Title Page

Historical family systems and lasting developmental trajectories in Europe: the power of the family?

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Authors full names:

Mikołaj Szoltysek (corresponding author), Institute of History, University of Warsaw, Krakowskie Przedmieście 26/28, 00-927 Warsaw, Poland, Ph.: 0048660717650; mszoltis@gmail.com; ORCID identifier: 0000-0003-4814-5586

Radosław Poniat, University of Białystok, Plac Niezależnego Zrzeszenia Studentów 1, 15-420 Białystok, Poland

Conflict of Interest

The authors declare that they have no conflict of interest.

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Historical family systems and lasting developmental trajectories in Europe: the power of the family?

ABSTRACT

Last years have witnessed a growing interest in economics and cross-cultural studies in the role of the historical family as the instigator of disparate developmental trajectories. This new emerging literature has already provoked a considerable amount of controversy, involving debates on the precise underlying mechanisms, the role of non-familial institutions and the possibility of reversed causality, as well as the quality of historical data. Using novel historical database of European family this paper reaffirms the hypothesis that historical family organization could be one of the intermediate factors associated with developmental and value disparities among European nations today pointed out in earlier scholarship. We show that countries starting out from more patriarchal family structures in the *past* exhibit more hierarchical gender relations, more collectivist mindsets, and lower levels of economic and human development *in the present*. These findings suggest that the criticism of the family role in comparative development may be premature, and that links between historical family organisation and developmental gradients merit further attention.

1. Introduction

Over the past few years historical family organization has become the object of extensive research in economics and cross-cultural studies as part of a broader agenda on the impact of culture on economic performance and comparative development (e.g., Greif, 2006; Duranton et al., 2009; De Moor & Van Zanden, 2010; Foreman-Peck, 2011; Carmichael et al., 2016a; Rijpma & Carmichael, 2016; Carmichael & Rijpma, 2017; Dennison & Ogilvie, 2014, 2016; Dilli, 2017; Le Bris, 2016; Szołtysek et al., 2017a; Santos Silva et al., 2017; Bertocchi & Bozzano, 2015; also Alexander & Welzel, 2015). This new emerging literature has already provoked a considerable amount of controversy, involving debates on the precise underlying mechanisms, the role of non-familial institutions and the possibility of reversed causality (Dennison & Ogilvie 2014, 2016; Carmichael et al., 2015; Carmichael et al., 2016b; Baten et al., 2016; Bertocchi & Bozzano 2016; Edwards & Ogilvie, 2018; Szołtysek et al., 2017a). A further recurring problem encountered in all previous studies was finding historical

demographic data that would be at the same time detailed and global to empirically anchor the authors' arguments (Szołtysek & Poniak 2018, 3-4). Recognising these challenges, Carmichael and co-authors (2016b, 200) suggested that the criticism of the family role in comparative development may well be premature, and should be re-examined using newly available historical demographic databases.

In line with this argument, most recently Szołtysek & Poniak (2018) argued that conclusions about the relationship between historical family organization and various societal outcomes may be sensitive to the data and measurement practices adhered to, and that that sensitivity has not received the attention it warrants (p. 3). As a way out of that conundrum, they suggested considering a broader range of indicators of historical family patterns than had been used thus far, in particular drawing on a recent explosion of historical census microdata. If results acquired using different measures of the focal predictor are complimentary to the wisdom of the earlier literature, then the generalisability of the family-development nexus can be strengthened.

The major contribution of this paper resides in expanding the on-going debate by testing the relationship between historical family patterns and contemporary gender inequality, value orientations, economic growth and human development using an alternative measure of family systems derived from historical microdata infrastructures recently developed through the Mosaic and the North Atlantic Population projects. Applying this novel composite measure of familial organisation in 26 European countries in the past (the Patriarchy Index; henceforth PI)¹, this paper reconfirms the hypothesis that historical family organization is associated with developmental and value disparities among European nations today pointed out in earlier scholarship. Based on cross-country correlations (bivariate) and multivariate regressions, we show that countries starting out from more patriarchal family

¹ The narrowing of our view to Europe is dictated by the scope of the datasets we employ. However, Emmanuel Todd's pioneering work (see below) has also started with Europe.

structures in the *past* exhibit more hierarchical gender relations, more collectivist mindsets, and lower levels of economic and human development *today*. Although we acknowledge that both family structures and developmental outcomes may have been influenced by some set of underlying factors which are not included in our analysis, a strong correlation between familial patterns today and in the past which we unravel, may suggest that the correlation is likely to run from persistent family patterns to developmental outcomes. By demonstrating that the major thread of the observations at the country level holds also when our analysis is pursued at a finer regional scale, we ascertain that the observed effect of historical family systems is not an artefact created by our aggregation procedures.

The remaining of this paper is divided into six parts. First, we review the state of research exploring how historical family patterns are linked to current developmental outcomes. We then present an alternative data framework and elucidate how this complements family system indicators used previously. In the core parts of the paper, we examine degrees of covariation between our historical country-level indicators of family organization (the PI) and selected contemporary indicators, using both correlational and multivariate framework. The next section discusses main causal paths that could account for the observed associations and provides evidence for the persistence of historical family patterns to the present. An assessment of the robustness of our results is then provided, including checks carried out at a lower aggregation scale (NUTS or ITAN level). We conclude by summarizing our findings.

2. Previous literature

The notion that family systems could have had an impact on wider societal outcomes has a long pedigree. For example, Max Weber alluded to it when he argued that strong family values do not allow for the development of individual forms of entrepreneurship, which are

fundamental to the formation of capitalist societies (Weber, 1904; also Banfield, 1958; Nimkoff, 1965, 61 ff). In the last few years, these ideas have regained prominence in the New Institutional Economics after scholars have re-discovered the importance of the family as the grassroot institution of society (see Alesina & Giuliano, 2010; earlier Sen, 1983; see also Mason, 2001; Carmichael et al., 2016a; Kok, 2017). For example, in a recent series of papers, Alesina and Giuliano (2010, 2014) have shown that the strength of family values across the world is negatively associated with a wide range of societal attitudes related to productivity and growth, as well as political participation, levels of trust and attitudes towards gender hierarchies (earlier, Bertrand & Schoar, 2006; also Kick et al., 2000; Daniele & Geys, 2016). Inspired by these insights, as well as by a growing recognition that human development can be affected by persistent historical traits (Nunn, 2009; Spolaore & Wacziarg, 2013), an increasing number of economic history works has incorporated past familial behaviour into explanations of developmental divergences within Europe and beyond (e.g. Greif, 2006; Duranton et al., 2009; De Moor & Van Zanden, 2010; Foreman-Peck, 2011; Dennison & Ogilvie, 2014, 2016; Bertocchi & Bozzano, 2015; Carmichael et al., 2016a; Rijpma & Carmichael, 2016; Carmichael & Rijpma, 2017; Dilli, 2017; Le Bris, 2016; Szołtysek et al., 2017a; Santos Silva et al., 2017; earlier also Reher, 1998; Therborn, 2004).

To date, an overwhelming majority of these attempts has relied on a world-wide classification of family systems inferred from anthropological evidence by Emmanuel Todd (1985, 1987)². Duranton and co-authors (2009) identified a significant association between Todd's classification of family types and regional disparities in educational attainment, social capital, labour participation, wealth, and inequality in Western Europe. Le Bris (2016) built the family score based on Todd's three family characteristics and found it to be significantly

² Todd himself argued that different patterns of family organization explain the diffusion of, or resistance to, a wide array of critical social changes on the continent, including the spread of Protestantism, secularism, and the acceptance and diffusion of communism.

(and robustly) associated with contemporary variation in economic outcomes as measured by GDP across 79 countries of the world. Building further upon this evidence, Carmichael and colleagues used a combination of Todd's classification with information compiled in George Murdock's *Ethnographic Atlas*³ to test the supposedly considerable influence the family patterns could have exerted on regional inequalities through their divergent impact on the status of women. In a series of analyses they found historical family constraints on female autonomy to be significantly and inversely associated with gender equality outcomes in contemporary world data (Rijpma & Carmichael, 2016, 38; Carmichael & Rijpma, 2017). They also established a strong and positive effect of "female friendly" family systems on the historical and contemporary estimates of per capita GDP for a range (25 to 52) of Eurasian countries. These analyses were complemented by Dilli (2017) who, relying on the same data, found that historical family institutions associated with higher female agency (such as equal inheritance practices, prevalence of nuclear household, late female marriage and absence of polygamy) were related to higher levels of economic development across 92 contemporary countries.

3. Materials and methods

3.1 Rationale for alternative data framework

Todd-like types of measures use dichotomous classifications to capture social norms or "cultural ideals" regarding family behaviours at the country level. For example, Carmichael's "Female Friendliness Index" strives to identify whether a country had historically preferred an extended household structure over a nuclear one, whether cousin marriage and patrilineal inheritance were practiced or not, whether polygamy was prevalent, and if postmarital residence was of a patrilocal or different type (Carmichael, 2014, 187–189; cf. Szołtysek & Poniat, 2018). However, considering that values, norms and beliefs may differ noticeably

³ Given the latter's meager coverage of Europe, the European part of their dataset was still based primarily on Todd.

from the grassroot practices (House et al., 2004; also Maseland & van Hoorn, 2009) and that they may evolve at their own pace, operationalizations of family systems based on the actual practices in family behaviours might present a viable alternative to look at when modelling family influences on wider societal outcomes.

In order to provide data framework for such an assessment of local family practices we drew on a massive repository of European historical census microdata recently made available through the North Atlantic Population (NAPP) and Mosaic projects⁴. All these databases were discussed in great detail elsewhere (Ruggles et al. 2011; Szołtysek & Gruber, 2016; Szołtysek et al., 2017b; Szołtysek & Poniati, 2018), and readers wishing to know more about them are encouraged to check the original sources.

Similar to Szołtysek and Poniati (2018), we use a combination of the two databases including 293 regional populations, with 15.3 million persons living in 3 million households, which outweighs all precedent infrastructure efforts in family history. Of the 293 regional datasets, a slight majority (56%) represents populations after 1850, while 44% cover populations before 1850, and 21% populations before 1800. The collection includes information on both rural and urban sites, although rural societies predominate (see Figure 1; also Supplementary material 1). Whereas Mosaic samples are of a varying level of representativeness (see Szołtysek and Gruber 2016), the NAPP components stand either for full-count census data or representative samples taken from them⁵.

Figure 1

⁴ See www.censusmosaic.org; <https://www.nappdata.org/napp/>. 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, as well as their sex, age, and marital status.

⁵ In collecting the NAPP data we gave preference to the oldest available censuses 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 England and Wales (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.

Table 1

Using a range of harmonized variables from the combined Mosaic and NAPP data, a composite measure of historical family organization known as the Patriarchy Index (see Szoltysek et al., 2017) was computed at the level of resolution of regions as presented in Figure 1⁶. Table 1 provides the complete list of the ten Index components, showing how they are defined and measured, and indicating the expected direction of their relationship with familial patriarchy levels (+/-) (for technical details and computation steps, see Gruber & Szoltysek, 2016).

3.2 Aggregation strategy

In order to be able to relate cross-country developmental measures to regional estimates of historical patriarchy, the latter were averaged for contemporary country borders and set at the average time when their relevant information was taken from, making allowance for multiple boundary changes wherever necessary. For the NAPP countries, we took an average PI value weighted by population size of their constituent regional groups. The same strategy was applied to a number of countries represented in Mosaic with several censuses of the same structure (i.e., either rural or urban)⁷. Where contrary was the case, the average PI value from various censuses was weighted by an index of urbanization characteristic of a given area at a respective time-period using data available from Malanima (2010). In this way, the weight of the overrepresented urban populations could be mitigated in favour of the rural areas. In cases

⁶ Cronbach's alpha for 10 components of the Index equals 0.87 (|0.85, 0.89|), suggesting that the items have a relatively high internal consistency. The alpha coefficient of reliability also suggests that there is no component variable that would decrease the Index's general reliability. Principal Component Analysis shows that the ten components of the Index load on two dimensions which altogether explain nearly 70 percent of the variance. This confirms that the notion of family systems we are trying to capture is not uni-dimensional (Wall, 1995). Because the construction of the Index was mainly theory-driven, the PCA results do not invalidate its composition.

⁷ The following discussion applies primarily to the Mosaic part of the dataset which is used for that paper. Since the NAPP regional data are derived from complete populations or 10% representative samples or full census counts, they are of no issue in this regard.

where our country values are derived from censuses of a different chronology, the above procedure was applied to each dataset separately, then followed by taking a grand mean (here, weights related to population size were no longer used). The only special case in this regard was that of Switzerland – for two periods the available populations were either rural or urban. Here, we still applied the urbanization weights (averaged for the periods in question) which enabled us to reduce the impact of a very low PI characterising the 19th-century Zurich. Finally, when it comes to countries for which there was only one census population available in Mosaic, the PI value derived from that population was taken to represent that whole country.

In the end, following those routines we were able to compute the historical PI values for 26 contemporary European countries. These values range from 12 to 29 index points, with Denmark scoring lowest on the patriarchy scale, and Albania situated at the top end of it (see Figure 2 and Table 2). The social prevalence of the PI may be taken to have variably influenced not only the cultural orientation of a given society in the past, but also largely the types of action which dominate thereby, including those prone to more rigid gender and age hierarchies, and enactment of loyalty to family, kin or lineage, filial piety, reverence for ancestors and obedience; and in-group assortative sociality.

Figure 2

Table 2

Although the outcomes of our aggregation strategy may be subject to some limitations due to the unequal coverage of contemporary countries with the Mosaic data, the dataset advanced here can be tentatively considered as accounting for cross-country variation in historical family patterns, i.e., as representing country-level generalizations derived from empirically assessed variation at the local/regional level. To this end it is perhaps worth

pointing out that except for Croatia, Bulgaria, Turkey and Spain (each represented by a single or highly clustered censal population; see Figure 1), the rest of our country means yield fairly reasonable representations of those countries' historical familial diversity at a certain moment in time (Szołtysek & Gruber, 2016)⁸. Controlling for time period in subsequent regression models also removes part of the evolution problem making it possible to account for variable effects of family systems captured at different times (see more in the robustness section).

3.3 Added value of the new measure

By combining a range of 10 variables related to familial behaviour rather than just focusing on marriage patterns or on the prevalence of nuclear families, the underlying structure of the PI moves us closer to a multi-stranded account of family organization in the past (cf. Mason 2001, 160–161). While Todd's scheme takes four (five in Carmichael's case) aspects and adds them up, implicitly assuming that each is equally important, the PI has the virtue of using more elements (10) with varying weights, thus reducing the chance that any strange component drives the index's variation. By choosing to use individual-level and age-specific measures instead of household-level variables we also ensure that our indicator of family structure is less sensitive to the latent variation in demographic conditions (Ruggles, 2012). Unlike Todd's scheme, the PI allows for making distinctions between various family types

⁸ A more qualitative assessment of the scattered data from Russia, customarily a troublesome case, also leads to safeguarding the reliability of our measures which were derived by pooling together data (altogether for 85 villages/settlements with 46,296 persons) from nearly 200 years of Russian history (1710-1897), and from such diverse contexts as the newly settled areas in the Urals, as well as established rural communities in the Central Black Earth Region; rural communities in sparsely populated rural regions with harsh ecological conditions; and those located in more environmentally friendly areas in the vicinity of urban centers. The five French data points in Figure 1 in fact refer to data for 25 villages in 25 départements of the French census of 1846. 11 villages are part of the collection of villages chosen by Louis Henry for the reconstruction of the population of France from 1670 to 1829 based on family reconstitution (all of them in Northern France). 14 villages are an extension to regions either not covered by this collection or a replacement for villages where no census data was preserved for 1846, or of too bad quality, or not available. They are the first villages beginning with the letter "B" in the respective départements. Altogether, the spatial distribution of these data captures the major distinction in family organisation between the north and south/southwest of the country (Flandrin, 1979). The share of rural regions in the French sample (79 percent) is also very close to Malanima's own estimations of the French urbanization rate from more or less the same period (Malanima, 2010, 246).

once account has been taken of gender relations through an explicit consideration of male domination among its diagnostic criteria. By capturing the inner architecture of generational and gender relations at the domestic level, it presents itself superior for the identification of the channels that – through their bearing on individual agency – may have affected economic behaviour and value formations. In particular, the PI has a virtue of capturing several elements of familial behaviour not accounted for in the measures used so far, such as living as non-kin or life-cycle servants, age at marriage, headship and seniority patterns, as well as female position in domestic organization. Not only have these long been used by historical demographers among the major measurement yardsticks to distinguish between different family systems (e.g., Wall, 1995), they also have a potential to capture important associations between the familial realm, on the one hand, and the economic performance and value orientation, on the other (de Moor & van Zanden, 2010; Foreman-Peck, 2011).

Furthermore, the continuous rating scale implemented in the PI provides a more sensitive metric for the assessment of the intensity of familial behaviour in a given population than fixed categories or binary measures. Finally, the scrutiny of historical record merged into our database allows to make the composition of our dataset fully transparent by providing explicit information about the sample size and the number of its constituent populations, the period of observation, and the share of urban–rural population in the data which came to constitute each country sample (Table 2).

In order to test the added value of the new measure we compared its predictive validity against an indicator of family organization derived from Todd’s typology and previously adhered to by economic historians (Carmichael’s “Female Friendliness Index”)⁹, by looking at how the focal variables associate with widely used measures of contemporary gender equality, value orientations, and economic output. Their response variables included: 1) two

⁹ While some of the FFI constituent measures correspond to the components of the PI, they were derived in a different manner by utilizing the information from Todd, and extrapolating it to entire countries.

contemporary gender inequality measures – Gender Inequality Index (GII) and the “Family” component of the Social Institutions and Gender Index (SIGI) (Van Staveren, 2013)¹⁰; 2) two country-level value indices – the Emancipatory Value Index (henceforth EVI), the most recent and complex representative of the “modernization-emancipation family” of value indices (Welzel, 2013; Sokolov, 2018, 3)¹¹; along with Hofstede’s *Individualism*, one of the dimensions of undoubtedly the most widely used national culture framework (Kirkman et al., 2006)¹²; finally, 3) the two most popular measures of development: contemporary *per capita* GDP (logged) and the Human Development Index (HDI)¹³.

Figure 3

Based on a series of unconditional correlations (with bootstrapped 95% confidence intervals; Figure 3), we found the association between historical “female-friendliness” and current outcomes to be generally weaker and less systematic compared to when contemporary country-level traits are predicted with the PI. For most of the variables, the relationship

¹⁰ Gender Inequality Index measures gender inequalities in: reproductive health (maternal mortality ratio and adolescent birth rates); empowerment (proportion of parliamentary seats occupied by females and the proportion of adult females and males aged 25+ years with at least some secondary education); and economic status (labour market participation of the female and the male populations aged 15+ years); see <http://hdr.undp.org/en/content/gender-inequality-index-gii>. Higher values indicate more gender inequality. The “Family” component of the Social Institutions and Gender Index measures a country’s “Discriminatory Family code” with regards to marriage, parental authority, and inheritance rights in 2014 (OECD Development Centre, Social Institutions and Gender Index 2014. Synthesis Report, 61–64; <https://www.genderindex.org/countries/?region=europe-and-central-asia>). We focus on this component because it yields global estimates for the largest number of countries from our collection, as well as for reasons of this component’s substantial affinity to our PI.

¹¹ The index represents a reformulation of Inglehart’s seminal measure of “survival versus self-expression values”. It is the 12-item index introduced by Welzel (2013, 57–104) based on responses taken from the World Values Surveys (WVS), measuring the belief in freedom of choice and equality of opportunities over four domains, each of which summarizes three items: (1) equity values: an orientation that prioritizes gender equality over patriarchy; (2) liberty values: an orientation that prioritizes reproductive freedoms over their restriction; (3) autonomy values: an orientation that prioritizes self-determination over obedience; (4) expression values: an orientation that prioritizes voice over order.

¹² It measures (at the national level) the strength of a preference for a loosely-knit social framework in which individuals are expected to take care only of themselves and their immediate families (its opposite being collectivism). See Hofstede et al., 2010; also <https://geert-hofstede.com>.

¹³ As is well known, the GDP and HDI –though partly overlapping – do not yield the same information and the correlation between them may well be far from absolute (Islam, 1995; Mukherjee & Chakraborty, 2010; cf. Bolt et al., 2014).

between the FFI and contemporary outcomes is either less predictive, more often only marginally significant, or the 95% confidence intervals are usually considerably wider than those calculated for the PI. The fact that respective correlations for the PI are much stronger and generally more reliable suggests that it is somewhat better suited for the purpose of predicting cross-country differences in present-day outcomes than its counterpart (see more in Szołtysek & Poniati, 2018, table 4, p. 20).

4. Empirical analysis

4.1 Hypotheses and stylized facts

Our general hypothesis is that centuries-old patterns of familial patriarchy—either directly, through their survival over time, or indirectly, through their internalization in values, customs, and culture—are strongly associated with current regional disparities across Europe in the areas we considered, i.e. in gender hierarchies, value orientations and developmental gradients. Before running more complex statistical models, we made some more specific predictions about the effects of historical family forms on the selected societal outcomes and proposed three *a priori* hypotheses.

Hypothesis 1. *Countries that had elevated PI levels in the past tend to preserve more rigid gender hierarchies in the present, whereas those with low scores on the PI scale are more likely to be associated with more egalitarian and gender-balanced outcomes*

[**Alternative:** *There is no, or a reverse correlation between past elevated PI levels and today's egalitarian and gender-balance characteristics*]

It has been increasingly recognized that while underdevelopment may enhance gender and other forms of inequality, societies hold certain cultural views, historically embedded, that may cast a long shadow on beliefs and values regarding the appropriate role of women

(Duflo, 2012; Jayachandran 2015; Alesina et al., 2013; Hansen et al., 2015; Olsson & Paik, 2016; also Bisin & Verdier, 2015). Deeply rooted patriarchal values from the last three hundred years are likely to persist, and hence become the long-lived sources of male and female identity. Because patriarchal family structures are characterised by customs and attitudes that collectively serve to maximise fertility (e.g., Dyson & Moore, 1983), they create incentives for women to remain subordinate in exchange for receiving support in raising their children, thus decreasing the former's bargaining power (Galor & Klemp, 2014). Patriarchal pre-adult socialisation practices and parental concerns about exposing daughters to outside influences (female virginity) further inhibit women from exercising agency outside of the domestic sphere, thus limiting their access to education, employment, and training (Caldwell, 1981; De Baca et al., 2014; Grogan, 2007). Finally, in many patriarchal societies, family honour depends on the sexual purity of women, and girls tend to marry at very young ages (Gruber & Szoltysek 2016; Caldwell, 1981, 10-11; Cain, 1988).

Hypothesis 2. *Countries characterized historically by strong patriarchal traits tend to preserve a more collectivist mindset and exhibit a weaker social prevalence of emancipatory values in the present*

[**Alternative:** *There is no, or a reverse correlation between past elevated PI levels and today's prevalence of emancipatory values*]

Olsson and Paik (2016) provided links between hierarchical value systems emerging after the Neolithic Revolution and current beliefs. By the same token, specific patterns of family organization may lead to a development of lasting individual psychological and behavioural dispositions affecting societies through bottom-up aggregation. Family loyalty and interdependence characteristic of patriarchal societies are generally considered the defining components of the cross-cultural dimension of collectivism-individualism, with collectivism

positively associated with strong family ties in multiple cross-cultural studies (e.g., Vandello & Cohen, 1999; Gelfand et al., 2004; Fincher & Thornhill, 2012; also Alesina & Giuliano, 2010). Evidence from traditional family- and kinship-based agrarian economies has shown that the underlying patriarchal authority structures are often consciously safeguarded by child-rearing practices that oppose fostering or even allowing competitiveness and individual initiative in children during their upbringing (Caldwell, 1981, 15). Because of its strong emphasis on loyalty to family, lineage, and kin, a patriarchal family structure restricts interactions between the family and the public sphere, and discourages family members from forming cooperative relationships with non-relatives. By this, it limits potentially stimulating “peer group effects” on development of more open value systems, instead rewarding conformity and normative behaviour and reduced trust in strangers, which may be fairly immune to top-down policy interventions (Ermisch & Gambetta, 2010; Alexander & Welzel, 2015).

Hypothesis 3. *Historically more patriarchal countries will more likely yield weaker indicators of economic growth and general human development in the present*

[Alternative: *There is no, or a reverse correlation between past elevated PI levels and today’s economic growth and general human development*]

In the male- and adult-centred patriarchal *milieu* the mobility of offspring is strongly limited (Caldwell, 1982), human capital acquisition by younger and female family members is discouraged by familial pressures, and women tend to become “overspecialised” in reproductive, child-rearing, and domestic work. These hamper the acquisition of knowledge and skills through training or apprenticeship that would enable offspring (females in particular) to enter into labour and contractual relationships in the public sphere and to accumulate other forms of human capital and transmit it intergenerationally, thus stifling

economic development in the longer run (Galor & Klemp, 2014). Because patriarchal societies place a high premium on family loyalty and reverence for ancestors (i.e., a collectivist mindset), their members (especially females and the youngsters) are less prone to engage in the types of entrepreneurship, collaboration with non-kin, and risk-taking that are prerequisites for economic success (Whyte, 1996). By rewarding conformity and normative behaviour, and by reducing trust in strangers, patriarchal societies are also likely to be characterized by a high avoidance of uncertainty and generally a less individualist culture (Thornhill & Fincher, 2014; also Ermisch and Gambetta, 2010; Gorodnichenko & Roland, 2011; Mokyr, 2016; also Ang, 2015). Because patriarchal structures hinder the ability of women to provide information, education, and cognitive skills to their offspring (Kambhampati & Rajan, 2008; Grogan, 2007), their effects tend to accumulate over time. In contrast, constraints on the power of the patriarch or the parents would improve incentives and property rights of women (and young men) and therefore the quality of decision making at that level (Dilli, 2017), leading towards more egalitarian achievements in human development.

4.2 Bivariate correlations

Figure 3 (above) reports the Pearson's r with 95% CI (bootstrapped) for unconditional correlations between our variable of interest – the PI – and selected measures of contemporary outcomes following the approach set out in Szoltysek and Poniat (2018). Each of the correlations involving the PI (indicated with a black point) is based on 1,000 simulations, and their confidence intervals were obtained with the use of the bias-corrected and accelerated (BCa) method (Efron, 1987).

In the first place, we find strong positive and statistically significant relationships between historical patriarchy and gender patterns today. The good match is most perceptive

with GII ($r = .64$; 95% CI $[0.43, 0.78]$), implying that along with the increase in historical patriarchy levels the probability that a given present-day country will be characterized by more gender inequality also increases pretty much linearly. The relationship between the PI and the “Discriminatory Family Code” component of SIGI looks similar, though it is somewhat weakened.

The correlations with measures of value orientation yield even more suggestive results. In both cases a country’s historical patriarchy profile is strongly and inversely related to the strength of emancipatory and individualistic values today. High patriarchy in the past strongly increases the probability that a country will have a low score on the EVI. With the Pearson r of -0.73 and plausible values of the true correlation as expressed by a 95% confidence interval ranging from $-.55$ to -0.83 , the estimate is very satisfactory from the accuracy point of view. Accordingly, we found an inverse and equally strong association of the PI with one of Hofstede’s *Individualism*. Also here, there is a fair amount of certainty in the true magnitude of the effect.

The PI is also associated with the chosen measures of development (*per capita* GDP (logged) and the HDI). The observed negative correlations are in line with theoretical expectations and previous research (e.g., Bertrand & Schoar, 2006; Dilli, 2017); they both appear to be very general and downward sloping. In the case of both variables the observed relationship is very strong, as even the lower boundaries of 95% confidence intervals (above -0.70 point estimate) would represent a genuinely large (negative) effect.

4.3 Multivariate analysis

A lurking threat to the observations above is the possibility that the PI may be correlated with unobserved country characteristics resulting in biased estimates of the effect of patriarchy on current developmental outcomes. In order to determine whether the associations predicted by

historical family systems are not spuriously driven we examined them against a host of covariates. Our choice of controls was based on a respective hypothesis under investigation, whenever possible taking the same set of baseline control variables as used in previous research to ensure comparability. The chosen covariates were grouped into four major categories. *Historical data characteristics* consider the average timing of the patriarchy score for each country, assuming that the relationship between the historical family and contemporary developmental outcomes may be weaker the further back in time the PI estimates go¹⁴. *Historical development factors* account for main alternative channels through which historical persistence could impact contemporary outcomes. Backward GDP reconstructions (corresponding to the time of patriarchy score in our database)¹⁵, and the state antiquity (Bockstette et al., 2002) were hypothesized as strong predictors of a country's prospective social and economic development (de Pleijt & van Zanden, 2016; Borcan et al., 2018)¹⁶. We also included the time elapsed since the first transition to agriculture (*Neolithic Revolution*; Putterman & Weil, 2010) because it has been shown to be positively related to the levels of economic development in recent times (Hibbs & Olsson, 2004; Putterman, 2008; Olsson & Paik, 2016), as well to contemporary gender equality (Hansen et al., 2015).

By the same token, two *bio-geographic endowment factors* were considered as having impact on various attitudinal outcomes today. Firstly, we included the climatic configuration of a country called the “Cool Water” (CW-) condition. Captured by an indexed measure

¹⁴ For example, the Albanian patriarchy score refers to 1918, meaning that the time span between the reference period and the presence is “only” one hundred years; on the other hand, the Croatian PI levels are from 1674, hence nearly two and a half centuries earlier.

¹⁵ Historical GDP data are derived from the data file created by the Maddison Project. In the case of some countries (e.g. Belarus), with GDP data only from the 20th century we have used data describing their nearest neighbours or former rulers (like the Russian Empire, in this case). Data source: <https://www.rug.nl/ggdc/historicaldevelopment/maddison/releases/maddison-project-database-2018>.

¹⁶ The effects of some further potential confounders – such as the percent rural datasets in the sample and historical urbanization (the latter commonly used in the literature) – have been reduced already at the stage of computing the PI, and hence they were not considered in the model specification.

proposed by Welzel (2013; henceforth: CWI)¹⁷, it has been shown to predict differences in pre-industrial female marriage ages, which in turn explain differences in gender equality today (Santos Silva, 2017). Secondly, we considered the historical pathogens prevalence (Murray & Schaller, 2010) based on Fincher et al. (2008) and Cashdan and Steel (2013) who demonstrated that people in regions historically more exposed to infectious diseases tended to develop more collectivist norms.

Finally, we account for two great socio-cultural, political and institutional transformations by including dummies for the *Protestant* heritage of the country and its experience of *Communism*. Following Becker and Woessmann (2009), as well as de Pleijt and van Zanden (2016), we hypothesize that Protestantism may have a significant effect on economic growth indirectly via its effect on human capital formation. We also presumed that a Protestant heritage would be an important objective factor favouring strong emphasis on secular-rational and self-expression values in the present, and a circumstance improving the status of women (Inglehart & Norris, 2003). In turn, institutional and cultural traits produced by communism are expected to continue retarding growth and human development today (e.g., Sookias et al., 2018). Furthermore, given that the *Communism* variable has spatial distribution corresponding neatly to the distribution of the countries' PI values, it also presents a major challenge to the explanatory power of our focal predictor.

A natural framework in which to test our hypotheses is a regression where a selected contemporary trait is the dependent variable, as in the form:

$$y_c = \alpha + \beta P_{cT} + T_c + X_{cI_4} + u_c$$

¹⁷ The indicator measures the prevalence of relatively cool temperatures in each season combined with abundant fresh water resources throughout the year, on a country's historically most populated area (see Welzel, 2013).

where y_c denotes respective current developmental or value measure at the country level; P_c is this country's level of historical family patriarchy; T is the historical period that varies depending on the country; X_c represents a vector of a country's other controls (historical-developmental; biogeographic; social transformations), and u_c is the error term.

In each model the PI serves as the main explanatory variable. However, since the PI estimates are time-dependent, all of our specifications also control for the period of patriarchy “observation” in each country allowing it to differ accordingly¹⁸.

In all subsequent models the selection of additional controls was decided depending on their theoretical relevance to the hypothesis under test. In each case, the presumption was that a given set of covariates could affect respective response variables through channels other than the one we are interested in identifying. Given the small N , the number of regressors in a single model was restricted to the most relevant ones and they were entered one at a time to avoid over-fitting. Crucially, in all regressions the bootstrap approximation was applied to the distribution of the least squares estimates and standard errors (involving 1,000 random samplings with replacement for each model; the bias-corrected and accelerated (BCa) method was used to obtain confidence intervals)¹⁹.

Table 3

Table 3 presents the results of OLS estimations for the relationship between the PI and gender inequality measures. It shows that historical family data are fairly good predictors of contemporary cross-country differences in inequalities between the sexes. In the model specification with PI and *Time period*, a one-standard deviation increase of the PI value is associated with an increase of GII by 0.61 standard deviation. This is a substantively

¹⁸ To some extent this also helps mitigate potentially confounding effects of the unobserved historical characteristics and/or processes (economic change, institutional and legal reforms, etc.).

¹⁹ We have used procedures implemented in the library *boot* of *R* language.

important effect equal to a change by 95 index points, equivalent to a difference in gender inequality levels between the Netherlands and Croatia. Controlling for the timing of the Neolithic Revolution or *Protestantism* (columns 2 and 3) – both of which enter the regression as expected – does not eliminate the role of the PI as the most significant covariate in the models. Its increase by one-standard deviation still results in an increase by 0.53 and 0.45 standard deviation in GII, respectively – i.e., in magnitudes only a little smaller than those of the basic model. Each of these three models produces R-squared of about .40, thus accounting for a non-negligible part of the variation in contemporary gender inequalities. Historical patriarchy stops being statistically significant only after the CWI is included as a control variable (column 4), which is not surprising given that both variables are strongly correlated ($r=-0.708$, significant at the 0.01 level). It stands to reason to argue that patterns of household formation and marriage, by being an important component of our index, provide a plausible, if not major, transmission channel through which biogeographic endowments have impacted contemporary gender relations (cf. Santos Silva et al., 2017).

The PI also appears to be a passable predictor of contemporary gender inequalities with regards to legal standing in marriage, parental authority, and inheritance rights as captured by SIGI's Discriminatory Family Code variable, whereby higher levels of historical patriarchy are associated with a more deterred position of women in the present, holding all else constant. Both in the basic specification with *Time period* (column 5) as well as in the model conditioning on the timing of the Neolithic Revolution (column 6), the effect of the PI is substantial and fairly stable. The latter model also suggests (in line with Hansen et al. 2015) that societies with longer histories of agriculture may have less equality in gender roles in the present, and that this may happen in part independently from their historical patriarchy levels. By accounting for the effect of *Protestantism* in column (7) we obtain results whereby historically high patriarchy levels still appear to be positively associated with contemporary

gender inequality, but the precision of this model becomes more uncertain. Meanwhile, controlling for CWI (column 8) does not yield any significant associations.

In turn, the inspection of Table 4 shows a strong and statistically significant association between the PI and emancipatory values (EVI) in nearly all five model specifications. In each model, higher patriarchy levels correspond to a weaker prevalence of emancipatory values in a country holding all else constant. To obtain an impression of the economic significance, we note that the result of our basic model (with *Time period* only) implies that a one-standard deviation increase in the PI is associated with a decrease in the prevalence of emancipatory values by nearly 0.20 index points in EVI, amounting to a shift in the individualism scale from the levels of Austria to those of Albania. The importance of the PI decreases somewhat after the effects of *Protestantism* are controlled for (column 3), with Protestant countries scoring higher on the emancipatory value scale, as expected. Even in this case, however, one standard deviation increase in the PI leads to a decrease in EVI by approximately 0.11 points. Though this is still a statistically significant result, it may contain both substantively meaningful as well as negligible effects ([0.64, -0.03]). Contrary to expectations, neither the CWI nor historical pathogen loads enter the emancipatory values regression with significant coefficients (columns 4 and 5), while patriarchy holds still. This is consistent with the view that the PI could act as a transmission channel for those variables. In the last model for EVI (column 5), the effect size of patriarchy becomes particularly strong: the estimated coefficient suggests an economically large relationship with the spread of emancipatory values that is about six times the effect of historical pathogen prevalence.

Table 4

Similar results are achieved when one of Hofstede's "dimensions of culture" is chosen as our response variable (columns 6-10 in Table 4). Also here, historically higher patriarchy

levels are invariably associated with lower contemporary individualism scores, but the models' statistical and substantive significance and their explanatory power are greater compared to regressions with EVI. For example, in four out of five specifications, an increase in PI by one-standard deviation leads to a decrease in individualism by over 0.8 standard deviation, i.e., by 20 index points – roughly the difference between Poland and the Netherlands. Moreover, the effect of historical patriarchy is fairly robust to the inclusion of other covariates, of which only the CWI enters the model with a significantly positive and strong coefficient (as expected; see Welzel, 2013).

Table 5

As shown in Table 5, historical PI also tends to be significantly and negatively associated with contemporary (2016) GDP *per capita* in all seven model specifications. In all models, standardized coefficients of PI are generally above 0.6, and in two cases close to 0.8, indicating an economically important effect of historical family systems. For example, in column (1), the PI's increase by a one-standard deviation results in a difference in real GDP values by nearly 100% (equivalent to the difference between Slovakia and Spain). The relationship between patriarchy and economic output would tend to be economically substantial even if the true correlation fell near the lower boundaries of the 95% confidence intervals (point estimate of -0.70). Noteworthy, the PI turns out to be a better predictor of contemporary cross-country differences in economic performance than Maddison's estimates of historical GDP (column 2), or the time elapsed since the Neolithic Revolution (column 3), and is also robust to controlling for the effects of *CWI*, *Protestantism*, or *State Antiquity* (columns 4, 5, and 7, respectively). The effects of the PI are also not wiped out after accounting for cross-country differences in the experience of Communism (column 6), which

exert strong and independent impact on contemporary GDP levels²⁰. A one-standard deviation increase in PI in this model is still associated with a decrease of predicted real GDP by 0.43 standard deviation, thus implying a shift in a country's economic output by roughly 40 percent.

Table 6

Similar results are obtained when the same covariates, including the PI, are regressed on the values of the HDI from 2015 (Table 6). Given that the GDP index in the HDI is based on GDP *per capita*, this can hardly be surprising. Interestingly, however, the HDI's greater accountability of non-economic elements of development results in revealing a stronger and more robust relationship with PI. Same as in the models with GDP, the effect of historical patriarchy is invariably negative and statistically significant, even after controlling for selected covariates. Of the latter, *CWI*, *Protestantism* and the experience of Communism enter regressions with significant coefficients and expected direction, but only the former surpasses the effect size of the PI. In this worst scenario, where the standardized Beta coefficient for the PI falls below 0.5, a one standard deviation increase in historical patriarchy still represents an economically substantial change in the value of HDI (by 62 index points).

5. Discussion

In general, our results appear to confirm that historical family patterns in Europe have a significant and strong association with current regional disparities in gender inequality, value orientations, and economic and human development. Four main causal paths could account for these associations. First, family structures may persist to the present and continue to affect

²⁰ Partial *r*-squared for PI in this model equals to 0.26, and for Communism to 0.34.

contemporary outcomes directly. This form of persistence seems intuitively likely given that family behaviour and values represent “deep” cultural layers which are transmitted from generation to generation and move slowly over time (Bisin & Verdier, 2000). Studies in family history have substantiated this observation by revealing a strong continuity of demographic behaviour, at least in some areas of Europe (Reher, 1998; Therborn, 2004; Wall, 2001; Dalla Zuanna & Micheli, 2004).

Second, the persistence may develop through intermediate factors, such as the nature of political or economic institutions that have been shaped first by family structures and have continued to influence contemporary societies in a path-dependant manner. For example, in the 1980s, after Laslett showed that the nuclear family structure had been the dominant family type in England long before the Industrial Revolution, some scholars (including Laslett himself) argued that the dominance of the nuclear family was among the necessary preconditions for modernisation and industrialisation, which could then spill over further developmental outcomes irrespective of subsequent family change.

Third, reverse causation is also plausible in that the differences in economic and institutional factors across countries might have themselves caused the variation in family structures. Greif (2006), for example, pointed out a potential feedback mechanism where on the one hand, nuclear families facilitate the development of corporate institutions, but on the other the socioeconomic transformations related to that development encourage the formation of nuclear families across Europe (also Dennison & Ogilvie, 2014, 672-685).

Finally, another possibility is that the links between family structures and developmental gradients are an outcome of a deeper underlying determinant. In this case, family structures would be endogenous to the causal process and of little explanatory value.

There is nothing about our own results that makes causality automatic from past to present outcomes, and our analyses are not legitimate to rule out entirely the possibility of

bidirectional or circular causation, or to assert that the links between family structures and developmental gradients are not an outcome of a deeper underlying determinant²¹. Although we cannot address the endogeneity question explicitly, an important step towards elucidation of the temporal order of the relationships between historical family and our developmental measures can be offered by examining the extent to which contemporary familial behaviours are related to historical family structures²².

For the purpose of this exercise we focus on contemporary accounts of actual (experienced or current) practices related to family life and cohesion²³. Our first measure of contemporary family was Welzel's measure of "contemporary patriarchy" based on a series of responses from the World Values Surveys which combined scores for percent females who declared being married before the age 20, and for the percent married men (aged 30-34 years) who declared living with parents. In addition, we used the percent 30-39 years old living with parents, based on data from *European Values Study 2008*. We assume that earlier marriage may enhance more patriarchal conjugal relations (Gruber & Szołtysek, 2016; Therborn, 2004), and that the family ties might be considered stronger and more aligned with a complex hierarchy of authority patterns based on age, if the individual lives with his or her parents (Reher, 1998; Todd, 1985).

²¹ Note, however, that our identification strategy allowed us to capture potential influences on our response variables other than those of historical patriarchy, including religion and historical development. In addition, we offered a number of theoretically informed guesses for why familial institutions might have played an exogenous role by suggesting that patriarchal structures from the last three hundred years may be prone to strong persistence (and even tend to accumulate over time), and hence remain fairly immune to top-down policy interventions (Alexander & Welzel, 2015).

²² Neither Todd's own work, nor papers by Greif, De Moor and Van Zanden, and Duranton and co-authors, provided any empirical tests of these potential trajectories. While Duranton only conjectured the long-term persistence of medieval family types (Duranton et al., 2009), Rijpma and Carmichael observed that while their measures of historical family had some predictive power for today's measures of family characteristics, they were far from perfect (Rijpma & Carmichael, 2016, 35).

²³ We have chosen to focus on practices rather than on subjective measures of the strength of family ties (such as those of Alesina & Giuliano, 2010 or Thornhill & Fincher, 2014, 131) because values scores tend to be noticeably different from practices scores across many recent cross-cultural studies (e.g., the GLOBE project; see House et al. 2004). For example, using data for 15 Mosaic/NAPP countries covered by the GLOBE project we established that *Gender Egalitarianism* and *In-group Collectivism* tended to be valued more than they were practiced (at the organizational level). Furthermore, for both variables, a clear linear positive relationship between Societal Values and Societal Practices could not be established (own calculations based on http://globeproject.com/study_2004_2007#data; available on request).

As a preliminary test of the degree of covariation between those measures, unconditional correlations between historical and contemporary data were plotted in Figures 4A-B (again, bootstrapping methods and 95% confidence intervals were used). The two correlations suggest that our data contain a strong signal. For both the *Contemporary Patriarchy* and the EVS data the relationship is fairly linear and has no particular outliers. In both cases, countries with higher patriarchal values in the past display higher frequencies of intergenerational co-residence in the present (and earlier marriage for women). The Pearson r of 0.80 or above, and plausible values of the true correlation as expressed by a 95% confidence interval ranging from 0.61 to 0.89 for the former variable, and from 0.74 and 0.94 for the latter one, imply very substantial effects.

Figures 4A-B

The associations from Figures 4A-B were found to be robust to controlling for a host of potentially confounding factors in the multivariate framework (Tables 7-8). Out of sixteen regression specifications for the two response variables of interest, the PI turns out to be significantly associated in fifteen models, suggesting that countries with heightened historical patriarchy quite invariably tend to display stronger intergenerational bonding, and hence stronger family ties, today. Furthermore, in most specifications the size effects of the PI are large or even very large. The only exception is the model for *Contemporary Patriarchy* indicating that the negative and significant relationship between past and present family patterns is not robust to including the current log GDP estimates (column 7 in Table 7). Potential concerns over this issue are, however, partly allayed by the fact that similar qualification does not apply to regression with our second measure of contemporary behaviour (see below). The association between historical and current patriarchy manifests itself most strongly in the model controlling for time period and historical pathogens (ibid.,

column 2), in which an increase by one-standard deviation in PI corresponds to increase in the *Contemporary Patriarchy* scores by 0.90 standard deviation, roughly equal to the difference between Switzerland and Albania. The PI's impact on current family practices remains genuinely robust to the inclusion of *Protestantism* (again, negatively correlated with the outcome variable), CWI, and the experience of Communism²⁴.

Tables 7-8

The regressions with the EVS data as a dependent variable provide further evidence which shows that higher patriarchy is related to contemporary tighter intergenerational bonds (Table 8). The results remain significant and strong across all specifications, even after controlling for the effect of Communism and the “cool-water” effect (both significant). Most noteworthy, the result is also robust to the inclusion of the contemporary log GDP, despite its otherwise strong effect on the outcome variable (an increase by one-standard deviation in GDP results in the decrease in the rates of contemporary co-residence by 0.56 standard deviation, equal to 5.3 index points). Same as in Table 7 above, historical pathogens, the experience of Protestantism, membership in the Soviet Union, and the variation in historical GDP do not matter, while the PI holds prominently in these models. For example, the results from column 2 purport to saying that a 4.6 point difference on the patriarchy scale could make a difference of about 9 percent points in contemporary co-residence, like, for example, between Austria (10.6%) and Lithuania (19.2%)²⁵.

²⁴ Interestingly, earlier membership in the Soviet Union does not enter the regression with significant sign, which might suggest that the specific Soviet-type of Communism had more impact on normative systems rather than on actual behaviour.

²⁵ Leverage tests revealed that these results are not subject to significant change with the removal of outlying cases (data available on request). Similar exercises performed for response variables measuring the normative aspect of family ties (Alesina/Giuliano's and Thornhill/Fincher's measures of the strength of family ties) established that the relationship between the PI and both indexes is positive and linear, although it tends to be wiped out by controlling for other covariates. The important exceptions were models indicating a strong positive linear relationship of historical patriarchy with contemporary family ties controlling for a country's experience

6. Robustness checks

Because the number of observations that are used for all the above-presented regressions is relatively small, we used the procedure for identifying the influence of particular observations on our model estimates. We then re-estimated all model specifications on subpopulations of Mosaic/NAPP countries omitting some potentially problematic observations. The identification of potentially “problematic” cases was carried out in two different ways. First, we have excluded four countries for which Mosaic data were the least representative (Bulgaria, Croatia, Spain and Turkey), on the grounds that their PI estimates may be biased and hence create spurious correlations in the models. Our exclusion criteria were concerned either with a country’s meagre representation in our database (e.g., only one very early population corresponding to modern day Croatia; exclusive focus on Istanbul as regards Turkey); or on the presumption that a country’s collection of data was too regionally clustered (two Bulgarian populations based on numerous small censuses from different periods; Spain represented by two populations only from Catalonia)²⁶. All the models we have presented above were then re-run on the basis of such a curtailed sample.

Additionally, we have run residuals and leverage tests based on the Cook’s distances derived from the full (based on the full countries sample) models. Highly influential cases observed separately for each model were then removed from the sample one at a time, each time refitting the regression model on the remaining countries. Since Cook’s D doesn’t have a one-standard and common accepted cut-off point, we have decided to use the formula $D_i > 4/n$ and discard values above 0.15. The new models were rerun only for those instances when at

of belonging to the USSR (of an equally strong but adverse effect). In this case, the PI’s increase by one-standard deviation led to increase in the strength of family ties by over 0.62 standard deviation.

²⁶ As mentioned before, the small number of mapped points in France in fact hides a much wider territorial coverage (25 departments) and hence does not present an additional case for exclusion in this regard.

least one observation was above that point, and in due course we found that the number of discarded observations was never greater than 3.

The results of these tests showed that even after removing particular observations our findings remained stable, both in terms of estimation and inference. The new models obtained by curtailing the dataset were very similar to those based on the full country sample (see Appendix 1, Tables 1A-1F). The PI effects, their direction and sizes are in almost every case identical with those from the full models. Important differences were identified only in the models with SIGI_Family as the dependent variable, and PI, *Time period* and *Protestantism* as the predictors (columns 5 and 6, Table 1B in Appendix 1), which turned out to be decidedly weak. Given that even the whole model based on the 25 countries instead of 21 and 23 was relatively weak and the PI influence was significant only at the 90% level, this finding does not seem to undermine our core findings.

6.1 The aggregation problem

Another potential threat to the robustness of our findings is that the observed effect of historical family systems might be artefactually created by our aggregation procedures. To mitigate this concern, the main body of our analyses has been replicated at the smaller geographical scale. In order to do so, the NAPP/Mosaic regional populations (see Figure 1) have been assigned to contemporary administrative units of the European Union (NUTS 2, 2016) or to their equivalents in the neighbouring countries (ITAN), and the PI estimates obtained with this procedure were linked to available developmental statistics²⁷.

Embarking on this approach faced a number of challenges, the most severe of which were that (1) part of our chosen developmental indicators was available only at the country

²⁷ Because one such unit usually encompassed 1-2 historical populations, weighting by an index of urbanization was omitted in this spatial join. In cases where a NAPP/Mosaic region spread over several contemporary statistical areas (especially in case of the NAPP data), it was assigned to that NUTS-2 unit where its centroid weighted by the population size was located.

level; (2) many existing survey data do not provide a division into NUTS; (3) the available data do not cover all our areas of interest; and (4) regional statistics were often computed using inconsistent methodologies and spatial divisions, especially in the non-EU countries. In consequence, although we were able to compute the historical PI scores for 124 NUTS/ITAN regions, in hardly any regression could all these observations be used conjointly. Accordingly, in our analyses we were able to consider only some of the dependent variables used previously (GDP, HDI, the percent 30-39 years old living with parents)²⁸.

In the face of the lack of lower-scale gender inequality measures, regional differences between female and male life expectancy at birth in 2016 (*de0*; Eurostat) were computed and used as a proxy for gender gaps in one particular dimension of human development achievements. Secondly, as a more attitudinal measure, we took the question C001 from the EVS data from 2008 asking whether respondents *strongly agreed*, *agreed*, *disagreed*, or *strongly disagreed* with the statement that “men should be preferred when jobs are scarce”.

Severe limitations applied also to the usage of the control variables, the number of which had to be kept small. Besides our focal predictor, in each model we control for the share of rural populations in the regional PI sample and for the average year of the patriarchy “observation” in each NUTS/ITAN unit. In addition, relying on theoretical considerations and findings of the main analysis (see Section 4.3), in each model we also accounted for whether the particular administrative unit had any experience of the communist regime. Same as in the main analysis, in all regressions the bootstrap approximation was applied to the distribution of the least squares estimates and standard errors (involving 1,000 random samplings with replacement for each model). Table 1G in Appendix 1 presents the results of the regressions.

²⁸ Data on regional GDP and the information used to compute *de0* are from Eurostat (<http://ec.europa.eu/eurostat/web/regions/data/database>). Regional HDI was obtained from Hardeman and Dijkstra (2014). The variable RHDI represents the region’s HDI rank among all EU regions. Data on question C001 and on the share of Dane on c001 and on the share of 30-39 years old living with parents, are from EVS 2008.

In 4 out of 5 models, the PI has turned out to have significant and theoretically plausible impact on the response variables. A historically higher PI is associated with a bigger female-to-male difference in life expectancy, lower GDP in 2012, and a lower rank in the ranking of the EU-RHDI, and these relationships are not explained away by accounting for the experience of communism. A historically higher PI is also associated with a stronger propensity among the EVS respondents to favour men on the labour market, overriding the effect of communism which ceased to be a significant predictor in that model. Although PI significantly predicts the percent 30-39 years old living with parents after controlling for basic historical characteristics (rural and time period; $\beta=.63$, $p\text{-value} < 0.001$), this effect was not robust to the inclusion of the experience of communism variable²⁹.

Altogether, the models in Table 1G largely corroborate the results at the country level obtained above and suggest that the effects of historical family systems observed in the main analyses were not an artefact created by our aggregation procedures.

7. Conclusions

Overall, the results of this paper add to the evidence that the principles of family organization at the grassroots of society could be one of the intermediate variables accounting for contemporary cultural and developmental divergences within Europe. An alternative measure of historical family systems (the PI) generates results similar to previous dispersed findings using different methodologies, and therefore reconfirms the importance of the historical family as the potential instigator of disparate and lasting developmental trajectories on the continent. In an effort toward understanding what drives these correlations, we have made informed suggestions that family patterns are very persistent. Accordingly, we have tested

²⁹ We have also run a series of separate regression models on the same data but including basic geographical controls (longitude/latitude; available on request). Although in basic models latitude and especially longitude were usually highly significant, with standardized coefficients often above 0.5, in most cases their importance (and statistical significance) were wiped out by the effect of Communism.

these presumptions empirically showing that familial behaviour today and family patterns in the past are, indeed, correlated, and that this persistence is not likely to be a sole by-product of differences in current economic development.

Our contribution also demonstrates potential gains of family and economic history working together to expand the existing data frameworks towards better understanding of the family-development nexus. Before we pause with putting effort in exploring what place the family occupied in the ‘horse race’ between various deep causes of human development (Dennison & Ogilvie, 2014), it might be worth considering what is to be gained from taking into account a broader range of indicators than has been used thus far, in particular by drawing on a recent explosion of historical-demographic data. To this end, our findings illustrate potential benefits of further efforts to mobilize historical data on family practices rather than crude measures of social norms or “cultural ideals” regarding family behaviours used before. This study is also the first to test the relationship between historical family patterns and contemporary gender inequality, value orientations, economic growth and human development within a unified empirical framework. Increased research efforts along each of these lines may help to strengthen the generalisability of the family-development nexus both in Europe, and beyond.

However, a number of limitations of this study must be clearly acknowledged. By necessity, our analyses were limited to 26 European countries. Given the small N, the number of regressors in a single model must have also been restricted, so the obtained coefficients cannot be very precise. Our sample size – which still represents more than a half of current European countries - can be placed at the bottom end of the sample size distributions that we have encountered in the output growth analyses, such as those of Levine and Zervos (1998; the N-range from 32 to 45 countries) or Knack and Keefer (1997; 29 countries). More recently, Santos Silva et al. (2017) presented regressions based on as small a number of

countries as 25 (p.59); and Olsson and Paik (2013) regressions based on 21 to 31 observations (see tables 3 and 5-6). It should also be noted that the main body of our analyses has been replicated with larger number of observations at the smaller geographical scale.

Regarding the data, we are aware that there are important areas of Europe not yet covered by the Mosaic/NAPP database (e.g., Italy), and that those covered sometimes are represented by spatially sparse data points. However, several attempts to remedy this problem have been attempted in the course of analysis showing that the general thrust of our results remains unaltered after removing the most susceptible observations and pursuing the lower-scale analyses. A major way to move forward in the future would be to create higher density “samples” for some countries and to further expand the regional analysis.

Finally, evidencing causality based on regressions with 26 observations and 7-8 explanatory variables is difficult unless one has an exogenous source of variation. Because there is a possibility that both family structures and development patterns are an outcome of a deeper, underlying influence, our regression estimates should be read as indicating association, rather than “causal influence” of family organization on numeracy. Although at this stage we cannot address the endogeneity question explicitly, we nevertheless provided a number of theoretical and empirical considerations in support of the view that such causal influences might indeed exist. A clear identification of the direction of these relationships presents a veritable task for the future.

Overall, future research should strive to test if the main insight of this paper remains intact by employing our measure of historical family to a wider assembly of settings (be those countries or their subdivisions) and taking into account other correlates (including spatial autocorrelation) and the potential for reverse causation. Should the results of those future endeavors indeed prove to suggest that historical family patterns cast a very long shadow on recent developmental outcomes, then certain recommendations can be made for policy-

makers to show greater sensitivity to historical path dependencies in the familial and demographic realm, and to bear in mind the specific characteristics of family systems when designing social policies. Bringing together policy-making with the type of work conducted by economic and family historians might be particularly important for the societies which don't have the growth enhancing family institutions historically.

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Figures

Figure 1: Spatial distribution of the combined Mosaic/NAPP data

Notes: for sources of the Mosaic and NAPP data, see Supplementary material 2.

Figure 2: Cross-country variation in historical patriarchy (combined Mosaic/NAPP data)

Notes: for sources of the Mosaic and NAPP data, see Supplementary material 2. Grey colour indicates no data.

Figure 3: Comparisons of the Pearson's r with 95% CI (bootstrapped) for unconditional correlations between indicators of family organization (PI vs FFI) and selected measures of gender inequality, value orientations, economic growth and human development

Source: for PI – Mosaic/NAPP database; for FFI – data courtesy of S. Carmichael.

Figures 4a-b: The unconditional correlations between contemporary family ties, patriarchy and intergenerational coresidence, and historical patriarchy levels (PI)

Notes: sources as in Tables 7-8.

Table 1: Components of the Patriarchy Index

Domain/ component	Component	Abbreviation	Definition/measurement	Relationship with patriarchy	Specification
Male domination	Proportion of female household heads	<i>Female heads</i>	the proportion of all female household heads among all adult (20+ years) household heads of family households	negative	age-standardized
	Proportion of young brides	<i>Young brides</i>	the proportion of ever-married women in the age group 15-19 years	positive	
	Proportion of wives who are older than their husbands	<i>Older wives</i>	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	<i>Female non-kin</i>	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	<i>Younger household head</i>	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	<i>Neolocal</i>	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	<i>Lateral</i>	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	<i>Married daughter</i>	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	<i>Boy as last child</i>	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	<i>Sex ratio</i>	the sex ratio (boys to 100 girls) in the youngest age group (0-4 years old).	positive	Only family households

Table 2: Basic characteristics of Mosaic/NAPP country samples

Country	PI	IQR	No. of regions	No. of settlements (parishes/villages)*	Pop.	% rural	Time (average)	Time (range)
Albania	29,32	4,0	14	291	140,611	58	1918	–
Austria	14,9	5,25	4	53	20,036	74	1910	–
Belarus	25,1	–	2	90(p)/340	44,508	100	1793	1768–1804
Belgium	14	–	1	5 (p)	13,666	100	1814	–
Bulgaria	24,68	–	2	8(p)	8,373	78	1908	1877–1947
Croatia	22	–	1	24	1,880	100	1674	–
Denmark	12,2	1,0	19	n/a	838,623	94	1787	–
France	16,37	3,0	5	33	27,745	79	1845	1831–1901
Germany	14,24	3,25	44	690	384,049	62	1824	1700–1900
Hungary	19	2,25	4	20 (p)	13,724	100	1869	–
Iceland	14	–	1	n/a	51,003	100	1703	–
Lithuania	22,5	–	2	480	19,917	100	1847	–
Netherlands	13,6	1,0	5	33	40,037	53	1811	1810–1811
Norway	16,3	1,0	21	n/a	878,073	95	1801	–
Poland	16,66	3,0	10	88(p)/440	83,276	65	1791	1666–1809
Romania	20,1	1,25	8	71	14,514	100	1844	1781–1879
Russia	22,59	2,75	11	85	46,296	92	1768	1710–1897
Slovakia	20,5	–	3	21	8,410	100	1869	–
Spain	20,9	–	3	13 (p)	23,997	29	1889	1880–1890
Sweden	12,9	2,25	24	n/a	4,624,825	92	1880	–
Ukraine	22,16	2,0	11	231	90,022	81	1815	1765–1897
U.K.	11,6	1,0	83	n/a	7,859,626	57	1881	–
Serbia	24,73	–	3	18	19,180	86	1870	1863–1884
Turkey	20	–	2	1	8,354	0	1896	1885–1907
Switzerland	17,28	–	3	18 (p)	16,140	68	1787	1671–1870
Latvia	19,1	2,25	4	67 (p)	35,807	100	1797	–

Source: Mosaic and NAPP.

* Numbers indicated with „p” stand for „parishes/or estates”; otherwise refer to settlements/villages.

Table 3: Cross-country variation in gender inequality

	GII	GII	GII	GII	SIGI Family	SIGI Family	SIGI Family	SIGI Family
Historical PI	0.61*** (0.14)	0.53*** (0.17)	0.45* (0.19)	0.13 (0.18)	0.53*** (0.16)	0.42** (0.15)	0.46^ (0.26)	0.25 (0.29)
Time Period	0.16 (0.14)	0.04 (0.27)	0.12 (0.15)	-0.03 (0.14)	0.16 (0.20)	0.04 (0.21)	0.15 (0.20)	0.05 (0.22)
Neolithic Revolution		0.38^ (0.22)				0.61** (0.22)		
Protestantism			-0.25 (0.19)				-0.09 (0.26)	
Cool Water effect				-0.71** (0.24)				-0.41 (0.32)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	26	24	26	26	25	23	25	25
Adj. R- squared	0.39	0.40	0.40	0.61	0.29	0.56	0.27	0.34
F-statistic	9.03**	6.19**	6.61**	13.92***	5.99**	10.42**	4.05*	5.27**
AIC	65.70	61.57	66.06	55.10	67.06	52.22	68.51	65.88

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses.

In columns 1-4, the dependent variable is Gender Inequality Index from 2015; in columns 5-8 - the “Discriminatory Family Code” component of the Social Institutions and Gender Index (SIGI) as of 2014.

^ p-value <0.1

* p-value <0.05 ** p-value <0.01

*** p-value <0.001

Table 4: Cross-country variation in value orientations

	EVI	EVI	EVI	EVI	EVI	Ind.	Ind.	Ind.	Ind.	Ind.
Historical PI	-0.73*** (0.12)	-0.79*** (0.17)	-0.38** (0.16)	-0.46* (0.23)	-0.88*** (0.14)	-0.87*** (0.10)	-0.88*** (0.13)	-0.83*** (14)	-0.53*** (0.14)	-0.88*** (0.14)
Time Period	0.16 (0.16)	0.31 (0.30)	0.25 (0.13)	0.20 (0.16)	0.14 (0.23)	0.10 (0.11)	0.22 (0.16)	0.11 (0.10)	0.24* (0.11)	0.14 (0.23)
Neolithic Revolution		-0.14 (0.26)					-0.23 (0.22)			
Protestant.			0.59** (0.19)					0.08 (0.14)		
Cool Water effect				0.39 (0.28)					0.53** (0.17)	
Pathogen load					0.14 (0.33)					0.14 (0.34)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	24	22	24	24	22	25	23	25	25	23
Adj. R- squared	0.51	0.48	0.62	0.52	0.50	0.67	0.78	0.67	0.80	0.64
F-statistic	12.87***	7.49**	13.45***	9.61***	8.15**	24.88***	27.02***	19.92***	32.56***	14.22***
AIC	55.88	55.27	50.57	55.66	54.01	48.37	37.99	49.21	36.62	43.79

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses.

In columns 1-5, the dependent variable is Emancipatory Value Index (time-pooled cross section country aggregates from WVS Waves 1 to 6); in columns 6-10 – one of Hofstede’s “dimensions of culture”

(*Individualism*), measuring the degree of interdependence a society maintains among its members (time-pooled cross section country aggregates from Hofstede et al., 2010; also <https://geert-hofstede.com>).

^ p-value <0.1

* p-value <0.05 ** p-value <0.01

*** p-value <0.001

Table 5: Cross-country variation in economic performance

	GDP (ln)	GDP (ln)	GDP (ln)	GDP (ln)	GDP (ln)	GDP (ln)	GDP (ln)
Historical PI	-0.83*** (0.10)	-0.68*** (0.19)	-0.79*** (0.11)	-0.60*** (0.12)	-0.69*** (0.13)	-0.43** (0.13)	-0.65*** (0.13)
Time period	0.01 (0.09)	-0.15 (0.18)	0.09 (0.22)	0.08 (0.10)	0.02 (0.10)	-0.10 (0.08)	-0.08 (0.11)
Hist. GDP (ln)		0.23 (0.30)					
Neolithic Revolution			-0.10 (0.12)				
Cool Water effect				0.38^ (0.18)			
Protestantism					0.22 (0.15)		
Communism						-0.51** (0.16)	
State Antiquity							0.31^ (0.15)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	26	25	24	26	26	26	26
Adj. R-squared	0.68	0.68	0.64	0.72	0.70	0.78	0.73
F-statistic	27.76***	18.29***	14.88***	22.91***	20.26***	30.92***	24.28***
AIC	48.84	48.50	48.60	45.29	48.31	39.82	44.14

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses. In all columns, the dependent variable is per capita gross domestic product of 2016 (transformed by taking the natural logarithm; source: the World Bank).

^ p-value <0.1

* p-value <0.05 ** p-value <0.01

*** p-value <0.001

Table 6: Cross-country variation in human development

	HDI	HDI	HDI	HDI	HDI	HDI	HDI
Historical PI	-0.82*** (0.10)	-0.71*** (0.17)	-0.77*** (0.12)	-0.45*** (0.09)	-0.65*** (0.12)	-0.50** (0.18)	-0.71*** (0.14)
Time period	-0.07 (0.10)	-0.11 (0.16)	0.05 (0.22)	0.08 (0.09)	-0.03 (0.10)	-0.14^ (0.08)	-0.11 (0.10)
Hist. GDP (ln)		0.14 (0.25)					
Neolithic Revolution			-0.23 (0.17)				
Cool Water effect				0.56*** (0.14)			
Protestantism					0.26^ (0.15)		
Communism						-0.39^ (0.21)	
State Antiquity							0.18 (0.17)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	26	25	24	26	26	26	26
Adj. R-squared	0.67	0.70	0.66	0.80	0.70	0.72	0.68
F-statistic	26.38***	19.66***	15.87***	35.03***	20.26***	22.79***	18.59***
AIC	49.77	45.26	48.32	37.16	48.31	46.03	49.93

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses. In all columns, the dependent variable is Human Development Index of 2015 (source: United Nations Development Program (2016). *Human development report 2016*. New York, 198-201).

^ p-value <0.1

* p-value <0.05

** p-value <0.01

*** p-value <0.001

Table 7: Cross-country variation in contemporary patriarchy

	1	2	3	4	5	6	7	8
Historical PI	0.83*** (0.13)	0.90*** (0.17)	0.57*** (0.16)	0.30* (0.14)	0.78*** (0.17)	0.50** (0.19)	0.18 (0.18)	0.74** (0.23)
Time Period	-0.03 (0.13)	0.00 (0.14)	-0.08 (0.14)	0.06 (0.09)	-0.01 (0.15)	-0.16 (0.13)	0.02 (0.08)	0.05 (0.25)
Pathogen load		-0.13 (0.28)						
Protestantism			-0.46** (0.16)					
Communism				0.68*** (0.15)				
Postsoviet					0.13 (0.17)			
Cool Water effect						-0.59* (0.24)		
GDP ln							-0.75*** (0.19)	
Historical GDP ln								-0.10 (0.42)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	21	21	21	21	21	21	21	20
Adj. R-squared	0.61	0.59	0.63	0.65	0.60	0.75	0.72	0.65
F-statistic	16.51***	10.77***	12.38***	13.52***	10.99***	20.74***	18.22***	12.66***
AIC	44.69	46.21	44.25	42.96	45.92	36.25	38.35	40.35

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses.

The dependent variable is Welzel's measure of contemporary patriarchy (time-pooled cross section country aggregates from WVS Waves 1-6; data: courtesy of C. Welzel).

^ p-value <0.1

* p-value <0.05 ** p-value <0.01

*** p-value <0.001

Table 8: Cross-country variation in contemporary residential proximity

	1	2	3	4	5	6	7	8
Historical PI	0.98*** (0.10)	0.95*** (0.15)	0.74*** (0.21)	0.49* (0.20)	1.08*** (0.14)	0.67*** (0.16)	0.50** (0.19)	0.90*** (0.20)
Time Period	0.05 (0.10)	0.01 (0.11)	0.00 (0.13)	0.10 (0.09)	0.03 (0.11)	-0.03 (0.11)	0.06 (0.10)	0.14 (0.20)
Pathogen load		0.10 (0.25)						
Protestantism			-0.30 (0.19)					
Communism				0.56** (0.20)				
Postsoviet					-0.15 (0.15)			
Cool Water effect						-0.42* (0.21)		
GDP ln							-0.56* (0.23)	
Historical GDP ln								-0.13 (0.27)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	23	22	23	23	23	23	23	22
Adj. R-squared	0.76	0.74	0.78	0.83	0.76	0.79	0.85	0.75
F-statistic	36.46***	21.27***	26.80***	36.13***	24.37***	28.35***	41.55***	22.58***
AIC	36.92	37.78	36.19	30.48	37.94	35.13	27.72	37.78

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses. The dependent variable is percent 30-39 years old living with parents, based on EVS (2016): *European Values Study 2008: Integrated Dataset (EVS 2008)*. GESIS Data Archive, Cologne. ZA4800 Data file Version 4.0.0, [doi:10.4232/1.12458](https://doi.org/10.4232/1.12458).

^ p-value <0.1

* p-value <0.05 ** p-value <0.01

*** p-value <0.001

Appendix 1

Table 1A: Cross-country variation in gender inequality (robustness tests with selected countries omitted)

	1	2	3	4	5	6
Historical PI	0.67*** (0.16)	0.63** (0.21)	0.63** (0.20)	0.51* (0.21)	0.16 (0.27)	0.17 (0.16)
Time Period	0.11 (0.15)	-0.07 (0.31)	-0.18 (0.31)	0.05 (0.15)	0.03 (0.16)	-0.11 (0.14)
Neolithic Revolution		0.33 (0.27)	0.31 (0.27)			
Protestantism				-0.27 (0.19)		
Cool Water effect					-0.65* (0.30)	-0.75*** (0.22)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes
N	22	20	22	22	22	25
Adj. R-squared	0.44	0.41	0.38	0.46	0.58	0.66
F-statistic	9.28**	5.50**	5.28**	6.90**	10.56***	16.76***
AIC	54.41	52.05	57.93	54.57	49.07	49.37
Omitted countries	Bulgaria, Croatia, Spain, Turkey	Bulgaria, Croatia, Spain, Turkey	Croatia, Turkey	Bulgaria, Croatia, Spain, Turkey	Bulgaria, Croatia, Spain, Turkey	Great Britain

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses. The dependent variable is Gender Inequality Index from 2015. Model specifications in columns (3) and (6) based on Cook's distances.

^ p-value <0.1

* p-value <0.05 ** p-value <0.01

*** p-value <0.001

Table 1B: Cross-country variation in gender inequality (robustness tests with selected countries omitted)

	1	2	3	4	5	6	7	8
Historical PI	0.47* (0.21)	0.45** (0.16)	0.42* (0.21)	0.32* (0.15)	0.47 (0.38)	0.23 (0.34)	0.28 (0.42)	0.04 (0.39)
Time Period	0.18 (0.20)	0.30^ (0.17)	0.19 (0.33)	0.16 (0.28)	0.18 (0.23)	0.25 (0.22)	0.15 (0.24)	0.13 (0.22)
Neolithic Revolution			0.48^ (0.27)	0.56* (0.27)				
Protestantism					0.01 (0.35)	-0.13 (0.36)		
Cool Water effect							-0.23 (0.40)	-0.44 (0.40)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	21	24	19	21	21	23	21	22
Adj. R-squared	0.22	0.34	0.38	0.46	0.18	0.18	0.21	0.19
F-statistic	3.96*	7.09**	4.67*	6.84**	2.52	2.58	2.79	2.66
AIC	58.92	62.71	51.24	52.51	60.84	66.37	60.17	63.34
Omitted countries	Bulgaria, Croatia, Spain, Turkey	Croatia	Bulgaria, Croatia, Spain, Turkey	Albania, Croatia	Bulgaria, Croatia, Spain, Turkey	Albania, Croatia	Bulgaria, Croatia, Spain, Turkey	Albania, Croatia, Turkey

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses. The dependent variable is the "Discriminatory Family Code" component of the Social Institutions and Gender Index (SIGI) as of 2014. Model specifications in columns (2), (4), (6), and (8) based on Cook's distances.

^ p-value <0.1

* p-value <0.05 ** p-value <0.01

*** p-value <0.001

Table 1C: Cross-country variation in value orientation (robustness tests with selected countries omitted)

	1	2	3	4	5	6	7	8	9
Historical PI	-0.77*** (0.14)	-0.70*** (0.11)	-0.77*** (0.17)	-0.86*** (0.20)	-0.49** (0.17)	-0.51*** (0.14)	-0.48* (0.23)	-0.79*** (0.16)	-0.81*** (0.16)
Time Period	0.06 (0.15)	0.09 (0.16)	0.17 (0.25)	0.42 (0.40)	0.17 (0.11)	0.41*** (0.11)	0.10 (0.15)	0.15 (0.20)	0.01 (0.30)
Neolithic Revolution			-0.11 (0.23)	-0.20 (0.40)					
Protestantism					0.47* (0.21)	0.53*** (0.14)			
Cool Water effect							0.36 (0.23)		
Pathogen load								-0.11 (0.32)	0.11 (0.41)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	22	23	20	20	22	21	22	20	20
Adj. R-squared	0.53	0.46	0.49	0.41	0.64	0.81	0.55	0.54	0.47
F-statistic	12.72***	12.53***	7.15**	5.48**	13.72***	28.81***	9.70***	8.62**	6.60**
AIC	52.72	55.71	50.45	52.99	45.23	30.65	50.25	48.02	50.94
Omitted countries	Bulgaria, Croatia, Spain, Turkey	Sweden	Bulgaria, Croatia, Spain, Turkey	Great Britain, Sweden	Bulgaria, Croatia, Spain, Turkey	Great Britain, Latvia, Norway	Bulgaria, Croatia, Spain, Turkey	Bulgaria, Croatia, Spain, Turkey	Spain, Sweden

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses. The dependent variable is Emancipatory Value Index (time-pooled cross section country aggregates from WVS Waves 1-6). Model specifications in columns (2), (4), (6), and (9) based on Cook's distances.
^ p-value <0.1 * p-value <0.05 ** p-value <0.01 *** p-value <0.001

Table 1D: Cross-country variation in value orientation (robustness tests with selected countries omitted)

	1	2	3	4	5	6	7	8
Historical PI	-0.88*** (0.13)	-0.90*** (0.19)	-0.91*** (0.11)	-0.84*** (0.19)	-0.38* (0.18)	-0.43** (0.15)	-0.81*** (0.16)	-0.74*** (0.18)
Time Period	0.10 (0.14)	0.30 (0.29)	0.24 (0.20)	0.11 (0.14)	0.16 (0.13)	0.25 (0.11)	0.01 (0.31)	0.25 (0.18)
Neolithic Revolution		-0.21 (0.22)	-0.36* (0.16)					
Protestantism				0.06 (0.17)				
Cool Water effect					0.63*** (0.17)	0.60*** (0.16)		
Pathogen load							0.11 (0.41)	-0.22 (0.26)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	21	19	21	21	21	24	19	22
Adj. R-squared	0.62	0.73	0.88	0.62	0.79	0.81	0.56	0.69
F-statistic	17.66***	17.25***	51.96	11.74***	25.55***	33.16***	8.56*	16.48***
AIC	43.76	36.48	21.77	44.99	32.73	34.19	40.33	39.20
Omitted countries	Bulgaria, Croatia, Spain, Turkey	Bulgaria, Croatia, Spain, Turkey	Hungary, Sweden	Bulgaria, Croatia, Spain, Turkey	Bulgaria, Croatia, Spain, Turkey	Turkey	Bulgaria, Croatia, Spain, Turkey	Austria

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses. The dependent variable is of Hofstede's "Individualism" (time-pooled cross section country aggregates from Hofstede et al., 2010; also <https://geert-hofstede.com>). Model specifications in columns (3), (6), and (8) based on Cook's distances.
^ p-value <0.1 * p-value <0.05 ** p-value <0.01 *** p-value <0.001

Table 1E: Cross-country variation in economic output (robustness tests with selected countries omitted)

	1	2	3	4	5	6	7	8	9	10	11	12	13	14
Historical PI	-0.84*** (0.11)	-0.87*** (0.08)	-0.70** (0.22)	-0.69*** (0.20)	-0.82*** (0.15)	-0.83*** (0.15)	-0.50** (0.17)	-0.62*** (0.11)	-0.69*** (0.15)	-0.73*** (0.13)	-0.34* (0.15)	-0.38** (0.13)	-0.63*** (0.18)	-0.64*** (0.15)
Time Period	-0.02 (0.12)	-0.04 (0.10)	-0.13 (0.19)	-0.19 (0.20)	0.18 (0.17)	0.25 (0.25)	0.03 (0.12)	0.11 (0.10)	0.03 (0.12)	-0.01 (0.09)	-0.05 (0.07)	-0.10 (0.07)	-0.05 (0.11)	-0.16 (0.11)
Historical GDP (ln)			0.19 (0.29)	0.28 (0.25)										
Neolithic Revolution					-0.16 (0.16)	-0.18 (0.16)								
CWI							0.44* (0.21)	0.36** (0.14)						
Protestant.									0.25 (0.16)	0.21 (0.15)				
Communism											-0.59*** (0.16)	-0.61*** (0.13)		
State Antiquity													0.32^ (0.19)	0.31^ (0.16)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	22	25	22	23	20	23	22	23	22	25	22	24	22	24
Adj. R-squared	0.70	0.74	0.68	0.77	0.66	0.65	0.75	0.79	0.71	0.76	0.81	0.89	0.74	0.78
F-statistic	24.28***	35.85***	15.95***	26.32***	13.11***	14.70***	21.99***	28.94***	18.31***	26.24***	30.23***	63.05***	20.74***	28.55***
AIC	41.50	41.70	42.87	37.08	41.10	46.47	37.52	34.75	40.63	40.98	31.85	20.75	38.53	37.14
Omitted countries	Bulgaria, Croatia, Spain, Turkey	Ukraine	Bulgaria, Croatia, Spain, Turkey	Great Britain, Ukraine	Bulgaria, Croatia, Spain, Turkey	Croatia	Bulgaria, Croatia, Spain, Turkey	Great Britain, Turkey, Ukraine	Bulgaria, Croatia, Spain, Turkey	Ukraine	Bulgaria, Croatia, Spain, Turkey	Turkey, Ukraine	Bulgaria, Croatia, Spain, Turkey	Russia, Ukraine

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses. The dependent variable is per capita gross domestic product of 2016 (transformed by taking the natural logarithm; source: the World Bank). Model specifications in columns (2), (4), ..., and (14) based on Cook's distances.

^ p-value <0.1

* p-value <0.05 ** p-value <0.01

*** p-value <0.001

Table 1F: Cross-country variation in human development (robustness tests with selected countries omitted)

	1	2	3	4	5	6	7	8	9	10	11	12
Historical PI	-0.85*** (0.11)	-0.75*** (0.21)	-0.57*** (0.14)	-0.84*** (0.14)	-0.82*** (0.12)	-0.43** (0.14)	-0.46*** (0.08)	-0.68*** (0.14)	-0.36* (0.15)	-0.36*** (0.10)	-0.68*** (0.18)	-0.74*** (0.15)
Time Period	-0.06 (0.1)	-0.14 (0.17)	-0.25 (0.15)	0.09 (0.09)	0.06 (0.19)	0.00 (0.10)	0.13 (0.09)	0.00 (0.09)	0.08 (0.07)	-0.10 (0.06)	-0.09 (0.10)	-0.08 (0.11)
Historical GDP (ln)		0.13 (0.26)	0.39 (0.25)									
Neolithic Revolution				-0.16 (0.18)	-0.09 (0.15)							
CWI						0.52** (0.17)	0.59*** (0.12)					
Protestantism								0.27^ (0.16)				
Communism									-0.57*** (0.16)	-0.60*** (0.13)		
State Antiquity											0.24 (0.19)	0.19 (0.14)
Intercept	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	22	22	23	20	23	22	25	22	22	25	22	24
Adj. R-squared	0.71	0.70	0.77	0.68	0.67	0.81	0.82	0.75	0.82	0.84	0.74	0.76
F-statistic	27.42***	17.66***	25.80***	14.76***	16.04***	30.79***	37.41***	21.72***	33.95***	42.99***	20.65***	25.50***
AIC	39.55	41.22	35.85	40.05	45.82	31.51	33.74	37.74	29.70	30.78	38.60	39.32
Omitted countries	Bulgaria, Croatia, Spain, Turkey	Bulgaria, Croatia, Spain, Turkey	Great Britain, Norway	Bulgaria, Croatia, Spain, Turkey	Turkey	Bulgaria, Croatia, Spain, Turkey	Great Britain	Bulgaria, Croatia, Spain, Turkey	Bulgaria, Croatia, Spain, Turkey	Turkey	Bulgaria, Croatia, Spain, Turkey	Turkey, Ukraine

Notes: cross-country OLS estimates, bootstrapped standard errors (based on 1,000 replications) in parentheses. The dependent variable is Human Development Index of 2015 (**source:** United Nations Development Program (2016). *Human development report 2016*. New York, 198-201). Model specifications in columns (3), (5), (7), (10), and (12) based on Cook's distances.

^ p-value <0.1

* p-value <0.05 ** p-value <0.01

*** p-value <0.001

Table 1G: Cross-regional (NUTS/ITAN-level) variation in contemporary developmental outcomes

	de0	logGDP_2012	RHDI_2012	C001	30-39 with parents
PI	0,37*** (0,09)	-0,26* (0,13)	0,48*** (0,11)	-0,4* (0,18)	0,19 (0,13)
Rural	-0,05 (0,05)	-0,17* (0,07)	0,05 (0,1)	-0,01 (0,15)	0,13 (0,09)
Time period	0,01 (0,06)	-0,19** (0,07)	-0,01 (0,05)	-0,12 (0,13)	0,08 (0,07)
Communism	0,74*** (0,06)	-0,74*** (0,08)	0,61*** (0,1)	0 (0,14)	0,62*** (0,12)
N	105	102	95	80	98
Adj. R-squared	0,788	0,658	0,643	0,115	0,546

Notes: cross-regional OLS estimates for NUTS/ITAN administrative units; bootstrapped standard errors (based on 1,000 replications) in parentheses.

“de0” is the difference between female and male life expectancy at birth in 2016 (*source*: Eurostat); “C001” is question C001 from the EVS data from 2008 asking whether respondents strongly agreed, agreed, disagreed, or strongly disagreed with the statement that “men should be preferred when jobs are scarce”.

RHDI_2012 represents the region’s HDI rank among all EU regions (*source*: Hardeman and Dijkstra 2014)

* p-value <0.05 ** p-value <0.01 *** p-value <0.001

Supplementary material 1

Table 2A: Mosaic datafiles (as of December 2017).

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
Poland-Lithuania, 1791-1792 census	3 urban regions	28,975
Moldavia, 1781-1879 status animarum	2 rural regions	5,291
Wallachia, 1838 census	4 rural regions	21,546
Russia, 1710 enumeration	6 rural and 2 urban regions in the Ural area	27,736
Russia, 1765 census	4 rural regions in the Hetmanate area (Ukraine)	18,391
Russia, 1795 revision lists	1 rural region in Ukraine	8,050
Russia, 1797 revision lists	4 rural regions in Courland (Latvia)	35,807
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; 4 rural regions and 1 urban region around Berdychiv in Ukraine	34,441
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
Switzerland, 1671-1762 church listings	2 rural regions in canton Zürich	10,988
Switzerland, 1870 census	Zürich	8,152
Ottoman Empire, 1885 census	Istanbul	3,408
Ottoman Empire, 1907 census	Istanbul	4,946
Mosaic data overall	142 regions (109 rural and 33 urban)	1,085,136
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

Supplementary material 2

Table 3A: References to the data

Mosaic data:

Karl Kaser, Siegfried Gruber, Gentiana Kera, Enriketa Pandelejmoni. *1918 Census of Albania, Version 0.1* [SPSS file]. Graz, 2011.

Laboratory of Historical Demography (MPIDR). *1869 Census of Hungary, Version 1.0* [Mosaic Historical Microdata File]. www.censusmosaic.org, 2014.

Laboratory of Historical Demography (MPIDR). *1910 Census of Austria, Version 1.0* [Mosaic Historical Microdata File]. www.censusmosaic.org, 2014.

Familiekunde Vlaanderen and Laboratory of Historical Demography (MPIDR). *1814 Census of Western Flanders, Version 1.0* [Mosaic Historical Microdata File]. www.censusmosaic.org, 2014.

Ulf Brunnbauer. *Household registers of Rhodope region, Version 1.0* [Mosaic Historical Microdata File]. www.censusmosaic.org, 2014.

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