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Original Article

Does Online Dating Challenge Gendered Divisions of Household Labor?

Gina Potarca[1] and Jennifer Hook[2]

[1]University of Liverpool
[2]University of Southern California

Abstract

Online dating has modified how people find and select partners. In addition to outcomes already observed (e.g., exogamy), we argue that by subverting normative dating scripts, online courtship practices may set the course for partnerships that display more egalitarian divisions of routine household labor. This may be particularly true for the married and for lower-educated women, who generally report the least egalitarian allocation of domestic work. Furthermore, we posit that the relationship between meeting context and household labor will be partially explained by the selectivity of those who search for partners online but also by mechanisms specific to online dating that allow for greater relationship quality. We use 2008–2019 German Family Panel (*pairfam*) data in random-effects regression models to predict sharing of routine housework among women in marital and cohabiting opposite-sex unions ($N = 3305$). We find that meeting online is associated with greater sharing of housework for married women with lower-education, and that the link is robust even after accounting for observed selection into online dating via entropy balancing weights. Contrary to expectations, partnership quality has no mediating effect. Much of the positive association remains unexplained, suggesting that the different ways men and women negotiate power in the dating phase in digital versus non-digital partner markets may indeed play a role in how gender is enacted later on.

Introduction

Dating online through websites, phone apps, or social networks is now common and socially endorsed (Smith 2016). In certain contexts (e.g., United States), it has even become the number one way of meeting partners (Rosenfeld, Thomas, and Hausen 2019). Social scientists have already documented several social transformations brought about by online dating. In the United States, for instance, internet-matched heterosexual couples are less socially homogamous (Thomas 2020) and transition faster into marriage than couples who met face-to-face (Rosenfeld 2017). In addition to these effects, scholars also expected online dating to transform gendered relationship practices (Hardey 2002). Despite advances in women's education, labor force participation, and earning power (England, Levine, and Mishel 2020), and nearly universal preferences for equity in the allocation of housework (Auspurg, Iacovou, and Nicoletti 2017), women still do more housework than men (Geist and Ruppanner 2018). Women's continued responsibility for unpaid labor is at the heart of gender inequality in rich countries, limiting women's employment and

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career progression with immediate and long-term consequences for their economic security and well-being (Coltrane 2000; Goldscheider, Bernhardt, and Lappegård 2015).

Research has yet to scrutinize the link between how people select partners and later sharing of domestic tasks within marriage or cohabitation. Through its unique features (e.g., wider dating pool, anonymity, refined searches, and compatibility matching), and a more liberal selection space, online dating might allow individuals, especially women, to better negotiate dating scripts, redefine relationship expectations, and ultimately create more egalitarian partnerships. Specific ways digital dating may help women deconstruct gender are by increasing access to egalitarian men and by affording greater bargaining power in a partner market with an oversupply of men. Even though gendered courtship patterns, such as men often making the first move, get replicated online (Dinh et al. 2022; Huber and Malhotra 2017), digital spaces of interaction also provide women the freedom to more assertively communicate with potential partners than in real life (Dwyer, Hookway, and Robards 2021). As the enactment of gender in the initial stages of dating echoes later in committed relationships (Lamont 2020), more equal interactions early on could reverberate into how evenly partners share responsibilities down the line. Moreover, online dating selectivity (e.g., on gender role attitudes, psychological or demographic profile) and partnership quality may also contribute to the link between meeting online and how partners divide household labor. By providing increased information on potential candidates than usually available in face-to-face meeting contexts, the internet could assist women in selecting partners on non-economic considerations, such as communication skills or ability to provide emotional support (Lawson and Leck 2006). Since high levels of emotional intimacy and communication quality predict equal sharing of domestic work (Carlson, Miller, and Rudd 2020; Komter 1989), online daters would eventually craft more egalitarian relationships than those selecting partners in conventional ways.

The association between digital mate selection and the division of household labor, however, likely varies by union type and across social groups (here defined by education). Given already less gendered practices of dividing domestic work within cohabitation than within marriage (Domínguez-Folgueras 2013), the online premium may be particularly visible among the married. Likewise, though highly educated women may be particularly skilled at capitalizing on digital spaces of mate selection (Potarca 2021) to initiate marriage-conducive partnerships, online dating may generate a more visible improvement for women with lower education, who typically experience a more traditional allocation of housework (Hook 2015). Furthermore, the gendered division of routine housework is resilient, especially in gender-conservative contexts (Hook 2017). Routine housework is one of the least enjoyed activities for both men and women and is more resistant to gender equity than other types of domestic labor, such as childcare (Sullivan 2013). Online dating may therefore assist certain groups (e.g., lower-educated married women) in catching up with the already more egalitarian (e.g., cohabiters and the highly educated), but it would not affect the latter, who may use the same strategies both offline and online. Any benefits of digital dating would thus not cumulate into additional advantage.

To test these assumptions, we use 2008–2019 panel data on cohabiting and married women from Germany and inspect the link between meeting context and the division of routine housework.[1] The analysis centers on women since men tend to be less accurate in reporting time spent in housework (Kan 2008) and because partnership quality predicts the division of housework for women, whereas for men, the direction of causality is reversed (Carlson et al. 2020). Germany represents a suitable context to examine whether meeting online disrupts the traditional allocation of housework as the male breadwinner and one-and-a-half earner models still persist (Hook 2015), and gender gaps in the division of domestic labor still prevail (Procher, Ritter, and Vance 2018).

This study provides the first quantitative and nationally representative evidence of how online dating transforms gendered practices of housework in cohabiting and marital unions. With the backdrop of a stalled gender revolution (England 2010; Goldscheider et al. 2015), it is important to inquire whether the internet as alternative partner market assists adults in deconstructing traditional partnering in general and gender-asymmetric scripts of dividing unpaid work within

committed unions in particular. In other words, it is important to identify whether advances in the digital revolution can serve as tools in advancing the gender revolution. Below, we outline the case for internet dating, mindful of the null hypothesis of no online effect given the "stickiness" of the gendered division of routine housework.

Background and Theory

How the Internet Promised to Challenge Gender

Initial claims about the transformative power of the internet envisioned cyberspace as an arena with great potential to subvert normative constructions of gender and intimate life. Given the disembodied nature of communication, individuals would be free to create less socially constrained identities and establish contacts on the basis of rapport and mutuality of interests (Cooper and Sportolari 1997). An unprecedentedly wide choice set and the ability to selectively choose whom to initiate or pursue interactions with was thought to revolutionize the way people select a partner (Thompson 2018).

Online dating certainly did not come without problematic aspects. It promoted an inherent culture of disposability (Newett, Churchill, and Robards 2018), reinvented new forms of gendered racism (Curington, Lundquist, and Lin 2021), and proliferated body surveillance (Strubel and Petrie 2017) and sexual harassment (Cooper and Sportolari 1997). Beyond these risks, however, the internet does afford a larger set of alternatives and diversity of options (Thomas 2020), which allow access to candidates that singles would not typically encounter in day-to-day interactions. Especially for women, extending the structural bounds of their partner market increases the chances of finding more egalitarian-minded men.

An additional structural advantage for women dating online is the greater bargaining power afforded by unequal sex ratios. Dating platforms and apps typically accommodate a greater number of men than women (Potarca 2020; Skopek, Schulz, and Blossfeld 2011). According to Guttentag and Secord (1983), unequal sex ratios impact the distribution of power between women and men. An undersupply of women in online spaces tilts the gender power balance in their favor, affording an advantage in how relationships are negotiated both in the beginning and later in marriage. Based on these arguments and empirical evidence of the link between imbalanced sex ratios and intra-household decisions (Ogasawara and Komura 2022; Stauder and Röhlke 2022), we contend that a more favorable partner market position helps women better assert preferences for gender equity and thus establish more egalitarian relationships in online than in conventional settings.

Furthermore, in digital spaces of interaction, increased anonymity, lower costs of rejection, and the freedom to create less socially constrained identities (Cooper and Sportolari 1997) may prompt a greater willingness to take risks and purposefully deviate from conventional dating (Chen and Liu 2019). Gendered courtship typically assigns men and women to specialized roles (e.g., men as providers and pursuers and woman as passive, non-assertive communicators), paving the way for male-dominated relationships (Lamont 2017). Certain traditional courtship tropes are reinscribed in how men and women interact in digital settings, such as men more often initiating conversations (Dinh et al. 2022; Huber and Malhotra 2017) or women using digital expression tools (e.g., emojis) meant to emphasize femininity (Pinsky 2019). Nevertheless, there is also evidence that online, women adopt more open sexual scripts and more assertively claim power in interactions with men by, for instance, being more forthcoming in communication (Cooper and Sportolari 1997; Lawson and Leck 2006) or more overtly expressing relationship needs (Dwyer et al. 2021). In qualitative studies of digital dating, women attest to an increased sense of agency and empowerment (Chen and Liu 2019), and the freedom to more proactively communicate with men than face-to-face (Dwyer et al. 2021). Research shows that individuals who purposefully endorse reciprocity (e.g., both partners actively advance the relationship) and an overt negotiation of needs early on (Lamont 2017) display more egalitarian relationships in the long term (Lamont 2020).

Based on this evidence, we propose that by providing women greater bargaining power, easier access to diverse options, and opportunities to engage in more open and assertive conversations with men, digital dating might encourage romantic partnerships that deviate from typical gendered constraints and later on display more egalitarian relationship practices[2] (*hypothesis 1*). Since cohabiters are commonly more progressive in their division of housework (e.g., Domínguez–Folgueras 2013), online effects may only apply to marriage (*hypothesis 2*). Beyond selection on gender role attitudes or age (which the analysis accounts for, as described later in the text), couples that eventually marry may benefit more from setting up a gender-disruptive relationship culture early on, which could help withstand the inertia of slipping into traditional roles later in marriage.

The survey data at hand do not allow us to examine the "black box" of these early-stage interactions but do allow us to directly test two mechanisms articulating the link between online matching and the division of housework. These include selection into online dating and relationship quality.

Selection into Online Dating

Several selection processes may confound the relationship between partners' equal participation in domestic work and online dating. First, individuals who generally have a more open personality (Timmermans and De Caluwé 2017) and attach greater value to egalitarian work–family arrangements may specifically choose internet dating to more easily search for partners willing to share domestic work. Qualitative evidence shows that women using online dating reject the idea of recreating a traditionally gendered union and instead seek romantic connections built on companionship and principles of equality and independence (Dwyer et al. 2021). Given the direct link between women's egalitarian gender attitudes and their partners' greater involvement in housework (Greenstein 1996), couples who met through online dating might exhibit more equal practices of domestic work than others by virtue of having a different gender ideology to begin with.

Second, online daters' specific life-course stage (Presser 1994) might play an important role in the link between meeting online and less gendered patterns of housework. Men and women who use online dating often have previous marital experience, but they are less likely to have resident children (Robnett and Feliciano 2011). Individuals who experience divorce and remarry tend to have more egalitarian practices (Kleider 2015). Furthermore, women without children spend less time doing housework than mothers (Yavorsky, Dush, and Schoppe-Sullivan 2015), and many couples transition from an egalitarian to a traditional gender regime after becoming parents (Katz-Wise, Priess, and Hyde 2010). Thus, *online daters may show a more egalitarian division of housework given an inherently more open personality, more progressive gender attitudes, or a greater likelihood of having been previously married and having no resident children (hypothesis 3)*.

Partnership Quality—Emotional Support and Partner Communication

Beyond selection effects, an important way in which internet dating may act as gender-emancipatory technology (Lopes and Vogel 2017) is that digital modes of partner selection, particularly dating platforms with built-in compatibility-based matching algorithms (Finkel et al. 2012), may facilitate a greater number of partnerships built on emotional connection and open communication. An early study found that couples who met online in the United States had slightly greater marital satisfaction than couples formed elsewhere (Cacioppo et al. 2013). Another study revealed meeting through dating websites (but not through phone apps or online social networks) in Switzerland to be positively associated with relationship satisfaction (Potarca 2020). In the absence of face-to-face interaction, attraction is built through non-physical attributes, such as shared interests (usually explicitly stated on dating profiles), affinity, and mutual self-disclosure, potentially leading to connections based on greater relationship satisfaction and emotional intimacy (Lawson and Leck 2006).

According to the gender power perspective (Carlson et al. 2020; Komter 1989), open communication between partners lays the ground for a more equal sharing of domestic tasks. Especially for women, the ability to clearly communicate, negotiate, and assert preferences (i.e., manifest power) may assist in enacting more egalitarian arrangements (Sassler and Miller 2017). Women who met their partner online would therefore have a more egalitarian division of housework than those who met their partner offline because of *potentially better communication quality and greater emotional support from their partner* (*hypothesis 4*).

Educational Differences

So far, we have formulated our expectations in broad terms, but whether internet dating disrupts gendered practices of housework is likely to be uneven across social groups, insofar as the perpetuation of doing gender in marriage intersects with hierarchies of class (Hook 2015). Here, we define social groups by educational attainment, recognized as a reliable indicator of both economic and cultural capital, and one of the most important commodities on the marriage market (Becker 1993; Oppenheimer 1988).

There are two reasons why *the anticipated positive effect of meeting online on the egalitarian division of housework could mainly apply to highly educated women (hypothesis 5a)*. First, class differences in digital skills (Ollier-Malaterre, Jacobs, and Rothbard 2019) might mean that, compared with the less educated, university-educated female daters are more adept at navigating partner selection online and thus make better use of its advantages. Well-educated women in Germany, for instance, more likely convert online dating into marriage, partly because of better matching on attitudes toward marriage and maternal employment (Potarca 2021). Second, in line with the relative resource perspective (Blood and Wolfe 1960), which claims that the partner with the most earning power uses their advantage to bargain reduced contributions to domestic work, women with more economic resources might dare to disrupt gendered conventions when dating online to a higher extent than lower-educated women. Empirical evidence shows that well-educated women are more assertive in voicing preferences regarding the allocation of housework, whereas lower-educated women are more often ignored by their partners (Miller and Carlson 2016).

Nevertheless, precisely because they are commonly more disadvantaged, lower-educated women with a more progressive mindset may resort to digital dating tools to seek more egalitarian partners. Even beyond potential selectivity on openness and gender ideology, the structural advantages of online partner markets (i.e., an oversupply of men and a more diverse dating pool) would ensure that less educated women encounter more opportunities to establish unions in which male partners contribute more to domestic tasks and experience a more substantive improvement in partnership quality than the already advantaged highly educated. Even though online dating may not significantly increase the odds of marriage for lower-educated women (Potarca 2021), it may ultimately improve relationships by allowing for more freedom to subvert gender hierarchy during courtship and then later in unions. As earlier stated, online dating may not afford a significant premium to the already more egalitarian highly educated, who may perform gender bargaining in a similar way online as offline. It is thus possible that *less (and not highly) educated women experience an increase in the sharing of housework if they met their partner online versus offline* (hypothesis 5b).

Data and Method

We use the German Family Panel (*pairfam*) 2008–2019, release 11.0 (Brüderl et al. 2020), a longitudinal survey dataset containing detailed yearly information on individuals' sociodemographic profile, preferences and values, and the context in which they met their partner (if in a union). *pairfam* contains information on the partnership trajectories of an initial sample of 12,402 randomly selected men and women who are nationally representative of cohorts born in 1971–73, 1981–83, and 1991–93. *DemoDiff* consists of an oversample of 1489 Eastern German respondents born in the years 1971–73 and 1981–83 (Kreyenfeld et al. 2012). *DemoDiff* was initiated

in parallel with *pairfam*'s Wave 2 and was separately conducted for three waves until Wave 5, when it was fully integrated into *pairfam*. In Wave 11, a refreshment sample was added, including approximately 5000 respondents from birth cohorts 1981–83 and 1991–93, as well as a new, younger cohort (2001–03). The total sample included 18,912 respondents. A detailed description of the study and its cohort-stratified random sample can be found in Huinink et al. (2011). The data ideally suit the objectives of this study, as they record information on young and established adults, typically more technologically savvy than their older counterparts (van Deursen and Helsper 2015).

For the current study, we selected cohabiting or married female respondents. To avoid considering cohabitations that eventually transition to marriage as two separate unions, we removed observations of premarital cohabitation. This also ensured getting close to comparing marital couples to those who choose long-term cohabitation as alternative to marriage. Additional analysis based on data where we also removed cohabiters with strong intentions to marry (and thus eliminated would-be premarital cohabitation spells) revealed similar results (available upon request). After removing male respondents ($n = 9161$), participants younger than 18 years ($n = 1615$), women who are single ($n = 1185$), in same-sex ($n = 104$) or non-residential unions ($n = 1108$), in relationships formed before the age of 18 ($n = 1091$) or before 1995 ($n = 577$), and cases with missing or irrelevant information on key variables ($n = 766$), our final sample consisted of 13,330 observations for 3305 women aged 18–48, in cohabiting ($n = 1378$) or marital couples ($n = 2016$). Eighty-nine cases reported both type of unions.

Measures

The outcome of interest refers to the intra-household division of routine housework (washing, cooking, and cleaning). The item is derived from the Relative Share of Labor Scale (Baxter 2002) measuring each partner's contribution to core household tasks. Participants answered the following question: "To what extent do you and [name of current partner] share duties in the following domains?" Responses range from (1) "(almost) completely my partner," (2) "for the most part, my partner," (3) "split about 50/50," (4) "for the most part, me," (5) "(almost) completely me," (6) "another person," to (7) "doesn't apply to our situation." The last two options contain few cases ($n = 23$ and 32 observations, respectively) and are discarded. The variable is reverse-coded so that higher values indicate greater contributions by the male partner. We treated the variable as a linear measure of sharing, as seen in previous studies (e.g., Nitsche and Grunow 2016), but also considered it as ordinal in additional random-effects ordered logit models, with similar results (reported in Supplementary Table D1 in the supplementary file). Histograms of the outcome variable by educational level are provided in Figure E1 in the supplementary file. Although "for the most part, me" is the modal response for both tertiary and non-tertiary-educated women, "split about 50/50" is the second most common response for tertiary-educated women, whereas "(almost) completely me" is the second most common response for non-tertiary-educated women.

To capture *meeting setting*, we used information on how respondents met their partners. The measure allows for a single answer from the following options: (1) "school, training, work," (2) "hobby, club, sports," (3) "bar, night club," (4) "friends or acquaintances," (5) "relatives," (6) "a personal ad," (7) "the internet," (8) "vacation," (9) and "others." All non-digital settings are grouped under a single (0) offline category. *pairfam* began to distinguish between two online settings (e.g., internet partner-finding service versus online social networks, chat rooms, etc.) only in Wave 4; starting with Wave 10, it added the possibility of couples having met through dating apps. To maximize the data at hand (i.e., only seven married women met their spouse on dating apps in Wave 11), and in line with earlier theoretical arguments, we used a broad online category for all 11 waves.

Educational level was measured on the following two-category scheme: (0) non-tertiary, which includes persons with no degree, first- and second-stage basic education, secondary-education, and postsecondary training that prepares for labor market entry and/or tertiary education, and (1) tertiary education, which includes a bachelor's degree and/or postgraduate studies. The measure

only distinguishes between respondents with high-level education and those without, given a small subsample of respondents with the lowest level (i.e., basic) education in unions initiated online (33 women).

To gather information on the quality of the relationship, pairfam adapted the four scales of *intimacy*, *admiration*, *dominance*, and *conflict* from the Network of Relationships Inventory (NRI) (Furman and Buhrmester 1985) and used them in an abbreviated form (i.e., shortened from three to two items each). The response format has five stages that reflect frequency: (1) "never" to (5) "always." For *intimacy*, we used the following items: "How often do you tell [name of current partner] what you're thinking?" and "How often do you share your secrets and private feelings with [partner]?" For *admiration* (i.e., esteem support): "How often does [partner] express recognition for what you've done?" and "How often does [partner] show that he/she appreciates you?" For partner's *dominance*: "How often does [partner] get his/her way when you can't agree on something?" and "How often does [partner] make you do things his/her way?' For *conflict*: "How often do you and [partner] disagree and quarrel?" and "How often are you and [partner] annoyed or angry with each other?" To create the four scales of partnership quality, we constructed four mean scores based on corresponding items. Cronbach's alpha was above 0.700 for each scale.

To account for the potential selectivity of online daters and balance the covariate distribution of our treatment and control groups (as described in the next section), we introduced measures of *gender role attitudes*, *personality, previous marital experience, and presence of children at* t_1 *(i.e., relationship onset)*. For *gender role attitudes*, similar to Nitsche and Grunow (2016), we used the first-recorded measurement of attitudes to minimize endogeneity and convergence in attitudes over time. Attitudes were measured every other year starting with Wave 1. When missing, we used information from the nearest wave recorded. Since most respondents were already partnered when first asked about attitudes, we only observed them pre-partnering in 21.6% of cases. We specifically looked at attitudes regarding men's involvement in the domestic sphere, which asked respondents for agreement on whether "Men should participate in housework to the same extent as women" on a five-point scale ranging from (1) "disagree completely" to (5) "agree completely." The five constructs of *personality* (neuroticism, extraversion, agreeableness, conscientiousness, and openness) were measured on a validated 21-item version of the Big Five Inventory (Rammstedt and John 2005). For each item, respondents rated their agreement using a five-point Likert-type scale ranging from (1) "absolutely incorrect" to (5) "absolutely correct." Personality was measured three times during the panel (i.e., in Waves 2, 6, and 10). For respondents who remained in the panel long enough to be surveyed more than once, we only considered the first measurement. We identified *previous marital experience* based on the number of marital unions respondents reported, and *presence of children* at t_1 based on information on children's date of birth and year the relationship started.

Selectivity-assessing analyses also included the following covariates likely associated with choice of online dating: respondent's age in years at t_1, relationship duration (in years), an indicator of whether the relationship started in the post-dating app period (i.e., after 2012), residence in former East Germany, migration background, and tertiary education at t_1. With few exceptions, most covariates are time-constant or pertaining to the pre-partnering stage.

In addition to moderating and mediating factors, the main analysis also included control measures that commonly impact the division of housework. Past research revealed that meeting online was associated with greater exogamy on education, race (Thomas 2020), or national origin (Potarca 2017), as well as greater homogamy in terms of age (Potarca 2020; Thomas 2020). Particularities in the sociodemographic composition of couples formed online may influence the economic dependence and power dynamics between partners and ultimately how egalitarian their division of housework is. For instance, educationally heterogamous couples may have a less egalitarian division of housework than homogamous couples, generally better at reaching agreement regarding the allocation of time and resources (Pesando 2021). Similarly, partners that differ in terms of national origin might experience a less egalitarian division of housework given lower levels of cultural cohesion (Dribe and Lundh 2011).

Based on both partners' education, we constructed an *educational matching* variable, distinguishing between homogamy (partners share the same level of education), hypergamy (the husband is better educated), and hypogamy (the wife is better educated). *Matching on origin* is coded based on partners' migration status, which differentiates between (0) native Germans and (1) respondents with a migration background, comprising all persons of non-German origin (e.g., ethnic Germans, half-Germans, Turks, and others). We distinguish between (1) homogamous couples where both partners are native Germans, (2) homogamous couples where both have a migration background, and (3) heterogamous couples. For sorting on age, similar to other studies (Dribe and Nystedt 2017), we distinguished between (1) partners close in age (the male partner is 0–2 years older than the female partner), (2) age hypergamy (the man is at least 3 years older), and (3) age hypogamy (the woman is at least 3 years older).

Along with accounting for couple composition, our analyses included the following controls: presence and age of resident children at survey year, distinguishing between respondents with no children, those for whom the youngest resident child is 2 years old or younger, and those for whom the youngest child is older than 3; and the division of paid work, based on both partners' employment status, and differentiating between (1) male breadwinner households (the male partner is full-time employed whereas the female partner is not working), (2) 1.5 breadwinner model (the man works full-time, the woman part-time), (3) dual earner (both work full-time), and (4) other arrangements.

Method

Entropy Balance Weighting

First, as previously mentioned, women who met their partner online versus offline may differ in ways that are systematically related to both meeting context and how housework is shared. To explore and account for such selection effects, we use entropy balancing (EP) (Hainmueller 2012), a multivariate matching technique that reweights the control group to match the covariate distribution of the treatment group, and in so doing closely emulating the statistical equivalence of random assignment. As opposed to reweighting based on propensity scores, which requires a time-consuming back-and-forth between selecting covariates, weighting, and balance checking, EB directly estimates weights that satisfy imposed balance constraints. Especially when dealing with modest-size samples, another advantage of EB is that it avoids trimming outliers, such as long-duration partnerships, overrepresented among offline-formed couples. Using the *ebalance* command (Hainmueller and Xu 2013) in Stata 17 (StataCorp 2021), we generate weights with the imposed condition that the mean and the variance of all variables become similar across the two groups. When computing weights, we additionally include interaction terms between each covariate and type of union to balance the distribution of variables across cohabiting and married subgroups, which also tackles the issue of observed selection into marriage. Indeed, after using weights in our main analysis, cohabiting effects were also reduced.

To explore differences in observed characteristics between women in offline- vs. online-formed unions, Figure 1 plots standardized differences in means for each covariate and interaction term. Full information on means and variances for each group is provided in Table A1 in the supplementary file. Focusing on unadjusted values, particularly those greater than ±0.2, we see that observations for women who met their partner online (hereafter, for brevity: online observations) were more often associated with cohabitation, shorter partnership duration, older age at t_1, previous marital experience, and having children pre-partnering[3]. Online observations were additionally linked to more progressive gender attitudes and certain personality traits (e.g., less extraversion and more openness), but differences compared with offline observations were not large. Figure 1 also shows that post-weighting differences shrank to zero, indicating that covariate balance was achieved.

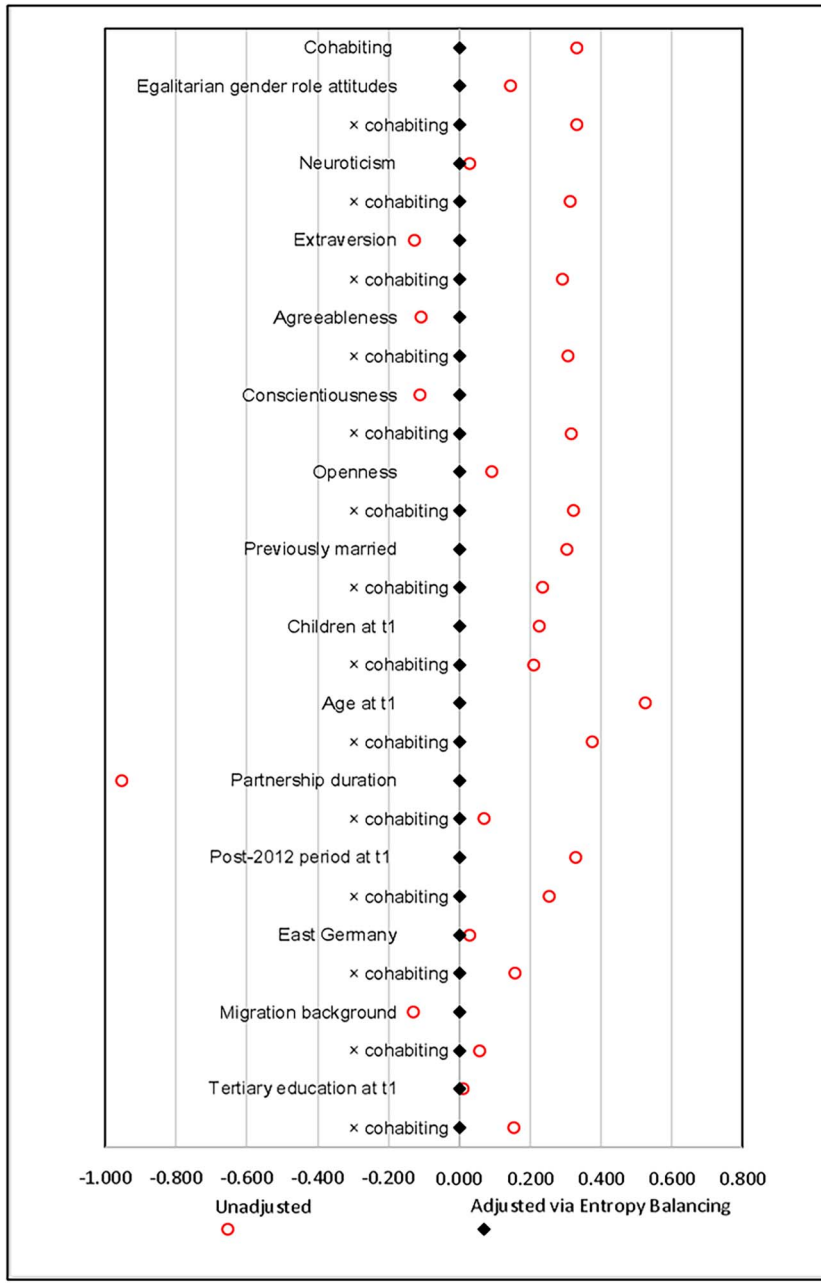

Figure 1. Standardized differences in means between the treated and control groups for selected covariates, before and after entropy balancing. Note: Positive (negative) values indicate that the mean for online-formed couples is higher (lower) than for offline-formed couples.

Random-Effects Regression

Using Stata's *mixed* command with the *pweight* specification to implement EB weights, we then conducted multilevel (i.e., random-effects) linear regression models to examine the extent to

which online-formed couples share routine housework. Besides the advantage of accounting for the clustering of responses within individuals, random-effects models use the optimal combination of between- and within-individual variability in the data, estimating a random intercept capturing unobserved effects for each respondent.

The specification of a first model estimating the main effect of meeting online took the following mathematical form:

$$y_{ij} = \beta_0 + \beta_1 \textbf{\textit{meeting online}}_{ij} + \beta_2 \boldsymbol{X}_{ij} + \beta_3 \boldsymbol{Z}_i + u_j + \varepsilon_{ij}, \quad (1)$$

where y_{ij} is the sharing of routine housework for respondent i in wave j, modeled as a linear function of meeting context. β_0 is the intercept, whereas β_1 represents the difference in sharing for women who met their partner online versus those who met theirs offline, adjusted for time-varying (vector $\boldsymbol{X}_{ij}$ with coefficient vector $\boldsymbol{\beta}_2$) and time-constant controls or mediating factors (vector $\boldsymbol{Z}_i$ with coefficient vector $\boldsymbol{\beta}_3$). The specification also includes between-individual (u_j) and within-individual (ε_{ij}) error terms. A second model, which includes an interaction between meeting online and cohabiting (versus being married), takes the following form:

$$y_{ij} = \beta_0 + \beta_1 \textbf{\textit{meeting online}}_{ij} + \beta_2 \boldsymbol{X}_{ij} + \beta_3 \boldsymbol{Z}_i + \beta_4 \textbf{\textit{cohabiting}}_{ij} + \beta_5 \textbf{\textit{meeting online}}_{ij} \times \textbf{\textit{cohabiting}}_{ij} + u_j + \varepsilon_{ij}, \quad (2)$$

where β_1 then indicates the difference in sharing for married women who met their partner online versus those who met theirs offline.

We fit models including all women and then for each educational group separately, given that the online dating effect and other previously suggested mechanisms may uniquely affect the housework supply of women with different education. We also considered fixed-effects models, which would have had the advantage of accounting for potential unobserved fixed traits, but we did not have sufficient variation in our main independent variable, with only 59 within-subject changes in treatment and only 20 of which involving marriage.

Results

Descriptive Results

Table 1 provides EB-weighted descriptive statistics by meeting setting for the general sample and for each educational group separately. Regardless of educational attainment and the method of meeting, women report a share of routine housework that is closer to "for the most part, me" (value = 2) than to "split about 50/50" (value = 3). We notice, however, that women with non-tertiary education who met their partner online report significantly more sharing of housework than those who met their partner offline. The latter report the lowest level of sharing.

Contrary to expectations, none of the partnership quality measures significantly varied across meeting context. The advantage that online-formed couples had in this respect (as additional unweighted analyses revealed) was accounted for when controlling for online dating selectivity via weights. Nevertheless, additional tests looking at married unions only showed that lower-educated women who met their spouse online did report significantly lower levels of partner dominance than those who met theirs offline (see Table B1 in the supplementary file). As anticipated, the data revealed fewer, albeit non-significant, homogamous unions among couples formed online versus offline but no significant difference in sorting on cultural origin across meeting context. Though similarity in partners' age was greater among online- than offline-formed couples, the difference was again non-significant.

Multivariate Results—Pooled Sample

We now turn to multivariate results from random-effects regression models of sharing routine housework (Table 2). We hypothesized that women who met their partner online would report

Table 1. Weighted Sample Descriptive Statistics, by Education and Meeting Context; *pairfam* Waves 1–11 (2008–2019)

		Pooled sample		Non-tertiary		Tertiary	
		Offline % Mean (SD)	Online % Mean (SD)	Offline % Mean (SD)	Online % Mean (SD)	Offline % Mean (SD)	Online % Mean (SD)
Sharing of routine housework		2.20 (0.86)	2.27 (0.89)	2.13 (0.86)	2.25 (0.90)	2.32† (0.84)	2.31 (0.88)
Emotional intimacy (range: 1–5)		3.88 (0.77)	3.85 (0.78)	3.83 (0.81)	3.84 (0.80)	3.97 (0.67)	3.88 (0.74)
Esteem support (range: 1–5)		3.74 (0.79)	3.70 (0.82)	3.70 (0.83)	3.70 (0.86)	3.81 (0.70)	3.70 (0.76)
Dominance of partner (range: 1–5)		2.89 (0.63)	2.87 (0.58)	2.88 (0.67)	2.84 (0.62)	2.91 (0.55)	2.91 (0.53)
Conflict (range: 1–5)		2.58 (0.70)	2.62 (0.71)	2.59 (0.73)	2.61 (0.74)	2.57 (0.65)	2.62 (0.66)
Educational matching	Homogamy	66.3	65.6	67.7	68.5	63.7	60.7
	Hypergamy	16.6	16.8	25.7	27.1	—	—
	Hypogamy	17.1	17.6	0.7	0.4	36.3	39.3
Origin exogamy		11.4	9.9	9.8	7.8	14.4	13.2
Age matching	Similar age	36.8	42.0	35.7	39.8	38.9	45.6
	Man older >2 years	50.3	47.8	50.0	47.5	50.8	48.4
	Woman older >2 years	12.9	10.2	14.3	12.7	10.3	6.1
Age of resident children	No children	37.9	39.9	33.0	35.9	46.8	46.6
	Youngest ≤3	26.7	27.0	25.4	24.5	29.1	31.0
	Youngest >3	35.4	33.1	41.6	39.7	24.1	22.3
Tertiary education		35.4	37.9	—	—	—	—
Division of paid work	Male breadwinner	19.2	20.5	20.2	22.0	17.3	18.2
	1.5 breadwinner	22.3	24.5	24.4	27.1	18.6	20.2
	Dual earner	26.6	24.4	26.4	20.2	27.1†	31.2
	Other	31.9	30.6	29.0	30.7	37.0	30.4
N (observations)		12113	1217	7399	756	4714	461
N (individuals)		2973	385	1789	245	1184	140

Note: †$p < 0.10$.

greater sharing (hypothesis 1) and that marital status would moderate this effect such that differences between online and offline couples would be visible when women are married rather than cohabiting (hypothesis 2). We further suggested that online dating selectivity (hypothesis 3) and greater partnership quality (hypothesis 4) may partly explain the association between meeting online and the way couples share housework.

Model 1 estimates the main effect of meeting online on the unweighted data while adjusting for several control variables. We refer to positive coefficients as indicating greater sharing and negative coefficients as lesser sharing; although higher values indicate greater contributions by the male partner, means and predicted values are uniformly lower than a 50/50 split, indicating that coefficients illustrate movements toward and away from equal sharing. Results indicate that women in couples that formed online reported greater sharing of routine housework than women in couples that met offline ($p < .05$), confirming hypothesis 1. Overall, online dating generated a modest shift of 0.109 on a five-point scale toward sharing. This is approximately 13% of a standard deviation and similar in size to the effect for tertiary education, long associated with more equal

Table 2. Results from Random-Effects Regression Models Predicting Sharing of Routine Housework among Partnered Women ($N = 3305$); *pairfam* Waves 1–11 (2008–2019)

	Unadjusted sample		Adjusted sample	
	Model 1	**Model 2**	**Model 3**	**Model 4**
Meeting online	0.109**	0.162**	0.144*	0.142*
	(0.035)	(0.052)	(0.062)	(0.060)
Cohabiting	—	0.156***	0.136***	0.144***
	—	(0.026)	(0.038)	(0.037)
Meeting online × cohabiting	—	−0.137†	−0.114	−0.111
	—	(0.070)	(0.083)	(0.082)
Emotional intimacy	—	—	—	0.005
	—	—	—	(0.027)
Esteem support	—	—	—	0.131***
	—	—	—	(0.024)
Dominance of partner	—	—	—	−0.061*
	—	—	—	(0.027)
Conflict	—	—	—	0.013
	—	—	—	(0.026)
Educational matching (ref.: homogamy)				
Hypergamy	−0.015	−0.002	−0.022	−0.024
	(0.032)	(0.032)	(0.051)	(0.051)
Hypogamy	−0.041	−0.051	−0.109*	−0.091†
	(0.031)	(0.031)	(0.050)	(0.049)
Origin exogamy	−0.060†	−0.048	−0.021	−0.012
	(0.035)	(0.035)	(0.061)	(0.059)
Age matching (ref.: similar age)				
Man older >2 years	0.019	0.022	0.031	0.030
	(0.022)	(0.022)	(0.037)	(0.037)
Woman older >2 years	0.054	0.044	0.052	0.047
	(0.041)	(0.041)	(0.067)	(0.068)
Age of resident children (ref.: no children)				
Youngest ≤3	−0.287***	−0.244***	−0.274***	−0.235***
	(0.023)	(0.024)	(0.045)	(0.046)
Youngest >3	−0.340***	−0.294***	−0.260***	−0.224***
	(0.023)	(0.024)	(0.046)	(0.046)
Tertiary education	0.115***	0.124***	0.097*	0.086*
	(0.027)	(0.027)	(0.044)	(0.044)
Division of paid work (ref.: male breadwinner)				
1.5 breadwinner	0.115***	0.117***	0.077†	0.088*
	(0.020)	(0.020)	(0.043)	(0.043)
Dual earner	0.312***	0.309***	0.329***	0.327***
	(0.024)	(0.024)	(0.047)	(0.048)
Other	0.347***	0.338***	0.390***	0.395***
	(0.021)	(0.021)	(0.047)	(0.047)
Constant	2.103***	2.014***	2.036***	1.644***
	(0.031)	(0.034)	(0.063)	(0.184)
Random effects				
σ_u	0.302***	0.300***	0.095***	0.091***
	(0.011)	(0.011)	(0.016)	(0.016)
σ_e	0.361***	0.360***	0.570***	0.562***
	(0.005)	(0.005)	(0.022)	(0.022)
Log pseudolikelihood	−14,247.0	−14,229.2	−2937.1	−2914.8
N (observations)	13,330			
N (individuals)	3305			

Note: (Robust) standard errors are in parentheses. σ_u = standard deviation of u_j (between-person heterogeneity), σ_e = standard deviation of ε_{ij} (within-person heterogeneity). †$p < .10$; *$p < .05$; **$p < .01$; ***$p < .001$.

sharing of housework (Hook 2015). To determine whether the shift exclusively applied to married women, Model 2 includes an interaction of meeting context and the type of union. Results show that for women in marital couples, meeting online was linked to a 0.162 increase in sharing. As expected, cohabiting women who met their partner offline reported significantly higher levels of sharing than those in marital unions (i.e., 0.156, $p < .001$); in line with hypothesis 2, their online premium, however, was minimal and non-significant, as additional contrasts of marginal effects revealed ($0.162 - 0.137 = 0.025$, $p = .603$). Supplementary contrasts also showed that among couples who met online, the marriage disadvantage became small and non-significant (−0.019, $p = .774$).

Further, to assess the extent to which online effects among the married can be attributed to observed selection into digital dating, we replicated the estimation of the previous model on data adjusted with EB-derived weights. As mentioned earlier, the weighting balances the two comparison groups on selection variables, including covariates measured at relationship onset, leading to results unaffected by confounding due to (observed) selection bias. The analysis essentially estimates average treatment effects given that the study population is re-weighted to assess the online dating effect in the scenario that this way of meeting was valid for everyone within the population. Though the main effect remained statistically significant ($p < .05$), adjusting for online dating selectivity in the matched sample (Model 3) reduced the coefficient for meeting online by 11.1%. Selection playing a part, albeit not large, in explaining the online dating effect on the sharing of housework lends support to hypothesis 3. Post-weighting, meeting online was associated with a 0.144 significant increase ($p < .05$) in men's share of routine housework for married women and only a 0.030 non-significant premium ($p = .590$) for cohabiting women.

Model 4, which adds indicators of partner's emotional support and communication quality, was also fit on weighted data because we wanted to estimate the mediating effect of partnership quality independent of selection effects. We found that when relationship quality was accounted for, meeting online was still linked to a 0.142 significant premium ($p < .05$) of more equally sharing housework for married women. Given the very small reduction in the meeting online effect in Model 4, as well as additional mediation analyses based on generalized structural equation modeling (not shown) indicating no indirect effect of online dating on sharing via partnership quality, hypothesis 4 is refuted by the data. The main effects of partnership quality confirmed that women in unions rated higher in esteem support (admiration) reported sharing housework more equally, whereas women in unions rated higher in partner dominance shared less equally.

Multivariate Results—By Educational Level

To test whether the association between meeting context and the sharing of routine housework differs across educational groups, we replicated the analyses above by education. As reported in Table 3, Model 2 indicates that the meeting online effect is only statistically significant among married (and not cohabiting) women with non-tertiary education, confirming hypothesis 5b. Accounting for online dating selectivity (Model 3) reduced the online effect by 22.3%, but lower-educated women who met their spouse online still reported 0.195 more equal sharing of housework than those who met their spouse offline ($p < .05$). This is about 22% of a standard deviation and nearly as large as the effect sizes of residing with children. Among the highly educated, this advantage was much smaller and non-significant.

Results for the lower-educated further showed that adding indicators of partnership quality (Model 4) attenuated the online dating effect by 4.1%. In terms of the main effect of partnership quality, results indicated a significant and positive association between esteem support and sharing of housework for both educational groups, though additional mediation analyses (not shown) showed no indirect effect of meeting online on sharing via esteem support. Higher levels of partner dominance significantly predicted less equal sharing for the highly educated only.

Next, to directly test whether educational differences in online dating effects are statistically significant, we replicated the estimation of Model 2 on the pooled sample but added an interaction between meeting context, type of union, and education. Figure 2 graphs predicted means of

Table 3. Results from Random-Effects Regression Models Predicting Sharing of Routine Housework for Partnered Women with Non-Tertiary ($N = 2001$) and Tertiary ($N = 1304$) Education; *pairfam* Waves 1–11 (2008–2019)

	Non-tertiary			Tertiary		
	Unadj. sample	Adjusted sample		Unadj. sample	Adjusted sample	
	Model 2	Model 3	Model 4	Model 2	Model 3	Model 4
Meeting online	0.251***	0.195*	0.187*	0.045	0.072	0.085
	(0.069)	(0.077)	(0.075)	(0.078)	(0.103)	(0.099)
Cohabiting	0.151***	0.114*	0.123*	0.164***	0.170**	0.168**
	(0.034)	(0.048)	(0.048)	(0.040)	(0.062)	(0.060)
Meeting online × cohabiting	−0.192*	−0.119	−0.116	−0.086	−0.129	−0.136
	(0.092)	(0.104)	(0.103)	(0.110)	(0.134)	(0.132)
Emotional intimacy	—	—	0.021	—	—	−0.036
	—	—	(0.035)	—	—	(0.036)
Esteem supports	—	—	0.114***	—	—	0.160***
	—	—	(0.030)	—	—	(0.038)
Dominance of partner	—	—	−0.032	—	—	−0.132**
	—	—	(0.034)	—	—	(0.044)
Conflict	—	—	0.020	—	—	−0.004
	—	—	(0.034)	—	—	(0.039)
Educational matching (ref.: homogamy)						
Hypergamy	−0.003	−0.024	−0.026	—	—	—
	(0.034)	(0.052)	(0.051)	—	—	—
Hypogamy	−0.029	−0.080	−0.062	−0.059	−0.115*	−0.104†
	(0.058)	(0.088)	(0.089)	(0.036)	(0.058)	(0.058)
Origin exogamy	−0.100*	−0.133†	−0.116	0.015	0.119	0.123
	(0.049)	(0.080)	(0.078)	(0.051)	(0.093)	(0.090)
Age matching (ref.: similar age)						
Man older >2 years	0.024	0.053	0.045	0.021	−0.002	0.010
	(0.029)	(0.047)	(0.047)	(0.033)	(0.060)	(0.059)
Woman older >2 years	0.002	0.042	0.035	0.133†	0.116	0.113
	(0.051)	(0.081)	(0.082)	(0.069)	(0.120)	(0.118)
Age of resident children (ref.: no children)						
Youngest ≤3	−0.186***	−0.236***	−0.206***	−0.311***	−0.322***	−0.270***
	(0.033)	(0.060)	(0.059)	(0.036)	(0.070)	(0.074)
Youngest > 3	−0.257***	−0.214***	−0.184**	−0.336***	−0.343***	−0.305***
	(0.032)	(0.056)	(0.056)	(0.037)	(0.078)	(0.077)
Division of paid work (ref.: male breadwinner)						
1.5 breadwinner	0.118***	0.075	0.078	0.123***	0.085	0.112†
	(0.025)	(0.057)	(0.056)	(0.032)	(0.061)	(0.063)
Dual earner	0.316***	0.352***	0.340***	0.306***	0.307***	0.319***
	(0.031)	(0.060)	(0.061)	(0.037)	(0.076)	(0.078)
Other	0.331***	0.421***	0.423***	0.350***	0.343***	0.353***
	(0.027)	(0.059)	(0.058)	(0.032)	(0.078)	(0.076)
Constant	1.985***	1.985***	1.505***	2.159***	2.196***	2.074***
	(0.043)	(0.077)	(0.234)	(0.048)	(0.098)	(0.267)
Random effects						
σ_u	0.321***	0.087***	0.084***	0.265***	0.100***	0.091***
	(0.015)	(0.015)	(0.015)	(0.015)	(0.031)	(0.031)
σ_e	0.367***	0.599***	0.592***	0.349***	0.518***	0.508***
	(0.007)	(0.030)	(0.030)	(0.008)	(0.029)	(0.029)
Log pseudolikelihood	−8810.4	−1887.8	−1876.2	−5403.2	−1041.4	−1028.7
N (observations)	8155			5175		
N (individuals)	2001			1304		

Note: (Robust) standard errors are in parentheses. Unadj. = unadjusted. σ_u = standard deviation of u_j (between-person heterogeneity), σ_e = standard deviation of ε_{ij} (within-person heterogeneity). †$p < .10$; *$p < .05$; **$p < .01$; ***$p < .001$.

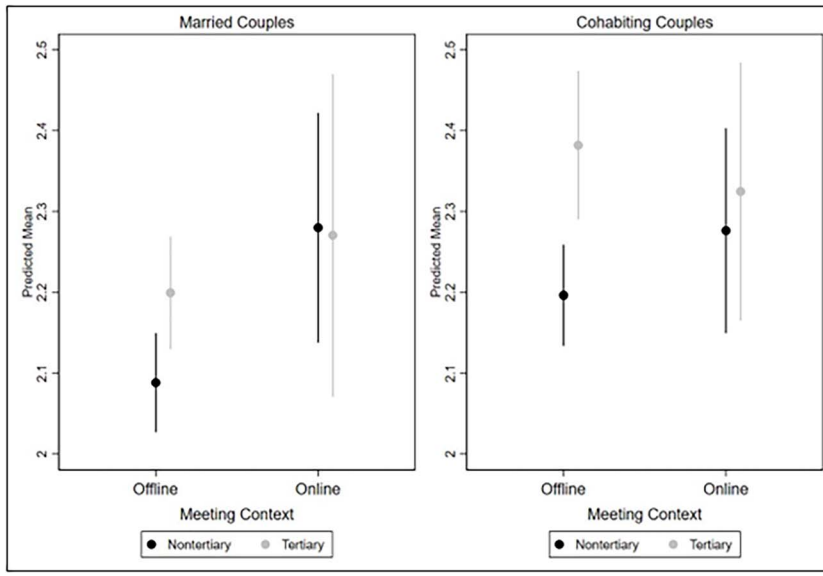

Figure 2. Predicted means of sharing routine housework, by union type, educational level and meeting context. Note: A value of 3 indicates equal sharing. A value of 2 indicates that women report "for the most part, me."

sharing routine housework based on this interaction (coefficients reported in Table C1 in the supplementary file) and shows that lower-educated women report significantly less sharing than the higher-educated when meeting their partner offline (the educational contrast is 0.111, $p < .05$ for married couples; and 0.186, $p < .01$ for cohabiting couples), but indistinguishable from tertiary-educated women when meeting online (the contrast value is −0.009, $p = .940$ for married couples; and 0.048, $p = .639$ for cohabiting couples).

Finally, to see if online-formed marital unions indeed start out with greater sharing, as we argued early on, we additionally investigated how practices of routine housework evolve across time, throughout the marriage trajectory. We distinguished between four stages in the trajectory: first year of marriage, second year, third to fifth year, and more than 5 years after marrying. The analysis included an interaction between stage in the marital trajectory and meeting online (coefficients included in Table 4). Predicted means graphed in Figure 3 (based on Model 3) showed no significant online dating (dis)advantage at either stage for high-educated women. For lower-educated women, however, results confirmed an online dating premium that is significant in the first (contrast of 0.339, $p < .05$), second (contrast of 0.237, $p < .05$), and third to fifth year (contrast of 0.161, $p < .10$) of marriage, and subsequently decreases, albeit not to zero, in later years. We also found that the inclusion of relationship quality measures in Model 3 increased the main effect of online dating by 2.9%, suggesting that men in unions initiated online contribute more to housework in the first year of marriage, despite lower levels of marital quality during that stage.

In summary, we find that meeting online is associated with greater sharing of routine housework. This association pertains to married women with non-tertiary education and is robust to accounting for online dating selectivity via EB weights. Relationship quality plays no role in offsetting the online dating effect for lower-educated women, with much of the positive association remaining unexplained.

Table 4. Results from Random Effects Regression Models Predicting Sharing of Routine Housework for Married Women with Non-tertiary ($N = 1219$) and Tertiary ($N = 779$) Education, by Stage in the Marital Trajectory; *pairfam* Waves 1–11 (2008–2019)

	Non-tertiary			**Tertiary**		
	Unadj. sample	**Adjusted sample**		**Unadj. sample**	**Adjusted sample**	
	Model 2	Model 3	Model 4	Model 2	Model 3	Model 4
Meeting online	0.372**	0.339*	0.349*	−0.041	−0.054	−0.037
	(0.122)	(0.136)	(0.137)	(0.127)	(0.153)	(0.153)
Marital trajectory (ref.: 1st year of marriage)						
2nd	0.038	0.073	0.111†	−0.086	−0.088	−0.083
	(0.051)	(0.060)	(0.058)	(0.060)	(0.077)	(0.078)
3rd–5th	0.051	0.075	0.119†	−0.148**	−0.102	−0.071
	(0.046)	(0.070)	(0.070)	(0.053)	(0.075)	(0.077)
6th+	−0.024	−0.015	0.049	−0.207***	−0.151	−0.096
	(0.047)	(0.068)	(0.069)	(0.056)	(0.096)	(0.097)
Meeting online × marital trajectory						
Meeting online × 2nd	−0.099	−0.102	−0.106	0.098	0.113	0.100
	(0.138)	(0.138)	(0.138)	(0.154)	(0.157)	(0.154)
Meeting online × 3rd–5th	−0.209†	−0.178	−0.209	0.012	0.056	0.079
	(0.126)	(0.142)	(0.138)	(0.135)	(0.169)	(0.168)
Meeting online × 6th+	−0.168	−0.177	−0.204	0.132	0.165	0.160
	(0.131)	(0.156)	(0.153)	(0.141)	(0.196)	(0.188)
Emotional intimacy	—	—	0.013	—	—	−0.077
	—	—	(0.040)	—	—	(0.049)
Esteem support	—	—	0.166***	—	—	0.198***
	—	—	(0.033)	—	—	(0.050)
Dominance of partner	—	—	−0.052	—	—	−0.115†
	—	—	(0.038)	—	—	(0.060)
Conflict	—	—	0.023	—	—	−0.017
	—	—	(0.047)	—	—	(0.062)
Educational matching (ref.: homogamy)						
Hypergamy	−0.044	−0.103	−0.105	—	—	—
	(0.042)	(0.067)	(0.065)	—	—	—
Hypogamy	−0.055	0.034	0.077	−0.029	−0.164†	−0.137
	(0.081)	(0.131)	(0.124)	(0.049)	(0.084)	(0.089)
Origin exogamy	−0.146*	−0.108	−0.084	−0.065	−0.016	−0.013
	(0.061)	(0.091)	(0.083)	(0.065)	(0.114)	(0.110)
Age matching (ref.: similar age)						
Man older >2 years	−0.043	0.011	−0.000	−0.017	−0.098	−0.091
	(0.037)	(0.066)	(0.064)	(0.044)	(0.082)	(0.080)
Woman older >2 years	−0.118†	−0.114	−0.111	−0.033	−0.174	−0.172
	(0.070)	(0.107)	(0.107)	(0.097)	(0.160)	(0.160)
Age of resident children (ref.: no children)						
Youngest ≤ 3	−0.215***	−0.356***	−0.328***	−0.267***	−0.261*	−0.200†
	(0.045)	(0.090)	(0.088)	(0.048)	(0.102)	(0.108)
Youngest > 3	−0.267***	−0.293**	−0.262**	−0.292***	−0.345**	−0.288*
	(0.045)	(0.095)	(0.093)	(0.051)	(0.126)	(0.126)
Division of paid work (ref.: male breadwinner)						
1.5 breadwinner	0.120***	0.022	0.029	0.165***	0.131†	0.141*
	(0.028)	(0.064)	(0.065)	(0.037)	(0.071)	(0.071)
Dual earner	0.383***	0.423***	0.413***	0.381***	0.507***	0.514***
	(0.040)	(0.086)	(0.087)	(0.046)	(0.083)	(0.081)

(*Continued*)

Table 4. Continued

	Non-tertiary			Tertiary		
	Unadj. sample	Adjusted sample		Unadj. sample	Adjusted sample	
	Model 2	Model 3	Model 4	Model 2	Model 3	Model 4
Other	0.309***	0.365***	0.371***	0.401***	0.421***	0.430***
	(0.032)	(0.073)	(0.073)	(0.039)	(0.093)	(0.088)
Constant	2.058***	2.105***	1.466***	2.280***	2.316***	2.166***
	(0.065)	(0.110)	(0.293)	(0.071)	(0.115)	(0.336)
Random effects						
σ_u	0.316***	0.081***	0.072***	0.273***	0.113***	0.101***
	(0.018)	(0.021)	(0.021)	(0.020)	(0.050)	(0.051)
σ_e	0.352***	0.573***	0.562***	0.347***	0.516***	0.504***
	(0.007)	(0.035)	(0.036)	(0.009)	(0.038)	(0.037)
Log pseudolikelihood	−5941.8	−986.2	−973.6	−3623.8	−560.8	−552.7
N (observations)	5645			3500		
N (individuals)	1219			779		

Note: (Robust) standard errors are in parentheses. Unadj. = unadjusted. σ_u = standard deviation of u_j (between-person heterogeneity), σ_e = standard deviation of ε_{ij} (within-person heterogeneity). $\dagger p < .10$; $^{*}p < .05$; $^{**}p < .01$; $^{***}p < .001$.

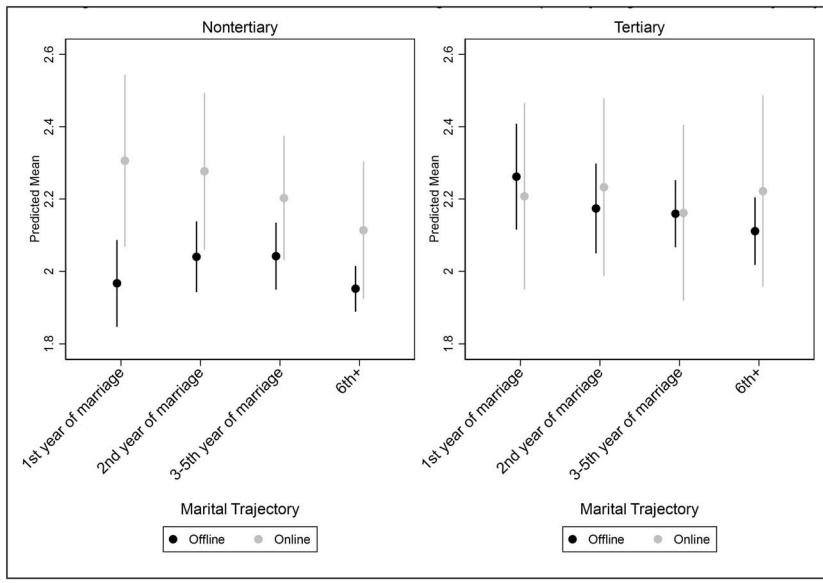

Figure 3. Predicted means of sharing routine housework, by stage in the marital trajectory and meeting context, for married women with non-tertiary and tertiary education.

Supplementary Analyses

We also conducted several supplementary analyses to examine robustness to alternative specifications. First, results from additional analyses (not shown) controlling for employment schedules (partners' share of overtime), cohort (instead of age), the presence of stepchildren, or social background (mother's education) as proxy for exposure to egalitarian gender norms in early life, are similar to those previously reported. Second, supplementary analyses (results presented in

Table F1 in the supplementary material) controlling for relative income as indicator of economic dependence between partners instead of educational sorting and division of paid work, revealed similar results. We did not include relative income in the main analysis due to a significant reduction in sample size had we done so. We also considered testing the effect of sorting on religion but could not include it given that religion was not measured on a yearly basis for either partner.

Furthermore, we aimed to see if previous patterns could also be observed for other outcomes of unpaid work. Results presented in Table G1, in the supplementary file revealed an even stronger positive association (than for routine housework) between meeting online and the sharing of shopping among lower-educated women. As was the case for housework, the association was only slightly offset by the inclusion of partnership quality measures. For childcare, results in Table G2 also indicated an online effect, albeit smaller, for less educated women. Finally, we included an additional indicator of gender role attitudes in the estimation of EB weights, namely, a mean score measuring attitudes toward women's paid work, based on the following two items measured on a five-point scale ranging from (1) "disagree completely" to (5) "agree completely," reverse-coded so that higher values indicate attitudes in support of women's employment: "Women should be more concerned about their family than about their career", and "A child aged under 6 will suffer from having a working mother" (Cronbach's alpha of 0.604). The results (Table H1 and Figure H1) are similar to those previously reported, with a reduction of 25.9% (instead of 22.3%) in the meeting online effect for lower-educated married women when accounting for selection bias via weights.

Discussion

In this study, we set out to investigate whether the way people select partners, particularly if through online channels, affects the division of routine housework, which remains a highly gendered practice in most of the Western world (Hook 2017). Through its many affordances (e.g., an enlarged and diverse field of eligible partners, a surplus of men), internet dating confers women greater bargaining power and the possibility to engage in more empowered, gender-equal interactions during courtship, which lay the groundwork for more progressive subsequent relationship practices (Lamont 2020), particularly within the more traditional confines of marriage. The willingness to challenge socially sanctioned conventions regarding the sharing of domestic work may originate in the self-selection of online daters, whose particularities (e.g., more gender-progressive and a more open personality) may determine how gender-disruptive their union will be. Beyond the influence of selection, we proposed that through online dating, women could select more supportive and emotionally compatible partners, which could contribute to establishing a more egalitarian division of household labor (Carlson et al. 2020; Komter 1989). To test these hypotheses, as well as scrutinize differences by educational level, we analyzed 2008–2019 panel data on 3305 cohabiting or married women from Germany. For the full sample, and then by education, we estimated random-effects regression models predicting couples' sharing of routine housework based on where they met.

Results show that internet dating seems to allow married women with non-tertiary education, who usually report the most traditional division of routine housework (Coltrane 2000; Domínguez-Folgueras 2013), to select partners with whom they more evenly share household tasks and consequently achieve parity with typically more advantaged groups (e.g., highly educated women). We found that controlling for observed selection into online dating via EB weights explained part of the effect, but did not fully account for it. Moreover, partnership quality had no mediating effect. Much of the online dating advantage remaining unexplained suggests that the different ways men and women negotiate power in the dating phase (Davis and Greenstein 2013) online versus offline may indeed play a role in how gender is enacted later on.

Alternative explanations may also relate to online dating allowing for better matching on attributes such as personality or gender role attitudes, which we could not include in the analysis

due to small partner samples (for women with non-tertiary education, for instance, their inclusion would have reduced the sample by 60.8%, from $n = 245$ to 96). Where information on partners was available, descriptive analyses (available upon request) did indicate that among couples who met in-person, highly trained women were more likely to be part of unions where both partners had egalitarian gender role attitudes than the less educated. If partner selection occurred online, the two educational groups showed no difference.

Another potential explanation lies in the fact that the internet facilitates relationships between partners with non-overlapping social networks, fewer friends, and who also are geographically distant from family. In her seminal work, Bott (1957) posited that spouses who have a less dense social network tend to share tasks, such as working, cleaning, and childrearing, more equally. We could not directly test this hypothesis given that measures related to social network and support were only included in Waves 4, 6, 8, and 10. Descriptive analyses (not shown) of other network indicators that were included in most waves, however, revealed that, irrespective of education, women who met their partner online were less satisfied with their friends and social contacts and lived further away from their mother than those who met their match elsewhere.

The different digital tools of partner search that women with different levels of education use may also be of relevance. We could not control for specific type of online meeting due to reasons explained above, but a look at the meeting contexts in Waves 10–11 indicates that lower-educated women were more likely to have met their spouse on social networks, which potentially erode traditional pairing to a greater extent than other online partner markets. Furthermore, additional analyses (not shown) indicated that the online effect was especially robust among those we met their partner in the pre-Tinder age. As data availability improves, distinguishing between different digital dating modalities and singling out the long-term impact of dating apps on questions of gender is a potential opportunity for future research. It is an open question, for instance, as to whether digital dating will continue to provide non-tertiary-educated women an advantage. As online modalities evolve toward greater popularity and more equal take-up, women's structural advantage and bargaining strength may fade over time.

One explanation for null findings among cohabiters and tertiary-educated married women is that the gendered division of routine housework is generally resistant to change, as previously noted. Observing no detrimental effect for these groups, however, indicates that the more negative aspects of internet dating, including the replication of traditional courtship rituals such as men's tendency to make the first move (Dinh et al. 2022), bear no effect on exacerbating gender inequality in the division of routine housework.

Another possible explanation for null findings among tertiary-educated women is the use of domestic help. Some women who declare taking up most of the housework may employ domestic help to assist with these tasks. In Waves 10 and 11, *pairfam* provides a measure of whether respondents regularly or occasionally employ a cleaner or other domestic help in their household. Cross-tabulations by education and meeting context (available from authors) revealed that high-educated women who met their partner online were significantly more likely to employ domestic help than were women who met their match in non-digital ways. We found no such difference among the lower-educated. The reason why we did not observe an online dating effect for women with tertiary education may therefore also have to do with women reporting about share of housework responsibility (i.e., mental charge) and not actual hours spent on housework. Future studies should compare both outcomes.

We also imagine the influence of period effects. Family policy reforms around parental leave and early childhood education and care were implemented and gained traction in Germany over this period (2008–2019), encouraging a less traditional division of household labor (Bünning 2015). These trends could have complimented what we observe for less educated women, yet there is no obvious reason why period effects would only be observed for couples that formed online.

Our study has several noteworthy limitations. First, even though we dealt with observed selection into digital dating via EB weights, we cannot rule out residual confounding due to

unobserved heterogeneity (that also varies over time). For example, additional descriptive analysis revealed an overrepresentation of those with previous marital experience among low-educated women who met their partner online. Though we include previous marital experience in the construction of EB weights, certain unobserved characteristics of lower-educated women who met their partner online may drive both the tendency to enact a more egalitarian division of chores and the propensity to divorce. Second, we acknowledge the issue of observing gender attitudes pre-partnering for only a part of respondents.

Despite limitations, this study suggests that, in light of the growing importance of online instruments for mate selection (Rosenfeld et al. 2019), internet matching might indeed play a part in accelerating the adoption of more egalitarian housework practices for those least likely to attain them offline—married women with lower education. Although online dating seems to function as a tool for the most disadvantaged to catch up to the more advantaged, it does not fully revolutionize gender as we are yet to see internet-matched couples in which partners equally share both paid and unpaid work[4] as the norm.

Endnotes

1. We do not include childcare in our main analysis for both theoretical and analytical reasons. In contrast to childcare, routine housework shows trends of disinvestment over time and is more resistant to gender equity (Sullivan 2013). However, we believe that the housework children create—cleaning up after children, doing their laundry, and preparing their meals—is included in the main analysis. Similar to Köppen and Trappe (2019), we also do not examine outcomes such as making repairs or handling administrative matters, given the different ways gender is enacted when completing occasional tasks as opposed to routine housework chores.
2. Irrespective of how fast partners move from online to offline communication, we argue that the different way selection is set up online already sets the tone for how the relationship unfolds over time.
3. Pre-partnering, women who date online are more likely to have children than those who do not. Unweighted descriptive statistics (available from authors) show that once in a union, as predicted, women in couples formed online are less likely than women in unions formed offline to have resident children. We account for both: the first when constructing weights and the second as a control in main analyses.
4. Additional analysis (results available upon request) shows that most online-formed couples with an egalitarian division of housework were less often dual earners.

About the Authors

Gina Potarca is a lecturer of Sociology at the University of Liverpool. Her research interests revolve around the application of digital, multilevel and longitudinal methods for the study of assortative mating, online partner selection, gender, and health. She has recently published in journals such as *Demography* and the *European Sociological Review*.

Jennifer L. Hook is a professor of Sociology at the University of Southern California. Her research examines how social contexts affect gendered divisions of work and care. Her recent work examines the influence of country context on women's employment, fathers' time with children, and the division of household labor. Her recent publications appear in the *American Sociological Review*, *European Sociological Review*, and *Journal of Marriage and Family*, among others.

Supplementary Material

Supplementary material is available at *Social Forces* online.

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Data Availability

The data underlying this article were provided by German Family Panel *pairfam* by permission. The *pairfam* datasets and materials can be downloaded by scientists who meet the criteria for access via the following link: https://www.pairfam.de/en/data/data-access/. Please contact the *pairfam* user support team for data access requests: support@pairfam.de.

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