However Far Away? The Spatial Contingencies of Assortative Mating

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Abstract

This study reconsiders a classic sociological question: how does space shape intimate ties? Specifically, it examines how the spatial segregation of ethnic groups contributes to patterns of ethnic endogamy. It does so by applying conditional logit models to Swedish population registers describing couples who began cohabiting between 1990 and 2017. The models compare observed unions to counterfactual unions drawn from available singles, distinguishing the effects of ancestry assortativity—based on country of origin—from residential propinquity, while controlling for matching on nativity, education, and age. Proximity strongly predicts partnering, but assortativity matters too, particularly for non-Western groups. Mediation analysis shows that failing to account for propinquity overstates endogamy by 20–40 percent for these groups, with stronger mediation for the most segregated groups. The findings suggest that segregation complements ethnic boundaries in the short term, but also suggest how integration may undermine group boundaries in the longer term.

Introduction

Patterns of endogamy and residential segregation serve as dual signals of boundaries between racial and ethnic groups (Kalmijn 1998; Kim and White 2010; White et al. 2005). Across times and places, and for socially salient racial and ethnic boundaries, people are more likely to form marital and cohabiting relationships within racial and ethnic boundaries and are more likely to reside in neighborhoods where their own group has greater representation. Classic sociological theories have posited a close link between assortative mating and segregation, emphasizing how the geographic distribution of groups shapes mating opportunities (Blau 1977a, 1977b; Blau et al. 1982). Indeed, marriages in early 20th century cities were often contracted between people living within walking distance (Abrams 1943; Bossard 1932; Clarke 1952; Davie and Reeves 1939; Marches and Turbeville 1953). But while the effect of neighborhood-to-neighborhood variations in segregation patterns received considerable attention in the mid-20th century (Kennedy 1943; Koller 1948; Morgan 1981; Peach 1974; Ramsøy 1966; Schnepp and Roberts 1952), contemporary analyses of assortative mating typically focus on "macro" segregation across larger sub-national geographies, e.g., provinces and counties (Eckhard and Stauder 2019; Harris and Ono 2005; Kalmijn and Van Tubergen 2010; Lichter et al. 2015; Lievens 1998). However, there has been a resurgence of interest in the role of smaller-scale geographies in patterns of racial and ethnic assortative mating (Choi and Tienda 2017b; Feng et al. 2010; Jarvis et al. 2023; Kalmijn and Flap 2001; Logan and Shin 2012; Mood and Jonsson 2025). The study presented in this paper contributes to this renewed line of research, examining how spatial distances between groups with cities relate to endogamy by immigrant background in Sweden.

The spatial distribution of populations, even at small geographic scales, conditions three proximate drivers of assortative mating: opportunities, preferences, and third-party influences (Blau 1977a; Kalmijn 1998). Or, if we use the metaphor of a partner market, geography shapes the accessible supply of partners with specific sets of traits, the demand for partners dictated by preferences, and the activities of intermediaries who facilitate meeting and matching (Schwartz 2013).

On the supply side, people typically conduct their lives in activity spaces centered around a place of residence (Jones and Pebley 2014). They meet others in social foci—workplaces, schools, places of worship, bars, etc.—that draw upon local catchments (Feld 1981; Puur et al. 2022), increasing interactions between singles who live near each other (Blau 1977a). On the demand side, geography matters because people often prefer partners who live nearby, not least because it simplifies the logistical challenges of dating. Finally, geography factors into intermediaries' recruitment of clientele and the choices they make on their behalf. This includes online dating platforms' explicit use of proximity in their recommendation algorithms (Tinder 2022). In short, the partner market is also a partner landscape.

In many Western contexts, this partner landscape is characterized by persistent racial and ethnic segregation (Elbers 2021; Jarvis et al. 2017; Malmberg et al. 2018; Musterd 2005). Singles live in closer proximity to those from their own racial and ethnic groups than to other singles. Because space conditions opportunities to meet and match, spatial *propinquity* between co-ethnics can induce endogamy even absent any explicit preferences to partner within ethnic boundaries. Our study asks how strong this propinquity effect is, whether propinquity can account for a portion of observed ethnic endogamy, and whether this changes our view of the relative strength of endogamy for different ethnic groups.

Our study addresses these questions using discrete choice models of union formation (Gullickson 2021; Haandrikman and van Wissen 2012; Jepsen and Jepsen 2002; Nielsen and Svarer 2009). Our approach builds on existing research examining how estimates of endogamy are affected by the sub-national and sub-regional geographic distribution of potential partners (Choi and Tienda 2017b; Harris and Ono 2005; Järv et al. 2015; Lievens 1998; Logan and Shin 2012; Mood and Jonsson 2025; Peach 1980). However, unlike studies that assess the spatial availability of co-ethnic partners primarily at the macro-level, we assess availability at the micro-level based on distances to extant potential partners. Our multinomial models treat each distinct, opposite-sex single as a prospective partner. We estimate the probability of forming a union with a particular opposite sex single as a function of the similarity between the focal partner and all other potential partners, including the realized partner. This modeling approach allows us to account for (1) the joint, marginal distributions of partner characteristics; and (2) the matching between egos and prospective partners on multiple social dimensions, especially ethnicity and proximity, but also age, education, and nativity. We then assess whether and to what extent propinquity mediates endogamy using methods developed for non-linear models (Karlson et al. 2012).

We perform this analysis using 1990–2017 Swedish population registers. These longitudinal data allow us to identify individual immigration backgrounds, socio-economic statuses, the timing of their entry into cohabitation or marriage, and, critically, the precise residential locations of (potential) partners. Beyond furnishing detailed data, Sweden is also a compelling place to study assortative mating because it is a microcosm of the social and demographic changes affecting many mature, post-industrial Western democracies. In terms of family formation, Sweden has a reputation as a "family forerunner", typically experiencing changes—e.g., the rise of non-marital cohabitation—earlier than its Western peers (Ohlsson-Wijk et al. 2020). In terms of ethnicity, Sweden's accommodation of refugees and political asylum seekers has substantially increased its diversity since the 1980s. Approximately 30% of Sweden's population now is either foreign born or has at least one foreign born parent, many with non-European backgrounds (Statistics Sweden 2023). Alongside increasing diversity, ethnic segregation has also reached levels similar to those of European peers (Jarvis et al. 2017; Malmberg et al. 2018; Musterd 2005, 2020). We expect that whatever relationship between propinquity and endogamy we uncover will resemble what we would discover in other contexts.

To preview, we find an independent propinquity effect for all groups in our study, where we distinguish people not by ethnicity, but by the closely related concept of ancestry, coded based on an individual's own or their parent's country of origin. The propinquity effects are somewhat weaker for Non-Western immigrants and their descendants. While we find strong endogamy for non-Western groups, both before and after controlling for propinquity, we also find that models without propinquity overstate the odds of endogamy by 20-40% on average, and by more than 70% for some groups. For those of Swedish and Western ancestry, we find a different pattern: endogamy is weakly positive, but appears stronger after we control for the proximity of different partners. Our results imply that while segregation—i.e., the spatial distances between groups—partly mediates assortative mating for non-Western groups, Westerners are more likely to partner with out groups in part because of (limited) spatial integration. Once we account for this spatial integration, their tendency to partner with their own group becomes more evident. However, compared to the effects of assortativity *per se*, the effect of propinquity on endogamy *rates* is only modest, accounting for less than 10% of endogamy

In the remainder of the paper, we review theoretical models of assortative mating and empirical evidence about the extent of ethnic endogamy, patterns of segregation, and the relationship between the two, with special attention paid to the Swedish case. We then outline our modeling framework and analytic strategy, present the results, and discuss their implications for understanding the spatial foundations of ethnic boundaries.

Background

Homogamy is a recurring feature of cohabiting and marital unions. Setting aside the fact that a preponderance of unions are different-sexed, for many other indicators of group membership and social statuses—including education (Blossfeld 2009), income (Schwartz 2010), ethnicity (Kalmijn 1993), and religion (Kalmijn 1991)—unions between similar people are more likely than unions between dissimilar people. This positive assortativity is remarkably robust across times and places (Kalmijn 1991; Mare 1991; Schwartz 2013). Three groups of factors are often invoked to explain these patterns of homogamy: individual preferences, opportunity structures, and third-party influences (Kalmijn 1998; Lichter and Qian 2019). Empirically, third party influences are often unobserved and are analytically subsumed by preferences and opportunities. In any case, these are integrally related to geography.

While people express preferences for numerous partner attributes—among them age, social class, religion, and physical attractiveness (South 1991)—our focus here is on ethnic and racial preferences. Attitude surveys provide the clearest evidence for ethnically and racially homogamous preferences. White Europeans have tended to be resistant to or lukewarm about forming unions with immigrants, particularly those from predominantly Islamic sending countries (Carol 2013; Huijnk et al. 2010; Storm et al. 2017). To the extent that immigrants—primarily Muslim immigrants from the Middle East—have opposed intermarriage, it has been largely on religious grounds (Carol 2013; Huijnk et al. 2010). In Sweden, Whites have expressed openness to intermarriage, but have been slightly less open to unions with non-Western groups, particularly those from Africa, the Middle East, and Eastern Europe (Osanami Törngren 2016). Behavioral evidence from online and offline dating studies mirrors these attitudes: people tend to initiate and reciprocate contact within their own group, and show an adherence to racial and ethnic hierarchies similar to those observed in survey data (Feliciano and Kizer 2021; Fisman et al. 2008; Lewis 2016; Potârcă and Mills 2015).

Across Western contexts, evidence based on observed unions largely conforms to preference data: people are more likely to form unions with partners from their own group, but also conform to an ethnic hierarchy with native Whites perched at the top (Choi and Tienda 2017a; Kalmijn 1994; Kalmijn and van Tubergen 2006; Qian 1997; Qian et al. 2012; Qian and Lichter 2007). In Sweden, exogamy appears to be least likely among those of African, Middle Eastern, and Central Asian backgrounds, and this is partly explained by cultural differences between Sweden and these sending contexts (Dribe and Lundh 2009, 2011). Exogamy is more likely among those with Western backgrounds with countries of origin that are more culturally similar to Sweden (i.e., immigrants or descendants of immigrants from countries in the EU/EEA, North America, or Oceania). Given that our study will use the same register data used in prior Swedish studies, we formulate our expectations in two hypotheses:

> *Hypothesis 1A:* All groups will be endogamous after controlling for the opportunity structure.

> *Hypothesis 1B:* Endogamy will be strongest for immigrant groups tracing origins to countries that are most culturally different from Sweden.

However, inferring endogamous behaviors based on observed unions requires accounting for the opportunity structures that constrain union formation (Qian 1997; Qian and Preston 1993; Schoen 1981). Opportunity structures are the social conditions that shape the likelihood of encountering partners with different social attributes and backgrounds, before people can exercise choice (Blau 1977a; Schwartz 2013). Even strong preferences for exogamy will fail to produce intergroup unions if opportunity structures preclude encounters between partners from different groups.

Relative groups sizes are key features of the opportunity structure (Blau 1977a). Majority and minority group members alike are more likely to encounter those from the majority purely by virtue of its numerical superiority. All else equal, this should increase rates of majority-majority endogamy and majority-minority exogamy. Minority groups, conversely, face a mathematical constraint—they have fewer co-ethnic partners available—decreasing the likelihood of endogamy. Empirical research addresses group size imbalances in a variety of ways. This includes controlling for group sizes at the level of national or regional marriage markets directly, or using related indicators, like group specific sex ratios or measures of heterogeneity (Blau et al. 1982; Blau and Schwartz 1984; Choi and Tienda 2017b; Fossett and Kiecolt 1991; Harris and Ono 2005; Okamoto 2007; Schoen 1981). This research consistently shows that differences in group sizes are key variables that predict endogamous and exogamous union formation. For minority groups, smaller relative sizes are associated with lower rates of endogamy and higher rates of exogamy.

Marriage market opportunities also depend on how people are spatially distributed within and across regions. In fact, many studies of assortative mating disaggregate analyses to the level of provinces, counties, or cities, controlling for group sizes within these (dis)aggregates in explicit recognition of the importance of local, rather than national, marriage market compositions (Abramitzky et al. 2011; Blau et al. 1982; Choi and Tienda 2017b; Harris and Ono 2005; Stier and Shavit 1994). This conforms to the principle of propinquity, or the tendency of people to interact more with those who are geographically closer (Small and Adler 2019). It also aligns with early studies that established a strong link between geographic proximity and partner choice (Bossard 1932; Davie and Reeves 1939; Ellsworth 1948; Marches and Turbeville 1953), a finding consistently supported across different contexts and time periods (Clegg et al. 1998; Coleman 1979; Coleman and Haskey 1986; Ekamper et al. 2011; Haandrikman et al. 2008; Haandrikman et al. 2011; Küchemann et al. 1974; Peach 1974). Even with the rise of online dating, which has purportedly loosened geographic constraints, most couples still form within relatively short distances (Bruch and Newman 2019; Haandrikman 2019), highlighting propinquity's continued importance in partner markets.

Propinquity shapes partner choices primarily because people face logistical constraints and live geographically circumscribed lives, but also because of subjective feelings of attachment and belonging. People move within activity spaces that are often centered on a place of residence, and within those activity spaces transit between social foci—including neighborhoods, schools, workplaces, restaurants, etc.—that draw on local catchments (Feld 1981; Jones and Pebley 2014; Oldenburg and Brissett 1982; Small and Adler 2019; Whyte 1980). People are more likely to meet when their activity spaces overlap and their social foci coincide (Feld 1981; Huckfeldt 1983; Kalmijn and Flap 2001; Puur et al. 2022). At the same time, proximity reduces the logistical costs of dating and relationship maintenance, making geographically close partners more viable and preferable (Haandrikman and van Wissen 2012; Lewis 2016; Zhang et al. 2022). Finally, proximity can reflect a sense of place identity and attachment among in-groups of "locals". This can manifest as affinities between local singles sharing the same place identity (Haandrikman and Hutter 2012; Hummon 1992; Lewicka 2011). This leads us to our Hypothesis 2:

> *Hypothesis 2:* Proximity increases the likelihood of union formation. Conversely, the further away a potential partner is, the lower the likelihood of forming a union with that partner.

Beyond its independent role in union formation, propinquity can also be expected to mediate the relationship between ethnic segregation and endogamy. The residential segregation of racial and ethnic groups is an aspect of the spatial opportunity structure that is particularly salient to patterns of intermarriage and endogamy. This segregation manifests both between (macro) and within (micro) cities (Lichter et al. 2015). Macro-segregation, captured by regional or metropolitan differences in relative group sizes, has been an enduring concern in studies of intermarriage. But studies are increasingly also adjusting for intra-regional measures of segregation (Choi and Tienda 2017b; Logan and Shin 2012; Okamoto 2007). This acknowledges that not only do different racial and ethnic groups live in different cities, but also they live in different neighborhoods within these cities. Measures of segregation are integrally related to distance: high levels of segregation imply greater spatial distances between groups (Massey and Denton 1988; Reardon et al. 2008; White 1983). Combined with the principle of propinquity, this implies that segregation will increase intra-group exposure and reduce inter-group contact. Correspondingly, higher levels of racial or ethnic segregation are associated with lower levels of intermarriage (Choi and Tienda 2017b; Logan and Shin 2012; Okamoto 2007).

However, segregation statistics are macro-level measures that offer noisy assessments of opportunity structures at the micro-level. Even in highly segregated populations, there will be some minority individuals who are spatially integrated with the majority, and some from the majority will live in or near neighborhoods with high minority representation. This implies measurement error at the micro-level that may lead us to understate the importance of spatial propinquity vis-à-vis ethnicity or race. Controlling for relative groups sizes at the level of the neighborhood (e.g., Feng et al. 2010) is an imperfect patch. Salient marriage market boundaries need not coincide with the boundaries of administrative neighborhoods.

An alternative is to do away with neighborhood boundaries and measures of segregation entirely, and instead to consider how each single is spatially situated relative to other singles. The spatial availability of *each potential partner* can be assessed directly using geographic distance. This treats distance as one explanatory variable among many shaping who partners with whom. Because singles with differing racial and ethnic backgrounds are segregated—and hence more spatially distant from each other—we expect proximity to mediate the relationship between ethnicity and partner choice. We formalize this in a third hypothesis:

> *Hypothesis 3A:* For all groups, estimates of ethnic endogamy will decline when controlling for proximity to potential partners.

We expect propinquity to mediate patterns of ethnic homogamy only to the extent that groups are residentially segregated. Compared to those who are more residentially integrated, members of more segregated groups will have fewer opportunities to meet singles from out-groups. As a result, propinquity will reinforce ethnic endogamy more strongly among highly segregated groups. In Sweden, immigrants from Africa and the Middle East, especially from Eastern Africa, are the most segregated, immigrants from Western countries (i.e., the EU/EEA, North America, and Oceania) are the least segregated, and immigrants from Asia, Latin America, and Eastern Europe fall somewhere in between (Jarvis et al. 2017; Malmberg et al. 2018). This leads us to a corollary to our third hypothesis.

> *Hypothesis 3B:* The mediation of endogamy by propinquity will be strongest for the most segregated groups.

Complementary to the mediation, it is also important to understand how propinquity contributes to ethnic endogamy at the population scale. Even if propinquity exerts strong effects at the micro-level, aggregate endogamy rates depend on population structure and the relative importance of matching on other traits. While propinquity and segregation may enhance endogamy, their contribution to overall rates of endogamy may be small if the endogamous "forces of attraction" *per se* are sufficiently large. And existing evidence suggests these forces of attraction are quite strong in Sweden (Mood and Jonsson 2025). This leads us to our final hypothesis:

> *Hypothesis 4:* Propinquity increases rates of ethnic endogamy at the aggregate level, but its influence is smaller than that of assortativity *per se*.

To test the above hypotheses, we combine detailed register data with the methods of discrete choice modeling. The register data allow us to define bespoke opportunity structures within which we can assess the distance between potential partners and their similarities in terms of immigration backgrounds. The discrete choice models allow us to consider the relative importance of propinquity and immigrant backgrounds in driving union formation at the level of dyads. We detail these data and methods below.

Data

Our study uses comprehensive data from the Swedish full population registers provided by Statistics Sweden (SCB), covering 1990–2022. During this period we observe precisely where people are living, including property identifiers and positions on a $100\,m \times 100\,m$ grid. We use a retrospective approach to identify the onset of cohabitation among those who eventually marry or have children together. Starting from the full population of couples that are ever observed to marry or co-reside with common children, we backtrack through residential records to identify the first year when both partners were registered as living on the same property. We say that a couple forms a union in the interval $[t, t+1]$ when the partners are living at different properties in year t, but living on the same property in year $t + 1$.

We restrict the analysis to unions that began cohabiting between 1990 and 2017. We use 1990 as the starting year because it is the earliest that we observed locations geocoded at $100\,m$ resolution. We use 2017, rather than 2022, as the endpoint because of our retrospective approach. Eligible couples are defined based on childbearing and marriage that can occur up to and including 2022. But many cohabitations initiated after 2017 will not have had sufficient time to progress to marriage or childbearing by 2022. And what cohabitations we can identify after 2017 are likely to be highly selected on attributes—e.g., religiosity—that predict fast or even concurrent transition into marriage or childbearing. Excluding cohabitations initiated after 2017 addresses this selective right-censoring problem.

Our approach to identifying cohabitation involves a number of omissions. We do not identify couples who never go on to marry or have common children. We also exclude marriage migrants, i.e., when one or both partners migrates to Sweden to initiate co-habitation (Niedomysl et al. 2010). We also exclude unions when partners lived in different counties (Län) at time t, since our analysis targets intra-metropolitan union formation—i.e., between people already living in the same county, covering 80% of new unions. Sensitivity analyses indicate that the omission of between-county unions does not substantially affect our results. Lastly, we exclude same-sex unions because of their relative rarity and the difficulty in defining the relevant partner market. This leaves us with $1,030,231$ unions formed in Sweden between 1990 and 2017.

We prepare our data and define the outcome for statistical modeling by constructing two dyadic, person-year-alternative data sets, one from the perspective of women initiating cohabiting relationships, and one from the perspective of the men who they partner with, following Gullickson (2021). Each woman (man) contributes an observation corresponding to the year, t, immediately preceding the onset of cohabitation at $t + 1$. This assumes that the start of cohabitation aligns with the conditions of the partner market at the time of their meeting. Within each person-year, we then add rows representing different alternative or counterfactual partners. We define potential, alternative partners in a way complementary to how we identify cohabitation. Potential partners must be (1) of a different sex, (2) living in the same Swedish county at t, and (3) single (i.e., not living on the same property as any person they ever marry or cohabit with). We further exclude individuals younger than 16 years old in any given year to ensure that all potential partners are of eligible partnering age. Additionally, we exclude potential partners who are more than 20 years younger or older than the focal woman (or man), and those younger than 16. We fill out each person-year with 101 potential partners, including the chosen partner and 100 randomly sampled singles. The outcome variable is then a dummy that's coded 1 for the observed partner and 0 for all other partners.

Independent variables

Our independent variables are either alternative or dyad specific. They represent the features of potential partners and their similarities or differences from the focal woman or man entering into a cohabiting union in the t to $t+1$ interval. Our main independent variables relate to ethnic endogamy and propinquity, while our control variables relate to age and educational assortativity.

Our register data lack information about subjective ethnic identifications. We therefore use *ancestry*, derived from country-of-birth data, as a proxy for ethnicity, acknowledging that this approach imperfectly captures the social meaning of ethnic identity. In any case, we know the specific country of birth for people from large sending countries, while less common origins are aggregated by Statistics Sweden into regional categories. Anyone born in Sweden to exclusively Sweden-born parents is coded as having Swedish ancestry. Anyone with a foreign-born parent is assigned a non-Swedish ancestry. We assign ancestry based on the mother's country-of-birth if parents have different foreign origins. If we lack information on parents, then we assign ancestry based on the person's own country of birth.

This results in a set of 50 "micro" ancestry categories, that we also aggregate, based on geographic proximity and cultural similarity, into 16 meso-ancestries, and then six macro-ancestries. The latter includes (1) Sweden; (2) Western/Northern/Southern Europe, North America, and Oceania; (3) Eastern Europe and the Balkans; (4) Asia; (5) Latin America and Caribbean; (6) the Middle East and Africa. To capture national, pan-national, and pan-ethnic assortativity (Okamoto and Mora 2014), we code each dyad based on whether the (potential) partner shares the same micro- or meso-ancestry as the focal ego. Shared meso-ancestry is treated exclusive of shared micro-ancestry: It is coded as one only when partners have the same meso-ancestry but not the same micro-ancestry. The full coding scheme is provided in Appendix Table A-1. Table 2 reports the frequencies of unions between macro-ancestry groups together with corresponding odds ratios, calculated following the approach in Mood and Jonsson (2025).

To measure propinquity, we calculate straight-line distances between residences at time t using $100\,m \times 100\,m$ coordinates from the geographic registers. We transform distances using a natural logarithm, reflecting our expectations that relative rather than absolute differences in distance will be most relevant to partnership choices. Non-parametric approaches suggest that the log-transformation provides a reasonable and parsimonious fit (Appendix Figure D-2).

We also include several dyadic controls. Because ancestry is often based on parental place of birth, we can distinguish ancestry from nativity. We therefore include (1) a dummy for whether the potential partner is foreign-born and (2) a dummy for whether both partners are foreign-born, capturing higher rates of union formation among and between first-generation immigrants. We also control for partner similarity in terms of age and education at time t. Age differences are coded into four categories (man younger; man 0–2 years older; 3–6 years older; 7+ years older), with 3–6 years as reference. Educational homogamy is classified into three categories based on whether (1) both partners have some tertiary education (regardless of degree completion), (2) neither partner has any tertiary education, and (3) partners differ in their educational attainment (the reference).

Missing values account for less than 1% of observations for most variables. These cases are removed from the analysis. For variables with higher rates of missingness, such as education, we include a separate category to preserve cases and maintain comparability across models.

Table 1 presents descriptive statistics for both the sampled counterfactual couples and the actual realized couples. Compared to counterfactual couples, partners in realized couples were living closer together at t, were more likely to have the same ancestry and nativity, were closer in age, and were more likely to have similar levels of education. So the descriptive statistics provide first-order evidence for homophily in our study population. To obtain estimates of endogmay and propinquity net of each other, and net of controls, we turn to discrete choice models.

Method

We use a discrete choice, conditional logistic modeling framework to assess the relationship between propinquity, ancestry assortativity, and union formation at the micro-level. This approach is increasingly used in studies of mate choice (Gullickson 2021, 2022; Haandrikman and van Wissen 2012; Qian and Lichter 2018). A key strength of the framework is the ability to account for variations of realized and potential partners in multiple dimensions. The models represent the structure of opportunities because only actually existing singles are considered as possible partners, reflecting the empirically observed, joint distributions of attributes among available partners. Conditional on this opportunity structure, the models also allow for assortativity in multiple dimensions. This feature allows us to flexibly model assortativity by ancestry and residential proximity, while simultaneously controlling for other sources of assortativity, like age and education. The models can also be repurposed to assess the relationship between segregation and endogamy at the macro-level, which we explore later on in this section.

The conditional logit model predicts the probability of choosing a specific, observed partner over a set of counterfactual potential partners. Specifically, our model stipulates a set of decision-makers, indexed by i, who are conceptualized as opting into a union. Each decision-maker faces a choice set of alternative partners, C_{it}, and forms a union with potential partner, j, with probability:

$$P_{ijt} = \frac{e^{\beta_i X_{ijt}}}{\sum_{k \in C_{it}} e^{\beta_i X_{ikt}}}$$

where X_{ijt} is a vector representing attributes of potential partners and how they compare to the focal decision-maker. β_i is then a vector of coefficients indicating the relative importance of the corresponding attributes for union formation, which could, in principle, differ between decision-makers. A convenient property of the conditional logit model is that coefficients obtained from models estimated using samples of alternative partners are unbiased and asymptotically consistent (Ben-Akiva and Lerman 1985; Bruch and Mare 2012). For each observed union, we thus construct a reduced, individual-specific choice set, C_{it}, comprising the actual partner and 100 randomly selected potential partners from the same county and year. Within these sampled choice sets, X_{ijt} is made up of dyadic indicators of endogamy by micro- and meso-ancestry, spatial proximity (log-distance), whether the focal and alternative partner are both immigrants, age differences, and educational similarity. To allow for heterogeneity in sorting patterns, we estimate separate models by ego's macro-ancestry. All models are estimated in Stata 18 using the clogit function with the ego's identifier specified as the grouping variable, spanning all of the potential partners in each individual's choice set, (StataCorp 2023).

To assess the extent to which residential propinquity mediates ethnic endogamy, we estimate nested conditional logit models with and without a distance term. Because coefficients in nonlinear models are not directly comparable across nested specifications, we use the KHB method (Karlson et al. 2012; Kohler et al. 2011) to assess mediation. This allows us to distinguish changes in the endogamy parameters that result from (1) including propinquity as a mediator from (2) decreases in the variance of error terms that result from adding more explanatory variables to a model (i.e., changes in scaling factors, see also Mood 2010, Williams and Jorgensen 2023 and Williams 2009).

We also compare these micro-level mediation effects to segregation at the macro-level. If segregation between groups is captured by spatial proximity terms in our micro-models, then the spatial mediation effects we have found should be strongest for the most segregated groups and weakest for the least segregated groups. To test this expectation, we narrow our investigation to Stockholm County. We disaggregate the ancestry groups and produce endogamy and mediation estimates for all 50 micro-ancestries by fitting a separate model and applying the KHB method for each group. Additionally, we compute the 1990-2017 average residential dissimilarity index for each group in relation to all other groups, based on the same population from which potential partners are drawn, that is, all single individuals, aged 16 and above residing in Stockholm County. This quantifies levels of segregation experienced by each group in Stockholm. We calculate dissimilarity indices using DeSOs, which are administrative neighborhoods established by SCB, each containing between 700 and 2,500 residents (Sweden 2024). Higher values of the dissimilarity index indicate more segregation, with values in the 0.3-0.6 range considered moderate and values above 0.6 considered high. We plot our 50 mediation estimates against these segregation measures and asses the strength of the relationship using a bivariate linear regression.

In a final post-estimation analysis, we use the model results to test how propinquity shapes observed *rates* of endogamy. The KHB method identifies the mediated odds of endogamy, but not how endogamy changes at the population level when propinquity is absent—that is, if distances between potential partners were irrelevant to union formation. To do this, we use our conditional logit models to predict new unions and then calculate the share of endogamous unions (i.e., the endogamy rate) for each macro-ancestry group. We do so under four scenarios: (1) a base scenario with assortativity by age, education, and nativity, but with propinquity and endogamy coefficients set to zero; (2) a scenario that adds endogamy effects, but sets the propinquity coefficients to zero; (3) a scenario that adds the propinquity effects, but sets the endogamy coefficients to zero; and (4) a scenario with both distance and endogamy coefficients included as estimated. Comparing endogamy rates across these scenarios allows as to assess the degree to which propinquity vs. ancestry assortativity *per se* contributes to endogamy.

Results

Before turning to the statistical models, we describe the spatial availability of partners and the potential for mediation for select ancestry groups, drawing on a subset of our data from Stockholm County, 2010–2015. Figure 1 shows cumulative distance distributions for realized partners, same-ancestry singles, and Swedes (non-Swedes for Swedish women). Separate panels are shown by women's ancestry, focusing on a set of major sending groups that illustrate variation across macro groups. Across all groups, realized partners lived closer than other available singles: median distances were typically 6–8 km, compared to 15 km for Swedish partners. Two groups stand out: Somali women, whose unions formed at exceptionally short distances ($\sim 2.5\ km$), and Thai women, whose realized distances were somewhat longer ($\sim 10\ km$). The figure also suggests that propinquity can contribute to endogamy. For many groups (e.g., Somali, Syrian, Iranian), same-ancestry singles were more proximate than Swedes, while for others (e.g., Finns, Germans) same-ancestry singles and Swedes lived at similar distances. Where people sharing the same ancestry are especially close, we expect that propinquity will be most likely to contribute to endogamy. But while these descriptive results are suggestive of a relationship between propinquity and ethnic endogamy, they ignore other drivers of assortativity, such as age and education. Our conditional logit models address this.

Conditional Logit Models and Mediation Analysis

Our conditional logit models give us estimates of propinquity effects, ancestry assortativity, and the mediation of the latter by the former, estimated separately for different macro-ancestry groups and controlling for sorting by age, education, and nativity. In the remainder of this section, we focus on ancestry and propinquity estimates. Full regression results are presented in Appendix Table B-1.

We begin with the untransformed propinquity coefficients for each macro-ancestry group, displayed in Figure 2. Across all ethnic groups, there is a negative association between residential distance and the likelihood of forming a union. The coefficients for the natural logarithm of distance before cohabitation range from -0.8 to -0.6, meaning that women were less likely to partner with more distant singles. Propinquity matters for all ancestry groups, but is most important for those with Swedish and Western European ancestry and least important for those with African and Middle Eastern ancestry, with the other groups falling in between. Unified models with interactions between the ego's ancestry and propinquity allow statistical tests of differences between groups (Appendix table D-1). For women, there are significant differences in propinquity between Western versus non-Western groups, as well as Africa and the Middle East versus all other groups. For men, significant differences emerge between Western groups, Eastern European groups, and all others.

Substantively, our models suggest that a woman with Swedish or Western European background had a 1.73 times larger probability of forming a union with a more proximate partner than with an otherwise equivalent partner living twice as far away ($\exp(\beta \cdot (\ln(d) - \ln(2d))) = \exp(-0.79 \cdot \ln(2)) \approx 1.73$). African and Middle Eastern women were over 1.5 times more likely to form a union with a nearby partner compared to an otherwise equivalent partner living twice as far away. Women with Eastern European, Asian, and Latin American backgrounds had 1.6 times higher probabilities of partnering with a nearby single compared to one living twice as far away. While these difference between groups are non-trivial, it should be emphasized that all groups are substantially more likely to partner with those living nearby. This confirms Hypothesis 2.

Our models also provide evidence of assortativity by ancestry. Figure 3 Panel A presents the odds of endogamy for women, stratified by macro-ancestry, comparing models with and without controlling for propinquity. That is, the point estimates represent the odds (i.e., proportional differences in probability) that women partnered with someone sharing the same micro-ancestry relative to an otherwise equivalent partner sharing neither the woman's micro- nor meso-ancestry. Odds from the baseline model (without propinquity) are adjusted using the KHB method to account for scaling (Karlson et al. 2012).

Prior to controlling for propinquity, our models suggest that women of African and Middle Eastern ancestry were most likely to be endogamous. They had over a 70 times greater probability of partnering with someone from the same micro-ancestry group compared to an otherwise equivalent single from a different ancestry group. They are followed by women with Asian, Latin American, and then Eastern European ancestry who were, respectively, nearly 70, over 25, and over 20 times more likely to partner endogamously. The endogamy effects are weakest for those with Swedish and Western European backgrounds. They had 1.5 and nearly 3 times higher probabilities of partnering with someone with the same micro-ancestry, respectively. This provides support for Hypothesis 1A. We also find support for Hypothesis 1B: Those born in or born to immigrants from countries that are most culturally different from Sweden (i.e., African and Middle Eastern countries) had the highest endogamy rates, while those with ancestry in from contexts that are culturally similar to Sweden (i.e., Western Europe), had the lowest endogamy rates.

Once we control for propinquity, our endogamy estimates decline substantially for all non-Western groups, but increase slightly for those with Western European and Swedish ancestry. In other words, without accounting for spatial segregation, we appear to overestimate endogamy for non-Western groups, but *underestimate* endogamy for the Swedish and Western European groups. For example, in Figure 3 Panel A, controlling for propinquity reduces the endogamy odds for women with African and Middle Eastern backgrounds from over 70 to just above 50. But for women with Swedish ancestry, endogamy odds *increase*—from 1.5 to 1.8 times more likely to partner with another person with Swedish ancestry—after accounting for propinquity.

Panel B of Figure 3 quantifies and offers statistical tests for the extent of mediation. For all non-Western groups, endogamy is significantly overstated when not controlling for propinquity. In the African and Middle Eastern ancestry group, the endogamy effect is 41% larger in the model without a propinquity term. The figures are 31%, 30%, and 18% for the Asian, Latin American, and Eastern European groups, respectively. Conversely, for those of Western and Swedish ancestry, we observe significant effect suppression. The endogamy effects are 5% and 18% smaller, respectively, when we fail to account for propinquity. The results offer qualified support for Hypothesis 3A:

propinquity mediates homogamy for non-Western groups, but not for Western groups.

Post-estimation

Figure 4 shows the association between the residential segregation of micro-ancestry groups, measured by the dissimilarity index, and the mediation of endogamy by propinquity. Mediation is expressed as the ratio of endogamy odds from models without versus with a propinquity term. The relationship is positive and approximately linear: the more segregated a group is, the greater the overestimation of endogamy when propinquity is ignored. The mediation is strongest for African and Middle Eastern groups, particularly women of Somali and Syrian ancestry, though less so among those of Iraqi origin. By contrast, significant suppression appears for groups with low segregation, such as Finnish and Norwegian women, whose dissimilarity indices fall in the 0.1–0.2 range and who are culturally close to Swedes. Taken together, the figure supports Hypothesis 3B: endogamy is overstated more for groups with higher levels of segregation.

While propinquity increases the likelihood of union formation and mediates part of the ancestry endogamy effect, its implications for overall endogamy rates remain unclear. To clarify the relationship between micro-level behaviors and macro-level endogamy rates, we use our model estimates to produce counterfactual predictions. Table 3 compares predicted endogamy rates for women under four counterfactual scenarios representing different combinations of including or suppressing endogamy and propinquity effects. Substantively similar results for men are presented in Table C-2. In each scenario, we hold assortativity by age, education, and nativity fixed.

We find that, in terms of endogamy rates, propinquity matters but ancestry assortativity effects are stronger. For all non-Swedish groups, endogamy rates are very low in the base scenario that includes only the effects of assortativity by age, education, and nativity. Endogamy rates are notably higher when we include the effects of ancestry assortativity. In the highly endogamous Middle East and Africa group, endogamy rates are nearly 40 percentage points higher when including ancestry assortativity. In contrast, including propinquity increases the endogamy rate by only up to three percentage points. In absolute terms, the propinquity effect is stronger for groups from

Eastern Europe, the Middle East and Africa. For Swedes, the endogamy rate even decreases slightly (−0.4%) with propinquity, suggesting that neighborhood-level mixing between Swedish singles and those with immigrant backgrounds masks in-group tendencies. Relative to differences between the base and full model scenarios, propinquity accounts for only between 4 and 14% of endogamy rates. In relative terms, propinquity explains a larger fraction of endogamy in the Western European group (13.8%), followed by the Eastern European (7.3%) and Middle East and Africa group (6.6%). Overall, the results support Hypothesis 4: propinquity contributes modestly to aggregate endogamy, but its influence is small compared to ancestry assortativity *per se*.

Discussion

This study revisits a timeless sociological question: how space shapes the formation of intimate ties. In doing so, we recover a thread in the sociology literature linking propinquity, segregation, and assortative mating (Bossard 1932; Burgess and Wallin 1943; Catton and Smircich 1964; Clarke 1952; Davie and Reeves 1939; Kennedy 1943; Koller 1948; Marches and Turbeville 1953; Peach 1974; Ramsøy 1966; Schnepp and Roberts 1952). Nearly a century after the first systematic studies of residential propinquity, our analysis demonstrates that both social boundaries and geographic distances matter for union formation. Using Sweden as a case, we find that individuals in all ancestry groups are more likely to partner within their ancestry groups and with those who live nearby. Crucially, we show that residential propinquity accounts for a substantial share of the *odds* of endogamy, particularly among the most segregated groups. For aggregated non-Western groups, failing to consider propinquity leads us to overstate the odds of endogamy by 20–40 percent. For detailed groups, the overstatement of endogamy can be as high as 75 percent. In contrast, for Swedish- and Western-origin groups, ignoring propinquity masks some of the relatively modest tendency toward in-group partnering. In Blau's (1977a) terms, propinquity specifies a micro-structural mechanism through which macro-structural segregation shapes opportunities for interaction. Physical distance remains a key conduit between social structure and social relations.

The strength and ubiquity of the propinquity effect underline this point. Across all ancestry groups, individuals were far more likely to partner with someone living nearby than with an otherwise equivalent person living farther away. Before cohabitation, the median distance between actual partners was less than 6 kilometers, compared to more than 33 kilometers among randomly paired singles. This stark contrast shows that the geography is not a statistical footnote but a defining feature of how relationships form. The conditional logit models confirm that this pattern persists even when we control for education, age, nativity, and ancestry. In other words, propinquity exerts its own force, belying claims that technological development has led to a "death of distance" (Cairncross 1997). Moreover, the log-distance functional form performs well in capturing the relationship between proximity and union formation. This indicates that *relative* distance matters. The difference between a potential partner who is 1 km away and one who is 2 km away is comparable to the difference between potential partners who are 10 km versus 20 km away. The fact that the functional form holds so well strengthens the conclusion that fine-grained geographies—i.e., short-distance variations from block to block or district to district—structure opportunities for meeting and partnering, consistent with other findings on propinquity (Small and Adler 2019) as well as research on social ties in schools (Mouw and Entwisle 2006).

On the endogamy front, we observe a gradient that largely confirms prior research. Individuals of non-Western origin, particularly those with African, Middle Eastern, and Asian ancestry, are most likely to form in-group unions even when controlling for education, age, nativity, and propinquity. Those of Western ancestry, by comparison, show only modest in-group tendencies, though they too are more likely to partner within their own ancestry than outside of it. And while accounting for propinquity reduces our estimates of endogamy for non-Western groups, it does not affect the relative standing of groups with respect to endogamy. For women, the African and Middle Eastern groups have the highest endogamy odds, and the Western groups the lowest, both before and after controlling for propinquity; among men, the Asian group shows the highest endogamy before and after controlling for propinquity. These patterns echo earlier findings from Sweden (Dribe and Lundh 2011) and other Western contexts (Kalmijn and van Tubergen 2006), suggesting that cultural homophily and/or discriminatory constraints continue to sustain ethnic boundaries. While important for understanding the social landscape of union formation, this gradient is less surprising; what our results add is evidence that propinquity systematically mediates these differences.

Our mediation analysis reveals that propinquity explains a larger share of endogamy among the more segregated groups. This aligns with our Hypothesis 3B: assuming propinquity causally affects partnering—whether through increased chances of meeting or through attraction to those living nearby—residentially clustered ethnic groups will form disproportionately more endogamous unions due to the higher share of co-ethnics in residential proximity. In practical terms, this means that spatial segregation augments in-group partnering: even if individuals had no preferences for co-ethnic partners, their local environments would still yield more co-ethnic unions than preferences alone would predict. Conversely, for the least segregated groups, we observe a suppression effect: once we control for propinquity, the endogamy coefficients, while weak, become slightly stronger.

This suggests a misalignment between preferences for co-ethnic neighbors and co-ethnic partners that differs between the most segregated and the least segregated groups. The least segregated groups—i.e., those with Western immigrant backgrounds—may be more tolerant of out-groups as neighbors than as romantic partners. This would lead to a degree of residential integration that results in a local surfeit of desired co-ethnic partners. In essence, distance must be surmounted to find a suitable partner, and our estimates of endogamy increase when we control for proximate partners. The opposite may be the case for the most segregated groups. They may find themselves in neighborhoods where co-ethnic potential partners are locally over-represented even compared to strong endogamous preferences. In this case, propinquity will drive higher rates of endogamy beyond what is preferred, yielding the mediation effects we find.

However, our results also highlight the continuing importance of social boundaries for *rates* of endogamy at the macro-level, even as propinquity shapes the *odds* of endogamy at the micro-level. After controlling for propinquity, the odds of endogamy remain quite high for most groups. Because of this, propinquity accounts for only a fraction of endogamy when translated into predicted endogamy *rates* at the population level. The African and Middle Eastern, which is the most segregated and has the highest odds of endogamy, presents an instructive example: Propinquity contributes about three percentage points to overall endogamy rates, compared to roughly forty from ancestry assortativity. This accords with other, recent research in Sweden (Mood and Jonsson 2025). The contrast between large micro-level mediation and small macro-level change suggests that propinquity complements rather than drives endogamy. Spatial integration would not be a silver bullet for surmounting social boundaries around ancestry groups, at least in Sweden. But in other contexts with larger-scale patterns of segregation—as it the case with Black-White segregation in some US cities (Lee et al. 2008)—we might expect propinquity to have a stronger impact on endogamy rates. In other words, the aggregate effects of propinquity on endogamy will depend on underlying segregation patterns, which are small-scale in Sweden but larger-scale elsewhere.

Our results also revealed an unexpected gradient of the propinquity effect across ancestry groups. While all groups were more likely to form unions with nearby partners than more distant partners, individuals of Western ancestry were most likely to partner with someone living nearby. The propinquity effects were weaker for non-Western groups, and were weakest among those with African or Middle Eastern ancestry. These differences have several different interpretations.

One possibility is that groups vary in their willingness or capacity to bridge distance. Western-origin individuals may exhibit stronger local attachment or a preference for geographically convenient partners (Haandrikman and Hutter 2012; Lewicka 2011), whereas non-Western groups may extend their search radius to find partners who share language, religion, or cultural background. Differences may also reflect how social life is spatially organized. If Western groups' activity spaces and social foci are concentrated close to home (Feld 1981; Jones and Pebley 2014), local encounters will translate more directly into unions. Non-Western groups, by contrast, may have more scattered social foci and far-flung networks—religious communities, cultural associations, or transnational ties—that loosen the link between residence and meeting opportunities (Järv et al. 2015). Exploring such variation in activity spaces represents a promising direction for future research.

Alternatively, variation in propinquity effects may reflect unequal constraints on the partner market. For non-stigmatized groups, a strong distance effect could stem from a local abundance of willing partners. Stigmatized groups, facing higher rejection or limited choice nearby, may have to search further afield, forming unions over greater distances. In this sense, weaker propinquity among African and Middle Eastern groups likely signals barriers rather than preferences. Local abundance and social acceptance make nearby unions feasible for Western-origin groups, while for others, spatial and social constraints require extending the geography of partner search.

Although the Swedish population registers offer exceptional coverage and precision, and our modeling framework captures partner-market constraints at an unusually fine scale, several limitations remain. First, similar to studies of neighborhood effects, residential location and partner choice may be jointly determined (Choi and Tienda 2017b). Individuals with particularly strong in-group preferences—both for partners and neighbors—may self-select into neighborhoods where co-ethnics are more abundant, creating a correlation between co-ethnic proximity and preferences for co-ethnic partners that reflects selection rather than mediation. This would bias our propinquity and mediation estimates upward. But triangulating our findings with existing literature leads us to doubt that selection alone explains our results. Swedes have a stronger position in the housing market than immigrants and are most likely to translate preferences into residential attainments (Weber and Vogiazides 2023). But rather than finding a mediation effect for Swedes, we find a suppression effect. This is consistent with empirical research on intra-ethnic relations, which suggests that people tolerate the prospect of out-group neighbors more than out-group romantic partners (Weinfurt and Moghaddam 2001). If immigrants were to take the same view, with out-groups tolerated more as neighbors than as romantic partners, and were able to enact their preferences just as well as Swedes, this would imply a similar suppression effect. But that is not what we find. Of course it is possible that immigrants in Sweden prefer in-group neighbors more than they prefer in-group neighbors, but this seems unlikely. But setting aside this line of argument, there may yet be technical solutions to the selection problem. Estimating a joint, mixed effects residential choice and assortative mating model, using longitudinal data describing both prior residential choices and the partner choice, could help (e.g., Ben-Akiva et al. 2002; Brownstone et al. 2000; Pinjari et al. 2008). Developing this technical solution is outside the scope of the present paper, and we leave it as a topic for future research.

Second, our analysis focuses on realized unions but does not account for individuals who remain single when faced with limited access to co-ethnic or otherwise preferred partners. If some individuals choose non-partnership over exogamy in constrained local markets, we may underestimate the influence of opportunity structures and propinquity on both endogamy and overall marriage rates. In other words, the observed mediation by propinquity likely captures only the behavior of those who do partner, not the deterrent effects of segregation on union formation itself. Moreover, our data exclude cohabiting couples who never go on to have common children, which likely omits a segment of relationships that are shorter, less formalized, or more likely to be ethnically mixed. To the extent that such couples differ systematically in their residential proximity or partner choices, this omission could bias our results toward higher observed endogamy and stronger propinquity effects.

Third, our analysis defines the partner market as those who are single the year before cohabitation. But relationships often develop over longer periods. Partners may have met earlier or moved closer before moving in together, which could inflate the estimated propinquity effect. Future research should take a longer-term perspective to capture how proximity evolves before partnership formation, offering insights into when and with whom individuals are most likely to interact.

Finally, we exclude couples in which one partner immigrated the preceding year to focus on unions likely formed within Sweden. This removes marriage migrants, which by construction will inflate estimates of endogamy among Swedes and understate endogamy for immigrants already residing in Sweden. It also ignores the possibility that segregation, or even integration, may push some into the international marriage market. Future research could examine how transnational unions and marriage migration reflect preferences and constraints in the extended partner market (Elwert 2020; Niedomysl et al. 2010; Östh et al. 2009).

While these limitations should encourage cautious interpretation and serve as fodder for future research, we believe they do not diminish our findings and our underlying message: Partner markets are also partner landscapes. The distribution of people across space structures the encounters in which social and ethnic boundaries play out. In revisiting the role of residential proximity, this study extends a long line of sociological work that has treated propinquity as a central mechanism linking spatial structure to social interaction (Blau 1977b; Bossard 1932; Kennedy 1943; Morgan 1981). In this sense, our analysis constitutes a contemporary reassessment of classic questions about how space shapes intimacy and social integration, coming after dramatic social and technological transformations of dating and the marriage market over the last few decades.

More broadly, our results suggest how racial and ethnic boundaries themselves are complicated by the demographic processes of residential mobility and union formation. People who cross an ethnic boundary when moving into a neighborhood or forming a union violate the boundary in the short term, but also create an ongoing challenge to it. Inter-ethnic couples' can affect patterns of segregation across neighborhoods and ties between each partner's network alters over the long term (Ellis et al. 2012; Jarvis et al. 2023). This can induce inter-group contacts that have difficult-to-predict effects on racial and ethnic beliefs, attitudes, and preferences (Pettigrew 1998; Quillian 1996). Interethnic couples can also challenge racial and ethnic boundaries by bearing children. Multiethnic children have idiosyncratic relationships to existing racial and ethnic group categories (Harris and Sim 2002; Quillian and Redd 2009) and act as living examples of how boundaries that are often portrayed as primordial and durable are, in fact, contingent and fragile. Because of these endogeneities, small, short-term effects of propinquity could accumulate into large effects over the long term. While our data give us no special insights into subjective racial and ethnic identifications or long-term dynamics, we hope that our research contributes to a literature that recognizes how demographic processes can mediate relationships between the material boundaries of geographic space and the symbolic boundaries of social space.

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Tables and Figures

Table 1: Descriptive statistics for actual and counterfactual unions formed within Swedish counties, 1991-2017.

	Counterfactual Couples		
	Male Egos	Female Egos	Actual Couples
Residential distance (m)			
Median	32,921	33,536	5,728
IQR	51,237	51,673	14,717
Endogamy (%)			
Same ancestry	59.3	58.4	70.1
Pan-ethnic Endogamy (%)			
Same regional ancestry	13.2	13.1	13.1
Generation (%)			
Partner foreign-born	14.8	14.8	9.3
Both partners foreign-born	1.5	1.6	3.7
Age Difference (%)			
Man 0-2 years older	10.7	11.5	33.6
Man 3-7 years older	18.9	15.2	31.5
Man > 7 years older	22.6	29.4	12.8
Woman older	47.8	44.0	22.1
Educational Difference (%)			
Both tertiary education	9.1	8.7	19.2
Neither tertiary education	48.6	47.9	53.6
Different education	38.6	39.1	25.7
Either missing education	3.7	4.3	1.4

Note: "Same ancestry" and "Same regional ancestry" are mutually exclusive indicators. "Same ancestry" refers to partners sharing the same detailed (micro) ancestry, while "Same regional ancestry" refers to partners from the same broader meso-ancestry category but different micro origins (e.g., both Middle Eastern, but from different countries). The generation variables are not mutually exclusive: "Partner foreign-born" captures cases where the partner (alter) is first-generation, regardless of the ego's nativity, while "Both partners foreign-born" captures cases where both ego and alter are first-generation immigrants. *Source:* Authors' calculations using 1991–2017 Swedish population registers.

Table 2: Counts of new unions by men's and women's ancestry and corresponding endogamy odds ratios, Sweden 1990–2017

	Women's Ancestry					
Men's Ancestry	Sweden	Western Europe	Eastern Europe	Asia	Latin America	Middle East & Africa
Sweden	685,259	85,971	20,818	12,682	6,782	8,648
Western Europe	85,143	20,014	4,332	2,200	1,614	1,956
Eastern Europe	18,468	3,477	11,636	546	597	1,033
Asia	4,881	789	330	2,778	131	281
Latin America	6,705	1,580	705	252	2,820	447
Mideast & Africa	10,884	2,697	2,051	601	735	20,018
Total	811,340	114,528	39,872	19,059	12,679	32,383
Endogamy Odds Ratio	3.4	1.8	16.5	26.7	29.7	93.6

Note: The top part of the table shows the observed number of unions between men (rows) and women (columns) across six macro-ancestry groups, with marginal totals in the bottom row. Endogamy odds ratios are calculated following Mood and Jonsson 2025, p.11, where the numerator is the odds of endogamy vs. exogamy for the group identified in the column, and the denominator is the odds of exogamy with the group identified in the column for all other groups. Values above one indicate the degree to which endogamy exceeds what would be expected under random mixing. *Source:* Authors' calculations using Swedish population registers, 1990–2022.

Table 3: Counterfactual predictions of endogamy rates for Swedish women with different ancestries

	Sweden	Western Europe	Eastern Europe	Asia	Latin America	M.E. & Africa
Endogamy Rates (%)						
Baseline	79.2	4.7	1.9	0.5	1.2	3.4
Baseline + Ancestry	84.6	8.6	17.9	12.2	14.0	43.2
Baseline + Propinquity	78.8	5.3	3.1	1.0	1.8	6.2
Full Model	84.1	9.1	19.3	13.0	15.2	46.2
Δ (Full – Base)	4.9	4.4	17.4	12.5	14.0	42.8
Ancestry Contribution						
Rate (p.p.)	5.4	3.9	16.1	11.7	12.8	39.9
% of Δ	110.3	89.3	92.4	93.6	91.5	93.1
Propinquity Contribution						
Rate (p.p.)	-0.4	0.6	1.3	0.5	0.6	2.8
% of Δ	-7.8	13.8	7.3	4.2	4.5	6.6

Note: Predicted shares of ancestry-endogamous unions are computed from conditional logit models estimated separately for women of each macro-ancestry group. The *Baseline* includes assortativity by age, education, and nativity, but with propinquity and ancestry assortativity coefficients set to 0. *Baseline + Ancestry* includes ancestry assortativity, but sets the coefficients for propinquity to 0. *Baseline + Propinquity* includes propinquity effects, but sets the ancestry assortativity coefficients to 0. Predictions for the *Full Model* are based on leaving all coefficients as estimated. *Source:* Authors' calculations using Swedish population registers, 1990–2022.

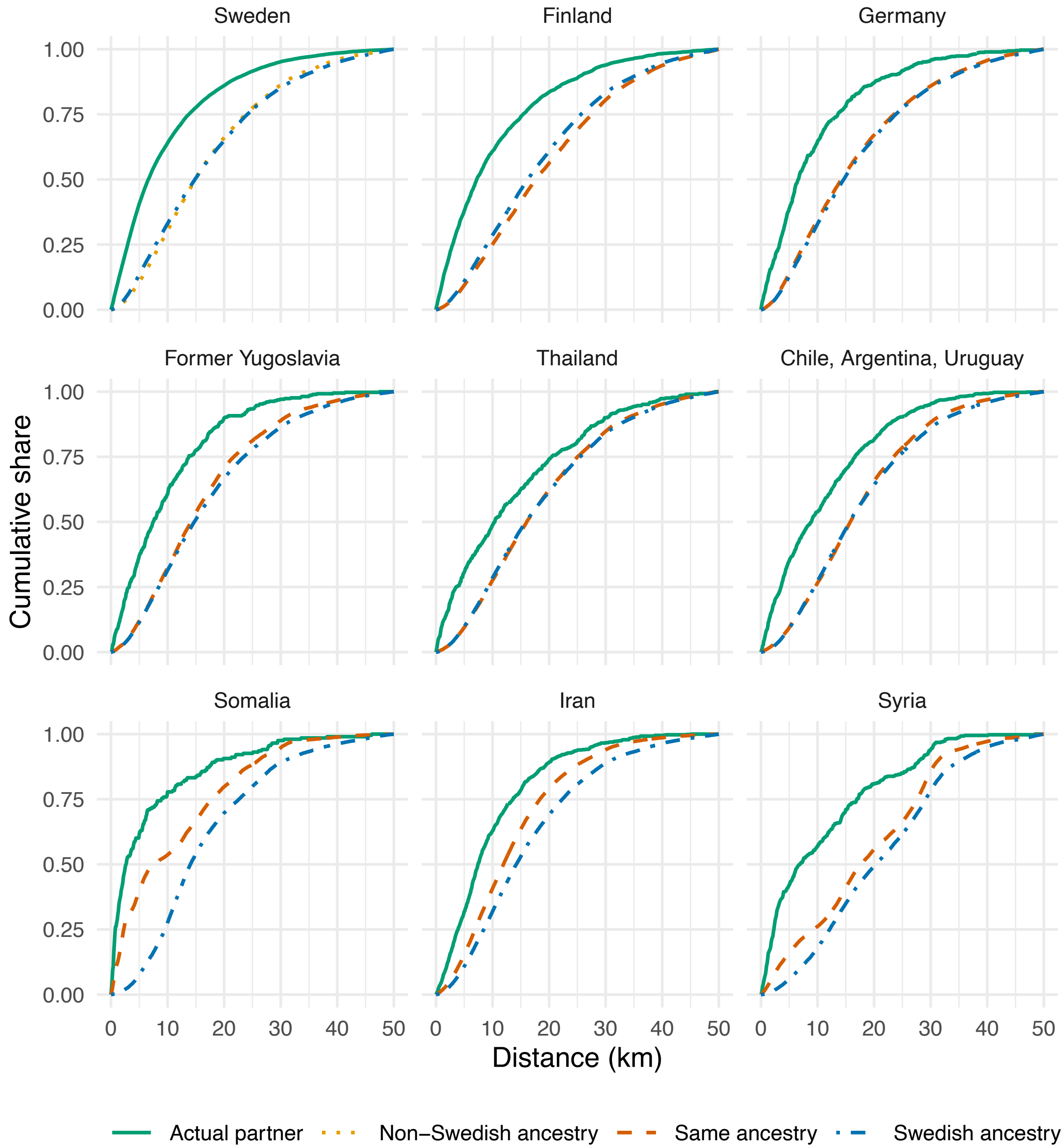

Figure 1: Cumulative shares of partners by residential distance for women in Stockholm County, 2010–2015. Each curve shows the proportion of actual partners, potential partners of the same ancestry, and potential partners of Swedish ancestry (non-Swedish ancestry for women of Swedish origin) encountered within a given distance. Plots are stratified by women's ancestry, with a selected set of larger origin groups included to illustrate contrasts across regions that vary in their degree of residential segregation. Source: Authors' own calculations based on Swedish register data.

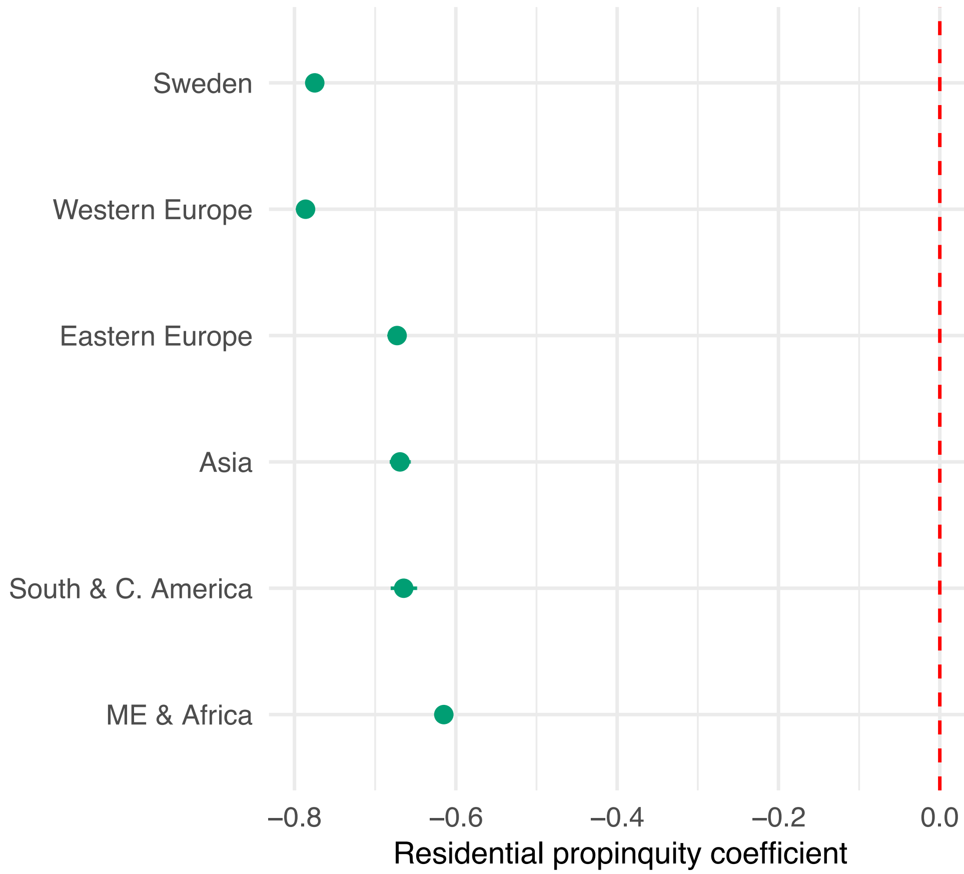

Figure 2: Propinquity coefficients from conditional logit models of women's partner choices in Sweden, by macro-ancestry group. Sample is restricted to couples living in the same county at the beginning (t) and end ($t + 1$) of each year. Models include controls for assortativity by ethnicity, age, nativity and education. Complete coefficients for women and men are listed in Appendix Tables B-1 and C-1. Source: Authors calculations using 1990-2022 Swedish register data.

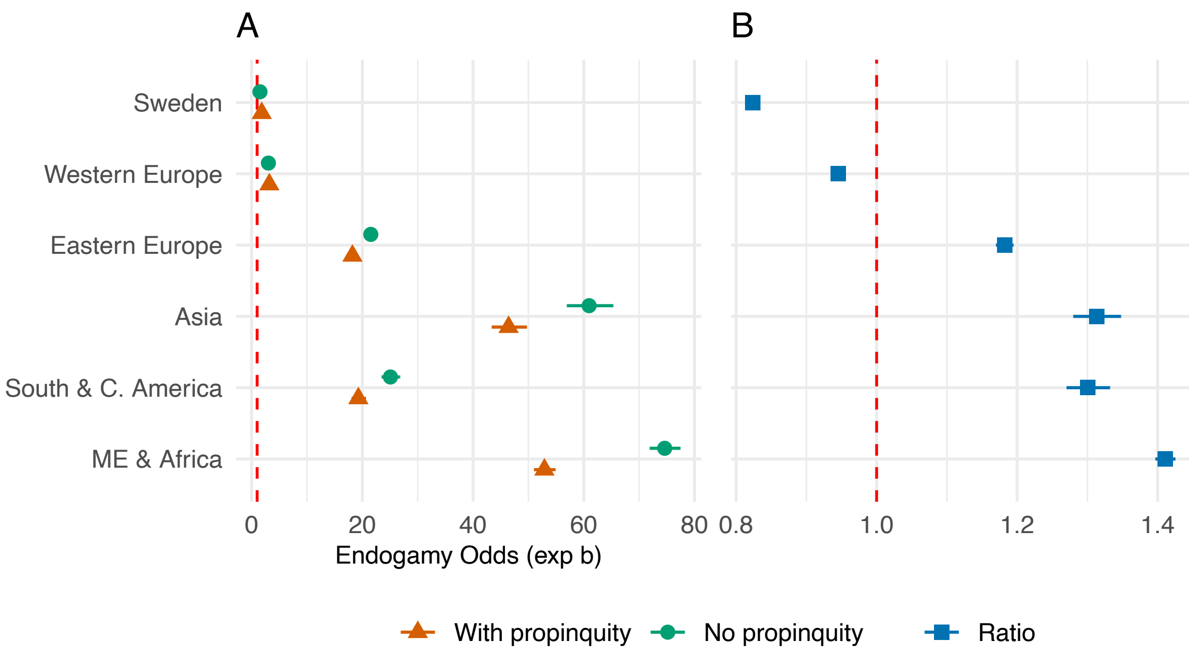

Figure 3: Endogamy and mediation estimates from conditional logit models of union formation. **(A)** Odds of endogamy relative to exogamy for women in Sweden, 1990-2022, by ancestry group. First model controls for propinquity and the other does not. Both models control for age differences, educational differences, and nativity differences. The coefficients in the model without propinquity term is adjusted using the KHB-method. **(B)** Ratio of coefficients from models in Panel A. Source: Authors' calculations using 1990-2022 Swedish register data.

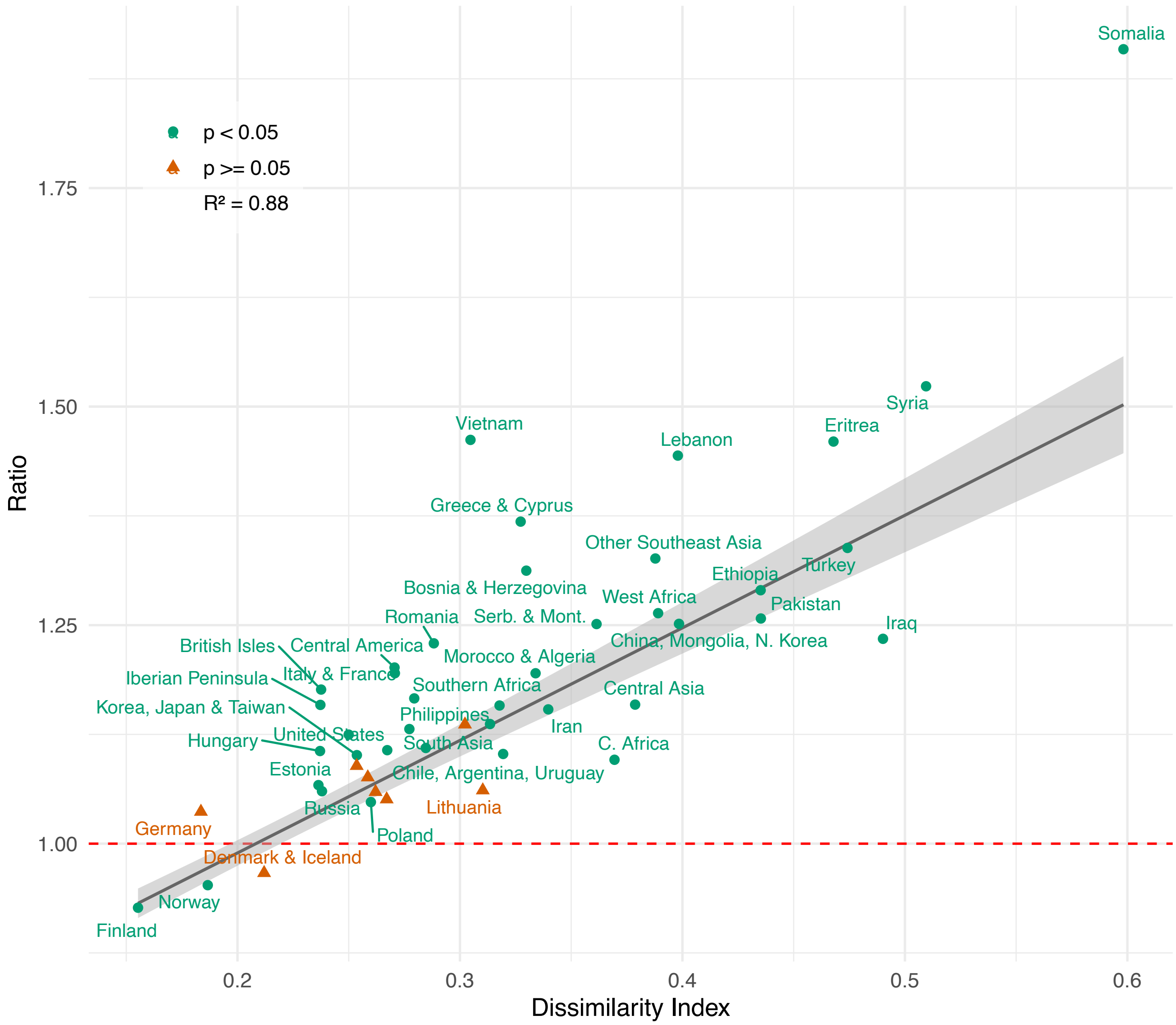

Figure 4: Association between mediation of endogamy by propinquity and residential segregation across ancestry groups for women in Stockholm County. The y-axis shows the ratio of endogamy coefficients from partner choice models without versus with a propinquity term—indicating how many times larger the estimated endogamy coefficient is when spatial proximity is not adjusted for. The coefficients from the model without the propinquity term are adjusted using the KHB method to ensure comparability between models. The x-axis shows each group's mean dissimilarity index between 1991 and 2017, measuring residential segregation from all other groups. The fitted line depicts an inverse-variance weighted regression.

However Far Away? The Spatial Contingencies of Assortative Mating

Appendices

Appendix A: Ancestry coding

Table A-1: Nested macro-, meso-, and micro-ancestry coding scheme

Macro-ancestry	Meso-ancestry	Micro-ancestry
African + Middle East + Central Asia	Afghanistan, Iran & Central Asia	Afghanistan
		Central Asian (Armenia, Uzbekistan etc.)
		Iran
	Horn of Africa	Eritrea
		Ethiopia
		Somalia
	North Africa + Middle East	Iraq
		Lebanon
		Morocco and Algeria
		Other Middle East
		Syria
		Tunisia, Libya, Egypt
		Turkey
	Sub-Saharan Africa	Central African Nations
		South African Nations
		West Africa
Asia and Southeast Asia	East Asia	China, Mongolia, North Korea
		Korea, Japan, Taiwan, Hong Kong, Singapore
	South and Southeast Asia	Pakistan
		Phillipines and Pacific Islands
		South Asia (India etc.)
		South East Asia
		Thailand
		Vietnam
Eastern Europe and Balkans	Balkans	Bosnia-Herzogovina
		Kosovo, Croatia, Albania etc.
		Serbia and Montenegro
		Yugoslavia
	Eastern Europe	Bulgaria, Ukraine, etc.
		Estonia
		Hungary
		Lithuania
		Poland
		Romania
		Russia and the Soviet Union
N+S+W Europe	Britain and Former Colonies	British Isles and Colonies
		United States and Canada

Continued on next page

Table A-1: (continued)

Macro-ancestry	Meso-ancestry	Micro-ancestry
	Germany, Belgium, Netherlands	Belgium and the Netherlands
		Germany etc.
	Other Nordic	Denmark and Iceland
		Finland
		Norway
	Spain, France, Italy, Greece	Greece and Cyprus
		Iberian Peninsula
		Italy and France
South and Central America	South and Central America	Central American and Carribean
		Chile, Argentina, Uruguay
		Colombia
		Other South America
Sweden	Sweden	Sweden

Note: The micro-ancestry groups are coded by Statistics Sweden to maintain anonymity for individuals from small groups. We are unable to break out countries of origin below the micro-level of aggregation.

Appendix B: Full regression table for women

Table B-1: Coefficients from conditional logit models of women's partner choices in Sweden, by macro-ancestry group.

Covariate	Sweden	Western Europe	Eastern Europe	Asia	Latin America	ME & Africa
Age difference						
Man 0-2 yrs older	0.36	0.34	0.27	0.30	0.39	0.17
	(0.35, 0.36)	(0.32, 0.35)	(0.24, 0.30)	(0.26, 0.34)	(0.34, 0.44)	(0.14, 0.21)
Man 3-6 yrs older (ref)	–	–	–	–	–	–
Man 7+ yrs older	-1.66	-1.55	-1.31	-1.03	-1.33	-1.09
	(-1.67, -1.65)	(-1.57, -1.53)	(-1.34, -1.28)	(-1.07, -0.98)	(-1.39, -1.27)	(-1.12, -1.05)
Woman older	-1.55	-1.43	-1.60	-1.58	-1.31	-1.74
	(-1.55, -1.54)	(-1.44, -1.41)	(-1.63, -1.57)	(-1.62, -1.53)	(-1.37, -1.26)	(-1.78, -1.70)
Education						
Both have tertiary	0.85	0.88	0.87	0.90	0.78	0.96
	(0.84, 0.85)	(0.85, 0.90)	(0.83, 0.90)	(0.85, 0.95)	(0.72, 0.85)	(0.91, 1.00)
Neither have tertiary	0.74	0.71	0.52	0.47	0.56	0.38
	(0.73, 0.74)	(0.69, 0.73)	(0.49, 0.56)	(0.42, 0.52)	(0.50, 0.62)	(0.34, 0.42)
Different levels (ref)	–	–	–	–	–	–
Education missing	-1.22	-0.63	-0.58	-0.66	-0.84	-0.61
	(-1.25, -1.18)	(-0.69, -0.56)	(-0.66, -0.51)	(-0.78, -0.54)	(-1.00, -0.69)	(-0.69, -0.53)
Endogamy						
Same micro ancestry	0.61	1.17	2.90	3.84	2.96	3.97
	(0.59, 0.62)	(1.13, 1.20)	(2.87, 2.94)	(3.77, 3.91)	(2.89, 3.02)	(3.93, 4.00)
Same meso ancestry	0.52	0.59	1.79	1.18	1.68	1.89
	(0.50, 0.53)	(0.56, 0.62)	(1.75, 1.84)	(1.03, 1.34)	(1.56, 1.80)	(1.84, 1.94)
Nativity						
Alter is 1st gen	-0.77	-0.58	-0.75	-0.94	-0.29	-0.47
	(-0.78, -0.75)	(-0.60, -0.55)	(-0.80, -0.70)	(-1.07, -0.81)	(-0.39, -0.20)	(-0.53, -0.41)
Both partners 1st gen		0.92	1.50	0.56	0.59	1.28
		(0.87, 0.96)	(1.45, 1.56)	(0.42, 0.69)	(0.48, 0.69)	(1.22, 1.35)
Residential Propinquity						
Log distance (m)	-0.78	-0.79	-0.68	-0.67	-0.67	-0.62
	(-0.78, -0.77)	(-0.79, -0.78)	(-0.68, -0.67)	(-0.68, -0.66)	(-0.68, -0.65)	(-0.63, -0.61)
N person-year-alternatives	**81,981,700**	**11,566,924**	**4,026,971**	**1,923,747**	**1,279,973**	**3,271,693**

Note: This table corresponds to Figure 2 in the main text. Parentheses show asymptotic 95% confidence intervals. Sample is restricted to couples living in the same county at the beginning (t) and end ($t + 1$) of each year. Source: Authors calculations using 1990-2022 Swedish register data.

Appendix C: Results for Men

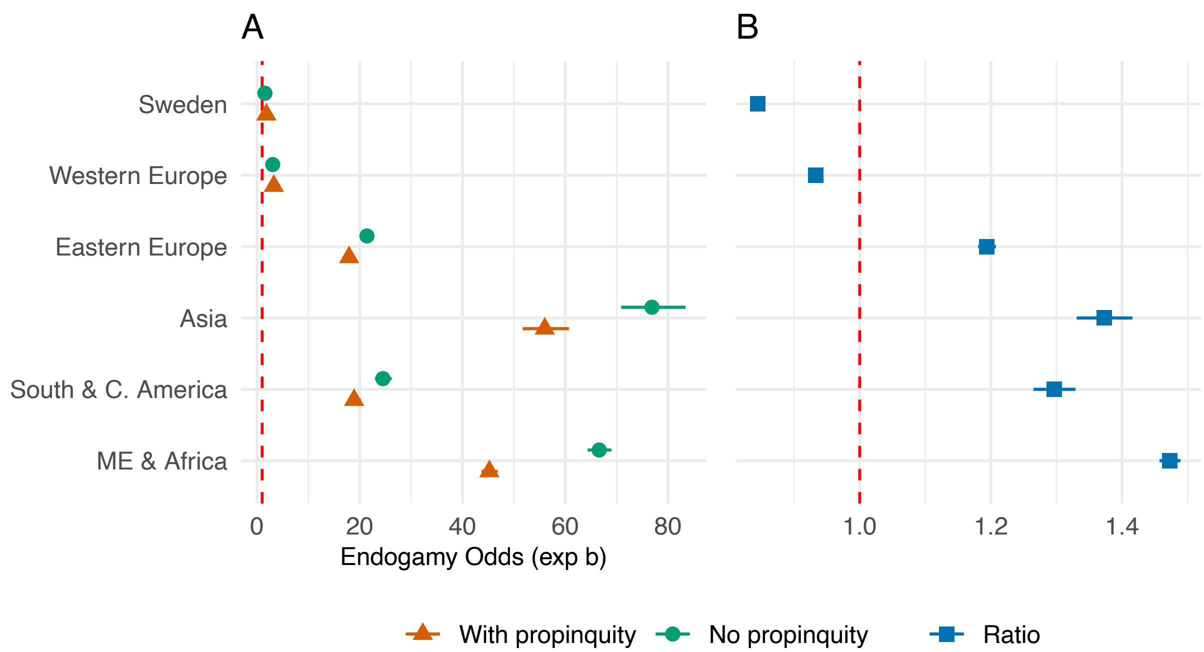

Figure C-1: Endogamy and mediation estimates from conditional logit models of union formation. **(A)** Odds of endogamy relative to exogamy for men in Sweden, 1990-2017, by ancestry group. First model controls for propinquity and the other does not. Both models control for age differences, educational differences, and nativity differences. The coefficients in the model without propinquity term is adjusted using the KHB-method. **(B)** Ratio of coefficients from models in Panel A. Source: Authors' calculations using 1990-2022 Swedish register data.

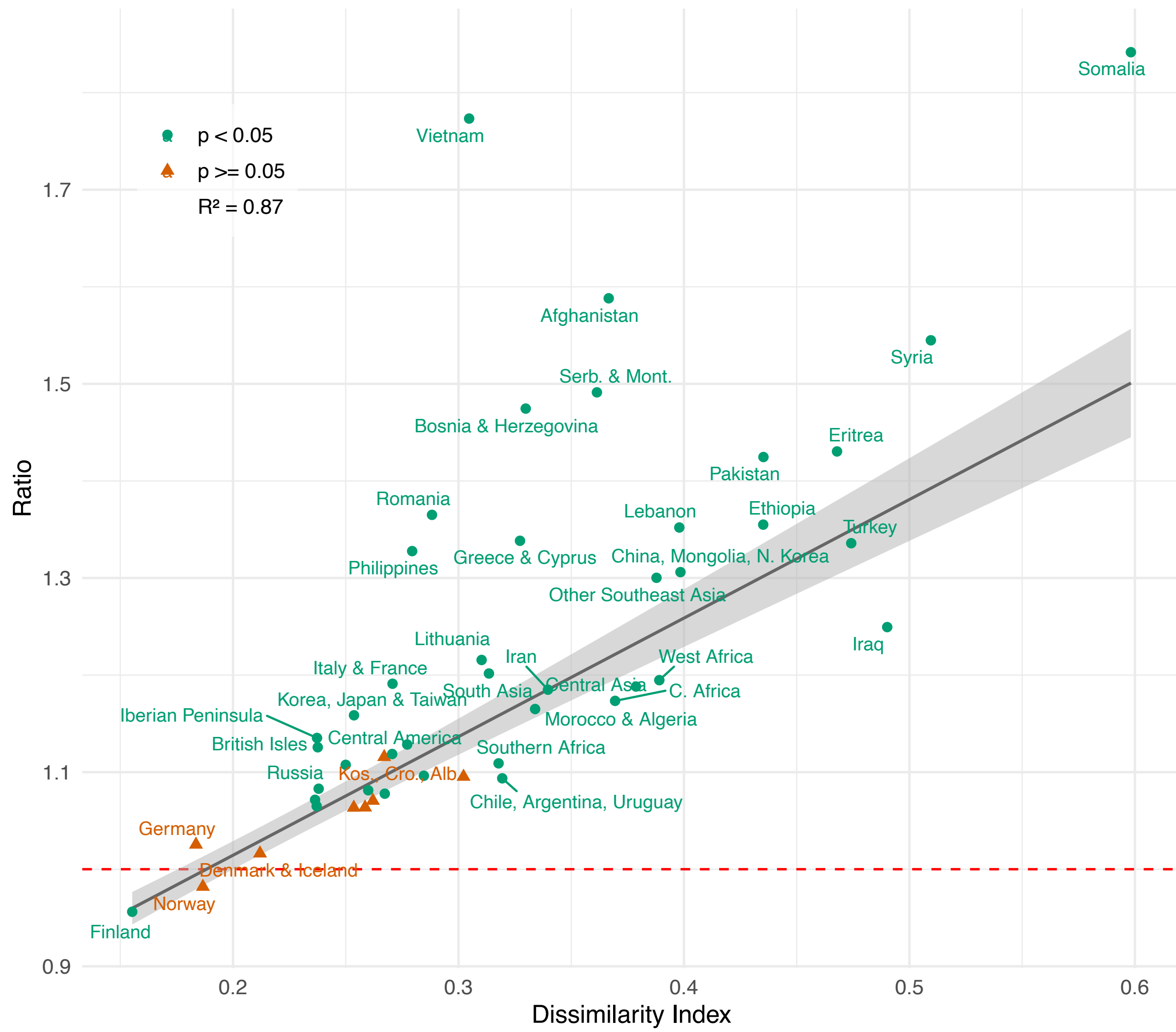

Figure C-2: Association between mediation of endogamy by propinquity and residential segregation across ancestry groups for men in Stockholm County. The y-axis shows the ratio of endogamy odds from models without the propinquity term and models with the propinquity term, while the x-axis shows the mean of each group's dissimilarity index between 1990-2017. The fitted line depicts an inverse-variance weighted regression to account for differences in group sizes.

Table C-1: Coefficients from conditional logit models of men's partner choices in Sweden, by macro-ancestry group.

Covariate	Sweden	Western Europe	Eastern Europe	Asia	Latin America	ME & Africa
Age difference						
Man 0-2 yrs older	0.64	0.50	0.58	0.59	0.52	0.49
	(0.63, 0.65)	(0.49, 0.52)	(0.55, 0.61)	(0.53, 0.65)	(0.47, 0.57)	(0.46, 0.52)
Man 3-6 yrs older (ref)	–	–	–	–	–	–
Man 7+ yrs older	-1.26	-1.18	-1.13	-1.10	-0.92	-0.81
	(-1.27, -1.26)	(-1.21, -1.16)	(-1.17, -1.09)	(-1.18, -1.02)	(-0.98, -0.86)	(-0.85, -0.78)
Woman older	-1.36	-1.43	-1.40	-1.20	-1.37	-1.43
	(-1.37, -1.36)	(-1.44, -1.41)	(-1.43, -1.37)	(-1.26, -1.14)	(-1.43, -1.32)	(-1.47, -1.40)
Education						
Both have tertiary	1.18	1.09	1.08	1.01	0.93	0.84
	(1.17, 1.19)	(1.07, 1.12)	(1.04, 1.12)	(0.92, 1.09)	(0.85, 1.00)	(0.79, 0.88)
Neither have tertiary	0.35	0.50	0.45	0.52	0.46	0.44
	(0.35, 0.36)	(0.48, 0.52)	(0.41, 0.48)	(0.45, 0.59)	(0.40, 0.51)	(0.40, 0.47)
Different levels (ref)	–	–	–	–	–	–
Education missing	-1.47	-0.91	-0.41	-0.31	-0.84	-0.26
	(-1.51, -1.43)	(-0.99, -0.84)	(-0.49, -0.33)	(-0.46, -0.16)	(-1.00, -0.67)	(-0.32, -0.19)
Endogamy						
Same micro ancestry	0.60	1.19	2.89	4.03	2.94	3.81
	(0.59, 0.61)	(1.16, 1.22)	(2.85, 2.92)	(3.95, 4.11)	(2.87, 3.00)	(3.78, 3.85)
Same meso ancestry	0.57	0.53	1.78	1.53	1.82	1.75
	(0.56, 0.58)	(0.51, 0.56)	(1.73, 1.83)	(1.37, 1.69)	(1.70, 1.93)	(1.71, 1.80)
Nativity						
Alter is 1st gen	-0.50	-0.55	-0.99	-1.08	-0.69	-0.74
	(-0.52, -0.49)	(-0.58, -0.52)	(-1.05, -0.94)	(-1.23, -0.93)	(-0.82, -0.57)	(-0.82, -0.67)
Both partners 1st gen		0.98	1.87	1.43	0.95	1.61
		(0.93, 1.02)	(1.80, 1.93)	(1.28, 1.59)	(0.82, 1.08)	(1.54, 1.69)
Residential Propinquity						
Log distance (m)	-0.80	-0.80	-0.71	-0.65	-0.68	-0.64
	(-0.81, -0.80)	(-0.81, -0.79)	(-0.72, -0.70)	(-0.67, -0.63)	(-0.70, -0.67)	(-0.64, -0.63)
N person-year-alternatives	**82,836,968**	**11,641,462**	**3,612,063**	**928,190**	**1,263,409**	**3,736,192**

Note: Parentheses show asymptotic 95% confidence intervals. Sample is restricted to couples living in the same county at the beginning (t) and end ($t + 1$) of each year. Source: Authors calculations using 1990-2022 Swedish register data.

Table C-2: Counterfactual predictions of endogamy rates for Swedish men with different ancestries

	Sweden	Western Europe	Eastern Europe	Asia	Latin America	M.E. & Africa
Endogamy Rates (%)						
Baseline	78.1	4.4	2.2	1.2	1.2	2.9
Baseline + Ancestry	83.5	8.5	19.8	24.7	14.0	38.0
Baseline + Propinquity	78.0	5.0	3.8	2.5	1.9	5.5
Full Model	83.2	9.1	21.3	25.7	15.2	40.9
Δ (Full – Base)	5.1	4.7	19.2	24.4	14.0	38.0
Ancestry Contribution						
Rate (p.p.)	5.4	4.2	17.6	23.4	12.8	35.1
% of Δ	106.0	88.6	92.0	95.9	91.0	92.5
Propinquity Contribution						
Rate (p.p.)	-0.1	0.7	1.6	1.3	0.7	2.7
% of Δ	-1.8	14.2	8.3	5.3	4.8	7.0

Note: Predicted shares of ancestry-endogamous unions are computed from conditional logit models estimated separately for women of each macro-ancestry group. The *Baseline* includes assortativity by age, education, and nativity, but with propinquity and ancestry assortativity coefficients set to 0. *Baseline + Ancestry* includes ancestry assortativity, but sets the coefficients for propinquity to 0. *Baseline + Propinquity* includes propinquity effects, but sets the ancestry assortativity coefficients to 0. Predictions for the *Full Model* are based on leaving all coefficients as estimated. *Source:* Authors' calculations using Swedish population registers, 1990–2022.

Appendix D: Alternate model specifications

We tested a number of alternate model specifications related to between group endogamy comparisons, nativity, and the (non-linear) effect of distance.

First, while the KHB method can address scaling factor issues for nested models estimated on the same sample, it cannot resolve issues of different scaling factors across samples, making it potentially problematic to compare, for example, endogamy coefficients across models estimated on different samples. To address this, we estimated pooled models that constrain the age and education assortativity coefficients to be the same across groups. The results are shown in Table D-1. With these coefficients constrained to be the same for all groups, it is then possible to compare endogamy and propinquity coefficients across groups using interactions. These results correspond to what we present in the main text.

We also could be concerned that different ancestry groups have different shares of first- vs. second-generation immigrants, and our endogamy results could be partly driven by this heterogeneity within groups. To address this, we considered interactions between nativity and endogamy. These results are presented in Table D-1. We find that second-generation immigrants are less endogamous across all ancestry groups, and that the patterns we present in the main text are largely driven by first generation immigrants.

Finally, we examined whether and to what degree the log-distance specification provides a reasonable fit for the data by breaking distances into a series of dummy variables. Figure D-2 shows results taken from the perspective of Swedish women, but the pattern is similar across sexes and ancestry groups (full results are available upon request). The non-parametric specification, based on doubling the bounds of the distance intervals, follows an almost linear decline in coefficients up until about 12 km, indicating that the log-distance specification captures the propinquity effect well.

Two deviations from the general pattern are worth noting. First, the category representing 0–100 meters departs from the otherwise smooth trend. Because residential coordinates in the registers are measured on a 100 × 100 m grid, any two individuals within 100 m are effectively registered in the same square but on different properties. This category may therefore include atypical residential setups——such as adjacent buildings, multi-household properties, or institutional housing——that are differently spaced compared to properties in different squares on the grid. In essence, the distribution of inter-property distances within squares may differ qualitatively from the inter-property distances between squares. Second, the decline in coefficients becomes steeper beyond roughly 12 km. This break point is not far from the typical diameter of a Swedish municipality, implying that the propinquity effect operates differently within and between municipalities within counties.

Table D-1: Coefficients from pooled conditional logit models of partner choices in Sweden, with macro-ancestry interactions

Covariate	Women	Men
Age difference		
Man 0-2 years older	0.35	0.62
	(0.34, 0.35)	(0.61, 0.62)
Man 3-6 years older (ref)	–	–
Man 7+ years older	-1.59	-1.23
	(-1.60, -1.58)	(-1.23, -1.22)
Woman older	-1.54	-1.37
	(-1.54, -1.53)	(-1.38, -1.37)
Education		
Both have tertiary	0.85	1.15
	(0.85, 0.86)	(1.14, 1.16)
Neither have tertiary	0.71	0.38
	(0.70, 0.71)	(0.37, 0.38)
Different educations (ref)	–	–
Either missing	-0.97	-1.11
	(-1.00, -0.95)	(-1.14, -1.08)
Endogamy		
Different micro ancestry (ref)	–	–
Same micro-ancestry	1.14	1.19
	(1.11, 1.17)	(1.16, 1.22)
Different meso ancestry (ref)	–	–
Same meso-ancestry	0.56	0.56
	(0.53, 0.58)	(0.54, 0.59)
Same micro-ancestry × Ego's ancestry		
Sweden	-0.51	-0.60
	(-0.54, -0.48)	(-0.63, -0.57)
Western Europe	–	–
Eastern Europe	1.86	1.74
	(1.82, 1.91)	(1.69, 1.78)
Asia	2.20	2.69
	(2.13, 2.27)	(2.61, 2.77)
Latin America	1.80	1.60
	(1.73, 1.88)	(1.53, 1.67)
Middle East & Africa	3.05	2.82
	(3.00, 3.10)	(2.77, 2.86)
Same meso-ancestry × Ego's ancestry		
Sweden	-0.02	-0.00
	(-0.05, 0.00)	(-0.03, 0.02)
Western Europe	–	–
Eastern Europe	1.32	1.22
	(1.27, 1.37)	(1.17, 1.28)

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Table D-1: (continued)

Covariate	Women	Men
Asia	0.16	0.81
	(0.00, 0.32)	(0.65, 0.97)
Latin America	1.10	1.10
	(0.98, 1.22)	(0.98, 1.22)
Middle East & Africa	1.50	1.32
	(1.45, 1.56)	(1.27, 1.37)
Nativity		
Alter is 1st gen	-0.72	-0.55
	(-0.73, -0.71)	(-0.56, -0.54)
Both partners 1st gen	1.19	1.20
	(1.17, 1.21)	(1.18, 1.22)
Residential Propinquity		
Log distance (m)	-0.79	-0.80
	(-0.80, -0.78)	(-0.81, -0.79)
Log distance × Ego's ancestry		
Sweden	0.01	-0.00
	(0.01, 0.02)	(-0.01, 0.00)
Western Europe	–	–
Eastern Europe	0.11	0.09
	(0.10, 0.12)	(0.08, 0.10)
Asia	0.12	0.15
	(0.10, 0.13)	(0.13, 0.17)
Latin America	0.12	0.12
	(0.10, 0.14)	(0.10, 0.13)
ME & Africa	0.16	0.16
	(0.15, 0.17)	(0.15, 0.17)
N person-year-alternatives	**104,051,008**	**104,018,284**

Note: Parentheses show asymptotic 95% confidence intervals. Models are estimated for all ancestry groups simultaneously, with ancestry by endogamy and ancestry by propinquity interactions. Sample is restricted to couples living in the same county at the beginning (t) and end ($t + 1$) of each union formation interval. Source: Authors calculations using 1990-2022 Swedish register data.

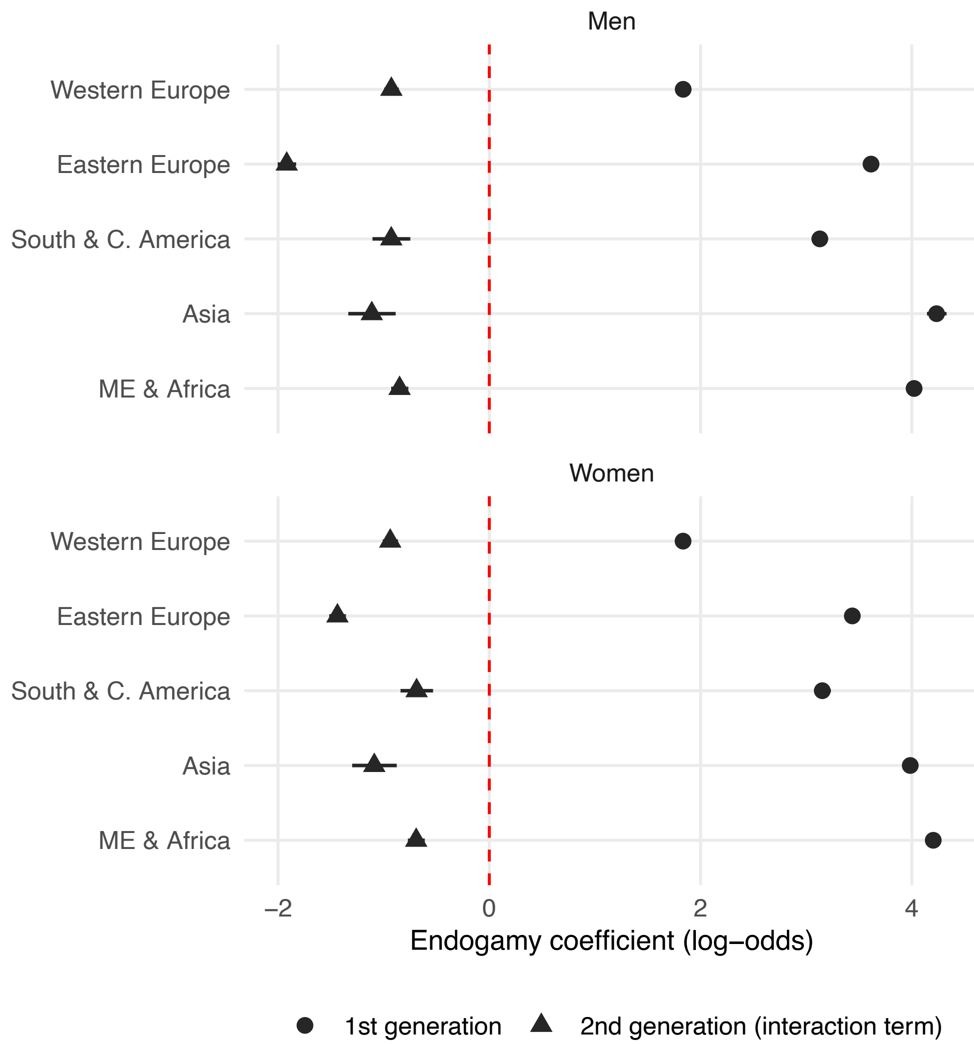

Figure D-1: Endogamy coefficients by nativity from conditional logit models of partner choices in Sweden. For each ancestry group, the figure shows the estimated endogamy coefficient for first-generation individuals and the corresponding interaction term for second-generation; the predicted coefficient for the second generation equals the sum of these two terms. The sample is restricted to couples living in the same county at the beginning (t) and end (t+1) of each year. Models include controls for assortativity by age, education, nativity, and residential propinquity. Source: Authors' calculations using Swedish register data, 1991–2022.

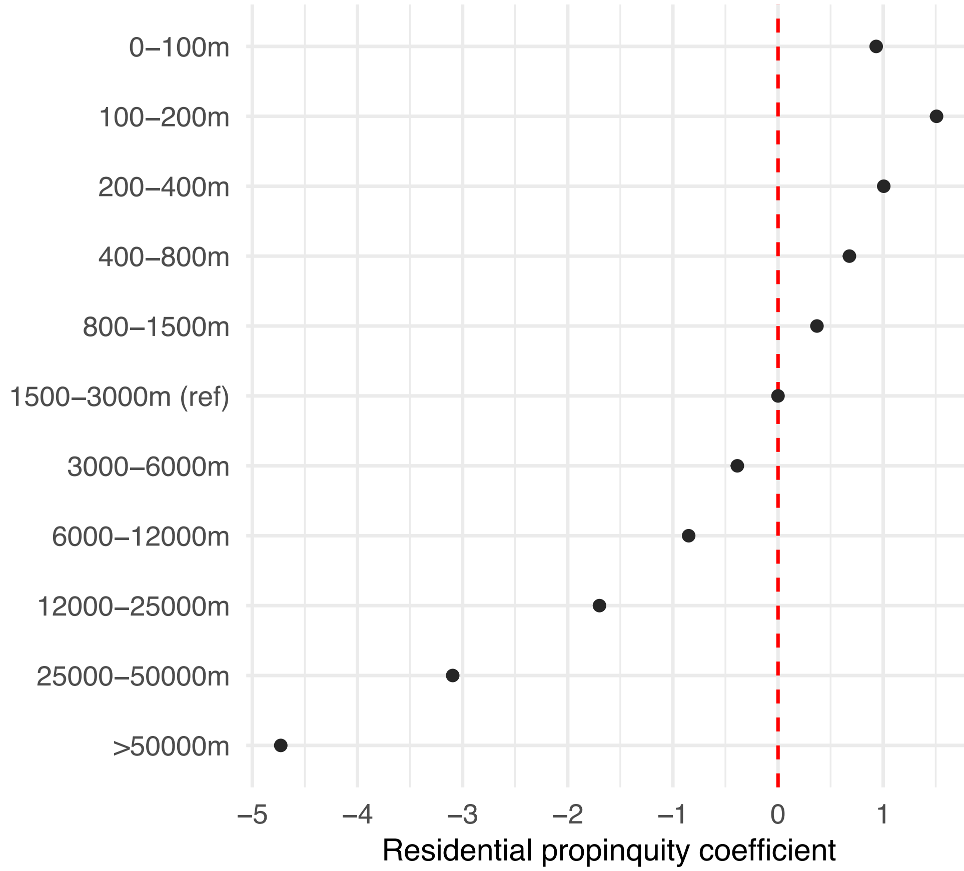

Figure D-2: Categorical propinquity coefficients from conditional logit model of Swedish women's partner choices in Sweden. Sample is restricted to couples living in the same county at the beginning (t) and end ($t + 1$) of each year. Model include controls for assortativity by ethnicity, age, nativity and education. Source: Authors calculations using 1990-2022 Swedish register data.