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LABOR, GLOBALIZATION AND INEQUALITY: ARE TRADE UNIONS STILL REDISTRIBUTIVE?

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ABSTRACT

Purpose – *Ascertaining the extent to which the generalized decline in union density, as well as the erosion in centralized bargaining structures and developments in other labor institutions, have contributed to rising within-country inequality.*

Methodology – *Econometric analysis of a newly developed dataset combining information on industrial relations and labor law, various dimensions of globalization, and controls for demand and supply of skilled labor for 51 Advanced, Central and Eastern European, Latin American, and Asian countries from the late 1980s to the early 2000s, followed by an analysis of 16 advanced countries over a longer time frame (from the late 1970s to the early 2000s).*

Findings – *In contrast to previous research, which finds labor institutions to be important determinants of more egalitarian wage or income distributions, the chapter finds that trade unionism and collective bargaining are no longer significantly associated with within-country inequality, except in the Central and Eastern European countries. These findings are interpreted as the result of trade unionism operating under*

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more stringent structural constraints than in the past, partly as a result of globalization trends. In addition, despite much talk about welfare state crisis, welfare states, historically the result of labor's power and mobilization capacity, still play an important redistributive role, at least in advanced countries.

Practical implications – *Union attempts at equalizing incomes by compressing market earnings seem ineffective and impractical in the current day and age. Unions should seek to increase the workers' skill levels and promote an egalitarian transformation of the workplace. This type of "supply-side" egalitarianism is not a new strategy for unions, but is very much embedded in the unions' DNA.*

Keywords: Inequality; trade unions; collective bargaining; labor law; industrial relations; globalization

Globalization promises to increase standards of living for all by bringing about greater specialization and higher productivity, cheaper goods and services, better access to credit and capital, and quicker diffusion of technological innovation. At the same time, there is growing concern in international policy circles, as well as among citizens, that globalization as we have seen it so far is not working (Goldberg & Pavcnik, 2007, p. 39; Wade, 2004), and increasing suspicion that its benefits only accrue to a small portion of the population, the very rich, while others gain little from it, except greater anxiety and a growing sense of precariousness (Luebker, 2004). It is also feared that the adverse distributional consequences of globalization may lead to a political backlash against it, and even to its undoing (Berger, 2000; OECD, 2007; Scheve & Slaughter, 2007; Scheve & Slaughter, 2004), and that this may bring back some form of economic protectionism, if not worse.

These concerns are not to be taken lightly: the first wave of globalization created an ever greater degree of economic integration across countries than is currently the case, e.g. as far as migration flows were concerned (Berger, 2003; O'Rourke, 2001). Yet it collapsed and subsequently gave way not only to economic protectionism but, more important, to fascist regimes in some countries. One of the reasons why the first globalization failed was the inability of governments to solve the "Polanyi problem," namely the problem of adequately managing the social disruptions associated with unfettered economic competition and a global free-market economy (Munck, 2004; Polanyi, 1957).

It has been argued repeatedly, including by the International Labor Organization (ILO) (2004), that to be sustainable and to bring positive outcomes for all, globalization needs a new international regulatory framework, and that this requires the introduction of an appropriate governance structure at the international level. However, as there is no consensus on how exactly to proceed on these matters, few concrete steps have been taken in this direction, and few will be taken, in all likelihood, in the foreseeable future. As a consequence, the international governance regime is and probably will remain for some time under-institutionalized, and the task of protecting societies from the potentially undesirable consequences of globalization will still fall heavily, if not exclusively, on national-level institutions – however weakened these may be at the moment. This chapter focuses on some of these institutions, those that have to do with workers' rights, trade unionism and collective bargaining.

The research question driving the chapter is whether industrial relations institutions contribute to reducing inequality in the current globalization era, and, if so, to what extent. We know from previous research on advanced countries that trade unionism and collective bargaining have redistributive effects. The chapter seeks to ascertain whether such inequality-reducing effects are still present when the analysis considers a more recent time period (the 1990s and early 2000s) than in previous studies, and includes developing countries as well.

There is reason to suspect that the same institutions that once improved earnings and income distributions may have recently become much less apt at doing so. Indeed, if one of the effects of globalization is to increase competition among firms and workers, for example by increasing product and labor demand elasticities (OECD, 2007, pp. 130–7; Rodrik, 1997; Scheve & Slaughter, 2004), such that firms cannot afford to deviate from market outcomes without running a serious risk of going out of business, and workers (particularly low-skilled) cannot push for wages much different from the ones that would prevail in a competitive equilibrium without jeopardizing their jobs, the impact of unions and collective bargaining on distributional outcomes is likely to be reduced.[1]

A number of developments suggest that labor institutions may have lost much of their redistributive potential in recent times (Pinto & Beckfield, 2010). One of these developments is the emergence in several countries, predominantly but not exclusively European, of a particular kind of centralized collective bargaining, known as social pacts, apparently similar to past models of centralized bargaining as far as institutional form is concerned, but rather different in outcomes, and, specifically, much more focused on national competitiveness than on redistribution (Baccaro & Lim, 2007; Berger & Compston, 2002; Fajertag & Pochet, 1997; Fajertag & Pochet, 2000; Hassel, 2003; Rhodes, 1996; Rhodes, 2001; Streeck, 2000). Other suggestive evidence comes from the recent shift in union wage policies: in several advanced countries trade union confederations no longer explicitly seek the compression of wage differentials as they did in the past, but have moved to more distributionally neutral wage policies (Baccaro & Locke, 1998; Edin & Holmlund, 1995; Schulten, 2002). Even in a country like Sweden, often considered a beacon of egalitarian capitalism, very high trade union density and a relatively centralized collective bargaining structure despite a recent shift from the national to the industry level (Pontusson & Swenson, 1996 ; Swenson & Pontusson, 2000) have not prevented inequality from growing considerably in the past few years (Atkinson, 2008; Bjorklund & Freeman, 2008 ; Gustavsson, 2007, pp. 85–7; Smeeding, 2002).

Addressing the question whether labor institutions still reduce inequality in the current era presents considerable empirical challenges. Country estimates of inequality are often based on different income concepts, population coverage, age coverage, etc., thus making both cross-sectional and longitudinal comparisons problematic (Atkinson & Brandolini, 2001). In addition, unlike advanced countries for which full time series data on union density and collective bargaining structures are available,[2] data on labor rights and industrial relations institutions for nonadvanced countries are sparse to say the least. In this chapter I collect the available evidence from various sources and make an effort to fill some data voids. Based on the availability of trade union, inequality, and other data, I end up focusing on 51 Advanced, Asian, Central and Eastern European, and Latin American countries. The time frame of the analysis is 1989–2005.

The main findings of the chapter are as follows: wile trade union density has been declining in almost all countries since the late 1980 at the same time as inequality has been increasing, the former does not seem causally associated to the latter. Controlling for various dimensions of trade and financial globalization, as well as human capital stock and a proxy for technologically induced shifts in the demand for skilled labor, union density and other institutional features are never significantly associated with the within-country variation in inequality from the late 1980s to the early 2000s, with the exception of the Central and Eastern European countries, where the sudden collapse of state-controlled trade unionism seems to have been one of the determinants of growing inequality (see Golden & Wallerstein, 2010 for similar conclusions). However, while there is no longitudinal association, there is a strong and robust cross-sectional association between labor institutions and inequality, indicating that historically two pillars of labor power – a higher proportion of wage and salaried workers organized by trade unions and a more centralized or coordinated collective bargaining structures – have produced societies that are on average more equal than others. A more in-depth analysis of advanced countries, conducted over a longer time frame, suggests that beginning with the 1990s labor institutions like trade union density, collective bargaining coverage, and particularly centralized collective bargaining, have become less effective in reducing inequality than they once were. Different from industrial relations institutions, a large welfare state remains instead highly redistributive, at least in advanced countries, and its redistributive effect does not seem to have changed over time.

The remainder of the chapter is organized as follows. I begin by reviewing the literature on the impact of labor institutions on inequality. I then move to a data section, which provides descriptive trends. The fourth section investigates, through an econometric analysis including all 51 countries, the linkages between income inequality and various labor institutions (ratification of core ILO conventions, respect of freedom of association and collective bargaining, unionization rates, and a more or less centralized or coordinated collective bargaining structure). The fifth section focuses on advanced countries and examines whether the impact of industrial relations institutions on inequality has changed over time by considering a longer time frame (1978–2002). I then summarize key findings from the analysis. I conclude by discussing policy implications.

TRADE UNIONS, COLLECTIVE BARGAINING, AND INEQUALITY: A REVIEW OF THE LITERATURE

In a recent literature review Richard Freeman, one of the key scholars in this field, argues not only that unions and collective bargaining improve the income distribution, but also that this is the only robust finding in the large literature on the effects of labor institutions on outcomes: ''For all of the difficulties in pinning down the impact of institutions on aggregate economic performance across countries, analyses have found that institutions have a major impact on one important outcome: the distribution of income'' (Freeman, 2007b, pp. 19–20).

Yet, what now seems (almost) received wisdom was a controversial statement only a few years ago. In his influential *Capitalism and Freedom*, Milton Friedman (1962, p. 124), for example, articulated a powerful argument why unions, far from acting as a "sword of justice" (Flanders, 1970; Metcalf, Hansen, & Charlwood, 2001), had anti-egalitarian distributional consequences:

> If unions raise wage rates in a particular occupation or industry, they necessarily make the amount of employment available in that occupation or industry less than it otherwise would be – just as any higher price cuts down the amount purchased. The effect is an increased number of persons seeking other jobs, which forces down wages in other occupations. Since unions have generally been strongest among groups that would have been high-paid anyway, their effect has been to make high-paid workers higher paid at the expense of lower-paid workers.

According to Friedman's argument, unions create inequality between two identical workers by pushing up wages in the union sector, and by depressing wages in the non-union sector (due to the increased supply of those who cannot find jobs in the unionized sector). If the workers are not identical, but, as Friedman believes, those organized in unions are more highly skilled, then unions contribute further to inequality by pushing up the skill premium relative to what it would be.

In a classic study on the effect of unionism in the United States using microdata, Freeman and Medoff (1984, chapter 5) reversed this argument. They showed that the effect of unions was theoretically ambiguous (see also Gottschalk & Smeeding 1997, p. 647), as unions, as argued by Friedman, did push up the wages of their members relative to nonmembers, but that this "monopoly" (or "between") effect was empirically dominated by three additional inequality-reducing effects: the dispersion of earnings within establishments was lower in union than non-union establishments, the dispersion across establishments was also lower (due to coordinated wage policies implemented by unions in collective bargaining), and the skill premium (between blue-collar and white collar workers) was lower in unionized establishments. Because the union wage premium benefited blue-collar workers more than others, the monopoly effect operated in the opposite direction from the one hypothesized by Friedman: it reduced inequality rather than increase it. As to mechanisms, the authors pointed to two in particular: (1) unions are democratic organizations, whose policy decisions may be expected to reflect the preferences of the median union member. If the median member is less skilled, and therefore less well-paid, than the average worker, the union will implement redistributive wage policies that reduce the skill premium; (2) union wage policies attach wages to occupations, not to workers based on supervisors' assessments, and since the distribution of occupations is probably less disperse than the distribution of supervisors' assessments of workers, union establishments have lower within-group dispersion than non-union establishments.

Twenty years after Freeman and Medoff (1984), these empirical findings still appeared very solid, having been corroborated by numerous subsequent studies (see Freeman 2007c for a review). For example, Card, Lemieux, and Riddell (2007) conducted a similar analysis to Freeman and Medoff (1984) based on microdata for three countries: the United States, Canada and the United Kingdom, all characterized by a divide between union and non-union sectors. They found that the dispersion of wages was lower for union workers than non-union ones within narrowly defined skill categories, thus confirming one of Freeman and Medoff's key results, and that, for male workers but not for female workers, unions also contributed to reducing the skill premium. The net effect was inequality-decreasing for men but not for women. For women, the inequality-increasing ''monopoly'' (or ''between'' effect) prevailed over the inequality-decreasing ''within'' effect. This divergence was due the different distribution of union membership between the two gender groups: while male union members were concentrated in the middle of the skill distribution, such that the ''monopoly'' effect boosted their wages relative to more skilled workers, female union members were positioned towards the top – also due to the fact that a higher proportion of female union members was in the public sector (Card, Lemieux & Riddell, 2007, p. 134). Interestingly, this analysis also revealed that the wage premium enjoyed by unionized workers over their nonorganized counterpart had declined between the early 1980s and early 2000, and that the ability of unions to compress the distribution of wages had also been declining over time (Card, Lemieux & Riddell 2007, p. 137 and 149–50). Overall, this analysis suggests that the impact of unionism on inequality is empirically dependent on whether the equalizing within-group effect prevails over the disequalizing between-group effect, which in turn depends on who the unions represent: if they predominantly represent the most skilled workers the net effect could be (as in Friedman's passage above and in the case of women in the United States, Canada, and the United Kingdom) to increase the dispersion of wages (for a recent analysis along similar lines see Becher & Pontusson, 2010). Also, according to this analysis the union impact on wages seems to be declining over time. In other words, unions seem less capable to affect both the level and the distribution of wages relative to a competitive scenario than they once were. We will return to this theme in the later analysis.

The work of Blau and Kahn (1996) has an important place in the literature because (to my knowledge) theirs is the only study in which the comparison relies on microdata on workers rather than on aggregate cross-section time-series data at the country level. The data they use come from various sources, but especially from the International Social Survey Program (ISSP). The authors examined 10 advanced countries in the mid- to late 1980s, and focused on differences between the United States and other countries. They found that the most important determinants of the greater dispersion in the bottom half of the wage distribution in the United States relative to other countries were institutional differences in wage-setting, and not demand and supply conditions. Focusing on the wage gap between two workers at the 50th and 10th percentile of the wage distribution, respectively, they found that while the difference in dispersion between the United States and the rest was not so great for the unionized sectors (union workers in the United States had almost the same degree of wage compression as in other countries), the dispersion of wages for non-union workers was much greater in the United States than in other countries. The authors interpreted this difference as due to institutional differences in the structure of collective bargaining which allowed unions to influence the wage structure of non-union workers to a much greater extent than in the United States, through various mechanisms like extension clauses, industry floors, or (given the greater power of unions outside of the United States) spontaneous adoption of union-contract provisions by non-union companies. In other words more centralized wage setting institutions in other countries brought about more wage compression than in the United States not so much among union members, but among workers that were not affiliated to trade unions. Consistent with these results, the authors also found that the union/non-union gap was greater in the United States than in other countries.

Partly as a result of the difficulty of collecting and standardizing microdatasets for a large number of countries, most comparative research on the determinants of inequality (especially the portion produced by noneconomists) takes the country/year as the unit of analysis. This approach exploits variation in union density rates and degrees of collective bargaining centralization across countries and/or within time to identify the effect of industrial relations institutions. In most cases it finds that institutions matter for inequality, but does not entirely agree as to exactly which institutions play the larger role. The major problem with this approach – which is also the one adopted in this chapter – is that, while it makes it possible to estimate net effects, it does not allow for an analysis of the different and possibly contradictory causal mechanisms by which unionization and collective bargaining impact inequality.

Wallerstein (1999) examined the effect of wage-setting institutions on earnings inequality in 16 OECD countries between (roughly) 1980 and 1992. This study used a rich dataset of institutional indicators pertaining to industrial relations features (measuring e.g. locus of bargaining, degree of government involvement in wage bargaining, degree of union confederation involvement in wage bargaining, internal concentration of union confederations, concentration across union confederations, etc.), which was developed by the author and two of his colleagues, and, repeatedly updated afterwards, were to become a *sine qua non* for quantitative comparative studies on industrial relations systems (Golden, Lange & Wallerstein, 2006). The author pooled observations across countries at three points in time, and estimated a model that had a measure of wage dispersion from the OECD Earnings Database as the dependent variable, several institutional predictors as independent variables (level of wage-setting, concentration between confederations, concentration between confederations, union density, and collective bargaining coverage), controlled for additional political and institutional determinants which might affect the distribution of earnings (political party orientation of government, government employment, government spending) and included a limited number of economic controls like trade exposure and measures of human capital supply. He found that the degree of collective bargaining centralization was by far the most important predictor of cross-country within-time differences in wage inequality, so much so that "it [was] difficult to find other variables that matter[ed] once the institutional variation in wage-setting [wa]s controlled for" (Wallerstein 1999, p. 650).

A similar study was performed by Rueda and Pontusson (2000), who examined the determinants of earnings inequality in the period between 1973 and 1995 in 16 OECD countries by using a dynamic model with country fixed effects and an instrumental variable approach (Anderson and Hsiao's estimator) to address the problem of the endogeneity of the lagged dependent variable. The model tested the effects of union density and collective bargaining centralization. Compared with Wallerstein's (1999) specification, this model went further in the attempt to control for economic conditions: it included controls for unemployment, trade with least developed countries and female labor force participation, and also included share of government employment, and government partisanship as institutional predictors. The choice of a fixed effects estimator allowed an exclusive focus on within-country changes in earnings inequality, controlling for time-unchanging differences in the average level of inequality across countries. The theoretical set-up assumed that the effects of both economic and institutional effects varied systematically across different "varieties of capitalism" (Hall & Soskice, 2001), and were hypothesized to be potentially very different in "liberal" (the United States and other Anglo-Saxon countries) vs. "coordinated" market economies (Germany and Nordic countries). The econometric results suggested that trade union density was the only predictor whose within-country variation was unconditionally negatively correlated with earnings dispersion, whereas the effects of all other variables varied across regimes. Bargaining centralization, for example, had a much stronger negative effect on inequality in coordinated economies than in liberal ones.[3] In the end, Rueda and Pontusson (2000) agreed with Wallerstein (1999) that labor institutions reduced inequality, but put a greater emphasis on trade union density than collective bargaining structure.[4]

In a recent paper Koeniger, Leonardi, and Nunziata (2007) improved on previous analyses by considering the impact of a larger array of labor market institutions: not just collective bargaining structure and trade union density rates, but also employment protection, replacement rates of unemployment insurance, duration of unemployment insurance, and size of the tax wedge. For data on labor market institutions they relied on a database assembled by Stephen Nickell and Luca Nunziata, and used previously to analyze the impact of labor market institutions on unemployment in OECD countries (Nickell, Nunziata, Ochel, & Quintini, 2001). The data on earnings inequality came from the OECD database on earnings. Greater richness in institutional detail came at the expense of a smaller number of advanced countries included in the analysis: 11. The time frame was 1973–1998. The analysis sought to build on the Wallerstein (1999) analysis, which the authors referred to (incorrectly) as "the only previous longitudinal study of wage inequality and institutions" (Koeniger et al., 2007, p. 341).[5] As in Rueda and Pontusson (2000), the analysis focused on within-country changes. The basic theoretical intuition was that labor market institutions reduced wage inequality by improving the bargaining position of unskilled workers more than that of skilled workers, thus leading to compression of wage differentials. The models also controlled for trade- and technology-induced demand shocks, and for the supply of skills. The theoretical predictions were largely confirmed by econometric results, which showed that all institutional variables were negatively associated with wage dispersion, except collective bargaining coordination, which, depending on specification, often had a positive sign. The authors concluded that changes in institutions explained the trajectory of wage inequality within countries at least as well as economic variables. Some of the econometric results were counterintuitive, however. For example, the proxy for labor demand shifts favoring the more highly skilled appeared to reduce, not increase, wage inequality, whereas greater supply of skilled labor seemed associated with an increase, not a reduction, in inequality. As acknowledged by the authors, these unexpected coefficients may signal specification problems.

Within this literature, the work of Bradley and co-authors (Bradley, Huber, Moller, Nielsen, & Stephens, 2003) while similar in style and methodological approach to others, stands out because, unlike the studies reviewed earlier, which focus on earnings inequality only, it investigates both the determinants of inequality in market income, and the determinants of inequality in post-tax and transfer income. The dependent variables (market income and disposable income) are measured using aggregate microdata from the Luxembourg Income Study (LIS).[6] The LIS is a collection of country-based microdatasets, which are harmonized to increase their comparability both across countries and over time.[7] In this study the sample included 14 advanced countries. Most data points used in the analysis were between the early 1980s and the mid-1990s, placed at approximately five-year intervals from one another. The specifications included a number of controls for economic conditions. The institutional variables considered were the union density rate and collective bargaining centralization. Since the chapter's main focus was on partisan effects, the cumulative shares of Social Democratic and Christian Democratic parties in government were included among the predictors. Like Rueda and Pontusson (2000) the authors found that trade union density was a more important determinant of inequality in market earnings than collective bargaining centralization, and that while redistribution through taxes and transfer was substantial in all countries, including those, like the Anglo-Saxon countries, characterized by a minimalist welfare state (Esping-Andersen, 1990), it was greatest in countries where governments were dominated by social democratic parties. Interestingly enough, trade union density and collective bargaining coverage did not just determine market incomes, but were also statistically associated with the extent of redistribution through taxes and transfer. Indeed, the authors argued that, due to collinearity among institutional and political indicators, a model in which redistribution was a function of the partisan composition of governments was statistically indistinguishable from models in which the main institutions considered were trade union density and collective bargaining centralization, respectively. However, comparative historical institutional considerations (in Australia, for example, a strong labor movement did not managed to reduce inequality because of the lack of social-democratic political dominance) lead the authors to privilege the political specification. Based on the results of this chapter one may hypothesize that the effect of trade unions is not just on market earnings but also, indirectly, on post-tax and transfer redistribution. Strong trade unions may proxy for other political variables (e.g., social democracy and associated policies) which reduce inequality through other means than compression of market earnings.

All cross-country longitudinal studies on the relationship between industrial relations institutions and inequality reviewed so far are based a limited number of advanced countries. I was able to find only one exception to this exclusive focus on advanced countries: a paper by Calderon, Chong, and Valdés (2004) on the impact of labor market regulation on income inequality in 121 countries between 1970 and 2000. This chapter relies on various indexes of labor regulations, both de jure (by counting the cumulative number of ILO core conventions ratified by the country/year in question) and de facto. Most institutional information is drawn from an unpublished database assembled by Rama and Artecona of the World Bank (2002).[8] Another source of information used in this chapter is the cross-sectional dataset of Botero et al. (2003) on the legislative protection of employment, industrial relations and social security. Due to a concern that, given the long time period, labor institutions may respond endogenously to income inequality, the authors use a dynamic GMM estimator controlling for country and time effects. Despite the much larger sample size and inclusion in the analytical framework of a number of developing countries, the econometric results are in line with other studies. In particular, trade union density is found to improve income inequality. The number of core ILO conventions ratified does not seem to have an impact on inequality.

The research reviewed so far (see Table 1 for a summary) suggests that industrial relations institutions are important determinants of cross-country differences in inequality. Several studies find that high trade union density rate is associated with lower inequality. A centralized collective bargaining structure also seems associated with greater equality, but this effect seems less robust across studies. The effect of trade unions and collective bargaining is a net effect, i.e. the resultant of various forces some of which may operate at cross purposes. Indeed, as shown by microstudies, whether trade unions reduce or increase inequality depends

Table 1. Cross-Country Time-Series Studies of the Relationship between IR Institutions and Inequality.

	Dependent Variable	Country Coverage	Time Coverage	Estimator Used	Impact of IR Institutions
Wallerstein (1999)	Earnings Inequality	16 Advanced countries	1980–1992	FGLS, error correction model, with and without country effects	Significant negative coefficient for Level of Wage Setting
Rueda and Pontusson (2000)	Earnings Inequality	16 Advanced countries	1973–1995	Anderson and Hsiao, dynamic model with country effects	Significant negative coefficient for union density
Bradley et al. (2003)	Market Income Inequality; Post Transfer and Taxes Reduction in Inequality	14 Advanced countries	Early 1980s-mid-1990s (for most countries)	Pooled OLS with cluster-robust standard errors, no country effects	Significant negative coefficient for union density
Calderon, et al. (2004)	Income Inequality	121 Countries	1970–2000	System GMM (dynamic model with country and time effects)	Significant negative coefficient for union density; insignificant coefficient for ILO core conventions ratifications
Koeniger et al. (2007)	Earnings Inequality	11 Advanced countries	1973–1998	Panel-Weighted Least Squares, with country and time effects	Significant negative coefficient for union density

strongly on who the unions represent, and particularly on whether union members are on average more skilled than other workers (Becher & Pontusson, 2010). Also, trade unions not only directly affect market earnings, by compressing the wage distribution, but also indirectly affect final incomes by being associated with other institutional and political variables (e.g., employment protection and unemployment insurance institutions, social-democratic regimes and associated economic policies), whose effect is to redistribute disposable incomes through more progressive taxes and transfers. The analysis that follows examines whether these conclusions hold when one focuses on the most recent period (from the 1990s on), and covers not just advanced countries, but also Latin American, Central and Eastern European, and (some) Asian countries. I begin by discussing the data.

INCOME INEQUALITY DATA

Unlike most cross-country time-series studies, the analysis below examines the impact of industrial relations institutions on income, not earnings, inequality. This choice was dictated both by the available data and by a theoretical choice: whereas data on average earnings by industrial categories are available for a large number of countries through the United Nations Industrial Development Organization (UNIDO various years), these data only refer to the manufacturing, and a fortiori, formal sector. Since the net effects of institutions like trade unions and collective bargaining may vary depending on the size of the informal sector (see Heckman & Pagés, 2000 for an argument along these lines), I prefer a broader measure of the dependent variable to a measure of inequality based on between-industry dispersion of formal wages in the manufacturing sector.[9]

I rely on estimates of Gini coefficients from secondary databases, which collect national statistics. The problems of secondary databases have been discussed in an influential article by Atkinson and Brandolini (2001). Focusing on OECD countries, i.e. on those countries for which data should at least in theory be more reliable, Atkinson and Brandolini (2001) show that differences underlying the data collected in secondary databases (having to do with different income concepts, area coverage, population coverage, etc.) negatively affects the robustness of not just cross-sectional analyses, but also longitudinal analyses focusing on the evolution of inequality within countries.

To increase comparability of the data, not only within countries, but also as much as possible across countries, I adopted the following strategy:[10]

(1) I relied primarily on the World Institute for Development Economics Research of the United Nations University's World Income Inequality Database Version 2.b (UNU-WIDER, 2007).[11] This is the largest secondary database available, and also includes the latest update of World Bank's Deininger and Squire dataset (2004). The WIID2b database often has multiple observations for a given country/year. The criteria that were adopted to extract data from this database aimed to maximize the within-country comparability of data. For each country I generally extracted data which came from the same survey instrument. When, in rare cases, I selected data from different surveys within the same country, there was a clear indication in the country notes that the two instruments were compatible. This also means that the income concept was kept constant within countries.
(2) When, for a given country, data from multiple surveys were available, I selected the survey that maximized coverage of the 1989–2005 period. When there was a tie, the survey with the higher data quality assessment in WIID2b was selected.
(3) I complemented the WIID2b database with data from a limited number of regional databases. For some Latin American countries, I used the Socio-Economic Database for Latin America and the Caribbean [SEDLAC] (2008).[12] For Central and Eastern European countries I used UNICEF's TransMONEE database (UNICEF, 2008).[13] For a limited number of advanced countries, I used data from the LIS website (LIS, 2008),[14] which seemed to cover the selected time period better than the LIS data included in WIID2b. For two countries, Hong Kong and Turkey, I used data from the World Bank's PovCalNet database (World Bank, 2008).[15]
(4) Although it was impossible to make sure that all estimates referred to the same income concept, I sought to maximize cross-country comparability by selecting data that referred to the distribution of incomes.[16]

Fig. 1 plots the average demeaned Gini coefficients (subtracting the country means, in order to focus on the within-country variation) against time. Not surprisingly, the graph shows that inequality has been growing considerably in the countries considered in this analysis. Table 2 reports the distribution of data sources for the 51 countries included in the analysis.[17]

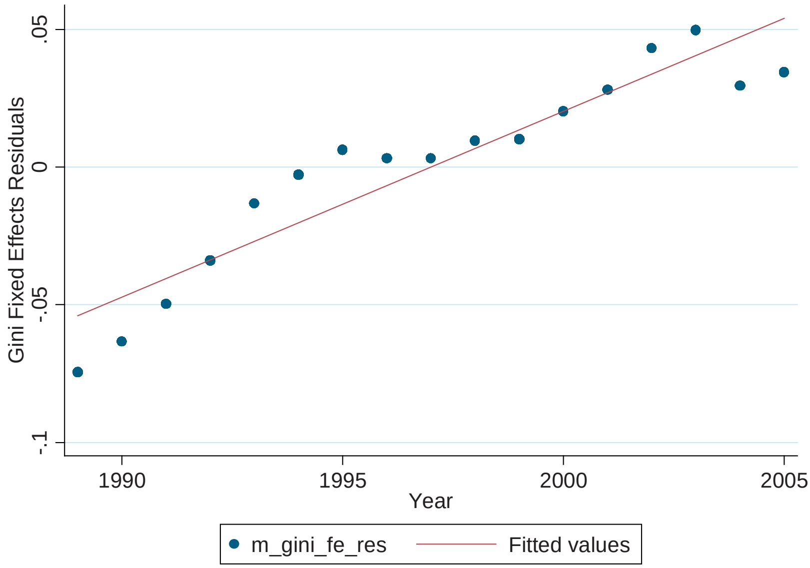

Fig. 1. The Trajectory of (Demeaned) Gini Coefficients Over Time.

Table 2. Sources of Data on Income Inequality (Gini Coefficients).

Database	Countries	Frequency	Percent
LIS	Australia, Austria, Belgium, Canada, Denmark, France, Italy, Spain, and Switzerland	105	16.88
PovCal	Hong Kong and Turkey	16	2.57
SEDLAC	Argentina, Brazil, Chile, Costa Rica, El Salvador, Honduras, Mexico, Paraguay, Uruguay, and Venezuela	155	24.92
TransMONEE	Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Slovak Republic, and Slovenia	123	19.77
WIID2b	China, Dominican Republic, Finland, Germany, Greece, India, Ireland, Jamaica, Japan, Korea, the Netherlands, New Zealand, Norway, Pakistan, Peru, Philippines, Portugal, Singapore, Sweden, Taiwan Province of China, the United Kingdom, and theUnited States	223	35.85
Total	51	622	100

MEASURES OF LABOR INSTITUTIONS

The labor institutions considered in the analysis are three: (1) trade union density, i.e. the percentage of wage and salaried workers that are affiliated to trade unions in a country in a given year; (2) collective bargaining structure, and particularly the degree to which collective bargaining is centralized or coordinated, i.e. either takes place at levels above the enterprise (e.g., at the industry or national level), or is coordinated through other mechanisms, including powerful and internally cohesive employer and worker organizations; and (3) labor law; specifically the degree this complies with international labor standards.

I relied extensively on the database assembled by Jelle Visser (2009) for Advanced and Central and Eastern European countries, which I complemented with data from various sources for Latin American and Asian countries.[18] Table 3 reports the sources of union density data. With the exception of seven countries (Singapore, Paraguay, China, Hong Kong, Spain, India, and Brazil) in which union density increased, and of three countries in which it did not change between 1989 and 2005 (Finland, Belgium, and Pakistan), in all other countries union density declined. The decline was dramatic in Central and Eastern European countries (Martin & Kaya, 2010): it was more than 50 percent in Hungary, Latvia, Czech Republic, Lithuania, and Estonia, which started from almost universal union affiliation in the Soviet years.[19]

For the index of collective bargaining structure I again rely extensively on the Visser database, which I complement with own research for other countries.[20] Visser's database provides an index of collective bargaining coordination, which in turn updates a previous index elaborated by Lane Kenworthy (2003). This 1-to-5 index is coded as follows:

1 = Fragmented wage bargaining, confined largely to individual firms or plants.
2 = Mixed industry- and firm-level bargaining, with little or no pattern-setting and relatively weak elements of government coordination such as setting of basic pay rate or wage indexation.

Table 3. Sources of Trade Union Density Data.

	Frequency	Percent
OECD.Stat	26	3.22
Visser 2008	438	54.21
Own research	344	42.57
Total	808	100

3 = Industry-level bargaining with somewhat irregular and uncertain pattern-setting and only moderate union concentration.
4 = Centralized bargaining by peak confederation(s) OR government imposition of a wage schedule/freeze, without a peace obligation OR informal centralization of industry- and firm-level bargaining by peak associations OR extensive, regularized pattern-setting coupled with a high degree of union concentration.
5 = Centralized bargaining by peak confederation(s) OR government imposition of a wage schedule/freeze, with a peace obligation OR informal centralization of industry-level bargaining by a powerful, monopolistic union confederation.

For the nonadvanced countries, however, often there was not enough information on the degree of coordination brought about by institutional features other than the structure of wage-setting. Therefore, for these countries the index is really an index of collective bargaining centralization, and the coding is simplified as follows (Golden et al., 2006):

1 = Plant-level wage-bargaining
2 = Mixed industry- and firm-level wage bargaining
3 = Industry-level wage bargaining
4 = Centralized wage bargaining without sanctions
5 = Centralized wage bargaining with sanctions.

It should also be added that most of the variation in this index is cross-sectional. This is not surprising, as the institutional structure of collective bargaining tends to be resilient over time, but may signal measurement error. Also, most of the within-country, longitudinal variation in the index is provided by the Advanced Countries. The index is entirely time-invariant for the Asian countries. Ireland emerges as the most coordinated country in the sample, closely followed by Norway. For 31 countries there is no apparent change in collective bargaining structure. For eight countries (Slovenia, Italy, Belgium, Finland, Hungary, Ireland, Portugal, and Spain) collective bargaining seems to become more coordinated/centralized. These are the countries that saw the emergence in the 1990s of social pacts. For 12 countries (Argentina, Peru, Slovak Republic, Sweden, Switzerland, Uruguay, Australia, Japan, Czech Republic, Estonia, Latvia, and Lithuania) the index signals a trend toward more decentralized/uncoordinated bargaining. There is a small trend toward bargaining decentralization in the whole sample.[21]

The third dimension of labor institutions considered in this analysis has to do with respect to international labor standards. I use three indicators: (1) the number of core ILO conventions ratified by a country in a given year;[22] (2) the number of Freedom of Association and Collective Bargaining core conventions (C87 and C98) ratified; and (3) unpublished violation severity scores elaborated and kindly made available to me by the OECD Secretariat.[23] The latter are based on the ILO Committee of Experts on the Application of Conventions and Recommendation's (CEACR) biannual reports on C87 and C98, the two core conventions on Freedom of Association and Collective Bargaining, respectively. For the countries that have ratified either convention, the CEACR writes a report every two years, which measures the distance between the norms contained in the convention and the de jure (and, to a lesser extent, also de facto) situation in each country. The OECD Secretariat coded the CEACR reports for a number of countries between 1990 and 1999 and elaborated two Violation Severity Indexes for C87 and C98, respectively.[24] Compared with the number of ratifications, these indexes (which are not available for all countries in the sample) tell us not just if a convention has been ratified, but also the extent of a country's compliance with the convention itself.[25] For C87 there is an increase in the severity of violations in the early 1990s and then a decrease. For C98 there seems to be a constant increase over time in the severity of violations.

GLOBALIZATION MEASURES

For all globalization measures and other economic controls (human capital and technologically induced demand for skilled labor), I rely on a database made available to me by the IMF Secretariat. This database has been used by the IMF for a recent report on globalization and inequality (IMF, 2007).[26]

The measures distinguish between trade and financial globalization. Trade globalization is operationalized through two indicators, one de facto and the other de jure: (1) trade openness, i.e., the sum of imports and exports (excluding oil-related transactions) over GDP and (2) de jure tariff openness, equal to 100 minus the tariff rate.[27] Financial globalization is also operationalized through a de facto and a de jure measure: (1) the ratio of inward FDI stock over GDP (Lane & Milesi-Ferretti, 2006);[28] (2) Menzie D. Chinn and Hiro Ito's measure of capital openness, capturing the extent of capital controls and based on the coding of information from the IMF's

Annual Report on Exchange Arrangements and Exchange Restrictions (AREAER) (Chinn & Ito, 2008).[29]

All measures of economic globalization display a marked growing trend in the period in question. On average, the countries in the sample have become more exposed to international trade, reduced the average tariff rate, increased the stock of FDI as percentage of GDP, and lowered de jure capital controls. Simultaneously, as shown in Fig. 1, union density rates have declined. The analysis below uses multiple regressions to parse out the respective contribution of the various factors to growing inequality.

THEORETICAL PREDICTIONS

Although the focus of the analysis is on the impact of labor institutions, and the globalization measures are used as controls, it is helpful to review briefly their expected effects (for recent reviews see Berger, 2000; Brady, Beckfield, & Zhao, 2007; Goldberg & Pavcnik, 2007; Guillen, 2001). According to the Stolper–Samuelson theorem, the consequences of trade openness should systematically differ across countries depending on their relative endowment of skilled vs. unskilled labor.[30] With trade openness, countries that are relatively rich in skilled labor should specialize in skilled-intensive productions, and vice versa. This should increase the effective demand for skilled labor and depress the demand for unskilled workers in skilled-endowed countries, and vice versa for countries rich in unskilled labor. To the extent that unskilled labor is the abundant factor in developing countries, and skilled labor in advanced countries, Stolper–Samuelson predicts that trade openness will reduce inequality (by compressing skill differentials) in developing countries, and increase inequality (by widening skill differentials) in advanced countries. This pattern is, however, incompatible with available evidence. Indeed, inequality has been growing in various developing countries exactly as their exposure to trade increased (Goldberg & Pavcnik, 2007, p. 55).

One argument about the effects of trade which is compatible with the current trend of growing inequality in both advanced and developing countries is the one advanced by Feenstra and Hanson (2001). This argument emphasizes that international trade does not just pertain to finished products, but also to intermediate products, and that one of the main features of globalization is the current international restructuring of production processes in global supply chains (Barrientos, 2007;

Gereffi, Humphrey, & Sturgeon, 2005). According to this model, firms in advanced countries outsource particular phases of the production process in developing countries. The outsourced activities are less skill-intensive from the point of view of developed countries, but relatively skill-intensive from the point of view of receiving countries. Thus, the effect of global production sharing is to shift labor demand away from unskilled workers and toward skilled workers in both developed and developing countries. Still another linkage between globalization and inequality has to do with the complementarity between capital and skilled labor (Acemoglu, 2002). To the extent that capital liberalization facilitates access to capital, it should cause an increase in the relative demand for skilled workers.

In theory, the impact of FDI on inequality should be similar to the Stolper–Samuelson prediction for trade: if FDI is attracted to a country because of relative abundance of a particular factor of production, then FDI in developing countries should increase demand for unskilled labor (the abundant factor) and lead to a more equitable distribution (Cornia, 2004; Vivarelli, 2004). However, there are also various channels by which FDI may worsen the distribution: one has been articulated by Feenstra and Hanson (2001): FDI may increase the demand for skilled labor in both advanced and developing countries, even if the transferred technology is neutral; the second is what Cornia (2004, p. 197) calls ‘‘systemic effect’’: in order to attract a greater share of FDI a country may relax a series of policy and regulatory constraints (e.g., concerning working conditions and taxation) which are associated with a more compressed income distribution.

Another channel by which globalization may affect inequality is by facilitating the transmission of skill-biased technological change from advanced to developing countries (Lee & Vivarelli, 2006, p. 7). Skilled-bias technical change increases both the relative price and relative quantity of skilled labor (Berman & Machin, 2004). If increased international competition forces companies to restructure and upgrade to defend themselves against competitors, or if the technology transferred with FDI is itself skill-biased, trade and financial liberalization may augment relative demand for skilled labor and increase inequality. In this case technological change would be an endogenous consequence of globalization.

Thus there are multiple channels by which different features of economic globalization may lead to greater within-country inequality. Some of these channels may operate at cross-purposes – for example if trade openness reduces inequality in a developing country by Stolper–Samuelson effects, while capital openness increases it – and net effects may vary country-by-country (Goldberg & Pavcnik, 2007). In a recent analysis of the impact of globalization on inequality, the IMF (2007) finds that while trade liberalization has contributed to reducing within-country inequality, financial globalization, and particularly a growing share of FDI liabilities over GDP, has increased it.

In addition to measures of the trade and financial dimensions of globalization, the econometric analysis reported below also controls for the degree of development of the credit market,[31] for human capital supply,[32] and for the technological intensity of production.[33] A more developed credit market may reduce income inequality by relaxing liquidity constraints on the less wealthy, i.e. by facilitating their access to credit. Similarly, a greater relative supply of skilled labor is likely to reduce inequality by reducing skill premia. Finally, the higher the (technologically induced) demand for skills, the higher the inequality, all other things being equal.

Among the institutional predictors, in addition to the ones discussed earlier (trade union density and collective bargaining coverage), various labor law-related indicators are also included: core convention ratification, and compliance with rights of association and collective bargaining. Although there is no guidance in the literature concerning their effects, they should operate in the same way as other institutions: to the extent that they strengthen the bargaining position of less skilled workers, or proxy for the government's favorable attitude toward redistribution, they should be associated with a more equal distribution. The econometric analysis below also controls for political regime, and specifically for political rights violations, by using the Freedom House indicator.[34] This is for two reasons: it is more than likely that the effects of trade unionism and collective are contingent on political regime: trade unions in nondemocratic countries (where membership may be compulsory or quasi-compulsory) may not redistribute as much as in democratic countries, and may not redistribute at all. Also, to the extent that in democratic regimes political parties are pushed by the logic of electoral competition to compensate increasing market inequality with redistributive taxes and transfers (Meltzer & Richard, 1981), countries with fewer political rights violations should have lower income inequality than others.[35] Table 4 summarizes the list of predictors included in the econometric analysis and theoretical expectations about their effects.

Table 4. List of Predictors and Expected Impact on Inequality.

Variable	Description	Expected Sign
Globalization measures		
FDI	Ratio of inward FDI stock over GDP	?
Tariff openness	100 minus the tariff rate	?
Capital account openness	Index capturing the extent of de jure capital controls	?
Trade openness	Sum of imports and exports (excluding oil-related transactions) over GDP	?
Other controls		
Average education	Average number of schooling years in the population aged 15+	–
ICT share	Stock of information and communication technology capital over total capita	+
Financial sector development	Ratio between private credit by deposit money banks and other financial institutions over GDP	–
Institutional measures		
Trade union density	Union membership over total wage and salary earners	–
Collective bargaining structure	Growing index of collective bargaining coordination/centralization	–
Core convention ratification	Number of ILO core conventions ratified	–
C87 severity index	Index capturing compliance with provisions in C87	–
C98 severity index	Index capturing compliance with provisions in C98	–
Reversed democracy index	Freedom House political liberty index	–

DO LABOR INSTITUTIONS REDUCE INEQUALITY? AN ECONOMETRIC ANALYSIS

The previous sections have argued that there is a clear growing trend in inequality in the countries considered, and simultaneously a clear declining trend in unionization, as well as a modest trend toward collective bargaining decentralization. The purpose of this section is to establish whether underneath this temporal coincidence between union decline and increasing inequality also lays a causal relationship. I begin by examining bivariate relationships.[36]

Bivariate Correlations

Fig. 2 plots demeaned Gini coefficient scores against demeaned union density scores. For each country/year the data have been expressed as deviations from country means. This allows one to focus on whether the change in union density is related to the change in Gini within countries.

The graph shows a negative association: the greater the decline in union density the greater the increase in inequality and (more rarely) vice versa. However, one should not conclude from this graph that the relationship is necessarily a causal one. This could be a spurious relationship, due to the fact that both variables are trended over time, and in opposite directions.[37]

The next graph focuses on the cross-sectional variation in the data: it abstracts from the way countries change over time and only considers their average values in the period. The graph shows again a clear negative correlation between unionization and inequality: the countries in which income inequality is on average lower in 1989–2005 tend to be the countries in which a greater proportion of wage and salaried workers is affiliated to

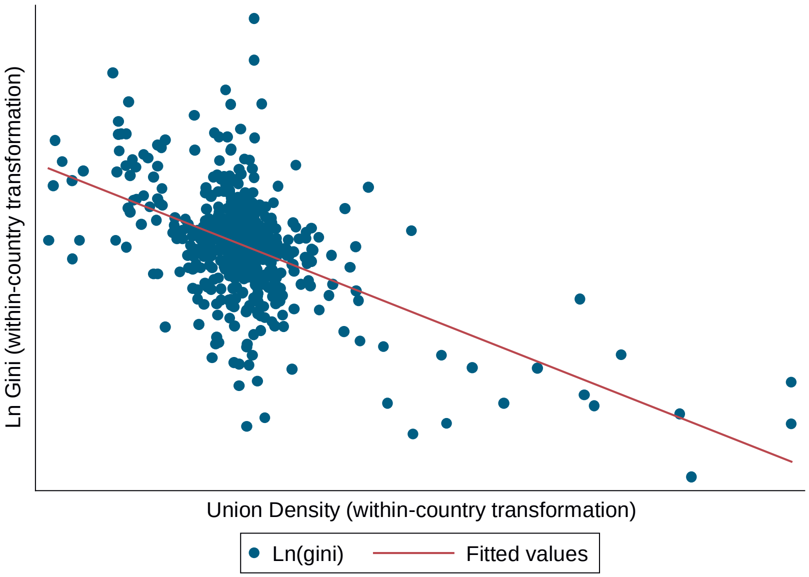

Fig. 2. Bivariate Relationship between Change in Gini Coefficient and Change in Union Density (Controlling for Country Averages).

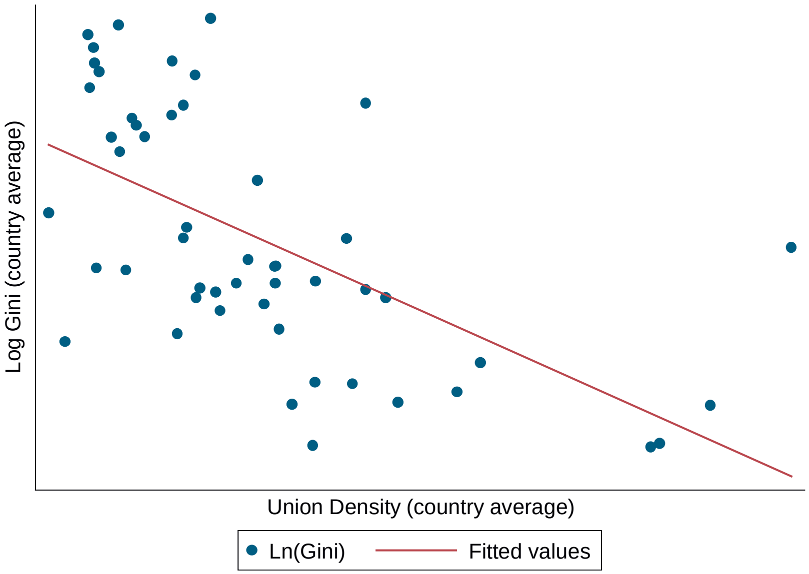

Fig. 3. Bivariate Relationship between Average Gini Coefficient and Average Union Density (1989–2005).

trade unions. Thus, bivariate graphs suggest that union density is negatively related to income inequality, both cross-sectionally and longitudinally (Fig. 3).

Another graph (not reported here) focuses on the structure of collective bargaining (combining cross-sectional and longitudinal variation in the data), and finds again a negative relationship with inequality: the more collective bargaining takes place at levels above the enterprise, the less unequal the distribution of income. If one distinguishes between cross-sectional and longitudinal variation in bargaining structure, one finds a strong negative cross-sectional but not longitudinal association. In other words, it is not the case that the more (less) collective bargaining becomes decentralized or uncoordinated, the more (less) inequality grows within a country. In fact the correlation between the two measures seems to be zero. As argued earlier, the countries in which the indicator of collective bargaining structure changes the most are the advanced countries. If, as argued earlier, historically centralized collective contributes to reducing inequality by limiting wage dispersion across sectors and skill levels and through other means, it seems that it may have forfeited these characteristics in recent years.

Concerning the relationship between core convention ratification and inequality, bivariate graphs not shown here suggest the following: there seems to be a small, negative relationship between the average number of core conventions ratified by a country and its average Gini coefficient. However, when one looks at the longitudinal relation (between change in core convention ratifications and change in inequality within countries over time), the slope of the curve is surprisingly positive.[38] This relationship is not only statistically very weak, but also in all likelihood spurious. It is probably due to the fact that both indicators, ratifications and inequality, tend to grow over time for unrelated reasons. At any rate, the bivariate associations suggest that ratification of core international labor conventions does not reduce income inequality.

More important seems the degree of compliance with the specific norms contained in C87 and C98. Figs. 4 and 5 plot average C87 and C98 severity index scores against average inequality, respectively, and reveal for both conventions a positive relationship: the more serious, on average, the violation of fundamental norms concerning freedom of association and collective bargaining, as assessed by the CEACR, the greater the average

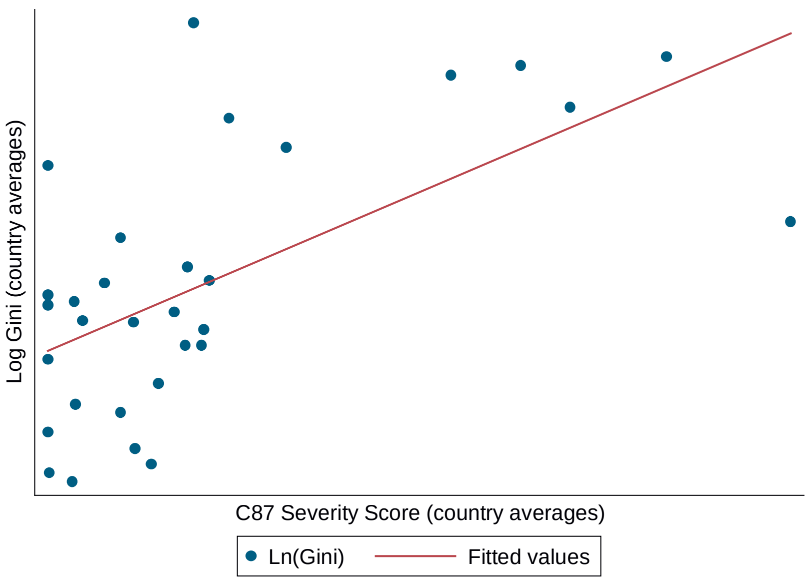

Fig. 4. Bivariate Relationship between Average Gini Coefficient and Average Convention 87 Violation Severity Score (1989–2005).

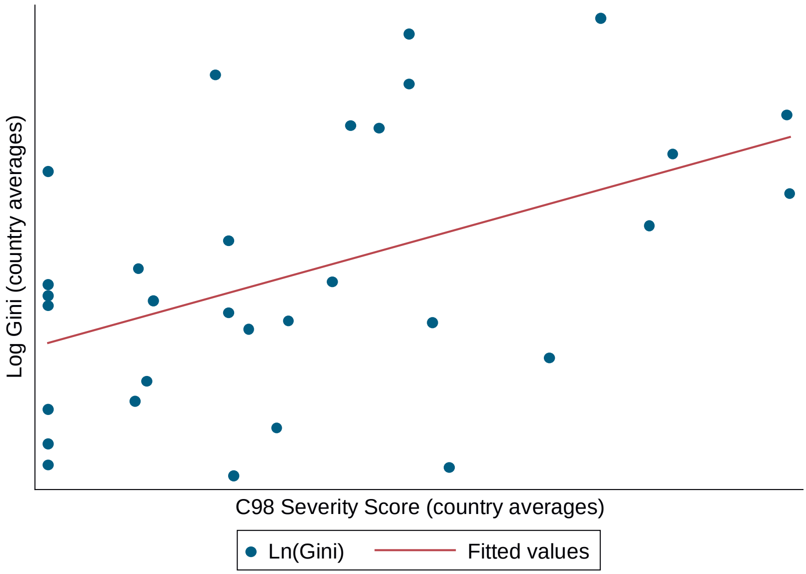

Fig. 5. Bivariate Relationship between Average Gini Coefficient and Average Convention 98 Violation Severity Score (1989–2005).

level of inequality in the country in question. The positive association seems steeper for C87 than for C98.

If one were to plot changes in Convention severity scores against changes in inequality within countries (not shown here), one would see for both C87 and C98 severity scores a much smaller positive relationship (larger for C98). Once again cross-sectional differences in institutions seem better associated with Gini coefficients than time changes.

The simple bivariate correlations discussed earlier suggest that labor institutions are important determinants of inequality, not so much across time (with the possible exception of the union density rate), as across countries. Cross-sectional differences in institutions are likely to reflect a constellation of factors that historically have led, either directly or indirectly, to a more compressed distribution of incomes. Indeed, labor institutions tend to come together as parts of a system: countries in which union density rates are higher are simultaneously countries in which where welfare states are more developed, taxation levels higher and more progressive, collective bargaining more centralized, labor law closer to international labor standards and better implemented.[39] What seems more surprising is that changes in these institutions seem less clearly associated with the increase in inequality. The next section examines whether this tentative conclusions holds when controlling for other potential determinants of income inequality through multivariate regression analysis.

Within-Country Regression Analysis

This session focuses on within-country changes. I estimate the following model:[40]

$$\ln(\text{gini}_{i,t}) = a + X_{i,t}\beta + Z_{i,t}\gamma + \delta_i + \tau_t + \varepsilon$$

where ln(gini) is the natural logarithm of the Gini coefficient in country i at time t; X is a vector of labor institutions variables, including the trade union density rate, the index of collective bargaining centralization/coordination, the number of core convention ratifications, the number of Freedom of Association and Collective Bargaining Core Conventions (C87 and C98), and the OECD indexes of C87 and C98 severity violations described earlier; Z a vector of economic and social controls, which includes the aforementioned measures of trade (trade openness, tariff liberalization) and financial globalization (FDI stock as percentage of GDP, capital account openness), as well as the average number of years of education in the country/year, credit by banks and other financial institutions. In separate specifications I also control for the share of ICT investment in total capital stock (a proxy for relative labor demand). The insertion of the δ_i (country dummies) allows for an exclusive focus on the time variation within countries. The time dummies (τ_t), capturing shocks affecting all countries simultaneously, seek to capture cross-sectional dependence in the errors and to account for the cyclical behavior (around a growing trend) of all the economic series presented earlier. Since the series are trended, it seems implausible that a shock (captured by the error term) is absorbed in only one year. For this reason I allow for first-order serial correlation in the errors:

$$\varepsilon_{i,t} = \rho\varepsilon_{i,t-1} + v_{i,t}$$

where $v_{i,t}$ is assumed to be i.i.d. and $|\rho| < 1$.

The econometric analysis reported below includes the following 42 countries for which there are data on all variables: Latin America and Caribbean (13 countries): Argentina, Brazil, Chile, Costa Rica, Dominican Republic, El Salvador, Honduras, Jamaica, Mexico, Paraguay, Peru, Uruguay, and Venezuela; Advanced Countries (21 countries): Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Turkey, the United Kingdom, and the United States; Central and Eastern European (CEE) Countries (2 countries): Hungary and Poland; and Asian Countries (6 countries): China, India, Republic of Korea, Pakistan, Philippines, and Singapore (6).[41] The time frame is 1989–2003, as the capital openness indicator is never available for 2004–2005. All variables, except Tariff Liberalization, Capital Openness, Union Density, and Collective Bargaining Structure are transformed to natural logarithms to increase the normalcy of their distribution.[42]

Columns 1–4 in Table 5 present the results of estimations in which the within-country variation in the Gini coefficient is solely a function of economic variables (globalization measures and controls). Column 1 includes FDI, the index of tariff liberalization, the index of capital account openness, average number of years of education, and a measure of development of the financial sector. Column 2 replaces the tariff-based measure of trade liberalization with a measure of trade openness. Column 3 tests whether trade openness has different impacts in advanced vs. developing countries, as suggested by the Stolper–Samuelson theorem (see Perry & Olarreaga, 2007), by introducing an interaction between the trade openness variable and a dummy that captures whether a country is advanced or developing. Column 4 estimates a Kuznets (1955) type of model by checking whether the trajectory of within-country inequality is different depending on levels of income and by introducing for this reason GDP and its square.

Of all economic controls, the only one that seems robustly associated with inequality is FDI stock as percentage of GDP: the greater the growth in FDI, the greater the increase in inequality within a country. FDI may play its effects through at least two channels: it may increase demand for skills in the receiving country at the same time as it decreased the relative demand for semi-skilled in the sending country (Feenstra & Hansonm 2001) – this is based on the assumption that FDI is low-skill for the sending country, for example in sectors like textile and apparel, while it is skill-intensive for the receiving country (IMF, 2007, p. 45). Also the need to attract FDI may induce a country to reduce taxes and adopt less redistributive social policies (Cornia, 2004). Of the other economic variables, tariff liberalization seems positively associated with inequality, while capital account liberalization, average education years and credit to the private sector are negatively signed. Generally, however, one can not reject the hypothesis of zero coefficients for these variables, with the exception of the tariff liberalization

Table 5. Determinants of Gini (Fixed Effects Models with AR(1) Errors), Intercept and Time Dummies Not Reported.

Dep. Var	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)
FDI	0.0243**	0.0209**	0.0215**	0.0275***	0.0237**	0.0260**	0.0263**	0.0263**	0.0264**	0.0266**
	(0.0101)	(0.0104)	(0.0105)	(0.0105)	(0.0104)	(0.0107)	(0.0106)	(0.0107)	(0.0106)	(0.0106)
Tariff liberalization	0.00133			0.00130	0.00147	0.00150	0.00183*	0.00190*	0.00184*	0.00196*
	(0.00102)			(0.00102)	(0.00106)	(0.00107)	(0.00108)	(0.00112)	(0.00109)	(0.00109)
Capital account openness	−0.00342	−0.00341	−0.00347	−0.00326	−0.00331	−0.00337	−0.00408	−0.00413	−0.00429	−0.00376
	(0.00338)	(0.00337)	(0.00338)	(0.00339)	(0.00348)	(0.00351)	(0.00351)	(0.00354)	(0.00352)	(0.00354)
Education years (average)	−0.256	−0.238	−0.239	−0.207	−0.212	−0.201	−0.194	−0.197	−0.203	−0.200
	(0.186)	(0.188)	(0.189)	(0.190)	(0.192)	(0.193)	(0.187)	(0.187)	(0.187)	(0.185)
Credit private Sector	−0.0118	−0.0108	−0.0107	−0.0123	−0.00956	−0.0102	−0.0106	−0.0109	−0.0115	−0.0106
	(0.0111)	(0.0111)	(0.0111)	(0.0111)	(0.0114)	(0.0115)	(0.0114)	(0.0114)	(0.0114)	(0.0113)
Trade openness		0.00831	0.0121							
		(0.0195)	(0.0209)							
Trade opennes advanced			−0.0208							
			(0.0409)							
GDP				−0.0412						
				(0.0432)						
GDP squared				0.00532						
				(0.00421)						
Union density (UD)					−0.0159	−0.0203	0.0526	0.0515	0.0529	0.0513
					(0.0628)	(0.0633)	(0.0749)	(0.0764)	(0.0751)	(0.0752)

Reversed democracy index						0.000228 (0.00473)	0.00154 (0.00479)	0.00163 (0.00484)	0.00168 (0.00480)	0.00129 (0.00481)
UD advanced							0.0123 (0.190)	0.0212 (0.190)	0.00715 (0.190)	0.0325 (0.190)
UD CEE							−0.356** (0.152)	−0.357** (0.150)	−0.353** (0.151)	−0.346** (0.151)
UD Asia							−0.231 (0.359)	−0.222 (0.362)	−0.218 (0.360)	−0.220 (0.361)
Collective bargaining structure								−0.00114 (0.00521)		
Core conventions ratified (#)									0.00295 (0.00425)	
C87 & C98 Ratified (#)										−0.0130 (0.0138)
Time dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	435	441	441	435	422	417	417	416	417	417
Number of countries	43	44	44	43	43	42	42	42	42	42
R^2 (within)	0.158	0.147	0.147	0.165	0.156	0.157	0.188	0.193	0.191	0.194
Estimated Rho	0.633	0.643	0.643	0.628	0.621	0.621	0.592	0.583	0.589	0.585

Notes: Standard errors in parentheses; $^{***}p<.01$, $^{**}p<.05$, and $^{*}p<.1$.
Dep. Var. – Ln(Gini) = Natural Logaritm(Gini Coefficient).

index which is often significant at the 10 percent level. Thus there is some evidence that the reduction of tariffs has contributed to increase income inequality in the past 15 years.

Columns 4–10 examine the impact of labor institutions controlling for other economic determinants. Column 5 examines the effect of union density. Column 6 adds the reversed democracy score (the higher the score, the more undemocratic the country/year in question). Column 7 examines possible heterogeneity in the impact of the unionization variable, and for this reason introduces specific terms for union density in Advanced Countries, Central and Eastern European Countries, and Asian Countries (the reference category is Unionization in Latin American Countries). Indeed, it is conceivable that in an economy characterized by a large informal sector, a high degree of organization of formal sector workers may increase income inequality, especially if trade union represent predominantly skilled workers (Heckman & Pagés, 2000). Column 8 controls for the impact of collective bargaining structure, the assumption being that a more centralized/coordinated collective bargaining structure tends to reduce inequality.[43] Column 9 checks whether an increase in the number of core convention ratifications has a significant impact on income inequality. Column 10 repeats the same analysis but with an exclusive focus on the two core conventions on Freedom of Association and Collective Bargaining (C87 and C98).[44]

The results of the analysis suggest that, generally speaking, changes in union density are not significantly associated with changes in income inequality in the period under investigation. However, if one distinguishes by region, one finds that in the Central and Eastern European countries, the (precipitous) decline in unionization after the collapse of the Berlin Wall seems to have significantly contributed to the increase in inequality.[45] Interestingly enough, while they are not significantly different from zero, the coefficients for unionization in Latin American and Advanced Countries are positive, not negative. The political freedom index is positive (indicating that the more political rights are violated, the greater inequality), but statistically insignificant. Also, there is no inequality-reducing effect of collective bargaining centralization/coordination: the coefficient is negative but statistically insignificant.[46] Finally, ratification of core conventions or FACB conventions is not significantly associated with inequality.[47]

Table 6 probes the previous results by re-estimating Model 8 in Table 5 after excluding one country at a time and examining the robustness of regression coefficients. The analysis confirms the previous conclusions: there seems to be a robust positive association with FDI, and a negative

Table 6. Jacknife Analysis of the Determinants of Gini Coefficients.

Variable	Full Model	Insignificant at 10% if Following Countries are Excluded	No. of Countries Insignificant	Max Value	Excluded Country	Min Value	Excluded Country
FDI	0.0263** (0.0107)	Korea	1	0.0367*** (0.00963)	China	0.0129 (0.0110)	Korea
Tariff liberalization	0.00190* (0.00112)	Austria, Belgium, Canada, China, Costa Rica, Ireland, Italy, Korea, Mexico, Norway, Pakistan, Poland, Singapore, Sweden, and the United Kingdom	15	0.00284** (0.00139)	Argentina	0.00124 (0.00100)	Pakistan
Capital account openness	−0.00413 (0.00354)	All	42	−0.00258 (0.00392)	Argentina	−0.00518 (0.00360)	Mexico
Education years (average)	−0.197 (0.187)	All except Venezuela	41	−0.0789 (0.200)	El Salvador	−0.755*** (0.271)	Venezuela
Credit private Sector	−0.0109 (0.0114)	All except China	41	−0.00397 (0.0114)	Finland	−0.0180* (0.0101)	China
Union density (UD)	0.0515 (0.0764)	All	42	0.0970 (0.126)	Mexico	0.0237 (0.0827)	Honduras
Reversed democracy index	0.00163 (0.00484)	All	42	0.00296 (0.00500)	Uruguay	−0.00131 (0.00505)	Mexico
UD advanced	0.0212 (0.190)	All	42	0.109 (0.193)	Finland	−0.0960 (0.168)	China
UD CEE	−0.357** (0.150)	Hungary	1	−0.274* (0.152)	Venezuela	−0.421** (0.185)	Poland
UD Asia	−0.222 (0.362)	All except China	41	0.352 (0.339)	Pakistan	−0.664* (0.391)	China
Collective bargaining structure	−0.00114 (0.00521)	All	42	0.000708 (0.00532)	Argentina	−0.00289 (0.00558)	Mexico

association with union density in Central and Eastern European Countries (Hungary and Poland). Tariff liberalization is less robustly associated with positive income inequality than FDI. Interestingly enough, if one removes Venezuela from the sample, a negative and highly significant association between education years and inequality emerges: the greater the supply of human capital (measured by average education years) the lower inequality, which is a priori what one would expect (see also Mahler, 2010).

Table 7 presents additional specification checks: since the Capital Openness and Education variables are not available for a number of CEE Countries, and they seem insignificant according to the previous analysis, they are removed from the econometric model in Table 7, Column 1, in order to appreciate the impact of union density for a greater number of CEE countries. The UD CEE coefficient now refers to a much larger sample of countries: Czech Republic, Estonia, Latvia, Lithuania, Slovak Republic, and Slovenia, in addition to Hungary and Poland. It remains negative, approximately of the same magnitude as before, and highly significant. Columns 2 and 3 introduce an important additional control: the share of information technology investment in the capital stock. This proxy captures technology-induced demand for skilled labor, and is only available for a subset of countries: specifically, for none of the CEE countries (so that estimation of the CEE-specific effect of union density now becomes impossible).[48] This proxy turns out to be a significant predictor of inequality: the higher the share of ICT, the higher inequality. Interestingly enough, the coefficient of FDI does not change much, while the coefficient of Tariff Liberalization becomes insignificant, suggesting that tariff liberalization operates by increasing the demand for skilled labor. Also, years of education emerges as a significant negative predictor of inequality. Column 3 distinguishes between the effect of collective bargaining structure in advanced countries and the rest, and finds an insignificant coefficient, which is, interestingly enough, positive and not negative, thus suggesting that increases in collective bargaining centralization in advanced countries tend to be associated with more inequality rather than less.[49]

Table 8 examines possible endogeneity on the right-hand side of the Gini equation, and specifically whether the reason why there is no significant effect of union density on income inequality, controlling for globalization forces, is that union density itself is affected by these globalization forces, so that its impact is captured by them. The results of two fixed effects model with AR(1) errors, where the dependent variable is unionization and within-country changes in unionization are regressed on globalization variables, suggest that countries in which the FDI stock increased as a percentage of

Table 7. Additional Specifications: Determinants of Gini (Fixed Effects Models with AR(1) Errors), Intercept and Time Dummies Not Reported.

Dep. var	(1)	(2)	(3)
	Ln(Gini)	Ln(Gini)	Ln(Gini)
FDI	0.0242**	0.0229**	0.0220*
	(0.0108)	(0.0115)	(0.0116)
Tariff liberalization	0.00184*	0.00167	0.00176
	(0.00111)	(0.00134)	(0.00135)
Capital account openness		−0.00651*	−0.00649*
		(0.00375)	(0.00375)
Education years (average)		−0.325*	−0.323*
		(0.187)	(0.187)
Credit private sector	−0.00684	−0.00693	−0.00682
	(0.0108)	(0.0120)	(0.0120)
ICT share capital (%)		0.163***	0.164***
		(0.0312)	(0.0311)
Union density (UD)	0.0351	0.0177	0.00400
	(0.0803)	(0.0765)	(0.0807)
Reversed democracy index	0.00328	−0.000932	−0.000876
	(0.00500)	(0.00501)	(0.00502)
UD advanced	−0.0199	−0.163	−0.157
	(0.194)	(0.194)	(0.194)
UD CEE	−0.373***		
	(0.138)		
UD Asia	−0.218	−0.169	−0.152
	(0.387)	(0.360)	(0.362)
Collective bargaining structure	−0.00190	−0.000725	−0.00722
	(0.00517)	(0.00541)	(0.0128)
Collective bargaining structure advanced			0.00794
			(0.0140)
Time dummies	Yes	Yes	Yes
Observations	485	355	355
Number of countries	49	35	35
R^2 (within)	0.176	0.231	0.234
Estimated Rho	0.564	0.575	0.572

Standard errors in parentheses; $^{***}p<.01$, $^{**}p<.05$, and $^{*}p<.1$.

GDP over the period are countries in which union density declined significantly. There is case study evidence on Ireland (a country in which FDI plays a key role) suggesting that as FDI flew to this country in the 1990s, MNCs (particularly American) increasingly refused to recognize trade unions (as they had done previously) and the public agency responsible for attracting FDI waived the union recognition requirement

Table 8. The Impact of Globalization on Union Density Rates (Fixed Effects Models with AR(1) Errors, Intercept and Time Dummies not reported).

Dep. Var	(1)	(2)
	Union Density	Union Density
FDI	−0.000930***	−0.000966***
	(0.000358)	(0.000359)
Tariff liberalization	0.000348	0.000385
	(0.000570)	(0.000572)
Capital account openness	0.00111	0.00108
	(0.00240)	(0.00240)
Trade openness		0.00680
		(0.0135)
Time dummies	Yes	Yes
Observations	564	564
Number of countries	43	43
R^2 (within)	0.0919	0.0950
Estimated Rho	0.714	0.708

Notes: Standard errors in parentheses; $^{***}p<.01$, $^{**}p<.05$, and $^{*}p<.1$.

for location grants (Gunnigle & McGuire, 2001 ; Roche & Geary, 1997). These examples suggest possible channels through which an increase in FDI may lead to lower unionization. Other facets of globalization (tariff liberalization, capital openness, and trade openness) do not seem to impact unionization significantly.

Table 9 re-estimates some of the models in Table 5 by dropping the FDI term and thus allowing the union density role to have potentially a greater impact on inequality, not mediated by FDI. Results do not change much, however. Both trade union density and other institutional variables remain insignificant predictors of inequality, again with the exception of trade union density in CEE countries.

In synthesis, the within-country econometric analysis suggests the following:

(1) An increase in the stock of FDI as percentage of GDP tends to be associated with greater inequality in the countries considered.
(2) Trade liberalization in the form of tariff reduction also seems to increase inequality, but less robustly than in the formed case.
(3) Other facets of globalization (capital openness and trade openness) do not seem to be significant predictors of income inequality.

Table 9. Determinants of Gini Coefficient, Excluding FDI (Fixed Effects Models with AR(1) Errors), Intercept and Time Dummies Not Reported.

Dep. Var	(1)	(2)	(3)
	Ln(Gini)	Ln(Gini)	Ln(Gini)
Tariff liberalization	0.00116	0.00150	0.00155
	(0.00105)	(0.00108)	(0.00112)
Capital account openness	−0.00287	−0.00359	−0.00359
	(0.00353)	(0.00353)	(0.00356)
Education years (average)	−0.197	−0.194	−0.194
	(0.200)	(0.190)	(0.190)
Credit private sector	−0.00980	−0.0107	−0.0113
	(0.0117)	(0.0115)	(0.0115)
Union density (UD)	−0.0106	0.0634	0.0589
	(0.0636)	(0.0750)	(0.0766)
Reversed democracy index	−0.00119	0.000303	0.000345
	(0.00469)	(0.00478)	(0.00483)
UD advanced		−0.00201	0.0109
		(0.192)	(0.192)
UD CEE		−0.380**	−0.385**
		(0.155)	(0.153)
UD Asia		−0.255	−0.245
		(0.360)	(0.363)
Collective bargaining structure			−0.00274
			(0.00520)
Time dummies	Yes	Yes	Yes
Observations	417	417	416
Number of countries	42	42	42
R^2 (within)	0.135	0.170	0.175
Estimated Rho	0.564	0.575	0.572

Notes: Standard errors in parentheses; $^{***}p<.01$, $^{**}p<.05$, and $^{*}p<.1$.

(4) Technology-induced shifts in the demand for skilled labor (captured by the share of ICT investment in the capital stock) tend to increase inequality.
(5) Changes in labor institutions within countries (trade union density, collective bargaining centralization/coordination, ratification of core conventions, respect of core labor standards) do not seem responsible for growing inequality over time, with the exception of trade union decline in the CEE countries, which seems to have contributed to the growth in inequality in that region.

Having examined how the change in labor institutions within countries has affected the change in inequality (within countries) in the past few years, the next section complements the previous analysis by looking at the cross-sectional association among variables.

Between-Country Regression Analysis

The goal of this section is to examine whether more institutionally dense countries (i.e. with a greater unionization rate, a more centralized collective bargaining system, a greater respect for political rights and core labor rights, etc.) have lower average levels of inequality controlling for various features of globalization.

Table 10 estimates essentially the same specifications as in Table 5, but focusing on the cross-sectional variation in the data. Columns 1 and 2 only contain economic controls. Columns 3–8 check for the impact of institutional predictors, allowing for a regionally differentiated impact of trade unionism (Columns 5–6), of collective bargaining structure (Column 7) and of both (Column 8).

The results of the between estimator are rather different from those of the within estimator. Differences in average levels of income inequality across countries seem to depend entirely on institutional differences, whereas the economic predictors are hardly ever statistically different from zero. The two exceptions are the measure of human capital, which (as expected) is negatively associated with inequality in the model with economic controls only (Table 10, Column 1), but whose coefficient declines dramatically in absolute value and becomes statistically insignificant once the institutional predictors are inserted, and the measure of FDI, which is positive but rarely significantly different from zero.[50]

Cross-sectionally, the countries in which trade union density is higher are those in which the income distribution is less unequal on average. Consistent with results from the within analysis, there seem to be regional differences in the impact of unionization. Greater union density in Latin American countries is not associated with lower inequality; on the contrary, the coefficient is positive, albeit insignificant. This may be due to the historical corporatist nexus linking trade unions to the state in some Latin American countries (Murillo, 2001 ; Zapata, 1998). Also, if trade unions represent predominantly skilled (e.g., public sector) workers, then the ''monopoly'' effect may empirically dominate the ''within'' effect, thus leading to a more unequal income distribution. Relative to Latin American countries, union

Table 10. Determinants of Gini Coefficients: Between Effects (Constant Not Reported).

Dep. Var	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)
FDI	0.0639*	0.0467	0.0648**	0.0397	0.0326	0.0231	0.0270	0.0253
	(0.0377)	(0.0458)	(0.0312)	(0.0295)	(0.0226)	(0.0231)	(0.0253)	(0.0240)
Tariff liberalization	0.00647	0.00735	0.00466	0.00446	0.00207	0.00257	0.00434	0.00214
	(0.00769)	(0.00819)	(0.00641)	(0.00560)	(0.00442)	(0.00434)	(0.00491)	(0.00468)
Capital account openness	−0.0470	−0.0725	−0.0469	−0.00747	−0.0192	−0.0103	−0.0237	−0.00575
	(0.0419)	(0.0501)	(0.0347)	(0.0320)	(0.0266)	(0.0266)	(0.0284)	(0.0279)
Education years (average)	−0.0566***	−0.0297	−0.0308	−0.00429	0.00288	−0.00240	−0.00342	−0.00667
	(0.0207)	(0.0275)	(0.0183)	(0.0176)	(0.0141)	(0.0142)	(0.0157)	(0.0149)
Credit private sector	−0.0862	−0.120	−0.0691	−0.0748	−0.00655	−0.0000386	−0.0109	0.0160
	(0.0666)	(0.0751)	(0.0559)	(0.0487)	(0.0554)	(0.0545)	(0.0518)	(0.0574)
ICT share capital (%)		−0.00865						
		(0.0129)						
Union density (UD)			−0.660***	−0.822***	0.348	0.462	−0.598***	0.421
			(0.161)	(0.149)	(0.322)	(0.323)	(0.143)	(0.489)
Reversed democracy index				0.0909***	0.0638**	0.0542*	0.0542**	0.0535*
				(0.0262)	(0.0268)	(0.0270)	(0.0261)	(0.0281)
UD advanced					−1.137***	−1.152***		−1.067**
					(0.283)	(0.277)		(0.490)
UD CEE					−1.707***	−1.716***		−3.835**
					(0.345)	(0.338)		(1.824)
UD Asia					−0.964**	−1.016***		−0.980*
					(0.372)	(0.366)		(0.535)
Collective bargaining structure						−0.0317	0.0715*	−0.0164
						(0.0213)	(0.0392)	(0.0545)
CB advanced							−0.114***	−0.0200
							(0.0358)	(0.0558)

Table 10. *(Continued)*

Dep. Var	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)	Ln(Gini)
CB CEE							−0.226***	0.395
							(0.0613)	(0.326)
CB Asia							−0.103	−0.00743
							(0.0625)	(0.0760)
Year	−0.0216	−0.0172	−0.000853	0.0257	0.0226	0.0220	0.0184	0.0229
	(0.0266)	(0.0338)	(0.0230)	(0.0217)	(0.0164)	(0.0161)	(0.0180)	(0.0170)
Number of countries	43	35	43	42	42	42	42	42
R^2	0.449	0.487	0.627	0.730	0.860	0.870	0.845	0.879

Notes: Standard errors in parentheses; $^{***}p<.01$, $^{**}p<.05$, $^{*}p<.1$.

density is associated with lower inequality in Advanced, CEE, and Asian countries.

The effects of collective bargaining structure also seem regionally specific: in Latin America a more centralized collective bargaining is associated with greater inequality, while the association is negative (relative to Latin American countries) in Advanced, CEE, and Asian countries. Overall, collective bargaining coefficients seem less robustly significant than union density rates. Interestingly, the more politically illiberal the government, the greater the inequality on average. This is not surprising and may be due to the fact that illiberal governments may be less disposed to correct inequality through redistributive policies than democratic ones (Meltzer & Richard, 1981 ; Sen, 1999). Other institutional measures having to do with labor law (core labor conventions, severity of violations of international norms, C87 and C98 severity scores) do not seem to have a significant cross-sectional association with inequality.[51]

In synthesis, the econometric analysis conducted so far suggests that despite a bivariate association between changes in union density and changes in inequality, the pronounced fall in trade union density in the past two decades, or the more modest trend toward collective bargaining decentralization, do not seem to have caused a rise in income inequality. There seems to be no robust within-country statistical association between changes in inequality and changes in labor institutions when other possible determinants of inequality are controlled for. The increase in inequality in the past 15 years seems mostly due to economic forces. In particular, a technologically induced shift in the demand for skilled labor and the increase in FDI stock over GDP (as well as tariff liberalization, although less robustly than other predictors) appear to have contributed to increase inequality.

When it comes to explaining differences in average levels of inequality across countries, however, one does find that labor institutions matter a lot. On average, the countries in which trade unions are stronger have lower levels of inequality than others. Less robustly, one also finds that a more centralized or coordinated structure of collective bargaining and more extensive political rights are associated with greater equality of incomes. These results do not seem very surprising: labor institutions are parts of social systems, and high trade union density and centralized collective bargaining structures are likely to be associated with other features (e.g., socialdemocratic governments and redistributive social policies), which in turn are likely to be conducive to a more egalitarian distribution of incomes. Interestingly enough, the econometric results suggest that labor institutions may function differently in different regions of the world, and that high trade union density and a more centralized collective bargaining structure may be conducive to greater inequality in Latin American countries, unlike other countries.

There may be several reasons why labor institutions fail to significantly affect recent changes in inequality while they do significantly impact average levels of inequality. One explanation could be measurement error: since the institutional variables are not measured very precisely, and probably less precisely than the economic variables, their impact may be attenuated. Another explanation may be that changes in institutions take a long time to affect the income distribution, and given the short time frame of the analysis here, one is unable to appreciate their effects. A third explanation may be that labor institutions may have begun to function differently from the past: whereas stronger trade unions and a more centralized structure of bargaining once led to a more compressed income distribution, recently they no longer do so, or do so to a much lesser extent. The next section explores this last hypothesis by focusing on 16 advanced countries.

IS THE INEQUALITY-DECREASING EFFECT OF IR INSTITUTIONS WITHERING AWAY?

This section addresses the question of whether the impact of labor institutions has been changing over time by taking a closer look at 16 advanced countries for which longer time-series data on institutions and other variables are available (Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, the Netherlands, Norway, Spain, Sweden, the United Kingdom, and the United States). I begin the analysis by re-estimating essentially the same within-model as in Table 5, over the same time frame as before, but controlling for the share of ICT investment in total capital – a measure which is available for all the above countries (Columns 1 and 3 in Table 11). The Reversed Index of Democracy is not included as it is entirely time-invariant for the 16 countries in question.[52] I also add a new predictor, the percentage of total public social expenditures over GDP, for which time series data are available (Column 3).[53] In so doing, I focus on the effects that labor institutions exert directly on income inequality. Those that these institutions exert indirectly, by being associated with a more generous welfare state, are now controlled for.

There are some interesting changes in the globalization variables when the focus is on advanced countries: FDI comes out as a significant predictor

Table 11. Determinants of Gini in 16 Advanced Countries (Fixed Effects with AR(1) Errors, Time Dummies and Constant Not Reported).

Dep. Var.	(1)	(2)	(3)
	Ln(Gini)	Ln(Gini)	Ln(Gini)
FDI	0.0157	0.0293*	−0.00214
	(0.0151)	(0.0152)	(0.0147)
Tariff liberalization	0.00271	0.00498	0.00397
	(0.00402)	(0.00413)	(0.00385)
Capital account openness	0.0192*	0.0132	0.0229**
	(0.0105)	(0.0109)	(0.00963)
Education years	−0.707	−1.124**	−0.838**
	(0.460)	(0.475)	(0.404)
Credit private sector	−0.0154	−0.0197	−0.00426
	(0.0138)	(0.0143)	(0.0132)
ICT share	0.197***		0.0922*
	(0.0554)		(0.0540)
Union density	−0.283	−0.226	−0.230
	(0.179)	(0.186)	(0.169)
CB coordination	−0.00312	−0.000705	−0.000978
	(0.00541)	(0.00555)	(0.00525)
Public social expenditures			−0.0113***
			(0.00261)
Time dummies	Yes	Yes	Yes
Observations	175	175	174
Number of countries	16	16	16
Adjusted R^2	0.168	0.0894	0.292
Estimated Rho	0.595	0.611	0.532

Notes: $^{***}p<.01$, $^{**}p<.05$, and $^{*}p<.1$; standard errors in parentheses.

only when technologically induced demand for skilled labor is not explicitly controlled for (Column 2): this signals that the FDI term probably proxies for this omitted variable, and that FDI in developed counties probably leads to greater demand for skilled labor (Feenstra & Hanson, 2001). Also, an increase in de jure Capital Openness seems to lead to greater income inequality in these countries. A greater supply of skills is associated with lower inequality. The greater the share of ICT investment (signaling greater relative demand for skilled labor) the more inequality increases. Total public social expenditures emerge as a highly significant predictor: the greater social expenditures decline in a country, the more inequality (measured by the Gini coefficient of equivalized net household disposable income) increases.[54]

Consistent with previous results, changes in unionization and collective bargaining coordination are both negatively signed, but not significantly different from zero. This is in contrast with previous research findings (all relative to a previous period) reported in Section 2, which suggest that industrial relations institutions have an equalizing effect on earning and hence income distributions. To check whether the effects have changed compared to the past, I now consider a longer time frame – 1978–2002 – for the 16 advanced countries in question.

The analysis that follows is (freely) inspired by a theoretical paradigm known as Power Resource Theory (PRT), which was elaborated to explain the historical trajectory of Scandinavian countries, i.e. advanced capitalist countries characterized by a highly egalitarian distribution of incomes (Esping-Andersen, 1990 ; Esping-Andersen & Korpi, 1984 ; Korpi, 1983; Korpi & Shalev, 1979; Stephens, 1979). According to PRT, there are durable differences in the organization of capitalist societies, which ultimately determine different levels of equality or inequality in the distribution of incomes (Korpi, 2006). The crucial factor determining these differences is the power of organized labor. The argument is that at a crucial moment in history – the period between WWI and WWII and then in the early post-war years – in some countries, but not in others, the labor movement and its political allies were able through mobilizations and industrial action to force capital into a historical compromise, whereby labor accepted the capitalist organization of the economy, but obtained in exchange not only a recognition of its prerogatives as labor market intermediary (through protective regulations on trade unionism and collective bargaining), but also protection against all sorts of social risks, and a growing expansion of social rights.

Over time, this historical compromise crystallized into a peculiar type of organized capitalism, best characterized by contrast with the model prevailing in the United States and (later) in other Anglo-Saxon countries: a highly institutionalized structure of the labor market, a large percentage of the workforce organized by trade unions, wages and working conditions determined through collective bargaining at the national level, an extensive welfare state whose provisions were a matter of citizenship rights, not of the individual's ability to pay, and, consequently, a more equitable distribution of incomes.

In brief, according to PRT, labor power is responsible *both* for the establishment of a large welfare state and for a highly institutionalized structure of the Industrial Relations system, and affects inequality through both channels: it contributes to compress market earnings directly (the

Industrial Relations channel) because trade unionism is historically associated with egalitarian wage policies ("equal pay for equal work"), and centralized wage bargaining further contributes to wage compression by reducing inter-establishment and inter-sector dispersion; it also contributes to reduce inequality indirectly by contributing to establish, and by reproducing over time, a large redistributive welfare state, which corrects market-generated inequality through redistributive taxes and transfers. The PRT argument incorporates an element of path-dependency (Pierson, 2004; Thelen, 1999): the events that shaped organized capitalism took place far back in history. However, since institutions are resilient and tend to change little and slowly over time, those formative events still shape cross-national differences in industrial relations and welfare systems.

Here I test the plausibility of the theoretical framework summarized earlier for cross-national differences in inequality through a simple empirical strategy: I compare cross-sectional regressions at two points in time: the period before 1990 (1978–1989) and the period from 1990 on (1990–2002).[55] The year 1990 was selected as a cut-off point because it divides the sample more or less in two. Substantially, the 1990 decade is the one in which the economic processes associated to globalization started to become most visible, and when the whole debate on globalization started.

Fig. 4 provides a pictorial representation of the hypothesized relationships between labor power, welfare state, and inequality, as well as the indicators used to operationalize these three constructs (discussed later). I hypothesize that the construct I refer to as "Labour Power" is positively related to the "Size of the Welfare State" and that it contributes to reduce Societal Inequality both directly and indirectly, through the size of the welfare state.

One obvious shortcoming of the empirical approach adopted here is that the sample size is very small ($n = 16$ at each of the two points in time). Thus the estimated models are necessarily highly parsimonious. Relying on the previous econometric analysis, which suggests that only the institutional variables are significant predictors of cross-sectional differences in Gini coefficients, I focus on these. As hypothesized by Power Resource Theory, institutions are likely to be parts of a system. Empirically, this implies that the institutional measures tend to be highly correlated and that it is difficult to parse out their respective contribution to inequality patterns. I rely on principal component analysis (PCA) to summarize the information provided by multiple indicators. Principal component analysis assumes that the data are visible manifestations of underlying hidden constructs, to which they are correlated, and seeks to express the hidden constructs as linear combinations of the (standardized) observed variables. I use multiple indicators to capture three hidden constructs: Labour Power, Welfare State Size, and Inequality.

To operationalize Labour Power, I use three correlated indicators: (1) the bargaining coordination index described earlier (BargCoord); (2) the collective bargaining coverage rate (BargCov), namely the percentage of workers covered by collective bargaining agreements (Ochel, 2001, p. 3) and the trade union density rate (TUDens). These indicators are all positively correlated and the pair-wise correlation coefficient is always higher than 0.5 as well as highly significant.

The results of the PCA analysis, reported in Table 12, suggest that the three indicators belong together: only one component has higher eigenvalue than 1 and captures about 63 percent of the total variance. The composite indicator of Labour Power uses the factor loadings of the first component, all positively signed, as weights, with bargaining coordination counting a little more than collective bargaining coverage and trade union density in determining the country score. Thus, Labour Power is high in countries with more coordinated bargaining, higher collective bargaining coverage, and higher trade union density.

Principal Component Analysis is used for the two other constructs, too (Tables 13 and 14). For Welfare State Size two indicators are used: (1) the total tax wedge as percentage of GDP, including social security and indirect taxes, which proxies for state intervention by measuring the extent to which a state is capable of extracting resources from its citizens for its activities;[56] (2) total public social expenditures as percentage of GDP, which directly

Table 12. Principal Component Analysis of Labour Power.

Number of Components	3	Components Retained	1	Number of Observations	366
	Component	Eigenvalue	Difference	Proportion of variance	Cumulative
	Comp1	1.90205	1.21004	0.6340	0.6340
	Comp2	0.692015	.286083	0.2307	0.8647
	Comp3	0.405932	.	0.1353	1.0000
Eigenvector	Variable	Comp1			
	BargCoord	0.6235			
	BargCov	0.5897			
	TUDens	0.5133			
Formula	Labour Power = 0.6235std(BargCoord) + 0.5897std(BargCov) + 0.5133std(TUDens)				

Table 13. Principal Component Analysis of Welfare State Size.

Number of Components	2	Components Retained	1	Number of Observations	352
	Component	Eigenvalue	Difference	Proportion of variance	Cumulative
	Comp1	1.8608	1.7216	0.9304	0.9304
	Comp2	0.139199	.	0.0696	1
Eigenvector	Variable	Comp1			
	TaxWedge	0.7071			
	SocExp	0.7071			
Formula	Welfare State Size = 0.7071std(TaxWedge) + 0.7071std(SocExp)				

Table 14. Principal Component Analysis of Inequality.

Number of Components	3	Components Retained	1	Number of Observations	90
	Component	Eigenvalue	Difference	Proportion of variance	Cumulative
	Comp1	2.77657	2.56678	0.9255	0.9255
	Comp2	0.209797	0.196167	0.0699	0.9955
	Comp3	0.01363	.	0.0045	1
Eigenvector	Variable	Comp1			
	D9D1	0.5964			
	D9D5	0.5605			
	PovRatio	0.5746			
Formula	Inequality = 0.5664std(D9D1) + 0.5605std(D9D5) + 0.5746std(PovRatio)				

captures social transfers. Here the first principal component captures almost the totality of variance (93%). The two variables are weighted equally in the composite indicator: the greater the percentage of total taxes and of public social expenditures, the greater the Size of the Welfare State.

The third principal component analysis captures how unequal a country is (Table 14). For this purpose, it uses three highly correlated indicators from the Luxembourg Income Studies database: (1) the D9/D1 ratio of Net Disposable Income, (2) the D9/D5 ratio of Net Disposable Income, which captures inequality in the upper part of the distribution, where, according to some analyses (Atkinson, 2007; Atkinson, 2008) inequality has grown the most; (3) the Poverty Ratio as percentage of people with less than 50% of the median Net Disposable Income. Once again, the first principal component captures most of the information in the data (93%). All three factor loadings are positive with approximately the same weight. A more unequal country is one in which the D9/D1, D9/D5, and Poverty ratios are higher.

I begin by examining the bivariate correlation between Labour Power and Welfare State Size before and after 1990 (Fig. 6). The relationship is positive in both periods. The countries with lower degrees of Labour Power, *in primis* the United States, tend to be characterized by a smaller Welfare State, and vice versa for countries with high Labour Power (the Scandinavian and Central European countries). The relative position of some countries changes over time – Australia, for example, is clearly an outsider in the former period (in the sense that it has a smaller welfare state than would be allowed by the measured strength of its labor movement) and less so in the second, whereas the United Kingdom shifts toward the US pole in the second period – but the shapes of the two curves remain remarkably similar across periods.[57]

Fig. 7 then examines the relationship between Welfare State Size and Inequality in the two periods. This relationship is negative as expected: the greater the Size of the Welfare State, the lower Inequality. The two opposite poles are once again the United States, on the one hand – a country with a residual welfare state and high levels of inequality – and Sweden on the other, where extensive social protections are accompanied by a much more

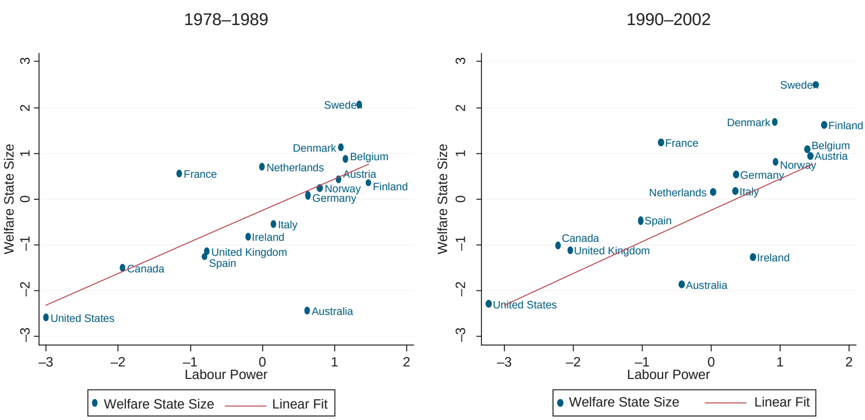

Fig. 6. Relationship between Labor Power and Welfare State Size: 1978–1989 vs. 1990–2002.

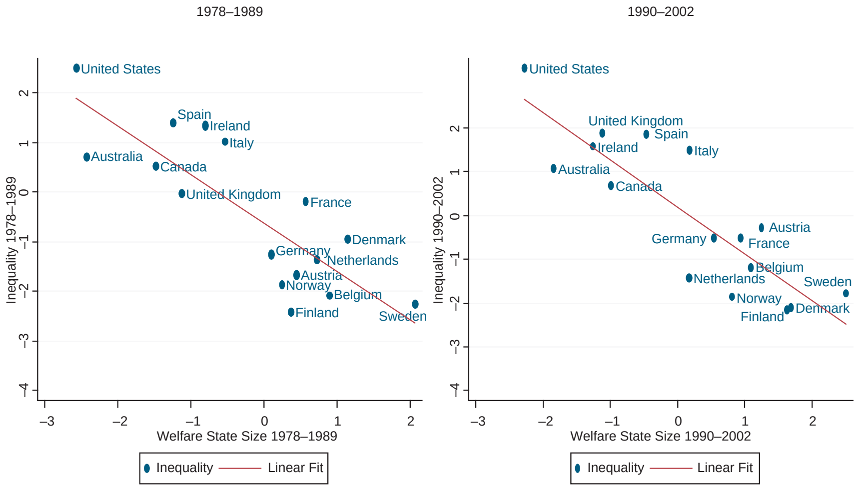

Fig. 7. Relationship between Welfare State Size and Inequality: 1978–1989 vs. 1990–2002.

egalitarian distribution of incomes. The slope of the two curves remains similar over time. However the second graph seems to have shifted rightwards compared to the first: both the size of the welfare state and inequality have grown on average in the 1990–2002 period. The increase in the size of the welfare state is due to well-known phenomena of population aging and the coming to maturity of various social programs (see Pierson, 2001). Also, the graphs in Fig. 6 confirm that over time the United Kingdom has shifted its relative position in the direction of the United States.

I now estimate the impact of both Labour Power and Welfare State Size on Inequality, controlling for each other, through regression analysis (Table 15). I also control for the power of left-oriented parties (measured through the proportion of seats in the lower chamber), which has been argued to affect the redistributive stance of governments (Bradley et al., 2003; Stephens, 1979), as well as for other economic determinants. The main goal of the analysis is to see whether the coefficients of the two main predictors change over time, and, if so, in which direction.

The parsimonious model with only two predictors in Table 15, Column 1 – Welfare State Size and Labour Power – performs remarkably well in explaining cross-country differences in Inequality in the 1978–1989 period, and accounts for almost 75 percent of the variance in the dependent variable. All regression coefficients are beta coefficients and are therefore

Table 15. Determinants of Inequality in 16 Advanced Countries (1978−1989), between Regressions (Constant Not Reported).

Dep. Var.	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality
Welfare state	−0.650***	−0.659***	−0.844***	−0.845***	−0.716***	−0.642***	−0.887***	−0.643***	−0.707***	−0.576***	−−0.646***	−0.647***
	(0.0993)	(0.101)	(0.144)	(0.160)	(0.153)	(0.122)	(0.265)	(0.111)	(0.150)	(0.139)	(0.108)	(0.136)
Labour power	−0.492***	−0.522**				−0.488***	−0.384*	−0.620***	−0.516***	−0.595***	−0.472***	−0.617***
	(0.131)	(0.198)				(0.127)	(0.201)	(0.130)	(0.133)	(0.165)	(0.116)	(0.139)
Left power		0.00418										
		(0.0203)										
CB coverage			−0.0180									
			(0.0109)									
Union density				−1.571								
				(1.109)								
CB coordination					−0.446**							
					(0.180)							
FDI						0.0605						
						(0.336)						
Tariff liberalization							0.136					
							(0.104)					
Capital account openness								−0.428**				−0.415*
								(0.141)				(0.227)
Education years									−0.209*			−0.0152
									(0.102)			(0.146)
Credit private sector										−0.442		
										(0.459)		
ICT share											0.252	
											(0.497)	
Observations	16	16	16	16	16	16	16	16	16	16	16	16
Adjusted R^2	0.748	0.729	0.685	0.682	0.732	0.728	0.760	0.846	0.781	0.741	0.730	0.832

directly comparable. The most important determinant of cross-country differences in inequality between 1978 and 1989 is the size of the welfare state. A one-standard deviation increased in the size of the welfare state reduces inequality by 0.65 standard deviations. Another predictor that is robustly different from zero is the Labour Power indicator. A one-standard deviation increase in the latter is associated with lower inequality of about 0.5 standard deviations. The electoral strength of the parliamentary left is insignificant when controlling for both the size of the welfare state and the power of labor (Column 2). The models in Columns 3–5 estimate separately the impact of different elements in the Labour Power indicator. The coefficient of the Collective Bargaining Coordination term is significantly different from zero (Column 5), while the others are not. The models in Columns 6–11 control for the same economic and globalization factors as examined earlier (FDI stock, Tariff liberalization, Capital openness, Years of Education, Credit to the Private Sector, Share of ICT investment in capital stock), one by one due to the small sample size. Both Capital openness and Education Years are negatively signed and significant. When both are entered in the specification simultaneously in Column 12, while the Welfare State Size and Labour Power terms remain highly significant, and the coefficient of the latter even increases in absolute value, the human capital control (Education years) becomes insignificant. These regression results suggest that institutional features of both the welfare state (captured by the Welfare State Size indicator) and of the labor market (captured by the Labour Power term) are the most important predictors of cross-country differences in inequality levels in the 1978–1989 period. Since, as we saw, Labour Power and Welfare State size are positively correlated, the regressions capture the direct effect of Labour Power on Inequality, net of its indirect effect through the welfare state. This direct effect is linked to the ability of labor to compress earnings in the market, before redistributive taxes and transfer are factored in[58] (Figs. 8 and 9).

Next I move to the period between 1990 and 2002 and re-estimate the same models as before (Table 16). The most important difference is that now Labour Power is much less robustly associated with Inequality than in the previous period.[59] The coefficient of Labour Power is still negative, but its magnitude is smaller in absolute value and often not significantly different from zero. Conversely, the Welfare State Size variable now plays a greater role in explaining cross-country differences. If one looks at the various components of Labour Power separately, one notices that the biggest change pertains to the Collective Bargaining Coordination index, whose coefficient is practically halved and no longer significant (Column 5). Thus,

Table 16. Determinants of Inequality in 16 Advanced Countries (1990–2002), between Regressions (Constant Not Reported).

Dep. Var.	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality	Inequality
Welfare state	−0.816***	−0.800***	−0.873***	−0.899***	−0.948***	−0.851***	−0.886***	−0.753***	−0.857***	−0.799***	−0.823***
	(0.196)	(0.210)	(0.180)	(0.162)	(0.157)	(0.240)	(0.170)	(0.181)	(0.200)	(0.231)	(0.208)
Labour power	−0.314	−0.306				−0.303	−0.365**	−0.411**	−0.348**	−0.393	−0.340
	(0.187)	(0.187)				(0.214)	(0.135)	(0.168)	(0.143)	(0.248)	(0.201)
Left power		−0.00365									
		(0.0144)									
CB coverage			−.00287								
			(.0135)								
Union density				−1.678							
				(1.026)							
CB coordination					−0.219						
					(0.158)						
FDI						−0.228					
						(0.387)					
Tariff liberalization							0.244				
							(0.182)				
Capital account openness								−0.765			
								(0.607)			
Education years									−0.317**		
									(0.117)		
Credit private sector										−0.595	
										(0.966)	
ICT share											−0.341
											(0.849)
Observations	16	16	16	16	16	16	16	16	16	16	16
Adjusted R^2	0.739	0.718	0.744	0.739	0.731	0.726	0.741	0.751	0.821	0.730	0.720

Notes: $^{***}p<.01$, $^{**}p<.05$, and $^{*}p<.1$; robust standard errors in parentheses.

it looks as though beginning with the 1990s coordinated bargaining no longer reduces inequality. When one controls for economic determinants one by one as was done before, one notices that Capital Openness is no longer significantly associated with lower inequality (Column 8). The effect in the previous period was probably due to small open countries like the Scandinavian countries which simultaneously had high capital openness and an egalitarian structure of incomes. As more countries open up their capital markets the effect disappears in the later period. The human capital control (Average Years of Education) remains significantly negative (Column 9). Even controlling for human capital, however, the impact of Labour Power is lower than in the previous period.[60]

Another way of looking at the changing impact of institutions is by visually inspecting the partial correlation of the Inequality indicator and the Labour Power indicator, controlling for Welfare State Size. Figs. 8 and 9 plot the residuals of a regression of inequality on welfare state size against

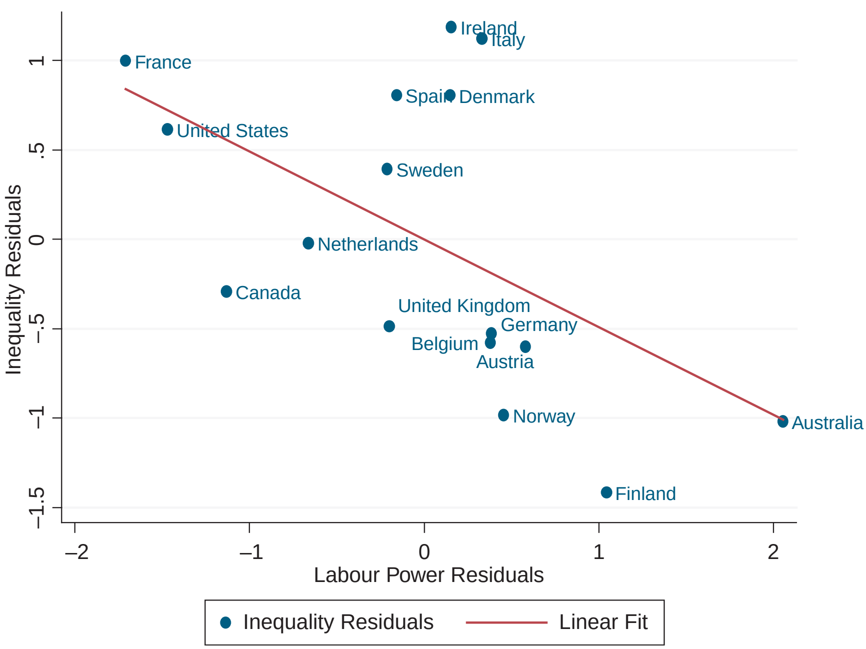

Fig. 8. Partial Correlation between Inequality and Labor Power Controlling for Welfare State Size between 1978 and 1989.

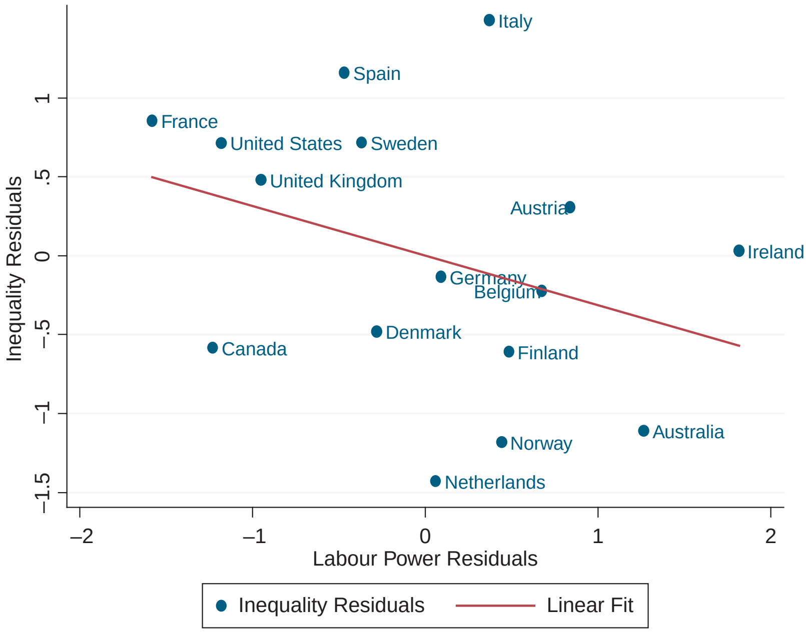

Fig. 9. Partial Correlation between Inequality and Labor Power Controlling for Welfare State Size between 1990 and 2002.

the residuals of a regression of labor power on welfare state size. The linear fit becomes much less steep in the period between 1990 and 2002 than in the previous period between 1978 and 1989.

These results suggest that, from the early 1990s on, the institutions associated with Labour Power – high trade union density, high collective bargaining coverage, a coordinated bargaining structure – and particularly coordinated bargaining, largely forfeited their capacity to directly reducing inequality and only kept an indirect effect on inequality through the size of the welfare state. This is consistent with micro evidence suggesting that the ability of unions to compress the distribution of wages has been declining over time (Card, Lemieux & Riddell, 2007, pp. 137 and 149–50). It is also consistent with case study evidence on recent developments in some of the countries included in this analysis. Some time ago unions participating in national collective bargaining engaged in explicit attempts at compressing skill differentials through various means: requests for lump-sum wage increases (which tend to favor the low-paid), tapered percentage wage increases (higher for the low-paid), skewed indexation mechanisms (like the Italian *scala mobile*) which assured those with low earnings a greater degree of protection from inflation, and which, particularly in times of double digit inflation, contributed to compress earnings, etc. (Baccaro & Locke, 1998; Edin & Holmlund, 1995; Erickson & Ichino, 1995; Schulten, 2002).

Over time, these strategies and institutional arrangements were largely dismissed. In Sweden and other Scandinavian countries, the egalitarian wage policies pursued by the unions from the late 1960s on created considerable problems for employers, who found it difficult to recruit and motivate highly skilled labor (Pontusson & Swenson, 1996; Swenson & Pontusson, 2000). They also generated problems for unions. For example, in the early 1980s the Swedish blue-collar union Metall found itself losing many members to the white-collar union SIF, and had eventually to drop the policy of wage compression as well as the whole model of national bargaining associated with it, which was replaced by sectoral bargaining (Thelen, 1993, p. 39).

Centralized bargaining did not die in the 1990s but largely lost its redistributive function. Centralized bargaining used to be one of the key institutions in ‘‘social corporatist’’ countries (Korpi, 1978; Pekkarinen, Pohjola, & Rowthorn, 1992; Pontusson, 2005; Rowthorn, 1992), where unions negotiated at the national level and exchanged wage moderation for both a more equitable distribution of earnings and more extensive social protection networks (Mares, 2006; Pizzorno, 1978).

In the 1990s, after a temporary decline in the 1980s, centralized bargaining surprisingly resurfaced in a number of countries, primarily but not exclusively European (Baccaro & Lim, 2007; Berger & Compston, 2002; Fajertag & Pochet, 1997 ; Fajertag & Pochet, 2000 ; Hassel, 2003). However, the social outcomes of these new forms of centralized bargaining, also known as ‘‘social pacts,’’ were considerably different and markedly less redistributive than in the past. These pacts seemed much more concerned with increasing country competitiveness than on redistribution (Rhodes, 1996 ; Rhodes, 2001 ; Streeck, 2000). In Ireland, for example, the collective bargaining system was strongly recentralized in the past two decades, yet there is little evidence that this may have contributed to reduce wage differentials (Baccaro & Simoni, 2007 ; Barrett, Gerald, & Nolan, 2000). In Italy, the *scala mobile* was abolished in 1992, and the unions negotiated with employers and the government a new architecture of nationally coordinated sectoral bargaining, which did not prevent increases in wage and income inequality (Baccaro, 2002; Brandolini, Cipollone, & Sestito, 2001; Erickson & Ichino, 1995).

In brief, faced with new and more stringent market constraints – more elastic labor demand, particularly for the low skilled, and high skill premia as a result of skill-biased technological change – union behavior seems to have became more market conforming over time and to have lost much of its redistributive features. Large welfare states, instead, continued to play an important redistributive role well into the 1990s. Indeed, an even greater proportion of the cross-country variation in Inequality was explained by differences in Welfare State size in this period than in the previous. This may seem surprising, given the debate on the crisis of the welfare state, but is in line with the findings of other scholars as well (Bradley et al., 2003; Brady, 2003; Brady, 2009; Kenworthy & Pontusson, 2005; Pontusson, 2005, chapter 7).

SUMMARY OF KEY FINDINGS

The analysis presented in this chapter suggests that there has been a considerable decline in unionization from 1989 onwords. Union density declined in almost all countries considered in this analysis. The decline was dramatic for Central and Eastern European countries, which started from very high levels (Martin & Kaya, 2010). Changes in collective bargaining structure were less spectacular, at least according to the available measures, which may overlook processes of erosion within formally stable structures (Streeck & Thelen, 2005). In most countries, the main level of collective bargaining did not change. There was, however, a modest trend toward decentralization in others.

Although income inequality increased in almost all countries in the sample, this increase does not seem to have been caused by the deterioration in industrial relations institutions (trade union decline and collective bargaining decentralization). Specifically, one cannot argue that union decline led to growing inequality. Barring the Central and Eastern European countries, there is no statistical association at standard levels of confidence between changes in union density and other labor institutions, and changes in inequality within countries, controlling for other determinants.

The recent increase in inequality seems better predicted by economic factors than by industrial relations institutions. For example, technology-induced shifts in the demand for skilled labor (captured by the share of ICT investment in the capital stock) are associated with greater inequality. An increase in the stock of FDI as percentage of GDP also tends to be associated with greater inequality (in the advanced countries this seems to happen because FDI appears to increase the demand for skilled labor). An increase in the supply of human capital (average years of education) lowers income inequality. The inequality-increasing effect of FDI (possibly linked to shifts in labor demand) seems the only robust effect of globalization trends on inequality according to the analysis. Tariff liberalization also seems associated with greater income inequality, but its impact appears less robust.

The econometric analysis conducted earlier also suggests that labor institutions may have different effects in different regions of the world. For Latin American countries (i.e., countries characterized by a large share of the informal economy and a tradition of corporatist unionism with close linkages to the State), there is no evidence that high trade union density and a more centralized collective bargaining structure lead to less inequality, and some evidence that the opposite may be true.

As far as advanced countries are concerned, high trade union density, a more coordinated collective bargaining structure, and greater coverage of collective bargaining agreements tend to be associated with a larger welfare state. This relationship does not change over time. Large welfare states, in turn, reduce inequality in advanced countries. This relationship, too, does not change over time. This may clash with recent debates about welfare state crisis but in consistent with other research findings (Bradley et al., 2003; Brady, 2009; Kenworthy & Pontusson, 2005; Pontusson, 2005, ch. 7). Cross-sectionally, an increase of one standard deviation in the size of the welfare state was associated with a decrease of 0.65 standard deviations in inequality before 1990 (controlling for labor power). From 1990 on this same effect was even higher: 0.8 standard deviations (controlling for labor power).

What changes from the 1990s on in advanced countries is the capacity of industrial relations institutions to reduce inequality directly by compressing market earnings. In particular, centralized collective bargaining seems to have become less redistributive than in the past. To the extent that industrial relations institutions continue to support and reproduce the welfare state, they reduce inequality indirectly though this channel. However, their direct equality-enhancing effect seems largely to have disappeared.

CONCLUDING REMARKS

The introduction to this chapter argued that, since the international economic governance framework is, and is likely to remain for some time, under-institutionalized, the task of protecting societies from the potentially undesirable consequences of globalization (the "Polanyi problem") falls largely on country-level institutions. Based on the analysis reported above, we can conclude that welfare state institutions still have strong redistributive effects (at least in advanced countries), but industrial relations institutions may have lost them, or may be in the process of losing them.

Trade unionism currently operates under more stringent structural constraints than in the past: more elastic labor demand curves, particularly for the low skilled, and greater wage premia for the high-skilled as a result of skill-biased technical change. To the extent that these constraints are produced or magnified in their effects by current globalization trends, these have a double effect: on the one hand they weaken trade unions – for example, the analysis above shows that FDI is associated with lower density rates; on the other hand they also reduce the space available to trade unions for redistribution.

Yet trade unions – historically a key actor in equalizing social conditions – can still contribute to reduce income disparities, and in ways that do not clash with current economic realities. The analysis reported earlier suggests that much of the current increase in inequality is due to a mismatch between demand and supply of skills. For various reasons (some related to globalization, some to technical change) the demand for skilled labor has increased more than its supply. If this is true, then trade unions do have an important role to play, not through policies seeking to compress wage differentials across skill levels (these seem to have become more difficult if not utterly unfeasible in current circumstances), but through supply-side policies aimed at increasing the workers' skill levels and at promoting an egalitarian transformation of the workplace, such that all jobs are challenging and stimulating, and workers have the skills needed to take them up. Examples of what trade unions might do reduce inequality in the current day and age include participating in vocational training programmes, or pushing management to adopt work restructuring schemes that enhance workers' abilities. This type of "supply-side" egalitarianism is not a new strategies for unions, but is very much part and parcel of the unions' cultural heritage (Baccaro & Locke, 1998). Reactivating this heritage seems not only possible but also desirable.

NOTES

1. To use the words of Richard Freeman: "When firms do not have 'rents' to share with workers, institutions cannot affect redistribution" (Freeman, 2007a, p. 15).
2. This is thanks to the data collection efforts of Jelle Visser over the years.

3. These results concerning heterogeneity of institutional effects across models of capitalism do not seem very robust. For example, Wallerstein (1999, p. 670) too tested for different effects in coordinated vs. liberal market economies (albeit with a smaller sample size), but could not reject the hypothesis of no differences.

4. However, in a related paper relying on very similar data and specification, Pontusson, Rueda, and Way (2003) found that both union density and bargaining centralization were important determinants. These slightly different findings may be due to the different estimator used in the second study: a least square dummy variable estimator, which is inconsistent with a dynamic model and whose bias can be sizeable with a short time dimension (*T* was considerably smaller than 30 in this analysis) (Judson & Owen, 1999).

5. This statement is incorrect, but only slightly: I, too, found very few longitudinal cross-country studies.

6. Market income includes wages and salaries, self-employment earnings, property income, and private pension income. Disposable income is market income after cash transfers and taxes. The unit of analysis was the household, not the individual, and the analysis was limited to households where the head was of working age, i.e. between 25 and 59.

7. For information, see: http://www.lisproject.org/

8. Many thanks to Martin Rama of the World Bank for making this database available to me as well. I did not use the information on trade union density therein for two reasons: (1) the data were aggregated in five-year averages and (2) they were expressed as percentage of total labor force and not as percentage of wage and salaried earners.

9. The University of Texas Inequality Project (http://utip.gov.utexas.edu/) has used the UNIDO data, as well as data from the World Bank's Deininger and Squire database, to estimate time series on household inequality for a number of countries (1996). These data (http://utip.gov.utexas.edu/data/EHIIv23.xls) are imputed data and are only available until the late 1990s. On the method used see Conceiçao and Galbraith (1998) and Galbraith and Kum (2004).

10. I collected data on inequality for 128 countries, but ended up focusing on 51 of them due to limited availability of labor institutions indicators and other controls.

11. The UNU/WIDER database can be downloaded on-line at the following address: http://www.wider.unu.edu/research/Database/en_GB/database/. A new version of the database (WIID2c) became available on May 31, 2008. The analyses below rely on the previous version (WIID2b).

12. The SEDLAC database is available at: http://www.depeco.econo.unlp.edu.ar/cedlas/sedlac/statistics.htm#inequality/. Data from SEDLAC are included in UNU/WIDER. However, the on-line version often had more recent estimates.

13. This, too, is available on-line: http://www.unicef-irc.org/databases/transmonee/#TransMONEE/. I used the data on net incomes.

14. Seehttp://www.lisproject.org/keyfigures/full_kf.xls/. The LIS estimates are based on microdata which are top coded and top coded to eliminate the extreme portions of the income distribution, where measurement error is more likely. Also, the unit of analysis is the household, ''equivalized'' to account for possible economies of scale within the household.

15. See http://iresearch.worldbank.org/PovcalNet/jsp/index.jsp/

16. Only about 6 percent of the data I use refers to concepts other than income: consumption (4.5 percent) and gross earnings (1.6 percent).

17. The Gini coefficient estimates were linearly interpolated. This increased the number of data points from 409 to 622.

18. I initially collected union density data for 139 countries from various sources, but ended up focusing on only 51 countries, those with a meaningful time variation and for which information on other variables was available. For Asian countries an important source was Kuruvilla, Subesh, Hyunji, and Soonwon (2002). I am very grateful to Pascal Annycke and Melissa Luongo for the excellent work they did in assembling some of the data, and, in the case of Melissa, for her research on a number of countries. The data from the Visser database are adjusted density rates: the number of union affiliates who are not wage and salary workers is subtracted from the numerator, and the number of wage and salary workers who do not have the right to organize (e.g., public sector workers in some countries) is subtracted by the denominator. For the other countries these adjustments were not possible. However, the denominator was kept as much as possible constant.

19. The union density variable was linearly interpolated. This increased the number of data points from 719 to 808.

20. Again, many thanks to Melissa Luongo for collecting the sources from which information needed for the coding was drawn.

21. A previous version of the chapter included a series of bivariate graphs. These have been omitted in this version for reasons of space.

22. The ILO core conventions are eight and pertain to: forced labor (C29 and C105), freedom of association and collective bargaining (C87 and C98), equality and nondiscrimination (C100 and C111), and prohibition of child labor (C138 and C182).

23. I am very grateful to Douglas Lippoldt of the OECD Secretariat for providing these data.

24. The index weights the perceived severity of the labor violation (based on the OECD Secretariat's assessment) by the severity of the CEACR evaluation of the situation. For more information on the construction of the index, see OECD (2000, pp. 85–87). The data have been linearly interpolated.

25. The number of countries for which the C87 severity score is available is 30 in 1990 and 32 in 2000. For the C98 severity score, these numbers are 29 and 32, respectively. It bears emphasizing that several countries in the sample have not ratified either or both conventions. For these countries the severity scores are obviously not available.

26. I am very grateful to Patrick Hettinger and Subir Lall of the IMF Secretariat for providing these data.

27. The tariff rate is an average of the effective tariff rate (tariff revenue/import value) and of the average unweighted tariff rate; see IMF (2007, p. 57).

28. The Lane and Milesi–Ferretti dataset on gross foreign asset and liability positions for 145 countries is available on-line at: http://www.tcd.ie/iiis/pages/people/planedata.php/

29. The Chinn-Ito de jure measure of capital openness is available on-line at: http://www.web.pdx.edu/~ito/kaopen_2006.xls/

30. The paragraphs that follow draw on Goldberg and Pavcnik (2007).

31. The measure of financial sector development is the ratio between private credit by deposit money banks and other financial institutions over GDP. See Beck, Demirguc-Kunt, and Levine (2007). The measure is available on-line at: http://siteresources.worldbank.org/INTRES/Resources/469232-1107449512766/FinStructure_60_06_final.xls/

32. The measure of human capital is Barro and Lee's average number of schooling years in the population aged 15 + (Barro & Lee, 2000). The Barro and Lee data are available every five years and until 2000. They have been interpolated and extrapolated to cover the 2001-2005 period. The Barro and Lee's database is available on-line at: http://www.cid.harvard.edu/ciddata/barrolee/appendix_data_tables.xls/

33. The proxy used is the ratio of the stock of information and communication technology capital to total capital. For more information on this variable, see IMF (2007, p. 58).

34. The Freedom House scores are available at: http://www.freedomhouse.org/uploads/FIWAllScores.xls. The Political Rights index is a 1-to-7 index, where higher scores indicate more serious violations of political rights.

35. Due to lack of data, I am unable to include additional institutional predictors: the minimum wage, which is likely to pull up the lower tail of the distribution, and labor market institutions like employment protection and unemployment insurance generosity, which are likely to improve the position of less skilled workers. However to the extent that the latter are correlated with unionization and collective bargaining, these may proxy for the missing institutions as well. Data on 18 advanced countries between 1960 and 1998 suggest that this may be the case: the correlation between union density rates and/or collective bargaining coordination scores, on the one hand, and measures of employment protection, unemployment benefit replacement and unemployment benefit duration, on the other hand, is always significantly different from zero (Baccaro & Rei, 2007).

36. The software used for all analyses is Stata 10 SE. Descriptive statistics and correlation matrix are in Appendices 1 and 2, respectively.

37. The time series are too short for meaningful tests of stationarity and co-integration. However, while the series are certainly long-memoried (De Boef, 2001), a unit-root problem is unlikely. I estimated a dynamic model including labor institutions, globalization variables, other economic controls, but not the country fixed effects (the right-hand variables were the same as in Column 1 of Table 5 except the lagged dependent variable was also included and the fixed effects excluded. Such model is known to bias the coefficient of the lagged dependent variable upwards. Yet not even with this estimator did the 95 percent confidence interval of the lagged dependent variable cover the value of 1, i.e. a unit root could not be detected (Bond, 2002).

38. Similar conclusions (both cross-sectionally and longitudinally) are reached if one focuses on ratification of core C87 and C98 only.

39. Statistically this phenomenon manifests itself as positive correlation among the labor institutions indicators.

40. The econometric model assumes that there is no reversed causation (and hence endogeneity) from income inequality to the right-hand side predictors. This assumption seems warranted as far as institutional variables are concerned: institutions are highly path-dependent, and to the extent that they change, the motivation is often more political than economic. It also seems unlikely that

inequality causes globalization, especially the more de jure dimensions of it like tariff and capital account liberalization. One possible source of endogeneity is with human capital supply. For this reason, the measure used is average years of education, and not the percentage of population with higher education (which is more likely to depend on skill differentials). There could be endogeneity on the right-hand side of the model: some of the predictors may be causally related to one another. For example, technology may depend on availability of skills (Acemoglu, 2002). Below I test explicitly for the possible endogeneity of union density to globalization. Endogeneity on the right-hand side of the statistical model is likely to manifest itself as multicollinearity, and to make it more difficult to reject hypotheses about zero coefficients.

41. For several Central and Eastern European countries: Czech Republic, Estonia, Latvia, Lithuania, Slovak Republic, Slovenia, as well as Taiwan, data on the Capital Account Openness Index are not available in our database. Data on Average Number of Education Years are also unavailable for Estonia, Latvia, Lithuania, and Taiwan. The Credit by Bank and Other Financial Institutions as a percentage of GDP variable is not available for Taiwan. The Reversed Democracy Index is not available for Hong Kong.

42. The Stata command used for estimation is *xtregar, fe*. This routine estimates time-series cross-section regressions when the error term is first-order autoregressive. It is based on Baltagi and Wu (1999) and is appropriate for unbalanced panels and for observations which are unequally spaced over time. The option *onestep* is used to estimate the autoregressive parameter ρ. This option implements the method proposed by Baltagi and Wu (1999). After ρ is estimated, the data are transformed a first time to remove the first-order autoregressive component, and then a second time to remove the fixed effects (within transformation). In this second transformation the first observation of each panel is dropped (see Stata Corporation, 2007, pp. 421–7). Note that the AR(1) component is around 0.6 in all specifications, i.e. sizeable. This implies that ignoring serial correlation of the errors, especially in the presence of heavily-trended independent variables, is likely to severely underestimate the standard errors of the coefficients (and overestimate the R^2) and lead to overly generous significance levels (see Gujarati, 2003, pp. 449–60). Indeed, when one estimates fixed effects models identical to the one reported in Table 5, but neglects the (first order) serial correlation in the error term, many more economic variables appear significantly different from zero and the R^2 is higher by more than 20 percent.

43. The Collective Bargaining Structure index is entirely time-invariant for Asian countries. Most of its time variation is due to variation within the Advanced Countries. An analysis of regional heterogeneity (similar to the one conducted for trade union density) makes little sense in this case.

44. The regression coefficient on the FACB variable depends on three countries only: Hong Kong, the Netherlands, and New Zealand. These are the only countries for which the 0–2 index of ratifications of C87 and C98 changes in the period under consideration. Instead, the number of core conventions ratified has greater time variation.

45. The coefficient of Unionization in CEE countries depends entirely on two countries: Hungary and Poland – the only CEE countries for which data on Capital Openness are available.

46. As argued above, the CB structure coefficient largely depends on developments in advanced countries, which are the only regional groups with considerable within-country variation.

47. Additional models have been estimated to assess the impact of variation of the C87 and C98 severity index scores on inequality, controlling for other determinants. None of these additional institutional variables seems to have a significant impact on (changes in) inequality.

48. Data on IT investments over capital stock are unavailable for the following countries: Czeck Republic, Dominican Republic, Estonia, Greece, Hong Kong, Hungary, Jamaica, Latvia, Lithuania, New Zealand, Poland, Portugal, Slovak Republic, Slovenia, Switzerland, and Taiwan.

49. The main pattern of econometric results reported earlier holds if different estimators are used. For example, rather than correcting for the AR(1) component by transforming the data, I also estimated a fixed effects dynamic specification in which the Gini coefficient depended on its level in the previous year, as well as on the (assumedly) exogenous variables. To correct for the bias created by the correlation between the lagged dependent variable and the fixed effects (Kiviet, 1995; Nickell, 1981), I used the Stata procedure *xtlsdvc*, which implements the method described in Bruno (2005) for bias correction, and uses bootstrapping to estimate the variance of the coefficients. With this estimator, the coefficients on FDI (positive), Tariff Liberalization (positive), and Capital Account Openness – all to be interpreted as short-term coefficients – emerged as significantly different from zero, while the coefficient of Trade Union Density in CEE Countries was marginally insignificant. Finally, rather than using an inconsistent estimator and correcting the bias, I also used an IV approach to the problem of the endogeneity of the dependent variable. I used the one-step system GMM estimator with cluster-robust standard errors (Bond, 2002, p. 3; Roodman, 2007) – Stata command: *xtabond2*. With this estimator, it seemed that the only robust predictor of the Gini coefficient was its value in the previous year, i.e. the lagged dependent variable. This was possibly the result of the inefficiency of IV estimators in finite samples.

50. Since the sample is unbalanced, and the countries are observed at different points in time, the variable YEAR checks whether the period in which the countries are observed affects the assessment of their average inequality.

51. I also estimated additional models checking for the impact of number of core conventions ratified, number of freedom of association and collective bargaining conventions ratified, C87 severity index, and C98 severity index, respectively, but found insignificant results. Also, I re-estimated the model in Table 10, Column 6, by excluding one country at a time and the pattern of results reported earlier held: the most unstable estimate was that of the Reversed Index of Democracy, whose magnitude was similar across specifications, but whose standard errors and significance levels seemed to vary depending on exclusion of particular countries. Given the absence of a standardized measure of income inequality, as discussed earlier, cross-sectional differences may reflect different ways to measure inequality in the various countries.

52. All these countries score 1 (minimum level of political rights violation) throughout the period.

53. The data come from the OECD Social Expenditure Database and are available until 2003.

54. One legitimate concern about the Social Expenditures variable has to do with possible reversed causation (from inequality to social expenditures) and hence endogeneity. However, if high inequality leads governments to increase social expenditures, then the correlation between the two should be positive, not negative as in Column 3 of Table 11. Thus, the coefficient of the social expenditure term can be considered a lower bound.

55. The reason why I do not estimate a time series cross-sectional model (TSCS) with annual data as I did above, even though annual data are available for some of the indicators, are multiple: (1) all indicators of inequality (from the LIS database) are available at best every five years; some institutional indicators, too, like collective bargaining coverage, are annual interpolations from five-year data; (2) a TSCS approach is more than likely to require fixed effects to control for time-invariant omitted variable. This is a problem, however, because the labor institutions I am interested in do not vary much over time but mostly across countries; and (3) the series are long-memoried and seem highly serially correlated. However, given the short duration of the series, no reliable tests of stationarity and co-integration are available.

56. Many thanks to Andrea Bassanini of the OECD Secretariat for providing this variable.

57. With a collective bargaining system characterized by compulsory arbitration, generally considered a functional substitute for centralized bargaining (Lansbury & Wailes, 2004), Australia scored almost as high as Central and Northern European countries on the Labour Power Index before 1990, but the Welfare State Size was similar to other Anglo-Saxon countries.

58. These results hold if the dependent variable is the Gini coefficient of net disposable income. The main differences with these alternative specifications is that the union density rate coefficient is significantly different from zero, and that the capital openness and education years variables are both insignificant in Column 12. Also, results hold if the equation in Column 1 is re-estimated after taking out one country at a time.

59. It is worth mentioning that a previous analysis had found that the impact of bargaining centralization in reducing wage dispersion was ‘‘virtually identical’’ in 1973 and 1985 (Rowthorn, 1992, p. 111).

60. Again, these results hold using the Gini coefficient as dependent variable. The main peculiarity is that union density does have a significant negative association with Gini, and its magnitude is only slightly smaller than in the previous period. As suggested earlier, it is Bargaining Coordination that seems to have lost its inequality-reducing effects, not so much Union Density. Also, results hold overall if the equation in Column 1 is re-estimated after taking out one country at a time. Interestingly, the Labour Power term is significant if Canada, Ireland, and Italy are taken out of the sample. This suggests that in the aforementioned countries Labour Power is less conducive to redistribution than elsewhere. Ireland and Italy experienced a marked increase in collective bargaining coordination in the 1990s, with the establishment of ‘‘social pacts,’’ but in both countries inequality did not decline or even increased.

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APPENDIX 1. DESCRIPTIVE STATISTICS, TOTAL AND BY REGION

Table A1

		Gini	FDI	Tariff Opennness	Capital Account Openness	Education Years	Credit Private Sector	Union Density	Reversed Democracy Index	Collective Bargaining Structure
Latin America and Caribbean	*N*	185.00	208.00	221.00	195.00	221.00	217.00	194.00	221.00	220.00
	Mean	0.51	22.24	88.43	0.31	6.15	0.27	0.19	2.38	1.56
	Standard deviation	0.05	14.42	5.53	1.51	1.31	0.14	0.10	1.04	0.76
	Minimum	0.41	2.96	54.30	−1.71	3.91	0.03	0.06	1.00	1.00
	Maximum	0.63	76.19	96.08	2.68	9.20	0.73	0.47	6.00	4.00
Advanced countries	*N*	245.00	352.00	363.00	323.00	374.00	374.00	368.00	374.00	373.00
	Mean	0.31	25.77	95.35	2.07	9.24	0.97	0.37	1.12	3.03
	Standard deviation	0.06	24.15	4.51	1.04	1.91	0.47	0.20	0.55	1.33
	Minimum	0.20	0.31	54.02	−1.07	4.06	0.13	0.08	1.00	1.00
	Maximum	0.46	151.31	98.94	2.68	12.21	3.45	0.87	5.00	5.00
Central and Eastern Europe	*N*	123.00	107.00	113.00	29.00	88.00	109.00	130.00	136.00	135.00
	Mean	0.29	22.65	92.66	−0.29	8.87	0.28	0.37	1.69	2.69
	Standard deviation	0.05	17.72	4.24	0.91	0.99	0.15	0.21	1.26	1.17
	Minimum	0.19	0.37	81.30	−1.07	6.59	0.02	0.11	1.00	1.00
	Maximum	0.40	87.95	100.00	2.12	10.04	0.69	0.99	6.00	5.00
Asia	*N*	69.00	128.00	119.00	105.00	136.00	117.00	116.00	119.00	136.00
	Mean	0.39	43.93	85.65	0.22	7.25	0.77	0.31	3.65	1.38
	Standard deviation	0.07	68.48	15.15	1.62	2.16	0.48	0.28	1.90	0.49
	Minimum	0.29	1.12	33.95	−1.71	3.75	0.18	0.03	1.00	1.00
	Maximum	0.53	275.44	100.00	2.68	11.12	1.77	0.95	7.00	2.00
Total	*N*	622.00	795.00	816.00	652.00	819.00	817.00	808.00	850.00	864.00
	Mean	0.38	27.35	91.69	1.14	8.04	0.66	0.32	1.89	2.34
	Standard deviation	0.11	34.05	8.20	1.59	2.19	0.50	0.21	1.39	1.30
	Minimum	0.19	0.31	33.95	−1.71	3.75	0.02	0.03	1.00	1.00
	Maximum	0.63	275.44	100.00	2.68	12.21	3.45	0.99	7.00	5.00

APPENDIX 2. CORRELATION MATRIX

Table A2

	Gini	FDI	Tariff Opennness	Capital Account Openness	Education Years	Credit Private Sector	Union Density	Reversed Democracy Index	Collective Bargaining Structure
Gini	1								
($N=$)	622								
FDI	0.1571*	1							
($N=$)	593	795							
Tariff	−0.3092*	0.4130*	1						
($N=$)	580	760	816						
Capital account openness	−0.3786*	0.3489*	0.6037*	1					
($N=$)	488	652	644	652					
Education	−0.6109*	0.1738*	0.5570*	0.5916*	1				
($N=$)	576	757	784	652	819				
Credit	−0.3619*	0.2068*	0.5044*	0.5782*	0.5746*	1			
($N=$)	584	768	801	647	780	817			
Union density	−0.5746*	−0.077	0.1322*	0.0981	0.3298*	0.1857*	1		
($N=$)	596	749	760	615	761	764	808		
Democracy	0.3625*	−0.1110*	−0.3701*	−0.4536*	−0.4936*	−0.2738*	0.0826	1	
($N=$)	616	779	799	637	802	802	797	850	
Collective bargaining structure	−0.5679*	−0.0828	0.3155*	0.3952*	0.3656*	0.2822*	0.4115*	−0.3548*	1
($N=$)	619	794	814	651	817	815	805	847	864

Notes: All correlation coefficients significant at 5%; starred correlation coefficients significant at 1%.

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