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The Impact of Economic Integration on Income Inequality in the EU: A Panel Data Analysis of the EU Members from 2002-2020

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Abstract This paper investigates the relation between economic integration and income inequality for the 27 current EU members from 2002 to 2020. The Gini coefficient is the dependent variable in a panel data regression analysis, and is expressed as a function of indicators related to economic integration (total trade, intra-EU trade, foreign direct investment (FDI) inflow and outflow) and a set of control variables. Although much research points to a significant relationship between inequality and integration, there is no consensus on the sign and magnitude of the effect. In this paper, four random effects panel models are estimated with robust standard errors to uncover this relationship. The results show that in the context of the EU, intra-EU trade is associated with decreased income inequality, while overall trade and FDI in- or outflow seem to have no direct effect on the income gap. In addition, the level of economic development seems to moderate the effect of intra-EU trade: for countries with a below-average gross domestic product (GDP) per capita, the reducing effect of integration on Gini is weaker than for richer countries. Additionally, market capitalisation, the presence of natural resources and government spending on social benefits are associated with reduced inequality, while unemployment and population size seem to drive income disparities. These results are consistent with a major part of the existing literature and lead to interesting conclusions for policy makers. The originality of this work differentiating it from prior research is twofold: (1) the region of examination is the EU, which is not often the subject of similar analyses, and (2) an interaction effect is examined that differs from the conventional measures for less developed economies.

Keywords: random effects panel models, economic integration, income inequality, European Union

JEL Classifications: C33, F15, D63

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I. Introduction

Inequality in the income distribution is a key feature of almost all economies in the world (Mankiw & Taylor, 2020), and has been rising for many countries across the globe over the

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past two decades (Allison et al., 2014; Dollar, 2005, UNDP, 2019). Moreover, in spite of the fact that Western European nations are generally associated with lower levels of economic inequality (Deininger & Squire, 1996), income disparities have increased in many of these countries as well (Hoffmeister, 2009; Huber & Stephens, 2014; Kuhn et al., 2016; UNDP, 2019). It is therefore vital for researchers and policy makers to understand the mechanisms that drive income inequality, if they aspire to alleviate it. On top of that, the inclusion of reducing inequality as one of the Sustainable Development Goals of the UN (United Nations, 2022) underscores its global relevance.

The accelerating pace of global economic integration, broadly understood as the continuous reduction of trade and investment costs (Akhmetova et al., 2017), drives efficiency and promotes economic growth around the world (Dreher, 2006). Despite having a positive impact on per capita income in many countries, the real effect of economic integration on the income distribution remains under debate (Tung et al., 2020). Even though fighting inequality is listed as one of the core values of the European integration project (European Union, 2022), many studies find a positive connection between economic integration and inequality of income (Alili & Adnett, 2018; Beckfield, 2009; Tridico, 2017). Nevertheless, the sign of this relation is not clear, as another group of works finds that economic integration is linked to reduced inequality (Furceri & Ostry, 2019; Jaumotte et al., 2013; Mundell, 1957; Ravinthirakumaran & Navaratnam, 2018; Tian et al., 2009). Others claim however that the connection is more complex, in the sense that whether integration augments or mitigates inequality depends on other moderating factors such as economic development (Cesaroni et al., 2019; Couto, 2018; Dorn et al., 2022; Ean et al., 2020; Tung et al., 2020). This ambiguity results in a lack of consensus among economists on the precise nature of the relationship (Bertola, 2010). Regardless, an essential characteristic of economic integration is the quest for competitiveness and economic efficiency, which has a large impact on the labour markets and consequently inequality (Baldwin & Wyplosz, 2020).

The purpose of this paper is to investigate this gap in the existing literature by examining the precise relationship between economic integration and income inequality in the context of the European Union (EU). The paper is divided into two main segments. First, a literature review is conducted with the aim of giving an extensive overview of the relevant concepts and previous research on the relationship between economic integration and income inequality. Thereafter, an empirical analysis is performed in order to test how these two concepts are connected in the 27 current EU members from 2002 to 2020. This work aims to contribute to the existing literature by inquiring about the relationship between economic integration and income inequality in the EU, and in addition by exploring the moderating role of economic development on this relationship. The principal findings of this study indicate the presence of a negative relation between intra-EU trade and income inequality. Furthermore, it is observed that this mitigating effect on inequality is less pronounced in countries with below-average

GDP per capita compared to those with GDP per capita exceeding the EU27 average. No substantial evidence for the significant effects of total trade, foreign direct investment (FDI) inflow or outflow on income inequality is found. Finally, the results are discussed and suggestions for further research are provided.

II. Literature Review

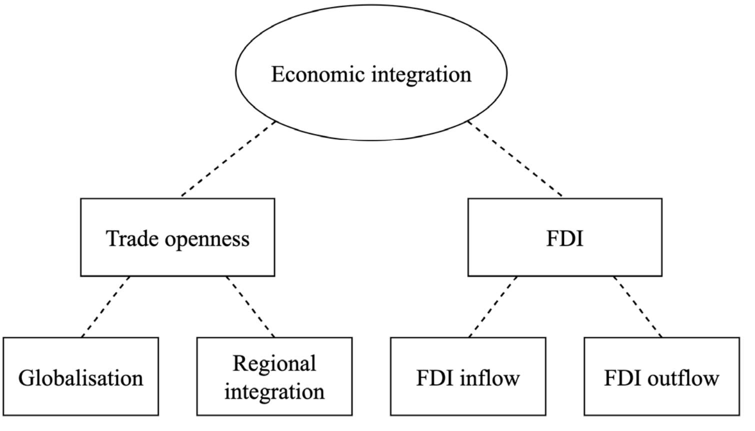
Piotrowska (2008) defines economic integration as "the expansion of markets from the national to the regional or world level". This expansion can be divided into two channels: regional integration and globalisation. As stated by Shangquan (2000), economic globalisation relates to the rising interdependence between global economies due to increased international trade, financial flows and the spread of technology. The concept can be viewed as a multi-dimensional phenomenon, caused by a variety of processes (Asteriou et al., 2014). Previous research has identified that trade liberalisation, technological innovation, migration and financial openness are some of the main drivers of globalisation (Begg et al., 2008; Jaumotte et al., 2013; Kose et al., 2009; Mills, 2009). On the other hand, regional integration alludes to the same concept but at a smaller scale, and often entails harmonisation of economic policies with negotiated regions (Beckfield, 2006).

However, there exists no precise definition of economic integration that is widely accepted in the literature (Brahmbhatt, 1998), and the distinction between regional and global integration is not always made by researchers. On top of that, a multitude of types of integration can be identified. Next to economic integration, political, monetary and market integration are distinguished. Although these different types are tightly connected with each other, this paper focuses specifically on the role of economic integration which should not be confused with these other forms.

As a consequence of the great complexity involved in its different types and stages, economic integration can be viewed as a latent construct that in itself is not directly observable. One way to measure economic integration is to assess trade barriers such as the tariffs imposed by governments. However, these measures are often inaccurate and hard to quantify, which results in limited data availability (Brahmbhatt, 1998). Therefore, many researchers instead use outcomes of integration in order to capture the concept. Popular outcome measures are international patterns of prices and trade flows, such as trade intensity, capital flows, foreign direct investment (FDI), and the flow of people (Preepremmote et al., 2018). One major difficulty with using outcome measures is, however, to disentangle the separate effects of other complex economic factors on these indicators. Researchers should therefore include the right control variables in their models to investigate the intrinsic effects of the indicators on each other.

Nevertheless, many studies that aim to map the relation with inequality commonly use trade openness and FDI as proxy variables for economic integration (Arribas et al., 2007; Tian et al., 2009). Both trade openness and FDI can be further divided into separate subaspects. As discussed earlier, trade openness consists of globalisation and regional integration, and FDI is composed of FDI inflow and outflow. These subdimensions are visualised in Figure 1, in which the round shape indicates that economic integration is an unobservable, latent concept, while the rectangular forms represent variables that are directly observable.

Figure 1. Conceptual representation of economic integration and its subdimensions



Economists often use trade openness or trade intensity as a proxy measure for economic integration (Baek & Shi, 2016), which is often defined as the amount of trade with cross-border economies divided by GDP. A major advantage of using trade openness is that it can differentiate between regional integration (intra-EU trade for example) and globalisation (overall international trade). The difficulty lies, however, in the fact that some analyses distinguish between these two forms of integration while others do not.

Several studies report an incremental relationship between trade intensity and income inequality (Alili & Adnett, 2018; Beckfield, 2009; Tridico, 2017). An explanation for this is that intensified trade is often associated with a rise in the relative demand for highly skilled labour, which causes a greater skill premium (Neagu et al., 2016). On top of that, international trade could also intensify income inequality as a consequence of disparities in returns to education and skills (Stiglitz, 1998). Analogously, research from Beckfield (2006) confirms that regional integration causes workers to be exposed to more international competition, which leads to higher wage inequality. In addition, the free movement of foreign factors of production creates an enormous labour pool in which domestic workers can more easily be substituted (Busemeyer & Tober, 2015). These factors can be external labour forces as well as non-labour factors, such as capital in the form of technology or new production methods.

However, the empirical evidence on the precise relation between the two variables is mixed (Baek & Shi, 2016; Winters et al., 2004). While some studies claim that globalisation only increases inequality within but not between nations (Flaherty & Rogowski, 2021; Hung, 2021), others state that it has greatly magnified both forms of inequality (Mazur, 2000). The decline in income dispersion between countries is often referred to as sigma convergence (Monfort, 2008), which is associated with economic integration according to Beckfield (2009). Evidence from Central and Eastern European (CEE) countries shows that although globalisation significantly increases income inequality among these countries, regional integration does not contribute to changes in income distribution (Piotrowska, 2008). The two main reasons for this mentioned in the study are (1) changes in the employment structure and (2) a rise in wage competition among workers. In addition, some researchers even argue that economic integration has narrowed the income gap worldwide because it has improved overall incomes (Mills, 2009). Likewise, in spite of the aforementioned amplifying effect of economic integration on income inequality, another group of studies reports the opposing finding that economic integration would decrease within-country income inequality (Furceri & Ostry, 2019; Jaumotte et al., 2013; Mundell, 1957; Ravinthirakumaran & Navaratnam, 2018; Tian et al., 2009).

The Heckscher-Ohlin model is the conventional theoretical framework that is used to study the distributional effects of trade on market outcomes (Dorn et al., 2022). It offers an explanation for how economic integration influences inequality by looking at productivity differences and the relative factor endowment in countries (Ohlin, 1933). Specifically, the model predicts that nations will export goods that are made with their abundant and cheap production factors, and import goods that use the countries' scarce factors of production. Therefore, industrialised or developed economies will tend to specialise in the production of goods and services which are skill- and capital-intensive, while developing economies will focus on the production of products which require less skilled labour (Baek & Shi, 2016). This leads to an increased demand for unskilled labour in developing countries, which raises their wages and thus narrows the gap between skilled and unskilled labour. The opposite is true for industrialised economies, in which the relative demand for low-skilled labour drops and consequently income accumulates towards high-skilled workers (Neagu et al., 2016). Owners of the abundant production factors in a country will therefore gain from trade openness, while the owners of scarce factors will be disadvantaged. This phenomenon is known as the skill-premium hypothesis of the Stolper-Samuelson theorem (Dorn et al., 2022; Stolper & Samuelson, 1941). Essentially, this means that the effect of economic integration on income inequality is moderated by the level of economic development of the countries involved. Other frameworks such as New Trade Theory contend that trade will first lead to more inequality, and then later to smaller income differentials as economies further integrate (Krugman & Venables, 1995). However, New Trade Theory is less explicit than the Heckscher-Ohlin model about the direct link between integration and within-country income inequality, since it focuses more on the effects of trade on inequality between nations in relation to costs of trade.

Much empirical evidence confirms the theoretical prediction that economic development moderates the relationship between economic integration and income inequality. For developing countries, higher trade intensity is associated with lower income inequality, and the reverse is true for industrialised nations (Cesaroni et al., 2019; Dorn et al., 2022; Ean et al., 2020; Tung et al., 2020). Nevertheless, these theoretical predictions do not always hold because the assumptions are not always met (Mankiw & Taylor, 2020). For instance, Daumal (2013) finds that trade openness increases inequality in some developing countries, while it decreases inequality in others. Other works even claim that globalisation drives income inequality for all emerging economies (Kahai & Simmons, 2005; Meschi & Vivarelli, 2009). Similarly, studies by Ametoglo et al. (2018) and Asteriou et al. (2014) find that integration reduces inequality in advanced economies, while it increases inequality in developing countries. An additional layer of complexity arises from the lack of precise criteria for classifying countries into developing and industrialised economies. For instance, most countries are neither completely developing nor fully industrialised, but are somewhere in between. Although the International Monetary Fund (IMF) and the UN have precise definitions of developing economies, these are not necessarily relevant for the EU nations. It could for instance be more sensible to classify them into low- versus high-wage countries, or to use scalar measures such as gross domestic product (GDP) per capita to measure the interaction effect. Despite this, economists assert that trade is only a contributing factor to the increasing income inequality around the world (Mankiw & Taylor, 2020). Ultimately, a last collection of analyses finds no significant relationship between trade openness and income inequality (Akyuz et al., 2022; Ali et al., 2015; Edwards, 1997; Lee & Vivarelli, 2006; Li et al., 1998).

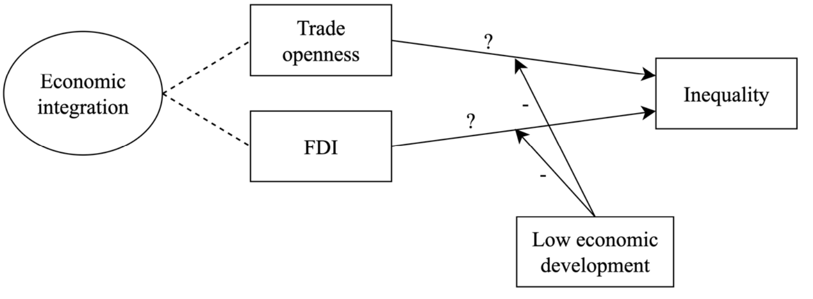
Next to trade intensity, the other major subdimension of economic integration is foreign direct investment. While much research on the impact of FDI on income inequality has been done, the theoretical framework behind this relationship is still unclear (Couto, 2018; Herzer & Nunnenkamp, 2013). Inward FDI can be defined as the value of foreign investors' equity in and net loans to enterprises in an economy (OECD, 2021), and is associated with economic growth in developing economies (de Mello Jr., 1997). Although FDI creates strong and long-lasting ties between economies (OECD, 2021), it has been linked to increased inequality of income within nations (Beckfield, 2009; Feenstra & Hanson, 1997; Mahutga & Bandelj, 2008; Mugeni, 2015; Neagu et al., 2016; Tsai, 1995; Tung et al., 2020).

However, preceding analyses have also shown that the impact of FDI on income inequality could be dependent on economic development, which is analogous to the interaction effect described above. Couto (2018) reports that FDI tends to decrease the Gini coefficient among low-income countries, but increases income inequality for middle- and high-income economies. Since FDI outflows from advanced industrialised countries tend to be concentrated in industries with low-skilled labour in the partnering country (Lee, 1996), low-skilled wages in the partnering economy will increase which reduces inequality of income (Couto, 2018; Deng & Lin, 2013; Tian et al., 2009). Conversely, from the standpoint of the industrialised country, a fast increase in FDI outflows often lessens the demand for this low-skilled labour and as a consequence widens its income gap (Leamer, 1996; Wood, 1994). Similarly, research from Ametoglo et al. (2018) shows that increased FDI inflow reduces income inequality in the Economic Community of West African States (ECOWAS). These results are consistent with the international trade patterns predicted by the Heckscher-Ohlin model and the Stolper-Samuelson theorem. The same reasoning as for trade openness in section 2 applies here: the accumulation of FDI in low-skilled labour industries would increase the demand for that labour and its productivity, and consequently the wages (Mundell, 1957). This finding is particularly important for developing economies, considering that they are often heavily dependent on inward FDI to drive economic growth (Couto, 2018).

Others come to the opposite conclusion that FDI increases inequality in developing economies, while it decreases inequality in industrialised economies (Basu & Guariglia, 2007; Daumal, 2013; Figini & Görg, 2011; Tsai, 1995; Wu & Hsu, 2012). The argument here is that FDI mainly increases demand for skilled, and not for unskilled labour, which consequently expands the wage gap between the two groups (Feenstra & Hanson, 1997). On top of that, this phenomenon could be explained by the fact that FDI stimulates the use of advanced technologies, which requires more highly skilled personnel (Nguyen et al., 2017). Consequently, the rise in demand for skilled labour might reduce the need for and accordingly the wages of low-skilled workers. Other studies add that FDI on average widens the income gap in the short-run, but reduces inequality in the long-run (Chintrakarn et al., 2012; Herzer & Nunnenkamp, 2013; Ravinthirakumaran & Navaratnam, 2018). Especially for developing countries, FDI inflow allows domestic companies to compete with multinational corporations, resulting in a more equal income distribution in the long-run (Ravinthirakumaran & Navaratnam, 2018). In contrast, some studies challenge this finding and claim that FDI reduces inequality both in the short and long term (Ucal et al., 2016). Also in this case, several works find no significant relationship between the two variables (Bhandari, 2007; Franco & Gerussi, 2013; Sylwester, 2005; te Velde & Morrissey, 2004).

On the basis of the literature discussed above, the following hypotheses can be formulated. Economic integration, measured by trade intensity and FDI, significantly influences income inequality. However, due to the mixed evidence found in the literature, the sign of the expected relationship is not specified. In addition to that, the relationship between economic integration and inequality could very well be dependent on the level of economic development of a given country. In less developed economies, it is expected that trade openness and FDI decrease inequality, while they would increase inequality in industrialised countries. This interaction effect is one of the central theoretical mechanisms in this analysis and is schematically represented in Figure 2.

Figure 2. Interaction between economic integration and economic development



Again, the round shape of economic integration indicates that it is a latent concept, comprised of two sub-aspects: trade openness and FDI (indicated by the dashed lines). The solid arrows portray the hypothesised directional relationship between the observable variables. In addition, it is expected that a set of control variables have a significant impact on inequality as well. These hypotheses are formalised below and tested against empirical data.

III. Methodology

A. Data and model

To test the hypotheses described above, panel econometric models are often used since they take time into account and control for individual heterogeneity, which increases the efficiency of econometric estimation (Ravinthirakumaran & Navaratnam, 2018). Nonetheless, as a result of the extremely large number of possible model configurations, model uncertainty is a widespread problem in econometric research (Furceri & Ostry, 2019).

In panel analysis, three types of models are commonly considered: pooled ordinary least squares (OLS), fixed effects (FE) and random effects (RE) models. The key difference between these models is in how they assume the individual error components behave. The pooled OLS model is a homogeneous model that assumes that the unobservable characteristics are constant across countries, and that there are no unobservable country-specific effects. This however is a rather strong assumption, which implies that all the observations within groups are independent of each other. For this reason, pooled OLS is seldom an appropriate model for panel data and is merely used as a baseline to compare other models to.

Besides that, two types of heterogeneous panel data models exist. They allow for the regression coefficients to vary across countries, meaning that there can be country-specific effects. The first type are fixed effects models, which focus only on variation within individual observations. In other words, the fixed effects model accounts for the time-invariant unobserved country-specific covariates by including country-specific intercepts. The FE model can be represented as:

$$y_{it} = \alpha + \boldsymbol{\beta}' \boldsymbol{X}_{it} + \epsilon_{it} \tag{1}$$

where α is a scalar representing an intercept that is constant across all countries and time periods, the vector $\boldsymbol{X}_{it}$ denotes the observable characteristics of the countries which may be constant or vary across time and $\boldsymbol{\beta}$ are the effects of $\boldsymbol{X}$ on y which are constant across all countries and time periods. The error component consists of three terms: $\epsilon_{it} = \mu_i + \lambda_t + v_{it}$. Here, μ_i is the individual-specific error component which captures any unobserved effects that are different across countries but are fixed across time, λ_t are the unobservable time-specific effects and v_{it} is the remainder stochastic disturbance term which is independent and identically distributed ($v_{it} \sim IID(0, \sigma_v^2)$). This means that μ_i and λ_t represent country- and time-specific intercepts respectively. When λ_t is assumed to be 0, the model is referred to as a one-way fixed effects model, and when $\lambda_t \rangle 0$, it is a two-way fixed effects model. The fixed effects models assume that both μ_i and λ_t are fixed parameters that have to be estimated, while v_{it} is stochastic. Additionally, they assume that the individual-specific effects (μ_i) are correlated with the observed characteristics ($\boldsymbol{X}_{it}$), thus $cov(\boldsymbol{X}_{it}, \mu_i) \neq 0$ (Baltagi, 2021).

In contrast, the random effects model assumes that the units of analysis have different intercepts, which follow a certain distribution. This implies that μ_i and λ_t are not fixed parameters, but are assumed to be randomly distributed: $\mu_i \sim IID(0, \sigma_\mu^2)$, $\lambda_t \sim IID(0, \sigma_\lambda^2)$ and $v_{it} \sim IID(0, \sigma_v^2)$. Consequently, countries do not have fixed country-specific intercepts, but rather these intercepts follow a given distribution. On top of that, $\boldsymbol{X}_{it}$ is independent of μ_i, λ_t and v_{it} for all i and t, thus $cov(\boldsymbol{X}_{it}, \mu_i) = 0$ (Baltagi, 2021). One major advantage of using random over fixed effects is that it is more efficient than fixed effects estimation, given that the assumptions are met. However, an endogeneity problem might arise because there might be nonzero correlation between the independent variables and the variance of the random intercept. This assumption can be checked by performing the Hausman test.

In light of the above, multiple models are estimated to assess the relationship between economic integration and income inequality. In spite of the fact that Gini has its limitations as a measure of inequality, it is a widely used concept in empirical economic research (Alili &

Adnett, 2018). Therefore, with the aim of augmenting comparability with other research, Gini is included as the dependent variable in this study. For ease of interpretation, it is expressed as a percentage, ranging from 0 % representing a perfectly equal income distribution, to 100 % portraying a completely unequal society. Below, a general panel model is shown that expresses Gini as a function of indicators related to economic integration (I_{it}, which represents overall trade, intra-EU trade, inward FDI or outward FDI), and a set of control variables that are common in the literature ($\boldsymbol{\Theta}_{it}$). From both a theoretical perspective (cf. Aizenman & Noy, 2006; Chen et al., 2020; Gopinath & Echeverria, 2004), as well as from an empirical standpoint (see section B below), there is potential for correlations between indicators of trade and FDI. Therefore, to avoid potential issues of multicollinearity by combining all indicators into one model, the general model is estimated four times, each time using only one indicator of economic integration. In addition, the natural logarithm is taken of the Gini coefficient and of certain independent variables (total trade, intra-EU trade, population, social benefits and technology exports). This is done to make skewed (log-normal) distributions approach normality, to account for non-linear relationships and to reduce the variance of the variables which increases precision (Gelman & Hill, 2017).

The countries of interest in the context of this paper are the 27 current EU member states, which generally tend to be more developed compared to other regions in the world (United Nations, 2019). Because of that, the models include interaction terms between the integration indicators and a dummy variable that takes a value of unity if the GDP per capita of the country is lower than the EU average for a given year, indicating that its economic development was below-average (represented by $I_{it} \cdot D_{it}$). The vector of control variables contains unemployment (%), population size, market capitalisation (% of GDP), natural resources (% of GDP), social benefits (% of GDP), and technology exports (% of total exports). The general model can thus be expressed by the following equation:

$$\ln(Gini)_{it} = \alpha + \beta I_{it} + \gamma I_{it} \cdot D_{it} + \boldsymbol{\delta}' \boldsymbol{\Theta}_{it} + \mu_i + \lambda_t + v_{it} \tag{2}$$

with $i \in \{1, \ldots, N\}$, $t \in \{1, \ldots, T\}$ and where $\ln(Gini)_{it}$ is the natural logarithm of the Gini coefficient for country i at time t. The overall Y-intercept is denoted by the constant α, and β is the coefficient of the indicator of economic integration (I_{it}, i.e. overall trade, intra-EU trade, FDI inflow or FDI outflow). The coefficient γ denotes the interaction effect between I_{it} and the low-development dummy variable D_{it} that equals 1 if the country has a below-average GDP per capita (in current USD) in year t (relative to the EU27 average), and 0 if the GDP is above-average. Furthermore, the model contains a vector of control variables $\boldsymbol{\Theta}_{it}$ with coefficients $\boldsymbol{\delta}'$. The residual term consists of an unobserved country-dependent error μ_i, an unobserved time-dependent error λ_t, and the idiosyncratic error term which is assumed to be $v_{it} \sim N(0, \sigma_v^2)$ for all i and all t. Based on equation (2), the hypotheses formulated above can be formalised as follows:

$H_{a1}: \beta \neq 0$, i.e. economic integration (in terms of overall trade, intra-EU trade, FDI in- and outflow) significantly affects Gini,

$H_{a2}: \gamma < 0$, i.e. in countries with below-average GDP per capita, economic integration decreases Gini,

$H_{a3}: \delta \neq 0$, i.e. the control variables have a significant impact on Gini.

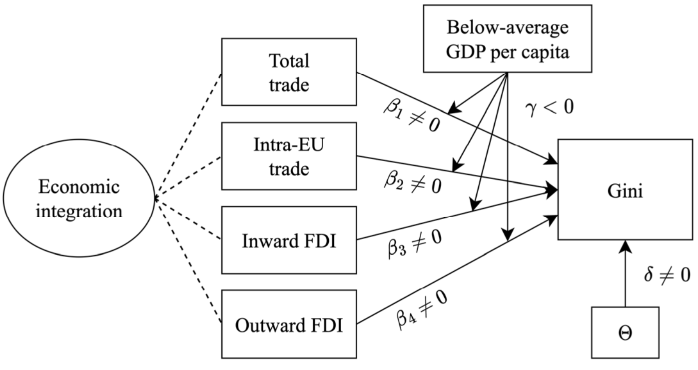
As noted earlier, due to the inconsistent empirical evidence on the direction of the relationship between economic integration and income inequality, alternative hypothesis H_{a1} is two-sided. The anticipated effects are summarised in Table 1 below, and Figure 3 visually portrays these expected relationships.

Table 1. *Variables Used in the Regressions and Their Expected Effects*

	Variable	Expected effect
I	Total trade (% of GDP)	$\beta_1 \neq 0$
	Intra-EU trade (% of total trade)	$\beta_2 \neq 0$
	Inward FDI (% of GDP)	$\beta_3 \neq 0$
	Outward FDI (% of GDP)	$\beta_4 \neq 0$
D	Below-average GDP per capita (dummy)	$\gamma_1 > 0$
I×*D*	Interaction economic integration & development	$\gamma_2 < 0$
Θ	Unemployment (%)	$\delta_1 > 0$
	Population size	$\delta_2 \neq 0$
	Market capitalisation (% of GDP)	$\delta_3 > 0$
	Natural resources (% of GDP)	$\delta_4 < 0$
	Social benefits (% of GDP)	$\delta_5 < 0$
	Tech exports (% of total exports)	$\delta_6 \neq 0$

Note. expected effect < 0 means that it is expected that the variable decreases inequality; > 0 that it increases inequality; and ≠ 0 that the relationship is ambiguous or that its direction is not specified.

Figure 3. Conceptual scheme of expected relations



B. Data exploration and descriptive statistics

Now that the general model is specified, the analysis can be conducted using empirical evidence. Data from the World Bank (2022) and Eurostat (2022) are used to perform the analyses and make the plots below. Table 2 lists all variable definitions and variable sources.

Table 2. *Variable Definitions and Data Sources*

Variable	Measurement definition	Source
Gini (%)	The relationship of cumulative shares of the population arranged according to the level of equivalised disposable income, to the cumulative share of the equivalised total disposable income received by them.	Eurostat
Total trade (% of GDP)	Total trade as % of GDP.	World Bank
Intra-EU trade (% of total trade)	Share of imports plus exports to other EU members as % of total trade.	Eurostat
Inward FDI (% of GDP)	Net inflows of investment to acquire a lasting management interest (10 % or more of voting stock) in an enterprise operating in an economy other than that of the investor. It is the sum of equity capital, reinvestment of earnings, other long term capital and short term capital as shown in the balance of payments.	World Bank
Outward FDI (% of GDP)	Net outflows (idem supra).	World Bank
Below-Average GDP per capita (dummy)	Dummy that equals 1 if the country has a below-average GDP per capita (in current USD) for a given year (compared to the EU27 average for that year).	Calculated from the World Bank
Interaction integration indicator & economic development	Economic integration indicator (e.g. total trade, intra-EU trade, FDI inflow or FDI outflow) × the below-average GDP per capita dummy.	Calculated
Unemployment (%)	The share of the labour force that is without work but available for and seeking employment.	World Bank
Population size	All residents of the country regardless of legal status or citizenship (midyear estimate).	World Bank

Table 2. *Continued*

Variable	Measurement definition	Source
Market capitalisation (% of GDP)	The share price times the number of shares outstanding (including their several classes for listed domestic companies as a % of GDP). Investment funds, unit trusts, and companies whose only business goal is to hold shares of other listed companies are excluded.	World Bank
Natural resources (% of GDP)	The sum of oil rents, natural gas rents, coal rents (hard and soft), mineral rents, and forest rents as a % of GDP.	World Bank
Social benefits (% of GDP)	Government expenditure on social benefits as % of GDP.	Eurostat
Tech exports (% of total exports)	Exports of products with high R&D intensity, such as in aerospace, computers, pharmaceuticals, scientific instruments, and electrical machinery (i.e. high-technology) as a % of total manufactured exports.	World Bank

The appendix contains the descriptive statistics of all the relevant variables, as well as plots illustrating the bivariate relationships between them, and density plots to visualise their distributions. Figure A1 shows the overall relationships, and Figure A2 makes a distinction between the two categories of the interaction variable (below- vs. above-average GDP per capita). Both plots are briefly discussed in the appendix. Figure 4 below displays the Gini coefficient for the 27 EU members in 2020, and Figure 5 shows its temporal patterns over time for certain selected countries.

Figure 4. Gini coefficients for the 27 EU members in 2020

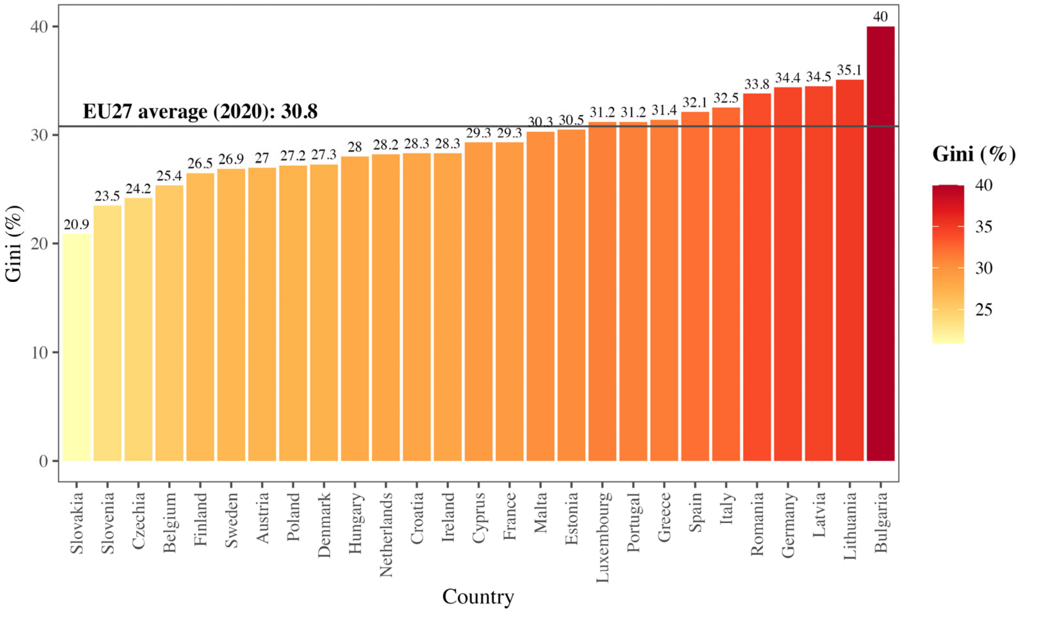
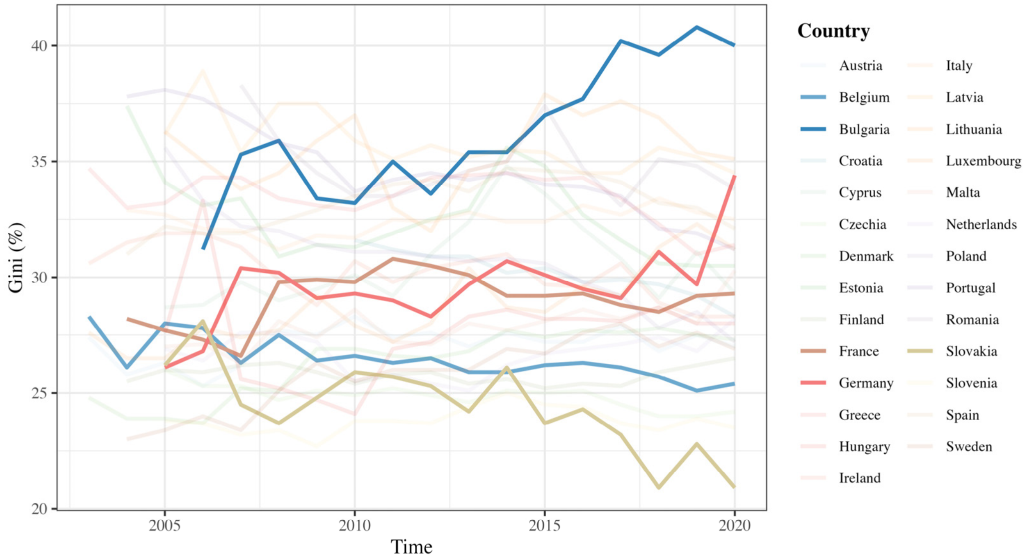


Figure 5. Temporal patterns of Gini for selected EU members from 2002-2020



As shown in the bar plot in Figure 4, the most recent data points out that income inequality is by far the highest in Bulgaria, followed by Lithuania and Latvia, while Slovakia, Slovenia and Czechia are the countries with the lowest Gini coefficients for 2020. Figure 5 illustrates the evolution of some selected member states over time (2002-2020), showing the steep rise for Bulgaria starting around 2012. Also, Germany has shown a fast increase in Gini over the last few years. However, no clear increasing or decreasing pattern is present for the EU27 as a whole.

For the whole period of analysis, the Gini coefficient ranges between 20.90 % and 40.80 %, with a mean of 29.68 % and a standard deviation of 3.98 %. As seen in Table A1 and Figure A1 in the appendix, only a few variables are unimodally and rather normally distributed (Gini and social benefits), while the others are very skewed. This justifies the logarithmic transformations performed on them when included in the models. Figure 6 displays the correlation matrix of all the variables that are used and shows which variables covary together. No correlations are higher than 0.5 (in absolute value), except for the correlations between intra-EU trade and population size, and between inward and outward FDI which are 0.68 and 0.81 respectively. This could be of concern regarding multicollinearity, which is further discussed below.

Figure 6. Correlation matrix

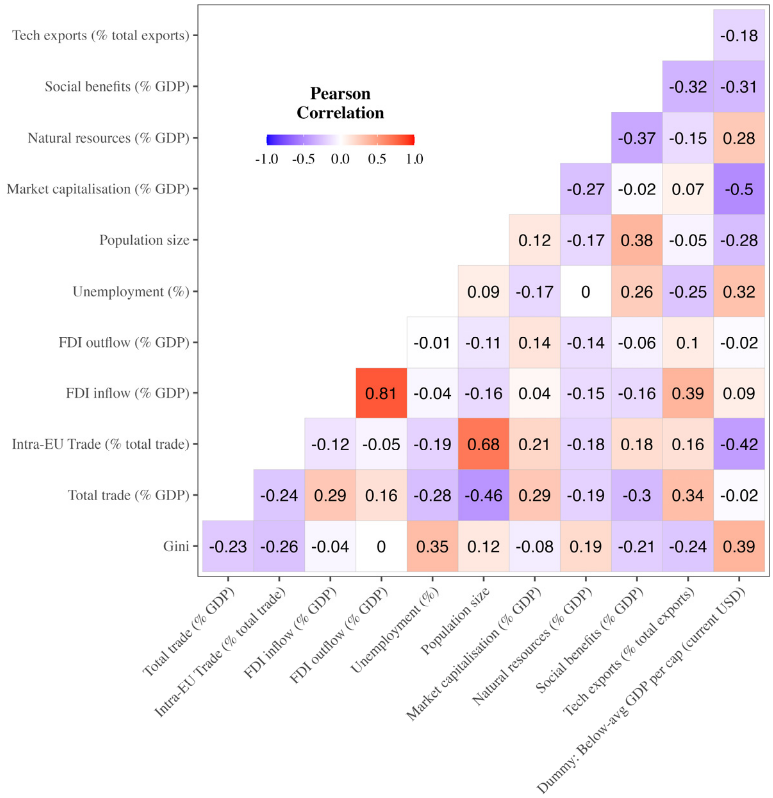
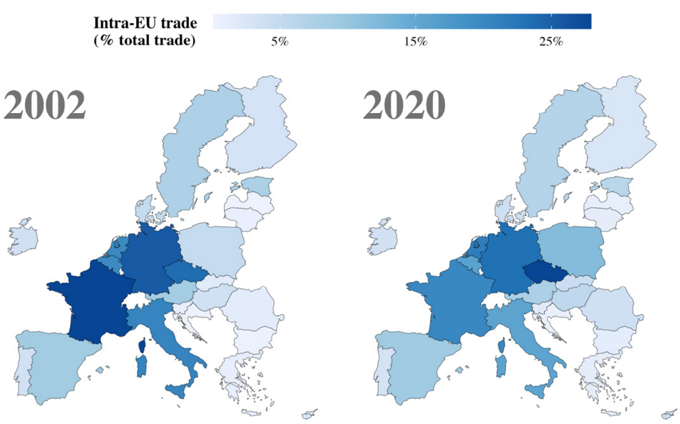


Figure 7 shows the geographical distribution of regional (intra-EU) trade intensity in 2002 versus in 2020. A darker shade of blue represents a greater amount of intra-EU trade as a percentage of total trade.

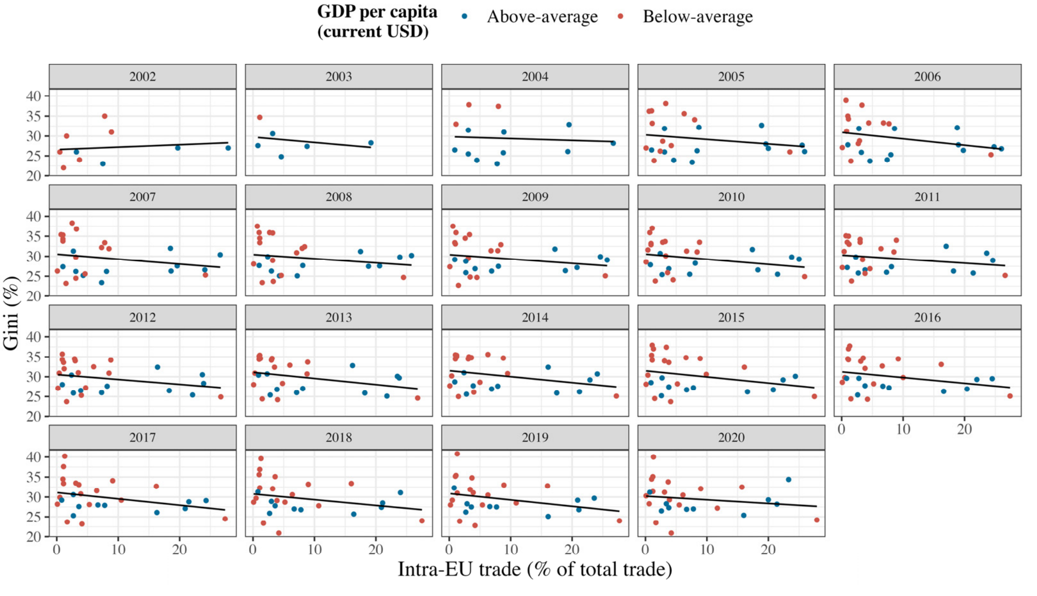
Figure 7. Geographical distribution of intra-EU trade in 2002 versus 2020



While the geographical pattern of intra-EU trade has changed somewhat over time, it remains clear that the countries which engage most in trade with other EU member states are The Czech Republic, Germany, The Netherlands, France and Belgium, relative to their total trade. Notably, these countries are situated at the centre of the EU's geographical layout, aligning with the predictions outlined in gravity models of international trade. It should however be noted that the intra-EU trade variable represents the proportion of total trade conducted with other EU members.

Next, the relationship between intra-EU trade and the Gini coefficient for each year in the sample (2002 to 2020) is visualised in Figure 8. The colour of the dots represents the categories of the interaction dummy: countries with a below-average GDP per capita are shown in red, and those with above-average GDP per capita are in blue.

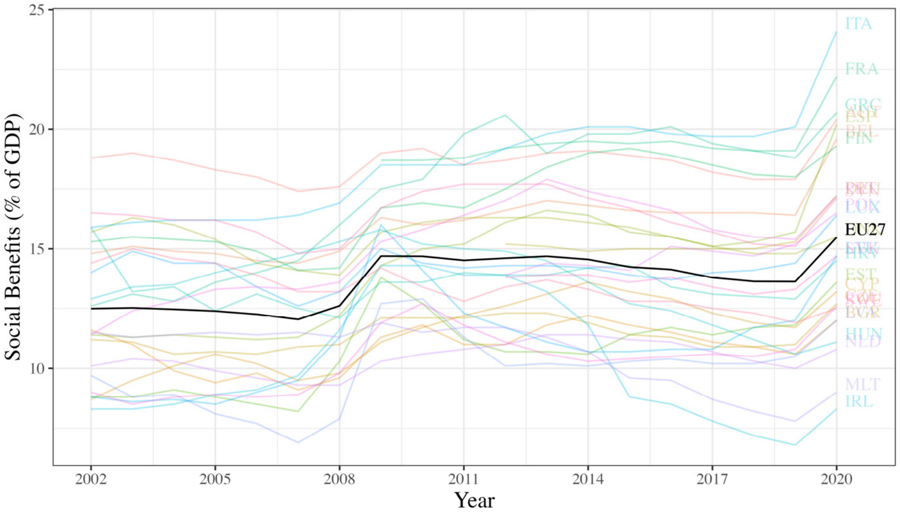
Figure 8. Evolution of the relation between intra-EU trade and Gini over time



The plot seems to show a slightly negative interrelationship between intra-EU trade and inequality for almost all years in the sample. Countries with below-average GDP per capita are concentrated on the left side of the plots, indicating less intra-EU trade as % of total trade. Additionally, these countries generally have higher Gini values than the EU members with higher GDP per capita. The question remains whether this apparent negative relationship between intra-EU trade and income inequality will persist after controlling for the relevant covariates in the panel regressions below.

A last interesting descriptive statistic is that government spending as % of GDP has on average slightly increased over the years. This is illustrated by the rise in social benefits of the EU27-average from 2002 to 2020 in Figure 9, and could be an indication of the willingness of the EU members to combat inequality. A notable pattern is that the average social spending of governments has risen during times of crisis, such as around the 2008 financial crisis and around 2020 during the COVID-19 pandemic.

Figure 9. The rise of government expenditure on social benefits over time



C. Panel regression

As discussed, multiple models are estimated, each containing a different indicator of economic integration. The first model uses total trade (as % of GDP) as a measure of overall trade integration, while the second model uses intra-EU trade (as % of total trade) to capture regional trade integration. To measure the effect of financial integration on inequality, models three and four include FDI inflow and FDI outflow (both as % of GDP), respectively. Besides this, all four models include interaction effects between the integration variables and a dummy variable indicating below-average GDP per capita (in current USD). This allows to explicitly test whether the effect of economic integration on income inequality is moderated by the level of economic development of the country.

As noted earlier, three model specifications should be tested in panel analyses: pooled OLS, fixed effects and random effects models. In order to decide which of these specifications is appropriate in the context of this analysis, the following tests are performed on the four models described above. First, an F-test is performed to guide the choice between the pooled and the fixed effects panel model. This test reaches significance for all four models (see Table A2 in the appendix), which favours the alternative hypothesis that the fixed effects model is preferred over the pooled OLS (Croissant & Millo, 2018). Thereafter, a Lagrange multiplier test is executed to determine whether there are individual and/or time effects based on the results of the pooled model (Breusch & Pagan, 1980). The tests are significant for individual (country) but not for time effects for all models (see Table A3), revealing that a one-way panel model (cf. $\lambda_t = 0$) is more appropriate for the data than the two-way specification. Lastly, the Hausman specification test is carried out to decide whether the fixed or random effects model specification is more suitable. This is a test of endogeneity, stating in the null hypothesis that the unique errors are not correlated with the regressors ($corr(\mu_i, X_i) = 0$). The output shows that the Hausman tests are insignificant for the four models (see Table A4), implying that although both models are consistent, the random effects model is more efficient (Hausman, 1978). Taking into account the considerations described above allows to reduce the general model to the following equation:

$$\ln(Gini)_{it} = \alpha + \beta I_{it} + \gamma I_{it} \cdot D_{it} + \boldsymbol{\delta}' \boldsymbol{\Theta}_{it} + \mu_i + \nu_{it} \tag{3}$$

This one-way random effects model assumes that $\mu_i \sim IID\left(0, \sigma_\mu^2\right)$ and $v_{it} \sim IID\left(0, \sigma_v^2\right)$, and can be thought of as a random intercept multilevel model since it accounts for country-specific random intercepts.

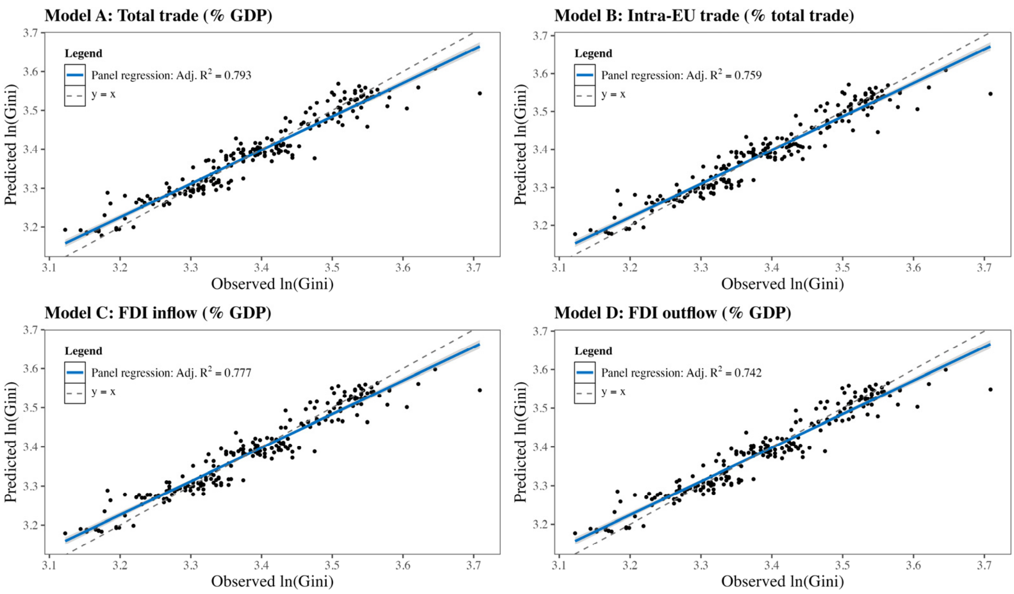
Table 3 contains the summaries of the four models, which are first estimated without robust standard errors (results with robust standard errors are later discussed in section 3.4). The coefficients of determination (R^2) of the models range from 0.75 for model D to 0.80 for model A, which means that the independent variables in the models explain about 75 to 80 % of the variation in Gini. Since adding explanatory variables automatically increases R^2, the adjusted R^2 is often a better measure of the explanatory power since it applies a 'penalty' for adding more predictor variables. These adjusted R^2-values range from 0.74 to 0.79. Notably, random effects panel models only compute predictions for Gini (and intercepts) for years when data for all the indicators is available. For this reason, only 218 of the 513 observations are used (for 21 countries), since the rest contain at least one missing value.

Table 3. *Model Summaries*

Model	n	T	N	R^2	Adj. R^2	χ^2	*df*	*p-value*
REM A	21	3-13	218	0.8019	0.7934	76.2482	9	0.0000
REM B	21	3-13	218	0.7692	0.7592	80.8229	9	0.0000
REM C	21	3-13	218	0.7864	0.7771	68.8947	9	0.0000
REM D	21	3-13	218	0.7528	0.7421	73.3784	9	0.0000

The fit of the models is illustrated in Figure 10, which plots the predicted values for $\ln(Gini)$ for each model against the observed data.

Figure 10. Model fit of the four models



The figures show that the predicted values for $ln(Gini)$ fit the observed values fairly well, although the models tend to slightly overestimate for small values, and underestimate for large values of Gini. This is illustrated by the deviation of the blue lines from the first bisector ($y = x$), which represents a perfect relationship. The estimated coefficients of the four models and their corresponding standard errors (shown in parentheses) are displayed in Table 4 and Table A7 in the appendix contains the country-specific random intercepts. As mentioned before, due to missing data only 218 observations were used, resulting in the fact that there are no intercepts calculated for Denmark, Estonia, Finland, Latvia, Lithuania and Sweden.

Table 4. *Random Effects Model Parameter Estimates (Non-Robust Standard Errors)*

	Dependent variable:			
	ln(Gini)			
	(A)	(B)	(C)	(D)
ln(Total trade (% GDP))	-0.070* (0.039)			
ln(Intra-EU trade (% total trade))		-0.099*** (0.036)		
FDI inflow (% GDP)			-0.00004 (0.0002)	
FDI outflow (% GDP)				-0.00001 (0.0002)
Unemployment (%)	0.005*** (0.001)	0.004*** (0.001)	0.005*** (0.001)	0.005*** (0.001)

Table 4. *Continued*

	Dependent variable:			
	ln(Gini)			
	(A)	(B)	(C)	(D)
ln(Population)	0.009 (0.016)	0.070*** (0.022)	0.026* (0.015)	0.027 (0.017)
Market capitalisation (% GDP)	-0.0003** (0.0001)	-0.0002* (0.0001)	-0.0002** (0.0001)	-0.0002* (0.0001)
Natural resources (% GDP)	-0.036*** (0.011)	-0.035*** (0.011)	-0.031*** (0.011)	-0.031*** (0.011)
ln(Social benefits (% GDP))	-0.131*** (0.031)	-0.127*** (0.031)	-0.121*** (0.031)	-0.122*** (0.031)
ln(Tech exports (% total exports))	-0.054*** (0.013)	-0.049*** (0.013)	-0.054*** (0.013)	-0.054*** (0.013)
Dummy: Below-average GDP per capita (current USD)	-0.073 (0.224)	-0.136 (0.086)	0.014 (0.047)	0.015 (0.052)
ln(Total trade (% GDP)) * Dummy	0.015 (0.045)			
ln(Intra-EU trade (% total trade)) * Dummy		0.059 (0.036)		
FDI inflow (% GDP) * Dummy			0.0001 (0.0002)	
FDI outflow (% GDP) * Dummy				0.0001 (0.0003)
Constant	4.031*** (0.400)	2.897*** (0.325)	3.389*** (0.270)	3.380*** (0.298)

Note. *p<0.1; **p<0.05; ***p<0.01

In light of the logarithmic model specification, the parameters of the independent variables should be done accordingly. Coefficients of the predictors are partial derivatives with respect to that particular predictor, i.e. the slopes of a k-dimensional hyperplane (where k is the number of explanatory variables). The coefficients of the variables which are logarithmically transformed can be expressed by:

$$\beta = \frac{\partial Y}{\partial X} = \frac{\frac{\Delta Y}{Y}}{\frac{\Delta X}{X}} \Leftrightarrow \Delta \% Y = \beta \Delta \% X$$

which should be understood as an increase in X by 1 % corresponds to an increase in Y by β %, holding all other variables constant. For the predictors that are not in log-form, the parameters are given by:

$$\beta = \frac{\partial Y}{\partial X} = \frac{\frac{\Delta Y}{Y}}{\Delta X} \Leftrightarrow \Delta\% Y = 100 \cdot \beta \Delta X$$

which implies that a unit change in X is associated with a $100 \cdot \%$ change in Y, ceteris paribus.

First of all, the coefficient of *ln*(total trade) in model A is negative and significantly different from zero, indicating that overall trade openness is associated with less income inequality for the 27 EU members during the period of inquiry. Concretely, an increase in trade as a % of GDP of 10 % on average leads to a reduction in Gini by 0.70 %, all else being equal. Similarly, the parameter estimates of model B show that more intra-EU trade is significantly negatively related to income inequality. When a country increases its intra-EU trade as % of total trade by 10 %, a 0.99 % decrease in Gini is expected. These findings are central to this study and refute the conclusions of Alili & Adnett (2018), Beckfield (2009), and Tridico (2017), which claim that trade openness increases income inequality. The results however do agree with the findings of Furceri & Ostry (2019), Jaumotte et al. (2013), Mundell (1957), Ravinthirakumaran & Navaratnam (2018), and Tian et al. (2009), which state that regional integration reduces the income gap. Notably, no significant interaction effects between total trade or intra-EU trade and economic development are found. This implies that the negative relationships between both overall trade openness and intra-EU trade openness, and Gini are consistent across all countries, regardless of their GDP per capita. Next, the coefficients of the indicators of financial integration (FDI in- and outflow) are very small and not significantly different from zero. This means that no evidence is found for an intrinsic relationship between financial integration measured by FDI and income inequality in the EU. Additionally, the parameters of the interaction terms with economic development are also found to be insignificant.

Lastly, the coefficients of most of the control variables seem to match the initial expectations. Higher unemployment tends to increase Gini, which corresponds to previous research from Ametoglo et al. (2018), Busemeyer and Tober (2015), Cesaroni et al. (2019), Furceri and Ostry (2019), Monnin (2014), and Tridico (2017). The results show that when the unemployment rate increases by 10 %, Gini will on average rise by 4 to 5 %, ceteris paribus. Also population size is a factor that is associated with more inequality in models B and C but not in A and D. This contradicts the findings from Ametoglo et al. (2018) and Dorn et al. (2022), but is consistent with Ali et al. (2015). Market capitalisation, a measure included to capture financialisation is significantly and negatively associated with Gini in all four models, with coefficients ranging from 0.0002 to 0.0003. Hence, a rise of market capitalisation as % of GDP of 10 % is expected to decrease Gini by 0.30 %. This is similar to the findings from Baek and Shi (2016), but is in contrast with the majority of the works discussed in the literature review (cf. Furceri and Ostry (2019), Neagu et al. (2016), and Tridico (2017)). Contrary to the results from Ametoglo et al. (2018) but in line with those of Ali and Sami (2016), the presence of natural resources tends to decrease inequality in the context of the EU: a rise in natural resources of 10 % of GDP on average reduces Gini by about 0.31 to 0.36 % in this sample. Government expenditure on social benefits is, as expected, also negatively related to inequality. On average, if a government decides to spend 10 % more on social security, the Gini coefficient will drop by 1.21 to 1.31 %, all else being equal. Similar results are reported in Beckfield (2006), Beckfield (2009), Busemeyer and Tober (2015), Kenworthy (1999), Piotrowska (2008), Tian et al. (2009), and Tridico (2017). This conclusion was of course anticipated, since one of the major objectives of public social spending is the reduction of inequality (Cook & Kabeer, 2009). Finally, the amount of technology exports is another factor that is associated with reduced inequality, a result that is confirmed by Cesaroni et al. (2019) and Tung et al. (2020), but contradicted by Asteriou et al. (2014) and Baek and Shi (2016).

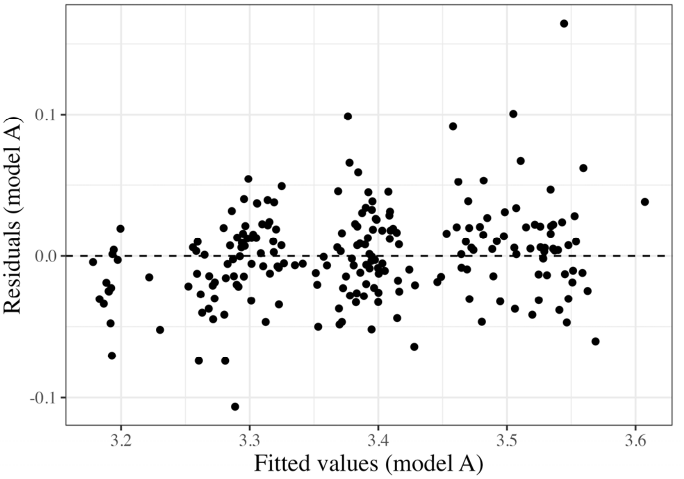
In conclusion, most of the coefficients of the control variables seem to match the hypotheses that were specified in advance. Similarly, the parameters of the trade integration indicators match the group of works that claim that trade reduces inequality, while no evidence has been found for a relationship between FDI and Gini within the bounds of the EU. Potential explanatory mechanisms for these effects are addressed in the general discussion below (section 4).

D. Model assumptions and diagnostics

Lastly, the standard assumptions of the model are checked in order to meaningfully interpret the results. First, it is checked whether the variables in the model are stationary. This can formally be tested with the Levin-Lin-Chu unit-root test. The test is significant ($z = -2.28$, $p = 0.01$), which means that the data is stationary and thus does not contain a unit root. Consequently, the joint probability distributions of the variables are constant in time, hence their variance and mean remain the same over the years of the panel (Gagniuc, 2017; Levin et al., 2002). Because of this, the variables tend to 'return to their mean', which is beneficial for inference.

Next, the variance-covariance matrix of the idiosyncratic error term ϵ_{it} implies a homoskedastic variance $\left(var(\epsilon_{it}) = \sigma_\mu^2 + \sigma_v^2\right)$ for all i and all t (Baltagi, 2021). This assumption is checked with the Breusch-Pagan test for heteroskedasticity and the residual plot in Figure 11. The test is significant for all four models (see Table A5 in the appendix) which indicates that the null hypothesis of constant variance (i.e. homoskedasticity) can be rejected, and thus $var(\epsilon_{it}) \neq \sigma_\mu^2 + \sigma_v^2$ (Breusch & Pagan, 1979). Hence, heteroskedasticity is present in the models, meaning that the residual variances are not constant for larger values of *ln*(Gini). This result is confirmed by the residual plot of model A below, which clearly shows a pattern of increasing variation in the residuals for higher predicted values. Similar patterns are present in the residual plots of the other models.

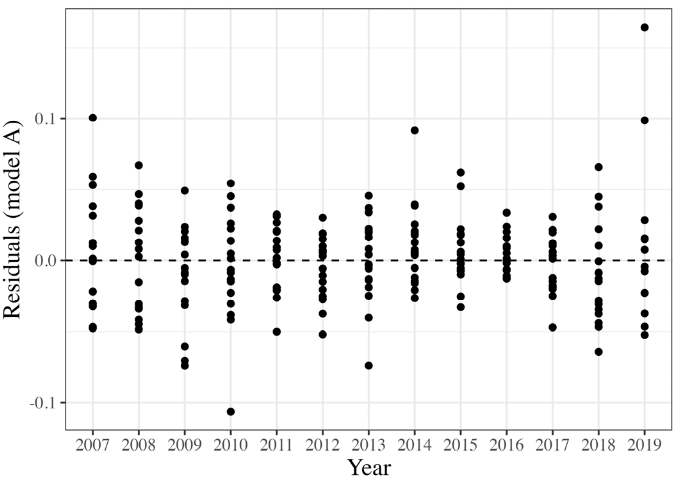
Figure 11. Residual plot of model A: residuals against fitted values



Although heteroskedastic estimators will still be unbiased, their standard errors will often be incorrect which makes inference based on their confidence intervals or p-values problematic. Therefore, the models will be re-estimated with robust standard errors, since they allow for heteroskedasticity.

To assess whether there is serial correlation present in the models, the Durbin-Watson test is performed. The null hypothesis of this test is that there is no autocorrelation in the idiosyncratic errors, hence: $cov(\epsilon_{t_1}, \epsilon_{t_2}) = 0$ (Croissant & Millo, 2018; Durbin & Watson, 1951). The test also reaches significance for all four models (see Table A6 in the appendix), which indicates that serial correlation is present, and thus different residuals over time are not independent. Graphically, this can be checked with a residual plot that shows the residuals against time (see Figure 12).

Figure 12. Residual plot of model A: residuals against time



The 'wave' pattern seen in Figure 12 shows that positive autocorrelation is present ($\rho\left(\epsilon_{t_1},\epsilon_{t_2}\right)\rangle 0$). This pattern is also present in the residual plots of the other models. This autocorrelation has the same consequences as the problem of heteroskedasticity and could even be a cause of heteroskedasticity.

To resolve the problems of heteroskedasticity and autocorrelation, the models are re-estimated using robust standard errors. Table 5 shows the regression estimates with robust standard errors by using the Arellano method (Arellano, 1987).

Table 5. *Random Effects Model Coefficients with Robust Standard Errors*

	Dependent variable:			
	ln(Gini)			
	(A)	(B)	(C)	(D)
ln(Total trade (% GDP))	-0.070 (0.051)			
ln(Intra-EU trade (% total trade))		-0.099** (0.038)		
FDI inflow (% GDP)			-0.00004 (0.0001)	
FDI outflow (% GDP)				-0.00001 (0.0002)
Unemployment (%)	0.005*** (0.001)	0.004*** (0.001)	0.005*** (0.001)	0.005*** (0.001)
ln(Population)	0.009 (0.014)	0.070** (0.028)	0.026** (0.012)	0.027** (0.012)
Market capitalisation (% GDP)	-0.0003*** (0.0001)	-0.0002** (0.0001)	-0.0002*** (0.0001)	-0.0002** (0.0001)
Natural resources (% GDP)	-0.036*** (0.011)	-0.035*** (0.012)	-0.031** (0.015)	-0.031** (0.015)
ln(Social benefits (% GDP))	-0.131*** (0.048)	-0.127** (0.052)	-0.121** (0.054)	-0.122** (0.055)
ln(Tech exports (% tot exports))	-0.054** (0.023)	-0.049** (0.024)	-0.054** (0.023)	-0.054** (0.022)
Dummy: Below-average GDP per capita (current USD)	-0.073 (0.304)	-0.136** (0.064)	0.014 (0.042)	0.015 (0.042)
ln(Total trade (% GDP)) * Dummy	0.015 (0.064)			
ln(Intra-EU trade (% total trade)) * Dummy		0.059** (0.029)		
FDI inflow (% GDP) * Dummy			0.0001 (0.0001)	
FDI outflow (% GDP) * Dummy				0.0001 (0.0002)
Constant	4.031*** (0.435)	2.897*** (0.425)	3.389*** (0.214)	3.380*** (0.224)

Note. *p<0.1; **p<0.05; ***p<0.01

Compared to the results in Table 4, a few things have changed. Although the estimated coefficients remain the same, the standard errors have changed slightly. For instance, the coefficient of *ln*(total trade) in model A is rendered insignificant. Moreover, the interaction effect between *ln*(intra-EU trade) and the below-average development dummy has become significantly different from zero. This suggests that economic development does indeed moderate the relationship between trade integration and income inequality. However, the sign of the coefficient is opposite to what is expected from the theoretical predictions of the Heckscher-Ohlin model and the Stolper-Samuelson theorem. While these theoretical frameworks anticipated that less developed economies would benefit more from trade openness in terms of income inequality than above-average countries, this finding suggests that the opposite is true. Specifically, the result shows that for EU members with a below-average GDP per capita, the reducing effect of intra-EU trade on Gini is weaker (-0.099 + 0.059 = -0.04) than for countries with above-average GDP per capita, ceteris paribus. However, this moderation effect is not strong enough to reverse the relation between trade and inequality for less developed economies, but rather reduces the beneficial effect of trade. Particularly, an increase in intra-EU trade as % of total trade by 10 % is expected to reduce Gini by 0.99 % for countries with above-average GDP per capita, while this will only reduce Gini by 0.40 % for economies with below-average GDP per capita. The significance levels of both the FDI variables as well as the covariates did not change substantially as a result of the robust standard errors.

Ultimately, it is checked whether there is a problem of multicollinearity, i.e. high correlations between the independent variables. As seen before in Figure 6, the correlations between the independent variables do not exceed 0.50, except for the correlations between intra-EU trade and population size, and between inward and outward FDI. To measure the severity of multicollinearity in the models, the variance inflation factors (VIFs) are calculated for each model. The *VIF*s of the predictors are displayed in Table 6 and should be interpreted as follows: the square root of the VIF indicates how much the standard error of an explanatory variable has increased compared to if it were completely uncorrelated to the other independent variables.

None of the VIFs are higher than the threshold value of 5, except for the variables that are included in the interaction term with the development dummy. This is of course anticipated, since the interaction term reintroduces the integration variables into the models for a second time. While multicollinearity could inflate the standard errors of the parameter estimates, thereby reducing their precision and statistical power of the models, it has no impact on the values of the estimates themselves. Considering that the VIF-values are high because of structural multicollinearity, that is, high correlations caused by model specification (i.e. the interaction terms), no further steps are undertaken to avoid it. Moreover, aiming to reduce the correlations by using centred versions of the integration variables results in a computationally singular predictor matrix, making it impossible to estimate the models.

Table 6. *Variance Inflation Factors*

Predictor	VIF (A)	VIF (B)	VIF (C)	VIF (D)
ln(Total trade (% GDP))	3.623			
ln(Intra-EU trade (% total trade))		7.571		
FDI inflow (% GDP)			12.549	
FDI outflow (% GDP)				15.841
Unemployment (%)	1.951	1.906	1.890	1.920
ln(Population)	1.312	1.902	1.098	1.095
Market capitalisation (% GDP)	1.134	1.085	1.110	1.318
Natural resources (% GDP)	1.097	1.095	1.035	1.034
ln(Social benefits (% GDP))	1.993	1.928	1.930	1.968
ln(Tech exports (% tot exports))	1.146	1.170	1.151	1.150
Dummy: Below-average GDP per capita (current USD)	28.096	3.331	1.107	1.094
ln(Total trade (% GDP)) * Dummy	28.942			
ln(Intra-EU trade (% total trade)) * Dummy		5.525		
FDI inflow (% GDP) * Dummy			12.684	
FDI outflow (% GDP) * Dummy				15.689

IV. General Discussion

The panel analysis performed in this analysis yields interesting results. When using robust standard errors to estimate the coefficients, evidence is found for a significant and negative relationship between intra-EU trade and income inequality. Moreover, this relation is moderated by the level of economic development of a country. For countries with a below-average GDP per capita (compared to the EU27 average in the same year), the reducing effect of regional trade is weaker than for richer countries. While this interaction effect with economic development was expected based on economic theory, the sign of the coefficient contradicts the conventional predictions of the Heckscher-Ohlin model. Namely, instead of below-average developed economies benefiting more from trade openness in terms of reduced inequality than richer countries, the opposite trend is observed. Nonetheless, this moderation effect is not strong enough to reverse the effect of regional integration for countries with below-average GDP per capita. Instead, intra-EU trade lowers the income gap for all countries in the EU, but this beneficial effect is weaker for poorer member states. Specifically, a 10 % increase in intra-EU trade as % of total trade is associated with a decrease in Gini of 0.99 % for countries with above-average GDP per capita, but only with a 0.40 % decrease for countries with below-average GDP per capita. In other words: intra-EU trade is associated with reduced inequality for both rich and poorer EU members.

The main line of reasoning found in the literature that provides a potential explanation for the negative coefficient of intra-EU trade is based on the skill-premium hypothesis of the Heckscher-Ohlin model. More regional trade increases the demand for production factors that are abundant in the export sectors of the exporting countries, and in the import sectors of the importing countries. It could be reasoned that, because most EU members are rather advanced economies, these abundant production factors predominantly include high-skilled labour (cf. Hoftijzer & Gortazar, 2018). As a consequence, the EU economies might both import and export goods that are largely made with high-skilled labour when they engage in intra-EU trade. This increases the demand for high-skilled workers which are relatively dominant in the EU labour markets, and in turn reduces inequality. This could also explain why the negative effect of regional trade is weaker for countries with a lower GDP (cf. the interaction effect): because the proportion of high-skilled employees in their economies could be smaller. Given that these assumptions about the factor endowments of the EU member states are true, these results are consistent with the theoretical predictions of the Heckscher-Ohlin model. Nevertheless, it remains unclear whether this presumption is valid in the context of the EU. Therefore, further analyses specifically using data on educational attainment and skill level of the labour markets, as well as data on the precise products and services that are traded between the countries are required in order to conclusively shed light on the skill-premium hypothesis. On the other hand, total trade, FDI inflow and FDI outflow do not have a significant impact on Gini at the 0.05 level when using robust standard errors. While Herzer & Nunnenkamp (2013) have found evidence for a negative long-run relation between FDI and income inequality, there is not a lot of clear evidence for this relationship for advanced economies. The authors also contend that the theoretical basis for the relation between both FDI inflow and outflow is theoretically ambiguous. In addition, this analysis uses inward and outward FDI flows to quantify financial integration, instead of using FDI stocks (see e.g. Herzer & Nunnenkamp, 2013; Neagu et al., 2016). Given the inherently volatile nature of FDI flows when compared to FDI stocks, they might appear to be unrelated to the Gini coefficient, which is itself quite stable. On top of this, no distinction was made between different types, such as vertical and horizontal FDI, which might relate differently to income inequality (cf. Herzer & Nunnenkamp, 2013). In the same manner, overall trade openness can be quantified in multiple ways. Using more elaborate trade metrics in forthcoming analyses could shine light on the precise relationship, or lack thereof, between global trade integration and inequality of income (see further below).

Next, the effects of the control variables are in line with most expectations as well. A higher unemployment rate drives inequality, a result that is expected since it increases the wage gap between the active and the unemployed (Cesaroni et al., 2019). Also population size is linked to more inequality, which could be explained by the fact that it increases the labour supply, resulting in lower wages (Claus et al., 2012). By contrast, the other covariates are associated with a smaller income differential. When economies are more financialised (measured by market capitalisation), income inequality is expected to fall. This is however different from the expectation of Tridico (2017), who declares that financialisation leads to financial capitalism, which in turn incentivises governments to compete by instantiating policies that allow for social dumping and labour market deregulation. Next, more natural resources present within a country translates into a lower Gini coefficient. A similar result is found by Hartwell et al. (2022), suggesting that the high institutional quality of the EU member states equips them with the ability to more effectively manage and equitably distribute the wealth generated from natural resources. As discussed before, the effect of social benefits is coherent with the expectations. Evidently, societies which invest more in social protection will have fewer inequalities. Lastly, more technological exports are also linked to less income inequality. Analogously to the explanatory mechanism underlying the skill-premium hypothesis, more technological exports could increase the wages of high-skilled labourers, which are more prevalent in the EU.

These results are relevant beyond the domain of econometric research and should be considered by governmental institutions and policy makers. The ambiguous relationship between inequality and integration is often subject of political debate, which can lead to Euroscepticism and pessimistic views towards the EU integration project (Burgoon, 2013; Kuhn et al., 2016). Notwithstanding, based on the current analysis, a recommendation for political leaders would be that instantiating or retaining trade barriers should not be done for the sake of reducing within-country income inequality. In fact, the reverse seems to be true: regional economic integration on average decreases Gini for the current EU members.

All things considered, this work differentiates from the existing literature in two ways: (1) the relationship between economic integration and income inequality was assessed for the European Union, a region that is not often subjected to similar analyses, and (2) an interaction effect is examined that differs from the conventional measures for less developed economies.

Regardless, this analysis also faces certain limitations and weaknesses. A first shortcoming is that from the results of this analysis, the presence of a causal relationship between the variables cannot be deduced. This is because proving causality is not strictly achievable in econometric research, since omitted variable bias can never be completely ruled out. It can for example only be inferred that intra-EU trade explains variation in Gini over time, but it cannot be asserted with absolute certainty that regional trade is the causal driver of the changes in inequality. Nevertheless, considering the estimated model is a random effects panel model, the coefficients could be carefully interpreted as potential causal effects for most practical purposes. Since the relationships between the integration indicators and Gini remain after controlling for the relevant covariates, it can be assumed with a fairly high probability that the effects are intrinsic and thus not mediated by other variables. Despite that, estimating more advanced models such as lagged panel models or random growth curve models (cfr. Berry & Willoughby, 2017; Duncan, 1969; Muthén & Curran, 1997) could be an interesting approach in novel research on this subject, since they are better suited for establishing causal inference. Moreover, as argued by Akyuz et al. (2022 and 2023), due to potential cross-sectional dependence in the data, subsequent studies should conduct cointegration analyses to explicitly explore possible long-run relationships between inequality and integration. It could also be of interest to conduct further investigation on alternative explanations and potential mediators for the relationships established in this work.

Another limitation is that the effects of the interaction terms should be interpreted with caution, since the moderation dummy - indicating below-average GDP per capita relative to the EU27 in the same year - differs substantially from the ones used in prior research. Instead, many other studies performing similar moderation analyses use 'true' developing countries, in the sense that they are classified as developing or low-income according to organisations such as the World Bank, the IMF or the UN. Another popular classification criterion is the Human Development Index (HDI), which takes many indicators into consideration such as life expectancy, education, income, etc. (Stanton, 2007). Yet, in the context of the EU, the relevance of these measures could be questioned, considering that all but five of the 27 EU members (Bulgaria, Croatia, Hungary, Poland and Romania) are classified as 'advanced economies' according to the IMF (2020), and that the HDIs of all EU members are very high (above 0.80; UNDP, 2022). Other cut-off values than below- and above-average GDP per capita, or even other metrics of economic development could also be tested in the interaction effect. Similarly, future studies could consider alternative indicators of economic integration such as the economic freedom index (see Bergh & Nilsson, 2010), the composite KOF (Konjunkturforschungsstelle) integration indices (see Dreher, 2006), or the openness typology described in Gräbner et al. (2021). In the same manner, other metrics to capture income inequality such as the Theil or Atkinson indices could be used as the dependent variable in further analyses. Obtaining comparable outcomes with these alternate measures would offer strong support for the inherent connection between economic integration and income inequality.

Moreover, it is necessary to perform further analyses to specifically investigate the skill-premium hypothesis of the Heckscher-Ohlin model as the explanatory mechanism underlying the relation between economic integration and inequality. By using data on educational and skill disparities within labour markets, it becomes feasible to inquire whether the effect of trade integration is mediated by variation in the ability of the workforce. Including these variables in the interaction terms with the integration variables would explicitly allow to test the skill-premium hypothesis. Additionally, utilising bilateral trade data such as the Comtrade database of the United Nations (2022) would allow examination of the specific goods and services that are traded between economies. This could provide insight into the validity of Heckscher-Ohlin assumptions on relative factor endowments and corresponding trade patterns.

Besides that, due to limited data availability, the regression was performed on a panel data set with missing values. Especially in country-level panel data, incompleteness is not always a result of randomly missing values, but is often caused by the data collection methods. This means that the incompleteness could be related to the idiosyncratic error term (v_{it}), which can potentially cause biased estimates. The missingness caused only 218 of the 513 observations to be used in the regression, which leads to having no estimated country-specific intercepts for certain member states. Although the estimation methods used in this analysis can handle incomplete panels, it could be beneficial to allocate more time and resources to the data collection with the aim of reducing incompleteness as much as possible. Additionally, alternative strategies to handle missing data such as multiple imputation could be employed in further analyses.

V. Conclusion

In conclusion, the ambiguity in the existing literature makes it difficult to assess the precise relationship between economic integration and income inequality. Nevertheless, the results of this empirical analysis shed light on this connection within the bounds of the EU. To answer the initial research question posed in the introduction: regional economic integration is associated with reduced income inequality for the EU member states during the period of inquiry, while total trade and the FDI variables appear to have no direct connection to inequality. In addition, for below-average economies in terms of GDP per capita, the relation between intra-EU trade and inequality is also negative, but weaker in absolute value. These relations hold, even after controlling for the relevant covariates, which assigns a high probability of these effects to be genuine and not spurious.

Overall, these findings agree with a significant part of previous economic research, yet are in contrast with others. Importantly, the conclusions of the analysis should be interpreted in consideration of its limitations, but could be guiding for future research. Regardless, understanding the link between economic integration and inequality of income should be of primary interest to those who aim to understand the consequences of the European integration project. Considering its multitude of negative effects, this understanding is vital in the global fight against inequality.

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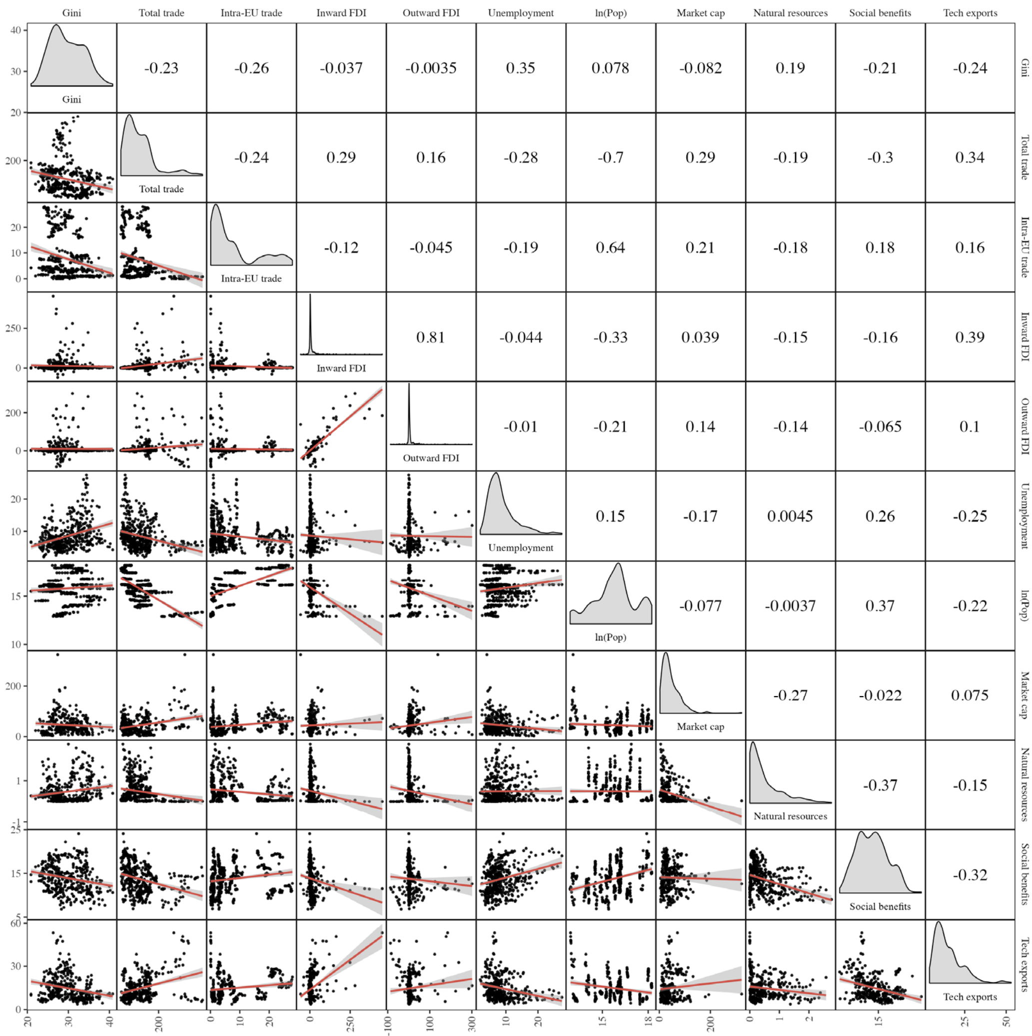
Appendix A. Additional Descriptive Statistics

Table A1. *Summary Table of Descriptive Statistics*

Variable	*n*	Min.	Max.	Median	Mean	Std. dev.	Skewness	Kurtosis
Gini (%)	453	20.90	40.80	29.20	29.68	3.98	0.31	-0.71
ln(Gini %))	453	3.04	3.71	3.37	3.38	0.13	0.08	-0.81
Total trade (% of GDP)	513	45.42	380.10	104.55	120.58	63.06	1.68	3.11
Intra-EU trade (% of total trade)	513	0.10	27.90	3.80	7.55	8.17	1.18	-0.02
Inward FDI (% of GDP)	513	-57.61	449.08	3.25	12.63	40.51	6.55	52.39
Outward FDI (% of GDP)	513	-87.23	301.25	1.94	9.27	36.04	4.46	26.41
Unemployment (%)	513	2.01	27.47	7.52	8.60	4.29	1.58	2.93
Population	513	395969.00	83240525.00	8391643.00	16329499.76	21476857.21	1.80	2.08
Market capitalisation (% of GDP)	361	2.64	322.34	35.35	44.22	34.83	2.42	12.10
Natural resources (% of GDP)	486	0.00	2.76	0.30	0.49	0.56	1.55	1.96
Social benefits (% of GDP)	476	6.80	24.10	13.60	13.71	3.22	0.26	-0.60
Tech exports (% of total exports)	378	3.77	53.02	12.24	14.78	8.33	1.53	3.10
Below-average GDP (dummy)	513	0.00	1.00	1.00	0.60	0.49	-0.39	-1.85

The plot below shows the bivariate relationships between the variables and displays their distributions by means of density plots on the diagonal. However, the scatter plots and their correlations should be interpreted in view of the fact that the data is a panel data set. The plots below aggregate all years, which could give seemingly inconsistent results with the findings of the panel analysis described in the analysis.

Figure A1. Correlation matrix with density plots



This plot shows the same correlation matrix, but distinguishes the two categories of the interaction term: countries with a below-average GDP per capita in red, and with above-average GDP per capita in blue. Interestingly, the density plots reveal that countries with a below-average GDP per capita generally have higher Gini coefficients, less intra-EU trade, more unemployment, lower populations, less market capitalisation, more natural resources, less social benefits and less technology exports.

Figure A2. Correlation matrix with density plots split over high vs low GDP per capita

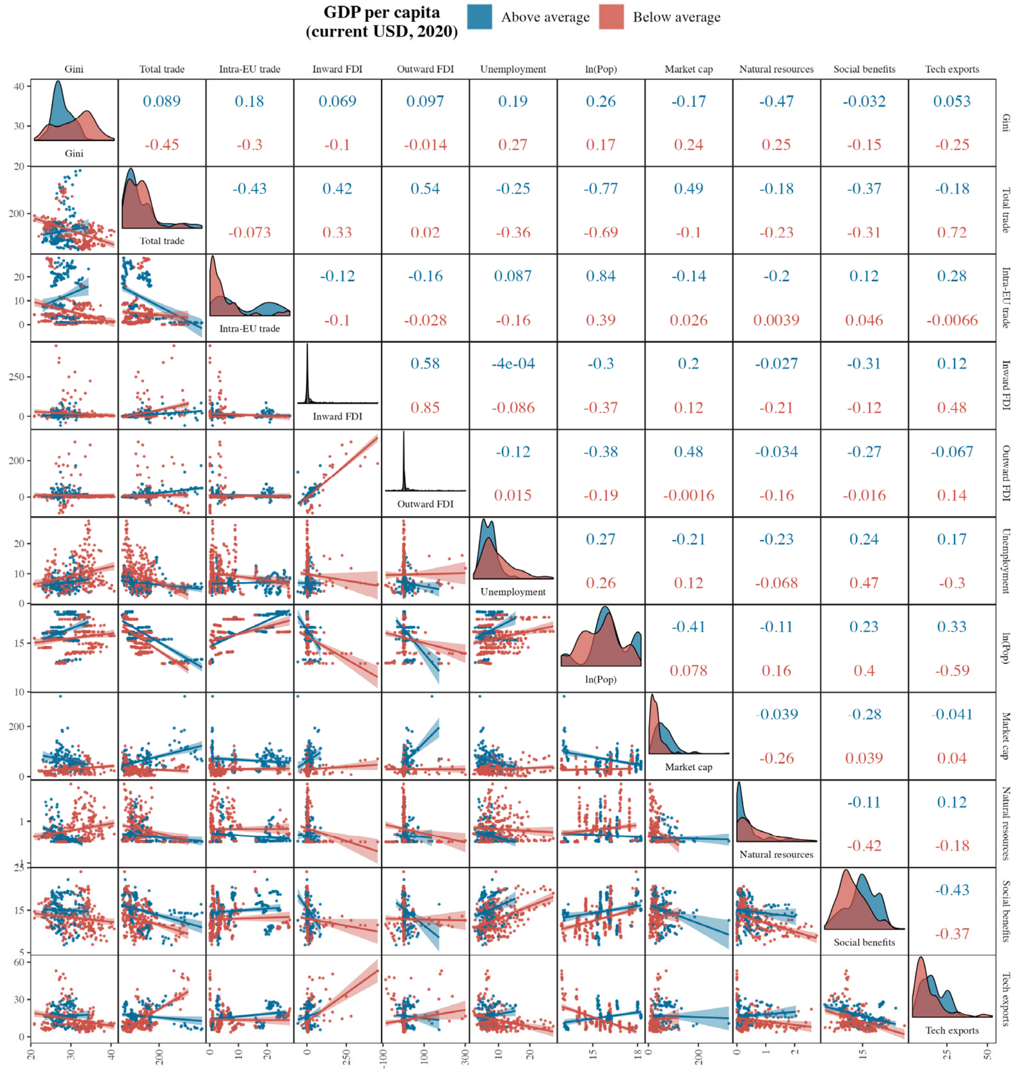


Table A2. *F-test of Poolability*

Model	F	df_1	df_2	*p*-value
A	53.466	19	189	$\langle 2.2 \cdot 10^{-17}$
B	59.659	19	189	$\langle 2.2 \cdot 10^{-17}$
C	69.785	19	189	$\langle 2.2 \cdot 10^{-17}$
D	72.867	19	189	$\langle 2.2 \cdot 10^{-17}$

Table A3. *Lagrange Multiplier Test for Individual Effects*

Model	χ^2	df	*p*-value
A	530.51	1	$\langle 2.2 \cdot 10^{-17}$
B	745.5	1	$\langle 2.2 \cdot 10^{-17}$
C	581.12	1	$\langle 2.2 \cdot 10^{-17}$
D	639.98	1	$\langle 2.2 \cdot 10^{-17}$

Table A4. *Hausman Specification Test*

Model	χ^2	df	*p*-value
A	8.9965	8	0.3426
B	7.5715	8	0.4764
C	0.44336	8	0.9999
D	1.6186	8	0.9906

Table A5. *Breusch-Pagan Test for Heteroskedasticity*

Model	BP	df	*p*-value
A	36.2307	9	0.0000
B	81.1224	9	0.0000
C	29.1558	9	0.0006
D	30.8530	9	0.0003

Table A6. *Durbin-Watson Test for Serial Correlation*

Model	DW	*p*-value
A	1.4620	0.0000
B	1.4352	0.0000
C	1.4412	0.0000
D	1.4514	0.0000

Table A7. *Country-Specific Intercepts* (μ_i)

Country	REM A	REM B	REM C	REM D
AUT	-0.0330	-0.0099	-0.0163	-0.0155
BEL	-0.0852	-0.0278	-0.1005	-0.1005
BGR	0.1644	0.1262	0.1592	0.1607
CYP	0.0547	0.1694	0.0770	0.0676
CZE	-0.1422	-0.0918	-0.1601	-0.1620
DEU	0.0093	0.0146	0.0006	-0.0007
ESP	0.0021	-0.0231	0.0077	0.0072
FRA	0.0150	0.0523	0.0345	0.0341
GRC	0.0647	0.0387	0.0919	0.0933
HRV	0.0035	-0.0220	0.0239	0.0253
HUN	-0.0426	-0.0615	-0.0690	-0.0699
IRL	0.0520	-0.0421	0.0434	0.0438
ITA	0.0089	0.0408	0.0401	0.0400
LUX	0.0858	-0.0242	0.0684	0.0704
MLT	0.0463	0.0284	0.0395	0.0480
NLD	-0.0529	-0.0037	-0.0703	-0.0715
POL	0.0063	-0.0366	-0.0188	-0.0197
PRT	0.0782	0.0884	0.0943	0.0949
ROU	0.1332	0.0988	0.1327	0.1326
SVK	-0.1687	-0.1580	-0.1855	-0.1860
SVN	-0.1999	-0.1571	-0.1928	-0.1919