European Journal of Population (2019) 35:825–849
https://doi.org/10.1007/s10680-018-9506-0

Regional Sex Ratio and the Dissolution of Relationships in Germany

Monika Obersneider[1] · **Jan-Christoph Janssen**[2] · **Michael Wagner**[2]

Received: 19 February 2017 / Accepted: 14 November 2018 / Published online: 28 November 2018

Abstract

The aim of this study is to analyse the connection between relationship stability and attractive alternatives, which is stressed in micro-level theories on union dissolution. The stability of relationships can be influenced by the availability of alternative partners, whereby the probability that a person will meet these alternatives is determined by the distribution of individuals with specific characteristics in the contextual setting the person is embedded in. Research on this macro–micro connection is sparse in Europe. The availability of alternatives on the contextual level is operationalised through varying sex ratios between and within German districts. The estimation of the union dissolution risk as a function of individual and contextual predictors is based on a discrete-time multilevel event-history analysis using *pairfam* data and data from official statistics. The main hypothesis, which asserts that there is a positive connection between unbalanced sex ratios and union dissolution, is not supported. This result calls into question the robustness of previous findings.

Keywords Sex ratio · Union dissolution · Divorce · Partner market · Multilevel event-history · *pairfam*

✉ Monika Obersneider
monika.obersneider@uni-due.de

Jan-Christoph Janssen
janssen@wiso.uni-koeln.de

Michael Wagner
mwagner@wiso.uni-koeln.de

[1] University of Duisburg-Essen, Duisburg, Germany

[2] University of Cologne, Cologne, Germany

1 Introduction

According to theories of partnership dissolution, the availability of attractive alternatives poses a threat to a couple's current relationship. The destabilising effect of having access to attractive alternatives is stressed in both major theories on union dissolution—exchange and family economic theory (Becker et al. 1977; Lewis and Spanier 1979). The existence of alternatives can jeopardise a couple's current relationship in different ways: by enabling the partners to be unfaithful, by leading the partners to consider the higher utility of a concrete alternative, or by encouraging the partners' subjective perception that a large number of attractive alternatives are available (South et al. 2001: 745). Thus, the stability of relationships can be influenced by the availability of alternative partners, whereby the probability that a person will meet these alternatives is determined by the distribution of individuals with specific characteristics in the contextual setting the person is embedded in (Blau 1977a). Hence, there is a theoretical reason to expect that there is a connection between the availability of alternative partners on the contextual level and union dissolutions on the individual level.

The contextual environment can impose restrictions on the dissolution of the relationship. Living in a community with strict divorce laws, conservative values, or a highly religious moral code can act as barriers to divorce, as a person considering separating is likely to anticipate negative consequences (Lewis and Spanier 1979: 287). However, the social context may also jeopardise the stability of the relationship by, for example, providing favourable contextual conditions for meeting new partners (Lyngstad 2011). Sex ratios in geographically distinct regions can be seen as contextual conditions that influence the availability of alternative partners in those places. These varying regional sex ratios may influence the risk of union dissolution (Klein and Stauder 2008: 78).

The examination of sex ratios in a wide range of sociological studies has provided valuable insights. These studies have, for example, looked at the role of sex ratios in assortative matching (Lichter et al. 1995; Abramitzky et al. 2011), women's roles in society (Guttentag and Secord 1983; South and Trent 1988), entering into marriage (Lichter et al. 1991; Lloyd and South 1996; Martin 2001), partner search (Klein and Stauder 2008; Uecker and Regnerus 2010), fertility (Schoen 1985), and remaining single (Lengerer 2011; Eckhard 2014). There are, however, only a handful of articles that focused on district-level sex ratios and their connection with individual dissolution risks (e.g. South 1995; South and Lloyd 1995; South et al. 2001; McKinnish 2004; Lyngstad 2011). Moreover, the main findings of these studies seem to differ based on the units of measurement and the method of analysis used. In the German context in particular, there is very little existing research on local sex ratios and their potential impact on relationship stability (Stauder 2015: 429; Wagner et al. 2015: 224); though there is evidence of significant differences in the divorce rates of the districts in Germany (BBSR 2014). The key research question of this study is whether district-level sex ratios are associated with relationship stability in Germany.

When investigating data using both individual- and contextual-level variables, the multilevel structure of dependencies in the data must be accounted for (Hox 2002: 3ff.). This factor has been largely neglected in the previous contributions that estimated sex ratios on the district level (see an exception: Lyngstad 2011). Thus, in the present study, we have chosen to use multilevel event-history analysis as an instrument for generating additional findings in the research area of separation and divorce.

The present paper is structured as follows. First, we discuss the current research on the relationship between sex ratios and various partner markets. Second, we explore the complex theoretical basis for the assumption that the partner market affects relationship stability and derive our hypotheses from the existing theoretical and empirical findings on this potential association. In the subsequent sections, we present the data sources, statistical method, and operationalisation approach we use in our analysis. Finally, we provide an empirical evaluation of the data, a discussion of the results, and, in the conclusion, suggestions for further analysis.

2 Previous Research: Sex Ratio and Union Dissolution

Since the beginning of 1990s, research on sex ratios and partner markets has been broken down into regional entities (South 1995; South and Lloyd 1995; South et al. 2001; Trent and South 2003; McKinnish 2004; Lyngstad 2011; Eckhard and Stauder 2016). This shift to a regional focus was made primarily because it was assumed that a balanced sex ratio at the national level does not indicate whether the real availability of alternative partners is equally balanced across industrial, rural, and urban areas. In a given country, different regions are likely to have different sex, age, educational, or religious affiliation structures; which may in turn affect the availability of attractive alternatives (Lyngstad 2011). Lengerer (2001: 142) emphasised the relevance of analysing local partner market opportunities, as partner-finding processes mainly take place within rather small-scale local entities.[1]

One of the most cited studies with regard to sex ratio and marriage stability is by South and Lloyd (1995), who examined labour market areas as regional units. This study is one of the first to use official statistics on the district level to measure the availability of alternative partners (for an application of subjective measures, see, for example, Udry 1981). The authors concluded that the "relative supply" of alternative partners has a U-shaped connection with the divorce rate. Specifically, they found that in regions with low sex ratios (more single women than single men) or high sex ratios (more single men than single women), the divorce rate is higher than in regions with balanced sex ratios.[2] Thus, the study's findings are in line with the

[1] According to their regional study, approx. 85% of couples in Germany meet in a radius of 20 km. The formation of a relationship in a radius of over 100 kilometres is rare (2–5% of the relationships; Lengerer 2001: 142).

[2] The effect remains significant after including various individual- and contextual-level variables. Female labour force participation and geographic mobility are likely to increase the divorce risk, which is not the case for the urban–rural variable (South and Lloyd 1995: 31; South 1995).

assumptions of family economics (Becker et al. 1977) and show—unlike, for example, Guttentag and Secord (1983)—an effect of regional opportunity structures for both sexes. It should be noted that only unmarried people were used in the study to define sex ratios and that the study sample consisted of relationships in their early stages. Since relationships are more fragile at the beginning (Esser 2000: 324ff.), the effect of sex ratios found in South and Lloyd may be overestimated and is not generalisable to all married couples.

South et al. (2001) and McKinnish (2004) also found support for the U-shaped relationship between regional sex ratio and divorce risk in the US. South et al. (2001) included "classical" variables from divorce studies, such as children, marriage duration, and age at marriage. They found that the effect of sex ratios remained significant, but had no interaction effects with the above-named micro-level variables. Moreover, because the study neglected the multilevel structure of the data, the standard errors may have been underestimated (Hox 2002: 3ff.). This risk of generating biased results was tackled by Lyngstad (2011). He is one of the pioneers in the development of a correct model specification for divorce studies investigating macro–micro linkages. He estimated multilevel models with time-variant contextual variables, while including regional dummy variables into the model to account for unobservable time-constant characteristics of the different regions (see also Kulu 2012).

Contrary to theories on union dissolution and the existing regional-level results, Lyngstad (2011) showed using register data from Norway that unbalanced sex ratios have a small negative effect on the divorce risk. This means that the high availability of alternative partners seems to stabilise marriages on the micro-level. Lyngstad (2011) reasoned that this stabilising effect might be explained by the higher levels of commitment and investments of the partner in a couple whose partner market opportunities are disadvantageous, which in turn prevents a dissolution ("commitment mechanism"; Lyngstad 2011: 71). This finding emphasises the importance of applying methods that can appropriately test the empirical connection between relationship stability and the local sex ratio in other settings.

We are not aware of any other empirical studies that have investigated the connection between district-level sex ratios and the dissolution of relationships in Germany. While a cross-sectional study by Stauder (2011) and a longitudinal study by Eckhard and Stauder (2016) were indeed focused on the district level, they used a more differentiated concept of "availability ratio"[3] than the aforementioned studies to measure features of the partner market. The results derived from their aggregated data analyses are in line with Lyngstad's (2011) controversial findings: i.e. an unbalanced district sex ratio is associated with a lower number of divorces on the district level. However, these studies were exposed to the ecological fallacy because they did not control for individual risk factors. Moreover, because they relied on district-level divorce rates, these studies captured legal marriages only. Hence, there is a gap

[3] The availability ratio is computed by contrasting a weighting factor of the potential partners with the weighting factor of competing individuals in one district (Eckhard and Stauder 2016: 126f.; Goldman et al. 1984).

in the research on regional-level sex ratios and their potential connection with individual separation rates in Germany.

3 Theoretical Considerations

The theoretical frame of this study follows a multilevel perspective. Our focus is on estimating the potential link between partnership stability on the micro-level and a socio-structural characteristic on the contextual level. Such macro-micro linkages are not new in sociology and are often used to explain the diverging outcomes of a dependent variable in different regions (e.g. Coleman 1986; Esser 1988). Hence, in this study, we try to build a bridge between the micro- and macro-theories on union dissolution and to offer concrete explanations of the mechanism of interest (South et al. 2001).

The micro-sociological exchange theory postulates the importance of the cost-benefit ratio of exchanging resources in (intimate) relationships. Hence, the stability of relationships depends on the relative benefits for both partners; i.e. on the exchange of immaterial and material resources (Hill and Kopp 2006: 108). It has, however, been hypothesised that the factors that influence marital stability are not just the partners' resources on the micro-level (*intradyadic factors*), but two other external factors (*extradyadic factors*): namely (1) the availability of attractive alternatives outside the relationship; and (2) the societal barriers to leaving the relationship (*e.g. strict divorce laws or a highly religious context;* Lewis and Spanier 1979: 286). Lewis and Spanier summarised the influence of these factors as follows (1979: 278): "[…] marriages of high quality tend to have high stability. This relationship is mitigated at times by more attractive alternatives, but conversely, it may be strengthened by external pressures to remain married, such as normative or institutional constraints". It therefore appears that high barriers to relationship dissolution could discourage a spontaneous decision to end a relationship due to attractive alternatives. In sum, it is necessary for theoretical reasons to include extradyadic factors into models of relationship stability, and specifically the opportunities on the partner market and the potential barriers to relationship dissolution on the societal level.

Another rational-choice theory on dissolution of relationships can be found in the realm of family economics (Becker et al. 1977: 1154). Following this approach, we can assume that relationships come to an end when the total gains from being in the current relationship are smaller than the anticipated benefits from being in alternative relationships or being single. Accordingly, when a new partner with more suitable characteristics—i.e. a partner who offers greater anticipated gains to the household—is available, the dissolution of a relationship is more likely (Becker et al. 1977: 1153).

Blau's macro-sociological structural theory (1977a, 1994) is a milestone for the contextual approach and is used here as a framework to analyse dissolution risks with respect to contextual factors. Blau conceptualised the social structure as the distribution of individuals in a population according to their social positions in a "multidimensional space of positions" (Blau 1977b: 28). He argued that population distribution "[…] exerts independent effects on social relations by circumscribing

the opportunities and limiting the choices in a population" (Blau 1994: 28). The closer individuals are positioned to each other in Blau's "multidimensional space", the more likely they are to interact and form of a relationship with each other. Hence, it is assumed that individuals with similar characteristics in terms of age, religion, and education are more likely to have similar life stories, roles, and experiences and thus to have higher chances of forming long-term relationships (Stauder 2008: 268; Blau 1977a).

The stability of relationships can be influenced by the availability of alternative partners, whereby the probability that a person will meet these partners is in turn a product of the social structure (Blau 1977a; see also South et al. 2001). Blau operationalises these structural frames through spatial and relational quantitative measures, such as the ratio of individuals in different positions or relational group sizes (Blau 1977b: 33). Hence, the chances of meeting attractive alternative partners depend on the distribution of the sexes (sex ratio) on a contextual level (Blau 1977a, b; Klein and Stauder 2008: 77, Stauder 2015: 404). For example, the chance that a man will meet a potential female partner is smaller on a partner market with more men than women than in a partner market with the reverse sex ratio.[4] In their contribution to the macrostructural opportunity theory of marital dissolution, South et al. (2001) stressed the importance of the sex ratio in different partner market settings: "The population sex distribution (and, to a lesser extent, race and age distributions) determines the availability of spousal alternatives. The relevant 'population' may be at the societal level, or at some subunit, such as the local marriage market or the workplace" (South et al. 2001: 745). Blau (1977a, b, 1994) emphasised the influence of the *social structure* on individual decision-making processes. Other potential factors of social interaction, like cultural characteristics, social norms, or institutional structures, were excluded from Blau's conceptualisation. Furthermore, the influence of individual social networks is not examined. Those are major points of criticism of Blau's narrow definition of social structures. However, the possible influence of partner market subunits does not diminish the importance of measuring the social structure on the regional level when estimating the potential contextual effect on union dissolution. In line with this theoretical focus, micro-level partner markets, such as individual friendship networks ("foci" see Feld 1981), are excluded from the following contextual analysis.

The contextual classification of Germany into districts and cities without districts (*Kreise und kreisfreie Städte*) is in line with Blau's macrostructural definition of available alternatives, yet they are even smaller units than labour market areas. The district units have been shown to correspond "to a high degree with the daily action radius of individuals" (Eckhard and Stauder 2016: 122; translated by the authors).[5]

[4] It is important to point out that the potential effect of sex ratios is stochastic: it is not assumed that individuals are perfectly informed about sex distributions in their environment and partner market opportunities. What matters is that sex ratios influence a person's real chances of meeting an (alternative) partner in his or her everyday life (Guttentag and Secord 1983: 162).

[5] Furthermore, Stauder (2015) found that the social structure of districts shapes micro-level partner markets in Germany. Hence, the district-level seems to be an appropriate regional unit to investigate the association between regional sex ratio and individual-level union dissolution. Nevertheless, it must be pointed out that strong variations in intermediate-level sex ratios, i.e. in sex ratios within one district, are

Thus, we concentrate our analysis of the regional sex ratio on the level of German districts and district-free cities. Based on these research findings and theoretical considerations, we assume a U-shaped effect of district-level sex ratios for both sexes (H_1)*: Imbalanced district-level sex ratios are positively related to the risk of union dissolution.*

The connection between sex ratios and the probability of union dissolution may be moderated through interaction effects. For example, couples in cohabitating unions have a higher risk of separating than legally married couples (Boyle et al. 2008; Rupp and Blossfeld 2008: 151). The decision to dissolve a union can be influenced by the greater exit costs of ending a legal union rather than a cohabiting union (Kopp et al. 2010: 47; Hradil and Masson 2008: 213). It is therefore relevant to examine whether sex ratios show a different interrelation with co-residing married and unmarried partners. We assume that cohabitating partnerships are more fragile than marriages, and we expect to find an interaction effect between sex ratios and relationship type (H_2)*: The risk of union dissolution is more strongly connected with imbalanced sex ratios for cohabiting than for married couples.*

According to exchange theory, the relationship quality is the main factor that explains that diverging dissolution rates between individuals, in addition to the availability of attractive alternatives and societal barriers. Relationship quality is a set of subjective evaluations by both partners of, for example, marital happiness, communication, and conflict strategies (Lewis and Spanier 1979: 269). The utility from the exchange of resources in the current relationship is compared to the anticipated alternative utility, whereby it can be assumed that relationships with a lower relationship quality are more likely to be affected by the availability of alternative partners. Thus, we hypothesise first that a high level of relationship satisfaction has a negative effect on the risk of dissolution. Second, we expect to observe a cross-level interaction effect of sex ratios on the district level and of relationship satisfaction on the individual level (H_3): *The risk of union dissolution is less strongly connected with sex ratios in relationships with higher relationship satisfaction levels.*

Following the argumentation of family economic theory, a large number of studies have shown that relationship-specific investments have a stabilising effect on a relationship (Becker et al. 1977; Wagner and Weiß 2006). The value of these investments, such as having children together or owning a joint property, is linked to the persistence of a relationship, as when the partners separate, the time they can spend with their children is reduced, high transaction costs can arise, and their housing quality may decrease (Brüderl and Kalter 2001: 407f). Thus, in assessing the likelihood of dissolution, the extent of "marital-specific investments" (Becker et al. 1977: 1152) has the opposite effect of attractive alternatives (Kalter 1999: 256). These investments may hold a partnership together even if attractive alternatives have a high anticipated utility. According to Esser (2001: 123ff.), an underinvestment in relationship-specific capital promotes the awareness of alternatives and can lead to

Footnote 5 (continued)
likely to diminish the connection between the district-level and the individual-level risks of union dissolution.

a redefinition of the current partnership ("reframing"). We assume that higher relationship-specific investments in, for example, children and mutual property reduce the dissolution risk. We thus hypothesise an interaction effect as follows (H_4): *The risk of union dissolution is less strongly linked to the sex ratio for couples with higher relationship-specific investments.*

4 Data and Method

4.1 Data

The analyses are based on micro- and macro-datasets. For the micro-data, the first six waves of the German *pairfam* ("Panel Analysis of Intimate Relationships and Family Dynamics") are used (Release 6.0; Brüderl et al. 2015; detailed overview in Huinink et al. 2011). *Pairfam* is a panel study starting in the year 2008 with 12,402 anchors from three birth cohorts: 1971–1973, 1981–1983, and 1991–1993. The longitudinal data are available on an annual basis, which allows us to analyse partnerships prospectively over a period of 6 years. In addition to its emphasis on detailed relationship characteristics and their changes over time, a major advantage of the *pairfam* dataset is that it allows us to merge individual data with official statistics on the regional level. Accurate geographic municipality identifiers are provided for every anchor at every wave, which makes the linkage with regional characteristics possible (Schmiedeberg 2015: 3).

The regional statistics on the 402 German districts and cities without districts (*Kreise und kreisfreie Städte*)[6] are gathered from three official databases. First, the ratio of men to women is available in the German Regional Database, which is updated annually on the basis of the population census (Statistische Ämter des Bundes und der Länder 2016). Second, the INKAR database provides time-variant indicators, such as the female labour force participation rate on the district level (BBSR 2014). The third source is the national census 2011, which incorporates a time-constant measure of the regional differences in the proportions of the population who belong to different religious groups (Statistische Ämter des Bundes und der Länder 2014).

Although *pairfam* provides retrospective information on past relationships, important covariates (e.g. relationship satisfaction, norms) can only be analysed for relationships in the observation window. Due to this restriction, 14,101 of the initial 27,504 relationships in the *pairfam* dataset are included in the following analysis. Furthermore, recurrent events (here: anchors' multiple relationships) will be excluded from the analysis.[7] Moreover, due to the usage of the sex ratio, only

[6] Of those, 107 are classified as cities without districts and 295 are classified as districts. The average population in cities without districts is 242,598 individuals and approx. 1392 people per square kilometre. In rural districts, the average population is 187,255 individuals and about 204 people per square kilometre (Statistisches Bundesamt 2015).

[7] Only the first relationship in the observation window is analysed. It is likely that some people have shorter relationships and are more likely to change relationships—characteristics that are difficult to control for (Kulu 2012: 886; Allison 1984: 54). This unobserved heterogeneity would require the integration

heterosexual relationships are analysed in this study. Living apart together (LAT) relationships are also excluded from the analyses because the geographic municipality identifiers are not available for the LAT partner. After these exclusions, 10,792 relationships remain in our sample.

4.2 Statistical Method

Multilevel event-history analysis is used to capture both individual relationship dynamics and possible regional influences on dyadic relationship stability (Steele 2011). On the micro-level, each relationship is split up into a person-year format. Only when a relationship ends with a separation or divorce—i.e. an event takes place—this relationship year is coded to one; otherwise, a relationship is right-censored and remains in the risk set coded with zero (Allison 1984: 16). Right-censored relationships include unions that have not experienced a dissolution before the end of the observation window. While it would be possible to use time as a continuous variable, the regional data and many important covariates are gathered annually. A discrete-time hazard function is specified in this study (Singer and Willett 2003: 330). This allows us to estimate logit models with time-variant and time-invariant predictors. Furthermore, the binomial Bernoulli distribution can be generalised to hierarchical data structures, and can be analysed with multilevel models (Steele 2011: 4; Hox 2002: 107ff.).

Hierarchical models take the potential nesting of individuals (level 1) in their environment (level 2) into account. To the extent that the individual-level variation is due to clustering into higher-level units, individuals can no longer be regarded as independent from each other, and the use of multilevel models becomes mandatory (Hox 2002: 3ff.). If the regional differences in sex ratio have the hypothesised influence on the dyadic risk of union dissolution, both the coefficient estimates as well as the standard errors would be biased with traditional logistic regression. The multilevel approach corrects for standard errors and yields more robust coefficients that would otherwise be underestimated (Guo and Zhao 2000: 444ff.). In multilevel analysis, a distinction is made between fixed coefficients (effect parameters constant over all units of investigation) and varying coefficients (effect parameters varying across groups). In the first step, an empty model with varying intercepts is estimated. This allows us to determine the level-specific variance decomposition of the dependent variable. In the second step, the individual-level fixed coefficients are included in the model. In the next model specification, the fixed coefficients of the level 2 units on the variable of interest are added to the equation. In the last step, both variation in intercepts and variation in regression coefficients between the contexts is allowed (random-coefficient model—RC). In this model, predictors on both levels and their interactions can be controlled for (cross-level interactions, Snijders and Bosker 2012: 41ff.).

Footnote 7 (continued)
of another (here: third) level of analysis (Steele 2011), which leads to additional model restrictions and convergence problems (Snijders and Bosker 2012: 210).

In the analysis, controlling for regional-level variables assumed to correlate with the variable of interest and other covariates helps us to reduce unobserved heterogeneity on the district level (Lyngstad 2011: 64). As both time-variant and time-invariant regional variables are included in the analysis, we cannot control for all of the unobserved stable differences between the districts, as we could when using a fixed-effects approach (Allison 2009: 78ff.). Furthermore, there is very little to no within-district variation in sex ratios in the observed 6 years, which causes estimation and severe efficiency problems when estimating the variable of interest using fixed-effects modelling. Consequently, the possible violation in the model of the assumption of uncorrelated district-level error terms and regional variables must be kept in mind when interpreting the results. Nevertheless, this method is more likely to generate valid results than studies estimating effects of sex ratios using single-level event-history models (e.g. South et al. 2001; South and Lloyd 1995), aggregate data (e.g. South and Trent 1988; Eckhard and Stauder 2016), or descriptive time-series analysis (e.g. Guttentag and Secord 1983; Grossbard-Shechtman 1993).

4.3 Operationalisation

All predictor variables are modelled as time-lagged predictors. Independent variables from the prior wave (relationship year$_{t-1}$) are used to estimate the risk of union dissolution in the subsequent wave (relationship year$_t$), which makes the sequential ordering of the data more accurate and accounts for the time lag in social processes (Singer and Willett 2003: 441f.). Macro-level indicators are grand mean centred by setting the zero points of the indicators on their means, which allows for a more straightforward interpretation of the intercept, its variance and cross-level interactions (Hox 2002: 49ff.).

4.3.1 Regional Sex Ratio

As shown above, the measurement of sex ratio is possible on different levels and in different age groups. On the one hand, rather narrow sex ratios have been analysed in previous research capturing alternative spouses in the age range of 2–3 years (Klein 1994; Guttentag and Secord 1983) or in the age range of 10–15 years (Grossbard-Shechtman 1993; South 1995; Lyngstad 2011). On the other hand, a relatively broad age span of 15–49 (South and Trent 1988, 1989) or 18–44 (South et al. 2001, 2003) years is used in most studies.

Although every measurement of macro-level sex ratios with regard to age spans is in some sense arbitrary (Lyngstad 2011: 66), Fossett and Kiecolt (1991) have shown that there is no apparent benefit in using narrow age spans in sex ratio studies. Instead, it is important to capture the age span in which most relationships are formed and reformed (South et al. 2001: 747). Using a broader sex ratio allows us to analyse all relationships in the *pairfam* study without excluding cases or estimating different models. Second, relationships in which one partner is significantly older than the other are not per se neglected. Third, to compare the results of sex ratios on the district level with international studies, good coherence in the measurement

is needed. To ensure that the selected instrument is valid, we conducted robustness checks with different models using the three cohorts in the *pairfam* study, and narrower sex ratios as well as non-metric sex ratios with dummy variables indicating a surplus of men or women. These modifications did not change our results. Therefore, we have chosen to use a marginally modified age range (in line with the cohorts in *pairfam*) of 16- to 49-year-old men in a specific district divided by 16- to 49-year-old women in the same district.[8] Hence, a sex ratio above one indicates a surplus of men in this region, while a sex ratio smaller than one indicates a higher share of women in the population in this age group. The sex ratio is measured in time-variant and in logarithmic form to ensure a symmetrical comparability of the surplus of men and women from the parity point (South and Trent 1989: 395). To control for the hypothesised effect of an *unbalanced* sex ratio, the quadratic term for sex ratio is also estimated (South and Lloyd 1995: 27).

4.3.2 Individual-Level Variables

For an accurate estimation of the effect of the regional sex ratio, individual- and couple-level variables must be included in the analysis. *Cohabitation duration* and its quadratic form are modelled to estimate the changing union dissolution risks through time (Kulu 2014). As marriages are shown to be more stable than non-marital unions, and relationships with biological children living in the household and high relationship-specific investments are associated with lower dissolution rates (Lyngstad and Jalovaara 2010), we include in the model *relationship status*, *number of biological children* living currently with both partners, and owning a *joint property*. The number of children is shown to have a nonlinear influence on the dissolution rate (Wagner and Weiß 2003: 44). Hence, the number of children is controlled for with a quadratic polynomial term. The expected negative influence of *relationship satisfaction* on the risk of union dissolution is measured as a continuous variable on a 10-point scale ranging from "very dissatisfied" to "very satisfied".

In contrast to having access to attractive alternatives, having individual (religious) norms, emotional closeness, or an idealistic-traditional view of the marriage act as barriers to union dissolution (Levinger 1979). It follows that the evaluation of a statement on *relationship norms* is included in the model: "Marriage is a lifelong union that should not be broken". The variable measuring *religious affiliation* differentiates between Christians, Muslims, and members of other religious groups, while individuals without religious affiliation form the reference category. According to microeconomic theory, relationships in which substitutive characteristics diverge are more stable; i.e. in which one of the partners is better educated and specialises in paid work (Becker et al. 1977). However, newer studies have shown that educational homogamy has a stabilising effect (Kaplan and Herbst 2015; Mäenpää and Jalovaara 2014). To account for both effects, the *educational levels* (ISCED 2011 with eight categories) of men and women, as well as the *relative educational levels* and the

[8] No limitations are made regarding the married population (South and Lloyd 1995), as a marriage is no guarantee of "retirement" from the available partner market (Stauder 2006: 617).

relative employment levels of the partners are included in the model. Relative educational level in a relationship is constructed through contrasting the ISCED level of the male partner with the ISCED level of the female partner. If the educational level is the same, it will be coded as "educational homogamy", otherwise either the man or the woman will be categorised as the one with the higher educational level. The same strategy is used to contrast dyadic full-time employment levels with resulting categories of "neither full-time", "only man full-time", "only woman full-time", and "both full-time".

As it has been widely shown that *age at the beginning of the relationship* is positively correlated with relationship stability (Lyngstad and Jalovaara 2010: 259), this age is estimated for both men and women. Intimate relationships that were entered into before age 14 are excluded from this study. To control for potential *cohort* effects, three *pairfam* cohorts are compared in the analysis. All individual-level characteristics, except age at the beginning of the relationship and cohort, are measured as time-variant variables.

4.3.3 District-Level Variables

Since German unification, substantial migration to western Germany from structurally weaker regions in the east has been taking place. Kröhnert and Vollmer (2012) found in their study that gender-specific migration from eastern to western Germany is related to women in eastern Germany having higher educational levels than men. The lack of jobs in eastern Germany for people with higher qualifications in the tertiary sector has led to large waves of east–west migration, particularly among young and highly educated women. These women have tended to move to economically stronger regions in the west and to larger cities (Häring et al. 2012). Consequently, a surplus of men, especially in the 18-29 age group, has emerged in eastern Germany (Kröhnert et al. 2006: 24). These developments show how disparities in structural characteristics influence age- and gender-specific migration patterns, which in turn influence the dynamics of regional sex ratios in Germany. Furthermore, there are persisting differences between the divorce rates in western and eastern Germany (e.g. Grünheid 2013: 22f.); and the proportion of couples living in non-marital unions remains higher in eastern than in western Germany (Huinink et al. 2012: 15ff.). Hence, when estimating sex ratios in the German context, *living in eastern or western Germany* is an important control variable. The region of residence is included as a dummy variable, with living in western Germany as the reference category. This variable captures cultural and normative characteristics, demographic and economic trends, and compositional effects.

A further variable measured on the regional level is the *proportion of highly educated persons in a district*. Lyngstad (2011) found a negative effect of well-educated ratios on the district level (controlling for sex ratios) on marriage dissolution in Norway. He stressed that this mechanism can be explained through "social imitation" resulting from lower divorce rates among highly educated couples. It is, however, important to keep in mind that this variable is correlated with economic factors, population density, and other unobserved community characteristics, which makes a causal interpretation of this relationship impossible (Lyngstad 2011: 62ff.). As we

stated above, one of the reasons for the male surplus in structurally weak regions in Germany is the emigration of women with high levels of education. Thus, sex ratios can be correlated with the proportion of the highly educated on the district level and should be included in the model. In this study, the aggregated education variable measures the proportion of individuals with tertiary education who contribute to social insurance in one calendar year (BBSR 2014).

Another variable associated with higher divorce rates (McKinnish 2004; South and Lloyd 1995) and sex ratios (Grossbard-Shechtman 1993: 94f.; Diekmann 1992) is the proportion of *female labour force participation*, operationalised here as an annual proportion of women in the total labour force who are contributing to social insurance (BBSR 2014). To measure potential external barriers (Lewis and Spanier 1979), we include the *proportion of Catholics* in a region, as Catholicism is the largest religious affiliation in Germany, but the share of Catholics varies considerably on the district level (Statistische Ämter des Bundes und der Länder 2014). These variables could (additionally) explain the community variation in union dissolution rates.

5 Results

First, the main independent variable sex ratio is described (see a cartographic illustration in Appendix 1). According to the map, a relatively balanced sex ratio in the analysed age group of 16- to 49-year-old individuals can be identified (median = 103; mean = 104) in Germany. Thus, the distribution of sex ratios is approaching the theoretical mean of balanced sex ratios (100 men to 100 women).[9] Regional units that tend to have more female than male residents (SR < 100) are more likely to be found in western Germany and in larger cities (e.g. Wiesbaden (92); Starnberg (92); Hamburg (93); Bonn (95); Düsseldorf (97)). A shortfall of women (SR > 100) can be observed mainly in eastern Germany and in rural areas (e.g. Hildburghausen (119); Ilm-Kreis (118); Oberspreewald-Lausitz (118); Schmalkalden-Meiningen (118)). The map shows that there are indeed differences in the explanatory variable between German districts and cities: there is a slight surplus of men in this age group, but the total variance of the independent context variable is rather small (Table 2).

Next, bivariate statistics of categorical independent variables with respect to the dependent variable relationship stability ("event occurred" and "no event or right-censoring") are presented (Table 1).[10] The final sample includes 4766 individuals. Whereas 4474 of these respondents do not show an event during the observation period, 292 (6%) separated or their relationship ended in divorce. The proportion of union dissolution is significantly higher for cohabitating couples than for married couples (10 vs. 4%; Pearson's $\chi^2 = 67$, $p < 0.001$). Couples who invested in

[9] Hence, the distribution of sex ratios is comparable with sex ratio distributions from other international studies (see South et al. 2001: 748).

[10] The adjusted sample for the final analysis consists of valid records in all variables (listwise exclusion). For this reason, 305 out of 402 German districts could be analysed, and 4766 relationships remain in the final sample. The event data (person-period format) are transformed into a person-based dataset (wide format). Time-variant variables are averaged over individual-specific means.

Table 1 Contingency table of relationship stability and categorical predictors in Germany. *Source*: Own computation with *pairfam* 6.0 (2008/09-2013/14) data (Brüderl et al. 2015)

	All		*No dissolution*		*Dissolution*	
	N	%	*N*	%		%
Level of rel. institutionalisation						
Married	2878	100	2768	96	110	4
Cohabitating	1888	100	1706	90	182	10
Relationship-specific investments						
No joint property	3355	100	3102	92	253	8
Joint property	1411	100	1372	97	39	3
Religiosity						
No affiliation	1338	100	1220	91	118	9
Catholic or protestant	2971	100	2807	94	164	6
Muslim	278	100	274	99	4	1
Other affiliation	179	100	173	97	6	3
Relative employment level						
Neither full-time	520	100	470	90	50	10
Only man full-time	2541	100	2416	95	125	5
Only woman full-time	666	100	625	94	41	6
Both full-time	1039	100	963	93	76	7
Educational homogeneity						
Educational homogeneity	2186	100	2066	95	120	5
Man higher educated	1536	100	1433	93	103	7
Woman higher educated	1044	100	975	93	69	7
Cohort						
1991–1993	126	100	99	79	27	21
1981–1983	1867	100	1711	92	156	8
1971–1973	2773	100	2664	96	109	4
Region						
Western Germany	3891	100	3666	94	225	6
Eastern Germany	875	100	808	92	67	8
N persons	4766	100	4474	94	292	6
N person years	13,989		13,697		292	

% Row percentages; Pearson's χ^2 in text

relationship-specific capital have lower dissolution rates than couples in relationships without those specific investments (Pearson's $\chi^2 = 39$; $p < 0.001$).

Furthermore, statistically significant differences with regard to relationship stability can be found between couples who differed in their religious affiliations (Pearson's $\chi^2 = 32$; $p < 0.001$), relative employment levels (Pearson's $\chi^2 = 20$; $p < 0.001$), and cohorts (Pearson's $\chi^2 = 91$; $p < 0.001$). A weak connection between the region of residence and union dissolution can be observed (Pearson's $\chi^2 = 4.4$; $p = 0.04$). But when we look at differences in educational homogeneity, we see no clear differences with regard to dissolution shares (Pearson's $\chi^2 = 2.8$; $p = 0.23$).

Table 2 Descriptive statistics of relationship quality and continuous predictors in Germany. *Source*: Own computation with *pairfam* 6.0 (2008/09-2013/14) data (Brüderl et al. 2015); Statistische Ämter des Bundes und der Länder (2014, 2016) and BBSR (2014)

	Min.	Max.	*No dissolution*		*Dissolution*		*T*
			Mean	SD	Mean	SD	
Individual level							
Cohabitation duration (years)	0	22	7.20	5.39	4.88	4.97	7.2
Relationship norms	1	5	3.45	1.31	2.94	1.33	6.5
Relationship satisfaction	0	10	8.01	1.83	6.83	2.27	10.5
Mutual children	0	10	1.19	1.06	0.67	0.98	8.2
Education: woman (ISCED)	0	8	4.93	1.67	4.61	1.66	3.2
Education: man (ISCED)	0	8	5.10	1.69	4.79	1.68	3.0
Age at rel. beginning: woman	14	53	22.30	5.06	21.85	5.20	1.5
Age at rel. beginning: man	14	61	25.07	6.01	24.79	6.15	0.8
District level							
Sex ratio	90.78	117.24	101.67	5.96	101.82	5.96	0.4
Proportion of highly educated	6.00	78.20	18.45	13.42	17.02	12.41	1.8
Female labour force participation	30.30	57.03	45.53	3.88	45.75	3.97	0.9
Proportion of Catholics	1.80	81.80	30.68	23.52	28.18	23.87	1.8
N persons	4766		4474		292		
N person years	13,989		13,697		292		

Mean, arithmetic mean; SD, standard deviation; *t*, *t* value from two-tailed *t* test; district-level variables not centred and not logarithmised

Table 2 shows the bivariate relationships of the continuous independent variables with regard to dissolution proportions in the final sample. The respondents who separated during the observation window had, on average, lived together for a significantly shorter period of time (7.20 vs. 4.88 years). In addition, they had on average weaker relationship norms, lower relationship satisfaction levels, and lower numbers of biological children. Furthermore, the couples who separated during the study were more likely to have low levels of education. This difference is shown to be statistically significant for both men and women ($p < 0.001$). However, when we look at the age at the beginning of the relationship, we see no substantial mean differences with respect to the dependent variable for men or for women. The results are similar for all predictors on the regional level; hence, we find no statistically significant difference in the mean sex ratio on the district level between the respondents who separated and the respondents who did not separate.

Our finding that there is no bivariate connection between the independent variable sex ratio and union dissolution is in line with our main hypothesis (H_1) about a U-shaped connection between sex ratios and union dissolution. As we hypothesised that a surplus of *both* men and women has a negative influence on relationship stability, we had expected to find that the connection is not identifiable *on average*. This finding requires further investigation by applying more advanced statistical methods.

As a starting point for a multilevel data analysis (Table 3), it is necessary to determine how much heterogeneity exists between contexts (level 2 units) with regard to the dependent variable, and whether the implementation of multilevel analysis is justified. The intra-class correlation coefficient (ICC; see Hox 2002: 117) in the present analysis is 0.049. This means that about 5% of the differences in the risk of union dissolution can potentially be attributed to the differences between the districts. Accordingly, the contextual-level differences in relationship stability are indeed fairly low. However, the Chi-square test (Snijders and Bosker 2012: 292) reveals that these differences are statistically significant ($\chi^2 = 415$, $p < 0.001$). The question of interest is whether these moderate but significant differences in relationship stability between the districts are related to diverging regional sex ratios.

Model 1 includes all independent variables on the micro-level in order to estimate the dyadic risk of union dissolution. Although non-marital relationships have higher shares of dissolution than married couples (Table 1) and display a positive effect on the hazard of union dissolution (Model 1), the effect is not statistically significant. This may be due to the relatively comprehensive control of exit costs and relationship-specific investments, which are in turn associated with the degree of institutionalisation of the relationship (Kopp et al. 2010). Accordingly, it cannot be clearly shown that cohabitating unions are per se more unstable than marital unions. In fact, this connection should be treated with caution if the relevant variables are not included in the model.[11]

The hazard of union dissolution decreases significantly with increasing relationship satisfaction. Hence, relationship satisfaction in the previous year negatively affects the relationship dissolution risk in the following relationship year. Furthermore, the model shows that the level of relationship-specific investments is related to union dissolution. Specifically, the presence of mutual children in the household has a significant stabilising effect on relationships. Compared to couples without joint property, couples with shared capital have a lower dissolution rate. These findings underpin the hypothesised negative effects of high relationship quality and high relationship-specific investments on union dissolution risks and are in line with exchange and family economic theories.

Control variables such as cohabitation duration, educational level, relative educational level, and relative employment level are not found to be statistically significant. However, strong relationship norms are shown to have a negative effect on relationship dissolution, with more religious couples having a lower dissolution risk than respondents with no religious affiliation. Furthermore, older cohorts are found to have a lower hazard of union dissolution than younger cohorts (Table 3).[12]

In Model 2, the contextual-level variables are included in the equation. The main independent variable sex ratio shows no statistically significant connection with

[11] A model without relationship-specific investments shows a significant influence of the degree of institutionalisation on union dissolution, which levels out after the inclusion of the relationship-specific investment variables.

[12] With this model, we cannot distinguish whether the effect is a pure age effect or a cohort effect on union dissolution.

Table 3 Discrete binomial multilevel event-history analysis for estimating relationship dissolution in Germany (logit coefficients). Source: Own computation with *pairfam* 6.0 (2008/09-2013/14) data (Brüderl et al. 2015); Statistische Ämter des Bundes und der Länder (2014, 2016) and BBSR (2014)

	Model 0	Model 1	Model 2	Model 3
Intercept	−3.93***	1.42**	1.40*	1.47*
	(0.001)	(0.56)	(0.55)	(0.58)
Individual variables: level$_1$				
Cohabitation duration		−0.02	−0.01	−0.01
		(0.04)	(0.04)	(0.04)
Cohabitation duration^2		0.00	0.00	0.00
		(0.00)	(0.00)	(0.00)
Cohabitating (*ref.:* married)		0.13	0.15	0.31
		(0.17)	(0.17)	(0.22)
Relationship satisfaction		−0.21***	−0.21***	−0.23***
		(0.02)	(0.02)	(0.03)
Mutual children		−0.70***	−0.71***	−0.77***
		(0.12)	(0.12)	(0.15)
Mutual children^2		0.09***	0.09***	0.09***
		(0.02)	(0.02)	(0.02)
Joint property (*ref.: no joint property*)		−0.49*	−0.54**	−0.47$^+$
		(0.20)	(0.20)	(0.28)
Relationship norms		−0.26***	−0.27***	−0.27***
		(0.05)	(0.05)	(0.05)
Catholic or Protestant (*ref.: no religion*)		−0.35**	−0.33*	−0.33*
		(0.13)	(0.15)	(0.15)
Muslim (*ref.: no religion*)		−1.29*	−1.15*	−1.1*
		(0.53)	(0.53)	(0.53)
Other religion (*ref.: no religion*)		−0.77	−0.72	−0.69
		(0.47)	(0.47)	(0.48)
Education: woman (ISCED)		−0.09	−0.08	−0.08
		(0.08)	(0.08)	(0.08)
Education: man (ISCED)		−0.04	−0.03	−0.02
		(0.08)	(0.09)	(0.09)
Education: man higher (*ref.: homogeneous*)		0.10	0.10	0.10
		(0.22)	(0.22)	(0.22)
Education: woman higher (*ref.: homogeneous*)		−0.03	−0.03	−0.02
		(0.24)	(0.24)	(0.24)
Only man full-time (*ref:. neither full-time*)		−0.13	−0.16	−0.15
		(0.17)	(0.18)	(0.18)
Only woman full-time (*ref:. neither full-time*)		0.22	0.21	0.20
		(0.24)	(0.25)	(0.25)
Both full-time (*ref.: neither full-time*)		−0.22	−0.26	−0.26
		(0.20)	(0.20)	(0.20)
Age at rel. beginning: woman		−0.00	−0.00	−0.00
		(0.02)	(0.02)	(0.02)
Age at rel. beginning: man		−0.00	−0.00	−0.00
		(0.01)	(0.01)	(0.01)
Cohort 1981–1983 (*ref.: 1991–1993*)		−1.05***	−1.08***	−1.08***
		(0.29)	(0.30)	(0.30)

Table 3 (continued)

	Model 0	Model 1	Model 2	Model 3
Cohort 1971–1973 (*ref.: 1991–1993*)		−1.63*** (0.41)	−1.67*** (0.41)	−1.66*** (0.41)
District variables: level$_2$				
ln sex ratio			0.36 (2.73)	−0.70 (5.13)
ln sex ratio^2			−19.32 (18.12)	−23.04 (26.35)
Region: eastern Germany(*ref.: western Germany*)			0.15 (0.28)	0.15 (0.28)
Proportion of highly educated			−0.02* (0.01)	−0.02* (0.01)
Female labour force participation			0.01 (0.02)	0.01 (0.02)
Proportion of Catholics			−0.00 (0.00)	−0.00 (0.00)
Interaction effects: level$_1$ *and level*$_2$				
Sex ratio*cohabitating				−3.89 (3.31)
Sex ratio*relationship satisfaction				0.49 (0.49)
Sex ratio*children				1.31 (1.79)
Sex ratio*joint property				−1.95 (5.02)
AIC	2837	2524	2528	2536
BIC	2852	2705	2754	2808
σ^2_{u0j}	0.170	0.147	0.129	0.171
σ^2_{u1j}				1.06
L_2 $\sigma_{u0j}{}^2$ in relation to model 0 σ^2_{u0j}	1	86	76	
N-level$_1$	13,989	13,989	13,989	13,989
N-level$_2$	305	305	305	305

***$p \leq 0.001$; **$p \leq 0.01$; *$p \leq 0.05$; $^+p \leq 0.1$; standard errors in brackets

relationship stability. Moreover, neither the separate modelling of a potential linear effect (e.g. Guttentag and Secord 1983; not shown in tables) nor the modelling of the expected nonlinear U-shaped relationship (e.g. South et al. 2001) indicates a significant connection between sex ratios on the district level and the dissolution of relationships. Thus, the main hypothesis that there is a negative relationship between unbalanced sex ratios and union stability cannot be supported with this sample on the level of German districts (H_1).

Differences between western and eastern Germany with regard to union dissolution risks are not found to be significant. This is also the case for the employment rate of women on the regional level, which shows a positive but insignificant logit coefficient. Furthermore, the role of the external barriers, as measured by the share

of Catholics, cannot be identified. Nonetheless, the proportion of highly educated employees in a district is shown to have a negative connection with the transition rate to union dissolution on the micro-level. Hence, the hazard of union dissolution decreases with a rising proportion of highly educated employees on the regional level. This effect is remarkable because the individual level of education of both partners and the educational homogamy of the dyad are controlled for in this multilevel model. Since the proportion of highly educated employees on the regional level is positively correlated with economic indicators (Lyngstad 2011: 62), this connection is an indication that the structural strength of the regional units can be related to the risk of union dissolution. In line with the results of Lyngstad (2011) and South and Lloyd (1995), we find that in this context, the variable "regional population density" is not significantly related to the hazard of union dissolution. We excluded this variable from the final analysis due to the risk of model over-specification.

The inclusion of contextual variables in Model 2 additionally explains the variance in the dissolution risks on the regional level. However, the remaining 76% of unexplained variance is based on unobserved heterogeneity between the German districts. It therefore appears that further research on the contextual explanatory factors with regard to relationship stability is needed. In addition, this district-level error term might be correlated with the sex ratio restricting our interpretations. As we cannot make any causal claims about the estimated "effect" of the sex ratio on union dissolution, our interpretation of our results is limited to purely correlational statements. An extension of the statistical model through varying slopes (random-coefficient model) does not improve the model fit criteria AIC and BIC. In addition, none of the expected interactions of the main variable sex ratio with the individual characteristics of the respondents can be detected. The availability of alternatives at the regional level does not interact with the degree of institutionalisation of the relationship (H_{2b}), the level of relationship satisfaction (H_{3b}), or the relationship-specific investments (H_{4b}). Thus, our hypotheses regarding cross-level interactions are not supported.

6 Discussion and Conclusion

The aim of this study was to analyse the stability of relationships as a function of the social structure the couples are embedded in. Our main focus was on the availability of alternative partners at the level of German districts and its potential connection with relationship stability. This availability of alternatives was indicated by the sex ratio in each district. It has been argued that the availability of alternative partners can promote the dissolution of an existing relationship. Hence, we investigated the macro-micro link between sex ratios at the district level and relationship stability at the micro-level.

Our theoretical modelling was based on theories from divorce and separation research. Micro-level theories, exchange theory, and family economic theory are based on the rational-choice approach and point to the relevance of the distribution of resources and relationship-specific investments for the stability of the relationship. The influence on union dissolution of having access to attractive alternatives

posited in these theories was of a particular relevance for the present work. According to Blau's structural theory, the societal context influences the likelihood of meeting (alternative) partners. Accordingly, the distribution of gender at the contextual level should have an influence on the stability of relationships. Based on these theoretical considerations and on previous research on sex ratios and union dissolution (e.g. South and Lloyd 1995; South et al. 2001; Lyngstad 2011), we derived our hypotheses. Union dissolution risks as a function of individual and contextual predictors were estimated by discrete-time multilevel event-history analysis using *pairfam* data and data from German official statistics. This method of analysis was found to adequately model the individual relationship histories and to estimate time-variant and time-invariant contextual parameters and their unbiased standard errors (Hox 2002).

The main hypothesis, which asserted that there is a positive connection between unbalanced sex ratios and union dissolution, could not be supported. The effect of district-level sex ratios was not found to be significantly related to the risk of union dissolution on the micro-level. Hence, the robustness of previous findings about the relationship-destabilising effect of unbalanced sex ratios on the regional level (e.g. South and Lloyd 1995; South et al. 2001) is called into question by these results (see also Lyngstad 2011).

Our failure to find a statistically significant connection can be explained in several ways. First, it is possible that the district level does not represent the actual opportunity structures of the individuals and that the relevant partner market may be found at more compressed levels of social interaction. To investigate whether this is the case, further research that looks at smaller partner markets, or the social "foci" at the workplace or social network level, is needed (Feld 1981; Lengerer 2001). Second, it can be assumed that the relatively small differences in the sex ratio distribution between the German districts do not have enough variance to allow us to discern a clear effect of the unbalanced sex ratio.[13] Third, it is possible that the sex ratios are too undifferentiated to allow us to detect the availability of attractive alternatives. A fourth explanation mechanism can also be suggested: i.e. since the sampling of the data is proportional to the number of inhabitants in a region (Suckow and Schneekloth 2009) and the sex ratios are rather unbalanced in regions with a lower population density, respondents from regions with unbalanced sex ratios may be underrepresented in this sample. Finally, unobserved heterogeneity on the contextual level can suppress the connection between sex ratios and union dissolution. Some omitted variables on the district level, such as the main industry sector, the unemployment rate, and the educational infrastructure, are likely to be correlated with sex ratios on the district level. This violates the assumption of uncorrelated district-level error terms and sex ratios. When the estimation of time-invariant contextual variables is not theoretically decisive, we suggest addressing this limitation through the use of fixed-effects methods for controlling confounders (Kravdal 2011: 291).

[13] The distribution of current gender relations between the German regional levels cannot be compared with a so-called marriage squeeze, which was, for example, observed in the post-war period (e.g., Martin 2001). Thus, the analysed sex ratios may not represent a substantial male or female surplus.

Nevertheless, the study showed that the micro-level variables are far more robust predictors of the risk of union dissolution than contextual predictors. Thus, it can be stated that a high level of relationship satisfaction in the previous year reduces the dissolution risk in the following year. Our findings also demonstrated that relationship-specific investments have a positive influence on the stability of the relationship over time. Moreover, stabilising effects were found for strong family-related norms and having a religious affiliation.

Looking at the contextual predictors, we see that only the proportion of highly educated employees in a district turned out to be a significant predictor. Accordingly, the risk of union dissolution decreases with an increasing proportion of highly educated people in a district. This finding indicates that relationship stability is linked to different economic and structural conditions on the regional level. Therefore, education-specific sex ratios and possible cross-level interaction effects between the educational levels of individuals and the educational levels of districts can serve as a starting point for further research contributions. In addition, a portion of the differences between the German districts remains unexplained in relation to the stability of the relationship, which underlines the relevance of the multilevel perspective for future analysis on union dissolutions.

For detecting the influences of smaller partner markets on dissolution rates, it would be appropriate to use micro-geographic data at the neighbourhood level. As this issue has so far been neglected, sex ratios in the workplace should be included in efforts to capture partner market opportunity structures in Germany. A continuation of research on partnership stability that looks at the role of alternative partners seems to be both necessary and promising.

Acknowledgements This paper uses data from the German Family Panel pairfam, coordinated by Josef Brüderl, Karsten Hank, Johannes Huinink, Bernhard Nauck, Franz Neyer, and Sabine Walper. Pairfam is funded as a long-term project by the German Research Foundation (DFG).

Compliance with ethical standards

Conflict of interest The authors declare that they have no conflict of interest.

Appendix

See Fig. 1.

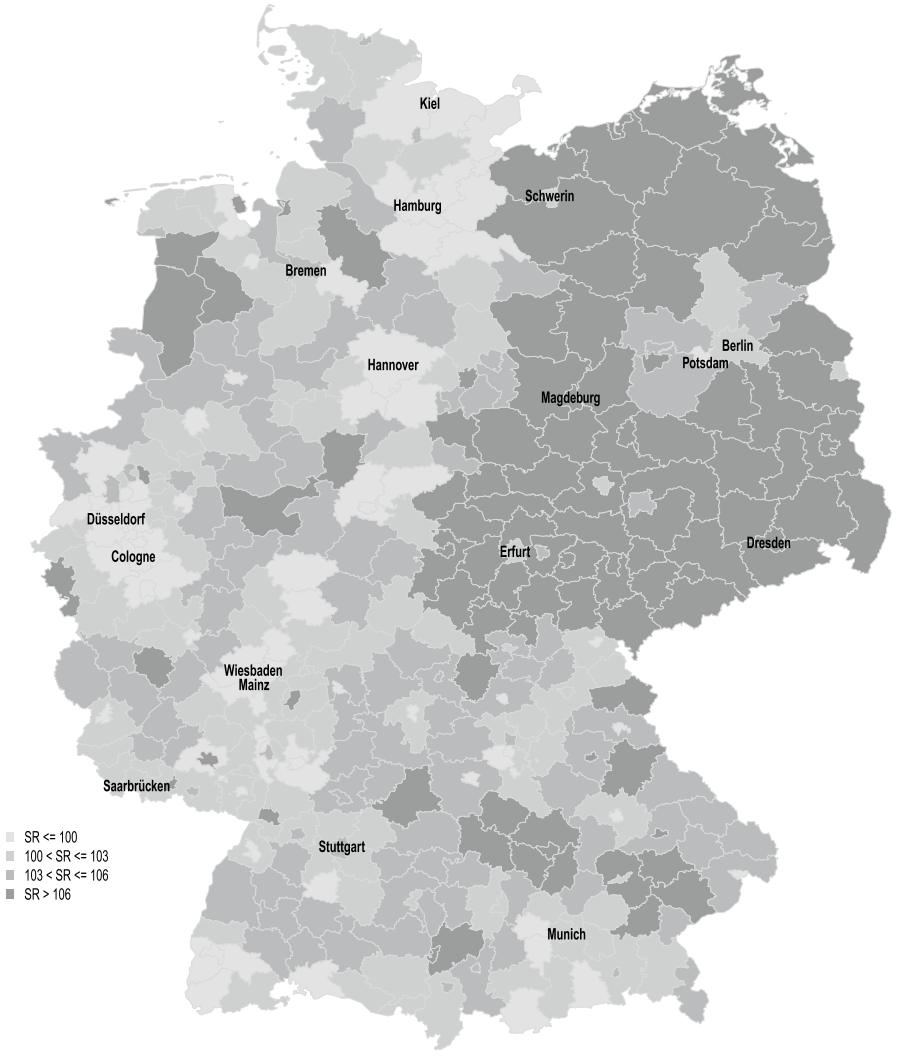

Fig. 1 Sex ratios in German districts (2012). Age group: 16–49-olds; *SR* sex ratio. *Source*: Own computation with regional data (Statistische Ämter des Bundes und der Länder 2014, 2016) in R version 3.3.2 from 2016-10-31

References

Abramitzky, R., Delavande, A., & Vasconcelos, L. (2011). Marrying up: The role of sex ratio in assortative matching. *American Economic Journal: Applied Economics, 3*(3), 124–157.

Allison, P. D. (1984). *Event history analysis. Regression for longitudinal event data.* Sage University papers. Quantitative applications in the social sciences, no. 07-046. Beverly Hills, CA: Sage Publications.

Allison, P. D. (2009). *Fixed effects regression models.* Sage University papers. Quantitative applications in the social sciences, no. 07-160. Beverly Hills, CA: Sage Publications.

BBSR (Bundesinstitut für Bau-, Stadt- und Raumforschung im Bundesamt für Bauwesen und Raumordnung). (2014). *INKAR. Indikatoren und Karten zur Raum- und Stadtentwicklung*. Ausgabe 2014. Bonn. http://www.inkar.de/. Accessed December 06, 2016.

Becker, G. S., Landes, E. M., & Michael, R. T. (1977). An economic analysis of marital instability. *The Journal of Political Economy, 85*(6), 1141–1187.

Blau, P. M. (1977a). *Inequality and heterogeneity. A primitive theory of social structure*. New York: Free Press.

Blau, P. M. (1977b). A macrosociological theory of social structure. *American Journal of Sociology, 83*(1), 26–54.

Blau, P. M. (1994). *Structural contexts of opportunities*. Chicago/London: The University of Chicago Press.

Boyle, P. J., Kulu, H., Cooke, T. J., Gayle, V., & Mulder, C. H. (2008). Moving and union dissolution. *Demography, 45*(1), 209–222.

Brüderl, J., Hank, K., Huinink, J., Nauck, B., Neyer, F. J., Walper, S., Alt, P., Buhr, P., Castiglioni, L., Finn, C., Hajek, K., Herzig, M., Huyer-May, B., Lenke, R., Müller, B., Peter, T., Salzburger, V., Schmiedeberg, C., Schubach, E., Schütze, P., Schumann, N, Thönnissen, C., & Wilhelm, B. (2015). *The German Family Panel* (*pairfam*). GESIS Data Archive, Cologne. ZA5678 Data file Version 6.0.0.

Brüderl, J., & Kalter, F. (2001). The dissolution of marriages: The role of information and marital-specific capital. *Journal of Mathematical Society, 25,* 403–421.

Coleman, J. S. (1986). Social theory, social research, and a theory of action. *American Journal of Sociology, 91*(6), 1309–1335.

Diekmann, A. (1992). Sex-ratio, divorce, and labor force participation—An analysis of international aggregate data. In G. Haag, U. Mueller, & K. G. Troitzsch (Eds.), *Economic evolution and demographic change. Formal models in social sciences* (pp. 283–293). Berlin, Heidelberg, New York: Springer VS.

Eckhard, J. (2014). Der sozialstrukturelle Kontext der zunehmenden Partnerlosigkeit in Deutschland. *Zeitschrift für Soziologie, 43*(5), 341–361.

Eckhard, J., & Stauder, J. (2016). Partnermarkt und familiendemografische Prozesse. In J. Stauder, I. Rapp, & J. Eckhard (Eds.), *Soziale Bedingungen privater Lebensführung.* Festschrift zum 60. Geburtstag von Thomas Klein (pp. 121–144). Wiesbaden: Springer VS.

Esser, H. (1988). Sozialökologische Stadtforschung und Mehr-Ebenen-Analyse. Soziologische Stadtforschung. *Kölner Zeitschrift für Soziologie und Sozialpsychologie. Sonderheft, 29,* 35–55.

Esser, H. (2000). *Opportunitäten und Restriktionen. Soziologie - spezielle Grundlagen* (Vol. 4). Frankfurt am Main: Campus Verlag.

Esser, H. (2001). Das „Framing" der Ehe und das Risiko zur Scheidung. In J. Huinink, K. P. Strohmeier, & M. Wagner (Eds.), *Solidarität in Partnerschaft und Familie. Zum Stand familiensoziologischer Theoriebildung. Familie und Gesellschaft,* Band 7 (pp. 103–129). Würzburg: Ergon Verlag.

Feld, S. L. (1981). The focused organization of social ties. *American Journal of Sociology, 86*(5), 1015–1035.

Fossett, M. A., & Kiecolt, J. K. (1991). A methodological review of the sex ratio. Alternatives for comparative research. *Journal of Marriage and the Family, 53*(4), 941–957.

Goldman, N., Westoff, C. F., & Hammerslough, C. (1984). Demography of the marriage market in the United States. *Population Index, 50,* 5–25.

Grossbard-Shechtman, S. (1993). *On the economics of marriage. A theory of marriage, labour, and divorce*. Boulder, CO, Oxford: Westview.

Grünheid, E. (2013). *Ehescheidungen in Deutschland: Entwicklungen und Hintergründe*. BiB Working Paper 1/2013. Bundesinstitut für Bevölkerungsforschung.

Guo, G., & Zhao, H. (2000). Multilevel modeling for binary data. *Annual Review of Sociology, 26,* 441–462.

Guttentag, M., & Secord, P. F. (1983). *Too many women? The sex ratio question*. Beverly Hills, London: Sage.

Häring, A., Stoye, K., Klein, T., & Stauder, J. (2012). 20 Jahre nach der Wende. Der Partnermarkt junger Erwachsener in Ost- und Westdeutschland. In J. Huinink, M. Kreyenfeld, & H. Trappe (Eds.), *Familie und Partnerschaft in Ost- und Westdeutschland. Ähnlich und doch immer noch anders.* 9. Sonderheft (pp. 257–275). Opladen: Budrich.

Hill, P. B., & Kopp, J. (2006). *Familiensoziologie. Grundlagen und theoretische Perspektiven.* 4. überarbeitete Auflage. Wiesbaden: VS Verlag für Sozialwissenschaften.

Hox, J. J. (2002). *Multilevel analysis. Techniques and applications*. New Jersey: Lawrence Erlbaum Associates.

Hradil, S., & Masson, S. (2008). Familie und Sozialstruktur. In N. F. Schneider (Ed.), *Lehrbuch Moderne Familiensoziologie. Theorien, Methoden, empirische Befunde* (pp. 197–218). Opladen, Farmington Hills, MI: Budrich.

Huinink, J., Brüderl, J., Nauck, B., Walper, S., Castiglioni, L., & Feldhaus, M. (2011). Panel analysis of intimate relationships and family dynamics (*pairfam*): Conceptual framework and design. *Zeitschrift für Familienforschung, 23,* 77–101.

Huinink, J., Kreyenfeld, M., & Trappe, H. (2012). Familie und Partnerschaft in Ost- und Westdeutschland. Eine Bilanz. In J. Huinink, M. Kreyenfeld, & H. Trappe (Eds.), *Familie und Partnerschaft in Ost- und Westdeutschland. Ähnlich und doch immer noch anders*. 9. Sonderheft (pp. 9–29). Opladen: Budrich.

Kalter, F. (1999). „The Ties that Bind"—Wohneigentum als ehespezifische Investition. In T. Klein & J. Kopp (Eds.), *Scheidungsursachen aus soziologischer Sicht*. 2. Band. (pp. 255–270). Würzburg: Ergon Verlag.

Kaplan, A., & Herbst, A. (2015). Stratified patterns of divorce: Earnings, education, and gender. *Demographic Research, 32,* 949–981.

Klein, T. (1994). Marriage Squeeze und Ehestabilität: eine empirische Untersuchung mit den Daten des sozio-ökonomischen Panels. *Zeitschrift für Familienforschung, 6*(3), 177–196.

Klein, T., & Stauder, J. (2008). Partnermärkte in Deutschland im Spiegel eines neuen Erhebungsinstruments. In M. Feldhaus & J. Huinink (Eds.), *Neuere Entwicklungen in der Beziehungs- und Familienforschung: Vorstudien zum Beziehungs- und Familienentwicklungspanel (PAIRFAM)* (pp. 77–113). Würzburg: Ergon Verlag.

Kopp, J., Lois, D., Kunz, C., & Becker, O. A. (2010). *Verliebt, verlobt, verheiratet. Institutionalisierungsprozesse in Partnerschaften* (1st ed.). Wiesbaden: Springer Fachmedien.

Kravdal, Ø. (2011). The fixed-effects model admittedly no quick fix, but still a step in the right direction and better than the suggested alternative. *Journal of Epidemiology and Community Health, 65*(4), 291–292.

Kröhnert, S., Medicus, F., & Klingholz, R. (2006). *Die demografische Lage der Nation. Wie zukunftsfähig sind Deutschlands Regionen?*. München: Deutscher Taschenbuch Verlag.

Kröhnert, S., & Vollmer, S. (2012). Gender-specific migration from eastern to western Germany: Where have all the young women gone? *International Migration, 50*(5), 95–112.

Kulu, H. (2012). Spatial variation in divorce and separation. Compositional or contextual effects? *Population, Space, Place, 18*(1), 1–15.

Kulu, H. (2014). Marriage duration and divorce: The seven-year itch or a lifelong itch? *Demography, 51*(3), 881–893.

Lengerer, A. (2001). Wo die Liebe hinfällt – ein Beitrag zur „Geographie" der Partnerwahl. In T. Klein (Ed.), *Partnerwahl und Heiratsmuster. Sozialstrukturelle Voraussetzungen der Liebe* (pp. 133–162). Opladen: Leske + Budrich.

Lengerer, A. (2011). *Partnerlosigkeit in Deutschland. Entwicklung und soziale Unterschiede*. Wiesbaden: VS Verlag für Sozialwissenschaften.

Levinger, G. (1979). A social psychological perspective on marital dissolution. In G. Leviger & O. C. Moles (Eds.), *Divorce and separation. Context, causes, and consequences* (pp. 37–66). New York: Basic Books.

Lewis, R. A., & Spanier, G. B. (1979). Theorizing about the quality and stability of marriage. In W. R. Burr, R. Hill, F. I. Nye, & I. L. Reiss (Eds.), *Contemporary theories about the family* (Vol. I, pp. 268–294). New York: Free Press.

Lichter, D. T., Anderson, R. N., & Hayward, M. D. (1995). *Marriage markets and marital choice*. Working paper 95-02. University Park, PA.: Population Research Institute, Pennsylvania State University.

Lichter, D. T., LeClere, F. B., & McLaughlin, D. K. (1991). Local marriage markets and the marital behavior of black and white women. *American Journal of Sociology, 96*(4), 843–867.

Lloyd, K. M., & South, S. J. (1996). Contextual influences on young men's transition to first marriage. *Social Forces, 74*(3), 1097–1119.

Lyngstad, T. H. (2011). Does community context have an important impact on divorce risk? A fixed-effects study of twenty Norwegian first-marriage cohorts. *European Journal of Population, 27*(1), 57–77.

Lyngstad, T. H., & Jalovaara, M. (2010). A review of the antecedents of union dissolution. *Demographic Research, 23,* 257–292.

Mäenpää, E., & Jalovaara, M. (2014). Homogamy in socio-economic background and education, and the dissolution of cohabiting unions. *Demographic Research, 30,* 1769–1792.

Martin, F. O. (2001). Marriage squeeze in Deutschland – aktuelle Befunde auf der Grundlage der amtlichen Statistik. In T. Klein (Ed.), *Partnerwahl und Heiratsmuster. Sozialstrukturelle Voraussetzungen der Liebe* (pp. 287–316). Opladen: Leske + Budrich.

McKinnish, T. G. (2004). Occupation, sex-integration, and divorce. *The American Economic Review, 94*(2), 322–325.

Rupp, M., & Blossfeld, H.-P. (2008). Familiale Übergänge: Eintritt in nichteheliche Lebensgemeinschaften, Heirat, Trennung und Scheidung, Elternschaft. In N. F. Schneider (Ed.), *Lehrbuch Moderne Familiensoziologie. Theorien, Methoden, empirische Befunde* (pp. 139–166). Opladen, Farmington Hills, MI: Budrich.

Schmiedeberg, C. (2015). *Regional data in the German family panel* (*pairfam*). Technical Paper. Nr. 07.

Schoen, R. (1985). Population growth and the birth squeeze. *Social Science Research, 14*(3), 251–265.

Singer, J. D., & Willett, J. B. (2003). *Applied longitudinal data analysis. Modeling change and event occurrence*. Oxford: Oxford University Press.

Snijders, T. A. B., & Bosker, R. J. (2012). *Multilevel analysis. An introduction to basic and advanced multilevel modeling* (2nd ed.). Los Angeles, London: Sage.

South, S. J. (1995). Do you need to shop around? Age at marriage, spousal alternatives, and marital dissolution. *Journal of Family Issues, 16*(4), 432–449.

South, S. J., & Lloyd, K. M. (1995). Spousal alternatives and marital relations. *American Sociological Review, 60*(1), 21–35.

South, S. J., & Trent, K. (1988). Sex ratios and women's roles: A cross-national analysis. *American Journal of Sociology, 93*(5), 1096–1115.

South, S. J., & Trent, K. (1989). Structural determinants of the divorce rate: A cross-societal analysis. *Journal of Marriage and the Family, 51*(2), 391–404.

South, S. J., Trent, K., & Shen, Y. (2001). Changing partners. Toward a macrostructural-opportunity theory of marital dissolution. *Journal of Marriage and Family 63, 3,* 743–754.

Statistische Ämter des Bundes und der Länder. (2014). *Anteil der Bevölkerungsgruppe mit Zugehörigkeit zur römisch-katholischen Kirche*. Ergebnis des Zensus 2011 zum Berichtszeitpunkt 9 Mai 2011. https://ergebnisse.zensus2011.de/#MapContent:00,R55,m. Accessed December 06, 2016.

Statistische Ämter des Bundes und der Länder. (2016). *Regionaldatenbank Deutschland*. https://www.regionalstatistik.de/genesis/online. Accessed December 06, 2016.

Statistisches Bundesamt. (2015). *Kreisfreie Städte und Landkreise nach Fläche, Bevölkerung und Bevölkerungsdichte am 31.12.2014*. https://www.destatis.de/DE/ZahlenFakten/LaenderRegionen/Regionales/Gemeindeverzeichnis/Administrativ/Aktuell/04Kreise.html. Accessed December 06, 2016.

Stauder, J. (2006). Die Verfügbarkeit partnerschaftlich gebundener Akteure für den Partnermarkt. *Kölner Zeitschrift für Soziologie und Sozialpsychologie, 58*(4), 617–637.

Stauder, J. (2008). Opportunitäten und Restriktionen des Kennenlernens. *Kölner Zeitschrift für Soziologie und Sozialpsychologie, 60*(2), 266–286.

Stauder, J. (2011). Regionale Ungleichheit auf dem Partnermarkt? Die makrostrukturellen Rahmenbedingungen der Partnerwahl in regionaler Perspektive. *Soziale Welt, 62*(1), 45–73.

Stauder, J. (2015). Durchdringende Sozialstruktur? Der Einfluss makrostruktureller Rahmenbedingungen auf den Partnermarkt. *Kölner Zeitschrift für Soziologie und Sozialpsychologie, 67*(3), 401–432.

Steele, F. (2011). Multilevel discrete-time event history models with applications to the analysis of recurrent employment transitions. *Australian & New Zealand Journal of Statistics, 53*(1), 1–20.

Suckow, J., & Schneekloth, U. (2009). *Beziehungen und Familienleben in Deutschland* (2008/2009). Welle 1. München: TNS Infratest.

Trent, K., & South, S. J. (2003). Spousal alternatives and marital relations. *Journal of Family Issues, 24*(6), 787–810.

Udry, J. R. (1981). Marital alternatives and marital disruption. *Journal of Marriage and Family, 43*(4), 889–897.

Uecker, J. E., & Regnerus, M. D. (2010). Bare market: Campus sex ratios, romantic relationships, and sexual behavior. *The Sociological Quarterly, 51,* 408–435.

Wagner, M., Schmid, L., & Weiß, B. (2015). Exploring increasing divorce rates in West Germany. Can we explain the iron law of increasing marriage instability? *European Sociological Review, 31*(2), 211–229.

Wagner, M., & Weiß, B. (2003). Bilanz der deutschen Scheidungsforschung. Versuch einer Meta-Analyse. *Zeitschrift für Soziologie, 32*(1), 29–49.

Wagner, M., & Weiß, B. (2006). On the variation of divorce risks in Europe. Findings from a meta-analysis of European longitudinal studies. *European Sociological Review, 22*(5), 483–500.