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# INCOME DYNAMICS IN COUPLES AND THE DISSOLUTION OF MARRIAGE AND COHABITATION*

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*Several studies have shown that a wife's strong (socio)economic position is associated with an increase in the risk of divorce. Less is known about such effects for cohabiting relationships. Using a unique and large-scale sample of administrative records from The Netherlands, we analyze the link between couples' income dynamics and union dissolution for married and cohabiting unions over a 10-year period. We find negative effects of household income on separation and positive effects of the woman's relative income, in line with earlier studies. The shape of the effect of the woman's relative income, however, depends on the type of union. Movements away from income equality toward a male-dominant pattern tend to increase the dissolution risk for cohabiting couples, whereas they reduce the dissolution risk for married couples. Movements away from income equality toward a female-dominant pattern (reverse specialization) increase the dissolution risks for both marriage and cohabitation. The findings suggest that equality is more protective for cohabitation, whereas specialization is more protective for marriage, although only when it fits a traditional pattern. Finally, we find that the stabilizing effects of income equality are more pronounced early in the marriage and that income equality also reduces the dissolution risk for same-sex couples.*

There is growing evidence on the importance of wife's economic independence for the dissolution of marriage. Studies have generally found that the risk of divorce is increased when the wife is working for pay and when the wife works more hours (Becker, Landes, and Michael 1977; Blossfeld and Muller 2002; Bracher et al. 1993; Cherlin 1979; De Rose 1992; Heckert, Nowak, and Snyder 1998; Jalovaara 2003; Manting and Loeve 2004; Poortman and Kalmijn 2002; South 2001; Von Gostomski, Hartmann, and Kopp 1998; Wagner 1993). The evidence is found both in the United States and in Europe. Theoretically, these effects are most often interpreted in terms of reduced specialization gains on the one hand and lower financial exit costs on the other hand (Becker 1981; Cherlin 1979, 1992; Oppenheimer 1997; Schoen et al. 2002).

Although the number of studies examining the link between women's socioeconomic position and divorce has increased greatly in recent years, important gaps in our knowledge remain. First, most of the evidence applies to the effect of the wife's labor force participation and—largely because of longitudinal data limitations—fewer studies have examined effects of partners' dynamic income levels. Hence, the evidence for relative income effects on separation so far is less consistent than the evidence for the effect of wife's employment (Rogers 2004). Some studies have found a clear positive effect of the wife's income share on divorce (Heckert et al. 1998; Jalovaara 2003; Liu and Vikat 2004); some authors have found an inverted U-shaped effect, with high levels of divorce occurring when husbands and wives have equal incomes (Rogers 2004); yet other authors have found a U-shaped effect of wife's income on separation, with high levels of divorce occurring when the wife has a low or a very high income (Ono 1998).

Second, we know little about economic influences on the dissolution of cohabiting relationships. An increasing number of unmarried couples live together, and these unions are

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known to be very unstable (Brines and Joyner 1999; Manting 1994b). Knowledge about the causes of this instability is limited, and theories about economic specialization and independence have rarely been tested for these relationships. A related gap is that we know little about whether economic theories of divorce apply to all couples or whether they are true only under certain conditions. This is especially important in light of recent and older criticisms of the economic theory. Several authors, for example, have emphasized the benefits of similarity and role sharing as opposed to the benefits of specialization and differentiation (Brines and Joyner 1999; Nock 1995; Oppenheimer 1997; Rogers 2004; Simpson and England 1981). Following these criticisms, one could argue that economic specialization in marriage is positive for some relationships but actually detrimental to other relationships.

In line with these criticisms, a few recent studies have suggested that the validity of economic theories is indeed conditional. Brines and Joyner have compared cohabiting and married unions and have shown that the number of hours the woman works has a positive effect on divorce—in line with economic theory—but a negative effect on the dissolution of cohabiting unions (Brines and Joyner 1999). Several studies have also shown that the effects of wife's employment are more modest in more recent periods than in older periods (Bracher et al. 1993; Poortman and Kalmijn 2002). In contrast to these findings, South (2001) found that for the United States, the effect of wife's employment on divorce increases over time. Studies have also interacted wife's economic characteristics with couple's value orientations, and in particular with women's gender norms, the idea being that in more egalitarian couples, the effects of wife's labor force participation on separation would not be detrimental to marriage. This idea has been corroborated in some studies (e.g., Kalmijn, De Graaf, and Poortman 2004) but not in others (e.g., Sayer and Bianchi 2000).

In this article, we examine whether the relative income position of the wife has a different effect for the dissolution risk of cohabiting couples than for the dissolution risk of married couples. More specifically, we test the hypothesis that specialization is stabilizing for married couples, whereas an egalitarian income pattern is stabilizing for cohabiting couples. In doing this, we replicate an earlier study on this topic for the United States (Brines and Joyner 1999).

Our work is not only a replication of this study; it also introduces a number of new elements. First, our work makes methodological progress. We analyze the hypothesis with a much larger sample of cohabiting unions, and we use better income data than have been used before, that is, data from annual tax records in The Netherlands over a period of 10 years. Our study includes 3,417 marriages and 9,725 cohabiting relationships, whereas the important study by Brines and Joyner included 1,855 marriages and 337 (long-term) cohabiting relationships. This means that we have more statistical power in estimating interaction effects of marriage and cohabitation on the one hand and income variables on the other hand.

Second, our data set allows us to look at heterogeneity within the cohabiting population. For example, we include same-sex cohabiting couples in our analyses. A well-known hypothesis is that egalitarian income patterns would be especially stabilizing for same-sex couples (Blumstein and Schwartz 1983). In a sense, the advantages of income equality that are believed to work for cohabiting couples may be especially important for same-sex cohabiting couples. In addition, we can distinguish between short- and long-term cohabiting relationships. For many couples, cohabitation is a stage before marriage, so that long-term cohabiting relationships may differ more from marital unions than do short-term unions (Brines and Joyner 1999).

A third contribution is that we aim to generalize the evidence to a different country, The Netherlands. This is important because cohabitation is a more accepted option in The Netherlands than in the United States. For example, data from the 2002 survey Family and Changing Gender Roles show that the number of people who disagree or strongly disagree with "couples living together without being married" is 36% in the United States but only

5% in The Netherlands.[1] Hence, there are very few moral objections to cohabitation in The Netherlands, at least in contemporary times. Cohabitation is also more common. The share of unmarried couples in all couple households is 9% in the United States (Simmons and O'Connell 2003) and 16% in The Netherlands.[2] The percentage of recently formed unions that start out as unmarried is 55% in the United States and about 70% in The Netherlands (Liefbroer and Dykstra 2000; Smock and Gupta 2002). Moreover, there are indications that cohabitation more often is a stage before marriage in the United States than in The Netherlands. The percentage of cohabiting couples who are married after three years is almost 50% in the United States but only about one quarter in The Netherlands (Latten and Cuyvers 1994; Oppenheimer 2003). Whether the thesis suggested by research on the United States can be generalized to other contexts is therefore an important empirical question.

## HYPOTHESES

In our view, there are economic and cultural ways of looking at the effect of wife's relative income on marriage and separation. The economic approach argues that different income arrangements in marriage change the financial costs and benefits of marriage and divorce. The cultural approach argues that different income arrangements in marriage have different meanings to couples, depending on the couples' value orientations and normative expectations. These two arguments can both be true at the same time. A given income arrangement in marriage can have a certain financial benefit, but this advantage can be counteracted by the normative disapproval that husbands and wives have of such an arrangement. In other words, we need to consider economic and cultural arguments simultaneously.

There are two microeconomic arguments about the effect of wife's relative income on separation (Becker 1981; Brines and Joyner 1999; Oppenheimer 1997; Rogers 2004; Schoen et al. 2002). First, when women have an independent income, they are better able to leave a poor marriage. In other words, the financial exit costs are lower when women have a higher income. This argument applies especially to the wife's income vis-à-vis the husband, that is, the wife's *relative* income. In The Netherlands, every person has the right to a minimum level of income through welfare so that financial independence is guaranteed, whether or not the wife works for pay. More relevant is the standard of living to which the wife has become accustomed. For the wife of a rich husband, receiving only welfare benefits after divorce will be experienced as downward mobility. For the wife of a low-earning husband, being on welfare will not be such a negative experience. The perceived exit costs will thus depend on how the wife's income level after divorce compares to the level of affluence that she experienced in marriage. For that reason, it is her income relative to that of the husband that matters for her perceived exit costs.

The argument of exit costs applies to the husband as well. Recent studies have shown that men may also experience a financial deterioration after divorce, especially in cases in which the woman brought in a large part of the household income (McManus and DiPrete 2001). If the husband is to a large extent dependent on the wife for his economic well-being, his exit costs will be high, and this implies a low risk of separation as well. As a result, we would expect the effect to be symmetric and the probability of separation to be lower when the income shares of husband and wife are unequal.

A second economic argument focuses on specialization (Becker 1981; Oppenheimer 1997). When both partners work for pay, there is less specialization in marriage, which reduces the gains to marriage. That specialization is beneficial is often illustrated by the finding that men can invest more in their careers when the wife is not working for pay. Such

1. These figures are based on our own calculations of the data. The data were collected for the International Social Survey Programme and distributed by the Zentralarchiv für Empirische Sozialforschung at the University of Cologne, Germany.

2. These figures are based on our own calculations from data made available online by Statistics Netherlands (see http://statline.cbs.nl).

investments lead to an increase in income, which in turn increases the household utility. Another benefit of specialization (or cost of nonspecialization) can be seen in the time pressure that dual-earner couples experience in their day-to-day lives. Especially when both partners work and when there are young children in the household, life may be experienced as more stressful and less comfortable (Schor 1991). In principle, the argument about specialization is gender-neutral. Hence, specialization occurs not only when men are the sole earners but also when women are the sole earners.

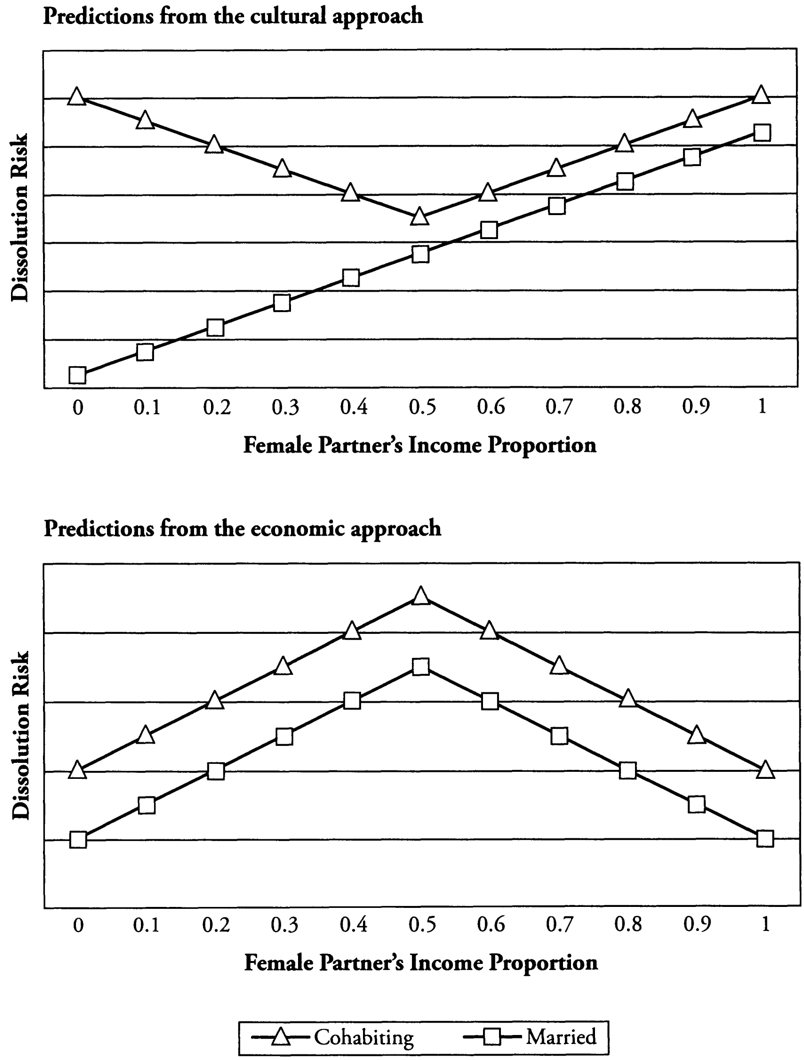
The microeconomic arguments outlined above are based on costs and benefits and ignore the preferences that couples have. In reality, different couples have different preferences, and these preferences are to a large extent based on their value orientations. Cultural arguments about the effect of wife's relative income on separation focus on these value orientations. The income arrangement that is chosen in marriage is directly related to the gender roles of men and women in marriage, and these roles are based on norms and values. It can therefore be argued that the evaluations of a certain income arrangement in marriage are different for couples with a traditional value orientation to gender than for couples with an egalitarian value orientation to gender.

Couples who have a traditional orientation to gender tend to prefer a situation in which the husband earns more, and they tend to disapprove of a situation in which the wife earns more. A husband may consider a situation in which his wife earns more to be a threat to his male breadwinner identity and may therefore disapprove of such an arrangement (Komarovsky 1962). This leads to the hypothesis that the more income the wife contributes, especially when she contributes more than her husband, the more unstable the marriage will be. For couples who have an egalitarian orientation to gender roles, we would expect to find quite different effects. Because such couples value equality between men and women, they will prefer to have (near) equal income contributions in marriage and will disapprove of situations in which either the husband or the wife earns more. In other words, in egalitarian couples, role collaboration and role sharing are preferred in relationships, rather than a division of labor along gender lines (Brines and Joyner 1999; Ono 1998; Rogers 2004).

We use the economic and cultural approaches for developing hypotheses about differences between cohabiting and married couples. We first note that there are important differences in value orientations depending on the legal status of the union. Several studies have shown that people who cohabit have a less traditional outlook on life. Cohabitors are less religious, more individualistic, and less dedicated to traditional family values (Lesthaeghe and Surkuyn 1988; Liefbroer 1991b; Rindfuss and Van den Heuvel 1990; Thomson and Colella 1992; Thornton, Axinn, and Hill 1992; Waite 1995). Attitudes toward gender roles also differ between cohabitors and married persons (Blumstein and Schwartz 1983; Smock 2000). Studies both in The Netherlands and in the United States show that people with liberal gender role attitudes are more likely to choose cohabitation over marriage (Clarkberg, Stolzenberg, and Waite 1995; Liefbroer 1991a).

If the assumption that cohabiting couples have more egalitarian values than married couples is correct, we can expect differences in the effect of relative income on dissolution, depending on whether the couple is married or unmarried (see Figure 1). According to the cultural approach, married couples will have higher divorce risks the higher the income share of the wife. For cohabiting couples, the cultural approach implies a model of equality: the more equal the incomes of the two partners, the *lower* the risk of dissolution. The economic approach has only one implication for both cohabiting and married couples. It predicts that the more equal the incomes, the *higher* the dissolution risk. Note that for cohabiting couples, the two approaches have opposite predictions, whereas for married couples, the predictions are different only for cases of reverse specialization (Figure 1). When the wife brings in most of the income, the marriage is stable according to the economic approach (due to specialization gains) but unstable according to the cultural approach (a violation of traditional gender norms).

Figure 1. Predicted Relation Between Female Partner's Income Proportion and the Dissolution Risk




It is useful to make further distinctions among the married and cohabiting unions in our data. First, we examine the role of union duration in combination with the role of the wife's income share. Previous research has shown that several determinants of union stability change over the course of the union (Manting 1994a; Morgan and Rindfuss 1985; South and Spitze 1986). For the effects of the wife's relative income, only few previous findings exist. In the United States, South (2001) found that the effect of the wife's working hours on divorce is absent early in the marriage and becomes more positive over the course of the marriage. For cohabiting unions, no such effects are yet found. In this article, we focus on hypotheses about the differences between married and unmarried unions. We expect that long-term cohabiting unions will differ more from marriage than do short-term cohabiting unions. Many short-term cohabiting unions convert to marriage, and it is therefore plausible that primarily the long-term cohabiting unions really reflect a value difference (Brines and Joyner 1999). To put it somewhat simply, cohabiting in the short run is a stage before marriage, and cohabiting in the long run reflects a choice for an alternative arrangement. As a result, one would expect the stabilizing effect of income equality to hold especially for long-term cohabiting couples.

Second, we formulate a hypothesis about same-sex cohabiting couples. Previous research has shown that cohabiting same-sex unions are more egalitarian in the way they divide paid and domestic labor than are heterosexual couples (Blumstein and Schwartz 1983; Ciano-Boyce and Shelley-Sireci 2002; Solomon, Rothblum, and Balsam 2005). Because clear gender roles are lacking, or at least more difficult to enact in same-sex couples, the more equal division of labor in such couples is not surprising. These differences further suggest that the advantages of income equality that are believed to work for cohabiting couples would work especially for same-sex couples. We therefore expect that an equal division of paid labor reduces dissolution chances for same-sex couples as well.

## DATA, METHOD, AND VARIABLES

### Data

In this article, we use a special and rather unique source of data: tax record data from The Netherlands, the so-called Income Panel Study (IPO). Some European divorce studies have used register data, but these studies typically did not include cohabiting relationships (Jalovaara 2003; Liu and Vikat 2004). The IPO contains a 0.6% sample from the population. Currently, information is available for about 115,000 persons who were in the sample between 1989 and 2000. The IPO data are based on a sample of individuals who file income taxes. To obtain information on union status, the IPO respondents were matched to data from the central population register. In doing this, information can be obtained about whether the person in the tax register data is part of a (married or unmarried) couple. Moreover, this procedure also makes it possible to identify the partner and thus to match the tax data from the partner to the IPO sample respondents. Even if partners file separately or if only one partner files income taxes because the other has no income, this information is added.

The IPO is an excellent source for studying the relationship between income and union dissolution. First, the data contain longitudinal information on income for each year, and because these income data are obtained from tax records, they are highly reliable. Second, in comparison with earlier studies, our sample is larger and includes more cohabiting relationships. Third, an advantage of a register panel is that there is almost no panel attrition. Of the respondents that we consider, we lose about 2%, mostly due to mortality or emigration. Fourth, the income data are at the level of the couple, which is obviously attractive for analyzing divorce and separation.

There are also disadvantages of our data. First, because the data were not collected for the purpose of studying divorce, they contain few possible divorce determinants. Second, the duration of observation is rather short. In any panel analysis without retrospective information, one needs to look at relationships that were formed during the panel window, and this means that we can look at only the first 1 to 10 years of the union. The average number of years we observe a union in our data is about four years. Third, the exact timing of union formation and dissolution is unknown. Events can be estimated only by comparing the situation at the end of a year to the situation at the end of the following year. This implies that the formation and dissolution of short unions will be underestimated because unions that began and ended in the same calendar year are simply not recorded.

We make the following selections from the data. First, we select unions formed between 1989 and 1999 because we want to study newly formed unions in the 1990s and because we want to avoid left-censoring. Second, we exclude persons who were divorced or widowed before the start of the union they formed during the panel period. Note that we are unable to exclude persons who ended an unmarried cohabiting relationship before the time we observe them during the panel.[3] Third, we exclude cases with missing data on central variables and cases with negative income variables. Fourth, we exclude unions that lasted only one year because in this case, no income information is present for the year that precedes the breakup. Finally, we include same-sex couples in the analyses. We limit same-sex couples to households in which the respondent is 30 years or older. This makes it unlikely that nonsexual unions, such as two same-sex students living together, are included (there are very few students older than 30). Mother-daughter households and other family households were never counted as couples because this information was available in the population register data. The remaining number of unions in the person-period file is 13,142 opposite-sex and 731 same-sex unions. Of the opposite-sex unions, 74% (9,725) began as unmarried cohabitations (3,417 were married directly). This corresponds well to other estimates for The Netherlands (see above).

Dissolution in a given year occurs when a person was part of a couple at the end of the previous year but is not part of a couple at the end of the current year. The end of living together is counted, regardless of whether there was an official divorce. The number of breakups of marriage in our data is 739, the number of dissolutions of opposite-sex cohabiting relationships is 2,931, and the number of dissolved same-sex couples is 544. Right-censoring occurs in the case of death or emigration of the respondent or at the end of the observation period.[4]

## Method and Variables

We use discrete-time event-history analysis to analyze the data. The dependent variable is the conditional log odds of the dissolution of a union at the end of a calendar year. Duration dependency is taken into account by a set of dummy variables for each duration of the union. The duration of the union starts at the beginning of cohabitation or marriage, whichever came first. The clock is *not* reset when a cohabiting union changes into a marriage. The duration effect is initially assumed to be equal for cohabiting and married respondents. Later models include interaction effects of duration and marital status.

The central independent variable is the legal status of the union measured at the end of the previous year. We include two dummy variables: cohabiting (cohabiting = 1, directly married = 0, married after cohabitation = 0) and married after cohabitation (cohabiting = 0, directly married = 0, married after cohabitation = 1). These are time-varying variables. The respondents who married directly are the reference category.

The income variables were obtained at the end of the year preceding the risk year. The information was obtained from both partners. We consider all income, including income from labor, social security, pensions, and other legal sources. Income reported on the individual tax record is considered individual income. If there were income sources that can be shared (e.g., welfare), we relied on the way it was reported by the respondent.

From this information, various income variables were constructed. The first variable is *total household income* (after taxes and corrected for inflation, using 2000 as the reference point, and scaled in 1,000€, with 1,000€ roughly equal to USD$1,200). This variable is the sum of the two personal incomes. To estimate the effects of the female partner's income, we calculated the *share of the female partner's income* of the total income, expressed as a

3. Because we do not have marital-status information for the spouse, some of the relationships may be second or higher-order marriages for the spouse.

4. To assess the possible death of a spouse, we look at the change in marital status to "widowed." We cannot assess the possibility that a cohabiting partner's death is erroneously included in the separation category. Detailed cross-checking on a subsample with vital statistics suggests that this occurred in 0.4% of the cohabitation separations.

proportion. To estimate the effect of this proportion, called $fpp_i$, alternative specifications were used:

$$\text{linear: } fpp_i \tag{1}$$

$$\text{curvilinear: } fpp_i \text{ and } fpp_i^2 \tag{2}$$

$$\begin{aligned}\text{splines: } & [fpp_i - 0.5] \text{ if } fpp_i \geq 0.5 \text{ (0 otherwise) and} \\ & [0.5 - fpp_i] \text{ if } fpp_i \leq 0.5 \text{ (0 otherwise)}\end{aligned} \tag{3}$$

$$\begin{aligned}\text{categorical: } & fpp_i = 0,\ 0 < fpp_i \leq 0.2,\ 0.2 < fpp_i \leq 0.4,\ 0.4 < fpp_i \leq 0.6, \\ & 0.6 < fpp_i \leq 0.8, \text{ and } 0.8 < fpp_i \leq 1.0.\end{aligned} \tag{4}$$

The curvilinear specification and spline specifications allow us to assess the degree of symmetry in the effect of relative income. The curvilinear specification allows for a U-shaped effect on separation, but it does not force the minimum or maximum level to be at a point of equal sharing of income. The spline function forces the midpoint to be at equal sharing but allows for asymmetry because the effect of relative income for women whose share is less than 50% may be weaker than the effect of relative income for women with shares over 50%. The last specification allows each category to have its own dissolution risk and is the least restrictive.

The selected demographic control covariates are the sex of the respondent, the presence of children in the household (time-varying), foreign background (whether or not one or both parents of the person were born abroad), age at the start of the union, degree of urbanization of the current residence (time-varying), and the age difference between partners. We are especially interested in large age differences because these are believed to be destabilizing and can also be associated with differences in income between partners. Age differences are measured with two dummy variables: the man is seven or more years older and the woman is seven or more years older. Small changes in this age definition did not alter the results. When same-sex couples are included, these two dummy variables are replaced by a single variable (age difference of seven or more years).

For the construction of the children variable, we made some special arrangements. The tax record data do not allow us to assess whether these are children of the respondent *and* the partner. This is potentially problematic because children from previous relationships may have a different impact on the stability of a union than children of the couple (Morgan and Rindfuss 1985). To solve this problem, we compared the ages of the children to the length of the union. We consider only children whose ages were lower than or equal to the duration of the union. Means and standard deviations of the independent variables can be found in Table 1.

## ANALYSES

### Descriptive Analyses

We start with a number of descriptive figures. To allow for comparisons with earlier studies, the descriptive figures do not include same-sex couples. In Figure 2, we look at the separation risk of marriages and cohabiting relationships by the duration of the union. For marriages, the duration is the duration of the union, including the possible prior cohabiting years. The figure shows that the risk of separation is much higher for cohabiting relationships than for marriages. We also see that the risk declines with the length of the union. This may be attributed to the increasingly selective nature of the survivors (an increasingly stable group remains). The decline in the risk is especially sharp for the first five years of cohabiting relationships. After that, the risk continues to decline, but the line is flatter. This shows that the first five years of cohabitation are the real "weeding" years.

**Table 1. Means and Standard Deviations (in parentheses) of Independent Variables**

| Variable | First Year | Person-Year File |
|---|---|---|
| Relationship Type | | |
| Married directly | .25 | .31 |
| Cohabiting | .75 | .47 |
| Married after cohabiting | .00 | .22 |
| Same-sex couple | .05 | .03 |
| Income Variable | | |
| Household income (× 1,000) | 26.4 (9.5) | 28.6 (10.4) |
| Female partner's proportion of household income | 42.3 (17.6) | 37.3 (18.3) |
| Control Variables | | |
| Duration | | 3.57 |
| Union before age 25 | .30 | .33 |
| Union after age 39 | .06 | .04 |
| First- and second-generation immigrant | .17 | .14 |
| City residence | .46 | .42 |
| Child in union aged 0–5 | .05 | .29 |
| Child in union aged 6–17 | | .00 |
| Age difference is seven or more years | .16 | .13 |
| Male is seven or more years older | .12 | .11 |
| Female is seven or more years older | .03 | .02 |
| Female respondent | .47 | .48 |
| *N* | 13,873 | 58,283 |

*Notes:* Income measures and age differences apply to opposite-sex couples only. The number of divorces is 739, the number of breakups of opposite-sex cohabiting unions is 2,931, and the number of break-ups of same-sex unions is 544.

**Figure 2. Separation Risk of Unions, by Type of Union**

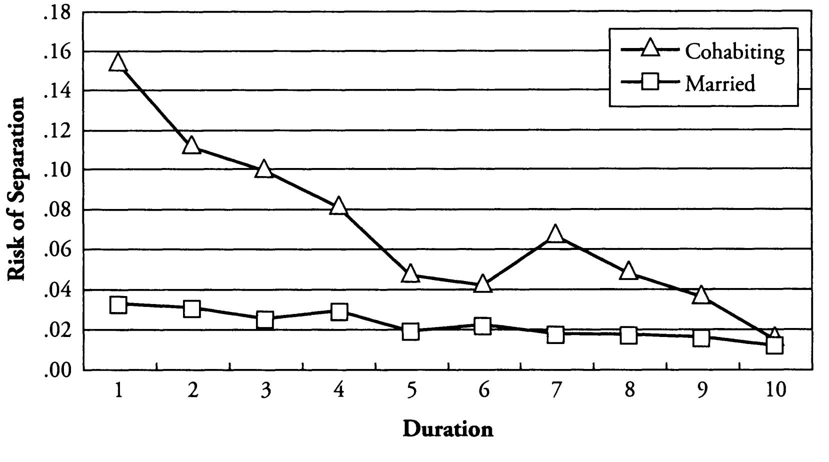


Figure 3. Female Partner's Income Share Over the Course of the Relationship

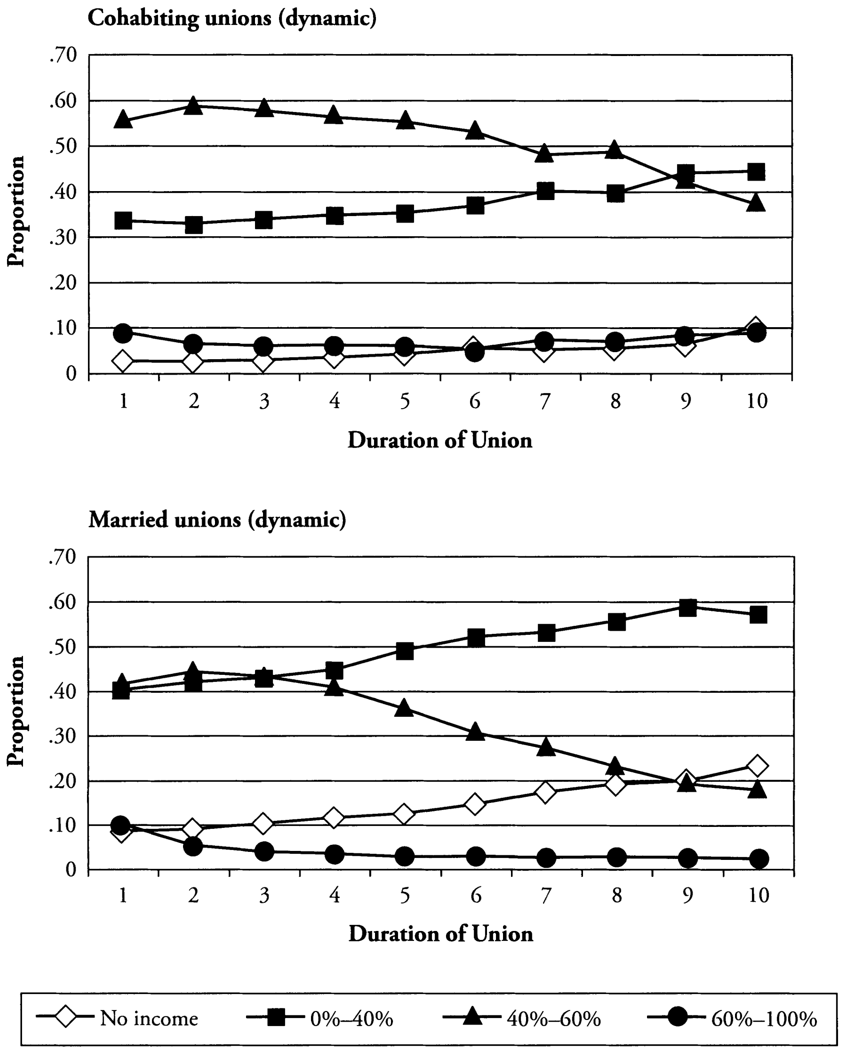



Figure 3 shows how the income contribution of the female partner changes with the duration of the union. Let us first focus on what happens in marriage. At the start of marriage, we see that the situation of equality is common. In about 40% of the marriages, the wife contributes 40%–60%. In another 40% of the marriages, we see that the wife makes a small contribution. The changes we see during marriage are quite dramatic. The condition of equality drops from 40% to 20% in 10 years time, whereas the situation in which the wife makes a small contribution increases from 40% to 60%. Although it is well known that gender inequality in income increases during marriage (especially during the childbearing and childrearing period), our data provide powerful evidence of these household income dynamics.

Next, we look at cohabiting relationships. We first compare these relationships to marriage, thereby focusing on the first years of the union. Overall, we see that gender income inequality is smaller in cohabitation than in marriage. The share of equal-income households is 15 percentage points higher in cohabiting relationships than in married relationships, and the share of couples in which women contribute no income or little income is lower. Changes are also somewhat different. Although gender inequality increases in cohabiting unions as well, this increase is more modest in cohabiting couples than in married couples.

## Event-History Models of Union Dissolution

**Baseline model.** Table 2 presents a baseline model without income variables. It includes same-sex couples. The most important result is the contrast between marriage and cohabitation. We see that cohabiting relationships have, on average, a 3.7 times higher odds of separation than do marital relationships. It is interesting to note that there is no significant

**Table 2. Baseline Discrete-Time Event-History Model of Union Dissolution**

| Variable | *b* | *p* | exp(*b*) |
|---|---|---|---|
| Duration Dependence (vs. 1 year) | | | |
| Duration 2 years | –0.356 | .00 | 0.700 |
| Duration 3 years | –0.498 | .00 | 0.608 |
| Duration 4 years | –0.611 | .00 | 0.543 |
| Duration 5 years | –1.071 | .00 | 0.343 |
| Duration 6 years | –0.980 | .00 | 0.375 |
| Duration 7 years | –0.915 | .00 | 0.401 |
| Duration 8 years | –1.005 | .00 | 0.366 |
| Duration 9 years | –1.122 | .00 | 0.326 |
| Duration 10 years | –1.681 | .00 | 0.186 |
| Type of Union | | | |
| Cohabiting (vs. married directly) | 1.296 | .00 | 3.655 |
| Married after cohabiting (vs. married directly) | –0.068 | .29 | 0.934 |
| Same-sex couple | 1.147 | .00 | 3.149 |
| Control Variables | | | |
| Union before age 25 (vs. ages 25–39) | 0.144 | .00 | 1.155 |
| Union after age 39 (vs. ages 25–39) | 0.005 | .95 | 1.005 |
| First- and second-generation immigrant | 0.604 | .00 | 1.829 |
| City residence | 0.511 | .00 | 1.667 |
| Child at home aged 0–5 | –0.181 | .01 | 0.834 |
| Child at home aged 6–17 | 0.300 | .46 | 1.350 |
| Age difference is seven or more years | 0.503 | .00 | 1.654 |
| Constant | –3.444 | .00 | 0.032 |
| Model Chi-Square | | 4,200 | |
| *df* | | 20 | |
| Number of Person-Years | | 58,283 | |
| Number of Dissolutions | | 4,214 | |

*Note:* All models include a control for the sex of the respondent.

effect of prior cohabitation on the odds of marital dissolution. Some earlier studies found a destabilizing effect of prior cohabiting experience, an effect that is often attributed to the cultural differences between respondents who enter marriage directly and those who enter via cohabitation (Hall and Zhao 1995; Manting 1994a). More recently, however, studies have shown that the effect of premarital cohabitation disappears when the duration of the union is modeled correctly (Brüderl, Diekmann, and Engelhardt 1997; Brüderl and Kalter 2001). More specifically, Brüderl and his colleagues found that when the premarriage years are included in the duration effect, couples who cohabited before marriage were no more likely to divorce than couples who married directly without cohabiting first.

We further see that same-sex unions are more unstable than opposite-sex unions. The coding relative to married couples is cumulative since same-sex couples also have a score of 1 on the cohabitation variable (only a handful of the same-sex couples were married). This means that same-sex couples have 3.1 times higher dissolution odds than opposite-sex *cohabiting* couples and $3.66 \times 3.15 = 11.5$ times higher dissolution odds than *married* couples. Note that same-sex couples are limited to couples in which the respondent was 30 or more years old at the start of the union. Because this age limit may also affect the dissolution risk for opposite-sex couples, we estimated the model again, also excluding opposite-sex couples in which the respondent was younger than 30 at the start of the union. The effect of same-sex couples in this model is 1.054, which is very close to the effect shown in Table 2.

Effects of the control variables are generally as expected. Higher dissolution risks are more common among couples who married young, who live in cities, who have no young children at home, who are first- or second-generation immigrants, and who have large age differences. More detailed analyses show that for opposite-sex couples, the effect of age differences is asymmetric. Age differences of seven years or more are more unstable than smaller age differences, but this effect is stronger when the female partner is older ($b = 1.104, p < .01$) than when the male partner is older ($b = 0.410, p < .01$).[5] In the models in which same-sex unions are not included, we control for age differences by including these two dummy variables (see Tables 3 and 4).

**Models with income variables.** In Table 3, we add income variables to the baseline model. We evaluate the alternative specifications with the Bayesian Information Criterion (BIC) measure because the models are not nested (Raftery 1996).[6] The more negative the BIC, the better the model. We first notice that in all models, there is a negative effect of household income. In the first model, the effect is –0.036, which means that for every 1,000€ (≈ USD$1,250) increase in disposable annual household income, the odds of separation decline by about 4%. When moving from a median level of income (for couples) to the welfare level, this implies nearly a doubling of the odds of separation, which is a strong effect.[7] The direction of the effect is in line with the literature, which has attributed union dissolution to the financial problems and resulting stress that may arise in lower-income families (Ross and Sawhill 1975; Voydanoff 1990).

Model 1 includes a linear variable for the female partner's income share, which shows a positive and significant effect. The higher the income share of the female partner, the higher the risk of separation. Model 2 includes a quadratic effect and is a significant improvement over Model 1. The BIC decreases, and the quadratic term is statistically significant. The pattern is U-shaped, and the minimum lies at a proportion of .25, showing that the effect is not symmetric. Model 3 replaces the quadratic specification with spline functions. The first spline coefficient is the effect of the female partner's income share for shares over 50%; the second spline coefficient is the effect of the female partner's income share for shares lower

5. These effects were obtained from an additional model that is not included in the table.

6. BIC = $-\chi^2 + df \ln(N)$, where $\chi^2$ is the likelihood ratio test for comparing the model to the null model.

7. The median annual disposable household income for couple households in 2000 was 29,500€. The level of welfare for couples is 12,000€ per year. Hence, the implied change in the odds is $e^{-0.036 \times (12 - 29.5)}$.

**Table 3. Discrete-Time Event-History Model of Union Dissolution With Income Variables**

| | Model 1 | | Model 2 | | Model 3 | | Model 4 | |
|---|---|---|---|---|---|---|---|---|
| Variable | *b* | *p* | *b* | *p* | *b* | *p* | *b* | *p* |
| Type of Union | | | | | | | | |
| Cohabiting (vs. married directly) | 1.358 | .00 | 1.376 | .00 | 1.384 | .00 | 1.376 | .00 |
| Married after cohabiting (vs. married directly) | –0.017 | .84 | –0.005 | .95 | –0.001 | .99 | –0.006 | .95 |
| Income Variables | | | | | | | | |
| Total household income | –0.036 | .00 | –0.033 | .00 | –0.031 | .00 | –0.032 | .00 |
| Female partner's income proportion | 0.379 | .00 | –0.486 | .07 | | | | |
| Proportion squared | | | 0.977 | .00 | | | | |
| Spline: proportion – .5 (for proportion > .5) | | | | | 1.313 | .00 | | |
| Spline: .5 – proportion (for proportion < .5) | | | | | 0.185 | .20 | | |
| Proportion 0 (reference group) | | | | | | | — | |
| Proportion 0–.2 | | | | | | | 0.379 | .00 |
| Proportion .2–.4 | | | | | | | 0.119 | .38 |
| Proportion .4–.6 | | | | | | | 0.088 | .82 |
| Proportion .6–.8 | | | | | | | 0.643 | .00 |
| Proportion > .8 | | | | | | | 0.426 | .00 |
| Constant | –3.550 | .00 | –3.399 | .00 | –3.465 | .00 | –3.560 | .00 |
| Model Chi-Square | 3,398 | | 3,410 | | 3,424 | | 3,460 | |
| *df* | 22 | | 23 | | 23 | | 26 | |
| BIC | –3,157 | | –3,158 | | –3,172 | | –3,175 | |
| Number of Person-Years | 56,707 | | 56,707 | | 56,707 | | 56,707 | |
| Number of Dissolutions | 3,670 | | 3,670 | | 3,670 | | 3,670 | |

*Notes:* Models include duration, age at start union, immigrant status, city residence, age differences, sex, and children. Sample excludes same-sex couples. Intercept is calculated after centering all variables except income share, cohabitation, and duration.

than 50%. The former effect reflects what happens when couples move away from equality in the direction in which the female partner has more income; the latter effect reflects what happens when couples move away from equality in the "traditional" direction. Two positive spline effects imply a V-shaped effect.

Based on the BIC statistic shown in Table 3, Model 3 is better than the null model and the quadratic model. Graphic inspection (not shown) shows that the implied forms of the relationship between separation and women's relative income under Models 2 and 3 are rather similar. Both spline coefficients are positive, showing that the form is V-shaped. However, the effect of the female partner's income share below equality is much smaller than the effect of the female partner's income share above equality. Hence, gender inequality in income increases the risk of separation when the female partner has more income, but the effect is much weaker when the female partner has less income than the male partner.

The final model (Model 4) is a more exploratory model and includes dummy variables for separate categories of income shares. The BIC suggests that this model is only a modest improvement in fit over the spline function model. One exception that comes out of this model is the contrast between women who have no income at all versus women who have

**Table 4. Discrete-Time Event-History Model of Union Dissolution With Interaction Effects**

| | Model 1 | | Model 2 | | Model 3 | | Model 4 | |
|---|---|---|---|---|---|---|---|---|
| Variable | *b* | *p* | *b* | *p* | *b* | *p* | *b* | *p* |
| Long Duration | –0.125 | .13 | –0.093 | .26 | 0.206 | .06 | 0.116 | .41 |
| Cohabiting | 1.570 | .00 | 1.463 | .00 | 1.558 | .00 | 1.525 | .00 |
| Cohabiting × long duration | –0.703 | .00 | –0.750 | .00 | –0.869 | .00 | –0.665 | .00 |
| Household Income | –0.033 | .00 | –0.039 | .00 | –0.041 | .00 | –0.043 | .00 |
| Female Partner's Proportion – .5 (> .5) | 1.346 | .00 | 1.069 | .01 | 1.238 | .01 | 1.099 | .02 |
| .5 – Female partner's Proportion (< .5) | 0.166 | .25 | –0.427 | .09 | 0.114 | .69 | –0.094 | .76 |
| Cohabiting × Household Income | | | 0.008 | .15 | 0.009 | .11 | 0.012 | .06 |
| Female partner's proportion – .5 (> .5) | | | 0.388 | .42 | 0.246 | .62 | 0.412 | .44 |
| .5 – Female partner's proportion (< .5) | | | 0.888 | .00 | 0.458 | .15 | 0.733 | .04 |
| Long Duration × Household Income | | | | | 0.006 | .38 | 0.012 | .19 |
| Female partner's proportion – .5 (> .5) | | | | | –0.337 | .63 | –0.061 | .95 |
| .5 – Female partner's proportion (< .5) | | | | | –1.546 | .00 | –0.981 | .06 |
| Cohabiting × Long Duration × Household Income | | | | | | | –0.013 | .33 |
| Female partner's proportion – .5 (> .5) | | | | | | | –0.647 | .64 |
| .5 – Female partner's proportion (< .5) | | | | | | | –1.606 | .07 |
| Constant | –3.808 | .00 | –3.724 | .00 | –3.818 | .00 | –3.799 | .00 |
| Model Chi-Square | 3,398 | | 3,407 | | 3,427 | | 3,430 | |
| *df* | 15 | | 18 | | 21 | | 24 | |
| Change in Chi-Square | | | 9.2 | | 19.9 | | 3.7 | |
| *p* Value of Change | | | 0.03 | | 0.00 | | 0.30 | |
| Number of Person-Years | 56,707 | | 56,707 | | 56,707 | | 56,707 | |

*Notes:* Models include duration, age at start union, immigrant status, city residence, age differences, sex, children. Intercept is calculated after centering all variables except wife's income share, cohabitation, and duration. Sample excludes same-sex couples.

some income. The effects suggest that women with a small income have a higher separation risk than women who have no income, whereas there is no difference between women with no income and women with intermediate income. Another finding from Model 4 is that couples with reverse specialization—in which women have most of the income—are not more stable than couples with a more or less equal income division. This goes against specialization theory, which suggests that specialization is beneficial regardless of which partner is earning more.

**Interactions with duration and union type.** To examine interactions with union type, we continue with the spline model, which fits slightly better than the quadratic model and not much worse than the more complex model with dummy variables. This model is also more flexible in that we can test for differences in slopes below and above equality. To simplify the interaction terms, we dichotomize duration into less than five years and five years or more. We think five years is a reasonable cutoff point because the real weeding years for cohabiting couples are in the first five years (Figure 2). Because of our design, however, most of the unions are relatively young. Hence, the number of dissolutions in the second interval is smaller (431) than the number of dissolutions in the first interval (3,239). This will limit our statistical power to discern interaction effects with duration and makes it difficult to further increase the cutoff point to six or seven years. The results are presented in Table 4. Because the models are nested, we use chi-square tests to compare the improvement in fit of the models.[8]

The first model contains only an interaction of duration and union type. This interaction effect is significant and negative, confirming the convergence of dissolution risks over the course of the union. The effect of cohabitation (i.e., the difference between cohabitation and marriage) is 1.570 in the first five years and 1.570 – 0.703 = 0.867 in the second five years.

Model 2 adds interactions between income variables and union type. The addition of these variables improves the fit of the model significantly according to the change in chi-square. The effect of the female partner's share for income levels above 50% does not interact with union type, whereas the effect for income levels below 50% interacts significantly. In other words, the effect of the female partner's income share below equality is significantly different for marriage and cohabitation. More specifically, for marriage, the implied effect of the wife's proportion is negative ($b = -0.427$), whereas for cohabitation, the implied effect is positive ($b = -0.427 + 0.888 = .461$). This means that moving away from equality toward a pattern of male dominance *decreases* the dissolution risk for marriage, whereas it *increases* the dissolution risk for cohabitation. In other words, inequality is destabilizing for cohabitation, much in contrast to what we see for marriage.

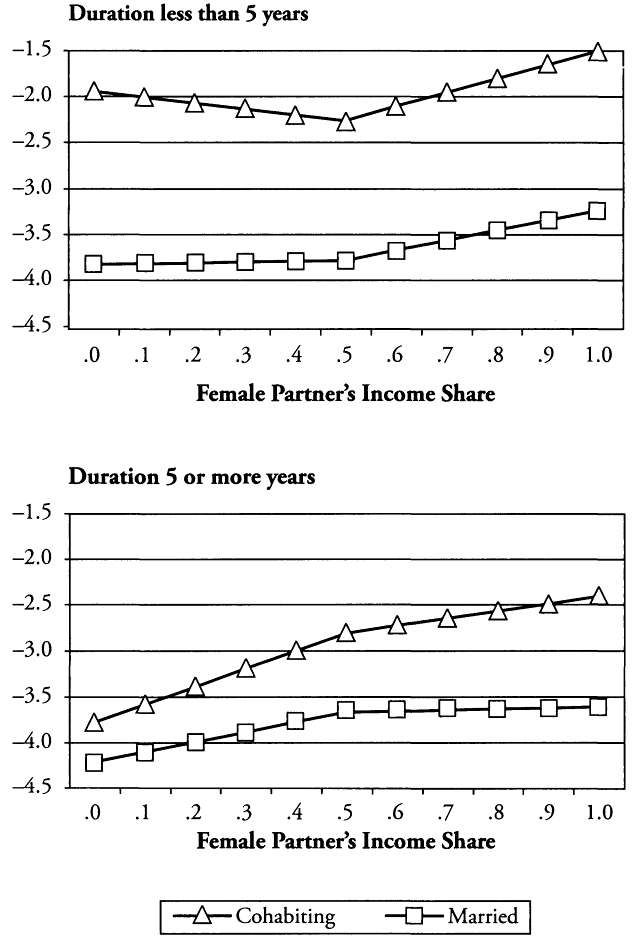
Model 3 adds interactions between income variables and relationship duration. The chi-square test for comparing nested models shows that this is a significant improvement in the fit of the model. The interaction effects show that the effect of the female partner's share below 50% interacts significantly with duration. The effect of the female partner's share above 50% does not interact with duration. The implied effects for marriages of short durations are 0.114 (below equality) and 1.238 (above equality). For marriages of longer durations, the effects are 0.114 – 1.546 = –1.432 (below equality) and 1.238 – 0.337 = 0.901 (above equality). In other words, in the short term, the effect is more or less flat below equality and increasing above equality. For longer durations, the spline coefficients have opposite signs so that the pattern becomes linear, with increasing shares of the female partner increasing the risk of separation across the full range of the possible income shares.

In Model 4, we add the three-way interaction between duration, income measures, and union type. This does not yield a significant improvement in fit. We do, however, see several (marginally) significant interaction effects. In line with Model 2, the effect of the female partner's share below 50% is more positive for cohabiting relationships. The interaction is $b = 0.733$ and statistically significant ($p = .04$). This interaction effect is consistent with the hypothesis that equality is more favorable for cohabiting couples. However, the interaction effect is reduced at longer durations ($p = .07$). The resulting interaction effect at longer durations is opposite: $b = 0.733 - 1.606 = -0.873$. Hence, the evidence for the hypothesis that equality is more favorable for cohabiting couples than for married couples applies only to the first years of the union.

To see more clearly what the interactions imply, we plot expected log odds of Model 4 in Figure 4. The figure shows how the conditional log odds of separation depend on the female partner's income share for each of the two types of relationships and the two durations. The control variables are set at their means. For short durations, the cultural hypothesis is confirmed. The pattern follows a model of equality for cohabiting couples and a model of specialization for married couples. In other words, in married unions, higher

8. We also considered interaction effects with premarital cohabitation, but these did not turn out to be statistically significant.

**Figure 4. Female Partner's Relative Income Effects on the Log Odds of Dissolution and Cohabiting Status**




income shares of the female partner increase the risk of divorce. This effect, however, is not symmetric—it is stronger when women have more than half of the income. In cohabiting unions, movements away from equality in either direction increase the risk of dissolution. The effect is not fully symmetric in cohabiting couples either, but it comes closer to a V-shaped pattern than the flatter effect observed for married couples. The difference in the slope of the line below equality for married and cohabiting couples is significant ($b$ = 0.733, $p$ = .04; Table 4).

For longer durations, the pattern becomes more "traditional" for both types of unions. For marriages, we see less asymmetry in slopes before and after the 50% point. More important, the pattern for cohabiting couples begins to resemble the pattern for married couples. The V-shaped pattern for cohabiting couples at short durations becomes a more or less linear pattern for cohabiting couples at longer durations, just like the pattern for married couples.

Finally, we assess how income shares affect the dissolution risk of same-sex cohabiting couples (Table 5). We define a variable as 1 minus the absolute difference in income of the two partners divided by the total income. This measure ranges from 0, for the case in which one partner brings in all the income, to 1, for couples in which partners have the same income. The effect of this variable on dissolution is negative but not statistically significant (Model 1). A closer inspection of the income patterns shows that a sizable group of couples are single earners (11%). Because this group may deviate from the rest, we add a dummy variable for this group. The results change when we add this variable (Model 2). The effect of the single-earner variable is negative and marginally significant, whereas the effect of equality becomes significant ($b = -0.722$, $p = .02$). This shows that single-earner couples are more stable than other couples with a low level of equality in income distribution (an equality score of near 0). Apart from this effect, higher levels of equality reduce the risk of dissolution. In the last model (Model 3), we delete the single-earner couples from the sample to check this result. This model confirms the negative and significant effect of income equality. These findings are in line with the hypothesis that equality is also a pattern that is stabilizing for same-sex couples.

Several of the control variables have an effect as well. First, same-sex couples living in urban areas have a higher dissolution risk than other couples, in line with the results for opposite-sex couples. Second, household income has a negative effect, as expected, and in line with the results for opposite-sex couples. Third, female couples are more stable than male couples. Finally, age differences do not increase the dissolution risk for same-sex couples, in contrast to the results for opposite-sex couples.

**Table 5. Discrete-Time Event-History Model of Union Dissolution for Same-Sex Couples**

| | Model 1 | | Model 2 | | Model 3 | |
|---|---|---|---|---|---|---|
| Variable | *b* | *p* | *b* | *p* | *b* | *p* |
| Duration 2 Years (vs. 1 year) | –0.543 | .00 | –0.535 | .00 | –0.651 | .00 |
| Duration 3 Years (vs. 1 year) | –0.922 | .00 | –0.920 | .00 | –0.912 | .00 |
| Duration 4 or More Years (vs. 1 year) | –2.107 | .00 | –2.109 | .00 | –2.127 | .00 |
| First- and Second-Generation Immigrant | 0.142 | .28 | 0.145 | .27 | 0.150 | .30 |
| City Residence | 0.323 | .01 | 0.320 | .02 | 0.363 | .01 |
| Two Women (vs. two men) | –0.351 | .01 | –0.334 | .01 | –0.294 | .04 |
| Age Difference Is Seven or More Years | –0.107 | .38 | –0.114 | .35 | –0.042 | .75 |
| Household Income | –0.035 | .00 | –0.037 | .00 | –0.038 | .00 |
| Relative Equality of Partners' Income | –0.302 | .13 | –0.722 | .02 | –0.705 | .02 |
| Single Earner | | | –0.528 | .08 | | |
| Constant | –0.010 | .96 | 0.330 | .23 | 0.271 | .34 |
| Model Chi-Square | 315 | | 319 | | 272 | |
| *df* | 9 | | 9 | | 9 | |
| Number of Person-Years | 1,555 | | 1,455 | | 1,378 | |
| Number of Dissolutions | 535 | | 535 | | 440 | |

*Note:* Model 3 excludes single-earner couples.

## DISCUSSION AND CONCLUSION

In this article, we have examined the effect of women's relative income on the risk of separation in detail, using a large and reliable longitudinal data set from The Netherlands. Our first finding is that there is a moderately positive effect of the female partner's relative income on separation. The higher the female's share of the household income, the higher the risk of separation. Although the general tendency of the effect is positive, the shape of the effect depends on the legal status of the relationship. Specifically, we find that the effect of the woman's income has a more or less continuous form for marriages, whereas it is more V-shaped for cohabiting unions. For income shares above equality—in which women have more income than men—the effects are the same: higher income shares of the female partner are associated with higher dissolution risks for both marriage and cohabitation. For income shares below equality, which is the most common range, the effects are different: higher shares of the husband reduce the divorce risk for marriages (although only weakly), but higher shares of the male partner *increase* the dissolution risk for cohabiting couples.

What do these patterns tell us about the different theoretical approaches? Our findings are more in line with the cultural approach than with the economic approach. According to the cultural approach, male dominance is stabilizing for marriages because it concurs with traditional gender values, whereas male dominance is destabilizing for cohabiting unions because it conflicts with preferences for gender equality. Economic theory would predict a stabilizing effect for both. Moreover, economic theory predicts that deviations away from equality toward female dominance are stabilizing (reverse specialization), but we find that female dominance is destabilizing. This, too, is in line with the cultural approach. Female dominance is at odds with both traditional gender values in marriage and notions of equality in cohabitation. Note that these conclusions obviously rest on the assumption that there are important (gender) value differences between married and cohabiting couples. While this assumption is plausible, it cannot be tested with register data. Register data are statistically powerful and have reliable measures, but they measure a limited number of concepts and do not include direct measures of cultural and social characteristics.

How do our results compare with earlier findings? The number of studies analyzing dynamic income effects is not large, and most studies have focused on the effects of women's labor force participation. Our evidence is consistent with the European study by Jalovaara (2003), who found that higher income levels of wives are associated with higher divorce risks when husbands' income levels are controlled for. Our evidence is also consistent with the analysis of U.S. data from the Panel Study of Income Dynamics by Heckert, Nowak, and Snyder (1998), who found a general positive effect of the wife's income share on the probability of divorce. Hecker et al. also found a different pattern for couples in which wives earn most of the income, but this deviation was not statistically significant due to the small number of cases with such an unusual pattern. Our data set is much larger and does not show that these cases of reverse specialization are significantly different. Our findings are less consistent with Rogers (2004), who found an inverted U-shaped effect. Even in Rogers' analysis, however, the positive effect of wife's income share dominated, which is consistent with our results.

Few authors have yet investigated how relative income effects differ between married and cohabiting couples. The most important exception can be found in the U.S. analysis of Brines and Joyner (1999). Although Brines and Joyner used a smaller sample of cohabiting relationships than we do, the general pattern of effects that they found is the same as in our work. We therefore concur with Brines and Joyner that the principles of stability in personal relationships are conditional rather than universal. Different relationships have different ideals and expectations, and this results in differential effects of income arrangements on their stability.

Although the differences between married and cohabiting couples are significant, additional analyses also suggest that the differences are especially marked at earlier durations. More important, we find that the effect of the female partner's income share in cohabiting unions becomes stronger and more linear—more "traditional"—as the union progresses. This latter finding is unexpected and makes our evidence for the interaction hypothesis weaker than it is in the United States. The interaction with duration itself was found earlier, however. For marriages, South (2001) found an increase in the effect of the wife's working hours on divorce with marital duration. One explanation of this interaction is that the negative side effects of an equal division of labor—for instance, the burden that women experience when doing both domestic and paid work—become more apparent over time.

Another new finding lies in our analyses of same-sex couples. We found that more-equal income shares in same-sex couples were associated with a lower risk of dissolution. This finding again shows that the theory of specialization does not hold for all types of couples. The norm of equality appears to be valid not only for opposite-sex cohabiting couples, but also for same-sex couples.

Finally, our work has shown that the evidence that was found for the United States can in part be generalized to a different setting. Cohabitation is more common and much more accepted in The Netherlands than in the United States. Thus, one would expect the evidence for the hypothesis to be weaker in The Netherlands than in the United States, but this does not turn out to be the case. Obviously, we need evidence from more than two countries to examine such speculations. Our results call for more systematic cross-national research in which differences between cohabitation and marriage are studied. For example, it is important to examine how differences in sex-role attitudes between married and cohabiting couples vary across countries with different levels of cohabitation. Similarly, more systematic cross-national research is needed on the stability of cohabiting unions.

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