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China's WTO accession and income inequality in European regions: External pressure and internal adjustments

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ABSTRACT

Exports from China have surged substantially since its accession to the World Trade Organization in 2001. We investigate how this expansion affected income inequality within European regions by separating the trade pressure experienced in external and domestic markets, as well as exploring the importance of several economic mechanisms. Despite some intermediate adjustments, softening the influence of Chinese pressure and even facilitating European exports, we establish a significant increase of inequality that is concentrated mostly in the lower part of regional income distributions. We determine a significant channeling of the trade pressure to income inequality through the shrinking manufacturing sector, the increasing unemployment rate, and the technological upgrade of manufacturing exports, together with an increasing demand for better-qualified labor.

1. Introduction

Income inequality tended to increase in most countries in the past decades, even though the global income inequality decreased due to catching-up by emerging economies (see, e.g., Alvaredo et al., 2017, 2018). This inequality increase in individual countries was driven by various intensive and intertwined processes, including globalization and skill-biased technological changes (see, e.g., ibidem, Dabla-Norris et al., 2015, OECD, 2015). The increase in income inequality in Europe was relatively moderate but quite uneven. As we will show later, inequality (of household equivalized income) even tended to decrease in most European Union (EU) regions during 1995–2000, while the trend reverted afterward. China's accession to the World Trade Organization (WTO) in 2001 was a major globalization event connected with trade liberalization. This paper reveals how this episode has contributed to the increase in income inequality in the (former) EU15 countries.[1]

Our research contributes to the understanding of the impact of China's WTO accession on income inequality in the EU – at the EU15, national, and regional levels – using EU15-wide regional data. We believe that, as (Alvaredo et al., 2017, p. 36) pointed out, "it is complementary to study inequality dynamics at the national, regional, and global levels". To

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[1] As we consider the period before the Brexit, the EU15 countries are Austria, Belgium, Denmark, Germany, Greece, Spain, Finland, France, Ireland, Italy, Luxembourg, the Netherlands, Portugal, Sweden, and the United Kingdom.

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our knowledge, the uniqueness of our analysis of China's impact on inequality in the EU lies not only in our EU-wide consideration and the employed income concept. We also reveal the important specificity of the EU case and overcome the limitation of previous studies with a narrow focus. Such studies considered either the domestic pressure or the very few sectors and goods having experienced direct changes in tariff and non-tariff trade barriers with China after its WTO accession. We explain why this narrow focus applied to the EU can be misleading, in contrast with many other countries, including the United States (US), where substantial changes in bilateral trade conditions with China took place after its WTO accession.

China's accession to the WTO by the end of 2001 and the ensuing acceleration of its foreign trade activity constitute a unique trade liberalization episode, whose potential effect on industrialized Western countries has already attracted considerable attention. Related research on labor market effects has taken off with the seminal regional study of Autor et al. (2013), which found falling wages and employment in affected local labor markets in the United States (US). More recently, the sector-level analysis of Pierce and Schott (2016) found adverse employment effects reinforced through input–output linkages, which (Acemoglu et al., 2016) confirmed using a combined regional and sectoral approach. Autor et al. (2016) provides a comprehensive view of the further corroborating evidence, emphasizing increasing wage inequality.

Fewer studies show negligible or even positive impacts on non-Chinese labor markets. For instance, Feenstra and Sasahara (2018) and Feenstra et al. (2019) found that the increased demand from China and cross-border service provision compensate for losses experienced in goods traded on the domestic and export markets. Nevertheless, even if there were certain compensations in terms of quantity leading to the same amount of employed labor, the change in the composition of goods and services produced and traded by an economy can have nontrivial implications for the demanded labor skills and income inequality. Consequently, next to the beneficial effects through increased product variety and cheaper products for consumers, more efficient firms, and decreasing global inequality between countries,[2] it is relevant to study the consequences of trade liberalization also for income inequality within countries, as these economic changes can have drastic socio-political implications. For instance, Colantone and Stanig (2018) showed that citizens of the United Kingdom (UK) regions affected more severely by the China shock tended to vote more to leave the European Union (EU), and at the same time, the percentage who voted to leave the EU was systematically larger among people having lower wages (Bell and Machin, 2016).

Europe-wide evidence on China's trade expansion's effect on income inequality is lacking, whereas the country studies available yield mixed results. In Norway, the expansion has decreased manufacturing employment's share through pressure on both the global and the local markets, as the study of Balsvik et al. (2015) demonstrates using rich regional data. In Denmark, the shock resulted in substantial inter-sectoral shifts with a still significant negative long-run effect on earnings (see Utar, 2018). In France, according to the study of Basco et al. (2017) (based on the idea of Autor et al., 2014, using employee–employer data), the effect was more pronounced on the lower end of the income distribution, manifesting itself in more churn-off and fewer hours worked, whereas (Malgouyres, 2017) finds that the wage distribution is uniformly negatively affected in manufacturing while the nontraded sector experienced wage polarization. Trade pressure has also accelerated technological change within and reallocation between firms, which have also had adverse labor market effects (see Bloom et al., 2016, using firm-level data). However, according to the study of Breemersch et al. (2019) on nineteen European Union (EU) countries using sectoral data, while China does seem to have some effect on job polarization, the primary driver behind it seems to be ongoing technological change, although this itself might have been induced by the China shock (Bloom et al., 2016). In Germany, it is mostly a secular trend that drives decreasing manufacturing employment, while globalization, and rising trade with China in particular, did not seem to speed up the manufacturing decline there, as (Dauth et al., 2017) find using sectoral–regional aggregated administrative data from Germany.

It is thus still a matter of active debate whether, in Europe, the labor market losses or benefits of the expansion of China dominate, and this is particularly true for inequality. The mere possibility of the China WTO accession's impact on the EU can be doubted based on the fact that, because of the preexisting bilateral trade agreements between China and the EU before the end of 2001, there were few changes in their bilateral trade conditions.[3] On the other hand the trade conditions between China and the US have been substantially affected, thus inspiring numerous research concentrating on it.

We point out that this fact does not eliminate the possibility of an impact on EU markets for at least the following two reasons. First, the change of the conditions of trade between China and third markets (i.e., outside the EU) affects the demand for goods (and/or services) exported from the EU to those third markets. Second, if the production of goods in China were connected with some fixed/sunk costs, the increase in output due to the global expansion of Chinese exports to the third countries would reduce the unit cost of production. Hence, it would lead to a competitive improvement of Chinese goods in terms of lower prices, even in markets where there were no changes of formal trade conditions in terms of tariff and/or non-tariff barriers.

[2] As shown, e.g., by Broda and Weinstein (2006), Halpern et al. (2015), and Lakner and Milanovic (2016), respectively.

[3] "7. The only obligation for WTO Members is that they must accord China so-called permanent MFN ('most favoured nation') status, entitling it to be treated in the same way as every other WTO Member, unless exceptions are specified in the protocol of accession. As the EU has always accorded China this status in any event, there will be virtually no practical impact". (see the Proposal for a Council Decision establishing the Community position within the Ministerial Conference set up by the Agreement establishing the World Trade Organization on the accession of the People's Republic of China to the World Trade Organization in European Commission, 2002 (url: https://eur-lex.europa.eu/LexUriServ/LexUriServ.do?uri=CELEX:52001PC0517:EN:HTML) or Snyder, 2009, p. 1069).

Two further predictions stem from the discussed arguments about the relative importance of and interconnections between the external pressure faced in export markets and the internal pressure experienced domestically by the EU countries due to the increasing imports from China.[4] First, the Chinese trade pressure faced directly in external markets is likely to be more significant than in the domestic EU market, because of larger changes of foreign trade conditions between China and third countries in comparison to those between China and the EU. This may also result in a failure to find a significant impact if only the domestic pressure were considered. Second, the domestically experienced pressure is likely to be highly correlated with the external pressure, as the former one stems from the scale effects of the increased total Chinese production and exports. This would also suggest that, at least in the EU's case, the best strategy for identifying the pressure would be the extraction of a common component from the two. Otherwise, high correlation between the domestic and external pressures can lead to a multicollinearity-induced increase in variance of standard estimators, and thus, potentially, in the apparent 'insignificance' and/or incorrect signs of some components. In our analysis, the discussed predictions will be corroborated by the data, whereas the previous EU studies did not take into account at least some of these aspects.

The initial China WTO accession shock to trade is not guaranteed to translate into further impacts, as it might be mitigated by various intermediate adjustments that reduce the original pressure. Hence, next to the pressure indicator, we consider various adjustment mechanisms. The first is import substitution, where imports from third markets are just substituted by Chinese goods, thus crowding out those from the third markets.[5] The second is export facilitation, where the supposedly cheaper intermediate products from China can facilitate EU production and exports. The third is export reallocation, where former exports from the EU to the markets of third countries can be replaced by EU exports to the fast-growing Chinese market if conditions there become more beneficial than elsewhere.

Given that these simple market-switching adjustments were insufficient to alleviate the pressure, further economic processes start taking place through various channels. The first of these is connected with the inter-sectoral shifts due to the Heckscher–Ohlin mechanism and factor price adjustments (see Leamer, 1995, Stolper and Samuelson, 1941, and, for a broader positioning and discussion, Helpman, 2017). The second is related to the intra-sectoral shifts due to the switch to skill-intensive products, economies of scale, and vertical specialization, as in Feenstra and Hanson (1996), Epifani and Gancia (2008), and Krugman (2008), and/or to the globalization-induced higher demand for skilled labor whenever only the most efficient firms export, as in Melitz (2003), Harrigan and Reshef (2015), and Furusawa et al. (2019). The first concentrates on changes between sectors—in our case, a decrease in the share of manufacturing. The second emphasizes vertical specialization, the importance of intermediate goods, and the selection of firms within the exporting manufacturing sector. Both result in forces that ultimately imply lower wages and potentially lost jobs for employees of firms that cannot or fail to adjust, but higher wages for higher qualifications and new jobs for those with firms that can take advantage of the ensuing changes. Dismissed workers that cannot be absorbed by more productive manufacturing firms or other sectors thus would become unemployed.

Even the inter-sectoral labor shift from manufacturing to other sectors, which would alleviate the increase in unemployment, is likely to raise income inequality for the following two reasons. Workers dismissed from the manufacturing sector are more likely to accept lower wages in this stressful situation with abnormal supply of labor due to the increased number of layoffs and the increased direct competition between them. In addition, there is also a simple composition effect, at least in the EU's case. As the right panel of Figure F3 reveals (see Appendix F.3 in the Supplementary Material), manufacturing and industry in general are among the sectors in the EU that typically have low intra-sectoral wage inequality. Hence, a random shift of labor from manufacturing to other (nonindustrial) sectors is likely to augment income inequality by itself, although in a less drastic and deterministic manner than a shift to unemployment. Consequently, the wage rate differentiation, inter-sectoral shifts, and (un)employment effects have the potential to increase income inequality, and the net of these is an empirical question.

In this paper, we study the effect of increasing trade pressure from China on regional income inequality[6] within the regions of the former EU15 countries (currently, EU14 and the UK) by constructing from different sources the comparable inequality and trade pressure measures covering the pre- and post-Chinese accession periods. Besides the fact that the number of observations with only fifteen cross-sections at the country level was simply insufficient for our econometric analysis,[7] the choice of the regional level is motivated by the fact that the dominant part of the pressure from China hit the manufacturing sector, whose employment share varies only modestly by countries but quite substantially by regions.[8]

4 Besides the external pressure on third markets and the domestically faced pressure, we also separated the China market as a particular EU export markets, in order to see if it has some specific importance, but it was insignificant in our sample.

5 See, e.g., Greenaway et al. (2008), Pham et al. (2016), and Baiardi and Carluccio (2019) for such evidence in exports of goods of varying technological intensity.

6 The precise concept of (net household equivalized) income under consideration is defined in Section 2. As inequality metrics, we will use the log-variance of income, the Gini index, and the income percentile ratios to measure income inequality. The main reported results rely on the log-variance of income because its empirical models outperform those with the Gini index in terms of adequacy, especially whenever a smaller number of instruments is employed.

7 We employ dynamic panel models estimated with the generalized method of moments, which relies on the asymptotics of the increasing number of cross-sections. A large number of cross-sections is therefore essential for our empirical application.

8 As our preliminary analysis shows, in the former EU15 regions to be considered, the labor share of 25 to 60-year-old workers working in manufacturing ranges from 3% to 34% at the NUTS1 level.

This variability, together with a much larger number of observations available at the regional level, increase the power of inference. At the same time, the level of (dis)aggregation needs to remain large enough that inequality metrics can be meaningfully applied and be sufficiently represented by the survey data (see the discussion in Section 2.1). Therefore, we do not go beyond the NUTS1[9] level of the EU's territorial disaggregation.[10]

Our investigation was motivated by the following stylized facts characterized in Appendix A. First, not only did major European exporting countries' shares of the total world exports fall along with that of the US after China's entry to the WTO in 2001 (see the left Figure A1 in Appendix A), but also the growth rate of exports decelerated substantially in most of the former EU15 countries (see Figure A3 in Appendix A). At the same time, inequality has increased in more than two-thirds of the regions over the same years, as indicated on the right panel of Figure A1. Whether this relationship is causal is not self-evident, but inequality tended to decrease in most EU regions before China's entrance to the WTO (see Figure F2 in Appendix F.2). At the same time, China's relative performance in terms of manufacturing exports has drastically improved in practically all (not only EU) markets and relative to the exports of practically any other country (see Figure A2 in Appendix A).

Our study has a broad scope. In addition to considering the EU-wide approach by constructing the respective comparable dataset, our main contribution to the current literature is threefold. First, we consider broad measures of household-equivalized income inequality that account for a wide array of individual, family, general equilibrium, and public policy adjustments. Unlike studies that focus on partial aspects or industries, our approach captures all of these adjustments and quantitatively characterizes China's WTO accession impact on regional income inequality using the aggregate measures of inequality. Second, most existing EU studies only examine either the trade pressure from China faced in domestic markets or the trade balance with China. Therefore, these studies cannot account for the primary pressure faced in export markets, and they underestimate its potential impact. The EU's case is specific, as there were few changes in the direct trade conditions between the EU and China after its accession to the WTO. However, ignoring the changes in third markets or concentrating exclusively on the European industries (like textile and apparel) that have directly experienced changes after China's accession to the WTO may lead to significantly biased estimates.[11] Even in methodologically refined discussions such as (Balsvik et al., 2015), the domestic and foreign market pressure indicators are considered separately despite the potential bias caused by omitting a variable. Using both the direct estimations and the analysis based on the loadings of the common factor of both pressure sources, we demonstrate that, in the EU's case, the pressure on export markets is predominant. We also show that only considering domestic pressure is insufficient and might lead to misleading results.[12] This result challenges some previous insignificant impact findings obtained by only using the domestic pressure or the import/export trade patterns with China. Third, by concentrating on the impact on income inequality and presenting the evidence in a unified framework, we show the potential relevance of the intermediate trade structure adjustments in EU countries. These adjustments mitigated the initial shock of Chinese trade pressure. We also identify the dynamic patterns of the changing significance of different channels of economic adjustment. To our knowledge, previous studies, even those that focus on separate EU countries, do not consider these dynamics. We also provide some additional insights on several other aspects of China's impact on the economy. First, we show that the lower and upper parts of regional income distributions were affected differently. Second, we demonstrate and explain the presence of a significant interaction between the shock caused by China and the financial crisis. Third, we introduce and explore several alternative instrumentations of trade pressure at the regional level. Our study points out the potential inadmissibility of instruments based on third countries' trade with China if the countries are geographically distant from China. Fourth, we quantify the importance of China's accession to the WTO on income inequality across European regions.

From the econometric point of view, our study differs from most of the previous literature because we use a dynamic panel data model estimated using the Generalized Method of Moments (GMM) and instrument the Chinese trade pressure along with additional regular instruments. This dynamic approach allows us to make distinctions between short- and long-term effects by deriving the respective paths of the impact. The approach also allows us to identify several other dynamic aspects, including the changing impact before and after the financial crisis and the changing importance of different channels of economic adjustment. This framework also allows us to explicitly control for other intermediate events and changes such as the EU's expansion, the introduction of the euro, and changing business cycle states and technologies, among others. The rest of the literature often relies on the instrumental variable (IV) estimator, using long differences. Here, the limited variability over time makes it more complicated to establish or account for the discussed effects. However, the identification is mainly achieved in the cross-sectional dimension with potentially weaker exclusion restrictions than in the panel framework. It should be pointed out that, other than in the discussed empirical literature

[9] Here, NUTS abbreviates the Nomenclature of Territorial Units of Statistics.

[10] Most of the regions are at the NUTS1 level with only a few exceptions — see the next section and Appendix F.2 for details.

[11] The increase in the competitiveness of Chinese products in third markets with changing conditions (particularly in the US) puts pressure on the competitiveness of EU products in those foreign markets, regardless of whether direct EU-China trade conditions have changed. Also, the increase in total Chinese exports and production can lead to scale effects, reducing unit costs and making Chinese products more competitive worldwide. In turn, this will put pressure on countries' domestic markets, regardless of whether direct EU-China trade conditions have changed or not.

[12] For instance, if researchers only considered the domestic pressure and used the long difference estimator that is widely applied in the related literature, China's impact would seem insignificant (for more details, see the discussion in the Robustness Section 5 and Table D8 in Appendix D).

on long differences, we do not assume that the instrumenting is correct in the GMM but explicitly test the admissibility of instruments in terms of over-identifying orthogonality conditions. Nevertheless, to be more confident in our findings, we also present a robustness analysis with the IV using long differences.

Being broad in its scope, our analysis has much in common with previous studies. Namely, we consider the trade pressure from China faced by EU producers in various markets, separating the external export markets, such as in Autor et al. (2013) and Balsvik et al. (2015), the domestic market, such as in Acemoglu et al. (2016), and also the potential gains from entering the Chinese market (as in Autor et al., 2013, or Balsvik et al., 2015). Given the previously highlighted importance of the external pressure for the EU's case, we start the presentation of findings from it while gradually introducing various other markets and adjustment effects afterward.

Our study finds a statistically significant and economically important increase in income inequality in the EU15 regions due to China's accession to the WTO, which is concentrated at the lower part of the income distribution. We use methods that are appropriate for a dynamic, EU-wide approach. We employ various micro- and international trade datasets to create several inequality measures of net household equivalized income adjusted for purchasing power differences, as well as various trade pressure indicators. For our empirical analysis, we employ the GMM/IV-based econometric estimations in panels and long differences. We establish that the most significant impact on inequality stems from the external pressure – which companies face in the export markets in which conditions changed substantially after China's accession – and not from the pressure encountered in the domestic market. Consequently, most previous studies are likely to underestimate the severity of the impact in the EU as they concentrated on the domestic pressure and considered only those sectors or goods that experienced direct changes in tariff and non-tariff barriers after China's WTO accession. We discover that the increase in inequality within EU regions was dominated initially by an inter-sectoral shift, manifesting itself largely through the shrinking shares of labor employed in manufacturing. According to our estimations, the unemployment rate became a significant contributor to inequality only after the global financial crisis. This change might indicate either that the initial shock due to the WTO accession was augmented further by the financial crisis or that the trade pressure's impact on inequality varies along the business cycle. Over the longer period, the intra-sectoral adjustments (within manufacturing) became the largest contributor to income inequality through the technological upgrading of exports (and production) together with the upskilling of the employed labor. We find some evidence that certain intermediate adjustments softened the initial pressure on income inequality. In particular, it appears that in the EU15, not only importers substituted goods from third countries with Chinese ones but also exporters used more affordable intermediate products from China. We also find that UK regions are among the most affected ones, which supports the potential relevance of the China shock for Brexit, as established by Bell and Machin (2016) and Colantone and Stanig (2018).

The structure of the paper follows the line of thought outlined previously. We delegate much of the technical detail to the Appendices, which are provided as electronic supplementary material. Appendix A presents the motivating plots discussed in the Introduction. Appendix B discusses a couple of econometric issues. It includes the state-space representation used to derive the country-specific dynamic common factors from the joint process of external and domestic Chinese trade pressure indicators together with its empirical characterization. It also presents the semi-reduced representation of the impact channels together with the calculation of their relative importance. Appendix C contains the detailed empirical analysis underlying Section 4, which summarizes the findings on the importance of other markets, intermediate adjustments, and impact channels. Appendix D reports the estimations of robustness analysis, including the results when varying the composition and number of instruments, with different definitions of pressure indicators, as well as controlling for the previous enlargement of the EU with countries from Central and Eastern Europe (CEE), migration, technological change, share of service exports, etc. Appendix E presents a stylized calculation of the implied inequality increase based on the simplest reduced-form specification of the impact measurement. Finally, Appendix F describes the trade, inequality, and employment data used in the paper, including the related stylized facts and respective plots.

2. Data and the econometric approach

2.1. Data

Our analysis rests on an unbalanced panel of 61–65 regions[13] covering the former EU15 countries (currently, EU14 and the UK) for the period between 1994 and 2014, created from multiple data sources (we discuss a few exceptions in Appendix F.2). The dataset brings together our impulse, outcome and intermediary variables, that is, measures of trade, inequality and employment structure, as well as other factors. A brief description of the dataset follows shortly, with further details provided in Appendix F and summary statistics of employed variables reported in Table D10 of Appendix D.

Our data of the trade drivers rely on the outcome of the joint OECD–WTO Trade in Value Added (TiVA) initiative.[14] Although qualitatively similar results were also obtained using the OECD Structural Analysis Statistics (STAN) database and, in particular, employing the STAN bilateral trade database by industry and end-use category, the TiVA approach has

13 Varying by different years due to the data (un)availability.

14 We use the 2016 version of it, which covers the largest span of the relevant period, whereas the 2018 version of the database initiates only in 2005.

a clear advantage in our case. For certain trade pressure indicators to be defined succinctly, we need direct comparability of exports, imports, production, and/or value added. This is granted in the TiVA database, where various issues of different pricing, data sources, etc. are already resolved in a unified framework, whereas they would be faced using primary data.

TiVA resulted in a database which includes, among others, bilateral trade statistics based on flows of value added embodied in final domestic demand, which is harmonized relying on the inter-country input–output tables. The employed 2016 edition of the TiVA database covers data from 1995 until 2011 and contains indicators for various economies, including those from the OECD, EU28 and G20, most East and South-east Asian economies and a selection of South American countries. The database also specifies indicators for various industrial sectors. Since we are interested mostly in manufacturing activities, we used multiple indicators (exports, imports, output, value added, etc.) with the Total Manufactures identifier (C15T37) from this database. We later aggregated these indicators to obtain exports from a given country to a given market and to create various trade pressure measures. These synthetic indicators became our key variables to measure the impulse of expanding Chinese trade activity.

We created our regional inequality database from first principles, as no ready-made dataset included the inequality measures we needed. To be able to span the mid-1990s to mid-2010s period, which was necessary to account for the pre- and post-Chinese accession periods, we used two sets of microdata as a basis: the European Community Household Panel (ECHP) for the 1990s and the European Union Statistics on Income and Living Conditions (EU-SILC) for the 2000s. Regional identifiers were missing for many years in the cases of Germany and the United Kingdom, so we also used their local household surveys, the German Socio Economic Panel (GSOEP) and the British Household Panel Survey (BHPS).

Year by year, we calculated our statistics using regional cross-sections of data and appended these to form a panel. We have chosen NUTS1 as the default regional unit, although we are forced to use country-level data in the case of the Netherlands, Portugal and Finland — see details in Appendix F.2 in Appendix F. From now on, we refer to these as 'regions'. Our choice of a regional unit is a pragmatic one, but it is also in line with the opinion of Boldrin and Canova (2001), which we share: for a region to be a meaningful unit of analysis, it 'has to be large enough to "convexify" undeniable human indivisibilities and micro fixed costs' – NUTS1 regions are just such entities. They are, on the other hand, sufficiently smaller than countries to make visible within-country heterogeneity regarding manufacturing and thus exposure to trade shocks.

Our data refer to individuals, and our target population is 25 to 60-years-olds to separate (otherwise important) peculiarities regarding the beginning and the end of a labor market career. We have used net household equivalized income, corrected for between-country price differences, as our income concept.[15] It absorbs the most shocks among all income types (see Benczúr et al., 2017, for recent EU-wide evidence) and includes components related to a wide array of events,[16] thus capturing a more complete, final effect. We have calculated the variance of the natural logarithm (log-variance) of income as our primary outcome, along with the Gini index and the 90/50 and 50/10 income percentile ratios.

Although we think that general household surveys are fit for supporting regional analysis of income inequality,[17] their sample size does not allow us to study the additional sectoral decomposition of employment at the regional level. Therefore, we have also used the EU Labour Force Survey (EU-LFS), which has larger survey samples, to calculate complementary statistics related to the labor market, in particular, the split of labor by industries and its various qualifications. We have worked similarly as with inequality-related data: we used individual data on the 25 to 60-year-old population to calculate several indicators of unemployment, employment and employment share of different subgroups of the employed. We shall also use one of these indices, the region's share of total national manufacturing employment, for the construction of our trade pressure indicators.

Ultimately, we obtain an unbalanced panel of 65 regions for the years 1994–2014, yielding an overall 1,348 observations on inequality. The statistics calculated rely on cells containing an average of 2,000 observations, 91 being the first percentile, 51 the minimum, 12,364 the 99th percentile and 15,018 the maximum number.[18]

Even though our database on inequality ends later than the one on trade statistics (in 2014 and 2011, respectively), these additional years on inequality turn out to be useful. As revealed in Section 5, the impact of trade pressure on income inequality can lag by up to three years.

15 The consideration of net income is relevant, as economic redistribution of income tends to increase when income inequality is rising (see Gozgor and Ranjan, 2017). Either automatic stabilizers or active economic policy reacting to shocks could counteract the initial impact increasing inequality of gross income.

16 These include changes in the level of wages, hours worked, employment status, the amount of taxes paid and transfers received by all household members, many of which are an adjustment margin for the individual.

17 See our discussion of 'representativity' in Appendix F.2.

18 At the country level, we have from about 3,500 (Belgium, Luxembourg) to 10,000 (Italy, Germany) observations, with an average of 5,500 per country before 2000, and 5,000 (Austria, Ireland, Luxembourg) to 24,000 observations (Italy), with an average of 10,000 per country, after 2000. Note that sample sizes are not proportional to population, e.g., the samples for the United Kingdom range from 5,500 to 10,000, while those for Luxembourg range from 3,300 to 8,000.

2.2. Econometric specification and estimation

Although stylized facts based on the raw figures presented in Appendix A already suggest a positive relationship between Chinese trade expansion and inequality in EU regions, there might be critical confounding processes and potential endogeneity at work, and controlling for them is essential to increase our confidence in the existence and strength of the effect. This subsection thus presents the econometric specifications under estimation together with various controls, whereas the next subsection discusses instrumental variables (IV) used to identify the Chinese trade pressure effect.

Concerning the basic functional form of the estimating equation, we follow regional analyses such as (Autor et al., 2013) and Balsvik et al. (2015) and rely mostly on simple (log-)linear specifications. In order to start discussing the estimation of the trade pressure from China, we express inequality in a region r in year t in the following general dynamic panel form:

$$\alpha(L)I_{r,t} = f\left(\boldsymbol{\mathcal{P}}_{r,t}, \boldsymbol{Z}_{r,t}\right) + \xi_{r,t}, \tag{1}$$

where:

$\alpha(L) = 1 + \alpha_1 L + \cdots + \alpha_k L^k$ stands for a lag polynomial with $k \in \mathbb{N}$ chosen based on the significance of lags;
$I_{r,t}$ is a measure of regional inequality;
f denotes a generic linear function that can also include lags, various transformations and interactions of its arguments while, potentially, dropping some of them, too;
$\boldsymbol{\mathcal{P}}_{r,t}$ is a vector of trade pressure indicators including the external pressure indicator $\mathcal{P}^{(X)}_{r,t}$, to be defined in Eq. (5); the domestic pressure indicator $\mathcal{P}^{(M)}_{r,t}$, to be defined in eq. (14); and their dynamic common factor $\mathcal{P}^{(F)}_{r,t}$, to be discussed shortly;
$\boldsymbol{Z}_{r,t}$ is a vector of additional variables, controlling for intermediate adjustments next to the initial pressure, accounting for economic channels of the impact, as well as controlling for various other region- and time-specific effects; and
$\xi_{r,t}$ is a zero mean error term satisfying the usual regularity conditions, but not necessarily uncorrelated with the pressure indicators in $\boldsymbol{\mathcal{P}}_{r,t}$ (or even with some components of $\boldsymbol{Z}_{r,t}$).

The log-variance of net household equivalized income will be the base measure of inequality. Since its distribution is skewed, we take its natural logarithm, using it as $I_{r,t}$, which also reduces the heteroskedasticity of errors in Eq. (1). We will also consider (logarithms of) the Gini index and the 90/50 and 50/10 income percentile ratios as the dependent variables. This is not only an additional robustness check, but also a study of whether the lower and upper parts of income distributions were affected similarly.

In Eq. (1), we use a generic linear function f instead of a fixed structure to allow for a flexible linear representation with many potential combinations of pressure, explanatory and control variables, some derivative indicators, as well as their lags. The most important cases are connected with the following alternative structures. First, we allow for different combinations of pressure indicators, including their common factor used as a single joint pressure indicator. For each region, it is derived from the state-space representation characterized in Appendix B.1. Second, we separate between the reduced and semi-reduced representations of the impact.

The reduced-form analysis aims at projecting (through instruments) inequality directly onto the pressure indicators (some or all components of $\boldsymbol{\mathcal{P}}_{r,t}$). In the basic form, it includes only the pressure indicator with individual and/or period effects,[19] but it can be extended to contain some additional controls accounting, among others, for general macroeconomic conditions such as the state of the business cycle and/or international competitiveness, some intermediate adjustments through trade reallocation between different markets, as well as other variables capturing economic structure, technological change, previous EU enlargements, intensity of migration, etc.

The semi-reduced representation aims instead at projecting (through the same set of instruments) inequality not directly onto the pressure indicators, but onto variables that could represent the economic channels of the impact. We aim at the structural characteristics, avoiding nominal variables, and use the regional unemployment rate, the regional labor share of those employed in manufacturing (out of all employed in a region), the share of high- and medium-high–tech goods in manufacturing exports, as well as the share of white collar workers and workers with higher education in manufacturing relative to analogous shares in all sectors in a region.[20] The unemployment rate and the manufacturing share measure the out-of-employment and inter-sectoral shift effects, whereas the last three indicators aim at capturing the pressure on inequality stemming from the changes within the manufacturing sector (intra-sectoral adjustments). We do not claim that this list of variables is exhaustive, but it is informative about the importance of economic processes

[19] It should be noted that individual effects are always included and compensate for the time-invariant regional characteristics influencing inequality, but which are difficult or impossible to quantify. The period effects were included only in the minimal specifications containing the trade pressure indicator alone. As the number of period effects is noticeable (nearly twenty), their presence substantially reduces the degrees of freedom and weakens the power of statistical inference. This was not problematic in the simplest specification with the pressure indicator alone (as will be presented in Table D2 in Appendix D), but became an issue whenever many other control variables were present in the specifications. Similar concerns prevented (Breemersch et al., 2017) from considering the period effects or trends altogether.

[20] The normalization with respect to all sectors intends to remove the potential general trends in an economy, e.g., because of the supply-side-induced changes connected with some general trends in education.

discussed in the Introduction, and it is admissible: the pressure indicators are insignificant if added to such specification.[21] This semi-reduced equation of channels, together with the corresponding separate equations determining the reaction of each of these variables to the initial pressure, will allow us to evaluate their importance in contributing to the total impact. A formal econometric characterization of the semi-reduced representation with its usage to derive the relative importance of the channels is explicated in Appendix B.2.

Our estimation method seeks to account for the dynamic features and, from that perspective, to improve upon previous ones that mostly used a variant of a multi-year difference estimator of (stacked) cross-sections. Our own approach involves estimating a dynamic panel regression model.[22] As the theoretical discussion of Bellon (2017) shows, the impact of trade liberalization over a shorter period might differ substantially from the longer impact because inequality can overshoot its steady-state level. It therefore becomes essential to allow for dynamic adjustments.

We estimate the model using a GMM estimator. As a base, we follow the two-step estimation strategy in first differences[23] of Arellano and Bond (1991), with second and third lags of the first differences of the dependent variable acting as the GMM instruments. The two-step approach relies on the estimator of the asymptotic variance–covariance matrix proposed in Windmeijer (2005), which both includes a small sample correction that benefits the second-stage estimates, e.g., by making the standard errors more precise, and is robust to arbitrary heteroskedasticity. Note that the two-step estimator is natural in our case, because the regional data we use have a known element of heteroskedasticity; we calculated them from individual data aggregating up to the regional level from different sample sizes, which leads to a varying precision of the estimates in different regions. Even though most of the results in the literature come from regressions weighted simply by the regional sample sizes underlying the aggregate data, we follow the advice of Dickens (1990) not to weight blindly, but to study the heteroskedasticity structure. Our results show that the assumption of heteroskedasticity driven only by the size of a region does not hold, with its contribution to the realized actual heteroskedasticity being relatively small. Thus, weighting only with the sample size would in fact be inefficient. Since we use a robust variance–covariance estimator along with two-step GMM, we do not apply such pre-weight.[24]

Aside from a few marginal cases, the estimated specifications seem to be sufficiently adequate at standard significance levels from the econometric point of view. First, the instruments' admissibility in terms of over-identifying restrictions based on orthogonality conditions (Sargan, 1958, and Hansen, 1982) is not rejected practically for all specifications of main interest. Second, the hypothesis of the absence of serial correlation of errors (Arellano and Bond, 1991) at higher lag orders than the first one was also not rejected, as required in the GMM estimation of dynamic panels with first differences. In addition to these statistics, both the number of cross-sections (regions) and the number of instruments will be reported in the tables, together with the total number of observations. In the baseline estimations, we set the number of instruments to about 70% of the number of cross-sections in order to comply with the rule of thumb (see, e.g., Roodman, 2009a,b) that the number of instruments should not exceed the number of cross-sections. In the robustness analysis in Section 5, it will be further reduced to about 33% and to below 10% of the number of cross-sections. Standard errors of estimated coefficients will be reported below each coefficient in brackets.

Our main interest in Eq. (1) lies in the sign, size, and empirical significance of the pressure indicators. To identify the impact of China's WTO accession, reducing also the potential endogeneity problem, we furthermore use regular instruments besides the previously discussed GMM instruments, as presented next. The instrumental approach is further necessary to get consistent parameter estimates under the presence of explanatory variables observed with errors; in our case, this is the estimated dynamic common factor.

2.3. Regular instruments

The GMM instruments defined previously are used to instrument for the lagged dependent variable in the dynamic panel. To deal with potential endogeneity issues of other determinants, we further used conventional instruments besides the GMM ones. Although (Balsvik et al., 2015) argues that pressure in the export markets is exogenous in the current setting of Chinese trade expansion, it is not guaranteed that inequality does not reflect, e.g., certain structural patterns of an economy that simultaneously determine both the level of inequality and the adjustment to the trade pressure at the same time, thus inducing endogeneity between the two.[25] In addition, the usage of instruments can also be helpful in reducing the bias due to omitted variables, provided that they and the instruments are orthogonal. There are many mechanisms driving inequality, such as autonomous changes in demand for final goods or labor; changes in the population structure in terms of age, sex, education; etc. However, they are less likely to be correlated with

[21] If a substantially significant channel, through which the pressure indicator affects inequality, were missing, the included pressure indicator would be expected to become significant.

[22] We use the `pgmm()` function from the plm package for R (see Croissant and Millo, 2008).

[23] The usage of purely first-differences–based estimates instead of the system GMM in our case has a specific advantage: as compared to the levels, the changes are less likely to be affected by the fact that we are using two different surveys to construct the panel of income inequality. Furthermore, the system GMM would substantially increase the number of instruments, creating pressure on the admissible number of instruments.

[24] Although using the estimated weights for heteroskedasticity pre-correction, obtained from a regression of squares of residuals on survey sample sizes and including an intercept, in the two-step GMM produced similar estimates to those obtained without such pre-weighting.

[25] Even in a simple framework with long differences, the results of the Wu–Hausman test indicate that the null hypothesis of exogeneity of trade pressure indicators is strongly rejected (see Table D8 in Appendix D).

the instruments that will be characterized shortly. Therefore, the IV strategy is already likely to reduce the potential acuteness of such misspecification. Nevertheless, in the robustness analysis in Section 5, we will also control for a number of additional factors.

Following the general idea of Acemoglu et al. (2016), we construct the regular instruments using China's trade with other countries. The following three types of instruments will be employed in the main estimations, exploring further the results' sensitivity to their variation in Section 5.

The first group of instruments of the Chinese trade pressure compares China's export performance to exports from countries that are geographically close to China and, at the same time, are the EU's three largest Asian trade partners besides China. It comprises the (logarithms of) ratios of China-to-India, China-to-Japan, and China-to-Korea exports, separating their total exports (WORLD) from their exports to countries belonging to the Organization for Economic Co-operation and Development (OECD). The total exports are best suited to capture total scale effects, while the OECD market, covering the most developed economies, could additionally allow for the potential specificity of trade with such countries, e.g., in terms of types and structure of products.

We use this simple split of markets, as refinement to finer groups (ASEAN, NAFTA, etc.) or even individual countries did not yield a sizable gain while creating pressure on the admissible number of instruments. The considered Asian exporters have faced the same changes (if any) in trade and shipping conditions from Asia to the rest of the world as China, which makes them a natural reference group.[26] These instruments thus correlate with the Chinese expansion of exports and, consequently, the external pressure, but not with the economic structure and inequality of the EU countries and their regions which we are interested in.

The just-discussed instruments of the first group capture general trends but are neither EU country-specific nor region-specific. By assuming that geographic similarity between countries/regions is to some extent informative about the similarity of pressure, we design the second and third groups of instruments to be specific to a particular region even though trade statistics are registered only at the country level. To achieve this, we exploit the varying distances of regions to their neighboring countries, as identified by distance between the two nearest regions of different countries. The main results reported hereafter will be based on the geodesic distances, but switching to distances by road or distances that take further economic factors into account did not change the qualitative picture (all these distances are taken from Persyn et al., 2019). The reciprocals of distances, normalized to add up to one[27] are used as region-specific weights to aggregate the simple relative Chinese trade pressure indicators observed in other EU15 countries (currently, EU14 and the UK).

The distance-based weighting is the same in the second and third groups of instruments, but the trade pressure indicators differ. The second group of instruments weights Chinese exports relative to the exports of each country from the (former) EU15 countries and thus aims at capturing the pressure in external markets. Meanwhile, the third group of instruments is obtained by weighting the ratios of imports to each of the (former) EU15 countries from China and India, China and Japan, and China and Korea. It thus evaluates whether imports to EU countries from China have increased relative to imports from the other Asian exporters that are the EU's three largest Asian trading partners besides China. In all cases, the logarithms of the defined quantities are employed to remove the dependence on scale differences and to reduce heteroskedasticity.

Formally, for each region indexed by r and period by t, the values of the three groups of instruments are defined as follows[28]:

$$W^{(1)}_{\cdot,t}(k,m) = \log\frac{X^{(m)}_{CN,t}}{X^{(m)}_{k,t}},\ k\in\{\text{IND, JPN, KOR}\},\ m\in\{\text{OECD, WORLD}\}, \tag{2}$$

$$W^{(2)}_{r,t}(m) = \log\left(\sum_{j=1,\ j\neq i^*(r)}^{N} w_{j,r}\frac{X^{(m)}_{CN,t}}{X^{(m)}_{j,t}}\right),\ m\in\{\text{OECD, WORLD}\}, \tag{3}$$

$$W^{(3)}_{r,t}(k) = \log\left(\sum_{j=1,\ j\neq i^*(r)}^{N} w_{j,r}\frac{M^{(j)}_{CN,t}}{M^{(j)}_{k,t}}\right),\ k\in\{\text{IND, JPN, KOR}\}, \tag{4}$$

where:

$N(=15)$ denotes the total number of countries from the former EU15;

$i^* := i^*(r)$ is a country index identified by a region index r, i.e., all region indices of a particular country are mapped to that index i^* of the country;

26 The results using China-to-US indicators were also explored (see Section 5 and Table D2 in Appendix D).

27 Note that the country of an analyzed region is forced here to have zero weight, i.e., it is omitted by construction.

28 It should be pointed out that we also considered further splitting of instruments, e.g., using a finer disaggregation of imports by countries in the second group of instruments, or further disaggregation of exports into additional sub-markets in the first and third groups of instruments. However, the results remained similar while creating pressure on the admissible upper limit of instruments (see the related discussion in Section 5 on the sensitivity of estimates to the number of instruments).

$w_{j,r} = \frac{\lambda_{j,r}}{\sum_{i=1, i \neq i^*(r)}^{N} \lambda_{i,r}}$, $\lambda_{i,r} = \frac{1}{D(i,r)}$, $i, j \in \{1, 2, \ldots, N\}$ are weights with $D(i, r)$ standing for the distance between a region indexed by r (from $i^*(r)$ country) and the closest region from a foreign country indexed by i;

$X_{CN,t}^{(m)}$ and $X_{k,t}^{(m)}$ are exports to the OECD countries or WORLD (as indexed by m) from China and either India, Japan, or South Korea (as indexed by k), correspondingly;

$X_{j,t}^{(m)}$ denotes exports to the OECD countries or WORLD from a former EU15 country indexed by $j \in \{1, 2, \ldots, 15\}$; and

$M_{CN,t}^{(j)}$ and $M_{k,t}^{(j)}$ are imports of the former EU15 country indexed by j from China and either India (IND), Japan (JPN), or South Korea (KOR), respectively.

In the baseline estimations, all three groups of instruments will be employed, whereas their partial use is also explored in Section 5 while studying the reduction of the total number of instruments and the projections using a single type of instruments. As we will also show in Section 5, the pressure impact can lag significantly by up to three periods; therefore, we allow up to three lags of the defined instruments too.

3. Inequality effects of the pressure in export markets

At the core of analyses looking at the effect of China's trade expansion on foreign labor markets is a pressure indicator. It is this indicator that summarizes and identifies the impact of trade expansion. In a regression framework, the parameter attached to this variable is the researchers' ultimate interest. In this section, we concentrate on its sign and significance, whereas Appendix E provides a stylized evaluation of its economic size by isolating the effect of China's trade pressure within the observed increase in regional inequality.

Trade pressure indicators have been in use since the seminal paper of Autor et al. (2013) in various forms. They usually sum the effect of growing Chinese exports on different markets, weighted by the importance of these markets. This importance is usually approximated by the share of the given market in the total trading activity of a sector within a region. The indicators increase with Chinese trade activity and decrease with the country's or region's output or exports. The markets considered can be external (i.e., export markets), such as in Autor et al. (2013) and Balsvik et al. (2015), or domestic, such as in Acemoglu et al. (2016). The indicator can incorporate both the gains in the Chinese market and the potential losses in others, such as (Autor et al., 2013), or look only at the latter, such as (Balsvik et al., 2015). It can rely on the level of Chinese exports, as did (Autor et al., 2013), or their value relative to the given country's own exports, as in Acemoglu et al. (2016). Finally, trade data are usually available at the national level and are projected to the regional level proportionally to employment, as in Autor et al. (2013). See Topalova (2010) and Kovak (2013) for an explanation of this practice in the general case of trade liberalization. In rare cases, such as in Balsvik et al. (2015), genuinely regional data are available, requiring no projection at all.

Besides the external (exports) and domestic (country's own) markets, where much of the literature has found negative effects, we will also distinguish the Chinese market, which could be expected to compensate for the pressure faced elsewhere. We do so using separate indicators in each case, starting, in this section, from the external pressure faced in export markets. We initiate from here for several reasons. First, as was highlighted in the Introduction, it is expected a priori to be the most important pressure component in the EU's case, and it will indeed turn out empirically to be the most relevant one. Second, even though we apply consistent parameter estimators, one can still doubt if results based on specifications that include the estimated dynamic common factors are sufficiently precise in our finite sample situation, thus preferring a more directly observed series. Third, it has a very clear break after 2001 (to be revealed in Fig. 1, which will be discussed shortly), which will be utilized to get a stylized prediction of the economic significance of the impact of China's accession on income inequality in Appendix E. Finally, it will allow us to highlight a number of important aspects in a simpler manner that will also be relevant for other indicators to be used later on.

Consequently, starting from a specification with a single (and the most important) determinant, we intend to show in this section that even the simplest specification points to the presence of a significant impact.

3.1. A viable pressure indicator for European regions

To develop our pressure indicator, we consider the best practice of the literature and constraints imposed by the data at hand. Because of its crucial role in Chinese trade expansion, we concentrate on the manufacturing sector. As our trade data are at the country level, we shall project them to the region level proportionally to the region's share in national manufacturing employment.[29] Thus we define our measure of the external pressure in export markets as follows:[30]

$$\mathcal{P}_{r,t}^{(X)} = \log\left(P_{i,t}^{(XC)} \underbrace{R_{i,r,t}}_{\text{projection}}\right), \quad \underbrace{P_{i,t}^{(XC)}}_{\text{country-pressure}} = \sum_{g} s_{i,g,t_0} \frac{X_{CN,g,t}}{X_{i,g,t}}, \tag{5}$$

[29] We prefer the region's share of national manufacturing as a projection variable because we did not detect any statistically significant influence on it from Chinese trade expansion, i.e., on the distribution of national manufacturing employment among regions—not to be confused with the share of manufacturing employment out of total employment in a region.

[30] Note that a region index r uniquely defines the respective EU15 country indexed by i; hence, we omit the country index i from the regional-level pressure indicator to simplify the notation of the variable on the left side of Eq. (5).

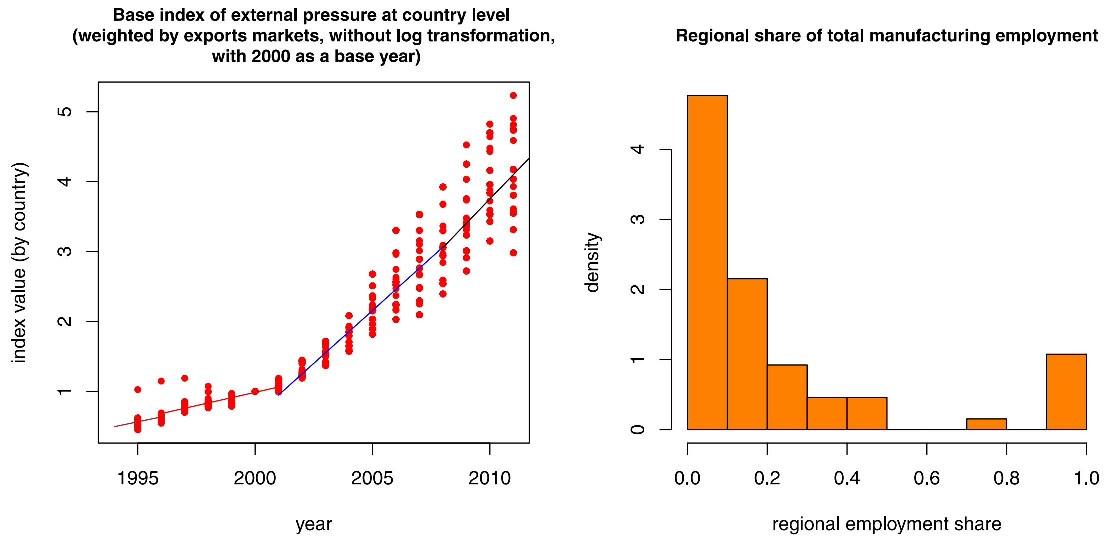

Fig. 1. Evolution of the country-level pressure in export market (left panel) and histogram of the regional projection term (right panel).

where:

$R_{i,r,t}$ denotes region r's share of total manufacturing employment in country i, year t;
s_{i,g,t_0} stands for the share of exports to market g out of total exports by country i in the pre-accession year t_0;
$X_{CN,g,t}$ represents exports from China to market g in year t, excluding country i if it belonged to region g;
$X_{i,g,t}$ are exports from country i to market g in year t.

The indicator defined in Eq. (5) thus has two parts. The country-level external pressure ($P_{i,t}^{(XC)}$) relates manufacturing exports from China to market g and exports from an EU15 country indexed by i to the same destination market, the resulting ratios being finally summed across all markets.[31] The index increases with the total value of China's exports and increases even more if such growth happens in a market that was important in the pre-accession year. Using China's exports in relative (rather than absolute) terms, we can take into account the changing export activity in a market: decreasing activity in country i will make the relative pressure larger, *ceteris paribus*. The usage of relative exports here has several advantages. First, levels cannot be fully informative about the pressure as, hypothetically, domestic exports could have been growing much faster than the Chinese exports, for instance, because the cheaper Chinese intermediate products might have been making domestic exports of final goods even more competitive in global markets. Second, the usage of ratios avoids the need to choose the denomination currency, which otherwise could have an effect on econometric estimations, although proper controlling and/or instrumenting might soften this aspect. Third, the ratio of nominal exports can be thought of as a stylized index representing the ratio of preference parameters in a hypothetical Cobb–Douglas utility function of goods from different markets,[32] which can be linked to the marketing literature on the importance of country of origin (see, e.g., Agrawal and Kamakura, 1999, Bloemer et al., 2009, and Magnusson and Westjohn, 2011, for overviews).

The left panel of Fig. 1 shows the evolution of country-level pressure in exports market $P_{i,t}^{(XC)}$ scaled with values at the base year 2000 for each country for visual comparability. The average trend line has a clear structural break after 2001.[33]

[31] The splitting of markets is defined taking into account the intensity of trade with the EU as well as the geographic or socioeconomic proximity of countries (see Section F.1).

[32] For a stylized motivation, suppose that the utility function at moment $t \in \mathbb{Z}$ is given by $U(t) = AQ_{CN,t}^{\alpha(t)}Q_{1,t}^{\beta_1(t)}, \ldots, Q_{n,t}^{\beta_n(t)}$, where $Q_{j,t}$ represents the quantity of products used from different markets $j \in \{CN, 1, 2, \ldots, n\}$, $n \in \mathbb{N}$, and the index CN identifies the Chinese origin of goods. Then, the ratio of nominal terms $\frac{P_{CN,t}Q_{CN,t}}{P_{j,t}Q_{j,t}} = \frac{\alpha(t)}{\beta_j(t)}$ can be informative about the development of preferences in terms of the origin of goods. In our case, we compare exports from various countries to a particular geographic entity (its imports from those countries), thus getting an insight about the developments in such hypothetical preferences. It is clear that this is a highly stylized motivation, as, besides the roughness of the aggregation level and ignorance of many other determinants, goods might be used not only by consumers, since there are intermediate products as well as re-exports. Nevertheless, it gives some intuition of what is behind the potential meaning of the employed ratios.

[33] As is shown in Figure E1 in Appendix E, the same break is also evident in the logarithmically transformed demeaned data of the country pressure without using any year-specific normalization.

Table 1
Base estimates with indicators of external trade pressure.

	Dependent variable: Income inequality (LVAR)							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Lag of the dependent variable	0.609*** (0.087)	0.507*** (0.063)	0.504*** (0.063)	0.507*** (0.064)	0.497*** (0.061)	0.496*** (0.064)	0.553*** (0.065)	0.489*** (0.065)
External pressure $\mathcal{P}_{i,t}^{(XC)}$ (country-level)		0.049** (0.022)						
External pressure $\mathcal{P}_{r,t}^{(X)}$			0.050** (0.022)	0.052** (0.023)	0.052** (0.023)	0.049** (0.024)	0.046* (0.025)	0.064** (0.031)
Real effective exchange rate (REER)				0.005 (0.021)		−0.005 (0.027)		0.002 (0.027)
Economic sentiment index (ESI)					−0.107 (0.075)	−0.110 (0.092)		−0.155 (0.113)
External pressure $\mathcal{P}_{r,t}^{(X)} * \mathbb{1}\{t \geq 2008\}$								0.040** (0.018)
$\mathbb{1}\{t \geq 2008\}$								0.017 (0.035)
P-val.(Hansen)	0.202	0.274	0.274	0.243	0.247	0.21	0.561	0.378
P-val.(AR1)	0	0	0	0	0	0	0	0
P-val.(AR2)	0.908	0.767	0.794	0.795	0.756	0.754	0.53	0.855
Number of regions	65	65	65	65	65	65	65	65
Number of instruments	46	46	46	46	46	46	46	46
Number of observations	1037	785	785	785	774	774	561	774

*p < 0.1.
**p < 0.05.
***p < 0.01.

To obtain region-specific external pressure in the export markets, we use a projection term $R_{i,r,t}$ defined as region r's share of total manufacturing employment in country i in year t.[34] Regional variation thus comes from the projection term, distributing 100% of the country-level pressure to the regions within a country[35] (in Section 5 we will also explore the sensitivity of results to fixing $R_{i,r,t} = R_{i,r,t_0}$ for $t > t_0 = 2000$). These data change only slowly over time[36] and have a skewed distribution – see the right panel of Fig. 1. Shares are typically below 50% – the only higher value belongs to Flanders (BE2), and values at 100% belong to countries aggregated to or defined as a single NUTS1 unit.

As a final step, we take the (natural) logarithm of the product of the country-level pressure and the projection term in Eq. (5) to account for the skewed distribution of the projection term[37] and to reduce the heteroskedasticity over time, which would otherwise translate to errors in the equations under estimation.

3.2. Base estimates and their sensitivity

Based on our motivating stylized facts and available evidence, we want to estimate the effect of the trade pressure in export markets defined in Eq. (5) on inequality. Table 1 contains the results of the basic estimations,[38] where the dependent variable is the logarithm of log-variance (LVAR) of income and only the external pressure indicator is under consideration. The results with other inequality metrics will follow shortly, and other pressure indicators will be considered in Section 4, whereas various robustness checks will be explicated in Section 5.

We find a significant first-order lag of the dependent variable across all specifications, which indicates quite substantial persistence in regional inequality that is somewhat larger in Column (1), when the pressure indicator with other factors is absent, and thus the conditioning here is only on the fixed effects.

34 One should take note that this is the distribution of manufacturing labor among the regions and not the distribution of total labor in a region among sectors.

35 We have also considered using the regional composition of manufacturing by its sub-industries as well using pressure indicators differentiated by exports linked to specific types of these particular industries, thus lending further variation to across-region pressure. This extension, however, was not viable due to the resulting massive number of missing observations at the detailed regional–sectoral level.

36 Regional labor share of total employment in manufacturing in a country is driven by factors other than the Chinese pressure indicators, which were insignificant in explaining them in a panel framework.

37 We have also explored other alternatives, including the projection of the logarithm of the country pressure using regional labor share, i.e., $log(P_{i,t}^{(XC)}) \cdot R_{i,r,t}$, but it had much worse adequacy in terms of the Hansen–Sargan test in the econometric estimations.

38 The tables were created using the stargazer package for R (see Hlavac, 2018).

Next, we first of all evaluate whether the significance of the impact erroneously stems only from the usage of the regional projection. Columns (2) and (3) thus present the results both with only the (logarithm of) country-level external pressure $\mathcal{P}_{i,t}^{(XC)} := \log(P_{i,t}^{(XC)})$ without any regional projection ($R_{i,r,t}$), and with the region-specific external pressure $\mathcal{P}_{r,t}^{(X)}$ (the term on the left in Eq. (5)), correspondingly. Given that the pressure indicator is significant in Column (2) already at the country level, it is clear that the significance obtained in the remaining columns with the regional-level indicator included does not stem from the employed projection alone.[39] Nevertheless, the usage of regional-level pressure will become beneficial in further cases containing many explanatory variables.

China's accession to the WTO took place about the same time when the euro was introduced: the euro was adopted on January 1, 1999, with coins and notes introduced into circulation in 2002 in twelve Euro Area countries. Although the euro was depreciating until 2001, later the trend reverted until 2008. Decreasing competitiveness due to the appreciation of the domestic currency could be an alternative explanation of the decreasing performance of European producers and exports. Therefore, in Columns (4) and (6), we extend the basic specification with the (logarithm of) real effective exchange rates of the former EU15 countries.

Another important aspect is that economic activity was accelerating during the 2001–2007 period in most of the former EU15 countries, which has certain implications for inequality (see, e.g., Barlevy and Tsiddon, 2006, and Sherman and Sherman, 2015). Allowing for such possibility, we use the (logarithm of the) Economic Sentiment Index (ESI) to proxy the country-specific state of a business cycle (see Columns (5) and (6) in Table 1).

However, both these variables turned out to be insignificant, barely changing the coefficient of the external pressure indicator. Given the instrumental nature of our estimation, this result is not unexpected. On the contrary, the parameter estimate of the external pressure $\mathcal{P}_{r,t}^{(X)}$ is highly significant in all specifications, indicating that a 10% increase in regional trade pressure (as given inside the brackets of Eq. (5)) yields about a 0.5% increase in the log-variance of income immediately (in the same year) and about a 1% increase in the long run.

An additional concern might be that these findings could be driven by the financial crisis and not the Chinese trade expansion.[40] To evaluate this concern, we present the results estimated with data only up to 2007 in Column (7) and, in Column (8), allowing for a structural break in 2008 using a dummy variable taking a value of one since 2008 and zero before that, i.e., introducing an indicator function $\mathbb{1}\{t \geq 2008\}$. We can convincingly see that the established effect of the external trade pressure is not driven only by the financial crisis. If anything, there is an interaction between the China pressure and the financial crisis, i.e., the financial crisis seems to augment the impact of Chinese trade pressure on inequality. First, this might be explained by the fact that the initial Chinese trade pressure shock resulted in a number of firms that were still active but barely surviving. The second shock, connected with the financial crisis, reinforced the pressure, raising it above a certain critical level, thus forcing the exit of such marginally functioning firms from the market. Second, during the upswing of economic activity before the crisis, workers who lost their jobs because of Chinese pressure could have been absorbed by other booming sectors, whereas they were likely to be among the first to be dismissed when the business cycle turned around after the financial crisis erupted.

3.3. Estimates with other inequality measures

The estimates so far have used the log-variance of income as the dependent variable, but there are other income inequality measures to consider. The Gini index is a popular aggregate metric of inequality that puts more weight on the middle of the distribution. In the current setup, it can be understood as a robustness check. Percentile ratios, on the other hand, can characterize different parts of the income distribution. We separate the lower and upper parts of the income distribution by considering the 50/10 and 90/50 income percentile ratios. Looking at the former is essential, as most employees in manufacturing are likely to be in the lower part of the income distribution.

Table 2 presents the results using alternative inequality measures as the outcome variable. LVAR stands for the already known results using the (logarithm of the) log-variance, and GINI represents the logarithm of the Gini index of inequality, while P5010 and P9050 denote the logarithms of the 50/10 and 90/50 income percentile ratios, respectively.[41] The left part of the results in Columns (1)–(4) is analogous to the one in Column (6) in Table 1, but with a varying dependent variable.[42] It is furthermore expanded in Columns (5)–(8) with additional specifications where insignificant control variables of macroeconomic conditions have been removed. Their comparison with Columns (1)–(4) reveals that, practically, there is no change in the coefficient of the pressure indicator. Hence, up until the robustness analysis, we will drop these control variables from further specifications.

[39] We should also point out that, in an analogous dynamic panel framework, there was no significant impact of the pressure indicator on the regional share of manufacturing labor itself.

[40] Although, at the EU level, the financial crisis resulted in convergence of income and income inequality across the EU countries (see Cabral and Castellanos-Sosa, 2019, and Kvedaras and Cseres-Gergely, 2020), the situation within countries and their regions might differ (see e.g. Martinez Turegano, 2020).

[41] Similar results also hold with variables without the logarithmic transformation, but we prefer the transformed ones, because the distributions of the initial variables were positively skewed.

[42] We use Column (6) of Table 1 as a baseline because the structural change of the impact connected with the financial crisis, considered previously in Column (8) of Table 1, was insignificant with other inequality indicators.

Table 2
Base estimates, various inequality indicators.

	Dependent variable:							
	LVAR (1)	GINI (2)	P5010 (3)	P9050 (4)	LVAR (5)	GINI (6)	P5010 (7)	P9050 (8)
Lag of dependent variable	0.496***	0.588***	0.375***	0.285***	0.504***	0.597***	0.376***	0.297***
	(0.064)	(0.064)	(0.051)	(0.098)	(0.063)	(0.066)	(0.051)	(0.089)
External pressure $\mathcal{P}_{r,t}^{(X)}$	0.049**	0.023**	0.034***	0.004	0.050**	0.023***	0.035***	0.008
	(0.024)	(0.009)	(0.010)	(0.007)	(0.022)	(0.008)	(0.008)	(0.006)
... its beta coefficient	[0.328**]	[0.377**]	[0.532***]	[0.093]	[0.330**]	[0.382***]	[0.510***]	[0.178]
	(0.162)	(0.154)	(0.161)	(0.167)	(0.146)	(0.128)	(0.142)	(0.148)
Real effective exchange rate	−0.005	−0.002	−0.002	−0.007				
	(0.027)	(0.008)	(0.009)	(0.007)				
Economic sentiment index	−0.110	0.021	−0.050*	0.010			−0.049*	
	(0.092)	(0.038)	(0.030)	(0.024)			(0.029)	
P-val.(Hansen)	0.21	0.368	0.721	0.565	0.274	0.43	0.78	0.579
P-val.(AR1)	0	0	0	0.001	0	0	0	0.001
P-val.(AR2)	0.754	0.247	0.292	0.612	0.794	0.234	0.292	0.59
Number of regions	65	65	65	65	65	65	65	65
Number of instruments	46	46	46	46	46	46	46	46
Number of observations	774	774	774	774	785	785	774	785

*p < 0.1.
**p < 0.05.
***p < 0.01.

Besides the usual statistics, including the parameter estimates and their standard errors provided in regular brackets, the square brackets in Table 2 report beta coefficients – the coefficients from the standardized regression – which are informative about the relative response of each variable in terms of standard deviations. The findings do seem to suggest that the effect does not come equally from the whole income distribution. First, the external pressure indicator has a positive effect in all cases but is insignificant for the upper percentile ratio P9050. Second, even compared with the significant LVAR and GINI coefficients, the beta coefficient of the external pressure alone is seemingly larger in the case of P5010, which is connected to the lower part of income distribution. This asymmetry suggests that the impact of the Chinese trade pressure on inequality seems to be concentrated at the lower part of income distribution. At the same time, the results that we see in summary inequality measures like LVAR and GINI are likely to be driven mostly by the lower half of the distribution. Nevertheless, the cut at the 50th and 10th income percentiles might not be ideal (or, potentially, time varying), because the pressure was found to be insignificant using P5010 with data before 2008 (not reported), whereas it was significant using the same data with the LVAR and GINI indices.

4. Importance of other markets, intermediate adjustments, and impact channels

So far, we have looked at the external pressure faced in export markets due to the expansion of Chinese exports, excluding both the Chinese market itself and the country of interest's own domestic market. Also, we have taken a look only at the basic equations without considering trade adjustment by firms in the affected country/region that can take place through various re-allocations of imports and exports between different markets, which, potentially, can reduce the pressure on firms and, therefore, also inequality. Finally, we have considered only the reduced-form specifications without evaluating the potential relevance of different impact channels. Now we sequentially lift these three restrictions, summarizing in this section the main findings of the detailed analyses presented in Appendices C.1–C.3.

Motivated by an idea similar to those of Autor et al. (2013) and Feenstra et al. (2019), in Appendix C we define, besides the external pressure, the additional indicators aimed at capturing the pressure faced in the domestic market and the potential expansion of exports in the Chinese market. The latter is insignificant in our data sample (see Table C1 in Appendix C.1). The domestic pressure has a (mildly) significant positive coefficient when considered individually but becomes insignificant (and of the incorrect sign) when included together with the external pressure, which is always significant. This points to the potential presence of multicollinearity, and indeed, the correlation coefficient between the external and domestic pressure indicators is almost 0.9. This empirical feature, together with the economic arguments explicated in the Introduction, motivated us to introduce a (dynamic) common factor of the external and domestic pressure, as defined in Appendix B.1, which becomes necessary to identify some additional intermediate adjustment effects to be discussed next. The estimations with the common factor component reiterate the previous findings obtained using the external pressure that the increase of income inequality in the EU15 regions due to Chinese pressure is significant (see Appendix C.1 for the empirical results with a detailed discussion).

Next, we take a look at the intermediate adjustments connected with simple reallocation of trade between the Chinese and third markets, which could potentially alleviate the pressure before any further economic adjustments were required in terms of amounts of production, labor, technology, wages, etc. Firstly, note that increased imports from China imply increased domestic competition with the country's total imports only if imports from other (third) markets remain at their previous level. If they are just substituted (crowded out) by the Chinese imports, *ceteris paribus*, the competitive pressure might remain the same from local firms' point of view. We label this as 'imports substitution'. Note that such substitution effects are connected solely with the domestic pressure and can moderate its effect, but cannot soften the external trade pressure faced globally in export markets.

The second intermediate adjustment we consider is related to the fact that more affordable imports from China can also induce local firms to rely more heavily on cheaper intermediate products from China. This can make the final goods produced in the EU more competitive globally, and we call such effect 'exports facilitation'.[43] Finally, if the fast-growing Chinese market becomes more attractive than markets of other countries, firms might want to reallocate their exports from those third markets to China. Although this is connected with the China option discussed previously, here we stress the switch from third (non-Chinese) export markets towards exports to China, instead of looking at exports to China from its market/demand perspective, and we call such effect 'exports reallocation'.

We find no significant influence on inequality stemming from the changing structure of exports in connection with 'exports reallocation' towards the Chinese market (see Table C2 in Appendix C.2). However, there are some indications of the presence of imports substitution and exports facilitation adjustments reducing the initial China trade pressure. Namely, when added individually, the coefficients of the imports substitution and exports facilitation indicators are significant and have the expected signs, although they become insignificant when included jointly, whereas the coefficient of the pressure indicator represented by the dynamic common factor of the external and domestic pressure stays significant.

So far, we have looked at the impact in a reduced form which can be connected with many economic mechanisms. Based on the discussion presented in the Introduction, we aim hereafter at evaluating the relative importance of the following three potential channels through which inequality changes: inter-sectoral shifts, intra-sectoral adjustments taking place within manufacturing, and increasing unemployment. For that purpose, we turn to a semi-reduced form and think about the impact of China's expansion as a two-stage process.[44] First, the initial trade pressure impulse affects some intermediate variables connected with different channels of the impact. Second, these transmission channels affect the outcome (inequality) itself, and we will take a look hereafter at the specific contribution of the components we can measure (see Appendix C.3 for details).

Namely, we regress inequality on the following 'channel' variables capturing the structural patterns of an economy: manufacturing employment's share of total regional employment, the regional unemployment rate, the share of medium and high technological intensity of exports, and two relative measures of the changing skills of the manufacturing workforce (see details and stylized facts in Section F.3 in the Appendix). The first relates the share of higher education degree holders among all workers in manufacturing to the same share in all sectors of the region, while the other relates the share of employees in white-collar jobs in manufacturing to the same share in all sectors of the region.

The inter-sectoral shift is connected with workers who are dismissed from manufacturing but find jobs in other sectors. *Ceteris paribus*, such a shift would be identified by a reduced share of workers in manufacturing without an increase in unemployment. Hence, the manufacturing share and the unemployment rate jointly identify the inter-sectoral shift and/or an increase in unemployment. We use the upgrading of the technological intensity of exports and the up-skilling of workers to identify the intra-manufacturing adjustments. Note that in all the cases to be considered next, we use the same instruments as previously in order to identify the changes in all of these variables connected with the trade pressure and not with some domestic developments.

Using the estimations presented in Tables C3–C4 in Appendix C.3, together with the methodological framework defined in Appendix B.2, we show that, initially, the largest adjustment took place through the inter-sectoral channel (the shrinking manufacturing sector). The unemployment channel became important only after the financial crisis, and though it increased the total impact, its share among the transmission channels is still smaller than that of the inter-sectoral shift. As time passed, the intra-sectoral adjustments became not only significant but also dominant, further increasing the total long-run impact of the trade pressure on inequality. Thus, over a longer span of time, the intra-sectoral adjustments became the main source of income inequality, overwhelming the initially predominant inter-sectoral shift.

5. Robustness checks and economic significance

In the sequel, we perform various robustness checks, which cover the estimations with varying numbers and types of instruments, lagging and period effects, and the usage of different regional projections, alternative estimators, inequality measures, intermediate adjustments, and a number of additional control variables. Here, we only discuss the main findings, referring further to Appendix D for the respective particular tables containing the estimation results.

First, we explore the importance of the number of instruments. In the basic estimation, the number of instruments made up approximately 70% of the number of cross-sections (regions). Shrinking the number of instruments by more than half (to about 33%) is considered in Table D1 (even further reduction to 9% will be discussed shortly). Practically,

[43] Something similar can also hold in terms of 'domestic facilitation', but we find no evidence of its significance when exports are subtracted from total production (see Appendix C.2).

[44] The methodology of such two-stage estimation-based evaluation of the importance of different variables is presented in Appendix B.2.

this leaves all the previously established results intact. The noticeable change in terms of significance is present only in Column (11) of Table D1, where the number of parameters under estimation is larger and only manufacturing's share and the technical intensity of exports remain significant, while other terms become insignificant. Even in this case, the point estimates of coefficients are very similar to those reported in Column (8) of Table C3 (see Appendix C.3).

Second, we investigate the importance of the composition of instruments in Table D2, at the same time reducing further the total number of instruments to just six. To achieve this we not only use the collapsed GMM instruments, but also include only one type out of the previously defined three variants of instruments given by $W^{(i)},\ i \in \{1, 2, 3\}$, as defined in Eqs. (2)–(4), and using a single third country for comparison with China's trade performance. First, Japan is employed as the largest economy and the EU trade partner of the considered three (Japan, Korea, and India). Next, although we expected theoretically in Section 2.3 that Asian countries would be more fit to form the instruments, Table D2 shows that similar results are obtained also when US trade data are used to form the instruments. In all cases, the coefficient of China's impact remains positive and significant. Thus, the finding of the increase in income inequality in the (former) EU15 regions due to the pressure from China's WTO accession does not seem to depend either on the number of instruments or on its particular composition.

Third, we check if the results are robust to the particular regional projection that we employed in Eq. (5). There, relying on the fact that the country-level China pressure was insignificant in explaining the variation in the regional share of manufacturing employment out of total national employment in manufacturing, we used the actually observed $R_{i,r,t}$ for the distribution of the pressure. In Table D3, we show further that our main findings are robust to imposing $R_{i,r,t} = R_{i,r,t_0}$ for $t > t_0 = 2000$ for projections. Namely, we used the shares in the pre-accession year for projections of the country-level trade pressure. The results here again remain very similar to the previously reported ones.

Fourth, as presented in Table D4 using the fully reduced form, the size and significance of the impact coefficient would even increase if one allowed for the lagging impact of the trade pressure and/or period effects.[45] Hence, the impact calculations that rely on the specification with the contemporaneous pressure can be considered more as a lower bound. Nevertheless, as discussed in Section 4 and Appendix C.1, the sharp increase in the coefficient of the pressure when the period effects are taken into account might also be induced by multicollinearity.

Fifth, we augment the reduced-form specification with a number of other control variables in Tables D5 and D6. In the former, we control for potential effects of technological development. For that purpose, the regional (log) levels and growth rates of patent applications (per million inhabitants) as well as linear and quadratic trends are included in addition to the trade pressure indicator. Furthermore, to check if there is some interaction between patents or their growth rate and Chinese trade pressure, we also include the respective interaction terms. Although the level of patents and the quadratic trend are also significant in Columns (1) and (6) of Table D5, the trade pressure indicator remains significant and even somewhat higher than without these controls. The patent growth rates are insignificant in Columns (3) and (4). A separate word is needed about the result presented in Column (2), where the interaction between the number of patents and Chinese pressure is added. The pressure indicator here becomes apparently insignificant. However, the pressure and the interaction terms are highly correlated by construction (with a correlation coefficient of 0.95). Furthermore, the coefficient of the interaction term is tiny, with a much higher standard deviation, which would suggest its irrelevance. If it is dropped, one gets back to the situation depicted in Column (1).

Table D6 checks the sensitivity of the main results when other control variables are added. First, the real effective exchange rate and economic sentiment index are added in Columns (1) and (2), which previously were considered in Tables 1 and 2 only with the external pressure ($\mathcal{P}^{(X)}_{r,t}$), while here they are used with the dynamic common factor ($\mathcal{P}^{(F)}_{r,t}$). These additional controls are again insignificant. In Column (3), the (logarithm of the) share of service exports is included for a rough evaluation of whether the potential reorientation to service exports could have softened the pressure faced in goods markets. Although the sign is negative, it is insignificant. Next, Columns (4)–(6) briefly investigate the hypothesis that the EU's expansion with the CEE countries in 2004 caused the inequality increase and not China's WTO accession: first, by creating a similar trade pressure, and, next, because the free movement of labor increased migration flows that created downward pressure on the wages of the poorest in the relatively richer EU15 countries, thus increasing inequality there. In Columns (4) and (5), we indeed see a significant impact, with the latter confirming the just-explicated hypothesis, but the Chinese trade pressure indicator remains highly significant. It apparently loses significance when the interaction term is introduced between the trade pressure with the year 2004 in Column (6). However, it is not because of the EU's expansion in 2004 with the CEE countries, but because the Chinese pressure's increasing impact on income inequality in the (former) EU15 countries only began to realize significantly since 2002 (see Column (7) in Table D6), i.e., after China's accession to the WTO. Namely, when an indicator of post-Chinese accession is included, taking a value of one since 2002 and zero otherwise, the indicator of post-CEE accession, which takes a value of one only since 2004, becomes insignificant. Furthermore, the size of the estimated impact here, i.e., in the aftermath of China's accession, is again larger than that witnessed using the whole sample without taking this structural break into account.

Sixth, in Table D7, we evaluate the robustness of the findings regarding trade openness, allowing for heterogeneity of the relationship in terms of country exposure to international trade. First, we divide the sample of countries into two

[45] Note that to keep the number of instruments at the admissible level when period effects are included, we proportionally shrink the number of other GMM and regular instruments.

groups, with trade openness below or above the median international trade-to-GDP exposure in the early nineties.[46] As the respective columns (1) and (2) reveal, the point estimate is larger in the group with openness above the median, but it is insufficiently significant due to the sizeable reduction in the number of observations.[47] As these results might further depend on a particular split level, we next include an interaction term of openness with the pressure indicator in columns (3)–(5). In all specifications, the interaction term is highly significant, while all other terms are quite insignificant.[48] To see the effect of different cut-offs, we look at the marginal impact of pressure on inequality at various levels of trade openness. We base our calculations on the last specification in column (5) and use the actual quartiles of openness observed in EU15 countries during the analyzed period since China's WTO accession, that is, 2001–2011; see the sub-table presented just below Table D7. The estimates are typically somewhat larger than those in the baseline of Table 1[49] thus suggesting again that the basic valuation might be somewhat conservative.

Seventh, Table D8 reports the IV estimation results based on long differences. This method has been applied to several previous studies on individual countries in which hundreds or even thousands of cross-sectional observations were available. For comparability, the GMM results are produced that make estimates using panel data with the same instrument of China's trade pressure, as in the case with IV estimations with long differences. The GMM estimates of the long-run impact remain similar. The long differences also reconfirm the previously observed qualitative pattern: the domestic pressure is occasional, whereas the external pressure and the common factor of the two pressure indicators are significant in all considered variants of the long difference estimations. However, the IV estimations with long differences produce substantially larger point estimates that also have much larger standard errors.

These observations are in full agreement with our additional simulations that replicate the stylized patterns observed in the trade pressure indicator in the left panel of Figure E1 and the regional dynamics of income inequality in the top-left panel of Figure E2.[50] They reveal a large variance of long difference estimators in situations where the number of observations and the coefficient of the lagged dependent variable are rather moderate.[51]

Nevertheless, apart from a single case of domestic pressure, the estimates based on the long differences are significant and could suggest that China's impact on inequality might be even larger than the impact derived using the panel GMM estimates. We still favor our results based on the dynamic panel GMM estimator because large-point estimates obtained from long differences imply that the inequality in EU countries would have sharply decreased after 2001 had China not acceded to the WTO. This implication does not seem very plausible.

We finalize the discussion in this section with a brief summary of the economic significance of the impact, while further details and simulation results are provided in Appendix E. To quantify the impact, we exploit the structural break observed in the country-level pressure indicator, extending the trend-line observed before 2002 to after China's WTO accession period (see Figure E1 in Appendix E), which we use as a counter factual baseline. Then, exploiting the fully reduced estimate given in Column (8) of Table 1, we calculate the dynamic path of the impact as defined by eqs. (25)–(26) in Appendix E. The dynamics of the predicted absolute increase in inequality are plotted in the top-right panel of Figure E2 using a box-plot of regional values, whereas the ranking of all regions by the predicted increase in 2011 is presented in the top panel of Figure E3. The comparison of the calculated increase that we assign to China's impact with the actual change in inequality since China's WTO accession is presented in the bottom panel of Figure E2: the left side plots the box-plots of the two, whereas the right side plots the ratio of the respective medians. The median of the actual increase in regional inequality since China's WTO accession in 2001 was about 0.03 and 0.06 in 2007 and 2011, respectively.[52] The median of the predicted increase in inequality levels due to Chinese trade pressure is 0.013 and 0.027 in 2007 and 2011, correspondingly. This constitutes about 40% of the actual increase. If compared to a hypothetical counter-factual of 'no-accession baseline', the log-variance of income is typically (the median across regions) larger by 5% in 2007 and by 9% in 2011, ranging, in 2007, from 2% in Spain to 7% in the UK, and, in 2011, from 5% in Spain to 13% in Ireland.

These results rely on the reduced-form specification that takes into account the presence of the structural break in the impact due to the financial crisis, as presented in Column (8) of Table 1. If one disregarded it, using the simplest specification as in Column (3) of Table 1, the impact up until 2007 would remain about the same (just marginally smaller), whereas that in 2011 would be smaller by about half of the currently reported difference between impact values in 2011 and 2007.

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46 The median is calculated from all EU15 countries using the 1990–1992 data, with similar results obtained when other base periods are used to calculate the median. We thank an anonymous referee for the idea of considering the heterogeneity in terms of openness.

47 The group of more open economies consists mainly of small countries (Denmark, Ireland, Luxembourg, etc.) that have only a few or just one NUTS1 region. Similar results are obtained with a single equation in which the pressure parameter is allowed to change for the group of more open economies.

48 Analogous results hold for both the split and the interaction analyses, also when the common factor of internal and external pressure is used instead of only the external pressure indicator employed here.

49 The countries with extreme openness (with the maximum openness in Luxembourg) might also fail to be informative for our purpose not only because of the large share of re-exports but also because of specific exports, such as financial services in Luxembourg.

50 Both the simulation results and the R code are available upon request from the authors.

51 The total number of observations in our long difference estimations ranges between barely 54 and 124, whereas the number of panel observations ranges between 296 and 1037. Also note that Hahn et al. (2007) turn to the long difference estimator to deal with situations when the coefficient of the lagged dependent variable is close to unity (in our case it is about only 0.5). In contrast, their simulations show no improvement in the moderate values of these parameters, even when the cross-sectional dimension (N=100) is larger than in our study.

52 Here, we look at the pre-crisis year and the last available year, where we can measure the impact.

6. Summary and conclusions

''The only obligation for WTO Members is that they must accord China so-called permanent MFN ('most favoured nation') status, entitling it to be treated in the same way as every other WTO Member, unless exceptions are specified in the protocol of accession. As the EU has always accorded China this status in any event, there will be virtually no practical impact'' (see Snyder, 2009, p. 1069).

This prediction turned out to be incorrect, as it did not take into account the fact that the change in conditions of trade between China and third markets outside the EU affects the demand of goods and services exported from the EU to those third markets. Furthermore, the large increase in the amount of total output due to the global expansion of Chinese exports reduced the unit cost of production, thus allowing for a competitive improvement in terms of lower prices, even in the markets where there were no changes of formal trade conditions in terms of tariff and/or non-tariff barriers.

China's accession to the WTO exerted substantial pressure on producers in the EU. They faced intense Chinese competition in not only the domestic but also the export markets globally. The induced adjustments of exports and domestic production also have implications for the labor markets, partly because of directly changing total demand for labor, especially in the manufacturing sector, partly because of changing demand for different skills needed in the new environment. This structural change may create winners and losers, potentially resulting also in higher income inequality.

Using net household equivalized income adjusted for purchasing power differences, we show that China's accession indeed had a statistically significant positive impact on income inequality in the (former) EU15 regions. The estimated impact is concentrated in the median to lower tail of regional income distributions, which is similar to findings obtained by Basco et al. (2017) for France.

We find that the most significant impact on inequality stems from the external pressure faced in export markets and not the domestic one. As a consequence, we would not recommend evaluating the impact of Chinese trade pressure in individual EU countries by considering solely the domestic pressure. In our sample, there is no significant indication that the growing Chinese market would compensate for the consequences of the pressure faced elsewhere. However, we obtained some evidence that the substitution of imports from third countries by Chinese goods and the use of more affordable intermediate products from China relative to the EU's exports softened the initial pressure on income inequality.

Initially, the established increase in inequality within the EU regions is dominated by the inter-sectoral shift, manifesting through the shrinking shares of labor employed in manufacturing (without a significant increase in unemployment). The unemployment rate became a significant contributor to inequality only after the financial crisis, which might indicate either that the initial shock due to the WTO accession was augmented further by the financial crisis or that the trade pressure's impact on inequality varies along the business cycle. Over the longer period, the intra-sectoral (intra-manufacturing) adjustments became the largest contributor to income inequality through the technological upgrading of exports (and production) together with the up-skilling of the employed labor.[53]

As a simple simulation reveals, the established impact is not only statistically significant but also economically important. In relative terms, as compared with a hypothetical 'no-accession baseline', the log-variance of income is typically 5% larger in 2007 and 9% larger in 2011. In absolute terms, the median of the actual increase in regional inequality since China's WTO accession in 2001 was about 0.03 and 0.06 in 2007 and 2011, respectively. The predicted median absolute increase in inequality due to Chinese trade pressure was 0.013 and 0.027 in 2007 and 2011, correspondingly. This constitutes about 40% of the actual increase. It should be pointed out that inequality, in terms of the considered household equivalized income, tended to decrease in the EU regions before 2001, while the trend reverted afterward.

The largest absolute increase was determined for regions from the UK, Belgium, Italy, Austria, and Ireland. The largest impact relative to the no-accession baseline was observed in Ireland, Denmark, Belgium, and the UK. Consistently with (Dauth et al., 2017), the predicted impact for German regions is typically among the smallest ones.[54] We find the UK regions to be among the most affected ones, which reinforces the likelihood of the results in Bell and Machin (2016) and Colantone and Stanig (2018) on the potential importance of the China shock for Brexit.

Our findings yield some insights that are potentially relevant for economic policy. First, the China case highlights that economic policy decisions cannot be based on bilateral considerations and that their proper evaluation should take into account a broader context in the globalized world. It is necessary to account for changes in all relevant markets, thereby evaluating the respective interactions and repercussions. Second, despite the theoretical suggestions and empirical findings by Gozgor and Ranjan (2017) that economic redistribution tends to increase when income inequality is rising, it appears that income redistribution was insufficient during the post-China accession period to compensate for its pressure on inequality in the EU. Therefore, similar shocks might call for additional ex-post compensating mechanisms or transition funds[55] and actions that smooth out and facilitate the absorption of laid-off workers. Such instruments are important not only for providing direct income support but also for facilitating job matching. Distributing the job search over a longer period reduces the initial pressure on wages and eases the retraining, requalification, and, therefore, the inter-sectoral shift of workers, which we established to dominate the increase in inequality initially. A more homogenous (and high-skill-oriented) composition of labor would further soften the skill premium increase. This effect, at least partially,

53 The upskilling of workers is consistent with negative employment effects for low-skilled workers observed in Norway (see Balsvijk et al. 2015).

54 Although, contrary to our results, these authors found no significant increase in inequality in Germany due to China's impact.

55 The Just Transition Fund supporting the EU Green Deal is an example of a specific fund of a similar kind already in place.

underlies the intra-sectoral contribution to inequality that we established to dominate the long-term impact. Finally, ex-ante policies that would increase the resilience of firms to any kind of external shocks on competitiveness through, for example, continuous innovation and labor upskilling would allow absorbing such shocks more smoothly and flexibly.

Exploring several other questions is left for future research. The impact on income inequality can be heterogeneous due to structural differences of economies, their institutions, and the economic policies put in place. Dissecting the contribution of each of these is key from a policy point of view. Deeper analysis of the cumulative impact of different shocks (e.g., China's accession, a financial crisis, Covid-19) is needed for a similar reason. Precise identification of the relevance of interactions between the business cycle state, sectoral composition, and sequences of various shocks can help to anticipate future shocks and target policy interventions better.

Declaration of competing interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper. The opinions expressed are those of the authors only and should not be considered as representative of the European Commission's official position.

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Supplementary material: Appendices A–F

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