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*Marcos A. Rangel*

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# Alimony Rights and Intrahousehold Bargaining: Evidence from Brazil*

Marcos A. Rangel†
Department of Economics
University of California at Los Angeles, UCLA

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**Abstract**

While theoretical models of family and household decision-making that highlight the role of the individual provide sharp empirical predictions, testing lags far behind. This paper provides a robust assessment of how shifts in the within-household balance of "decision power" affect family-level behavioral choices regarding labor supply and investments in the human capital of children. Using an exogenous source of variation provided by the adoption of a law (extension of alimony rights to cohabitants), this paper provides empirical evidence that (intra-household) empowerment of women resulted in the reduction of hours worked by female heads, and in the redistribution of household resources towards schooling of first-born girls. The results reveal heterogeneous characteristics of the parental preferences that are not compatible with the "unitary" representations of the household.

**Keywords**: Intrahousehold bargaining, cohabitation, alimony rights, time allocation, investments in children.

**JEL codes: O12, D13, J12, J22**

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"...the marital couple is not an independent entity with a mind and heart of its own, but an association of two individuals each with separate intellectual and emotional makeup."

**The US Supreme Court**
**Griswold *vs.* Conneccticut, 1965**.

# 1 Introduction

In recent years the study of decision-making within the household has received considerable attention in the social sciences literature. The extent to which members of the same household individualistically pursue their self-interest as opposed to being characterized by a group who share a common goal (i.e., common preferences) is an important question for many areas of scientific inquiry. Households are the institution within which life's basic decisions are made - decisions about having and raising children, about time spent on work, about consumption and investment. Therefore, accounting for the potential role of intra-familial bargaining is key in understanding issues which have importance in the measurement of individuals' well-being. Paralleling a line of demographic literature,[1] theoretical models of family and household decision-making that highlight the role of the individual have been developed since seminal contributions by McElroy & Horney (1981), Manser & Brown (1980), and Chiappori (1988). While these models provide sharp empirical predictions, testing lags far behind. This paper provides a robust assessment of those models by focusing on how exogenous shifts in the within-household balance of "decision power" affect families' behavioral choices.

At the core of all models of household decision-making is the idea that individuals share the benefits of forming a household in some way. Their premise is that observed behavior is a result of negotiations among family members, reflecting each individual's perception of costs/benefits as well as their relative "power" in asserting their own preferences at the household-level. A key difficulty in this interpretation is the fact that power is not clearly defined and is expected to (endogenously) vary across time and social contexts. In order to circumvent the endogeneity problem of other power measures used in early contributions to the literature, such as the non-labor income distribution within households,[2] this paper exploits a change in the Brazilian legislation as a source of exogenous variation. In particular, the causal inference is based on the extension of alimony rights (previously exclusive to formally married couples) to a large fraction of the adult population living in consensual unions.[3] The hypothesis in the present study is that the

[1]See Stark (1984).

[2]See Thomas (1990, and 1993), Schultz (1990), and Bourguignon et al. (1993).

[3]Throughout the text, **cohabitation** and **consensual union** are used interchangeably. They are short for "marital cohabitation without legal marriage", "marriage performed with less solemnity than a true formal marriage", or even "marriage by habit and reputation".

concession of alimony rights improves the women's outside-option (that is, their welfare level in case of relationship dissolution) and should strengthen their bargaining position within their partnerships, increasing their ability to appropriate a bigger share of the gains from the marital interaction. Holding other factors that influence the household decision making constant, changes in the allocation of resources associated to the new legislation's implementation should reflect the effects of power reallocation only if individual preferences within the household are heterogeneous.

The empirical analysis in the present study is based on data from the Brazilian Household Survey (*Pesquisa Nacional de Amostra de Domicilios*-PNAD) from 1992, 1993 and 1995. The PNAD yearly sample consists of approximately 45 thousand observations on adult couples, and 65 thousand observations on "children" (ages 5 to 17). The large samples allow various stratification strategies, which ensure the robustness of the empirical results. The econometric analysis investigates the compatibility of the data with the bargaining power hypothesis raised above by focusing on household-level decisions regarding: i) adults' time allocation to market and housekeeping activities, and; ii) investments in the human capital of children (schooling).

The empirical evidence uncovered in the present study indicates that (intra-household) empowerment of women caused an increase in the female consumption of leisure and, most importantly, reallocation of resources towards schooling of first-born daughters (relative both to their younger sisters). In general, first-born daughters of cohabiting couples have a probability of staying in school 9% higher than their younger sisters. Interestingly, this effect was shown to be stronger when considering a stratum of the households where the mother would be more dependent on alimony in case of relationship's dissolution (i.e., the less educated). These changes in allocation of investments in children's human capital in response to outside-option changes are not predicted by models that treat the household as a "common preferences" unit. Therefore, the results reveal heterogeneous characteristics of the parental preferences underlying the within-household allocation of adults' time and of investments in children. In a country where poverty has been shown to be pervasive and strongly correlated with low education and child labor,[4] such findings regarding the intrahousehold decision-making process governing investments in the human capital of children also shed light on intricate social mobility issues, even though these are not directly explored in this paper.

The remaining sections of the paper are organized as follows. Section 2 describes demographic trends of marriage and cohabitation, summarizing some sociological interpretations. In Section 3, the economic theory of the household and its empirical implications are presented. Section 4 focuses on the institutional environment. The description of the data sets and the econometric identification strategy are presented in Section 5. Section 6 presents the estimation results. Section 7 concludes with a discussion about the relation between the present paper's results and other contributions to the literature.

[4]See Barros (2001), for example.

## 2 Marriage, cohabitation, and the family

Cohabitation is reshaping family life all over the world. Among European countries from the mid-1970's to the mid-1990's, the percentage of young-adult women (ages 20 to 24) in a marital union without being formally married increased from 11% to 49% in France, 57% to 78% in Sweden, 48% to 75% in Denmark, and from 11% to 55% in Great Britain.[5] In the United States, even though cohabitation is considered a recent phenomenon, Casper et al. (1999) reports that the proportion of young women (ages 15 to 24) in cohabitation scaled up by a factor of three between 1977 and 1997 (2.5% to 7.5%). Finally, according to Westoff et al. (1994)'s analysis of late 1980's figures, in a handful of African countries and in all countries in Latin America & the Caribbean at least 15% of women in a union were in consensual relationships.

Brazil is not an exception to the cohabitation popularization trend. Even though the ratio of women in marital unions has remained stable (1960-1995),[6] the proportion of cohabitations among all nuptial arrangements has rapidly increased since the mid-1980's, as seen in the figures reproduced in Table 1. The change in stocks reflects the reduction in legal-marriages and the acceleration of cohabitation, divorce, and separation rates started. They coincide with cultural changes that culminated with a generalized emancipation of women, a period "during which women's opportunities outside home increased substantially", as phrased by Thomas (1994). According to Brazilian demographic studies, as Goldani (1997), those were the most significant changes observed in the Brazilian social and microeconomic environment between 1975 and 1995. The pattern of marital status in the first half of the 1990's (reproduced in Table 2) indicate that cohabitation not only corresponds to a large fraction of marital unions among latter cohorts, but also to a significant share among the earlier ones.

If on one hand the fact that divorce, separation and cohabitation represent instability of marital unions and arguably explain the increased proportion of female single-headed households observed in the late 1990's,[7] on the other, the rise in cohabitation rates can also be interpreted as offsetting the upward(downward) trend in divorce(marriage). In a vision shared by a growing number of sociologists, the indication is that, even though cohabitation had reshaped family life, it did so in a subtle way. In other words, cohabitation "should not be considered an increase in singlehood" or a "retreat from familism", but rather it should be seen "very much like a family status", as expressed in Treas & Lawton (1999) and Bumpass et al (1991), respectively. This seems to the true for Brazil, where half of the consensual unions endure for more than 5 years.[8] Moreover, anecdotal evidence presented by Rao & Greene (1996) indicates that Brazilian cohabitors generally refer to themselves as "married", and use the words "husband" and "wife" to refer to their partners. Most importantly, the informal nature of the marital relationship does not prevent the formation of families.

[5]See sources cited in Emrisch & Fracescone (2000).

[6]See Lazo (2002).

[7]See Medeiros et al (2002) and Medeiros & Osorio (2002).

[8]Author's calculations based on raw data from the World Bank's Living Standards Measurement Study (1996/1997), co-organized by the Brazilian Institute of Geography and Statistics (IBGE).

Lazo (1994) reports that "total marital fertility" is higher among cohabiting couples than among formally married ones, even after controlling for age, education and duration of union.[9]

Given the Brazilian context, two major considerations should be underscored: First, as cohabiting couples represent a large proportion of the population, investigating the effects of law changes which directly impact adults and children in such informal marital arrangement is an important issue in itself. Second, the similarities of cohabitation and marriage in terms of familial arrangements makes the latter a comparable group regarding the effects of shifts in the decision power over intra-household allocation of resources.[10]

Inference about shifts in bargaining power works as a reverse engineering process: from observed behavior, structural characteristics of the decision-making process are validated or not, and, in some cases, even identified. With this objective, the next section briefly introduces relevant household-economics theory, emphasizing its potential for empirical examination.

## 3 Conceptual framework

Due to its simplicity and convenience, the most well known model of household behavior is based on the assumption that a representative individual can serve as the element of study.[11] In the so-called "unitary model", the implicit assumption is that for a coalition of two or more individuals to act together with a common purpose, each individual has to set aside separate preferences while the group creates something completely new - a set of common preferences for defining their collective behavior. Using this shortcut, individualistic behavior can be directly projected to household units and to households' interactions at the community level, automatically fitting the treatment of consumer choices in the microeconomics literature. Such model describes the operation of households as "black-boxes", leaving unexplored interesting elements of the intrahousehold interaction. It is silent, for example, with respect to the formation of alternative (informal) marital arrangements, and to the increased rates of households' dissolution. In essence, as posed by Thomas et al. (2002), one can say that, in certain domains, the unitary model is based on "assumptions that are neither theoretically appealing nor likely to be innocuous".

In an attempt to fill these gaps, intra-household bargaining and collective decision-making models have lately been developed in the economics literature. These models attempt to relax the above assumptions, while restricting observed behavior in a way that specific nuances of the household decision process could be empirically verified.[12]

[9]See the briefing paper "Research on Today's Issues" (September, 2002), by the National Institute of Child Health and Human Development (NICHD) for a North-American perspective on off-marriage fertility. See also Seltzer (2002).

[10]See Sprey (1999) for a general discussion.

[11]See review papers in Rosenszweig & Stark (1997) and in Haddad et al (1997).

[12]See Doss (1996).

### 3.1 Modeling

The static household welfare evaluation corresponds to a function or a specific weighted aggregation over the "felicity functions" of both heads (decision-makers in this context):

$$W^h = W[U^m(X, q^c; k^h, \varepsilon^h), U^f(X, q^c; k^h, \varepsilon^h)]$$

Where $X$ represents the consumption vector (including leisure), and the superscripts on the individual felicity functions indicate association with the male ($m$) or female ($f$) head. $q^c$ represents the "quality" of child $c$ (a good produced in the household), and $k^h$ and $\varepsilon^h$ represent observed and unobserved characteristics of the household, respectively. Observed and unobserved characteristics of the household are combinations of heads' characteristics determined at the time of the matching in the marriage market. However, except for their implication for the empirical strategy, this study abstracts from matching issues.[13] In this formulation, children are assumed to be passive, with their consumption and time allocations decided by the parents.

The child's quality is represented by the production function $q^c$, which summarizes an implicit and unknown relation between child outcomes and childrearing inputs (breast-feeding, externalities of adult consumption of cigarettes, school enrollment, caloric composition of diet, etcetera):[14]

$$q^c = q\left(X; k^h, k^c, \varepsilon^h, \varepsilon^c\right)$$

The fact that the consumption vector influences the production output rules out separability assumptions for the felicity functions. Child's individual (observed and unobserved) characteristics not captured in the household characteristics are represented by $k^c$ and $\varepsilon^c$. It is implicitly assumed, therefore, that each parent uses the see the production process of child quality in the same way - i.e., they assign the same weights to different types of consumption goods and investments. They are, however, allowed to disagree in the valuation of the child's quality relative to their own consumption.

Households are constrained in time ($\overline{L}$) and by their budgets. The latter, taking prices $\left[P = (p; w^m; w^f; w^c)\right]$ as exogenous parameters, is represented by:

$$P \cdot X = (w^m + w^f + w^c) \cdot \overline{L} + y^m + y^f$$

Where $w$ is the market wage, and $y$ represents the individual-level non-labor income.

*a) the unitary model*

[13] See Foster (1998) and Bergstrom (1993) for a discussion on the topic.

[14] It is fair to assume that the parents evaluation of their children 's quality is (simultaneously) based on the observation of outcomes and parental effort towards their improvement (inputs).

In the unitary model, either by the power of consensus (Samuelson, 1956) or by the emergence of a dictator (Becker, 1991), the decision process is summarized by a representative individual utility function. This means that either the household agrees on the utility derived from each choice to be made (therefore, the felicity functions in $W$ are identical) or the choices will be based on the dictators' view of the world (the non-dictator felicity gets zero weight in $W$).

In the model language described above, the unitary model assumptions correspond to the following implications in terms of demand functions:

$$X = X\left(P; Y, k^h, k^c, \varepsilon^h, \varepsilon^c\right) \tag{1}$$

where $Y = \left(y^m + y^f\right)$. This implies that only the household's pooled income influences the choices. The household is a single and monolithic decision unit, and the income allocation across household members does not matter for time and expenditure decisions, nor does any other parameter affecting individual welfare if the relationship were to dissolve.

*b) the cooperative bargaining models*

Cooperative bargaining models are essentially an attempt to model cooperative behavior without abandoning the individual decision-theoretic foundations of microeconomic theory. Most of these models are based on Nash's solution to a two-person game, where the final payoff (excess utility obtained under agreement among household members) depends exclusively on the reservation utility of each agent when the disagreement takes place. Each individual has, therefore, specific preferences that are pursued while bargaining takes place. As the welfare under the disagreement condition increases(decreases), individuals are able to appropriate a greater(smaller) share of the surplus derived from their interaction.

Two applications of this concept have been used in household economics: Manser & Brown (1980) and McElroy & Horney (1981) explore reservation utilities determined by household (marriage) dissolution (divorce); Lundberg & Pollack (1993) argue in favor of "threat points" being the result of a non-cooperative game among the household members. In the divorce-threat case, individual power is determined by the attractiveness of the divorced-status option. In the no-cooperation threat case, the bargaining position is determined by individual roles in the provision of household-level public goods affected by the lack of cooperation.[15]

In both cases, however, the axioms of cooperative bargaining correspond to the solutions of the maximization, subject to the same constraints as above, of an implicit welfare function represented by the "Nash product":

[15]See Lundberg & Pollack (1993) and Pollack (2002).

$$W = \left[U^m(X, q^c; k^h, \varepsilon^h) - V^m\left(P^0, P^m; \lambda^m\right)\right] \times \left[U^f(X, q^c; k^h, \varepsilon^h) - V^f\left(P^0, P^f; \lambda^f\right)\right]$$

Where $V$ reflects the reservation indirect utility function, $\lambda^i$ summarizes (non-price) characteristics that influence the welfare under the outside option of each of the heads, and $P$ is partitioned into public ($P^0$) and private good ($P^i$) prices. The outcome in terms of a general reduced-form demand function is represented by:

$$X = X\left(P, Y, \lambda^m, \lambda^f; k^h, k^c, \varepsilon^h, \varepsilon^c\right) \tag{2}$$

Notice that demands are now affected by the parameters that affect the balance of decision power in the household.

*c) collective models*

Collective models were developed in Chiappori (1988, 1992) and further explored by the same author in collaboration with other scholars. The basic reasoning in such models is to appeal to the rationality of individuals in a context that abstracts from specific bargaining rules. It axiomatically requires that the actors in the process reach Pareto-efficient agreements. The justification of this assumption is almost invariably based on the repeated nature of the interaction between married individuals, which promotes the exhaustion of efficiency enhancement possibilities.

In the collective model's basic form, the efficiency assumption is equivalent to the following implicit structure for the household maximization problem:

$$\begin{aligned}
& \max_{X^f, X^m, X^c} W = \mu U^m\left(X, q^c; k^h, \varepsilon^h\right) + (1-\mu) U^f(X, q^c; k^h, \varepsilon^h) \\
s.t. \quad : \quad & P \cdot X = \left(w^m + w^f + w^c\right) \cdot \overline{L} + y^m + y^f \\
: \quad & q^c = q\left(X; k^h, k^c, \varepsilon^h, \varepsilon^c\right),
\end{aligned}$$

where the newly introduced $\mu$ is known as the Pareto weight, and reflects the relative importance of each of the heads in the aggregated household utility. Taking these weights as exogenously fixed and interior (bounded away from zero and one), the objective described above reproduces the consensus model *a la* Samuelson (1956). If exogenously set at one or zero, they reproduce the "Beckerian" dictator model. For the purpose of this model, they correspond to bargaining positions derived from the social and cultural environment. Therefore, they are assumed to be a function of prices, the income distribution between head and spouse, and the previously discussed environmental parameters: $\mu\left(P, \lambda^m, \lambda^f\right)$.

The assumptions of this model imply the following reduced-form demand functions:

$$X = X\left(P^0, P^i, Y, \mu; k^h, k^c, \varepsilon^h, \varepsilon^c\right)$$

or:

$$X = X\left(P, Y, \lambda^m, \lambda^f; k^h, k^c, \varepsilon^h, \varepsilon^c\right) \tag{3}$$

Notice that they correspond, therefore, to the outcome of the bargaining models discussed above. This is a result of the fact that, under symmetric information, the collective model is a generalized version of the cooperative bargaining ones.

## 3.2 The empirical content of the models

When compared with the unitary model, the most basic difference emerging from models that consider the bargaining processes is that elements that exclusively influence the individuals' outside-option (what determines the bargaining power) should also affect the allocation of resources within an intact relationship. In terms of the reduced-form demands described in the previous subsection, it corresponds to the effects of $\lambda$'s on the allocation of household resources. Therefore, in theoretical terms, the unitary model is a valid simplified characterization only if $\frac{\partial X}{\partial \lambda^i} = 0$ for both heads ($i = m, f$), or, in terms of relative bargaining power, simply if:

$$\frac{\partial X}{\partial e} = 0$$

where $e = f(\lambda^f, \lambda^m)$ is assumed to be an indicator of relative bargaining power of female heads.[16]

The challenge is, therefore, the identification of meaningful empirical counterparts for those parameters that influence the distribution of bargaining power. The empirical literature on the topic has focused on two versions of those restrictions:

a) The unitary model would be a valid representation of the decision making regarding allocation of resources towards $X$ if redistribution of non-labor income between the heads, holding total non-labor income constant, did not affect expenditure/investment decisions. In other words, the effect of male and female non-labor income should be identical if demand only depends on the pooled non-labor income (income-pooling hypothesis):

$$\left[\frac{\partial X}{\partial y^m}\right]_{dY=0} = \left[\frac{\partial X}{\partial y^f}\right]_{dY=0} = 0$$

[16]As $e = \frac{\lambda^f}{\lambda^m}$ or $e = \lambda^f - \lambda^m$, for example.

b) variables that exclusively affect the welfare outside the relationship should not influence outcomes if bargaining among members were not a significant component of the decision process taking place within the household (the unitary model is either a dictatorship or a group with complete agreement about preference ordering), with decision power being either completely concentrated (on the dictator's hands) or unimportant:

$$\frac{\partial X^i}{\partial e} = 0$$

The characteristics summarized in $e$ reflect what McElroy (1990) calls "extrahousehold environmental parameters". They are an allusion to elements that influence the status under the threat point condition (and the decision process as a whole) but do not affect preferences or the budget constraint. Also known as "distribution factors" they encompass, for example, indicators of the (re)marriage market competitiveness, welfare programs conditional on marital status, laws governing divorce and marital property division, child support, social networks and religion, or even general "sociocultural norms".

In the case of income-pooling restrictions (a), it is somewhat difficult to argue, however, that the cross-sectional differences in the amount of non-labor income assignable to each particular household member (explored in early contributions) are exogenous to the expenditure/time allocation decision itself, or to individuals' unobserved characteristics. In this context, endogeneity can bias the results against the unitary model. To avoid such endogeneity problems, a small number of recent contributions following the seminal work by Lundberg, Pollack & Wales (1997) have exploited "exogenous variations" promoted by changes in government transfers programs. Based on the re-arrangement of household income generated by changes in the child allowances payments from husbands' pay checks to wives' bank accounts, for example, those studies have uncovered important aspects of the intrahousehold decision making.[17]

Rubalcava & Thomas (2000) and Gray (1998), alternatively, interpret exogenous variation in the income prospects of non-married women as changes in "extra-household environmental parameters" for intact households, in the spirit of restriction (b). Both studies design "quasi-experiments" by exploring across-households heterogeneity in the exposition to those changes. Rubalcava & Thomas (2000) focuses on changes in (potential) welfare payments to single-mothers (AFDC) over time and across US-states, showing that they influence the composition of expenditures in poor households headed by married couples. Gray (1998) explores US state-specific divorce settlements and property division legislation to proxy for changes in the bargain power of married women, investigating their effects over female labor-supply. These studies can be subject to some limitations, however. Rubalcava & Thomas (2000)'s estimates may be biased if there is welfare-motivated migration amongst poor families, which has recently been shown to be significant by Gelbach (2002). Gray (1998)'s analysis, on the other hand, may be confounded by state-specific trends in

[17] See Ward-Batts (2000), for example. Rubalcava et al (2002), on the other hand, explores random allocation of income-transfers to women in the context of the Mexican PROGRESA. Importantly, in such formulation, control groups are explored in order to net-out effects of confoundifng factors.

economic conditions.

In conclusion, the rejection of unitary model restrictions are more robust if tests can be based on "exogenous variation", and, even better, on "natural-experiments". It is with this experimental design in mind that this paper "takes the models to the data". The next section draws a general picture of the changes in the Brazilian legal system that are explored in the empirical exercises.

# 4 Institutional environment

The long-term demographic trends briefly described in Section 2 motivated courts and legislators to rethink the legal basis of the family. In Brazil, this manifested itself as a slow evolution of the laws from guaranteeing rights of a specific "ideal" family unit towards the protection of alternative familial arrangements, and led to the definition of rights and responsibilities of individuals within the family (with respect to the State and to each other).

The first steps taken were in terms of the rights and responsibilities with respect to the State and other third parties (Labor Law, the Social Security Law, and some aspects of the Civil Code). Important legislative measures were: i) extension of inheritance and alimony rights of biological sons born out-of-wedlock (1942, 1949, and 1977); ii) cohabiting spouses were allowed to be the beneficiary of partner's workplace-accidents pension (1944); iii) extension of real state rental contracts to cohabitant spouses after the death of the partner (1961); iv) the cohabitant partner could be listed as dependent for income tax purposes (1963, and 1965), and; v) the cohabitant partner could be listed as the beneficiary of a social security pension (1962, 1969, 1974, and 1976).

Legislation regarding property division and alimony following the relationship dissolution was, however, much slower to change. Jurisprudence in these areas essentially followed an outdated Civil Code (1942) and consistently denied cohabitants alimony and property division rights.[18] It was only in 1988, with the new Federal Constitution (Article 226, paragraph 3), that new guidelines were set. In its text, the Constitution universalized the spirit of the (topic-specific) laws that ruled the relation between the State and cohabiting couples, establishing that **stable consensual unions** should be recognized as a legal family entity "for purposes of protection by the State".[19] Additionally, religious marriages, considered until then as informal unions for legal purposes, were granted the recognition by the State - provided that they were confirmed in a "Public Registry Office".[20]

[18]See Azevedo (1997).

[19]See Appendix A, for a complete version of the Constitutional Article 226.

[20]Matielo (1998) describes that, due to lack of information about the new Constitution or about the registration requirement (not explicit in the Constitutional text), most of religious marriages can still be considered consensual unions for the effects of the law.

In the context of individual rights and obligations between partners, the Constitution had distinct effects in terms of property division and alimony. Jurisprudence on property division, for example, adopted the Federal Supreme Court's Recommendation 380 ("*Sumula 380*" of the *Supremo Tribunal Federal*) to the cohabitation context. Until then, S*umula 380* exclusively ruled rights over property after business partnerships' break up. In practice, it meant that, starting in the early 1990's, informally married spouses were given the right over their partners' property when able to prove collaboration for its accumulation.[21]

In the context of alimony rights, the Constitutional text was ineffective, however. Jurisprudence held that cohabitation would not yield maintenance rights to the cohabitants under any circumstances, as it had before 1988.[22] The justification was, according to Ribeiro (2002) and Branco (1994), that the Constitution had only established the responsibility of the State with respect to cohabitants, but not the responsibility of one partner to the other in the context of maintenance. Pessoa (1997) and Matielo (1998) further emphasize that, although it had recognized cohabitations as a family unit, "consensual union" and marriage were still distinct entities. In effect, the Constitutional text suggested the necessity of legal facilitation to convert such an arrangement into marriage. In the understanding of the Brazilian courts, this meant that maintenance obligations were only possibly derived from biological paternity/maternity and formal marriages. Post-dissolution alimony rights and obligations pertinent to the latter should, therefore, not be extended to cohabitation. Based on this argument, alimony petitions were normally dismissed without a hearing, and state-level Superior Courts usually corroborated these decisions.[23]

Diverging from the *status-quo*, the Law 8971, of December 1994, introduced alimony rights to cohabitants fulfilling certain criteria.[24] It worked as an extension of the applicability of the Law 5478/1968 ("Alimony Law") to the dissolution of cohabitations. The Law 8971/1994 established that the cohabitant partners of divorced, legally separated, and widowed individuals could require alimony upon the relationship's dissolution. Alimony requests would be valid if the cohabitation were publicly known and had **endured for five years or more** (waived in the event of **common offspring**). In other words, the partner requesting alimony must prove with the testimony of neighbors, building managers, or renters, the existence of a stable union for more than 5 years. Alternatively, the registration of a child in the name of both partners would be sufficient to waive the duration requirement. The beneficiary has the right until commencing a new (stable) relationship and so long as they can prove **financial necessity** - almost invariably defined by courts in terms of potential earnings (or the individual's stock of human capital).[25] The alimony amounts were established by a judge according with the debtors' financial capabilities (normally 25 to 33% of monthly

[21] This evolution in the jurisprudence was latter consolidated in the form of law (Law 9278/1996) - if no specific contract had been written, cohabitants were allowed 50% of the assets purchased after the start of the relationship.

[22] Some legal texts call attention to specific cases in which the Superior Court of Rio Grande do Sul (the state in the extreme South of Brazil), granted cohabiting spouses compensation for "domestic services", or "palimony", immediately after 1988.

[23] See, for example, the alimony legal request procedure and rulings described in the Suppreme Court's Recurso Especial 36040-RJ - *Joana Darc Andrade versus Joao Rafael da Costa.*

[24] Additionally the law established the right of cohabitants to participate on the partners' inheritance process.

[25] "Alimony is based on necessity, what is not the case for a woman educated and able to work" [author's translation]. See Agravo de Instrumento # 596030, 8a Camara Civil do TJRGS (Superior Court of Rio Grande do Sul), June-27-1996.

income converted into minimum wage units for indexation purposes).

The new law also established a process for the alimony request (which reduces the transaction cost) and strong enforcement mechanisms. First, individuals requesting the ex-partner payment of maintenance could make use of a lawyer assigned by the Judicial System (free of charge). Second, the expected litigation duration was substantially reduced, with the possibility of provisional payments before the final ruling. Third, the law made the information about the legal apparatus surrounding alimony rights publicly available. Finally, enforcement was dramatically strengthened - failure to meet the required monthly payment would result in imprisonment.

As it induced an exogenous shift in the balance of power within households (cohabitant couples), mostly favoring women, the implementation of Law 8971 is explored as an exogenous variation in the empirical exercises below. Moreover, the fact that married couples were not affected by law allows the construction of a natural-experiment.

# 5 Data and econometric identification strategy

## 5.1 The data on cohabitation and marriage

The data set used in this study is from the Brazilian Household Survey (*Pesquisa Nacional de Amostra de Domicilios*-PNAD) undertaken in the years 1992, 1993 and 1995 by the Brazilian Census Bureau (*Instituto Brasileiro de Geografia e Estatistica, IBGE*).[26] These waves of the survey are appropriate for before-after exercises due to the fixed questionnaire design. The sampling scheme is based on a three-level multi-stage procedure - a successive selection of municipalities, census sectors and households. PNAD collects information on household demographic characteristics, income, labor supply and human capital investments. For the selected years, it includes information on cohabitation and formal marriages, which is common to decennial demographic censuses, but only sporadically investigated by household surveys.[27]

Some shortcomings of PNAD are particularly important for this study: a) the absence of data on the duration of marital relationships; b) no data on the official registry of religious marriages, and; c) no information on the biological father of each child in the sample. Without such information, the group of individuals to which the law adoption is immediately effective (those in informal marriages for more than 5 years, with common offspring, or unregistered religious marriages) cannot be recovered. This means that an unknown proportion of the group that is eligible to the treatment ends up not being treated. The empirical identification discussion below is based, therefore, on the "potentially" exposed group - cohabitation and religious marriages - restricting the analysis to an intent-to-treat investigation.

[26]Due to budgetary problems, the PNAD was not conducted in 1994.

[27]See Medeiros & Osorio (2002) and Lazo (2002) for a presentation of the basic structure of the surveys and general outcomes regarding nuptiality and household arrangements.

The exercises below are based on information about households where heads and spouses were older than 15 and younger than 55. Even though there can be married or cohabitant couples living with parents/in-laws, couples that were not on the condition of head/spouse of the household were dropped from the sample. In the case of the outcomes for children, due to age-limits in the availability of schooling and child-labor information (the "investment allocation variables), the sample was further restricted to sons and daughters between ages 5 and 17.[28]

Table 3 presents general statistics for individual characteristics of women and men in the sample. These figures indicate that, compared to their married counterparts, cohabiting female spouses are (on average) younger and less educated. In all the years, cohabitants also have smaller probability of being from the white population, but higher probability of living in a metropolitan area. Total per-capita income is higher amongst married-by-law couples, but non-labor income is slightly higher amongst female cohabitors.[29] In all the cases there is sufficient overlap of characteristics across these groups. In other words, each of the female characteristics (a certain age or education level, for example) appears in both cohabitant and married groups. Finally, notice that the comparison of men's individual characteristics in cohabitation or marriage follow, except for individual non-labor income level, the female case.

Table 4 focuses on general statistics (single variable analysis) for adults' behavioral choices regarding extensive and intensive margin of time-allocation. The outcomes in terms of labor supply (market or housework) for adults are, generally, not dramatically different for cohabitants and married couples. When the changes between 1993 and 1995 are investigated, the impression is that cohabitant women have not changed hours worked, and have increased labor force participation. For formally married women there is a bigger increase in both categories. Except for the increased participation in housekeeping activities, no significant changes were observed amongst males.

Finally, in Table 5 the outcomes and basic characteristics of children in the sample are presented. The figures indicate higher education attainment and higher proportion of school enrollment among children of formally married parents. At the same time, the labor force participation is also higher for them than for the children of cohabiting parents. These gaps may be the result of differences in the average age of parents in each of these arrangements. - indicating the presence of couples on different sections of their life cycle. Although not explored in Table 5, it is important to notice that the patterns of investment in children are naturally staggered, and a comparison of changes in the amount of investment should take that into account. Figure 1 presents the descriptive statistics of human capital investments (flow and stock) for children of both formally and informally married parents, according with their age and gender. Notice that in terms of the investments investigated in this paper (schooling and labor supply), parents face different trade-off's as their

[28]The analysis were reproduced for a subset of the sample (mothers ages 18 to 55, and children ages 7 to 14) without significantly altering the results.

[29]This latter finding in particular indicates that women may take decisions about accumulation of labor income in order to improve their options outside the marital relationship, indicating that the masure of non-labor income can be an imperfect measure of power distribution (i.e., non-exogenous).

children grow old. Between ages 5 and 10 the decisions regarding schooling of children are basically decisions about entry (or timing of entry). After ten years of age, those decisions are more likely about (timing of) exit. Therefore, child labor can have different cost/benefit considerations for those two age-groups - before 10 as a hazard to the child's health and/or schooling quality (high grade repetition, for example), while for a teenager it can also (but not necessarily do) represent acquisition of labor market experience.

### 5.2 Identification strategies

In order to test the compatibility of household decisions with the restrictions of the unitary model, this section discusses identification of the causal impact of alimony law on cohabiting couples' observed behavior. The reasoning is that alimony rights favoring female spouses (in case of dissolution) should correspond to an improvement in their bargaining position , affecting allocation decisions on intact households only in case of heterogeneous preferences. In other words, the estimation of the (theoretical) bargaining power redistribution effects is equivalent, in practice, to the causal inference of the law implementation.

The objective of the estimation strategy is the identification of $\frac{\partial X}{\partial e}$ for the demands of cohabiting couples. Considering the discrete change in the extra-environmental parameters explored in this paper, the estimation of such derivative correspond to the estimation of differences in the levels of $X$ under distinct legal regimes. For the ease of notation, let $e = 0$ in case of no alimony rights for cohabitants and $e = 1$ otherwise. Let $Z$ be the vector with community, household and individual observed characteristics. Therefore, from equation (3):

$$\frac{\partial X}{\partial e} = \frac{\Delta X}{\Delta e} = X\left(Z, e = 1; \varepsilon^h, \varepsilon^c\right) - X\left(Z, e = 0; \varepsilon^h, \varepsilon^c\right) \tag{4}$$

Adopting the notation used in the causal-inference literature, let $X(h, t)$ represent the observed outcome in household $h$ at the time-period $t$. Let $T(h)$ be an indicator function, assuming the unit value in the year after the law implementation (post-treatment period). Define $X_{e=1}(h, t)$ and $X_{e=0}(h, t)$ as the theoretical outcomes of household $h$ if exposed and non-exposed to the treatment, respectively. Consequently, the effect of the treatment in household $h$ can be represented by the difference $X_{e=1}(h, t) - X_{e=0}(h, t)$. In addition, let $D(h)$ be the indicator of membership in the eligible-to-treatment group. The expected impact of the law over households in that group can be theoretically represented by:

$$\tau(Z) = E[X_{e=1}(h, t) - X_{e=0}(h, t) \mid Z(h), D(h) = 1, T(h) = 1]$$

Notice that the most basic issue in this evaluation is a missing data problem: the counter-factual is not available for each household in the sample of eligible couples. In other words, for each individual in the eligible group at $T(h) = 1$, the outcome is observed under the treatment but not under the non-exposition status:

$$X(h,t) = D(h)T(h)\cdot X_{e=1}(h,t) + [1 - D(h)T(h)]\cdot X_{e=0}(h,t) \tag{5}$$

For the eligible group in the post-treatment period it corresponds to $X(h,1) = X_{e=1}(h,1)$. In this way, the effect of the treatment on the treated can be re-written as:

$$\tau(Z) = E\left[X(h,t) \mid Z(h), D(h)=1, T(h)=1\right] - E\left[X_{e=0}(h,t) \mid Z(h), D(h)=1, T(h)=1\right]$$

The identification of such model requires, therefore, a specific modeling of the counter-factual demand function $X_{e=0}(h,t)$ for the eligible group in period $T(h)=1$. While one alternative would be the use of pre-treatment outcomes for the treated group,[30] the relevance of time-trends and the influence of other concurrent events aside from the law adoption may confound the inference. Interestingly, if only part of the observed population is subject to the bargaining power shifter (the treatment), a sample of non-treated observations that are directly affected by the alternative factors can be used to net out effects of any confounding event. This improved strategy can be designed under the following sufficient identification condition:

**Condition 1** *:*

$$\begin{aligned} & \{E[X_{e=0}(h,t) \mid Z(h), D(h)=1, T(h)=1] - E[X_{e=0}(h,t) \mid Z(h), D(h)=1, T(h)=0]\} \\ = \; & \{E[X_{e=0}(h,t) \mid Z(h), D(h)=0, T(h)=1] - E[X_{e=0}(h,t) \mid Z(h), D(h)=0, T(h)=0]\} \end{aligned}$$

*Which states that, conditional on covariates, the treatment and the control groups would have followed parallel paths in the absence of the treatment.*

*Phrased in a slightly different way, modeling of the counter-factual demand has to take into account time-effects. Under this condition, such effects are assumed to be common across eligible and non-eligible groups so that:*

$$\begin{aligned} E[X_{e=0}(h,t) \;\mid\; & Z(h), D(h)=1, T(h)=1] = E[X_{e=0}(h,t) \mid Z(h), D(h)=1, T(h)=0] \\ & + \{E[X_{e=0}(h,t) \mid Z(h), D(h)=0, T(h)=1] - E[X_{e=0}(h,t) \mid Z(h), D(h)=0, T(h)=0]\} \end{aligned}$$

Taking into consideration that all theoretical measures (given Condition 1) can be observed, the causal-inference can be rewritten as a so-called "difference-in-differences parameter" represented by:

$$\begin{aligned} \tau(Z) \;=\; & \{E\left[X(h,t) \mid Z(h), D(h)=1, T(h)=1\right] - E\left[X(h,t) \mid Z(h), D(h)=1, T(h)=0\right]\} \\ & - \{E\left[X(h,t) \mid Z(h), D(h)=0, T(h)=1\right] - E\left[X(h,t) \mid Z(h), D(h)=0, T(h)=0\right]\} \end{aligned} \tag{6}$$

[30]This strategy is explored by Lundberg, Pollack & Wales (1997) and Ward-Batts (2000).

*a) Parametric treatment-effect identification*

In order to simplify the connection of the theory with the data, a parametric version of the outcomes derived from the reduced-form demands discussed in Section 3 can be implemented.[31] Assuming a linear version of the demand on the non-exposition status and under Condition 1:

$$X_{e=0}(h,t) = \alpha_0 + T(h)\alpha_1 + D(h)\alpha_2 + Z(h)\left[\beta_0 + T(h)\beta_1\right] + v(h,t)$$

where the effects of time ($\alpha_0$) and observed covariates ($\beta$) are common to all individuals (eligible or not), the effects of covariates are and where $v(h,t)$ collapses all unobservable characteristics. Considering a constant treatment effect $\tau(Z) \equiv \tau$, the re-arrangement of equation (**??**) yields the following empirical model:

$$X(h,t) = \alpha_0 + T(h)\alpha_1 + D(h)\alpha_2 + Z(h)\left[\beta_0 + T(h)\beta_1\right] + \left[D(h)T(h)\right]\cdot\tau + v(h,t) \tag{7}$$

From this parametric formulation, it is clear that the **sufficient identifying assumption** described in Condition 1 can be represented as:

$$E[v(h,t) \mid T(h), Z(h), D(h)] = 0$$

This condition would allow the consistent estimation of $\tau$ by standard **ordinary least squares** on the pooled cross-sectional samples of the treatment and control populations. In other words, the identification and consistent estimation of the treatment effect are subject to the exogeneity of the group indicator - they require absence of (across-groups) selection bias in response to the law. What one needs is that the difference in the expected effects of unobservables over the outcomes are either constant across groups (over time) or constant over time (within groups).

*b) Parametric differential treatment-effect identification*

There are reasons to believe that selection bias might play a role in the empirical exercise when using repeated cross-sectional samples. In that event, the selection bias generated would be a result of changes (between cross-sectional sample draws) in the composition of unobservable characteristics in either the treatment or the control group due to the effects of the law over marriage markets. In the context

[31] Semi-parametric inference on the difference-in-differences framework is still incipient. See Abadie (2003). See also Athey & Imbens (2002) for discussions on non-linear specifications.

of treatment effect evaluation, selection on those unobservables, which may directly affect the outcomes of interest, hinders the attempt to isolate the effect exclusively attributable to the treatment.

A simple strategy to work around this issue is to focus on differential effects of the treatment. Exploiting a particular characteristic of the data set (the panel structure introduced by the existence of multiple-child households) and **assuming that the unobserved characteristics can be decomposed into child and parental-level additive components** ($v(h,t) = \varepsilon^c(h,t) + \varepsilon^h(h,t)$), siblings' fixed-effects can be used to wash-out household-level unobserved characteristics. In this formulation household-level observed characteristics are also extracted from the model, preventing the direct inference of the treatment-effect (a household-level variable) on the demand function associated with a specific child (individual). The idea is to explore heterogeneity in child characteristics (gender, age and birth order, for example) and in their influence over the effect of the treatment. The existence of heterogeneity (in both outcomes and characteristics) guarantees consistent inference of the *differential treatment-effect* within each household.

In practical terms, let $Z(h)$ be partitioned into parental (household) and child level characteristics, respectively: $\left[\begin{array}{cc} Z^h(h) & Z^c(h) \end{array}\right]$. Re-writing the model in (7) accounting for the influence of child characteristics over trends, over group characteristics and, most importantly, over the treatment ($\tau(Z) \equiv \tau_0 + Z^c(h) \cdot \tau_1$):

$$\begin{aligned} X^c(h,t) \quad = \quad & \alpha_0 + T(h)\alpha_1 + D(h)\alpha_2 + Z^h(h)\left[\beta_0 + T(h)\beta_1\right] + Z^c(h)\left[\gamma_0^c + T(h)\gamma_1^c\right] \\ & + Z^c(h)\alpha_0^c + \left[Z^c(h)T(h)\right]\alpha_1^c + \left[Z^c(h)D(h)\right]\alpha_2^c \\ & + \left[Z^c(h)Z^h(h)\right] \cdot \left[\beta_0^c + T(h)\beta_1^c\right] \\ & + \left[D(h)T(h)\right] \cdot \left[\tau_0 + Z^c(h)\tau_1\right] \\ & + \varepsilon^c(h,t) + \varepsilon^h(h,t) \end{aligned} \tag{8}$$

The elimination of household-specific effects reduces the empirical exercise to the ordinary least squares estimation of:

$$\begin{aligned} X^c(h,t) - \overline{X}(h,t) \quad = \quad & \left[Z^c(h) - \overline{Z}(h)\right]\left[\gamma_0^c + T(h)\gamma_1^c\right] + \left[Z^c(h) - \overline{Z}(h)\right]\alpha_0^c \\ & + \left\{\left[Z^c(h) - \overline{Z}(h)\right]T(h)\right\}\alpha_1^c + \left\{\left[Z^c(h) - \overline{Z}(h)\right]D(h)\right\}\alpha_2^c \\ & + \left\{\left[Z^c(h) - \overline{Z}(h)\right]Z^h(h)\right\} \cdot \left[\beta_0^c + T(h)\beta_1^c\right] \\ & + \left[D(h)T(h)\right] \cdot \left[Z^c(h) - \overline{Z}(h)\right]\tau_1 \\ & + \eta^c(h,t) \end{aligned} \tag{9}$$

where $\eta^c(h,t)$ represents demeaned individual-level unobserved characteristics. From equation (9) the estimation of the partition $\tau_1$ related to observed child characteristics (under the linear-additive assumption for unobserved characteristics) is consistent if:

$$E[\eta^c(h,t) \mid T(h), Z(h), D(h)] = 0$$

or, in other words, if there is no across-group selection (or marriage market effects) related to the characteristics of children.

# 6 Regression results

## 6.1 Adults: "Intent-to-treat" Level-Effects

Estimation of the level-effects of alimony law implementation is based on equation (7). The outcome of interest is the division of labor (housekeeping and market work) between marital partners. The results of the estimation of the difference-in-differences parameter using least-squares regressions are presented in Panel A of Table 6. All regressions include education, age, and non-labor income of both men and women, as well as controls for household demographics and geographic location. Results indicate that working women favored by the extension of the alimony law (cohabitors) were shown to significantly reduce, relative to their married counterparts, total hours supplied to the labor market (3.5%) or hours supplied to their primary job (3%). Additionally, once again relative to married females, cohabitant women perform housekeeping activities less frequently than before (0.5%).[32] Combined, these results indicate that the redistribution of power towards women was associated with increased consumption of leisure.

A couple of concerns related to the strategy above are addressed in Panel B of Table 6: i) the possibility that individuals may have anticipated implementation of the law, adapting their behavior before the administration of the treatment, and; ii) the possibility that imbalances (between control and treatment) in the distribution of covariates that directly influence trends in the outcome variables may have driven the results. Both concerns are investigated using a "placebo" treatment, which consists of the reproduction of the treatment effect calculations as if the law change had occurred between the 1992 and 1993 waves of PNAD, and not in 1994. Since no significant difference between cohabitant and married outcomes were observed before the law implementation, the pre-treatment experiment yields some robustness to the inference on the actual treatment.

The pre-treatment experiment strategy just described does not address another potential problem: shocks occurring between 1993 and 1995 that affected individuals in the treatment group differently than individuals in the control group. The inflation stabilization plan adopted in July of 1994 could be considered a confounding "treatment", for example. It is well known that exchange-rate based stabilizations, as it was the case in Brazil, promote a decrease in the relative prices of tradable with respect to non-tradable

[32]This corresponds to approximately 1/5 of the women **not** housekeeping in 1993.

goods. This suggests that economic sectors (rural and urban), regions, and individuals with particular wealth levels and portfolio compositions, may fare better than others. While the experimental design allows for the possibility that factors aside from the law change influence both the treatment and the control groups, if the imbalance in characteristics across-groups is not correctly controlled, the observed treatment impact may capture both the stabilization and the power redistribution effects.

To verify that covariates included in the model were able to sufficiently control for across-groups differences in observed characteristics, the second column of Panel A in Table 7 re-estimate the model excluding covariates used as controls for characteristics that potentially influence the stabilization effects but (arguably) not the ones originated by the power redistribution experiment. The idea is that if changes in macroeconomic conditions were solely responsible for the estimated effects because individual characteristics were not correctly controlled for due to wrong functional form or omission, the exclusion of additional variables would reinforce the estimated parameters in a particular direction. If, for example, cohabitant and married couples were confined to two geographical locations with different macroeconomic prospects (the former in a recession region), the reduction in worked hours could be exclusively reflecting the contraction of labor demand and not the alimony law implementation (a federal law change affecting both regions). The indicators of geographic location are intended to control for such factors and, if that is the case, their exclusion would drive up the misleading negative effect over worked hours. A similar reasoning applies to time-varying effects of assets' ownership (interaction of the wealth variables and the year dummy). The results indicate that exclusion of controls does not significantly affect the estimated parameters.

In a second attempt to address the same issue, this paper proposes an alternative strategy. Since the objective is to disentangle effects of the treatment (alimony law) from the concurrent stabilization shock, the paper explores a "pseudo" treatment experiment. The strategy is based on the fact that the control group (married couples) can be subdivided into categories that differ in terms of most individual characteristics in such a way that approximately reproduces differences between cohabitant and married couples. Taking married-by-law as the "pseudo-treatment" group and the married-by-law & religion as the "pseudo-control" group, the exercise calculates difference-in-differences parameters for this unreal experiment. The results presented in Table 7 (Panel B) corroborate the existence of treatment effects that are exclusive to the cohabitant couples (in particular, female outcomes), reinforcing the idea that the alimony law itself impacted couples' behavior.

Another important consideration is that inframarginal labor-supply changes could be reflecting the fact that women and men have, on average, attachment to the labor force regulated by different types of contracts. Formal work is more common among men, while women have higher proportion of informal participation (possibly due to its flexibility). It would be specially troublesome if those types of contractual arrangements were unevenly distributed across informally and formally married individuals. Table 8 reproduces results of actual and "placebo" treatments stratifying by "stayers" (individuals that did not change job in the one year period preceding the survey) and by their degree of contractual formality. The results confirm previous findings, indicating that the burden of market work was relatively shifted towards male cohabitors. In a sample where both men and women are either self-employed or informal wage-workers, cohabiting women, relative to their married counterparts, reduced hours supplied to the market by 7.4%, while the ratio of hours worked by male and female partners increased 5.5%.[33]

Importantly, models that investigate the level effects of alimony law implementation are potentially subject to selection biases originated by changes in the marriage market. In other words, the adoption of the law may have altered the attractiveness of cohabitation for both men (negatively) and women (positively), possibly changing the composition of the treatment and control groups between 1993 and 1995.[34] This is the case when one considers the effect of divorce costs on the likelihood of marriage formation/dissolution, for example.[35] Similar reasoning would suggest compositional effects resulting from the alimony law adoption. The problem emerges if the composition of the groups change in terms of variables that are unobservable (from the econometrician's perspective) but intrinsically connected to the outcomes of interest (ability or "comparative advantage", in market work or in childrearing, are good examples).

Even though the focus on differential effects amongst children proposed in the next subsection works around this concern, a simpler alternative can be explored on the investigation of outcomes for adults. It consists on considering that selection is expected to be less of an issue for couples living together for 5 years or more. Although information on duration of relationships is not available, results presented in Table 9 assume that couples with children older than 5 would be living together for that period or longer. Table 9 also explores the fact that Brazilian courts have associated the requirements of the law to the female educational attainment - the sample was divided into less and more educated women (low and high potential earnings).[36] The less educated women ought to closely correspond to the stratum of the cohabitant population that was (or had higher probability of being) subject to the treatment, while the more educated would not be affected. The estimated effects were significant for exactly the sub-group for which selection was expected to be unimportant and the law impact expected to be stronger. Less educated cohabiting women (in households with all children older than 5) reduced labor supply by 9.3% relative to their married counterparts. No significant impact emerged from the sample of more educated women. The results corroborate the previous indications of inframarginal labor-supply effects attached to the law implementation.

Before investigating child outcomes it is interesting to note that the selection itself, if relevant, is also a result of the law implementation. In this sense, it may reflect the fact that matching decisions are

[33]See Tiefenthaler (1999) for additional discussion on labor contracts and testing of unitary model restrictions.

[34]In fact, the 1993-1995 growth in the average age of both men and women cohabiting was significantly smaller than for married couples, suggesting some differential compositional change.

[35]See Bougheas & Geogellis (1999) for a theoretical model, and Peters (1986) and Frieberg (1998) for the US-based empirical evidence.

[36]The trends in educational level observed between 1993 and 1995 were the same for both informal and formally married individuals, assuring no differential compositional change in terms of education. Moreover, in regressions relating the probability of cohabitation to high and low educational level, no education-group differential effects were found significant.

also related to relationship security and/or power distribution within the household. Interpreting the results with this perspective would require, however, a better knowledge of how the concurrent stabilization shock influenced marriage/cohabitation decisions. Moreover, the cross-sectional data used in this paper (without retrospective history of marital relationships) is not appropriate to the investigation of this specific topic.

### 6.2 Children: "Intent-to-Treat" Differential-Effects

The analysis in this subsection exploits within family variation in school enrollment, and its connection to the alimony law implementation. This outcome can be considered as the core of a broader decision regarding the allocation of resources towards investments in the human capital of children. Investigation of outcomes amongst children of married and cohabiting couples is not complete, however, without the investigation of differential effects. The idea is that the pattern of investments in the schooling and labor market activities of children evolves differently depending on child gender, age and/or birth-order. Without allowing for these patterns, the aggregation can cancel out heterogeneous effects.

In the empirical specification investigated in this subsection, the assumption underlying the included covariates in equation (8) is that the intent-to-treat effects can be parametrized as a function of gender (male child) and birth-order (oldest child in the household) indicators. The effects are also interacted with sibship gender-composition dummies. The main objective of the empirical exercises is, therefore, to uncover the gender-specific birth-order effects. All regressions estimated in this section include controls for additional child-level characteristics, namely age (second-order polynomial) and years of education attainment splines. The controls for parental and household characteristics are the same as the ones used in the calculation of outcomes for adults (age, education, income, assets and geographic location). Finally, the sample is restricted to children between 5 and 17, families with no children outside the household, and to households with at least two children within that age range.

Table 10 reports the results of the ordinary least-squares regressions on both the actual and the pre-treatment experiments. Significant results of the comparisons between cohabitant and married couples behavior were observed in the schooling investments on girls and in the differential effects with respect to investments in first-born girls. The results were found significant for the households with mixed gender-composition. In other words, the effects were observed when a brother was present. In order to assess potential selection (on unobservables) problems, Table 11 reproduces results of ordinary-least squares and household fixed-effects estimations.

The results in Table 11 indicate that selection does not alter the main conclusion with respect to differential birth-order effects amongst girls. The idea that resources were disproportionately allocated towards schooling of first-born girls is reinforced. For this sample of cohabiting couples, which includes the ones that were not actually treated (but that cannot be identified), the intent-to-treat effects over children indicate an increase of 9% in education investments for first-born daughters (relative to latter born girls). On the other hand, first-born boys were not treated differently from younger boys in the school enrollment context, even though weak evidence of gender-differentiated investments among latter born children was observed. These results corroborate the findings against the restrictions of the unitary model for the labor supply decisions of adult cohabitors presented in the previous subsection.

Table 12 investigates if the results above hold for different child-age stratifications of the data. In Panel A the differential effects of the law over children of cohabiting couples (relative to children of married couples) are calculated for children 11 and younger, while in Panel B the same calculation is performed for children ages 11 and older. The results indicate that the reallocation of resources towards first-born girls is pertinent both for the younger and older groups. Once again, strongly significant results indicate that, in families of informally married parents, first-born girls were treated differently from their younger sisters, but first-born boys did not receive any differential schooling investment relative to their younger brothers.[37]

Table 13 explores characteristics of the law requirements that identify a stratum of the cohabitant mothers' population that was (or had higher probability of being) subject to the treatment. As in the case of outcomes for adults, the idea is to stratify the sample into less and more educated women (low and high potential earnings), in an attempt to extract the most out of the law's "necessity clause". The results in Panel A indicate striking differences going in the expected direction: bigger effects among women with less then elementary school, and no significant effect among the more educated ones. First-born daughters of less educated cohabiting mothers were 16.4% more likely to be enrolled in school than younger girls, relative to the differential effects amongst daughters of married couples. No significant pattern of changes emerges from the pre-treatment experiment in Panel B.

Table 14 calculates the significance of the difference between the strata of less and more educated mothers. The effects are significant for the actual treatment, being also significantly different from the education differential in the pre-treatment experiment. Compared with their married counterparts, first-born girls of cohabiting couples with less educated women receive disproportionately more schooling investment (relative to their younger sisters) than in cohabiting couples with more educated women (13.4%). This differential is also 22.6% bigger when the law is implemented than in the pre-treatment years. The results reconfirm the existence of treatment effects of the alimony law, with the shift in power having birth-order and gender differential effects for investments in children within households headed by cohabiting couples.

[37]Note that the differential effects are stronger in the stratified samples than in the general one, suggesting that sibship density may have an influence in the outcomes. See Powell & Steelman (1990) for a discussion on sibling density.

# 7 Conclusions

This paper investigates the effect of a shift in the balance of decision power within households. Using an exogenous source of variation provided by the adoption of a law (extension of alimony rights to cohabitants), and the similar "family status" of cohabitant and married couples, this paper has provided empirical evidence that (intra-household) empowerment of women resulted in: i) reduction in hours worked by female heads, including reduced housekeeping activities, and; ii) redistribution of household resources towards schooling of first-born girls. Consistent with the theory and aspects of the law, these effects were shown to be stronger when considering a stratum of households where the mother is likely to be more dependent on alimony in case of relationship dissolution.

The time-allocation results confirm previous findings (based on US data sets) reported in the literature. The contraction in female labor supply following increases in women's bargaining power was also observed in the study by Rubalcava & Thomas (2000), while the reduction in hours dedicated to home-production were shown pertinent by Gray (1998). The combination of both results suggests that, for Brazilian women in consensual unions, bargaining power upgrades were associated with increased consumption of leisure.

The differential pattern of investment in the human capital of children within families has received attention of many areas of scientific inquiry. In economics, traditional contributions are based on the quantity-quality trade-off of fertility and investments in children.[38] Differential investments within the household were also object of investigation, with explanations to systematic gender bias in the allocation of resources revolving around either arguments of differential costs/returns to investments in girls and boys, as in Rosenzweig & Schultz (1982),[39] or arguments based on the possibility that parents exhibit "unequal concern" for identical outcomes of sons and daughters.[40] Thomas (1994), challenging assumptions of the unitary model, explores the idea that not only parents are differently concerned with outcomes of sons and daughters, but also that mothers and fathers have heterogeneous preferences regarding their offsprings. His empirical evidence for families in Brazil, Ghana and the United States indicates that maternal resources impact more strongly the accumulation of human capital of their daughters, while paternal resources have a bigger impact on sons'. With parental resources measured by differences in education and in non-labor income, the author suggests that the indication is that parents have different preferences regarding investments in their children.

Birth-order differentials, on the other hand, have received more attention in the sociology and psychology literatures.[41] Psychologists have given special attention to the association between birth-order and aspects of a child's personality and intellectual development.[42] Two studies are of special interest to the evidence presented here. Hetherington et al. (1978) suggests that boys are more affected than girls by

[38]See Becker & Lewis (1973). See also Butcher & Case (1994).
[39]The same reasoning is also developed in the evolutionary biology literature. See Trivers (1972), for example.
[40]See Behrman et al. (1986).
[41]See Parish & Willis (1993), Hauser & Kuo (1998), Powell & Steelman (1990) and references therein.
[42]See review sections in Hertwig et al. (2002), and Sulloway (2001), for example.

an unstable relationship between their parents (probability of divorce), therefore, it can be the case that male children of cohabiting couples are in an unfavorable environment relative to their sisters. Second, findings presented in Mekos et al. (1996) indicates differential treatments are strongly associated to behavioral problems in remarried families, in particular among teenagers living with non-biological fathers. Combined, these findings indicate that teenage boys would be receiving less investments due to their slower development (grade repetition, for example) and to the non-biological ties with their mothers' partner.

Due to the natural experiment formulation adopted here, explanations of differential treatment to boys and girls related to time-changes in the cost or return to the investments are ruled out because they should influence both treatment and control groups. Moreover, the measurement of bargaining power based on an exogenous change in the institutional environment bypass potential criticisms regarding the use of education (may reflect childrearing technology) and non-labor income (may be endogenous) as proxies for "power". Regarding the findings in the psychology literature, unless unobserved child characteristics actually affect the marriage-market decisions (selection of cohabitation *versus* marriage), the experimental design also rules out explanations based on channels related to personality-formation differences between boys and girls living with cohabiting or married "parents". In particular, time differences are expected to eliminate those characteristics that are expected to be constant within each of those marital arrangements.[43]

The evidence that more familial resources were allocated to the schooling of first-born and teenage girls presented in the current study is also compatible with Brazilian social norms related to the "formulation" of intergenerational contracts. As briefly discussed in Goldani (1999), the indication is that, in most cases, daughters are the ones that provide functional and financial help to their senior parents. As women live longer than men, and are expected to live longer spells in widowhood, it is intuitive that the investment in "care takers" is more attractive from the mother's than from the father's perspective. Alternatively, the fact that the effects are observed between sisters that have a male sibling indicates that preferences of mothers and fathers may differ in what each one perceives as an "equitable distribution" of investments amongst daughters and sons, but not amongst a same-sex sibship.

In conclusion, changes in the allocation of time or of investments in children's human capital in response to outside-option attractiveness are not predicted by models that treat the household as an unit. The results of this study yield insight into "the nature of parental preferences" which underlie the within-household allocation of resources. They *strongly suggest* that the unitary formulation of the household is an imprecise description of the familial decision process regarding the allocation of time and schooling investments. In essence, they reveal that such allocations are outcomes of an elaborate process, where observed decisions result from bargaining and negotiations amongst parents with different preferences and varying ability to assert their "vision of the world" at the household level.

[43] In other words, the evidence of no differential effects in the pre-treatment period invalidates the vision that differences in the investment in boys and girls are intrinsic to informal marriages.

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## Appendix A - Brazilian Federal Constitution of 1988

CHAPTER VII

FAMILY, CHILDREN, ADOLESCENTS AND THE ELDERLY

Article 226. The family, which is the foundation of society, shall enjoy special protection from the State.

Paragraph 1 - Marriage is civil and the marriage ceremony is free of charge.

Paragraph 2 - Religious marriage has civil effects, in accordance with the law.

Paragraph 3 - For purposes of protection by the State, the stable union between a man and a woman is recognized as a family entity, and the law shall facilitate the conversion of such entity into marriage.

Paragraph 4 - The community formed by either parent and their descendants is also considered as a family entity.

Paragraph 5 - The rights and the duties of marital society shall be exercised

equally by the man and the woman.

Paragraph 6 - Civil marriage may be dissolved by divorce, after prior legal separation for more than one year in the cases set forth by law, or after two years of proven de facto separation.

Paragraph 7 - Based on the principles of human dignity and responsible parenthood, family planning is a free choice of the couple, it being within the competence of the State to provide educational and scientific resources for the exercise of this right, any coercion by official or private agencies being forbidden.

Paragraph 8 - The State shall ensure assistance to the family in the person of each of its members, creating mechanisms to suppress violence within the family.

## Appendix B - Tables and Figures

Attached

**Table 1: Marital Arrangements (all ages, as % of all marital relationships)**

| | 1960 [1] | 1970 [1] | 1976 [2] | 1980 [1] | 1984 [2] | 1991 [2] | 1995 [3] | 2000 [3] |
|---|---|---|---|---|---|---|---|---|
| **Formal Relationships** | 73.4 | 78.6 | 82.9 | 81.2 | 83.5 | 76.5 | 72.1 | 67.3 |
| **Informal Relationships** | 26.6 | 21.4 | 17.1 | 18.8 | 16.6 | 23.6 | 28.0 | 32.6 |

Sources: [1] Goldani & Wong (1980), [2] Lazo (1994), and [3] IBGE, PNAD and Censo Demografico, respectively

Formal = Married by law and by law & religion. Informal = Cohabitation and married by religion

**Table 2: Marital Arrangements by Cohort of Birth, women ages 15 to 55**

| | All Women | | | All Women in a Realtionship | | |
|---|---|---|---|---|---|---|
| | **1992** | **1993** | **1995** | **1992** | **1993** | **1995** |
| ***Cohort Born in the 1970's*** | | | | | | |
| Cohabitation | 9.7 | 11.1 | 14.0 | 44.9 | 46.4 | 48.6 |
| Marriage (religion) | 1.0 | 1.1 | 1.2 | 4.6 | 4.6 | 4.2 |
| Marriage (law+religion) | 6.4 | 7.0 | 8.4 | 29.6 | 29.3 | 29.2 |
| Marriage (law) | 4.5 | 4.7 | 5.2 | 20.8 | 19.7 | 18.1 |
| Single | 75.1 | 72.3 | 66.8 | | | |
| Divorced | 3.1 | 3.6 | 4.2 | | | |
| Widowed | 0.1 | 0.1 | 0.1 | | | |
| *Observations* | *21,022* | *24,246* | *31,102* | *4,541* | *5,795* | *8,957* |
| ***Cohort Born in the 1960's*** | | | | | | |
| Cohabitation | 17.8 | 18.8 | 21.2 | 27.4 | 27.6 | 29.9 |
| Marriage (religion) | 2.2 | 2.2 | 2.4 | 3.4 | 3.2 | 3.4 |
| Marriage (law+religion) | 31.0 | 32.3 | 33.0 | 47.7 | 47.4 | 46.6 |
| Marriage (law) | 14.0 | 14.8 | 14.2 | 21.5 | 21.7 | 20.1 |
| Single | 26.2 | 23.0 | 18.7 | | | |
| Divorced | 8.1 | 8.2 | 9.6 | | | |
| Widowed | 0.8 | 0.8 | 1.0 | | | |
| *Observations* | *27,277* | *27,026* | *27,525* | *17,730* | *18,405* | *19,488* |
| ***Cohort Born in the 1950's*** | | | | | | |
| Cohabitation | 15.2 | 15.4 | 16.1 | 19.9 | 20.2 | 21.3 |
| Marriage (religion) | 2.5 | 2.3 | 2.5 | 3.3 | 3.0 | 3.3 |
| Marriage (law+religion) | 43.3 | 43.7 | 42.6 | 56.7 | 57.2 | 56.4 |
| Marriage (law) | 15.3 | 15.0 | 14.3 | 20.1 | 19.6 | 18.9 |
| Single | 10.2 | 9.3 | 8.5 | | | |
| Divorced | 10.9 | 11.3 | 12.4 | | | |
| Widowed | 2.6 | 3.0 | 3.5 | | | |
| *Observations* | *21,998* | *21,935* | *22,118* | *16,784* | *16,758* | *16,699* |
| ***Cohort Born in the 1940's*** | | | | | | |
| Cohabitation | 11.1 | 10.4 | 11.1 | 15.1 | 14.5 | 15.9 |
| Marriage (religion) | 3.6 | 3.5 | 3.1 | 4.9 | 4.9 | 4.4 |
| Marriage (law+religion) | 46.9 | 46.2 | 44.3 | 64.0 | 64.6 | 63.4 |
| Marriage (law) | 11.7 | 11.4 | 11.4 | 16.0 | 15.9 | 16.3 |
| Single | 6.3 | 6.7 | 6.4 | | | |
| Divorced | 12.2 | 13.0 | 13.3 | | | |
| Widowed | 8.3 | 8.7 | 10.4 | | | |
| *Observations* | *14,611* | *14,625* | *14,594* | *10,710* | *10,457* | *10,202* |
| ***Cohort Born in the 1930's*** | | | | | | |
| Cohabitation | 7.5 | 7.8 | 9.1 | 11.6 | 11.9 | 14.1 |
| Marriage (religion) | 4.0 | 4.6 | 4.5 | 6.2 | 7.0 | 7.0 |
| Marriage (law+religion) | 43.2 | 44.8 | 43.0 | 66.6 | 68.2 | 66.7 |
| Marriage (law) | 10.2 | 8.5 | 7.9 | 15.7 | 12.9 | 12.2 |
| Single | 6.1 | 6.6 | 5.8 | | | |
| Divorced | 12.1 | 11.7 | 11.8 | | | |
| Widowed | 16.8 | 15.9 | 17.8 | | | |
| *Observations* | *4,556* | *3,619* | *1,367* | *2,957* | *2,378* | *882* |

Source: Pesquisa Nacional de Amostra de Domicilios, PNAD (1992-1995)

**Table 3: Descriptive Statistics - Men and Women ages 15 to 55 in a marital relationship**

| | Informal Relationships | | | Formal Relationships | | |
|---|---|---|---|---|---|---|
| | **1992** | **1993** | **1995** | **1992** | **1993** | **1995** |
| *Individual & Matching Characteristics* | | | | | | |
| **Male Age** | 34.73 | 34.63 | 34.74 | 37.88 | 38.17 | 38.67 |
| | *(0.09)* | *(0.08)* | *(0.08)* | *(0.05)* | *(0.05)* | *(0.05)* |
| **Female Age** | 31.15 | 31.04 | 31.30 | 34.54 | 34.81 | 35.42 |
| | *(0.08)* | *(0.08)* | *(0.08)* | *(0.05)* | *(0.05)* | *(0.05)* |
| **Male-Female Age Gap** | 3.59 | 3.58 | 3.44 | 3.34 | 3.36 | 3.25 |
| | *(0.07)* | *(0.07)* | *(0.06)* | *(0.03)* | *(0.03)* | *(0.03)* |
| **Male Education** | 4.59 | 4.74 | 4.81 | 5.96 | 6.09 | 6.21 |
| | *(0.04)* | *(0.04)* | *(0.03)* | *(0.02)* | *(0.02)* | *(0.02)* |
| **Female Education** | 4.59 | 4.75 | 4.94 | 6.12 | 6.24 | 6.44 |
| | *(0.04)* | *(0.03)* | *(0.03)* | *(0.02)* | *(0.02)* | *(0.02)* |
| **Male-Female Educ. Gap** | 0.00 | -0.02 | -0.13 | -0.15 | -0.15 | -0.24 |
| | *(0.03)* | *(0.03)* | *(0.03)* | *(0.02)* | *(0.02)* | *(0.02)* |
| **Male White** | 40.40 | 40.90 | 41.40 | 59.20 | 58.90 | 58.80 |
| | *(0.50)* | *(0.40)* | *(0.40)* | *(0.30)* | *(0.30)* | *(0.30)* |
| **Female White** | 42.10 | 42.90 | 42.40 | 61.70 | 61.70 | 61.10 |
| | *(0.50)* | *(0.40)* | *(0.40)* | *(0.30)* | *(0.30)* | *(0.30)* |
| **Interracial Couple** | 29.80 | 28.40 | 26.80 | 22.70 | 22.20 | 21.30 |
| | *(0.40)* | *(0.40)* | *(0.40)* | *(0.20)* | *(0.20)* | *(0.20)* |
| *Household Geographic Location* | | | | | | |
| **Urban Sector** | 81.60 | 81.40 | 82.10 | 81.80 | 82.20 | 82.30 |
| | *(0.40)* | *(0.40)* | *(0.30)* | *(0.20)* | *(0.20)* | *(0.20)* |
| **Metropolitan Sector** | 42.10 | 43.30 | 41.90 | 38.90 | 38.90 | 38.40 |
| | *(0.50)* | *(0.50)* | *(0.40)* | *(0.30)* | *(0.30)* | *(0.30)* |
| *Household and Individual Wealth* | | | | | | |
| **House Ownership** | 54.50 | 53.90 | 55.80 | 62.00 | 62.80 | 64.50 |
| | *(0.50)* | *(0.50)* | *(0.40)* | *(0.30)* | *(0.30)* | *(0.30)* |
| **HH per capita Income** | 248.81 | 265.29 | 291.95 | 361.26 | 397.79 | 418.61 |
| | *(3.67)* | *(4.71)* | *(4.77)* | *(3.47)* | *(4.30)* | *(3.54)* |
| **HH Non-Labor Income** | 45.05 | 47.38 | 49.35 | 72.95 | 88.44 | 84.50 |
| | *(2.33)* | *(3.26)* | *(2.74)* | *(2.54)* | *(2.86)* | *(2.38)* |
| **Male Non-Labor Income** | 21.62 | 27.31 | 25.49 | 52.51 | 61.71 | 62.17 |
| | *(1.73)* | *(2.86)* | *(2.25)* | *(2.21)* | *(2.38)* | *(1.91)* |
| **Fem. Non-Labor Income** | 16.78 | 15.39 | 17.62 | 11.28 | 14.86 | 13.46 |
| | *(1.11)* | *(1.16)* | *(1.12)* | *(0.70)* | *(1.11)* | *(0.83)* |
| *Observations* | *11,633* | *12,109* | *14,410* | *33,156* | *33,241* | *33,204* |

Notes: Standard-errors in (parentheses) and t-stats in [brackets]. Income in real values of Sep-1999 (deflators are region-specific).

Source: PNAD (1992-1995).

## Table 4: Descriptive Statistics for Time Allocation - Men and Women ages 15 to 55 in a marital relationship

| | Informal Relationships | | | Formal Relationships | | |
|---|---|---|---|---|---|---|
| | **1992** | **1993** | **1995** | **1992** | **1993** | **1995** |
| *General Sample* | | | | | | |
| **Fem. Housekeeping** | 96.97 | 97.95 | 97.67 | 97.12 | 97.81 | 98.02 |
| | *(0.16)* | *(0.13)* | *(0.13)* | *(0.09)* | *(0.08)* | *(0.08)* |
| **Male Housekeeping** | 39.59 | 46.50 | 51.34 | 39.81 | 45.80 | 50.52 |
| | *(0.45)* | *(0.45)* | *(0.42)* | *(0.27)* | *(0.27)* | *(0.27)* |
| **Female in Labor Force** | 51.94 | 51.94 | 54.30 | 53.79 | 54.28 | 57.53 |
| | *(0.46)* | *(0.45)* | *(0.42)* | *(0.27)* | *(0.27)* | *(0.27)* |
| **Male in Labor Force** | 97.08 | 96.92 | 96.91 | 96.48 | 96.43 | 96.25 |
| | *(0.16)* | *(0.16)* | *(0.14)* | *(0.10)* | *(0.10)* | *(0.10)* |
| *Working Couples Sample* | | | | | | |
| **Fem. Hours Worked (all jobs)** | 33.43 | 33.40 | 33.39 | 31.53 | 31.65 | 32.31 |
| | *(0.23)* | *(0.23)* | *(0.21)* | *(0.13)* | *(0.13)* | *(0.12)* |
| **Male Hours Worked (all jobs)** | 46.69 | 45.93 | 46.08 | 46.83 | 46.34 | 46.53 |
| | *(0.18)* | *(0.17)* | *(0.16)* | *(0.10)* | *(0.10)* | *(0.10)* |
| **Female Real Hourly Wage** | 1.69 | 1.74 | 1.96 | 2.28 | 2.58 | 2.73 |
| | *(0.04)* | *(0.05)* | *(0.05)* | *(0.03)* | *(0.06)* | *(0.04)* |
| **Male Real Hourly Wage** | 2.60 | 2.90 | 3.02 | 3.96 | 4.50 | 4.63 |
| | *(0.06)* | *(0.11)* | *(0.07)* | *(0.06)* | *(0.11)* | *(0.06)* |
| *All Observations* | *11,633* | *12,109* | *14,410* | *33,156* | *33,241* | *33,204* |
| *Working Couples Observations* | *5,051* | *5,295* | *6,656* | *15,859* | *16,248* | *17,132* |

Notes: Standard-errors in (parentheses). Wages in real values of Sep-1999 (deflators are region-specific).

Source: PNAD (1992-1995).

**Table 5: Descriptive Statistics - Childen ages 5 to 17 in couple-headed households**

| | Informal Relationships | | | Formal Relationships | | |
|---|---|---|---|---|---|---|
| | **1992** | **1993** | **1995** | **1992** | **1993** | **1995** |
| **Male** | 51.40 | 50.70 | 51.17 | 51.27 | 51.11 | 51.11 |
| | *(0.42)* | *(0.41)* | *(0.38)* | *(0.23)* | *(0.23)* | *(0.23)* |
| **Education** | 1.53 | 1.57 | 1.73 | 2.52 | 2.61 | 2.84 |
| | *(0.02)* | *(0.02)* | *(0.02)* | *(0.01)* | *(0.01)* | *(0.01)* |
| **Age** | 10.00 | 9.98 | 10.08 | 10.65 | 10.73 | 10.93 |
| | *(0.03)* | *(0.03)* | *(0.03)* | *(0.02)* | *(0.02)* | *(0.02)* |
| **School Enrollment Rate** | 72.17 | 74.75 | 77.71 | 81.71 | 83.62 | 86.58 |
| | *(0.38)* | *(0.36)* | *(0.32)* | *(0.17)* | *(0.17)* | *(0.16)* |
| **Labor Force Participation** | 18.79 | 18.40 | 18.18 | 21.01 | 20.40 | 20.76 |
| | *(0.33)* | *(0.32)* | *(0.29)* | *(0.18)* | *(0.18)* | *(0.19)* |
| *Observations* | *14,011* | *14,520* | *17,119* | *49,201* | *49,218* | *47,747* |

Notes: Standard-errors in (parentheses).

Source: PNAD (1992-1995).

**Table 6: "Intent-to-Treat" Level-Effects of the Alimony Law, Time Allocation of Adults (OLS Estimations)**

| | PANEL A: Actual Treatment (1995-1993) | PANEL B: Pre-Treatment (1993-1992) |
|---|---|---|
| *Outcomes* | **All Controls** | **All Controls** |
| **Fem. Housekeeping** | -0.519 | 0.202 |
| | *[2.39]** | *[0.83]* |
| **Male Housekeeping** | 0.524 | 0.291 |
| | *[0.72]* | *[0.39]* |
| **Fem. In Labor Force** | -0.137 | -0.685 |
| | *[0.19]* | *[0.93]* |
| **Male In Labor Force** | 0.157 | -0.246 |
| | *[0.64]* | *[0.97]* |
| **Both In Labor Force** | 0.233 | -0.836 |
| | *[0.33]* | *[1.14]* |
| *Sample restricted to working couples* | | |
| **Fem. Hours Worked - Prim. job (log)** | -3.042 | -1.000 |
| | *[2.11]** | *[0.66]* |
| **Male Hours Worked - Prim. job (log)** | -0.984 | -0.078 |
| | *[1.39]* | *[0.11]* |
| **Fem. Hours Worked - All jobs (log)** | -3.485 | -0.686 |
| | *[2.28]** | *[0.43]* |
| **Male Hours Worked - All jobs (log)** | -0.860 | -0.313 |
| | *[1.19]* | *[0.42]* |
| **Male-Female Diff. In Worked Hours** | 0.402 | -0.129 |
| | *[1.01]* | *[0.30]* |
| **Male-Female Diff. In Tot. Worked Hours** | 0.433 | -0.197 |
| | *[1.08]* | *[0.46]* |
| **Male-Female Ratio of Worked Hours (log)** | 2.049 | 0.952 |
| | *[1.33]* | *[0.60]* |
| *Observations* | *92,940* | *90,107* |
| *Observation Working Couples* | *45,311* | *42,425* |

Notes: a) Coefficients are reported after multiplied by 100 (except for hours differences). t-statistics in [brackets] are calculated with jacknifed standard-errors.

b) "All controls" includes education splines (male and female), age polynomial (male and female), household demographics, non-labor income and assets (house ownership) for male and female, and geography indicators. The income, assets and geography effects are allowed to vary across years.

* Classical significance (p-value> 5%)

**Table 7: "Intent-to-Treat" Level-Effects of the Alimony Law, Time Allocation of Adults (OLS Estimations)**

| | PANEL A: Actual Treatment (1995-1993) | | PANEL B: Pre-Treatment (1993-1992) |
|---|---|---|---|
| *Outcomes* | **All Controls** | **No Geog. and Asset Controls** | **All Controls** |
| **Fem. Housekeeping** | -0.519 | -0.535 | 0.026 |
| | *[2.39]** | *[2.54]** | *[0.10]* |
| **Male Housekeeping** | 0.524 | 0.055 | -0.981 |
| | *[0.72]* | *[0.08]* | *[1.12]* |
| **Fem. In Labor Force** | -0.137 | -0.529 | -0.913 |
| | *[0.19]* | *[0.71]* | *[1.07]* |
| **Male In Labor Force** | 0.157 | -0.811 | 0.260 |
| | *[0.64]* | *[0.33]* | *[0.85]* |
| **Both In Labor Force** | 0.233 | 0.042 | -0.428 |
| | *[0.33]* | *[0.06]* | *[0.50]* |
| *Sample restricted to working couples* | | | |
| **Fem. Hours Worked - Prim. job (log)** | -3.042 | -2.789 | -0.584 |
| | *[2.11]** | *[1.92]* | *[0.34]* |
| **Male Hours Worked - Prim. job (log)** | -0.984 | -0.614 | -1.461 |
| | *[1.39]* | *[0.88]* | *[1.75]* |
| **Fem. Hours Worked - All jobs (log)** | -3.485 | -3.224 | -0.048 |
| | *[2.28]** | *[2.09]** | *[0.03]* |
| **Male Hours Worked - All jobs (log)** | -0.860 | -0.552 | -1.597 |
| | *[1.19]* | *[0.78]* | *[1.84]* |
| **Male-Female Diff. In Worked Hours** | 0.402 | 0.441 | -0.446 |
| | *[1.01]* | *[1.10]* | *[0.93]* |
| **Male-Female Diff. In Tot. Worked Hours** | 0.433 | 0.470 | -0.513 |
| | *[1.08]* | *[1.16]* | *[1.06]* |
| **Male-Female Ratio of Worked Hours (log)** | 2.049 | 2.159 | -0.897 |
| | *[1.33]* | *[1.39]* | *[0.49]* |
| *Observations* | *92,940* | | *66,445* |
| *Observation Working Couples* | *45,311* | | *33,556* |

Notes: Coefficients are reported after multiplied by 100 (except for hours differences). t-statistics in [brackets] are calculated with jacknifed standard-errors.

* Classical significance (p-value> 5%)

**Table 8: "Intent-to-Treat" Level-Effects of the Alimony Law, Time Allocation of Working Couples (OLS Estimations)**

| | PANEL A: Actual Treatment | | | PANEL B: Pre-Treatment | | |
|---|---|---|---|---|---|---|
| | **Working Couples** | **Working Couples** | **Working Couples** | **Working Couples** | **Working Couples** | **Working Couples** |
| *Outcomes* | | **same job** | **same informal job** | | **same job** | **same informal job** |
| **Fem. Hours Worked - Prim. job (log)** | -3.042 | -5.239 | -7.371 | -1.000 | 0.588 | -0.040 |
| | *[2.11]** | *[3.03]** | *[3.02]** | *[0.66]* | *[0.33]* | *[0.16]* |
| **Male Hours Worked - Prim. job (log)** | -0.984 | -1.402 | -1.829 | -0.078 | 0.028 | -0.059 |
| | *[1.39]* | *[1.62]* | *[1.46]* | *[0.11]* | *[0.03]* | *[0.05]* |
| **Male-Female Diff. In Worked Hours** | 0.402 | 0.748 | 1.199 | -0.129 | -0.379 | -0.467 |
| | *[1.01]* | *[1.55]* | *[1.84]* | *[0.30]* | *[0.74]* | *[0.67]* |
| **Male-Female Diff. In Tot. Worked Hours** | 0.433 | 0.875 | 1.355 | -0.197 | -0.518 | -0.647 |
| | *[1.08]* | *[1.80]* | *[2.05]** | *[0.46]* | *[1.00]* | *[0.92]* |
| **Male-Female Ratio of Worked Hours (log)** | 2.049 | 3.828 | 5.531 | 0.952 | -0.526 | 0.383 |
| | *[1.33]* | *[2.08]** | *[2.16]** | *[0.60]* | *[0.28]* | *[0.15]* |
| *Observations* | *45,311* | *31,110* | *18,001* | *42,425* | *29,155* | *16,519* |

Notes: See Notes for Table 6. "Same job" refers to individuals not changing jobs in a 12 months period.

* Classical significance (p-value> 5%).

**Table 9: "Intent-to-Treat" Level-Effects of the Alimony Law, Time Allocation of Working Couples, by Female Education and Family Composition**

| | PANEL A: Less Educated Women (less than 4 years) | | | | PANEL B: More Educated Women (4 years or more) | | | |
|---|---|---|---|---|---|---|---|---|
| | **No Children** | **All younger** | **At least one** | **All older** | **No Children** | **All younger** | **At least one** | **All older** |
| *Outcomes* | | **than 5** | **younger than 5** | **than 5** | | **than 5** | **younger than 5** | **than 5** |
| **Fem. Hours Worked - Prim. job (log)** | 4.175 | 16.008 | -9.985 | -7.865 | -2.603 | 0.274 | -2.293 | -2.284 |
| | *[0.41]* | *[1.45]* | *[1.65]* | *[1.96]** | *[0.78]* | *[0.07]* | *[0.55]* | *[0.97]* |
| **Male Hours Worked - Prim. job (log)** | 0.646 | -3.705 | -2.907 | -2.233 | 3.501 | -1.772 | -1.540 | -1.471 |
| | *[0.13]* | *[0.98]* | *[1.15]* | *[1.23]* | *[1.76]* | *[0.85]* | *[0.77]* | *[1.15]* |
| **Fem. Hours Worked - All jobs (log)** | 4.558 | 17.655 | -9.267 | -9.335 | -3.150 | -2.407 | -1.903 | -1.760 |
| | *[0.43]* | *[1.51]* | *[1.39]* | *[2.18]** | *[0.87]* | *[0.61]* | *[0.43]* | *[0.71]* |
| **Male Hours Worked - All jobs (log)** | 0.650 | -3.850 | -2.319 | -1.997 | 3.060 | -1.688 | -1.473 | -1.349 |
| | *[0.12]* | *[0.99]* | *[0.86]* | *[1.08]* | *[1.51]* | *[0.86]* | *[0.71]* | *[1.03]* |
| **Male-Female Diff. In Worked Hours** | -1.826 | -3.599 | 1.005 | 1.174 | 2.113 | -0.132 | -0.034 | -0.184 |
| | *[0.69]* | *[1.49]* | *[0.72]* | *[1.15]* | *[1.93]* | *[0.12]* | *[0.03]* | *[0.25]* |
| **Male-Female Diff. In Tot. Worked Hours** | -1.480 | -3.561 | 1.352 | 1.177 | 2.005 | -0.066 | -0.167 | -0.190 |
| | *[0.55]* | *[1.47]* | *[0.96]* | *[1.14]* | *[1.82]* | *[0.06]* | *[0.14]* | *[0.26]* |
| **Male-Female Ratio of Worked Hours (log)** | -3.831 | -19.713 | 6.956 | 5.731 | 6.067 | -2.048 | 0.763 | 0.805 |
| | *[0.35]* | *[1.74]* | *[1.11]* | *[1.36]* | *[1.64]* | *[0.49]* | *[0.17]* | *[0.32]* |
| *Observations* | *2,304* | *3,341* | *6,984* | *14,654* | *7,948* | *13,507* | *12,386* | *31,840* |

Notes: See Notes for Table 6.

* Classical significance (p-value> 5%).

**Table 10: "Intent to Treat" Effects of Alimony Law by Sibship Gender-Composition (OLS Estimations**

**Dependent Variable: School Enrollment for Children ages 5 to 17**

| | Son | | Daughter | | Birth-Order Differential | | Gender Differential |
|---|---|---|---|---|---|---|---|
| | Oldest | Latter | Oldest | Latter | Sons | Daughters | among Latter Born |
| | [1a] | [1b] | [2a] | [2b] | [1a] – [1b] | [2a] – [2b] | Sons–Daughters |
| **PANEL A: Actual Treatment Effects - Informal versus formal marriages, 1995-1993** | | | | | | | |
| **Two or more children, boys and girls** | -2.83 | -0.99 | 4.99 | -4.17 | -1.84 | 9.16 | 3.18 |
| | *[1.53]* | *[0.63]* | *[3.05]** | *[2.72]** | *[0.82]* | *[4.37]*** | *[1.67]* |
| **Two or more children, only boys** | 0.50 | 1.82 | - | - | -1.31 | - | - |
| | *[0.24]* | *[0.75]* | | | *[0.52]* | | |
| **Two or more children, only girls** | - | - | -2.52 | -0.89 | - | -1.63 | - |
| | | | *[1.31]* | *[0.40]* | | *[0.67]* | |
| **PANEL B: Pre-Treatment Effects - Informal versus formal marriages, 1993-1992** | | | | | | | |
| **Two or more children, boys and girls** | 1.53 | 3.01 | 0.80 | 2.27 | -0.73 | -1.47 | 0.75 |
| | *[0.80]* | *[1.82]* | *[0.42]* | *[1.43]* | *[0.34]* | *[0.84]* | *[0.38]* |
| **Two or more children, only boys** | 2.04 | 3.35 | - | - | -1.30 | - | - |
| | *[0.87]* | *[1.27]* | | | *[0.49]* | | |
| **Two or more children, only girls** | - | - | 2.59 | 1.51 | - | 1.08 | - |
| | | | *[1.23]* | *[0.60]* | | *[0.41]* | |

Note: absolute-value t-statistics in [brackets] calculated with robust standard-errors (clustered at household level). Children sample sizes are 68,486 (Panel A) and 68,714 (Panel B).
Regressions further control for children age (second order polynomial) and education splines (plus dummies for illiteracy and completion of elementary and primary schools), second-order polynomial of parental age, parental education splines, parental race dummies, household demographics, non-labor income, asset ownership (dwelling), and geographica
Finally, the effects of assets, non-labor income, geographical region and urban location are allowed to be time-dependent.
* Classical significance (p-value >5%). ** Schwarz criterion significance (critical-value 3.34)

**Table 11: "Intent to Treat" Effects of Alimony Law by Sibship Gender-Composition**

**Dependent variable: School Enrollment for Children ages 5 to 17**

| | OLS Estimations | | | Household Fixed-Effects Estimations | | |
|---|---|---|---|---|---|---|
| | Birth-Order Differential | | Gender Differential for Latter Born (male-female) | Birth-Order Differential | | Gender Differential for Latter Born (male-female) |
| | Sons | Daughters | | Sons | Daughters | |
| **PANEL A: Actual Treatment Effects - Informal versus formal marriages, 1995-1993** | | | | | | |
| **Three or more children, boys and girls** | -1.84 | 9.16 | 3.18 | -3.25 | 9.03 | 3.34 |
| | *[0.82]* | *[4.37]*** | *[1.67]* | *[1.18]* | *[3.46]*** | *[1.56]* |
| **Two or more children, only boys** | -1.31 | - | - | -0.59 | - | - |
| | *[0.52]* | | | *[0.25]* | | |
| **Two or more children, only girls** | - | -1.63 | - | - | -2.57 | - |
| | | *[0.67]* | | | *[1.11]* | |
| **PANEL B: Pre-Treatment Effects - Informal versus formal marriages, 1993-1992** | | | | | | |
| **Three or more children, boys and girls** | -0.73 | -1.47 | 0.75 | -0.76 | -1.27 | 0.34 |
| | *[0.34]* | *[0.84]* | *[0.38]* | *[0.27]* | *[0.46]* | *[0.15]* |
| **Two or more children, only boys** | -1.30 | - | - | -2.51 | - | - |
| | *[0.49]* | | | *[1.00]* | | |
| **Two or more children, only girls** | - | 1.08 | - | - | 2.35 | - |
| | | *[0.41]* | | | *[0.93]* | |

Note: absolute-value t-statistics in [brackets] calculated with robust standard-errors (for OLS also clustered at household level). See Notes on Table 10.

* Classical significance (p-value >5%). ** Schwarz criterion significance (critical-value 3.34)

**Table 12: "Intent to Treat" Effects of Alimony Law, 1995-1993**

**School Enrollment for Children ages 5 to 17 (Fixed-Effects Estimation)**

| | Birth-Order Differential | | Gender Differential for |
|---|---|---|---|
| | Sons | Daughters | Latter Born (male-female) |
| **PANEL A: Children ages 11 and under** | | | |
| **Three or more children, boys and girls** | -4.60 | 16.43 | 9.04 |
| | *[0.83]* | *[3.03]** | *[1.76]* |
| **Two or more children, only boys** | 1.30 | - | - |
| | *[0.40]* | | |
| **Two or more children, only girls** | - | -5.54 | - |
| | | *[1.59]* | |
| **PANEL B: Children ages 11 and above** | | | |
| **Three or more children, boys and girls** | -9.26 | 13.37 | 9.45 |
| | *[1.57]* | *[2.32]** | *[1.81]* |
| **Two or more children, only boys** | 2.33 | - | - |
| | *[0.58]* | | |
| **Two or more children, only girls** | - | 2.92 | - |
| | | *[0.85]* | |

Note: absolute-value t-statistics in [brackets]. See Notes on Table 10.

Sample sizes are: 23,441 (PANEL A); 17,148 (PANEL B).

* Classical significance (p-value >5%).

## Table 13: "Intent to Treat" Effects of Alimony Law – School Enrollment for Children ages 5 to 17 (Fixed-Effects)

| | Less Educated Mothers (less than 4 years) | | | More Educated Mothers (4 years or more) | | |
|---|---|---|---|---|---|---|
| | Birth-Order Differential | | Gender Differential for | Birth-Order Differential | | Gender Differential for |
| | Sons | Daughters | Latter Born (male-female) | Sons | Daughters | Latter Born (male-female) |
| **PANEL A: Actual Treatment Effects - Informal versus formal marriages, 1995-1993** | | | | | | |
| **Three or more children, boys and girls** | -5.81 | 16.43 | 5.31 | -0.37 | 3.02 | 1.13 |
| | *[1.33]* | *[3.85]*** | *[1.61]* | *[0.11]* | *[0.94]* | *[0.41]* |
| **Two or more children, only boys** | -0.48 | - | - | 0.12 | - | - |
| | *[0.11]* | | | *[0.04]* | | |
| **Two or more children, only girls** | - | -3.40 | - | - | -2.37 | - |
| | | *[0.74]* | | | *[0.89]* | |
| **PANEL B: Pre-Treatment Effects - Informal versus formal marriages, 1993-1992** | | | | | | |
| **Three or more children, boys and girls** | 2.73 | -6.52 | -2.95 | -5.87 | 2.99 | 3.33 |
| | *[0.63]* | *[1.48]* | *[0.89]* | *[1.59]* | *[0.87]* | *[1.13]* |
| **Two or more children, only boys** | 1.74 | - | - | -5.30 | - | - |
| | *[0.40]* | | | *[1.69]* | | |
| **Two or more children, only girls** | - | 7.88 | - | - | -1.12 | - |
| | | *[1.64]* | | | *[0.38]* | |

Note: absolute-value t-statistics in [brackets]. Regressions estimated separatedly for each quadrant. See Notes on Table 10.

Sample sizes are 22,362 (Less Educated), 46,124 (More Educated); 23,670 (Less Educated), 45,044 (More Educated) for Panel A and Panel B, respectively.

* Classical significance (p-value >5%). ** Schwarz criterion significance (critical-value 3.28)

## Table 14: "Intent to Treat" Effects of Alimony Law

### School Enrollment for Children ages 5 to 17 (Fixed-Effects Estimation)

| | Less Educated vs. More Educated Mothers | | |
|---|---|---|---|
| | Birth-Order Differential | | Gender Differential of |
| | Sons | Daughters | Birth-Order Differential |
| **PANEL A: Actual Treatment Effects - Informal versus formal marriages, 1995-1993** | | | |
| Three or more children, boys and girls | -5.34 | 13.40 | 4.04 |
| | *[0.94]* | *[2.48]** | *[0.93]* |
| Two or more children, only boys | -0.51 | - | - |
| | *[0.10]* | | |
| Two or more children, only girls | - | -0.81 | - |
| | | *[0.15]* | |
| **PANEL B: Pre-Treatment Effects - Informal versus formal marriages, 1993-1992** | | | |
| Three or more children, boys and girls | 8.67 | -9.20 | -6.05 |
| | *[1.50]* | *[1.63]* | *[1.35]* |
| Two or more children, only boys | 7.50 | - | - |
| | *[1.36]* | | |
| Two or more children, only girls | - | 8.96 | - |
| | | *[1.56]* | |
| **PANEL C: Actual versus Pre-Treatment Effects** | | | |
| Three or more children, boys and girls | -14.01 | 22.61 | 10.09 |
| | *[1.40]* | *[2.35]** | *[1.32]* |
| Two or more children, only boys | -8.01 | - | - |
| | *[0.85]* | | |
| Two or more children, only girls | - | -9.77 | - |
| | | *[1.00]* | |

Note: absolute-value t-statistics in [brackets]. Joint estimation. Sample size 102,948. See Notes on Table 10.

* Classical significance (p-value >5%).

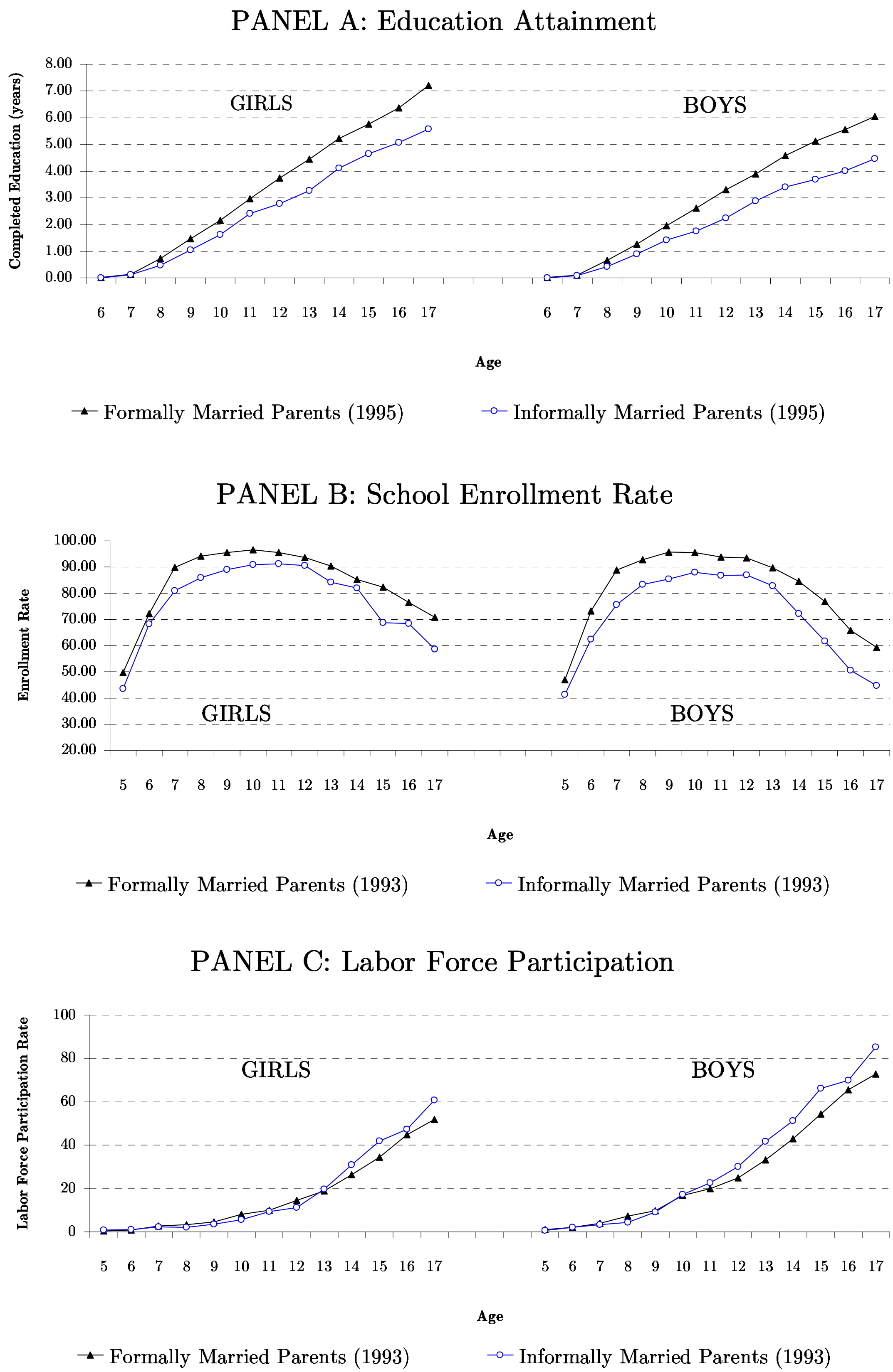


**Figure 1: Patterns of Investments in Children, by age and gender**