Essays on Complementarities in Bipartite Matching and in Policy Combination

by

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I, João Herculano de Carvalho Montalvão Machado, confirm that the work presented in this thesis is my own. Where information has been derived from other sources, I confirm that this has been indicated in the thesis.

Chapters 1 and 3 are sole authored papers. Chapter 2 is based on conjoint work with Nirmala Nair (Ekjut), Rajendra Mahapatra (Ekjut), Rajesh Sinha (Ekjut), Audrey Prost (Institute of Child Health UCL), Anthony Costello (Institute of Child Health UCL), Prasanta Tripathy (Ekjut), and Jolene Skordis-Worrall (Institute of Child Health UCL).

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Introduction

This dissertation contains three essays on the implications of complementarities on the equilibrium sorting in the marriage market, and on the optimal bundling of different development policies.

This first chapter develops and tests a model of marital sorting on gender-role attitudes and intrahousehold time allocations with search frictions in the marriage market, and endogenous intrahousehold bargaining power. It is shown that individuals develop a marital taste for similar gender culture partners in order to avoid conflict in decision-making within their future households. This incentive for matching assortatively is stronger for individuals anticipating little say in intrahousehold decision-making. Using data from the British Household Panel Survey, it is shown that the ability that a woman has to guide the extensive margin of her labor market supply according to her own gender-role attitudes, is entirely driven by her search for a same-attitudes partner while in the marriage market.

The second chapter provides empirical evidence on whether health education and microfinance act as substitutes or complements in reducing neonatal mortality. Identification exploits the randomized placement of a health educational intervention in rural India, stratified by the presence of a pre-existing microfinance intervention, together with the longitudinal dimension of our dataset. We find that the two interventions substituted each other: both were more effective when offered in isolation then when offered together. Further analysis shows that these interventions operated through different and substitutable channels. The health education intervention increased the adoption of hygienic health behaviours in home deliveries, whereas the microfinance intervention increased payments made to traditional birth attendants. These findings challenge the preconceived policy notion that complementarities between these two ingredients for development call for their joint supply. In contrast, they suggest that policy makers may get more out of each by offering them in isolation to their communities.

The final chapter analyses a decentralized two-dimensional marriage market model with transferable utility, where individuals' attributes are uniformly distributed on the unit square. I first show that matching of likes along both dimensions is the competitive equilibrium when the geometric average within-attribute complementarity is greater than the geometric average between-attribute complementarity. A finding that nests, as a special case, Becker's assortative matching result, and is in contrast to previous literature suggesting that the concept of assortative matching is not well defined in multi-dimensions. I then show that away from their optimal (similar-type) partners, individuals are willing to compensate mismatches on one of the attributes with opposite mismatches on the other attribute. A finding that in turn sheds new light on the trade-offs that individuals make in less than perfectly competitive multidimensional marriage markets, such as those plagued by search frictions.

1

Gender Ideology, Marriage, and Intrahousehold Specialization: Theory and Evidence from British Couples

1.1 Introduction

Women still spend substantially more time at home and less time in the market than men, despite the dramatic narrowing of the educational and wage gender gaps in the past decades.[1] Researchers are thus increasingly looking for explanations beyond the classical insight developed by Becker (1965) based on comparative advantage.[2] Within this new agenda, Fernández and Fogli (2009), and Farré and Vella (2009), provide causal

[1]For evidence on the persistent pattern of intrahousehold specialization, see e.g. Browning, Chiappori and Weiss (2011, chap. 1.2). For evidence on the narrowing of the educational and gender wage gaps, see e.g. Goldin, Katz and Kuziemko (2006) and references therein.

[2]Becker (1965) informally stated that household members who are relatively less efficient in the market sector would spend relatively more time in the domestic sector. Becker (1981, chap. 2) formally proved this proposition.

empirical evidence that culture, by shaping both the wives and the husbands preferences, plays a decisive role in explaining this persistent phenomenon of gendered intrahousehold specialization. Yet little is known on the theoretical mechanisms driving these empirical patterns.

This paper formalises the conjecture that positive assortative matching on gender-role preferences in the marriage market acts as a substitute for future intrahousehold decision-making power on the allocation of spousal times to the market and the domestic sectors. To the extent that the balance of power between spouses depends on their relative contributions to household income, and therefore on who does what within the household, this hypothesis then translates into a specific prediction that will be taken to the data. This prediction says that the extent to which the impact of a woman's attitude on the allocation of time within her household is channeled through the the choice of her partner's attitude in the marriage market is greater among more traditional households, wherein she specialises in the domestic sector and her partner specialises in the market sector.

To understand the intuition behind this idea, imagine a woman wishing to have a traditional family where she stays at home and specialises in domestic production, whereas her husband specialises in market production. Suppose further that this woman, while in the marriage market, expects to have little bargaining power *vis-a-vis* her future husband over such intrahousehold time allocation decision. Then, of course, it will be in her best interest to match with a man who shares with her a preference for traditional gender roles in the family. Matching with a similar partner will compensate for her lack of say over this decision, giving her an indirect ability to translate her preference into the intrahousehold allocation of time. From her perspective, thus, it is as if her husband will be deciding on her behalf.

But under such intrahousehold time allocation scenario, this woman's contribution to the household income will effectively pale in comparison to her husband's, and so will her decision-making power. Her expectation, while in the marriage market, of not having sufficient say within her future household, will thus turn out to be correct, justifying her effort there to find a similar partner. Conversely, her husband, having anticipated to hold most of the intrahousehold decision-making power, has had a much weaker incentive for matching assortatively while in the marriage market. In contrast to his wife therefore, his ability to translate his gender-role preference into their time allocations will be, mostly, a direct one.

Using data from the British Household Panel Survey, we find a very substantial causal effect of female gender-roles attitudes on the degree of intrahousehold specialization, which on average is *entirely* driven by the fact that more progressive females are matched to more progressive partners. To the extent that females can choose the attitudes of their partners, while searching for one in the marriage market, this result provides strong evidence in favour of our hypothesis. Females, anticipating the need to be backed by their partners over their labour market participation decisions, appear to strategically develop a (horizontal) preference for same-attitude partners in the marriage market. A rational strategy since the more similar they are, the more likely they will be to agree on the optimal decision, regardless of the balance of power between them.

Moving beyond this finding at the average to analyze its heterogeneity across the distribution of intrahousehold specialisation distribution, adds furthers support to our hypothesis. As it turns out, this pattern is exclusively concentrated among highly specialized households, in fact applying only to the extensive margin of female labour market supply. A finding that is consistent with the idea that females in the vicinity of this margin are the ones with the lowest degree of bargaining power in household decision making (due to their reduced contribution to household income), and therefore the ones who have had the most to gain from matching assortatively on attitudes once in the marriage market.

On top of its contribution to understanding the pathways through which culture, and more specifically gender ideology, determines specialisation within the household, this paper therefore also sheds light on the perceived, but mostly atheoretical, notion that individuals with similar non-productive psychological traits are attracted to each other.[3] Here, people care about the gender-role preference of the partner because in equilibrium it affects the intrahousehold time allocation, over which their own gender-role preferences are defined. Specifically, a "horizontal" taste for mating with likes on this dimension, emerges endogenously in response to the fact that who does what within the household is

[3]Dupuy and Galichon (2012) are the first to provide econometric evidence on assortative matching on personality traits in the marriage market.

the outcome of a bargaining process where both spouses' gender-role preference weigh in.

The organisation of this paper as follows. Section 1.2 develops its main idea formally with a intrahousehold bargaining model over time allocations with endogenous bargaining power embedded in a wider marriage market model. Section 1.3 introduces the dataset used, the British Household Panel Survey, and provides motivating descriptive statistics. Section 1.4 discusses the econometric framework used to test the model in the data, and presents the results. Section 1.5 discusses additional issues and concludes.

1.2 Theoretical Analysis

In this section I develop a stylised model of marital sorting on gender-role preferences followed by intrahousehold bargaining over time allocations with endogenous bargaining power. The goal is to make precise why and when, the choice of a marriage partner with compatible gender-role preferences, is a mechanism through which prospective partners improve their ability to influence the subsequent intrahousehold decisions over which partner does what, according to their own preferences.

While in the marriage market, individuals anticipate the outcome across potential partners of the subsequent intrahousehold bargaining over time allocations. The model allows for search frictions in the marriage market and assumes that partners' utilities are non-transferable. Both these features are necessary for the household decision to depend on the balance of power within the household, which is turn assumed to be endogenous.

Following Basu (2006), the model further allows for the intrahousehold bargaining power to be endogenously determined by the intrahousehold time allocation. This feature generates the novel prediction that the incentive that an individual has to match assortatively while in the marriage market is endogenously determined by the future intrahousehold time allocation, which in turn feeds into the balance of power between the spouses.

My model can thus be seen as an extension of the original static model of endogenous intrahousehold bargaining power developed by Basu (2006), in the sense that here individuals anticipate the equilibrium of that model while searching for a partner in the marriage market. Search frictions play a crucial role because they generate imperfect positive assortative matching on gender-role attitudes. In their absence, perfect assortative matching would unfold, and the model would then collapse to the standard unitary model.

The setup of the model is organized as follows. First, a simple model of intrahousehold bargaining over time allocations is introduced, where the balance of power between the spouses is itself a function of their time allocation decisions. After deriving the intrahousehold equilibrium as function of partners' preferences, I then move to the marriage market, which is populated by a continuum of men and women differing on their preferences. Anticipating the future intrahousehold outcome as a function of their own and their potential partners' preferences, each individual chooses a partner to maximize their expected marital payoff. In order to be able to empirically allow for independent variation in partners' preferences, I then explicitly allow for mismatch to occur in equilibrium due to search frictions. Finally, I derive implications of the model that will guide the empirical analysis.

1.2.1 Intrahousehold Bargaining

A household comprises a male m and a female f who can spend their time in the market sector or in the domestic sector to produce a household public good denoted by $c > 0$. Other uses of time are held fixed and the total amount of working time is normalized to unity. The household production function is given by $c = z(h_m + h_f)$, where $h_i \in [0, 1]$ denotes partner i's time spent on home production, with $i = \{m, f\}$, and $z > 0$ denotes purchased good. The household budget constraint is given by $z = w_m n_m + w_f n_f$, where $n_i = 1 - h_i$ denotes partner i's time spent in the market, and w_i his or her market wage.[4]

Who does what within the household in order to produce c is captured by the female's share of time allocated to the market sector

$$s = \frac{n_f}{n_m + n_f} \in [0, 1]. \tag{1.1}$$

If $s = 1/2$, then the partners spend the same amount of time in each sector.[5] The further

[4]This is the standard textbook framework of intrahousehold time allocations (see Browning, Chiappori and Weiss, 2011).

[5]Observe that since $n_i + h_i = 1$, there is a one-to-one relationship between the female's share of market work and her share of domestic work.

away s is from $1/2$, the more specialisation there is within the household. A $s < 1/2$ captures a 'traditional' household where the female spends relatively less time in the market and more time at home than the male. A $s > 1/2$ captures a household where these gender-roles are reversed.

Partners derive utility from two sources. First, from the amount of the household good c consumed. Second, from the level of intrahousehold specialisation s underlying the production of c. Specifically, partner i's utility function is given by

$$U_i = c - \alpha \frac{(s-\theta_i)^2}{2}. \tag{1.2}$$

where $\theta_i \in [0,1]$ and $\alpha \geq 0$. This utility function says that the quality of the household good c consumed, as perceived by partner i, is decreasing in the distance between the *realized* level of intrahousehold specialization sustaining its production, s, and this partner's *ideal* level of intrahousehold specialization, θ_i. For example, an individual with $\theta_i < 1/2$ wishes c to be produced in a way that is consistent with traditional gender-roles in the family. The closer θ_i is from 0 the more traditional this individual is. The strength of this preference over the intrahousehold specialisation is captured by α, which for simplicity is assumed to be the same for both partners.

The household chooses (n_m, n_f) in order to solve the following problem

$$\begin{aligned} \max_{n_m,n_f} \quad & \lambda\, U_f + (1-\lambda)\, U_m \;\equiv\; z(2-n_m-n_f) - \alpha\left[\lambda \frac{(s-\theta_f)^2}{2} + (1-\lambda)\frac{(s-\theta_m)^2}{2}\right] \\ \text{s.t.} \quad & z = w_m n_m + w_f n_f, \;\text{ and }\; s = n_f/(n_m+n_f), \end{aligned} \tag{1.3}$$

where $\lambda \in [0,1]$ captures the balance of power within the household. The greater λ is, the greater the female's say is in the household's decision over the choice of (n_m, n_f). Following Basu (2006), this balance of power is endogeneized. Specifically, λ is assumed to be an increasing function of the female's contribution to the household income $y = w_f n_f/(w_f n_f + w_m n_m)$, i.e. $\lambda'(y) > 0$.[6] For tractability, the following parameterisation is

[6]This departs from initial models of intrahousehold decision-making where bargaining power was assumed to be independent of the intrahousehold outcomes themselves (see e.g. Manser and Brown, 1980; McElroy and Horney, 1981; Sen, 1983; Chiappori, 1988)

adopted,

$$\lambda = a + by, \tag{1.4}$$

with $a \in (0, \frac{1}{2})$ and $b = 1 - 2a$, implying that $b \in (0, 1)$.[7] Observe that (1.4) implicitly assumes that the balance of power between the partners tilts at $y = 1/2$, i.e. $\lambda > (<)1/2$ when $y > (<)1/2$.[8]

Intrahousehold Equilibrium

To start things off, suppose for a moment that $\alpha = 0$, i.e. partners only care about the amount of c consumed.[9] In such case, there is nothing to bargain over and the solution to problem (1.3) is entirely driven by efficiency reasons, reprising Becker's result of intrahousehold specialization based on comparative advantage (Becker, 1981, chap. 2). That is, if $w_i > w_j$, then the higher wage partner fully specialises in market work, $n_i = 1$, and the lower wage partner fully specialises in domestic work, $n_j = 0$. If instead both partners have the same wage w, then from an efficiency perspective who does what within the household becomes irrelevant, as long as n_m and n_f are such that

$$n_f = 1 - n_m \equiv s. \tag{1.5}$$

In the absence of a comparative advantage thus any level of intrahousehold specialisation $s \in [0, 1]$ is equally efficient and generates the same amount of household good $c = w$. In the rest of the analysis, I will assume away comparative advantages in order to focus on the partners' gender role preferences.

Suppose now that $\alpha > 0$, i.e. partners also care about how the household good c is produced. In such case, the solution to problem (1.3) still satisfies (1.5), but now it further

[7] This parameterization of the bargaining power function rules out multiplicity of equilibria within the household. To be specific, sufficient conditions in my setting for uniqueness are $a > 0$ and monotonicity of the slope (i.e. at most a quadratic function). For multiplicity of equilibria scenarios see Basu (2006).

[8] Naturally, the exact location of this threshold level of relative income that tilts the balance of power within the household is an empirical question. What is important to retain here is that monotonicity of λ with respect to y implies that such threshold exists.

[9] In such case the framework would collapse to the standard textbook setting used to explain Becker's comparative advantage theory (see Browning, Chiappori and Weiss, 2011).

entails the following condition

$$s = \lambda\theta_f + (1 - \lambda)\theta_m. \tag{1.6}$$

That is, when the partners directly care about their relative roles in the family, they choose a specific level of intrahousehold specialisation along the efficiency schedule $s \in [0, 1]$, which is an average of their ideal levels, θ_f and θ_m, weighted by their relative bargaining powers.

But the balance of power λ depends itself on the outcome s. Specifically, given that $w_m = w_f$, the power function (1.4) simplifies to

$$\lambda = a + bs = \frac{1 + b(2s - 1)}{2}. \tag{1.7}$$

Hence, the intrahousehold equilibrium must simultaneously satisfy (1.7) and (1.6), entailing the equilibrium level of intrahousehold specialisation

$$s_{mf} = \frac{\theta_f + \theta_m - b(\theta_f - \theta_m)}{2(1 - b(\theta_f - \theta_m))}, \tag{1.8}$$

and the equilibrium balance of power

$$\lambda_{mf} = \frac{1 - b(1 - 2\theta_m)}{2(1 - b(\theta_f - \theta_m))}. \tag{1.9}$$

It can be readily verified from (1.8) and (1.9) that the ability that each partner has to translate his or her own preference into the chosen level of intrahousehold specialisation depends itself on this outcome, and therefore on their preferences. If the partners have sufficiently traditional preferences, then the male will be more influential than the female. Specifically, if $\theta_m + \theta_f < 1$, then $\partial s_{mf}/\partial\theta_m > \partial s_{mf}/\partial\theta_f$. The more traditional the couple is, the greater this imbalance in power is. This is because s is increasing in both partners' preference parameters, and therefore so it will be the female's decision-making power. If the couple is sufficiently 'progressive', then she becomes more influential than her. Specifically if $\theta_m + \theta_f > 1$, then $\partial s_{mf}/\partial\theta_m < \partial s_{mf}/\partial\theta_f$ since $s, y, \lambda > 1/2$.

Figure 1.1 illustrates the intrahousehold equilibrium (1.8) and (1.9) for an hypothetical couple (θ_m, θ_f). The horizontal axis measures the level of specialization s. The vertical axis measures the balance of power λ. The steeper line represents the specialization function (1.6). The flatter curve represents the power function (1.7). The intrahousehold equilibrium is given by the intersection of these two lines. Observe that since these partners' preferences are such that $\theta_m+\theta_f < 1$, the equilibrium s is closer to θ_m than it is from θ_f, i.e. the male has a greater ability to impose his own preference. This is a consequence of s being less than 1/2 within this household, allowing the male to be more influential than the female, i.e. $s < 1/2$.

1.2.2 Marriage Market

The marriage market is populated by a continuum of women and men whose types, indexed by θ_f and θ_m respectively, are both assumed to follow a uniform distribution on $[0, 1]$. I assume that these preference parameters are publicly observable, and hence, while in the marriage market, individuals perfectly anticipate the intrahousehold equilibrium associated with each potential choice of a partner.

Individual θ_i's marital payoff from matching with partner θ_j, V_{ij}, is thus obtained by plugging the equilibrium intrahousehold specialisation s_{ij}, given by (1.8), and the associated amount of household good $c = w$ produced, into his or her utility function, given by (1.2),

$$V_{ij} = w - \beta_{ij}\frac{(\theta_j - \theta_i)^2}{2}. \tag{1.10}$$

After simple algebra it can be shown that $\beta_{fm} = \alpha[(1 - \lambda_f^\star)/(1 - (\lambda_f^\star - \lambda_m^\star))]^2$ and $\beta_{mf} = \alpha[\lambda_m^\star/(1 - (\lambda_f^\star - \lambda_m^\star))]^2$, where $\lambda_f^\star = a + b\theta_f$ captures the female θ_f's bargaining power when matched with a male with a preference similar to her own, i.e. when her partner's type is $\theta_m = \theta_f$; and $1 - \lambda_m^\star = 1 - a - b\theta_m$ captures the male's bargaining power when matched with a female with a preference similar to his own, i.e. when his partner's type is $\theta_f = \theta_m$.

Marital Preferences

Specification (1.10) allows to fully characterise both men and women marital preferences across all potential matches in the marriage market. Three characteristics are of particular interest. First, both have an endogenous mating preference for partners with gender-role preferences similar to their own.[10] Specifically, marital payoffs (1.10) are maximized at $\theta_j = \theta_i$. To see this let $\Delta = \theta_f - \theta_m$ measure the mismatch between a female θ_f and a male θ_m. Differentiating (1.10) with respect to Δ, it follows that the marginal cost of a mismatch for a female is given by $MC_f = \Delta(1-\lambda_f^\star)^2/(1-b\Delta)^3$, and for a male it is given by $MC_m = \Delta{\lambda_m^\star}^2/(1-b\Delta)^3$, which are both equal to zero at $\Delta = 0$.[11] The intuition behind this result is that whomever they end up matching with, the resulting equilibrium level of intrahousehold specialisation will be an average of their ideals', and as a result matching with a similar partner ensures their most preferred outcome.

Second, the strength of this mating preference for similarity that women have in the marriage market is increasing in their own degree of traditionality, i.e. $-\partial MC_f/\partial\theta_f > 0$. The intuition for this result is clear. More traditional women will have in equilibrium lower bargaining power, which will reduce their ability to pull s towards θ_f, thus magnifying their loss from a mismatch. The opposite holds for men, whose cost of mismatching is decreasing in their own degree of traditionality, as confirmed by $\partial MC_m/\partial\theta_m > 0$. Traditional men will have greater bargaining power in household decision-making, making them less concerned with matching with women whose gender-role preferences are different than their own.

Third, women particularly dislike men who are more traditional than themselves, i.e. men such that $\theta_m < \theta_f$. Conversely, men particularly dislike women who are less traditional than themselves, i.e. women such that $\theta_f > \theta_f$. To confirm that this is indeed the case for women, just compare MC_f when $\Delta = \varepsilon > 0$ to than when $\Delta = -\varepsilon < 0$. It is readily verified that the latter is greater than the former. The reason for this asymmetry is because the balance of power within the household changes differently across these two types of mismatches. Figure 1.2 explains the intuition behind this result. It depicts intrahousehold equilibria involving a female θ_f matched with three alternative men: (i) one as traditional as herself, $\theta_m = \theta_f$, (ii) one more traditional than herself, $\theta'_m < \theta_f$, and (iii) another one less traditional than herself, $\theta''_m > \theta_f$.

[10] In Chiappori and Oreffice (2008), individuals also develop an endogenous horizontal preference for matching with individuals with attitudes towards fertility similar to their own. However, in their paper the marriage market is assumed to be frictionless, utilities are transferable, and bargaining power is not endogenously determined by the intrahousehold outcomes (fertility choices in their case).

[11] In addition, observe that marital payoffs are globally concave in Δ, confirming that zero mismatches are optimal.

If the woman depicted in Figure 1.2 matches with the man $\theta_m = \theta_f$, she is able to experience her ideal degree of intrahousehold specialisation $s = \theta_f = \theta_m$, and her decision-making power in such relationship is given by λ. Suppose now that she matches with the more traditional male $\theta'_m < \theta_f$. The following effects will unfold. First, if her level of say within this new relationship were to stay the same as before, λ, then the new level of specialisation taking into account the fact that this man is more traditional than the previous one, would be $\tilde{s} = \lambda\theta_f + (1-\lambda)\theta'_m < s$. However, if her share of market work reduces, so will her relative contribution to household income, and therefore a transfer of power to her new partner will take place. Under $\tilde{s}$, her new bargaining power will be $\tilde{\lambda} < \lambda$, which will in turn feed into a new decrease in s, further away from her ideal and closer to her partner's ideal. This process eventually converges to (s', λ').

Suppose now that this woman instead matches with the more progressive man $\theta''_m > \theta_f$. If her bargaining power were to stay the same as when matched with $\theta_m = \theta_f$, at λ, the allocation of time within this new household would then be $\hat{s} = \lambda\theta_f + (1-\lambda)\theta''_m > s$. But under $\tilde{s}$, her bargaining power will increase to $\hat{\lambda} > \lambda$, enabling her to offset part of the increase in s, pulling it closer to her ideal $s = \theta_f$. This process eventually converges to (s'', λ''). Clearly, the distance between s' and s is larger than that between s'' and s, and as a result the loss in her surplus from a mismatch is greater when it involves a man more traditional than herself. We could use a similar line of argument to show that the exact opposite holds for men. Men would rather mismatch with traditional women than with a progressive women, since the latter is associated with a reduction in their own bargaining power.

Search Frictions

Given these mating preferences, endogenously driven by the anticipation of the intrahousehold equilibria in the marriage market, the natural question is then: who matches with whom? In the absence of any frictions in this market, the answer would also be quite natural: individuals would maximize their marital payoffs (1.10), and the marriage market equilibrium would entail perfect positive assortativeness on gender-role preferences. That is, and individual of type θ_i would match with a partner of type $\theta_j = \theta_i$. But such outcome would jeopardise the very purpose of this paper, as a perfect correlation in partners' preferences would render impossible the task of breaking down the effect of partners' preferences on their time allocations, into their direct component and their assortative matching component.

I thus need to generate some degree of imperfect assortativeness. Following the tradition in the marriage market literature, the route I take here is to assume the presence of search frictions in this marriage market. An elegant and parsimonious way to do this is to adopt the search technology developed by Atakan (2006), and further simplified by Eeckhout and Kircher (2011).[12] Specifically, this technology assumes that the marriage market involves two rounds. In the first round, men and women are randomly paired. They then have the option to either stay together for free or, at a cost $c > 0$, move into the second round where they will meet with certainty their payoff maximising partners as in the frictionless word.

A type θ_i will thus accept to match with a type θ_j, who has randomly met with in the first round, if the payoff he or she derives in such marriage is greater than the net payoff that he or she can obtain if instead decides to enter into the second round and marry with a type $\theta_j = \theta_i$. Using (1.10), this condition becomes

$$(\theta_j - \theta_i)^2 \leq 2c/\beta_{ij}, \tag{1.11}$$

which can be solved for θ_i in order to obtain this type's acceptance set in the marriage market, $\mathcal{A}_i \subseteq [0, 1]$, that is the set of types $\theta_j \in [0, 1]$ that he or she is willing to match with.

Women's Acceptance Sets

After lengthy algebra, it can be shown that the acceptance set of a female θ_f is given by

$$\mathcal{A}_f = \{\theta_m \in [0,1] \mid \theta_f + \sigma_f - \psi_f \leq \theta_m \leq \theta_f + \sigma_f + \psi_f\}, \tag{1.12}$$

[12] A crucial aspect that sets apart my analysis from theirs, is that here I explicitly endogeneize the balance power within a partnership, whereas they assume it to be exogenous. Another important way in which my analysis differs from theirs, is that here I am working with a setup where partners' utility are non-transferable (as in Shimer and Smith (2000)), whereas in their setup partners' utilities are fully transferable.

where $\sigma_f = 2bc/\alpha[(1-\lambda_f^\star)^2 - 2b^2c/\alpha] > 0$, and $\psi_f = [(1-\lambda_f^\star)\sqrt{2c/\alpha}]/[(1-\lambda_f^\star)^2 - 2b^2c/\alpha] > 0$. The interpretation of (1.12) follows closely the one used to characterise women's marital preferences, and the following two key features are in order.

First, this female θ_f accepts an interval of men's types θ_m symmetrically defined around $\theta_f + \sigma_f$, i.e. around her own type plus a small adjustment term. That her own type θ_f is anchoring the left and right-hand sides of her acceptance set, follows from the idea that her payoff maximising partner is a man with a gender-role preference similar to her own. That, however, she adjusts upwards the centre of her acceptance set by the term σ_f, follows from the idea that, conditional on mismatching, women have a preference for men who are more progressive than themselves. It can be shown however that for small c this term is negligible.

Second, more traditional women are choosier, as their acceptance sets are narrower than those of more progressive women, i.e. $\partial\psi_f/\partial\theta_f = (\partial\psi_f/\partial\lambda_f^\star) \times (\partial\lambda_f^\star/\partial\theta_f) > 0$. This follows from the idea that more traditional women, anticipating less decision-making power within their future households, have more to loose from a mismatch in preferences.

Figure 1.3 illustrates the acceptance sets of all females in the market, that is the females' acceptance region. The concentric filled curves departing from the 45^0 degree line are the contours of the marital payoff function (1.10) for the females. Among these, the contour coinciding with the 45^0 line indicates the payoff level for women matched with men with preferences similar to their own. Hence, this is the highest possible payoff. Contours further away from the 45^0 line indicate successively smaller marital payoffs for women that are increasingly mismatched. The dashed straight lines are the contours of the equilibrium intrahousehold specialisation function (1.8). Among these, the contour coinciding with the -45^0 line indicates couples wherein there is no specialisation, i.e. combinations of $(\theta_m, \theta_f) \in [0,1]^2$ such that $s_{mf} = 0.5$. Contours further to the left of the -45^0 line indicate increasingly specialised couples wherein females spend increasingly less time in the market and more at home than their partners, $s_{mf} < 0.5$. Contours further to the right of -45^0 line indicate increasingly specialised couples, but in the opposite way, i.e. couples wherein females spend increasingly more time in the market and less time at home than their partners, $s_{mf} > 0.5$. The absolute value of the slope of a given contour k, given by $|\partial\theta_m/\partial\theta_f|_k = \lambda^\star_m/(1-\lambda^\star_f)|_k$, indicates the balance of power within these couples as measured by their partners' level of decision-making power when these are matched with their outside-option payoff-maximising partners.

Women's acceptance sets in the marriage market are given by the shaded region in Figure 1.3, which has the following two basic primitives. First, the contours of the women's marital payoffs (1.10) given by the concentric curves departing from the 45^0 degree line. Among these, the contour coinciding with the 45^0 line indicates the payoff level for women matched with men with preferences similar to their own. Hence, this is the highest possible payoff. Contours further away from the 45^0 line indicate successively smaller marital payoffs for women within couples where partners are increasingly mismatched.

Second, the contours of the equilibrium intrahousehold specialisation function (1.8), which are given by the dashed straight lines. Each of these contours gives a set of couples, within each the equilibrium time allocation entails a particular constant degree of specialisation s_{mf}. Among these, the contour coinciding with the -45^0 line indicates couples wherein there is no specialisation, i.e. combinations of $(\theta_m, \theta_f) \in [0,1]^2$ such that $s_{mf} = 0.5$. Contours further to the left of the -45^0 line indicate increasingly specialised couples wherein females spend increasingly less time in the market and more at home than their partners, $s_{mf} < 0.5$. Contours further to the right of -45^0 line indicate increasingly specialised couples, but in the opposite way, i.e. couples wherein females spend increasingly more time in the market and less time at home than their partners, $s_{mf} > 0.5$. The absolute value of the slope of a given contour k, given by $|\partial\theta_m/\partial\theta_f|_k = \lambda^\star_m/(1-\lambda^\star_f)|_k$, indicates the balance of power within these couples as measured by their partners' level of decision-making power when these are matched with their outside-option payoff-maximising partners.

Taken together with the information contained by these two contour maps, the shape of the shaded region in graph depicted by Figure 1.3, which captures the acceptance sets of all women in the market, makes geometrically precise the analytical arguments just made. Specifically, the women's acceptance region is a band that is approximately centred along the perfectly positive assortative matching schedule, whose thickness is decreasing in the degree of traditionality of women. This narrowing effect is the result of more traditional women being more choosy in anticipation of their reduced say within their future households, as captured by the flattening pattern of the iso-specialization contours as we approach the origin of the graph.

Men's Acceptance Sets

The acceptance set of men θ_m is, in turn, given by

$$\mathcal{A}_m = \{\theta_f \in [0,1] \mid \theta_m - \sigma_m - \psi_m \leq \theta_f \leq \theta_m - \sigma_m + \psi_m\}, \tag{1.13}$$

where $\sigma_m = 2bc/\alpha[\lambda_m^{\star\,2} - 2b^2c/\alpha] > 0$, and $\psi_m = [\lambda_m^\star\sqrt{2c/\alpha}]/[\lambda_m^{\star\,2} - 2b^2c/\alpha] > 0$. The interpretation of (1.13) is the inverse of that used to understand women's acceptance sets. In particular, men accept an interval of women whose types θ_f are symmetrically defined around $\theta_m - \sigma_m$, i.e. around his own type minus an adjustment term. This adjustment term captures the idea previously discussed that men, conditional on mismatching, have a preference for women who are more traditional than themselves. Again, this adjustment term can be shown to be negligible. Furthermore, traditional men, i.e. those with lower θ_m, are less choosy. This is because these type of men anticipate a rather large level of decision-making within their future households, and therefore are prepared to accept a wider range of females.

Figure 1.4 depicts the acceptance sets of all men in the marriage market, together with the contours of their marital payoffs. It can be see, that these are the mirror-image around the -45^0 line of those of women, precisely capturing the asymmetric nature of marital preferences across gender.

Marriage Market Equilibrium

Marriage is a voluntary decision, hence individuals' acceptance sets are in principal not enough to define who matches with whom. It may well be that an individual may be willing to match with a particular partner, but not the other way round.[13] As a result, such marriage should never occur. We thus need to introduce an additional concept, that

[13] As it is well known, this is the result of non-transferability between partners' payoffs (see, e.g. Smith (2006)).

of opportunity sets. The opportunity set of an individual whose type is θ_i, is the set of potential partners that are willing to match with him or her, and therefore given by $\mathcal{O}_i = \{\theta_j \in [0,1] \mid \theta_i \in \mathcal{A}_j\}$. This type's matching set, i.e. who he or she actually ends up marrying with, is given by $\mathcal{M}_i = \mathcal{A}_i \cap \mathcal{O}_i$. These matching sets across the entire population in the marriage market, thus define the marriage market equilibrium.

The solution to the marriage market equilibrium is simple, and illustrated in Figure 1.5. Amongst the segment of the marriage market population such that couples formed therein are associated with an intrahousehold balance of power tilted in favour of men, i.e. the region in the graph below the -45^0 line, who matches with whom is entirely determined by womens' acceptance region, as this lies strictly within that of the men's. On the opposite side of the population, where women within any couple that forms there have greater say than their partners, the matching region is entirely driven by the men's acceptance region.

To confirm this result, observe that from the individuals' acceptance rules given by (1.11), the largest mismatch accepted by women satisfies $(\theta_f - \theta_m)^2 = 2c/\beta_{fm}$, whereas the largest mismatch accepted by men satisfies $(\theta_f - \theta_m)^2 = 2c/\beta_{fm}$. It thus follows that women (men) are choosier than men (women) when $\beta_{fm} > (<)\ \beta_{mf}$, that is when $\lambda_m^\star < (>)\ 1-\lambda_f^\star$. Naturally, they are equally choosy, when $\lambda_m^\star = 1-\lambda_f^\star$. That is, along the knife-edge set of potential couples (θ_m, θ_f) along the -45^0 line in Figure 1.5, the women's acceptance region coincides with that of men.

1.2.3 Mapping the Theory to the Data

Recall that the goal of this paper is to argue that positive assortative matching on gender-role preferences is a mechanism through which individuals, who anticipate little bargaining power within their future households, use in order to implement their own preferences.

According to the model just developed, the marginal impact of the females' (or males') gender-role preferences on the average degree of specialisation within their households, acts through the following two additive channels

$$\frac{dE(s(\theta_f,\theta_m)|\theta_f)}{d\theta_f} = \frac{\partial s(\theta_f, E[\theta_m|\theta_f])}{\partial \theta_f} + \frac{dE[\theta_m|\theta_f]}{d\theta_f} \times \frac{\partial s(\theta_f, E[\theta_m|\theta_f])}{\partial \theta_m}. \tag{1.14}$$

The first term on the right-hand side of (1.14) is the *direct* channel. It captures the average impact of a marginal increase in her preference parameter on the degree of intrahousehold specialisation, keeping fixed the average preference parameter of her partner $E[\theta_m|\theta_f]$. The second term, is the *assortative matching* channel. It captures the following sequence of effects. First, a marginal increase in the woman's preference parameter leads to an increase in the average preference parameter of her partner, through a different choice of a partner in the marriage market. Second, this will in turn further add to the change in the average degree of intrahousehold specialisation.

It can be shown that for small search costs c, individuals are *on average* approximately perfectly positively sorted on gender-role preferences, i.e. $E[\theta_m|\theta_f] \approx \theta_f$. As a result, using the solution to the intrahousehold equilibrium specialisation s_{mf}, given by (1.8), the direct channel is approximately equal to

$$\frac{\partial s(\theta_f, E[\theta_m|\theta_f])}{\partial \theta_f} = \frac{1 - b(1 - 2E[\theta_m|\theta_f])}{2[1 - b(\theta_f - E[\theta_m|\theta_f])]^2} \approx \frac{1 - b(1 - 2\theta_f)}{2}, \tag{1.15}$$

and the matching channel is approximately equal to

$$\frac{dE[\theta_m|\theta_f]}{\partial \theta_f} \times \frac{\partial s(\theta_f, E[\theta_m|\theta_f])}{\partial \theta_m} \approx \frac{1 + b(1 - 2\theta_f)}{2[1 - b(\theta_f - E[\theta_m|\theta_f])]^2} \approx \frac{1 + b(1 - 2\theta_f)}{2}, \tag{1.16}$$

whose sum (1.14) leads to a constant overall effect $dE[s(\theta_f, \theta_m)|\theta_f]/d\theta_f \approx 1$. Yet, (1.15) and (1.16) reveal that the relative weight of its two channels varies across the distribution of women's gender-role preferences. Specifically, the difference between the assortative matching channel and the direct channe, $b(1 - 2\theta_f)$, is decreasing in θ_f. The former dominates the latter when $\theta_f < 0.5$.

The intuition for this result is simple. Among sufficiently traditional women, their ability to directly influence the intrahousehold decision-making process is limited, and therefore the impact of their gender-role preferences on the time allocation within their households is mostly channeled through their choices of partners in the marriage market, with preferences on average similiar to their own.

Further observe that since $E[\theta_m|\theta_f] \approx \theta_f$, using (1.8) we have that $E[s(\theta_f, \theta_m)|\theta_f] \approx 2\theta_f$. An alternative way to put the above prediction is then that the average impact of women's gender role preferences on the degree of intrahousehold specialisation varies across the distribution of the latter. In particular, the assortative matching channel will dominates the direct channel among more traditional couples, that is, among couples where the woman specialises in the domestic sector and the man specialises in the market sector. It is this prediction that I will take to the data.

1.3 Data and Descriptive Statistics

The data used in this study comes from the British Household Panel Survey that began in 1991 with approximately 5,500 households in Great Britain. Every year since then, adult members of these households were asked to report about how many hours in a typical week they expected to spend in their market job and in domestic chores, and every other year they were asked to report their degree of agreement with a set of six statements about the relative roles of women and men in the family and in the market. Our sample is restricted to years where both time use and attitudes data are available, and to adult respondents who where either married or cohabiting and living together with their partners. We further dropped respondents that reported to be full-time students, unemployed, or retired.

1.3.1 Intrahousehold Specialization

Table 1.1 describes the main features of the data on intrahousehold time allocations. There are clear gender asymmetries within the household in the distributions of time allocated to the market and the domestic sectors. Nearly all men work full-time in the labour market ($\geq$ 35 hours/week) and allocate no more than 10 hours a week to the domestic sector. In contrast, only 41% of women work full-time in labour market, 18% do not participate at all there, and 62% work more than 10 hours a week in the domestic sector. Specifically, during a typical week men allocate an average 38 hours to the market and 5 hours to the household, whereas women allocate 24 hours to their market jobs and 16 hours to domestic chores.

There is also an important gender asymmetry in the degree of dispersion of these distributions. The standard deviations of the women's time allocated to the market and to the domestic sectors are more than twice as large as those of their partners. Women compensate compensate additional work in the labour market with less work in the domestic sector. More precisely, a one hour increase in their market work is on average associated with a 18.2 minutes decrease in the amount of time they allocate to domestic chores.[14] The scant variability in men's time allocations implies that spousal time uses are mostly independent from each other.[15]

This picture of gender asymmetry in intrahousehold time allocations, stemming from the joint analysis of both spouses' times allocated to the market and to the domestic sectors, can be summarized into a single measure capturing the degree of intrahousehold specialization

$$s_{mf} = \left(\frac{n_f}{n_f + h_f}\right) \times \left(\frac{n_m}{n_m + h_m}\right)^{-1} \tag{1.17}$$

where n_i and h_i denote partner i's number of hours per week allocated to the market and to the domestic sectors, respectively. If $s_{mf} < 1$, we then say that the female is specialised in the domestic sector relative to her partner, and vice-versa if $s_{mf} > 1$. If $s_{mf} = 1$, then there is no specialization within this household since the own shares of working time that each spouse allocates to the market sector are equal.[16] Figure 1.6 shows the sample distribution of this measure. In 88% of the households, women are specialized in the domestic sector relative to their partners. The large mass point at zero captures households where women do not participate at all in the labour market. In the typical household, the woman's own share of market work corresponds to approximately 63% to that of her partner's.

Appendix Figure 1.A1 plots changes in s_{mf} over the female's life-cycle for each of the four major cohorts in the data. Each curve represents a cohort. We can see that these curves link to each other on a fairly continuous and smooth fashion. If there were large cohort effects, these curves should be shifted away from each other. This series is thus dominated by life-cycle effects, rather than by cohort effects. The shape the life-cycle trend show the well known U-shaped profile of female participation in the labor market.

[14] This estimate was obtained from an OLS regression of the women's time allocated to the market sector to the time they allocate to the domestic sector.

[15] This stands in contrast with the conventional theoretical approach of modelling a correlation in partners' time uses stemming from comparative advantages arguments.

[16] See Pollak (2011) for a discussion on the different meanings of specialization in economic theory and in ordinary language.

This profile reflects the impact of child bearing on labour supply, especially during the child's early years.

1.3.2 Gender-Role Attitudes

One of the six statements on gender roles reads: "a husband's jobs is to earn money, a wife's job is to look after the home and family." The five answer categories were labelled from "strongly agree" to "strongly disagree". Basic descriptives reported in Appendix Table 1.A1 and in Appendix Figures 1.A2 and 1.A3, reveal a considerable degree of heterogeneity in the answers to all the statements, most of which driven by the cross-sectional dimension of the dataset, as seen in Columns 3 and 6 of Table 1.A1. This finding indicates that attitudes are a persistent subjective trait across time, which will motivate the use of cross-sectional approaches in our econometric analysis.

It is well understood that subjective data of this type may be measured with error (Bertrand and Mullainathan, 2001). As a first step to overcome this problem, rather than using the responses to these questions directly, I employ an Item Response Theory (IRT) framework in order to estimate the underlying common component driving all of the responses.[17] Specifically, I use the IRT semiparametric estimation procedure developed by Spady (2010), whose details are summarized in Appendix 1.A.[18] The resulting individual-level index is then taken to be the measure of gender role attitudes, whose sample distribution, split by gender, is shown in the Appendix Figure 1.A4. Higher values of this index are associated with more progressive views (or, equivalently, lower values are associated with more traditional views). Figures 1.A5 and 1.A6 in the Appendix in turn, show that the variation in attitudes in the data is mostly driven by heterogeneity across marriages within cohorts rather than across cohorts These figures plot changes in both males and females' attitudes over the life-cycle for the four key cohorts in the data. Similarly to we saw for intrahousehold specialization in Figure 1.A1, these curves link to each other on a fairly continuous and smooth fashion, thus indicating that they dominated

[17]This approach of reducing the dimensionality of these measures into their common component is supported by the large Cronbach's alpha reliability coefficients reported in the footnote of Appendix Table 1.A1.

[18]Despite the advantages of this method relative to standard factor and principal component analyses (Spady, 2010), our results are robust if we instead use those alternative methods.

by life-cycle effects, rather than by cohort effects.

Figures 1.7 and 1.8 show the cumulative distribution function of our measure of intrahousehold specialization (1.17) conditional on the quartiles of each partner's attitudinal distribution. These figures clearly indicate that both more progressive women and men are more likely to come from households where there is less specialisation.[19] Figure 1.A7 in the Appendix further shows that this relationship between gender role attitudes and intrahousehold specialization holds within the different cohorts.[20] It is however possible that the observed change across time in average allocation of time within the household may be in part driven by cohort level factors, other than changes in attitudes. One way to think about this is to realize that in fact each cohort represents a different marriage market, each possibly driving a different within-household distribution of bargaining power and therefore household resources. For example, it is well known that the increased female labor force participation has been accompanied by a reduction in the wage gender gap across these cohorts. In order to mitigate the presence of cohort level confounders, it will thus be important to control for cohort fixed effects in the econometric analysis.

At same time, Figure 1.9, which plots for each value of the female attitude the average attitude of her partner, shows a strong positive pattern of correlation in spousal attitudes.[21] To the extent that the choice of a partner in the marriage market precedes the intrahousehold decision-making process, but follows the formation of these attitudes, then the idea put forward in the theoretical section of this paper - that the impact of individuals' gender-role preferences on the allocation of time within their households is, in part, channeled through their choices of spouses with gender-role preferences similar to their own - may be true.

[19] A supplementary descriptive regression analysis on the relationship between intrahousehold time allocations and attitudes is presented in Appendix 1.B.

[20] This figure focus on women's birth cohorts, but men's cohorts generate a very similar pattern.

[21] The (OLS) correlation between partners' attitudes is 0.404 (s.e.=0.008), as reported in Column 8 of Appendix Table 1.A1.

1.4 Econometric Analysis

1.4.1 Strategy

My econometric strategy is as follows. First, I estimate the total causal effect of both women's and men's attitudes on the degree of specialization within their households, i.e. without controlling for their partners' attitudes. I then control for their partners' attitudes and see how those initial coefficients change. A decrease in their magnitude will indicate that part of that total effect is channeled through their behaviour in the marriage market whereby they match with a similar-attitude partner. To incorporate the idea that this assortative matching channel depends on the balance of power within the household, and this in turn on the intrahousehold time allocations themselves, I perform this exercise across different points of the intrahousehold specialization distribution.

Clearly, the very first econometric concern is that the descriptive evidence of Section 1.3, indicating a correlation between partners' attitudes and the degree of intrahousehold specialisation, may not be necessarily causal. This is so due to the classical endogeneity concerns: reverse causality, measurement error, and omitted variables bias. Reverse causality is a strong possibility in our setting since individuals' attitudes may be partially rationalizing the actual gender roles' setting that they face within their homes in order to avoid cognitive dissonance problems (Akerlof and Dickens, 1982). Alternatively, reverse causality can be the result of a self-inference process that individuals go though if they are uncertain about their own "deep values" (Bénabou and Tirole, 2011; Bertrand and Mullainathan, 2001).

Furthermore, despite the rigorous approach taken in Section 1.3 to estimate latent attitudes from observed self-reported attitudes, I remain doubtful that those became free of measurement error. For example, traditional individuals may fear being labelled "conservative" at a time where society tends to give support to sophisticated attitudes towards gender roles. It is well known that the presence of measurement error may attenuate causal inference, a problem that is magnified if the inference is based on within-variation and the regressor of interest is persistent across time (see e.g. Griliches and Hausman, 1986). At a minimum then, in this context, preference should be given for an estimation procedure based on cross-sectional variation.

Finally, attitudes may also be correlated with unobservables that are themselves correlated with intrahousehold time allocations. An obvious example of such an omitted variable could be partners' relative productivities in the market and in the domestic sectors. Inference that ignores such possibilities may then be biased. Observe that these may also be coupled with measurement error issues. For example, females lacking labour market skills may feel a discomfort with it, and as result adopt a wishful-thinking coping mechanism where they pretend to adhere to traditional views about gender roles.

I tackle these three concerns simultaneously by employing a (cross-sectional) Instrumental Variable (IV) approach, whereby motivated by the theoretical contribution of Fernández, Fogli and Olivetti (2004) on the intergenerational transmission of gender-role preferences, I instrument respondents' attitudes with information on whether their mothers were participating in the labour market when they were growing up.[22] This approach will identify the causal effects of the spousal gender-role attitudes, if the association between these childhood experiences and the degree of intrahousehold specialisation takes place *only* through these attitudes.

One however may be skeptical about the validity of this "exclusion restriction". Indeed, one could argue that the partners' mothers labour market statuses could have created a chain of events, parallel to the formation of their attitudes, that ultimately also contributed to explain the current time allocations within their household. For example, the labour market participation decision of their mothers could have been influenced by the amount of labour market income brought home by their fathers. This background socio-economic status could have in turn directly influenced the respondents' educational investments and career aspirations. Alternatively, one could argue that a working mother provides less parenting support to her children, which again could have had a direct influence on respondents' educational performance and career prospects (see e.g., Belsky et al., 2007). In order to mitigate these threats to the identification ability of our instrument, it will therefore be important to condition the ensuing analysis on observable measures capturing

[22] A strategy also followed by Farré and Vella (2009) using the National Longitudinal Survey of Youth, to estimate the (total) causal effect of spousal gender-role attitudes on the extensive margin of female labour market supply.

these potential confounding channels through which the instrument can also affect the degree of intrahousehold specialisation.

1.4.2 The Total Effect of Spousal Gender-Role Attitudes

My first task is to estimate the total causal effect of both the wives' and the husbands' attitudes on the degree of specialization within their households. In practice, the IV approach followed here is implemented with a Two-Stage Least Squares (2SLS) model, ran separately for wives and husbands, whose second-stage equation is given by

$$s_i = \boldsymbol{\alpha}'\boldsymbol{x}_i + \gamma a_i + \varepsilon_i, \tag{1.18}$$

where s_i is the degree of intrahousehold specialisation within the respondent i's household, defined in (1.17), and $\boldsymbol{x}_i$ is a vector of control variables. These include basic controls, namely both the respondent's and his or her spouse's age, and cohort dummies; plus a set of variables addressing the previously discussed concerns on the validity of the exclusion restriction. Specifically, these additional controls are the socio-economic status of the respondent's father, the type of school he or she attended (namely, whether comprehensive or secondary modern), and the age he or she left school. The parameter of interest is γ, which captures the causal effect of respondent i's gender-role attitude a_i.[23] Both here and in the first-stage equation below, standard errors are clustered both at the individual and at the time level (Thompson, 2011; Cameron, Gelbach and Miller, 2011). Clustering at the individual level allows for error terms to be correlated within individual across time. This is particularly important given the persistency of attitudes. Clustering at the year level allows for shocks that induce correlation across individuals in a moment in time.

The first-stage equation for this 2SLS model is

$$a_i = \boldsymbol{\delta}'\boldsymbol{x}_i + \pi z_i + \eta_i \tag{1.19}$$

where z_i is a dummy variable that equals 1 if the mother of individual i worked when he

[23]Observe that none of the variables in specification (1.18) has a time subscript because the instrumental variable used only varies across, and not within, individuals. Furthermore, a cross-sectional analysis in our context seems appropriate in light of our previous descriptive finding indicating that most of the sample variation in attitudes is cross-sectional.

or she was 14, i.e. the instrumental variable. Standardized estimates of π, the first-stage effect of the instrument on the respondent's attitude, are reported in Panels A and B of Table 1.2, for wives and husbands respectively. These estimates show a substantial and significant positive relationship between the instrument and the respondent's attitude. On average, a woman whose mother was a labour market participant when she was aged 14 is .165 standard deviations (hereafter, σ) more progressive than a female whose mother was fully specialised in the domestic sector (s.e.=0.018). The instrument is rather strong as suggested by the very large F-statistic reported at the bottom of the table. A similar result holds for men. On average, men whose mothers worked when they were 14 years of age are a substantial $.125\sigma$ more progressive than men whose mothers did not work.

The 2SLS results are reported in Panels A and B of Table 1.3, for wives and husbands respectively. Column 1 shows that a σ increase in the wife's attitude leads to an increase of 26 percentage points (s.e.=4.3) in her own share of market work relative to her spouses' own shares of market work. That is, it virtually eliminates the sample mean degree of intrahousehold specialisation of 62.2%, which is reported at the top of the column. A similar effect, both in terms of magnitude and significance, is found for the husband's attitude. [24]

Moving beyond this average effect, Columns 2-5 focus on the total effects of both the wives' and husbands' attitudes on the likelihood that degree of specialisation within their households exceeds some particular threshold. That is, each of these columns takes $1[s > q]$ for consecutively higher choices of q as the outcome variable in the second-stage Eq.(1.18). Column 2 presents estimates for the case where $q = 0$, i.e. for the total effect of both spouse' attitudes on the extensive margin of the wife's labour market supply. We can see that the likelihood of the wife entering the labour market is significantly estimated to increase by 21.6 percentage points in response to a σ increase in own attitude (s.e.=4.6), and by 32% in response to a σ increase in her partner's attitude (s.e.=5.8).[25]

In the remaining columns, the threshold values of q are chosen in order to break

[24] These are estimates of the causal effect of female attitudes for females whose attitudes were influenced by whether their mothers worked when they were aged 14.

[25] The result found here that the impact of the husband's attitude is about 10 percentage points higher than that of his wife's, on the extensive margin of her labour market supply, departs from from that obtained by Farré and Vella (2009) for the US context. There they find no substantial difference between these two effects, both fluctuating around the effect I find here for the wife's attitude.

down the sample distribution of the intrahousehold specialisation measure, s, across its four quartiles. That is, Columns 3, 4, and 5, focus on the likelihood that s is above its 25th, 50th, and 75th percentile, respectively. Overall, the results suggest that the total causal effect both spouses' attitudes are not limited to the extensive margin of the wife's labour market supply. They appear to be persistent across the entire distribution of intrahousehold specialisation.

1.4.3 The Assortative Matching Channel

My next task is to understand the extent to which these total causal effects are being channeled through the respondents' choices of spouses with gender-role attitudes similar to their own. In order to isolate the *direct* component of those effects, that is the part that takes place over and above the one that occurs through positive assortative matching on attitudes in the marriage market, I use a multiple IV strategy. Specifically, I employ a 2SLS procedure, whose second-stage equation is given by

$$s_i = \boldsymbol{\beta}'\boldsymbol{w}_i + \gamma_1 a_{1i} + \gamma_2 a_{2i} + \mu_i, \tag{1.20}$$

where s_i is defined as before; and $\boldsymbol{w}_i$ now includes the previous vector of control variables but for both spouses. The parameters of interest are γ_1 and γ_2, where γ_1 captures the (direct) causal effect of the wife's attitude a_{1i} conditional on her husband's attitude, and γ_2 captures the (direct) causal effect of the husband's attitude a_{2i} conditional on his wife's attitude.

The two first-stages of this new 2SLS procedure are given by

$$a_{1i} = \boldsymbol{\xi}'\boldsymbol{w}_i + \pi_1 z_{1i} + \pi_2 z_{2i} + \upsilon_i \tag{1.21a}$$

$$a_{2i} = \boldsymbol{\zeta}'\boldsymbol{w}_i + \psi_1 z_{1i} + \psi_2 z_{2i} + \nu_i \tag{1.21b}$$

where z_{1i} and z_{2i} are, respectively, dummies capturing the wife's and the husband's mother work status when they were aged 14.[26] Estimates of π_1, π_2, ψ_1, and ψ_2, are reported in

[26] Observe that if sorting on attitudes was perfectly positive assortative then partners' attitudes in the second-stage regression would be perfectly multicolinear, and as a result the second-stage regression could not be estimated. Also, if sorting on family experiences is very strong our instruments will be highly

Panel C of Table 1.2. The two instruments are positively and significantly correlated with both partners' attitudes. That an individual's attitude is correlated with her or his partner's instrument is a symptom of positive assortative matching on attitudes.[27] The instruments, both in isolation and jointly, appear to be sufficiently strong as suggested by the Angrist-Psichke and the Kleibergen-Paap F-statistics, respectively, reported at the bottom of the table.

The 2SLS results are reported in Panel C of Table 1.3. Column 1 shows that when controlling for both spouses' attitudes, the previously estimated coefficients measuring the average total effects of the attitudes of each spouse, reported in Panels A and B, decrease and become statistically insignificant. This result seems to provide some initial support to the conjecture that part of those *raw* effects are intermediated by positive assortative matching on gender-role attitudes. The remaining columns, however, provide the strongest piece of evidence that this conjecture indeed seems to hold. We can see that the extensive margin of the wive's labour market supply is entirely driven by their husbands' attitudes. However, as we look at the impacts across more egalitarian patterns of intrahousehold time allocations, the impact of the wife's attitude increasingly gains dominance over that of her partner. Indeed, on the likelihood that her own share of market work is above 75% of the couples, is entirely driven by her attitude.

1.4.4 Robustness

Dealing with Censoring

The accuracy of our main results can be criticized on the basis that the linear models from which they were obtained are inappropriate given the censoring nature of the distribution of intrahousehold specialisation, our main outcome (but see Angrist and Pischke, 2009, pg. 75). However, as we now show, the essence of our results is unchanged when we employ the censored quantile instrumental variable (CQIV) procedure developed by Chernozhukov, Kowalski and Fernández-Val (2011), which is specifically tailored to deal simultaneously with censoring and endogeneity.

correlated and as a result they would be weak.

[27]In fact if we regress each partner's attitude on both instruments plus the partner's attitude, this "cross"-correlation between an individual's attitude and his or her partner's instrument vanishes.

As before, we first use this alternative econometrics framework to estimate the total effect of female attitudes, to then confront them with those obtained conditional on her partner's attitudes. Specifically, the first set of results are obtained from the following model

$$Q_\tau(y_i|\boldsymbol{x}_i, a_{1i}) = \max\{0, \boldsymbol{\alpha}'_\tau \boldsymbol{x}_i + \gamma_\tau a_{1i} + \rho_\tau \eta_i\}, \tag{1.22}$$

where η_i is replaced by the residuals of the following regression

$$a_{1i} = \boldsymbol{\delta}' \boldsymbol{x}_i + \pi z_{1i} + \eta_i, \tag{1.23}$$

and all the remaining variables and parameters are the same as before.[28] The second set of CQIV estimates, i.e. those pertaining to the causal effect of female attitudes over and above positive assortative matching are obtained from the following extension of (1.22)

$$Q_\tau(y_i|\boldsymbol{x}_i, a_{1i}, a_{2i}) = \max\{0, \boldsymbol{\beta}'_\tau \boldsymbol{w}_i + \gamma_{1\tau} a_{1i} + \gamma_{2\tau} a_{2i} + \boldsymbol{\rho}'_\tau \boldsymbol{\eta}_i\}, \tag{1.24}$$

where now $\boldsymbol{\eta}_i = (\upsilon_i \quad \nu_i)'$ is replaced by the residuals of the following regressions

$$a_{1i} = \boldsymbol{\xi}' \boldsymbol{w}_i + \pi_1 z_{1i} + \pi_2 z_{2i} + \upsilon_i \tag{1.25a}$$

$$a_{2i} = \boldsymbol{\zeta}' \boldsymbol{w}_i + \psi_1 z_{1i} + \psi_2 z_{2i} + \nu_i, \tag{1.25b}$$

and all the remaining variables and parameters are the same as in Section 1.4.3.

The results for the first and second set of estimates are reported in panels A and B, respectively, of Table 1.4. It shows estimates at the mean and at the bottom and top quartiles of the intrahousehold specialization distribution. Consistent with our previous findings, we can see that the total causal effect of female attitudes is solely driven by the fact that more progressive females are matched with more progressive males. This result appears to be mostly concentrated on the lower part of the distribution, which is in line with the theoretical prediction that females therein strategically develop a hor-

[28] Observe that in contrast to the 2SLS procedure, the CQIV method follows a "control function" approach. Yet their key identification assumptions are similar in spirit. By controlling for η_i in the censored quantile regression (1.22), the CQIV isolates the part of the variation in attitudes that is exogeneously driven by the instruments.

izontal matching preference for same-attitude partners to compensate for their reduced intrahousehold bargaining power.

Alternative Instruments

As explained in Section 1.4.1, mother's labour supply may affect current intrahousehold time allocations through channels other than the formation of gender role attitudes. For that reason, I have controlled for a number of covariates that can be expected to capture these additional channels: the socio-economic status of the respondent's father, the type of school he or she attended, and the age he or she left school. In this section, I examine how the results change when I use an alternative instrumental variable for gender-role attitudes. This alternative instrument is an index formed from answers to different questions in the survey that try to elicit respondents' attitudes towards homophobic and family values.[29]

The analysis follows the same structure as before. First, I estimate the total causal effect of both women's and men's gender-role attitudes on the degree of specialization within their households, i.e. without controlling for their partners' attitudes. I then control for their partners' attitudes and see how those initial coefficients change. The first and second stage results are reported in the Appendix Tables 1.A3 and 1.A4, respectively. As expected, the first-stage results shows that gender-role attitudes, the endogenous variables, are highly correlated with attitudes towards family values, the instruments. Now, however, the second-stage results are not consistent with my earlier findings. Specifically, we see that the coefficients on each partner's gender-role attitudes always have a significant and substantial impact independently of whether the other partner's gender-role attitude is kept constant.

How to interpret this? One hypothesis is that the theoretical model developed in Section 1.2 is not consistent with the data. Another hypothesis is that these alternative

[29]These questions ask the respondent's degree with the following six statements: (i) "homosexual relationships are wrong", (ii) "cohabiting is alright", (iii) "divorce better than unhappy marriage", (iv) "parents ought to stay together for children", (v) "marital status is irrelevant to children", and (vi) homosexual relationships are wrong". Possible answers were: "strongly agree", "agree", "neither agree nor disagree", "disagree", and "strongly disagree". I constructed an index capturing the common variation in the answers to these questions using the Item Response Theory framework utilized to construct the index of gender-role attitudes (explained in in Appendix 1.A). These questions were first asked in the eighth wave of the BHPS, and in alternate years to the questions eliciting gender-roles attitudes. As a result, for given year in our sample we use the respondents' answers to these questions that were given in the previous year, when available.

instruments are not appropriate. Indeed, one can make the case that it is harder for this alternative instrumental variable to pass the exclusion restriction assumption. This is because gender-role attitudes are likely not developed in isolation from views about homophobic and family values, but rather they are likely closely intertwined and stem from a common broad-based belief system. That is, this instrument may be measuring essentially almost the same thing as the endogenous variable itself. As a result, all the endogeneity concerns related to gender-role attitudes also hold for this alternative instrument.[30]

1.5 Conclusion

This paper has argued, theoretically and empirically, that an important component of the woman's overall labour market participation strategy is the choice of her partner's culture. The key insight is that a woman with greater career aspirations chooses to match with a more progressive man in order to overcome possible conflict within the household on time allocation decisions. This in turn explains the observed pattern of positive assortative matching on gender-role attitudes in the marriage market.

On a more general level, my analysis emphasis the importance that the reduction of search frictions in the marriage market may have in the prevention of intrahousehold conflict, when prospective partners have heterogeneous preferences over future marital outcomes. In a perfectly competitive marriage market, spouses would be perfect assortatively matched with respect to their preferences, and as a result they would decide and behave according to the standard unitary model where partners share the same preferences over intrahousehold economic outcomes.

My model assumes non-transferable utility, implying that there is no marriage market price. An alternative modeling strategy could be to explicitly allow for transfers within the household. In such setting, given their types, partners would jointly choose their time allocations, which would in turn determine the match surplus in utility terms. While in the marriage market, agents would then chose with which type to match with based on the match surplus when there are transfers that are agreed upon. In this alternative

[30] In fact, the results obtained with these alternative instruments are very similar to the ones obtained without instrumenting gender-role attitudes.

setting, the sharing rule would be also endogenously determined but now as a transfer that clears the marriage market. An interesting research strategy within this setup would be to devise an empirical framework that would allow identification and estimation of preferences towards gender roles based on observed intrahousehold time-allocations.

1.A Appendix: Estimation of Latent Gender-Role Attitudes

This section briefly describes the Item Response Theory method used to estimate latent gender role attitudes using the observed responses to the six attitudinal statements presented in Table 1.A1 and Figures 1.A2 and 1.A3. The starting point of this method, is the intuitive and plausible assumption that more progressive respondents are more likely to give more progressive responses. Formally, this can be stated as follows. Let θ denote an individual's underlying latent attitude, where higher values capture more progressive views about gender roles, and consider two individuals i and j such that $\theta_j > \theta_i$. Then, letting $r_k \in \{1, ..., 5\}$ denote a possible response to a particular statement $k = \{1, ..., 6\}$, this assumption says that $p(r \leq z|\theta_j) \leq p(r \leq z|\theta_i)$ with $z \in \{1, ..., 5\}$.

Now, assuming independence of the responses to the different statements conditional on θ, we have that the probability of the joint response $\boldsymbol{r} = (r_1, ..., r_6)$ is given by $p(\boldsymbol{r}) = \int \prod_{k=1}^{6} p_k(r_k|\theta) f(\theta) d\theta$. From this, we can use Bayes' Theorem to infer the posterior distribution of attitudes conditional on the observed responses, $f(\theta|\boldsymbol{r}) = f(\theta, \boldsymbol{r})/p(\boldsymbol{r}) = p(\boldsymbol{r}|\theta) f(\theta)/p(\boldsymbol{r})$. Setting $f(\theta)$ to follow a $N(0, 1)$ distribution, this posterior object is then estimated with a Sieve Maximum Likelihood procedure (as derived in Spady, 2010). Finally, taking the average value of this estimated distribution for each individual in the data we obtain the attitude measure that is used throughout the paper, and whose density conditional on gender is graphed in Figure 1.A4.

1.B Appendix: Descriptive Regression Analysis

This section performs a descriptive estimation analysis of the relation between different measures of time use within household i surveyed in year t, denoted by y_{it}, and the attitude of the female in that household, a_{1it}. Specifically, we report estimates of variants of the following equation

$$y_{it} = \boldsymbol{\beta}' \boldsymbol{x}_{it} + \gamma a_{1it} + \varepsilon_{it} \tag{1.26}$$

where vector $\boldsymbol{x}_{it}$ contains partners ages and cohort dummies. The parameter of interest is γ, which depending on the dependent variable used, captures the correlation between female attitudes and a particular intrahousehold time-use variable.

Appendix Table 1.A2 report the results. Panel A shows estimates obtained from pooled Ordinary Least Squares (OLS) applied to Eq.(1.26), thus capturing the overall sample correlation between female attitudes and intrahousehold time allocations. Standard errors in this panel are adjusted for heteroskedasticity and serial correlation. Panels B and C break down this overall correlation into its cross-sectional and longitudinal dimensions. Specifically, Panels B and C report estimates obtained from a between effects model and a fixed effects model, respectively, applied to Eq.(1.26).

Female attitudes are significantly correlated with their own allocation of time, positively so to the market sector and negatively so to the domestic sector. The magnitude of these correlations is substantial. In contrast, there is very little evidence of a substantial correlation between females' attitudes and their partners' time allocations. Finally, but not surprisingly, given that most of the sample variation in attitudes is driven by cross-sectional heterogeneity, as shown in Columns 3 and 6, the overall pattern of correlation between attitudes and intrahousehold time allocations is also mostly driven by its cross-sectional dimension. A finding that justifies the use of cross-sectional approaches in Section 1.4.

The results obtained from this descriptive analysis are virtually identical if instead of conditioning on the attitude of the female we condition on her partner's attitude (not shown).

1.C Appendix: Tables and Figures

Table 1.1: Intrahousehold time allocations (hours/week)

	Wives			Husbands			Total (household)			Wives' shares		
	Mean (1)	SD (2)	BV (%) (3)	Mean (4)	SD (5)	BV (%) (6)	Mean (7)	SD (8)	BV (%) (9)	Mean (10)	SD (11)	BV (%) (12)
Market work*	23.97	14.50	.656	38.09	6.53	.622	62.06	16.01	.721	.350	.197	.712
Domestic work	15.95	9.68	.672	5.03	4.52	.635	20.98	10.38	.664	.739	.197	.714
Total	39.92	13.28	.647	43.11	7.67	.638	83.04	15.89	.659	.471	.113	.636
Market Share	.546	.306	.723	.879	.131	.607	.742	.129	.702			

Notes: Data pooling waves 1, 3, 5, 7, 9, 11, and 12 of the BHPS. All sample members for whom data on the partner is also available in any wave, totaling 12,381 observations. Number of observations is 13,183. *Fraction of wives that participate in the labour market is 0.823 (s.d.=0.381). The coefficient of between variation in Columns 3, 6, 9, and 12, for each variable is the sum of squares of differences between individual means and the whole-sample mean, divided by the sum over all individuals and years of the square of the difference between each observation and the mean. P-values on the mean differences between spouses' time allocated to both activities not reported but all equal 0.000.

Table 1.2: First-stage estimates

Dependent variable:	Wife's attitude	Husband's attitude
	(1)	(2)
A. First-stage for the total effect of the wife's attitudes		
Wife's mother worked when she was 14	.165***	-
	(.018)	
A-P F test	87.09	-
B. First-stage for the total effect of the husband's attitudes		
Husband's mother worked when he was 14	-	.125***
		(.017)
A-P F test	-	78.95
C. First-stage for the partial effects of both spouses' attitudes		
Wife's mother worked when she was 14	.163***	.099***
	(.018)	(.018)
Husband's mother worked when he was 14	.094***	.121***
	(.019)	(.018)
A-P F test	13.25	12.02
K-P W F test	9.05	
S-Y weak id 10% critical value	7.03	

Notes: *** denotes significance at 1%; ** at 5%; * at 10%. 12,381 observations in all regressions. Standard errors robust to arbitrary heteroskedasticity and arbitrary autocorrelation. The A-P F statistic is the weak identification Angrist-Pischke F statistic. The K-P F statistic is the weak identification Kleibergen-Paap rk Wald F statistic. The three regressions control for both spouses' ages and cohort dummies. On top of these basic controls, the regression in Panel A also controls for the wife's father socio-economic status, the type of school she attended, and the age at which she left school; the regression in Panel B does the same but for the husband; the regression in Panel C does the same for both spouses.

Table 1.3: 2SLS estimates

Dependent variable	s_{mf}	$1[s_{mf} > 0]$	$1[s_{mf} > .5]$	$1[s_{mf} > .75]$	$1[s_{mf} > .9]$
Mean	*.625*	*.823*	*.703*	*.460*	*.243*
	(5)	(1)	(2)	(3)	(4)
Panel A: Total effect of women's attitudes					
Woman's attitude	.260***	.216***	.258***	.328***	.248***
	(.043)	(.046)	(.055)	(.061)	(.050)
Panel B: Total effect of men's attitudes					
Man's attitude	.263***	.320***	.293***	.270***	.165***
	(.053)	(.058)	(.066)	(.069)	(.057)
Panel C: Partial effects of spousal attitudes					
Woman's attitude	.092	.026	.132	.330**	.293**
	(.101)	(.110)	(.131)	(.148)	(.130)
Man's attitude	.257***	.312***	.222	.021	-.050
	(.100)	(.147)	(.157)	(.124)	(.132)

Notes: *** significant at 1%; ** significant at 5%; * significant at 10%. 12,381 observations in all regressions. The underlying dependent variable, s_{mf}, is the female's own share of market work divided by her partner's own share of market work. Standard errors robust to arbirtraty heteroskedasticity and arbirtrary autocorrelation. All regressions control for partners' ages, cohort dummies, each partner's father socio-economic status, type of school attended by each partner, and the age at which each partner left school. Robust standard errors are reported in parentheses.

Table 1.4: Censored quantile IV estimates

Quantile:	.25	.5	.75
	(1)	(2)	(3)
Sample quantile	*.425*	*.719*	*.896*
Panel A: Total effect of female attitudes			
Female's attitude	.481***	.251***	.141***
	(.088)	(.045)	(.028)
Observations		12,381	
Panel B: Partial effects			
Female's attitude	-.023	.050	.044
	(.398)	(.216)	(.095)
Partner's attitude	.512***	.253	.134
	(.171)	(.432)	(.111)
Observations		12,381	

Notes: *** denotes significance at 1% level, ** at 5% level, * at 10% level. The dependent variable is the female's own share of market work divided by her partner's share of market work. Standard errors, reported in parentheses, are bootstrapped using 200 replications.

Table 1.A1: Self-reported attitudes

	Wives			Husbands			Difference	Correlation
	Mean	SD	BV (%)	Mean	SD	BV (%)	p-value	OLS estimate
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Husband should earn, wife stay at home	3.887	.962	.704	3.665	.977	.685	[.000]	.275*** (.009)
Pre-school child suffers if mother works	3.186	1.060	.700	2.901	1.037	.670	[.000]	.341*** (.008)
Family suffers if mother works full-time	3.221	1.105	.712	3.134	1.047	.678	[.000]	.325*** (.009)
Woman and family happier if she works	2.886	.750	.570	2.855	.757	.546	[.000]	.129*** (.009)
Husband and wife should both contribute	3.510	.909	.649	3.461	.898	.614	[.000]	.218*** (.009)
Full time job makes woman independent	3.082	.988	.637	3.157	.913	.595	[.000]	.192*** (.010)
Attitude (Estimated index)	.087	1.564	.768	.0093	1.543	.730	[.177]	.404*** (.008)

Notes: Possible answers to all statements were 1 “strongly agree”, 2 “agree”, 3 “neither agree nor disagree”, 4 “disagree”, and 5 “strongly disagree”. Answers to the first three statements were reversed such that higher answers to every question indicate less traditional views about gender roles. The Cronbach alpha statistic 0.7068 for females and 0.6910 for males. The attitude index is the estimated latent obtained from the IRT procedure. The coefficient of between variation BV in Columns 3 and 6 for each variable is the sum of squares of differences between individual means and the whole-sample mean, divided by the sum over all individuals and years of the square of the difference between each observation and the mean. The pairwise correlation measures reported in Column 8 are obtained from OLS regressions of the female’s answers to each question against those of her partner, and the corresponding standard errors are robust to arbitrary heteroskedasticity and arbitrary auto-correlation.

Table 1.A2: Descriptive regression estimates on intrahousehold time allocations

	Wife's Time Use				Husband's Time Use			Intrahousehold Specialization
Dependent variable	Works in the market [=1 if yes] (%)	Market time (hours/week)	Domestic time (hours/week)	Market share [=(2)/[(2)+(3)]] (%)	Market time (hours/week)	Domestic time (hours/week)	Market share [=(5)/[(5)+(6)]] (%)	Female relative market specialization [=(4)/(7)] (%)
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Sample mean	*.823*	*24.043*	*15.952*	*.547*	*38.667*	*5.027*	*.891*	*.626*
[standard deviation]	*[.381]*	*[14.493]*	*[9.666]*	*[.306]*	*[4.653]*	*[4.505]*	*[.086]*	*[.363]*
				Panel A: Pooled OLS				
Wife's Attitude	.089***	4.315***	-1.885***	.087***	-.054	.412***	-.008***	.110***
	(.003)	(.122)	(.084)	(.003)	(.042)	(.040)	(.001)	(.003)
R-squared	.0645	.0921	.0758	.0900	.0065	.0171	.0173	.0898
				Panel B: Between Effects				
Wife's Attitude	.101***	4.797***	-2.152***	.100***	-.062	.470***	-.009***	.125***
	(.006)	(.216)	(.140)	(.005)	(.078)	(.067)	(.001)	(.006)
"Between" R-squared	.0745	.1114	.1193	.1149	.0131	.0190	.0167	.1142
				Panel C: Fixed Effects				
Wife's Attitude	.041***	1.415***	-.527***	.032***	-.024	.063	-.002*	.037***
	(.005)	(.159)	(.118)	(.003)	(.056)	(.057)	(.001)	(.004)
"Within" R-squared	.0113	.0103	.0046	.0109	.0028	.0018	.0026	.0100

Notes: *** denotes significance at 1% level, ** at 5%, * at 10%. Each column corresponds to one regression. Number of observations in all regressions is 12,381. All regressions control for spousal ages and cohort dummies. Standard errors reported in parentheses. In panel A, standard errors are adjusted for household clustered heteroskedasticity and autocorrelation.

Table 1.A3: First-stage estimates with alternative instruments

Dependent variable:	Wife's attitude	Husband's attitude
	(1)	(2)
A. First-stage for the total effect of the wife's attitudes		
Wife's family values	.306***	-
	(.007)	
A-P F test	2051.86	-
B. First-stage for the total effect of the husband's attitudes		
Husband's family values	-	.326***
		(.014)
A-P F test	-	524.50
C. First-stage for the partial effects of both spouses' attitudes		
Wife's family values	.286***	.059**
	(.009)	(.012)
Husband's family values	.033*	.293***
	(.012)	(.018)
A-P F test	758.40	234.33
K-P W F test	396.06	
S-Y weak id 10% critical value	7.03	

Notes: *** denotes significance at 1%; ** at 5%; * at 10%. 5,686 observations in all regressions. Standard errors robust to arbitrary heteroskedasticity and arbitrary autocorrelation. The A-P F statistic is the weak identification Angrist-Pischke F statistic. The K-P F statistic is the weak identification Kleibergen-Paap rk Wald F statistic. The three regressions control for both spouses' ages and cohort dummies. On top of these basic controls, the regression in Panel A also controls for the wife's father socio-economic status, the type of school she attended, and the age at which she left school; the regression in Panel B does the same but for the husband; the regression in Panel C does the same for both spouses.

Table 1.A4: 2SLS estimates with alternative instruments

Dependent variable	s_{mf}	$1[s_{mf} > 0]$	$1[s_{mf} > .5]$	$1[s_{mf} > .75]$	$1[s_{mf} > .9]$
Mean	*.625*	*.823*	*.703*	*.460*	*.243*
	(5)	(1)	(2)	(3)	(4)
Panel A: Total effect of women's attitudes					
Woman's attitude	.133***	.087***	.143***	.142***	.128***
	(.009)	(.007)	(.023)	(.008)	(.011)
Panel B: Total effect of men's attitudes					
Man's attitude	.132***	.099***	.126***	.149***	.127***
	(.005)	(.010)	(.013)	(.011)	(.007)
Panel C: Partial effects of spousal attitudes					
Woman's attitude	.084***	.051***	.098***	.084***	.072***
	(.009)	(.010)	(.024)	(.006)	(.024)
Man's attitude	.080***	.060***	.066***	.099***	.090***
	(.006)	(.011)	(.005)	(.016)	(.016)

Notes: *** significant at 1%; ** significant at 5%; * significant at 10%. 5,686 observations in all regressions. Standard errors robust to arbirtraty heteroskedasticity and arbitrary autocorrelation. All regressions control for partners' ages, cohort dummies, each partner's father socio-economic status, type of school attended by each partner, and the age at which each partner left school. Robust standard errors are reported in parentheses.

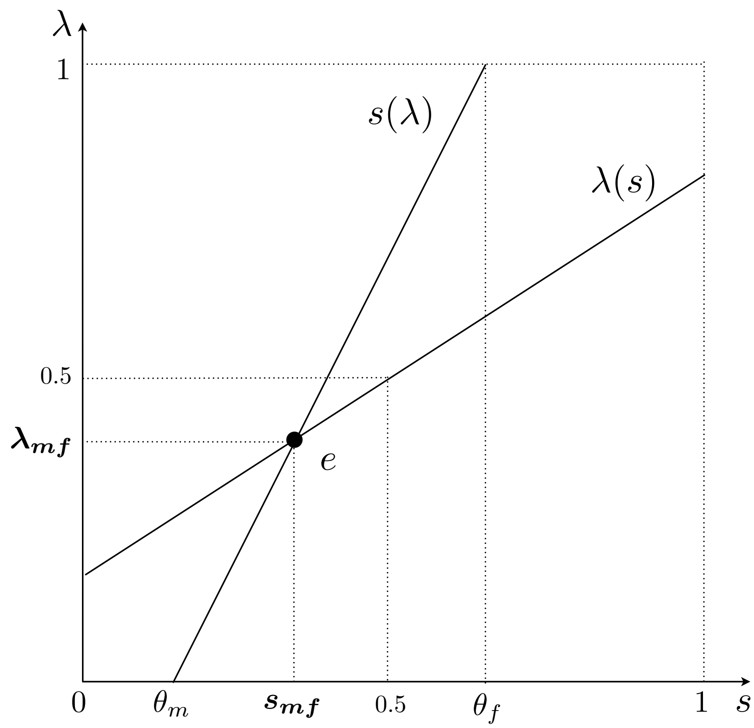

Figure 1.1: Intrahousehold equilibrium

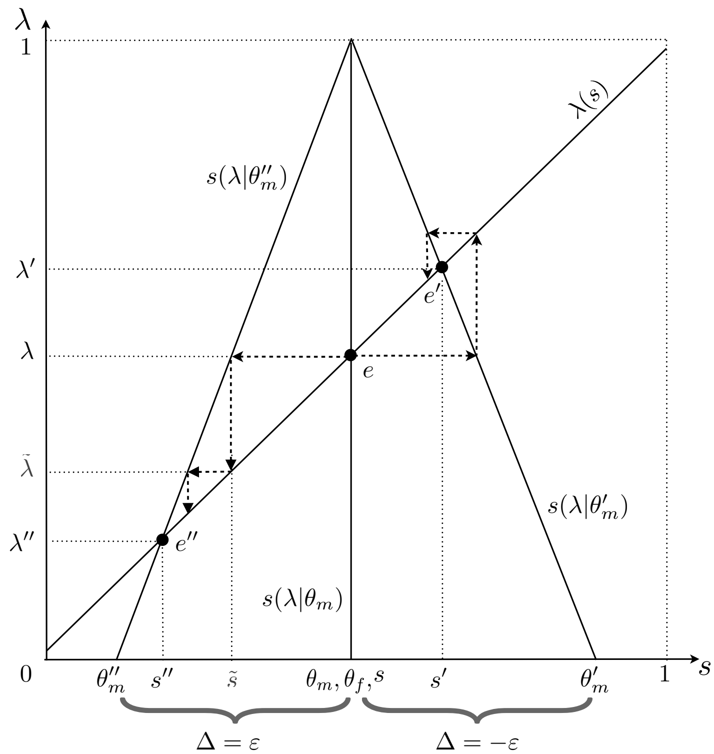

Figure 1.2: Change in intrahousehold equilibria induced by a positive and a negative mismatch

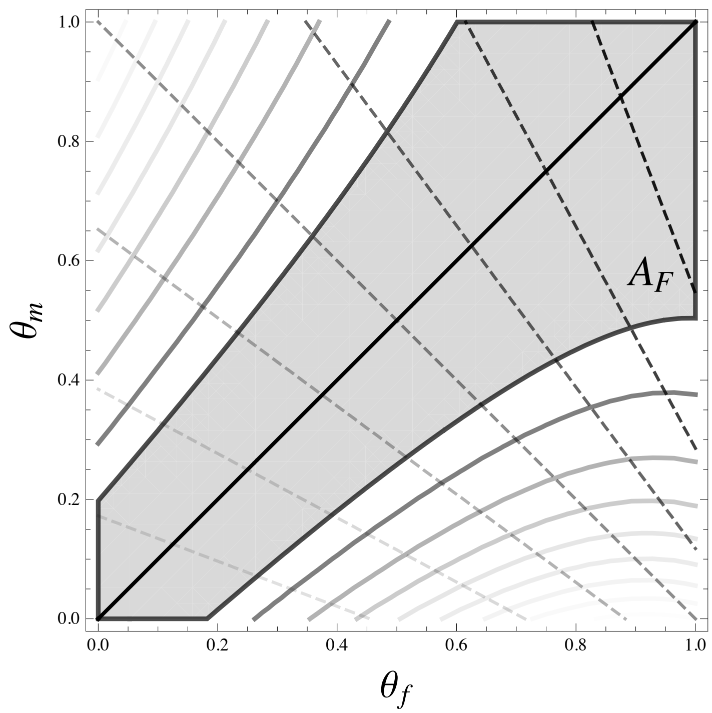

Figure 1.3: Women's acceptance sets

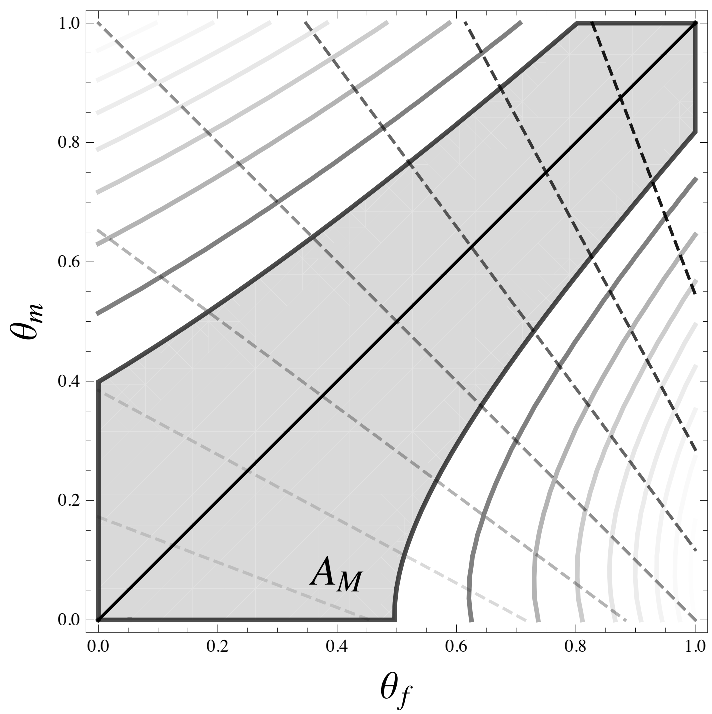

Figure 1.4: Men's acceptance sets

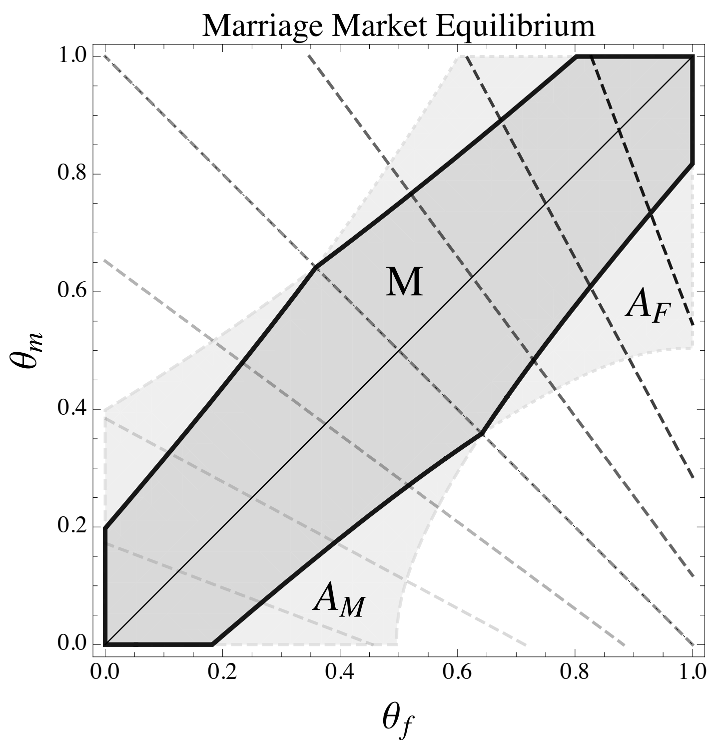

Figure 1.5: Matching sets

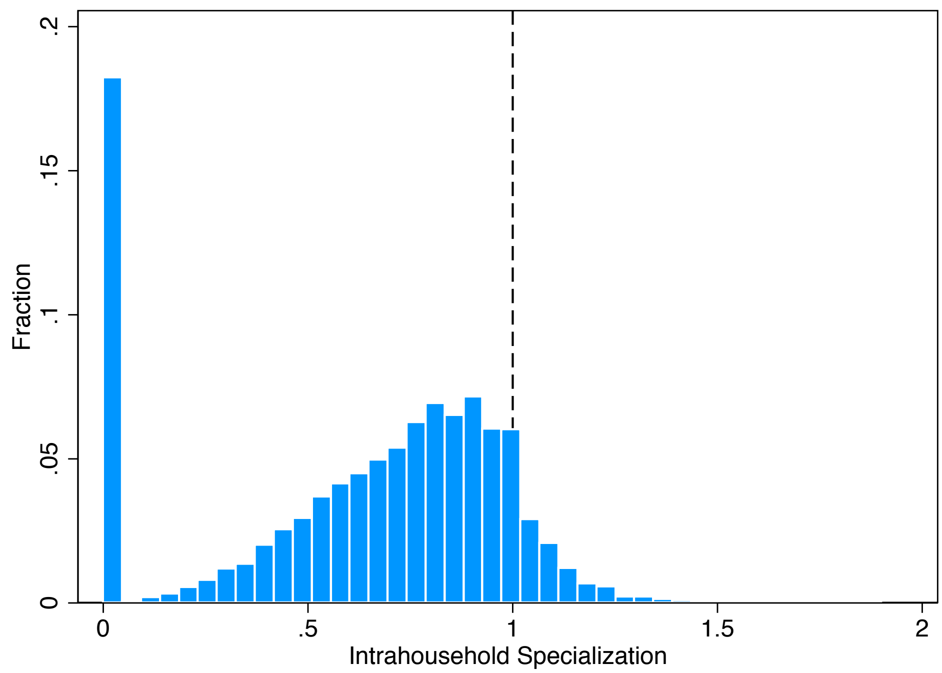

Figure 1.6: Distribution of intrahousehold specialization

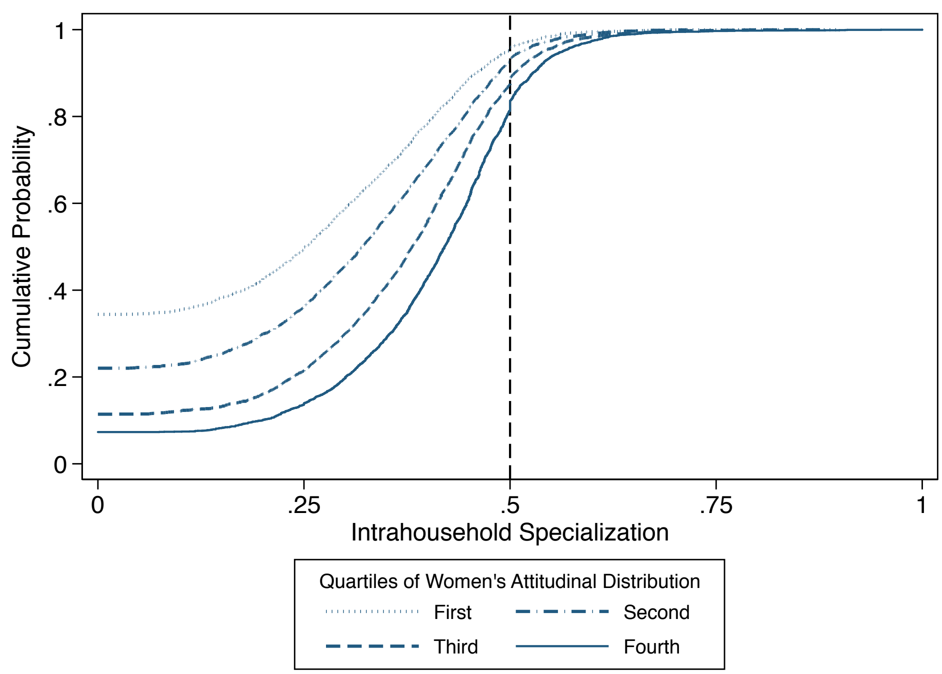

Figure 1.7: Intrahousehold specialization and women's gender-role attitudes

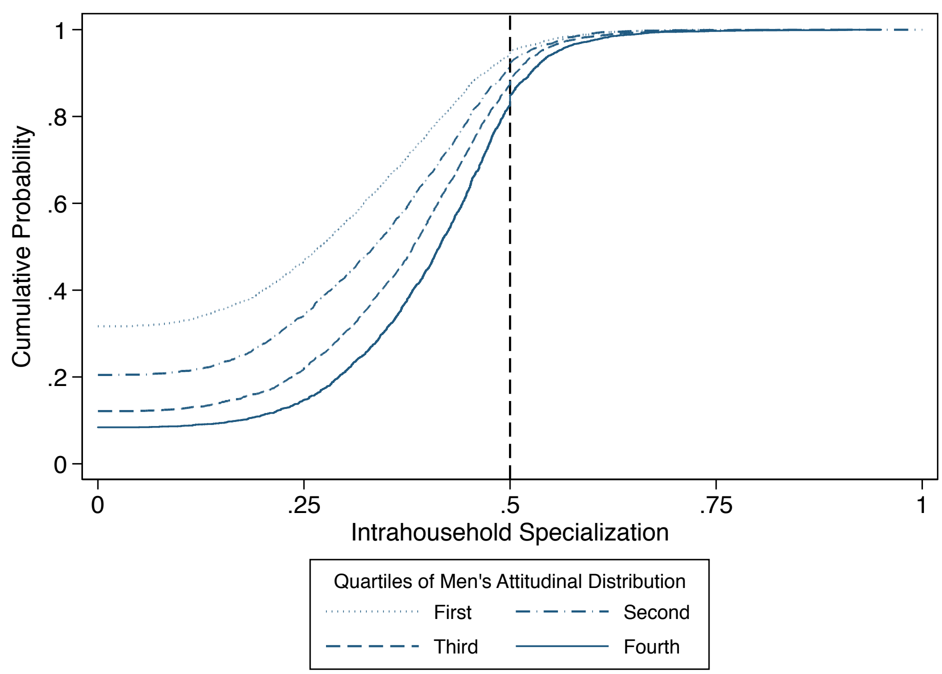

Figure 1.8: Intrahousehold specialization and men's gender-role attitudes

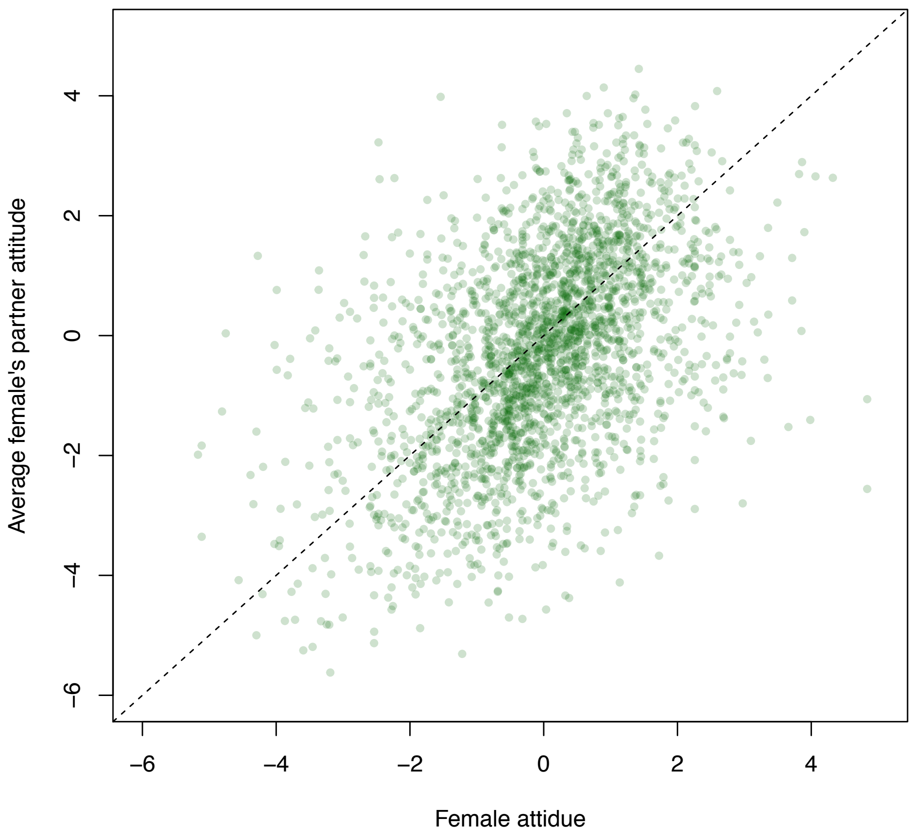

Figure 1.9: Positive assortative matching on gender-role attitudes

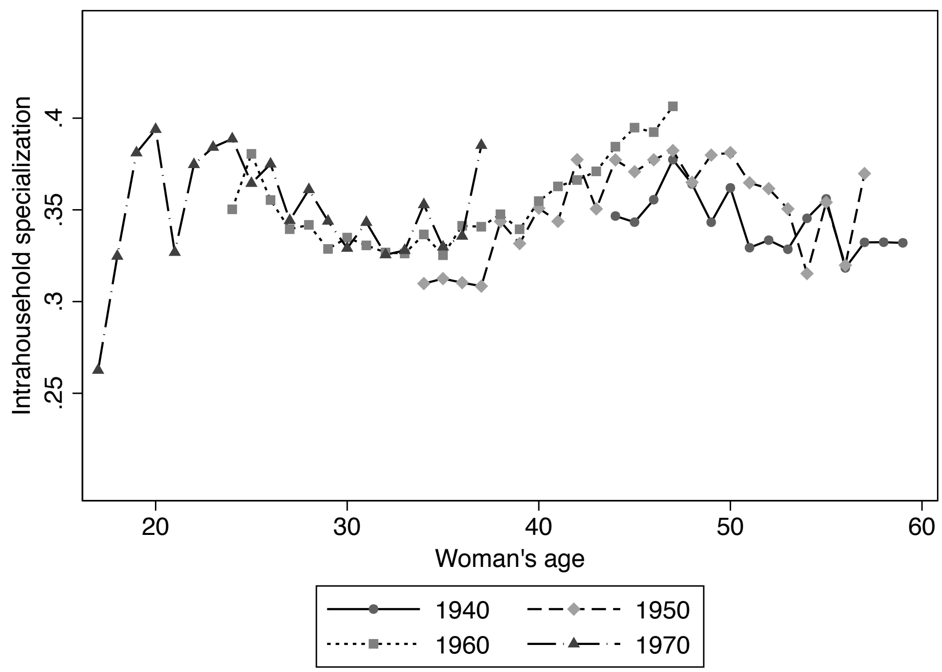

Figure 1.A1: Intrahousehold specialization by women's birth cohorts

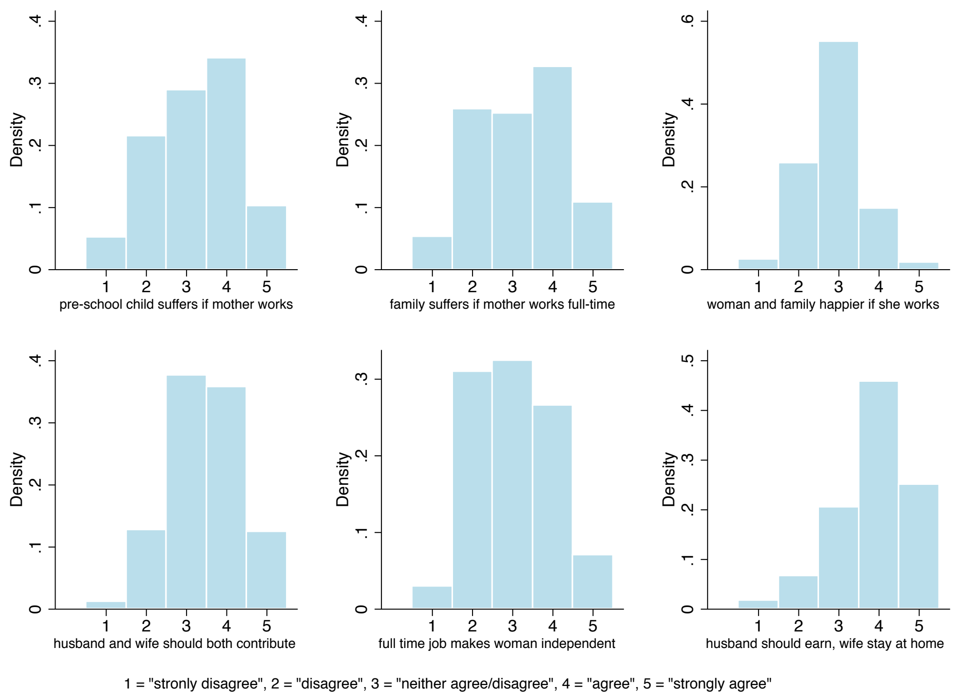

Figure 1.A2: Histograms of wives' self reported attitudes

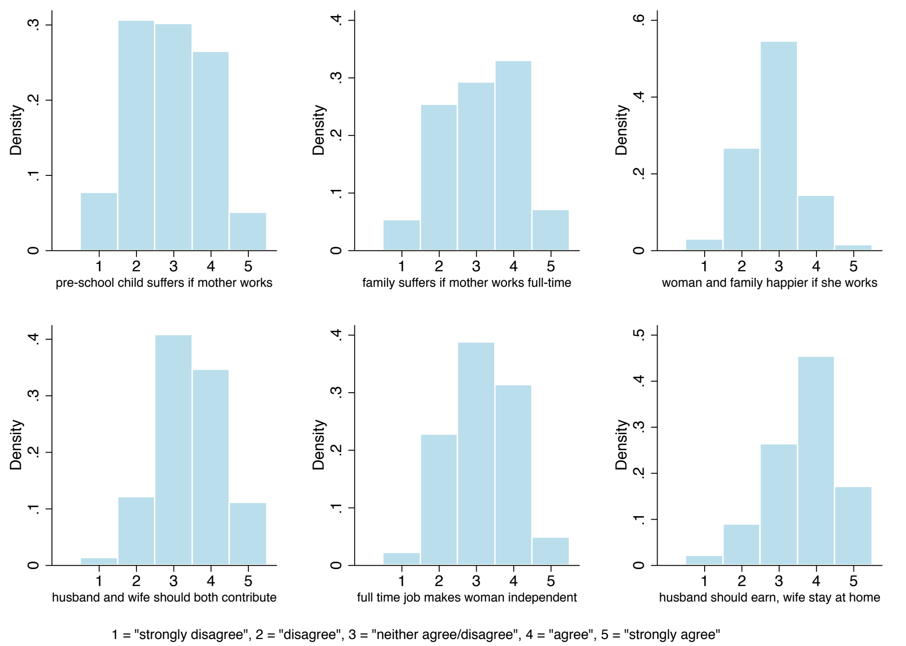

Figure 1.A3: Histograms of husbands' self reported attitudes

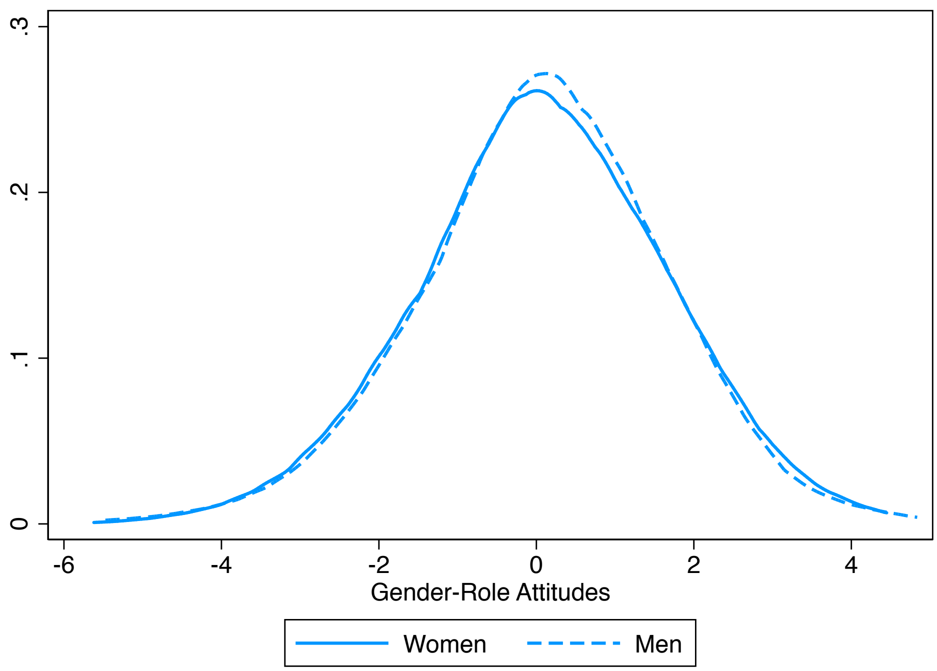

Figure 1.A4: Kernel density estimates of attitudes, by gender

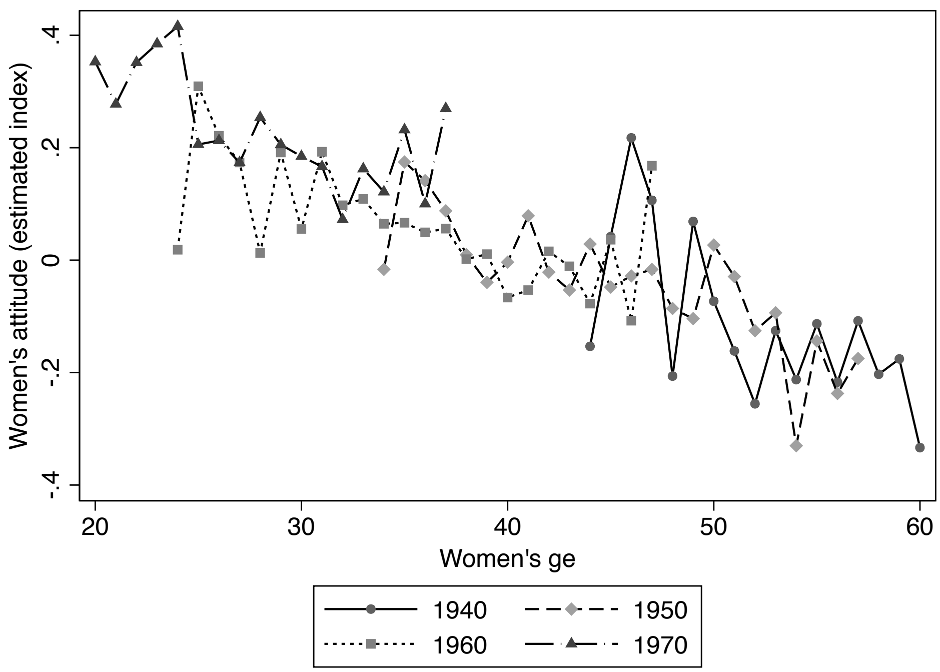

Figure 1.A5: Women's gender role attitudes by birth cohorts

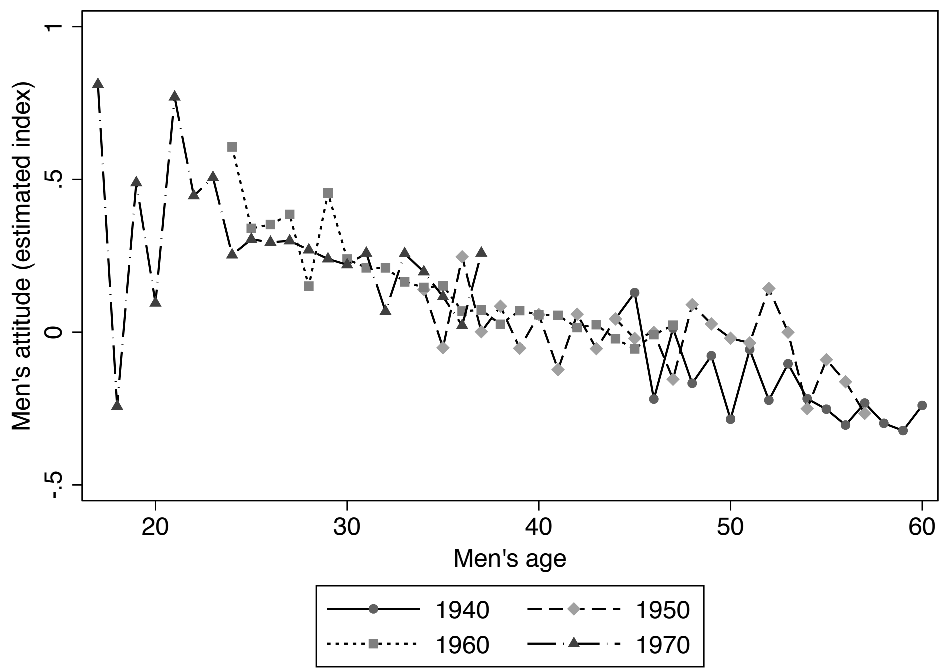

Figure 1.A6: Men's gender role attitudes by birth cohorts

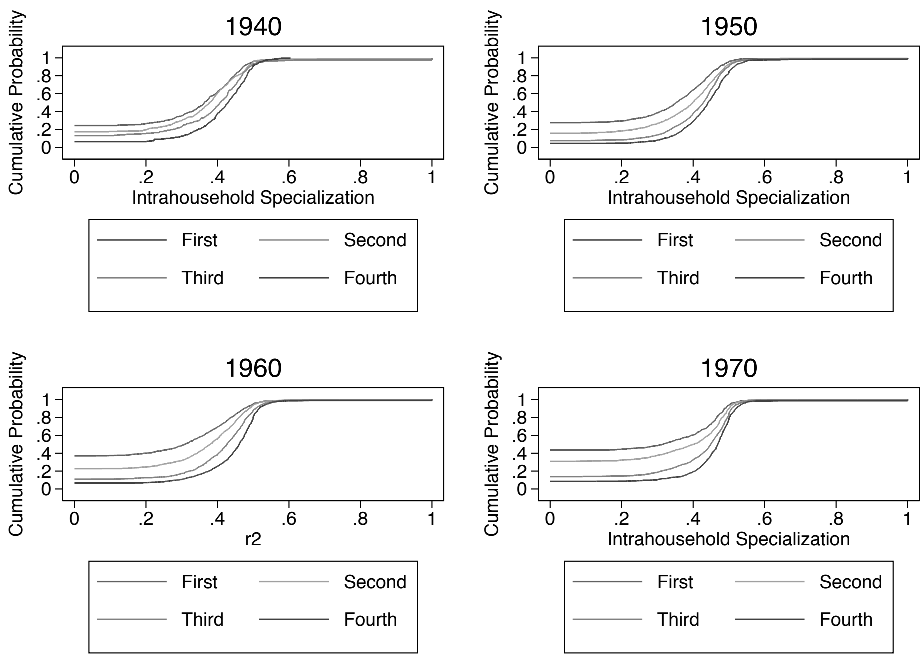

Figure 1.A7: Intrahousehold specialization and women's attitudes by birth cohorts

2

Health Education and Microfinance: Complements or Substitutes in Reducing Neonatal Mortality?

2.1 Introduction

Multiple NGO interventions, offering different development services, coexist in space and time. Yet little is known about whether spillover effects across them exist. Using longitudinal data from a randomised health education intervention in rural India, stratified by the presence of a pre-existing microfinance intervention, this paper finds a significant and substantial negative interaction between the effects of these two interventions in reducing neonatal mortality. Both are found to be more effective when offered in isolation than when offered together. This finding is surprising given popular claims that health education and microfinance complement each other. It is also an important finding for policy makers and donors alike who wish to maximize the effectiveness of their development strategies in a world of tough budgetary decisions.

In theory, health education is expected to relax knowledge constraints on the mapping from health behaviours into health outcomes (Grossman, 1972; Kenkel, 1991). Microfinance, in turn, is expected to relax financial constraints that limit an individual's ability to invest in healthy behaviours (Banerjee et al., 2010). The rationale that is typically offered for a possible complementarity between these two inputs for improving health outcomes is as follows. The increased income provided by microfinance is expected to increase the individuals' ability to act on new health knowledge. Conversely, new health knowledge is expected to improve the individuals' ability to more effectively allocate the additional income obtained from microfinance to health care expenditures.

Crucial to this complementarity argument is the assumption that the health behaviours outside the individuals' information sets, over which health education sheds light, are more costly than those that are already well known to them. If the opposite holds, however, then health education and microfinance should be expected to substitute each other. This is because in such case, financial constraints increases the value of knowledge about alternative and more affordable health behaviours; and, conversely, lack of knowledge about these alternative behaviours raises the need for the relaxation of financial constraints in order to invest in known, yet costlier, health behaviours. This alternative hypothesis has seldom been put forward both by researchers and development practitioners. This is surprising given that many health education interventions in developing countries aim at disseminating knowledge on simple and affordable health behaviours amongst their financially constrained populations.

In this paper we provide evidence in favour of this alternative hypothesis by exploiting longitudinal data from the Ekjut cluster-randomized control trial.[1] This intervention formed groups of women in remote villages located in a rural part of India, aimed at increasing awareness on the causes of neonatal mortality in the community, as well as on affordable preventive health behaviours to tackle such causes. The randomised placement of these groups across different communities was stratified by the presence of a

[1] The Ekjut trial has been recognised as Trial of of the Year by the Society for Clinical Trials.

pre-existing microfinance intervention. Specifically, in communities where microfinance was not in place the Ekjut implemented new groups of women, whereas in communities where microfinance was in place, it piggy-backed on the pre-existing self-help groups of women involved in credit and savings activities. This design feature, together with the longituindal dimension of our dataset, provides a particularly suitable econometric platform for the analysis of whether health education and microfinance act as substitutes or complements in reducing neonatal mortality.

Our central finding, under weak identifying assumption, is that the health education intervention and the microfinance intervention substituted each other in reducing neonatal mortality. When offered in isolation, both led to a strong reduction in neonatal mortality; but when offered together, their joint effect was smaller than the sum of their individual effects. That is, the presence of one of the interventions caused a reduction in the effectiveness of the other.

Additional analysis on the mechanisms through which these two interventions acted sheds further light on this negative interaction effect between the two interventions. Confirming earlier results by Tripathy et al. (2010), we find that the health education intervention when offered alone led to an increase in the adoption of affordable hygienic behaviours during home deliveries. Though more difficult to pin down the exact channels through which the microfinance intervention reduced neonatal mortality, the data suggests that it did so by increasing payments made to traditional birth attendants. These channeling effects of both interventions when offered alone are substantially weakened when they were jointly offered.

These additional findings suggest that (*i*) in communities not receiving the health education intervention, informational constraints on alternative self-help preventive behaviours during home deliveries raised the need for the relaxation of financial constraints in order to attract the most successful and experienced birth attendants, thus making the microfinance intervention more potent there; and (*ii*) in communities not receiving the microfiance intervention, financial constraints raised the need for knowledge on those behaviours in order to compensate for the households' inability to match with the most valued birth attendants, thus making the health education intervention more potent there.

Two earlier studies that have tested for spillover effects between health education and microfinance are worth mentioning. First, Smith (2002) compares an intervention that bundled credit and health education with an intervention that only offered credit to its clients using a non-experimental dataset in urban Honduras and rural Ecuador. His results are mixed. For the Honduras subsample he finds that only the integrated package significantly reduced childhood diarrhoea probability. The opposite result however was found for the Ecuador subsample, where only the credit-only intervention was found to significantly reduce diarrhoea incidence.

Second, Hamad, Fernald and Karlan (2011) evaluate the impact of a randomized health information intervention on a sample of microcredit clients in Peru. The authors do not find evidence of complementarities nor substituabilities between microfinance and health education on child health. A drawback of their study is that it lacks the appropriate counterfactual, which would consist of a parallel experiment on a sample of non-microcredit clients. In this respect, an advantage of our study is that the health intervention analysed has been randomly assigned both to a sample of communities involved in microfinance activities and to a sample of communities not involved in microfinance activities. Our study further differs from these two earlier studies in the sense that we focus on the effects that the interaction of health and the microfinance interventions can have at the community level, not at the individual level. That is, we focus on *intention-to-treat* effects. These are the parameters of interest in our case due to the community-based nature of the health intervention that we analyse.

More generally, this paper can be placed within the growing trend of research that analyses interactions across different development ingredients as well as different development interventions. Duflo, Dupas and Kremer (2011) show that providing HIV information to teenage girls in Kenya has stronger effects on sexually transmitted infections, but weaker effect on early fertility, when accompanied by education subsidies. Ashraf, Jack and Kamenica (2013) show that providing additional information about a health product in Zambia significantly increases the impact of a price subsidy on purchases of the product. Bandiera et al. (2010) show that past experience with NGO projects discourages participation in adolescent training programs in Uganda.

This paper is organised as follows. Section 2.2 describes the research design and the data set used in this paper. Section 2.3 formally discusses the identification concerns and strategies used to uncover the triple of causal effects of interest: the independent effect of the health education intervention, the independent effect of the microfinance intervention, and the interaction effect between the two interventions. Section 2.4 checks for treatment-control balance, and identifies observable differences across microfinance strata. Section 2.5 examines participation decisions in the health education intervention, separately across microfinance strata. Section 2.6 presents the main results, performs a robustness analysis, and uncovers differential mechanism through which each intervention operated. Section 2.7 concludes.

2.2 Research Design and Data

This paper uses data collected to evaluate the Ekjut cluster-randomised controlled trial in rural India. This intervention, which began in 2005, randomly implemented 244 women's groups in half of 36 clusters located in three contiguous districts of Jharkhand and Orissa: Seraikela-Karshwan, West Singhbhum, and Keonjhar. Each of these groups, of around 15-20 women, held a participatory learning and action cycle of 20 monthly meetings. During these meetings, under the guidance of trained local facilitators, group members diagnosed the causes underlying maternal and neonatal deaths in their communities, and identified simple, affordable, and effective health behaviours in order to avoid those deaths.[2]

The randomised allocation process of the 36 clusters into treatment and control was stratified according to both their pre-existing assignment to an on-going microfinance intervention as well as their district location. Table 2.1 reports the distribution of clusters by their district location and by their assignment statuses to both the health education and the microfinance interventions. Although by design treatment and control clusters are balanced with respect to these two dimensions, we can see that a cluster's district location is not orthogonal to its assignment to the microfinance intervention. This is not surprising

[2]These facilitators, though formally not health educators, received basic training to discuss health problems during pregnancy and childbirth in addition to participatory communication techniques. For a more detailed discussion on the operations of these groups see Tripathy et al. (2010) and the YouTube video `http://youtu.be/7joI2G67M_0`.

given that the placement of the mircofinance intervention was not itself randomised, and it readily illustrates the identification problem we will have to deal with. To the extent that district-specific characteristics can interact with the effectiveness of the Ekjut intervention, attributing any difference in its effectiveness across microfinance strata to the presence of microfinance *only* may be misleading. These and other identification concerns will be formally discussed in Section 2.3.

In the non-microfinance stratum, the Ekjut implemented new groups of women, whereas in the microfinance stratum it used the pre-existing self-help groups of women involved in credit and savings activities. Initially these self-help groups were closed, but with the addition of the Ekjut cycle of meeting, these groups opened up to new members. Members of the self-help groups were allowed to stay for the Ekjut meetings, but new Ekjut members were not given access to the pre-existing microfinance services. One particular concern that can thus emerge from this design feature, and that we will be addressing, is that the self-selection mechanism driving individual participation in the Ekjut meetings within the microfinance stratum may have been different than that within the non-microfinance stratum. For example, one may hypothesise that participation within the microfinance stratum could have been partially driven by the expectation of gaining access to the pre-existing self-help groups.

In order to collect data to evaluate the effectiveness of the Ekjut intervention, a key informant surveillance system was implemented during the 3 years of the intervention, as well as during a 8 month baseline period immediately prior to the start of the intervention. Key informants, essentially traditional birth attendants and active village members, identified all births, neonatal and maternal outcomes in the study location. They then met with interviewers, who in turn verified all births and interviewed all identified mothers at around 6 weeks after their delivery.[3] The interviewer collected detailed information about the women's pregnancy, delivery, and post-partum periods up to that period, as well as about their socio-economic and demographic characteristics.

[3]These informants did not have incentives for under- or over-reporting these events as they were allocated to familiar and manageable geographical areas (their allocated area corresponded to approximately 250 households), and their payments were only made after the interviewers verified their identified births and deaths. To minimise errors, all interviewed women were 'snowballed' to identify any other women in the study area who had given birth recently (see Barnett et al. (2008) for a detailed description of this surveillance methodology).

We supplemented this dataset with information on village-level amenities from the 2005 Indian census for the districts of Seraikela-Karshwan and West Singhbhum.[4] We did this because we wanted understand the extent to which villages with and without microfinance differed in ways that could interact with the effectiveness of the health education, and thus put at risk identification of the interaction causal effect between the two interventions.

2.3 Identification

2.3.1 Triple of Causal Effects, and Potential Sources of Bias

We borrow Rubin's notion of potential outcomes in order to define the causal effects of interest. Following the time structure of our main source of data, let Y_{ikt}^{00}, Y_{ikt}^{10}, Y_{ikt}^{01}, and Y_{ikt}^{11}, denote respectively neonatal survival in household i in cluster k at time period t (where $t = 0$ for the *pre*-health education intervention period, and $t = 1$ for the *post*-health education intervention period), across the following four potential scenarios: (i) this cluster k receives neither intervention, (ii) it receives the health education intervention only, (iii) it receives the microfinance intervention only, and (iv) it receives both interventions.

We focus on causal effects averaged across births within clusters sharing the same assignment status with respect to the health education and the microfinance interventions. That is, on the (the policy-relevant) average intent-to-treat (ITT) effects. Let e be a dummy indicating the health education treatment group, and m a dummy indicating the microfinance stratum. Thus, for a given group of clusters, identified by both e and m, at a given time period t, we can define the triple of causal effects

$$\boldsymbol{\tau}_{emt} = (\alpha_{emt}, \beta_{emt}, \gamma_{emt}), \tag{2.1}$$

where $\alpha_{emt} = E[Y_{ikt}^{10} - Y_{ikt}^{00}|e, m, t]$ is the average ITT independent causal effect of the health education intervention; $\beta_{emt} = E[Y_{ikt}^{01} - Y_{ikt}^{00}|e, m, t]$ is the average ITT independent causal effect of the microfinance intervention; and $\gamma_{emt} = E[(Y_{ikt}^{11} - Y_{ikt}^{10}) - (Y_{ikt}^{01} - Y_{ikt}^{00})|e, m, t]$ is the average ITT interaction causal effect between the two interventions.

[4]Unfortunately, this census did not have information for most of the villages located in the district of Keonjhar studied in this paper. As a result, when village level data is used, our analysis is restricted to the districts of Seraikela-Karshwan and West Singhbhum only. This sample restriction does not raise serious concerns, apart from potentialy reducing the precision of our estimates, because the randomisation of the health education was stratified both across district and by the presence of microfinance.

Realized outcomes, Y_{ikt}, can be linked to the the triple of causal effects $\boldsymbol{\tau}_{emt}$ given by (2.1). Letting E_{kt} and M_{kt} be dummy variables indicating whether cluster k *actually* receives the health education intervention and the microfinance intervention, respectively, at time period t, we can then write

$$Y_{ikt} = \lambda_{emt} + \alpha_{emt}E_{kt} + \beta_{emt}M_{kt} + \gamma_{emt}(E_{kt} \cdot M_{kt}) + \varepsilon_{ikt}, \tag{2.2}$$

where $\lambda_{emt} = E[Y_{00ikt}|e, m, t]$ and $E[\varepsilon_{ikt}|e, m, t] = 0$, which says that in the absence of both interventions, neonatal mortality is determined by a possibly time-varying cluster effect. This section will make precise the assumptions required to identify the parameters of interest. Specifically, it will show that in order to identify the stand-alone impact of microfinance we will have to assume that λ_{emt} equals the sum of a time-invariant difference between microfinance and non-microfinance strata, and a time effect that is common across microfinance and non-microfinance clusters. Furthermore, in order to identify the impact of microfinance on the effectiveness of the health education intervention, we will have to assume that clusters in the microfinance stratum do not differ systematically from those in the non-microfinance stratum in ways that may drive a wedge in the effectiveness of the health education intervention. To see this, let us first use Eq.(2.2) to define the parameters of interest.

Differencing Eq.(2.2) both across e and t, for $m = 0$, gives

$$\Delta_t \Delta_e E[Y_{ikt}|e, m = 0, t] = \alpha_{101} + \Delta_t \Delta_e \left(\lambda_{e0t}\right), \tag{2.3}$$

where α_{101} is the average independent ITT effect of the health education intervention among clusters treated by this intervention only; and $\Delta_t \Delta_e \left(\lambda_{e0t}\right)$ is a basing term capturing a possible underlying time-varying treatment-control difference in expected neonatal mortality, within the non-microfinance stratum.

If, instead, Eq.(2.2) is differenced both across m and t, for $e = 0$, we obtain

$$\Delta_t \Delta_m E[y_{ikt}\,|\, e = 0, m, t] = \Delta_t\left(\beta_{011}\right) + \Delta_t \Delta_m \left(\lambda_{0m1}\right), \tag{2.4}$$

where $\Delta_t(\beta_{011})$ is the independent ITT causal effect of microfinance, between time periods $t = 0$ and $t = 1$, among clusters receiving microfinance only. We call this effect, the *continuation* causal effect of microfinance since it nets outs its effect up to $t = 0$. The term $\Delta_t \Delta_m (\lambda_{0m1})$ captures a potentially biasing underlying difference in average neonatal mortality between microfinance strata, within the health education control group.

Finally, differencing Eq.(2.2) across e, m, and t, gives

$$\Delta_t \Delta_m \Delta_e E[y_{ikt}|e, m, t] = \gamma_{111} + \Delta_m (\alpha_{1m1}) + \Delta_t \Delta_e (\beta_{e11}) + \Delta_t \Delta_m \Delta_e (\lambda_{em1}), \tag{2.5}$$

where γ_{111} is the average ITT interaction between the health education and microfinance among clusters that actually host both interventions (the interaction effect on the treated). There are three possible sources of bias in the triple difference given by Eq.(2.5). First, $\Delta_m (\alpha_{1m1})$ captures the possibility that the average independent effectiveness of the health education intervention at $t = 1$ may interact with underlying characteristics that are specific to the microfinance stratum. Second, $\Delta_t \Delta_e (\beta_{e11})$ captures the possibility that the independent 'continuation' effectiveness of the microfinance intervention, may interact with time-varying underlying characteristics that are specific to the health education treatment group. Third, $\Delta_t \Delta_m \Delta_e (\lambda_{em1})$ captures a possible time-varying interaction between underlying characteristics specific to the health education treatment group, on one hand, and to the microfinance stratum, on the other hand, that can in turn affect expected neonatal mortality.

2.3.2 How Much Can We Identify?

Stratified randomisation of the health education intervention implies that on average the only difference between treatment and control groups, within both microfinance strata, is their assignment statuses to the health education intervention. This has two implications. First, in the absence of the health education intervention, expected neonatal mortality should be the same in treatment and control groups, i.e. $\Delta_e (\lambda_{emt}) = 0$. As a result, the biasing terms $\Delta_t \Delta_e (\lambda_{e0t})$ and $\Delta_t \Delta_m \Delta_e (\lambda_{em1})$ in Eqs. (2.3) and (2.5), respectively, vanish. Second, in the absence of the health education intervention, the effectiveness of the microfinance intervention should be the same in treatment and control groups, i.e. $\Delta_e (\beta_{e1t})$. As a result, the biasing term $\Delta_t \Delta_e (\beta_{e11})$ in Eq.(2.5) vanishes.

It thus follows that the stratified randomized nature of the health education intervention enables Eq.(2.3) to fully identify the independent causal effect of the health education intervention, α_{101}, and it substantially eases the ability of Eq.(2.5) to identify the interaction causal effect between the health education and the microfinance interventions, γ_{111}. For Eq.(2.5) to fully identify this latter parameter, however, it has to be the case that $\Delta_m(\alpha_{1m1}) = 0$.

H 1. *It is assumed that $\Delta_m(\alpha_{1m1}) = 0$, i.e. that in the absence of microfinance, clusters in the microfinance stratum do not differ from those in the non-microfinance stratum in ways that may drive a wedge in the effectiveness of the health education intervention.*

This identification assumption may be unrealistic given the non-random placement of the microfinance intervention. Indeed, and anticipating on our own analysis, we observe that villages located in the microfinance stratum are better served by maternal and child health care centres than villages located in the non-microfinance stratum. If these health care facilities either complement, $\Delta_m(\alpha_{1m1}) > 0$, or substitute, $\Delta_m(\alpha_{1m1}) < 0$, the health education intervention, then Eq.(2.5) will either over- or under-estimate, respectively, the true interaction causal effect between heath education and microfinance, γ_{111}.

Hence, in the empirical section of this paper we will relax H1, and will proceed as follows. First, we will identify observable differences in individual, household, and village level characteristics, between the microfinance stratum and the non-microfinance stratum. Second, when estimating the triple of causal effects $\boldsymbol{\tau}_{emk}$, we will control for possible heterogeneous effects of the health education intervention with respect to these characteristics. The resulting estimate for the interaction causal effect, $\Delta_t\Delta_m\Delta_e E[y_{ikt}|e, m, t]$, will thus identify the differential causal effect of the health education intervention between the microfinance and the non-microfinance strata, over and above its heterogeneous effectiveness with respect to these characteristics. To the extent that explicitly controlling for these observable sources of heterogeneity, implicitly controls of unobserved sources of heterogeneity, this procedure will identify the true interaction causal effect between the two interventions, γ_{111}.

Finally, the longitudinal dimension of our dataset, in turn, implies that any underlying time-invariant difference in expected realized neonatal mortality between the microfinance and the non-microfinance stratum can be eliminated. That is, under the assumption that there are no time-varying differences in expected neonatal mortality between these two strata, the biasing term $\Delta_t \Delta_m (\lambda_{0m1})$ in Eq.(2.4) vanishes, enabling it to identify the causal effect of the microfinance intervention between $t = 0$ and $t = 1$, i.e. its continuation casual effect $\Delta_t (\beta_{011})$.

H 2. *It is assumed that* $\Delta_t \Delta_m (\lambda_{0m1}) = 0$*, i.e. that there are no time-varying differences in expected neonatal mortality between clusters with and without microfinance.*

2.4 Balance Checks

In order to gauge the potential sources of bias discussed in the previous section, we now identify observable differences across the microfinance strata, and assess whether the stratified randomisation of the health education intervention successfully balanced treatment and control groups. To do so, we estimate regressions of the following form using data from the baseline period

$$X_{ihvk} = \alpha_0 + \alpha_1 E_k + \alpha_2 M_k + \alpha_3 [E_k \times M_k] + \boldsymbol{\alpha_4'} \boldsymbol{d}_k + \varepsilon_{ihvk}, \tag{2.6}$$

where X_{ihvk} is a particular observable characteristic, either at the respondent mother level i, at the level of her household h, or at the level of her village v; E_k is a dummy indicating whether the cluster k from where this mother comes from is located in the treatment group; M_k is a dummy indicating whether the cluster is located in the microfinance stratum; and $\boldsymbol{d}_k$ is a vector of dummies indicating the cluster's district location. Inference is made robust to heteskodestaticity by clustering the standard errors at the cluster level - the level at which the randomisation of the health education intervention took place.

For each characteristic on the left-hand side of Eq.(2.6), α_1, α_2, and α_3, measure, respectively, its expected treatment-control difference in the non-microfinance stratum, its expected difference across the microfinance strata within the control group, and the difference in its expected treatment-control difference across microfinance strata.[5] Under the null hypothesis that the stratified randomisation of the health education intervention

[5]The parameter α_3 equivalently measures for each characteristic the expected treatment-control difference in its difference across microfinance strata.

successfully balanced treatment and control groups within both microfinace strata, our estimates for both α_1 and α_3 should not significantly differ from zero.[6] However, since the microfinance intervention was not itself randomly placed, there is a priori no reason to expect γ_2 to be equal to zero. In order to net out our estimates of γ_2 from district specific components that may simultaneously predict the availability of microfinance and the characteristics measured on the left-hand side of Eq.(2.6), we throughout control for district fixed effects.

The results are reported in Tables 2.2, 2.3, and 2.4. Table 2.2 focus on the main outcome of interest, the likelihood of the baby dying within the first six weeks of life, and on different care seeking behaviours during pregnancy and hygienic practices at home deliveries. Table 2.3 focus on individual and household background level characteristics. Table 2.4 focus on village level characteristics. In each of these three tables, Column 1 shows the means and standard deviations of these variables for the subsample of control clusters in the non-microfinance stratum. Column 2 reports estimates for α_1, Column 3 for α_2, and Column 4 for α_3. Column 5, in turn, reports the p-value associated with the F-test for the null hypothesis that the triple $(\alpha_1, \alpha_2, \alpha_3)$ is jointly equal to zero.

Overall it appears that the stratified randomisation of the health education intervention successfully balanced treatment and control groups, within both microfinance strata. Table 2.2 shows that neonatal mortality at baseline is statistically indistinguishable between treatment and control clusters. Within the non-microfinance stratum, the likelihood that a newborn baby dies within the first six weeks of life is 4.5% in control clusters. A figure that raises by a statistically insignificant 1.7 percentage points (s.e.=1.2) among treated clusters within that stratum. The hypothesis that this treatment-control balance extends to the microfinance stratum cannot be rejected. There, the treatment-control difference in neonatal mortality is a statistically insignificant 1.4 percentage points smaller than that in the non-microfinance stratum (s.e. = 1.4). The rest of Table 2.2 indicates that this pattern of balance between treatment and control clusters, within both microfinance strata, cannot be rejected for different care-seeking behaviours during pregnancy and hygienic practices during home deliveries

[6] Observe that an interesting corollary of the null that $\alpha_3 = 0$ is that any difference across microfinace strata detected in the control group should also be detected in the treatment group (and vice-versa).

Tables 2.3, and 2.4 provide further evidence suggesting that the randomisation of the health education intervention has been successful. Column 2 of these tables, shows that out of 22 treatment-control differences on individual, household, and village level characteristics, within the non-microfinance stratum, only one - just below 5% of the total - is statistically significant at the 5% level. Column 4, in turn, shows that this pattern of balance between treatment and control clusters within the non-microfinance is significantly indistinguishable from that in the microfinance stratum.

A similar set of comparisons, but now across microfinace strata, tells a different story. Although no significant differences across this strata were detected on neonatal mortality and other health related measures, we do find that mothers coming from the microfinance stratum appear to be better off than mothers coming from the non-microfinance stratum. Specifically, Column 3 of Tables 2.3 and 2.4 shows that within the control group of clusters the presence of microfinance is associated with a 0.26 decrease in the number of times the respondent mother was pregnant (s.e.=0.15), a 19.4 percentage points increase in the likelihood that she has a say within her household (s.e.=9.5), a 0.23 increase in the number of assets held by her household (s.e.=0.14), and a 38 percentage points increase in the likelihood that her village has a maternal or a child health centre (s.e.=11.7). As a consequence of the successful randomisation of the health education intervention, Column 4 indicates that we cannot reject the hypothesis that these differences across microfinance strata among control clusters extend to the treated clusters.

To the extent that the observed differences across microfinance strata can interact with the effectiveness of the health education intervention, it will therefore be important to control for these observable sources of variability in its effectiveness, when estimating its differential causal effect across microfinance strata in Section 2.6. Doing this will help us isolate the part of that overall differential effectiveness that is only due to the presence of the microfinance intervention.

2.5 Participation

This subsection examines whether participation rates in the Ekjut intervention, as well as observable individual determinants of participation, vary across microfinance strata.

Whether the presence of microfinance increases or reduces the overall participation rate in the Ekjut intervention, depends on whether microfinance and health education complement or substitute each other. If the relaxation of financial constraints induced by the microfinance intervention raised the need for health education, we should then expect it to have increased. Under the alternative substitutability hypothesis, we should expect it to have reduced.

The presence of microfinance may have also affected the composition of the participants in the Ekjut intervention. A particular concern here is the possibility that participation within the microfinance stratum may have been partially driven by the expectation of gaining future access to the pre-existing self-help groups involved in credit and savings activities. If this was the case, it is then possible that the health education groups in these communities failed to attract the mothers who could have benefited the most from it, leading to a reduction in its effectiveness.

To shed light on these questions, we estimate versions of the following specification with a linear probability model, using data from the intervention period in the treated clusters,

$$P_{ihk} = \beta_0 + \beta_1 M_k + \boldsymbol{\beta}_2' \boldsymbol{X}_{ih} + +\boldsymbol{\beta}_3' [M_k \times \boldsymbol{X}_{ih}] + \boldsymbol{\beta}_4' \boldsymbol{d}_k + \varepsilon_{ihk} \tag{2.7}$$

where P_{ihk} is a dummy equal to one if the respondent mother i participates in the health education intervention, $\boldsymbol{X}_{ih}$ is a vector of characteristics of this respondent and of her household h, M_k is a dummy variable that is equal to 1 if cluster k is located in the microfinance stratum, and $\boldsymbol{d}_k$ is a vector of district dummies. Standard errors are clustered at the cluster level.

In order to examine whether aggregate participation rate in the Ekjut intervention varies with the presence of the microfinance intervention, we start by assuming that the pattern of individual correlates of participation is homogenous across the microfinance strata. That is, we restrict the vector of parameters $\boldsymbol{\beta}_3$ in Eq.(2.7) to be equal to zero. Our estimate of β_1 will thus measure the difference across microfinace strata in average participation rates in the Ekjut intervention, controlling for both individual and household level characteristics, as well as for district fixed effects. In addition, our estimates for the vector $\boldsymbol{\beta}_2$ will measure the average individual level determinants of participation controlling for both the presence of microfinance in the community and district fixed effects.

The results, reported in Column 1 of Table 2.5, show that the presence of microfinance reduces average participation rates in the health education intervention. Specifically, a mother coming from a microfinance cluster is on average 10.9 percentage points (s.e.=5.7) less likely to participate compared to a mother residing in a cluster without microfinace. Though statistically significant only at the 10% level, the magnitude of this effect is substantial once compared against the average participation rate of 37%. This finding is consistent with the idea that the presence of the microfinace intervention reduced the demand for the health education intervention, favouring the hypothesis that these two interventions may have substituted each other.

Column 1 also shows that on average the only significant individual level predictors of participation are the respondent's number of past pregnancies and her literacy level. Specifically, one additional past pregnancy episode is associated with an increase in the likelihood of participating in a health education group by 5.2 percentages points (s.e.=1.0), and being able to read and write also increases this probability by 3 percentage points (s.e.=1.4). These two pieces of evidence seem to suggest that greater awareness of possible health complications during pregnancy increased the demand for health education.

In order to understand whether the presence of microfinace also affected the composition of the participants in the Ekjut intervention, we now drop the previous restriction that $\boldsymbol{\beta}_3 = 0$ in Eq.(2.7). Now the parameters of interest are the vector $\boldsymbol{\beta}_2$, which measures the correlates of participation within the non-microfinance stratum, the vector $\boldsymbol{\beta}_2 + \boldsymbol{\beta}_3$, which measures the correlates of participation within the microfinance stratum, and the vector $\boldsymbol{\beta}_3$, which measures the differential pattern of correlates of participation between clusters with and without microfinance. Under the null hypothesis that the presence of microfinance did not affect who decide attends the health education groups, our estimates of $\boldsymbol{\beta}_3$ should not significantly differ from zero.

The results are reported in Columns 2-4 of Table 2.5. We can see that the reviously identified significant predictors of participation *on average*, namely the number of past pregnancies and whether the respondent is literate, appear to be homogenous across microfinance strata. This finding supports the idea that microfinance does not appear to have fundamentally changed the self-selection mechanism driving participation in the Ekjut intervention. However we can also see that household asset ownership is a significant negative predictor of participation, but only within the non-microfinance stratum. Specifically, there a one standard deviation increase in the household asset index is associated with a 2.6 percentage points (s.e.=0.7) decrease in participation. In the microfinance stratum, however, we cannot reject the hypothesis that household asset ownership is not significantly correlated with participation. This differential pattern across microfianance is statistically significant at 5% level, as indicated by its p-value reported in Column 4.

Based on the literature suggesting that the poorest women are likely to be excluded from microfinance due to their increased risk of default (see e.g. Armendariz and Morduch, 2010), this finding goes against the idea that the presence of microfinance made participation in the Ekjut intervention to be partially driven by the expectation of future access to the pre-existing self-help groups involved in credit and savings activities. How then can we interpret this finding and how it may affect the effectiveness of the health education intervention?

One plausible interpretation is that wealthier women that were already participating in those self-help groups may have decided to stay for the health education sessions. In such case, whether this will enhance or reduce the effectiveness of the Ekjut intervention will again depend on whether microfinance and health education are complements or substitutes. If these women, who are relatively less financially constrained, are in better position to act on the information gleaned through the health education sessions, then we should expect this to increase the effectiveness of the Ekjut intervention. If, alternatively, this relaxation of financial constraints reduced the need for health education, we should then expect this to further contribute to a reduction in the effectiveness of the Ekjut intervention.

2.6 Results

2.6.1 Main Estimates

Equation (2.2) in Section 2.3, which related realized outcomes and the triple of causal effects of interest, is estimated using the following linear probability model

$$Y_{ikt} = \delta_0 + \delta_1 E_k + \gamma_2 M_k + \delta_3 (E_k \cdot M_k) + \delta_4 T_t + \delta_5 (T_t \cdot E_k) + \delta_6 (T_t \cdot M_k) + \delta_7 (T_t \cdot E_k \cdot M_k) + \varepsilon_{ikt}, \quad (2.8)$$

where i indexes births, k clusters, and t time periods; Y_{ikt} is a dummy that is equal to one if the newborn baby i died in the first six weeks of life, and zero otherwise; E_k is a dummy for clusters assigned to health education; M_k is a dummy for clusters assigned to microfinance; and T_t is a dummy for the *post* health education intervention period.

The parameters of interest are δ_5, δ_6, and δ_7, which under the assumption (soon to be relaxed) that $E[\varepsilon_{ikt}|E_k, M_k, t_t] = 0$, they are respectively equal to the average ITT independent causal effect of health education among clusters receiving the health education intervention only, α_{101}, the average ITT independent continuation causal effect of microfinance among clusters receiving the microfinance intervention only, $\Delta_t(\beta_{01t})$, and the average ITT interaction effect between health education and microfiance among clusters receiving both interventions, γ_{111}. A positive estimate of δ_7, will indicate that health education and microfinance complement each other, whereas a negative estimate of δ_7, will indicate that health education and microfinance substitute each other.

Standard errors are adjusted to account for the clustering nature of the data. To mitigate concerns about the possibly small number of clusters that may emerge in response to the cutting nature of our estimation, across cluster-level assignment status to both interventions, standard are estimated using Bell and McCaffrey's (2002) Biased Reduced Linearization (BRL) estimator.

The result are reported in Table 2.6. Panel A reports estimates for $(\delta_5, \delta_6, \delta_7)$. It also reports estimates for $\delta_5 + \delta_7$, the causal effect of the health education education intervention when built on top of the microfinance intervention; $\delta_6 + \delta_7$, the causal effect of the microfinance intervention when used as a platform for the implementation of the health education intervention; and $\delta_5 + \delta_6 + \delta_7$, the joint effect of both interventions.

Column 1 indicates that (i) both interventions are positive inputs for reducing neonatal mortality, as they both independently reduce the likelihood that a newborn baby dies within the first six weeks of life, and that (ii) they substitute each other, as their joint effect when offered together is smaller then the sum of their independent effects. Specifically, it shows that neonatal mortality is reduced by 3.96 percentage points (s.e.=1.04) when health education is offered alone, by 1.69 percentage points (s.e.=0.70) when microfinance is offered alone, and by 2.79 percentage points (s.e.=0.82) when both interventions are offered simultaneously. That is, their interaction approximately halves the sum of their individual impacts. An effect that is statistically significant at the 1% level.

Columns 2 and 3 present estimates from a Logit specification adapted to Equation (2.8). This serves two purposes. First, as a basic robustness check to the previous baseline OLS estimates, and second, to better gauge their relative magnitudes. The estimated marginal effects in Column 3 are only slightly smaller than the OLS estimates. The estimated odds ratios in Column 4 show that neonatal mortality reduces by 58% in response to the health education intervention, and by 31% in response to the microfinance intervention. Offering both interventions together reduces their individual effectiveness by 85%.

To isolate the impact of the presence of microfinance per se on the effectiveness of the health education intervention, we need to purge the previous estimates of dimensions that both drive the presence of microfinance and responses to the health education intervention. We thus now allow the impact of the health education intervention to be heterogeneous across districts and along each of the characteristics that in Section 2.6 were found to differ across microfinance strata at baseline.[7]

We thus estimate the following specification

$$\begin{aligned} Y_{ihvkt} &= \delta_0 + \delta_1 E_k + \delta_2 M_k + \delta_3 (E_k \cdot M_k) + \delta_4 T_t + \delta_5 (T_t \cdot E_k) + \delta_6 (T_t \cdot M_k) + \delta_7 (T_t \cdot E_k \cdot M_k) \\ &+ \boldsymbol{\phi}_1' \boldsymbol{X}_{ihvk} + \boldsymbol{\phi}_2' (\boldsymbol{X}_{ihvk} \cdot E_k) + \boldsymbol{\phi}_3' (\boldsymbol{X}_{ihvk} \cdot T_t) + \boldsymbol{\phi}_4' (\boldsymbol{X}_{ihvk} \cdot T_t \cdot E_k) + \varepsilon_{ihvkt}, \qquad (2.9) \end{aligned}$$

where the vector $\boldsymbol{X}_{ihvk}$ includes district dummies, the respondent's number of past pregnancies, a dummy of whether she has a say in household decision-making, an index of

[7]Though not reported, these variables were not found to have been affected by neither intervention, either when offered in isolation or when offered jointly.

ownership of household durable assets, and a dummy for whether her village has a maternal or a child health care centre. Continuous variables in $\boldsymbol{X}_{ihv}$ are appropriately rescaled to have a mean of zero and a standard deviation of one. Now, the parameter δ_7 identifies the differential effectiveness of the Ekjut intervention across microfinance strata, over and above its heterogeneity along the dimensions captured by the vector $\boldsymbol{X}_{ihv}$.

The results are reported in Table 2.7. To benchmark the analysis, Column 1 shows the baseline estimates derived in the previous subsection. Column 2 controls for heterogeneity of the Ekjut intervention with respect to the respondent's number of past pregnancies, Column 3 with respect to whether she has a say in household decision-making, Column 4 with respect to household asset ownership, Column 5 with respect to district location, and Column 6 with respect to all these dimensions simultaneously. Column 7, in turn, controls for heterogeneity with respect to the presence of a maternal or a child health care centre in the village, using the restricted sample of clusters located in the districts of Seraikela-Karshwan and West Singhbhum for which we have census data on village-level amenities.

Overall, the results convincingly show that our previous estimates are robust to the inclusion of these interactions. The estimate on the interaction effect between the two interventions not only does not decrease when allowing for these different sources of heterogeneity, but it actually appears to increase. This indicates that our initial estimate was not picking up a spurious positive interaction of the Ekjut intervention with respect to the observed improved characteristics found to hold in the microfinance stratum.

2.6.2 Mechanisms

Having ruled out that the substitutability is driven by the fact that mothers in villages located in the microfinance stratum are better off than mothers in villages located in the non-microfinance stratum, we now re-estimate Eq.(2.8) replacing the dependent variable with different care-seeking behaviours during pregnancy and home-care practices. The goal is to understand the mechanisms through which each intervention acted on, both when offered in isolation and when offered together. Again, the parameters of interest are $(\delta_5, \delta_6, \delta_7)$, whose interpretation is analogous to that in sub-Section 2.6.1, though appropriately adapted of the dependent variable in question.

The results are reported in Table 2.8. For each of these behaviours, Columns 1 to 3 report estimates of the independent causal effect of the health education intervention, the independent causal effect of the microfinance intervention, and the interaction causal effect between the two interventions, respectively. Columns 4 to 6, in turn, report the causal effects of the health education intervention when built on top of the microfinance intervention, the causal effects of the microfinance intervention when used as platform for the health education intervention, and the joint effects of both interventions, respectively.

Overall the results suggest that the two interventions essentially acted through substitutable channels in reducing neonatal mortality. Consistent with earlier results by Tripathy et al. (2010), Column 1 shows that the health education intervention when offered alone led to an increase in the birth attendant's use of a clean delivery kit and the adoption of clean delivery practices, such as hand washing, use of gloves, use of boiled thread, and use of plastic sheet for a clean delivery surface. Column 4 shows that some of these channels appear to be substantially muted in the presence of the microfinance intervention. For example, in communities without microfinance the health education intervention increased the utilisation of safe-delivery kits by 22 percentage points (s.e.=6.9), however in communities with microfinance this effects more than halves and is not significantly different than zero.

Column 2 shows that the microfinance intervention, in turn, when offered in isolation led to a 10.4 percentage points increase in payments made to birth attendants (s.e.=4.8). In contrast, Column 5 shows that when offered together with the health education intervention, we cannot reject the hypothesis that these payments were unaffected. Finally, Column 4 shows that when offered on top of the microfinance intervention, the health education intervention fully undoes its impact on these payments.

An exception to this pattern of substitutability is the uptake of antenatal checkup. When offered in isolation the microfinance intervention reduced the likelihood of the mother having a checkup during pregnancy by 10.2 percentage points (s.e.=5.9). Yet, when the health education intervention was offered on top of the microfinance intervention, the likelihood of this behaviour increased by 18.5 percentage points (s.e.=5.8). That is, health education not only curbed this apparent pernicious effect of microfinance, but it even appears to have reversed it. Overall, though, the pattern of substitutability between microfinance and health education in improving health behaviours reported in Table 8 clearly outweighs this particular source of complementarity.

These results suggest that the presence of microfinance in a community, by allowing households to pay for healthcare, reduces the need for information in alternative affordable health behaviors. Essentially, the presence of microfinance is thus thought to have a similar role to that of variation in household assets. We can therefore check whether in non-microfinance communities, the impact of the health education intervention is smaller among wealthier households. Specifically, we estimate the following specification using data from clusters not hosting the microfinance intervention

$$\begin{aligned} Y_{ihkt} &= \delta_0 + \delta_1 E_k + \delta_2 T_t + \delta_3 (T_t \cdot E_k) + \phi_1 A_{ihk} + \phi_2 (A_{ihk} \cdot E_k) + \phi_3 (A_{ihk} \cdot T_t) \\ &+ \phi_4 (A_{ihk} \cdot T_t \cdot E_k) + \varepsilon_{ihkt}, \end{aligned} \tag{2.10}$$

where A_{ihk} is a demeaned index of ownership of household durable assets. The parameter δ_3 measures the average stand-alone ITT effect of the health education intervention. The parameter δ_4 measures the heterogeneity of this effect with respect to household asset ownership. Table 2.9 reports the results. Consistent with the hypothesis, the impact of the health education intervention is greater among poorer households.

2.7 Conclusion

We have presented evidence from a stratified randomised controlled trial in rural India on whether a health education intervention and a microfinance intervention complemented or substituted each other in reducing neonatal mortality. Our central finding is that both interventions were significantly and substantially more effective when offered in isolation then when offered together.

We interpret this result as suggestive evidence that increased financial constraints in communities without access to microfinance, raised the need for knowledge on simple and affordable hygienic practices at home deliveries - thus making the health education intervention more effective there. Conversely, increased knowledge constraints in communities without access to the health education intervention, raised the need for the relaxation of financial constraints in order to contract the services provided by the more expensive, and hence potentially more skilled, birth attendants - thus making the microfinance intervention more potent there.

A plausible alternative explanation that cannot be ruled out is the presence of diminishing returns to scale in the underlying health production function. It is generally understood that health production is subject to the law of diminishing returns, even though in his seminal theoretical work Grossman (1972) assumed constant returns to scale (see e.g. Galama et al., 2012). Concavity of the health production function would precisely predict that both interventions are more effective when offered in isolation then when offered together.

Our key finding that health education and microfinance are more effective when offered in isolation, then when offered jointly, is in stark contrast with conventional wisdom suggesting that these two types of development interventions complement each other, and should therefore be offered together to poor communities. They are good news for policy makers and donors alike whose budgets for development interventions are typically constrained. Although these findings hold under weak identifying assumptions, we cannot rule the possibility of a bias in the estimation of this substitutability effect driven by unobservables that are both correlated to the presence of the microfinance intervention and cause heterogenous responses to the health education intervention. Given the important policy implications of our results, it is critical to further scrutinise them in future work.

2.A Appendix: Tables and Figures

Table 2.1: Randomization of clusters into treatment and control groups, stratified by the presence of microfinance

	No Microfinance			Microfinance			
Districts	Control	Treatment	Total	Control	Treatment	Total	Total
West Singbhum	3	2	5	3	4	7	12
Sarakila-Kharswan	4	5	9	2	1	3	12
Kehonjar	2	2	4	4	4	8	12
Total	9	9	18	9	9	18	36

Table 2.2: Neonatal mortality, care-seeking behaviour, and home-care practices

	Mean	OLS coefficients			P-value on F-test for (2), (3), and (4)
	No intervention	Health education	Microfinance	Interaction	
	(1)	(2)	(3)	(4)	(5)
Neonatal death	.045	.017	.011	-.014	[.291]
[=1 if infant died in first 6 weeks]	(.207)	{.012}	(.007)	(.014)	
Had health checkup	.538	-.027	-.001	-.062	[.802]
[yes =1]	(.499)	{.093}	(.095)	(.146)	
Took iron tablets	.654	-.009	.026	.012	[.945]
[yes =1]	(.476)	{.085}	(.087)	(.116)	
Had a tetanus injection	.643	-.048	.040	-.015	[.732]
[yes =1]	(.479)	{.096}	(.080)	(.116)	
Institutional delivery	.128	-.029	.034	-.020	[.253]
[yes =1]	(.335)	{.025}	(.037)	(.046)	
Made payments to BA	.651	-.001	-.022	-.070	[.511]
[yes =1]	(.477)	{.080}	(.057)	(.098)	
BA used safe-delivery kit†	.098	-.005	.021	-.024	[.932]
[yes =1]	(.297)	{.041}	(.044)	(.066)	
BA washed hands with soap†	.276	.035	-.012	.009	[.933]
[yes =1]	(.447)	{.115}	(.088)	(.140)	
BA used plastic sheet†	.061	-.023	.049	-.007	[.884]
[yes =1]	(.240)	{.058}	(.072)	(.093)	
BA used gloves†	.040	-.017	-.020	.013	[.414]
[yes =1]	(.198)	{.013}	(.015)	(.017)	
BA tied cord with boiled thread†	.144	.001	.033	-.010	[.981]
[yes =1]	(.351)	{.042}	(.050)	(.069)	

Notes: *** denotes significance at the 1% level, ** at the 5% level, * at the 10% level. Sample used is all observations at the baseline year (number of observations = 4,437, number of clusters = 36). Means, OLS estimates, standard deviations reported in parentheses, standard errors in curly brackets, p-values reported in squared brackets † Excludes institutional deliveries. For each row, OLS coefficients and corresponding standard errors in Columns 2-4, as well as p-values reported in Column 5, are obtained from an OLS regression of the corresponding characteristic on a dummy for whether the observation comes from the treatment or the control group, on a dummy for whether it comes from the microfinance or the non-microfinance stratum, and on the interaction between these two dummies, plus district dummies. Standard errors are clustered at the cluster level.

Table 2.3: Individual and household level characteristics

	Mean	OLS coefficients			P-value on F-test for (2), (3), and (4)
	No intervention	Health education	Microfinance	Interaction	
	(1)	(2)	(3)	(4)	(5)
Age at interview [years]	25.57 (5.56)	.361 {.705}	-.764 {.655}	.051 {.848}	[.199]
Age at first pregnancy [years]	18.93 (2.36)	-.084 {.344}	-.535 {.359}	.306 {.452}	[.327]
Age at marriage [years]	19.73 (2.54)	.204 {.423}	-.312 {.421}	.006 {.514}	[.474]
Past pregnancies [number]	2.63 (1.21)	.165 {.135}	-.256** {.148}	.193 {.230}	[.071]
Literate [yes =1]	.270 (.444)	-.074 {.049}	.071 {.055}	-.019 {.085}	[.190]
Tribal member [yes =1]	.746 (.435)	.028 {.045}	-.116 {.057}	.078 {.075}	[.099]
Hindu [yes =1]	.492 (.500)	-.042 {.068}	.052 {.056}	-.021 {.079}	[.403]
Say over health care decisions [yes =1]	.254 (.436)	-.008 {.085}	.194** {.095}	-.061 {.159}	[.153]
Household assets [score 1-10]	1.394 (1.196)	-.423*** {.122}	.232* {.139}	.186 {.213}	[.002]
Household owns at least 2 bighas of land [yes =1]	.435 (.496)	-.072 {.087}	-.047 {.073}	.089 {.114}	[.845]
Household has a BPL card [yes =1]	.654 (.473)	.027 {.047}	-.011 {.053}	.002 {.074}	[.882]

Notes: *** denotes significance at the 1% level, ** at the 5% level, * at the 10% level. Sample used is all observations at the baseline year (number of observations = 4,437, number of clusters = 36). Means, OLS estimates, standard deviations reported in parentheses, standard errors in curly brackets, p-values reported in squared brackets † Excludes institutional deliveries. For each row, OLS coefficients and corresponding standard errors in Columns 2-4, as well as p-values reported in Column 5, are obtained from an OLS regression of the corresponding characteristic on a dummy for whether the observation comes from the treatment or the control group, on a dummy for whether it comes from the microfinance or the non-microfinance stratum, and on the interaction between these two dummies, plus district dummies. Standard errors are clustered at the cluster level. Whether the respondent was considered to be literate or not was decided by the interviewer as a function of the respondent's ability to read a passage that was presented to her. The asset score is the the first principal of of dummy variables capturing ownership of the following household assets: bicycle, motorcycle, radio, fridge, tv, fan, tape, battery, generator, and electricity. Say over health care decisions is a dummy that equals one if the respondent mother reports to have a say in the household decision on whether she could seek care in case of illness.

Table 2.4: Village level characteristics

	Mean	OLS coefficients			P-value on F-test for (2), (3), and (4)
	No intervention	Health education	Microfinance	Interaction	
	(1)	(2)	(3)	(4)	(5)
Area	341.19	76.26	-110.55	23.97	[.056]
[in hectares]	(306.30)	{77.87}	{87.92}	{119.91}	
Households	136.30	34.66	38.42	-43.02	[.196]
[number]	(78.63)	{31.69}	{22.81}	{44.75}	
Household size	5.089	-.060	-.060	.237	[.560]
[number of persons per household]	(.569)	{.208}	{.109}	{.237}	
Distance to nearest town	24.59	3.75	-8.34	-8.55	[.398]
[in kilometers]	(21.10)	{10.26}	{8.84}	{12.34}	
General health care centre	.081	.012	.080	-.075	[.862]
[yes =1]	(.273)	{0.63}	{.093}	{.131}	
Maternal/child health centre	.058	.003	.380***	-.110	[.008]
[yes =1]	(.235)	{.079}	{.117}	{.199}	
Primary school	.732	.079	.000	-.036	[.719]
[yes =1]	(.443)	{.085}	{.060}	{.102}	
Middle/secondary school	.232	-.044	.011	.168	[.756]
[yes =1]	(.423)	{.097}	{.082}	{.177}	
Electricity for domestic use	.108	-.056	.016	.094	[.467]
[yes =1]	(.311)	{.050}	{.058}	{.084}	
Wet land	.041	.018	-.003	.078	[.175]
[% of irrigated agricultural land]	(.103)	{.027}	{.025}	{.052}	
Dry land	.462	.058	.090	-.094	[.619]
[% of unirrigated agricultural land]	(.214)	{.081}	{.077}	{.117}	

Notes: *** denotes significance at the 1% level, ** at the 5% level, * at the 10% level. Sample used is all observations at the baseline year in the districts of Seraikela-Karshwan and West Singhbhum (number of observations = 3,046, number of clusters = 24). Means, OLS estimates, standard deviations reported in parentheses, standard errors in curly brackets, p-values reported in squared brackets. For each row, OLS coefficients and corresponding standard errors in Columns 2-4, as well as p-values reported in Column 5, are obtained from an OLS regression of the corresponding characteristic on a dummy for whether the observation comes from the treatment or the control group, on a dummy for whether it comes from the microfinance or the non-microfinance stratum, and on the interaction between these two dummies, plus district dummies. Standard errors are clustered at the cluster level.

Table 2.5: Correlates of participation in the Ekjut meetings

		Microfinance versus non-microfinance strata		
	Full Sample	No microfinance stratum	Microfinance stratum	Difference (3) - (2)
	(1)	(2)	(3)	(4)
Microfinance intervention [=1 if cluster located in microfinance stratum]	-.107* (.057)			
Age at interview [years]	-.001 (.002)	.002 (.002)	-.003 (003)	[.214]
Age at marriage [years]	.012 (.007)	.014 (.010)	.006 (.005)	[.512]
Age at first pregnancy [years]	-.010 (.007)	-.012 (.007)	-.002 (008)	[.120]
Number of past pregnancies [number]	.052*** (.010)	.043*** (.011)	.060*** (.014)	[.382]
Literate [yes =1]	.030** (.014)	.030 (.023)	.020 (.016)	[.719]
Tribal member [yes =1]	.019 (.046)	.001 (.075)	.046 (.042)	[.609]
Hindu [yes =1]	-.037 (.042)	.002 (.055)	-.055 (.051)	[.974]
Say over health care decisions [yes =1]	.023 (.039)	-.021 (.058)	.034 (.039)	[.443]
Household assets [first principal component]	-.008 (.009)	-.026*** (.007)	.010 (.014)	[.032]
Household owns at least 2 bighas of land [yes =1]	.038 (.029)	.072 (.046)	-.001 (.024)	[.172]
Household has a BPL card [yes =1]	.025 (.016)	.039 (.023)	.008 (.016)	[.826]
District effects	Yes		Yes	
R-squared	.1063		.1232	
Number of observations (clusters)	9249 (18)		9249 (18)	

Notes. *** denotes significance at the 1% level, ** at the 5% level, * at the 10% level. Standard errors clustered at the cluster level. Dependent variable is a dummy =1 if the respondent mother participates in an Ekjut group, 0 otherwise. Standard errors reported in parentheses. Column 1 presents estimates obtained from a single regression. Column 2-4 presents estimates from a single regression.

Table 2.6: Estimates of intent-to-treat effects on neonatal and fetal mortality

	OLS	Logit	
	Marginal effect (1)	Marginal effect (2)	Odds ratio (3)
Main coefficients			
Health education only	-.040***	-.039***	.535**
⟨TE⟩	(.010)	(.009)	[.326-.878]
Microfinance only	-.017**	-.016**	.525***
⟨TM⟩	(.007)	(.007)	[.369-.746]
Interaction	.028**	.027**	1.625
⟨TEM⟩	(.012)	(.011)	[.881-2.999]
Additional coefficients			
Health education with microfinance	-.012*	-.012*	.996
⟨TE+TEM⟩	(.007)	(.006)	[.986-1.007]
Microfinance with health education	.011	.012	.995
⟨TM+TEM⟩	(.010)	(.008)	[.980-1.010]
Health education and microfinance	-.029***	-.027***	.978***
⟨TE+TM+TEM⟩	(.008)	(.008)	[.966-.990]
{Pseudo} R-squared	.0017	{.0041}	
Clusters	36	36	
Observations	22,370	22,370	

Notes: *** denotes significance at 1% level, ** at 5% level, * at 10% level. Mean marginal effects, odds ratios, BRL standard errors clustered at the cluster level reported in parentheses, 95% confidence intervals reported in brackets. Mean baseline probability of the infant dying in the first six weeks of life is 5.37%. Mean baseline probability of the infant being stillborn is 3.78%.

Table 2.7: Robustness

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Health education only	-.042***	-.043***	-.043***	-.041***	-.048**	-.046**	-.041**
⟨TE⟩	(.011)	(.011)	(.011)	(.010)	(.021)	(.022)	(.015)
Microfinance only	-.018**	-.020**	-.018**	-.019**	-.020***	-.023***	-.022**
⟨TM⟩	(.007)	(.007)	(.008)	(.007)	(.007)	(.008)	(.011)
Interaction	.029**	.031**	.027**	.029**	.037**	.036**	.040**
⟨TEM⟩	(.013)	(.007)	(.013)	(.012)	(.016)	(.017)	(.011)
Controlling for Ekjut treatment effect heterogeneity with:	None	Past pregnancies	Intrahousehold say	Household asset score	District effects	All individual/ household factors and district effects	Maternal/child health care centre
R-squared	.0017	.0027	.0018	.0025	.0032	.0050	.0041
Observations (clusters)	23,189 (36)	22,370 (36)	22,370 (36)	22,370 (36)	22,370 (36)	22,370 (36)	15,370 (24)

Notes: *** denotes significance at 1% level, ** at 5% level, at * 10% level. OLS estimates. Dependent variable is a dummy that equals one if the baby died within the first six weeks of life. BRL standard errors clustered at the cluster level reported in parentheses. Each column presents results obtained from a single regression. Columns 1-6 use the full dataset. Column 7 uses data from the districts of Seraikela-Karshwan and West Singhbhum, for which census information on village level amenities is available.

Table 2.8: Mechanisms

	Main parameters			Additional parameters				
	Health Education only ⟨TE⟩	Microfinance only ⟨TM⟩	Interaction ⟨TEM⟩	Health Education with Microfinance ⟨TE+TEM⟩	Microfinance with Health Education ⟨TM+TEM⟩	Health Education and Microfinance ⟨TE+TM+TEM⟩	R-squared	Observations [clusters]
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Dependent variable:								
Had health checkup [yes =1]	-.007 (.075)	-.102* (.059)	.192** (.095)	.185*** (.058)	.090 (.074)	.084 (.076)	.0179	22,370 [36]
Took iron tablets [yes =1]	-.007 (.048)	-.127** (.059)	.078 (.073)	.071 (.055)	-.049 (.043)	-.056 (.049)	.0101	22,370 [36]
Had tetanus injection [yes =1]	.023 (.041)	-.053 (.042)	.037 (.053)	.060* (.034)	-.016 (.032)	-.064 (.117)	.0119	22,370 [36]
Institutional delivery [yes =1]	-.040 (.030)	-.013 (.039)	.034 (.047)	-.006 (.036)	.021 (.026)	-.019 (.038)	.0097	22,370 [36]
Made payments to BA† [yes =1]	-.026 (.041)	.104** (.048)	-.087 (.069)	-.113** (.055)	.017 (.049)	-.009 (.050)	.0150	22,370 [36]
BA used safe-delivery kit† [yes =1]	.220*** (.069)	.056 (.065)	-.130 (.092)	.090 (.060)	-.074 (.065)	.146** (.066)	.0516	22,370 [36]
BA washed hands with soap† [yes =1]	.237** (.104)	-.025 (.084)	-.181 (.136)	.056 (.087)	-.206* (.106)	.031 (.087)	.0545	22,370 [36]
BA used plastic sheet† [yes =1]	.183*** (.051)	.021 (.057)	-.008 (.082)	.175*** (.064)	.013 (.059)	.196*** (.057)	.0631	22,370 [36]
BA tied cord with boiled thread† [yes =1]	.248*** (.086)	-.059 (.048)	-.070 (.125)	.177* (.092)	-.129 (.116)	.119 (.095)	.0753	22,370 [36]

Notes: *** denotes significance at 1% level, ** at 5% level, * at 10% level. OLS estimates. BRL standard errors clustered at the cluster level reported in parentheses. † Excludes institutional deliveries. Each row presents results obtained from a single regression.

Table 2.9: Heterogeneity of ITT of health education intervention with respect to household asset ownership

	Baseline	Household assets
Health education	-.040***	-.037***
	(.010)	(.010)
Health education X Household asset score	-	.019***
		(.007)
R-squared	.0017	.0026
Observations	11,293	11,293

Notes: *** denotes significance at the 1% level, ** at 5% level, * at 1% level. Dependent variable is a dummy that equals one if the baby died within the first six weeks of life. BRL standard errors clustered at the cluster level reported in parentheses. Each column presents results obtained from a single regression. Both columns use the full dataset. The asset score is the demeaned sum of dummy variables capturing ownership of the following household assets: bicycle, motorcycle, radio, fridge, tv, fan, tape, battery, generator, and electricity.

3

Multidimensional Assortative Matching

3.1 Introduction

In a unidimensional marriage market, Becker (1973) has shown that complementarity between partners' attributes in the joint output function leads to positive assortative matching, i.e. better (say) educated females match with better educated males. Chiappori, McCann and Nesheim (2010) argue that this much celebrated notion of positive assortative matching does not extend naturally to multi-dimensions. Yet, empirical evidence strongly suggests that individuals tend to match with, and have a preference for, similar partners in a variety of traits (Hitsch, Hortaçsu and Ariely, 2010; Banerjee et al., 2011).

In a multi-dimensional marriage market, I derive a simple and intuitive condition on the joint output function that ensures positive assortative matching on multiple traits. This condition naturally nests Becker's result for the unidimensional setting.[1] I then show

[1] However, in the multidimensional setting, supermodularity on a given trait, though necessary, is not sufficient for positive assortative matching on that trait.

that matching preferences are such that, away from their optimal partners, agents are willing to compensate mismatches on one attribute with opposite mismatches on other attributes. That is, it is all right for females (males) to be a little less good looking if they are also a little more educated, or a little less educated if they are also a little more good looking.

More broadly, this paper contributes to the literature on multidimensional matching in the following ways. First, I develop a transferable utility model that extends the unidimensional notion of assortative matching to a multidimensional setting. Using techniques from the mathematical literature on optimal transportation, Chiappori, McCann and Nesheim (2010) have analyzed the existence and uniqueness of multidimensional assignments under transferable utility, and have in particular established purity of the assignment. Yet, purity simply means that the assignment matching males to females is deterministic (i.e. one-to-one), and not necessarily assortative. In fact, the authors argue that the concept of (positive) assortative matching in the unidimensional setting is not well defined in a multidimensional setting. In this paper, I show that when agents traits are jointly uniformly distributed in the unit square, multidimensional positive assortative matching is well defined.

Second, I solve in closed form a model of matching along multiple continuous dimensions under transferable utility. Chiappori, Oreffice and Quintana-Domeque (2012*b*) and Banerjee et al. (2011) also solve in closed form a model of matching along two traits - the former within a transferable utility setting and the latter within a non-transferable utility setting - but they both consider the case where one of the traits is continuous and the other is discrete.

Third, I characterize the sorting trade-offs across different traits that agents face in the marriage market both along and outside the equilibrium path. This is a first attempt to derive iso-attractiveness curves in a "truly" multidimensional marriage market. Chiappori, Oreffice and Quintana-Domeque (2012*a*) also examine iso-attractiveness curves but in a context where multiple matching dimensions are collapsed into a single index. Furthermore, in this paper the equilibrium trade-offs that agents make in the marriage market can be thought as being the outcome of random search frictions a la Atakan (2006). Our analysis thus also departs from Galichon and Salani (2012), who explicitly model multidimensional matching but in a frictionless framework.

I acknowledge that after having independently discovered the main finding of this paper, which generalizes the unidimensional notion of assortative matching to a multidimensional setting, I learnt that Lindenlaub (2013) also independently made the same discovery.

Throughout the paper I assume that attributes are uniformly distributed on the unit square. Clearly this is a strong restriction, which would be easily be rejected in the data. Yet, this condition is sufficient, but not necessary, for two-dimensional positive assortative matching to hold. Although I do not explore this route in this paper, I believe that more complex distributions would still be possible as long there is a similar correlation structure between attributes across gender, i.e. symmetric copulas across gender. Otherwise some sort of non-assortative matching will follow trivially. However, leaving aside distributional concerns, this paper isolates in a precise, and hopefully intuitive away, the role of technology, namely of multidimensional complementarities in the joint output function, in explaining multidimensional assortative matching and mating preferences.

This paper is organized as follows. Section 3.2 sets out the model, and introduces two key objects: the affinity matrix and the sorting matrix. Section 3.3 derives the main result of this paper, which extends unidimensional assortative matching to a multidimensional setting. Section 3.4 counterfactually moves beyond the perfectly competitive environment, allowing me to derive mating preferences over the entire set of potential partners in the marriage market. Section 3.5 provides a parametric example, which allows me to solve the model in closed form. Section 3.6 concludes.

3.2 The Model

3.2.1 Agents and Marriage Market

The marriage market is populated by a continuum of males and females. A male is indexed by a two-dimensional vector of traits $\boldsymbol{x} = (x_1, x_2) \in [0, 1]^2$. The first and second traits, x_1 and x_2, are (say) his level of education and his looks, respectively. Similarly, a female is indexed by $\boldsymbol{y} = (y_1, y_2) \in [0,1]^2$, where y_1 is her level of education and y_2 is her looks. I assume that both $\boldsymbol{x}$ and $\boldsymbol{y}$ are uniformly distributed. Males and female match pairwise in a perfectly competitive marriage market environment.

3.2.2 Technology

When male $\boldsymbol{x}$ and female $\boldsymbol{y}$ form a match, they produce positive output $f(\boldsymbol{x}, \boldsymbol{y}) \geq 0$. It is assumed that f is twice continuously differentiable, with $D_{\boldsymbol{x}} f(\boldsymbol{x}, \boldsymbol{y}) > \mathbf{0}$ and $D_{\boldsymbol{y}} f(\boldsymbol{x}, \boldsymbol{y}) > \mathbf{0}$.

Here, as in the unidimensional setting studied by Becker (1973), complementarities across the partners' traits will be shown to play a critical role in determining who matches with whom in equilibrium. I now thus introduce the matrix of second-order cross-partial derivatives of $f(\boldsymbol{x}, \boldsymbol{y})$ with respect to $\boldsymbol{x}$ and $\boldsymbol{y}$,

$$D^2_{\boldsymbol{xy}} f(\boldsymbol{x}, \boldsymbol{y}) = \begin{bmatrix} \dfrac{\partial^2 f(\boldsymbol{x}, \boldsymbol{y})}{\partial x_1 \partial y_1} & \dfrac{\partial^2 f(\boldsymbol{x}, \boldsymbol{y})}{\partial x_1 \partial y_2} \\ \dfrac{\partial^2 f(\boldsymbol{x}, \boldsymbol{y})}{\partial x_2 \partial y_1} & \dfrac{\partial^2 f(\boldsymbol{x}, \boldsymbol{y})}{\partial x_2 \partial y_2} \end{bmatrix}, \tag{3.1}$$

which Dupuy and Galichon (2012) have termed the *affinity* matrix. The main diagonal isolates the *within*-traits complementarities. Specifically, the $(1,1)^{th}$ entry measures the degree of complementarity between partners' education levels, and the $(2,2)^{th}$ entry measures the degree of complementarity between partners' looks. The anti-diagonal, in turn, isolates the *between*-traits complementarities. Specifically, the $(2,1)^{th}$ entry measures the degree of complementarity between males' education and females' looks, and the $(1,2)^{th}$ entry measures the degree of complementarity between males' looks and females' education.

3.2.3 Assignment

The assignment matching males with females is governed by a measure-preserving mapping $\mu(\boldsymbol{x}) = \boldsymbol{y}$, which means that male $\boldsymbol{x}$ is married with female $\boldsymbol{y}$. Specifically, this mapping is a vector field $\mu(\boldsymbol{x}) = (\mu_1(\boldsymbol{x}), \mu_2(\boldsymbol{x})) = (y_1, y_2) = \boldsymbol{y}$, where for a given male $\boldsymbol{x}$ the first component gives his partner's education level, i.e. $\mu_1(\boldsymbol{x}) = y_1$, and the second component gives his partner's looks, i.e. $\mu_2(\boldsymbol{x}) = y_2$.

In this paper, I am interested in studying (positive) assortative assignments. An assignment is said to be assortative if (i) it is pure (one-to-one or deterministic), i.e. a male is matched with one and only one female, and (ii) monotonic. In the unidimensional setting, where agents are heterogeneous only with respect to (say) their education levels, a deterministic monotonic assignment means that the function $\mu_1(x_1) = y_1$ is either monotonically increasing (or decreasing). That is, more educated males are married with more (less) educated females.

In my two-dimensional setting there are two possible positive assortative assignments.[2] The first entails positive assortativeness within both traits, i.e. better educated males marry better educated females, and better looking males marry better looking females - which I term by *within-attribute* positive assortative matching. The second entails positive assortativeness between the two traits, i.e. better educated males marry better looking females, and better looking males marry better educated females - which I term by *between-attribute* positive assortative matching.

These two alternative two-dimensional positive assortative assignments can be equivalently defined in terms of the gradient of $\mu(\boldsymbol{x})$

$$D_{\boldsymbol{x}}\mu(\boldsymbol{x}) = \begin{bmatrix} \dfrac{\partial \mu_1(\boldsymbol{x})}{\partial x_1} & \dfrac{\partial \mu_2(x)}{\partial x_1} \\ \dfrac{\partial \mu_1(\boldsymbol{x})}{\partial x_2} & \dfrac{\partial \mu_2(x)}{\partial x_2} \end{bmatrix}, \tag{3.2}$$

which I term the *sorting* matrix. If the assignment is within-attribute positively assortative, then this matrix is the identity matrix. If, alternatively, the assignment is between-attribute positively assortative, then this matrix has ones on its anti-diagonal and zeros on its main diagonal.[3] Figure 3.1 illustrates these two alternative sorting patterns.

[2]Though I focus on positive assortative matching, my analysis and results apply to negative assortative matching.

[3]If a given component $\mu_i(\boldsymbol{x})$ is strictly increasing in x_j, then due their measure-preserving nature, it has to be the case that $\mu_i(\boldsymbol{x}) = x_j$. That is, a given male (female) attribute can be monotonically related with at most one female (male) attribute. As a result, $\mu(x_1, x_2) = (\mu_1(x_1), \mu_2(x_2)) = (x_1, x_2)$ under *within*-traits positive assortative matching, and $\mu(x_1, x_2) = (\mu_1(x_1), \mu_2(x_2)) = (x_2, x_1)$ under *between*-traits positive assortative matching.

3.2.4 Competitive Market Equilibrium

Matching takes place in a perfectly competitive environment. Men's attributes are observable and 'priced', with $p(\boldsymbol{x}, \boldsymbol{y})$ denoting the transfer received by a man with attributes $\boldsymbol{x}$ matched with a woman with attributes $\boldsymbol{y}$. Women's attributes are observable and priced, with $q(\boldsymbol{x}, \boldsymbol{y})$ denoting the transfer received by a woman with attributes $\boldsymbol{y}$ matched with a man with attributes $\boldsymbol{x}$. The output generated by match between a male $\boldsymbol{x}$ and a female $\boldsymbol{y}$ is thus given by $f(\boldsymbol{x}, \boldsymbol{y}) = p(\boldsymbol{x}, \boldsymbol{y}) + q(\boldsymbol{x}, \boldsymbol{y})$.

Taking the transfer schedules (or prices) as given, each woman chooses the man with whom she wishes to match. That is, female $\boldsymbol{y}$ solves

$$\max_{\boldsymbol{x}} f(\boldsymbol{x}, \boldsymbol{y}) - p(\boldsymbol{x}, \boldsymbol{y}). \tag{3.3}$$

The first-order condition equates marginal prices to marginal productivities along the equilibrium path

$$D_{\boldsymbol{x}} f(\boldsymbol{x}, \mu(\boldsymbol{x})) - D_{\boldsymbol{x}} p(\boldsymbol{x}, \mu(\boldsymbol{x})) = 0. \tag{3.4}$$

The second-order condition is that the Hessian matrix evaluated along the equilibrium path is negative definite

$$\mathcal{H} = D^2_{\boldsymbol{xx}} f(\boldsymbol{x}, \mu(\boldsymbol{x})) - D^2_{\boldsymbol{xx}} p(\boldsymbol{x}, \mu(\boldsymbol{x})) \prec 0. \tag{3.5}$$

3.3 Multidimensional Assortative Matching

This section presents the main result of this paper. It derives sufficient conditions on the technology, namely on the affinity matrix (3.1), in order to obtain within-attribute positive assortative matching in equilibrium, i.e. for the sorting matrix (3.2) to be the identity matrix.

Following Eeckhout and Kircher (2010), if I totally differentiate the first-order condition (3.4) with respect to $\boldsymbol{x}$, and use (3.4), yields

$$D^2_{\boldsymbol{xx}} f(\boldsymbol{x}, \mu(\boldsymbol{x})) - D^2_{\boldsymbol{xx}} p(\boldsymbol{x}, \mu(\boldsymbol{x})) = -D^2_{\boldsymbol{xy}} f(\boldsymbol{x}, \mu(\boldsymbol{x})) \times D_{\boldsymbol{x}} \mu(\boldsymbol{x}). \tag{3.6}$$

The second-order condition (3.5) is thus equivalent to positive definiteness of the product between the affinity matrix (3.1) and the sorting matrix (3.2)[4]

$$D^2_{\boldsymbol{xy}} f(\boldsymbol{x}, \mu(\boldsymbol{x})) \times D_{\boldsymbol{x}}\mu(\boldsymbol{x}) = \begin{bmatrix} \dfrac{\partial^2 f(\boldsymbol{x}, \mu(\boldsymbol{x}))}{\partial x_1 \partial y_1} & \dfrac{\partial^2 f(\boldsymbol{x}, \mu(\boldsymbol{x}))}{\partial x_1 \partial y_2} \\ \dfrac{\partial^2 f(\boldsymbol{x}, \mu(\boldsymbol{x}))}{\partial x_2 \partial y_1} & \dfrac{\partial^2 f(\boldsymbol{x}, \mu(\boldsymbol{x}))}{\partial x_2 \partial y_2} \end{bmatrix} \begin{bmatrix} \dfrac{\partial \mu_1(\boldsymbol{x})}{\partial x_1} & \dfrac{\partial \mu_2(\boldsymbol{x})}{\partial x_1} \\ \dfrac{\partial \mu_1(\boldsymbol{x})}{\partial x_2} & \dfrac{\partial \mu_2(\boldsymbol{x})}{\partial x_2} \end{bmatrix} \succ 0. \tag{3.7}$$

I am now ready to characterize a set of sufficient conditions on the joint production function $f(\boldsymbol{x}, \boldsymbol{y})$ in order to obtain within-attribute positive assortative matching. To arrive at these conditions, simply observe that since under such assignment the sorting matix is the identity matrix, we have that $D^2_{\boldsymbol{xy}} f(\boldsymbol{x}, \mu(\boldsymbol{x})) \times D_{\boldsymbol{x}}\mu(\boldsymbol{x}) = D^2_{\boldsymbol{xy}} f(\boldsymbol{x}, \mu(\boldsymbol{x}))$. Hence, the second-order condition in (3.7) boils down to symmetric positive definiteness of the affinity matrix. As a result, within-attribute positive assortative matching obtains when: (*i*) partners' educational levels are complementary, $\partial^2 f(\boldsymbol{x}, \boldsymbol{y})/\partial x_1 \partial y_1 > 0$, (*ii*) partners' looks are complementary, $\partial^2 f(\boldsymbol{x}, \boldsymbol{y})/\partial x_1 \partial y_1 > 0$, and (*iii*) the determinant of the affinity matrix is positive, which given monotonicity of the square root is equivalent to

$$\sqrt{\frac{\partial^2 f(\boldsymbol{x}, \boldsymbol{y})}{\partial x_1 \partial y_1} \times \frac{\partial^2 f(\boldsymbol{x}, \boldsymbol{y})}{\partial x_2 \partial y_2}} > \sqrt{\frac{\partial^2 f(\boldsymbol{x}, \boldsymbol{y})}{\partial x_1 \partial y_2} \times \frac{\partial^2 f(\boldsymbol{x}, \boldsymbol{y})}{\partial x_2 \partial y_1}},$$

i.e. the geometric average within-attribute complementarity is larger than the geometric average between-trait complementarity.

Proposition 1 (Matching of likes)**.** *In a two-dimensional marriage market with transferable utility, where individuals' attributes are distributed uniformly on the unit square, the equilibrium matching will be positively assortative along both dimensions if the joint output function is supermodular along both attributes, and the geometric average within-supermodularity is greater than the geometric average between-supermodularity.*

Also observe that the equilibrium further requires the matrix in (3.7) to be symmetric. To see this, observe that positive definiteness of $D^2_{\boldsymbol{xy}} f(\boldsymbol{x}, \mu(\boldsymbol{x})) \times D_{\boldsymbol{x}}\mu(\boldsymbol{x})$ only ensures that the assignment $\mu(\boldsymbol{x})$ is a maximum of the objective function in (3.3). But for this

[4] Negative definiteness of the Hessian matrix $\mathcal{H}$ means that its leading principal minors $|\mathcal{H}_k|$, for $k = \{1, 2\}$, are negative for odd k and positive for even k. But, since $|\mathcal{H}_k| = \left|-(D^2_{\boldsymbol{xy}} f(\boldsymbol{x}, \boldsymbol{y}) D_{\boldsymbol{x}}\mu(\boldsymbol{x}))_k\right| = (-1)^k \left|(D^2_{\boldsymbol{xy}} f(\boldsymbol{x}, \boldsymbol{y}) D_{\boldsymbol{x}}\mu(\boldsymbol{x}))_k\right|$, this is equivalent to the 2 leading principal minors of $D^2_{\boldsymbol{xy}} f(\boldsymbol{x}, \boldsymbol{y}) D_{\boldsymbol{x}}\mu(\boldsymbol{x})$ being both positive.

assignment to be a competitive equilibrium, the price schedule $p(\boldsymbol{x})$ sustaining it must exist. Standard integrability conditions from multivariate calculus (see, e.g., Theorem 3.4 in Lang (1997)), tell us that a necessary and sufficient condition for such function to exist is symmetry of the matrix of its second-order partial derivatives $D^2_{\boldsymbol{xx}}p(\boldsymbol{x})$, i.e. $\partial^2 p(\boldsymbol{x})/\partial x_1 \partial x_2 = \partial^2 p(\boldsymbol{x})/\partial x_2 \partial x_1$. From (3.6), this matrix is equal to

$$D^2_{\boldsymbol{xx}}p(\boldsymbol{x}) = D^2_{\boldsymbol{xx}}f(\boldsymbol{x}, \mu(\boldsymbol{x})) + D^2_{\boldsymbol{xy}}f(\boldsymbol{x}, \mu(\boldsymbol{x})) \times D_{\boldsymbol{x}}\mu(\boldsymbol{x}).$$

Since by continuity of $f(\boldsymbol{x}, \boldsymbol{y})$, $D^2_{\boldsymbol{xx}}f(\boldsymbol{x}, \mu(\boldsymbol{x}))$ is symmetric, then symmetry of $D^2_{\boldsymbol{xx}}p(\boldsymbol{x})$ requires symmetry of $D^2_{\boldsymbol{xy}}f(\boldsymbol{x}, \mu(\boldsymbol{x})) \times D_{\boldsymbol{x}}\mu(\boldsymbol{x})$.

At this point it is useful to relate the main result of this paper to Chiappori, McCann and Nesheim (2010), who argue that the notion of positive assortative matching cannot be extended to a multidimensional setting. These authors have shown that if $D_{\boldsymbol{x}}f(\boldsymbol{x}, \boldsymbol{y})$ is injective with respect to $\boldsymbol{y}$ (the *twist condition*), then the assignment is pure, i.e a male is matched with one and only one female, for any distributions of $\boldsymbol{x}$ and of $\boldsymbol{y}$.[5] My result is a special case of theirs since my condition that $D^2_{\boldsymbol{xy}}f(\boldsymbol{x}, \boldsymbol{y})$ is positive definite ensures that $D_{\boldsymbol{x}}f(\boldsymbol{x}, \boldsymbol{y})$ is injective with respect to $\boldsymbol{y}$, and the resulting positive assortative assignment that I obtain implies purity. What allows me to go beyond purity and extend the unidimensional notion of positive assortative matching to a multidimensional setting, are my simplifying distributional assumptions, namely that $\boldsymbol{x}$ and $\boldsymbol{y}$ are both uniformly distributed.

Finally observe that my main result nests Becker's unidimensional assortative matching result. To see this, suppose that agents are heterogeneous with respect to one attribute only, say education. In such case, our conditions would boil down to

$$\frac{\partial^2 f(x_1, y_1)}{\partial x_1 \partial y_1} \times \frac{\partial \mu_1(x_1)}{\partial x_1} > 0,$$

and positive assortative matching on education, i.e. $\partial \mu_1(x_1)/\partial x_1 = 1$, would emerge under complementarity in partners' educational levels, i.e. $\partial^2 f(x_1, y_1)/\partial x_1 \partial y_1 > 0$.

[5] See also Dizdar and Moldovanu (2013) and Dizdar (2013) who also use the twist condition to examine multidimensional assignments with transferable utility.

3.4 Mating Preferences Across Potential Partners

The goal of this section is to derive individuals' mating preferences across the entire set of potential partners. The previous section allowed us to derive payoffs only along the perfectly competitive equilibrium matching path. In order to derive individuals' payoffs under alternative matches, I thus need to adopt some non-competitive procedure that splits the joint output between partners in such matches.

To do so, I follow the standard approach that connects bargaining theory with the theory of competitive equilibrium in markets. Specifically, I assume that every *potential* couple Nash bargains the split of their joint output, with the disagreement points being given by their perfectly competitive payoffs. The word potential is emphasised because in a perfectly competitive market these couples will never form. The resulting payoffs can thus be interpreted as being *counterfactual*, in the sense that they give the payoff a person would have had if instead of matching with the perfectly competitive partner he or she had matched with someone else. Clearly, in real marriage markets that are subject to imperfections, such as search frictions, these non-ideal matches may actually take place.

Hence, for any male $\boldsymbol{x}$ and any female $\boldsymbol{y}$ that potentially matches together, they will Nash bargain their joint output, $f(\boldsymbol{x}, \boldsymbol{y})$, in excess of their perfectly competitive payoffs, which I denote by $p^\star(\boldsymbol{x})$ and $q^\star(\boldsymbol{y})$ respectively. Under symmetric (exogenous) bargaining power, each partner thus obtains half of this surplus, $(f(\boldsymbol{x}, \boldsymbol{y}) - p^\star(\boldsymbol{x}) - q^\star(\boldsymbol{y}))/2$, plus his or her perfectly competitive payoff. Anchoring the analysis around the latter, it follows that the marital payoff of any male $\boldsymbol{x}$ as a function of $\boldsymbol{y} = \boldsymbol{x} + \boldsymbol{\delta}$, with $\boldsymbol{\delta} = (\delta_1, \delta_2) = (y_1 - x_1, y_2 - x_2) \in [-1, 1]^2$, is given by

$$p^\diamond(\boldsymbol{x}, \boldsymbol{x} + \boldsymbol{\delta}) = \frac{1}{2}[f(\boldsymbol{x}, \boldsymbol{x} + \boldsymbol{\delta}) + p^\star(\boldsymbol{x}) - q^\star(\boldsymbol{x} + \boldsymbol{\delta})]. \tag{3.8}$$

3.4.1 Single-Peaked Mating Preferences

The first thing to note is that individuals' payoffs are single peaked at $\boldsymbol{\delta} = \mathbf{0}$, i.e. at their (perfectly competitive) optimal partners $\boldsymbol{y} = \boldsymbol{x}$. To see this, observe that a critical point of (3.8) is found by solving $D_{\boldsymbol{y}} f(\boldsymbol{x}, \boldsymbol{x} + \boldsymbol{\delta}) - D_{\boldsymbol{y}} q^\star(\boldsymbol{x} + \boldsymbol{\delta}) = \mathbf{0}$. But since from the previous section (see first-order condition (3.4)) we know that under the competitive equilibrium assignment, marginal transfers (or prices) equal marginal productivities, we have that $D_{\boldsymbol{y}} q^{\star}(\boldsymbol{x}+\boldsymbol{\delta}) = D_{\boldsymbol{y}} f(\boldsymbol{x}+\boldsymbol{\delta}, \boldsymbol{x}+\boldsymbol{\delta})$, and therefore a critical point must solve

$$D_{\boldsymbol{y}} f(\boldsymbol{x}, \boldsymbol{x}+\boldsymbol{\delta}) - D_{\boldsymbol{y}} f(\boldsymbol{x}+\boldsymbol{\delta}, \boldsymbol{x}+\boldsymbol{\delta}) = \mathbf{0}. \tag{3.9}$$

Naturally, equation (3.9) holds true at $\boldsymbol{\delta} = \mathbf{0}$. To confirm that the zero mismatch is the unique global maximiser, I check that the matrix of second-order derivates of the payoff function (3.8) with respect to $\boldsymbol{\delta}$ is negative definite at $\boldsymbol{\delta} = \mathbf{0}$. This matrix, obtained by differentiating the left-hand side of (3.9) with respect to $\boldsymbol{\delta}$, is given by

$$D^2_{\boldsymbol{yy}} f(\boldsymbol{x}, \boldsymbol{x}+\boldsymbol{\delta}) - D^2_{\boldsymbol{xy}} f(\boldsymbol{x}+\boldsymbol{\delta}, \boldsymbol{x}+\boldsymbol{\delta}) - D^2_{\boldsymbol{yy}} f(\boldsymbol{x}+\boldsymbol{\delta}, \boldsymbol{x}+\boldsymbol{\delta}), \tag{3.10}$$

which at $\boldsymbol{\delta} = \mathbf{0}$ collapses to $-D^2_{\boldsymbol{xy}} f(\boldsymbol{x}, \boldsymbol{x})$. Now, since $D^2_{\boldsymbol{xy}} f(\boldsymbol{x}, \boldsymbol{y})$ is positive definite, it follows that (3.10) is negative definite, thus confirming that $\boldsymbol{\delta} = \mathbf{0}$ is indeed the unique global maximiser. Hence, as in the unidimensional setting (see, e.g. Shimer and Smith, 2000), multidimensional positive assortative matching is associated with an endogenous preference for likes.

Moving now to individuals' marital payoffs with partners other than their soul mates, I focus on two types of such mismatches. First, same-directional mismatches: where one of the partners is more attractive than the other on both attributes ($\delta_1 > 0$ and $\delta_2 > 0$, or $\delta_1 < 0$ and $\delta_2 < 0$). Second, opposite-directional mismatches: where one of the partners is more attractive than the other on one of the traits, but less attractive on the other trait ($\delta_1 > 0$ and $\delta_2 < 0$, or $\delta_1 < 0$ and $\delta_2 > 0$). The question is thus: what is the differential reduction in martial payoffs induced by these two types of mismatches?

3.4.2 Opposite-Direction Mismatches

Let me first analyse the change in male $\boldsymbol{x}$'s payoff $p^{\diamond}(\boldsymbol{x}, \boldsymbol{x}+\boldsymbol{\delta})$ induced by an opposite-directional marginal mismatch. Specifically, and to simplify the algebra, consider the case where the magnitudes of the mismatches on both dimensions are equal in absolute value, i.e. $\delta_1 = -\delta_2 > 0$. Clearly, I am not interested in evaluating this difference at the critical point $\boldsymbol{\delta} = \mathbf{0}$, because there the curvature of the payoff function is flat. Instead, I evaluate it at $\tilde{\boldsymbol{y}} = (x_1 + k, x_2 - k)$ for a small $k > 0$. Thus, the change in $p^\diamond(\boldsymbol{x}, \tilde{\boldsymbol{y}})$ induced by an infinitesimally small increase in k, is given by the directional derivative of $p^\diamond(\boldsymbol{x}, \tilde{\boldsymbol{y}})$ along the vector $\hat{\boldsymbol{v}} = \langle 0, 0, 1, -1 \rangle$, i.e.

$$\nabla_{\tilde{\boldsymbol{v}}}\, p^\diamond(\boldsymbol{x}, \tilde{\boldsymbol{y}}) = \frac{\partial p^\diamond(\boldsymbol{x}, \tilde{\boldsymbol{y}})}{\partial y_1} - \frac{\partial p^\diamond(\boldsymbol{x}, \tilde{\boldsymbol{y}})}{\partial y_2} = \left(\frac{\partial f(\boldsymbol{x}, \tilde{\boldsymbol{y}})}{\partial y_1} - \frac{\partial f(\tilde{\boldsymbol{y}}, \tilde{\boldsymbol{y}})}{\partial y_1} \right) - \left(\frac{\partial f(\boldsymbol{x}, \tilde{\boldsymbol{y}})}{\partial y_2} - \frac{\partial f(\tilde{\boldsymbol{y}}, \tilde{\boldsymbol{y}})}{\partial y_2} \right). \tag{3.11}$$

Equation (3.11) indicates that the effect on $p^\diamond(\boldsymbol{x}, \hat{\boldsymbol{y}})$ induced by this type of mismatch, acts through the following channels. First, an increase in his partner's looks, over and above his own looks, will increase both the output they produce together (contributing to an increase in his payoff) as well as the output she produces with her outside option partner (contributing to a decrease in his payoff). Second, a decrease in his partners' educational level, over and below his own educational level, will decrease both the output they produce together (contributing to a decrease in his payoff) as well as the output she produces with her outside option partner (contributing to an increase in his payoff). To understand the sign and the overall magnitude of (3.11), let k be infinitesimally small. This implies that (3.11) becomes the directional derivative of $\nabla_{\hat{\boldsymbol{v}}}\, p^\diamond(\boldsymbol{x}, \tilde{\boldsymbol{y}})$ along the vector $\tilde{\boldsymbol{u}} = \langle -1, 1, 0, 0 \rangle$, i.e.

$$\begin{aligned} \nabla_{\tilde{\boldsymbol{u}}} \left(\nabla_{\tilde{\boldsymbol{v}}}\, p^\diamond(\boldsymbol{x}, \tilde{\boldsymbol{y}}) \right) &= \nabla_{\tilde{\boldsymbol{u}}} \left(\frac{\partial f(\tilde{\boldsymbol{y}}, \tilde{\boldsymbol{y}})}{\partial y_1} \right) - \nabla_{\hat{\boldsymbol{u}}} \left(\frac{\partial f(\tilde{\boldsymbol{y}}, \tilde{\boldsymbol{y}})}{\partial y_2} \right) \\ &= -\frac{\partial^2 f(\tilde{\boldsymbol{y}}, \tilde{\boldsymbol{y}})}{\partial x_1 \partial y_1} - \frac{\partial^2 f(\tilde{\boldsymbol{y}}, \tilde{\boldsymbol{y}})}{\partial x_2 \partial y_2} + \frac{\partial f(\tilde{\boldsymbol{y}}, \hat{\boldsymbol{y}})}{\partial x_1 \partial y_2} + \frac{\partial f(\tilde{\boldsymbol{y}}, \tilde{\boldsymbol{y}})}{\partial x_2 \partial y_1}. \end{aligned} \tag{3.12}$$

Equation (3.12) shows that an opposite-directional mismatch has two competing effects on male $\boldsymbol{x}$'s payoff. One is negative and driven by the within-attributes complementarities, the other one is positive and driven by the between-attributes complementarities. To understand these effects, it is crucial to observe that when male $\boldsymbol{x}$ matches with a female that is better good looking but less educated than himself, this female's outside option partner is also better good looking but less educated than him.

As a result, the positive impact of her improved looks on their joint output is, on one hand, smaller, and on the other hand, larger, than that on the output she produces together with her outside option partner. Smaller, because $\boldsymbol{x}$ is not as good looking as her outside option partner, and males' and females' looks complement each other. Larger, because $\boldsymbol{x}$ is better educated than her outside option partner, and males' educational levels and females' looks are complementary.

Similarly, the negative impact of her reduced education on their joint output is, on one hand, larger, and on the other hand, smaller, than that on the output she produces together with her outside option. Larger, because $\boldsymbol{x}$ is better educated than her outside option partner, and males and females' educational levels complement each other. Smaller, because $\boldsymbol{x}$ is not as good looking as her outside option partner, and males' looks complement females' educational levels.

Finally, it is easy to show that when the geometric average within-attribute complementarity is larger than the geometric average between-attribute complementarity, as assumed, the overall effect given by (3.12) is negative. To ease notation, let f_{ij} denote $\partial^2 f/\partial x_i \partial x_j$. The goal is then to show that if $f_{11}f_{22} > f_{12}^2$, then $f_{11} + f_{22} > 2f_{12}$. Suppose otherwise, i.e $f_{12} = (f_{11} + f_{22} + k)/2$ with $k > 0$. In such case, $f_{11}f_{22} > ((f_{11} + f_{22} + k)/2)^2$ would have to hold. An impossibility, since this inequality is equivalent to $(f_{11} - f_{22})^2 < -(2(f_{11} + f_{22}) + k)k$, which clearly cannot hold as a positive number cannot be smaller than a negative number.

3.4.3 Same-Direction Mismatches

I now analyse the change in $p^\diamond(\boldsymbol{x}, \boldsymbol{x}+\boldsymbol{\delta})$ induced by a same-directional marginal mismatch. Specifically, consider the case where the magnitudes of the mismatches on both dimensions are equal and positive, i.e. $\delta_1 = \delta_2 > 0$. For the same reason as before, I do not want to evaluate this difference at the critical point $\boldsymbol{\delta} = \mathbf{0}$, i.e. $\boldsymbol{y} = \boldsymbol{x}$. Instead, I evaluate it at $\hat{\boldsymbol{y}} = (x_1 + k, x_2 + k)$ for a small $k > 0$. The change in $p^\diamond(\boldsymbol{x}, \hat{\boldsymbol{y}})$ induced by an infinitesimally small increase in k, is thus given by the directional derivative of $p^\diamond(\boldsymbol{x}, \hat{\boldsymbol{y}})$ along the vector $\hat{\boldsymbol{v}} = \langle 0, 0, 1, 1 \rangle$, i.e.

$$\nabla_{\hat{\boldsymbol{v}}}\, p^\diamond(\boldsymbol{x}, \hat{\boldsymbol{y}}) = \frac{\partial p^\diamond(\boldsymbol{x}, \hat{\boldsymbol{y}})}{\partial y_1} + \frac{\partial p^\diamond(\boldsymbol{x}, \hat{\boldsymbol{y}})}{\partial y_2} = \left(\frac{\partial f(\boldsymbol{x}, \hat{\boldsymbol{y}})}{\partial y_1} - \frac{\partial f(\hat{\boldsymbol{y}}, \hat{\boldsymbol{y}})}{\partial y_1}\right) + \left(\frac{\partial f(\boldsymbol{x}, \hat{\boldsymbol{y}})}{\partial y_2} - \frac{\partial f(\hat{\boldsymbol{y}}, \hat{\boldsymbol{y}})}{\partial y_2}\right). \tag{3.13}$$

Equation (3.13) shows that an increase in his partner's educational level and looks, over and above his own educational level and looks, will increase both the output they produce together (contributing to an increase in his payoff) as well as the output she produces with her outside option partner (contributing to a decrease in his payoff). Again, making k infinitesimally small, equation (3.13) becomes the directional derivative of $\nabla_{\hat{\boldsymbol{v}}}\, p^{\diamond}(\boldsymbol{x}, \hat{\boldsymbol{y}})$ along the vector $\hat{\boldsymbol{u}} = \langle -1, -1, 0, 0 \rangle$, i.e.

$$\begin{aligned} \nabla_{\hat{\boldsymbol{u}}} \left(\nabla_{\hat{\boldsymbol{v}}}\, p^{\diamond}(\boldsymbol{x}, \hat{\boldsymbol{y}}) \right) &= \nabla_{\hat{\boldsymbol{u}}} \left(\frac{\partial f(\hat{\boldsymbol{y}}, \hat{\boldsymbol{y}})}{\partial y_1} \right) + \nabla_{\hat{\boldsymbol{u}}} \left(\frac{\partial f(\hat{\boldsymbol{y}}, \hat{\boldsymbol{y}})}{\partial y_2} \right) \\ &= -\frac{\partial^2 f(\hat{\boldsymbol{y}}, \hat{\boldsymbol{y}})}{\partial x_1 \partial y_1} - \frac{\partial^2 f(\hat{\boldsymbol{y}}, \hat{\boldsymbol{y}})}{\partial x_2 \partial y_2} - \frac{\partial f(\hat{\boldsymbol{y}}, \hat{\boldsymbol{y}})}{\partial x_1 \partial y_2} - \frac{\partial f(\hat{\boldsymbol{y}}, \hat{\boldsymbol{y}})}{\partial x_2 \partial y_1}. \end{aligned} \tag{3.14}$$

Equation (3.14) shows that the overall effect on $p^{\diamond}(\boldsymbol{x}, \hat{\boldsymbol{y}})$ induced by this mismatch is unambiguously negative and larger than that induced by an opposite-directional mismatch. This is because her higher attributes are more valuable when interacted with her soul mate's, also higher, attributes.[6]

In sum, the following can be concluded. Individuals have single-peaked mating preferences, which peak at their soul mate partners. Away from these, they particularly dislike partners who are either more or less attractive than themselves on both attributes. Thus, conditional on mismatching, they still aim for some sort of balance, in the sense that they prefer mismatches on a given attribute to be counterbalanced by an opposite mismatch on the other attribute.

Proposition 2 (Compensatory Mismatches)**.** *In a two-dimensional marriage market with transferable utility, where individuals attributes are uniformly distributed on the unit square, and the geometric average within-attribute complementarity is greater than the geometric average between-attribute complementarity: away from their ideal partners, individuals are willing to counterbalance a mismatch on one attribute with an opposite mismatch on the other attribute.*

[6] The results for $\delta_1 = \delta_2 < 0$ are exactly the same, only the interpretation changes. In that case, his payoff drops because the reduction in the joint output function from matching with a woman less attractive than him on both attributes more than outweighs the reduction in her outside option payoff.

3.5 Parametric Example

3.5.1 Multidimensional Assortative Matching

The main goal here is to derive closed-form expressions for the males' and females' payoffs along the competitive equilibrium path, i.e. $p^\star(\boldsymbol{x})$ and $q^\star(\boldsymbol{y})$ respectively. Consider thus the following quadratic joint production function

$$f(\boldsymbol{x},\boldsymbol{y}) = \boldsymbol{x}'\boldsymbol{A}\boldsymbol{y}, \tag{3.15}$$

where $\boldsymbol{x} = (x_1, x_2)$, $\boldsymbol{y} = (y_1, y_2)$, and $\boldsymbol{A}$ is a 2-by-2 matrix, with $A_{ij} = \partial^2 f(\boldsymbol{x},\boldsymbol{y})/\partial x_i \partial y_j > 0$. That is, $\boldsymbol{A}$ is the affinity matrix and both attributes are assumed to be complementary. I further assume that this matrix is positive definite, i.e. $A_{11}A_{22} > A_{12}A_{21}$.

A woman $\boldsymbol{y}$ chooses a male $\boldsymbol{x}$ in order to maximize her share of the marital payoff $f(\boldsymbol{x},\boldsymbol{y}) - p(\boldsymbol{x})$. The two first-order conditions of this problem are given by

$$p_{x_1}(\boldsymbol{x}) = A_{11}y_1 + A_{12}y_2 \tag{3.16a}$$

$$p_{x_2}(\boldsymbol{x}) = A_{22}y_2 + A_{21}y_1, \tag{3.16b}$$

which have to hold along the competitive equilibrium path $\mu(\boldsymbol{x})$. Since I have assumed that $A_{11}A_{22} > A_{12}A_{21}$, I know from the previous analysis that the competitive equilibrium entails likes sorting with likes along both attributes, i.e. $(y_1, y_2) = \mu(\boldsymbol{x}) = (x_1, x_2)$. Substituting this into the first-order conditions (3.16a) and (3.16b) gives

$$p_{x_1}(\boldsymbol{x}) = A_{11}x_1 + A_{12}x_2 \tag{3.17a}$$

$$p_{x_2}(\boldsymbol{x}) = A_{22}x_2 + A_{21}x_1. \tag{3.17b}$$

In order to exist, the system of prices $p(\boldsymbol{x})$ that decentralizes this assignment must thus simultaneously satisfy (3.17a) and (3.17b). Partially integrating (3.17a) with respect to x_1, yields

$$p(\boldsymbol{x}) = A_{11}\frac{{x_1}^2}{2} + A_{12}x_1x_2 + c_2(x_2), \tag{3.18}$$

which cannot yet be the true price-schedule as I still need to find an expression for the function $c_2(x_2)$. To do so, I can partially differentiate (3.18) with respect to x_2, to obtain

$$p_{x_2}(\boldsymbol{x}) = A_{12}x_1 + c_2'(x_2), \tag{3.19}$$

which, in turn, must be equal to (3.17b). Thus, I have that

$$c'(x_2) = A_{22}x_2 + (A_{21} - A_{12})x_1. \tag{3.20}$$

Now, clearly, I can only integrate $c'(x_2)$ back to $c(x_2)$ if $A_{21} = A_{12}$, i.e. if the affinity matrix is symmetric. Thus, letting $A_{12} = A_{21}$, I obtain

$$c(x_2) = A_{22}\frac{x_2^2}{2} + c, \tag{3.21}$$

where c is a constant of integration (which I assume to be equal to zero).

Substituting (3.21) into (3.18) yields the price schedule for males along the competitive equilibrium path

$$p(\boldsymbol{x}) = A_{11}\frac{{x_1}^2}{2} + A_{22}\frac{{x_2}^2}{2} + A_{12}x_1x_2. \tag{3.22}$$

An analogous argument shows that the price schedule for females along the competitive equilibrium path is given by

$$q(\boldsymbol{y}) = A_{11}\frac{{y_1}^2}{2} + A_{22}\frac{{y_2}^2}{2} + A_{12}y_1y_2. \tag{3.23}$$

3.5.2 Mating Preferences Across Potential Partners

I can now derive closed-form expressions for individuals' payoff functions across the entire set of potential partners. Substituting the assumed specification for the joint output function, given by equation (3.15), and the males' and females' payoffs along the competitive equilibrium path, given by equations (3.22) and (3.23) respectively, into the generic males' marital payoff function, given by equation (3.8), gives

$$p^{\diamond}(\boldsymbol{x}, \boldsymbol{x}+\boldsymbol{\delta}) = \frac{1}{2}\left[A_{11}\left(x_1^2 - \frac{\delta_1^2}{2}\right) + A_{22}\left(x_2^2 - \frac{\delta_1^2}{2}\right) + 2A_{12}\left(x_1x_2 - \frac{\delta_1\delta_2}{2}\right)\right]. \tag{3.24}$$

Expression (3.24) reveals that the contours of the marital payoff function of a given male x as a function of the distance to the female he potentially matches with, $(\delta_1, \delta_2) = (y_1 - x_1, y_2 - x_2) \in [-1, 1]^2)$, are concentric oblique ellipses around his optimal partner in the perfectly competitive market, i.e. with centre at $\boldsymbol{\delta} = \mathbf{0}$. Figure 3.2 illustrates these contours for a given male $\boldsymbol{x}$.

Each elliptical contour represents a set of equally attractive partners. The gain in the marital payoff function increases as the ellipse gets smaller. Individuals thus display a clear preference for likes, i.e. for mates with similar attributes, since their marital payoffs increase the smaller the distance between their potential partners' attributes and their own attributes. The rotated or oblique shape of the ellipses captures the finding summarized in Proposition 2 that individuals have a stronger aversion for partners who are either more or less attractive than themselves on both traits, than for partners who are more attractive than themselves on one attribute but less attractive on the other attribute. A mismatch of the former type (same-directional mismatch) entails a larger reduction in the marital surplus than a mismatch of the latter type (opposite-directional mismatch). Observe that this is so because there are non-zero between-attribute complementarities, i.e. $A_{12} \neq 0$.

If searching for a partner entails an explicit search cost as in Atakan (2006), then this male x will accept any partner whose attractiveness does not follow below some minimum threshold. In our example, this minimum quality is given by the red contour in Figure 3.2. Given transferable utility, this male will thus in equilibrium match with any female located in the region inside that contour.

3.6 Conclusion

Gary Becker (1973) has shown that in a unidimensional context if (say) partners' educational levels are complementary, then better educated females will match with better educated males. But suppose that (say) partners' looks also complement each other. Can we conclude than the equilibrium assigment entails positive assortative matching on both dimensions? That is, that better educated males match with better educated females, and more attractive males match with more attractive females?

This paper shows that the answer to this question is: not necessarily. This is because when multiple traits are *simultaneously* taken into account, it may well be the case that the gain of matching better educated individuals with each other and good looking individuals with each other, is outweighed by the gain of *cross* matching better educated individuals with good looking individuals. Hence, assortative matching on both education and looks only emerges if the within-traits complementarities (complementarity between partners' educational levels, and complementarity between partners' looks) outweighs the between-traits complementarities (complementarity between his education and her looks, and complementarity between his looks and her education).

This paper further examines in closed form the sorting trade-offs that agents face outside the equilibrium path, allowing me to characterize their iso-attractiveness curves in the marriage market. An interesting direction for further research would be to extend my model into an empirical framework that would allow the identification and estimation of agents' iso-attractivenes maps in the marriage market, and ultimately of complementarities in the marital output function, as Choo and Siow (2006), Galichon and Salani (2012), and Dupuy and Galichon (2012), do but without providing a closed form.

3.A Appendix: Figures

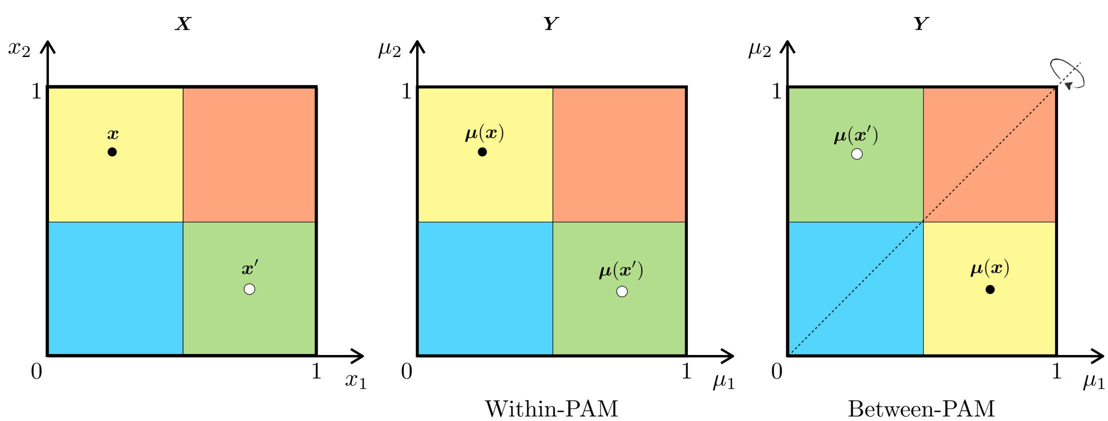

Figure 3.1: Within and between positive assortative matching

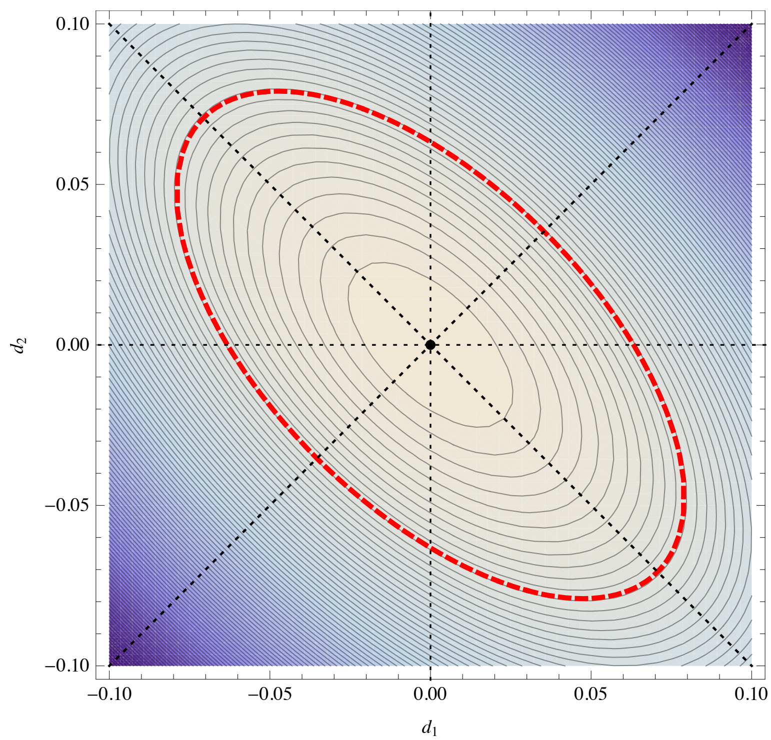

Figure 3.2: Iso-attractiveness curves

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