

Kvedaras, Virmantas; Cseres-Gergely, Zsombor

**Working Paper**

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

JRC Working Papers in Economics and Finance, No. 2020/1

**Provided in Cooperation with:**

Joint Research Centre (JRC), European Commission

*Suggested Citation:* Kvedaras, Virmantas; Cseres-Gergely, Zsombor (2020) : China's WTO accession and income inequality in European regions: External pressure and internal adjustments, JRC Working Papers in Economics and Finance, No. 2020/1, ISBN 978-92-76-18590-1, Publications Office of the European Union, Luxembourg, <https://doi.org/10.2760/752376>

This Version is available at:

<https://hdl.handle.net/10419/227664>

**Standard-Nutzungsbedingungen:**

Die Dokumente auf EconStor dürfen zu eigenen wissenschaftlichen Zwecken und zum Privatgebrauch gespeichert und kopiert werden.

Sie dürfen die Dokumente nicht für öffentliche oder kommerzielle Zwecke vervielfältigen, öffentlich ausstellen, öffentlich zugänglich machen, vertreiben oder anderweitig nutzen.

Sofern die Verfasser die Dokumente unter Open-Content-Lizenzen (insbesondere CC-Lizenzen) zur Verfügung gestellt haben sollten, gelten abweichend von diesen Nutzungsbedingungen die in der dort genannten Lizenz gewährten Nutzungsrechte.

**Terms of use:**

*Documents in EconStor may be saved and copied for your personal and scholarly purposes.*

*You are not to copy documents for public or commercial purposes, to exhibit the documents publicly, to make them publicly available on the internet, or to distribute or otherwise use the documents in public.*

*If the documents have been made available under an Open Content Licence (especially Creative Commons Licences), you may exercise further usage rights as specified in the indicated licence.*



<https://creativecommons.org/licenses/by/4.0/>

## JRC TECHNICAL REPORTS

# China's WTO accession and income inequality in European regions

*External pressure and  
internal adjustments*

Kvedaras, V.

Cseres-Gergely, Zs.

2020

*JRC Working Papers in Economics and Finance, 2020/1*



This publication is a Technical report by the Joint Research Centre (JRC), the European Commission's science and knowledge service. It aims to provide evidence-based scientific support to the European policymaking process. The scientific output expressed does not imply a policy position of the European Commission. Neither the European Commission nor any person acting on behalf of the Commission is responsible for the use that might be made of this publication. For information on the methodology and quality underlying the data used in this publication for which the source is neither Eurostat nor other Commission services, users should contact the referenced source. The designations employed and the presentation of material on the maps do not imply the expression of any opinion whatsoever on the part of the European Union concerning the legal status of any country, territory, city or area or of its authorities, or concerning the delimitation of its frontiers or boundaries.

**Contact information**

Name: Virmantas Kvedaras

Address: Via E. Fermi, 2749, I-21027 Ispra (VA), ITALY

Email: [virmantas.kvedaras@ec.europa.eu](mailto:virmantas.kvedaras@ec.europa.eu)

Tel.: + 39 0332 783517

**EU Science Hub**

<https://ec.europa.eu/jrc>

JRC120455

PDF

ISBN 978-92-76-18590-1

ISSN 2467-2203

doi:10.2760/752376

Luxembourg: Publications Office of the European Union, 2020

© European Union, 2020



The reuse policy of the European Commission is implemented by the Commission Decision 2011/833/EU of 12 December 2011 on the reuse of Commission documents (OJ L 330, 14.12.2011, p. 39). Except otherwise noted, the reuse of this document is authorised under the Creative Commons Attribution 4.0 International (CC BY 4.0) licence (<https://creativecommons.org/licenses/by/4.0/>). This means that reuse is allowed provided appropriate credit is given and any changes are indicated. For any use or reproduction of photos or other material that is not owned by the EU, permission must be sought directly from the copyright holders.

All content © European Union, 2020

How to cite this report: Kvedaras, V., and Cseres-Gergely, Zs. (2020) China's WTO accession and income inequality in European regions: External pressure and internal adjustments, *JRC Working Papers in Economics and Finance, 2020/1*, Publications Office of the European Union, Luxembourg, 2020, ISBN 978-92-76-18590-1, doi:10.2760/752376, JRC120455.

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

VIRMANTAS KVEDARAS<sup>1</sup> AND ZSOMBOR CSERES-GERGELY<sup>1</sup>

<sup>1</sup>*European Commission, Joint Research Centre (JRC), Ispra, Italy*

5 April 2020

## 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.

**JEL Classification:** D31, D63, F16, F61

**Keywords:** China, EU, globalization, income, inequality, regions, trade

**Disclaimer:** The opinions expressed are those of the authors only and should not be considered as representative of the European Commission's official position.

For fruitful discussions and constructive suggestions, we thank Péter Benczúr, Jose Manuel Rueda Cantuche, Claudio Deiana, Péter Harasztosi, Juan Manuel Valderas Jaramillo, Bálint Menyhért, Balázs Muraközy, Frank Neher, Elisa Tosetti, Alessandra Tucci, as well as participants of the MKE 2017 annual conference in Budapest and seminars at the JRC in Ispra and Seville. The usual disclaimer applies.

# 1 Introduction

China's accession to the World Trade Organization (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 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 (2017) and Feenstra *et al.* (2017) 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,<sup>1</sup> 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).

---

<sup>1</sup>As shown, e.g., by Broda and Weinstein, 2006, Halpern *et al.*, 2015, and Lakner and Milanovic, 2016, respectively.

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 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. 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.* (2017) 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.<sup>2</sup> On the other hand the trade conditions between

---

<sup>2</sup>“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).

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.

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.<sup>3</sup> 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

---

<sup>3</sup>Besides the external pressure on third markets and the domestically faced pressure, we also separated the China market as a particular EU exports market, in order to see if it has some specific importance, but it was insignificant in our sample.

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.<sup>4</sup> 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 and vertical specialization, as in Feenstra and Hanson (1996) 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) and Harrigan and Reshef (2015). 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.

---

<sup>4</sup>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.



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 A13 reveals (see Appendix E.3), manufacturing and industry in general are among the sectors in the EU that typically have low intra-sectoral wage inequality.<sup>5</sup> 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<sup>6</sup> within the regions of the former EU15 countries<sup>7</sup> (currently, EU14 and the UK) by constructing from different sources the comparable inequality and trade pressure measures covering the period from 1996 to 2014. Besides the fact that the number of observations with only fifteen cross-sections at the country level was simply insufficient for our econometric analysis,<sup>8</sup> 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

---

<sup>5</sup>Similar findings are also presented in Martinez Turegano (2020) using a different income concept.

<sup>6</sup>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.

<sup>7</sup>The countries are Austria, Belgium, Denmark, Germany, Greece, Spain, Finland, France, Ireland, Italy, Luxembourg, the Netherlands, Portugal, Sweden and the United Kingdom.

<sup>8</sup>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.

quite substantially by regions.<sup>9</sup> 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<sup>10</sup> level of the EU’s territorial disaggregation<sup>11</sup>.

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 A12 in Appendix E.2). At the same time, China’s relative performance in terms of 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 analysis of the initial pressure is closest to the regional approach of Balsvik *et al.* (2015) but also accounts for various intermediate adjustments and uses a dynamic panel estimation method with additionally constructed regional trade pressure instruments, discussed in Section 2. We consider the trade pressure from China faced by EU producers in various markets, separating the external exports 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

---

<sup>9</sup>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.

<sup>10</sup>Here, NUTS abbreviates the Nomenclature of Territorial Unites of Statistics.

<sup>11</sup>Most of the regions are at the NUTS1 level with only a few exceptions – see the next section and Appendix E.2 for details.

importance of the external pressure, we start the presentation of findings from it while gradually introducing various other markets and adjustment effects afterwards.

The structure of the paper follows the line of thought outlined above. We delegate much of the technical detail to the Appendices. Appendix A presents the motivating plots discussed in the Introduction. Appendix B 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 C 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, and an evaluation of some alternative econometric approaches to the measurement of the impact. Finally, Appendix D presents a stylized calculation of the implied inequality increase based on the simplest reduced-form specification of the impact measurement. Appendix E describes the trade, inequality, and employment data used in the paper. It also presents the stylized facts and respective plots revealing the basic relations between inequality and various explanatory variables.

## 2 Data and the econometric approach

### 2.1 Data

Our analysis rests on an unbalanced panel of 61–65 regions<sup>12</sup> 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 E.2). The dataset brings together our impulse, outcome and intermediary variables, that is, measures of

---

<sup>12</sup>Varying by different years due to the data (un)availability.

trade, inequality and employment structure, as well as other factors. A brief description of the dataset follows shortly, with further details provided in Appendix A.

Our data of the trade drivers rely on the outcome of the joint OECD–WTO Trade in Value Added (TiVA) initiative<sup>13</sup>. 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 a clear advantage in our case. For certain trade pressure indicators to be defined succinctly, we need a 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

---

<sup>13</sup>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.

Household Panel (EHP) 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 Section E.2 in Appendix A. 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. 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,<sup>14</sup> 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<sup>15</sup>, their sample size does not allow us to study the addi-

---

<sup>14</sup>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.

<sup>15</sup>See our discussion of ‘representativity’ in Appendix E.2.

tional sectoral decomposition of employment at the regional level. Therefore, we have also used the EU Labour Force Survey, 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 99<sup>th</sup> percentile and 15,018 the maximum number.<sup>16</sup>

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 6, the impact of trade pressure on income inequality can lag by up to three years.

## 2.2 Econometric specification and estimation

Although stylized facts based on the raw figures presented in Appendix A already suggest a positive relationship between the Chinese trade pressure and inequality, 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 used to identify the Chinese trade pressure effect.

---

<sup>16</sup>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.

Concerning the basic functional form of the estimating equation, we followed 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:

$$\alpha(L)I_{r,t} = f(\mathcal{P}_{r,t}, \mathbf{Z}_{r,t}) + \xi_{r,t}, \quad (1)$$

where:

$\alpha(L) = 1 + \alpha_1 L + \dots + \alpha_k L^k$  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;

$\mathcal{P}_{r,t}$  is a vector of trade pressure indicators including the external pressure indicator  $\mathcal{P}_{r,t}^{(X)}$ , to be defined in eq. (5); the domestic pressure indicator  $\mathcal{P}_{r,t}^{(M)}$ , to be defined in eq. (6); and their dynamic common factor  $\mathcal{P}_{r,t}^{(F)}$ , to be discussed shortly;

$\mathbf{Z}_{r,t}$  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  $\mathcal{P}_{r,t}$  (or even with some components of  $\mathbf{Z}_{r,t}$ ).

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 C.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  $\mathcal{P}_{r,t}$ ). In the basic form, it includes only the pressure indicator with individual and/or period effects,<sup>17</sup> 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,

---

<sup>17</sup>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 A2 in Appendix B), 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.



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.<sup>18</sup> 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 discussed in the Introduction, and we will show that it is admissible: the pressure indicators are insignificant if added to such specification. 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 C.2.

Our estimation method seeks to improve upon previous ones that mostly used a variant of a multi-year difference estimator of (stacked) cross-sections (see Appendix C.3 for a discussion of this approach with its empirical application to our case). Our own approach involves estimating a dynamic panel regression model.<sup>19</sup> As the theoretical discussion of Bellon (2015) shows, the impact of trade liberalisation 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 generalized method of moments (GMM) estimator. As a base, we follow the two-step estimation strategy in first differences<sup>20</sup> of Arellano

---

<sup>18</sup>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.

<sup>19</sup>We use the `pgmm()` function from the `plm` package for R (see Croissant and Millo, 2008).

<sup>20</sup>The usage of purely first-differences-based estimates instead of the system GMM in our case has

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.<sup>21</sup>

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)

---

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.

<sup>21</sup>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.

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 6, 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 the 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 on the exports market 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. 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 the instruments that will be characterized shortly. Therefore, the instrumental variable (IV) strategy is already likely to reduce the potential acuteness of such misspecification. Nevertheless, in the robustness analysis in Section 6, 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 6.

The first group of instruments of the Chinese trade pressure compares China’s exports 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.<sup>22</sup> 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

---

<sup>22</sup>The results using China-to-US indicators were also explored (see Section 6 and Table A2 in Appendix B)

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 the 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,<sup>23</sup> 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

---

<sup>23</sup>Note that the country of an analyzed region is forced here to have a zero weight, i.e., it is omitted by construction.

groups of instruments are defined as follows<sup>24</sup>:

$$W_{.,t}^{(1)}(k, m) = \log \frac{X_{CN,t}^{(m)}}{X_{k,t}^{(m)}}, \quad k \in \{\text{IND, JPN, KOR}\}, \quad m \in \{\text{OECD, WORLD}\}, \quad (2)$$

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

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

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;

$w_{j,r} = \frac{\lambda_{j,r}}{\sum_{i=1, i \neq i^*(r)}^N \lambda_{i,r}}$ ,  $\lambda_{i,r} = \frac{1}{D(i,r)}$ ,  $i, j \in \{1, 2, \dots, N\}$  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;

$X_{CN,t}^{(m)}$  and  $X_{j,t}^{(m)}$  are exports to the OECD countries or WORLD from China and from a former EU15 country indexed by  $j \in \{1, 2, \dots, 15\}$ , correspondingly; 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 6 while studying the reduction

---

<sup>24</sup>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 6 on the sensitivity of estimates to the number of instruments).

of the total number of instruments and the projections using a single type of instruments. As we will also show in Section 6, the pressure impact can lag significantly 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 D 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., exports 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 Figure 3.1, 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 D. 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.<sup>25</sup> Thus we define our measure of the external pressure in exports markets as follows<sup>26</sup>:

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

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 equation (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<sup>27</sup>. 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

---

<sup>25</sup>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.

<sup>26</sup>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).

<sup>27</sup>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 E.1).

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,<sup>28</sup> 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 Figure 3.1 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.<sup>29</sup>

To obtain region-specific external pressure in the export markets, we use a projection term  $R_{i,r,t}$  defined region  $r$ 's share of total manufacturing employment in country  $i$  in year  $t$ .<sup>30</sup> Regional variation thus comes from the projection term, distributing 100% of the country-level pressure to the regions within a country<sup>31</sup> (in Section 6, 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

---

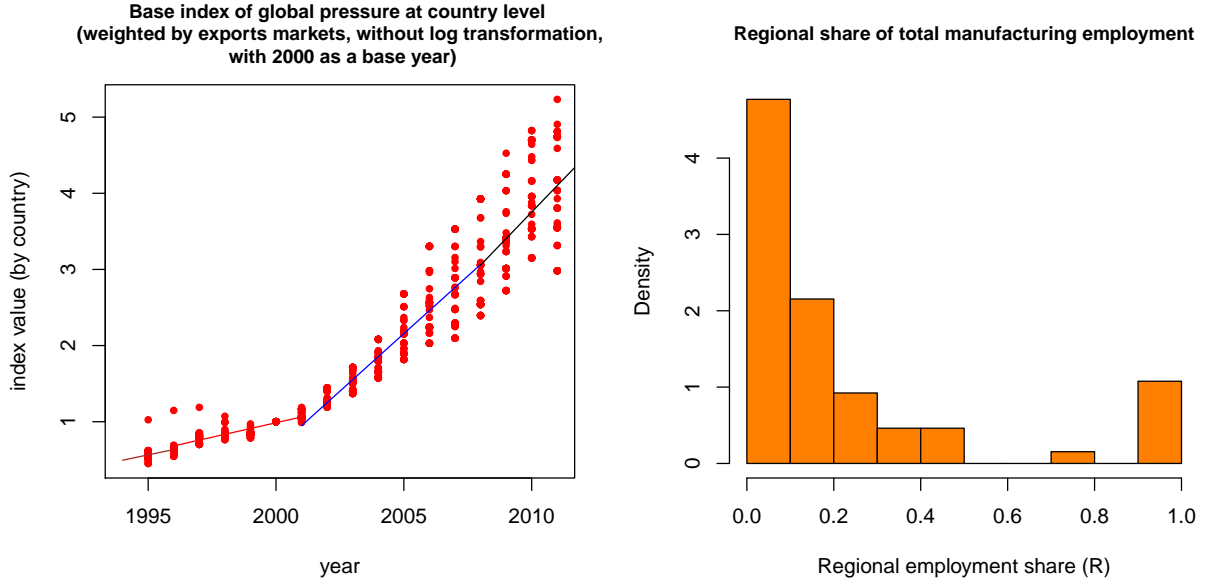
<sup>28</sup>For a stylized motivation, suppose that the utility function at a moment  $t \in \mathbb{Z}$  is given by  $U(t) = A Q_{CN,t}^{\alpha(t)} Q_{1,t}^{\beta_1(t)}, \dots, Q_{n,t}^{\beta_n(t)}$ , where  $Q_{j,t}$  represents the quantity of products used from different markets  $j \in \{CN, 1, 2, \dots, n\}$ ,  $n \in \mathbb{N}$ , 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.

<sup>29</sup>As is shown in Figure A7 in Appendix D, the same break is also evident in the logarithmically transformed demeaned data of the country pressure without using any year-specific normalization.

<sup>30</sup>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.

<sup>31</sup>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.

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



data change only slowly over time<sup>32</sup> and have a skewed distribution – see the right panel of Figure 3.1. 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 equation (5) to account for the skewed distribution of the projection term<sup>33</sup> and to reduce the heteroskedasticity over time, which would otherwise translate to errors in the equations under estimation.

<sup>32</sup>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.

<sup>33</sup>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.

### 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 exports markets defined in eq. (5) on inequality. Table 1 contains the results of the basic estimations<sup>34</sup>, 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 6.

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.

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.<sup>35</sup> 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

---

<sup>34</sup>The tables were created using the `stargazer` package for R (see Hlavac, 2018).

<sup>35</sup>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.

Table 1: Base estimates with indicators of external trade pressure

	<i>Dependent variable: Income inequality (LVAR)</i>							
	(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				0.005 (0.021)		-0.005 (0.027)		0.002 (0.027)
Economic sentiment index					-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
Reg.	65	65	65	65	65	65	65	65
Instr.no.	46	46	46	46	46	46	46	46
Obs.	1037	785	785	785	774	774	561	774

*Note:* \*p<0.1; \*\*p<0.05; \*\*\*p<0.01

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 (REER) 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 equation (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.<sup>36</sup> 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

---

<sup>36</sup>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.

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.<sup>37</sup> 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.<sup>38</sup> 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

---

<sup>37</sup>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.

<sup>38</sup>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.

variables from further specifications.

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. Third, the coefficient of the lagged dependent variable is much smaller for P5010 as compared with both LVAR and GINI, which implies a larger long-term impact on inequality measured by the 50/10 ratio of income percentiles. 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 50<sup>th</sup> and 10<sup>th</sup> 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.



Table 2: Base estimates, various inequality indicators

	<i>Dependent variable:</i>							
	LVAR	GINI	P5010	P9050	LVAR	GINI	P5010	P9050
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Lag of dependent variable	0.496*** (0.064)	0.588*** (0.064)	0.375*** (0.051)	0.285*** (0.098)	0.504*** (0.063)	0.597*** (0.066)	0.376*** (0.051)	0.297*** (0.089)
External pressure $\mathcal{P}_{r,t}^{(X)}$	0.049** (0.024)	0.023** (0.009)	0.034*** (0.010)	0.004 (0.007)	0.050** (0.022)	0.023*** (0.008)	0.035*** (0.008)	0.008 (0.006)
... its beta coefficient	[0.328**] (0.162)	[0.377**] (0.154)	[0.532***] (0.161)	[0.093] (0.167)	[0.330**] (0.146)	[0.382***] (0.128)	[0.510***] (0.142)	[0.178] (0.148)
Real effective exchange rate	-0.005 (0.027)	-0.002 (0.008)	-0.002 (0.009)	-0.007 (0.007)				
Economic sentiment index	-0.110 (0.092)	0.021 (0.038)	-0.050* (0.030)	0.010 (0.024)			-0.049* (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
Reg.	65	65	65	65	65	65	65	65
Instr.no.	46	46	46	46	46	46	46	46
Obs.	774	774	774	774	785	785	774	785

*Note:*

\*p<0.1; \*\*p<0.05; \*\*\*p<0.01

## 4 Importance of other markets and intermediate adjustment

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 domestic market of the country of the region of interest. Also, we have taken a look only at the basic equations without considering trade adjustment by firms of the affected country/region. This can happen through various re-allocations of imports and exports between different markets that, potentially, can reduce the pressure on firms and, therefore, also inequality. Now we lift these restrictions.

### 4.1 Other markets

Motivated by an idea similar to those of Autor *et al.* (2013) and Feenstra *et al.* (2017), we define the additional indicators aiming to capture the pressure faced in the domestic market and the potential expansion of exports in the Chinese market. The indicator for ‘domestic pressure’ is defined analogously to that for external pressure, just with one (local) market<sup>39</sup>:

$$\mathcal{P}_{r,t}^{(M)} = \log \left( \frac{M_{i,t}^{(\text{from China})}}{O_{i,t} - X_{i,t}} R_{i,r,t} \right), \quad (6)$$

where:

$M_{i,t}^{(\text{from China})}$  indicates imports from China to country  $i$ ;

$O_{i,t}$  is output of manufacturing goods in country  $i$ ; and

$X_{i,t}$  represents total exports of manufactured goods from country  $i$  to all countries.

The country-level pressure is equal to the import of manufacturing goods from China

---

<sup>39</sup>As in the definition of the external pressure (see Footnote 26), we omit the country index on the left side.

to country  $i$  divided by the local use of manufacturing goods.<sup>40</sup> The regional pressure index relies on the country-level one, distributing it to region  $r$  proportionally to the share of manufacturing in local employment. Local use of manufacturing goods is the difference between output and export of manufacturing goods. The pressure increases with imports, but decreases with local production, because that implies a larger domestic market share. One expects this effect to increase inequality, as increasing (relative) imports might imply either a decrease of employment in local manufacturing or a shift towards more advanced technology complemented by skilled employees who in turn might command higher wages.

To account for the potential to export to the growing Chinese market, we introduce the following two indicators. The first one uses the same principle as for the external and domestic pressure indicators and defines the ‘China option’ based on the ratio of exports from an EU country  $i$  to China ( $X_{i,t}^{(\text{to CN})}$ ) to China’s total imports of manufactured goods from all countries ( $M_t^{(\text{of China})}$ ):

$$\mathcal{C}_{r,t}^{(1)} = \log \left( \frac{X_{i,t}^{(\text{to CN})}}{M_t^{(\text{of China})}} R_{i,r,t} \right), \quad (7)$$

where, as previously, region  $r$ ’s share of total manufacturing employment in country  $i$  ( $R_{i,r,t}$ ) is used to distribute the impact to regions. This indicator represents the opportunity to exploit the growth of the Chinese domestic market and thus increases with exports to China while decreasing with all imports of China, as the opportunity is relative to the total market. The idea here is that entering the WTO had a two-way effect, making both Chinese goods and the Chinese market available to the rest of the world.

We also use an alternative ‘China option’ indicator, where, instead of the potential possibility and ability to export to China, as captured by the total Chinese imports in the denominator of eq. (7), the actual share of exports to China of a country’s total

---

<sup>40</sup>We also explored the total production alone, manufacturing value added, or further refinement of manufacturing output less exports without reexports in the denominator of eq. (6), which in all cases gave similar results to those reported in the sequel.

exports is evaluated by

$$C_{r,t}^{(2)} = \log \left( \frac{X_{i,t}^{(\text{to CN})}}{X_{i,t}} R_{i,r,t} \right), \quad (8)$$

where the definition of variables remains the same as in eqs. (6) and (7).

It should be pointed out that, other than for the external and domestic pressure indicators defined in eqs. (5) and (6), the China option's influence on inequality is less clear cut. Although in general it can be expected to reduce the trade pressure faced elsewhere, if European exporters were competitive in the Chinese market, the particular composition of products that are exported might matter as well.

Table 3 reports the estimation results, starting from the previously explored external pressure in Column (1). Columns (2)–(4) include the newly defined indicators of domestic and Chinese markets individually. The domestic pressure is significant and has a positive coefficient, whereas both Chinese market options are insignificant. Further columns augment the basic model of the external pressure with the newly added variables—first, individually in Columns (5)–(7), and then jointly in Columns (8)–(9)—by including the respective pairs of the domestic market and the two China options. The external pressure remains significant in all these cases, whereas the China option remains insignificant, while the domestic pressure becomes insignificant, too. However, the cases that include the external and the domestic pressure together, as in Columns (5), (8), and (9), have a very distinctive pattern that points to the collinearity of variables: the standard deviations (and the coefficient estimates) drastically increase (about four to five times), and the domestic pressure, which was significant in Column (2), not only becomes insignificant, but also changes the sign of its coefficient. And indeed, the correlation coefficient between the external and domestic pressure indicators is almost 0.9. Therefore, only a linear combination of these two variables can be precisely estimated with the actual data, while multicollinearity makes the estimates of their individual effects highly imprecise due to the increased variance of the parameter estimator.

This empirical feature, together with the economic arguments explicated in the In-

Table 3: Estimates with other markets and the dynamic common factor  $\mathcal{P}_{i,t}^{(F)}$  of the external and domestic pressure

Dependent variable: Income inequality (LVAR)													
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)*
Lag of dependent var.	0.504*** (0.063)	0.516*** (0.064)	0.548*** (0.078)	0.500*** (0.063)	0.477*** (0.072)	0.522*** (0.076)	0.513*** (0.064)	0.494*** (0.083)	0.483*** (0.073)	0.522*** (0.076)	0.515*** (0.066)	0.561*** (0.088)	0.290*** (0.071)
External pressure $\mathcal{P}_{r,t}^{(X)}$	0.050** (0.022)				0.172* (0.097)	0.047** (0.022)	0.092** (0.043)	0.175* (0.094)	0.219** (0.106)				
Domestic pressure $\mathcal{P}_{r,t}^{(M)}$		0.038* (0.021)			-0.113 (0.091)			-0.119 (0.089)	-0.116 (0.090)				
China option $\mathcal{C}_{r,t}^{(1)}$			-0.131 (0.126)			-0.048 (0.132)		-0.056 (0.144)		-0.048 (0.132)			
China option $\mathcal{C}_{r,t}^{(2)}$				0.040 (0.025)			-0.052 (0.048)		-0.056 (0.050)		-0.052 (0.050)		
Common factor $\mathcal{P}_{r,t}^{(F)}$ (of $\mathcal{P}_{r,t}^{(X)}$ and $\mathcal{P}_{r,t}^{(M)}$ )										0.017** (0.008)	0.032** (0.015)	0.023*** (0.007)	0.238*** (0.088)
P-val.(Hansen)	0.274	0.257	0.271	0.229	0.27	0.26	0.262	0.254	0.253	0.261	0.266	0.327	0.396
P-val.(AR1)	0	0	0	0	0	0	0	0	0	0	0	0	0
P-val.(AR2)	0.794	0.787	0.724	0.768	0.786	0.787	0.804	0.773	0.791	0.773	0.779	0.896	0.52
Reg.	65	65	65	65	65	65	65	65	65	65	65	65	65
Instr.no.	46	46	46	46	46	46	46	46	46	46	46	46	40
Obs.	785	785	785	785	785	785	785	785	785	785	785	1037	919

Notes:

\* p<0.1; \*\* p<0.05; \*\*\* p<0.01, \*  $\mathcal{P}_{r,t-3}^{(F)}$  is used here

troductioin, motivated us to introduce a (dynamic) common factor of the external and domestic pressure, as defined in Appendix C.1.<sup>41</sup> One can understand it as a composite pressure indicator extracting the joint component from the external and domestic pressures – it becomes necessary to identify some additional intermediate adjustment effects, as otherwise, their joint usage together with the external and domestic pressure indicators leads to the insignificance of all coefficients because of multicollinearity.

Columns (10)–(12) report the results of estimation with the dynamic common factor: alone in Column (12) and augmented with the China option indicators in Columns (11) and (12). Interestingly, across all the considered specifications, we find no evidence during the analyzed period of the Chinese market becoming significantly important, at least for changes of income inequality in the countries under consideration.

The presented estimations with the common factor component reiterate the previous finding that the increase of income inequality in the EU15 regions due to the Chinese pressure is significant. However, up until now, the estimations allowed everywhere for individual effects while excluding the potential importance of period effects. Column (13) reports the results with both the individual and the period effects being present; however, for it to be significant, a lagged pressure series is required that is more significant than the contemporaneous one (see the discussion in Section 6).<sup>42</sup> The pressure effect remains significant, but both the standard errors and the estimated coefficient of the pressure become much larger and quite similar to those in Columns (5), (8), and (9), where we previously witnessed the presence of multicollinearity. In fact, the common factor  $\mathcal{P}_{r,t}^{(F)}$ , or any other pressure indicator, and the period effects should be quite collinear: as Figure 3.1 revealed, a large part of the identification of the China

---

<sup>41</sup>We also considered the principal component analysis of the two pressure indicators and used the first principal component (PC1), explaining 92% of total variation of external and local pressure. In qualitative terms, this gave an analogous result to the results reported with the dynamic common factor but does not benefit from a clear interpretation of the joint index.

<sup>42</sup>We use contemporaneous series in the main specifications in order to simplify the presentation and modeling and to reduce the potential over-fitting. Otherwise, a large number of modeling choices emerge, e.g., which lags of other series to consider, whether to fix the same lag for all variables or allow for different ones, which lag selection method to employ, etc.

impact stems from the structural break in the trend of the pressure. Similar arguments prevented Breemersch *et al.* (2017) from considering the period effects and trends in their analysis altogether (in Section 6, we will also include the linear and quadratic trends, but aiming more at proxying the potential technological change, if it were better captured by trends than by the growth rate of innovations). In our opinion, the size estimate of the Chinese pressure is more precise in the specifications without multicollinearity, whatever the reason, whether the joint consideration of several correlated pressure effects or period effects.

## 4.2 Intermediate market adjustment effects

In the previous sections, we have considered the effect of trade pressure without taking into account any potential adjustments. In this subsection, we take a look at the intermediate adjustments connected with simple reallocation of trade between the Chinese and third markets, potentially alleviating the pressure before any further economic adjustments would be 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. If this change of import sources in the final use of goods were taking place, such effect needs to be evaluated complementarily together with the domestic pressure created by China. We label it as 'imports substitution'.

To allow for such a substitution in the equations, we introduce the relative weight of imports from non-Chinese sources ( $M_{i,t}^{(\text{non-CN})}$ ) compared to goods produced and used locally:

$$\mathcal{S}_{r,t}^{(1)} = \log \left( \frac{M_{i,t}^{(\text{non-CN})}}{O_{i,t} - X_{i,t}} R_{i,r,t} \right), \quad (9)$$

where other variables are defined as previously. Note that the denominator here is the

same as in the domestic pressure eq. (6), which allows the domestic pressure to be reduced by the counteraction of the terms in the numerators of eqs. (6) and (9). The domestic pressure and imports substitution defined by these equations are expected to have opposite signs if such an effect is relevant.

Instead of allowing for the cancellation of the terms as above, we could also look directly at the ratio by country  $i$  of imports from China ( $M_{i,t}^{(\text{CN})}$ ) to those from other countries ( $M_{i,t}^{(\text{non-CN})}$ ), in a similar way to eq. (8):

$$\mathcal{S}_{r,t}^{(2)} = \log \left( \frac{M_{i,t}^{(\text{CN})}}{M_{i,t}^{(\text{non-CN})}} R_{i,r,t} \right). \quad (10)$$

The significance of either  $\mathcal{S}_{r,t}^{(1)}$  or  $\mathcal{S}_{r,t}^{(2)}$  would point to the presence of some kind of substitution effects. Note that the substitution effects discussed above are connected solely with the domestic pressure and can moderate its effect, but cannot soften the external trade pressure faced globally in exports markets.

Secondly, more affordable imports from China can also induce 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’. To assess its importance, we introduce an indicator that increases with the value of imported intermediate goods from China ( $M_{i,t}^{(\text{int-CN})}$ ) relative to exports (or production) from a given country:

$$\mathcal{F}_{r,t}^{(1)} = \log \left( \frac{M_{i,t}^{(\text{int-CN})}}{X_{i,t}} R_{i,r,t} \right). \quad (11)$$

The results with total manufacturing production in the denominator instead of exports are analogous to those with  $\mathcal{F}_{r,t}^{(1)}$ ; therefore, we do not report them. However, we may also be interested in whether this result is driven by exports or more by the domestic use of production, as captured by  $(O_{i,t} - X_{i,t})$ . To investigate this, we will also consider



a ‘domestic facilitation’ indicator:

$$\mathcal{F}_{r,t}^{(2)} = \log \left( \frac{M_{i,t}^{(\text{int-CN})}}{O_{i,t} - X_{i,t}} R_{i,r,t} \right). \quad (12)$$

Finally, if the fast-growing Chinese market becomes increasingly attractive, firms might want to reallocate their exports from other markets to China. The previously considered China option indicator  $\mathcal{C}_{i,t}^{(1)}$  might not capture this, e.g., if exports to China grow at the same rate as Chinese imports. Therefore, similar to the pair of domestic pressure and the substitution effect, the China option measure needs a complementary indicator that would gauge the extent of exports to non-Chinese markets, which we call ‘exports reallocation’. This reallocation would decrease with the weight of export to countries other than China ( $X_{i,t}^{(\text{non-CN})}$ ) relative to imports from China ( $M_{i,t}^{(\text{of CN})}$ ), and to capture this, we include

$$\mathcal{R}_{r,t}^{(1)} = \log \left( \frac{X_{i,t}^{(\text{non-CN})}}{M_{i,t}^{(\text{of CN})}} R_{i,r,t} \right) \quad (13)$$

which shares the same denominator with the China option  $\mathcal{C}_{i,t}^{(1)}$  in order to allow for the potential counteraction of the numerators. If such exports reallocation were taking place, then the China option and exports reallocation would be expected to have coefficients of opposite signs.

Alternatively, and in a similar way to eqs. (8) and (10), we also consider a direct metric of export reallocation defined by the ratio of exports by country  $i$  to China ( $X_{i,t}^{(\text{to CN})}$ ) and other countries ( $X_{i,t}^{(\text{non-CN})}$ ):

$$\mathcal{R}_{r,t}^{(2)} = \log \left( \frac{X_{i,t}^{(\text{to CN})}}{X_{i,t}^{(\text{non-CN})}} R_{i,r,t} \right). \quad (14)$$

It should be noted that  $\mathcal{R}_{r,t}^{(2)}$  is closely connected with the  $\mathcal{C}_{r,t}^{(2)}$  defined in eq. (8), which has total exports in its denominator instead of exports to countries other than China in eq. (14).

In Table 4, we report the estimation results using the newly introduced indices. Column (1) presents the basic specification with the composite pressure indicator—the dynamic common factor of external and internal pressure—as in Column (12) of Table 3. As it accounts for variations in both the external and the domestic pressure, we use it hereafter together with additional indicators of intermediate market adjustments.

Subsequently, we first individually add the just-discussed import substitution, exports facilitation, and exports reallocation indicators in Columns (2)–(7), whereas Columns (8) and (9) report the results of the joint estimation using the (most) significant variables.

Column (2) indicates that import substitution, as measured by  $\mathcal{S}_{i,t}^{(1)}$ , has a negative and significant coefficient, reducing the increase in inequality. This is consistent with the fact that the domestic pressure induced by China’s WTO accession was dampened in the EU regions by the diversion of imports from other countries towards Chinese products in the post-accession period. The second measure of import substitution ( $\mathcal{S}_{i,t}^{(2)}$ ) is insignificant in Column (3) but makes the pressure coefficient insignificant, too.<sup>43</sup> The export facilitation in Column (4) is also significant at the 10% significance level (the same holds when the total production is used in eq. (11) instead of exports), whereas there is no significant change from the ‘domestic facilitation’ perspective presented in Column (5). Finally, none of the discussed indicators of exports reallocation is significantly different from zero (see Columns (6) and (7)). This is also consistent with the previous result that the indicators of the Chinese market option were insignificant in Table 3.

In summary, when added individually, the coefficients of imports substitution and exports facilitation indicators are significant and have the expected signs, thus indicating the presence of these market-switching adjustments. When the (most) significant substitution and facilitation effects are considered jointly in Columns (8) and (9), they become insignificant, whereas the coefficient of the pressure indicator remains significant.

---

<sup>43</sup> Although the coefficient of the substitution indicator is very small and has a much larger standard error than that of the pressure indicator.

Table 4: Intermediate market adjustment effects

Dependent variable: Income inequality (LVAR)									
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Lagged dependent variable	0.561*** (0.088)	0.455*** (0.061)	0.509*** (0.066)	0.494*** (0.068)	0.467*** (0.081)	0.486*** (0.066)	0.516*** (0.066)	0.462*** (0.064)	0.450*** (0.069)
Common factor $\mathcal{P}_{r,t}^{(F)}$ (of $\mathcal{P}_{r,t}^{(X)}$ and $\mathcal{P}_{r,t}^{(M)}$ )	0.023*** (0.007)	0.040*** (0.012)	0.019 (0.047)	0.077** (0.036)	0.068** (0.033)	0.044** (0.018)	0.031** (0.015)	0.064* (0.035)	0.108* (0.049)
Import substitution $\mathcal{S}_{r,t}^{(1)}$		-0.360** (0.148)						-0.277 (0.187)	-0.182 (0.207)
Import substitution $\mathcal{S}_{r,t}^{(2)}$			-0.006 (0.147)						
Exports facilitation $\mathcal{F}_{r,t}^{(1)}$				-0.189* (0.109)				-0.095 (0.119)	-0.156 (0.140)
Domestic facilitation $\mathcal{F}_{r,t}^{(2)}$					-0.126 (0.080)				
Exports reallocation $\mathcal{R}_{r,t}^{(1)}$						0.085 (0.058)			0.097 (0.066)
Exports reallocation $\mathcal{R}_{r,t}^{(2)}$							-0.048 (0.048)		
P-val.(Hansen)	0.327	0.372	0.303	0.313	0.49	0.254	0.265	0.35	0.337
P-val.(AR1)	0	0	0	0	0	0	0	0	0
P-val.(AR2)	0.896	0.537	0.782	0.732	0.798	0.844	0.781	0.567	0.675
Reg.	65	65	65	65	65	65	65	65	65
Instr.no.	46	46	46	46	46	46	46	46	46
Obs.	1037	785	785	785	677	785	785	785	785

\*p&lt;0.1; \*\*p&lt;0.05; \*\*\*p&lt;0.01

Note:

ant. Finally, we should recall that, by construction, the most individually significant substitution effect is connected solely with the domestic pressure through crowding out of other foreign imports by Chinese imports, but it cannot compensate for the external trade pressure faced globally in exports markets. This, besides a clearly observable inflection point in the trend since 2002 discussed in Figure 3.1, was an additional motivation to use only the external pressure indicator in the stylized simulation of the impact, presented in Appendix D.

## 5 Importance of impact channels: A semi-reduced form

So far, we have looked at the impact in a reduced form which can be connected with many economic mechanisms. In order to shed some light on their importance, we think about the impact of China’s expansion as a two-stage process. 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. For the clarity of discussion, we will reverse the process by starting from the second stage, i.e., the equation of impact channels.

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 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 E.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. Note that in all 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. We use the upgrading of the technological intensity of exports and the up-skilling of workers to identify the intra-manufacturing adjustments.

The estimation results are presented next in two tables. First, Table 5 reports the projection of inequality on the variables listed above to identify the significant ones. Second, Table 6 presents the results of projections of these channel/transmission variables on the trade pressure indicator, in connection with the first stage of the trade pressure impact. In both cases, we rely on the strategy already used in creating our base estimates and project the channel variables on the same set of instruments as in the fully reduced form to retain the correspondence of results explicated in Appendix C.2.

Columns (1)–(5) in Table 5 explore the individual significance of the transmission variables. All the coefficients here have the expected signs: negative for the manufacturing labor share and positive for the rest. Manufacturing share, unemployment rate, and white collar workers are also immediately significant at the usual significance levels. The individually considered contemporaneous variables of technological intensity of exports and the share of workers with higher education are insignificant, but their lags are significant (not reported). This might indicate that the pressure channeling through manufacturing reduction and increasing unemployment starts almost instantaneously, whereas the upgrading of technology and the up-skilling of labor might lag

somewhat after the initial trade pressure takes place. The estimation of joint effects corroborates this intuition.

Columns (6)–(8) explore the significance of variables employed jointly and not individually. Using the backwards selection strategy to select the most informative variables, i.e., starting from all variables and iteratively dropping the least significant one, it reports only the remaining variables that are significant at least at the 10% significance level. To evaluate the potentially lagging effects, we estimate the specifications not only with the contemporaneous relationship in Column (6), but also with variables lagged once and twice in Columns (7) and (8), respectively. Consistently with the previous prediction, only the manufacturing labor share and the unemployment rate are significant in the contemporaneous specification in Column (6), whereas the technological intensity of exports joins this combination after one lag, as reported in Column (7). Finally, higher education also becomes significant in Column (8) when we allow for its interaction with the technological intensity of exports. The share of white collar workers, despite being significant individually in Column (4), does not become significant in specifications with many variables.

Two additional aspects should be stressed. First, Column (7) is not only the best specification when we fix one lag of explanatory variables, but it is also produced by the backwards selection strategy from all considered lags when only linear terms are included. Second, the appropriate lags of the pressure indicator are insignificant if added to the specifications in Columns (6)–(8). If a substantially significant channel, through which the pressure indicator affects inequality, were missing, the pressure indicator would be expected to become significant under such specifications.

Table 5: Relevance of impact channels

Lag of explanatory variables:	<i>Dependent variable: Income inequality (LVAR)</i>							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	lag = 0				lag = 1		lag = 2	
Lag of dependent variable	0.470*** (0.067)	0.509*** (0.091)	0.455*** (0.073)	0.532*** (0.078)	0.576*** (0.084)	0.416*** (0.076)	0.309*** (0.081)	0.208** (0.092)
Manufacturing labor share	-0.346*** (0.089)					-0.369*** (0.106)	-0.358*** (0.096)	-0.262*** (0.113)
Unemployment rate		0.187*** (0.046)				0.179*** (0.044)	0.185*** (0.051)	0.195*** (0.070)
Techn. intens. of exports (TECH)			0.408 (0.263)				0.535** (0.253)	0.592* (0.329)
White-collar workers				0.402* (0.222)				
Workers with higher education (HEDU)					0.139 (0.215)			1.261** (0.127)
Interaction: HEDU * TECH								0.257** (0.127)
P-val.(Hansen)	0.447	0.241	0.197	0.254	0.218	0.431	0.638	0.483
P-val.(AR1)	0	0	0	0	0	0	0	0
P-val.(AR2)	0.438	0.359	0.608	0.42	0.351	0.584	0.659	0.766
Reg.	65	65	65	65	65	65	65	65
Instr.no.	46	46	46	46	46	46	46	46
Obs.	887	887	968	886	859	887	802	717

*Note:* \*p<0.1; \*\*p<0.05; \*\*\*p<0.01

In summary, the established significance of the considered channel variables is compatible with a combination of inter-sectoral and intra-sectoral mechanisms: inequality is affected not only by the shrinking manufacturing sector and increasing unemployment, but also by the technological upgrading of manufacturing exports together with upskilling of employed labor.

As the results provided above were identified through the projections on instruments, the significance of variables is informative about the relevance of different channels. Nevertheless, besides their statistical significance, we aim at evaluating their relative importance. Therefore, we additionally present not only the beta coefficients in Table A7 in Appendix B, which reveal the size of responses to standardized changes in the channel variables, but also the response of each variable employed above to the trade pressure itself. Table 6 next reports these results using the common factor ( $\mathcal{P}_{r,t}^{(F)}$ ) of the external and domestic pressure, with the following abbreviations used therein: MSHR – labor share in manufacturing, UNRT – unemployment rate, TECH – technological intensity of exports, WHCO – share of white-collar workers, HEDU – share of labor with higher education.

Apart from the unemployment rate variable to be discussed shortly, all other coefficients have the expected signs and are significant. It should be noted that, unlike the unemployment rate, the structural break connected with the financial crisis was insignificant for the remaining variables. The reported equations for the unemployment rate are problematic. First, whenever the structural break is introduced in Column (3), the pressure coefficient before the crisis becomes negative. Even though it is positive during the aftermath of the financial crisis, the unemployment-reducing trade pressure effect before the crisis is somewhat counterintuitive.<sup>44</sup> Second, and more importantly, the validity of over-identifying restrictions by the Sargan-Hansen test statistic is rejected, which casts doubt on the consistency of its estimation. We tried to solve this

---

<sup>44</sup>Although it is not impossible. For instance, consider the case where the trade pressure is connected with some ‘creative destruction’ process and induces the emergence of new firms to produce new products that penetrate broader markets. At the same time, the dismissed workers could also become more mobile in regional and/or sectoral terms. This would allow the booming sectors to extend or start new projects employing additional labor during the upswing of economic activity.



Table 6: Initial responses to primary trade pressure

	<i>Dependent variable:</i>					
	MSHR	UNRT	TECH	WHCO	HEDU	
	(1)	(2)	(3)	(4)	(5)	(6)
Lag of dependent variable	0.705*** (0.059)	0.816*** (0.046)	0.618*** (0.047)	0.410*** (0.152)	0.102 (0.160)	0.354* (0.196)
Common factor $\mathcal{P}_{r,t}^{(F)}$	-0.024*** (0.004)	0.003 (0.007)	-0.050*** (0.008)	0.008** (0.003)	0.020*** (0.004)	0.011** (0.005)
$\mathcal{P}_{r,t}^{(F)} * \mathbb{1}\{t \geq 2008\}$			0.103*** (0.040)			
y2008			-0.083 (0.090)			
P-val.(Hansen)	0.25	0.051	0.036	0.17	0.207	0.558
P-val.(AR1)	0	0	0	0.007	0.022	0.008
P-val.(AR2)	0.073	0.864	0.393	0.37	0.497	0.653
Reg.	65	65	65	65	65	65
Instr.no.	46	46	46	46	46	46
Obs.	1087	1087	1087	1158	1083	1042

Note:

\*p&lt;0.1; \*\*p&lt;0.05; \*\*\*p&lt;0.01

issue by increasing autoregressive lags, varying and considering more lags of the pressure indicator, looking for other non-linearities, and introducing additional controls, including trends, economic sentiment indicator, exchange rates, but without success<sup>45</sup>. We therefore report the decomposition results next with both specifications, hoping that the true impact is somewhere in between the two. However, one has to keep this potential caveat in mind when interpreting further results. The positive side is that, irrespective of the particular specification employed for the initial impulse equation of the unemployment rate, the decomposition pattern obtained with respect to other indicators remains the same.

As defined in Appendix C.2, the initial responses together with the channel equa-

<sup>45</sup>It should be noted that we do not change the set of instruments, because this would invalidate the decomposition defined in Appendix C.2.

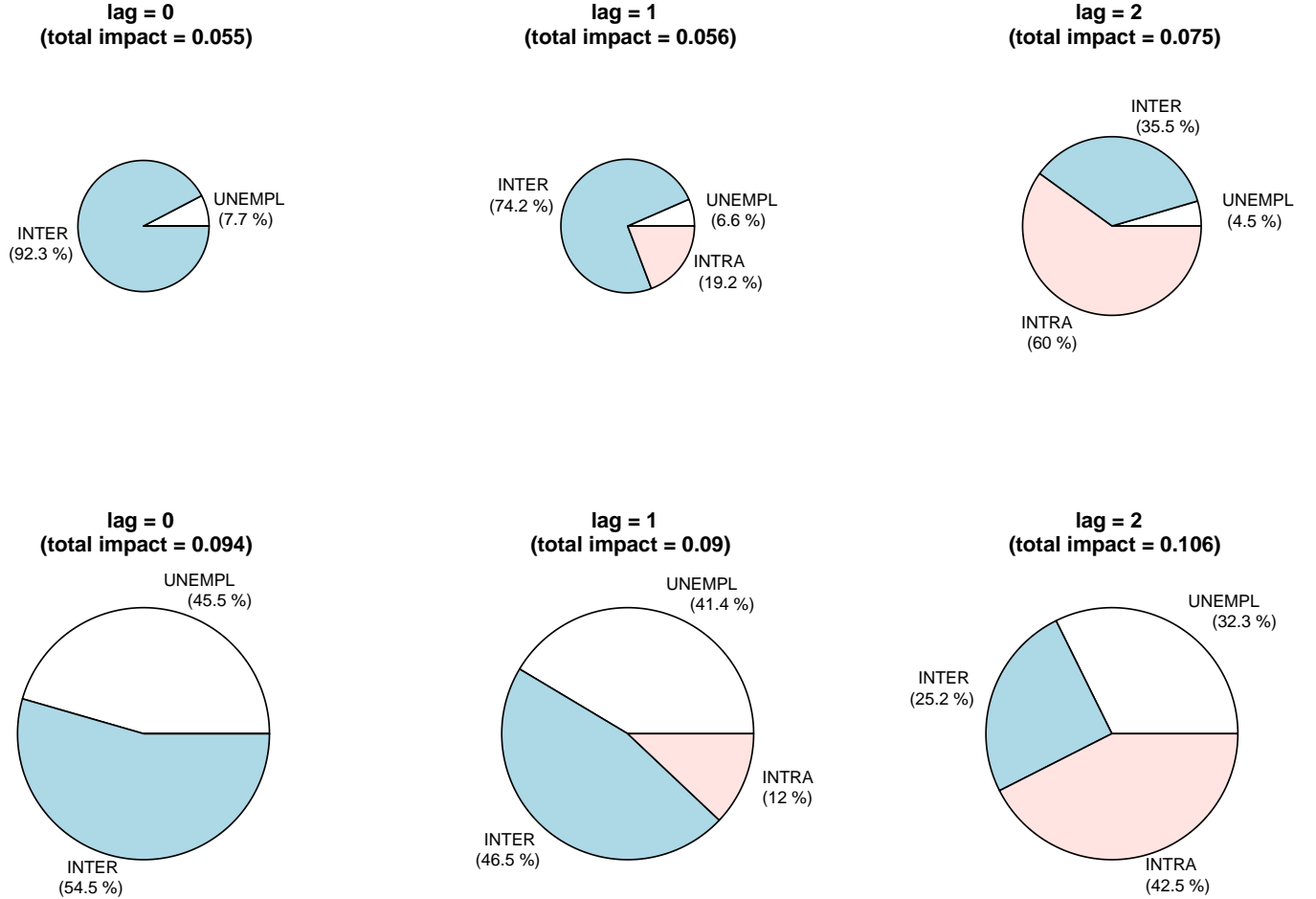
tion determine the decomposition, where the total long-run impact is split into terms associated with each channel variable. Based on it and the results reported in Tables 5 and 6, Figure 2 reports the estimated total long-run impact of the trade pressure on inequality, together with its share explained by the inter-sectoral shift (through the reduction of the share of labor in manufacturing), the intra-sectoral adjustments (through the increasing technological intensity of exports and/or upskilling of the employed labor force), and the increasing unemployment rate. In Figure 2, they are abbreviated as INTER, INTRA, and UNEMPL, respectively. In correspondence with Columns (6)–(8) of Table 5, the decomposition is provided for the impact without any lag (lag=0), with one lag (lag=1), and with two lags (lag=2). The upper and lower panels of the figure use the specifications of the primary impact on unemployment rate as in Columns (2) and (3) of Table 6, respectively. The radius of circles corresponds to the total long-run impact that is reported in the brackets below each lag.<sup>46</sup>

All decompositions are in accordance with the fact that, initially, i.e., at lag=0, the largest adjustment takes place through the inter-sectoral channel (the shrinking manufacturing sector). The unemployment channel became important only after the financial crisis and also increased the total impact, but its share among the transmission channels is still smaller than that of the inter-sectoral shift. As time passes, the intra-sectoral adjustments not only become significant at lag one but also become dominant at lag two, further increasing the total long-run impact of the trade pressure on inequality. Thus, over the span of two years, the intra-sectoral adjustments become the main source of income inequality, overwhelming the initially predominant inter-sectoral shift.

---

<sup>46</sup>It should be pointed out that, despite the sample variations induced by differences in data availability of various variables, the calculated total long-run impact corresponds well with that estimated directly from the reduced form. For instance, the top-left panel in Figure 2, linked to the case of lag=0 of the semi-structural approach, produces a total impact of 0.055. The reduced-form equation that corresponds to this case is given in Column (12) of Table 3, where the derived long-run impact is approximately 0.052 ( $\approx \frac{0.023}{1-0.561}$ ).

Figure 2: Importance of impact channels



Note: the upper and lower panels use the primary impulse specifications of unemployment rate given in Columns (2) and (3) of Table 6, respectively. The lower panel presents the results with the financial crisis effect, i.e, obtained under  $\mathbb{1}\{t \geq 2008\} = 1$ . The radius of circles corresponds to the total long-run impact that is reported in the brackets below each lag. INTER stands for inter-sectoral shift, INTRA for intra-sectoral adjustments, and UNEMPL for a move to unemployment. The total impact here represents the long-run impact, which takes into account both the immediate partial impact and adjustments through the lagged dependent variable (see Appendix C.2).

## 6 Robustness checks and economic significance

In the sequel, we perform five 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 B 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 A1 (even further reduction to 9% will be discussed shortly). Practically, this leaves all the previously established results intact. The noticeable change in terms of significance is present only in Column (11), 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 5.

Second, we investigate the importance of the composition of instruments in Table A2, 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 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 A2 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 neither 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 A3, 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. This seems to be admissible, as these regional shares vary slowly over time. The results here again remain very similar to the previously reported ones.

Fourth, as presented in Table A4 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.<sup>47</sup> 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, 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 A5 and A6. 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 A5, the trade pres-

---

<sup>47</sup>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.

sure indicator also 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 A6 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 only with the external pressure ( $\mathcal{P}_{r,t}^{(X)}$ ), while here they are used with the dynamic common factor ( $\mathcal{P}_{r,t}^{(F)}$ ). 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 a downwards 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 is introduced of 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 A6), 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 in to account.

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 D. 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 A7 in Appendix D), 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 D. The dynamics of the predicted absolute increase in inequality are plotted in the top-right panel of Figure A8 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 A9. 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 A8: 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 the China’s WTO accession in 2001 was about 0.03 and 0.06 in 2007 and 2011, respectively. 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. The relative ranking of countries is barely affected.

## 7 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 become the largest contributor to income inequality through the technological upgrading of exports (and production) together with the up-skilling of the employed labor.

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 afterwards.

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. 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.

## References

- Acemoglu, D., Autor, D., Dorn, D., Hanson, G. H. and Price, B. (2016) Import Competition and the Great US Employment Sag of the 2000s, *Journal of Labor Economics*, **34**, S141–S198.
- Agrawal, J. and Kamakura, W. (1999) Country of origin: A competitive advantage?, *International Journal of Research in Marketing*, **16**, 255–267.
- Arellano, M. and Bond, S. (1991) Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations, *Review of Economic Studies*, **58**, 277–97.
- Autor, D. H., Dorn, D. and Hanson, G. H. (2013) The China Syndrome: Local Labor Market Effects of Import Competition in the United States, *American Economic Review*, **103**, 2121–2168.
- Autor, D. H., Dorn, D. and Hanson, G. H. (2016) The China Shock: Learning from Labor-Market Adjustment to Large Changes in Trade, *Annual Review of Economics*, **8**, 205–240.

- Autor, D. H., Dorn, D., Hanson, G. H. and Song, J. (2014) Trade adjustment: Worker-level evidence, *The Quarterly Journal of Economics*, **129**, 1799–1860.
- Baiardi, D. and Carluccio, B. (2019) At the roots of China’s striking performance in textile exports: A comparison with its main Asian exporters, *China Economic Review*, **54(C)**, 367–389.
- Balsvik, R., Jensen, S. and Salvanes, K. G. (2015) Made in China, sold in Norway: Local labor market effects of an import shock, *Journal of Public Economics*, **127**, 137–144.
- Barlevy, G. and Tsiddon, D. (2006) Earnings inequality and the business cycle, *European Economic Review*, **50**, 55–89.
- Basco, S., Liégey, M., Mestieri, M. and Smagghue, G. (2017) Trade and Inequality: Evidence from Worker-Level Adjustment in France.
- Bell, B. and Machin, S. (2016) Brexit and wage inequality, in *Brexit Beckons: Thinking ahead by leading economists* (Ed.) R. Baldwin, VoxEU.
- Bellon, M. (2015) Trade liberalization and inequality: a dynamic model with firm and worker heterogeneity.
- Benczúr, P., Cseres-Gergely, Z. and Harasztosi, P. (2017) EU-wide income inequality in the era of the Great Recession, Working Paper 2017/14, Joint Research Centre, European Commission, Ispra.
- Bloemer, J., Brijs, K. and Kasper, H. (2009) The CoO-ELM model: A theoretical framework for the cognitive process underlying country of origin-effects, *European Journal of Marketing*, **43**, 62–89.
- Bloom, N., Draca, M. and Van Reenen, J. (2016) Trade induced technical change? The impact of Chinese imports on innovation, IT and productivity, *The Review of Economic Studies*, **83**, 87–117.

- Boldrin, M. and Canova, F. (2001) Inequality and convergence in Europe's regions: reconsidering European regional policies, *Economic Policy*, **16**, 206–253.
- Breemersch, K., Damijan, J. P. and Konings, J. (2017) Labour Market Polarization in Advanced Countries, OECD Social, Employment and Migration Working Papers 197, DOI: 10.1787/06804863-en.
- Broda, C. and Weinstein, D. E. (2006) Globalization and the Gains from Variety, *The Quarterly Journal of Economics*, **121**, 541–585.
- Cabral, R. and Castellanos-Sosa, F. (2019) Europe's income convergence and the 28 latest global financial crisis, *Research in Economics*, **73**, 23–34.
- Colantone, I. and Stanig, P. (2018) Global Competition and Brexit, *American Political Science Review*, **112**, 201–218.
- Croissant, Y. and Millo, G. (2008) Panel Data Econometrics in R: The plm Package, *Journal of Statistical Software*, **27**.
- Dauth, W., Findeisen, S. and Suedekum, J. (2017) Trade and Manufacturing Jobs in Germany, *American Economic Review*, **107**, 337–342.
- Dickens, W. T. (1990) Error Components in Grouped Data: Is It Ever Worth Weighting?, *The Review of Economics and Statistics*, **72**, 328–333.
- European Commission (1996) The European Community Household Panel: ECHP. Vol. 1: Survey Methodology and Implementation, Eurostat Theme 3, Population and social conditions Series F, Accounts and surveys, Off. for Off. Publ. of the Europ. Communities, Luxembourg, oCLC: 833152860.
- European Commission (2002) 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 /\* COM/2001/0517 final - CNS 2001/0218 \*//, *The Official Journal of the European Union*, **051 E**, 26 February 2002, 0314–0315.

- Eurostat (2017) Quality Report of the European Union Labour Force Survey 2015: 2017 Edition, Quality Report, Publication Office, Luxembourg, oCLC: 1032580929.
- Feenstra, R., Ma, H. and Xu, Y. (2017) US Exports and Employment, NBER Working Paper 24056, National Bureau of Economic Research, Inc.
- Feenstra, R. C. and Hanson, G. H. (1996) Globalization, Outsourcing, and Wage Inequality, *American Economic Review*, **86**, 240–245.
- Feenstra, R. C. and Sasahara, A. (2017) The ‘China Shock’, Exports and U.S. Employment: A Global Input-Output Analysis, Working Paper 24022, National Bureau of Economic Research, DOI: 10.3386/w24022.
- Frick, J. R. and Krell, K. (2010) Measuring Income in Household Panel Surveys for Germany: A Comparison of EU-SILC and SOEP, Tech. rep.
- Greenaway, D., Aruneema, M. and Milner, C. (2008) Has China Displaced Other Asian Countries Exports, *China Economic Review*, **19**, 152–169.
- Halpern, L., Koren, M. and Szeidl, A. (2015) Imported Inputs and Productivity, *American Economic Review*, **105**, 3660–3703.
- Hansen, L. (1982) Large sample properties of generalized method of moments estimators, *Econometrica*, **50**, 1029–1054.
- Hansen, L. P. and Hodrick, R. J. (1980) Forward exchange rates as optimal predictors of future spot rates: An econometric analysis, *Journal of Political Economy*, **88**, 829–853.
- Harrigan, J. and Reshef, A. (2015) Skill-biased heterogeneous firms, trade liberalization and the skill premium, *Canadian Journal of Economics/Revue canadienne d’économique*, **48**, 1024–1066.
- Helpman, E. (2017) Globalisation and wage inequality, *Journal of the British Academy*, **5**, 125–162.

- Hlavac, M. (2018) *stargazer: Well-Formatted Regression and Summary Statistics Tables*, Central European Labour Studies Institute (CELSI), Bratislava, Slovakia, R package version 5.2.2.
- Holmes, E., Ward, E. and Scheuerell, M. (2020) Analysis of multivariate time-series using the MARSS package, User Guide version 3.10.12.
- Kovak, B. K. (2013) Regional Effects of Trade Reform: What is the Correct Measure of Liberalization?, *American Economic Review*, **103**, 1960–1976.
- Kroh, M., Kühne, S., Siegers, R. and Belcheva, V. (2018) SOEP-Core – Documentation of Sample Sizes and Panel Attrition (1984 until 2016), Tech. rep. 480, DIW/SOEP, Berlin.
- Krugman, P. R. (2008) Trade and Wages, Reconsidered, *Brookings Papers on Economic Activity*, **2008**, 103–154.
- Kvedaras, V. and Cseres-Gergely, Z. (2020) Convergence of income distributions: Total and inequality-affecting changes in the EU, *Economics Letters*, **188**, 1–6.
- Lakner, C. and Milanovic, B. (2016) Global Income Distribution: From the Fall of the Berlin Wall to the Great Recession, *The World Bank Economic Review*, **30**, 203–232.
- Leamer, E. (1995) The Heckscher-Ohlin Model in Theory and Practice, *Princeton Studies in International Finance*, **no. 77**.
- Lynn, P. (2006) Quality profile: British Household Panel Survey – Version 2.0: Waves 1 to 13: 1991-2003, Tech. rep., Institute for Social and Economic Research, University of Essex.
- Magnusson, P. and Westjohn, S. A. (2011) Is There a Country-of-Origin Theory?, in *Handbook of Research in International Marketing* (Eds.) C. Subhash and D. Griffith, Elgar.

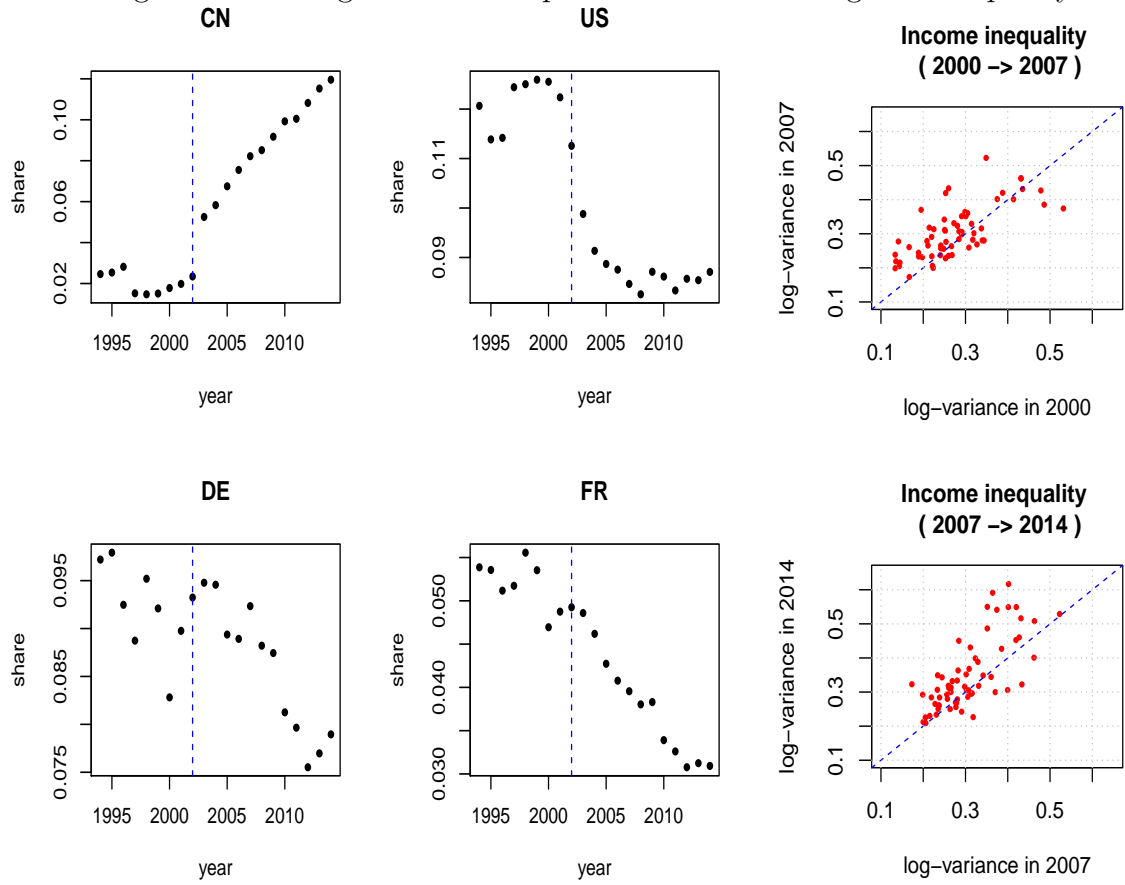
- Martinez Turegano, D. (2020) Employment composition and labour earnings within EU countries, JRC Technical Reports JRC120255, Publication Office, Luxembourg.
- Melitz, M. J. (2003) The Impact of Trade on Intra-Industry Reallocations and Aggregate Industry Productivity, *Econometrica*, **71**, 1695–1725.
- Persyn, D., Diaz-Lanchas, J., Barbero, J., Conte, A. and S., S. (2019) A new dataset of distance and time related transport costs for EU regions, Tech. rep., Territorial Development Insights Series, JRC119412, European Commission.
- Pham, C., Nguyen, X., Sgro, P. and Tang, X. (2016) Has China Displaced its Competitors in High-tech Trade, *The World Economy*, **40**, 1569–1596.
- Pierce, J. R. and Schott, P. K. (2016) The Surprisingly Swift Decline of US Manufacturing Employment, *American Economic Review*, **106**, 1632–1662.
- Roodman, D. (2009a) A Note on the Theme of Too Many Instruments, *Oxford Bulletin of Economics and Statistics*, **71**, 135–158.
- Roodman, D. (2009b) How to do xtabond2: An introduction to difference and system GMM in Stata, *Stata Journal*, **9**, 86–136.
- Rowley, J. C. R. and Wilton, D. A. (1973) Quarterly models of wage determination: Some new efficient estimates, *American Economic Review*, **63**, 380–389.
- Sargan, J. (1958) The estimation of economic relationships using instrumental variables, *Econometrica*, **71**, 393–415.
- Sherman, H. and Sherman, P. (2015) Inequality and the Business Cycle, *Challenge*, **58**, 51–63.
- Snyder, F. (2009) *The European Union and China, 1949-2008: Basic Documents and Commentary*, Oregon.
- Stolper, W. F. and Samuelson, P. A. (1941) Protection and Real Wages, *The Review of Economic Studies*, **9**, 58–73.

- Topalova, P. (2010) Factor Immobility and Regional Impacts of Trade Liberalization: Evidence on Poverty from India, IMF Working Paper WP/10/218.
- Windmeijer, F. (2005) A finite sample correction for the variance of linear efficient two-step GMM estimators, *Journal of Econometrics*, **126**, 25–51.



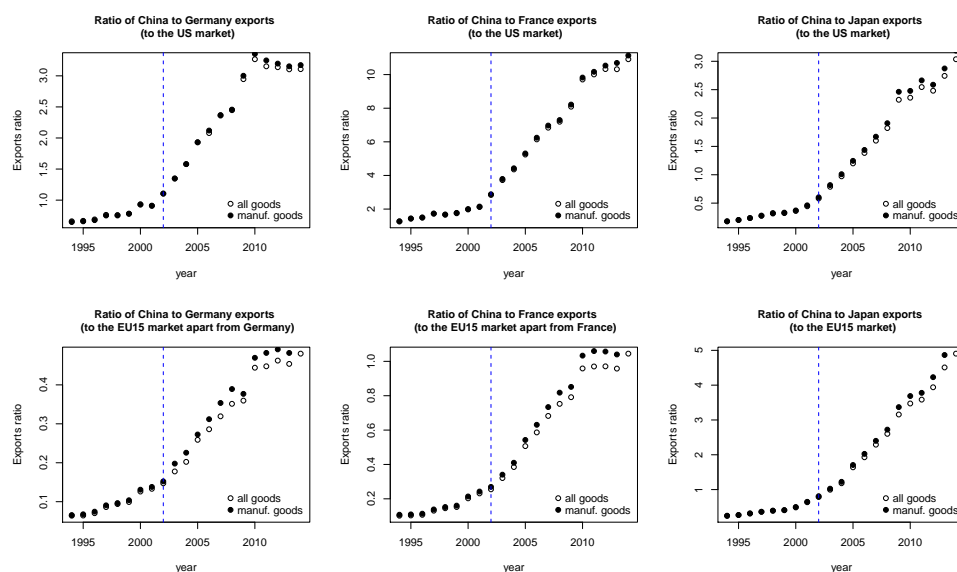
## A Appendix: Motivating stylized facts

Figure A1: Changes in world export shares and EU regional inequality



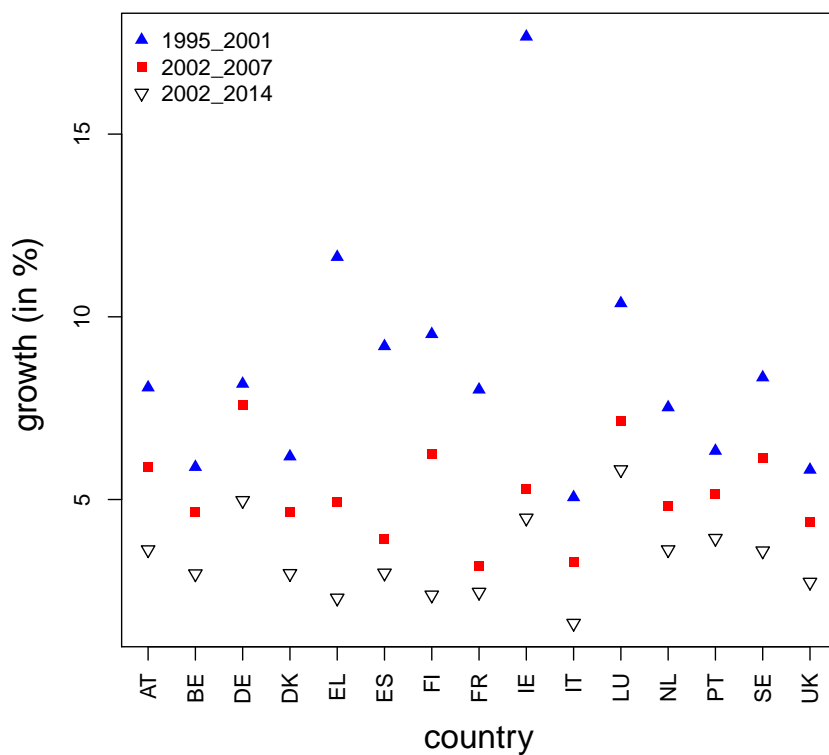
Source: own calculations using various data sources. The IMF BOP data were used to derive the export shares presented in the left panel. ECHP, EU-SILC, GSOEP, and BHPS were employed for inequality (see Appendix E for details).

Figure A2: Ratios of exports of goods from China and a selected country to the US (top panel) and the EU15 (bottom panel) markets



Source: OECD data.

Figure A3: Average growth rates of export of goods and services in EU15 countries



Source: own calculations based on the IMF BOP data.

## B Appendix: Robustness and other sensitivity checks

### List of tables in Appendix B:

Table A1: Results with a smaller number of instruments:  $\sim 33\%$  of the number of cross-sections

Table A2: Results with separate types and a smaller number of instruments:  $\sim 9\%$  of the number of cross-sections

Table A3: Results with regional projections fixed at  $R_{i,r,t} = R_{i,r,t_0}$ , for  $t > t_0 = 2000$

Table A4: Results with lagged pressure indicator and period effects

Table A5: Results with additional control variables: patents and trends

Table A6: Results with additional control variables: macroeconomic conditions, service exports, and the EU expansion with the CEE countries

Table A7: Relevance of the impact channels: beta coefficients from regressions with standardized variables

Table A1: Results with lower number of instruments:  $\sim 33\%$  of the number of cross-sections

	Dependent variable:										
	LVAR (1)	GINI (2)	P5010 (3)	P9050 (4)	(5) lag = 0	(6)	(7)	(8)	(9)	(10) lag = 1	(11) lag = 2
Lag of explanatory variables:											
Lag of dependent variable	0.528*** (0.082)	0.530*** (0.080)	0.639*** (0.089)	0.580*** (0.204)	0.647*** (0.078)	0.501*** (0.106)	0.648*** (0.092)	0.514*** (0.081)	0.477*** (0.114)	0.199 (0.137)	0.179** (0.089)
External pressure $\mathcal{P}_{r,t}^{(X)}$	0.069*** (0.022)	0.020** (0.008)	0.026*** (0.007)	0.006 (0.006)	0.072** (0.030)	0.283*** (0.106)					
Domestic pressure $\mathcal{P}_{r,t}^{(M)}$						-0.205** (0.096)					
China option $\mathcal{C}_{r,t}^{(1)}$						-0.087 (0.194)					
Common factor $\mathcal{P}_{r,t}^{(F)}$							0.027*** (0.006)	0.137* (0.077)			
Import substitution $\mathcal{S}_{r,t}^{(1)}$								-0.071 (0.261)			
Exports facilitation $\mathcal{F}_{r,t}^{(1)}$								-0.256 (0.234)			
Exports reallocation $\mathcal{R}_{r,t}^{(1)}$								-0.082 (0.069)			
Manufacturing labor share									-0.296** (0.122)	-0.440*** (0.126)	-0.396* (0.227)
Unemployment rate									0.206*** (0.050)	0.155** (0.061)	0.120 (0.100)
Techn. intens. of exports (TECH)									0.575** (0.223)	0.353* (0.199)	0.353* (0.199)
Workers with higher education (HEDU)										1.127 (1.521)	1.127 (1.521)
Interaction: HEDU * TECH										0.173 (0.278)	0.173 (0.278)
P-val.(Hansen)	0.136	0.099	0.342	0.34	0.074	0.196	0.277	0.32	0.236	0.191	0.085
P-val.(AR1)	0	0	0	0.002	0	0	0	0	0	0.001	0
P-val.(AR2)	0.842	0.217	0.112	0.382	0.603	0.783	0.952	0.704	0.488	0.976	0.428
Reg.	65	65	65	65	65	65	65	65	65	65	65
Instr.no.	22	22	22	22	22	22	22	22	22	22	22
Obs.	785	785	785	785	561	785	1037	785	887	802	717

\*p<0.1; \*\*p<0.05; \*\*\*p<0.01

Note:

Table A2: Results with separate types of instruments and their smaller number:  $\sim 9\%$  of the number of cross-sections

		<i>Dependent variable:</i>									
		LVAR									
Type of instrument:		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	$W^{(1)}$	$W^{(1)}$	$W^{(3)}$	$W^{(2)}$	$W^{(1)}$	$W^{(3)}$	$W^{(1)}$	$W^{(3)}$	$W^{(2)}$	$W^{(1)}$	$W^{(3)}$
Country of comparison:	JPN	JPN	JPN	-	USA	USA	JPN	JPN	-	USA	USA
Lag of dependent variable		0.137 (0.139)	0.142 (0.171)	0.315** (0.140)	0.191 (0.185)	-0.173 (0.240)	0.325*** (0.108)	0.261 (0.165)	0.468*** (0.108)	0.281** (0.131)	0.002 (0.229)
Common factor $\mathcal{P}_{r,t}^{(F)}$		0.022*** (0.007)	0.032*** (0.007)	0.022*** (0.007)	0.027*** (0.008)	0.029*** (0.010)					
External pressure $\mathcal{P}_{r,t}^{(X)}$							0.052** (0.021)	0.074*** (0.025)	0.043* (0.024)	0.060** (0.024)	0.076*** (0.029)
P-val.(Hansen)		0.706	0.127	0.477	0.234	0.299	0.337	0.089	0.222	0.173	0.131
P-val.(AR1)		0.004	0.012	0.001	0.018	0.483	0	0.001	0	0	0.144
P-val.(AR2)		0.332	0.419	0.628	0.485	0.135	0.537	0.499	0.752	0.466	0.191
Reg.		65	65	65	65	65	65	65	65	65	65
Instr.no.		6	6	6	6	6	6	6	6	6	6
Obs.		1037	1037	1037	1037	1037	785	785	785	785	785

*Note:*

\*p<0.1; \*\*p<0.05; \*\*\*p<0.01

Table A3: Results with regional projections fixed at  $R_{i,r,t} = R_{i,r,t_0}$ , for  $t > t_0 = 2000$

	Dependent variable:										
	LVAR	GINI	P5010	P9050	LVAR						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
Lag of explanatory variables:					lag = 0					lag = 1	lag = 2
Lag of dependent variable	0.505*** (0.063)	0.556*** (0.066)	0.598*** (0.072)	0.432*** (0.163)	0.553*** (0.065)	0.499*** (0.083)	0.553*** (0.087)	0.471*** (0.070)	0.416*** (0.076)	0.291*** (0.087)	0.189* (0.097)
External pressure $\mathcal{P}_{r,t}^{(X)}$	0.049** (0.022)	0.024*** (0.008)	0.024*** (0.007)	0.007 (0.006)	0.045* (0.025)	0.173* (0.094)					
Domestic pressure $\mathcal{P}_{r,t}^{(M)}$						−0.120 (0.088)					
China option $\mathcal{C}_{r,t}^{(1)}$						−0.077 (0.147)					
Common factor $\mathcal{P}_{r,t}^{(F)}$							0.022*** (0.006)	0.102* (0.054)			
Import substitution $\mathcal{S}_{r,t}^{(1)}$								−0.142 (0.242)			
Exports facilitation $\mathcal{F}_{r,t}^{(1)}$								−0.178 (0.164)			
Exports reallocation $\mathcal{R}_{r,t}^{(1)}$								−0.080 (0.065)			
Manufacturing labor share									−0.369*** (0.106)	−0.355*** (0.107)	−0.195 (0.143)
Unemployment rate									0.179*** (0.044)	0.184*** (0.060)	0.190** (0.083)
Techn. intens. of exports (TECH)										0.666* (0.396)	0.875 (0.539)
Workers with higher education (HEDU)											1.297* (0.673)
Interaction: HEDU * TECH											0.237 (0.145)
P-val.(Hansen)	0.268	0.383	0.942	0.584	0.558	0.258	0.335	0.373	0.431	0.632	0.424
P-val.(AR1)	0	0	0	0.003	0	0	0	0	0	0	0
P-val.(AR2)	0.773	0.214	0.124	0.468	0.535	0.772	0.888	0.709	0.584	0.426	0.32
Reg.	65	65	65	65	65	65	65	65	65	65	65
Instr.no.	46	46	46	46	46	46	46	46	46	46	46
Obs.	785	785	785	785	561	785	1037	785	887	802	717

\*p<0.1; \*\*p<0.05; \*\*\*p<0.01

Note:

Table A4: Results with lagged pressure indicator and period effects

Period effects:	<i>Dependent variable: LVAR</i>							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	No	No	No	No	Yes	Yes	Yes	Yes
Lag of dependent variable	0.310*** (0.067)	0.310*** (0.068)	0.300*** (0.067)	0.311*** (0.065)	0.295*** (0.070)	0.290*** (0.072)	0.271*** (0.079)	0.290*** (0.071)
Lag 0 of Common Factor $\mathcal{P}_{r,t}^{(F)}$	-0.002 (0.018)				-0.046 (0.106)			
Lag 1 of Common Factor $\mathcal{P}_{r,t}^{(F)}$		-0.001 (0.017)				0.017 (0.103)		
Lag 2 of Common Factor $\mathcal{P}_{r,t}^{(F)}$			-0.017 (0.021)				0.207 (0.129)	
Lag 3 of Common Factor $\mathcal{P}_{r,t}^{(F)}$	0.038** (0.016)	0.038** (0.017)	0.053** (0.021)	0.037*** (0.006)	0.216** (0.088)	0.230** (0.102)	0.182* (0.099)	0.238*** (0.088)
P-val.(Hansen)	0.392	0.39	0.403	0.428	0.378	0.341	0.428	0.396
P-val.(AR1)	0	0	0	0	0	0	0	0
P-val.(AR2)	0.889	0.894	0.81	0.899	0.54	0.595	0.682	0.52
Reg.	65	65	65	65	65	65	65	65
Instr.no.	46	46	46	46	40	40	40	40
Obs.	919	919	919	919	919	919	919	919

*Note:* \*p<0.1; \*\*p<0.05; \*\*\*p<0.01

Table A5: Results with additional control variables: patents and trends

	<i>Dependent variable: LVAR</i>					
	(1)	(2)	(3)	(4)	(5)	(6)
Lag of dependent variable	0.170 (0.157)	0.164 (0.165)	0.233* (0.127)	0.220 (0.158)	0.568*** (0.090)	0.291*** (0.109)
Common factor $\mathcal{P}_{r,t}^{(F)}$	0.049*** (0.010)	0.049 (0.079)	0.034*** (0.007)	0.036*** (0.008)	0.031* (0.018)	0.066*** (0.017)
Logarithm of number of patent applications (PAT)	-0.152** (0.075)	-0.163** (0.076)				
Interaction: $\mathcal{P}_{r,t}^{(F)} * \text{PAT}$		-0.0001 (0.018)				
Growth rate of patent applications (GRPAT)			-0.140 (0.097)	-0.093 (0.198)		
Interaction: $\mathcal{P}_{r,t}^{(F)} * \text{GRPAT}$				-0.016 (0.088)		
Linear trend term					-0.003 (0.006)	-0.046*** (0.011)
Quadratic trend term						0.001*** (0.0003)
P-val.(Hansen)	0.59	0.559	0.541	0.482	0.299	0.405
P-val.(AR1)	0.005	0.007	0	0.001	0	0
P-val.(AR2)	0.724	0.714	0.727	0.682	0.913	0.514
Reg.	65	65	65	65	65	65
Instr.no.	46	46	46	46	46	46
Obs.	667	667	650	650	1037	1037

\*p&lt;0.1; \*\*p&lt;0.05; \*\*\*p&lt;0.01

Note:



Table A6: Results with additional control variables: macroeconomic conditions, service exports, and the EU expansion with CEE countries

	<i>Dependent variable: LVAR</i>						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Lag of dependent variable	0.486*** (0.055)	0.551*** (0.082)	0.533*** (0.097)	0.483*** (0.066)	0.224** (0.093)	0.346*** (0.104)	0.168* (0.100)
Common factor $\mathcal{P}_{r,t}^{(F)}$	0.028*** (0.007)	0.024*** (0.007)	0.027*** (0.011)	0.048** (0.019)	0.045*** (0.011)	-0.018 (0.020)	-0.064*** (0.024)
Real effective exchange rate	0.008 (0.021)						
Economic sentiment index		-0.100 (0.070)					
Share of services in total exports			-0.051 (0.186)				
Logarithm of ratio of total exports: CEE to EU15				-0.239* (0.142)			
Growth rate of immigration					0.043* (0.024)		
lag(DCF_rmnf1 *y2002, 0)							0.153** (0.078)
y2002						0.170*** (0.051)	
lag(DCF_rmnf1 *y2004, 0)							-0.021 (0.074)
y2004						-0.005 (0.035)	-0.068 (0.042)
P-val.(Hansen)	0.368	0.364	0.289	0.303	0.655	0.274	0.418
P-val.(AR1)	0	0	0	0	0	0	0
P-val.(AR2)	0.578	0.88	0.975	0.799	0.708	0.615	0.201
Reg.	65	65	65	65	65	65	65
Instr.no.	46	46	46	46	46	46	46
Obs.	980	1022	983	785	446	1037	1037

*Note:* \*p<0.1; \*\*p<0.05; \*\*\*p<0.01

Table A7: Relevance of the impact channels: beta coefficients from regressions with standardized variables

Lag of explanatory variables:	<i>Dependent variable: LVAR</i>							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	lag = 0				lag = 1			
Lag of dependent variable	0.470*** (0.067)	0.509*** (0.091)	0.455*** (0.073)	0.532*** (0.078)	0.576*** (0.084)	0.416*** (0.076)	0.309*** (0.081)	0.208*** (0.092)
Manufacturing labor share	-0.414*** (0.107)				-0.443*** (0.127)		-0.429*** (0.115)	-0.314** (0.136)
Unemployment rate		0.318*** (0.078)			0.304*** (0.074)		0.314*** (0.086)	0.332*** (0.119)
Techn. intens. of exports (TECH)			1.484 (0.957)				1.945** (0.922)	1.782 (1.183)
White-collar workers				0.230* (0.127)				
Workers with higher education (HEDU)					0.126 (0.195)			0.435* (0.243)
Interaction: HEDU * TECH								0.283*** (0.140)
P-val.(Hansen)	0.447	0.241	0.197	0.254	0.218	0.431	0.638	0.483
P-val.(AR1)	0	0	0	0	0	0	0	0
P-val.(AR2)	0.438	0.359	0.608	0.42	0.351	0.584	0.659	0.766
Reg.	65	65	65	65	65	65	65	65
Instr.no.	46	46	46	46	46	46	46	46
Obs.	887	887	968	886	859	887	802	717

*Note:*

All variables standardized. \*p<0.1; \*\*p<0.05; \*\*\*p<0.01

## C Appendix: Econometric considerations

### C.1 Dynamic common factor: Extraction model and empirical characterization

Using the `MARSS()` function from the R package `MARSS` (see Holmes *et al.*, 2020), the dynamic common factor ( $\mathcal{P}_{r,t}^{(F)}$ ) is derived from the following state-space representation:

$$\begin{bmatrix} \tilde{\mathcal{P}}_{r,t}^{(X)} \\ \tilde{\mathcal{P}}_{r,t}^{(M)} \end{bmatrix} = \begin{bmatrix} a_{X,r} \\ a_{M,r} \end{bmatrix} \mathcal{P}_{r,t}^{(F)} + \begin{bmatrix} \varepsilon_{r,t}^{(X)} \\ \varepsilon_{r,t}^{(M)} \end{bmatrix}, \quad \left( \varepsilon_{r,t}^{(X)}, \varepsilon_{r,t}^{(M)} \right)' \sim N \left( \begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \sigma_{X,r}^2 & 0 \\ 0 & \sigma_{M,r}^2 \end{bmatrix} \right) \quad (15)$$

$$\mathcal{P}_{r,t}^{(F)} = \mathcal{P}_{r,t-1}^{(F)} + \zeta_{r,t}, \quad \zeta_{r,t} \sim N(0, 1), \quad \text{Corr}(\zeta_{r,t}, \varepsilon_{r,t}^{(i)}) = 0, \quad i \in \{X, M\},$$

where:

$\tilde{\mathcal{P}}_{r,t}^{(X)}$  and  $\tilde{\mathcal{P}}_{r,t}^{(M)}$  stand for the standardized external  $\mathcal{P}_{r,t}^{(X)}$  and domestic  $\mathcal{P}_{r,t}^{(M)}$  pressure series, respectively;

$\sigma_{X,r}^2$  and  $\sigma_{M,r}^2$  are the variances of the corresponding individual components  $\varepsilon_{r,t}^{(X)}$  and  $\varepsilon_{r,t}^{(M)}$ ; and

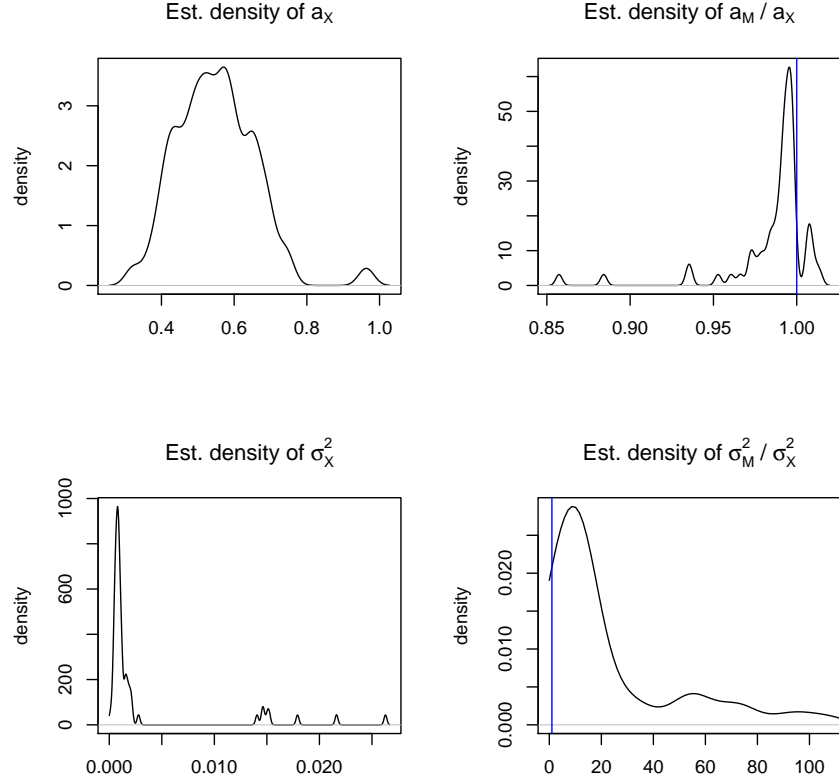
$a_{X,r}$  and  $a_{M,r}$  are the corresponding loadings of the common factor  $\mathcal{P}_{r,t}^{(F)}$ , specific to each separate region indexed by  $r$ .

Besides the extraction of the dynamic common factor of the trade pressure, system (15) can also be used to obtain a linear prediction for the missing observations, i.e., where some data of the external and/or domestic pressure are missing.

Figure A4 provides some intuition about the importance of the (estimated) loadings. The top-left panel presents the distribution of  $a_{X,r}$ , the top-right panel plots the distribution of  $a_{M,r}/a_{X,r}$ , and the bottom-left panel characterizes  $\sigma_{X,r}^2$  in an analogous manner, whereas the bottom-right panel depicts the distribution of  $\sigma_{M,r}^2/\sigma_{X,r}^2$ . Typically, the loading of the external pressure  $a_{X,r}$  is larger than that of the domestic pressure, as  $a_{M,r}/a_{X,r} < 1$  holds in most cases. The idiosyncratic component of the domestic pressure is also much larger, i.e.,  $\sigma_{M,r}^2 > \sigma_{X,r}^2$ . This also shows that the external

pressure is more informative about the common component than the domestic one in the stochastic process defined by system (15), just as was expected using economic arguments in the Introduction.

Figure A4: Density plots of regional values of the estimated parameters of system (15)



## C.2 The reduced and semi-reduced form with the calculation of the importance of channels

To simplify the presentation, we consider a stylized representation of the reduced-form equation with a single pressure indicator, say  $\mathcal{P}_{r,t}$ , and without individual effects or other control variables:

$$\alpha(L)I_{r,t} = \beta\mathcal{P}_{r,t} + \xi_{r,t}, \quad E(\xi_{r,t}|\mathbf{W}_{r,t}) = 0, \quad (16)$$

where  $\mathbf{W}_{r,t}$  comprises a vector of instruments. Assuming that the polynomial  $\alpha(L)$  has no unit roots, the long-run impact of the pressure  $\mathcal{P}_{r,t}$  on inequality is defined by  $\alpha^{-1}(1)\beta$  from

$$I_{r,t} = \alpha^{-1}(L)\beta\mathcal{P}_{r,t} + \alpha^{-1}(L)\xi_{r,t}, \quad (17)$$

with its conditional expectation with respect to the instruments given by

$$E(I_{r,t}|\mathbf{W}_{r,t}) = \alpha^{-1}(L)\beta E(\mathcal{P}_{r,t}|\mathbf{W}_{r,t}). \quad (18)$$

The employed semi-reduced system that evaluates the contribution of channels  $Z_{r,t}^{(1)}, Z_{r,t}^{(2)}, \dots, Z_{r,t}^{(k)}$  is given by

$$\begin{aligned} \alpha_1(L)I_{r,t} &= \theta_1 Z_{r,t}^{(1)} + \theta_2 Z_{r,t}^{(2)} + \dots + \theta_k Z_{r,t}^{(k)} + \zeta_{r,t}, \quad E(\zeta_{r,t}|\mathbf{W}_{r,t}) = 0, \\ \lambda_1(L)Z_{r,t}^{(1)} &= \gamma_1 \mathcal{P}_{r,t} + v_{r,t}^{(1)}, \quad E(v_{r,t}^{(1)}|\mathbf{W}_{r,t}) = 0, \\ \lambda_2(L)Z_{r,t}^{(2)} &= \gamma_2 \mathcal{P}_{r,t} + v_{r,t}^{(2)}, \quad E(v_{r,t}^{(2)}|\mathbf{W}_{r,t}) = 0, \\ &\vdots \\ \lambda_k(L)Z_{r,t}^{(k)} &= \gamma_k \mathcal{P}_{r,t} + v_{r,t}^{(k)}, \quad E(v_{r,t}^{(k)}|\mathbf{W}_{r,t}) = 0, \end{aligned} \quad (19)$$

where the lag polynomials  $\alpha_1(L)$  and  $\lambda_i(L)$ ,  $i \in \{1, 2, \dots, k\}$  are again assumed to have all roots outside the unit circle. Note that we call this system ‘semi-reduced’ from the economic point of view, as there is no simultaneity between the channel components  $Z_{r,t}^{(i)}$ ,  $i \in \{1, 2, \dots, k\}$ . Here, we just use simple projections for the decomposition.

Using the inversion of polynomials in all the equations of the system, placing the respective solutions of  $Z_{r,t}^{(i)}$ ,  $i \in \{1, 2, \dots, k\}$  into the first equation of system (19), and

taking the conditional expectation with respect to  $\mathbf{W}_{r,t}$  yields:

$$\begin{aligned}
E(I_{r,t}|\mathbf{W}_{r,t}) &= E\left(\alpha_1^{-1}(L)\left(\theta_1 Z_{r,t}^{(1)} + \theta_2 Z_{r,t}^{(2)} + \cdots + \theta_k Z_{r,t}^{(k)}\right) \middle| \mathbf{W}_{r,t}\right) \\
&= E\left(\alpha_1^{-1}(L)\left(\theta_1 \lambda_1^{-1}(L)\gamma_1 \mathcal{P}_{r,t} + \theta_2 \lambda_2^{-1}(L)\gamma_2 \mathcal{P}_{r,t} + \cdots + \theta_k \lambda_k^{-1}(L)\gamma_k \mathcal{P}_{r,t}\right) \middle| \mathbf{W}_{r,t}\right) \\
&= \alpha_1^{-1}(L)\left(\theta_1 \lambda_1^{-1}(L)\gamma_1 + \theta_2 \lambda_2^{-1}(L)\gamma_2 + \cdots + \theta_k \lambda_k^{-1}(L)\gamma_k\right) E(\mathcal{P}_{r,t}|\mathbf{W}_{r,t}).
\end{aligned} \tag{20}$$

Equating eqs. (18) and (20), canceling  $E(\mathcal{P}_{r,t}|\mathbf{W}_{r,t})$ , and evaluating at  $L = 1$  gives the decomposition of the total long-run impact  $(\alpha^{-1}(1)\beta)$  into the parts associated with the channel contributions:

$$\alpha^{-1}(1)\beta = \sum_{i=1}^k C_i. \tag{21}$$

Here, the  $i^{th}$  channel contribution  $C_i$  is given by

$$C_i = \alpha^{-1}(1)\lambda_i^{-1}(1)\gamma_i\theta_i. \tag{22}$$

Given that  $\alpha^{-1}(1)\beta \neq 0$  and  $C_i \geq 0$  for all  $i \in \{1, 2, \dots, k\}$ , eq. (21) allows for the decomposition into shares  $\sum_{i=1}^k \frac{C_i}{\alpha^{-1}(1)\beta}$  and therefore directly reveals the importance of different channels. The non-negativity holds in the cases underlying Figure 2. Here, the manufacturing share's case is of interest, because its gamma and theta coefficients are negative individually, but their product is positive and satisfies the stated requirement.

### C.3 Appendix: Multi-year difference versus the GMM estimation

Many of the key studies in the related literature (such as Autor *et al.*, 2013, and its followers) use long differences (taken over many years) in one way or another for estimation, whereas we are using a panel of yearly data with the GMM estimator in the first differences. In this subsection, we first discuss the choice of the yearly GMM estimator, comparing it to three alternatives: (i) panels formed from overlapping (moving) longer differences; (ii) an estimation based on differences, where the pre- and post-impact differences over more extended periods are stacked (or other versions of a panel structure where one uses non-overlapping subsample differences); and (iii), the simple difference, where only the post-impact differences are employed. After discussing why options (i) and (ii) are less efficient choices in our case, we still present the simple estimation result based on approach (ii) without controlling for potential endogeneity or the presence of other driving factors through the proper use of instruments, as we do in the GMM. Importantly, the obtained estimate with this simple multi-year difference approach is in between the GMM-based long-run estimates that take and do not take the financial crisis into consideration, in connection with Columns (8) and (3) of Table 1, respectively. Finally, we show how the simplistic approach in (iii) might be erroneous whenever substantial cross-sectional variation dominates the contribution of the pressure indicator.

The usage of a panel with overlapping samples over many years, as in the first option stated above, is well known to result in a moving average error term of about the same order as the number of years within the window (see Rowley and Wilton, 1973, Hansen and Hodrick, 1980). If one used moving averages of ‘longer differences’, the error term would be substantially correlated with a lengthy lag of the dependent variable, thus requiring the inclusion of a large number of autoregressive terms (to approximate a high-order moving average process) or an explicit modeling of the moving average component in the error process. Ignoring this problem would hinder

statistical inference. It is barely possible to address this issue in our case because the effective number of years is not large and has many gaps in the middle of the sample. This approach is highly inefficient at best relative to the standard yearly panel GMM approach, which does not suffer from serially correlated errors at higher-order lags. In our case, only the first lag was problematic, just as it is expected to be with the first differences approach in panels, whereas the second-order correlation was already insignificant. Thus, by using the GMM, we can expect to get not only consistent, but also more efficient estimates, as fewer data are lost due to lagged series/instruments, whereas more data would be lost in the moving average approach with ‘long multi-year differences’ by construction.

The second option that stacks the pre-impact and post-impact multi-year differences does not have the above problem because the two differences do not overlap. It does, however, reduce the number of observations quite substantially because only two (long) periods contribute to the estimates (before and after the impact). The total number of observations is only twice the number of the cross-sectional units, much less than what the GMM uses with all available periods included. This loss of observations makes the GMM more efficient. Furthermore, the dynamic panel specification allows us to explicitly separate the immediate and long-run impacts.

In order to show the difference using our data, we present stylized facts and estimates based on long differences. The left panel of Figure A5 plots the two multi-year differences over two periods on which the estimation can rely. Blue color identifies the difference before the impact; red and brown identify the second long difference. Because German regions do not have regional identifiers before 2002 in the original data from the LFS, the second difference is somewhat shorter.<sup>48</sup> We mark them in brown. Notice that dropping the first part of data and using only the second difference alone (see the right panel of the figure, which plots the same brown/red points, magnifying the picture) would result in erroneous inference of an absent relationship.<sup>49</sup>

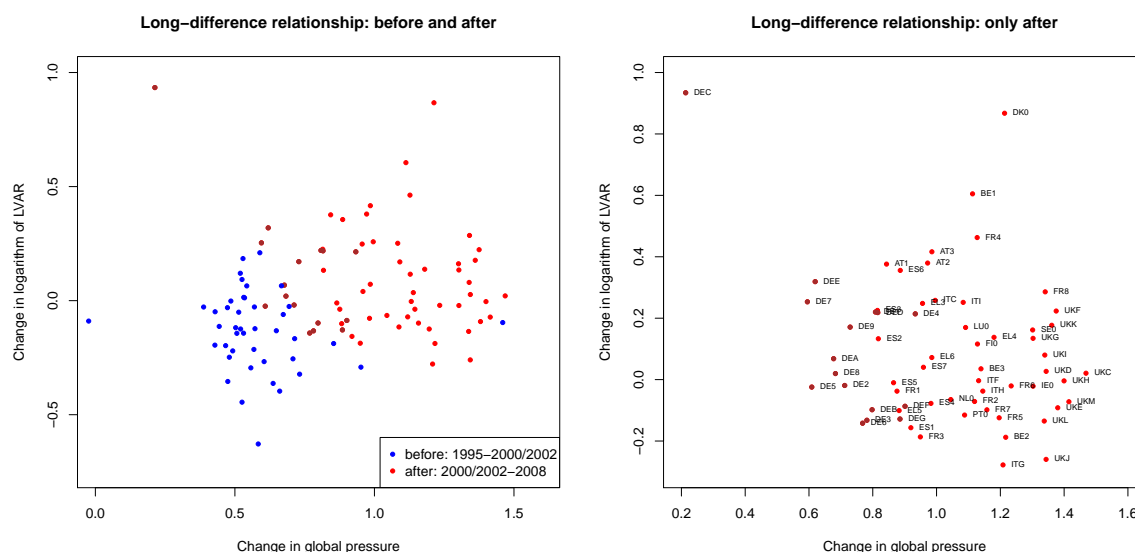
---

<sup>48</sup>We believe that this asymmetry should not create much difference, because right after China’s WTO entrance, there was little impact if any.

<sup>49</sup>As region ‘DEC’ is a clear outlier and highly influential, we have removed it from the estimation



Figure A5: Empirics with ‘long differences’



Column (1) of Table A8 presents simple ordinary least squares estimates using the stacked pre- and post-impact long differences (without controlling for any endogeneity or omitted variable bias). This specification is akin to the one producing based estimates in Column (3) of Table 1 with the GMM estimator. Observe that here the implied long-run impact is 0.11, whereas the parameter estimate from long differences is close, but somewhat larger (0.14). At the same, it is smaller than the long-run impact in the base GMM estimates when the potential impact of the financial crisis is taken into account, as in Column (8) of Table 1. In this case, the estimated long-run impact would constitute about 0.2.

Because only a small share of the total variance is explained by the model with multi-year differences (less than 7%; see Column (1) in Table A8) and it is dominated by the substantial cross-sectional variation, discarding information from before China's WTO accession is costly. Column (2) in Table A8 shows the long-difference estimate of the critical parameter using post-accession data only. As expected, such estimates are severely biased and very imprecise.

To show that this is not merely an artifact in our data, we performed a Monte  


---

sample. The same results hold if one removes the outlying region ‘DK0’, Denmark.

Table A8: Results with long differences: actual ( $\Delta\text{LVAR}$ ) and simulated ( $\Delta y$ ) data

	<i>Dependent variable:</i>			
	$\Delta\text{LVAR}$	$\Delta\text{LVAR}$	$\Delta y$	$\Delta y$
	(2 per.)	(1 per.)	(2 per.)	(1 per.)
	(1)	(2)	(3)	(4)
GLOB	0.141*** (0.051)	−0.072 (0.117)		
$p$			0.129*** (0.048)	−0.065 (0.110)
Constant	−0.118** (0.047)	0.154 (0.125)	0.086* (0.045)	0.355** (0.138)
Observations	108	63	130	65
$R^2$	0.068	0.006	0.053	0.005
Adjusted $R^2$	0.059	−0.010	0.046	−0.010
Residual Std. Error	0.235	0.215	0.417	0.404
F Statistic	7.739***	0.381	7.225***	0.347
<i>Note:</i>		*p<0.1; **p<0.05; ***p<0.01		

Carlo simulation that replicates the main features of the model discussed earlier.<sup>50</sup> For clarity, we keep the underlying data generating process simple (absent confounding factors, endogeneity, and sector-specific effects). We generate the ‘pressure’ variable  $p$  as

$$p_{i,t} = \exp\{\mathbb{1}_{\{t>5\}} \cdot 0.1(t-5)\} + e_{i,t}, \quad e_{i,t} \sim \text{i.i.d.N}(0, 0.1),$$

and the dependent variable as

$$y_{i,t} = 10 + 0.2 \cdot p_{i,t} + u_{i,t}, \quad u_{i,t} \sim \text{i.i.d.N}(0, 0.1),$$

where  $e_{i,t}$  and  $u_{j,s}$  are independent for all possible values of indices  $i, t, j, s$ . The number of cross-sectional units is fixed at 65, while the number of periods is fixed at 13, with the impact arriving in the sixth year.<sup>51</sup>

We summarize the simulation results in Figure A6. The pressure variable is in the top-right and the dependent variable in the top-left panel; colors identify different regions’ realizations over time. The lower panel replicates the situation observed earlier in Figure A5 with the two long differences (before and after the impact) as in approach (ii) and the single long difference using only the data since the initiation of the impact (left and right panel, respectively) as in approach (iii). Note that even though the true coefficient of the impact of  $p$  is 0.2, the latter suggests even a ‘negative tendency’. The simple ordinary least squares estimation results we report in Columns (3) and (4) of Table A8 are in line with what one can infer from these panels upon visual inspection. Column (3) shows a highly significant estimate<sup>52</sup>. The estimate in column (4) gives an insignificant coefficient with the wrong sign. Hence, option (iii) is not only inefficient but also likely to be severely biased in general.

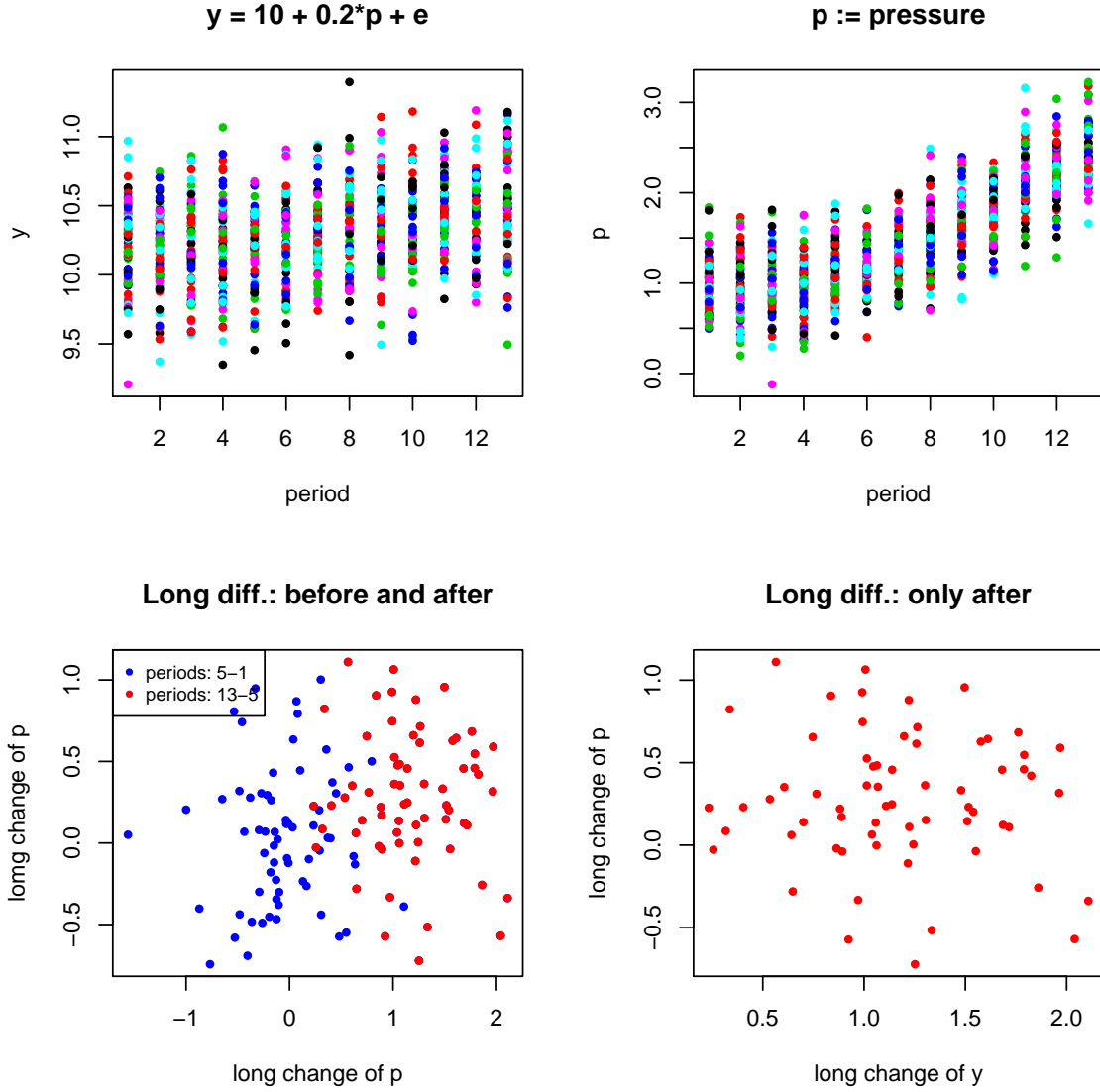
---

<sup>50</sup>The R code reproducing the results is available upon request. The discussion here relies on a single realization (with `set.seed(1)`), but it is the typical one.

<sup>51</sup>In the empirical counterpart, these would correspond to the period from 1997 to 2008 with the impact starting in 2002.

<sup>52</sup>Note that the hypothesis that the parameter equals the true 0.2 value cannot be rejected at the usual significance levels: the p-value of such test is 0.14.

Figure A6: Simulation with ‘long differences’



## D Appendix: Stylized calculation of the impact

To define the counterfactual scenario of ‘no accession’, we exploit the structural break observed in the country-level pressure indicator, extending the trend-line observed until 2001 also after China’s WTO accession. The left panel of Figure A7 plots (black circles) the dynamics of the logarithms of actual data, i.e.,  $\log(\mathcal{P}_{i,t}^{(XC)})$ , after demeaning them by country to remove the country-level fixed effects. The estimated baseline is presented as a red line, with the 95% confidence bands in blue. Note that the counter-factual

scenario is given by the scatter of country-specific values (in red), which is fixed at 2001 and extended following the trend-line for later years. The difference between the (logarithms of) actual data depicted in black circles and the baseline-based red points of the counter-factual is plotted in the middle panel of the figure and will constitute the primary country-level trade shock, which hereafter is denoted by  $\Xi_{i,t}^{(X)}$ . To get a glance at the dynamics of the shock with non-demeaned data, the right panel of Figure A7 plots the actual and the counter-factual (non-demeaned) logarithmically transformed data obtained by adding back the country-specific means. In our opinion, a visual fitting and extension by eye of the data before 2001, i.e., before China's accession to the WTO, would concur with the red dots of the counter-factual scenario rather than with the sudden increase in actual values depicted by the black circles.

Next, we turn to the calculation of the impact. For that purpose, we use the specification with the structural break in the impact due to the financial crisis, as presented in Column (8) of Table 1. Recall that  $I_{r,t}$  denotes the logarithm of log-variance of income and  $\mathcal{P}_{r,t}^{(X)} = \log(\mathcal{P}_{i,t}^{(XC)} R_{i,r,t})$ . Let us write the equation representing the logarithms of actual data and corresponding to Column (8) in Table 1 as

$$I_{r,t} = \alpha I_{r,t-1} + (\beta_0 + \beta_1 \cdot \mathbb{1}\{t \geq 2008\}) \cdot \mathcal{P}_{r,t}^{(X)} + u_{r,t}, \quad (23)$$

where the remainder  $u_{r,t}$  includes other terms of no interest for our calculations. Assuming that neither it nor the parameters of the equation are affected by China's accession, the difference between the logarithms of actual values of inequality and the values in the counter-factual scenario of 'no accession' (identified by a tilde) would be given by

$$\Xi_{r,t}^{(I)} := I_{r,t} - \tilde{I}_{r,t} = \alpha(I_{r,t-1} - \tilde{I}_{r,t-1}) + (\beta_0 + \beta_1 \cdot \mathbb{1}\{t \geq 2008\}) \cdot (\mathcal{P}_{r,t}^{(X)} - \tilde{\mathcal{P}}_{r,t}^{(X)}). \quad (24)$$

Hence, recalling that  $t_0$  denotes the pre-accession year, the dynamic impact over

the post-accession period is given, in log-difference terms, by

$$\Xi_{r,t_0+h}^{(I)} = \sum_{j=0}^{h-1} \alpha^j (\beta_0 + \beta_1 \cdot \mathbb{1}\{t_0 + h - j \geq 2008\}) \cdot \Xi_{i,t+h-j}^{(X)}, \quad h \in \{1, 2, \dots\}, \quad (25)$$

where it was exploited that

$$I_{r,t} - \tilde{I}_{r,t} = \begin{cases} 0, & t \leq t_0 \\ \neq 0, & t > t_0 \end{cases},$$

and

$$\begin{aligned} \mathcal{P}_{r,t}^{(X)} - \tilde{\mathcal{P}}_{r,t}^{(X)} &= \log(\mathcal{P}_{i,t}^{(XC)} R_{i,r,t}) - \log(\tilde{\mathcal{P}}_{i,t}^{(XC)} R_{i,r,t}) \\ &= \log(\mathcal{P}_{i,t}^{(XC)} / \tilde{\mathcal{P}}_{i,t}^{(XC)}) = \begin{cases} 0, & t \leq t_0 \\ \Xi_{r,t}^{(X)}, & t > t_0 \end{cases}. \end{aligned}$$

Consequently, using the log-difference defined in eq. (25), the ratio of the actual inequality to the hypothetical baseline (without a logarithmic transformation) is given by  $\exp(\Xi_{r,t_0+h}^{(I)})$ , whereas the change in the level of inequality (again without a logarithmic transformation) in each post-accession year is defined by

$$\Delta_{r,t_0+h} = \exp(I_{r,t_0+h}) - \exp(\tilde{I}_{r,t_0+h}) = \exp(I_{r,t_0+h}) \cdot \left(1 - \frac{1}{\exp(\Xi_{r,t_0+h}^{(I)})}\right), \quad h \in \{1, 2, \dots\}. \quad (26)$$

The top-right panel of Figure A8 presents the dynamics of inequality change due to China's WTO accession ( $\Delta_{r,t}$ ), as calculated in eq. (26) using the primary shocks to the external pressure  $\Xi_{r,t}^{(X)}$ , defined previously and plotted in Figure A7. Here, the regional distribution is captured by the box-plot of the calculated regional increase in inequality. The top panel of Figure A9 fixes the year to 2011 and shows the predicted change in inequality for all regions in decreasing order of magnitude.

To get a further sense of the economic significance, we compare the dynamics of

the predicted absolute increase with the actual changes. First, the top-left panel of Figure A8 plots the actual inequality levels, measured in terms of income log-variance. Each red dot represents the median of regional inequality values for a fixed year. The bottom-left panel of Figure A8 then compares the box-plots of the actual increase in inequality levels since 2000 with the change predicted by the previously described procedure. Note that the range in orange color in the bottom-left figure corresponds exactly to the interquartile range in the top-right panel of Figure A8, which also is of the same color. The median levels of actual changes are depicted by the central black dashes of each box. By taking the ratio, the bottom-right panel of Figure A8 relates the median of the predicted values due to China's WTO accession to the median of the actual change since 2000. It reveals that even before the financial crisis, the predicted Chinese pressure impact was sizable. It is also important to stress that the actual inequality levels tended to decrease before China's WTO accession in the considered EU15 regions.

The lower panel of Figure A9 presents additional scatter-plots of actual vs. calculated changes in the log-variance of income over the periods 2000–2007 and 2000–2011. It reveals that several UK regions are predicted to be among the most affected ones in both 2007 and 2011.

Figure A7: The estimated baseline of the counter-factual scenario of ‘no accession’ and the simulated initial shock

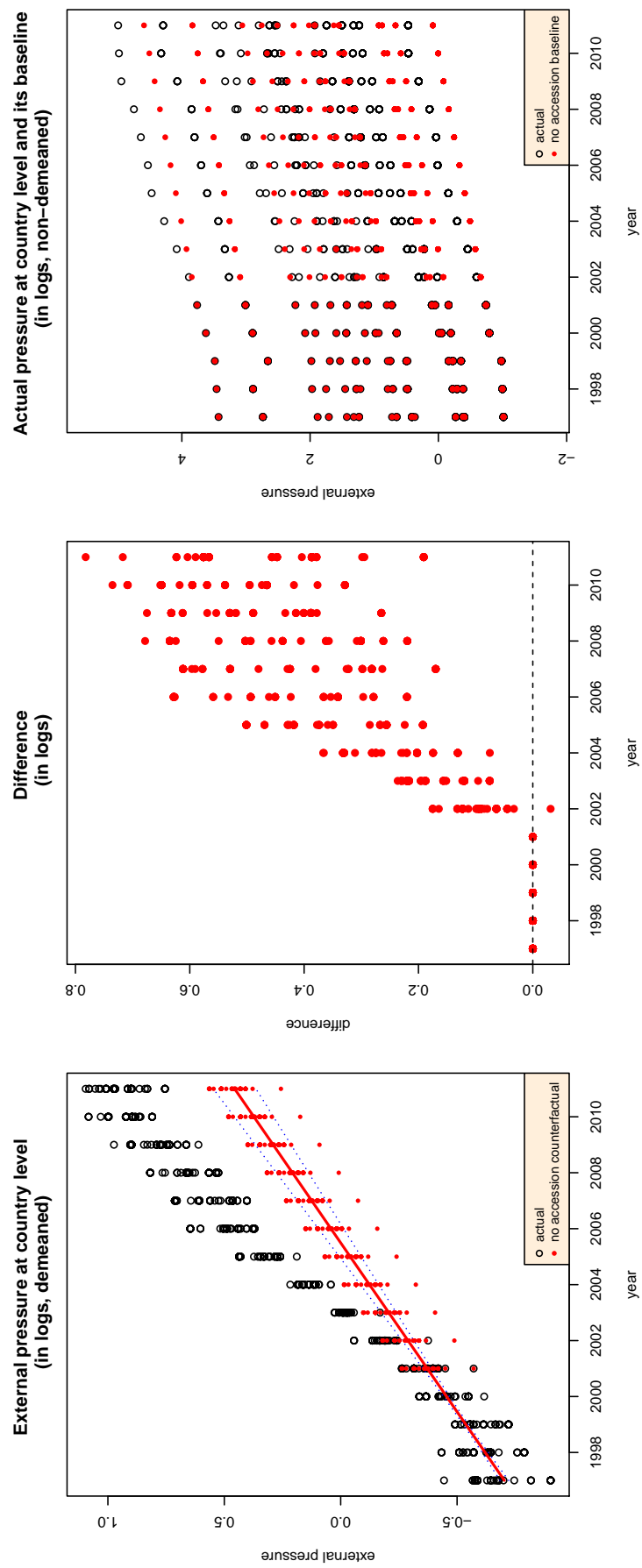




Figure A8: Dynamics of actual inequality (log-variance) and the predicted impact (increase) with their comparison

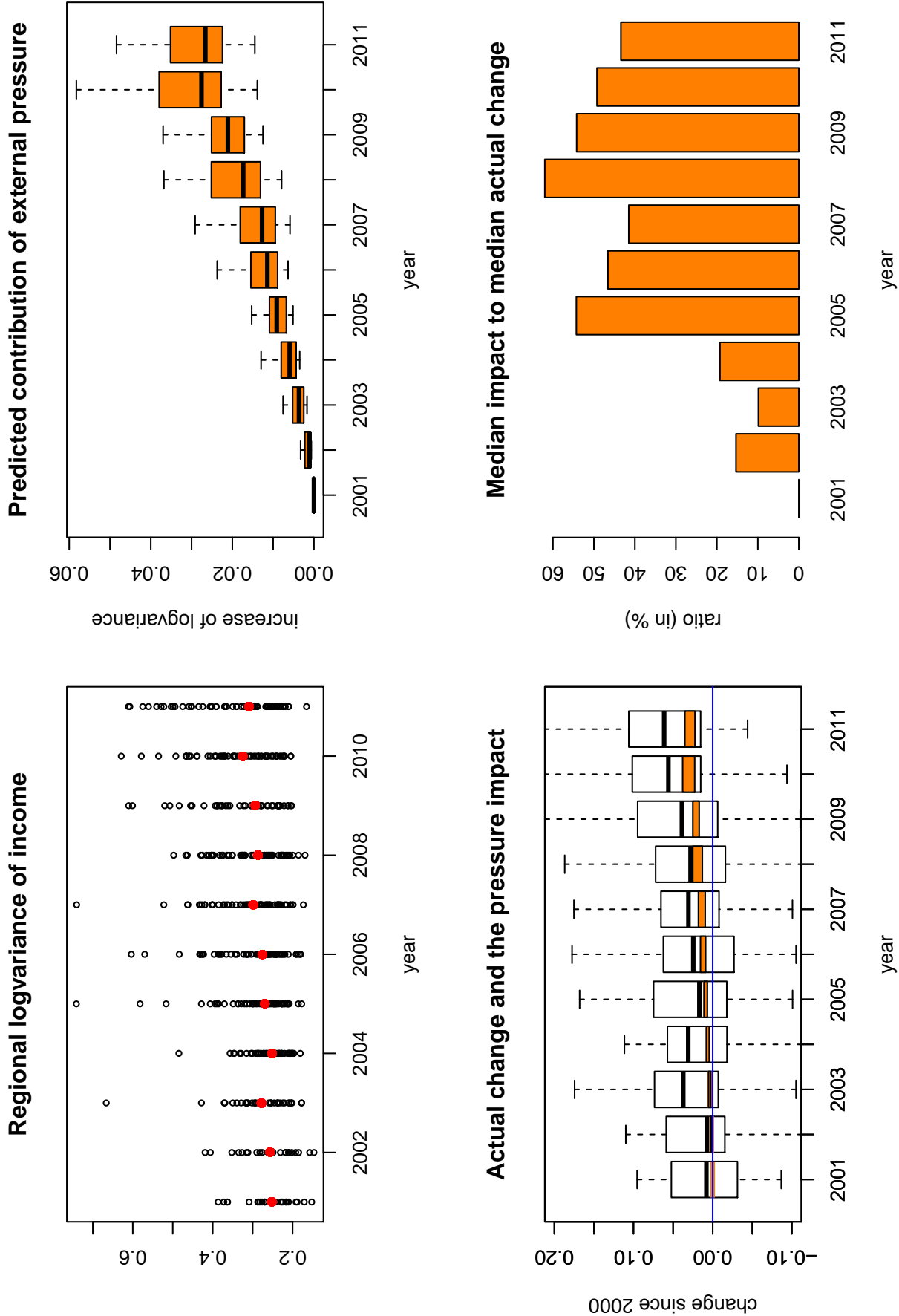
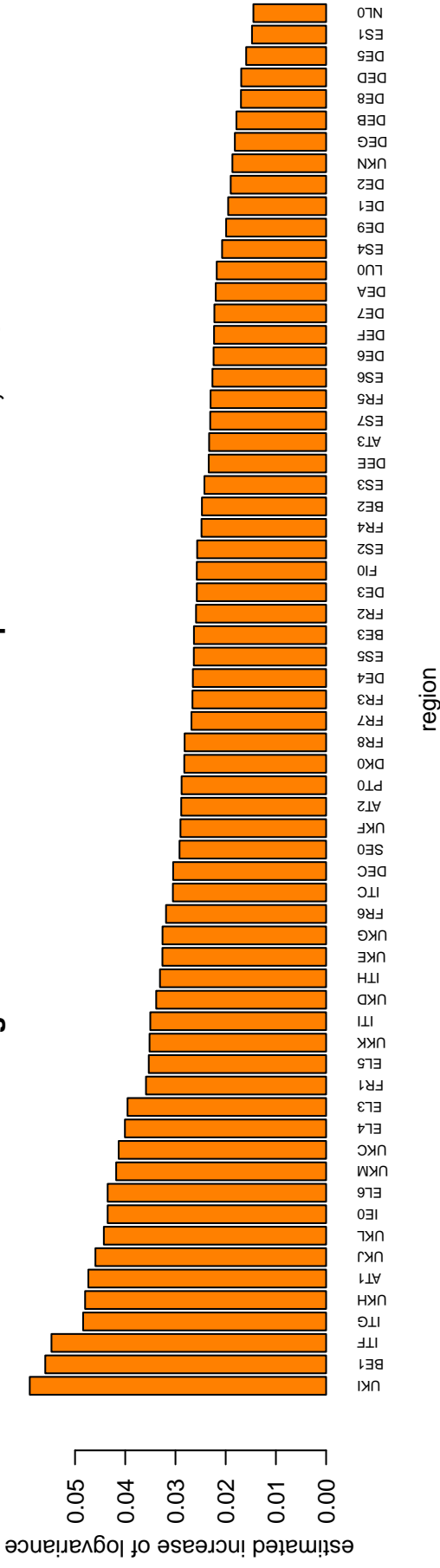
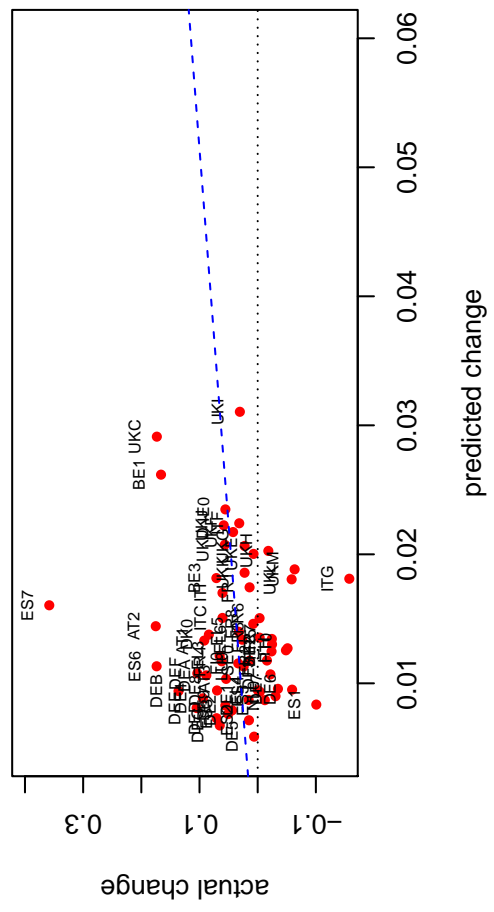


Figure A9: Predicted increase in the log-variance of income in regions in 2011 (upper panel) and its comparison with the actual increase since 2000 (lower panel)

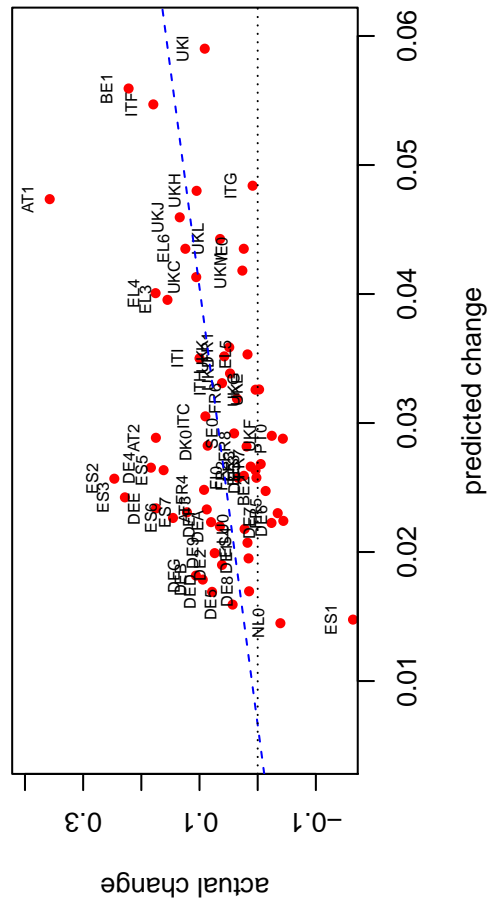
**Estimated China's trade pressure absolute impact  
on logvariance of net household equivalized income, in 2011**



**Actual vs predicted change since 2000, in 2007**



**Actual vs predicted change since 2000, in 2011**



## **E Appendix: Details on data and rich stylized facts**

### **E.1 Data on trade**

The basis for our analysis of trade pressure is the OECD-WTO Trade in Value Added (TiVA) database.<sup>53</sup> To calculate the external pressure indicator, we use some standard geographical aggregates considered in the OECD TiVA database, taking into account the intensity of trade with the EU as well as the geographic or socioeconomic proximity of countries. Consequently, we consider the following eight aggregate export destination markets: i) ASEAN countries, ii) Australia, India, Japan, Korea, iii) EU13 (EU28-EU15), iv) NAFTA, v) South America, vi) Russia and Turkey, vii) Switzerland, Norway, South Africa, Saudi Arabia, viii) Rest of World. The last one is defined by the difference between the total manufacturing exports and the sum of manufacturing exports to destination markets i)-vii). Groups i)-v) are based mostly on the geographical/socioeconomic proximity of countries, whereas groups vi) and vii) are based more on their trade intensity with the European countries.

For some descriptive statistics, we also use the IMF balance of payments (BOP) statistics database and the OECD Bilateral Trade in Goods by Industry and End-use (BTDIxE) database (ISIC\_Rev.4).<sup>54</sup>, which contains also the split of exports and imports by technological intensity

#### **E.1.1 The relative expansion of China's exports in US and EU15 markets**

The left panel of Figure A1 shows that China's export share of world exports has increased dramatically after its accession to the WTO by the end of 2001, while the shares of the US, Germany, and France dropped. The relative increase of Chinese exports persists also if considered in different markets. Figure A2 compares exports of goods (all goods and only manufactured goods) from China and Germany, France, and Japan to the US and the (former) EU15 destination markets. Since 2002, a structural

---

<sup>53</sup>To be found at <http://oe.cd/tiva>.

<sup>54</sup>To be found at <https://stats.oecd.org/index.aspx?queryid=64755>.

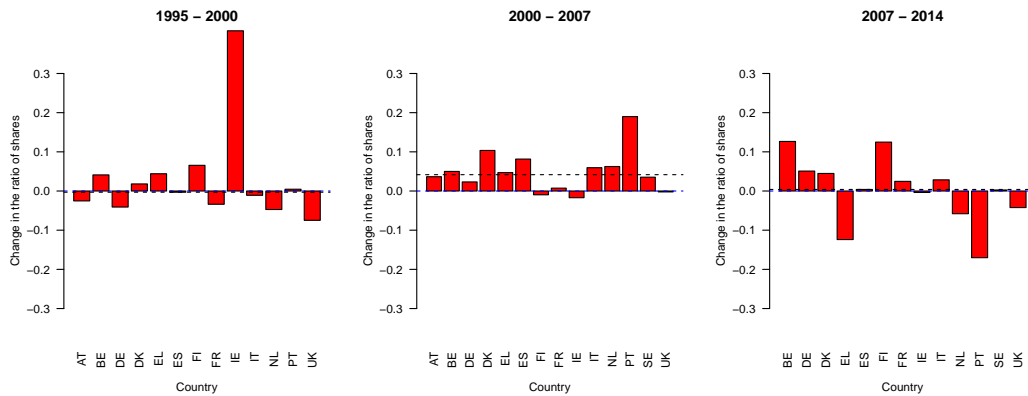
break appears, which is somewhat more pronounced for manufacturing than all goods. This pattern holds for many other countries (not reported).

It should be pointed out that not only the export shares of European and many other countries started to dropping after China's WTO accession, but also the growth rates of total exports decreased (see Figure A3).

### E.1.2 Changes in the technological (R&D) intensity of exports

Here, we provide some empirical evidence on the increasing technological intensity in exports of intermediate goods, which is consistent with the importance of vertical specialization induced by competition between developed and developing countries, stressed by Krugman (2008). In order to avoid the purely global demand-driven changes in exports structure, we consider the relative export shares, by comparing the share of exports with different R&D intensity of a country to the respective share of world exports of the same intensity.<sup>55</sup>

Figure A10: Changes in the share of high and medium-high technological intensity intermediate goods in exports of the EU15 countries relative to the respective share of world exports



Note: The dashed line represents the median change value.

Source: own calculations based on the OECD bilateral trade statistics.

The relative share of the high and medium-high technological intensity interme-

<sup>55</sup>Since we do not have actual data on the composition of world exports by technological intensity, we proxy these by those derived from all available countries in the OECD BTDiX bilateral trade statistics database.

diates has increased during 2000-2007 in almost all countries, whereas there was no such clear tendency during the 1995-2000 period and the median value (represented by a dashed horizontal line) was close to zero both in this and the 2007-2014 period – see Figure A10 on the preceding page. The dynamic pattern is thus consistent with a potential upgrade of manufacturing exports of intermediate goods induced by China’s accession to the WTO (see e.g. Bloom *et al.*, 2016). It is worth stressing that such a shift seems to be more connected with intermediates as we have not witnessed similar changes for other products.

## E.2 Data on inequality

In order to study inequality, we look at the target reference period of 1994-2014.<sup>56</sup> The primary basis of our database is the European Community Household Panel (ECHP, for 1995-2000) and the European Union Survey of Income and Living Conditions (EU-SILC, for 2003-2014).<sup>57</sup> Both collect yearly individual-level microdata on the region the household lives, demographic characteristics, income and living conditions for individuals in households. The two surveys share the same aim and try to be compatible with each other. There are some methodological differences, but none of these affects our analysis directly. Nevertheless, the timing of the two surveys implies that there are no data for years 2001 and 2002.

Although Germany and the UK are part of both surveys, regional identifiers are not present for them, and thus we decided to drop and substitute their data with those from the German Socio-Economic Panel (GSOEP) and the British Household Panel Survey (BHPS), respectively. Being the sources for the ECHP, both the BHPS and the GSOEP have similar sample designs, resolution and income concepts. The conceptual

---

<sup>56</sup>Note that from now on we use ‘year’ and ‘period’ where we mean ‘reference year’ and ‘reference period’. This distinction is important because almost all information on income and labor market participation relate to the reference year, which is the year preceding the interview for all countries except for the UK and Ireland (where it is the year of the interview).

<sup>57</sup>Note that the ECHP started in 1993, but we decided not to use years before 1994 as not all variables needed were present. Also, note that not all countries are present in either the ECHP or in the EU-SILC from the beginning.

similarity of the EU-SILC and the ECHP provides a bridge to the former too.<sup>58</sup> Unlike other countries, data for Germany and the UK are available also for 2001 and 2002, but they are still missing for the UK in 2009, where a new survey called Understanding Society carries over from 2010.

The extent to which regional identifiers are recorded in the datasets are different, some containing NUTS2 identifiers, some containing nothing at all. Besides using supplementary data for Germany and the UK, this forced us to treat the Netherlands, Portugal (islands merged) and Finland (Åland merged) as a single NUTS1 unit besides Denmark, Sweden, Ireland, Portugal mainland which are one unit by definition. Regional identifiers in all surveys are recoded to correspond to 2014 NUTS1 identifiers.<sup>59</sup>

It is instructive to compare the size of the resulting regions to those of US states. The 65 EU regions have a total implied population of 194 million persons in our sample in 2010, while the 50 US states had a total of 310 million. Minimum, maximum, mean and median were 0.5, 8.8, 3 and 2.4 in the EU and 0.5, 373, 6.2 and 4.4 million in the US. Taking away the seven largest states (California, Florida, Illinois, New York, Ohio, Pennsylvania and Texas), the US figures become 0.5, 9.9, 3.9 and 2.6, more similar to those of the EU.

The income concept we use is household net per capita income, adjusted for household size using the modified OECD equivalence scale and for purchasing power differences using the price level indexes of Eurostat. Although this measure of income is further away from the labor market than labor income and is thus subject to several smoothing mechanisms, this is the only one that, using different data sources/surveys, can be followed throughout all of the periods and in particular before and after China's accession to the WTO. Our sample units are individuals between the ages 25-60 to avoid

---

<sup>58</sup>Also some difficulties connected with German unification, apparent in the case of Germany are better handled in the GSOEP – see (Frick and Krell, 2010).

<sup>59</sup>Because of insufficient information, we only recode the appropriate regions, but do not treat boundary shifts in Greece between EL1 - EL5 and EL2 - EL6 respectively and in Italy between ITD - ITH and ITE - ITI. See the history of changes at <http://ec.europa.eu/eurostat/web/nuts/history>.

peculiarities regarding the beginning and the end of a labor market career.

The above household surveys are usually not used for computing regional statistics, but we argue that they are useful for this purpose. Household surveys seldom target to be representative at other than the national level, representativeness meaning that matching the sample averages of some key variables to an outside reference is an explicit target of the survey (often with some mechanism to ensure this). Nevertheless, the design of such surveys is complicated, often including multistage sampling and stratification of the sample, which is useful in the current case. Stratification in general and geographic stratification, in particular, is of help. It means that the national sample is, in fact, a collection of subsamples within geographic regions, which happen to be NUTS1 or an even lower level set of regions. We believe that if such sampling was used to collect the data and the resulting sample size is sufficiently large for calculating our indicator, we can trust that the regional inequality figures are useful. Even if not representative, they are ‘random enough’ within a region so that all we have to worry is the sampling and estimation precision.<sup>60</sup>

Because the exact sampling mechanism works differently in different countries and surveys, we have taken a look at the sampling plans of the surveys we use. Again, we have to keep in mind that Sweden, Ireland, Luxembourg and Denmark are by definition, while the Netherlands, Portugal and Finland are by necessity constitute a single NUTS1 region, therefore there is nothing to look at their case (there are more than enough observations, and they are by definition representative). We only have to worry about seven countries: Austria, Belgium, Greece, Spain, France, Italy and the UK before 2000. In the case of the ECHP, the description in section 4.3 and Table 7 on page 29 of European Commission (1996) makes it clear that sampling involved stratification at least at the NUTS1 level in all of the above cases, so we do not have to worry. In the case of the EU-SILC, one has to consider the Quality Reports country

---

<sup>60</sup>A commonly used rule of thumb for calculating averages is to have at least 30-50 observations, while less and less central distributional measures require more and more observations. In our case, we have a few regions with one hundred observations, but in the vast majority of cases we have thousands.

by country – we have chosen to look at the initial, 2006 ones.<sup>61</sup> The discussion above makes it clear that, except for Austria, sampling has involved stratification at the NUTS1 level or below, so again, we do not have to worry. The Austrian one is a simple random sample, which, in an extreme case can, but in a country with almost 9 million inhabitants is unlikely to produce a regionally highly unbalanced sample at the NUTS1 level. As a countermeasure, subsection 2.1.8.3 points out, that the ‘adjustments were carried out simultaneously on household and individual level and with reference to [...] region (nine categories: Nuts II level)’. Indeed: tabulating the population shares of the regions reveals smooth transitions in time with weighted data, but some nonlinearities (temporary drops/increases between 2004-2008). Still, the dispersion of weights across regions remains moderate. We have checked the GSOEP and the BHPS individually. In Germany, almost all subsamples are stratified by NUTS1 regions at least (that of the GDR at the NUTS3 level) – see section 2 ‘Sampling of SOEP Subsamples A to M3/4’ in Kroh *et al.* (2018). In the UK, stratification was based on postcodes (or if there were fewer than 500 delivery points/addresses, groups of them) – see subsection 3.2.1.1 in Lynn (2006). Because there is a 1:1 mapping between postcodes and NUTS3 regions in the UK (and in other EU countries too), this implies NUTS3 stratification.<sup>62</sup>

Because our database is a composite one, we inspect its time-series properties to see if there is reasonable continuity. Both income levels and the log-variance indicator of income inequality show that continuity is reasonable – see Figure A11 on the next page.

Broad stylized facts on the changes of inequality between the 1990s and the 2000s, before and after the crisis indicate that these periods are characteristically different. Between 1995 and 2000, inequality has not increased in most regions and has decreased in many regions in Spain, Greece and Germany, among other countries – see Figure A12. After 2000 however, inequality has increased in the majority of regions. Before the crisis, the increases affected a wide range of regions, but not in all of them. After

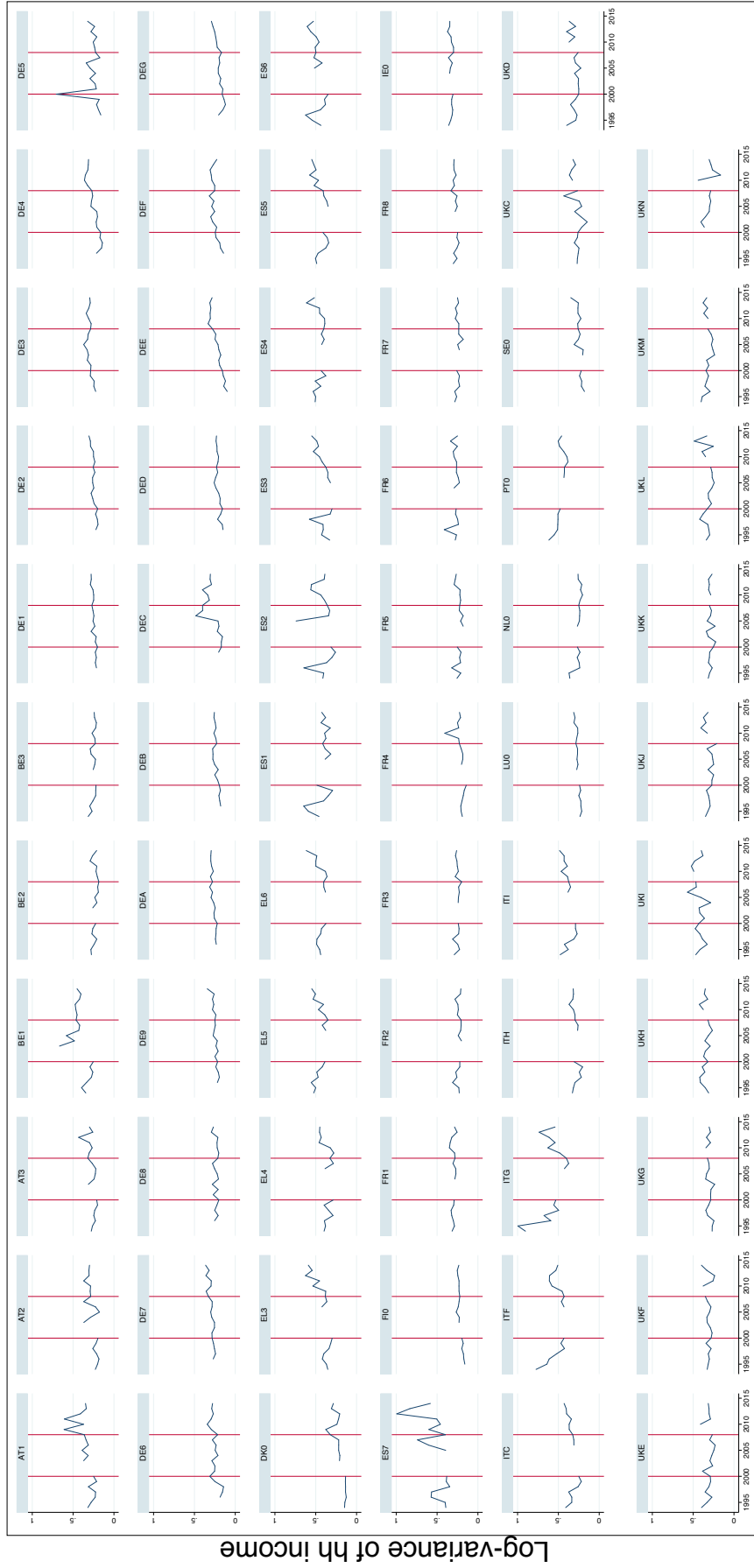
---

<sup>61</sup>Available at <https://circabc.europa.eu/> under Library / Quality assessment/quality reports.

<sup>62</sup>See <https://ec.europa.eu/eurostat/web/nuts/correspondence-tables/postcodes-and-nuts> on the underlying mapping.

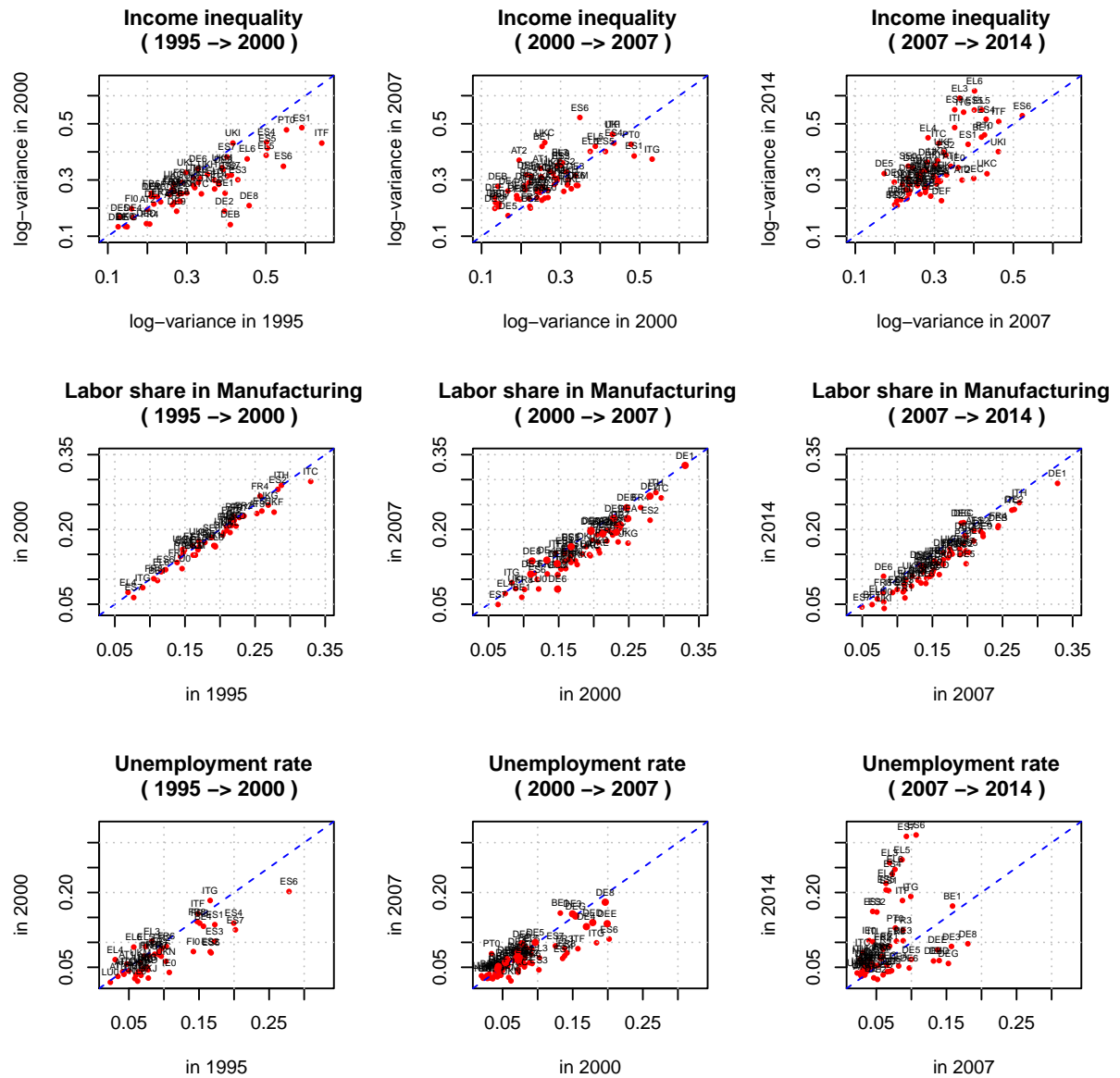


Figure A11: Time series of log-variance by NUTS1 regions



its outbreak, however, practically all of them were affected, and the increases were of greatest magnitude for those in the Mediterranean countries.

Figure A12: Changes in income inequality, Manufacturing labor share, and unemployment rate in NUTS1 regions



Source: own calculations from ECHP, EU-SILC, and EU-LFS microdata.

Note: due to the absence of previous years in the EU-LFS, the end- and starting year for Germany on Figures in column 1 and 2 is 2002 instead of 2000.

### E.3 Data on employment

Publicly available databases do not contain sufficiently detailed and long historical time series of employment levels and composition that we need, hence, we build these directly from microdata. We use the European Labour Force Survey (EU-LFS) to compile a panel data of regions (modified NUTS1, as defined and harmonized in case of data on inequality). Also here, we focus on employees aged 25 to 60 but do not restrict the sample in other ways. The only difference is that, for Germany, the regional identifier is available only since 2002 and we are forced to proxy the previous years with it<sup>63</sup>.

Similarly to the EU-SILC, we argue for the reliability of EU-LFS data at the NUTS1 level, but here the argument is even stronger. Firstly, ‘All countries but Denmark use the region, either at NUTS 2, NUTS 3 and NUTS 4 level or nationally defined areas, as stratification variable’ – see page 7 of the EU-LFS quality report Eurostat (2017). Secondly, the sample sizes are much larger than those in the EU-SILC, usually at the order of several thousand observations, except some regions in Germany. Thirdly, we are using the EU-LFS to calculate shares only, which does not have demanding data needs.

Based on the employment data, one can observe that the share of manufacturing employment in a region was quite stable in the second half of the 1990s, but had significantly dropped in the 2000s – see Figure A12 on the preceding page. Similarly to income inequality, this drop has deepened after the financial crisis, but from an already lower level and at a slower pace. Considering the whole observation period, the connection between inequality and the share of labor working in manufacturing/industry appears to be negative and rather strong – see the first and second panel of Figure A13. This seems to be connected with the fact that a broad Industry sector (B-E), which comprises Manufacturing<sup>64</sup>, has typically the smallest within-sector inequality in EU15 countries (see the rightmost panel of Figure A13 on the next page).

---

<sup>63</sup>This is rather unfortunate as we use values for the pre-WTO-accession year 2000, for a basis of comparison. We have tried to create the same data using the GSOEP, but because of its much smaller sample size, we decided not to pursue this method.

<sup>64</sup>We rely on this rather than on manufacturing only as only this coarser category is available in the ECHP and EU-SILC databases.

It should be pointed out that the transition to Rev. 2 of the NACE classification from 1 January 2008 onwards introduced a break in the time series of employment by sectors. As, apart from Sweden, the EU-LFS provides the decomposition for 2008 using both Rev.1 and Rev.2 of the classification, we use their ratio in 2008 to construct also a re-scaled series that intends to diminish the impact of the change. Nevertheless, this issue does not seem to have a substantial impact on our findings because the estimation results are similar using unscaled and rescaled data, using data only up until 2007, or allowing for the respective structural breaks.

Figure A13: Changes in regional income inequality and labor share in Industry (sectors B-E)



Source: own calculations from ECHP, EU-SILC, and EU-LFS microdata.

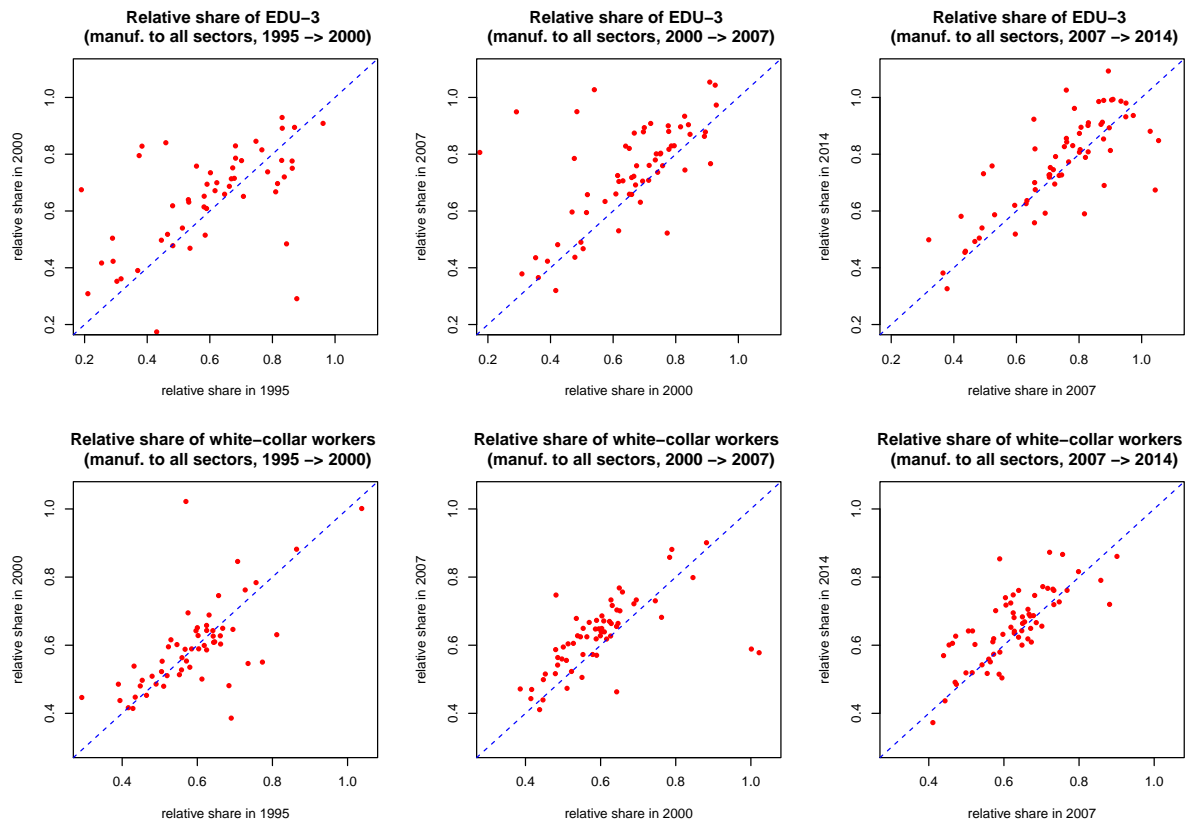
Finally, in Figure A12 on page 96 we take a look at the changes in unemployment rate. We again compare the three periods as before. While the unemployment rate was stable or decreasing in the first two periods, it has strongly increased in many regions during the last (post-crisis) period (see the bottom panel of Figure A12 on page 96).

### E.3.1 Skill upgrade in Manufacturing

A strive for higher productivity and/or a shift of production towards products of higher quality and/or technological intensity might require higher skills of workers. To explore

the potential presence of such behavior, we define two indicators using the share of workers in Manufacturing having tertiary education, relative to the share of all workers in a region having tertiary education,<sup>65</sup> and the share of white-collar employees in Manufacturing to the share of white-collar employees working in all sectors in the region. We again consider the relative shares to control for a potential general shift caused by other factors specific to the region (or even more globally<sup>66</sup>) and not only the manufacturing sector in that region.

Figure A14: Changes in the relative share of employees with tertiary education (top panel) and white-collar positions (bottom panel) in Manufacturing as compared to all employees in an EU15 region.



Note: EDU-3—tertiary education.

Source: own calculations based on the ELFS micro data.

The top panel of Figure A14 shows the relative share of employees with tertiary

<sup>65</sup>We also considered separation into lower-secondary and upper-secondary education, but these did not reveal a pattern that would be as clear as the one observed in case of tertiary education.

<sup>66</sup>For instance, the globally increasing share of labor with higher education.

education, while the bottom panel depicts the relative share of white-collar workers in the EU15 regions. Even after controlling for the general trend in a region, there is a clear shift towards workers with higher education and white-collar occupations, which is consistent with the general findings in Martinez Turegano (2020). Although the difference is less clear cut, a somewhat larger increase in these labor upskilling indicators seems to be present during 2000-2007.

## **GETTING IN TOUCH WITH THE EU**

### **In person**

All over the European Union there are hundreds of Europe Direct information centres. You can find the address of the centre nearest you at: [https://europa.eu/european-union/contact\\_en](https://europa.eu/european-union/contact_en)

### **On the phone or by email**

Europe Direct is a service that answers your questions about the European Union. You can contact this service:

- by freephone: 00 800 6 7 8 9 10 11 (certain operators may charge for these calls),
- at the following standard number: +32 22999696, or
- by electronic mail via: [https://europa.eu/european-union/contact\\_en](https://europa.eu/european-union/contact_en)

## **FINDING INFORMATION ABOUT THE EU**

### **Online**

Information about the European Union in all the official languages of the EU is available on the Europa website at: [https://europa.eu/european-union/index\\_en](https://europa.eu/european-union/index_en)

### **EU publications**

You can download or order free and priced EU publications from EU Bookshop at: <https://publications.europa.eu/en/publications>. Multiple copies of free publications may be obtained by contacting Europe Direct or your local information centre (see [https://europa.eu/european-union/contact\\_en](https://europa.eu/european-union/contact_en)).



## The European Commission's science and knowledge service

### Joint Research Centre

#### JRC Mission

As the science and knowledge service of the European Commission, the Joint Research Centre's mission is to support EU policies with independent evidence throughout the whole policy cycle.



**EU Science Hub**  
[ec.europa.eu/jrc](https://ec.europa.eu/jrc)



@EU\_ScienceHub



EU Science Hub - Joint Research Centre



EU Science, Research and Innovation



EU Science Hub



Publications Office  
of the European Union

doi:10.2760/752376

ISBN 978-92-76-18590-1