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Abstract

Variation in living standards across Europe, especially in income, has decreased over the last few decades, but the last recession brought convergence to a halt. This paper asks, first, whether sigma-convergence is found when other dimensions of inequality are taken into account, and second, whether the recent economic recovery led to renewed convergence. To assess sigma-convergence, I estimate transnational inequality in the euro area (EA-13) using a decomposable multidimensional inequality measure including income, occupational prestige, education, and employment status as key dimensions economic wellbeing and inequality. I quantify the contribution of factor shares to within- and between-group inequality across the euro area using a counterfactual decomposition method together with bootstrapped confidence intervals. The results show that, like income, multidimensional inequality increased significantly starting in 2008, mainly driven by income and employment status. Just two years later, in 2010, sigma-convergence started to decline, and in 2014 reached a level of divergence that had only been seen previously before the introduction of the euro. The income dimension best explains between-country divergence, but differences in employment status and the correlation between dimensions contributed substantially to within-country inequality. A formal club convergence test shows two of the European country clubs—Central Europe and Southern Europe—to be key drivers of divergence, with the exception of Spain as a potential outlier. The results show that the recent economic recovery in the euro area has brought about initial relief in multidimensional inequality, but that the level of transnational and between-country inequality as well as divergence remains high.

Keywords: multidimensional inequality, convergence, euro area, transnationalization, composite index

JEL classification: D3, F15, I30

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1 Introduction

Income inequality increased in most European countries after the onset of the European financial and economic crisis. At the same time, convergence of incomes across countries came to a halt. Before the recession, the so-called "convergence machine" of the European Union and the euro area (EA) succeeded in leveling cross-country differences, especially by increasing incomes and living standards in Southern and Eastern Europe (Goedemé and Collado, 2016). However, recent economic downturns as well as political events have shaken beliefs that the common currency alone can ensure ongoing economic and social convergence within the EA. The idea of gradual convergence appears to be contradicted by the experiences of Southern European countries, which have struggled with high unemployment, low economic growth, and drastically reduced public services. For the EA, convergence is not only a political goal, but also a necessary foundation for common macroeconomic policy within the monetary union.[1] Therefore, this paper asks whether the development of incomes was accompanied by similar changes in other dimensions of inequality, and whether multidimensional measures confirm the increasing divergence within the euro area.

To assess the level of inequality and trends in cross-country differences, looking at different indicators separately is not enough: individual living standards and inequalities are affected above all by correlations among different dimensions at the household level. Instead of using a dashboard approach, multidimensional inequality and convergence should be evaluated using a social welfare function that accounts for the relative importance and correlations of different dimensions (Stiglitz, 2009; Tsui, 1999).

The assumption that convergence takes place across multiple dimensions of economic wellbeing is closely related to the sociological concept of transnationality. According to this concept, Europeanization at the economic, political, and monetary level does not only lead to consolidated political and economic institutions (Heidenreich, 2016b, p. 30), which influence political decisions and change the distribution of national well-being, but also extends to the individual level by generating shared norms of equality and reference frames, which in turn determine perceptions of inequality (Poppitz, 2016), opportunities, and economic stress (Heidenreich, 2016c; Whelan and Maître, 2013).

As national inequality estimates preclude the existence of such extended reference frames, this paper estimates inequality by treating the EA as a transnational entity. To investigate transnational inequality and convergence within the EA empirically, I exploit the methodological link between the two. As inequality measures are nothing other than measures of variation, they have been used previously to describe σ-convergence, following Martin (1996). In line with this work, I use a multidimensional inequality index to assess transnational inequality. As I estimate transnational inequality using household data instead of national or regional

[1] A related argument for socioeconomic convergence is the political goal of social and political cohesion within Europe to prevent the recurrence of wars and other catastrophic historical events. However, the link between convergence and social cohesion is weak and has been the subject of numerous studies (Vergolini, 2011).

aggregates, I use a sub-group decomposition to investigate the degree of convergence by the contribution of between-country differences to overall inequality. Finally, to formally test for the existence of convergence clubs within the EA, I apply formal club convergence tests to the multidimensional inequality estimates.

By using an axiomatic welfare measure, I decompose sub-group inequality further into factor shares by constructing counterfactual distributions (Decancq et al., 2017). This allows me to evaluate the contribution of individual dimensions to overall convergence or divergence in the EA for the first time. I proceed in Section 2 by reviewing the relevant literature in each of these research strands and in Section 3 by selecting appropriate decomposition and weighting methods. Section 4 presents the data sources, and Section 5 discusses the results. Section 6 concludes and outlines possible directions for future research.

2 Literature Review

The formation of the European Union and especially the creation of the euro area have spurred research on transnationality and convergence for two reasons. First policy makers want to evaluate the effects of various policy initiatives such as the Lisbon treaties and the Horizon 2020 strategy. Second, the heterogeneous effects of the economic and financial crisis within the EA as well as popular movements against further Europeanization and the euro itself have challenged the idea that convergence and social cohesion are increasing continuously. To date, however, the convergence literature rooted in classical growth theory and the literature on transnationality and social cohesion have remained largely separate strands. This paper aims to bring together the most important theories and findings from both strands of research to identify the shortcomings of previous works.

Growth and convergence

According to the neoclassical growth model, countries eventually converge to the same level of economic wellbeing, conditional on a set of structural parameters. While the original growth model predicts a negative relationship between growth rates and initial income levels ($\beta-$convergence), the decrease in overall variation ($\sigma-$convergence) is a necessary condition (Young et al., 2008). Empirically, various works have documented strong $\beta-$convergence in the initial years of the EA and a subsequent halt since 1990 (Beckfield, 2009; Bouvet, 2010). As already mentioned, the process of economic equalization does not, of course, imply a simultaneous equalization of social and cultural identities between nations. The few empirical works to assess convergence within Europe including other dimensions than income have done so by analyzing each dimension separately (Otoiu and Titan, 2015; Sarracino and Mikucka, 2017). By design, this dashboard approach cannot account for the correlation between dimensions or assess the level of overall convergence. On the global level, the work of Jordá and Sarabia (2015) is a rare exception, as they evaluated convergence across income, education, and health based on the Human Development Index (HDI). The authors report overall $\sigma-$convergence on the world level to be driven by the education dimension, while income follows a twin-peak distribution. However, because of the simplicity of the HDI, their work ignores the sensitivity of the results with respect to normative decisions discussed in the axiomatic welfare measurement literature, in particular, aggregation order, substitution elasticity, weighting between dimensions (Greco et al., 2018), as well as heterogeneity within countries. Döpke et al. (2017) investigate to what extent the eligibility of European regions for convergence fund resources depends on the dimensions considered to measure divergence, while explicitly considering the impact of weighting dimensions. The authors emphasize the influence of weights, if the EU convergence policies would depend on a multidimensional inequality measure. However, a decomposition by dimensions and into within- and between-regional inequality is missing, because the analysis is based on aggregate regional data from the OECD. Despite the above mentioned shortcomings, the approach to measure $\sigma-$convergence using a subgroup decomposable inequality (Jordá and Sarabia, 2015) and the focus on heterogeneity across multiple dimensions within the EU (Döpke et al., 2017) constitute the starting point for this work and the link to the topic of transnationality.

Transnationality and convergence

Transnationality originates from the idea that relationships across borders emerge not only between states (internationalism), but also between individuals. In addition to economic and institutional integration, socio-economic spaces evolve and cultural identities can converge, for example, through migration and multinational citizenship (Berger and Weiß, 2008). Consequently, comparing national distributions is very different from comparing transnational inequality. Comparing national inequality estimates has its own merits, but refers to a status quo that is based on national entities. Transnational inequality, in contrast, refers to the distribution of achievement in important dimensions of economic wellbeing by individuals from different national entities, by acknowledging extended reference groups (Heidenreich, 2016a, p. 9) as well as the relevance of these groups for social policies (Atkinson, 1995, p. 71).

However, estimating transnational inequality introduces new conceptual problems that should be noted. First, transnational identities vary between individuals, and the extent to which they do so is positively correlated with individuals' socio-economic status (Mau and Mewes, 2008). Because of this correlation, assuming a unilateral degree of transnationality can bias inequality estimates. Second, the value of some outcomes such as educational titles depends on specific context in which they are evaluated, which can be local or transnational (Weiss, 2005) and thus, affect the level of measured inequality. Third, transnationality provides not only a perspective on inequality but potentially can be seen as an additional dimension of wellbeing and therefore socio-economic stratification, an aspect that is rarely taken into consideration in quantitative assessments of transnational inequality.

Early efforts to estimate transnational income or earnings inequality for the EU suffered from the lack of comparable household survey data.[2] In addition to the shortage of data, a number of methodological issues such as the need for a harmonized definition of available income, equivalization of household incomes, and adjusted price differences (Brandolini, 2007) have limited the validity of the results. With the availability of the EHCP and EU-SILC, things changed for the better, and the use of purchasing power parities and the new OECD household equivalization scale have now become standard practice.

Within the EU15, inequality of equivalized household incomes increased from 1996 to 2008, mainly driven by higher inequality at the bottom of the distribution, while top- and middle-income inequality stagnated (Papatheodorou and Pavlopoulos, 2014, p. 456). Based on similar methods, Heidenreich (2016c) showed that since 2008, income inequality increased again within the EU15 up to a Gini index of 0.3 in 2012. In the enlarged European Union (EU-27), income inequality is, of course, higher, but it declined in the same period from 0.354 to 0.338.[3] However, transnational income inequality in the EU-27 is still lower than in the US, with a Gini index of 0.382 (Heidenreich, 2016c, p. 29). By decomposing equivalized disposable household incomes into different income components, Vacas-Soriano and Fernández-Macías (2017) found increasing income inequality mainly driven by individuals being pushed out of the labor market, while labor income inequality remained stable. Whereas the aforementioned works discuss the level of economic integration in Europe and some refer to the idea of transnationality, they fail to include other dimensions than income when assessing convergence.

Similar to the literature on $\sigma-$convergence, multidimensional inequality estimates for transnational entities are rare. This is even more surprising since the development of multidimensional inequality measures based on social welfare functions has made significant process in recent years.[4] Based on social welfare functions, these multidimensional indices allow for consistent aggregation across different dimensions by explicitly including the normative decisions involved in the aggregation. Moreover, the methods available for decomposition into population subgroups and factor shares provide an analytical tool to analyze the interplay between different dimensions. So far, these methodological advances have only been used to measure subgroup inequality in emissions of four greenhouse gases at the global level. In this case, declining interregional inequality contributed to an overall decrease in emissions inequality, independent of normative parameters (Remuzgo and Sarabia, 2015; Remuzgo et al., 2016). Investigations of frequently discussed dimensions of inequality usually take national

[2]For a comparison of early transnational income inequality estimates, see Table 1 in Vacas-Soriano and Fernández-Macías (2017).

[3]Similar patterns are found by Boix (2004), Brandolini (2009), and Bönke and Schröder (2014).

[4]For extended surveys multidimensional inequality measures, see Aaberge and Brandolini (2015) and Chakravarty and Lugo (2016).

borders as given by analyzing cross-country differences, thus remaining within the realm of 'methodological nationalism' (Beck, 2008).

Finally, the present work contributes to the discussion of optimal policy and currency areas. In this literature to date, possible benefits of a joint policy or currency area resulting from economies of scale have been considered to be negatively related to the degree of cultural diversity (Alesina et al., 2017) as well as heterogeneity in economic development and idiosyncratic business cycles (Mundell, 1961). By estimating subgroup inequality, this work examines cross-country differences in relation to within-country inequalities in multiple dimensions of economic wellbeing. The smaller the contribution of a particular dimension to cross-country inequality, the less it can be expected to impede European integration. For dimensions that play a larger role in within-country inequality, the inequalities might be tackled more effectively at the European level if they are not caused by country-specific circumstances.

3 Methods

Measuring the distribution of various dimensions of inequality increases the degrees of freedom for normative choices. One has to decide not only on the level of inequality aversion but also on the order of aggregation, on the level of substitutability, and on the relative weights of dimensions. At the same time, high comparability between a multidimensional measure and unidimensional measures such as the Gini index for equivalized disposable household income is desirable to examine the results in the context of previous research and facilitate relevant policy conclusions.

Aggregation

To ensure comparability and explicit consideration of normative choices, this work relies on a combination of a CES-like aggregation function and classical inequality measures. First, the CES function aggregates outcomes for each individual while defining the degree of substitution and the relative weights and controlling for the correlation between dimensions. Second, aggregating across individuals using a Gini index makes it possible to set the degree of inequality aversion and maintains a certain degree of comparability with the unidimensional Gini index (Banerjee, 2010; List, 1999). Both steps can be merged into a single well-being function that fulfills most necessary axioms for inequality measures (Decancq and Lugo, 2012). However, the Gini index restricts the degree of inequality aversion and cannot be decomposed into additive subgroups and factor shares at the same time, which is essential to conduct the transnational analysis and to assess the degree of σ-convergence between countries.[5] Therefore, a second specification based on the Generalized Entropy (GE) indices complements the results of the

[5]The restrictive use of the Gini index has also been criticized based on the fact that it is relatively insensitive to changes at the top and bottom of the distribution (Osberg, 2017).

Gini index and allows for subgroup decomposition while fulfilling a similar set of axioms for inequality measures (Maasoumi, 1986). As Bosmans et al. (2015) showed, both two-step aggregation methods have a normative justification if measuring inequality is the only objective.

This leads to a universal CES-like aggregation function (1) aggregating individual achievement a^i across different dimensions j including respective weight w_j and the degree of substitution β as well as three inequality measures, which differ not only in decomposability but also in inequality aversion. The GE_0, also known as the mean log deviation, is more sensitive to changes at the lower end of the distribution, whereas the GE_1 or Theil index emphasizes changes at the top of the distribution. Together with the Gini index, which is most sensitive to changes at the middle of the distribution, the three indices provide a broader picture of distributional changes (Cowell, 2011).

$$x_i = \left(\sum_{j=1}^{m} w_j (a_j^i)^{1-\beta} \right)^{\frac{1}{1-\beta}} \qquad \text{if } \beta \neq 0, 1 \tag{1}$$

Parameter choices

The axiomatic approach highlights four normative criteria needed to measure multidimensional inequality: dimension selection, weighting, substitution elasticity, and inequality aversion.[6] Assessing the impact of all four parameters is beyond the scope of this work. To simplify the empirical analysis, dimension selection and substitution elasticity are based on established parameter choices. The dimensions of economic inequality are derived from Bourdieu's theory of socio-economic stratification, while the selection of proxies for each dimension closely follows Poppitz (2017). According to Bourdieu, stratification can be described by three types of capital: economic, cultural, and social. They are distinguished by their transferability between individuals and mode of accumulation (Bourdieu, 1983).

While the dimensions are selected based on Bourdieu's Capital Theory, the relative importance of each proxy is determined by hedonic weights, which makes the weighting procedure a mixture of statistical methods and normative criteria.[7] To derive the hedonic weights, the dimensions of inequality are regressed on a subjective measure that consistently represents the welfare rank or position within society of each individual. After controlling for the influence of other factors (Z_{it}), the estimates yield the relative importance of each dimension. Because the functional estimation function also determines the marginal rate of substitution between

[6] Of course, additional empirical problems can affect these normative parameters, such as the method of normalization of outcomes.

[7] Decancq and Lugo (2013) have surveyed weighting methods, while Brandolini (2009) and Poppitz (2017) discuss the method of hybrid weights in detail.

dimensions, the specified regression model resembles the functional form of the aggregation:

$$SSS_{it} = \alpha + \sum_{j=1}^{m} \beta_j x_{jit}^{\delta} + \gamma' Z_{it} + \upsilon_t + \epsilon_{it} \tag{2}$$

In this case, δ is equivalent to the degree of substitution β. Model (2) is estimated for a range of reasonable parameter choices ($0 < \delta < 2$) and the parsimonious model is selected based on the smallest log-likelihood. Replicating the CES functional form not only makes it possible to estimate the degree of substitution, but also, in the case of only one dimension, the functional form is equivalent to standard unidimensional measures of income inequality. The drawback is the assumption of constant and equal marginal rates of substitution for all dimensions. Subjective social status (SSS) is used as the subjective measure (S_{it}), which represents the individual self-reported position within society on a ten-point scale from top to bottom.[8] In contrast to other subjective measures such as life satisfaction, SSS depicts the relative position within society in the medium or long term (Evans and Kelley, 2004; Kelley and Evans, 1995).

Factor and subgroup decomposability of multidimensional inequality

Inequality measures based on generalized entropy indices are additively decomposable into subgroups (c) by equalizing the effect of the respective between-country and within-country component (Cowell, 2011):

$$GE_1^b = \frac{1}{N} \sum_{i=1}^{N} \left[\frac{\bar{x}_{ic}}{\bar{x}} \ln \frac{\bar{x}_{ic}}{\bar{x}} \right] \tag{3}$$

$$GE_0^b = \frac{1}{N} \sum_{i=1}^{N} \left[\ln \frac{\bar{x}}{\bar{x}_{ic}} \right] \tag{4}$$

In addition, to investigate the contribution of a particular dimension to total inequality GE_α it can be helpful to decompose inequality into factor shares. While Shorrocks (1982) provided a decomposition method for additive factor shares of the Gini index, Remuzgo and Sarabia (2015) showed how to decompose multiplicative factor shares of the Theil index (GE_0) by constructing counter-factual distributions for each dimension. Using a related approach, Decancq et al. (2017) showed how inequality can be decomposed for any GE index while controlling for the effect of correlation between dimensions. To control for the contribution of the correlation between different dimensions at the individual level, all outcomes within each dimension are repeatedly reshuffled at random. The average inequality estimate over all reshuffles yields the contribution of the correlation between dimensions $GE_\alpha(\tilde{L})$. Subsequently, achievement in one dimension is replaced stepwise by the average achievement before reshuffling again to

[8]The exact question respondents are asked is “In our society there are groups which tend to be towards the top and groups which tend to be towards the bottom. Below is a scale that runs from top to bottom. Where would you place yourself now on this scale?” (ISSP, 2016).

obtain the contribution of each dimension $GE_\alpha(\bar{L}_j)$. Together, total inequality is decomposed into $m+1$ components:

$$GE_\alpha = \left(GE_\alpha(L) - GE_\alpha(\tilde{L})\right) + \left(GE_\alpha(\tilde{L}) - GE_\alpha(\bar{L}_j)\right) \tag{5}$$

Due to the additive subgroup decomposability of GE indices, Decancq et al. (2017, p. 231) provide a solution to decompose the contribution of each factor share by population subgroups. The method yields factor shares for both within- and between components and thereby the contribution of each dimension to convergence or divergence within the euro area. However, to estimate the contribution of each factor share to between-country inequality requires reshuffling achievement levels, not only within dimensions but also within subgroups. Since the between-groups contribution is based on subgroup averages ($\bar{x}_{ic}$ in equations (3) and (4)), there is no contribution by correlation to the between-country inequality.

4 Data, sample, and estimation of weights

Multidimensional inequality is estimated using the European Survey of Income and Living Conditions (EU-SILC), the successor to the European Community Household Panel. EU-SILC covers all of the members and prospective members of the European Union since 2005, including harmonized sample selection and weighting criteria with between 5,000 and 30,000 observations per country and year (EU-SILC 2018).[9]

Economic capital is approximated by equivalized net household income as provided by EU-SILC, including imputed rents and transfers minus taxes (WINC). Education and occupational prestige aim to proxy cultural capital. Occupational prestige is derived from the ISCO occupational category, transformed into Standard International Socioeconomic Occupational Status (SIOPS) from Ganzeboom and Treiman (1996), while education is measured in years (EDUCYRS). In the absence of common proxies, social capital is approximated by the employment status (EMPLY). The argument is that once controlling for income loss in the case of unemployment, there is an additional effect on social capital due to the loss of recognition and social networks. Details on the empirical definition of the proxies can be found in Table A.3 and descriptive statistics in Table A.4. The correlation matrix (Table 1) reveals a positive, but relatively low correlation between WINC and EMPLY, suggesting that employment status contributes additional information on individuals, as income is shared within households by definition.

To ensure comparability of the proxies over time and across countries, monetary variables are converted into purchasing power standards (PPS) based on household final consumption as suggested by Brandolini et al. (2012). To prevent systematic missing variables for education

[9]None of the EA-13 countries in the sample use register data, minimizing a potential bias due to different survey methods (Krell et al., 2015). In addition, sampling information is used to estimate standard errors (Goedemé, 2013).

Table 1: Correlation among dimensions of inequality

	WINC	EDUCYRS	SIOPS	EMPLY
WINC	1			
EDUCYRS	.183	1		
SIOPS	.369	.284	1	
EMPLY	.203	.0536	.136	1

Note: Pairwise correlation coefficients using population-inflated cross-sectional weights. *Source:* EU-SILC (2018).

and occupational prestige, the target population consists of individuals between the ages of 18 and 64 who are not in education. Using personal cross-sectional weights (PB040), the sample has been re-weighted to match the target population. As EU-SILC surveys income in the previous calendar year, most studies using these data backdate income observations by one year. However, the reference year of all non-monetary variables equals the survey year, which is why income observations are not backdated in this case. Alternatively, the four-year rotating panel structure would make it possible to calculate the income in the reference year, but only for three quarters of the sample. Due to the delayed availability of EU-SILC panel data and the missing observations, this work acknowledges the conflict in reference years but ignores this aspect in the calculations reported below in order to use the latest waves including all available observations.

To estimate the aggregation weights for each dimension, the International Social Survey Program (ISSP) serves as a second data source as EU-SILC does not provide information on subjective social status. The ISSP consists of harmonized cross-sectional surveys from national general social surveys and covers topics similar to EU-SILC. However, the ISSP lacks the high level of harmonization, has substantially fewer observations per wave, and is not available annually for each country. The most significant issue, however, is that some European countries in the ISSP report gross instead of net household income. Therefore, the observed sample is restricted to 9 out of 13 euro-area members in 2007.[10] This reduced sample is not representative of the whole EA-13, representing only 93.4% of the total EA-13 population in 2016, but as I estimate hedonic weights fixed across countries and over time, I assume this effect to be minor. Even the potential effect on estimated weights of the missing countries Greece and Ireland, which saw massive economic transformations in the sample period, should be limited in a sample of nine countries and four time spells.[11]

In order to harmonize the available data sets and to minimize the selection bias due to missing country/year waves in the ISSP, only one wave per country and three-year time spell was selected. If more than one wave per time-spell was available, the wave with the most

[10] The four missing countries are Finland, Greece, Ireland, and Luxembourg.

[11] The alternative, omitting the four countries from the EU-SILC sample, yields lower transnational and between-country inequality estimates. This effect is mainly driven by Greece, while the other three countries barely affect overall results.

Table 2: Sample of country/year observations from ISSP

	AT	BE	DE	ES	FR	IT	NL	PT	SI
2005 - 2007		2005	2006	2007	2007		2006	2006	2005
		(920)	(1796)	(1142)	(1290)		(1152)	(690)	(342)
2008 - 2010	2008	2008	2008	2010	2009	2009	2008		2009
	(575)	(810)	(1649)	(1153)	(1752)	(268)	(1262)		(391)
2011 - 2013	2013	2011	2012	2012	2011	2011	2011	2012	2011
	(641)	(749)	(2103)	(2293)	(1900)	(550)	(801)	(563)	(360)
2014 - 2016	2016	2015	2014	2014	2016		2014		2015
	(545)	(629)	(2237)	(1436)	(895)		(987)		(453)

Note: The table shows for each country and three-year time span the selected ISSP wave and the number of non-missing observations in parentheses. *Source:* ISSP (2016).

observations was chosen.[12] After deleting missing observations row-wise, this leaves 32,224 observations in total and between 268 and 2293 observations per country and time spell (Table 2). Demographic and control variables have been transformed to harmonize changing variable definitions and survey methods over time and to match variable definitions of the EU-SILC.[13]

Based on regression model (2) hedonic weights were estimated using an OLS estimator, country/year fixed effects, and the ISSP data. Non-linear estimation models accounting for the ordered dependent variable yield similar results, but have been discarded due to lower efficiency (Bahamonde-Birke and Ortúzar, 2017). Besides the four dimensions, the model includes age, age squared, sex, household composition, and marital status as control variables. All covariates were z-standardized to ensure comparability.

Table 3 reports the estimation results for the total sample, which will be used for multi-dimensional inequality estimates, and for each three-year time spell separately. Across all models, estimates are positive and highly significant. Since all variables are z-standardized, the estimates indicate the predicted change in subjective social status due to a variable change by one standard deviation. From the size of the estimates, I conclude that income is the most important dimension (0.645) and employment status is the least important, with a still sizable estimate of 0.134 while education and occupational prestige are equally relevant with estimates of 0.217 and 0.208, respectively. Finally, the degree of substitution between the dimensions of inequality, derived from σ, is estimated to be 0.589, which suggests that there is considerable complementarity between dimensions of inequality. Over the observed sample period, the relevance of education and occupational prestige increased at the expense of income. For employment status, I find a greater variation over time without a clear trend. Overall, the adjusted $r^2 = 0.317$ is in line with previous works but highlights once again that a substantial part of subjective social status remains unexplained.

[12] Within each spell, the country-specific sample weights (w^s) were reweighted by countries' population share (pop_c) to correct for different sample sizes per country and time-spell: $w^p = w^s * \left(\frac{pop_c}{\sum pop}\right) / \left(\frac{N_c}{\sum N}\right)$.

[13] The variable definitions (Table A.1) and descriptive statistics (Table A.2) are reported in the appendix.

Table 3: Hedonic weights regression

dependent variable	subjective social status				
	2005–2007	2008–2010	2011–2013	2014–2016	total
income	0.589	0.548	0.545	0.567	0.645
	(0.027)***	(0.053)***	(0.026)***	(0.055)***	(0.021)***
education (years)	0.182	0.178	0.191	0.286	0.217
	(0.028)***	(0.028)***	(0.027)***	(0.035)***	(0.015)***
occupational prestige	0.196	0.215	0.197	0.239	0.218
	(0.024)***	(0.033)***	(0.026)***	(0.031)***	(0.014)***
employed (dummy)	0.117	0.140	0.067	0.217	0.134
	(0.023)***	(0.026)***	(0.019)***	(0.030)***	(0.012)***
age	0.074	0.062	0.012	0.127	0.068
	(0.049)	(0.053)	(0.048)	(0.061)*	(0.027)*
age^2	0.105	0.112	0.075	0.183	0.125
	(0.052)*	(0.051)*	(0.051)	(0.061)**	(0.028)***
female	0.007	-0.002	-0.045	0.008	-0.013
	(0.021)	(0.021)	(0.021)*	(0.025)	(0.011)
hh composition	Yes	Yes	Yes	Yes	Yes
marital status	Yes	Yes	Yes	Yes	Yes
δ	0.460	0.517	0.457	0.379	0.411
adjusted r^2	0.285	0.316	0.335	0.325	0.317
N	7332	7860	9960	7182	32334

Note: $^+$ p<0.10, * p<0.05, ** p<0.01, *** p<0.001. S.E.s in parentheses. The table reports the estimation results for each three-year time spell and the pooled sample. δ reports the parameter choice that minimizes the log-likelihood within the parameter range $0 < \delta < 1$ for each regression. All regressors are z-standardized. *Source:* Author's calculations based on ISSP (2016).

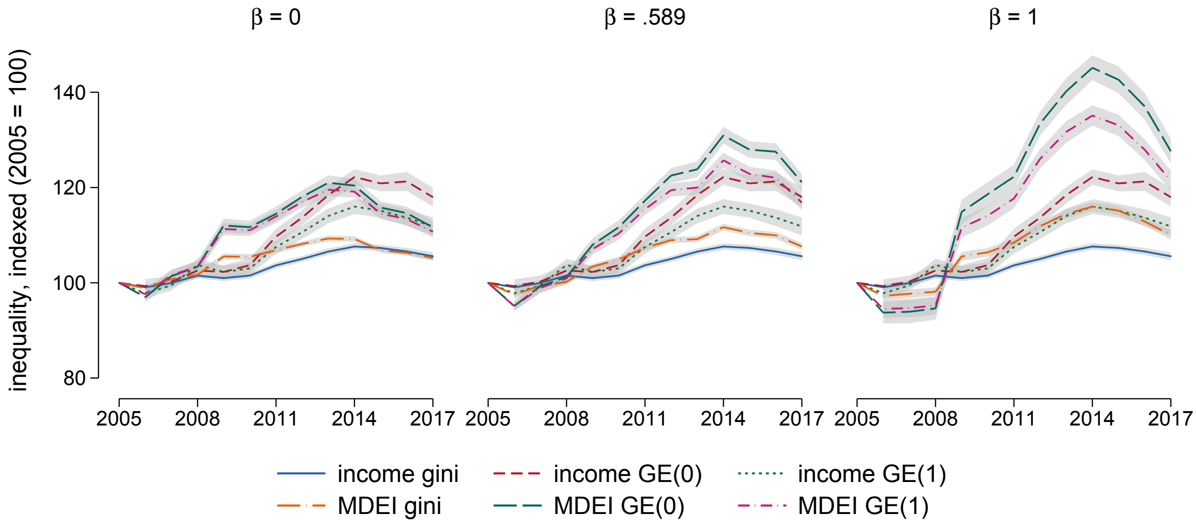
5 Results

As noted previously, this paper uses a variety of methods to investigate divergence in the euro area across multiple dimensions of inequality based on household survey data. Before discussing developments between countries, this section first presents the results from transnational inequality estimates and the contributions of different dimensions of inequality to transnational inequality overall. Second, maintaining the assumption of transnational well-being, divergence is assessed by comparing the contribution of inequality between countries to the inequality within countries. As before, the contribution of each dimension of inequality to the respective subgroup component is derived from counterfactual factor decomposition. Finally, the results section looks at national inequality estimates to examine whether convergence clubs among country have emerged during the economic and financial crisis using a *log t*-test and a clustering algorithm.

Transnational inequality over time

Irrespective of the specification, transnational inequality increased between 2006 and 2014 to previously unknown levels and has declined gradually since then. The increase in multidimensional inequality within the euro area timely parallels the economic recession in Europe and is at the same time contrasted by the gradual increase in income inequality between 2006 and 2014. Figure 1 illustrates the development of transnational MDEI and income inequality using inequality series indexed to 100 in 2005, compares inequality estimates from the Gini index and the Generalized Entropy indices, and distinguishes among three different levels of substitution elasticity (for absolute inequality estimates, see Table A.5).

Figure 1: Transnational income inequality and MDEI within the EA-13



Note: Income and MDEI inequality from 2005 to 2016 estimated by Gini and GE indices with $\alpha = \{0, 1\}$ and indexed to 2005 = 100. For multidimensional inequality, the degree of substitution varied $\beta = \{0, 0.589, 1\}$ using estimated dimension weights. The gray areas show 95% confidence intervals based on bootstrapped standard errors (512 rep.). Absolute inequality estimates are reported in Table A.5. *Source:* Author's calculations based on EU-SILC (2018).

The continuous growth in transnational income inequality starting in 2005 reached its peak in 2014, but the start of the economic recovery in the euro area reversed this trend. Income inequality in 2017 was slightly lower than in 2014, but still 5.57% higher than in 2005. Despite the strong increase, the Gini index for disposable household income in the euro area (0.300 in 2014) was still lower in 2017 than in the enlarged EU, at 0.336 (Vacas-Soriano and Fernández-Macías, 2017) or 0.377 (2013) in the US (LIS, 2018). The higher income inequality within the EU-28 comes as no surprise given the greater heterogeneity in the European Union. What is even more interesting is the downward trend within the EU-28 that came to a halt with the economic recession of 2009, while income inequality in the EA-13 continued to rise until 2014.

With respect to multiple dimensions of transnational inequality, the center graph in Figure 1 plots the preferred specification with an estimated substitution elasticity of $\beta = 0.589$. In direct comparison, multidimensional inequality has grown faster than income inequality, as revealed by the indexed time series and irrespective of the chosen inequality index. Although absolute levels of multidimensional inequality depend heavily on the substitution elasticity, the overall development during the crisis was the same for all degrees of substitutability except for one detail. Assuming that the dimensions of inequality are substitutes (left graph in Figure 1) inequality started to rise in 2008 and later increases were only gradual, compared to the center and right-hand graphs, according to which inequality increased substantially between 2009 and 2014 when assuming higher degrees of complementarity ($\beta > 0$). This sensitivity to substitutability suggests that before 2010, all dimensions of inequality increased, but that distributional changes across dimensions were uneven across households in the following years.

Comparing the results of the different inequality indices, two observations stand out. First, the stable difference between the GE_0 and GE_1 estimates across all levels of substitutability reveals that inequality rose even more sharply at the bottom than at the top of the multidimensional and income distribution. Second, individuals at the bottom of the distribution seem to have had more problems substituting low outcomes in one dimension with higher outcomes in another, as the gap between GE_0 and GE_1 widens with greater complementarity.

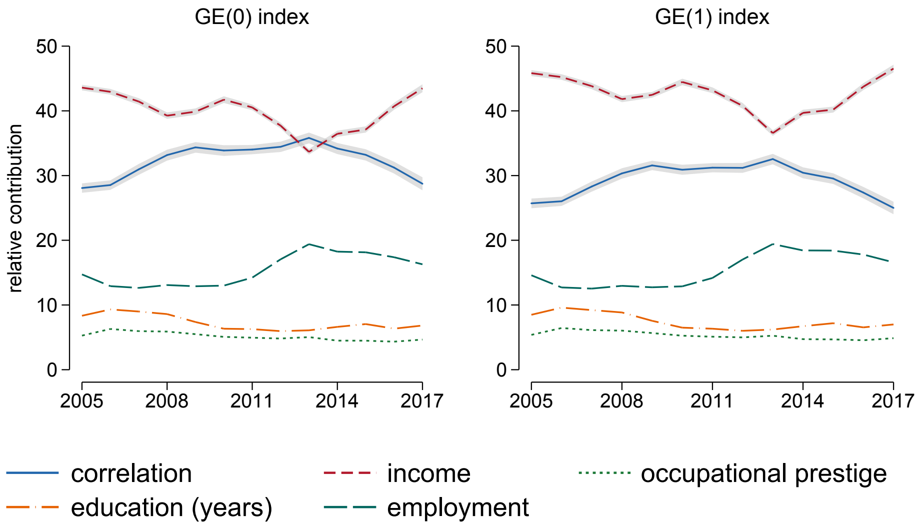
To summarize these results, non-monetary dimensions of transnational inequality increased more sharply and two years earlier than transnational income inequality. In addition, the financial crisis seemed to have only a limited effect on a single dimension of inequality, whereas the following economic recession had a sweeping effect on multiple dimensions of inequality. Clearly, a decomposition by dimension is warranted to understand the role of each of these dimensions and their joint development.

Factor decomposition and the role of employment status

Throughout the crisis, the contribution of different dimensions to total inequality changed substantially. At first, rising income inequality played a major role, but starting in 2010, employment status took over the central role. Based on the factor decomposition methods presented in Section 3, the absolute contribution of each dimension to total inequality is reported in Figure 2. Although all three of the inequality measures considered can be decomposed by factor shares, only the results from Generalized Entropy (GE) indices are presented, as the following subgroup decomposition is restricted to this class of indices.

To eliminate the effect of correlations between dimensions, outcomes are reshuffled by random across individuals. As a result, the mean of inequality estimates after reshuffling is subtracted from the original inequality estimate to derive the absolute contribution of the correlation, while the standard errors are obtained for the reshuffled results. On average, more than 32% of total inequality can be attributed to the correlation between dimensions. Therefore, ignoring the contribution of correlation by assuming perfect substitutability or aggregating across individuals first, as the HDI does, would seriously underestimate inequality. In every year since 2006, the contribution of the correlation to inequality increased until 2014. As such, one could describe the increase in multidimensional inequality as an increase in multiple deprivation. On average, the contribution is 3% lower for the GE_1 index compared to the GE_0 index. Intuitively, this difference suggests that low outcomes in multiple dimensions occur more often at the bottom of the distribution, which in turn leads the correlation component to increase together with inequality aversion.

Figure 2: Relative contribution of factor shares by inequality measure



Note: Relative contribution of factor shares and correlation to total inequality measured by GE_0 and GE_1 indices using estimated dimension weights and substitution elasticity (β = .589). Gray whiskers show bootstrapped 95% confidence intervals (512 rep.) *Source:* Author's calculations based on EU-SILC (2018).

According the the GE_0 index, the relative contribution of income is only slightly more important, at 39.8% on average, while employment status contributes 15.3% to total inequality on average. As expected, the contribution of income rises slightly as inequality aversion increases, but one would also expect that employment status is more important for individuals at the lower end of the distribution (GE_0). However, in light of the substantial increase of the employment dimension between 2010 and 2014, the difference in average contributions is negligible.

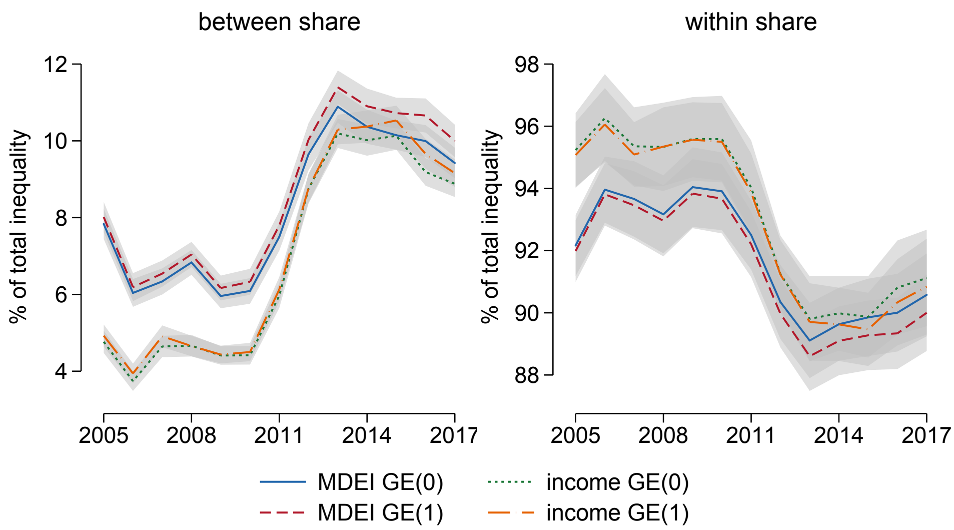
The most interesting result of the factor decomposition is how the interplay of income, employment, and correlation components contributes to overall inequality. Even before the financial crisis in 2008, the correlation between dimensions started to rise, thus amplifying the rise of income and employment status inequality in 2008 and 2011. In other words, transnational multidimensional inequality increased not only because of inequality in income and employment status, but also because more households, especially at lower end of the distribution, suffered from low outcomes in more than one dimension for which they could not compensate. Together, the income and correlation trend lead to rising multidimensional inequality, but only under the condition of some substitutability (Figure 1). Therefore, only when the economic recession hit the euro area and unemployment rates started to rise in 2010 did inequality begin to increase. This occurred irrespective of the degree of substitution, even though the contribution of income inequality did not grow further after 2011. In a similar vein, the decline in multidimensional inequality since 2014 is driven more by a decline in the correlation component and the employment dimension than by income inequality. Finally, occupational prestige and educational inequality do contribute to inequality, but their relative contribution to total inequality is relatively small.

In summary, in terms of levels, income is the major source of transnational inequality in the euro area, but employment status inequality and the correlation between dimensions substantially contributed to the increase in multidimensional inequality within the EA-13 between 2009 and 2014. After 2014, the contribution of the correlation between multiple dimensions of inequality decreased, but as income inequality increased further and unemployment recovered only slowly, multidimensional inequality in the EA-13 is still significantly higher than before the crisis.

Subgroup decomposition and between-country divergence

The fact that transnational inequality has risen over the last decade, as shown in Figure 1, also raises the question of whether this was driven by greater disparities within countries or by divergence between countries. Without giving up the transnational assumption, we can analyze the contribution of between-country differences using the additive subgroup decomposability of Generalized Entropy measures. Figure 3 illustrates the strong increase in σ-divergence by showing the percentage of total inequality, explained by between-country differences in income and multidimensional inequality.

Figure 3: Subgroup decomposition of inequality for EA-13



Note: Share of total income and multidimensional inequality in the EA-13 explained by between-country and within-country inequality. Measured by GE_0 and GE_1 indices using estimated dimension weights and substitution elasticity. Whiskers show 95% confidence intervals based on bootstrapped standard errors (512 rep.). *Source:* Author's calculations based on EU-SILC (2018).

In general, only a small fraction of total inequality in the EA-13 is explained by heterogeneity between countries, while more than 90% of the total inequality results from heterogeneity within countries. In absolute numbers, total and between-country income inequality are higher than multidimensional inequality (Table A.8), but before the crisis, the share of multidimensional inequality resulting from differences between countries was higher than for income alone. However, in the years leading up to the financial and economic crisis, the share of between-country inequality increased by 5.5 percentage points for income and by 4 percentage points for multidimensional inequality.

Within only three years, from 2010 to 2013, the between-country share roughly doubled. Since 2013, between-country inequality for income and MDEI have contributed more than 10% to total inequality. This level of cross-country divergence among the EA-13 was only reached previously prior to 1998, one year before the introduction of the euro (Papatheodorou and Pavlopoulos, 2014, p. 456). Therefore, both well-being concepts, income and MDEI, confirm previous results on σ-divergence within the EA-13 (Bönke and Schröder, 2014, p. 21). This development stands in contrast to that in the EU-28, where rising income inequality during the economic recession led to a halt of convergence between countries, but did not cause a trend reversion (Vacas-Soriano and Fernández-Macías, 2017).

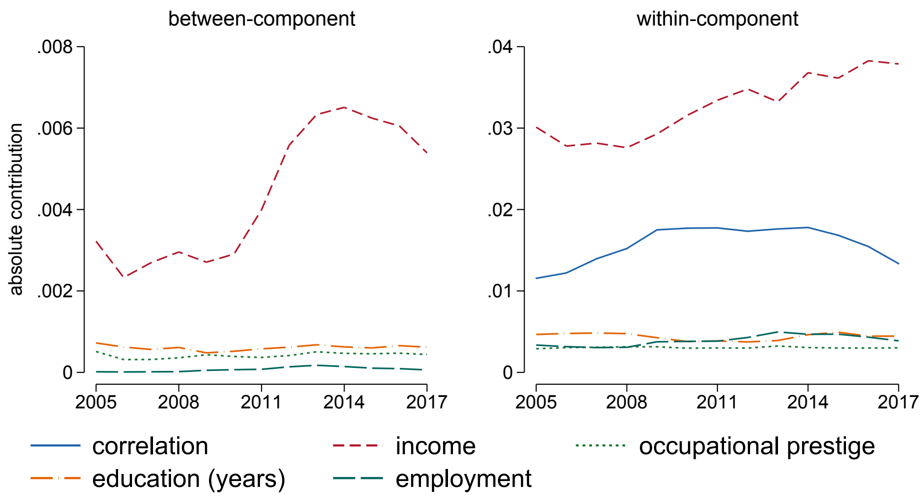
The timing deserves special attention, because divergence increased from 2010 onward, whereas transnational inequality already began to increase in 2008 when the financial crisis first hit. Because incomes are reported for the previous calendar year in the EU-SILC survey, in contrast to the other dimensions, the time series might lag behind real developments, but not by more than one year. Therefore, the transnational inequality estimates clearly confirm that the economic recession and not the financial crisis drove the euro area apart.

Drivers of divergence

Figure 2 suggested that income is the single most important dimension of economic wellbeing in transnational inequality within the EA-13, but which dimensions pushed the countries of the initial euro area apart during the economic recession? Conveniently, the subgroup contributions can be further decomposed by factor shares as outlined in Section 3, with the exception that the correlation among dimensions does not contribute to between-country inequality by definition.

Figure 4 plots the absolute contribution of each dimension to the respective subgroup inequality component. When comparing the respective contributions to within- and between-country inequality, the differences are again more substantial for income. While income contributes on average 54.6% to within-country inequality, the contribution to between-country inequality rose steadily from 72% in 2005 to 84% in 2014. Conversely, the contribution of non-monetary dimensions such as education and occupational status remained relatively stable over time. Only cross-country inequality in employment status increased slightly during the recession years, but the relative contribution to between-country inequality is still small with the factor share rising from 0.3% to 1.8%. The non-monetary dimensions are of greater relevance for within-country inequalities. Occupational prestige, education, and employment

Figure 4: Factor share decomposition of subgroup inequality



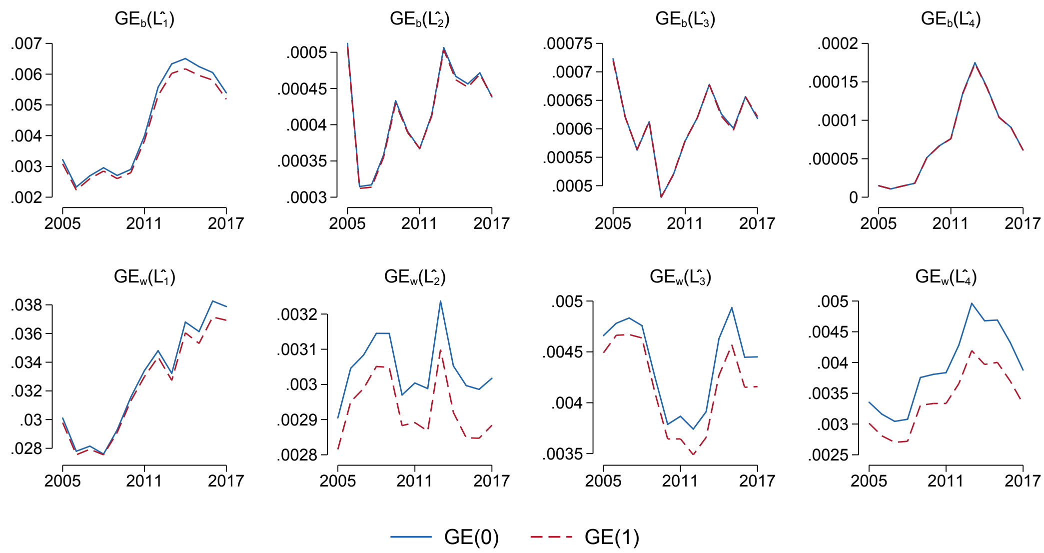
Note: Absolute contribution of factor shares to between-country and within-country inequality (GE_0) using the estimated dimension weights and substitution elasticity (β = .589). *Source:* Author's calculations based on EU-SILC (2018).

status make a relatively stable contribution to within-country inequality, at 5.13%, 7.43%, and 6.51% respectively on average.

In general, two important conclusions can be drawn from Figure 4. First, the rise in cross-country divergence was mainly caused by increasing income differences between EA-13 countries, since no other dimensions of economic inequality saw such a significant rise in heterogeneity across countries. Because cross-country income inequality has not decreased substantially since 2014, neither has total inequality between countries. Second, the short but persistent increase in between-country inequality was accompanied by a gradual increase in within-country inequality of income and employment status. After 2014, neither of the two dimensions saw a substantial decline, which makes the correlation between dimensions the major component contributing to the total decline in within-country inequality. This suggests that with the economic recovery, more households found it easier to compensate for lower achievement in one dimension with higher achievement in other dimensions, resulting in a lower number of households that were deprived in multiple dimensions of economic inequality, even though inequality in the separate dimensions remained high.

As a robustness check, Figure 5 presents the different distributional impact of each dimension to by comparing the absolute factor shares after varying inequality aversion. The upper row suggests that the distribution between countries is not sensitive to inequality aversion. However, factor shares of within-country inequality differ with respect to inequality aversion as the lower row of Figure 5 indicates. Inequality of employment status is more severe at the bottom of the distribution, which leads to a higher factor share of both dimensions when using the GE_0. To summarize, income disparities have driven the countries in the euro area apart, while the

Figure 5: Factor shares by varying inequality aversion of GE indices



Note: Comparison between absolute factor shares of between-country and within-country inequality for GE_0 and GE_1 indices using estimated dimension weights and substitution elasticity (β = .535). *Source:* Author's calculations based on EU-SILC (2018).

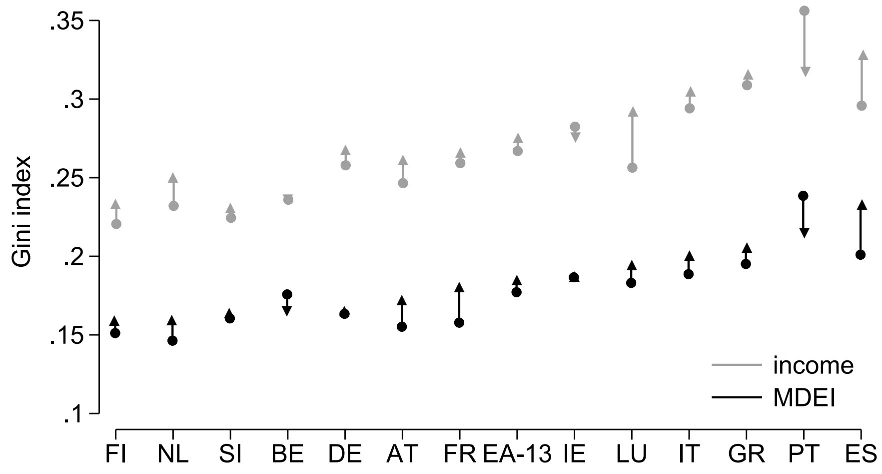
poor performance of labor markets and higher income inequality increased social stratification within countries, especially at the bottom of the distribution.

National inequality and convergence clubs

In a final step, I abandon the assumption of transnational inequality and thus also the determination of individual welfare relative to other households in the euro area. This makes it possible to depict the development of multidimensional and income inequality on a country level, to identify the country-specific contribution to divergence in the euro area, and to test for club convergence using established clustering methods (Phillips and Sul, 2009).

According to Figure 6, the number of countries that saw increases in income inequality (gray) varies widely, from relatively equal countries (Finland, Netherlands) to countries with average inequality (Austria, France) and those with high inequality (Spain, Greece, Italy). Outliers are Portugal, where income inequality declined from a very high level, and Luxembourg, where the opposite development occurred. For multidimensional inequality (black), we can observe a relatively similar development, with half of the countries showing a rise in inequality and the other half of countries showing only small changes in inequality. Again, Portugal is an outlier, with a significant reduction in multidimensional inequality, as is Belgium, where multidimensional inequality declined against the upward trend in income inequality. The unweighted average Gini indexes for income and multidimensional inequality reported in Figure 6, about 0.1 points lower than the respective transnational inequality estimates because they ignore by definition the cross-country inequality. What remains rather unclear from this graph is how the individual changes in within-country inequality have contributed to

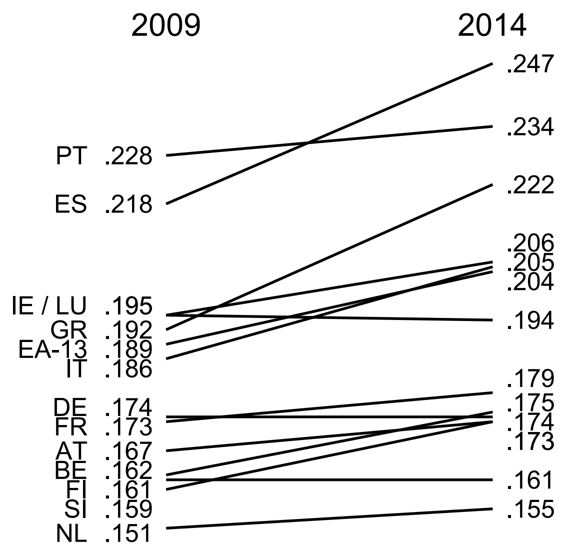
Figure 6: Country estimates of income inequality and MDEI, 2005-2017



Note: Change in income inequality and MDEI between 2005 to 2017 measured by the Gini index using estimated dimension weights and substitution elasticity (β = .535). Countries sorted by MDEI in 2017. *Source:* Author's calculations based on EU-SILC (2018).

the overall process of divergence in the euro area, or more specifically, whether individual countries or convergence clubs caused the overall divergence in the euro area.

Figure 7: Rank changes in multidimensional inequality, 2009-2014



Note: Changes in multidimensional inequality from 2009 to 2014 measured by the Gini index using estimated dimension weights and substitution elasticity (β = .535). *Source:* Author's calculations based on EU-SILC (2018).

Before investigating the question of club divergence statistically, Figure 7 illustrates the rank and level changes in multidimensional inequality during the most turbulent period, 2009 to 2014. Due to the considerable differences in levels, three country groups are intuitively identified based on Figure 7, with Portugal and Spain in the top group. Despite a lower level of inequality in the second group (Luxembourg, Ireland, Greece, and Italy), inequality grew on average by 7% over the five years. Only the last group, consisting mainly of central European countries, saw inequality growing by only 4.1% on average. At first glance, the graph suggests that three country clubs were driving σ-divergence, although the visual identification of convergence clubs is arbitrary by definition.

To formally test for the existence of convergence clubs, I use the method proposed by Phillips and Sul (2007, 2009). The *log t*-test as proposed by these authors makes no parametric assumption about the convergence process and is robust to common time series estimation problems. In addition, the clustering algorithm identifies convergence clubs endogenously, whereas in other algorithms, the number of clubs needs to be specified ex ante. The *log t*-test relies on the assumption that a balanced time series panel (country-year observations of inequality estimates) can be described by a transitory and a static component. If the former component tends towards the panel average, this implies σ-convergence. This relative transition is tested by a specific test regression, where the estimated transition coefficient is expected to be $\hat{b} \geq 0$ in the case of convergence with the null hypothesis of convergence (Phillips and Sul, 2007). Given that the previous results have suggested a process of divergence within the euro area, I expect to reject the null hypothesis of convergence for the full sample.

By using an iterative procedure as described in Phillips and Sul (2009), the *log t*-test makes it possible to identify the number, composition, and trend of convergence clubs endogenously without a prior assumption about the composition of the clubs. In short, the algorithm starts with an initial country and tests whether other countries can be added to the club without rejecting the null hypothesis of convergence. If no more converging countries are found, the algorithm repeats the exercise with the remaining countries, until every country either belongs to a convergence club or is found to be individually divergent. The original method suggests using the country with the highest outcome (GDP per capita) in the final year as the starting point of the identification procedure. In the case of inequality, this would make the procedure highly dependent on extreme cases, which is why convergence clubs are identified starting with the country with the lowest inequality in the last observed year. As a safeguard, I rely on an extended version of the algorithm to prevent of an over-identification of convergence clubs (Schnurbus et al., 2017). Similar to previous studies and as recommended by Phillips and Sul (2009), the observations from 2005 to 2008 ($k = 0.3$) are selected as the reference period to test for convergence. In contrast to the growth convergence literature, I refrain from using a smoothing algorithm to distinguish between transitory and static components of inequality, because the aim is to observe how inequality reacts to macroeconomic shocks. All estimations were carried out using the Stata package provided by Du (2017).

The null hypothesis of convergence for the Gini index is rejected using the *log t*-test for income ($\hat{b} = -1.0995$, $\hat{t}_b = -9.0853$) and multidimensional inequality ($\hat{b} = -1.0870$, $\hat{t}_b = -6.2514$). While these results reject the hypothesis of convergence across the EA-13, they leave open whether overall divergence or club convergence is the cause. According to the club convergence algorithm using the *log t*-test, two convergence clubs can be identified. However,

Table 4: Convergence clubs of inequality in the Euro area (EA-13)

	income			mdei		
club #	Gini	GE0	GE1	Gini	GE0	GE1
1	-0.338 (-1.079) AT BE FI FR DE IE NL SI	-0.283 (-0.836) AT BE FI FR DE IE NL SI	-0.328 (-1.005) AT BE FI FR DE IE NL SI	0.201 (0.608) AT BE FI FR DE IE LU NL SI	-0.037 (-0.093) AT BE FI FR DE IE LU NL SI	-0.152 (-0.425) AT BE FI FR DE IE LU NL SI
2	-0.292 (-0.614) GR IT LU PT ES	-0.586 (-1.164) GR IT LU PT	-0.086 (-0.207) GR IT LU PT ES	0.188 (0.997) GR IT PT ES	1.882 (3.247) GR IT PT	-0.139 (-0.860) GR IT PT ES
none		ES			ES	

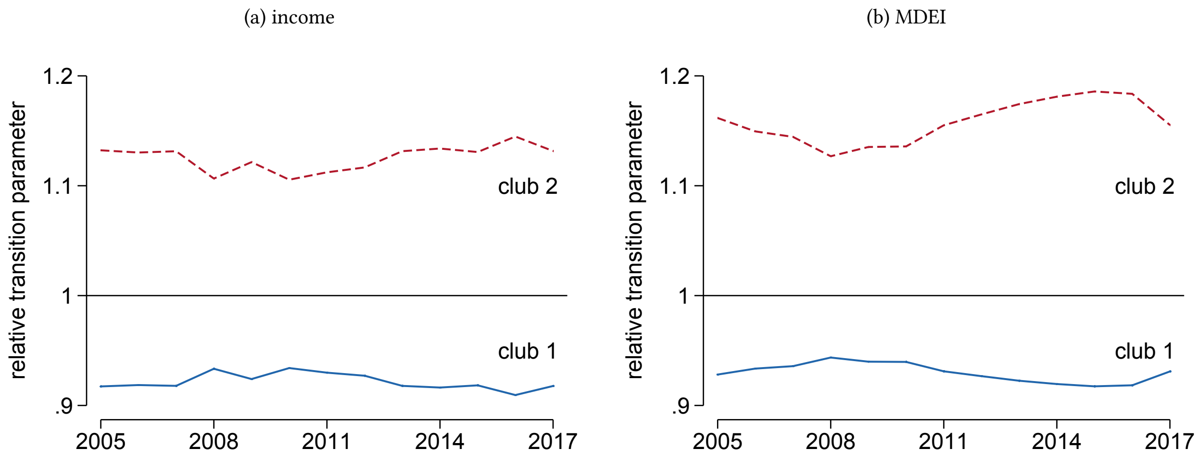
Note: Convergence clubs for income inequality and MDEI identified by a clustering algorithm based on *log t*-test for three different inequality indices (Gini, GE_0, and GE_1). Each cell reports $\hat{b}$ and $\hat{t}_b$ of the respective *log t*-test and the countries that belong to the club. The final row lists the group of non-converging countries. Clubs are identified by a four-step algorithm (Phillips and Sul, 2009) starting with the country with the lowest inequality in the final period. *Source:* Author's calculations based on EU-SILC (2018).

the exact club definition and the number of individually divergent countries are sensitive to the chosen inequality index and well-being concept (Table 4).

Across all specifications, the group of central European countries including Ireland and Finland turns out to be the first robust convergence club. The second club is again represented by a core group including Italy, Greece, and Portugal, which are sometimes joined by Spain or Luxembourg. Comparing the results for income and multidimensional inequality, no clear differences are evident. However, the affiliation of Luxembourg, which experienced the greatest increase in income inequality of any country in the sample, depends on the dimension selection. According to income inequality, Luxembourg belongs to the second club, whereas multidimensional inequality finds Luxembourg in the first club. If anything, then the lower point estimates of the *log t*-test for income suggest stronger divergence within clubs than for multidimensional inequality. Moreover, Spain is usually found to belong to the second group, but when using the GE_0 index, which is less inequality averse towards the top, Spain is found to be an individually divergent country, emphasizing the exceptional adverse effect of the economic recession on poor households in Spain.

Figure 8b shows that changes in multidimensional inequality of the countries in the second convergence club are not the only culprits behind the overall divergence in the EA-13 since 2010. The graphs plot relative transition curves, calculated from the cross-sectional averages of the relative transition parameters for both convergence clubs (Phillips and Sul, 2009, p. 1159). As the transition parameter is rescaled by the panel average, parameters below one indicate lower-than-average inequality and a movement towards one would indicate convergence. According to multidimensional inequality, divergence between convergence clubs was mainly driven by the southern European countries (club 2) as they drifted further away from the panel average than in the case of income inequality. Whether a continued economic recovery will

Figure 8: Relative transition paths of convergence clubs

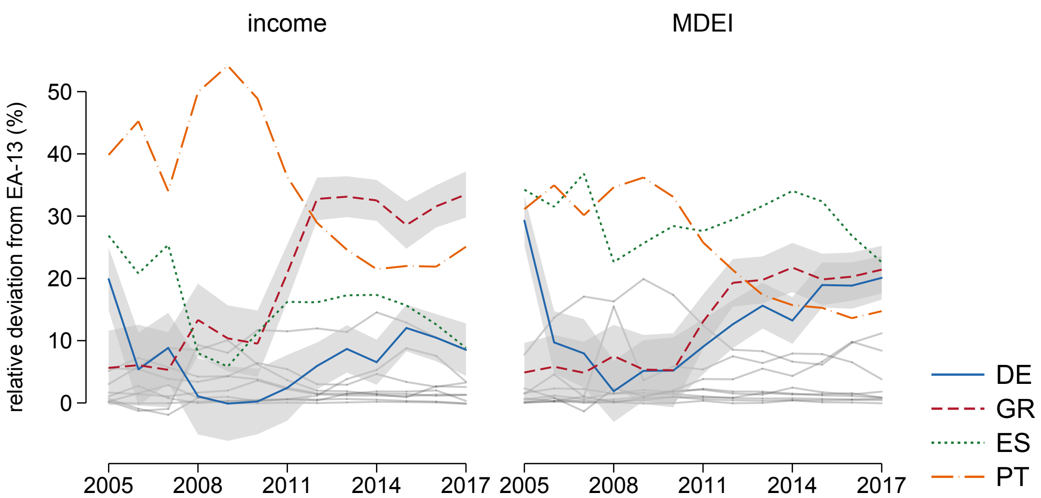


Note: Relative transition paths for convergence clubs derived from cross-sectional averages of each club. Based on Gini index (Columns 1 and 4 of Table 4) *Source:* Author's calculations based on EU-SILC (2018).

bring back a convergence in multidimensional inequality remains speculative, but the small downturn in 2017 gives hope.

One shortcoming of the club convergence test used above is that inequalities between countries are ignored by definition, because only cross-country differences in inequality levels are compared. However, the subgroup decomposition of transnational inequality estimates revealed that between-country differences contribute up to 10%. Therefore, a second method is used to assess the impact of individual countries on convergence in the euro area.

Figure 9: Country-specific contribution to subgroup inequality in EA-13



Note: Absolute difference between EA-13 between-country inequality and inequality estimate after replacing outcomes of each country with EA-13 average (without the country of interest). Measured by multidimensional GE(0) index using estimated dimension weights and substitution elasticity (β = .535). *Source:* Author's calculations based on EU-SILC (2018).

To this end, Figure 9 illustrates the contribution of each country individually by plotting the relative change of between-country inequality when the respective country outcomes are replaced with average outcomes of all other countries (see Table A.9). As expected, Greece, Spain, and Portugal contribute substantially to income divergence in the euro area in the period after 2010, when between-country inequality skyrocketed to 10% of total inequality (Figure 3).

Again, multidimensional inequality tells a slightly different story. Spain replaces Greece as the single most important country driving cross-country divergence. Without Spain, between-country inequality in the EA-13 would be more than 30% lower. Surprisingly, in 2017, Greece contributes as much as Germany to divergence within the EA-13, at 21.4% and 20.0%, respectively. To put it differently, the relatively strong increase in multidimensional inequality in Greece drives divergence in the EA-13 as much as the relatively positive development in Germany. Stagnating income inequality contributed to Germany's outlier position, but without its exceptional development in the other dimensions, Germany's between-country contribution would be only half this size.

Overall, the answer to the club convergence hypothesis remains ambiguous. Income and multidimensional inequality point towards divergence between two clubs, basically Central Europe and Southern Europe. Income, however suggests a greater contribution of Southern Europe to overall divergence, whereas multidimensional inequality shows the contributions of both clubs to be similar. In addition, the distinct contributions of Spain, Greece, and Germany also allow those countries to be seen as three individual contributors to the overall divergence in the EA-13.

6 Conclusion

The question of how inequality developed over the last decade is especially relevant for the euro area, where the recent financial and economic crisis underscored existing heterogeneity and structural differences. Previous works showed convergence of incomes in the initial years of the common currency area, but a reversal of this process has led to increasing income divergence after 2008. However, wellbeing and inequality or the distribution of welfare is best understood as a multidimensional concept consisting of both monetary and non-monetary dimensions. Therefore, this work estimated multidimensional inequality using income, education, occupational prestige, and employment status whereas a hedonic regression framework (Poppitz, 2017) was used to weight dimensions of inequality as well as the degree of substitution between dimensions. Following the literature on transnational income inequality, σ-convergence was then assessed by multidimensional inequality estimates for all member states of the euro area in 2007 (EA-13) treating them as one single country.

Within the EA-13, my estimations show a strong increase in both income and multidimensional inequality starting in 2008. Income inequality supersedes multidimensional inequality, independent of the degree of substitution, indicating that non-monetary dimensions do substitute income inequality to some extent. Consequently, income is the most important dimension, contributing 37.1% of total inequality. Among individuals at the lower end, the correlation between dimensions matters the most, especially since the onset of the euro crisis.

The crisis was also the starting point for a rise in between-country inequality, a measure of σ-convergence, which increased from 6.1% in 2010 to 10.4% in 2014. According to the subgroup decomposition, divergence started to increase two years later than total inequality, which correlates closely to the outbreak of the euro crisis and the following economic recession, whereas total inequality already started rising with the financial crisis. Similar to overall inequality, the increasing divergence between countries is mainly driven by income differences, as this dimension contributed 84% to total cross-country inequalities in 2014. Despite the gradual rise in income inequality within countries, the correlation between dimensions and employment status inequality have contributed to higher within-country inequality, even though the relative share declined. In the light of the macroeconomic recession, the increasing multidimensional inequality can be attributed to households that could not share risks either between dimensions or within households (Vacas-Soriano and Fernández-Macías, 2017, p. 18) and the rising differences between countries. In the short run, labor market policies are one important factor in mitigating crisis effects. The fact that many countries developed their labor market policies based on their fiscal capacity rather than demand might have amplified the heterogeneity within the euro area.

Overall, the multidimensional perspective confirms the divergence within the EA-13, adding little additional information to the findings from income inequality. However, when it comes to the question of what (clubs of) countries contributed to the overall divergence, multidimensional inequality provides somewhat different results. Both well-being concepts suggest that two convergence clubs—Central Europe and Southern Europe—have emerged, and that the total divergence is a result of differences between the two clubs. However, income inequality suggests that rising inequality in southern Europe is mainly to blame, whereas multidimensional inequality suggests that both clubs contributed to σ-divergence by similar means. In addition, when looking at country-specific contributions to total divergence in multidimensional inequality, some individual countries, including Germany, appear to have played a special role. No other country of the size of Germany experienced stagnating multidimensional inequality despite the overall recession in the euro area. In 2013, Germany's relative contribution to multidimensional σ-divergence (15.6%) was therefore slightly lower than that of Greece (19.7%) and the contributions of both countries converged to 20% in 2017.

In summary, the German success story, with small but positive economic growth rates, stagnating income inequality, and decreasing unemployment rates can be seen from a different angle. Assuming that this development was made possible by the slow growth of unit labor costs compared to labor productivity, Germany was able to utilize a comparative advantage within the currency union at the expense of other euro-area members. In this case, the surging export surplus and the considerable σ-divergence within the euro area might be interpreted as two sides of the same coin. Although favorable for Germany, not all euro-area members can adopt this strategy at the same time, which might put the future development of the euro area at risk.

Because divergence seems to be remaining high, European policies aimed at economic convergence and social cohesion are needed now more than ever. Otherwise, doubts as to the perspectives of the euro will continue to arise, and macroeconomic policies for the whole euro area will face increasing policy trade-offs amid the high heterogeneity within the monetary union. However, the cross-country differences in multiple dimensions found in this work could be also related to regional or cultural heterogeneities (Alesina et al., 2017) masked by cross-country differences. In such cases, policies would be better aimed at regional or occupational groups rather than specific countries. Unlike Döpke et al. (2017), who show that the eligibility of EU regions for convergence policies depends on the weights attributed to specific dimensions of economic wellbeing, this work documented a strong rise in divergence at the country level irrespective of dimension selection, weight decisions and inequality aversion. In order to design efficient EU convergence policies, future studies are needed to clarify the relevance of nation-states and regions and to compare them to other reference frames based on occupational or cultural criteria.

A Appendix

Table A.1: Variable definitions and transformation of ISSP (2016)

Variable	Definition
income	Disposable income of all household members (di08), equivalized by new OECD scale and deflated by consumer price index from Eurostat (base year: 2010). Top 0.1% incomes winsorized.
education	Years of completed education.
occupational prestige	Standard International Occupational Prestige Scale from Treiman (1977). Derived from four–digit International Standard Classification of 2008 (fisco08 and misco08) and recoded with updated tables provided by Ganzeboom and Treiman (1996).
employment status	Dummy variable treating students, pensioners voluntary unemployed and others as not employed (0) and only full- and part-time as employed (1).
household groups	Recode of household types (dh05) into six groups: single, single with children, couple, couple with children, three or more generations, and others.
marital status	Dummy variables distinguishing between five groups: married, widowed, divorced, seperated but married, single.

Note: Available ISSP waves from 2004 to 2016 have been merged and changing variable definitions have been harmonized accordingly.

Table A.2: Descriptive statistics of the pooled ISSP (2016) sample

	mean	sd	min	max
subjective social status	5.53	1.71	1	10
disposable household income (monthly)	1,625	1,247	0	14,479
education in years	12.9	4.07	0	40
occupations prestige (SIOPS)	42.6	13.3	5	78
employment status (dummy)	.915	.279	0	1
age	43.4	12.2	18	65
female (dummy)	.5	.5	0	1

Note: Descriptive statistics of pooled ISSP sample using survey and population weights used for the estimation of hedonic weights. Income in real purchasing power units.

Table A.3: Variable definitions and transformation of EU-SILC (2018)

Variable	Definition
income	Disposable income of all household members, equivalized by new OECD scale and deflated by consumer price index from Eurostat (base year: 2010). Top 0.1% incomes winsorized.
education	Years of completed education.
occupational prestige	Standard International Occupational Prestige Scale from Treiman (1977). Derived from four–digit International Standard Classification of 2008 (fisco08 and misco08) and recoded with updated tables provided by Ganzeboom and Treiman (1996).
employment status	Dummy variable treating students, pensioners voluntary unemployed and others as not employed (0) and only full- and part-time as employed (1).
household groups	Recode of household types (dh05) into six groups: single, single with children, couple, couple with children, three or more generations, and others.

Table A.4: Descriptive statistics of the pooled EU-SILC (2018) sample

	mean	sd	min	max
disposable household income (annual)	21,277	11,769	.0612	125063
education in years	15	7.46	0	40
occupations prestige (SIOPS)	40.2	12.9	5	69
employment status (dummy)	.911	.284	0	1
age	43.9	12.1	18	65
female (dummy)	.487	.5	0	1

Note: Descriptive statistics of pooled EU-SILC sample using survey and population weights used for the inequality estimation. Income in real purchasing power units.

Table A.5: Transnational inequality estimates for the EA-13

	income			MDEI		
	Gini	GE_0	GE_1	Gini	GE_0	GE_1
2005	0.279	0.140	0.130	0.183	0.057	0.054
	(0.0008)	(0.0009)	(0.0007)	(0.0005)	(0.0003)	(0.0003)
2006	0.276	0.139	0.127	0.178	0.054	0.051
	(0.0009)	(0.0010)	(0.0008)	(0.0005)	(0.0003)	(0.0003)
2007	0.279	0.141	0.129	0.181	0.057	0.053
	(0.0008)	(0.0009)	(0.0007)	(0.0005)	(0.0004)	(0.0003)
2008	0.283	0.144	0.135	0.183	0.058	0.054
	(0.0009)	(0.0011)	(0.0009)	(0.0005)	(0.0004)	(0.0003)
2009	0.282	0.144	0.133	0.189	0.062	0.058
	(0.0009)	(0.0010)	(0.0008)	(0.0005)	(0.0004)	(0.0003)
2010	0.283	0.146	0.134	0.192	0.064	0.060
	(0.0009)	(0.0011)	(0.0009)	(0.0006)	(0.0004)	(0.0004)
2011	0.289	0.154	0.140	0.196	0.067	0.062
	(0.0009)	(0.0012)	(0.0009)	(0.0005)	(0.0004)	(0.0003)
2012	0.293	0.160	0.144	0.199	0.070	0.064
	(0.0009)	(0.0011)	(0.0009)	(0.0006)	(0.0005)	(0.0004)
2013	0.297	0.166	0.148	0.199	0.071	0.065
	(0.0010)	(0.0012)	(0.0010)	(0.0006)	(0.0005)	(0.0004)
2014	0.300	0.172	0.151	0.204	0.075	0.068
	(0.0010)	(0.0012)	(0.0010)	(0.0007)	(0.0005)	(0.0004)
2015	0.299	0.170	0.149	0.202	0.073	0.066
	(0.0009)	(0.0012)	(0.0010)	(0.0006)	(0.0005)	(0.0004)
2016	0.297	0.170	0.147	0.201	0.073	0.066
	(0.0011)	(0.0014)	(0.0012)	(0.0006)	(0.0005)	(0.0004)
2017	0.294	0.166	0.145	0.196	0.069	0.063
	(0.0012)	(0.0014)	(0.0012)	(0.0007)	(0.0005)	(0.0004)

Note: Inequality estimates for income and MDEI based on the Gini index or the GE indices ($\alpha = \{0, 1\}$) using estimated weights and substitution elasticity ($\beta = 0.535$). Bootstrapped standard errors in parentheses. *Source:* Author's calculations based on EU-SILC (2018).

Table A.6: Absolute contribution of factor shares

	2005	2006	2007	2008	2009	2010	2011	2012	2013	2014	2015	2016	2017
GE_0 correlation	0.01602	0.01549	0.01754	0.01915	0.02117	0.02158	0.02275	0.02408	0.02530	0.02553	0.02424	0.02272	0.01986
GE_0 WINC	0.02487	0.02330	0.02348	0.02266	0.02457	0.02659	0.02711	0.02636	0.02378	0.02724	0.02709	0.02962	0.03004
GE_0 EDUCYRS	0.00301	0.00342	0.00338	0.00340	0.00338	0.00323	0.00331	0.00337	0.00356	0.00336	0.00327	0.00315	0.00322
GE_0 SIOPS	0.00476	0.00505	0.00509	0.00496	0.00454	0.00403	0.00420	0.00416	0.00430	0.00495	0.00515	0.00461	0.00472
GE_0 EMPLY	0.00840	0.00701	0.00716	0.00755	0.00795	0.00828	0.00952	0.01193	0.01370	0.01363	0.01325	0.01264	0.01125
GE_1 correlation	0.01388	0.01336	0.01512	0.01653	0.01826	0.01841	0.01946	0.02012	0.02109	0.02066	0.01959	0.01804	0.01577
GE_1 WINC	0.02474	0.02321	0.02341	0.02278	0.02456	0.02649	0.02691	0.02630	0.02369	0.02694	0.02665	0.02882	0.02935
GE_1 EDUCYRS	0.00291	0.00331	0.00327	0.00329	0.00328	0.00313	0.00318	0.00322	0.00340	0.00320	0.00311	0.00300	0.00307
GE_1 SIOP	0.00458	0.00493	0.00492	0.00482	0.00437	0.00387	0.00395	0.00388	0.00402	0.00456	0.00477	0.00431	0.00441
GE_1 EMPLY	0.00788	0.00653	0.00670	0.00706	0.00737	0.00768	0.00883	0.01098	0.01258	0.01251	0.01221	0.01172	0.01047

Note: Absolute contribution of each dimension to total inequality of the EA-13 using estimated weights and substitution elasticity ($\beta = 0.535$). *Source:* Author's calculation based on EU-SILC (2018).

Table A.7: Absolute contribution of factor shares to between-country inequality

	2005	2006	2007	2008	2009	2010	2011	2012	2013	2014	2015	2016	2017
GE_0^b	0.00447	0.00328	0.00359	0.00394	0.00367	0.00388	0.00502	0.00675	0.00769	0.00774	0.00741	0.00727	0.00651
GE_1^b	0.00433	0.00318	0.00349	0.00384	0.00357	0.00377	0.00486	0.00647	0.00738	0.00740	0.00711	0.00702	0.00631
GE_0^b correlation	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000
GE_1^b correlation	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000	0.00000
GE_0^b WINC	0.00322	0.00233	0.00270	0.00296	0.00271	0.00291	0.00399	0.00558	0.00633	0.00651	0.00625	0.00605	0.00539
GE_0^b EDUCYRS	0.00051	0.00031	0.00032	0.00036	0.00043	0.00039	0.00037	0.00041	0.00051	0.00047	0.00046	0.00047	0.00044
GE_0^b SIOPS	0.00072	0.00062	0.00056	0.00061	0.00048	0.00052	0.00058	0.00062	0.00068	0.00063	0.00060	0.00066	0.00062
GE_0^b EMPLY	0.00001	0.00001	0.00001	0.00002	0.00005	0.00007	0.00008	0.00014	0.00017	0.00014	0.00010	0.00009	0.00006
GE_1^b WINC	0.00308	0.00223	0.00260	0.00285	0.00261	0.00280	0.00384	0.00530	0.00603	0.00617	0.00596	0.00581	0.00518
GE_1^b EDUCYRS	0.00051	0.00031	0.00031	0.00036	0.00043	0.00039	0.00037	0.00041	0.00050	0.00046	0.00045	0.00047	0.00044
GE_1^b SIOP	0.00072	0.00062	0.00056	0.00061	0.00048	0.00052	0.00058	0.00062	0.00068	0.00062	0.00060	0.00066	0.00062
GE_1^b EMPLY	0.00001	0.00001	0.00001	0.00002	0.00005	0.00007	0.00008	0.00014	0.00017	0.00014	0.00010	0.00009	0.00006

Note: Absolute contribution of each dimension to between-country inequality in the EA-13 using estimated weights and substitution elasticity ($\beta = 0.535$). *Source:* Author's calculation based on EU-SILC (2018).

Table A.8: Between component of subgroup decomposition, EA-13

	income		MDEI	
	GE_0	GE_1	GE_0	GE_1
2005	0.00669	0.00639	0.00447	0.00433
	(0.00021)	(0.00019)	(0.00011)	(0.00011)
2006	0.00522	0.00500	0.00328	0.00318
	(0.00018)	(0.00017)	(0.00010)	(0.000095)
2007	0.00654	0.00634	0.00359	0.00349
	(0.00020)	(0.00019)	(0.000098)	(0.000093)
2008	0.00672	0.00627	0.00394	0.00384
	(0.00021)	(0.00018)	(0.000095)	(0.000091)
2009	0.00633	0.00588	0.00367	0.00357
	(0.00018)	(0.00016)	(0.00010)	(0.000095)
2010	0.00643	0.00602	0.00388	0.00377
	(0.00018)	(0.00016)	(0.00011)	(0.00010)
2011	0.00919	0.00855	0.00502	0.00486
	(0.00025)	(0.00023)	(0.00012)	(0.00012)
2012	0.0140	0.0126	0.00675	0.00647
	(0.00031)	(0.00027)	(0.00015)	(0.00014)
2013	0.0169	0.0152	0.00769	0.00738
	(0.00032)	(0.00029)	(0.00015)	(0.00014)
2014	0.0172	0.0156	0.00774	0.00740
	(0.00035)	(0.00032)	(0.00017)	(0.00016)
2015	0.0172	0.0157	0.00741	0.00711
	(0.00031)	(0.00029)	(0.00014)	(0.00014)
2016	0.0156	0.0143	0.00727	0.00702
	(0.00031)	(0.00029)	(0.00015)	(0.00015)
2017	0.0147	0.0133	0.00651	0.00631
	(0.00029)	(0.00027)	(0.00014)	(0.00014)

Note: Absolute contribution of between-country inequality to total income or multidimensional inequality measured by the GE indices ($\alpha = \{0, 1\}$) using estimated weights and substitution elasticity ($\beta = 0.535$). Bootstrapped standard errors in parentheses.
Source: Author's calculations based on EU-SILC (2018).

Table A.9: Country contribution to between component of subgroup decomposition, EA-13

	AT	BE	DE	ES	FI	FR	GR	IE	IT	LU	NL	PT	SI
2005	0.73	0.11	29.3	34.3	0.50	1.43	4.90	0.0028	2.35	1.60	7.75	31.1	0.12
	(0.018)	(0.0028)	(0.84)	(0.95)	(0.013)	(0.033)	(0.12)	(0.000068)	(0.059)	(0.044)	(0.22)	(0.90)	(0.0030)
2006	0.31	0.76	9.72	31.5	1.18	4.59	5.82	0.31	0.82	2.33	13.7	35.0	0.19
	(0.0092)	(0.022)	(0.28)	(1.04)	(0.033)	(0.13)	(0.16)	(0.010)	(0.024)	(0.066)	(0.45)	(1.26)	(0.0053)
2007	0.35	0.046	7.95	36.8	0.62	1.01	4.84	1.02	-1.35	2.24	17.1	30.1	0.11
	(0.0090)	(0.0014)	(0.24)	(1.10)	(0.018)	(0.027)	(0.13)	(0.031)	(-0.036)	(0.067)	(0.57)	(1.04)	(0.0032)
2008	0.58	0.072	1.88	22.6	1.55	15.5	7.53	0.14	2.35	1.48	16.3	34.6	0.26
	(0.014)	(0.0019)	(0.048)	(0.63)	(0.041)	(0.40)	(0.21)	(0.0036)	(0.061)	(0.040)	(0.45)	(1.14)	(0.0065)
2009	0.94	0.76	5.16	25.6	2.07	3.65	5.36	-0.087	1.11	1.60	19.9	36.2	0.43
	(0.023)	(0.021)	(0.14)	(0.72)	(0.050)	(0.10)	(0.16)	(-0.0024)	(0.030)	(0.042)	(0.60)	(1.19)	(0.012)
2010	1.88	0.98	5.19	28.4	1.74	5.88	5.25	-0.032	2.01	1.42	17.3	33.1	0.95
	(0.055)	(0.027)	(0.15)	(0.74)	(0.045)	(0.15)	(0.17)	(-0.00088)	(0.053)	(0.038)	(0.51)	(1.25)	(0.028)
2011	2.14	0.89	9.02	27.6	2.27	5.38	13.0	0.34	3.84	1.07	12.5	25.8	0.59
	(0.053)	(0.023)	(0.22)	(0.69)	(0.060)	(0.12)	(0.33)	(0.0079)	(0.093)	(0.025)	(0.31)	(0.69)	(0.014)
2012	1.53	0.85	12.6	29.4	1.84	7.49	19.3	-0.050	3.80	0.76	8.55	21.4	0.42
	(0.035)	(0.018)	(0.28)	(0.67)	(0.043)	(0.16)	(0.47)	(-0.0011)	(0.088)	(0.016)	(0.19)	(0.57)	(0.0087)
2013	1.34	1.61	15.6	31.6	1.81	6.41	19.8	0.0069	5.51	0.71	8.31	17.4	0.36
	(0.028)	(0.034)	(0.35)	(0.64)	(0.038)	(0.13)	(0.47)	(0.00014)	(0.12)	(0.014)	(0.19)	(0.44)	(0.0071)
2014	2.46	1.37	13.2	34.1	1.50	7.94	21.7	0.41	4.33	0.74	6.63	15.7	0.44
	(0.054)	(0.031)	(0.29)	(0.71)	(0.030)	(0.19)	(0.56)	(0.0087)	(0.10)	(0.016)	(0.14)	(0.38)	(0.0097)
2015	1.71	1.40	18.9	32.4	1.21	7.84	19.8	0.11	6.61	0.66	6.15	15.3	0.41
	(0.039)	(0.030)	(0.43)	(0.66)	(0.027)	(0.15)	(0.52)	(0.0022)	(0.14)	(0.014)	(0.14)	(0.32)	(0.0093)
2016	1.37	1.56	18.8	26.8	1.11	6.54	20.3	0.073	9.82	0.56	9.66	13.6	0.49
	(0.028)	(0.034)	(0.44)	(0.44)	(0.024)	(0.14)	(0.50)	(0.0015)	(0.21)	(0.011)	(0.20)	(0.30)	(0.0097)
2017	1.82	0.85	20.1	22.6	0.87	3.80	21.4	-0.058	8.39	0.68	11.2	14.7	0.45
	(0.037)	(0.017)	(0.45)	(0.43)	(0.017)	(0.079)	(0.53)	(-0.0011)	(0.19)	(0.014)	(0.26)	(0.34)	(0.0090)

Note: Percentage change of between-country component of multidimensional inequality when replacing individual outcomes for the respective country with average outcomes of all other countries. Measured by the GE_0 index using estimated weights and substitution elasticity ($\beta = 0.535$). *Source:* Author's calculations based on EU-SILC (2018).

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