Alimony Rights and Intra-Household Allocation of Resources: Evidence from Brazil

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Abstract

Traditional household economic theory has largely relied on the concept of a "unitary" family decision-making model. Exploiting an exogenous source of variation provided by the extension of alimony rights to cohabiting couples in Brazil, this paper presents robust empirical evidence challenging this long-held notion. The hypothesis investigated here is that alimony rights improve women's outside options (that is, the welfare level in case of relationship dissolution), thus strengthening their negotiating positions, and increasing their influence over the allocation of resources within intact partnerships. Econometric results indicate that the empowerment of women reduces hours worked by female adults and impacts the level of investment in the human capital of the next generation. These findings suggest that models of the family should take intra-household heterogeneity in preferences into account.

Keywords: Intra-household decision-making, cohabitation, alimony rights, time allocation, investments in children. **JEL codes:** D13, J12, J22, O12.

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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

Household economic theory has traditionally treated the household as the appropriate unit of decision-making. Since consumer demand theory is predicated on the preferences of individuals, it is clear that the assumptions of such "unitary models" impose restrictions over observed behavior. This paper investigates one of those restrictions empirically; that changes in the intra-household distribution of power in decision-making should have no effect over the allocation of resources. The present study exploits an arguably exogenous change in Brazilian laws regarding couples' alimony rights and obligations upon relationship dissolution. This change in the institutional environment proxies a redistribution of bargaining power within intact households that favors women. The evidence uncovered indicates that female empowerment is likely to result in changes to the allocation of time to market work and housekeeping activities, as well as to the level of investment in the human capital of the next generation, particularly for girls. These findings are not consistent with models of the family that ignore intra-household heterogeneity of preferences.

At the core of models of the household that highlight the role of individual actors is the idea that individuals share in some way the benefits of forming a household. The premise is that observed behavior is a result of negotiations amongst family members, reflecting each individual's perception of costs and benefits as well as his or her relative "power" in asserting private preferences at the household level.[1] While these alternative models provide empirical predictions in sharp contrast with those of the unitary model, testing lags far behind. A key difficulty in testing these models is the fact that "power" has no clear (exogenous) empirical counterpart. The present paper circumvents this problem by exploiting the December 1994 extension of alimony rights to a large fraction of Brazilian couples living in consensual unions.[2] The hypothesis is that these rights improve women's outside options (that is, their welfare level in case of relationship dissolution) and thus strengthen their relative negotiating position within intact partnerships. It follows that newly

[1]See Manser and Brown (1980), McElroy and Horney (1981), Chiappori (1988), and Ulph (1988).

[2]Throughout this text, cohabitation, informal marriages and consensual union are used interchangeably. These are short for "marital cohabitation without legal marriage" or "marriage by habit and reputation".

empowered women should be able to appropriate a bigger share of the gains from their marital interactions. Therefore, holding constant other factors that influence household decision-making, changes in the allocation of resources following the new legislation's implementation should reflect the effects of the power redistribution only if individual preferences within the household are heterogeneous.

The empirical analysis is based on data from three waves (1992, 1993, and 1995) of the Brazilian Household Survey (*Pesquisa Nacional de Amostra de Domicilios* or PNAD). The PNAD yearly random sample consists of approximately 45,000 working-age couples and 65,000 children ages 5 to 17. The large samples support analyses of several sub-samples in order to assess the robustness of results and the credibility of the identification strategy. In order to control for confounding factors and trends, this paper explores similarities between married and cohabiting couples in Brazil, assigning the former the role of "control group." The econometric analysis investigates the compatibility of the data with the restriction imposed by the unitary model (i.e., no effects of changes in the balance of power) 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, or schooling.

The evidence presented here indicates that the intra-household empowerment of women caused an increase in the female consumption of leisure, and a reallocation of resources towards the schooling of older girls. Relative to their formally married counterparts, adult cohabiting women reduce housekeeping activities by 0.5% and labor supplied to the market by 3.4%. Moreover, the probability of attending school increases by 5.2% for their older daughters (relative to younger brothers). This effect is stronger among households in which women would be more dependent on alimony if the relationship dissolved (i.e., households headed by less educated mothers). The focus on gender and birth-order is motivated by the increased interest of economists in the allocation of resources within the household that depend on the biological characteristics of children, an issue extensively discussed by sociologists and psychologists. The results are compatible with evidence on parent-child relationships documented in the psychology literature, with mothers being "closer" to their daughters and to their first-born offspring. The results are also compatible with ethnographic research into Brazilian social norms related to the "formulation" of intergenerational contracts: older daughters are typically the ones who provide functional and financial help to their senior parents. As women live longer than men, and are expected to experience spells of widowhood, investment in "care takers" may be more attractive from the mother's perspective. Above all, these changes in the allocation of time and resources in response to changes in outside options are not consistent with models that assume that the household is the appropriate unit of decision-making. In a country where poverty has been shown to be pervasive and strongly correlated with low education and child labor, these findings regarding investments in human capital of the next generation also shed light on intricate social mobility issues and on the development process in general.

The remainder of the paper is organized as follows. Section 2 presents the economic theory of the household and its empirical implications. Section 3 describes demographic trends in marriage and cohabitation, summarizing some sociological interpretations and highlighting aspects of the Brazilian institutional environment. Section 4 describes the data sets and the econometric identification strategy. Section 5 presents the estimation results and discusses their relationship to other contributions to the literature. Section 6 concludes.

2 Conceptual framework

The most well known and widely adopted model of household behavior is based on the assumption that a representative individual can serve as the element of study.[3] 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 sets aside his or her particular preferences and the group creates something completely new — a set of common preferences defining their collective behavior. Even though this model has played a key role in the understanding of behavioral choices of families, it treats the operation of households as "black-boxes," leaving unexplored potentially interesting elements of intra-household interactions.

In an attempt to fill these gaps, intra-household bargaining and collective decision-making models have been developed. These models generalize the unitary formulations while restricting observed behavior in a way that specific nuances of the household decision-making process can be empirically identified. This section describes, in simple theoretical terms, the major characteristic that differentiates unitary and individualistic models: the role played by negotiations amongst households members and by changes in the intrahousehold balance of decision power in influencing observed outcomes.

[3]See review papers in Haddad et al. (1997) and in Rosenzweig and Stark (1997).

a) a general characterization

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

$$W^h = W[U^m(X, q; \underline{k}, \underline{\varepsilon}), U^f(X, q; \underline{k}, \underline{\varepsilon}), \lambda] \tag{1}$$

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 represents the "quality" of children (a good produced in the household), and k and ε represent vectors of observed and unobserved characteristics of the household, respectively. Characteristics of the household are determined at the time of the matching in the marriage market.[4] λ is a parameter summarizing the balance of decision power, and does not affect either individual preferences or the total household budget. In the present paper it is generally defined in terms of the female head's relative influence over the decisions regarding the intra-household allocation of resources.

In this formulation, children are assumed to be passive, with consumption and time allocations decided by their parents. The children's quality is represented by the production function Q, which summarizes an implicit relationship between child outcomes and childrearing inputs (food intake, parental time, school enrollment, etc.):

$$Q\left(q, X; \underline{k}, \underline{\varepsilon}\right) = 0. \tag{2}$$

Each child's individual (observed and unobserved) characteristics are also in the household characteristics. It is implicitly assumed that each parent sees 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 on the valuation of the child's quality relative to their own consumption.

Households are constrained by their time endowment ($\overline{L}$) and by their budgets. Let w be the market wage, then, with all prices $\left[P = (p; w^m; w^f)\right]$ assumed to be exogenous, the full-income budget constraint is:

$$P \cdot X = (w^m + w^f) \cdot \overline{L} + Y, \tag{3}$$

[4]Except for their implication for the empirical strategy, this study abstracts from matching issues. See Bergstrom (1993) and Foster (1998) for a discussion on the topic.

where Y is the non-labor income aggregated at the household-level.

This general (non-unitary) characterization take the individual members (or decision-makers) as basic elements, implicitly treating household decisions as the outcome of interactions and negotiations amongst them. This description of the process can rely either on bargaining models (see Manser and Brown, 1980; McElroy and Horney, 1981; Ulph, 1988; and Lundberg and Pollak, 1993) or on repeated interactions that achieve some form of efficient outcome (Chiappori, 1988, 1992; Browning and Chiappori, 1998). The fact is that these alternative models share a common rationale: the distribution of "decision power" within a household plays a crucial role in molding decisions. Each individual seeks to allocate household resources towards a bundle of goods he or she prefers, and the final outcome depends on each one's capability to assert personal preferences at the household level.

In this case, the general reduced-form demand function is augmented by factors that reflect the power distribution within the household and can be represented by:

$$X = X\left(\lambda; Z, \underline{\varepsilon}\right) \tag{4}$$

where Z collapses the observed characteristics: household, individual and community-level.

b) the unitary-models' basic restriction and its empirical implication

Under the assumptions of the traditional unitary model, either by the power of consensus (Samuelson, 1956) or by the emergence of a dictator (Becker, 1991), the decision process can be 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).[5] This implies that the household can be treated as a single decision unit, and the notion of power plays no role either on time allocation or expenditure and investment decisions. This construct correspond to the following (reduced-form) demand for goods and services:

$$X = X\left(Z, \underline{\varepsilon}\right) \tag{5}$$

When compared with models that hold any of the unitarian assumptions, the most basic difference emerging from considering the negotiating process is that elements exclusively influencing the balance of decision power should also affect the allocation of resources.

[5] The unitary formulation can also be derived from a environment with complete set of state-contingent claims, and perfectly enforceable contracts within the household.

In terms of the reduced-form just described, it corresponds to non-negligible effects of λ over realized demands. Therefore, in theoretical terms, the unitary model requires a neutrality with respect to power measures, implying that:

$$\frac{\partial X}{\partial \lambda} = 0 \tag{6}$$

The challenge then is the identification of meaningful empirical counterparts for those parameters that exogenously influence the distribution of power. Individuals within a household may derive power from multiple sources, each of which reflecting different aspects of the current or of available alternatives to the current arrangement. Early contributions to the literature focus on the notion that shares of income would capture most of the relative power. The operationalization of such an idea is limited by endogeneity issues, however, because individual labor supply (and income from labor) can actually be considered an outcome of the negotiation amongst family members. Non-labor income, although appealing in a static framework,[6] can only be valid if one sets aside issues of inter-temporal allocation of household resources.

In order to avoid endogeneity problems, the seminal work by Lundberg et al. (1997) exploited "exogenous variations" promoted by changes in government transfers programs. Based on the re-arrangement of household income generated by changes in governmental child allowance payments from husbands' pay-checks to wives' bank accounts, that study uncovered important aspects of the intra-household decision making process.[7] One possible drawback of their study is, however, the fact that confounding trends or concurrent events (unrelated to the balance of power redistribution) may have blurred the inference. This issue is directly assessed in the empirical section below.

Alternatively, as suggested by McElroy (1990), variables that exclusively affect the welfare outside the relationship could be used as proxies for power, as they should not influence outcomes if negotiations amongst members were not a significant component of the decision process within the household. The so-called "extrahousehold environmental parameters" (EEP) are factors 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 directly. Also known as "distribution factors," these encompass, for example, indicators of (re)marriage market competitiveness, welfare programs conditional on marital status (Rubalcava and Thomas, 2000), and laws governing divorce and marital property division (Gray,

[6]See Thomas (1990) and Schultz (1990).

[7]See also Ward-Batts (2000).

1998 and Chiappori et al., 2002).

In the present paper, the EEP idea is explored through exogenous variation provided by institutional changes. Specifically, the empirical exercise exploits the extension of alimony rights to cohabiting women (informally married) in order to "take the models to the data." The reasoning is that alimony rights favoring women, in the event of relationship dissolution, should correspond to an improvement in their negotiating position, affecting allocation decisions of intact households only if preferences are heterogeneous. In other words, the estimation of theoretical power redistribution effects is equivalent, in practice, to the causal inference of the law implementation. The next section draws a general picture of demographic aspects and details changes in the Brazilian legal system that are explored in the econometric estimations.

3 Demographic trends and institutional environment

3.1 Marriage, cohabitation and the family

Cohabitation is reshaping family life all over the world. In Great Britain, the percentage of women ages 20 to 24 in a heterosexual marriage-like relationship without being formally married increased from 11% to 55% between the mid-1970's and the mid-1990's.[8] In the United States, where cohabitation is considered a recent phenomenon, the proportion of women ages 15 to 24 in cohabitation increased by a factor of three between 1977 and 1997 (from 2.5% to 7.5%; Casper et al., 1999). Finally, in a handful of African countries and in all countries in Latin America and the Caribbean at least 15% of women in a union were in consensual relationships in the late 1980's (Westoff et al., 1994).

Brazil is no exception. While the share of women in some form of marital union remained stable between 1960 and 1995, the proportion of informal marriages among all unions rapidly increased since the mid-1980's, from 16.6 percent in 1984 to 28.0 percent in 1995.[9] In order to illustrate the pattern of marital status in the first half of the 1990's captured in the data used in the present paper, Appendix Table A1 presents figures of marital status for females between 15 and 55 years of age. The indication is that although

[8]Figures in other European countries are equally impressive. See sources cited in Emrisch and Fracescone (2000).

[9]See Goldani and Wong (1980) and Lazo (1994). Preliminary data from the 2000 Census indicates that the rate of informal marriages has reached 32%.

cohabitation corresponds to a larger fraction of marital unions among later cohorts, it also amounts to a significant share among the older ones.

The rise in cohabitation in Brazil can be interpreted as offsetting the upward (downward) trend in divorce (formal marriage). In a vision shared by a growing number of sociologists, the indication is that, even though cohabitation reshapes family life, it does so in subtle ways. In other words, cohabitation "should not be considered an increase in singlehood" or a "retreat from familism" (Treas and Lawton, 1999), but rather it should be seen "very much like a family status" (Bumpass et al., 1991). This interpretation particularly fits the Brazilian reality, where half of these consensual unions last more than 6 years.[10] Different from the case of the United States, cohabitation in Brazil resembles the informal unions in Scandinavia, where it is a socially accepted form of living in a union and of forming families. Ethnographic evidence presented by Rao and Greene (1996) indicates that Brazilian cohabitants 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. Lazo (1994) reports that "total marital fertility" is even higher among cohabiting couples than among formally married ones (after controlling for age, education and duration of union).

Therefore, in the Brazilian context, two major considerations should be underscored: First, as cohabiting couples represent an increasingly 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 can make the latter a good comparison group when measuring the effects of shifts in decision power over the intrahousehold allocation of resources.

3.2 Institutional environment

The long-term demographic trends described above motivated Brazilian courts and legislators to rethink the legal basis of the family. This has promoted a slow evolution from laws that guaranteed rights of specific "ideal" family units towards the protection of alternative familial arrangements, and led to the definition of rights and responsibilities of individuals within the family (both with respect to the State and to each other).

[10] Author's calculations based on raw data from the World Bank's Living Standards Measurement Study (1996/1997) and from the Macro International Inc.'s Demographic and Health Survey (DHS III,1996).

The first steps taken were in terms of the rights and responsibilities with respect to the State and other third parties. Important legislative measures were taken between 1942 and 1977. Issues covered include inheritance and alimony rights of biological sons born out-of-wedlock (including ones born from cohabiting relationships), and the inclusion of cohabitant partners as beneficiaries in the income tax and social security pension systems. Legislation regarding property division and alimony following the relationship dissolution was, however, much slower to change. Jurisprudence in these areas essentially followed the Civil Code of 1942 and consistently denied cohabitants alimony and property division rights.[11]

In 1988, however, guidelines were set by a new Federal Constitution (Article 226, paragraph 3). In its text, the Constitution universalized the spirit of the laws that governed 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." Additionally, religious marriages, considered until then as informal unions for legal purposes, were granted recognition by the State — provided that they were confirmed in a "Public Registry Office."[12]

The Constitution had distinct effects in terms of property division and alimony rights, since its text was dubious with respect to the assignment of rights and obligations between partners. The jurisprudence on property division, for example, indicates that courts adapted the Federal Supreme Court's Recommendation 380 ("*Sumula 380*" of the *Supremo Tribunal Federal*) to the cohabitation context. Originally, *Sumula 380* exclusively ruled rights over property after dissolution of business partnerships. Under the new Constitution, starting in the early 1990's, informally married spouses were given the right over their partners' property when able to prove collaboration in its accumulation.[13]

The new Constitutional text was ineffective in the context of alimony rights, however. Courts continued to hold that cohabitation did not yield maintenance rights for cohabitants under any circumstances, as observed in the jurisprudence.[14] The justification was, according to Ribeiro (2002) and Branco (1994), that the Constitution had only established the respon-

[11]See Azevedo (1997).

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

[13]This evolution of the jurisprudence was later consolidated in Law 9278/1996 — if no specific contract had been written, cohabitants were allowed 50% of the assets acquired after the start of the relationship.

[14]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. These are considered isolated cases, and Federal Courts did not corroborate them.

sibility 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 the State recognized cohabiting couples as family units, "consensual union" and marriage were still distinct entities. In effect, the same Constitutional article's 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 formal marriages. Post-dissolution alimony rights and obligations pertinent to the latter should, therefore, not be extended to cohabiting couples. Based on this argument, cohabitants' alimony petitions were normally dismissed without a hearing, and state-level Superior Courts usually corroborated these decisions.[15]

Diverging from the *status-quo*, the Law 8971, of December 1994, introduced alimony rights to cohabitants fulfilling certain criteria. The new legislation worked as an extension of the applicability of the Law 5478/1968 ("Alimony Law") to the dissolution of cohabitations. 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. Beneficiaries have the right to alimony until the commencement of a new (stable) relationship and so long as they can prove financial necessity — almost invariably defined by the courts in terms of potential earnings (or the individual's stock of human capital).[16] The alimony amounts are established by a judge according to the debtors' financial capabilities (normally 25 to 33% of monthly income converted into minimum wage units for indexation purposes).

The new law also established a process for alimony requests, reducing transaction costs, and activated 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 from years to months, 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

[15] 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*.

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

monthly payment would result in imprisonment.[17] The imprisonment would not eliminate the debt, however.

Because it can be interpreted as a shift in the balance of power within households (specifically, within cohabitant couples) favoring women, in the empirical exercises below, the implementation of Law 8971 is explored as an exogenous variation. Moreover, the fact that married couples were not affected by law change allows the construction of a plausible comparison group.

4 Data and econometric identification strategy

4.1 The data on cohabitation and marriage

The data set used in this study is from three years of the Brazilian Household Survey (*Pesquisa Nacional de Amostra de Domicilios,* PNAD) conducted by the Brazilian Census Bureau (*Instituto Brasileiro de Geografia e Estatistica,* IBGE).[18] The 92, 93 and 95 waves of the survey are appropriate for before and 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. The 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 in Brazil.[19] The PNAD yearly random sample consists of approximately 45,000 observations on adult couples, and 65,000 observations on children ages 5 to 17. These large samples support analysis of several sub-samples, allowing for the assessment of both the results' robustness and the identification strategy's credibility. The empirical exercises below directly address some shortcomings of the data, such as the absence of information on the duration of marital relationships by exploring additional sources of information and strategies that bypass the major concerns.

The data base employed focuses on information about households where heads and

[17] Although the number of individuals being arrested was small, the few events happened to be widely broadcasted.

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

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

spouses were older than 15 and younger than 55 years old.[20] In the case of the outcomes for children, due to age-limits in the availability of schooling information (the "investment allocation" variable), the sample was further restricted to families with sons and daughters between ages 5 and 17. Couples that were either in consensual unions or married only by the church are considered informally married, while the formally married are the ones with legal certification (with religious certification or not).

Table A2 presents general statistics for individual characteristics of women and men in the sample. These figures indicate that, compared to their formally married counterparts, informally married individuals are younger and less educated. In both groups, the average male-female education gap has increased over time, reflecting a characteristic that distinguishes Brazil from other developing countries (in particular in Asia and Africa): women are more educated than men, and this difference in educational attainment has consistently widened since the 1980's in the overall population (Beltrao, 2002). For the three years in the sample, cohabitants also have a smaller probability of being from the white population. Labor force participation and housekeeping activities are very similar for both men and women in the two groups, with slightly higher labor force participation amongst cohabiting men and married women. The outcomes in terms of labor supply for adults are, generally, not dramatically different for cohabitants and married adults. Cohabitants have a higher probability of living in a metropolitan area, while home ownership and per-capita non-labor income is higher amongst married-by-law couples. The measure of child-labor history (one if started working before age 15), an attempt to measure social status of the heads of the household when in childhood, does not indicate the same socio-economic advantage of legally married couples, however. The demographics of the household differs in the direction expected due to differences in age patterns, with cohabitants in a earlier phase of their life cycles. On average, cohabitants are members of a more economically disadvantaged group, but each of the characteristics (a certain age, income or education level, for example) has representatives amongst both cohabitant and married individuals. Most importantly, the observed characteristics of these two groups do not seem to evolve differentially over time. The *changes* in the age gap between partners, in the educational attainment of men and women, and in the household-level non-labor income between 1993 and 1995 are not different for formally or informally married couples/individuals.

Table A3 shows general statistics for characteristics of the children of couples in both

[20] Even though there can be married or cohabitant couples living with parents/in-laws, couples that were not considered head/spouse of the household were dropped from the sample. The focus is on the household-level decision-makers, which are asummed to be the head and the spouse. A small proportion of the sample was lost and no differential reduction of the samples (married versus cohabiting) was observed.

nuptial arrangements. The figures indicate higher educational attainment and higher proportion of school enrollment among children of formally married parents. At the same time, labor force participation is also higher for them than for the children of cohabiting parents. The levels are different for boys and girls, although the differences between informally and formally married parents still holds for both strata. Importantly, there may be heterogeneity in enrollment rates within households. This possibility is explored in the econometric specifications below.

4.2 Identification strategies

The objective of the estimation strategy is the identification of $\frac{\partial X}{\partial \lambda}$ for the demands of cohabiting couples, as discussed in the theoretical section above. Considering the change in the extra-environmental parameters explored in this paper, the evaluation of such a derivative corresponds to the estimation of differences in the levels of X under changed and unchanged legal regimes. Adopting the notation used in the causal-inference literature, let X_{ht} represent the realized outcome for household h at time-period t. Let T_h be an indicator function, assuming the unit value in the year after the law implementation (the post-treatment period). In addition, let D_h be the indicator of membership in the eligible-for-treatment group. Define $X_{ht}(1)$ and $X_{ht}(0)$ as the theoretical outcomes of household h when exposed ($\Delta\lambda \neq 0$) and non-exposed ($\Delta\lambda = 0$) to the treatment, respectively. Consequently, the effect of the treatment in household h can be represented by the difference $X_{ht}(1) - X_{ht}(0)$. The expected or average impact of the law over households in the eligible group can be theoretically represented by:

$$\tau(Z) = E[X_{ht}(1) - X_{ht}(0) \mid Z_h, D_h = 1, T_h = 1] \tag{7}$$

where, once more, Z is the vector of observed community, household and individual characteristics.

The identification of such a parameter requires, therefore, a specific modeling of the (not directly observed) counter-factual demand function $X_{ht}(0)$ for the eligible group in period $T_h = 1$. While one alternative would be the use of pre-treatment outcomes for the treated group,[21] the relevance of time-trends and the influence of concurrent events aside from the law adoption might confound the inference. However, if only part of the observed population is subject to the power shift (the treatment), a sample of comparable non-treated

[21] This is the strategy explored by Lundberg et al. (1997) and by Ward-Batts (2000).

observations (which are equally affected by the alternative events) can be used to net out effects of any confounding factors. Phrased in a slightly different way, the modeling of the counter-factual demand for the treated group can be based on the demand of a control-group. In this way, as long as both groups were to follow parallel paths in the absence of the treatment, the causal-inference can be rewritten as the so-called "difference-in-differences parameter" represented by:

$$\begin{aligned} \tau(Z) \;=\; & \{E\left[X_{ht} \mid Z_h, D_h = 1, T_h = 1\right] - E\left[X_{ht} \mid Z_h, D_h = 1, T_h = 0\right]\} \\ & - \{E\left[X_{ht} \mid Z_h, D_h = 0, T_h = 1\right] - E\left[X_{ht} \mid Z_h, D_h = 0, T_h = 0\right]\} \end{aligned} \tag{8}$$

The present paper explores two strategies for identification of this treatment parameter, namely level (ordinary least squares) and differential effects (fixed-effects estimation). Both rely on flexible parametric specifications.[22]

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 2 can be implemented. Assume a linear version of the demand for household h in the absence of the intervention:

$$X_{ht}(0) = \alpha_0 + T_h\alpha_1 + D_h\alpha_2 + Z_h \cdot (\beta_0 + T_h\beta_1) + \eta_{ht} \tag{9}$$

where the effects of time (α_0) and observed covariates (β) are common to all individuals (eligible or not), the effects of covariates are captured in Z_h, and η_{ht} collapses all unobservable characteristics. Considering a constant treatment effect $\tau(Z) \equiv \tau$, the following empirical model (in terms of the realized outcomes) can be explored:

$$X_{ht} = \alpha_0 + T_h\alpha_1 + D_h\alpha_2 + Z_h \cdot (\beta_0 + T_h\beta_1) + (D_hT_h) \cdot \tau + \eta_{ht} \tag{10}$$

From this parametric formulation, it is clear that the sufficient identifying assumption implicit in the choice of a control group can be represented by:

$$E[\eta_{ht} \mid T_h, Z_h, D_h] = 0 \tag{11}$$

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

This condition would allow the consistent estimation of τ 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 treatment (time $\times$ group) indicator — they require an absence of across-groups selection bias in response to the law.

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 (see Marrufo, 2001). In the present paper one concern is that a selection bias could result from changes (between cross-sectional sample draws) in the composition of unobservable characteristics in either the treatment or the control group due to the law's effects over marriage markets. In the context of treatment effect evaluation, selection on unobservables that 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 (e.g., $\mu_{cht} = \varepsilon_{cht} + \eta_{ht}$), siblings' fixed-effects can be used to wash-out household-level unobserved characteristics. One remarkable feature of this specification is that, even when the effects of household-level (parental) unobservable characteristics are assumed to vary over time, the within-household differential effects can still be consistently estimated.

In this formulation household-level observed characteristics are also eliminated 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). However, the idea is to explore heterogeneity in child characteristics (gender 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: $\left[\begin{array}{cc} Z_h^H & Z_{ch} \end{array}\right]$. Re-writing the model in (10) accounting for the influence of child characteristics over trends, over the effects of group characteristics and over the treatment effects ($\tau(Z) \equiv \tau_0 + Z_{ch} \cdot \tau_1$) yields:

$$\begin{aligned} X_{cht} &= \alpha_0 + T_h\alpha_1 + D_h\alpha_2 + Z_h^H \cdot (\beta_0 + T_h\beta_1) + Z_{ch}\alpha_0^c \\ &+ (Z_{ch}T_h)\,\alpha_1^c + (Z_{ch}D_h)\,\alpha_2^c + \left(Z_{ch}Z_h^H\right) \cdot [(\beta_0^c + T_h\beta_1^c)] \\ &+ (D_hT_h) \cdot (\tau_0 + Z_{ch}\tau_1) + \varepsilon_{cht} + \eta_{ht} \end{aligned} \tag{12}$$

The estimation of τ_1, the interaction between observed child characteristics and the treatment's impact, is consistent if:

$$E[\varepsilon_{cht} \mid T_h, Z_h, D_h, \eta_{ht}] = 0 \tag{13}$$

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

5 Regression results

5.1 Adults: time allocation effects

The outcome of interest in this section is the allocation of time (to housekeeping and market work) of male and female heads. Table 1 presents the "raw data" levels, time differences, and difference-in-differences for informally and formally married males and females. Significant differential changes are observed for hours worked and housekeeping. Relative to the formally married women, cohabiting females not only reduced hours supplied to the labor market, but also were less likely to perform housekeeping activities after the passage of the law. No significant differential effects were observed either among men or with respect to female labor force participation.

Column I of Table 2 presents the estimated difference-in-differences parameter using least-squares regressions (with controls for observed characteristics). These estimates of the effects of the alimony law implementation are based on the model shown in equation (10). All regressions include education and age of both men and women, as well as controls for household-level per-capita non-labor income (square root), home ownership, demographic composition and geographic location. The results indicate that, relative to their married counterparts, cohabitant women performed housekeeping activities less frequently than before the law (0.5%), this amounts to approximately 1/4 of the share of women *not* housekeeping in 1993. Additionally, once again relative to married females, working women favored by the extension of the alimony law significantly reduced hours weekly supplied to the labor market (3.4% or approximately 2/3 of an hour). Combined, these results suggest that the redistribution of power towards women was associated with an increased consumption of leisure.[23] No significant effects were found amongst male cohabitants.

Column II of Table 2 addresses two issues or concerns related to the strategy above: i) 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, and ii) the possibility that individuals may have anticipated implementation of the law, adapting their behavior before the administration of the treatment. Both concerns are investigated by the reproducing the treatment effect calculations as if the law change had occurred between the 1992 and 1993 waves of PNAD, and not in 1994. Because no significant differential changes between cohabitant and married outcomes were observed before the law implementation, the pre-treatment experiment rules out the influence of differential trends or anticipation over the results.

The pre-treatment estimations just described do not address another potential problem: shocks that occurred between 1993 and 1995 and 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 was the case in Brazil, promote a decrease in the relative prices of tradable versus non-tradable goods. This suggests that some economic sectors, 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 for, then the observed treatment effect may capture both stabilization and power redistribution effects.

To verify that covariates included in the model sufficiently control for differences in observed characteristics across groups, an indirect argument can be presented. Column II of Table 3 presents a reestimate of the treatment effects excluding covariates used as controls for characteristics that potentially influence the stabilization effects but arguably not the ones originated by the power redistribution. The underlying hypothesis is that, if individual characteristics were not correctly controlled and changes in macroeconomic conditions were

[23]In principle, PNAD interviewers were instructed to consider child-care a housekeeping activity. It is not clear, however, that mothers and fathers consider all time spent with children as housekeeping rather than leisure.

solely responsible for the estimated effects, then the exclusion of additional variables would increase the magnitude of the effects in absolute value. If, for example, cohabitant and married couples were confined to separate geographical regions with different macroeconomic prospects (the former in a recession region), then the reduction in worked hours may simply reflect a local contraction of labor demand rather than the alimony law implementation (a federal law change affecting both regions). The indicators of geographic location control for such factors and, in this case, their exclusion would drive up the negative effect on worked hours. A similar argument applies to the time-varying effects of non-labor income and of asset ownership (interactions of the wealth variables with the year dummy). The results in Column II reveal that the exclusion of controls does not significantly affect the size of the estimated parameters.

Another important consideration is that changes in hours worked could reflect the fact that cohabiting and married individuals have, on average, differing contractual attachment to the labor force (for example, less formal jobs among cohabitants). One may consider, for example, that those with formal contracts are less vulnerable to macroeconomic shocks. Hence, Columns III and IV in Table 3 limit the sample to "stayers" (individuals who did not change jobs in the one year period preceding the survey) and to couples in which both partners had non-formal and formal employment, respectively. The results are consistent with those presented above, indicating that the burden of market work did indeed shift toward male cohabitants. In a sample where both men and women are neither formal wage-workers nor public sector employees, working cohabiting women (relative to their married counterparts) reduced hours supplied to the market by 7.9% (1.4 hours per week), while the ratio of hours worked by male and female partners increased 5.8%. The results of the subsample of formal workers indicates similar patterns, but with smaller and not significant outcomes. The flexibility of worked hours among non-formal workers is one possible explanation for such differences in the responses' size, however.

The investigation of effects associated with the alimony law implementation is potentially subject to selection biases originating from changes in the marriage market. In other words, the adoption of the law may have altered the attractiveness of cohabitation for both men and women, possibly changing the composition of the treatment and control groups between 1993 and 1995.[24] This is the case when one considers the effect of divorce costs on the likelihood of marriage formation/dissolution, for example.[25] Similar reasoning would

[24]The trends in educational level and age gap between partners observed between 1992, 1993 and 1995 were the same for both informal and formally married individuals. These results assure no differential compositional change in terms of age and education.

[25]See Bougheas and Geogellis (1999) for a theoretical model, and Peters (1986) and Friedberg (1998) for

suggest compositional effects resulting from the alimony law adoption. A problem emerges if the composition of the groups changes 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).

Setting inflation stabilization effects aside, or considering that it does not differently affect marriage or cohabitation rates, three hypotheses about the compositional changes can be discussed. First, couples on the margin of cohabitation and marriage, which in the steady-state would have married, may have been made indifferent between the two arrangements. As a result, the two groups are expected to become more similar: a larger proportion of cohabitants looks like married couples after the law (and behave accordingly). This is partially ruled out by the findings presented above. Even though fewer hours worked among cohabitants makes married and cohabiting women more similar, fewer housekeeping activities among the latter have actually made them dissimilar.[26] Second, single women may have had more incentives to start new cohabitation relationships. However, as newly formed couples are not affected by the law in terms of allocation of resources, they should not be responsive in terms of consumption decisions. The effects uncovered in the econometric estimations are likely to be downward biased if this is a major factor in the selection process.

Finally, cohabiting men may have had new incentives to terminate relationships that were not stable (in legal terms) at the time the law was passed. As a result, selection occurs because after the law the remaining cohabitant men are the ones who are willing to give their partners a larger share of the marital surplus anyway (i.e., even before or in the absence the law). In order to circumvent this selection hypothesis, estimations in Table 4 are based on two strata of the original sample. The first is limited to childless couples (Column I), which is the group that is both more likely to be unaffected by the law (non stable unions), and more likely to change in composition in response to the law. The second stratification (in Column II) explores the idea that such a selection mechanism should not to be an issue for the stratum of couples already living together for 5 years or more. Although information on duration of relationships is not available, a possible assumption is that couples with at least one child older than 5 are living together for that period or longer the estimation can be based on that group. Cohabiting women with children above 5 year of age reduced housekeeping participation by 0.9%, while cohabiting working women reduced labor supply by 5% (both relative to their married counterparts). No significant effects are found among

the U.S.-based empirical evidence.

[26]Investigation of formal marriages rates from yearly census data (all legal marriages are registered in the *Censo de Registro Civil* - IBGE) indicates no alteration in the trend after the passage of the law. Marriage rates continuously fall at the same rate throughout the 1990's.

childless couples. The results corroborate the effects over intact households posited to be associated with the law implementation.[27]

5.2 Children: school enrollment differential-effects

The analysis in this subsection focuses on variations in school enrollment within families. Enrollment decisions are at the core of a broader set of choices regarding investment in children's human capital. Analysis of these decisions is incomplete if it ignores within household differences. For example, if the mother's preferences regarding the distribution of education among children differ from those of the father, the empowerment of women would be expected to result in a reorientation of the allocation of resources. An analysis which relied solely on aggregated effects would not necessarily identify this effect.

The differential pattern of investment in the human capital of children within families has received attention in many areas of scientific inquiry. In economics, explanations of within-family differences in investments usually involve either differential costs and returns to investments (as in Rosenzweig and Schultz, 1982), or the possibility that parents exhibit "unequal concern" for identical outcomes amongst their children (as in Behrman et al., 1986). It has also been emphasized that these investment decisions cannot be studied in isolation from factors as family size and sibling sex composition (Lindert, 1977; Kessler, 1991; and Butcher and Case, 1994) — a vision shared by the extensive sociology literature on the topic, as Smelser and Stewart (1968), Powell and Steelman (1990), Conley (1991), Parish and Willis (1993), and Hauser and Kuo (1998). The major focus of the studies cited has been the relationship between intrahousehold allocation of resources and gender and birth-order of the children. None of these studies investigates heterogeneity in parental preferences regarding the distribution of education.

More recent contributions to the economics literature have dealt with some of these issues. Duflo (2000), for example, examines the effect of government transfers to elderly women and men in South Africa on investment in girls relative to boys. She estimates that girls' anthropometrics are improved (relative to their brothers) when grandmothers receive transfers from the government, but does not find evidence of gender differentials

[27] An additional concern is that selection may occur even in households with children older than five because it may be a result of divorced women's selection into (and out of) a second relationship. When further restricting to a sample of women with children ages 2 and younger, assuming that those are common offsprings to the current partnership, the overall pattern of results persists. However, while stronger effects are observed, only the housekeeping one remains significant.

when grandfathers are the beneficiaries. Her findings suggest that grandmothers "prefer" to invest on their granddaughters. Thomas (1994) also explores possible heterogeneity in parental preferences, finding empirical evidence in Brazil, Ghana and the United States which indicates that maternal resources have a stronger impact on the accumulation of daughters' human capital, while paternal resources have a stronger impact on sons. Similar results are also reported by Contreras and Rubalcava (2000), who examine the nutritional status of Chilean children and their birth-order and gender. They find some evidence of differentiation of children along both dimensions, with mothers' resources affecting the nutrition of younger girls disproportionately more than the nutritional status of younger boys.

Psychologists have also examined the association between birth-order and gender and child development. This literature most often focuses on differences in the interaction between children, parents, and siblings.[28] According to Sulloway (2001), siblings in competition for parental favor put in motion a family dynamic that leads to the formation of relatively stable "niches." These niches are shaped by individual characteristics like gender and birth-order. Niche differentiation within the household is then associated with parent-driven processes, generally involving differential parental investment of time and resources. Indicators of the quality of the relationship among family members are commonly used to assess favoritism — including, for example, children's self-esteem, perception of parental favoritism, closeness to kin and support received after distressing experiences. Recent empirical examinations include Salmon and Daly (1998) and Rohde et al. (2003), which present evidence that women are more likely than their brothers to name mothers as the closest kin. Focusing on birth-order differentials and exploring data from six countries (Austria, Germany, Israel, Norway, Russia, and Spain), Rohde et al. (2003) also present evidence that first-borns are more likely to develop stronger connections with their mothers.[29]

Overall, these findings suggest that the relationship between parental preferences and children's gender and birth-order may differ. The direction of the interaction of these effects is an empirical question. The exercises presented below examine a specific conjecture: if mothers and fathers depend on children's support as they age, they might have (insurance motivated) incentives to devote greater resources to earlier born children. This would be the case because these offspring become economically independent earlier, so that they offer returns while parents are still able to enjoy them.[30] A relevant qualification of this argument is that, as women live longer than men, and are expected to experience longer spells of

[28]See Sulloway (1996) and review section in Hertwig et al. (2002).

[29]On the other hand, middleborns are particularly more likely to name their fathers as the closest parent. See also Schachter (1982).

[30]This conjecture shared by Behrman and Taubman (1986), Horton (1988), and Kessler (1991).

widowhood, the investment in "care takers" should be more attractive from the mother's than from the father's perspective. Furthermore, since ethnographic evidence on Brazilian social norms suggests, older *daughters* are typically the ones that provide functional and financial help to their senior parents (Camarano and El Ghauri, 1999; Goldani, 1999; and Saad, 1999), these birth-order effects likely differ by gender.

The main objective of the empirical exercises based on equation (12) is, therefore, to identify differences in the effects of the alimony law implementation by birth-order and gender. The effects are calculated for households with boys and girls, and depending on the sex of the first-born child. All regressions estimated in this subsection include controls for child-level characteristics, namely age and educational attainment dummies. The controls for parental and household characteristics are the same as those used in the calculation of outcomes for adults (age, education, non-labor income, home ownership, household demographics and geographic location). 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.

Column I of Table 5 (PANEL A) reports the results of the ordinary least-squares regressions for each child, by birth order and gender. Significant differences between cohabitant and married couples behavior are observed. In particular, there are positive differential effects with respect to investments in daughters that are the first-born child of cohabiting couples, with negative differential change for non-first-born girls. In order to assess potential selection on unobservables problems, Column III in Table 5 reproduces gender-differential results using household fixed-effects estimations (solely based on within household variation). The figures indicate that selection does not alter the main conclusion with respect to differential gender and birth-order effects. For this sample of cohabiting couples, the effects over children indicate an increase of 5.2% in enrollment of oldest daughters (relative to their younger brothers). On the other hand, enrollment non-first-born daughters of cohabitant couples have grown in a slower pace than for non-first-born daughters of married couples (the relative change in the gender differential is -7.3%).[31] The latter result is compatible with a reallocation of time, with the mother consuming more leisure and investing in the schooling of the oldest daughter, while shifting some of the housekeeping burden to the younger daughters. However, these figures are also compatible with the idea that mothers invest more time in leisure activities that are consumed jointly by themselves and their younger daughters. Finally, as indicated in PANEL B of Table 5, no effects were found when a boy was the first-born child, suggesting that signs of favoritism depend both on gender and birth-order.

[31] The bulk of these effects correspond to entry in school being deterred for non-oldest daughters.

Table 6 explores characteristics of the law requirements that identify a stratum of the cohabitant mothers' population that is expected to be more affected by the treatment. The sample of couples with a female first-born child is stratified by the education of the mother, in an attempt to extract the most out of the law's "necessity" clause — by which women with less potential earnings were considered more dependent on their partners and, consequently, on alimony payments. The results in PANEL A indicate striking differences in the expected direction: stronger effects for children of women with less than elementary school education (Column I), and no significant effects for children of more educated women (Column II). The probability of attending school increased 12.6% for first-born daughters of less educated cohabiting mothers (relative to their younger brothers), compared to the differential effects amongst daughters of married couples. The effect for the oldest daughters is also shown to be significantly different for less and more educated mothers. The additional level of differentiation eliminates, however, the significance of the effects amongst non-oldest daughters (Column III). No significant pattern of changes emerges from the pre-treatment experiment in PANEL B. 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. In particular, the empowerment of women is shown to cause an increase on the investment in education of their first-born daughters.

PANEL A in Table 7 focuses on results of inframarginal changes in targeted educational attainment. The sample is restricted to children that have already been to school, and the dependent variable is measured as the logarithm of the completed years of education plus the enrollment (0 or 1) in the current year. Consistent with previous results, the estimated changes presented on Table 7 indicate that the investment in education of oldest daughters of cohabitant mothers increases relative to their brothers and in comparison with the case of married couples' children. The gender-differential targeted educational attainment was shown to increase 2.8% faster for children of cohabitant couples than for children of married mothers (in particular for the less educated group). There is no effect for the pre-treatment years, as shown in PANEL B.[32]

[32] As an exploratory check of the investment in old-age security hypothesis posed above, I have examined if it is the case that the favoring of first-born girls is stronger among older mothers. The idea is that discounting should attenuate the effects among younger mothers. Preliminary estimation indicates that, both for the overall enrollment and for the targeted educational attainment, the evidence is compatible with the proposed channel of causation.

5.3 Discussion

The time-allocation results found in Section 5.1 conform with previous findings (based on U.S. data sets) reported in the literature. The contraction in female labor supply following increases in women's power was also observed in a study by Rubalcava and Thomas (2000), while the reduction in hours dedicated to home-production were shown by Gray (1998).[33] The combination of both results suggests that, for Brazilian women in consensual unions, increased power was associated with increased consumption of leisure.

With respect to investments in children, differential effects of birth order and gender are observed in cohabiting versus married couples. It is possible to rule out explanations which are based on time-changes in the cost or return to the investments because those would influence both cohabitant and married couples. Moreover, the estimation based on an exogenous change in the institutional environment overcomes potential problems regarding the use of education (which may also reflect childrearing technology) and non-labor income (which may be endogenous) as proxies for "power." 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.[34] In particular, time differences are expected to eliminate those characteristics that are expected to be constant within each of the marital arrangements.

The evidence that more familial resources were allocated to the schooling of first-born girls presented in Section 5.2 above is compatible with Brazilian social norms related to the "formulation" of intergenerational contracts. The findings in the present paper suggest that, by increasing the investment in the education of first-born girls, mothers may be targeting better prospects for their daughters, either in the job or marriage markets (via assortative mating), but they may also allocate resources to maximize their private returns in terms of old-age security.

6 Conclusions

This paper investigates the effect of a shift in the balance of decision power within households over the allocation of time and investments in children. Using an exogenous source of variation provided by the adoption of a law (extension of alimony rights to cohabitants),

[33]See also Chiappori et al. (2002).

[34]See Hetherington et al. (1978) and Mekos et al. (1996).

and the similar "family status" of cohabitant and married couples in Brazil, this paper provides robust empirical evidence that empowerment of women results in: i) a reduction in hours of work by female heads, including housekeeping, and ii) a redistribution of household resources toward schooling of first-born girls. These effects were shown to be stronger when considering households headed by less educated women. This result is consistent both with the theory and with aspects of the law, since women with lower potential earnings are more likely to be dependent on alimony in the event of relationship dissolution.

These findings are not consistent with a characterization of the household that ignores heterogeneity of preferences, and they challenge the long-held notion that households can be studied through the lens of a representative-agent model. The evidence strongly suggests that the intra-household allocation of resources is the outcome of an elaborate process, where observed decisions result from bargaining and negotiations between males and females with different preferences and varying abilities to assert their "vision of the world" within the household.

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Table 1: Levels, changes and differential changes in observed outcomes of interest for men and women ages 15 to 55

	Informally Married Couples		Formally Married Couples		Informal-Formal
	Level	Difference	Level	Difference	Diff-in-Diffs
	1993	1993-1995	1993	1993-1995	1993-1995
Housekeeping Indicator (x 100)					
Male	46.50	4.84	45.80	4.72	0.12
		(0.62)		(0.39)	(0.73)
Female	97.95	-0.28	97.81	0.20	**-0.49**
		(0.18)		(0.11)	**(0.21)**
Labor force participation (x 100)					
Male	96.89	0.02	96.43	-0.19	0.20
		(0.21)		(0.15)	(0.26)
Female	51.94	2.36	54.28	3.25	-0.89
		(0.62)		(0.39)	(0.73)
Log-hours worked in the week (conditional on both partners participation), changes are x 100					
Male	3.80	0.13	3.81	0.73	-0.60
		(0.59)		(0.35)	(0.68)
Female	3.37	-0.34	3.30	3.10	**-3.44**
		(1.26)		(0.75)	**(1.47)**
Female-male difference	-0.44	-0.47	-0.50	2.36	-2.84
		(1.35)		(0.80)	(1.57)

Notes: Standard-errors in parentheses under estimated changes in means. Samples sizes are 92,964 for extensive margin and 45,177 for sub-sample of participants.

Table 2: Time allocation effects of the alimony law (Diff-in-diffs) informally versus formally married adults (ages 15 to 55)

	Column I: Pre-post treatment Diff-in-Diffs 1993-1995	Column II: Pre-pre treatment Diff-in-Diffs 1992-1993
Housekeeping Indicator (x *100)*		
Male	0.48	0.30
	(0.73)	(0.76)
Female	**-0.51**	0.20
	(0.22)	(0.24)
Labor force participation (x *100)*		
Male	0.30	-0.45
	(0.25)	(0.26)
Female	-0.15	-0.69
	(0.70)	(0.74)
Log-hours worked in the week (conditional on both partners participation), changes are x *100*		
Male	-0.89	-0.12
	(0.71)	(0.71)
Female	**-3.39**	-0.92
	(1.44)	(1.51)
Female-male difference	-2.50	-0.80
	(1.53)	(1.60)

Notes: Robust standard-errors in parentheses under estimated changes in means. Samples sizes are 92,964 for extensive margin and 45,177 for sub-sample of participants (Column I). For Column II samples sizes are 90,121 for extensive margin and 42,423 for sub-sample of participants.

Table 3: Time allocation intensive margin effects of the alimony law informally versus formally married adults (ages 15 to 55) - macroeconomics controls and conditional on labor contracts

	Column I	Column II	Column III	Column IV
	Pre-post treatment	Pre-post treatment	Pre-post treatment	Pre-post treatment
	General model	*No geography or assets controls*	*Non-formal workers only*	*Formal workers only*
	Diff-in-diffs	Diff-in-diffs	Diff-in-diffs	Diff-in-diffs
	1993-1995	1993-1995	1993-1995	1993-1995
Log-hours worked in the week (conditional on both partners' participation)				
Male	-0.89	-0.59	-2.04	-1.35
	(0.71)	(0.69)	(1.41)	(1.34)
Female	**-3.39**	-2.80	**-7.89**	-2.70
	(1.44)	(1.46)	**(2.88)**	(1.58)
Female-male difference	-2.50	-2.21	**-5.84**	-1.35
	(1.53)	(1.56)	**(2.98)**	(1.87)

Notes: Robust standard-errors in parentheses under estimated changes in means. Samples size is 45,177 (Columns I and II), 14,193 (Column III), and 7,223 (Column IV).

Table 4: Time allocation effects of the alimony law
informally versus formally married adults (ages 15 to 55) - family composition strata

	Column I	Column II	Column III
	Pre-post treatment	**Pre-post treatment**	**Pre-post treatment**
	Childless couples only	*Couples with at least one child > 4*	*Couples with at least one child > 4 and youngest child < 3*
	1993-1995	1993-1995	1993-1995
Housekeeping Indicator (x 100)			
Male	0.61	0.80	0.43
	(2.36)	(0.92)	(0.59)
Female	0.68	**-0.91**	**-1.21**
	(0.75)	**(0.27)**	**(0.56)**
Labor force participation (x 100)			
Male	0.50	0.24	0.43
	(0.75)	(0.34)	(0.59)
Female	0.23	-0.79	-0.21
	(2.33)	(0.90)	(1.92)
Log-hours worked in the week (conditional on both partners participation x 100)			
Male	3.76	-0.15	-4.00
	(2.05)	(0.09)	(2.08)
Female	-2.91	**-4.95**	-7.98
	(3.78)	**(1.79)**	(4.66)
Female-male difference	-6.67	-3.40	-3.98
	(4.12)	(1.90)	(4.97)

Notes: Robust standard-errors in parentheses under estimated changes in means. Samples for extensive margin are 7,475 (I), 65,340 (II), and 11,129 (III). For intensive margin samples are 4,162 (I), 32,957 (II), and 4,771 (III).

Table 5: Children's school enrollment effects of the alimony law (ages 5 to 17)
Linear probability models (enrolled=1), children of informally married versus children of formally married household heads

	Column I Informally versus formally married Diff-in-Diffs 1993-1995 OLS	Column II Gender Differential Diff-in-Diffs-in-Diffs 1993-1995 OLS	Column III Gender Differential Diff-in-Diffs-in-Diffs 1993-1995 FE - Within HH
PANEL A: First-born is a girl			
First-born	**5.66**	**5.56**	**5.20**
	(1.70)	**(2.42)**	**(2.10)**
Younger daughter(s)	**-8.15**	**-8.25**	**-7.30**
	(2.48)	**(3.03)**	**(2.95)**
Younger son(s)	0.10		
	(1.73)		
PANEL B: First-born is a boy			
First-born	-2.78	0.65	1.10
	(1.90)	(2.54)	(2.19)
Younger daughter(s)	-2.13		
	(1.68)		
Younger son(s)	-3.19	1.06	0.27
	(2.63)	(3.12)	(3.17)

Notes: Robust standard-errors in parentheses under estimated changes in means. Samples are 22,540 (female first-born), and 23,536 (male first-born)

Table 6: Children's school enrollment effects of the alimony law (ages 5 to 17) - mother's education strata
Linear probability models (enrolled=1), children of informally married versus children of formally married heads
Female first-born in mixed-sex sibship households (Within household differential effects' estimation).

	Column I **<u>Gender differentials</u>** **Diff-in-Diffs-in-Diffs** Less educated mothers (education < 4)	**Column II** **<u>Gender differentials</u>** **Diff-in-Diffs-in-Diffs** More educated mothers (education >= 4)	**Column III** **<u>Gender differentials</u>** **Quadruple-Diffs** Less-More educated mothers
PANEL A: Pre-post treatment (1993-1995)			
First-born	**12.57**	1.46	**11.11**
	(3.71)	(2.47)	**(4.53)**
Younger daughter(s)	**-9.56**	-4.23	-5.33
	(4.58)	(3.81)	(6.03)
PANEL B: Pre-pre treatment (1992-1993)			
First-born	-2.74	-0.97	-1.77
	(3.94)	(2.64)	(4.82)
Younger daughter(s)	6.84	-0.88	7.72
	(4.85)	(3.94)	(6.32)

Notes: Robust standard-errors in parentheses under estimated changes in means. Samples are 7,927 and 14,413 (PANEL A), and 8,214 and 14,506 (PANEL B).

Table 7: Children's inframarginal intended schooling effects of the alimony law (ages 5 to 17) - mother's education strata log (educational attainment + enrollment), children of informally married versus children of formally married heads Female first-born in mixed-sex sibship households (Within household differential effects' estimation).

	Column I Gender differentials Diff-in-Diffs-in-Diffs Less educated mothers (education < 4)	Column II Gender differentials Diff-in-Diffs-in-Diffs More educated mothers (education >= 4)	Column III Gender differentials Quadruple-Diffs Less-More educated mothers
PANEL A: Pre-post treatment (1993-1995)			
First-born	**2.78**	-0.78	**3.56**
	(1.08)	(0.48)	**(1.20)**
Younger daughter(s)	-1.23	-0.96	-0.27
	(1.33)	(0.61)	(1.48)
PANEL B: Pre-pre treatment (1992-1993)			
First-born	-0.12	-0.05	-0.07
	(1.24)	(0.55)	(1.38)
Younger daughter(s)	-0.91	-0.24	-0.67
	(1.58)	(0.79)	(1.80)

Notes: Robust standard-errors in parentheses under estimated changes in means. Samples are 6,152 and 13,641 (PANEL A), and 6,336 and 13,156 (PANEL B).

Table A1: Marital status of women ages 15 to 55 - by decade of birth

	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 A2: Descriptive statistics - men and women 15 to 55 by type of marital relationship

	Informally Married Couples			Formally Married Couples		
	1992	1993	1995	1992	1993	1995
Individual and matching characteristics						
Male age	34.73 (0.09)	34.63 (0.08)	34.74 (0.08)	37.88 (0.05)	38.17 (0.05)	38.67 (0.05)
Female age	31.15 (0.08)	31.04 (0.08)	31.30 (0.07)	34.54 (0.05)	34.81 (0.05)	35.42 (0.05)
Age gap	3.59 (0.07)	3.58 (0.07)	3.44 (0.06)	3.34 (0.03)	3.36 (0.03)	3.25 (0.03)
Male education	4.56 (0.04)	4.70 (0.04)	4.78 (0.03)	5.94 (0.02)	6.07 (0.02)	6.19 (0.02)
Female education	4.59 (0.04)	4.75 (0.03)	4.94 (0.03)	6.12 (0.02)	6.23 (0.02)	6.44 (0.02)
Education gap	0.00 (0.03)	-0.05 (0.03)	-0.16 (0.03)	-0.17 (0.02)	-0.17 (0.02)	-0.25 (0.02)
Male white (%)	40.39 (0.46)	40.89 (0.45)	41.39 (0.41)	59.19 (0.27)	58.92 (0.27)	58.81 (0.27)
Female white (%)	42.12 (0.46)	42.87 (0.45)	42.40 (0.41)	61.70 (0.27)	61.66 (0.27)	61.09 (0.27)
Individual activities						
Male housekeeping (%)	39.59 (0.45)	46.50 (0.45)	51.34 (0.42)	39.81 (0.27)	45.80 (0.27)	50.52 (0.27)
Female housekeeping (%)	96.97 (0.16)	97.95 (0.13)	97.67 (0.13)	97.12 (0.09)	97.81 (0.08)	98.02 (0.08)
Male in labor force (%)	97.08 (0.16)	96.89 (0.16)	96.91 (0.14)	96.47 (0.10)	96.43 (0.10)	96.24 (0.10)
Female in labor force (%)	51.94 (0.46)	51.94 (0.45)	54.30 (0.42)	53.79 (0.27)	54.28 (0.27)	57.53 (0.27)
Male hours worked - week*	47.58 (0.18)	46.84 (0.17)	46.98 (0.16)	47.65 (0.10)	47.10 (0.10)	47.32 (0.10)
Female hours worked - week*	34.43 (0.23)	34.40 (0.22)	34.42 (0.20)	32.52 (0.13)	32.62 (0.13)	33.29 (0.12)
Household location						
Urban sector (%)	81.58 (0.36)	81.36 (0.35)	82.07 (0.32)	81.85 (0.21)	82.19 (0.21)	82.31 (0.21)
Metropolitan sector (%)	42.06 (0.46)	43.29 (0.45)	41.94 (0.41)	38.89 (0.27)	38.94 (0.27)	38.37 (0.27)
Assets and non-labor income						
Home ownership (%)	60.29 (0.45)	60.14 (0.45)	61.67 (0.41)	71.28 (0.25)	72.08 (0.25)	73.46 (0.24)
Per-capita non-labor income**	13.49 (0.79)	14.04 (0.99)	14.87 (1.10)	19.07 (0.83)	22.81 (0.74)	21.78 (0.76)
Working history						
Male was child-laborer (%)	33.29 (0.44)	33.66 (0.43)	36.84 (0.40)	33.41 (0.26)	35.17 (0.26)	37.02 (0.27)
Female was child-laborer (%)	15.36 (0.33)	15.93 (0.33)	17.52 (0.32)	16.14 (0.20)	16.69 (0.20)	18.49 (0.21)
Observations	*11,633*	*12,109*	*14,410*	*33,156*	*33,241*	*33,204*
Observations (participants)	*5,051*	*5,295*	*6,656*	*15,801*	*16,216*	*17,060*

Notes: Standard-errors in parentheses next to estimated means. *conditional on participation of both partners on the labor market. ** in R$ per month - real values of Sep-1999.
Source: PNAD (1992-1995).

Table A3: Descriptive statistics - children 5 to 17 by type of marital relationship of the household heads

	Informally Married Couples			Formally Married Couples		
	1992	1993	1995	1992	1993	1995
All children						
Proportion male (%)	51.40 (0.42)	50.70 (0.41)	51.17 (0.38)	51.27 (0.23)	51.11 (0.23)	51.11 (0.23)
Age	10.00 (0.03)	9.98 (0.03)	10.08 (0.03)	10.65 (0.02)	10.73 (0.02)	10.93 (0.02)
Education	1.53 (0.02)	1.57 (0.02)	1.73 (0.02)	2.52 (0.01)	2.61 (0.01)	2.84 (0.01)
Enrollment rate (%)	72.17 (0.38)	74.75 (0.36)	77.71 (0.32)	81.71 (0.17)	83.62 (0.17)	86.58 (0.16)
Labor force participation rate (%)	14.35 (0.30)	14.21 (0.29)	13.98 (0.27)	16.90 (0.17)	16.19 (0.17)	16.56 (0.17)
Male children						
Age	10.11 (0.04)	10.05 (0.04)	10.15 (0.04)	10.70 (0.02)	10.78 (0.02)	10.99 (0.02)
Education	1.45 (0.02)	1.44 (0.02)	1.58 (0.02)	2.39 (0.02)	2.48 (0.02)	2.69 (0.02)
Enrollment rate (%)	69.93 (0.54)	71.83 (0.52)	75.51 (0.46)	80.38 (0.25)	82.39 (0.24)	85.04 (0.23)
Labor force participation rate (%)	19.55 (0.47)	19.53 (0.46)	18.76 (0.42)	22.56 (0.26)	20.87 (0.26)	21.61 (0.26)
Female children						
Age	9.89 (0.04)	9.91 (0.04)	10.01 (0.04)	10.60 (0.02)	10.67 (0.02)	10.88 (0.02)
Education	1.62 (0.03)	1.70 (0.03)	1.88 (0.02)	2.66 (0.02)	2.75 (0.02)	3.00 (0.02)
Enrollment rate (%)	74.55 (0.53)	77.76 (0.49)	80.01 (0.44)	83.11 (0.24)	84.90 (0.23)	88.19 (0.21)
Labor force participation rate (%)	8.86 (0.34)	8.73 (0.33)	8.97 (0.31)	10.95 (0.20)	11.29 (0.20)	11.27 (0.21)
Observations	*14,011*	*14,520*	*17,119*	*49,201*	*49,218*	*47,747*

Notes: Standard-errors in parentheses next to estimated means.
Source: PNAD (1992-1995).