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Diffusion of Childbearing Within Cohabitation

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Abstract The article analyzes the diffusion of childbearing within cohabitation in Norway, using municipality data over a 24-year period (1988–2011). Research has found substantial spatial heterogeneity in this phenomenon but also substantial spatial correlation, and the prevalence of childbearing within cohabitation has increased significantly over time. We consider several theoretical perspectives and implement a spatial panel model that allows accounting for autocorrelation not only on the dependent variable but also on key explanatory variables, and hence identifies the key determinants of diffusion of childbearing within cohabitation across space and over time. We find only partial support for the second demographic transition as a theory able to explain the diffusion of childbearing within cohabitation. Our results show that at least in the first phase of the diffusion (1988–1997), economic difficulties as measured by increased unemployment among men contributed to the diffusion of childbearing within cohabitation. However, the most important driver for childbearing within cohabitation is expansion in education for women.

Keywords Childbearing within cohabitation · Diffusion · Norway · Municipality · Spatial panel model

Introduction

Over the last few decades, cohabitation has gradually emerged as a new family form—in many cases not only being a precursor to marriage but also replacing marriage. The most important manifestation of the latter is perhaps that increasingly couples are

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having children without, or at least before, getting formally married (Perelli-Harris et al. 2012). As normative views about cohabitation have changed, institutions have also followed suit, and cohabiting couples in many countries now have almost the same rights as married couples. The underlying idea of the second demographic transition (SDT) is that in Western societies, spearheaded by the Nordic ones, the centrality of the family is declining and is being replaced by support for more liberal demographic behaviors, such as divorce, cohabitation, and nonmarital childbearing (Van de Kaa 1987). These new demographic behaviors are viewed as progressive independence of individuals who give growing importance to self-realization and their psychological well-being as well as to their personal freedom of expression (Van de Kaa 1987). A key aspect of the SDT is that it starts from new ideas first introduced by the "forerunners" and then spreads through a process of diffusion (Lesthaeghe and Neels 2002; Lesthaeghe and Surkyn 2008). Yet, as new demographic behaviors are spreading, Western societies have also experienced a process of massive expansion in education for women: more than ever before, women are able to level the playing field with men in terms of work and income (Casterline 2001), which has reduced the gains to marriage (Becker 1981).

An important debate in this literature regards the extent to which new behaviors can be seen as a diffusion process of new ideas spreading. Originally, SDT referred to cultural change, with age/cohort structures, education, religion, and urbanization being key elements. In particular, SDT predicts new ideas and behaviors to originate among young, highly educated, and nonreligious people in urban settings. However, the SDT disregards the gender dimension and, in particular, the importance of the changing role of women (Bernhardt 2004). Equally plausible, women's attitudes and values may have changed because of educational expansion and female labor force participation, thereby weakening the traditional male breadwinner model, and giving rise to more egalitarian societies, and causing new family forms to emerge. The latter argument alludes to the idea that new demographic behaviors are arising because of structural change (Bongaarts and Watkins 1996; Cleland and Wilson 1987). Rather than new ideas by themselves spreading, they might be the result of the diffusion of education and labor markets and, more generally, institutional rights for women. If this is the case, then a new demographic behavior may not be a result of the ideas spreading but rather, the result of new institutions spreading. Finally, the "pattern of disadvantage" hypothesis, as proposed by Perelli-Harris et al. (2010), suggests that childbearing within cohabitation is more common among the lower socioeconomic strata and that it originated among the more disadvantaged people.

In this article, we tackle this issue by using municipality data from Norway over a 24-year period between 1988 and 2011. We adopt a spatial-temporal perspective and investigate how childbearing within cohabitation has been spreading geographically across Norwegian municipalities. Our spatial panel model incorporates spatial autocorrelation on the dependent variable and, importantly, also on the explanatory variables. The spatial correlation on the dependent variable shows the extent to which neighboring municipalities correlate geographically in terms of childbearing within cohabitation. In other words, it measures whether the prevalence of childbearing within cohabitation is more similar to the prevalence observed in neighboring municipalities as opposed to the prevalence observed in municipalities that are farther away. This by itself is an indication of a process of diffusion (Baller et al. 2001; Messner and Anselin 2004). The spatial autocorrelation of the explanatory variables enables us to assess the drivers behind the diffusion of childbearing within cohabitation. This form of autocorrelation measures whether a change in a given independent variable in neighboring municipalities has an effect on childbearing within cohabitation in the reference municipality. We consider four sets of factors. The first set includes measures of religion, urbanization, and age, all of which were the initial key elements of the SDT idea (Van de Kaa 1987). The other factors concern expansion of education for women, women's empowerment, and financial difficulties or deprivation.

Results show that the characteristics of the municipality matter for the prevalence of childbearing within cohabitation (direct effect), but we also find evidence of diffusion originating from the characteristics of neighboring municipalities (indirect effects). However, the role of the explanatory variables in explaining diffusion is very different, depending on the period and, by extension, the stage of diffusion one considers.

Background

If marriage was considered as the main arena for family formation and childbearing in the past, we now increasingly observe children born within cohabiting unions (Perelli-Harris et al. 2012). The disconnection between marriage and childbearing represents one of the most dramatic family changes observed over the past decades. According to the SDT idea (Surkyn and Lesthaeghe 2004), new behaviors are rooted in unprecedented changes in lifestyle choices that in turn stem from individuals' ideational and value change. There has been (and still is) a lively debate about the merit of the SDT as a theoretical explanation for demographic trends (Cliquet 1991; Coleman 2004). Certainly, the SDT is not always correct in its predictions. For example, as the SDT is taking hold, one prediction is that fertility will decline. Clearly, however, fertility today is substantially higher in countries where the SDT has proceeded furthest (Sobotka 2008). There are objections to the way new ideas can simply spread and be the driving force behind new behaviors. Instead, one could argue that new behaviors are driven by structural and institutional change. The SDT theory has been criticized because it does not explicitly incorporate a gender perspective (Bernhardt 2004). Some researchers also doubt the existence of "forerunners"—or at least about who they actually are. According to the SDT, the forerunners of SDT behaviors would be highly educated individuals. However, studies have identified the forerunners of childbearing within cohabitation in the Nordic countries as the least-educated women (Thomson 2013; Trost 1978), and Perelli-Harris et al. (2010) showed a persistent negative educational gradient of first births within cohabitation in Norway, which is more consistent with a pattern of disadvantage than with the ideological shifts assumed by the SDT. Finally, because of the rather heterogeneous patterns of demographic behaviors observed across societies, some researchers doubt whether the emergence of these new behaviors indeed represents a transition (Coleman 2004).

Still, the idea of the SDT has come to dominate many of the arguments made by family demographers in recent decades and certainly has had strong appeal for many (for a debate about the concept, see Billari and Liefbroer 2004; Coleman 2004; Micheli 2004; Van de Kaa 2004). Despite its failings in some respects, SDT provides intuitive insights that are often supported empirically. For instance, new ideas being introduced by the forerunners may indeed have an impact on other individuals, leading to the adoption of new behaviors on a broad scale. This might certainly make sense: information spreads ever more quickly as the world becomes more globalized through media and digital communication (Hornik and McAnany 2001).

At the aggregate level, societies appear to be leading others in terms of new demographic behaviors, the Nordic countries being the typical example. In these countries, the proportion of nonmarital births started to increase in the 1970s; and, as of this writing, the majority of children are born within cohabitation. However, it is in the Nordic countries where the onset of mass education came first. Other than increasing the opportunity cost of children through increased labor force participation and wages (Becker 1981), education also gives rise to nonconformism, lowers the importance of religion, increases the tolerance of unconventional sexual behavior, and increases personal self-realization (Thornton et al. 2008). In addition, education is associated with more liberal attitudes with respect to the sphere of family ties (Aassve et al. 2013). In other words, rather than being a process of diffusion governed by religion, urbanization, and age, ideational and value change may instead spread because of educational expansion.

More recently, there has been increased focus on the role of institutions for which women's empowerment—or, in Esping-Andersen's (2009) terminology, "women's revolution"—plays a central role. Gender equality, women's empowerment, and female labor force participation are all important drivers of the modernization process. With empowerment, women have a stronger influence on the decision-making processes regarding childbearing, work, and union formation and dissolution, thereby becoming more independent from the social and institutional constraints that had limited their possibility of supporting modern demographic behaviors. However, the extent to which gender equality takes hold in a society also depends on the implementation of appropriate institutions, such as policies and infrastructures geared toward helping women to combine family and work. Whereas Sobotka (2008) showed that countries in which the gender revolution spread earlier also adopted norms and institutional features that allowed a faster acceptance of postmodernism, McDonald (2006, 2013) argued that a critical difference exists between gender equality and gender equity, with the latter reflecting perceptions about which roles are appropriate and just for men and women. Like McDonald (2013), other researchers (Arpino et al. 2015; Esping-Andersen 2009) argued that any mismatch between gender equity and equality may have an impact on demographic behaviors and that mismatches typically occur if societies are slow in adapting or expanding the necessary institutions. The key insight from this literature for our analysis is that diffusion of institutions also influences the extent to which women are able to break with the social position that they had in the past (Oppenheimer 1994).

In our empirical analysis, childbearing within cohabitation is taken as a function of several indicators that matter for demographic attitudes and behaviors, grouped into four main factors. The first concerns the original elements of the SDT idea: namely, that a new demographic behavior is driven by secularization, urbanization, and age structure. Strong ties to the church typically slow down or hold back nonconformist behaviors, and whereas Norway as a whole can hardly be considered as a very religious society, there are considerable geographical differences in how religion matters for individuals. Similarly, the prevalence of childbearing within cohabitation is predicted to be higher in urban rather than rural settings and among younger rather than older people. The second indicator concerns the spread of mass education for women, which other than increasing the opportunity costs of childbearing, also brings about modern attitudes. The third indicator concerns women's empowerment. The move away from the male breadwinner society (very much dominant in Norway during the 1960s and 1970s) to that of an egalitarian one, where the dual-earner couple becomes the norm, has certainly increased women's autonomy and bargaining power, consequently marginalizing marriage as an institutional protection for women wanting to have children. The last indicator considers the idea of a pattern of disadvantage (Perelli-Harris et al. 2010). This argument builds on the observation that historically, one often finds a positive correlation between material disadvantage and the rate of nonmarital childbearing. This pattern is documented for the United States, Russia, and also for several European countries (Billy and Moore 1992; Perelli-Harris and Gerber 2011; Perelli-Harris et al. 2010).

Common to the four aspects that we identify as possible explanations for the spread of modern family behaviors is the need for a spatial perspective to account for the process of diffusion (Palloni 2001). The theoretical perspective of the diffusion of innovations explains the spread of new ideas and behaviors as a function of factors associated with the successful adoption of innovations across people and places (Rogers 1995; Valente 1995). Behavioral innovations do not occur randomly in space and time but instead spread among people via social networks and kinships (Casterline 2001; Cleland and Wilson 1987; Montgomery and Casterline 1996). An important contribution of this kind of analysis is that of the Princeton European Fertility Project, which found that the fertility decline of the eighteenth and nineteenth centuries resulted from the diffusion of new attitudes (value and cost of children) and behaviors (availability and awareness of birth control techniques) across European provinces sharing similar cultural characteristics (Bongaarts and Watkins 1996; Coale and Watkins 1986). Demography has long been considered a "spatial social science" (Voss 2007) because of its attention to the role of geographic space in explaining similarities and differences in demographic behaviors across populations or groups; indeed, demographers have recently returned to this emphasis on the importance of the spatial aspects of demographic behaviors (Boyle 2003; Chi and Zhu 2008; de Castro 2007; Entwisle 2007; Lesthaeghe and Lopez-Gay 2013; Schmertmann et al. 2010; Voss 2007; Weeks 2004). Particularly relevant to the present study, Klüsener et al. (2012) described spatial patterns of nonmarital fertility across European states and regions. Despite these previous studies, it is fair to say that explanations of spatial patterns of demographic behaviors are relatively rare and tend to be limited to the fertility decline during the first demographic transition (FDT) across different regions of the world (Tolnay 1995; Van Bavel 2004) or the current fertility decline characterizing developing countries (Potter et al. 2002, 2010; Watkins 1987; Weeks et al. 2000). As concerns the SDT, only descriptive indicators give support to a process of spatial diffusion in Europe (Lesthaeghe and Lopez-Gay 2013; Lesthaeghe and Neels 2002; Valkonen et al. 2008) and in the United States (Lesthaeghe and Neidert 2006). Although spatial analysis of demographic patterns is rather limited, with the onset of more detailed data that allow for the spatial perspective and not least the introduction of appropriate multivariate statistical techniques, the literature is set to see an increase of spatial modeling of demographic phenomena in the years to come (e.g., Boyle 2003; de Castro 2007; Voss 2007).

The key novelty of the spatial analysis that we propose in this article is that we assess not only the *incidence* of nonmarital childbearing but also the extent to which the explanatory variables—and hence the theoretical perspectives—matter for the *diffusion* of childbearing within cohabitation across space and over time. By estimating the spatial autocorrelation parameter associated with our dependent variable, we will be able to evaluate whether childbearing within cohabitation spreads across neighboring municipalities (i.e., whether there is spatial autocorrelation in childbearing within cohabitation), or whether municipalities with high (low) nonmarital childbearing are randomly distributed in space and do not show any spatial pattern.

Furthermore, we will be able to evaluate whether childbearing within cohabitation in a given municipality depends only on the characteristics that we identified as explanatory variables (the three SDT indicators, female educational expansion, women's empowerment, and economic difficulty) measured in the municipalities, or whether characteristics of neighboring municipalities also have an effect. We assume that these characteristics can also spread, and we are able to relate their diffusion to the diffusion of childbearing within cohabitation in two possible ways. First, childbearing within cohabitation in a given municipality increases over time because certain characteristics (its drivers) are becoming more widespread in the same municipality (direct effect). We can assume, although we cannot explicitly test this assumption given the nature of our data, that this change involves a diffusion of these characteristics both within and across social groups in a given municipality. According to this first diffusion mechanism, the characteristics of the forerunners spread within municipalities, hence leading to an increase in the population at risk of adopting new ideas and behaviors. Second, childbearing within cohabitation in a given municipality increases over time because its drivers are becoming more widespread in neighboring municipalities (indirect effect). In other words, new ideas and behaviors can spread even if the characteristics of the carriers do not spread because they are diffused by the forerunners across people and social groups. For example, a strong indirect effect of education would imply diffusion of childbearing within cohabitation in the reference municipality, resulting from having many highly educated women living in neighboring municipalities. Thus, indirect effects indicate which characteristics are important for the social interaction process leading to the spatial diffusion of childbearing within cohabitation.

Our analysis focuses on Norway, which together with the other Nordic countries is considered a forerunner of SDT (Van de Kaa 1987). Although its population is not large compared with other countries, it does make for an interesting case study given that it currently has one of the highest proportions of childbearing within cohabitation. In 1988, which is the starting point of our data (see the following section for details), the diffusion of childbearing within cohabitation had already started to the point that the majority (about 52 %) of first births occurred outside of marriage (i.e., to cohabiting and single mothers; 30 % and 22 %, respectively), whereas first births within marriage accounted for 48 % of the total. Based on Norwegian survey data, the proportion of children born within cohabitation is estimated to be about 7 % for the period 1979–1983, which means that a rapid increase occurred in the years just prior to our starting point. By the late 1990s, the percentage of first births to cohabiting mothers accounted for approximately one-half of the total first births (50 % in 1998), and it remained fairly stable until the late 2000s (53 % in 2011); first births to single mothers remained stable, at approximately 17 %, over the whole period.[1]

Data and Descriptive Findings

Statistics Norway provides municipal-level information about first births by union status of the mother, starting in 1987, from which we compute a measure of first childbearing within cohabitation versus marriage. More specifically, we compute the percentage of first births to (unmarried) cohabiting mothers over the total number of first births to married or cohabiting mothers in each municipality. Our panel comprises all 435 Norwegian municipalities[2] during the period 1988–2011. The choice of indicators for explaining childbearing within cohabitation derives from the discussion presented in the Background section, although our choice is affected by data limitations in some cases. To measure importance of religion, we use the percentage of representatives from the Norwegian Christian Democratic Party (KrF) in the municipal council. This political party favors religious values and traditions more than any of the other political parties, and strong support for this party is taken as a proxy for religiosity of individuals living in the municipality. The level of urbanization is approximated by the observed population density, constructed by dividing the number of inhabitants in the municipality by its area (measured in km^2). The old-age dependency ratio is calculated as the ratio between the population aged 65 and older on the population aged 15–64 (multiplied by 100). Female educational expansion is measured by the percentage of women aged 16 and older who achieved high education (corresponding to International Standard Classification of Education (ISCED) values 5–6). We measure women's empowerment as the percentage of female municipal council representatives. Finally, the male unemployment rate is computed as the percentage of men aged 16–66 who are unemployed on the total male work force of the same age. This variable is measured with one-year lag to reflect the association between economic difficulties and childbearing within cohabitation versus marriage at the time of conception rather than at birth. For this reason, our analysis refers to the period 1988–2011.

Figure 1 shows the evolution of first births by union status of the mother. The continuous line represents the percentage of first births within cohabitation versus marriage (i.e., our dependent variable, computed excluding births to

[1] We calculated these percentages from our original sample of municipalities and weighted them using the number of inhabitants. Thus, our numbers differ slightly from the official figures from Statistics Norway. However, because the official statistics include numbers only for the period 2001–2011, we use the data from our original sample when presenting the time trend for the whole country.

[2] During the period 1988–2011, administrative changes were aimed at reducing the overall number of municipalities, which has changed slightly from year to year. To have a balanced panel (which is necessary for our statistical analysis), we referred to the administrative subdivision that was in place at the beginning of the period we study (i.e., a total of 435 municipalities in 1988).

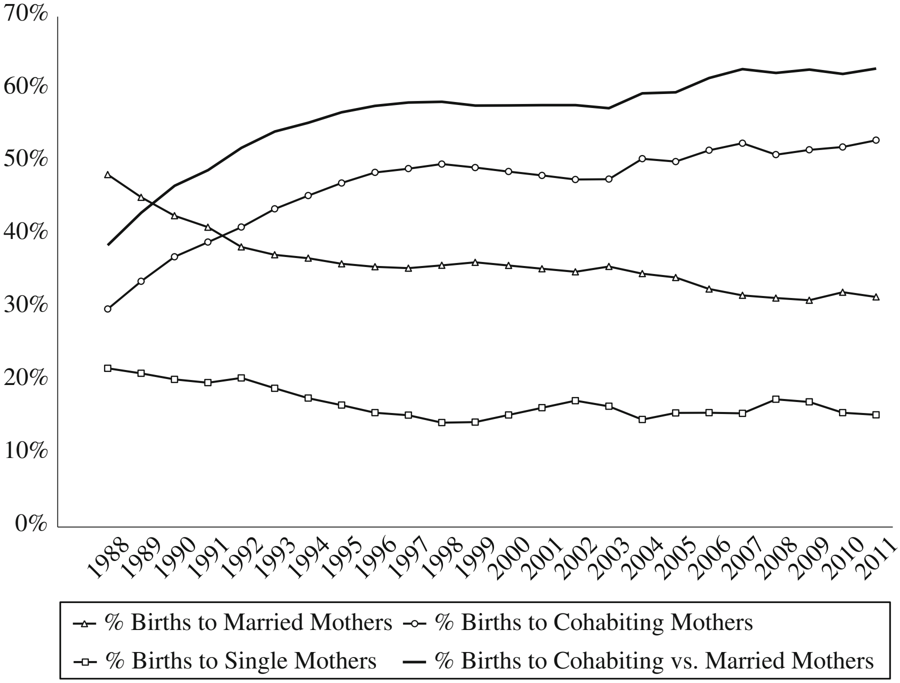

Fig. 1 Percentage of first births by union status of the mother (married, cohabiting, single) and percentage of first births within cohabitation versus marriage (dependent variable) across Norwegian municipalities, 1988–2011. Percentages are calculated using the number of inhabitants in the municipalities as weights

single mothers). Starting from an already high value of 39 % in 1988, first childbearing within cohabitation versus marriage reached 52 % in 1992 and continued to increase steadily until 1997 (58 %). It then continued to grow but at a much lower pace, reaching 63 % in 2011.

Whereas Norway and the other Nordic countries are often thought of as homogenous societies, this is not generally the case across the 435 municipalities, where substantial heterogeneity is evident. Figure 2 maps childbearing within cohabitation versus marriage in Norwegian municipalities in three periods: 1988–1991, 1998–2001, and 2008–2011. In the first period, the highest prevalence of childbearing within cohabitation is found in the northern part of the country. For several of the municipalities in the North, we find that more than 50 % of first births occurred within cohabitation rather than marriage beginning in the late 1980s. The South, instead, was characterized by a much lower prevalence (less than 25 %). Over time, however, several southern municipalities also reached a percentage of childbearing within cohabitation above 50 %; and by the last period (2008–2011), most of them had converged toward the national average.

The maps in Fig. 2 demonstrate the clusters of neighboring municipalities sharing high or low rates of first childbearing within cohabitation, which give rise to spatial dependence or spatial autocorrelation (Anselin 1988). Formally, the existence of spatial autocorrelation in childbearing within cohabitation can be tested using the Moran's I index, a global diagnostics tool for exploratory spatial data analysis that tests whether the value of a variable observed in a

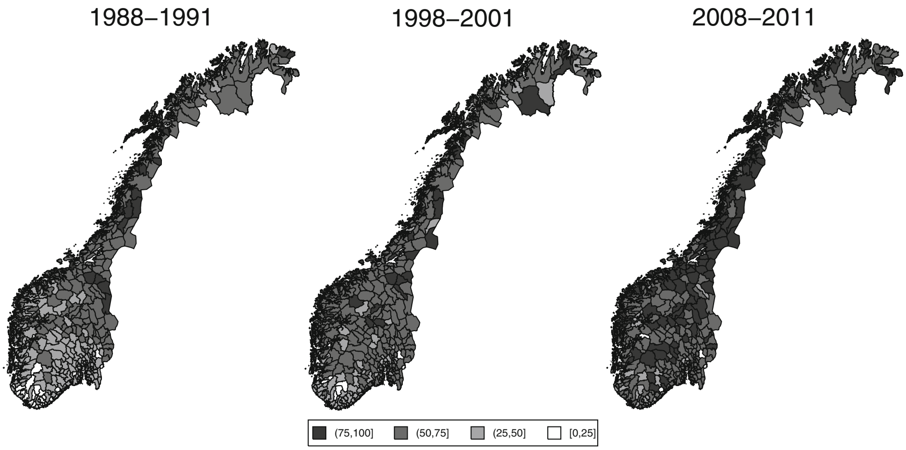

Fig. 2 Spatial distribution of childbearing within cohabitation versus childbearing in marriage. Descriptive statistics are calculated using the number of inhabitants in the municipalities as weights

given location is independent of the values observed in neighboring locations.[3] For our data, the Moran's I index equals .43 in the late 1980s and declines to .15 in the late 2000s. The fact that the Moran's I index declines over time is consistent with the innovation-diffusion perspective. Childbearing within cohabitation—that is, an innovation that was uncommon in the past—starts to emerge at some point in time but not simultaneously across the whole country. Rather, it is limited to selected clusters of "innovative" municipalities. Because our time series does not cover these first stages of the innovation-diffusion process, we cannot provide evidence of the diffusion since its onset. Nonetheless, in the late 1980s, the heterogeneity across municipalities was high, with clusters of municipalities where childbearing within cohabitation was already widespread as opposed to clusters where it was more rare (Fig. 2), yielding a high spatial autocorrelation. The spatial autocorrelation decreased over time as childbearing within cohabitation was adopted by more municipalities, consistent with the notion of a demographic transition (Lesthaeghe and Neels 2002; Lesthaeghe and Vanderhoeft 2001).

[3] The Moran's I index (Moran 1950) is formally described as follows:

$$I = \frac{n}{\sum_{i=1}^{n}\sum_{j\neq i}^{n} w_{ij}} \frac{\sum_{i=1}^{n}\sum_{j\neq i}^{n} w_{ij}(y_i-\bar{y})\left(y_j-\bar{y}\right)}{\sum_{i=1}^{n}(y_i-\bar{y})^2},$$

where y_i is the value assumed by the variable in the ith location, $\bar{y}$ is the sample mean, w_{ij} is the spatial weight assigned to the jth location, and n the number of spatial units (see the Methods section for a definition of spatial weight). Like the conventional correlation coefficient, the Moran's I index ranges between –1 (perfect negative spatial autocorrelation: e.g., a location with a high value of the variable is surrounded by locations with low values of the variable) and 1 (perfect positive spatial autocorrelation: i.e., similar values are clustered together in space). An index value close to 0 indicates random spatial distribution: that is, no spatial autocorrelation.

Figure 3 shows a scatterplot map of the local indicator of spatial association for 1988 (Anselin 1995).[4] This figure shows two distinct spatial regimes. Municipalities in the Center-North are defined as "high-high" areas in the sense that childbearing within cohabitation is persistently high across neighboring municipalities. Instead, municipalities in the South are defined as "low-low." The plot would suggest that the municipalities in the Center-North are the forerunners in the diffusion process of childbearing within cohabitation. The map also identifies *outliers*—that is, single municipalities with either high or low rates of childbearing within cohabitation but with neighboring municipalities that differ. In the legend, these are referred to as "high-low" or "low-high," respectively. This finding is consistent with Trost's (1978) results showing that the forerunners of childbearing within cohabitation in Norway were the least-educated women living in the northern regions, characterized by a low degree of urbanization. Also, results from cross-sectional spatial regression models (not shown) suggest a negative association of childbearing within cohabitation with education and urbanization but a positive association with economic difficulties during the late 1980s. Hence, the forerunner municipalities in childbearing within cohabitation were located in the North, were the most economically disadvantaged, and had the lowest educational attainment for women.

Figure 4 presents the spatial distribution and descriptive statistics of the explanatory variables that we use in the multivariate analyses. The values are the averages observed in the intermediate period of our panel (1998–2001). Figure 4 shows that spatial autocorrelation is also present in the explanatory variables: municipalities with high (low) prevalence of a certain indicator tend to be surrounded by municipalities with high (low) prevalence of the same indicator. For example, the Oslo area is characterized by high prevalence of women in politics, high educational attainment for women, low male unemployment rate, high population density, low importance of religion, and low old-age dependency ratio. Municipalities in the South of the country are characterized by high importance of religion.

Between 1988 and 2011, all these indicators changed. The average across all municipalities over time (weighting for the number of inhabitants, results not shown) shows that importance of religion declined (the percentage of representatives from the KrF in the municipal council declined from 9 % in 1988 to 6 % in 2011), as did the old-age dependency ratio (from 26 % in 1988 to 23 % in 2011), while population density increased (from 260 to 368 inhabitants per km^2). The percentage of women with high education and the percentage of women in politics increased substantially during the 24-year period, from 13 % to 31 % and from 35 % to 51 %, respectively. The time trend of male unemployment rate, instead, is dependent on the business cycle. It increased sharply from 1.3 % in 1988 to 5.2 % in 1994, after which it fell back to its original level and remained stable until 2011.

[4] Local indicators of spatial association allow the decomposition of global indicators, such as the Moran's *I*, into the contribution of each individual municipality. In this way, it is possible to identify local spatial clusters. To produce Fig. 3, we used the *spatlsa* command in Stata (Pisati 2001).

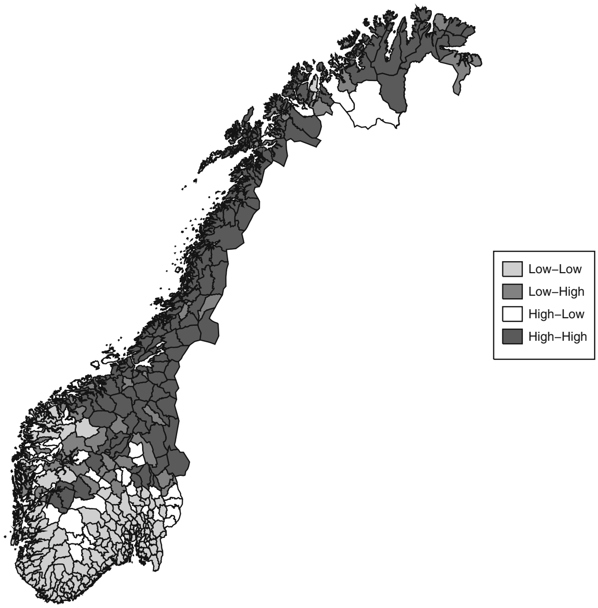

Fig. 3 Moran scatterplot map: Childbearing within cohabitation versus marriage, 1988

Method

The key novelty of our modeling scheme is that it allows for spatial autocorrelation in both the dependent and the key explanatory variables. Spatial autocorrelation in the dependent variable establishes the extent to which childbearing within cohabitation in any given municipality depends on childbearing within cohabitation in other neighboring municipalities, and consequently identifies whether the increase in childbearing within cohabitation is characterized by a process of diffusion. Assuming that significant spatial autocorrelation exists in the dependent variable, the autocorrelation on the explanatory variables enables us to disentangle the extent to which childbearing within cohabitation is driven directly from the municipality's own characteristics as well as indirectly from the characteristics of the neighboring municipalities.

We start by reviewing the fixed-effects panel model, which models a linear relationship between a dependent variable y_{it} (in our case, childbearing within cohabitation versus marriage) in municipality i and year t, and a vector of independent variables ($\mathbf{x}_{it}$), also measured in municipality i and year t. The model can be formally described as follows:

$$y_{it} = \mathbf{x}_{it}\boldsymbol{\beta} + \mu_i + \varepsilon_{it}, \tag{1}$$

where i indexes the municipalities ($i = 1, \ldots, N$) and t represents the periods ($t = 1, \ldots, T$). The vector $\mathbf{x}_{it}$ has dimension $1 \times k$, $\boldsymbol{\beta}$ is a matching vector of fixed unknown parameters, and μ_i denotes municipal-specific fixed effects, assumed to be constant over time and independent of the error term ε_{it}. Municipal-specific fixed effects control for unobserved time-invariant

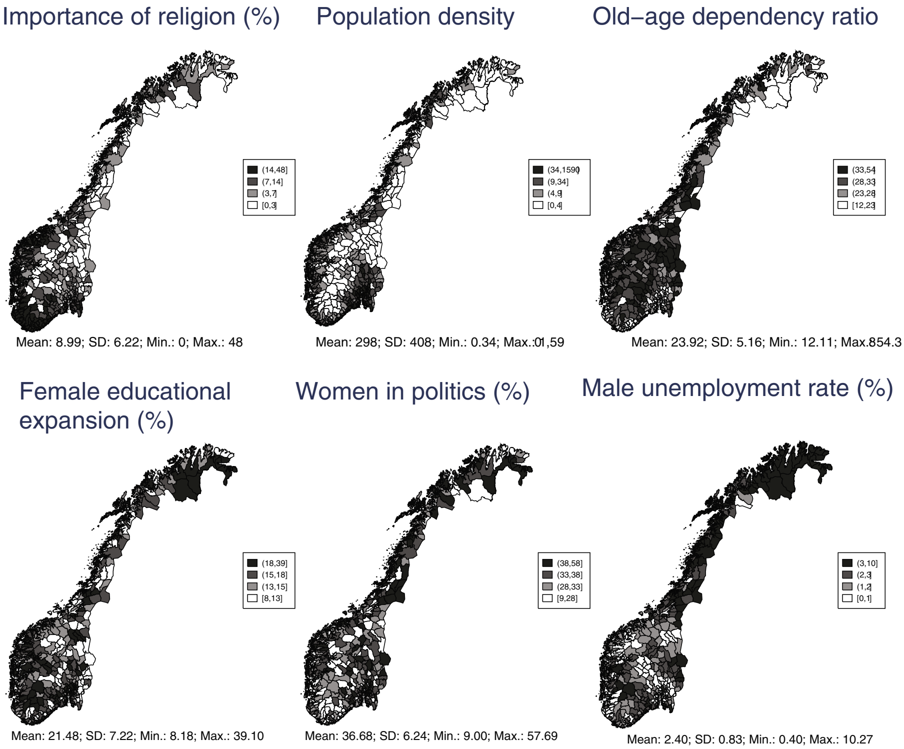

Fig. 4 Spatial distribution of explanatory variables, 1998–2001 (quintiles). Descriptive statistics are calculated using the number of inhabitants in the municipalities as weights

characteristics. The fixed-effects panel model presented in Eq. (1) assumes that the observations—that is, the municipalities—are independent. However, as the descriptive statistics in the Data and Descriptive Findings section show, this is an unrealistic assumption because the prevalence of childbearing within cohabitation is similar in neighboring municipalities. Consequently, the simple fixed-effects panel model in this setting would produce biased parameter estimates. Following the notation introduced by Elhorst (2010a), we expand Eq. (1) to account for spatial dependence in the dependent variable, obtaining Eq. (2):

$$y_{it} = \delta \sum\nolimits_{j=1}^{N} w_{ij} y_{jt} + \mathbf{x}_{it} \boldsymbol{\beta} + \mu_i + \varepsilon_{it}. \tag{2}$$

The resulting model is the spatial lag or spatial autoregressive (SAR) panel model (Anselin 1988). The key difference with respect to Eq. (1) is the introduction of the spatial lag ($\sum_{j=1}^{N} w_{ij} y_{jt}$) of the dependent variable that allows the percentage of first births in cohabitation versus marriage in municipality i and year t, y_{it}, to depend on the percentage observed in neighboring municipality j and year t, y_{jt}. The neighboring structure across municipalities is measured by the spatial weight, w_{ij}. The weight equals $1/\eta_i$ if $j \in N(i)$, and 0 otherwise, where $N(i)$ defines the set of all neighbors to municipality i, and η_i is the number of neighbors to municipality i. Neighbors are defined on the basis of a contiguity criterion, according to which two municipalities are neighbors if they share a border or an edge.

The scalar parameter δ is referred to as the spatial autocorrelation coefficient in the dependent variable. A positive estimate of δ means that childbearing within cohabitation in neighboring municipalities increases the likelihood of childbearing within cohabitation in the reference municipality. In other words, δ tells us whether a diffusion mechanism is in place, according to which childbearing within cohabitation spreads across municipalities, after we control for the characteristics of the carriers. In cross-sectional spatial models, δ portrays a spatial pattern reflecting a diffusion process (Baller et al. 2001; Messner and Anselin 2004). In spatial *panel* models, given the introduction of the time dimension, the interpretation of δ as reflecting diffusion is reinforced.

To account for spatial dependence in the independent variables, Eq. (2) can be further extended to include a spatial lag of the independent variables ($\sum_{j=1}^{N} w_{ij}\mathbf{x}_{ijt}\boldsymbol{\gamma}$). The resulting model is the spatial Durbin (SDM) panel model (Anselin 1988; Anselin et al. 2008):

$$y_{it} = \delta\sum_{j=1}^{N} w_{ij}y_{jt} + \mathbf{x}_{it}\boldsymbol{\beta} + \sum_{j=1}^{N} w_{ij}\mathbf{x}_{ijt}\boldsymbol{\gamma} + \mu_i + \varepsilon_{it}, \tag{3}$$

where $\mathbf{x}_{it}$ is the vector of independent variables measured in municipality i; $\mathbf{x}_{ijt}$ is the vector of independent variables measured in municipality j, both of dimension $1 \times k$; and $\boldsymbol{\beta}$ and $\boldsymbol{\gamma}$ are matching vectors of fixed unknown parameters. The key advantage of this extension is that the model now allows childbearing within cohabitation in each municipality i to depend on a set of independent variables measured in the same municipality ($\mathbf{x}_{it}\boldsymbol{\beta}$) as well as on an average of the same independent variables measured in neighboring municipalities ($\sum_{j=1}^{N} w_{ij}\mathbf{x}_{ijt}\boldsymbol{\gamma}$). In other words, the vector parameter $\boldsymbol{\gamma}$ reflects the extent to which childbearing within cohabitation in each municipality i is affected by characteristics averaged over its neighboring municipalities.

In opposition to the fixed-effects model, the parameter estimates in spatial models contained in the vector $\boldsymbol{\beta}$ cannot be interpreted as simply the marginal effect of a change in the explanatory variable on the dependent variable. Instead, as LeSage and Pace (2009) and Debarsy et al. (2012) observed, the total marginal effect is now a combination of direct and indirect effects, interpreted on the basis of the partial derivatives. To demonstrate this result, it is useful to express Eq. (3) in matrix notation (Elhorst 2010b):

$$Y_t = \delta\mathbf{W}Y_t + X_t\beta + \mathbf{W}X_t\gamma + v, \tag{4}$$

where $v = \mu + \varepsilon_t$.

Using the reduced form and leaving the subscript t aside for simplicity, we obtain the following expression:

$$Y = (I - \delta\mathbf{W})^{-1}(X\beta + \mathbf{W}X\gamma) + (I - \delta\mathbf{W})^{-1}v. \tag{5}$$

The matrix of partial derivatives of Y with respect to the kth independent variable in year t (Elhorst 2012), reads as follows:

$$\begin{aligned}\left[\frac{\partial Y}{\partial x_{1k}} \quad \cdots \quad \frac{\partial Y}{\partial x_{Nk}}\right] &= \begin{bmatrix} \frac{\partial y_1}{\partial x_{1k}} & \cdots & \frac{\partial y_1}{\partial x_{Nk}} \\ \vdots & \ddots & \vdots \\ \frac{\partial y_N}{\partial x_{1k}} & \cdots & \frac{\partial y_N}{\partial x_{Nk}} \end{bmatrix} = (I_N - \delta\mathbf{W})^{-1} \begin{bmatrix} \beta_k & w_{12}\gamma_k & \cdots & w_{1N}\gamma_k \\ w_{21}\gamma_k & \beta_k & \cdots & w_{2N}\gamma_k \\ \vdots & \vdots & \ddots & \vdots \\ w_{N1}\gamma_k & w_{N2}\gamma_k & \cdots & \beta_k \end{bmatrix} \\ &= (I - \delta\mathbf{W})^{-1}[\beta_k I_N + \gamma_k\mathbf{W}],\end{aligned} \tag{6}$$

where w_{ij} represents the (i,j)th element of the spatial weight matrix $\mathbf{W}$; and β_k and γ_k are the kth element of vectors $\boldsymbol{\beta}$ and $\boldsymbol{\gamma}$, respectively. The *own* derivative (i.e., the diagonal elements) for the ith municipality measures the impact of a change in the kth independent variable in municipality i on the dependent variable in municipality i. This impact, referred to as the "direct effect," includes feedback-loop effects. In fact, given that each municipality is considered its neighbors' neighbor, a change in the kth independent variable in municipality i affects the dependent variable in municipality i also through an effect going from municipality i to neighboring municipality j (and then back to i) via spatial autocorrelation on the dependent variable.

The important point here is that a change in the kth independent variable measured in a given municipality i can potentially affect childbearing within cohabitation in all other municipalities. In fact, a change in the kth independent variable in municipality i will also affect the dependent variable in neighboring municipality j as is expressed by the off-diagonal elements of the matrix of partial derivatives. This impact is referred to as the "indirect" or "spatial spillover effect." Indirect effects originate from neighboring municipalities when $\gamma \neq 0$, known as local effects, but also from municipalities that are not their own neighbors in so far as $\delta \neq 0$, known as "global effects." Because the direct, indirect, and total effects (i.e., the sum of the direct and indirect effects) differ across municipalities, summary measures are usually chosen to report their average effects (LeSage and Pace 2009). The average total effect is the average row sums of the elements of the matrix of partial derivatives in Eq. (6); the average direct effect is the average of the diagonal elements (own derivatives); and the average indirect effect is the average row sum of the nondiagonal elements (cross-derivatives; i.e., the difference between the average of all derivatives (the average total effect) and the average own derivative (the average direct effect).

More intuitively, the reported average direct effect measures the average impact of a change in a given explanatory variable in municipality i on childbearing within cohabitation in the same municipality, also taking into account feedback loops given that changes in nonmarital childbearing in municipality i influence nonmarital childbearing in its neighbors and vice versa. The average direct effect is the average impact on childbearing within cohabitation arising from a change in the explanatory variables: that is, what happens when the characteristics of the forerunners of the new demographic behavior spread among people in a given municipality. The average indirect effect instead measures the average impact on childbearing within cohabitation in municipality i ($i \neq j$) arising from a change in a given explanatory variable in *all neighboring* municipalities, and it gives a measure of the spatial spillover effect from neighboring municipalities. The indirect effect is a measure of the social interaction process occurring among people living in different municipalities. Finally, the average total effect is the sum of the direct and indirect effects; it measures the impact on childbearing within cohabitation in the average municipality resulting from changing a given independent variable by the same amount across all municipalities, taking into account both own-municipality effects and spatial spillover effects.[5] Alternatively, we can interpret the average total effect as the total cumulative impact on childbearing

[5] The model is estimated using the *xsmle* procedure in Stata (Belotti et al. 2013). Details about the estimation procedure can be found in Elhorst (2010a) and Lee and Yu (2010).

within cohabitation (on average) arising from a change in a given independent variable in municipality *j* (LeSage and Pace 2009).

We report standard errors and significance levels associated with the estimates. However, because our data refer to the whole population of Norwegian municipalities, they do not represent a random sample from a population unless we think of a super-population constituted of Norwegian municipalities in a longer period or of municipalities in different countries (Berk et al. 1995).

Results

Table 1 presents the results from the spatial Durbin panel model estimated for the entire period 1988–2011. The estimated spatial autocorrelation coefficient (δ) is reported at the bottom of the table. Without the explanatory variables, the spatial autocorrelation is estimated to be .25 (not reported), indicating spatial dependence of childbearing within cohabitation across Norwegian municipalities. In other words, an increase in childbearing within cohabitation in neighboring municipalities increases childbearing within cohabitation in the reference municipality. With explanatory variables, it is reduced to .14 (as reported in Table 1). Thus, for the given period, the explanatory variables explain approximately 50 % of the spatial autocorrelation of the dependent variable.

For each of the explanatory variables considered, we report the average *direct effect* (i.e., the average effect of a change in each indicator in a given municipality on childbearing within cohabitation in the same municipality, including feedback-loop effects), the average *indirect effect* (i.e., the average effect of a change in each indicator in all neighboring municipalities on childbearing within cohabitation in a given municipality), and the *average total effect* (i.e., the sum of the indirect and the direct effects).[6] These are all expressed as marginal effects. For completeness, we also report the estimated coefficients β and γ.

We start by commenting on the *direct* effects in Table 1. Importance of religion (as measured by the support to the KrF) has a negative but very small impact on childbearing within cohabitation, suggesting that in areas where religion is important, the new demographic behavior is less widespread. Population density (i.e., urbanization as measured by the number of inhabitants per km^2) does not appear to matter for childbearing within cohabitation: its coefficient is very small. This contrasts with the old-age dependency ratio, which has a high negative direct effect. The most likely explanation for this effect is that municipalities with a higher percentage of older people tend to hold traditional attitudes (also with regard to union formation), hence lowering the prevalence of childbearing within cohabitation. These municipalities may also be exposed to considerable selection in the sense that a high old-age dependency might be a result of high rates of out-migration of younger individuals, who may hold more modern attitudes toward union formation.

The percentage of women with higher education has a strong positive effect: a one-unit increase in the percentage of women with tertiary education in the reference municipality is associated with an increase of 0.616 in the percentage of women having

[6] Following the procedure described in LeSage and Pace (2009), we evaluate the statistical significance of the spatial direct and indirect effects using simulations to compute the standard errors.

Table 1 Results from fixed-effects spatial Durbin panel model, 1988–2011

	β	γ	Marginal Effects		
			Direct Effect	Indirect Effect	Total Effect
Importance of Religion	–0.069	–0.171	–0.074	–0.191†	–0.266*
	(0.058)	(0.109)	(0.057)	(0.114)	(0.126)
Population Density	–0.014	0.025	–0.013	0.025	0.012
	(0.017)	(0.021)	(0.017)	(0.021)	(0.018)
Old-Age Dependency Ratio	–0.205*	–0.378**	–0.219**	–0.429**	–0.648***
	(0.081)	(0.124)	(0.078)	(0.125)	(0.127)
Female Educational Expansion	0.597***	0.496***	0.616***	0.609***	1.225***
	(0.117)	(0.127)	(0.114)	(0.117)	(0.066)
Women in Politics	0.032	–0.104*	0.028	–0.104*	–0.076
	(0.025)	(0.047)	(0.024)	(0.049)	(0.055)
Male Unemployment Rate	–0.309	2.012***	–0.243	2.068***	1.825***
	(0.222)	(0.270)	(0.214)	(0.261)	(0.195)
Spatial Autocorrelation Coefficient (δ)	0.138***				
	(0.013)				

†p < .10; *p < .05; **p < .01; ***p < .001

a first child within cohabitation versus marriage in that same municipality. In other words, municipalities where women are, on average, highly educated tend to have a higher percentage of childbearing within cohabitation versus marriage. In contrast, a higher rate of women's involvement in politics at the municipality level has only a marginal positive effect on the rate of childbearing within cohabitation.

Leaving the discussion of direct effects of male unemployment rate for later, we next turn to *indirect* effects. The indirect effects tend to be larger than the direct effects. The indirect effects refer to the characteristics observed in other municipalities and measure the extent to which they matter for the dependent variable in any given municipality. To exemplify, the *direct* effect of –0.074 associated with importance of religion refers to the average effect of increasing the percentage of representatives from the KrF in a particular municipality. The indirect effect, on the other hand—which in this case is –0.191 and hence is stronger in magnitude—refers to what happens to the municipality-specific rate of childbearing within cohabitation, from increasing the percentage of representatives from the KrF *in all neighboring* municipalities (Parent and LeSage 2010). Hence, importance of religion has a small, negative indirect impact on childbearing within cohabitation. The indirect effect of population density (i.e., urbanization) is positive but very small. The indirect effect of the old-age dependency ratio, however, is negative and relatively strong.

As for education among women, we find that the indirect effect has about the same magnitude as the direct effect. Again, the estimate suggests that as education increases in the neighboring municipalities, it also increases the rate of childbearing within cohabitation for the reference municipality. The indirect effect of women in politics is negative. When the model is estimated without including female education and male unemployed rate (not shown), the indirect effect of women in politics is estimated to be positive. Similarly, the estimated direct effect of population density is estimated to be positive. The sign of these effects change when the additional independent variables are included in the model, suggesting possible multicollinearity.

We turn next to the rate of unemployment among men. The direct effect of men's unemployment is negative but small, but its indirect effect is positive and large. This means that as male unemployment increased in other municipalities, it also increased the rate of childbearing within cohabitation in the reference municipality. It is useful to consider more carefully the meaning of direct and indirect effects at this point. The large indirect effect can be interpreted as the effect of an overall increase (i.e., in all neighboring municipalities) in unemployment among men and therefore as a reflection of the effect from a more general economic recession. As we indicated earlier, childbearing within cohabitation versus marriage increased (and hence diffused) significantly in the period 1988–1997 (from 39 % to 58 %). From 1997 onward, the rate of childbearing within cohabitation stabilized apart from the very last few years of the period observed, where it again increased slightly. The period 1988–1997 was also a time in which unemployment fluctuated significantly: it rose substantially from 1988 onward, peaking at 1994, and then decreased again by the late 1990s. This is in contrast to the last period from the early 2000s to 2011, when the average country unemployment rate was stable. Given that these periods appear to differ in such important ways, it is useful to perform separate estimations for the two time segments. Tables 2 and 3 report the parameter estimates for the periods 1988–1997 and 1998–2011, respectively.

Table 2 shows that in the first phase of its diffusion, childbearing within cohabitation has two key drivers: higher education among women, and male unemployment rate. Both variables have a positive total effect on childbearing within cohabitation. Certainly, our estimates suggest that the period of increased unemployment among men as observed from 1988 to the mid-1990s indeed contributed to the diffusion of childbearing within cohabitation. The role of these two variables differs, however, in the sense that the former has a strong direct *and* indirect effect, whereas the male unemployment rate affects childbearing within cohabitation only through the indirect effect. Interestingly, before we include the explanatory variables in the model, the spatial autocorrelation in childbearing within cohabitation in this first period is estimated at 0.25 (not shown). After the variables are included, the spatial autocorrelation is close to 0, meaning that during this period in which the rate of childbearing within cohabitation increased, its diffusion is entirely explained by expansion of education among women and the male unemployment rate. In the full model, none of the other coefficients are significant for this period. Table 3, which presents estimates from the period 1998–2011, shows a different pattern. The effect of higher education among women still matters, but the coefficients are smaller in magnitude. Moreover, the indirect effect of the male unemployment rate is still positive albeit much smaller; the total effect, which is quite small, turns negative. Instead, we find importance of religion to matter more strongly through its indirect effect. This finding seems to indicate that very religious municipalities (i.e., municipalities with a high percentage of the population voting for the KrF) are the more resistant to the behavioral change. As before,

Table 2 Results from fixed-effects spatial Durbin panel model, 1988–1997

			Marginal Effects		
	β	γ	Direct Effect	Indirect Effect	Total Effect
Importance of Religion	–0.044	–0.243	–0.045	–0.236	–0.281
	(0.112)	(0.235)	(0.112)	(0.227)	(0.250)
Population Density	0.009	–0.001	0.009	–0.001	0.008
	(0.041)	(0.055)	(0.041)	(0.052)	(0.036)
Old-Age Dependency Ratio	0.105	–0.300	0.102	–0.284	–0.182
	(0.182)	(0.327)	(0.180)	(0.307)	(0.321)
Female Educational Expansion	1.701***	0.858**	1.709***	0.861***	2.570**
	(0.274)	(0.309)	(0.269)	(0.287)	(0.150)
Women in Politics	0.031	–0.116	0.030	–0.110	–0.080
	(0.039)	(0.076)	(0.038)	(0.071)	(0.079)
Male Unemployment Rate	–0.428	1.924***	–0.417	1.822***	1.405***
	(0.303)	(0.378)	(0.303)	(0.357)	(0.258)
Spatial Autocorrelation Coefficient (δ)	0.027				
	(0.022)				

$^{**}p < .01$; $^{***}p < .001$

we find no effect of population density and women in politics, whereas the old-age dependency ratio has both a direct and an indirect effect. The negative

Table 3 Results from fixed-effects spatial Durbin panel model, 1998–2011

			Marginal Effects		
	β	γ	Direct Effect	Indirect Effect	Total Effect
Importance of Religion	–0.140†	–0.601***	–0.160	–0.653***	–0.813***
	(0.085)	(0.155)	(0.084)	(0.163)	(0.182)
Population Density	–0.029	0.038	–0.028	0.035	0.007
	(0.035)	(0.042)	(0.034)	(0.042)	(0.043)
Old-Age Dependency Ratio	–0.389*	0.621*	–0.372*	0.596*	0.223*
	(0.171)	(0.263)	(0.166)	(0.266)	(0.275)
Female Educational Expansion	0.348	0.525	0.368*	0.598†	0.966***
	(0.216)	(0.235)	(0.206)	(0.222)	(0.131)
Women in Politics	0.030	–0.090	0.026	–0.090	–0.064
	(0.036)	(0.070)	(0.035)	(0.072)	(0.081)
Male Unemployment Rate	–0.774†	0.679	–0.757	0.594†	–0.163
	(0.403)	(0.509)	(0.395)	(0.497)	(0.405)
Spatial Autocorrelation Coefficient (δ)	0.135***				
	(0.017)				

$^{\dagger}p < .10$; $^{*}p < .05$; $^{***}p < .001$

direct effect of old-age dependency makes sense: municipalities with a higher dependency ratio will consist of individuals who, on average, have more conservative attitudes, as reflected by the lower rate of childbearing within cohabitation. The positive indirect effect of old-age dependency may instead be a reflection of migration across municipalities, which is in part a selection effect because migration of younger individuals toward neighboring municipalities will both increase the dependency ratio in the municipality of origin and increase childbearing within cohabitation in the receiving municipality. Because the indirect effect is stronger than the direct one, the total effect of the old-age dependency ratio in the second period is positive.

In addition to the models reported in Tables 1, 2, and 3, we estimate a range of alternative models as a means of robustness checks (not shown). These include models in which the 5 % smallest municipalities were excluded and models in which all islands were excluded. Several municipalities are contained within islands (or a group of islands), and as such they do not have a land border with other municipalities. The appropriate definition of what constitutes a bordering municipality in this case is not quite clear. Whereas most of the islands are close to the mainland, others are not. Results are robust when the small municipalities and islands are excluded. Likewise, the estimates remain robust when we exclude Oslo from the estimation. Oslo, the capital city, is by far the largest municipality in terms of population and has a much higher level of urbanization than any of the other municipalities. Results are also robust to a different specification of area units that considers 89 economic regions (Local Administrative Units, LAU-1) instead of municipalities.

Conclusion

In this article, we studied the diffusion of a new demographic behavior—childbearing within cohabitation—across Norwegian municipalities over a 24-year period (1988–2011). Our contribution lies in the way we model the spatial autocorrelation of childbearing within cohabitation, using a spatial panel Durbin model. The results give support to the underlying idea of the SDT that new behaviors spread over time and across space. The key benefit of our analysis, however, is seen through the spatial autocorrelation of the explanatory variables. Here we see more mixed support for SDT as a theory underlying behavioral change. During the last decades, Norway has clearly experienced a strong trend of secularization, but we find very limited support for the importance of religion being a driver of childbearing within cohabitation. Only in the last period (1998–2011) do we find an effect of religion, but this is a period in which childbearing within cohabitation did not increase any further. In this period, religion certainly explains part of the observed variation in childbearing within cohabitation across municipalities, suggesting that religiosity is an important hindrance for childbearing taking place within cohabitation in Norway. On the other hand, we do not have strong evidence to suggest that religion affected its diffusion, which by and large took place prior to 1998. We find even less support for urbanization being an underlying driver behind childbearing within cohabitation. Insofar as population density serves as a suitable proxy of SDT, we find no strong effect in either of the two periods considered. On the other hand, we do find an effect of the age structure. Those municipalities with a high old-age dependency ratio have lower rates of childbearing within cohabitation, consistent with a presumption that the older population holds more conservative attitudes in general. We do not find any effect of women's empowerment as measured by women in politics.

The key driver behind childbearing within cohabitation comes from increased education among women. We find that both the direct and indirect effects of women's education are very strong. Thus, childbearing within cohabitation is higher in municipalities where women have high average educational levels. More importantly for explaining the diffusion, however, is the very strong indirect effect of women's education: as education spreads (i.e., expands) among women in neighboring municipalities, childbearing within cohabitation in the reference municipality also increases. It is well established that education fosters modern and less-traditional attitudes (Aassve et al. 2013), and insofar as the new demographic behavior is a result of new ideas and attitudes penetrating through society, our analysis suggests that education is the principal driver. In addition, one cannot ignore the role of the male unemployment rate, which we find to have a very strong indirect effect in the period in which childbearing within cohabitation was diffusing. This result gives some support to the idea of a pattern of disadvantage in which nonmarital childbearing is argued to be more common among the lower socioeconomic strata (Perelli-Harris et al. 2010). Whether one can speak of a pattern of disadvantage in Norway is, however, not so clear. In the 1980s and 1990s, Norway experienced an increase in the unemployment rate, which, we have argued, contributed to the diffusion of childbearing within cohabitation. However, unemployment has no explanatory power in the second period we explore, indicating that childbearing within cohabitation is not necessarily more prevalent in poorer municipalities. One may, of course, question the extent to which the unemployment rate reflects a pattern of disadvantage. The key argument of the pattern of disadvantage is that there is an *educational gradient* with respect to cohabitation. For the United States, it is well established that cohabitation is more prevalent among those with lower education (see, e.g., Upchurch et al. 2002). Here, in contrast, we find very robust evidence for higher education (not lower education) being a driver for higher rates of childbearing within cohabitation. This is not to say that high educational levels do not correlate negatively with childbearing within cohabitation. Indeed, cross-sectional versions of our estimation, although not included in our results, show a negative correlation between educational level and childbearing within cohabitation; thus, at any point in time (although it weakens over time), childbearing within cohabitation is more prevalent among those with lower education. However, on the basis of cross-sectional analyses, we cannot say that lower education *leads to* more childbearing within cohabitation. The fixed-effect estimation, controlling for unobserved characteristics, shows that the opposite is the case. Our analyses suggest that whereas childbearing within cohabitation may have started with those having lower education, its diffusion resulted from an expansion of education.

The method adopted in this analysis is extremely useful in enabling identification of the factors (and quantification of their magnitude) explaining how diffusion of a phenomenon takes place. The present analysis does not come without caveats, however. Because new phenomena often materialize over a long period, research on these topics also requires long time series. Here, we have information over a relatively long period of 24 years, but childbearing within cohabitation appeared in some municipalities much earlier, and we consequently cannot observe the complete process. Instead, we have a snapshot (albeit a long one), and although we make some indication to the geographical location where the phenomenon may have started, we cannot be entirely sure about that. Moreover, one important finding of our analysis is that the covariates appear to play different roles according to the stage of the diffusion considered. This also means that the relevance of the theoretical arguments for explaining new behaviors also depends on the stage of the diffusion.

The spatial approach reflects the idea that a new behavior spreads geographically, and the phenomenon of interest is assumed to be influenced by the characteristics of the neighboring units. This may seem a strong assumption given that new behaviors may spread through several channels and that in the digital age, networks may no longer be limited to geographic neighbors. Also, with municipal-level data, we are not able to account for diffusion mechanisms taking place among socioeconomic groups residing in the same or in neighboring municipalities. Finally, we are considering individual behaviors, but our data are aggregated up to the municipal level, creating the potential for ecological fallacy; our analysis is not necessarily robust in this respect. Still, in the absence of network data on reproductive behaviors describing interactions among people located across a society, municipal-level data offer an appropriate approximation and offers the possibility to study the diffusion of a new demographic behavior and its drivers across space and over time.

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