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The Patriarchy Index: a new measure of gender and generational inequalities in the past

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Abstract

In this article, we present a new measure for use in cross-cultural studies of family-driven age- and gender-related inequalities. This composite measure, which we call the Patriarchy Index, combines a range of variables related to familial behaviour that reflect varying degrees of sex- and age-related social inequality across different family settings. We demonstrate the comparative advantages of the index by showing how 266 historical populations living in regions stretching from the Atlantic coast of Europe to Moscow scored on the patriarchy scale. We then compare the index with contemporary measures of gender discrimination, and find a strong correlation between historical and current inequality patterns. Finally, we explore how variation in patriarchy levels across Europe is related to the socio-economic and institutional characteristics of the regional populations, and to variation across these regions in their degree of demographic centrality and in their environmental conditions. Overall, the results of our study confirm previous findings that family organisation is a crucial generator of social inequality, and point to the importance of considering the historical context when analysing the current global contours of inequality.

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Introduction

Inequality is one of the most-discussed issues in contemporary social sciences, and in national and global politics (Milanović 2005). Over the past decade, the study of inequality has advanced considerably. Large quantities of data have been collected on a (nearly) global scale, and increasingly sophisticated analyses of these data have been conducted. The aim of these analyses has been to identify the different dimensions of 'inequality', the processes that led to divergence and convergence in these trends, and the consequences of existing disparities (Therborn 2006). Economists in particular have devoted a great deal of energy to conducting global studies of the contemporary distributional dynamics. Economic historians and demographers have also contributed to this discussion by generating findings that have improved our understanding of historical inequalities in the spatial distribution of wealth, income, and well-being around the globe (e.g., Van Zanden et al. 2014a, 2014b; Milanović et al. 2011; Klüsener et al. 2014).

Gender inequality has always been a crucial element in these debates. Interest in the issue of gender equality has been fuelled by the recognition that women play important roles in a wide range of development outcomes (World Bank 2011). While we have extensive scientific evidence on broad movements aimed at achieving gender equality in many parts of the world, whether we are able to monitor the impact of these efforts efficiently depends on our ability to measure forms of sex-related inequality across societies. Over the past three decades, specific measures have been developed that capture different aspects of gender inequality in outcomes, and that focus on the institutions that perpetuate gender disparities (for a review, see Malhotra et al. 2002; Klasen 2006; recently Carmichael et al. 2014; Dilli et al. 2015).

Although the body of literature on gender inequality is large, most of the existing studies on this issue have at least two main drawbacks. First, there is a striking absence of long-term perspectives in many quantitative gender inequality studies. None of the composite gender indices used in the developmental literature pre-date the 1990s, and while recent attempts to provide greater historical depth via the Historical Gender Equality Index (HGEI) undeniably further the current measurement spectrum, they represent only a moderate step forward in terms of providing a long-term perspective, as the gender measures they are based on go back only as far as the 1950s (see Carmichael et al. 2014; Dilli et al. 2015). When we move further back in time, the available indicators tend to narrow quite dramatically, and are non-existent for larger social groups and geographical population clusters in the more distant past

(Drwenski 2015). This lack of historical data is a potentially serious problem for scholars of contemporary trends in gender inequality, because variation in levels of gender inequality may have historical roots, and the processes through which women have gained greater equality in terms of rights and socio-economic standing have unfolded over a long period of time (Dorius and Firebaugh 2010).

Second, while various authors have stressed the multidimensionality of gender inequality (e.g., Whyte 1978; Schlegel 1972; Mason 1986; Young et al. 1994), most of the analyses have tended to investigate gender discrimination separately from other associated forms of discrimination. However, it has been suggested that gender inequality is inextricably intertwined with other systems of inequality (Coltrane and Adams 2000: x; cf. also Young et al. 1994:61), and especially with discrimination by age (seniority), or the institutionalised superiority of older family members relative to younger family members. According to Therborn (2004:13-14), age discrimination and gender inequality are the two 'basic intrinsic dimensions' of patriarchy. These two forms of expropriation are dialectically related, and often act to reinforce each other in fostering a complex hierarchy of authority patterns based on both age and gender (see Joseph 1996; Dyson and Moore 1983; also Halpern et al. 1996).

In this paper, we make a fourfold contribution to this on-going discussion. First, we propose the use of a new inequality measure: namely, the Patriarchy Index (later PI), which encompasses gender and its related discrimination dimension (i.e., seniority) (see Gruber and Szołtysek 2016). Second, we apply this new measure to examine regional patterns, thereby providing geographical richness and temporal depth to existing accounts of gender and generational inequalities in the European past. Third, we demonstrate that variation in the PI across Europe is highly correlated with spatial variation in contemporary measures of gender inequality, and suggest that variation in gender equality across Europe is subject to path dependencies. A fourth distinctive contribution of our paper is that we attempt to explore the relevance of specific socio-economic, institutional, and locational characteristics for explanations of historical variation in patriarchy across Europe.

The text is organised as follows. First, we present our data and explain how they were used for the construction of the PI. Next, we illustrate how the PI is applied to data for 266 regional populations of historic Europe, located from the Atlantic coast to Moscow. In two subsequent sections, we check for correlations between our measure and other gender inequality measures, and then present a spatially sensitive regression analysis of the relationship between patriarchy levels and broad variations in socio-economic, institutional, and locational characteristics across Europe. We conclude by highlighting the implications of our findings for research on historical levels of inequality and comparative development.

Data

Because historical measures of inequality are difficult to develop any attempts to study this issue on a larger scale are contingent upon the availability of relevant data (Johnston 1985). In constructing a composite historical measure of age and sex discrimination, we relied on census and census-like microdata. We chose these data because of their broad availability across historic Europe. Thanks to the Mosaic Project and the North Atlantic Population Project (NAPP), such data are publicly obtainable in the form of machine-readable, harmonised microdata samples that are relatively easy to process (Szołtysek and Gruber 2016; Szołtysek 2015a; Ruggles et al. 2011) (see Table 1 and Figure 1 below show the distribution of regions covered by Mosaic and NAPP across Europe)[2].

The Mosaic Project (Szołtysek and Gruber 2016) currently encompasses 115 regional populations of continental Europe captured through various kinds of historical census and census-like materials other than full-count national censuses (e.g., local fragments of censuses, church lists of parishioners, tax lists, local estate inventories). It contains data going back to 1700, or even earlier[3]. In order to cover Great Britain and Scandinavia, we also decided to draw upon historical national census public-use microdata from the North Atlantic Population Project (NAPP; see Ruggles et al. 2011)[4]. The Mosaic and the NAPP microdata samples are very similar in terms of structure, organisation, and the types of information they provide. All of the samples describe the characteristics of individuals in a given settlement or area grouped into households (co-resident domestic groups), and provide information on the relationships between co-resident individuals. All of the demographic variables stored in these two datasets

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[2] www.censusmosaic.org; https://www.nappdata.org/napp/.

[3] Even though the Mosaic data are based on various sampling schemes (which are in turn contingent upon data availability), they cannot be considered a probability sample of the historical European societies or of the cultures for which the Mosaic database provides information.

[4] In order to minimise the possible modernisation effects of the 19th century on patriarchal patterns, we gave preference to the oldest available NAPP data for north-western Europe. It was possible to obtain data for Iceland, Denmark, and Norway for the late 18th/early 19th centuries; while for Sweden (1880) and Great Britain (1881) we were forced to use NAPP data from the late 19th century (the data for Great Britain in 1851 were highly clustered, and were therefore not considered). Except in England, where we employed a 10-percent sample, we used 100-percent samples. All of the other data from Great Britain represent 100-percent samples.

are harmonised across space and time using common international standards, which allows us to generate historical localised gender and generational indicators across multiple locations.

Since we situate our approach at the meso level of comparative analysis, our units of analysis are ‘regions’. The regions in the NAPP data are the administrative units that were used in the respective census, and that were considered by NAPP. The Mosaic data are organised by separate locations, which in most cases also represent separate administrative units. However, as the Mosaic data for a given region are often not complete, and since we lack information on the exact administrative boundaries of many of them, we had to use more flexible approaches (see Szołtysek and Gruber 2016:44). As a rule of thumb, we ensured that each Mosaic region had at least 2,000 inhabitants, and that urban and rural settlements were separated. In a few cases, enumeration data from the same unit (usually urban) collected at different time periods are treated as independent regions. Overall, our analysis covers 266 regional populations (see Table 1).

We grouped these regions into seven larger territorial clusters designed to capture the range of institutional and socio-economic characteristics across Europe. The NAPP data were used in the Scandinavia and the Great Britain clusters. The Mosaic data were divided into the following clusters: Germany (German-dominated areas other than the Habsburg territories), West (areas west and south-west of Germany), Habsburg, East (east-central and eastern Europe, including the former Polish-Lithuanian Commonwealth and Russia) and Balkans (areas south and/or east of Croatia and Hungary). In the regression analysis, we further subdivided some of these regions when introducing regional dummies. Our motivation for this decision was a desire to account for within-region variation in the PI levels, which remained unexplained in the models with all of the covariates other than the regional dummies. The Great Britain cluster was subdivided into England, Scotland, and Wales; the East into Central-East (Poland) and East (locations further east); and the Balkans cluster into Albania and Southeast (Figure 1).

Table 1 somewhere here

Figure 1 somewhere here

The 266 regional populations cover large parts of Europe, and run across many – though not all[5] - important fault lines in the European geography of demographic regimes (Hajnal 1982; Szołtysek 2015a). Furthermore, our dataset covers a large share of the variation across Europe in terms of geographical features, populations, cultures, and socio-economic geography: i.e., plains, mountains, and coastal areas; the free and the un-free peasantries; a variety of ethnicities and religions; and a range of regional patterns of economic growth in the early modern and modern eras. Of the 266 regional datasets, a slight majority (59 percent) represents populations after 1850, while 41 percent cover populations before 1850, and 16 percent populations before 1800. The collection includes information on both rural and urban sites, although rural societies clearly predominate[6].

Patriarchy and its composite measure

In line with a number of recent theorists, we see patriarchy not as having a single form or site, but as encompassing a much wider realm (cf. Kandiyoti 1988; Joseph 1996). According to Therborn, patriarchy has two basic intrinsic dimensions: 'the rule of the father and the rule of the husband, in that order' (2004:13-14). Thus, patriarchy encompasses both stratification of social attainment by sex and the domination of men over each other based on the seniority principle (Joseph 1996). Halpern, for example, showed that the multifaceted nature of the Balkan patriarchy was historically anchored in the interlocking combination of the rule of the father, the eldest man in the family, and the husband (Halpern et al. 1996)[7].

Based on these considerations, we conceptualised 'patriarchal' elements as clustering in the four 'domains' that we believe capture the four major dimensions of the phenomenon under consideration: the domination of men over women, the domination of the older generation over the younger generation, the extent of patrilocality, and the preference for sons. Table 2 provides a list of the components we considered. The table also shows how we defined and

[5] The current scope of Mosaic does not cover the main Iberian and Mediterranean countries, like Portugal, Spain (except for Catalonia), Italy, and Greece. This gap in the data impedes our ability to explore the north-south dimension of variation in family systems across Europe, as has, for example, been discussed by Reher 1998.

[6] In general, the Mosaic data consist of regions formed by one or more locations that were either urban or rural, while the NAPP data were analysed according to the regional division in the census. This implies that the regions based on the NAPP data usually comprise both urban and rural populations. The definition of urban was not the same across all of the NAPP data, but we took the information provided in the microdata. As the censuses of Iceland in 1703 and of Norway in 1801 do not provide such information, we have assumed that these regions were predominantly rural. However, the Norwegian city region of Christiania was treated as urban.

[7] To the best of our knowledge, Malhotra's et al. (1995) remains the only formal specification of 'patriarchy'. However, the authors focused solely on the gender aspect of patriarchy.

measured these components, and it indicates the expected direction of their relationship with societal patriarchy levels (+/-) (for a comprehensive discussion of all components and age standardisation, see Gruber and Szołtysek 2016).

Table 2 somewhere here

We chose these components because we believe they capture the most essential aspects of particular domains, given the data constraints. Most of the component variables directly capture various forms of gender and generational biases at the household level. Other variables, like patrilocality, proxy behavioural patterns that could not be derived directly from our data (in this case, inheritance practices). For the most part, we chose to use individual-level age-specific measures instead of household-level variables, because the former tend to minimise the undesired influence of variation in demographic conditions on indicators of family structure (Szołtysek 2015b). For example, instead of using the incidence of three-generation households for the domain *generational domination*, we decided to use the generational patterns of headship, the age-specific patterns of household formation, and the residential patterns of the aged. Accordingly, we chose not to consider the proportion of the elderly living with a married son (another common demographic measure), because without the inclusion of information on headship it is a poor measure of the level of patriarchal behaviour in the domestic group.

From our component variables we derive the Patriarchy Index (later PI) as a single composite measure, following the strategy detailed elsewhere (Gruber and Szołtysek 2016). The PI characterises the situations of women, the aged, and young people according to the extent to which they had obtained socially valued resources (such as a desirable position or status); albeit without measuring the positions of these groups relative to certain normative standards or reference categories. The index values thus represent absolute, not relative measures of gender and age inequality (see Johnston 1985:233 ff.; Young et al. 1994:57-58). Table 3 presents a summary of the descriptive statistics for all of the variables considered for the computation of the index.

Table 3 somewhere here

The index's domains were shown to be positively correlated with each other at a significant level (although none of these correlations were exceptionally high) (Gruber and Szołtysek 2016). As we can see in Figure 2, similar results were found for the relationship between gender and generational domination; variables that we assume are intertwined. This point needs to be emphasised in the context of Todd's (1987) argument that high female status and strong parental authority over children were often present simultaneously in historic populations, and led to increased investments in human capital. However, the empirical evidence we provide in Figure 2 does not show the presence of a high degree of female agency combined with a high degree of parental power. Overall, the within-index relationships we found are reassuring, as they both validate the use of this variable as a measurement of patriarchy, and justify our claim that it is important to explore gender and seniority biases in conjunction with each other.

Figure 2 somewhere here

Deriving gender and generational biases from domestic co-residence data—like the data used in the Mosaic/NAPP format—has certain theoretical merits. The household not only played an essential role in the functioning of preindustrial economies and societies (Szołtysek 2015a); it represented the most basic arena in which kinship bonds were formed, socialisation occurred, and values were transmitted. Those values were concerned with issues of power and equality, justice and gender relations, age hierarchy, and the relationship between the individual and the authorities (Kok *forthcoming*). Because family and household organisation patterns affected the status of women, the level of investment in human capital, and the persistence of specific cultural norms and values, it is perhaps not surprising that there is a strong relationship between prevalent family structures and development levels of regions (e.g., Alesina and Giuliano 2014; Carmichael et al. 2016). It is thus clear that the household is a particularly meaningful site for measuring gender equity and discrimination (e.g., Folbre 1986; Malhotra et. al. 2002; also Narayan 2006; see also Carmichael and Rijpma, this issue).

Spatial distribution of the Patriarchy Index

We present the distribution of the PI across space in two ways. Figure 3 charts the data geographically, while Figure 4 shows the complete scale of index points arranged according to macro-regional membership and time period.

Figure 3 somewhere here

Figure 4 somewhere here

The observed PI values range from 8 to 35 points. In the context of the data we used, we found that while all of the regional populations had at least some patriarchal features, as defined above; none of the regional populations could be characterised as fully patriarchal (maximum PI: 40 points). At the most general level, the ranking of the regions is broadly consistent with previous findings from the historical demographic and sociological literature, and seems to confirm the well-known east-west pattern (Hajnal 1982; Therborn 2004). Western Europe was shown to be much less patriarchal than eastern and south-eastern Europe. If we look at the map (Figure 3), we can see that patriarchal features become increasingly prevalent as we move east and south of the Danube after it passes Vienna; and east of the Bug River, a tributary of the Vistula river, where Polish and Ukrainian ethnicities converge; and then farther into the territories of European Russia.

This generalisation is, however, subject to some qualifications. While it is indeed the case that the areas around the North Sea Basin had relatively low patriarchy levels, similarly low levels were also found in parts of Germany and the areas of Scandinavia near the Baltic Sea. Especially in the cities in today’s eastern Germany, which is adjacent to Scandinavia, the levels of patriarchy appear to have been low. Indeed, patriarchy levels were low in regions spread across a vast area of Europe, ranging from Iceland and Great Britain; through northern France, the Low Countries, and parts of Germany and Scandinavia; into Poland and Austria. Equally interesting is the long spread of medium patriarchy levels between our dataset’s western, eastern and southern opposites, thus linking Catalonia and southwestern France with various culturally and geographically disparate areas of Westphalia and Tyrol, and with a long vertical axis stretching from Lithuania to Wallachia (Romania) in Southeastern Europe. In contrast to the results presented in the mainstream literature, our findings indicate that areas with elevated PI values also existed in north-western Europe, such as in the “Bible Belt” in the south-western part of Norway, in north-western Germany, and on the Shetland Islands.

We also observed a considerable degree of variation within countries and across the macro-regions of Europe. The territories between the Baltic, the Adriatic, and the Black Seas seem to have been particularly diverse, as they appear to have encompassed areas with low levels of patriarchy (like the western and northern parts of historical Poland), as well as areas with moderate to high levels of patriarchy (like many parts of Hungary, Slovakia, and Romania). In fact, historical Poland-Lithuania (which included modern-day Poland, Lithuania, Belarus, and large parts of Ukraine) is the only historical region for which we found a combination of high-to-low patriarchy intensities, possibly indicating a transitory, intermediate pattern (Szołtysek 2015b). Our results also show that the German territories had highly diverse PI values, ranging from very low to medium levels.

Furthermore, Figure 4 suggests that the decline in patriarchy levels may not have been continuous or linear; i.e., that conservative patriarchal cultures did not necessarily evolve into ‘modern’, gender-egalitarian societies. For example, many regions of Sweden had higher patriarchy levels than Denmark, even though the Swedish census was taken almost a century later than the Danish census. Moreover, Iceland had much lower PI values than Norway, despite having been surveyed one hundred years earlier. Similarly, data from Germany show that the PI values improved little from the early to the late 19th century. While regions of eastern Europe are underrepresented in the data from later periods, there are no clear signs that the patriarchy levels in these regions declined drastically with the passage of time. Finally, the data for southeastern Europe indicate that Albania of the early 20th century was much more patriarchal than several populations of the Balkans in the early 19th century or earlier.

The Patriarchy Index and other family system measures

Given the character and the geographic distribution of the PI, a considerable overlap between this measure and some common measures of historical family systems is to be expected. This assumption seems to be confirmed by our attempt to match our findings on the distribution of patriarchal features with the results of Dennison and Ogilvie (2014). These authors created a Borda ranking of European societies based on what they called ‘the three European Marriage

Pattern criteria' (EMP): female marriage age, female celibacy, and household complexity (Dennison and Ogilvie 2014:669-670). To compare our results with those of Dennison and Ogilvie, we started with their division of Europe into 33 societies (some of which included country subdivisions), which they derived from 365 research studies. We then used our dataset on 266 regional populations to derive the average PI values for this division. This approach allowed us to cover 18 out of the 33 societies studied by Dennison and Ogilvie. To ensure that big cities did not dominate the outcomes, we gave each region the same weight regardless of its population size in obtaining these averages. However, when interpreting the results of this comparison, it is important to note that neither the values obtained by Dennison and Ogilvie (2014) nor our values are representative in a strict statistical sense. Nevertheless, we believe that this comparison provides us with an impression of the relationship between these two measures. The scatterplot derived from this comparison is presented in Figure 5 (a regression line was added to the scatterplot for orientation). The outcomes of the analysis suggest that there is a positive relationship between the PI and the EMP rankings (Pearson's r: 0.86): i.e., societies with a high PI are more likely to be characterised by early female marriage age, low rates of female celibacy, and high levels of household complexity. While this finding might not be particularly surprising, it provides support for the view that the PI is a useful measure of historic cross-cultural differentials in family organisation. Such a measure is more comprehensive than the usual triad of features commonly studied in family history research: namely, age at marriage, celibacy, and household structure (Hajnal 1982; Gruber and Szołtysek 2016).

Figure 5

While it is appropriate to use the PI for such purposes, it is important to note that the observed distributions of patriarchy levels may not necessarily overlap with the spatial patterning of the three main types of family systems that are commonly assigned to historic European societies: neolocal nuclear, patrilocal stem, patrilocal joint. While it has been argued that 'gender bias informs [the three main types of] family systems in contingent fashion' (Skinner 1997:58), and that this bias is most pronounced in joint family systems and is least pronounced in conjugal family systems, we believe that the family system-patriarchy relationship is more complex (Fig. 4 above). For example, our findings show that the regions in which the conjugal-neolocal family model was most prevalent (like northern France and Romania) can still be distinguished by their relative patriarchy levels; and that the European regions in which the joint family was prevalent (e.g., in Albania, Slovakia, Lithuania, central Ukraine, and Russia) had a range of patriarchal values. Although these observations are still tentative, they may invite family historians to reassess their conceptual apparatus.

The Patriarchy Index and other gender inequality measures

In evaluating the relevance of the PI for wider inequality studies, two potential caveats need to be addressed. First, sex-related or age-related social inequalities are frequently not limited to the realm of family, as they are in the PI, but encompass other dimensions of social life, such as politics and labour markets (see; Young et al. 1994:57, 59). However, as we argued above, we believe that the PI is relevant for social dimensions beyond those associated with the family, as it appears that household organization practises helped to uphold systematic forms of gender and generational biases within societies. Nonetheless, future research should compare our results with the findings of historical patriarchy studies that are based on other sources and that cover other spheres; provided such studies are conducted on a sufficiently large scale.

Second, it is important to remember that unlike most existing social science indicators of sex discrimination, the PI does not just reflect gender differences (or women's status), but merges the gender dimension with that of seniority. While this approach provides a more comprehensive account of the multidimensionality of empowerment and agency than most other measures of patriarchy offer (see Narayan 2006:74-75), it makes it more difficult to compare the PI with more gender-focused measures.

Given these ambiguities, it is interesting to explore the question of to what extent the variation in the combination of various historical family-related institutions and societal mechanisms that the PI captures is similar to the present-day spatial variation in macro indicators of gender inequality. We therefore decided to compare the historical variety in PI levels with today's (2013) spatial variation based on a well-established measure from inequality research: the Gender Inequality Index (GII)[8]. While some of the GII's components correspond

[8] The GII measures gender inequalities in three important aspects of human development: reproductive health, measured by the maternal mortality ratio and the adolescent birth rates; empowerment, measured by the proportion of parliamentary seats occupied by females and the proportion of adult females and males aged 25+ years

roughly to certain components of the PI, they were obtained in a different manner using contemporary data. In deriving the PI values for this assessment, we followed the procedure we used in the comparison of our data with the Dennison and Ogilvie data, except that in this case we were basing the divisions on the present-day countries for which GII data are available. As in the exercise above, we compared only those countries for which PI data are existing.

Figure 6 somewhere here

A comparison of the GII levels with our derived PI values is presented in Figure 6. Again, the comparison should be interpreted with caution, as our PI data for the present-day countries are not representative in a strict statistical sense. The graph we derived suggests that there is a rather strong positive relationship between historical patriarchy levels and the GII values (Pearson's r: 0.65). Thus, it appears that areas which had relatively high patriarchy levels in the past also tend to have relatively high gender inequality levels today. Although this comparison has some limitations, we believe that it merits attention.

A similar approach can be followed in comparing the PI with the Historical Gender Equality Index (HGEI) (Dilli et al. 2015). The HGEI represents the most recent methodological innovation in the measurement of gender inequality worldwide (Figure 7)[9]. For our comparison, we decided to use HGEI data for the years 2000-2010, as they are the most complete. Again, we found a clear relationship between the two measures: countries with high scores on the historical patriarchy scale tend to have low levels of gender equality today (Pearson's r: -0.72).

Figure 7 somewhere here

with at least some secondary education; and economic status, expressed as labour market participation of the female and the male populations aged 15+ years. See http://hdr.undp.org/en/content/gender-inequality-index-gii.

[9] This measure aims to detect gaps between men and women rather than absolute levels of achievement; and especially indicators of the unequal treatment of women. The HGEI captures gender differences in life expectancy, labour force participation, infant mortality, educational attainment, marriage age, and political participation. We thank S. Carmichael and A. Rijpma for sharing their data with us.

Beyond having a purely diagnostic dimension, the finding that historical and contemporary inequality patterns are linked suggests that gender disparities persist over the long term. The mere establishment of such associations does not, of course, allow us to posit the existence of direct causal links between the past and the present. If historical patriarchy levels influenced contemporary gender inequalities, they probably did so in a path-dependent manner. But even without making strong claims about the lasting impact of historical patterns on the patterns that exist today, we believe that the observed associations are of relevance for social scientists and researchers engaged in comparing developmental levels, as these links imply that the historical context is an important consideration when analysing the global contours of contemporary forms of social inequality.

Patriarchy in context

Why did some historical societies have higher levels of patriarchy, while others had lower levels? Referring to medieval England, Bennett (2007:78) suggested that ‘patriarchy was an effect of many institutions’, but did not explain this observation any further. Therborn argued that a process of ‘de-patriarchalization’ took place, and asserted that this development was influenced by legal changes, proletarianisation, and wider processes of urbanisation and industrialisation (2004:17-22; similarly, Moghadam 1992; Miller 1998). According to Rahman and Rao (2004), the key determinants of female inequity were cultural norms (especially regarding kinship), economic conditions, and state policies and legislation. Meanwhile, Alesina hypothesized that traditional agricultural practices influenced the historical gender division of labour and the evolution of gender norms (Alesina et al. 2013; earlier, Boserup 1970; see also Carmichael and Rijpma, this issue).

Using our data to address the question posed above is a challenging endeavour, especially given the inherent difficulties we face in obtaining comprehensive information on potential covariates from the surviving body of historical statistics. Since the PI captures a multidimensional phenomenon, it is associated with wide range of socio-demographic and cultural dimensions. Hence, any modelling attempt is likely to be confronted with multicollinearity issues. In addition, since many of the cultural and institutional traits that may influence patriarchy are likely to be influenced by patriarchy themselves, the relationship might go both ways. Thus, the investigation that follows merely represents a first attempt to explore a number of hypotheses related to conditions that are potentially relevant for understanding variation in patriarchy levels. In this analysis we derived spatially sensitive regression estimates[10] of the associations between patriarchy levels and a broad range of socio-economic and institutional characteristics of the regional populations, as well as regional variation in the degree of demographic spatial centrality and environmental conditions[11].

The Patriarchy Index serves as our dependent variable. In attempting to account for variation in living standards, we decided to include the proportion of the population who were elderly (aged 65+) and the child-woman ratio (CWR) as covariates. The latter indicator is the ratio of children under age five to the number of women between ages 15 and 49 (see Willingan and Lynch 1982:102-104)[12]. We assume that regions with a relatively high level of development also had relatively high proportions of elderly people and relatively low patriarchy levels (Rosset 1964:209-210, 231)[13]. To interpret the CWR, we have to take endogeneity concerns into account, as the CWR might be more than just a proxy for the level of development. It is also likely that the link between patriarchy and fertility levels was positive (Dyson and Moore 1983). Overall, however, we expect to find that in areas with relatively high levels of development both the CWR and the patriarchy values would have been low[14].

To explore the potentially 'depatriarchalising' role of urban (industrial) life, we included a covariate for the share of the population in each region who were living in rural areas (see ft. 6). Next, in order to account for whether the region was more centrally or more peripherally located within Europe, we derived a 'population potential' covariate (see Stewart and Warntz 1959). This provides information whether a specific region was situated close to important population centres of Europe, or rather in peripheral sparsely populated areas (see Appendix 1 for technical details). During our period of observation, the cost of transport was still an essential factor in the extent to which people had access to markets. Accordingly, the potential for economic growth was usually relatively low in peripheral areas that were situated

[10] We decided to use robust regression since it is less affected by violations of OLS assumptions, and it allowed us to minimise the effects of the outliers detected in our database. In every model, we used the MM-type regression estimator described by Yohai (1987) and Koller and Stahel (2011), which was implemented in the R library robustbase.

[11] For two out of the 266 regional populations covered by our dataset (on the English Channel Islands), we were unable to derive all of the geocovariates. These regions had to be excluded from the analysis.

[12] In the CWR the relationship between the number of children and the number of potential mothers is usually multiplied by 1,000. But to avoid small coefficient values in our regression results, we decided to use this ratio without such a multiplication.

[13] While it could be argued that less-developed outmigration regions might have had a higher proportion of elderly people, we believe that this pattern did not prevail during the observation period.

[14] It is possible that in peripheral regions with low levels of development children were under-recorded in the censuses. However, we do not believe that this issue affected the association between the CWR and patriarchy levels on a European scale.

far away from important population centres. In addition, the extent to which large numbers of people were living in close proximity was an important factor in determining the degree to which knowledge and skills were diffusing and being maintained (Goldin 2016:59). Overall, we expect to find a negative association between the population potential and the PI.

A control for terrain ruggedness was included to account for variation in the potential for economic and social development (Jimenez-Ayora and Ulubaşoğlu 2015; see Appendix 1 for technical details). Rugged topography may affect the ability of residents to engage in intensive agricultural activities, and their access to public infrastructure such as educational institutions or the transport system. This issue was of particular relevance in the period between 1850 and 1950, when having access to railways was an important determinant of the developmental prospects of a given region. Moreover, in regions with rugged terrain cultural anomalies may persist longer as such a terrain constrained at least in historical times people to communicate with individuals in nearby locations. However, the relationship between the ruggedness of a region and its development prospects is not clear-cut, as some of these areas had access to water energy and/or mineral deposits. Exploiting these resources would have provided the populations in these regions with opportunities to engage in proto-industrialisation; a process that is generally associated with the depatriarchalisation of family relations (Medick 1976:303). Overall, however, we expect to find a positive association between terrain ruggedness and PI levels across Europe.

We also added a covariate that indicates whether the populations were subjected to serfdom. According to our considerations there are three possible channels through which serfdom may have increased the patriarchal bias among these regions. First, the regions with serfdom tended to have more complex families that were more likely to exhibit high PI levels. Second, the Russian version of serfdom in particular provided conditions in which the authority of the household patriarch was institutionally endowed by the seigniors. Finally, because of its heavy reliance on coerced labour with draught animals (*corvee*), serfdom created structural conditions that devalued female labour (Szołtysek 2015b, vol. 1; Alesina et al. 2013). Thus, we can assume that serfdom had negative effects on women's status and agency levels.

Another important aspect that should be considered in this context is the classification according to the period in which all or most of the data for each of our 266 regional populations were collected. Based on the general consensus of the sociological and the historical literature (Therborn 2004), we expect to find that patriarchy levels decreased over time. For our models, we considered the following categories: pre-1800, 1800-1850, and after 1850 (reference category). Finally, we included dummies for 11 regions of Europe (as depicted in Figure 1) in an attempt to account for unobserved developmental effects, such as the efficiency of the bureaucracy, the role of the labour markets, and the legal system (with Germany used as a reference category). In addition, we considered several other variables, such as the share of cropland, the rules of descent, and the dominant religion. However, for a number of reasons we decided not to account for these variables in the analysis[15]. As the included attributes are limited, the regression models should not be interpreted as an attempt to establish causality. The main purpose is to explore the association between the PI and the available covariates in a multivariate framework.

To account for variation in the density of locations across various parts of Europe, we decided to apply weights to ensure that each of the seven large regions is given equal weight in the regressions (Balkans, East, Germany, Great Britain, Habsburg, Scandinavia, West). As we are analysing spatial data, it is likely that our regressions are influenced by spatial autocorrelation, which might introduce bias into both the coefficient estimates and the obtained significance levels (Bivand et al. 2013). To explore the degree to which our models are affected by spatial autocorrelation, we derived the Moran's I index of spatial autocorrelation for the model residuals. The Moran's I is very similar to the Pearson's correlation coefficient, except that it does not control for the correlation between two variables x and y, but rather for the correlation between a variable y in region i, and the value of the (weighted) mean of y in neighbouring regions j. In calculating the Moran's I, we considered the five nearest neighbouring regions[16]. In order to determine whether our models estimates are potentially biased due to multicollinearity, we applied variance inflation factors (VIF).

[15] Information on the share of cropland has been taken into account based on an argument made by Alesina et al. (2013; earlier, Boserup 1970) that levels of male dominance may have been particularly high in regions dominated by crop production. We derived these data from the HYDE database using an approach similar to the one we used to obtain the information on terrain ruggedness for the Mosaic and NAPP regions (see Appendix 1). In this case, however, we obtained the mean value in the share of cropland from the raster data. But because the model outcomes for this variable were not stable, and because this variable was highly correlated with the population potential variable, we decided to take it out. The rules of descent (Murdock 1949:15, 43-46), which is another potential determinant of patriarchy, has not been considered here because some of the PI components seem to be strongly related (if not epiphenomenal) to prevailing descent rules. Religion was not included in the models, as it was highly correlated with some of the regional dummies.

[16] Derived by calculating the spherical distances between the regions' coordinates. As the regions' coordinates for the Mosaic dataset, we used the population-weighted coordinates derived from our 1,692 Mosaic locations. For the NAPP data, we took as the coordinates for each region the location that had the highest population density within the region in 1800, according to the HYDE database (see Appendix 1).

Table 5 presents the results of our regressions. We would first like to provide an overview of the spatial clustering of the dependent variable. As we applied regional weights in the regression, we decided not to derive the Moran's I for the dependent variable, but instead to calculate a base model that simply takes into account the dependent variable, the intercept, and the weights. For the residuals of this model, we then derived the Moran's I. The results of this approach show that, in line with the visual impression in Figure 3, the patriarchy levels in our dataset are characterised by extremely high levels of positive spatial autocorrelation. In Model 1, we only controlled for our demographic proxies for development levels and the time period. In this model, only the proportion of elderly people is significant, and the coefficient estimate is in the expected direction. However, the Moran's I on the residuals indicates that there are high levels of positive spatial autocorrelation. This finding implies that in our regressions the independence assumption is probably violated, which could in turn cause bias in the coefficient estimates and increase the likeliness to obtain too high significance levels.

In Model 2 we added all of the other socio-economic and environmental covariates. This substantially increased the r-squared, but it did not reduce the high degree of spatial autocorrelation among the residuals. Thus, the estimates of Model 2 also have to be interpreted with caution. Compared to the outcomes of Model 1, the estimated coefficient for the elderly is attenuated, but remains highly significant. The CWR and the time dummies became significant in the expected direction. However, for the time dummies we do not find the expected negative gradient. The outcomes for all of the other significant variables are in the anticipated direction. Relative to other areas, the regions with serfdom had significantly higher patriarchy levels, and those with a high degree of terrain ruggedness had higher PIs. In addition, peripheral areas with low population potential were significantly more likely than the reference group (regions with high population potential) to have had high PI levels. The outcome for the variable for the share of the population living in rural areas is significant at the 0.1-level, and indicates that rural areas had higher PI levels.

In order to further reduce the spatial autocorrelation in our models, we included our regional controls in Model 3. Furthermore, because observations for only one or two periods were available for a number of our 11 macro-regions (which made tracking changes over time within these regions very difficult)[17], we also added interaction effects between the regions and the time periods. The introduction of these regional dummies and the interaction effects

[17] This was particularly challenging in the case of Albania, where we detected the highest PI levels, while all of the observations are from the last period.

allowed us to reduce the Moran's I on the residuals substantially, to 0.03. Although this value is still significant at the 0.01 level, it provides us with some confidence that Model 3 is much less biased due to spatial autocorrelation. However, introducing the regional dummies came at a price: namely, that the dummies might act as a proxy for some of the other covariates. Our outcomes for serfdom are likely affected by this problem, as the areas with serfdom were primarily concentrated in the east. Thus, it is not surprising that the coefficient for serfdom is not significant in the full model, while we obtained highly significant positive estimates for our dummy East. This finding is further corroborated by the VIF values, which are high for these two variables in Model 3.

We refrain from discussing the interaction effects, as they are in part difficult to interpret. Of the other covariates, all of the significant variables are in the expected direction. In Model 3 we also obtained the expected negative gradient for the time dummies. However, the VIFs for these two dummies are very high. Terrain ruggedness is no longer significant, possibly due to the fact that a large share of the populations who lived in rugged terrain is clustered in certain regions (e.g., in Albania). However, the VIF for terrain ruggedness is not high. Regional dummies seem to account for a large portion of the variation. It can also be noted that the share of the population living in rural areas and the population potential variable seem to be more relevant for understanding variation in patriarchy levels in Model 3 than in Model 2. The share of the elderly in the population, on the other hand, is not significant in Model 3.

When we excluded the CWR from Model 3 based on endogeneity concerns, the outcomes for the other variables did not change substantially. We also ran separate models on the Mosaic and the NAPP regions. The outcomes of these models differ especially with regard to the population potential variable. There is a much more pronounced association between high PI values and low population potential in the Mosaic than in the NAPP dataset. A possible explanation for this result is that remoteness might have played a much bigger role in areas located on the continent compared to Great Britain and Scandinavia, as this latter group of regions likely had easier access to the sea/international trade.

While our modelling attempts were subject to a number of limitations, the outcomes provide us with some confidence that there was a negative association between development levels and PI levels in historical Europe. In particular, our findings suggest that patriarchy levels in the past were especially high in rural and peripherally located areas. The role of geovariates for understanding variation seems to be particularly pronounced for continental Europe other than Scandinavia.

Conclusions

In this study, we sought to move the analysis of historical trends in gender inequality beyond the usual confines of a one-dimensional focus on sex-stratification/discrimination. With the Patriarchy Index we proposed a historical inequality measure that combines the power of the father and the power of the husband with other dimensions in a composite approach. By doing this, we demonstrated that limited but widely available historical data can be used to construct variables that allow to measure historic trends in gender and generational relations across Europe. Moreover, we presented the argument that by comprising localised indicators which combine both gender and age relations, the index allows to better account for the historical cross-cutting of gender bias with other forms of discrimination. We believe that these contributions will facilitate the historical reconstruction of the dynamics of power in preindustrial Europe, and enhance the current body of historical statistics on cross-societal inequalities.

We applied the PI to census microdata to provide an account of the regional prevalence of gender- and age-based authority patterns across Europe with a focus on the 18th and 19th centuries. This analysis showed that the complex societies of (western) Eurasia (Goody 1976) differed significantly in their patriarchy levels as conceptualised in the PI. As the spatial contours of this variation do not necessarily align with the corresponding spatial patterning of the main types of historical family systems, family historians may wish to further explore this line of research.

Our finding that the historical PI values are associated with values obtained by contemporary measures of gender inequality provides support for the argument that variation in historical conditions, structures, and institutions can be relevant for understanding contemporary spatial disparities in development, well-being, and wealth (e.g., Nunn 2009). It also reiterates the importance of the family and of the household as historically crucial sites for generating societal inequalities (Alesina et al. 2013).

Our regression results suggest that PI values tended to be higher in areas that were more remote and less well integrated. Can we assume that the more diversified social structures of cities and densely populated regions mitigated against higher PI levels? Or should we interpret elevated PI values as adaptation mechanisms that were triggered in response to challenges and constraints created by local geographic externalities? In other words, can these higher PI values be seen having arisen in response to low levels of state penetration, weak institutions, and poor access to public services and infrastructure; and hence in reaction to broader ‘spatial poverty traps’ (Bird et al. 2010)? Future research should attempt to explore these issues further. It is equally important that scholars continue seek to better understand the relationship between patriarchal structures and prospects for human development. The Mosaic and the NAPP data could be very well suited to that line of research.

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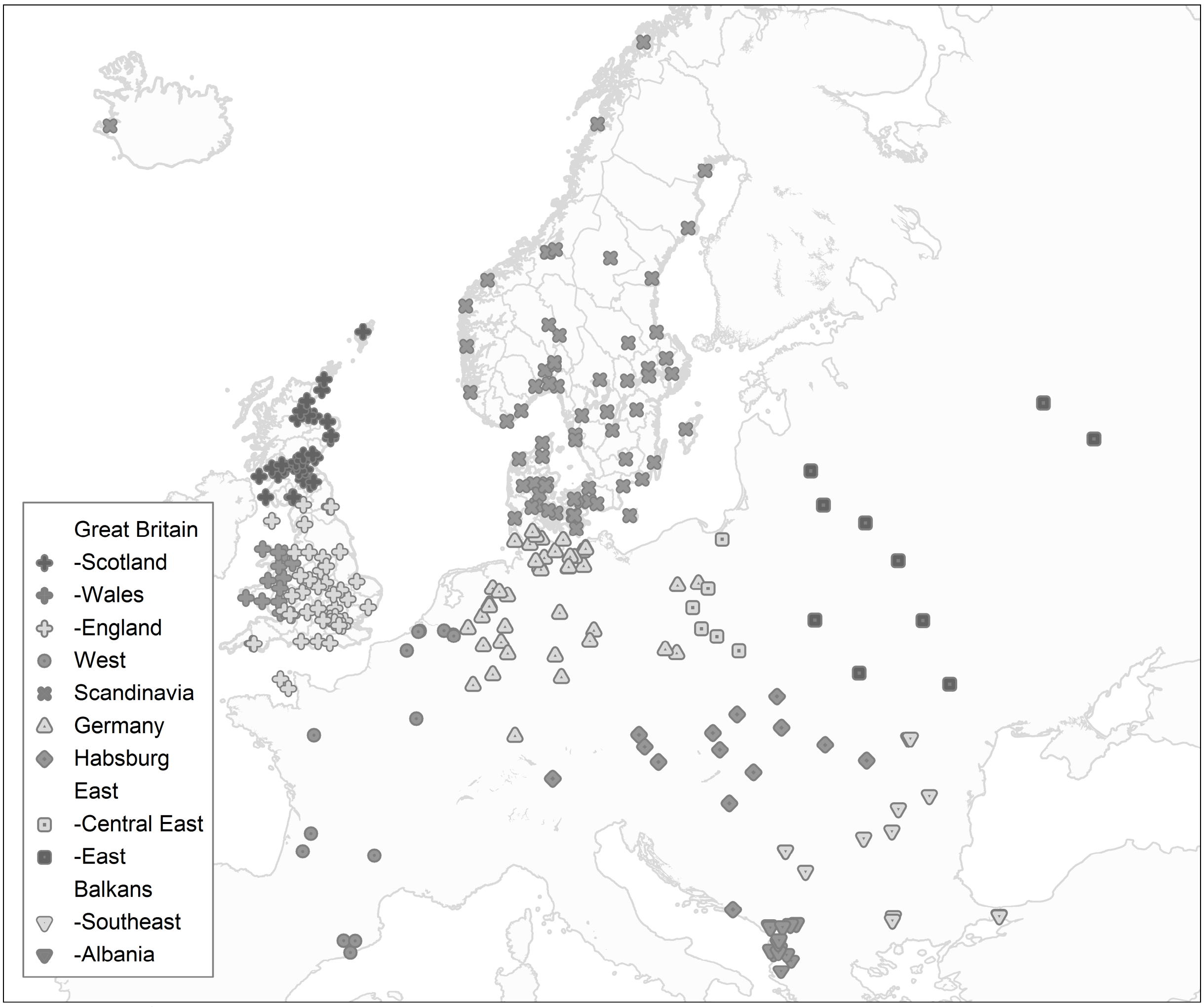
Tables and figures

Table 1: Data used for analysis

census	regions	N (=pop.)
Mosaic data:		
Albania, 1918 census	8 rural regions, 6 cities	140,611
Austria-Hungary, 1869 census	9 rural regions from Hungary, Romania, Slovakia	31,406
Austria-Hungary, 1910 census	3 rural regions and 1 city from Austria	20,036
Belgium 1814 census	1 rural region from Western Flanders	13,666
Bulgaria, 1877-1947 household registers	1 rural region and 1 city from the Rhodope area	8,373
Dubrovnik, 1674 status animarum	1 rural region from Dalmatia	1,880
Denmark, 1803 census	9 rural regions and 2 urban regions from Schleswig and Holstein	107,861
France, 1846 census	3 rural regions	16,967
France, 1831-1901 census	1 rural region from South-Western France	5,109
France, 1846-1856 census	1 city from South-Western France	5,669
German Customs Union, 1846 census	10 rural regions and 4 urban regions	36,760
German Customs Union, 1858 census	1 rural region from the East	3,468
German Customs Union, 1861 census	1 rural region from the Southwest	6,541
German Customs Union, 1867 census	4 rural regions and 1 city in Mecklenburg-Schwerin	66,938
Germany, 1900 census	1 city	55,705
Mecklenburg-Schwerin, 1819 census	3 rural regions and 1 city	37,332
Münster, around 1700 status animarum	3 rural regions in North-Western Germany	23,010
Münster, 1749 status animarum	3 rural regions in North-Western Germany	34,169
Netherlands, census 1810-1811	2 rural regions and 3 cities in the south	40,037

Poland-Lithuania, 1768-1804 listings	12 rural regions	155,818
Moldavia, 1781-1879 status animarum	2 rural regions	5,291
Wallachia, 1838 census	4 rural regions	21,546
Russia, 1795 revision lists	1 rural region in Ukraine	8,050
Russia, 1814 private enumeration	1 region in Central Russia	2,955
Russia, 1847 enumeration	2 rural regions in Lithuania and Belarus	19,917
Russia, 1897 census	1 rural region around Moscow	11,559
Serbia, 1863 census	1 rural region and 1 city	9,746
Serbia, 1884 census	1 rural region	9,434
Spain, 1880-1890 local census	1 rural and 2 urban regions in Catalonia	23,997
Ottoman Empire, 1885 census	Istanbul	3,408
Ottoman Empire, 1907 census	Istanbul	4,946
Mosaic data overall	115 regions (89 rural and 26 urban)	932,205
NAPP data		
Denmark, 1787 census (100%)	21 regions	838,623
Iceland, 1703 census (100%)	1 region	51,003
Norway, 1801 census (100%)	19 regions	878,073
Sweden, 1880 census (100%)	24 regions	4,624,825
United Kingdom, 1881 census:		
England (10% sample)	76 regions	2,926,374
Wales (10% sample)	13 regions	1,573,065
Scotland (10% sample)	32 regions	2,783,354
Islands (10% sample)	3 regions	139,614
NAPP data overall	151 regions	14,252,150

Figure 1: Spatial distribution of Mosaic and NAPP data by major territorial groupings



For sources of the Mosaic and NAPP data: see Online Appendix 2.

Table 2: Components of the Patriarchy Index

Domain/ component	Component	Abbreviation	Definition/measurement	Relationship with patriarchy	Specification
Male domination	Proportion of female household heads	*Female heads*	the proportion of all female household heads among all adult (20+ years) household heads of family households	negative	age-standardized
	Proportion of young brides	*Young brides*	the proportion of ever-married women in the age group 15-19 years	positive	
	Proportion of wives who are older than their husbands	*Older wives*	the proportion of all of the wives who are older than their husbands among all of the couples for whom the ages of both partners are known	negative	age-standardized
	Proportion of young women living as non-kin	*Female non-kin*	the proportion of women aged 20-34 years who live as non-kin, usually as lodgers or servants	negative	age-standardized
Generational domination	Proportion of elderly men coresiding with a younger household head	*Younger household head*	the proportion of elderly men (aged 65+ years) living in a household headed by a male household head of a younger generation	negative	Only family households;the elderly men must be relatives of the household head
	Proportion of neolocal residence among young men	*Neolocal*	the proportion of male household heads living without any relatives except spouses and children among ever-married men in the age group 20-29 years	negative	only family households; age-standardized
	Proportion of elderly people living with lateral relatives	*Lateral*	the proportion of elderly people (aged 65+ years) living with at least one lateral relative in the household	positive	Only family households
Patrilocality	Proportion of elderly people living with married daughters	*Married daughter*	the proportion of elderly people (aged 65+ years) living with at least one married daughter in the same household among those elderly people who live with at least one married child in the same household	negative	Only family households
Son preference	Proportion of boys among the last child	*Boy as last child*	the proportion of boys among the last children (if the last child is one of a set of siblings of both sexes, he or she will be excluded from the analysis).	positive	only children of household heads; only age group 10 to 14 years; family households
	Sex ratio of youngest age group	*Sex ratio*	the sex ratio (boys to 100 girls) in the youngest age group (0-4 years old).	positive	Only family households

Table 3: Descriptive statistics for the components of the Patriarchy Index (266 regional populations of Europe).

Component	Mean	Standard deviation	Minimum	Maximum
Female household heads	0.13	0.06	0.01	0.31
Young brides	0.06	0.11	0.00	0.66
Older wives	0.22	0.08	0.01	0.37
Females non-kin	0.21	0.12	0.00	0.57
Younger household head	0.13	0.11	0.00	0.68
Neolocal	0.60	0.23	0.03	0.97
Lateral	0.14	0.12	0.00	0.73
Married daughter	0.35	0.20	0.00	0.80
Boy as last child	0.50	0.05	0.34	0.81
Sex ratio	101.90	7.80	81.80	137.30

Source: Mosaic/NAPP projects; own calculations

Figure 2: Scatterplot of the Male Domination Index and the Older Generation Domination Index

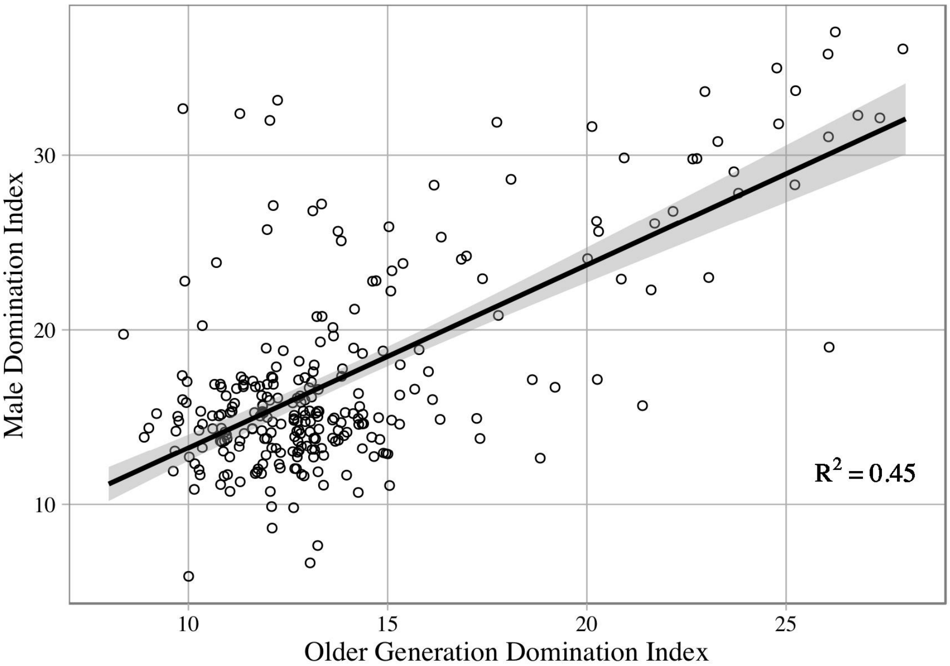


Figure 3: The spatial distribution of the Patriarchy Index

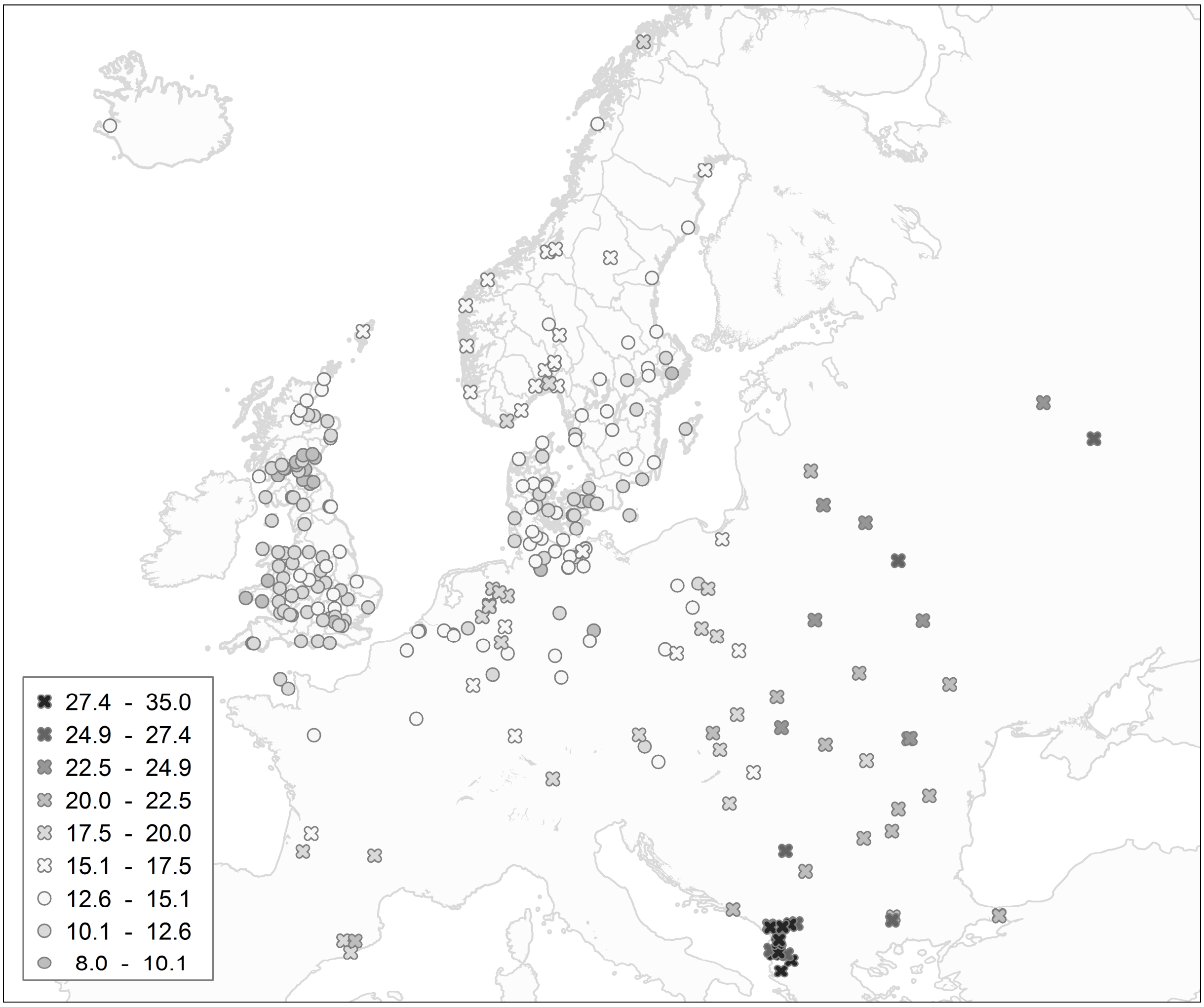


Figure 4: Regional values of the Patriarchy Index by time period and macro-geographical membership

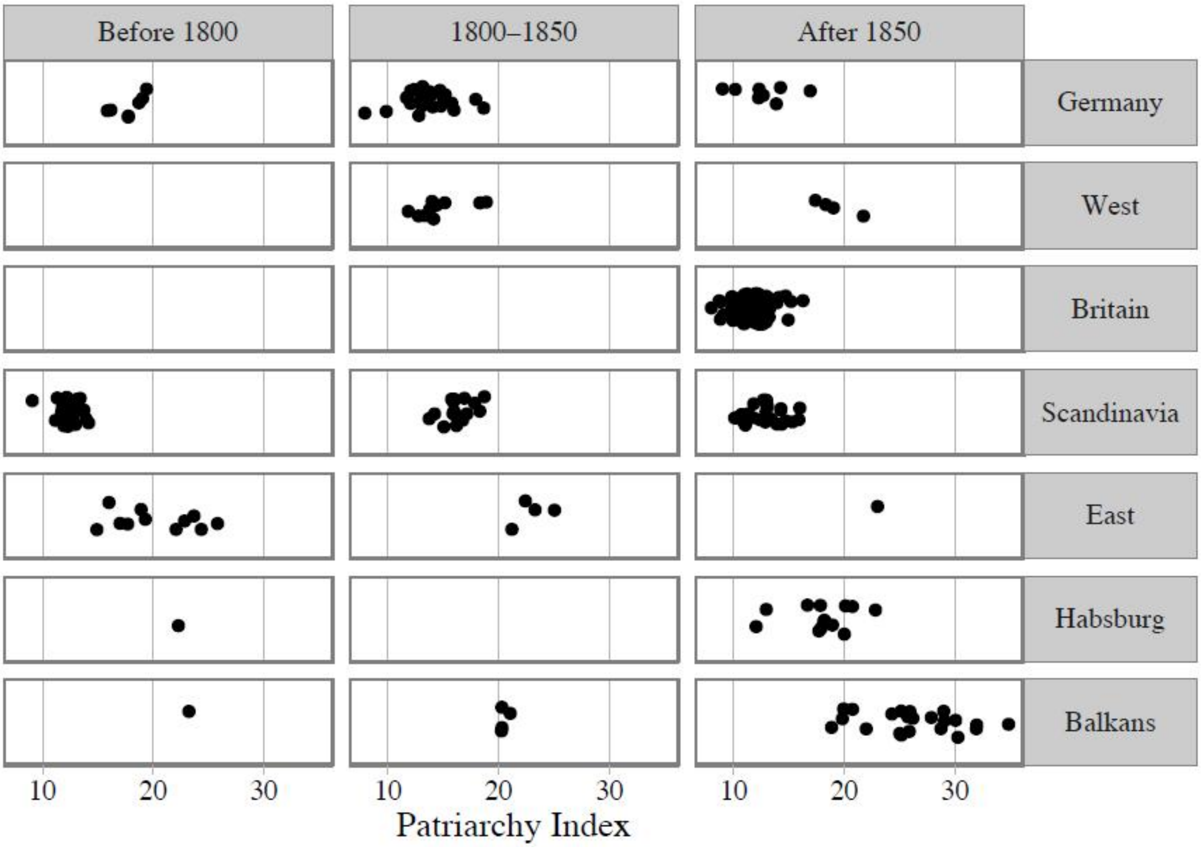


Figure 5: The relationship between the PI and a Borda ranking of European societies according 'the three European Marriage Pattern criteria' (combined Mosaic/NAPP)

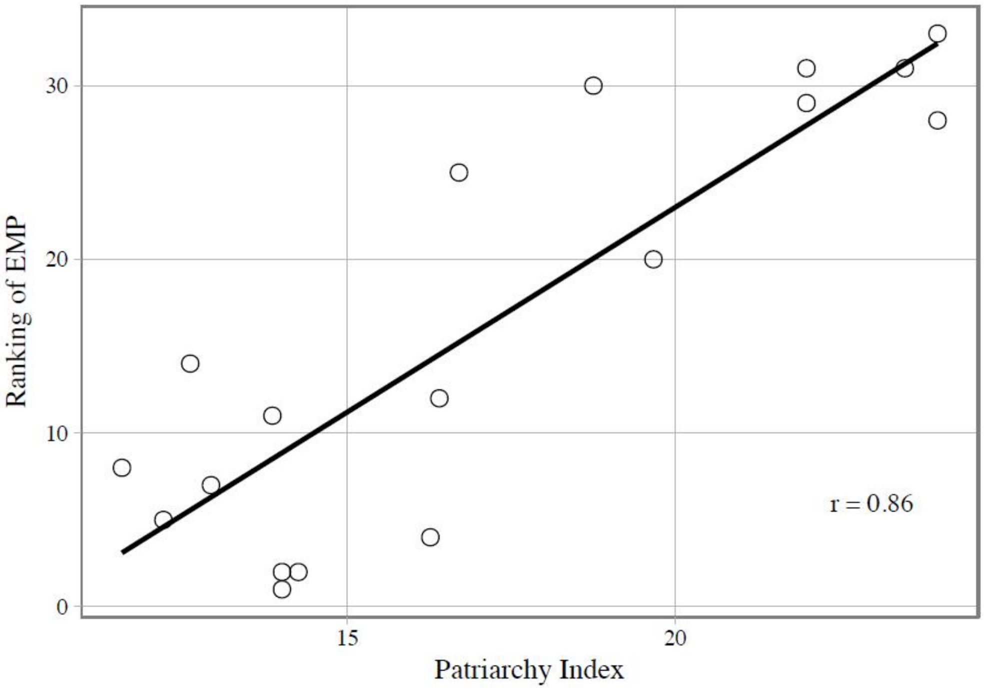


Figure 6: The relationship between the PI and the Gender Inequality Index [2013] (http://hdr.undp.org/en/content/gender-inequality-index-gii) for combined Mosaic/NAPP data

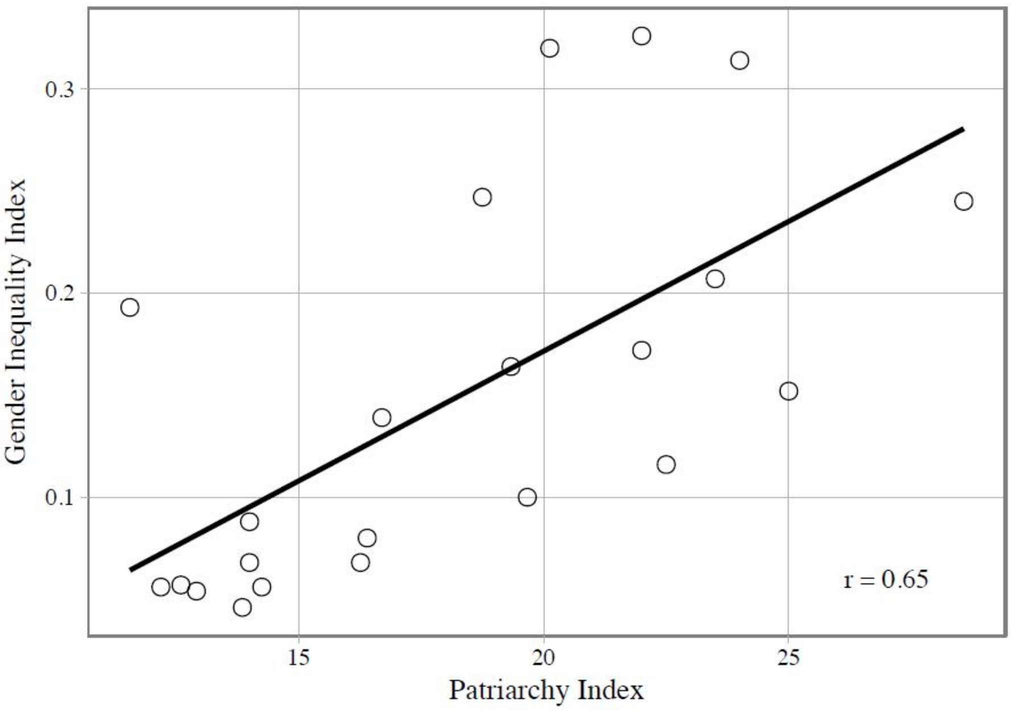


Figure 7: The relationship between the PI and the Historical Gender Equality Index for 2000 (Carmichael), for combined Mosaic/NAPP data

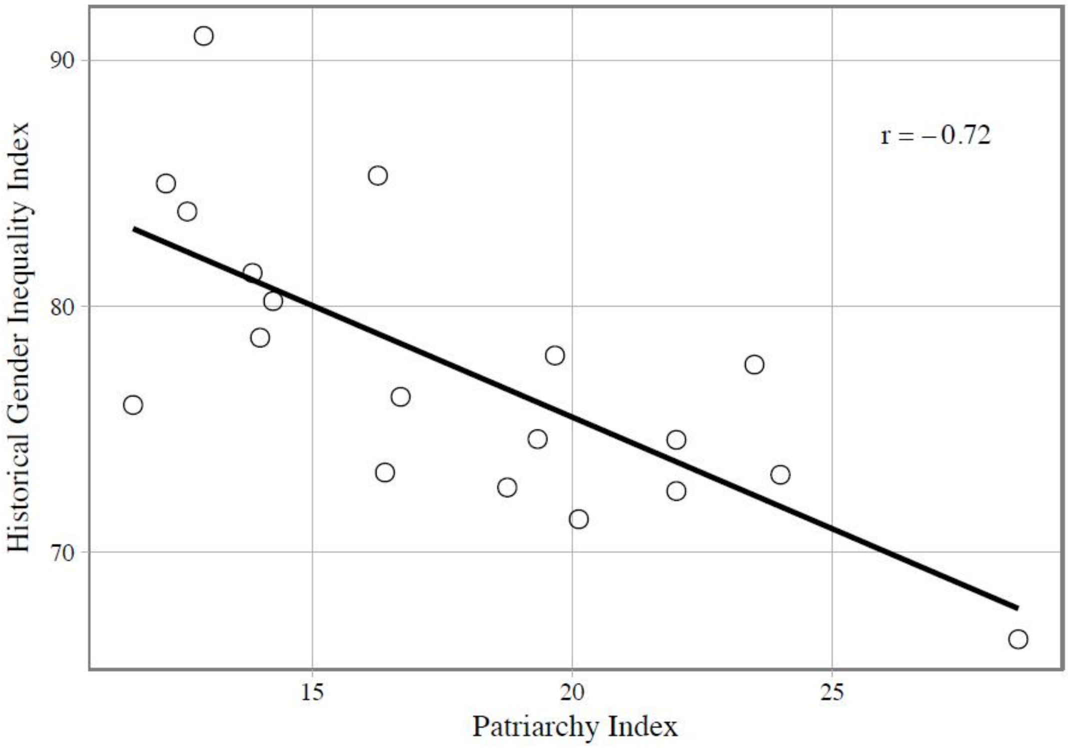


Table 5: Regression results

	Model 1					Model 2					Model 3				
	β	s.e.	std. β	p.-v.	VIF	β	s.e.	std. β	p.-v.	VIF	β	s.e.	std. β	p.-v.	VIF
(Intercept)	19.10	1.49	0.00	***	0.00	12.20	1.25	0.00	***	0.00	8.46	0.81	0.00	***	0.00
Proportion of elderly > 65	-1.09	0.13	-0.38	***	1.21	-0.82	0.12	-0.28	***	1.68	0.07	0.06	0.02		2.45
Child-woman ratio	2.97	2.12	0.06		1.18	6.18	1.81	0.12	***	1.30	5.00	1.09	0.10	***	2.41
Share rural						1.07	0.59	0.07	#	1.24	2.21	0.29	0.14	***	1.74
Population potential (ref.: High)															
- Low						1.23	0.45	0.12	**	1.42	1.94	0.45	0.18	***	6.91
- Medium						-0.53	0.41	-0.05		1.33	-0.19	0.28	-0.02		2.83
Terrain ruggedness (log-transformed)						0.43	0.14	0.10	**	1.27	-0.05	0.09	-0.01		2.81
Serfdom (yes)						3.33	0.65	0.17	***	1.86	0.30	0.55	0.02		5.90
Period (ref.: after 1850)															
- Before 1800	0.83	0.62	0.06		1.17	1.54	0.53	0.11	**	1.61	4.88	0.73	0.36	***	13.88
- 1800-1850	0.77	0.51	0.07		1.14	2.28	0.40	0.20	***	1.35	0.87	0.55	0.08		10.38
Region (ref.: Germany)															
- Albania (Balkans)											13.90	0.74	0.63	***	5.61
- Southeast (Balkans)											10.85	0.66	0.47	***	4.89
- Central East (East)											-1.11	0.74	-0.03		4.01
- East (East)											9.55	0.92	0.37	***	12.07
- England (Great Britain)											-1.52	0.60	-0.11	*	4.21
- Scotland (Great Britain)											-2.64	0.61	-0.17	***	3.59
- Wales (Great Britain)											-2.79	0.73	-0.12	***	1.97
Habsburg											5.24	0.58	0.24	***	5.60
Scandinavia											-2.08	0.68	-0.18	**	10.47
West											7.86	0.65	0.36	***	7.88
Interaction effects: see below															
N	264					264					264				
Adj. R squared	0.10					0.42					0.92				
Moran's I (with p-value)															
- Residuals of base model	0.89***					0.89***					0.89***				
- Residuals	0.81***					0.81***					0.03**				

continued…

	β	s.e.	std. β	p.-v.	VIF	β	s.e.	std. β	p.-v.	VIF	β	s.e.	std. β	p.-v.	VIF
Interaction effects															
Southeast (Balkans) before 1800											-7.77	1.35	-0.10	***	1.82
Southeast (Balkans) 1800-1850											-6.30	0.92	-0.16	***	3.33
East (East) before 1800											-5.51	1.24	-0.15	***	12.22
East 1800-1850											-2.73	1.20	-0.07	*	8.16
Habsburg before 1800											-3.65	1.21	-0.05	**	2.79
Scandinavia before 1800											-4.59	0.87	-0.26	***	6.51
Scandinavia 1800-1850											2.34	0.75	0.12	**	4.13
West 1800-1850											-7.57	0.80	-0.29	***	8.53

Significant at p<0.1; * p<0.05; ** p<0.01; *** p<0.001.
β: coefficient estimate; s.e.: standard error; std. β: standardized coefficient estimate; p.-v.: p-value; VIF: variance inflation factor

Online Appendix 1: Construction of GIS-based covariates

In deriving the population potential variable, we used global population count raster data derived from the History Database of the Global Environment (HYDE), Version 3.2. These data are available in 10-year intervals from 1700-2000. We chose the data for 1800. It is important to note that these data are estimates. We cut the file to ensure that we were only considering populations living in areas located between a longitude of 60° west and 60° east, and a latitude of 20° and 80° north. We then re-projected the raster data to a Lambert Azimuthal Equal Area projection. The population potential measure gives population situated nearby more weight than population further apart (see Stewart and Warntz 1959). Thus, locations surrounded by areas with high population numbers have a higher population potential compared to locations surrounded by sparsely populated areas. We derived this measure using the stewart-command in the R-library SpatialPosition with the following specifications: span=100000; b=2; typefct= exponential. As the location for which we performed the calculation, we used for the Mosaic dataset the coordinates for the 1,692 Mosaic locations from which we derived the data for our 115 Mosaic regions. We calculated the population potential for each location, and obtained from the outcomes a population-weighted value for the 115 Mosaic regions. For the NAPP regional data we used as coordinate the location of the raster point within the NAPP region that had the highest population density in 1800, according to the HYDE database. This was motivated by the fact that the population potential measure is very sensitive to the population in the immediate surroundings. Thus, we decided not to take the geographical or population-weighted centroids of the NAPP regions, which might have been situated in subareas of the regions that were sparsely populated.

The data on terrain ruggedness was obtained from the GTOPO30 elevation raster dataset, which is a global digital elevation model with a horizontal grid spacing of 30 arc seconds. To derive the information on terrain ruggedness, we used the Terrain Ruggedness Index (TRI) (Wilson et al. 2007). We did so by applying the focal function in the R-library raster (the TRI-formula is provided in the help function of 'terrain' in the raster library). For the Mosaic locations, we obtained the information for our set of 1,692 locations by considering the raster data within a circle with a diameter of 7.5 km centred on the location coordinates. Based on these data, we derived the population-weighted values for our 115 Mosaic regions. For the NAPP regions we had to use another approach, as we did not have location information for all of the settlements within a NAPP region. Here we faced the challenge that especially NAPP regions in Scandinavia were characterised by vast subareas with low population density. In order to avoid that our regional TRI values are dominated by information for such sparsely populated areas, we decided to only consider those areas of a NAPP region that had a population density above five people per km² in 1800 (determined through a mask operation using the HYDE 3.2 population density raster dataset).

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