Trade Liberalization and Labor Market Institutions*

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Abstract

While the distributional consequences of trade liberalization at the firm level are now well understood, the previous literature has paid limited attention to how variations of domestic institutions across countries impact the *reallocation effect* suggested by the New New Trade Theory. Building on the varieties of capitalism literature, we advance the argument that the distributional effects of trade liberalization are systematically different in "liberal market economies" (LMEs) and "coordinated market economies" (CMEs). This is because CMEs feature the presence of coordinated wage-settlement institutions, which pose a ceiling to the increase of wages, helping smaller firms to weather the raising competition triggered by tariff reduction. We test this hypothesis using a firm-level dataset on EU countries, which includes more than 800,000 manufacturing firms between 2003 and 2016. We rely on a novel measure of preferential tariff reductions to capture the occurrence of trade liberalization. We find that, for productive firms, gains from trade are twice as large in LMEs as they are in CMEs. We complement our analysis by showing that there is a weaker demand for redistribution in CMEs compared to LMEs in case of preferential liberalization. The results of our paper inform a growing literature on the winners and losers from trade liberalization at the firm level, pointing out the importance of institutions in taming market forces.

Keywords: trade liberalization, trade agreements, labor institutions, heterogeneous firms, wage.

JEL Classification: F13, F14, F16.

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1 Introduction

The notion that open international markets make societies better off has become increasingly contested in recent years in many advanced economies. The rhetoric and trade policy initiatives of the Trump Administration are perhaps the most visible examples of this trend. The imposition of unilateral trade-restrictive measures against China and the European Union, the withdrawal from the Trans-Pacific Partnership (TPP), and the ongoing boycott of the World Trade Organization, amongst others, are all part of a broader strategy explicitly designed with a view to making US trade policy more responsive to the frustration of American citizens who could not see clear benefits from international trade agreements (United States Trade Representative 2017). But the result of the popular referendum in the UK advocating for a Brexit and the vocal opposition in many European Union member states against the Transatlantic Trade and Investment Partnership and the Canada-Europe Trade Agreement also tell of the mounting disquiet that the prospect of trade liberalization generates among the wider public in many advanced economies.

Central to these discussions about the merits of trade liberalization seems to be a widespread popular perception that globalization has generated greater wealth for a small group of individuals and firms while making the majority of citizens worse off. These perceptions are in line with the recent research in international trade that has found that, indeed, the benefits of trade liberalization are highly concentrated in the hands of few "superstar" exporting firms. Differently from earlier theories of trade policy predicting distributive consequences affecting either entire classes or specific industries (Hiscox 2001), these studies highlight that trade liberalization may generate stark distributive conflicts, pitting firms against each other. These works show that firm-level differences in size and productivity can account for their heterogeneity in export performance, which in turn explains why trade liberalization can generate uneven reallocation of profits across firms (Melitz 2003; Baccini et al. 2017; Osgood et al. 2017; Kim 2014; Bernard et al. 2007). Thus, both critics of globalization in the political arena and the recent literature converge in depicting trade liberalization as a policy choice that contributes to the concentration of wealth in the hands of the few, at the expense of the many (Osgood 2016).[1]

At the same time, there are significant differences in the extent to which protectionist sentiments rise and affect the political debate across countries. While in some countries popular concerns over the welfare effects of trade liberalization are widespread and have generated marked protectionist responses by elected representatives, in other cases opposition to trade liberalization has been much less intense. This observation underscores the importance of assessing whether, and eventually how, domestic institutional factors influence how the welfare effects of trade liberalization are distributed in different societies. The comparative political-economy literature has long noted that different domestic institutional setups can affect the distributive consequences and politics of trade in systematic ways (Katzenstein 1985; Rogowski 1987; Milner and Kubota 2005; Kono 2009). Yet, the literature on firm-heterogeneity and trade politics has so far only investigated patterns of firm income redistribution within single countries, largely focusing on the welfare effects of trade liberalization on US firms (Baccini et al. 2017; Bradford Jensen et al. 2017; Osgood 2016; Kim, 2017; Osgood et al. 2017; Kim et al. 2017). In short, we know virtually nothing about how domestic institutional factors influence

[1]For a survey of this literature, see Kim and Osgood (2019).

which firms win or lose from trade liberalization. This is an important oversight, since this type of analysis could help shed light on the observed variation in patterns of interest aggregation and political action across different countries over the merits of trade liberalization.

This paper contributes to the expanding literature on firm heterogeneity and trade politics by incorporating domestic institutional differences into the analysis of the inter-firm competitive dynamics generated by trade liberalization. To do so, we draw on insights from the Varieties of Capitalism (VoC) literature (Hall and Soskice 2001) and develop an argument that highlights how variations in labor market institutions affect the distribution of welfare effects of trade liberalization across firms. In particular, we focus on how the configuration of labor unions and employers interactions, encapsulated in different wage bargaining institutions, affects the relative cost of employing one factor of production, i.e., labor, which, in turn, influences the ways in which the gains and losses of trade liberalization are distributed across firms at different productivity levels.

The building block of our argument is the distinction between wage bargaining systems in *liberal market economies* (LMEs) and *coordinated market economies* (CMEs). In the former, the relationships between workers and employers rely heavily on competitive market forces. In the latter, the logic of market competition is tamed by the existence of wage bargaining institutions that facilitate and reinforce strategic coordination between trade unions and employers associations. These differences have important implications for how the gains and losses of trade liberalization are distributed across firms. By equalizing wages at equivalent skill levels across an industry, coordinated wage bargaining institutions contribute to taming the inter-firm competitive dynamics generated by trade liberalization. More specifically, we argue that trade liberalization generates a stronger reallocation effect from least to most productive firms in countries with liberal labor market institutions, and a weaker reallocation effect in countries with coordinated labor market institutions.

We test our argument using a firm-level dataset on European Union (EU) countries, which includes more than 800,000 manufacturing firms between 2003 and 2016. To capture the occurrence of trade liberalization, we rely on tariff cuts implemented by the EU with trade partners in all preferential trade agreements (PTAs) signed after 1995. Our identification strategy boils down to a triple difference-in-differences in which the distributional effect of firms productivity and tariff cuts varies across labor market institutions. We find that, for productive firms, gains from trade are twice as large in LMEs as they are in CMEs. For instance, our results indicate that the most productive British firms increase their revenue 20 percent more than the least productive British firms do as a consequence of tariff cuts. On the contrary, the most productive German firms increase their revenue 10 percent more than the least productive German firms do in the presence of preferential trade liberalization. Moreover, we show evidence of the mechanism highlighted in the theory: wages increase significantly more in LMEs than in CMEs as a result of trade liberalization.

We complement our analysis at the firm level with evidence at the individual level, using questions from the European Social Survey and a novel geographical measure of trade liberalization weighted on share of workers employed in very productive firms, which we geo-located at the level of EU regions. By exploiting the heterogeneous impact of trade liberalization across European regions, we show that there is a weaker demand for redistribution in CMEs compared to LMEs in case of preferential liberalization, given that gains from trade are more uniform in the presence of labor market frictions.

Importantly, this effect is driven by low-income individuals, who are the likely losers from trade openness in developed economies.

Our paper speaks to three main streams of research. First, several empirical papers have documented selection and market share reallocation effects as a result of trade liberalization (Pavcnik 2002; Trefler 2004; Bernard et al 2006; Amiti and Konings 2007; Topalova and Khandelwal 2011). Recent studies have pointed out that a few large productive firms enjoy the lion's share of the benefits from trade liberalization at the expense of smaller, less productive firms (Osgood et al. 2016; Baccini et al. 2017). Our paper shows that domestic institutions affect gains from trade and that labor market frictions make benefits from trade liberalization more uniform among firms.

Second, our paper speaks to a large literature on the effect of globalization on individuals' preferences over policy (Scheve and Slaughter 2004; Walter 2010; Hainmueller and Hiscox 2006; Mansfield and Mutz 2009; Margalit 2012; Walter 2017). In particular, recent studies find that trade shocks trigger attitude in favor of economic nationalism and, in turn, affect voting behavior (Margalit 2011; Autor et al. 2016; Jensen et al. 2017; Ballard-Rosa et al. 2017; Colantone and Stanig 2018a, 2018b). Our paper shows that, in the case of trade liberalization, individual preferences over redistribution is heterogeneous among types of labor market institutions. In other words, domestic institutions may mitigate the backlash against globalization experienced by developed democracies.

Third, the findings of our paper informs the debate concerning the effect on inequality of globalization in general (Ruggie 1982, Katzenstein 1985, Rodrik 1998, Rudra 2002), and trade liberalization in particular (Hanson and Harrison 1999, Goldberg and Pavcnik 2004, Jensen and Rosas 2007, Topalova and Khandelwal 2011, Dix-Carneiro 2014). Much of this literature provides empirical evidence either at the macro level or from single countries. We are the first study to show the micro-foundation of the distributional effect of trade liberalization for a large number of developed economies. In doing so, we are able to unveil the mechanisms linking trade openness to inequality at the level of both firm and individual.

2 Theory

Our theoretical framework is built on three steps: (1) gains from trade are heterogeneous among firms; (2) different labor market institutions generate different labor market frictions; (3) labor market frictions affect the distributional effects of trade liberalization among firms. Below we explain each step in detail.

2.1 Trade liberalization and superstar firms

Traditional theories of trade politics have long predicted trade liberalization will generate distributive consequences affecting either entire classes or specific industries. The Stolper Samuelson model emphasizes class-based cleavages by showing how the easing of trade increases returns to the factor a country has in abundance, while harming returns to the scarce factor (Rogowski 1989). The Ricardo-Viner model predicts sectoral or industry-based political cleavages by showing how, under conditions of imperfect factor mobility, trade liberalization alters goods prices shared among all firms in the same industry (Hiscox 2001). While generating starkly different predictions about the political cleavages engendered by trade liberalization, both models share the view that firms play no significant role in determining preferences and patterns of political action over trade policy (Osgood et al. 2016).

In recent years, the role of firms has gained center stage in the study of trade politics (Madeira 2016). Trade politics theory's shift towards a better appreciation of the role of firms has been driven by the growing importance of intra-industry trade (IIT), i.e., trade of different varieties of the same product between countries with similar factor endowments (Gruber and Lloyd 1971; Brulhart 2009), and by the availability of new firm-level data on export engagement and sales showing that even in the most export-oriented sectors only a minority of large firms manages to enter and remain in exports markets (Bernard et al. 2007). Against this background, a new wave of studies has highlighted how trade liberalization can engender stark distributive conflicts between firms within an industry (see Bernard et al. 2012 for an overview). Building on the seminal work of Melitz (2003), these contributions show how firm-level differences in size and productivity can account for their heterogeneity in export performance, which in turn explains why trade liberalization can generate uneven reallocation of profits across firms (e.g., Osgood et al. 2017; Kim 2017; Bernard et al. 2007).

Exporters face trade costs, which include fixed costs of distribution and servicing, and variable costs such as transport, insurance, fees, and tariffs. Firm productivity plays a crucial role in selecting the firms that are able to access export markets: only the most productive firms can bear trade costs while remaining profitable. Trade liberalization, in turn, reduces the variable costs of trade, i.e., tariffs. This lowers the productivity threshold that firms must meet to access foreign markets, hence motivating new firms to engage in trade. At this stage two mechanisms allow more productive firms to gain disproportionally from trade liberalization. On the one hand, increased competition in international markets leads to a reduction of prices, which, in turn, lowers firms' profits. On the other hand, expanding foreign sales generate an increase in the demand for labor, which causes real wages to rise (Egger and Kreickemeier 2007; Helpman et al. 2010). Only large and productive firms, which in equilibrium have lower prices, can cope with the twofold need of charging lower prices in the face of increasing competition in foreign markets, and absorbing the higher wages generated by the growing demand for labor that expanding foreign sales stimulate. In contrast, low-productivity firms decrease their sales or exit the market altogether. These responses lead to a reallocation of resources toward high-productivity firms at the expense of low-productivity ones (Baccini et al. 2017; Bernard et al. 2012; Osgood et al. 2017).

2.2 Labor market institutions and wage bargaining systems

The comparative political economy literature has long noted that domestic labor market institutions have a systematic impact on the evolution of the countries' comparative advantages, on their trade profiles, and on the distributive consequences of trade liberalization (Hall and Soskice 2001a; Hancké et al. 2007; Manger and Sattler 2014). We draw on this literature to develop an argument that highlights how diverging labor market institutions affect the distributional impact of trade liberalization at the firm level. In particular, we focus on how the configuration of labor union and employers interactions, encapsulated in different wage bargaining institutions, affects the relative cost of employing one factor of production, i.e., labor, which, in turn, influences the ways in which the gains and losses of trade liberalization are distributed across different firms at different productivity levels.

The building block of our argument is the well-known distinction between wage bargaining systems in liberal and coordinated market economies. The VoC approach (Hall and Soskice 2001b) is based on the notion that the presence or the absence of mechanisms of strategic coordination between firms and employees is key to understanding how the political economies of advanced capitalist countries differ. This body of work relies on the conceptual distinction between LMEs, in which "firms coordinate their activities primarily via hierarchies and competitive market arrangements"; and CMEs, in which firms "depend more heavily on non-market relationships to coordinate their endeavors with other actors and to construct their core competencies" (Hall and Soskice 2001b: 8).[2] LMEs and CMEs are characterized by different institutional complementarities, which include several dimensions such as the relationships between firms and employees, the coordination of both firms and employees among themselves (in the form of employers associations and trade unions respectively), the issue of skill formation and training (Estevez-Abe et al. 2001; Culpepper and Thelen 2008), and corporate governance (Vitols 2001; Brsch 2007).

Differences in wage bargaining systems are key to this conceptual distinction. In LMEs, the relationships between individual workers and employers rely heavily on competitive market forces, which make it relatively easy for firms to hire labor to take advantage of new opportunities, as well as to fire workers in the face of economic downturns. Trade unions are less cohesive and encompassing, which makes economy-wide coordination difficult to achieve. In this context, the poaching of skilled workforce by competing firms is common practice, which leads wages to vary widely among different firms within a sector, depending on individual firms productivity. As a result, wages are largely determined by macro-economic policy and market competition (Hall and Franzese 1998).

In CMEs, the logic of market competition is still present but tamed by the existence of (formal and informal) wage bargaining institutions that facilitate and reinforce strategic coordination. Trade unions and employers' associations have the strength and the ability to negotiate deals that grant moderate (and in any case predictable) wage increases for firms in exchange for a commitment to maintain employment levels (and often investment in training) for the workforce, which "makes it possible for firms to retain a skilled workforce through economic downturns and to invest in projects generating returns only in the long run" (Hall and Soskice 2001b: 22). Thus, wages are set through industry-level bargains between encompassing trade unions and employer associations. This kind of strategic coordination makes it difficult for firms to poach workers, ensuring that wage increases develop in line with productivity gains, and thus ultimately limiting the inflationary effects of wage settlements (Hall and Franzese 1998; Hall and Soskice 2001b).

A brief description of the German and the British wage settlement systems helps us illustrate these differences. The German wage bargaining system is highly coordinated; negotiations are typically conducted at the regional level for each industry, involving a coalition between large employers and trade unions. What is termed a "pilot agreement" is usually reached for a particular bargaining district and then is transferred to other regions of the same industry across Germany (Addison et al. 2007). This means a range of companies are affected across the country when a wage agreement is settled in a particular sector (Hassel and Rehder 2001). Moreover, there is also coordination between unions

[2] Although VoC scholars have also identified other varieties of capitalism, conceptualized either as intermediate between the two main types (Molina and Rhodes 2007), or as clearly distinct from them (Thatcher 2007; Nlke and Vliegenthart 2009), the distinction between LMEs and CMEs seems to be empirically the most solid (Hall and Gingerich 2009).

and employers of different industries, which creates a highly uniform collective wage bargaining policy across different sectors of the economy. The system subsequently yields incremental changes in wages, which does not create incentives for employers to poach workers from other companies, or for workers to switch firms or even sectors. For instance, the German union of steelworkers *IG Metall* follows a pattern of collective wage bargaining with the employer association *Gesamtmetall* approximately every two years. In the latest round of wage bargaining in 2016, the union negotiated a wage increase for almost 4 million employees across the country (IG Metall 2016), following a strikingly similar increase of around 0 3% per year that it had negotiated for the last 20 years, with exceptionally little difference to firms profits or economic crises over that time period.

In comparison to Germany, the wage bargaining system in the United Kingdom (UK) is highly decentralized. Instead of multi-employer and sectoral lines, bargaining in the UK typically occurs at the local level, which is usually a particular plant or a company, between an employer and a group of workers. There are no national laws that govern collective bargaining rights, and freedom of contract at the plant or firm level prevails in industrial relations, especially after the extensive reforms imposed during late 1980s (Forth et al. 2013). A quick look at the website of the Wage Indicator Foundation and the Trades Union Congress (TUC), a British umbrella organization that represents different unions, reveals the existence of a list of *local* collective agreements filed, for instance, across several sectors in Britain through various competing trade unions.[3] Subsequently, the incentive for changing company or sector is higher for workers, and firms themselves have more tools available to increase or decrease labor costs.

We can illustrate this system by considering, again, the steel manufacturing sector. Unlike its German counterpart, the largest union of steelworkers in the UK, the Iron and Steel Trades Confederation (ISTC) attempts to bargain new wages for its workers by plant, rather than at the industry level. Over the last two years, the ISTC negotiated with several different companies, such as British Steel Corp. and Tata Steel, and agreed on different contracts with each of the negotiating parties.[4]

Figures 1 is illustrative of this trend. The graph shows the growth rate of household income in Germany and the UK between 1995 and 2016. It indicates that while the average income growth of German individuals fluctuated between -0.46% and 2.53%, for British individuals this figure was between -1.73% and 5.66%. We can observe that the earnings of German individuals have been much more incremental (and stable) in comparison to those of British individuals, whose incomes fluctuate more considerably and asymmetrically between different economic cycles.

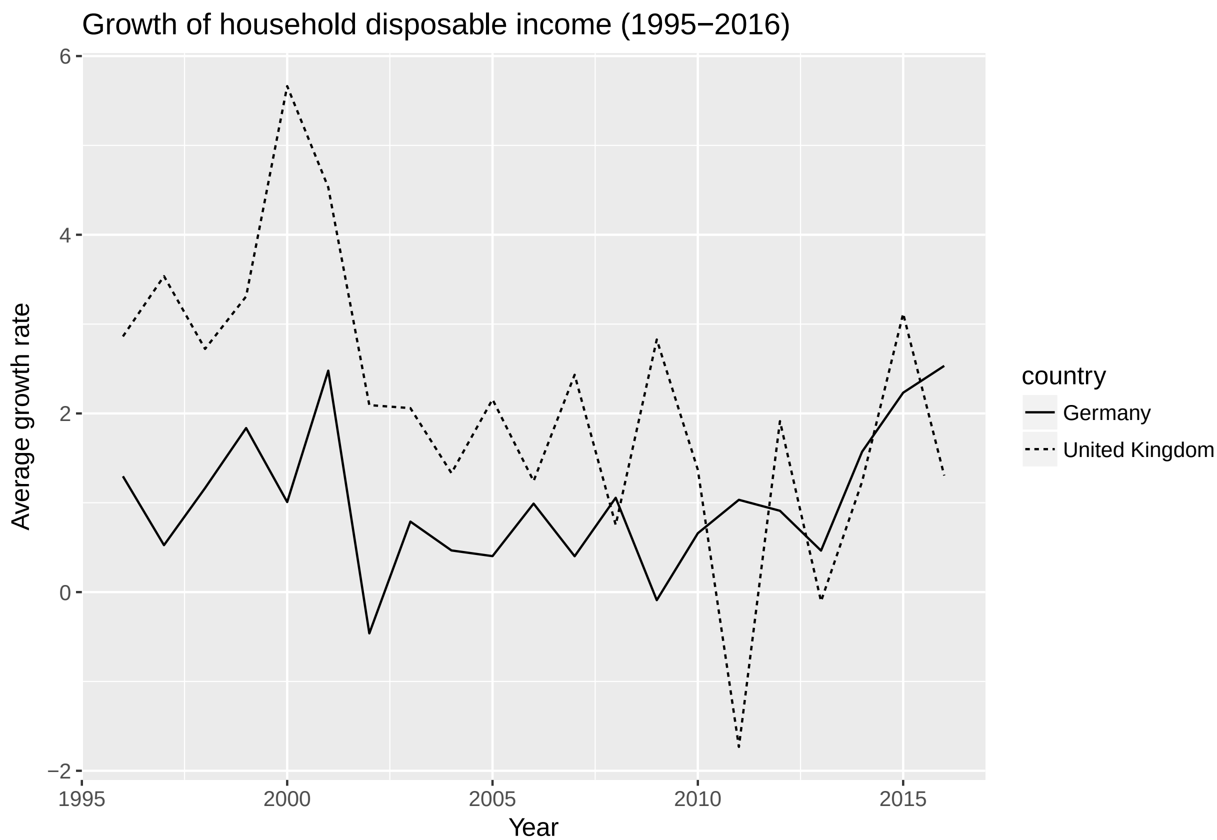
2.3 Gains from trade and the labor market

Differences in wage bargaining systems have important effects on how the gains and losses of trade liberalization are distributed across firms. In both liberal and coordinated market economies, a lowering of barriers generates a twofold effect: the most competitive firms will manage to secure their sales in the domestic market and increase their exports, while the least competitive ones will not be able

[3] See: Trade Union Congress (TUC) Collective Bargaining Agreements, available at: `https://www.tuc.org.uk/workplace-guidance/collective-bargaining` and Wage Indicator Foundation, Collective Agreement Database, available at: `https://wageindicator.co.uk/advice/collective-agreements-database`.

[4] See Iron and Steel Trades Confederation agreements at: `https://community-tu.org/tata-steel-workers-endorse-union-agreement/`.

Figure 1: Household income: Germany versus the UK



to export and will experience a contraction of their sales in the domestic market (due to increased competition from foreign firms). High-productivity firms will therefore increase their demand for labor and thus create an upward pressure on salaries which is the point where we expect LMEs and CMEs to diverge.

In LMEs, the flexibility of wage bargaining institutions allows the most productive firms to freely increase wages to attract new workers. Because of their high productivity, these firms will be able to recoup these increased labor costs and remain competitive in foreign markets. On the contrary, the least productive ones will not be able to absorb wage increases and remain profitable at the same time. Large and highly productive firms are therefore likely to increase their export sales, while small- and low-productivity firms will exit export markets and decrease domestic sales. In line with models of firm heterogeneity, trade liberalization can be expected to reallocate resources towards high-productivity firms and, as a result, to increase the industrys overall productivity.

In CMEs, coordinated wage bargaining systems will mitigate the selective and redistributive effects of trade liberalization. The most important difference here is that firms are constrained in their ability to poach workers from competing firms by the existence of coordinated wage bargaining institutions that set a ceiling on wage increases for all firms operating in a particular sector. This type of wage bargaining institution thus breaks one of the key channels through which trade liberalization generates uneven distributional consequences across firms. By equalizing wages at equivalent skill levels across an industry, these institutions reduce the ability of the most competitive firms to offer higher wages.

Rising labor costs effectively increase the zero-profit productivity cutoff, which indicates which firms will find it profitable to remain in the market. Formal and informal institutions that put a ceiling on wages thus stabilize the zero-profit productivity cutoff, softening the negative impact of trade liberalization for the firms that are at the bottom of the productivity scale. As a result of reduced pressure on the least productive firms, the reallocation effect generated by trade liberalization should be lower in CMEs than in LMEs.

The causal mechanisms identified so far leads us to formulate a specific hypothesis regarding the reallocation effect of trade liberalization and the growth of sector productivity. While we expect the reallocation effect of exposure to international trade to be present in all countries, the labor market institutions typical of CMEs should reduce their impact compared to LMEs. Therefore, our hypothesis is: *In the aftermath of trade liberalization, reallocation of revenues from the least to the most productive firms will be higher in liberal market economies than in coordinated market economies.*

We conclude by noting that the effect of labor market frictions on firms' performance is not unequivocal. There is indeed the other side of the coin of our argument on labor costs. A possible way through which low-productivity firms can hope to remain in the market is by reducing labor costs (e.g., lowering the minimum wage), as doing so might help them continue to serve the domestic market. Rigid labor market institutions that constrain the downward movements of wages may make it more difficult for low-productivity firms to navigate increasing competition, leading to the opposite effect postulated by our theory.

3 Data

We test our argument using a reduced form approach. Below we describe our sample and main variables.

Sample We test the empirical implications of our argument on a large number of firms from EU countries during the period between 2003 and 2016.[5] Firm-level data come from the Amadeus database provided commercially by the Bureau Van Dijk. The data gathering process has been performed following best practices in terms of downloading methodology and cleaning procedures (Kalemli-Ozcan et al 2015). Our baseline sample includes more than 800,000 firms operating in manufacturing in the (up to) 28 EU countries. To analyze the distributional consequences of trade liberalization, our unit of observation is the firm-industry-country-year.

The key advantage of using the Amadeus database is that includes firms from all the EU countries. This provides us with the variation in labor institutions that we need to test our argument. Moreover, the dataset includes a large number of firm-level characteristics, which is crucial to build measures of productivity and other important controls. Furthermore, the database includes a large number of firms of different sizes and productivity, which operate in a large number of industries at the NAICS 4-digit level. This heterogeneity allows us to exploit variation across firms and across tariff cuts.

The Amadeus database also has some shortcomings. First, while the sample of firms is generally proportional to the size of the economy, East European countries are more heavily represented (see

[5]Amadeus has firm-level data even prior to 2003, but only a handful of firms and a few countries are covered. The number of firms covered by Amadeus increases significantly over time.

Figure 5 in Appendix E). Second, the Amadeus database does not systematically collect longitudinal firm-level data. In particular, because the sample does not include the universe of firms, a firm f may be present in 2006 but not in 2007, either because it exited the market or because it was not surveyed. Hence, our data are repeated cross-sectionally.[6]

Dependent variable Our dependent variable is the logarithm of revenue of firm f in industry i in country c in year t. This variable is our proxy of gains from trade, which allows us to quantify the distributional consequences of trade liberalization. According to Melitz (2003), revenue increases proportionally to firm productivity in case of tariff reduction (the so-called reallocation effect). There are other proxies capturing the distributional consequences of trade liberalization. An obvious candidate would be firm exit, which captures the selection effect; i.e., productive firms should exit less frequently than unproductive firms in case of trade liberalization. However, our repeated cross-sectional data are not suitable to measure firm exit. Another option would be to rely on profit rather than revenue. We opt for revenue because its coverage is substantively better in our data.

Independent variables To test our argument, we rely on three main independent variables. To begin with, we need a measure of firm productivity. Measuring productivity is particularly challenging and the debate on the best way to do so is still unsettled in economics. Our approach is to rely on different measures of productivity and to show that our results remain largely unchanged. We use a standard measure of total factor productivity (TFPR), using the Solow residuals in the main analysis.[7] As we show below, results are robust to alternative measures of productivity. In particular, we rely on measures of productivity following Olley and Pakess (1996), Levinsohn and Petrins (2003), and Wooldridge's (2009) adaptations. All these measures can be easily estimated using firm-level characteristics available in the Amadeus database, e.g., working capital.[8] We label this variable *TFPR*.

The second key independent variable is a measure of trade liberalization. We rely on an original dataset containing preferential tariff concessions made by the EU in all PTAs signed between 1995 and 2014. For all PTAs, we extracted tariff schedules, each containing around 5,000 tariff lines at a highly disaggregated level. All PTAs contain at least two tariff schedules, one for the EU with its trade partner, and one for its trade partner with the EU. Our data are at the Harmonized Commodity Description and Coding System (HS) 6-digit level.

The data were compiled from two sources. We took tariff data for the year prior to entry into force of the PTA from the World Integrated Trade Solution (WITS) dataset, which relies on data reported by customs administrations. We then added information on tariff concessions from the officially negotiated tariff schedules listed in the appendices of the PTAs. The coverage of our tariff data is

[6] Financial data for companies within Amadeus are retained for a rolling period of eight years. When a new year of data is added, the oldest year is dropped, meaning only the most recent data for each company are available.

[7] TFPR is calculated using simple firm-level Solow residuals. We calculate TFPR for each firm-year by regressing the firm-level log of revenue on firm-level physical assets, employment, year, 4-digit industry, and country fixed effects. The residuals of this regression, which might also be negative, are our time-varying measures of firm productivity. We rescaled this variable so that it has only positive values.

[8] These measures of productivity are estimated using the Stata 15 command "prodest'.' The number of observations is significantly lower when we use these alternative measures due to missing data. That is the reason why we do not use them in the main analysis.

significantly better than the data from WITS as documented in Baccini et al. (2018). Moreover, our tariffs are *de jure* and not applied. Paired with the fact that all EU countries face the same tariffs, this should mitigate concerns about endogeneity. Think of *de jure* tariffs as instruments for applied tariffs in reduced form. To merge the tariff data with the Amadeus database, we use available crosswalks from HS 6-digit to NAICS 4-digit.

Our data include preferential tariffs (PRF) from the entry into force of a PTA to the end of the implementation period, since not all tariffs go to zero in the year of ratification. In other words, we are able to capture the phasing-out period for each product at the 6-digit level. Importantly, we collected data for the average tariff (MFN) existing before entry into force of the agreement for each PTA. That allows us to capture the (import) tariff cut (i.e., MFN-PRF) implemented by the EU in each 6-digit product in each year. In the main analysis, we rely on trade-weighted tariffs in industry i, i.e., tariffs divided by value of imports in industry i.[9] Details on how we build our measure of tariff reduction are explained in the Appendix A. We label this variable $\Delta\tau$.[10]

The third variable is a measure of the degree to which wages are coordinated in an economy. This variable, which we label *CME*, comes from the ICTWSS database[11] (Visser 2016) and is based on the index developed by Kenworthy (2001). This proxy measures "the degree of intentional harmonization observed in the wage-setting process" (Kenworthy 2001: 76), that is, the extent to which the wage settings determined by major players (peak-level union and employer confederations, unions, and employer associations of influential sectors, like metal manufacturing) are followed by the rest of the economic actors. The variable is ordinal, ranging from one ("Fragmented wage bargaining, confined largely to individual firms or plants" [Visser 2016]) to five ("Maximum or minimum wage rates/increases based on centralized bargaining" [Visser 2016]), and it captures the level of actual competition between firms on salaries. In countries scoring one (e.g., the UK), each firm can freely increase salaries to attract workers. The more a country has a wage-setting dynamic that limits (formally or informally) this tendency, the more we expect the relocation effect to be constrained. For instance, Germany scores four in this variable, implying that "wage norms are based on centralized bargaining by peak associations with or without government involvement" (Visser 2016).

The key independent variable is the triple interaction between firm productivity, tariff cuts, and coordinated wages. As is customary, we also include the double interaction terms and each variable alone in our model specification, unless these terms are absorbed by the fixed effects.

4 Empirical Strategy

Our approach to identification is a triple difference-in-differences.[12] We compare the evolution of (the log of) revenue across industries and firms according to the degree of trade liberalization and firm productivity in countries with different labor market institutions. Firm productivity varies across

[9] Weighted-trade tariffs correct for the bias created by simple tariffs, which can either underestimate or overestimate the level of protection in an industry. Results are similar if we rely on simple tariffs (available upon request). Similarly, results do not change if we use cumulative and proportional tariff cuts (see below for details).

[10] Figures 6 and 7 show the distribution of tariff cuts by industry and over time.

[11] Database on Institutional Characteristics of Trade Unions, Wage Setting, State Intervention and Social Pacts in 51 countries between 1960 and 2014, available at: `http://uva-aias.net/en/ictwss`.

[12] On the advantages of triple difference-in-differences for identification purposes, see Berck and Villas-Boas (2016).

firm, but does not vary over time. In other words, firms enter into the dataset with a given level of productivity, which is assumed exogenous and remains constant.[13] Tariff cuts vary across industries and over time, but not across countries as all the EU countries face the same preferential tariffs.[14] Finally, we let the variable *CME* vary across countries and over time, since we do not have reason to believe that EU PTAs affect labor market institutions.[15]

More formally, we estimate the following baseline model:

$$Revenue_{fict} = \beta_0 + \beta_1 TFPR_{fic} + \beta_2 \Delta\tau_{it} + \beta_3 CME_{ct} + \beta_4 TFPR_{fic} \times \Delta\tau_{it} + \beta_5 TFPR_{fic} \times CME_{ct} + \beta_6 \Delta\tau_{it} \times CME_{ct} + \beta_7 TFPR_{fic} \times \Delta\tau_{it} \times CME_{ct} + \beta_8 X_{fict} + \beta_9 W_{ict} + \delta_t + \delta_i + \delta_c + \epsilon_{fict}, \tag{1}$$

where revenue is the dependent variable, *TFPR*, $\Delta\tau$, *CME* , and their interactions are the main independent variables. β_0, β_1, ... β_9 are the coefficient. In particular, the key coefficient of interest is β_7, which we expect to be negative. δ_t, δ_i, and δ_c are year, industry, and country fixed effects respectively. Year fixed effects capture and control for overall trends in firms revenue. Industry and country fixed effects net out time-invariant differences across industries and countries.[16] ϵ_{fict} accounts for all residual determinants of the dependent variable.

Moreover, the matrix X includes standard firm-level controls. More specifically, we control for firm size as well as firm age and its squared value. Firm size is measured using the (log of) number of employees, whereas firm age is captured by the number of years in which a firm has operated in the market. The matrix W includes industry-level controls. In particular, we control for Most-Favored-Nation (MFN) tariffs, the (log of) labor-capital ratio, and market concentration, which we measure with the Herfindahl-Hirschman index of revenue.[17]

We run OLS regressions with standard errors clustered by firm, though we show below that our results are not sensitive to a specific clustering choice. Because our dataset includes more than 800,000 private and public firms for a period of over 10 years, we have more than 4 million observations in our baseline models. Note that the Amadeus database reports only the main industry in which firms operate, i.e., each firm compares only once in each year.

There are several concerns to our identification strategy. First, it may be that time-varying industry and country-level unobservables pose a threat to identifying the effect of our main independent variables and their interactions. Therefore, we include industry- (2-digit) year and country-year fixed effects in some models.[18] In doing so, we are unable to estimate the coefficient of *CME* due to perfect

[13]Our results hold if we allow productivity to vary over time (see Table 6 in Appendix E). These models require accepting a further identification assumption: firm productivity does not change differentially in countries with different labor market institutions as result of trade liberalization, a point that we address below.

[14]The interaction between firm productivity and tariff cuts is a type of Bartik instrument (Bartik 1991; Blanchard and Katz 1992). For a widely cited application of Bartik instruments, see Autor et al. (2013).

[15]To the best of our knowledge, there is no provisions that force EU countries to change their labor market regulations in PTAs.

[16]The variable *CME* varies enough over time that we are able to estimate both its coefficient and the country fixed effects.

[17]Our results remain unchanged if we interact MFN tariffs with *TFPR* and *CME* and include this triple interaction, which is never significant, on the right- hand side of our model.

[18]Because Amadeus reports only the main industry in which firms operate and because we use a time-invariant measure of productivity, results do not hold if we include industry- (4-digit) year fixed effects.

collinearity. Second, industries implementing trade liberalization may have been on a different trend compared to industries facing no tariff cuts. Support for the parallel trend assumption comes from the fact that our results are robust to the inclusion of industry-country specific time trends.[19]

Finally, and perhaps more importantly, the main threat to the identification strategy comes from variables that are correlated with *CME*. Indeed, it may be the case that there are country-level characteristics that are responsible for the differential effect of firm's productivity and tariff cuts on revenue. Take the presence of the Euro, for instance. If countries adopting the common currency are correlated with (say) CMEs, the monetary mechanism could be mediating the effect of firm productivity and tariff cuts on firm performance. To address this concern we identify a large number of country-level variables, we interact them with *TFPR* and $\Delta\tau$, and then we include them together with our key triple interaction terms in our models. The logic is that should our results remain unchanged, we can safely rule out that these confounders invalidate our identification strategy.

We identify the following confounders: corruption, level of unemployment, and access to credit. Moreover, we include other variables that capture the market structure and could also act as confounders: social welfare expenditure, size of the service sector, fiscal capacity, FDI outflows (and inflows), and the presence of the Euro.[20] In theory, each of these variables could mediate the effect of trade liberalization and firm productivity on firm performance (see Appendix B for more details). Before interacting these variables with *TFPR* and $\Delta\tau$, we begin with noting that their correlation with *CME* is generally quite low (see Table 7 in Appendix E). Then, we include each of these (potential) confounders on the right-hand side of our main model and, as a very conservative test, all triple interaction terms at the same time.[21]

5 Results

Main findings We begin with showing the results of a simple model including only the interaction term between *TFPR* and $\Delta\tau$ (Model 1). The coefficient of the interaction term is positive and significant, implying that revenue of productive firms increases more than revenue of unproductive firms in the case of tariff reduction.[22] This finding captures the reallocation effect predicted by the Melitz model (2003). Note also that the coefficient of *TFPR* is positive as expected (i.e., the more productive a firm, the higher the revenue even without trade liberalization), whereas the coefficient of $\Delta\tau$ is negative (i.e., unproductive firms face a reduction of revenue). This result identifies clear winners and losers from preferential trade liberalization and is in line with Baccini et al. (2017).

We now move to our main analysis. Models 2-6 include the triple interaction term among *TFPR*, $\Delta\tau$, and *CME*, whose coefficient is always negative and significant as expected. This implies that the reallocation effect is weaker in CMEs compared to LMEs. Putting it differently, productive firms increase their revenue less in CMEs than in LMEs as a result of preferential liberalization. These

[19]Results are robust to the inclusion of quadratic industry-country specific time trends.

[20]Description and sources of these variables are reported in Appendix B. Moreover, we include other potential confounders such as presence of state-owned companies in an economy and other-than-tariff barriers to trade and investment (not reported). The correlation between *CME* and each of these variables is quite low and our results remain unchanged (available upon request).

[21]Descriptive statistics are reported in Table 8 in Appendix E.

[22]Figure 8 in Appendix E reports the effect of the interaction term graphically.

Table 1: Main analysis

	(1)	(2)	(3)	(4)	(5)	(6)
	OLS					
	ln Revenue					
TFPR	0.33**	0.30**	0.30**	0.30**	0.31**	0.30**
	(0.002)	(0.006)	(0.006)	(0.006)	(0.007)	(0.006)
Δτ	-15.13**	-16.94**	-17.01**	-17.53**	-16.45**	-16.95**
	(0.683)	(1.961)	(1.969)	(1.983)	(1.960)	(1.961)
CME		-0.83**			-0.85**	-0.84**
		(0.087)			(0.086)	(0.087)
TFPR*Δτ	0.40**	0.45**	0.45**	0.46**	0.44**	0.45**
	(0.018)	(0.052)	(0.053)	(0.053)	(0.052)	(0.052)
TFPR*CME		0.02***	0.02***	0.03**	0.02**	0.02***
		(0.002)	(0.002)	(0.003)	(0.002)	(0.002)
Δτ*CME		1.85**	1.83**	1.84**	1.67**	1.85**
		(0.648)	(0.650)	(0.653)	(0.648)	(0.648)
TFPR*Δτ*CME		**-0.05****	**-0.05****	**-0.05****	**-0.05****	**-0.05****
		(0.017)	**(0.017)**	**(0.017)**	**(0.017)**	**(0.017)**
Constant	-8.31**	4.73**	5.33**	10.96**	6.47**	-80.05
	(0.069)	(0.032)	(0.201)	(1.139)	(0.239)	(154.60)
Observations	5,135,314	4,053,929	4,053,929	4,053,929	4,053,929	4,053,929
R-squared	0.754	0.765	0.766	0.763	0.768	0.792
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	No	No	Yes	Yes
Country FE	Yes	Yes	No	No	Yes	Yes
Industry FE	Yes	Yes	Yes	No	Yes	Yes
CountryYear FE	No	No	Yes	Yes	No	No
IndustryYear FE	No	No	No	Yes	No	No
CountryIndustry FE	No	No	No	No	Yes	No
Trends	No	No	No	No	No	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by firms in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. The outcome variable in all models is the log of revenue. Sources: Amadeus dataset, Baccini et al. (2018), and Visser (2016).

results remain virtually the same when we include country-year fixed effects (Model 3), country-year and industry-year fixed effects (Model 4), country-industry fixed effects (Model 5), and country-industry specific trends (Model 6). Importantly, the coefficient of the interaction term between $TFPR$ and $\Delta\tau$ remains positive and significant, adding plausibility to the results.

To help in grasping our results, we graph the effect of the triple interaction term in Figure 2. In particular, we report the marginal effect of tariff cuts for different levels of firm productivity across the five types of labor market institutions in EU countries. Figure 2 shows that the marginal effect is significantly more elastic for LMEs than for CMEs. In other words, the increase of revenue is significantly larger for productive firms in LMEs than for productive firms in CMEs as a result of preferential trade liberalization. Take Germany (CME=4) and the UK (CME=1), for instance: the elasticity of the marginal effect is twice as large for the UK as it is for Germany. This implies that the most productive British firms increase their revenue 20 percent more than the least productive British firms do as tariff cuts kick in. On the contrary, the most productive German firms increase their revenue 10 percent more than the least productive German firms do in the presence of preferential trade liberalization.

While we have provided evidence that labor market institutions mediate the distributional consequences of trade liberalization, there may be the concern that other country-level characteristics are responsible for this mediating effect rather than wage coordination. Given the nature of the triple difference-in-differences, these confounders are a threat to our identification strategy if and only if they impact firm performance differentially in industries facing tariff cuts *and* based on firm productivity. To sharpen our identification strategy, we include the triple interaction term of $TFPR$, $\Delta\tau$, and a large number of country-level (potential) confounders together with our key triple interaction term, i.e., $TFPR \times \Delta\tau \times CME$.

Table 2 shows the results of these tests. Model 1 includes the triple interaction term with the level of corruption.[23] Model 2 includes the triple interaction term with the level of unemployment. Model 3 includes triple interaction terms with country-level variables capturing the market structure: social welfare expenditure, size of the service sector, fiscal capacity, FDI outflows (and inflows), and the presence of the Euro. Model 4 includes interaction terms with variables capturing how easy access to credit is for firms. Model 5 includes all of the above triple interaction terms. Our main coefficient of interest remains negative and statistically significant. Importantly, its magnitude does not shrink and, if anything, increases in some model specifications. In sum, there is no evidence of confounding factors driving our results.

Mechanisms After having shown the effect of labor market institutions on the distributional consequences of preferential trade liberalization, we explore the mechanisms at play. A crucial channel through which the reallocation effect is weaker in CME is by setting a ceiling to the increase of wages. We explore this mechanism with two different tests. First, we use real wage data, collected by the ILO, for all EU countries between 1995 and 2008.[24] Importantly for us, the ILO wage data varies

[23] Results are similar if we use other measures of quality of institutions such as rule of law, government effectiveness, and regulatory quality. These variables are highly correlated with one another and that is why we do not include all of them at the same time.

[24] The ILO wage data covers the period between 1983 and 2008. However, we do not have data on preferential tariffs prior to 1995.

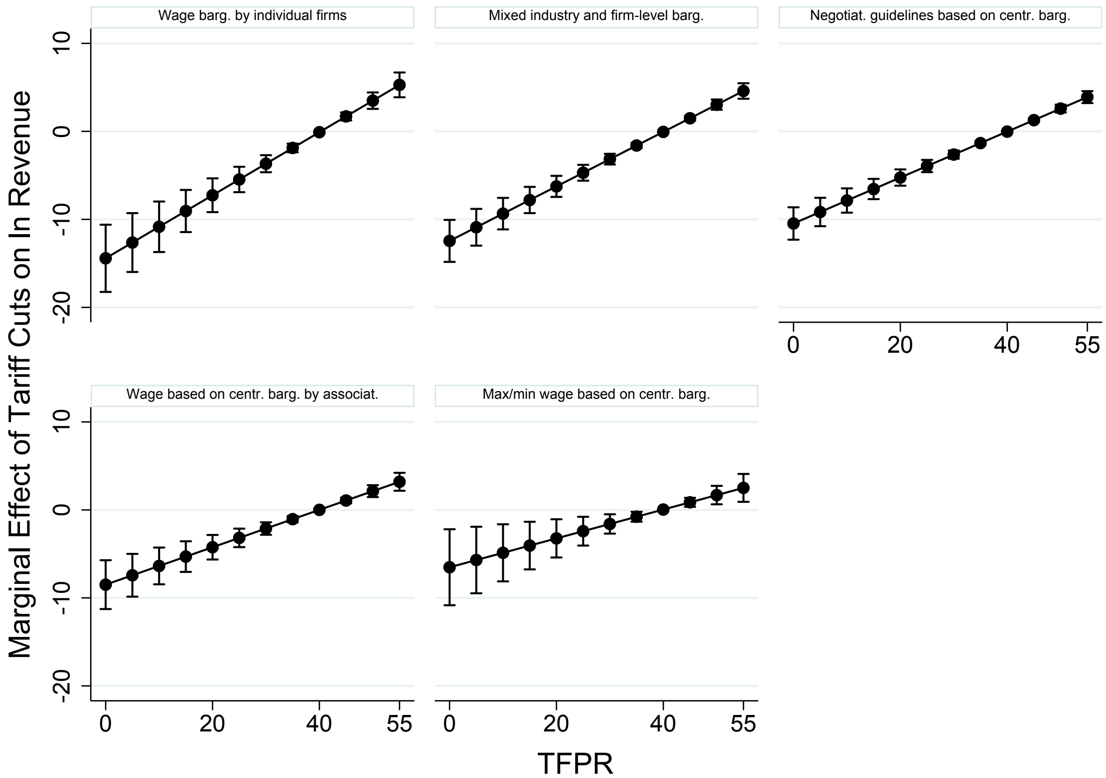
Table 2: Identification test

	OLS				
	ln Revenue				
	(1)	(2)	(3)	(4)	(5)
TFPR	0.30**	0.13**	-0.69**	0.26**	-1.19**
	(0.006)	(0.007)	(0.078)	(0.006)	(0.079)
$\Delta\tau$	-18.92**	-24.75**	-64.27**	-25.76**	-121.26**
	(2.143)	(1.966)	(18.672)	(2.539)	(0.241)
CME	-0.75**	-1.05**	4.99**	-0.82**	4.00**
	(0.084)	(0.087)	(0.174)	(0.080)	(0.171)
TFPR*$\Delta\tau$	0.50**	0.66**	1.69**	0.69**	3.24**
	(0.057)	(0.053)	(0.499)	(0.069)	(0.625)
TFPR*CME	0.02**	0.03**	0.13**	0.03**	0.11**
	(0.002)	(0.002)	(0.005)	(0.002)	(0.005)
$\Delta\tau$*CME	1.68**	2.72**	4.20**	-2.25**	5.41**
	(0.647)	(0.675)	(0.823)	(0.638)	(1.093)
TFPR*$\Delta\tau$*CME	**-0.04****	**-0.07*****	**-0.11****	**-0.06****	**-0.14****
	(0.017)	**(0.018)**	**(0.022)**	**(0.018)**	**(0.030)**
Constant	3.24**	4.80**	-4.75**	4.11**	14.74**
	(0.099)	(0.032)	(0.135)	(0.028)	(0.190)
Observations	4,053,929	4,053,929	3,217,585	4,044,630	3,212,608
R-squared	0.767	0.767	0.803	0.766	0.804
Controls	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes	Yes
Corruption	Yes	No	No	No	No
Unemployement	No	Yes	No	No	No
Market structure	No	No	Yes	No	No
Access to credit	No	No	No	Yes	No
All	No	No	No	No	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by firms in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. The outcome variable in all models is the log of revenue. Sources: Amadeus dataset, Baccini et al. (2018), Visser (2016), WGI, WDI, and ILO.

Figure 2: The effect of tariff cuts on firm revenue for different levels of productivity and for different labor markets



Note: The predictions are plotted from Model 2 in Table 1. CME=1: "Fragmented wage bargaining, confined largely to individual firms or plants." CME=2: "Mixed industry and firm-level bargaining, weak government coordination through MW setting or wage indexation." CME=3: "Negotiation guidelines based on centralized bargaining." CME=4: "Wage norms based on centralized bargaining by peak associations with or without government involvement." CME=5: "Maximum or minimum wage rates/increases based on centralized bargaining." 99% C.I.

across industries.[25] We run a model with the first differences of wages as the outcome variable and the interaction between $\Delta\tau$ and *CME* as key independent variables. We also include year, industry, and country fixed effects. Results are reported in Table 3 (Model 1). The coefficient of the interaction term is negative and significant as expected. This supports the argument that wages in CMEs increase less than wages in LMEs in case of tariff cuts.[26]

Second, we use firm-level data capturing the cost of employees over revenue as the dependent variable.[27] Assuming that workers (other-than-wage) benefits do not change differentially between CMEs and LMEs as a result of tariff reduction, this should be a good proxy for wages. We run models with this variable as outcome and the interaction between $\Delta\tau$ and *CME* as key independent

[25] We use available crosswalks to merge the ISIC 88 Rev 3 industries, which the ILO uses, with the NAICS 4-digit industries of our dataset.

[26] Figure 9 in Appendix E shows the effect of the interaction term graphically.

[27] We are unable to use the first differences of variable capturing the cost of employees over revenue, since our data are repeated cross-sectionally.

Table 3: Mechanism: wages

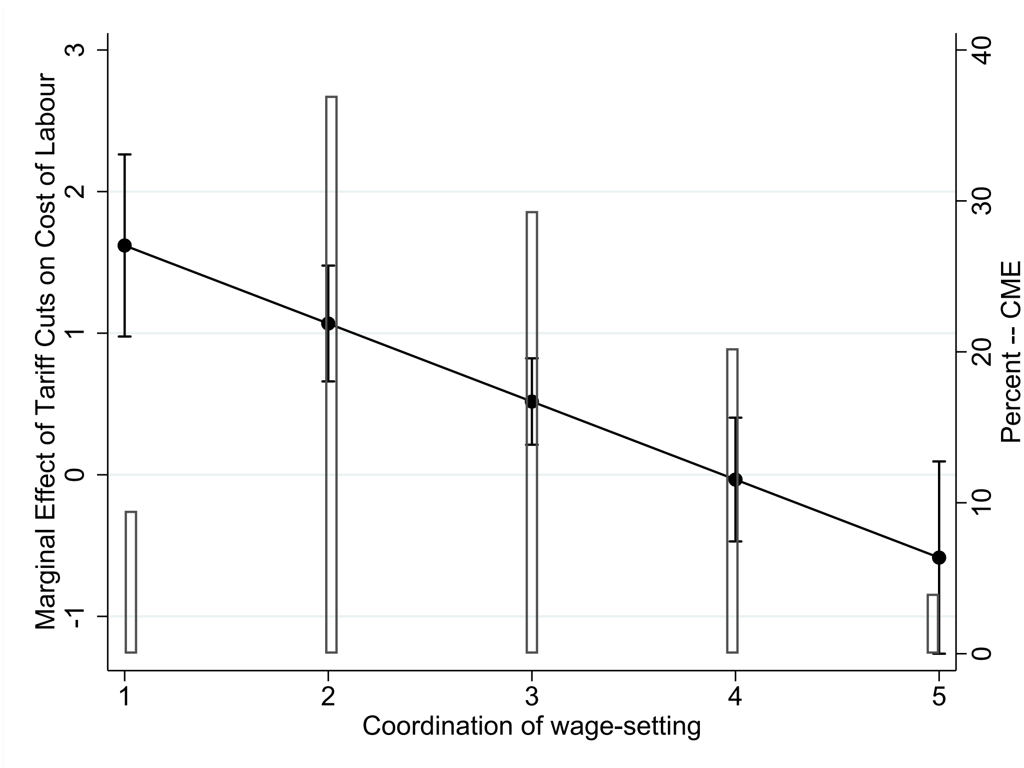
	OLS							
	Wage (f.d.)	Cost of employees/revenue						
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\Delta\tau$	4.45*	4.39**	2.17**	8.67**	6.89**	-1.84	8.44**	17.34**
	(1.875)	(0.369)	(0.360)	(0.402)	(0.390)	(4.249)	(0.447)	(5.978)
CME	-0.05**	-0.76**	5.23**	-0.78**	-0.74**	-0.87**	-0.83**	0.33**
	(0.009)	(0.020)	(0.191)	(0.020)	(0.020)	(0.027)	(0.020)	(0.030)
$\Delta\tau$*CME	**-1.13***	**-1.41****	**-0.60****	**-1.41****	**-0.28***	**-2.96****	**-0.61****	**1.94****
	(0.417)	**(0.118)**	**(0.116)**	**(0.117)**	**(0.136)**	**(0.206)**	**(0.129)**	**(0.313)**
Constant	0.19**	110.71**	95.30**	86.19**	121.19**	-47.21**	108.41**	55.65**
	(0.045)	(0.534)	(0.751)	(0.336)	(0.724)	(4.314)	(0.574)	(0.761)
Observations	1,202	3,629,212	3,629,212	3,629,212	3,629,212	2,903,748	3,628,568	2,903,105
R-squared	0.220	0.318	0.320	0.319	0.319	0.320	0.321	0.251
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Robust standard errors in parentheses ** $p<0.01$, * $p<0.05$

Note: OLS with robust standard errors clustered by firms in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. The outcome variables are wages and cost of employees over revenue. Sources: ILO, Amadeus dataset, Baccini et al. (2018), and Visser (2016).

variables. Results are shown in Table 3. In Model 2, we include firm size, capital-labor ratio, market concentration, and MFN tariffs as controls as well as year, industry, and country fixed effects. In Model 3, we include the interaction between each control and *CME*. In Models 4-8, we include the interaction term between $\Delta\tau$ and potential confounders in line with our previous identification tests. The coefficient of the interaction between $\Delta\tau$ and *CME* is negative and significant across these models. Figure 3 shows the effect of the interaction term graphically. As expected, the cost of labor increases more for LMEs than CMEs as a result of preferential trade liberalization.

Figure 3: The effect of tariff cuts on cost of labor for different labor markets



Note: The predictions are plotted from Model 3 in Table 3. 99% C.I.

Another mechanism implied by our theory is that wage coordination should reduce the reallocation effect, especially in differentiated industries. In differentiated industries, firms are not in direct competition with each other, since they do not sell the exact same product but a close substitute. Thus, firms should be able to keep their share of the market even in the presence of price differentials due to consumers love of variety and their attachment to specific brands. In this case, the reallocation effect should be driven almost exclusively by the labor market channel.[28] In short, after trade liberalization an economy experiences an increase in economic activity. This leads to an increase in workers demands, which in turn pushes wages up. In turn, as the result of increasing costs of production, the most productive firms, which are able to recoup these costs, erode the market share of unproductive firms that cannot sustain these rising wages.

On the contrary, the competition channel should trump the labor market channel in homogenous

[28]This mechanism comes directly from the original Melitz (2003) model, whose setup is a monopolistic market in which there is only one firm in each product line.

industries in which firms sell the exact same product. Simply put, as a consequence of trade liberalization, we should see consumers shifting their demand to the cheapest undifferentiated goods with or without an increase in wages.

Table 4: Mechanism: product differentiation

	OLS		
	ln Revenue		
	Differentiated	Referenced	Homogeneous
VARIABLES	(1)	(2)	(3)
TFPR	0.29**	0.36**	0.44**
	(0.008)	(0.014)	(0.031)
Δτ	-15.78**	-9.37*	-30.07*
	(2.495)	(3.961)	(12.794)
CME	-0.03**	-0.20	-0.02
	(0.002)	(0.211)	(0.476)
TFPR*Δτ	0.42**	0.25*	0.80*
	(0.067)	(0.106)	(0.341)
TFPR*CME	2.09**	0.01	-0.00
	(0.830)	(0.006)	(0.013)
Δτ*CME	0.04**	-0.00	1.22
	(0.008)	(0.017)	(4.091)
TFPR*Δτ*CME	**-0.05****	**-0.01**	**-0.03**
	(0.022)	**(0.035)**	**(0.109)**
Constant	4.67**	5.01**	4.21*
	(0.180)	(0.098)	(1.828)
Observations	2,532,064	790,678	115,223
R-squared	0.783	0.795	0.757
Controls	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Country FE	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by firms in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. The outcome variable in all models is the log of revenue. Sources: Amadeus dataset, Baccini et al. (2018), Visser (2016), and Rauch (1999).

To test this mechanism, we use Rauchs (1999) categorization of differentiated, referenced, homogenous industries.[29] We then run our main models for these three split samples. Table 4 shows the

[29]Rauch data use the SIC 4-digit trade categorization. We use available crosswalks to merge the data to our NAICS

results of this test. As expected, differentiated industries are the ones driving the results. Indeed, our key triple interaction terms is negative and significant only in Model 1, whereas it is not significant in Models 2 and 3. Importantly, the coefficient of the triple interaction term is substantively larger in Model 1 compared to Models 2 and 3. In sum, labor market institutions matter exactly in those industries in which trade models suggest that these institutions should matter.[30]

Additional evidence We test further implications of our theory by corroborating the robustness of our main findings with different model specifications. While we give the details in Appendix C, we summarize the main findings here. To begin with, we show that firm productivity increases more in LMEs than in CMEs in the short term, whereas it grows more in CMEs than in LMEs in the long term, a result in line with the VoC literature. Moreover, we show that our results are robust to the use of alternative measures of the labor market frictions. Furthermore, we provide further evidence on the mechanism at play. Two results stand out. First, labor flexibility does not seem to affect the reallocation effect. Second, the reallocation effect is driven by a surge of imports triggered by preferential tariff cuts as in Melitz (2003). Finally, we run a battery of robustness checks, which leave our main findings unchanged.

6 Gains From Trade and Attitude Towards Redistribution

The analysis at the firm level has shown that gains from trade are more uniform in CMEs than in LMEs. That is, productive firms in CMEs benefit from trade liberalization less than productive firms in LMEs do. In this section we explore whether more uniform gains from trade lead to a weaker demand for redistribution in CMEs compared to LMEs. The logic is that workers share the same destiny as their firms. When firms gain from trade liberalization, so do workers. On the contrary, when firms lose from trade liberalization, workers suffer in terms of being laid off and/or receiving lower income. Because winners and losers from trade are less clear-cut in CMEs than in LMEs, we expect a differential effect of trade liberalization on attitude towards redistribution across labor market institutions.

Model To test the effect of preferential liberalization on individual attitude toward redistribution, we make use of the European Social Survey (ESS), which covers all EU countries and several waves from 2002 to 2016. Importantly for us, the ESS reports the geographical location of each respondent at the level of NUTS-2 regions.[31] Following previous studies (Rehm 2009; Wren and Rehm 2013; Walter 2017), we rely on the following question to capture preferences over redistribution: *The government*

4-digit database.

[30] In the sub-sample of differentiated goods, results are robust to the inclusion of industry 4-digit-year fixed effects (available upon request).

[31] The classification of these regions changed throughout the ESS survey waves in terms of scale. For each country represented in the ESS surveys, we first identified how their region variable changed over time (becoming finer, or coarser, or remaining the same). For the ESS surveys between 2002 and 2008, we matched these region values to respective NUTS-1, -2, -3 2013 regions based on the official Eurostat names. In the case where region names changed, we used Eurostat tables that show the changes in NUTS regions over time to identify the relevant NUTS 2013 region (`http://ec.europa.eu/eurostat/web/nuts/history`). For the ESS surveys from 2010 onwards, the various region variables were collapsed into one, which contained the actual NUTS code rather than the official name of the region, and thus did not require additional matching.

should take measures to reduce differences in income levels. While this variable is originally a five-point-scale, we recode it in a dummy scoring one if respondents agree or strongly agree with the aforementioned question.[32]

Our main independent variable measures the magnitude of trade liberalization in a specific industry i weighted on the share of manufacturing of industry i in a NUTS-2 region r. In order to build our independent variable, we geo-coded all the firms used in the previous analysis at the level of a NUTS-2 region.[33] This Bartik instrument is similar to the one used by Autor et al. (2013) and by Colantone and Stanig (2018a). More formally, the instrument is given by the following equation:

$$Instrument\ for\ PRF\ Liberalization_{crt} = \sum_{j} \frac{L_{rjf}}{L_r} \times \frac{\Delta\tau_{jt}}{Import_{cj}}, \tag{2}$$

where c indexes countries, r NUTS-2 regions, j industries, f firms, and t years.

$\frac{\Delta\tau_{jt}}{Import_{cj}}$ is the yearly change in preferential tariff cuts in country c and industry j. These tariff cuts are normalized by imports in the same country and industry before the PTA is signed. In order to back out the region-specific trade shock, we take the weighted sum of the change in tariff cuts per worker across industries, where the weights capture the relative importance of each industry in a given region. Specifically, the weights are defined as the ratio of the number of workers in region r and industry j over the total number of workers in the region.

The important difference with respect to previous studies is given by the index *f*. Indeed, we are interested in the share of employees in region r and industry j working in very productive firms, which we expect to gain from trade liberalization proportionally more than less productive firms. Thus, the numerator of $\frac{L_{rjf}}{L_r}$ measures the share of workers in firms belonging to the upper quartile of the productivity distribution. The underlying logic is as follows: larger preferential liberalization shocks are attributed to regions characterized by larger shares of workers employed in very productive firms, who should gain disproportionally more from tariff cuts than workers employed in any other firms. [34]

The unit of analysis is respondent-NUTS-2 region-country-ESS wave. Since ESS waves are every other year, we take the bi-yearly sum of the equation 2. We drop the 2002 ESS wave, because we do not have firm-level data prior to 2003. Armed with these dependent and independent variables, we estimate the following baseline model:

$$\begin{aligned} Redistribution_{prcw} = \gamma_0 + \gamma_1 PRF\ Liberalization_{rcw} + \gamma_2 CME_{cw} + \gamma_3 PRF\ Liberalization_{rcw} \times \\ CME_{cw} + \gamma_4 X_{prcw} + \gamma_5 X_{prcw} \times CME_{cw} + \delta_w + \delta_r + \epsilon_{prcw}, \end{aligned} \tag{3}$$

where p indexes people responding to the ESS, r NUTS-2 regions, c indexes countries, and w waves. γ_0, γ_1, ... γ_5 are the coefficient. In particular, the key coefficient of interest is γ_3, which we expect

[32]We use a dummy variable because the relevant variation is between those in favor and those against redistribution policies, whereas there is limited variation across the five-point-scale. Our results are similar if we use the ordinal measure. Figure 10 in Appendix E shows the geographical distribution of this variable.

[33]More details are provided in Appendix D.

[34]Figure 11 in Appendix E shows the geographical distribution of this variable.

to be negative. The logic is the following: (1) very productive firms gain disproportionally more than less productive firms as a result of trade liberalization, and the same is true for workers employed in these firms; (2) uneven gains from trade among firms and workers trigger demand for redistribution; (3) the gains from trade are more uniform in CMEs compared to LMEs due to labor market frictions; (4) this demand is weaker in CMEs compared to LMEs, as trade from grains are more uniform.

Moreover, δ_w and δ_r are wave and country fixed effects respectively. Wave fixed effects capture and control for overall trends in respondents' demand for redistribution. Region fixed effects net out time-invariant differences across NUTS-2 regions. ϵ_{prcw} accounts for all residual determinants of the dependent variable. Since CME_{cw} varies across country and over time, we are able to estimate its coefficient even when we include region fixede effects.

Furthermore, the matrix X includes standard individual-level controls. More specifically, we control for the industry in which respondents are employed (NACE 2-digit), level of income, level of education, gender, whether respondents are unemployed, whether respondents are members of a trade union, and ideology. Each of these controls is interacted with *CME*.

Our empirical strategy is a triple difference-in-differences, in which the treatment (*PRF Liberalization*) varies in intensity across regions and over time and it is interacted with labour market institutions. We run OLS regressions with standard errors clustered by region.[35] We have about 25,000 respondents per wave for a total of (about) 180,000 observations.

Results The main results are reported in Table 5. The coefficient of the interaction between the instrument in the equation 2 and *CME* is always negative and significant as expected. This is true also when we add country-wave fixed effects to account for time-varying difference among countries (Model 3) and region specific trends to take care of pre-trend issues (Model 4). In a nutshell, we observe that the demand for redistribution in CMEs is weaker than in LMEs in the case of preferential liberalization and of a large share of workers employed in very productive firms. This is evidence that gains from trade are more uniform in CMEs than in LMEs.

To help in visualizing our results, we graph the effect of the interaction term in Figure 4. In particular, we report the marginal effect of instrument for PRF liberalization for the five types of labor market institutions in EU countries. Figure 4 shows that when *CME*=1 and *CME*=2, i.e., low market frictions, preferential trade liberalization triggers an increase in the demand for redistribution, though the effect is not significant for *CME*=2. On the contrary, when *CME*=3, *CME*=4, and *CME*=5, i.e., large market frictions, preferential trade liberalization triggers a decrease in the demand for redistribution. The demand for redistribution is significantly higher in a country like the UK (i.e., *CME*=2) than in countries like Germany (*CME*=4) and, in fact, it is twice as high in the UK as it is in Germany. On the contrary, there is no difference across coordinated-market economies, i.e., *CME*=3, *CME*=4, and *CME*=5.

We conclude with a couple of further tests, whose results are reported in Appendix E. First, in line with the previous analysis, we interact possible confounders at the country level with our instrument for PRF liberalization and include these interactions on the right-hand side of our main model. The

[35] We opt for OLS regressions because we are in a difference-in-differences setting, which requires a linear estimator. Results are similar if we use logistic regressions.

Table 5: Demand for Redistribution

	OLS			
	Support for Redistribution			
	(1)	(2)	(3)	(4)
Instrument for PRF Liberalization	0.05*	0.04*	0.08**	0.12**
	(0.020)	(0.021)	(0.022)	(0.038)
CME	-0.00	-0.00		0.01
	(0.010)	(0.010)		(0.018)
Instrument for PRF Liberalization*CME	**-0.02****	**-0.02***	**-0.04****	**-0.06****
	(0.010)	**(0.010)**	**(0.011)**	**(0.019)**
Constant	0.74**	0.74**	0.76**	0.49**
	(0.045)	(0.044)	(0.052)	(0.078)
Observations	176,209	176,209	183,800	157,028
R-squared	0.075	0.075	0.072	0.089
Controls*CME	Yes	Yes	Yes	Yes
Wave FE	No	Yes	No	Yes
Region FE	Yes	Yes	No	Yes
Country-Wave FE	No	No	Yes	No
Trends	No	No	No	Yes

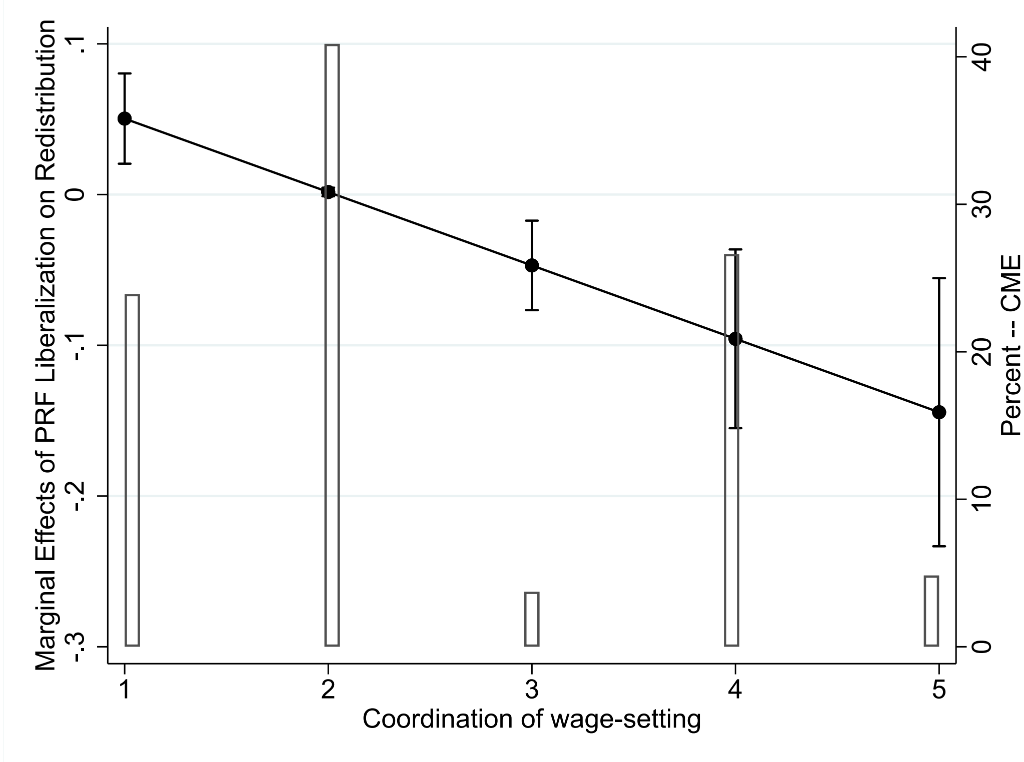
Robust standard errors in parentheses ** $p<0.01$, * $p<0.05$

Note: OLS with robust standard errors clustered by region in parentheses. Unit of observation is respondent-region-country-wave. The outcome variable in all models is a dummy scoring one if respondents answer strongly agree or agree to the following sentence: *The government should take measures to reduce differences in income levels.* Sources: Amadeus dataset, Baccini et al. (2018), Visser (2016), and ESS (2018).

main findings remain unchanged (see Table 16). Second, we show that our results are driven by low-income respondents, who are less likely to be employed in very productive firms and more likely to lose out from trade liberalization. To implement this effect heterogeneity, we split the sample into low- and high-income respondents and re-run our main model (see Table 17). The coefficient of the interaction between the instrument in equation 2 and *CME* remains negative and significant in Model 1 (low-income sub-sample), whereas it is not significant in Model 2 (high-income sub-sample). Finally, our results remain unchanged if we include other instruments for PRF liberalization with the lower quartile of firm productivity.

To sum up, our analysis at the individual level confirms the results of the analysis at the firm level. Distributional consequences of trade liberalization are more severe when labor market frictions are limited and as such demand for redistribution is weaker in CMEs than it is in LMEs.

Figure 4: The effect of instrument for PRF liberalization on individual attitude toward redistribution for different labor markets



Note: The predictions are plotted from Model 4 in Table 5. 95% C.I.

7 Conclusion

This paper explores the distributional consequences of trade liberalization across different types of labor market institutions. The main findings of the paper are twofold. In the analysis at the firm level, we show that the reallocation effect is weaker in CMEs than in LMEs. That is, revenue of productive firms increases proportionally less in CMEs compared to revenue of productive firms in LMEs. This effect is driven by smaller increases of wages in CMEs compared to LMEs due to labor market frictions, something that we documented in our analysis. In the analysis at the individual level, we find that, as a result of trade liberalization, the demand for redistribution is weaker in CMEs compared to LMEs. This effect is driven by more uniform gains from trade in CMEs compared to LMEs due to the aforementioned labor market frictions.

The policy implications of our analysis are important and timely. First, ours is the first paper to provide a micro-level analysis of the effect of preferential trade liberalization on a large number of firms across several countries. Our results indicate that even relatively minor tariff cuts implemented by developed economies with (usually) smaller and less developed markets have important distributional consequences and generate clear winners and losers. In this regard, our paper complements the results of recent studies on the China shock showing the severity of the economic damage produced by trade liberalization (Autor et al. 2013).

Second, our findings indicate that some labor market institutions mitigate the winner-take-all effect produced by trade liberalization, producing more uniform gains from trade. While trade liberalization is akin to increasing the market power of few large corporations (Osgood et al. 2016, Baccini et al. 2017), some countries are less prone than others to produce superstars, given the presence of labor market frictions. This is an unintended positive consequence of labor market frictions, which often have been blamed for high unemployment and sluggish economic growth.

Third, our results show that labor market institutions mediate the effect of trade liberalization on people's policy preferences. It is claimed that the current backlash against globalization in developed countries is also triggered by extensive job losses in manufacturing due to the competition from emerging economies (Colantone and Stanig 2018a, 2018b). This being the case, our findings indicate that not all countries are affected in the same way by trade openness. In particular, we document that variation in labor institutions leads to variation in levels of inequality once trade liberalization kicks in. In this regard, our paper helps explain why Brexit happened in the UK and not in Germany and why nationalist parties are in power in Hungary and Poland, but not in the Netherlands.

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Appendix A: Tariff Cuts

We build our tariff cut variable ($\Delta\tau$) following the steps below:

1. We have data on preferential (PRF) tariffs at the HS 6-digit level for all the PTAs signed by the EU post-1995. For each product, we know preferential tariffs in time zero, i.e., year of ratification, and for all subsequent years until preferential tariffs go to zero (up to 22 years). In other words, we know the phase-out tariff period for each product for each PTA.
2. For each product at the 6-digit level, we know the MFN tariff, which we use as baseline to calculate the tariff cut.
3. We create a variable *PRF* that captures the level of PRF tariff for each product for each PTA in each year. This variable takes into account the phase-out tariff period. For instance, if a PTA is ratified in 2000, *PRF* of product i includes the level of PRF tariff from 2000 to 2021.
4. We create a tariff cut variable for each product and for each PTA. Tariff cut is the difference between MFN and *PRF* in the year of ratification and it is the inverse of the first difference of *PRF*, i.e., *PRF lagged-PRF*, in subsequent years. In other words, to calculate the tariff cut, we use MFN as baseline for the first year in which PRF tariffs kick in and the PRF tariffs of the previous year in subsequent years in which a PTA is in force.
5. We create a variable capturing proportional tariff cuts, i.e., $\frac{MFN-PRF}{MFN}$, in the first year and $\frac{PRF\ lagged-PRF}{PRF\ lagged}$, following the same procedure as in 4.
6. We create weighted tariff cuts and weighted proportional tariff cuts dividing tariffs by import value. We then follow the same procedure as in 4 and 5.
7. We sum all the tariff cuts (weighted and not) across all EU PTAs for a given product i in a given year t. That gives us our measure of preferential trade liberalization.
8. We take the average value of proportional tariff cuts (weighted and not) across all EU PTAs for a given product i in a given year t.
9. We merge the dataset with a NAICS 4-digit variable to merge the tariff data with the Amadeus database.
10. We take the average value of all our measures of tariff cuts (proportional and not, weighted and not) in each year to move from HS 6-digit to NAICS 4-digit. Note that we did not sum the tariff cut in this case because there are different numbers of 6-digit products in 4-digit industries.

Appendix B: Confounders

The variables that we analyze as possible confounders in the empirical analysis are the following.

Corruption The logic is that corruption may create additional fixed or variable costs for firms, especially when competition increases due to trade liberalization. These additional costs are more likely to be supported by productive firms rather than unproductive firms. In turn, this creates uneven gains from trade. If corruption correlates with labor market fractions, it may be a confounder. We rely on a measure of control of corruption by the Worldwide Governance Indicators (WGI) of the World Bank (Kaufmann et al. 2010). The time span is between 1996 and 2016.[36]

Unemployment The logic is that (pre-trade-liberalization) high level unemployment reduces the increase of wages after trade liberalization. In turn, this may help unproductive firms in case of increasing competition due to tariff cuts. If unemployment correlates with labor market fractions, it may be a confounder. We rely on a measure of unemployment collected by the ILO and available through the WDI. The time span is between 1960 and 2016.

Market Structure We include social expenditure. Data are from the OECD and are available from 1990 to 2016. Moreover, we include the size of the service sector, amount of taxes over GDP, and amount of FDI outflows. Data are from the WDI and are available from 1960 to 2016. Finally, we include a dummy for countries that adopted the Euro. Data come from `https://europa.eu/european-union/about-eu/money/euro_en`. All these variables can mitigate (e.g., social expenditure) or magnify (Euro) the reallocation effect. Therefore, they are all potential confounders.

Access to Credit In countries in which access to credit is easy, firms can weather the increasing competition triggered by trade liberalization better than in countries in which firms face credit constraints. In particular, easy access to credit can help small, unproductive firms.[37] To capture access to credit we rely on the following variables: (1) domestic credit to private sector by banks (% of GDP); (2) domestic credit provided by financial sector (% of GDP); and (3) domestic credit to private sector (% of GDP). Data come from the WDI and are available from 1960 to 2016.[38]

[36] Results are similar if we use other variables capturing the quality of governance, e.g., rule of law and regulatory quality.

[37] For a review of the literature on trade liberalization and access to credit, see Folrey and Manova (2015).

[38] Results are similar if we use variables capturing access to credit from The Enterprise Survey of the World Bank. We do not rely on these variables in the main analysis because data start from 2006.

Appendix C: Additional Evidence

In what follows we show an auxiliary finding related to firm productivity, further results using alternative measures of the labor market, and a battery of robustness checks.

Auxiliary finding: firm productivity One of the main aspects in which we expect firms in LMEs and in CMEs to differ is the ways in which they pursue productivity increases. In LMEs, most of the workforce is characterized by generic skills that can be easily relocated in different sectors, and firms rely on "high volumes of production and low labor costs for competitive advantage" (Hall 2007: 46; Rubery 1994). In CMEs, the existence of institutions that make it difficult for firms to fire workers has incentivized companies to "base their competitive strategies more heavily on quality rather than cost considerations" (Hall 2007: 48). This has led firms in CMEs to develop sectoral vocational training schemes and to invest in high-skill workers. Simplifying, labor institutions in LMEs encourage firms to compete by lowering labor costs; labor institutions in CMEs encourage firms to compete by pursuing incremental technological innovation.

What can be the implications of these differentiated paths to productivity for the effects of trade liberalization? Hall and Soskice (2001) have stressed that LMEs are better equipped to support "radical innovation," while CMEs' institutions are more supportive of "incremental innovation." This means that LMEs will perform better in sectors like "biotechnology, semiconductors, and software development," while CMEs will perform better in "machine tools and factory equipment, consumer durables, engines, and specialized transport equipment (Hall and Soskice 2001: 39). No empirical analysis has, to our knowledge, tested which of the two models yields greater productivity increases, in the short- and in the long-term.

However, in light of the above, it makes sense to expect that productivity will increase more in LMEs than in CMEs in the *short term*, whereas productivity will increase more in CMEs than in LMEs in the *long term*. To test this argument, we run models in which the outcomes are *TFPR*. Our key variable of interest is the interaction between $\Delta\tau$ and *CME*. We also include standard controls like firm size, MFN tariffs, (log of) capital-labor ratio, and number of firms entering and exiting the industry. Moreover, we include industry, country, and year fixed effects.

We run error correction models to disentangle the short- from the long-term effect. As expected, the coefficient of the interaction term between the first differences of $\Delta\tau$ and *CME*, which estimates the short-term effect, is negative and significant (Table 9). On the contrary, the long-term multiplier of the interaction term is positive and significant across, implying that CMEs' productivity increases more than LMEs' productivity in the long term. These results represent a valuable insight for the VoC literature, in that they show how CMEs' focus on "incremental innovation" ensures greater productivity gains in the long term.

Other labor market institutions While wage coordination is among the most important institutional features of varieties of capitalism (see Hall and Gingerich 2009; Guardiancich and

Guidi 2016), there are other characteristics of the labor market that may be relevant to mediating the distributional consequences of trade liberalization. To address these concerns, we identify other variables from the ICTWSS database: union density, measure of centralization of wage bargaining, government intervention in wage bargaining, sectoral organization of employment relations, and authority of unions over affiliates. The variables that we analyze as alternative measures of labor market frictions are the following. All of them are taken from the ICTWSS database (Visser 2016).

Union Density The percentage of union members out of the total number of employed and salaried workers.

Centralization of wage bargaining A composite index that combines information about the predominant level at which wage bargaining takes place, the frequency or scope of additional enterprise bargaining, the possibility of renegotiation of contractual provisions at lower levels, the articulation of enterprise bargaining, and the possibility to derogate to national or sector-level agreements.

Government intervention in wage bargaining An ordinal variable ranging from 1 to 5, measuring the degree to which the government influences wage bargaining, where 1 means no intervention whatsoever and 5 means that the government imposes private sector wage settlements, places a ceiling on bargaining outcomes or suspends bargaining (Visser 2015).

Sectoral organization of employment relations An ordinal variable measuring how institutionalized are the relationships between employers and unions at the sectoral level. The possible values are 0 (no institutionalization), 1 (medium institutionalization), and 2 (strong institutionalization).

Authority of unions over their affiliates A proxy measuring the authority of confederations over sectoral or local branches. This variable combines information on whether the confederation is routinely involved in consultation with government, controls the appointment of affiliates' leaders, is involved in negotiation of the affiliates' wage agreements, has a fund for official strikes, and can veto strikes by affiliates.

We interact each of the aforementioned variables with $TFPR$ and $\Delta\tau$. Because these variables tend to be highly collinear, we do not include all of them at the same time and we do not include them together with our main triple interaction term. Results are reported in Table 10 in Appendix E. Three out of seven triple interactions are significant and have the expected negative sign. More specifically, government intervention in wage bargaining weakens the reallocation effect as well as authority of confederation over its affiliates and mandatory extension of collective agreements to non-organized employers. These results confirm that labor market frictions help unproductive firms to reduce uneven distributional consequences of trade liberalization through posing a wage ceiling.

Alternative mechanisms Table 11 in Appendix E shows that our results hold even if we use a measure of labor flexibility, which represents another type of labor friction (Model 1).[39] Moreover, our results hold if we include the triple interaction of *MFN*, *TFPR*, and *CME*, which is not significant (Model 2), and if we include the log of import from the rest of the World as control (Model 3). Importantly, when we include the interaction among the log of import, *TFPR*, and *CME*, this term has the expected negative sign and "kills" the significance of our main triple interaction terms (Model 4).

This last piece of empirical evidence is completely in line with our argument: (1) the reallocation effect is triggered by increasing competition from foreign firms as a result of preferential tariff cuts; (2) the reallocation effect is weaker for CMEs than LMEs due to labor market frictions, which help the unproductive firms to compete with the productive firms in case of preferential trade liberalization.

Robustness checks We implement several robustness checks, which we report in the Appendix. First, we show that our results hold if we use different measures of productivity (see Table 12 in Appendix E). Second, our results are unchanged if we double-cluster the standard errors by firms and industries. Moreover, our results are robust to the inclusion of firm fixed effects. We do not include firm fixed effects in the main model, because our data are repeated cross-sectionally. Furthermore, our results hold if we include a lagged dependent variable on the right-hand side. We do not include it in the main model to avoid the Nickell bias (1981). In addition, results hold if we use (the log of) profit instead of (the log of) revenue. All these tests are reported in Table 13 in Appendix E.

Furthermore, we show that our results are similar if we use cumulative tariff cuts (weighted and non-weighted). In addition, we rerun our main model with tariff cuts that will be implemented after 2016 as a placebo test. As expected, the coefficient of the triple interaction term is not significant because these tariff cuts have not kicked in yet. Moreover, we show that results are similar even if we use export preferential tariffs. We do not include export tariffs in the main analysis since many firms are not exporters or have missing export data in the Amadeus database. All these tests are reported in Table 14 in Appendix E.

Finally, we run our main model for each country to document transparently which countries drive our results (see Table 15 in Appendix E). For instance, the coefficient of the interaction between $\Delta\tau$ and *TFPR* is twice as large for the UK (i.e., CME=1) as it is for Germany (i.e., CME=4).[40] In fact, the coefficient of the interaction between $\Delta\tau$ and *TFPR* is not significant for countries like Germany, Denmark, and the Netherlands, implying that the reallocation effect is quite weak.[41]

[39] Data come from the OECD. Large values of *labor flexibility* implies that firms cannot easily lay off workers. The correlation between *CME* and *labor flexibility* is -0.01.

[40] Our results hold if we drop both Germany and the UK.

[41] Our results are not sensitive to the exclusion of any of these countries.

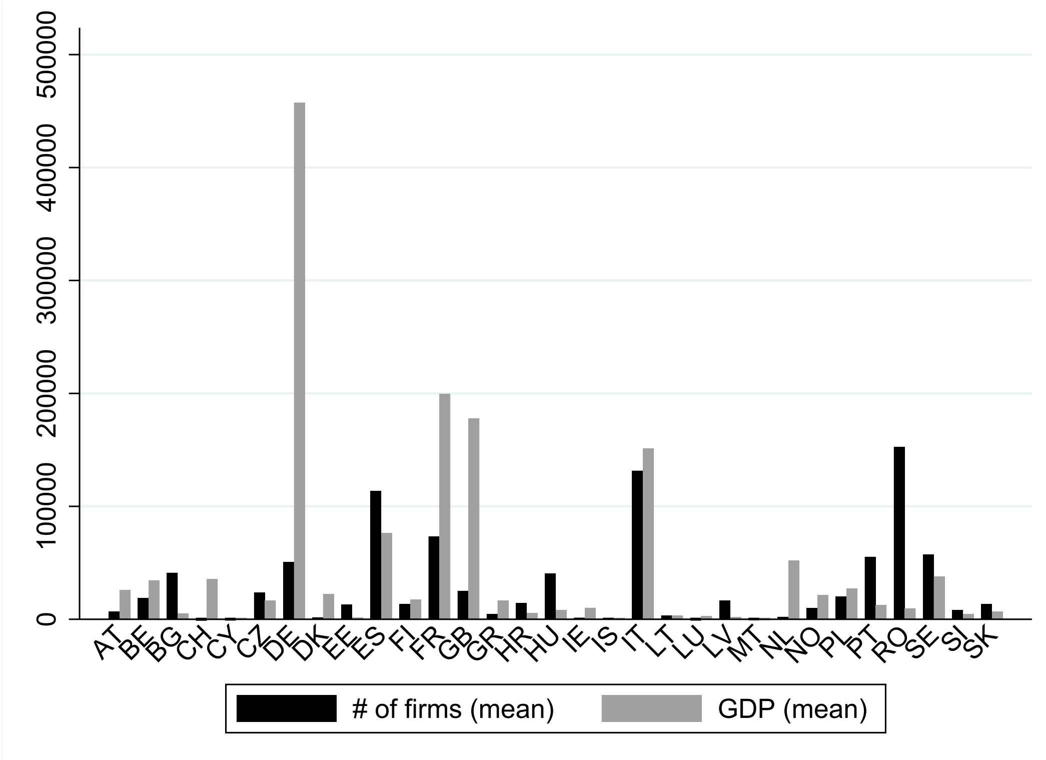
Appendix D: Geocoding Amadeus

Geocoding Amadeus was performed differently for each country. There is no standardized method, as each Amadeus dataset had different values in terms of the geographic variables. First, we looked at the postal code variable (zip code). Eurostat provides postcodes to NUTS region tables for each country in the European Union; however, in many cases the matches were geographically inaccurate. The postal code was still useful in some cases, especially in countries with relatively well-documented postal code systems. We then resorted to the region variable provided in Amadeus, which contains the general region in which a firm is located. The entries in the region variable often matched with a NUTS-2 or -3 level name. In most cases, if a country had NUTS-3 names within the region variable, a simple merge was performed. In other countries the region variable was finer in scale, corresponding to Local Administrative Units (LAUs), which are used by Eurostat to a lesser extent. Again, once the administrative level used in the region variable was identified, a merge was performed.

In the rare case where the region did not match any of the official Eurostat tables, we resorted to official country statistics websites to determine which administrative levels were used. Geocoding based on the region variable covered most of the Amadeus observations, and if a dataset was incomplete, we used a combination of the city and region variables to geocode. This combination was used to prevent any errors which may arise due to duplicate city names in certain countries. String matching based on city and region was performed with the help of data from Geonames, a free geographic database which covers all countries and place names (`https://www.geonames.org/`). These datasets contain the relevant administrative boundaries, which often matched Eurostat's NUTS-2 or -3 official names, and again a simple merge was performed.

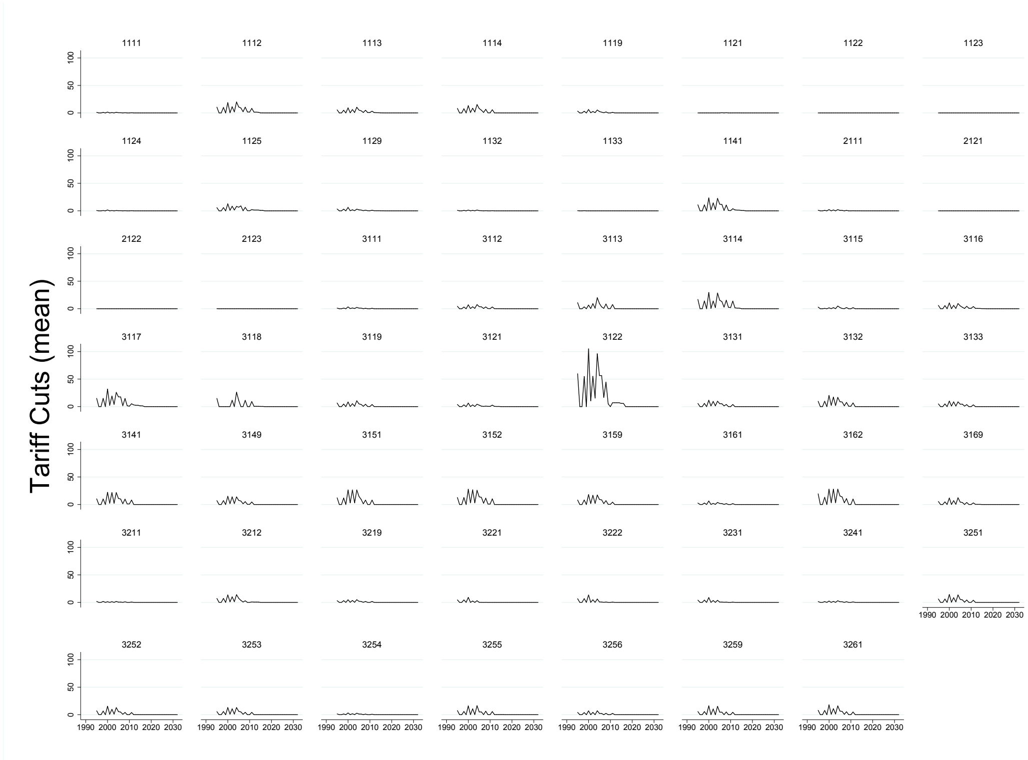
Appendix E: Figures and Tables

Figure 5: Sample: number of firms and GDP



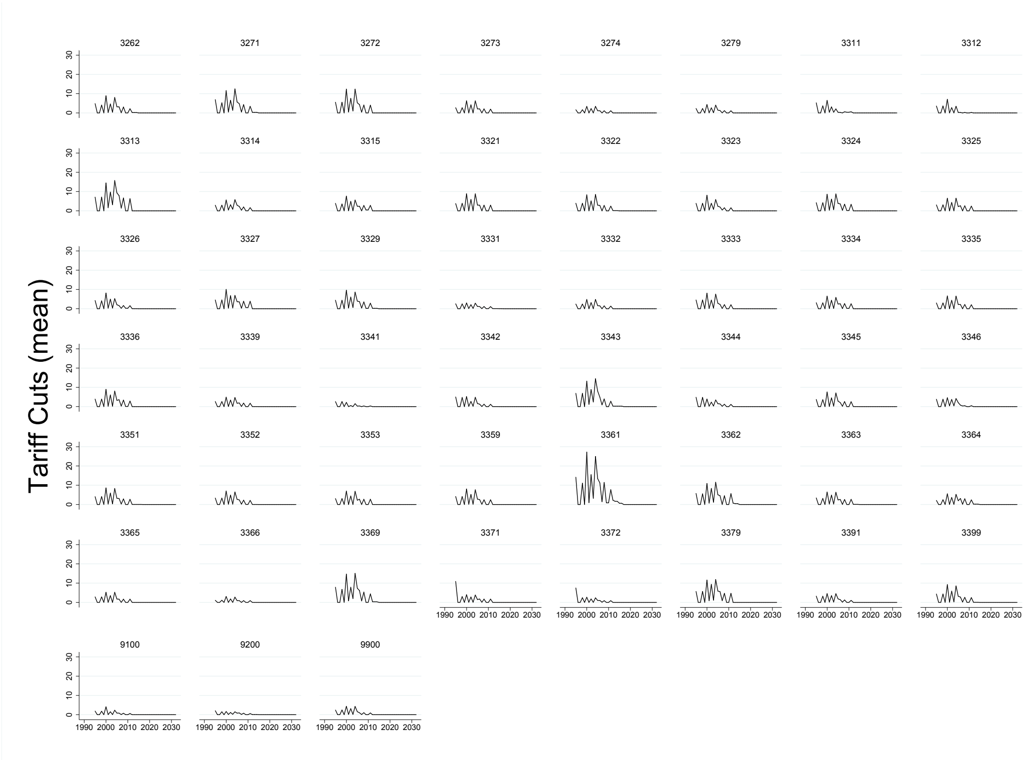
Note: Sources: Amadeus dataset and WDI.

Figure 6: Tariff cuts by industry and time (part 1)



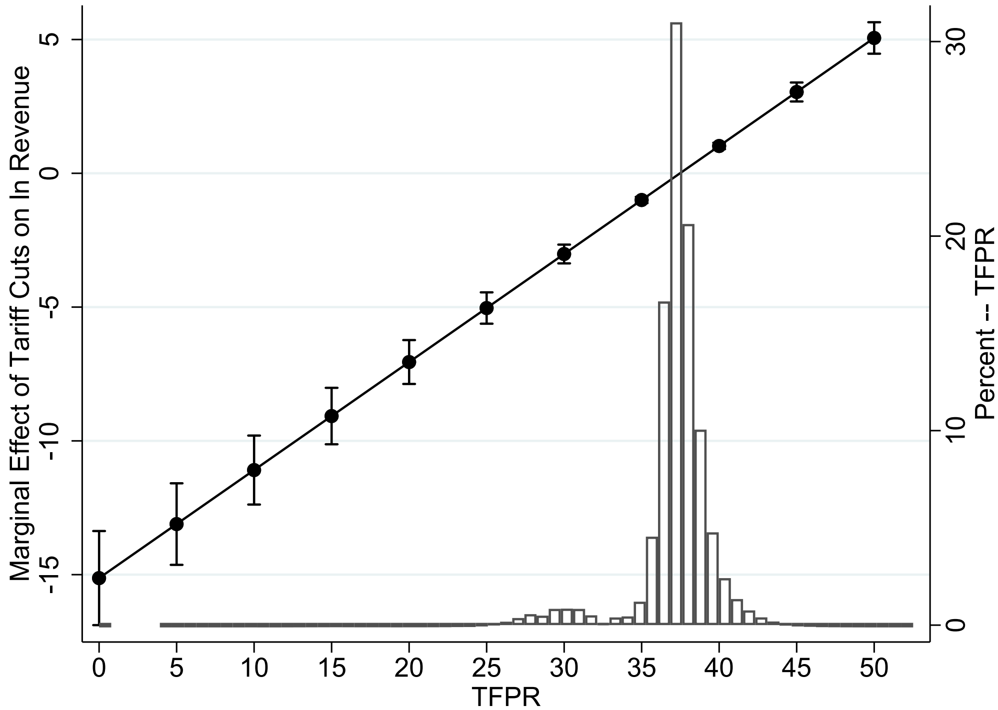
Note: Source: Baccini et al. (2018).

Figure 7: Tariff cuts by industry and time (part 2)



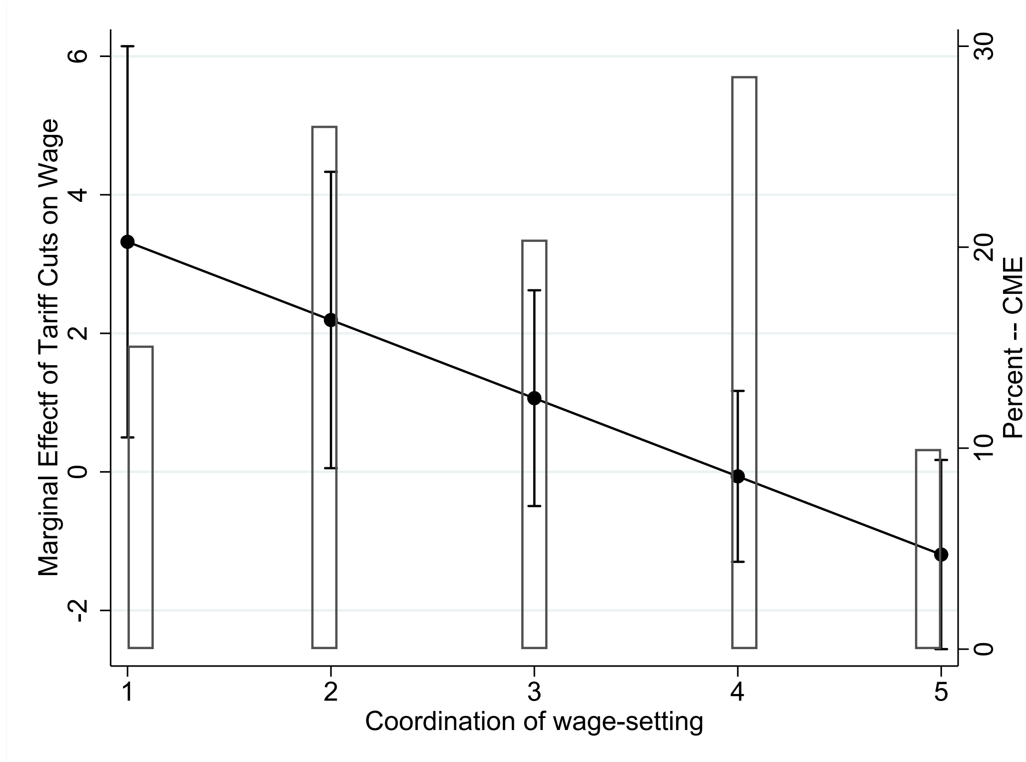
Note: Source: Baccini et al. (2018).

Figure 8: The effect of tariff cuts on firm's revenue



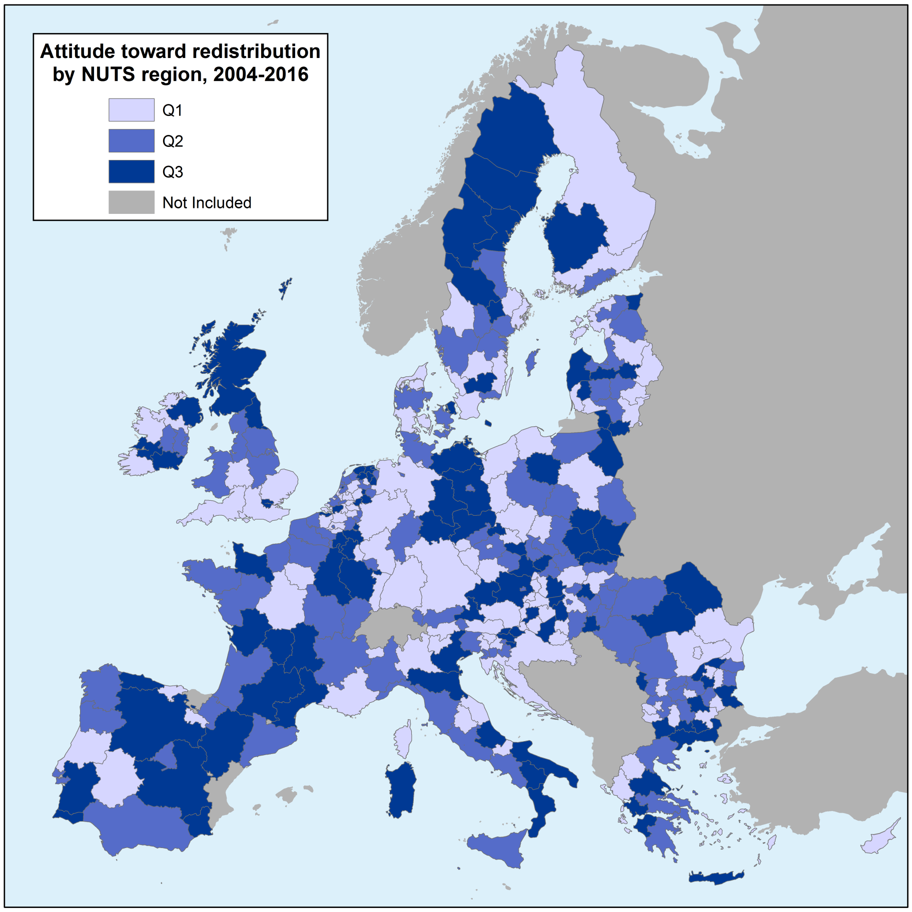
Note: The predictions are plotted from Column 1 in Table 1. 99% C.I.

Figure 9: The effect of tariff cuts on wage



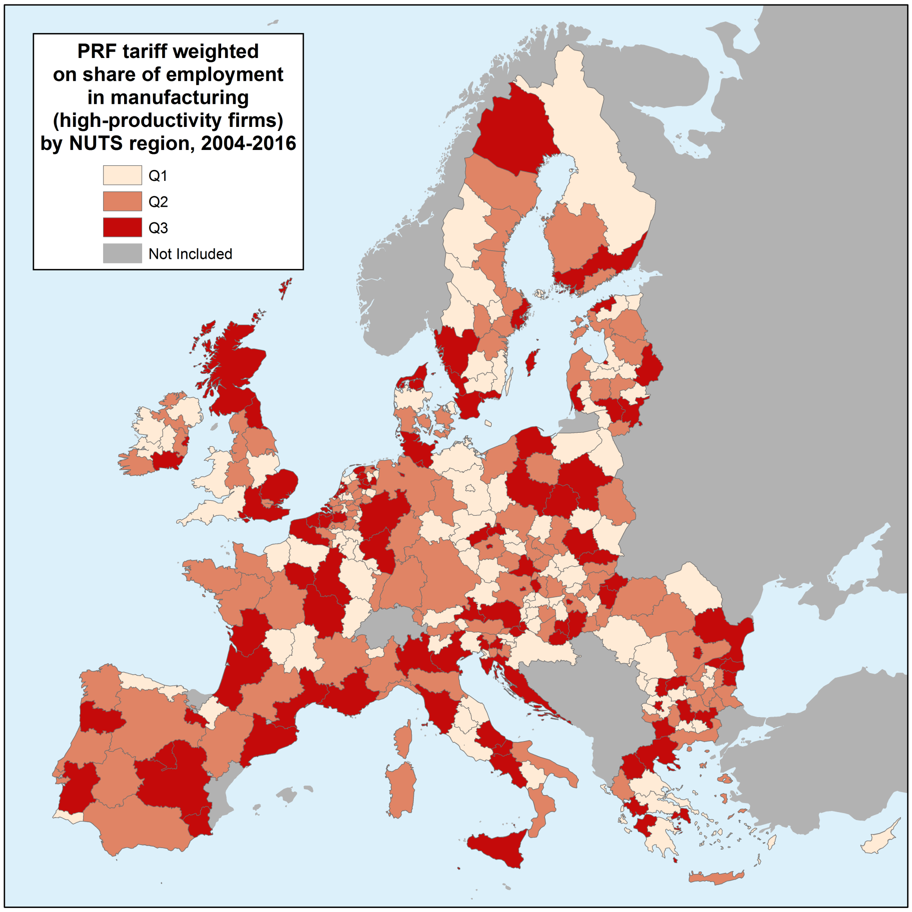
Note: The predictions are plotted from Column 1 in Table 3. 99% C.I.

Figure 10: Demand for redistribution



Note: The variable capturing individual attitude towards redistribution is a dummy scoring one if respondents answer strongly agree or agree to the following sentence: *The government should take measures to reduce differences in income levels.* Data are unavailable for ES21, ES53, ES70. Regions FRA1, FRA2, FRA3, FRA4, FRA5, ES63, ES64, PT20, PT30 are not shown on the map.

Figure 11: Instrument for PRF liberalization



Note: The variable *Instrument for PRF Liberalization* measures preferential tariff cuts weighted on the share of manufacturing workers employed in very productive firms. Data are unavailable for ES21, ES53, ES70. Regions FRA1, FRA2, FRA3, FRA4, FRA5, ES63, ES64, PT20, PT30 are not shown on the map.

Table 6: Main analysis (time-varying TFPR)

	(1)	(2)	(3)	(4)	(5)	(6)
	OLS					
	ln Revenue					
TFPR	0.86**	0.86**	0.89**	0.90*	0.86*	0.86**
	(0.001)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Δτ	-3.71**	-11.55**	-6.89**	-4.95**	-12.04**	-11.19**
	(0.230)	(0.752)	(0.682)	(0.702)	(0.755)	(0.012)
CME		0.21**			0.11**	0.29**
		(0.001)			(0.031)	(0.001)
TFPR*Δτ	0.09**	0.30**	0.17**	0.11**	0.32*	0.30**
	(0.006)	(0.019)	(0.017)	(0.017)	(0.019)	(0.019)
TFPR*CME		0.002**	-0.00	-0.002**	0.005**	0.00**
		(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Δτ*CME		2.90**	1.83**	1.27**	3.08**	2.08**
		(0.261)	(0.235)	(0.237)	(0.261)	(0.104)
TFPR*Δτ*CME		**-0.08****	**-0.05****	**-0.03****	**-0.08****	**-0.08****
		(0.007)	**(0.006)**	**(0.006)**	**(0.007)**	**(0.007)**
Constant	28.80**	-30.04**	-29.20**	-31.22**	-28.85**	13.77**
	(0.033)	(0.087)	(0.306)	(0.769)	(0.194)	(0.915)
Observations	5,148,266	4,053,929	4,053,929	4,053,929	4,053,929	4,053,929
R-squared	0.946	0.951	0.958	0.960	0.952	0.951
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	No	No	No	Yes
Country FE	Yes	Yes	No	No	No	Yes
Industry FE	Yes	Yes	Yes	Yes	Yes	Yes
CountryYear FE	No	No	Yes	Yes	Yes	No
IndustryYear FE	No	No	No	Yes	Yes	No
CountryIndustry FE	No	No	No	No	Yes	No
Trends	No	No	No	No	No	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by firms in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. The outcome variable in all models is the log of revenue. Sources: Amadeus dataset, Baccini et al, (2018), and Visser (2016).

Table 7: Correlations of confounders

	CME	Corruption	Unemployment	Social expenditure	Services (%GDP)	Tax (%GDP)	FDI outflows	Euro	Private credit	Bank credit	Financial credit
CME	1										
Corruption	**0.48**	1									
Unemployment	**0.11**	-0.12	1								
Social expenditure	**0.33**	0.27	-0.03	1							
Services (%GDP)	**0.25**	0.20	-0.01	0.77	1						
Tax (%GDP)	**0.20**	0.05	-0.32	0.37	0.16	1					
FDI outflows	**0.65**	0.72	-0.18	0.34	0.28	0.44	1				
Euro	**0.17**	-0.10	0.33	0.36	0.50	-0.27	-0.27	1			
Private credit	**0.21**	0.39	0.50	0.17	0.26	-0.22	0.13	0.33	1		
Bank credit	**0.21**	0.39	0.50	0.17	0.25	-0.22	0.13	0.33	1	1	
Financial credit	**0.30**	0.25	0.55	0.43	0.40	-0.18	0.07	0.55	0.89	0.89	1

Table 8: Descriptive Statistics

Variable	Obs	Mean	Std. Dev.	Min	Max	Source
ln Revenue	5,673,960	12.87	4.12	-21.93	29.35	Amadeus
$\Delta\tau$	5,673,960	0.02	0.09	0	3.82	Baccini et al (2018)
TFPR (baseline)	5,673,960	37.21	2.33	0	52.52	Amadeus
CME	4,452,775	2.75	1.03	1	5	Visser (2016)
TFPR*$\Delta\tau$*CME	4,452,775	2.96	11.31	0	585.05	
TFPR*$\Delta\tau$	5,673,960	0.82	3.18	0	146.26	
TFPR*CME	4,452,775	102.36	39.22	0	231.02	
$\Delta\tau$*CME	4,452,775	0.08	0.30	0	15.27	
MFN	5,673,960	3.70	3.89	0	47.26	Baccini et al (2018)
HHI	5,673,960	0.06	0.10	0	1	Amadeus
ln K/L	5,135,314	11.73	1.86	-17.74	26.07	Amadeus
Age	5,673,960	8.75	2.23	1	10	Amadeus
Age2	5,673,960	81.55	30.63	1	100	Amadeus
Firm Size	5,673,960	2.04	1.43	0.00	13.31	Amadeus
TFPR (time varying)	5,673,960	0.00	2.38	-39.20	15.49	Amadeus
ln Trade	5,673,960	8.97	7.28	0	19.16	Comtrade
Labour Flexibility	2,954,954	2.56	0.63	1.10	4.58	OECD
Union Density	3,061,883	27.20	17.26	6.53	77.71	Visser (2016)
Centralization	2,597,609	0.35	0.11	0.10	0.88	Visser (2016)
Government Intervention	4,430,996	2.99	0.95	1.50	5	Visser (2016)
Ext	4,442,412	1.43	1.28	0	3	Visser (2016)
Sector	4,341,650	1.32	0.71	0	2	Visser (2016)
Unauthority	4,319,871	0.34	0.14	0	0.80	Visser (2016)
Cfauthority	4,319,871	0.37	0.16	0	0.70	Visser (2016)
Corruption	5,673,960	0.73	0.84	-0.58	2.23	www.govindicators.org
Social Expenditure	4,195,067	25.13	3.74	11.00	31.90	WDI
Services (%GDP)	5,673,960	61.12	9.61	42.48	77.81	WDI
Taxes (%GDP)	5,673,960	19.53	4.33	1.50	51.11	WDI
FDI outflows	5,672,982	0.91	1.30	-0.06	7.68	WDI
Euro	5,673,960	0.57	0.50	0	1	https://europa.eu/
Private Credit	5,664,554	89.73	44.01	0.19	253.26	WDI
Bank Credit	5,664,554	89.67	44.00	0.19	253.15	WDI
Financial Credit	5,664,554	128.19	63.83	0.23	316.61	WDI
Unemployment	5,673,960	11.88	5.33	2.92	22.67	ILO
Wage f.d.	1,202	0.03	0.70	-7.85	0.62	ILO
Cost of employees/Revenue	4,617,749	27.33	19.24	0	100	Amadeus
$\Delta\tau$ (cumulative)	5,673,960	2.08	1.06	0	4.01	Baccini et al (2018)
$\Delta\tau$ (placebo)	5,019,723	0.00	0.00	0	0	Baccini et al (2018)
$\Delta\tau$ (export)	5,673,960	0.27	0.35	0	7.01	Baccini et al (2018)

Table 9: Productivity and trade liberalization

	ECM	
	TFPR (f.d.)	
VARIABLES	(1)	(2)
$\Delta\tau$	-0.00	0.01
	(0.012)	(0.009)
ΔCME	-0.09**	-0.17**
	(0.003)	(0.002)
$\Delta\tau$*ΔCME	**-0.13****	**-0.15****
	(0.012)	**(0.010)**
τ	-0.23**	-0.18**
	(0.029)	(0.025)
CME	-0.12**	-0.27**
	(0.003)	(0.002)
τ *CME	**0.08****	**0.08****
	(0.009)	**(0.008)**
TFPR (lagged)	-0.32**	-0.37**
	(0.001)	(0.003)
Long-term multiplier	**0.25****	**0.22****
	(0.01)	**(0.01)**
Constant	0.43**	0.90**
	(0.014)	(0.018)
Observations	3,326,937	3,012,646
R-squared	0.162	0.208
Controls	No	Yes
Year FE	Yes	Yes
Country FE	Yes	Yes
Industry FE	Yes	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by firms in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. The outcome variable in all models is TFPR. Sources: Amadeus dataset, Baccini et al. (2018), and Visser (2016).

Table 10: Alternative measures of labor frictions

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				OLS			
				ln Revenue			
TFPR*Δτ*Union Density	**0.00**						
	(0.001)						
TFPR*Δτ*Centralization		**-0.08**					
		(0.232)					
TFPR*Δτ*Govt. Intervention			**-0.07****				
			(0.017)				
TFPR*Δτ*Sectoral Organiz.				**-0.02**			
				(0.024)			
TFPR*Δτ*Authority of Union over Local Branches					**-0.05**		
					(0.158)		
TFPR*Δτ*Authority of Confederation over its Affiliates						**-0.48****	
						(0.138)	
TFPR*Δτ*Mandatory Extension of Collective Agreements to Non-organised Employers							**-0.15****
							(0.014)
Constant	5.06**	4.98**	4.07**	1.69**	3.87**	4.32**	4.27**
	(0.033)	(0.083)	(0.026)	(0.039)	(0.054)	(0.046)	(0.027)
Observations	2,897,046	2,470,583	4,032,150	3,956,669	3,934,890	3,934,890	4,043,566
R-squared	0.782	0.780	0.766	0.768	0.769	0.769	0.767
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by firms in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. The outcome variable in all models is the log of revenue. Double interaction terms and single terms are included, but not reported. Sources: Amadeus dataset, Baccini et al. (2018), and Visser (2016).

Table 11: Alternative mechanisms

	(1)	(2)	(3)	(4)
	OLS			
	ln Revenue			
TFPR*Δτ*CME	**-0.05***	**-0.05****	**-0.05****	**-0.003**
	(0.027)	**(0.018)**	**(0.017)**	**(0.017)**
TFPR*Δτ*Labour Flexibility	**-0.04**			
	(0.03)			
TFPR*MFN*CME		**0.0003**		
		(0.001)		
TFPR*Imports*CME				**-0.004****
				(0.0003)
Constant	-15.35**	-7.15**	7.05**	0.34
	(0.567)	(0.260)	(0.219)	(0.26)
Observations	4,053,929	4,053,929	4,053,929	4,053,929
R-squared	0.754	0.765	0.763	0.768
Controls	Yes	Yes	Yes	Yes
Control & Imports	No	No	Yes	No
Year FE	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by firms in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. The outcome variable in all models is the log of revenue. Double interaction terms and single terms are included, but not reported. Sources: Amadeus dataset, Baccini et al. (2018), Visser (2016), and Comtrade (2018).

Table 12: Alternative measures of productivity

	(1)	(2)	(3)	(4)	(5)
	OLS				
	ln Revenue				
Labour Product*Δτ*CME	**-0.002****				
	(0.001)				
TFP*Δτ*CME		**-0.01****			
		(0.003)			
TFPR*Δτ*CME (Olley and Pakes)			**-0.001***		
			(0.0003)		
TFPR*Δτ*CME (Levinsohn and Petrin)				**-0.001***	
				(0.0003)	
TFPR*Δτ*CME (Wooldridge)					**-0.001***
					(0.0003)
Constant	6.18**	4.89**	4.06***	3.70**	3.69**
	(0.006)	(0.037)	(0.042)	(0.043)	(0.046)
Observations	4,008,342	2,321,574	1,806,661	1,806,661	1,806,661
R-squared	0.993	0.818	0.876	0.838	0.877
Controls	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by firms in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. The outcome variable in all models is the log of revenue. Double interaction terms and single terms are included, but not reported. Sources: Amadeus dataset, Baccini et al. (2018), and Visser (2016).

Table 13: Alternative model specifications

	(1)	(2)	(3)	(4)
	OLS			
	ln Revenue			lnProfit
TFPR*Δτ*CME	**-0.05****	**-0.09****	**-0.05***	**-0.02****
	(0.016)	**(0.020)**	**(0.022)**	**(0.001)**
ln Revenue (lagged)			0.47**	
			(0.004)	
Constant			2.52**	6.32**
			(0.299)	(0.510)
Observations	4,053,929	3,941,169	1,900,636	2,275,573
R-squared	0.765	0.882	0.820	0.306
Controls	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Double clustering (firms and industry)	Yes	No	No	No
Firm FE	Yes	Yes	No	No

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by firms and industry (Model 1) and by firms (Models 2-3) in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. Double interaction terms and single terms are included, but not reported. The outcome variable in all models is the log of revenue. Sources: Amadeus dataset, Baccini et al. (2018), and Visser (2016).

Table 14: Alternative measures of tariff cuts

	(1)	(2)	(3)	(4)
	OLS			
	ln Revenue			
TFPR*Δτ*CME (cumulative & weighted)	**-0.004***			
	(0.002)			
TFPR*Δτ*CME (cumulative & non-weighted)		**-0.003****		
		(0.002)		
TFPR*Δτ*CME (placebo)			**2.33e+09**	
			(3.499e+09)	
TFPR*Δτ*CME (export)				**-0.002****
				(0.000)
Constant	-6.56**	-5.76**	-7.37**	-6.45**
	(0.339)	(0.354)	(0.215)	(0.248)
Observations	4,053,929	4,053,929	3,966,589	4,053,929
R-squared	0.765	0.765	0.764	0.765
Controls	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05,

Note: OLS with robust standard errors clustered by firms in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. The outcome variable in all models is the log of revenue. Double interaction terms and single terms are included, but not reported. Sources: Amadeus dataset, Baccini et al. (2018), and Visser (2016).

Table 15: Analysis by country

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)
	OLS													
	ln Revenue													
	Austria	Belgium	Bulgaria	Cyprus	Czech Rep.	Germany	Denmark	Estonia	Spain	Finland	France	UK	Greece	Croatia
	CME=4	CME=5	CME=2	CME=2	CME=2	CME=4	CME=4	CME=1	CME=3	CME=4	CME=2	CME=1	CME=3	CME=2
TFPR*Δτ	**1.76***	**0.38****	**1.03****	**1.91***	**0.59****	**0.12**	**0.45**	**0.12**	**0.19****	**0.31***	**0.23****	**0.25****	**0.44***	**0.53**
	(0.71)	**(0.16)**	**(0.10)**	**(0.92)**	**(0.10)**	**(0.08)**	**(0.24)**	**(0.08)**	**(0.02)**	**(0.13)**	**(0.09)**	**(0.07)**	**(0.22)**	**(0.46)**
Observations	24,332	146,189	232,351	810	165,732	182,317	4,158	73,765	848,850	74,545	344,662	143,098	36,857	44,747
R-squared	0.92	0.51	0.69	0.62	0.51	0.93	0.81	0.78	0.86	0.8	0.92	0.82	0.75	0.64
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	No	No	Yes	Yes	Yes	Yes	No	No	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	No	Yes	Yes	Yes	Yes	Yes	No	Yes	Yes	Yes	Yes
	(15)	(16)	(17)	(18)	(19)	(20)	(21)	(22)	(23)	(24)	(25)	(26)	(27)	(28)
	OLS													
	ln Revenue													
	Hungary	Ireland	Italy	Lithuania	Luxembourg	Latvia	Malta	Netherlands	Poland	Portugal	Romania	Sweden	Slovania	Slovakia
	CME=1	CME=2	CME=3	CME=1	CME=3	CME=1	CME=2	CME=4	CME=1	CME=2	CME=3	CME=4	CME=3	CME=2
TFPR*Δτ	**0.21**	**0.11**	**0.32****	**0.29****	**-0.03**	**1.09****	**1.79***	**0.06**	**-0.01**	**0.27****	**0.54****	**0.71****	**0.35****	**0.60****
	(0.16)	**(0.13)**	**(0.06)**	**(0.10)**	**(0.68)**	**(0.16)**	**(0.75)**	**(0.15)**	**(0.04)**	**(0.06)**	**(0.04)**	**(0.13)**	**(0.06)**	**(0.13)**
Observations	236,654	5,822	889,828	17,096	978	86,397	1,380	9,745	65,285	394,769	647,420	303,197	56,332	97,998
R-squared	0.77	0.91	0.7	0.84	0.69	0.65	0.77	0.82	0.86	0.6	0.53	0.73	0.9	0.62
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	No	No	Yes	Yes	Yes	Yes	No	No	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	No	Yes	Yes	Yes	Yes	Yes	No	Yes	Yes	Yes	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by firms in parentheses. Unit of observation is firm-industry (4-digit NAICS)-country-year. The outcome variable in all models is the log of revenue. *CME* reports the median value for each country between 2003 and 2016. Single terms are included, but not reported. Sources: Amadeus dataset, Baccini et al. (2018), and Visser (2016).

Table 16: Demand of Redistribution: Identification Test

	OLS				
	Support for Redistribution				
	(1)	(2)	(3)	(4)	(5)
Instrument for PRF Liberalization	0.05**	0.06**	0.07**	0.06**	0.06**
	(0.017)	(0.017)	(0.017)	(0.018)	(0.018)
CME	-0.01	-0.06*	-0.36	-0.01	-1.03
	(0.014)	(0.026)	(0.234)	(0.016)	(0.630)
Instrument for PRF Liberalization*CME	**-0.03****	**-0.03****	**-0.03****	**-0.03****	**-0.03****
	(0.008)	**(0.008)**	**(0.008)**	**(0.009)**	**(0.009)**
Constant	0.06	1.53**	0.75**	0.57**	-0.42
	(0.062)	(0.094)	(0.186)	(0.106)	(0.485)
Observations	189,847	189,847	141,833	184,877	137,883
R-squared	0.076	0.076	0.079	0.077	0.081
Controls*CME	Yes	Yes	Yes	Yes	Yes
Region FE	Yes	Yes	Yes	Yes	Yes
Wave FE	Yes	Yes	Yes	Yes	Yes
Corruption	Yes	No	No	No	No
Unemployement	No	Yes	No	No	No
Market structure	No	No	Yes	No	No
Access to credit	No	No	No	Yes	No
All	No	No	No	No	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by region in parentheses. Unit of observation is respondent-region-country-wave. The outcome variable in all models is a dummy scoring one if respondents answer strongly agree or agree to the following sentence: *The government should take measures to reduce differences in income levels.* Sources: Amadeus dataset, Baccini et al. (2018), Visser (2016), and ESS (2018).

Table 17: Demand of Redistribution: Mechanisms

	OLS		
	Support for Redistribution		
	Low Income	High Income	Whole Sample
VARIABLES	(1)	(2)	(6)
Instrument for PRF Liberalization	0.06*	0.00	0.03
	(0.024)	(0.026)	(0.023)
CME	0.02*	-0.02	-0.00
	(0.011)	(0.020)	(0.010)
Instrument for PRF Liberalization*CME	**-0.03***	**-0.00**	**-0.03***
	(0.012)	**(0.013)**	**(0.011)**
Constant	0.67**	0.82**	0.73**
	(0.046)	(0.080)	(0.044)
Observations	77,462	61,264	189,847
R-squared	0.057	0.080	0.076
Controls*CME	Yes	Yes	Yes
Including other instr. of PRF liberal.	No	No	Yes
Region FE	Yes	Yes	Yes
Wave FE	Yes	Yes	Yes

Robust standard errors in parentheses ** p<0.01, * p<0.05

Note: OLS with robust standard errors clustered by region in parentheses. Unit of observation is respondent-region-country-wave. The outcome variable in all models is a dummy scoring one if respondents answer strongly agree or agree to the following sentence: *The government should take measures to reduce differences in income levels.* Sources: Amadeus dataset, Baccini et al. (2018), Visser (2016), and ESS (2018).