there is no other province-specific shocks that affect relationships lasting more than a given number of years. Those regressions will be estimated in a similar way as above and standard errors will also be clustered by province.

4.2 The Data

The empirical framework presented above implies that, in order to classify whether or not a relationship was subject to alimony payments, information about the exact duration of the relationship is required. On this basis, we used the *General Social Survey* (GSS) data of 2001, from *Statistics Canada*. This is a survey that was performed once, but it asked very detailed retrospective questions on one's past relationships, fertility, education and labor market activities. Particularly useful for this study, the GSS compiled information on up to four past marriages and up to six common-law relationships (up to three that eventually transformed into marriages and up to three that did not). The total sample was 24,310 Canadians aged 15 and above in all provinces but not in the territories. This dataset is used to construct relationship-specific information regarding the year it began, its duration, the way the relationship ended as well as labor, schooling and fertility decisions of the respondent during the relationship. This survey does have a few less attractive features. First, it is retrospective and thus subject to recall bias. Nevertheless, as long as this bias is not altered by the new alimony rules, it should not affect our results. Furthermore, the labor market data are quite coarse as they include neither hours nor weeks worked.

The data collected from the GSS measure all the characteristics needed to classify a union as subject to the new alimony rules or not: the age at which the relationship began (which, taken together with the year of birth, identifies the year in which the relationship began), the age at which it ended (or whether it was still active at the moment of the survey) and the province of birth. The variables A and A^* can thus be computed for each relationship, as well as all the ingredients required as control variables in equations (35) through (38). In addition, self-reported demographic controls are also available for inclusion in X_i as additional regressors.

The data provide a variety of outcomes to measure the impact of the legislation on the behavior of spouses. First, one can determine whether the cohabitation eventually led to marriage (which can be seen as a substitute for cohabitation). The data also allow us to measure the overall duration of the relationship (including the years of marriage when relevant) and whether or not the relationship had ended at the time of the 2001 survey.

The key outcomes of interest for this paper, however, relate to the measures of relative welfare of each partner. As is common in this literature, we use labor supply as a proxy for the consumption of leisure and thus higher labor supply will be assumed to imply lower welfare. No information regarding the number of hours or weeks worked is provided in the GSS. Instead, the respondents were asked to detail the full retrospective history of their work and education lives. All work episodes are described including the year they began, the year they ended and whether they involved mostly full-time or part-time work. All interruptions, which are defined as periods of more than 3 months where the individual was not working for a variety of reasons, including lack of work, sickness, maternity/paternity leaves, retirement, job switches, etc, are also mentioned as well as any educational experience. Finally, any maternity or paternity leave is also compiled. The periods of work, hiatus and education are then matched to each relationship based on the years of each event.

Finally, the survey only provides limited information on the spouses for each of the relationships detailed by the respondents. The data available include the age difference between the respondent and his/her partner as well as the marital status of the partner before the relationship. Only for the current partner is more information gathered but there is a severe selection bias which prevents us from using these data. Also, most of the first generation of laws only applied to couples of the opposite sex. The data do not allow us to identify whether previous partners were of the opposite sex so all relationships in the dataset are treated as being heterosexual and thus subject to the law. Information on current partners, in any case, indicates that less than one percent of current cohabitation relationships are homosexual and thus the measurement error induced by this assumption appears to be minimal.

We do not use the full GSS sample. Immigrants and individuals born in the territories are excluded since their province of birth is outside the sample. Relationships that began before 1960 are also excluded so as to focus on relationships that are closer to the legislative changes observed. This gives a sample of 7,520 common-law relationships and 11,279 marriages. When estimating equations (37) and (38), one might be worried that longer relationships will be weighted more heavily than shorter relationships since the panel regressions include one observation per year in which the regression was active. This is particularly worrisome if the new legislation impacted the length of the relationship, thus implying that the selection into the sample is endogenous to the legislation. In order to alleviate this concern, the basic sample will include 10 years of observations for all relationships (that is, irrespective of whether the relationship lasted for more or less than 10 years). In this case, the sample will include 60,515 year-relationship observations on common-law relationships and 101,387 on marriages.

Summary statistics of the main variables of interest are presented in Table 2 for both marital and common-law relationships. In this table, every relationship is one observation and the summary statistics are computed using the person-specific weights provided by the survey. The first section documents the demographic characteristics of the respondents. In both marriages and cohabitations, relationships begin on average around the age of 26 and 48 percent of the sample have a male respondent. Respondents in common-law relationships differ from those in legal unions in many ways: they are more likely to be French-speaking; Catholics or atheists; less likely to have attended religious services as teens; more likely to have a high school diploma but less likely to have a college degree. About 37 percent of the common-law relationships in the sample eventually led to marriages. Cohabitations are, nevertheless, more likely to have ended by the time of the survey (despite the fact that this measure excludes cases in which cohabitation evolved into a legal union that is still alive). And they are much shorter-lived (16 versus 8 years) even after we take stock of the entire relationship (cohabitation plus eventual marriage). Cohabitations and marriages also differ in the characteristics of the partners they involve. While 86 percent of the respondents' partners were unmarried before their marriage, only 74 percent of them were so before cohabitation. Partners were also more likely to have larger differences in age among cohabitations than marriages. Finally, during the length of a common-law relationship, 22 percent of the respondents studied, 75 percent worked full-time at some point, 18 percent were on hiatus, and 8 percent took a parental leave (mostly maternity). Both full-time work and work interruptions were more common among marriages.

Table 3 presents the summary statistics for our panel data where each observation is a year-relationship. The first panel includes all years where the relationship was active, and the second imposes the restriction that 10 years after the relationship began must be included in the sample, whether or not the relationship was active at that point. Relationships that last more than 10 years are censored. A pattern similar to the one in Table 2 arises for the labor market variables although, here, there is less difference between cohabitations and marriages. In 12 percent of the years in the sample of cohabitations, the respondent was studying. In 75 percent of those years, he or she was working and in about 65 percent of the cases, on a full-time basis. Work interruptions and family leaves were more common in marriages. The difference between the two panel samples is mostly in the percentage of observations from unions formed before the legislative changes and the incidence of years when the couple was subject to alimony rules: While 62 percent of all year-relationship observations in the full sample were from cohabitations formed before the law had changed, that number falls to 38 percent when the relationships are constrained to include 10 years of data after the beginning of the relationship. Given the change in sample size, this implies that, when the sample is restricted to couples formed before the new rules were implemented, the sample size will not change tremendously. Obviously, this is not true of marriages.

5 The Impact of Granting Alimony Rights

5.1 Alimony Rights and Labor Supply

Alimony payments are usually made from the higher earning partner to the lower earning one. Since over the period in question, men were still more likely to earn higher incomes than their spouses, we will assume that legally requiring alimony payments favored females. Thus, we should expect that, when a relationship becomes eligible for these rules, female partners may decrease their labor supply provided that leisure is a normal good. Every year in the sample, we have information on whether an individual stopped working, went to school, worked full or part time, or used parental leave. This section explores changes in spousal labor supply using these various outcome measures as proxies for labor supply.

For our main results, we utilize the sample which includes at least ten years following the beginning of the relationship, irrespective of whether this relationship was still active. Relationships that last longer than 10 years are included until they end. The results of regression equations (38) are presented in Panel A of Table 4 while those of (37) are found in Panel B. Panel C shows the results of a simple difference-in-difference regression focusing only on couples formed after the legislation had been passed. The results imply that when a relationship is granted the right to petition for alimony, women are about 4.7 percent less likely to work full time and 2 percent more likely to be studying. They are also 5.3 percent more likely to have stopped working. Their likelihood of having taken maternity leave increases by 2.4 percent (although not statistically so) even if the probability of having a child is unchanged.[21] Males, on the other hand, responded to the legislation in a statistically different manner. They appear to have reduced their likelihood of being in school and having suffered work interruptions but more likely to have worked, particularly full-time. They appear to have been less likely to take paternity leave but this result is very small in magnitude and not significant.

What is much more striking is the difference between Panel A and the following two panels of the table. Once one measures the impact of a relationship being subject to alimony rules,

[21]Results about fertility are not presented but available upon request.

but only looking at its impact on relationships formed after the legal change to be part of the sample, the conclusions are very much different. This is true both in Panel B and C, despite the difference in the estimation strategy in these two samples. The coefficients are usually smaller in magnitude for couples formed after the legal change and of the opposite signs as the ones presented in Panel A. These results appear to suggest that relationships formed after the "rules of the game" were changed responded very differently to being subject to the alimony rules than those that were formed before such a legal change, and in a way that is perfectly consistent with the theoretical framework presented above.

These findings are explored in more details in the subsequent table where various robustness checks are performed. We, in particular, investigate the robustness of three outcomes: whether the individual studied in a given year; whether the individual worked full time; and whether the respondent had work interruptions. In the first column, the results are presented assuming that no exceptions are made for relationships with children and that the alimony rights are granted solely on the basis of the duration of cohabitation. The estimates are fairly consistent with those presented in the previous table. The next column uses all years when the relationship was active and thus excludes years after a short relationship terminated. Doing so does alter the results but more in significance than in magnitudes. This highlights the importance of having a comparison point for relationships that lasted longer which is impossible if we use only the active relationships in the sample. The next two columns compare the results restricting the sample to either older relationships, in column (3), or more recent ones, in column (4). Overall, the measured impact appears to be larger and more significant when focusing on recent legal changes than older ones even if some significant impacts are measured for working full-time in the older sample. The next column excludes all relationships from Quebec which shrinks the sample by almost half. The results on the probability of studying or suffering a work interruption are unchanged but the results on labor supply are much larger and more significant than previously. Column (6) repeats the exercise, but this time using married individuals as a placebo group. The coefficients in this case are much smaller in magnitude and rarely significant. When they are (as for the case of work interruptions) they are of the opposite sign for cohabitations. Thus, this suggests that the results obtained in the previous table are not driven by events contemporaneous to the legislative changes affecting all types of unions in a geographical location. The last column includes controls for forthcoming legislative changes. In all outcomes, the fact that an individual would become subject to alimony rights in 2 years has no significant effect on his or her contemporary labor supply. Furthermore, except in the case of work interruptions, the introduction of such an additional control does little to change the size and significance of the coefficients of interest.

Table 6 then explores who is more likely to respond to these new rules concerning alimony. The same three outcomes as in the previous table are presented and results are fairly similar for other outcomes The odd columns present the coefficients of the main effect of the law and the even ones, the interaction term between the legislative change and a dummy for the respondent being a male. The theoretical framework suggests that an individual with a higher income would be more likely to respond to the policy change if there exists some capacity to commit to allocations. While the data do not offer information regarding the income of the respondent at the time of the relationship, we use two proxies: the education level and the age at which the relationship began. The first panel contrasts the treatment effect by the education level of the respondent. In the case of whether the respondent worked full-time in a given year, the legislative change appears to have affected more directly individuals with higher levels of education as predicted by the model. While men with more education responded more strongly to the legislation in terms of work interruptions and studies (although not significantly so), it is women with lower levels of education who did so for work interruptions and women with a high school degree who did so for studying. The age at which one began the relationship also shows some evidence supporting the hypothesis of the model but not very strongly so. It does appear that couples formed before age 21 responded less in terms of labor supply but more in terms of schooling. There is no strong difference between individuals who were in their 20s at the beginning of the relationship and those who were older.

So far, we assumed that women are the lower earning spouses and thus would be the ones benefitting automatically from alimony payments. However, there are couples for whom this was not the case. While we have no information in the data on the relative income of partners, we can use a crude proxy given by their age difference. Those results are presented in the bottom panel of Table 6. The results are strongly in agreement with the hypothesis that the older spouse would be the one responsible for making alimony payments. The results we have found earlier appear to be concentrated among couples for whom the women was at least 5 years younger, thus making them more likely to obtain alimony in case of separation. On the other hand, results are of the opposite signs (although rarely significant except for the probability of studying) when women were much older than their partners, which is again consistent with the hypothesis that the older partner was more likely to be the one making the alimony payments.

Although the model justifies clearly the results obtained in Table 4, it also suggests that women who enter cohabitation relationships after the legal changes would need to compensate their partners for the obtention of these new rights, in particular during the earlier part of a relationship. This is explored in Table 7 where the sample is the same as that in Panel B of Table 4, that is all year-couple observations of a relationship (including at least 10 years following the beginning of a relationship if the relationship does not last long) but where relationships that were "caught" by the change of law are truncated at that point. The first panel simply compares the labor force participation of women and men constrasting relationships formed *before* and *after* the legal change. While the statistical significance of the result is limited, the pattern suggested is clearly in harmony with what our model would suggest: women who enter common-law relationships *after* the legal change are more likely to work and less likely to study or have work interruptions. The difference in the probability of having a work interruption is of about 7 percent. Panel B explores this pattern more carefully by interacting whether the relationship started before or after the passing of the law with a linear indicator for the number of years since the relationship began. As years go by, women in the new regime should be able to benefit from alimony payments, either because the relationship ends at some point or because it stays active and she receives the payoff she was anticipating within the relationship, which is higher under the new law due to its affect on the intertemporal path of allocations (as in Corollary 2). This is exactly what the results of the bottom panel indicate. We now see that, at the beginning of a relationship, men were more likely to study and less likely to work after the new rules were implemented. Women were less likely to suffer work interruptions. However, with the passage of time, the pattern is reversed as men's labor force participation increases and that of women decreases. The results are particularly marked for work interruptions where the probability of experiencing a work interruption would eventually favor women 11 years after the beginning of the relationship.

5.2 Alimony Rights and Relationship Stability

We have shown in our theoretical section that, when utility is transferable, the 'Becker-Coase' theorem applies resulting in no change in divorce likelihoods following changes in alimony laws. However, this might not be the case when utility is not fully transferable. Hence, one may wonder whether granting alimony rights to relationships change the duration and the stability of those relationships. This closely relates to the debate surrounding the impact of no-fault divorce laws on the incidence of divorce in the United States (see, for example, Wolfers, 2006). The threat of separation may be sufficient to alter the way the bargaining power is at play within the relationship but this threat may need to be exercised in some cases when there are frictions in making compensating spousal transfers within the union.

We consider four related outcomes here. First, many cohabitation relationships eventually lead to a formal marriage. However, if cohabitation becomes more similar to traditional legal unions in the eyes of the law, there are less incentives for such a transformation. One may thus expect that granting alimony rights to cohabiting couples may lead fewer of them to "tie the knot". Granting alimony rights also changes the "outside option" of each partner, making it more costly for one and less costly for the other. The overall impact of these rules on the stability of the union is, thus, unclear. We use whether the relationship was still alive at the time of the survey and the length of the cohabitation as our measures of stability. Since there is a potential trade-off with marriage, the total length of the relationship (including any subsequent marriage when relevant) is also included as an outcome.

Table 8 presents the results for these outcomes. The top panel includes the results from the estimation of equation (36), while Panel B presents those from the estimation of equation (35) and Panel C presents the results of a simple difference-in-difference among unions formed after the legal change. The even columns represent the results when males and females are pooled, and the odd columns, one where all control variables as well as the treatment variable are interacted with a dummy for the respondent being a male. The results suggest that the Becker-Coase theorem cannot be rejected in this sample: the likelihood of a couple being separated at the time of the survey in 2002 does not seem to be significantly related to the fact that the union became subject to alimony rights, at least in the case of existing couples. Similarly, receiving alimony rights does not appear to have changed the overall duration of the relationship for couples formed before the legal change. However, it does appear to have significantly increased the duration of the cohabitation fraction of the relationship by about 2 years and reduced the likelihood of transforming the relationship into a marriage by about 14 percent. There is little indication that these impacts differed between male and female respondents. On the other hand, for unions that formed after the laws were changed, being eligible for alimony payments appears to have shortened the duration of cohabitation (and maybe even of the overall duration of the relationship, depending on the estimation strategy used). For none of the outcomes, is the impact of being granted alimony rights among all relationships similar in size to the estimated impact in Panel A.

Table 9 then evaluates how robust these results are to some variations. Six different tests are performed for three different outcomes, each of them presented in a different panel. First, only rules regarding whether the relationship lasted the required number of years are used, ignoring the exceptions linked to the presence of children. The results are similar in magnitude in significance to the ones presented in Table 8. Column (2) then adds a number of covariates including the mother tongue, religious background and educational attainment of the respondent. The results, once more, are unaltered by this modification. Columns (3) and (4) then explore whether restricting the sample to relationships formed early or late influences the estimates. This distinction appears to have an impact on the conclusions we reach. When focusing on relationships that began before 1990, one finds that granting alimony rights not only lengthened the cohabitation but increased its stability, making it less likely to be over by 2001 and increasing its overall duration by about 1.7 years — a little less than half of the increase in cohabitation length. On the other hand, relationships formed after 1980 display a similar pattern as the one highlighted for the entire sample. This could be simply due to the fact that recent relationships are more likely to be censored and, thus, less likely to significantly demonstrate an impact on long-lasting measures such as relationship stability. Given that a large fraction of cohabitations in the sample include respondents born in Quebec, the only province which has not granted alimony rights, column (5) repeats the estimation but this time excluding all relationships from Quebec. In this sample, the Becker-Coase theorem appears to be violated again, as relationships that became subject to the alimony rights were 13 percent less likely to have ceased to exist at the time of the survey and would have lasted about 2.6 more years in total. Cohabitation is increased more significantly in this sample but the likelihood of it evolving into a marriage is not changed. Finally, the last column of the table introduces a lagged indicator which is equal to one if the couple was to become subject to alimony rules in 2 years. Comfortingly, the coefficient of that dummy variable is small and insignificant (except in the case of the overall duration), suggesting that the results presented in the previous table were not simply capturing a time trend.

Table 10 then explores whether these laws differentially affected individuals with distinct attributes. Each column represents a different outcome of interest as in Table 8. Each entry in the table corresponds to the treatment effect interacted with a given characteristic. All regressions are run solely on couples that were "caught" by the legislative change as in Panel A of Table 8. As in Table 6, the first panel explores whether couples with higher incomes, as proxied by age and education, reacted more or less to the policy change. The likelihood that a relationship was inactive by 2001 appears to be influenced by the alimony rights more strongly for individuals with less education, those who started the relationship at an older age and those foe whom the women was much older than her partner. In no subgroup is the total duration of the relationship significantly impacted by the legal change. However, the two variables that were significantly modified by the granting of alimony rights for the entire sample, the likelihood of eventual marriage and the duration of the cohabitation, do appear to be responding more strongly in those groups where the labor supply responses were more visible — that is, individuals with more education, older couples and those where the the man was much older than his partner.

6 Conclusions

We advanced in this paper a collective intra-household allocation model that incorporates both the process of spousal matching and the prospect of divorce. Doing so, we found elements of policy neutrality, as we identified that changes in alimony laws would affect existing couples and couples-to-be differently. For existing couples, legislative changes that favor women benefit them especially if the marriage match quality is low. For couples not yet married, however, they generate offsetting intra-household transfers and lower intra-marital allocations for women.

We then provided some empirical evidence from Canada by using differential rules in distinct Canadian provinces and the fact that those were phased out at different moments over the last 35 years. We found that being offered alimony rights led females to increase their likelihood of attending school and experiencing work interruptions and decreasing their likelihood of working full-time. Males, on the other hand, experienced the opposite pattern once they became subject to the new legislation. Moreover, being subject to the alimony rules decreased the probability that these unions lead to marriages, lengthen the duration of the cohabitation and, in some samples, also led to fewer separations.

These results held within a given relationship over time, but they did not apply to individuals who were *married*, as those already benefited from these rights and thus were unaffected by the new laws. All the more important, we also identified contrasting outcomes for the new alimony rights' impact on the behavior of cohabiting couples who got together *after* the alimony rights were granted: among such couples, men — and not women — were more likely to study, have more work interruptions, whereas they were less likely to work or work full time.

In a context where, all around the world, many couples choose to cohabitate before or in lieu of marriage, the issues we have explored here appear to be more and more relevant. Our results suggest that while government intervention in this "market" may have short-run impact on existing couples, they are unlikely to be able to alter outcomes more than a simple inter-temporary shift.

Furthermore, it is somewhat striking that couples may be influenced by a set of rules which they could have enacted themselves without any government involvement through a private contract. It is true that individuals involved in a personal relationship rarely itemize their respective rights and obligations in terms of a contract. In fact, Weiss and Willis (1993) document that few couples agree on the separation of assets and payment of alimonies through a pre-nuptial agreement. Understanding whether there are barriers that prevent these private agreements or if this type of legislation is necessarily welfare-reducing if partners are not willing to replicate it privately is left to further research.

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7 Tables and figures

Table 1: Summary of legislations granting alimony rights to cohabiting couples

Province	Legislation	Year	Length required	Special cases
Newfoundland	Family Law Act	2000	2 years	1 year with a child
PEI	Family Law Act	1995	3 years	1 year with a child
Nova Scotia	Maintenance and Custody Act	1989	1 year	
New Brunswick	Family Services Act	1980	3 years	
Quebec	No law			
Ontario	Family Law Act	1978	3 years	Automatic with a child
Manitoba	Family Maintenance Act	1983	5 years	1 year with a child
Saskatchewan	Family Maintenance Act	1990	3 years	Automatic with a child
Alberta	Domestic Relations Act	1999	3 years	1 year with a child
British Columbia	Family Relations Act	1979	2 years	

Table 2: Summary statistics (cross-section sample)

	Cohabitations						Marriages		
	before the law			after the law			all		
	N	mean	st.d.	N	mean	st.d.	N	mean	st.d.
Demographic characteristics									
Age relationship began	4025	25.98	8.46	3495	27.17	8.72	11279	26.18	7.69
Male	4025	0.47	0.50	3495	0.48	0.50	11279	0.48	0.50
English-speaking	4020	0.28	0.45	3484	0.88	0.32	11235	0.68	0.47
French-speaking	4020	0.70	0.46	3484	0.07	0.25	11235	0.26	0.44
Catholic	4001	0.70	0.46	3436	0.27	0.44	11077	0.44	0.50
Atheist	4001	0.13	0.34	3436	0.35	0.48	11077	0.15	0.36
Protestant	4001	0.14	0.35	3436	0.35	0.48	11077	0.39	0.49
Attended rel. services at age 15	3990	0.42	0.49	3432	0.36	0.48	11012	0.59	0.49
High school graduate	4024	0.81	0.39	3487	0.84	0.37	11255	0.81	0.40
College graduate	4024	0.19	0.39	3487	0.17	0.38	11255	0.20	0.40
Relationship characteristics									
Ended as marriage	4025	0.34	0.47	3495	0.40	0.49	11279	1.00	0.00
Relationship has ended	4025	0.49	0.50	3495	0.39	0.49	11279	0.28	0.45
Duration (years)	4025	5.07	5.46	3495	3.45	3.62	11279	16.03	11.36
Duration (total in years)	4025	8.90	7.90	3547	6.40	5.91	11279	16.03	11.36
Partner's characteristics									
Spouse was prev. unmarried	4019	0.76	0.43	3495	0.71	0.45	11266	0.86	0.35
Age difference (own-spouse)	3573	-0.61	4.87	3201	-0.50	4.84	11024	-0.28	4.21
Women at least 5 years older	3573	0.06	0.24	3201	0.08	0.26	11024	0.03	0.17
Male at least 5 years older	3573	0.24	0.43	3201	0.24	0.42	11024	0.20	0.40
Legislative status									
Subject to alimony rights	3967	0.06	0.23	3495	0.74	0.44	11279	0.61	0.49
... with no child rules	3967	0.06	0.23	3495	0.53	0.50	11279	0.60	0.49

All statistics weighted by person-specific weights.

Table 3: Summary statistics (panel sample)

	Cohabitations before the law		Cohabitations after the law		Marriages all	
	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev
Including all active relationships	N=24286		N=15513		N=186304	
Subject to alimony rights	0.05	0.21	0.49	0.50	0.42	0.49
Subject to alimony rights-no child rules	0.04	0.21	0.44	0.50	0.40	0.49
Studied	0.11	0.31	0.13	0.33	0.07	0.25
Worked	0.75	0.43	0.74	0.44	0.75	0.43
Worked full time	0.66	0.47	0.62	0.49	0.64	0.48
Stopped working	0.11	0.32	0.10	0.30	0.16	0.37
Number of children	0.87	1.15	0.86	1.16	1.68	1.27
Had a maternity leave	0.04	0.20	0.03	0.17	0.07	0.25
Had a paternity leave	0.00	0.05	0.00	0.04	0.00	0.03
Censored panel (at least 10 years)	N=39066		N=26379		N=194427	
Subject to alimony rights	0.04	0.18	0.34	0.47	0.40	0.49
Subject to alimony rights-no child rules	0.04	0.17	0.26	0.44	0.38	0.49
Studied	0.11	0.32	0.12	0.32	0.07	0.25
Worked	0.76	0.43	0.74	0.44	0.75	0.43
Worked full time	0.66	0.47	0.62	0.48	0.64	0.48
Stopped working	0.13	0.34	0.12	0.33	0.16	0.37
Number of children	0.90	1.12	0.93	1.14	1.66	1.27
Had a maternity leave	0.06	0.23	0.05	0.21	0.07	0.25
Had a paternity leave	0.00	0.06	0.00	0.04	0.00	0.03

All statistics weighted by person-specific weights.

Table 4: Impact of alimony rights on labor supply

	Studied	Worked	Worked full time	Work interruptions	Maternity leave	Paternity leave
	(1)	(2)	(3)	(4)	(5)	(6)
	Panel A: Only relationships formed before a law was passed					
Subject to alimony rights	0.020†	-0.030	-0.047†	0.053†	0.024	
	(0.011)	(0.022)	(0.024)	(0.027)	(0.019)	
Subject to alimony rights	-0.040†	0.086*	0.104*	-0.070†		-0.002
*male	(0.018)	(0.035)	(0.034)	(0.036)		(0.003)
R-square	0.475	0.701	0.722	0.445	0.375	0.101
N	39066	39066	39066	39066	22242	16824
	Panel B: All relationships except those "caught"					
Subject to alimony rights	0.009	0.030	0.058†	-0.059*	-0.054†	
	(0.013)	(0.021)	(0.030)	(0.025)	(0.026)	
Subject to alimony rights	0.019	-0.002	-0.018	0.038		-0.004
*male	(0.035)	(0.023)	(0.024)	(0.033)		(0.003)
R-square	0.485	0.716	0.731	0.445	0.375	0.094
N	63305	63305	63305	63305	35957	27348
	Panel C: Only relationships formed after the legal change					
Subject to alimony rights	-0.003	0.008	0.032†	-0.010	-0.002	
	(0.010)	(0.020)	(0.017)	(0.010)	(0.017)	
Subject to alimony rights	0.015	0.030	0.011	-0.017		-0.002
*male	(0.022)	(0.021)	(0.014)	(0.015)		(0.003)
R-square	0.498	0.738	0.741	0.434	0.343	0.123
N	26379	26379	26379	26379	14883	11496

All regressions include fixed effects for year of observation and the double-interactions between province, duration and year as specified in equation (38) for Panel A and equation (37) for Panel B. All regressions are weighted using person-specific weights. The sample includes at least ten years following the beginning of any cohabitation relationship or all the years if the relationship lasted longer.

Standard errors are clustered at the province level.

†: 10% significance, *: 5% significance, **: 1% significance

Table 5: Impact of alimony rules on labor supply-robustness checks

	No child rules (1)	Full panel (2)	Before 1990 (3)	After 1980 (4)	Without Quebec (5)	Married (6)	Lagged (7)
				Panel A: Studied			
Subject to alimony rights	0.020†	0.022	0.008	0.038*	0.013	0.002	0.030†
	(0.011)	(0.013)	(0.014)	(0.012)	(0.012)	(0.004)	(0.016)
Subject to alimony rights male	-0.038†	-0.034	-0.027	-0.075**	-0.034	-0.003	-0.067*
	(0.017)	(0.020)	(0.021)	(0.016)	(0.030)	(0.005)	(0.023)
Will be subject to alimony in 2 years							-0.008
							(0.033)
Will be subject to alimony in 2 years*male							0.066
							(0.040)
R-square	0.475	0.574	0.450	0.477	0.373	0.429	0.476
N	38793	24286	29212	27795	14848	150347	39066
				Panel B: Worked full time			
Subject to alimony rights	-0.044†	-0.028	-0.062†	-0.031	-0.125**	0.008	-0.072**
	(0.021)	(0.022)	(0.030)	(0.027)	(0.031)	(0.007)	(0.018)
Subject to alimony rights male	0.097*	0.076†	0.127**	0.108*	0.229**	-0.006	0.134**
	(0.030)	(0.037)	(0.035)	(0.041)	(0.065)	(0.007)	(0.036)
Will be subject to alimony in 2 years							0.045
							(0.041)
Will be subject to alimony in 2 years*male							-0.079
							(0.061)
R-square	0.720	0.775	0.717	0.727	0.703	0.723	0.723
N	38793	24286	29212	27795	14848	150347	39066
				Panel C: Work Interruption			
Subject to alimony rights	0.070*	0.054	0.060	0.068†	0.021	-0.035**	0.052
	(0.025)	(0.030)	(0.037)	(0.029)	(0.031)	(0.008)	(0.031)
Subject to alimony rights male	-0.080*	-0.051	-0.082	-0.097*	-0.054	0.054**	-0.070
	(0.032)	(0.044)	(0.048)	(0.033)	(0.042)	(0.014)	(0.039)
Will be subject to alimony in 2 years							-0.054
							(0.044)
Will be subject to alimony in 2 years*male							0.077
							(0.046)
R-square	0.435	0.471	0.423	0.443	0.439	0.523	0.446
N	38793	24286	29212	27795	14848	150347	39066

All regressions include fixed effects for year of the observation and their interaction with a male dummy and the double-interactions between province, duration and year as specified in Equation 38. All regressions are weighted using person-specific weights. The sample includes at least ten years following the beginning of any cohabitation relationship, except in column (2) where it includes only years where the cohabitation relationship was active. The first column only includes, as subject to alimony rules, couples who qualified because of the duration of their relationship, not because they had children. The third restricts it to relationships that began before 1990, the next one, to couples formed before 1980. The fifth column excludes all couples from Quebec. The sixth only includes married couples and the last includes an indicator for becoming subject to the law in 2 years.

Standard errors are clustered at the province level.

†: 10% significance, *: 5% significance, **: 1% significance

Table 6: Impact of alimony rights on labor supply-heterogenous effects

	Studied		Worked full time		Work Interruptions	
	Main effect	Interaction	Main effect	Interaction	Main effect	Interaction
	(1)	(2)	(3)	(4)	(5)	(6)
By education:						
Less than high school	0.021	0.000	0.029	0.014	0.096†	-0.109
	(0.022)	(0.031)	(0.036)	(0.037)	(0.044)	(0.062)
High school graduate	0.027*	-0.046	-0.053	0.107†	0.040	-0.045
	(0.010)	(0.030)	(0.044)	(0.054)	(0.039)	(0.051)
College graduate	-0.019	-0.210	-0.184*	0.335**	0.032	-0.147
	(0.048)	(0.127)	(0.073)	(0.091)	(0.078)	(0.136)
By age:						
Began rel. before age 21	0.063†	-0.025	0.034	0.043	-0.002	0.021
	(0.033)	(0.057)	(0.054)	(0.118)	(0.060)	(0.052)
Began rel. between 22-27	-0.005	-0.017	-0.086	0.142*	0.085	-0.111
	(0.017)	(0.026)	(0.047)	(0.047)	(0.064)	(0.069)
Began rel. after 28	0.008	-0.081	-0.079†	0.117	0.069	-0.100
	(0.044)	(0.061)	(0.037)	(0.067)	(0.073)	(0.078)
By age difference:						
Fem. at least 5 yrs younger	0.063	-0.109**	-0.090*	0.187**	0.158*	-0.211*
	(0.035)	(0.017)	(0.035)	(0.049)	(0.054)	(0.075)
Within 5 years of spouse	-0.012	-0.010	-0.024	0.031	0.016	-0.005
	(0.024)	(0.036)	(0.025)	(0.032)	(0.055)	(0.069)
Fem. at least 5 yrs older	-0.003	0.062*	0.043	-0.107	-0.032	-0.030
	(0.013)	(0.025)	(0.034)	(0.060)	(0.056)	(0.057)

All regressions include fixed effects for year of observation and their interactions with a male dummy and the double-interactions between province, duration and year as specified in equation (38). All regressions are weighted using person-specific weights. The sample includes at least ten years following the beginning of any cohabitation relationship. Each set of columns ((1), (2), (3) and (4), (5), (6)) and section of the table correspond to a regression. The table entries are the coefficients of an indicator for the couple being subject to alimony rules interacted with the characteristics as listed in the first column of the table (for columns (1), (3), (5)) and the interaction of that term with a dummy for the respondent being a male in columns (2), (4) and (6).

Standard errors are clustered at the province level.

†: 10% significance, *: 5% significance, **: 1% significance

Table 7: Policy neutrality of alimony rights rules

	Studied	Worked	Worked Full-Time	Work Interruption
	(1)	(2)	(3)	(4)
Panel A: Comparing new and old relationship (N=63493)				
Formed after a legal change	-0.076	0.003	0.017	-0.076**
	(0.208)	(0.034)	(0.029)	(0.020)
Formed after a legal change*male	0.001	0.003	-0.053	0.071**
	(0.453)	(0.036)	(0.037)	(0.021)
R-square	0.021	0.046	0.055	0.037
Panel B: Comparing new and long relationships over time (N=63493)				
Formed after a legal change	-0.013	0.004	0.020	-0.091**
	(0.010)	(0.033)	(0.031)	(0.021)
Formed after a legal change*male	0.033**	-0.005	-0.061†	0.083**
	(0.010)	(0.032)	(0.033)	(0.024)
Formed after a legal change	-0.002	-0.004*	-0.005**	0.008**
time since beginning	(0.001)	(0.001)	(0.002)	(0.001)
Formed after a legal change	0.000	0.004	0.006*	-0.007**
time since beginning*male	(0.001)	(0.003)	(0.002)	(0.001)
R-square	0.032	0.047	0.056	0.039

All regressions include fixed effects for province, year of observation and their interactions with a male dummy. Regressions in Panel B also include controls for the duration of the relationship and its interaction with being male. The sample includes all years of common-law relationships and the ten years following the beginning of one for those that lasted too short, except for years where the couple was subject to alimony rights when they had entered into the relationship before the legal change. All regressions are weighted using person-specific weights.
Standard errors are clustered at the province level.
†: 10% significance, *: 5% significance, **: 1% significance

Table 8: Impact of alimony rules on relationship stability

	Relationship is over in 2001		Cohabitation led to marriage		Duration		Duration (total)	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Panel A: Only relationships formed before a law was passed							
Subj. to alimony rights	-0.109	-0.117	-0.140**	-0.177**	2.604*	2.125*	0.530	0.505
	(0.081)	(0.085)	(0.025)	(0.047)	(0.806)	(0.705)	(0.762)	(1.180)
Subj. to alimony rights		0.063		0.064		0.468		-1.122
male		(0.063)		(0.098)		(0.517)		(1.643)
R-square	0.294	0.308	0.262	0.275	0.917	0.921	0.480	0.495
N	4025	4025	4025	4025	4025	4025	4014	4014
	Panel B: All relationships except those "caught"							
Subj. to alimony rights	0.051	0.029	-0.007	0.017	-0.211*	-0.347*	-1.262†	-0.587
	(0.038)	(0.035)	(0.020)	(0.029)	(0.084)	(0.114)	(0.620)	(0.692)
Subj. to alimony rights		0.063		-0.045		0.230†		-1.806
male		(0.043)		(0.060)		(0.108)		(1.010)
R-square	0.261	0.269	0.262	0.276	0.955	0.956	0.502	0.516
N	7163	7163	7163	7163	7163	7163	7147	7147
	Panel C: Only relationships formed after the legal change							
Subj. to alimony rights	-0.029	-0.059*	0.007	-0.036	-0.097†	-0.046	-0.210	-0.149
	(0.016)	(0.018)	(0.014)	(0.025)	(0.044)	(0.091)	(0.242)	(0.414)
Subj. to alimony rights		0.082†		0.092†		-0.108		-0.199
male		(0.038)		(0.044)		(0.103)		(0.627)
R-square	0.210	0.224	0.259	0.279	0.948	0.948	0.555	0.578
N	3495	3495	3495	3495	3495	3495	3489	3489

All regressions include fixed effects for province, beginning year of cohabitation, duration categories and the double-interactions as specified in equation (36) for Panel A and equation (35) for Panel B. All regressions are weighted using person-specific weights.

Standard errors are clustered at the province level.

†: 10% significance, *: 5% significance, **: 1% significance

Table 9: Impact of alimony rights on relationship stability-robustness checks

	No child rules (1)	With covariates (2)	Before 1990 (3)	After 1980 (4)	Without Quebec (5)	Lag (6)
	Panel A: Relationship is over in 2001					
Subj. to alimony rights	-0.085	-0.109	-0.172**	-0.105	-0.134**	-0.023
	(0.084)	(0.080)	(0.046)	(0.110)	(0.032)	(0.074)
Will become subject to						0.040
alimony rights in 2 years						(0.057)
R-square	0.290	0.300	0.219	0.311	0.135	0.302
N	4012	3985	2529	3076	1411	4025
	Panel B: Cohabitation led to marriage					
Subj. to alimony rights	-0.147**	-0.140**	-0.118*	-0.149**	-0.034	-0.258**
	(0.022)	(0.027)	(0.039)	(0.024)	(0.070)	(0.045)
Will become subject to						-0.011
alimony rights in 2 years						(0.052)
R-square	0.259	0.265	0.196	0.246	0.117	0.265
N	4012	3985	2529	3076	1411	4025
	Panel C: Duration					
Subj. to alimony rights	2.503*	2.581*	3.903**	1.592†	4.111**	3.238**
	(0.814)	(0.817)	(0.641)	(0.716)	(0.683)	(0.926)
Will become subject to						0.206
alimony rights in 2 years						(0.321)
R-square	0.920	0.918	0.914	0.942	0.765	0.945
N	4012	3985	2529	3076	1411	4025
	Panel D: Duration (total) in 2001					
Subj. to alimony rights	0.328	0.463	1.736*	-0.137	2.618*	1.045
	(0.743)	(0.764)	(0.571)	(0.633)	(0.938)	(0.698)
Will become subject to						-2.643*
alimony rights in 2 years						(1.090)
R-square	0.479	0.484	0.279	0.577	0.295	0.497
N	4001	3974	2518	3073	1402	4014

All regressions include fixed effects for province, beginning year of cohabitation, duration categories and the double-interactions as specified in equation (36). All regressions are weighted using person-specific weights. The first column only includes, as subject to alimony rules couples who qualified because of the duration of their relationship, not because they had children. The second column includes controls for language, religion and education. The third restrict it to relationships that began before 1990, the next one, to couples formed before 1980. The fifth column excludes all couples from Quebec. The sixth only includes married couples and the last includes an indicator for becoming subject to the law in 2 years.

Standard errors are clustered at the province level.

†: 10% significance, *: 5% significance, **: 1% significance

Table 10: Impact of alimony rights on relationship stability-heterogenous effects

	Relationship over in 2001 (1)	Cohabitation led to marriage (2)	Duration (3)	Duration (total) (4)
By education:				
Less than high school	-0.170†	-0.124*	3.236**	1.371
	(0.077)	(0.046)	(0.907)	(1.024)
High school graduate	-0.090	-0.111**	2.149*	0.345
	(0.083)	(0.018)	(0.768)	(0.715)
College graduate	-0.057	-0.358*	3.406*	-0.649
	(0.115)	(0.129)	(1.403)	(1.458)
By age:				
Began rel. before age 21	-0.105	-0.042	1.712*	0.174
	(0.075)	(0.066)	(0.697)	(1.130)
Began rel. between 22-27	-0.006	-0.166**	2.538*	-0.755
	(0.108)	(0.040)	(1.020)	(0.785)
Began rel. after 28	-0.185†	-0.192**	3.339*	1.703
	(0.090)	(0.058)	(1.037)	(1.306)
By age difference:				
Female at least 5 years younger	-0.078	-0.235**	3.149*	0.047
	(0.098)	(0.047)	(1.032)	(1.008)
Within 5 years of spouse	-0.075	-0.122**	2.284*	0.197
	(0.079)	(0.027)	(0.792)	(0.680)
Female at least 5 years older	-0.213†	-0.061	1.776†	0.898
	(0.106)	(0.089)	(0.830)	(1.118)

All regressions include fixed effects for province, beginning year of the cohabitation, duration categories and the double-interactions as specified in equation (36). All regressions are weighted using person-specific weights. Each column and section of the table correspond to a regression. The table entries are the coefficients of an indicator for the couple being subject to alimony rules interacted with the characteristics as listed in the first column of the table.
Standard errors are clustered at the province level.
†: 10% significance, *: 5% significance, **: 1% significance