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Is Your Spouse More Likely to Divorce You if You Are the Older Partner?

The authors assessed how the relative age of spouses affects whether men or women initiate a divorce, using data from the National Survey of Families and Households. Ex-spouses' reports of who left generally agreed, but not always, so the analysis used a latent class model embedded in an event-history model with competing risks that the woman leaves the man or the man leaves the woman. Support was not found for the hypothesis that age heterogamy itself increases the odds of divorce: Even large age differences did not make men more likely to leave younger wives, and women's exits were as likely when the marriage is homogamous as when she was older. The main conclusion is that both men and women are more likely to leave if their spouse is older than they are. The effects were stronger for men, but the gender difference in effect size was not statistically significant.

The relative age of spouses may affect how attractive one seems to the other, who is more likely to provide care for the other in old age, and whether a marriage violates social norms. Thus, one might expect relative age to affect the motivation of one partner to divorce the other. In this study we asked how the relative age of two spouses affects whether one partner leaves the other and whether having an older partner affects men's divorce decisions differently than it affects women's decisions. The second question is relevant because marriage is a quintessentially gendered institution, yet most past research has ignored the issue of whether the relative age configurations that lead men to leave marriages are distinct from those that encourage women to leave. Only one prior study, which used Dutch data, has explored effects of spouses' relative ages on both men's and women's initiation of divorces (Kalmijn & Poortman, 2006). Our analysis is the first to explore this question for the United States, and it will allow us to see if findings are similar across two affluent nations.

We used data from a longitudinal survey (the National Survey of Families and Households [NSFH]; see http://www.ssc.wisc.edu/nsfh/) that followed both ex-spouses if they divorced and asked each who had wanted the divorce more. Taking this response as an indicator of which spouse left the other or initiated the divorce, we examined how partners' relative age affects either partner's likelihood of divorcing the other. We examined whether the data are more consistent with perspectives claiming (a) that age homogamy stabilizes and age

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heterogamy weakens marriage, (b) that hypergamy (men being older) stabilizes marriage because of its gendered institutionalization, (c) that a double standard of aging leads men—but not women—to leave older spouses, or (d) that a more general and gender-neutral devaluation of aging leads younger spouses to leave spouses older than themselves.

Background

A prominent idea about marriage stresses that homogamy on almost any characteristic—for example, race, education, social class, religion, or age—promotes initial attraction and partnering as well as marital cohesion (Schwartz, 2013). The idea is that persons who have similar backgrounds will get along better because they share cultural values, lifestyles, or network connections. Thus, we would expect both husbands and wives to be more likely to initiate divorce in the presence of an age difference in either direction; the theory makes no prediction about whether a wife being older or a husband being older is worse for marital stability. A different view is that marriage is socially constructed as a gendered institution that favors *hypergamy*—that men are to be superior to women on such ordered dimensions as age, height, earnings, and education (Sayer, England, Allison, & Kangas, 2011). In this view, other things equal, either spouse is less likely to initiate divorce if the husband is older. Another perspective emphasizes *ageism* (Nelson, 2002), the devaluation of persons because they are older. In a gender-neutral version of this view, a single standard of ageism leads both men and women to be more likely to leave partners older than themselves. However, conventional wisdom suggests that ageism is not gender neutral, that there is a double standard of aging that arises in part from a cultural standard of beauty that privileges youthful looks, combined with an ethos in which women are judged by looks more so than men are (Carpenter, Nathanson, & Kim, 2006; England & McClintock, 2009; Gibson, 1996). Thus, men may be more likely to leave a marriage if their wives are older than themselves, but women may not devalue older men.

Studies that have examined effects of spouses' relative ages on divorce have produced mixed results. Clarkwest (2007) reported a marginally significant ($p < .10$, two tailed test) effect of the absolute value of husband minus wife's age, but the measurement scheme did not indicate whether the elevated divorce of heterogamous couples came entirely from the hypogamously married or equally from the hypergamously married. Phillips and Sweeney (2005) found that Blacks were least likely to divorce in the most hypergamous category (where the husband is more than 5 years older), but Whites' divorce rates did not differ across the three relative age categories. Heaton (2002) found the least divorce in the most hypergamous category (husband at least 2 years older). None of these studies divided the risk of divorce into the competing hazards of husband-initiated and wife-initiated divorces.

The only previous study to have examined the effect of spouses' relative age on the competing risk of divorces initiated by husbands and wives was conducted by Kalmijn and Poortman (2006), using Dutch data. They found that women were more likely to leave if their husbands were more than 5 years their senior ($p < .10$, two-tailed test, showing marginal significance), whereas men who were more than 5 years senior to their wives were significantly less likely to leave. The reference category was the relatively homogamous category in which men are no more than 5 years older and women no more than 1 year older than their spouse, and this homogamous category had no different odds of either men or women leaving than the category in which the wife was more than a year older. Kalmijn and Poortman also examined effects of relative age on whether there was a joint decision to divorce, a category containing only 10% of divorces, and found that the wife being more than 1 year older predicted such initiations, but the husband being more than 5 years older had no effect.

In sum, past studies hint that the husband being older than the wife stabilizes marriages, but the only study to have examined effects on divorces initiated by husbands versus wives used Dutch data and found that, other things equal, each gender exits more when they are the younger—and they are leaving an older—partner.

Method

We used data from Waves 1, 2, and 3 of the NSFH (*NSFH1*, *NSFH2*, and *NSFH3*, respectively). The NSFH comprises a national probability sample of 13,007 adults ages 19 and older first interviewed in 1987–1988. In married-couple households one adult was randomly selected as the primary respondent, and the spouse completed a self-administered questionnaire. Wave 2 was conducted in 1992–1994, and again data were obtained from both spouses. Analogously, Wave 3 was collected in 2001–2002. Because of funding, NSFH3 limited follow-up to respondents over age 45 at NSFH3 or those with an eligible focal child at NSFH1 (these children were at least 18 on January 1, 2001). At NSFH1, the response rate for primary respondents was 74%, and 83% of spouses of respondents completed questionnaires (Sweet, Bumpass, & Call, 1988). Of the 6,875 NSFH1 married couples, 86% had at least one partner interviewed at NSFH2; 71% of husbands and 80% of wives. Of couples married at NSFH1, the either-partner response rate at NSFH3 was 57%, with 37% of husbands and 47% of wives interviewed. The attrition rate, as well as the dropping of some respondents at Wave 3, are limitations of the NSFH. However, we used the NSFH because it is the only longitudinal U.S. data set on a national probability sample that both tracks marital dissolution and includes a measure of which spouse wanted the relationship to end. This measure is crucial to us because the novel contribution of our study was assessing the effects of spouses' relative ages on whether each of men and women leave marriages.

Of the original 6,875 married couples for whom one was interviewed at NSFH1, we used data from the 3,622 couples in which the selected adult and his or her spouse each completed the questionnaire at NSFH1, neither spouse was widowed by NSFH3, at least one of the two spouses (or ex-spouses) was reinterviewed at NSFH2 or NSFH3, the two persons did not disagree about the status of their marriage at NSFH1 (this eliminated only 12), and neither spouse was age 55 or older at NSFH1. Of these 3,622 couples in our analysis, 747 experienced a divorce or separation between NSFH1 and NSFH3.

We transformed the data on these 3,622 couples into a form in which couple-months are the units of analysis, appropriate to the event-history models described below. Each couple had a record for every month from the NSFH1 interview survey month to the month separation occurred, if it occurred, or to the NSFH3 survey month if the couple remained together. This yielded a data set of 413,650 couple-months, but we were unable to use the full set of couple-months because of the computational burden of estimating the latent class models. Thus, our working data set for model estimation was a disproportionate stratified random sample that included all 747 couple-months in which a divorce/separation occurred and an 8% random sample from the remaining couple-months, yielding a total of 33,738 couple-months. This kind of sampling does not bias the estimates from logistic regression (Allison, 2012, p. 103), and there is little loss of statistical power. In the original data set, 10% of the couple-months were contributed by couples who eventually divorced; because of oversampling by including all couple-months in which divorces occurred, in the final data set 12% of data came from couples who eventually divorced. For consistency, we present descriptive statistics as well from this sample of 33,738 couple-months; for descriptive statistics we weighted down the couple-months with a divorce to constitute the same proportion of the 33,738 that they constituted of the 413,650.

Our dependent variable is whether a breakup occurred and, if so, who initiated it. For couples who separated or divorced after NSFH1, the month in which the breakup occurred was ascertained at the next interview. Consistent with most past research (Teachman, 1983, p. 109), we considered marriages dissolved at the point of separation. In the wave after the divorce, each ex-spouse was asked which person wanted the breakup more. In the NSFH2, husbands and wives who had experienced a marital separation or divorce between NSFH1 and NSFH2 completed a self-administered module on the experience of relationship dissolution, including the following question, which ascertained which spouse most wanted the divorce: "Sometimes both partners equally want a marriage to end, other times one partner wants it to end much more than the other. Circle the number of the answer that best describes how it was in your case." Response categories included (1) "I wanted the marriage to end but my husband/wife did not," (2) "I wanted it to end more than my husband/wife did," (3) "We both wanted it to end," (4) "My husband/ wife wanted the relationship to end more than I did," or (5) "My husband/wife wanted the marriage to end but I did not." Whereas in the NSFH2 the question about who wanted the divorce was part of a self-administered paper module, in the NSFH3 the module was administered verbally by the interviewer. The question wording was changed slightly to collapse and simplify response categories:

> Some partners disagree about how much they want their marriage to end. In your case, who most wanted your marriage to end? Would you say that you wanted it most, you both wanted it equally, or that your husband (wife) wanted it to end most?

We used the collapsed categories for divorces between the NSFH1 and NSFH2 to make them compatible with these categories used for divorces between the NSFH2 and NSFH3.

We made the simplifying assumption that in most divorces one party wanted to break up more than the other, even if only slightly more. Our data reinforced the reasonableness of this assumption. Although 16% (women's report) to 25% (men's report) of divorces entailed both partners wanting it equally, partner agreement was least on the "both wanted equally" category, with fewer than half of men agreeing when their ex-wives gave this characterization. By contrast, when women said they wanted the divorce more, men agreed 67% of the time, and when women said the husband wanted the divorce more, 67% of the men agreed. The reasonableness of the decision was further buttressed by the finding from Dutch data that both partners initiated the separation jointly in only about 10% of divorces (Kalmijn & Poortman, 2006). However, a study that used Australian data showed a higher percent (25% by women's and 38% by men's report) in the "jointly initiated" category (Hewitt, Western, & Baxter, 2006). Our decision to assume that one party wanted to break up more than the other was also necessary as a practical matter; our attempts to estimate models that included a latent class for divorces equally wanted by both partners led to convergence failures.

We used both ex-partners' reports, together with the latent class analysis described below, to analyze who initiated the breakup and the determinants of such initiation. Although our data on the dependent variable consisted of reports of who wanted the divorce more, our discussion will interchangeably refer to these reports as indicating who wanted the divorce, who initiated the divorce, or who left whom.

A past analysis using these measures of who wanted the breakup (Sayer et al., 2011, Table 2) showed that, of women who reported on a marital dissolution between waves, 59% said they had wanted it more, 25% said that the man had wanted it more, and 16% said that both had wanted it equally; the analogous percentages for men are 46% (she wanted it more), 30% (he wanted it more), and 25% (they wanted it equally). The two parties in a couple agreed on which partner wanted the divorce more much more often than they disagreed. When they did disagree, it was usually a matter of one spouse saying that one wanted it more and the other saying they wanted it equally. There were very few cases in which one said the wife wanted the divorce more and the other said the husband wanted it more. There was, however, a problem of substantial missing data for the divorce-initiation question: Thirty-five percent (260 of 747) of ex-husbands and 22% (168 of 747) of ex-wives did not provide a report on which spouse wanted the divorce more. On this dependent variable or any of our independent variables, if we lacked one spouse's report, it could be either because the person was not interviewed at the survey wave or because, while interviewed, he or she did not answer the question.

Our key independent variables were based on the relative ages of the spouses. We present results from two models. One model treated spouses' relative ages as consisting of three categories: (a) wife was more than 3 years older than husband (hypogamy), (b) spouses' age difference was no more than 3 years in either direction (homogamy, our reference category), and (c) husband was more than 3 years older than the wife (hypergamy). One might argue that a 3-year difference in age is sufficiently trivial that a higher cutting point would be more meaningful. However, only 6% of wives were more than 3 years older than their husbands, only 3% were more than 5 years older, and only 0.6% were 10 years older. (Among those wives who divorced, only 7% were more than 3 years older than their husbands, 4% were more than 5 years older, and 0.5% were more than 10 years older.) Thus, to make the hypogamy category require a higher cutoff would render the group so small as to jeopardize statistical power. A second model simply used husband's age minus wife's age; this model presumed that relative age has a linear effect, with the wife being maximally older and the husband being maximally older as ends of an interval-level continuum. Although this measure is not useful for testing hypotheses about preference for homogamy, it is useful as an alternative test of ageist preferences for partners younger than oneself because it allows the magnitude of the age differences to affect divorce.

It is the relative age of the spouses that is posited to have an effect in theories maintaining that either homogamy or hypergamy stabilizes marriage. Our categorical measures tested these preferences, and our linear difference measure tested the effect of the magnitude of hypergamy. In the case of theories positing a single or double standard of ageism, it is less clear if what is prized is an absolutely young spouse or a spouse who is young relative to one's own age. Our strategy was to focus on predictions about and thus measure the relative age of spouses because two of the theories refer exclusively to this, and the others positing ageism may refer to relative age as well. Consistent with this strategy, relative youth was captured by our categorical and linear measures. We wanted to make sure that we were not also picking up effects of absolute age on our coefficients for relative age, but we were not able to come up with a strategy to do so that did not introduce troublesome collinearity. We controlled for marital duration and the two spouses' average age at marriage, but we could not also control for husband's and/or wife's absolute age. We thought it was unlikely that our measures of relative age will tap effects of absolute age, but we concede that it is possible.

Our models contained numerous control variables, chosen because they may be correlated with relative age and affect divorce and thus cause bias in our estimates if not included. Two indicators of spouses' potential earnings entered our models as controls: husband's and wife's recent employment. NSFH2 and NSFH3 respondents were shown a calendar and asked the month when they had started and ended any periods of employment between the previous and current wave; this allowed us to create a measure for whether each spouse was employed in each person-month of the marriage after the NSFH1. Our models entered the number of months each spouse was employed out of the 24-month period that ended 12 months before the month in question. (Spouses counted as employed if they were ever employed in the month, part or full time. However, a sensitivity analysis replacing this with only months when the work was full time produced very similar results.) Lagging the measure 12 months minimized possible endogeneity of employment with respect to separation or the intent to separate. We preferred measures of employment to earnings because the latter could only be updated once, at the NSFH2; with up to 7 years between NSFH1 (1987–1988) and NSFH2 (1992–1994) and up to 10 years between NSFH2 and NSFH3 (2001–2002), measures of earnings could precede the divorce by up to a decade, whereas employment was recorded for every month between waves. We controlled for a continuous measure of the duration of the marriage (expressed in years but accurate to the month). Results not shown indicated that when we added the square of duration to our models, only one model of four showed a significant effect of the square of duration, and models including the squared term produced virtually identical effects of relative age. We controlled for the presence of children and their ages using four variables: the number of children at each of ages 0–1, 2–5, 6–12, and 13–18. These measures were measured at the NSFH1 and updated after the NSFH2.

Other control variables were entered as unchanging characteristics of the couple or one of the spouses. For each spouse there was an indicator variable coded 1 if the person did not live with both biological or adoptive parents until age 19; in most cases a 1 indicates that the parents divorced, but it also taps situations in which one parent was never present, one parent died, or parents lived apart for some other reason. We controlled for cohabitation experience with dummy variables coded 1 if the husband (wife) had ever cohabited with anyone. We attempted to control for whether the bride was pregnant before the marriage with a dummy variable coded 1 if a birth occurred between 6 months prior to the marriage and 7 months after the marriage. We controlled for whether this was a second (or higher order) marriage with dummy variables for each spouse, coded 1 if the wife (husband) had ever been previously divorced (before the NSFH1). We controlled for race with a dummy variable coded 1 when either the husband or the wife was Black. (Most of these marriages contained two Black partners, but we did not include measures of each partner's race because it created high collinearity.) Because there may be more conflicts over stepchildren than children that spouses had or had adopted together, we also included two dummy variables: whether at the NSFH1 children were

Table 1. *Means and Standard Deviations of Independent Variables*

Variable	*M*	*SD*
Relative age dummy variables		
Wife > 3 years older than husband (0, 1)	.06	
Wife within 3 years (reference)	.65	
Wife > 3 years younger than husband (0, 1)	.29	
Husband's minus wife's age (years)	2.19	14.79
Mean of spouses' age at marriage	24.59	19.82
Wife's age at marriage (years)	23.49	20.35
Husband's age at marriage (years)	25.68	21.93
Either spouse Black (0, 1)	.10	
Wife's parents divorced (0, 1)	.21	
Husband's parents divorced (0, 1)	.20	
Wife ever cohabited (0, 1)	.27	
Husband ever cohabited (0, 1)	.30	
Wife previously divorced (0, 1)	.20	
Husband previously divorced (0, 1)	.21	
First birth conceived before marriage (0, 1)	.15	
Child(ren) present not husband's (0, 1)	.15	
Child(ren) present not wife's (0, 1)	.07	
Number and ages of children		
Number of children 0–1 years old	0.03	0.65
Number of children 2–5 years old	0.24	2.05
Number of children 6–12 years old	0.57	3.06
Number of children 13–18 years old	0.40	2.46
Wife's education		
< High school (reference)	.11	
High school graduate (0, 1)	.39	
Some college (0, 1)	.25	
College degree (0, 1)	.25	
Husband's education		
< High school (reference)	0.12	
High school graduate (0, 1)	0.33	
Some college (0, 1)	0.24	
College degree (0, 1)	0.30	
Number of months wife employed last 2 years[a]	16.27	36.23
Number of months husband employed last 2 years[a]	20.00	29.30
Duration of marriage (years)	18.01	35.97

[a]The 2-year period begins 12 months before the current month.

present who were not the biological children of the husband and analogously for the wife. Finally, we controlled for age at marriage by averaging the two spouses' ages at marriage. We did this rather than entering each spouse's age separately to avoid collinearity. In sensitivity tests not shown, models entering either the wife's or husband's age at marriage without the other produced the same substantive conclusions regarding our hypotheses. Means and standard deviations of independent variables are presented in Table 1.

We estimated models that combined latent class models with discrete-time event-history models. In these models, husbands' and wives' reports of divorce initiation were treated as fallible indicators of the true initiation status, which has three latent categories. In turn, the true initiation status was modeled in a competing-risks framework as a multinomial logistic regression.

We began with the latent class portion of the model. Data were arrayed with couple-months as the unit of analysis. In each couple-month, we assumed that couples could be classified into one of three states: (a) wife initiated a divorce/separation, (b) husband initiated a divorce/separation, or (c) neither initiated. Let C be a variable with values of 1, 2, and 0, corresponding to these three states. We did not directly observe C; instead, we observed the husband's report, H, and the wife's report, W. Both of these variables were coded 1 if the husband wanted the divorce more, 2 if the wife wanted the divorce more, 3 if both wanted the divorce equally, and 0 if neither partner initiated a divorce. For simplicity, the latent class variable did not allow for the possibility that both partners wanted the divorce equally, even though respondents may have chosen this option; we think it reasonable to assume that one spouse always wanted the divorce more, however slightly. Moreover, when we attempted to estimate a four-class model (results not shown) the model would not converge because only 18 of the 747 divorces were assigned to the latent class for "both wanted the divorce equally."

The relationship between the latent class variable and each spouse's report can be described by a set of conditional probabilities (the probability of falling into one of the observed classes given membership in one of the latent classes). The model specified the joint probability of the two observed variables and one latent variable as

$$\Pr(C = j, W = k, H = l) = \Pr(C = j) \times \Pr(W = k \mid C = j)\Pr(H = l \mid C = j)$$

This equation expresses the usual "local independence" assumption that H and W are statistically independent, conditional on the value of C.

The conditional probabilities were subject to the following constraints, which merely express the fact that if no separation occurred in a given month there is no discrepancy between the husband's report, the wife's report, and the value of the latent variable:

$$\Pr(W = 0 \,|\, C = 0) = \Pr(H = 0 \,|\, C = 0) = 1$$

$$\Pr(W = j \,|\, C = 0) = \Pr(H = j \,|\, C = 0) = 0\,, \quad j \neq 0$$

$$\Pr(W = 0 \,|\, C = j) = \Pr(H = 0 \,|\, C = j) = 0, \quad j \neq 0$$

The next step was to let the probability distribution of C depend on predictor variables in a multinomial logistic regression model, as described by Yamaguchi (2000). Let p_{ijt} be the conditional probability that couple i at time t is in state $C = j$, given that the couple has not already separated. Dependence on predictor variables is specified as follows:

$$\log\left(\frac{p_{ijt}}{p_{i0t}}\right) = \alpha_j(t) + \beta_j \mathbf{x}_{it}, \quad j = 1, 2$$

where $\mathbf{x}_{it}$ is a column vector of possibly time-varying covariates, $\boldsymbol{\beta}_j$ is a row vector of coefficients, and $\alpha_j(t)$ is some function of t. This multinomial logit model can be viewed as consisting of two simultaneous, binary logit models, one predicting initiation by the husband versus no separation, the other predicting initiation by the wife versus no separation.

The risk of separation/divorce began with the first month after the Wave 1 interview, although the dependence on duration $\alpha_j(t)$ was specified as the number of months since the couple was married. Couples in which both partners dropped out of the study by the NSFH3 without having separated by the NSFH2, or whose marriage remained intact from the NSFH1 to the NSFH3, were treated as right-censored observations.

The combination of the latent class and multinomial logit models can be viewed as a discrete analogue of the *multiple indicator–multiple cause* model for continuous variables (Hauser & Goldberger, 1971). We simultaneously estimated the two components of the model by maximum likelihood (with robust standard errors) using Mplus software (Muthén & Muthén, 2007). One of the attractions of this implementation of maximum likelihood is that it accommodates missing data on the dependent variable, under the assumption that the data are missing at random. This capability is especially important because a substantial fraction of couples were missing the report of either the husband or the wife. In these cases, the model assigned the case to either "husband wanted" or "wife wanted" divorce on the basis of the information available. About 29% of the couple-months had missing data on one or more predictor variables; multiple imputation was used to handle these missing data (Allison, 2001). Missing data were imputed under a multivariate normal model and the missing-at-random assumption using PROC MI in SAS; five imputed data sets were produced in SAS and input into Mplus. All model estimation analyses are unweighted. An intuitive explanation of what the latent class aspect of the model was doing is that, when one partner's report of who wanted the divorce more was missing or one partner stated that the divorce was equally wanted by both, the other partner's report was used by the model to assess which partner wanted it more, as were the values of other variables, insofar as the model showed them to be correlated to who wanted the divorce more.

We show results from two different versions of the latent class, event-history model; they differ in how the main independent variable of spouses' relative age was operationalized. Model 1 had two dummy variables to code three categories of relative age: (a) wife was more than 3 years older, (b) wife was more than 3 years younger, with a relatively homogamous category of (c) spouses within 3 years of each other as the reference category. Model 2 used a continuous measure of relative age: husband's minus wife's age. The first specification had the virtue of allowing the relationship to be nonlinear as, for example, if husbands were more likely to leave an older spouse but did not distinguish between homogamy and a younger spouse or if either husbands or wives were more likely to stay in homogamous than either hypergamous or hypogamous marriages. The second specification, while forcing linearity, had the advantage that it allowed the magnitude of age differences to matter. (We tested for nonlinearity by including the square of the age difference, but it was not significant, so we left it out.) We present conditional odds ratios (exponentiated coefficients, referred to below as *ORs*).

Results

Results from our latent class event-history models are presented in Table 2. Unless otherwise

Table 2. *Odds Ratios From Latent Three-Class Model Predicting Whether a Married Couple Stays Together, Husband Initiates Divorce, or Wife Initiates Divorce*

	Model 1		Model 2	
Predictor	Divorce, husband wanted more	Divorce, wife wanted more	Divorce, husband wanted more	Divorce, wife wanted more
Relative age dummy variables				
Wife > 3 years older than husband (0, 1)	1.87*	0.77		
Wife within 3 years (reference)				
Wife > 3 years younger than husband (0, 1)	0.50**	1.38**		
Husband's minus wife's current age (years)		0.92***	1.05***	
Mean of spouses' age at marriage	0.94*	0.92***	0.94*	0.92***
Either spouse Black (0, 1)	0.76	1.45*	0.76	1.49**
Husband's parents divorced (0, 1)	2.02**	1.41*	2.14***	1.32
Wife's parents divorced (0, 1)	0.88	1.60**	0.84	1.71***
Wife ever cohabited (0, 1)	1.35	1.44**	1.36	1.44**
Husband ever cohabited (0, 1)	0.97	1.06	0.96	1.06
Wife previously divorced (0, 1)	1.07	1.33**	1.11	1.31*
Husband previously divorced (0, 1)	1.03	1.05	1.08	1.02
First birth conceived before marriage (0, 1)	0.88	1.16	0.91	1.15
Number and ages of children				
Number of children 0–1 years old	0.33	0.54**	0.33	0.53**
Number of children 2–5 years old	0.76	0.76**	0.78	0.75***
Number of children 6–12 years old	0.84	0.92	0.85	0.91
Number of children 13–18 years old	0.98	1.13	1.01	1.12
Child(ren) present not husband's (0, 1)	0.75	0.85	0.75	0.86
Child(ren) present not wife's (0, 1)	1.45	0.67	1.43	0.65
Wife's education				
< High school (reference)				
High school graduate (0, 1)	1.02	0.79	0.95	0.82
Some college (0, 1)	0.86	0.81	0.81	0.85
College degree (0, 1)	1.22	0.73	1.17	0.76
Husband's education				
< High school (reference)				
High school graduate (0, 1)	1.05	1.16	1.15	1.14
Some college (0, 1)	0.95	1.08	1.06	1.05
College degree (0, 1)	0.93	0.96	1.03	0.93
Number of months husband employed last 2 years[a]	0.96***	0.98***	0.96***	0.98***
Number of months wife employed last 2 years[a]	1.01	1.01	1.01	1.01
Duration of marriage (years)	0.93***	0.91***	0.93***	0.90***

[a]The 2-year period begins 12 months before the current month.
*$p < .05$. **$p < .01$. ***$p < .001$, two-tailed tests.

noted, all effects that we discuss are statistically significant ($p < .05$) by a two-tailed test. In regard to Model 1, with categorical measures of relative age, if the wife was more than 3 years older, husbands had an 87% greater odds of initiating divorce (OR = 1.87), but wives had a 23% lower odds (OR = 0.77) of initiation (the latter effect is not significant at $p < .05$). Both are relative to the reference category of homogamy. Reestimation of the model changing the reference revealed that women were significantly more likely to leave if they were at least 3 years younger than their husband relative to if they were 3 years older (results not shown). If the wife was more than 3 years younger, husbands had a 50% lower odds (OR = 0.50) of initiating divorce, but wives had a 38% higher odds (OR = 1.38).

Of the theories discussed earlier, this pattern of effects is most consistent with the single standard of ageism: Both the husband and the wife were more likely to leave spouses who were older than they. The effects were stronger for husbands, which suggests a gendered double standard of aging, with men preferring a youthful spouse more than women do, but the differences between husbands and wives were not statistically significant (tests not shown). In other words, the following two effect difference contrasts were nonsignificant: (a) the effect of the wife being older on men leaving and the effect of the wife being younger on the women's leaving and (b) the effect of the wife being younger on men leaving and the effect of the wife being older on women's leaving.

The results were not consistent with a preference for homogamy by either men or women. The two findings inconsistent with that thesis are that (a) a man was only about half as likely to leave in the heterogamous case, where his wife was younger than he is, than in the homogamous case, where spouses were within 3 years of each other, and (b) women's exits were as likely when the marriage was homogamous as when she was older. The results were also not consistent with a uniform preference by both genders for hypergamy in divorce decisions because wives are more likely to leave husbands who were more than 3 years older than to leave those who were their age or younger.

The results for Model 1 suggest the simpler specification in Model 2. Here, relative age was represented by a single variable: husband's age minus wife's age. Model 2 fit the data better than Model 1 by two information criteria: (a) the Akaike Information Criterion (8,431.8 for Model 1 vs. 8,419.8 for Model 2) and (b) the Bayesian Information Criterion (8,987.8 for Model 1 and 8,959.1 for Model 2). Model 2 assumed a linear relationship, that is, that each additional year of difference between husband's age and wife's age has the same impact on the odds of initiating divorce. We tested for nonlinearity by including the square of the age difference in the model, but it was not significant in effect on either husbands' or wives' initiation of divorce.

In Model 2 the age difference had a significant effect on both husband's and wife's initiation of divorce (see Table 2). The OR of 0.92 implies that for each additional year that husband's age exceeds wife's age, the odds that he leaves go down by 8% and, equivalently, that for each additional year that her age exceeds his age the odds that he will leave increase by 8%. A 10-year difference would more than double the odds of leaving relative to no difference. For wives, the effect was in the same direction, but smaller; for each additional year of his age minus hers, she was 5% more likely to leave (OR = 1.05). The difference between the effects of 5% and 8% was not statistically significant, however. The findings from Model 2, like Model 1, seem most consistent with a single standard of ageism affecting both male and female spouses, in which the older one's spouse is relative to one's own age the more likely one is to leave.

Discussion

Our results show that the relative ages of spouses affect whether husbands leave wives as well as whether wives leave husbands and that, other things equal, the tendency is for the older of the two spouses to be left by the younger. The homogamy thesis has not been consistently supported in prior research that combined women's and men's exits, and our research casts further doubt on its relevance for predicting marital dissolution. Two findings speak loudly against the theory. First, a man is only about half as likely to leave in the heterogamous case where his wife is younger than he is than in the homogamous case where spouses are within 3 years of each other. Second, women's exits are as likely when the marriage is homogamous as when she is older.

Because our results show that difference in age has opposite effects on whether men or women leave, these effects will at least partially cancel each other out in models that fail to distinguish husbands' from wives' decisions to exit a marriage. Moreover, because women want the divorce more than men in approximately two thirds of divorces (Sayer et al., 2011), models assessing the effect of spouses' relative ages on the overall hazard of divorce without regard to who initiated it will reflect the wife's decisions more than the husband's. These considerations suggest that much past research on relative age on divorce has been misleading. Ours is the first study on U.S. data to show that both men and women are more likely to leave if they are the younger of the two partners. The similar finding from a high-quality Dutch study (Kalmijn & Poortman, 2006) suggests that the pattern may be more general.

Slight age hypergamy is the statistical norm in American marriages, with a typical husband being about 2 years older than his wife in the data used here (see Table 1). The findings on men's exits are consistent with a male preference for hypergamy, whereas women's exits showed no preference for hypergamy, as women are more likely to leave an older spouse. This suggests that the convention of hypergamy, through promoting later dissatisfaction by women, may contribute to the excess of women's over men's initiation of divorce found in many studies (Hewitt et al., 2006; Kalmijn & Poortman, 2006; Sayer et al., 2011). What is puzzling, though, is that people who were content with a particular age gap when they married would make that same age gap the basis of a divorce decision years later. In the case of men, perhaps those who ignored the social norm advocating hypergamy at marriage revert to the conventional view over time, thus affecting their divorce initiations. Alternatively, perhaps, for either sex the same spouse starts to look less appealing as he or she ages, and it is absolute age that is driving divorces. As we noted, our inability to control for absolute age in models including relative age makes this a possible interpretation of the findings if absolute and relative age are correlated net of covariates.

Our results show that having an older partner elevates men's odds of leaving more than it does women's, and Kalmijn and Poortman (2006) found this as well with Dutch data. However, in our sample these differences were not statistically significant. With only 747 divorces in our data and a very uneven split between the proportion of those initiated by men and women, our statistical power was limited. The most parsimonious conclusion thus is that the effect of age differences represents a single rather than double standard, affecting decisions of wives and husbands alike. It is quite possible, however, that a data set with more divorces and thus more statistical power would show that ageism in decisions about divorce is stronger in men than in women, and if this were found it would provide additional evidence for the gendered institutionalization of heterosexual marriage. This points to the need for a new survey that contains questions on which partner wanted a divorce more and that has a sample of sufficient size to allow us to ascertain whether the ageism in divorce decisions is gender neutral or a double standard.

References

Allison, P. D. (2001). *Missing data*. Thousand Oaks, CA: Sage.

Allison, P. D. (2012). *Logistic regression using SAS: Theory and application* (2nd ed.). Cary, NC: SAS Institute.

Carpenter, L. M., Nathanson, C. A., & Kim, Y. J. (2006). Sex after 40? Gender, ageism, and sexual partnering in midlife. *Journal of Aging Studies, 20,* 93–106. doi:10.1016/j.jaging.-2005.06.003

Clarkwest, A. (2007). Spousal dissimilarity, race, and marital dissolution. *Journal of Marriage and Family 69,* 639–653. doi:10.1111/j.1741-3737.2007.00397.x

England, P., & McClintock, E. A. (2009). The gendered double standard of aging in U.S. marriage markets. *Population and Development Review, 35,* 797–816. doi:10.1111/j.1728-4457.2009.00309.x

Gibson, D. (1996). Broken down by age and gender: The problem of old women redefined. *Gender & Society, 10,* 433–488. doi:10.1177/089124396010004005

Hauser, R. M., & Goldberger, A. S. (1971). The treatment of unobservable variables in path analysis. In H. L. Costner (Ed.), *Sociological methodology* (pp. 81–117). San Francisco: Jossey-Bass.

Heaton, T. B. (2002). Factors contributing to increasing marital stability in the United States. *Journal of Family Issues, 23,* 392–409. doi:10.1177/0192513X02023003004

Hewitt, B., Western, M., & Baxter, J. (2006). Who decides: The social characteristics of who initiates marital separation. *Journal of Marriage and Family*, *68,* 1165–1177. doi:10.1111/-j.1741-3737.2006.00321.x

Kalmijn, M., & Poortman, A. (2006). His or her divorce? The gendered nature of divorce and its determinants. *European Sociological Review, 22,* 201–214. doi:10.1093/esr/jci052

Muthén, L. K., & Muthén, B. O. (2007). *Mplus user's guide* (5th ed.). Los Angeles: Muthén & Muthén.

Nelson, T. D. (2002). *Ageism: Stereotyping and prejudice against older persons*. Cambridge, MA: MIT Press.

Phillips, J. A., & Sweeney, M. M. (2005). Premarital cohabitation and marital disruption among White, Black, and Mexican American women. *Journal of Marriage and Family, 67,* 296–314. doi:10.1111/j.0022-2445.2005.00117.x

Sayer, L., England, P., Allison, P. D., & Kangas, N. (2011). She left, he left: How employment and satisfaction affect women's and men's decisions to leave marriages. *American Journal of Sociology, 116,*1982–2018. doi:10.1086/658173

Schwartz, C. R. (2013). Trends and variation in assortative mating: Causes and consequences. *Annual Review of Sociology, 39,* 451–470. doi:10.1146/annurev-soc-071312-145544

Sweet, J. A., Bumpass, L. L., & Call, V. (1988). *The design and content of the National Survey of Families and Households*. Madison: Center

for Demography and Ecology, University of Wisconsin.
Teachman, J. D. (1983). Early marriage, premarital fertility, and marital dissolution: Results for Blacks and Whites. *Journal of Family Issues, 4,* 105–126. doi:10.1177/019251383004-001006
Yamaguchi, K. (2000). Multinomial logit latent-class regression models: An analysis of the predictors of gender-role attitudes among Japanese women. *American Journal of Sociology, 105,* 1702–1740. doi:10.1086/210470