Abstract

The overall goal of the present work is to determine the effect that unionization has over the distribution of income inequality. To do this, we use a subset of OECD countries that allow us to have the required time series data for unionization and income inequality between the years 1970 and 2006, including Chile. We use cointegrated panels to control for possible endogeneity between unionization and inequality. The results obtained show that, on average, for the countries in the sample, an increase in unionization improves income distribution, making it less unequal. However, these results are not homogeneous across countries. We can see that in countries with a high rate of unionization, high coverage of collective bargaining and relatively low levels of inequality, an increase in unionization increases inequality. However, for countries such as Chile, where these variables are low, we see that an increase in unionization reduces income inequality.
1 Introduction

In the discussion about how to combat inequality in income distribution, the issue of trade union action, and mainly collective bargaining, is usually controversial. From a most classical perspective, trade unionism is a rigidity that does not allow the labor market flexibility to operate in an efficient manner. This is due to the fact that collective bargaining tries to raise the wages of workers above their level of productivity, inflating the costs of production of companies that in the long run, will have to lay off workers in order to keep operating. This increase in unemployment, would lead to an increase in income inequality, because both variables are correlated.

The opposing point of view is that trade unionism is a mechanism to balance forces in the market where wages are determined. It is a tool that in the long run allows improvement in the situation of the most vulnerable groups of society. There is ample evidence in the literature regarding the importance of trade union action in the decrease in wage inequality through collective bargaining (Hayter, 2002; Kahn, 2000; Kostzer et. al., 2005; among others). The arguments that support the fact that trade unions are involved in this negotiation lie in the power asymmetry between employers and workers. Trade unions totally or partially balance this asymmetry because labor law allows them to use participatory mechanisms, such as protests or strike, to improve the position to negotiate worker’s wages (Trajtemberg, 2008). However, the effect that trade unions have on the distribution of income it is not sufficiently documented.

Another aspect that has not been addressed enough in the literature, as indicated by Herzer (2014), is that the effect of trade unionism on inequality does not have to be homogeneous. On the contrary, there are several reasons to think that the effects may be specific to each country. In this sense, the question that this work will attempt to answer is what is the impact that unionization has had on income inequality over the last four decades? Studying the case of Chile and other countries. The hypothesis of this investigation is that, at an international level, there is a negative relationship between trade union action and income inequality. The public policy implication is clear: if there is a significant relationship, this work would motivate the discussion of certain reforms to the labor market to promote trade union action as a mean to combat inequality.

The overall goal of the present work is to determine the effect that unionization has over the inequality of income distribution. To achieve this, we use panel data that includes 11 countries. This data base contains time series data of unionization and income inequality between the years 1978 and 2005.
Some efforts have to be done in order to include Chile in the analysis. As specific objectives we have (1) estimate a panel model of fixed effects to determine the influence of rate of unionization and centralization of collective bargaining over the level of income inequality; (2) analyze the relationship between union density and income inequality using econometric techniques for cointegrated panels, aiming to control for possible endogeneity between these variables; (3) determine the causality of the relationship if there is cointegration among the variables; (4) investigate which are the specific effects by country and (5) discuss the implications of the results obtained in the public policy.

Since we are using the rate of union density as dependent variable, instead the proportion of workers who negotiate collectively, this research aims to contribute to the debate about the relationship between unionization and inequality. With this analysis we seek to incorporate some new visions that affirm the impact of unions over inequality is not only through collective bargaining. For example, the work developed on Hacker and Pierson (2010) proposes that, beside collective bargaining, the role of trade unions is exercising political influence over the public sphere and social policies. This results is projection in both the design and implementation of public policies of a country. Therefore states where unions have more power tend to generate public policies more directed toward equality.

This paper is divided into five sections. In section 2 we present a literature review relevant to the topic. In section 3 we present the data and some descriptive statistics. Section 4 presents the econometric model and section 5 concludes.

2 Literature Review

The traditional analysis approach for the impact of unionization over income inequality generally led to conclude that an increase in the rate of unionization implies an increase in income inequality by the increase of unemployment rate. Theoretically, it is argued that by increasing the trade union density, and therefore the bargaining power of workers, they are in a better position to negotiate higher wages than those determined by competitive market. On one hand, better wages would lead to a reduction of the company profits; on the other hand, better wages would lead to a reduction in the level of utilization of the work force (Sachs and Larraín, 2002). Eventually, both effects would result in the need for businesses to hire fewer workers or dismiss some employees, thus increasing the rate of unemployment.

Empirical evidence indicates that an increase in the unemployment rate implies a rise in income inequality due to the fact that generally layoffs affect with greater force lower salary workers. Because
firms prefer to retain the more skilled workers, the wages of less-skilled workers fall in comparison to the first (Beyer, 1997; Saunders, 2002; Aboumoomi, 2003; Agenor, 2002; Mukoyama and Sahin, 2006). Under this perspective, wage bargaining by trade unions should increase income inequality.

However, the evidence for this relationship is not entirely clear. Some argue, that the coefficient that relates the union density with the unemployment rate is significantly greater than zero, like Chechhi and Garcia-Péñaloza (2010) studying the impact of labor market institutions on income distribution using a panel of OECD countries. This means that the greater level of unionized workers, the greater the unemployment rate. On the other hand, there are investigations such as Gustafsson and Johansson (1999) and OECD (2004), where there is no evidence of a relationship between these variables. Herzer (2014) presents the hypothesis that this is due to the fact that trade unions care both about wages and employment, therefore, in a market situation of imperfectly competitive labor there is an action margin for trade unions to raise wages without generating negative effects on the level of employment.

Related to this, there is a wide variety of studies advocated to investigate the relationship between trade union action and wage inequality. However, since a worker’s income includes her wage, there is a high positive correlation between wage inequality and income inequality (Chechhi & García Peñaloza, 2010). An investigation developed on ILO (2008) makes it clear that a high rate of trade union density, a high rate of collective bargaining coverage and a high level of coordination among bargaining structures are directly associated with lower levels of wage inequality.

Mishel (2012) attempts to study the effect of a decline in trade union action over the wage level of middle-class workers in the state of Michigan in United States. This for the period contained between the years 1973 and 2007. He argues that theoretically collective bargaining should reduce income inequality for three reasons:

1. The determination of wages in collective bargaining focuses on the establishment of a "standard wage" for similar jobs between different companies, and for individuals’ occupations within the same company. This benchmarking results in a smaller wage difference between workers and, consequently, less income discrimination against women and minorities.

2. Wage gaps between occupations tend to be lower when there is collective bargaining, so wages in occupations that tend to be poorly paid tend to be higher in a circumstance of trade union action.

3. Collective bargaining is more common among middle-class workers, therefore, collective bargaining reduces the wage gap between middle-class workers and high-income workers, who generally are
not benefited from collective bargaining.

In his study, Mishel finds that a decrease in collective bargaining explains slightly over a third of the rise in wage inequality among men in the state of Michigan (33.9%). The author concludes that the strengthening of collective bargaining and the recognition of its role in income distribution are necessary so that the benefits of economic growth can reach the majority of the population.

Card, Lemieux & Riddell (2004), in order to understand the trend of income inequality over time, studied the relationship between trade unionism and wage inequality in the case of United States, United Kingdom and Canada. Using microdata for two decades, the authors conclude that des-unionisation explains a substantial portion of the increase in wage inequality among men in the case of United Kingdom and the USA since the early 1980s.

In the same fashion, Antonczyk, Fitzenberger & Sommerfeld (2010) investigate the rising wage inequality, the decline of collective bargaining and the higher income inequality by gender between the years 2001 and 2006 in the case of Germany. Their results show that the sharp decline in the wage bargaining contributes substantially to the sharp increase in wage inequality, although they stress that there are other factors that have higher impact in this situation.

In the case of Latin America, Trajtemberg (2008) investigates the influence of trade unions over income distribution in Argentina. Using the survey, he can separate those worker who possess collective agreement from those who do not, and analyze variations in the wage level for both groups. The investigator concluded that the "strengthening of trade unions and the revival of labor institutions that protect the rights of workers are key pieces to promote more equal conditions". This is because average wage is greater in that group not ruled by the agreement, but also in that group are concentrated the higher wage differences.

In the case of Uruguay, Alves et al. (2012) studies the dynamics of inequality. They find that inequality starts to decrease in the beginning of 2007, and conclude that this is due to the fact that two years prior to this, centralized collective bargaining was restored after being suspended starting from the years 1990 and 1991, changing the labor market in a significant way. However, the authors suggest that because of the difficulty to isolate the effects of such a measure, it is not possible to verify formally the impact of the return of collective bargaining in the reduction of income inequality. According to international evidence, it might have compressed wage inequality at least among formal workers.
Among the few efforts that have been made to document the relationship between trade unionism and income inequality, one of the most important is that presented on Duran (2011). He estimates three models: the first, using cross section data for 49 countries, concludes that increasing the level of workers who negotiate collectively by 10 percentage points produces an improvement in the Gini index of 4.8%. For second model, they perform an analysis of time series for the case of Chile and New Zealand. This one concludes that an increase of 10 percentage points in the rate of unionization improves the Gini coefficient by 4.1%. Finally, the last model performs a simulation of the impact that variation in the structure of collective bargaining has over the Gini coefficient for Chile. The results indicate that an increase in the level of centralization of collective bargaining results in a rise the Gini coefficient. This means that the structure of collective bargaining, and therefore the structure of trade unions, has an effect over income inequality.

A more current research, Herzer (2014), performs an analysis of cointegration techniques of heterogeneous panels for a group of 20 countries. He finds that there is a negative long-term causal relationship between unionization and income distribution. However, he stresses that the effects do not have to be necessarily homogeneous across countries and believes that in some specific cases, as in the countries with high trade union density and high coverage of collective bargaining such as Belgium, Denmark, Finland, Norway or Sweden, the relationship is even positive, i.e. the greater the level of trade unionism, the greater the level of inequality. However, the author ignores the fact that its conclusion applies in the opposite direction, i.e. in countries with low union density and low coverage of collective bargaining, the effects of trade unionism on the inequality are negative. In spite of this, Herzer is emphatic that there are differences in other labor market institutions that may also explain the high heterogeneity of the effects.

In conclusion, we find ample evidence in the literature supporting the fact that trade union movement has an effect on wage inequality, and that this kind of inequality is highly correlated with inequality[EvP1]. However, according to mainstream economic theory, trade unionism and collective bargaining have a negative effect on the employment rate, and this should be related with a worst distribution of income. Therefore, the final effect of unionization on inequality is not clear and it will depend on which effect predominates. As indicated earlier, the main goal of this research is to study what is the final effect.
3 Methodological Approach

3.1 Empirical model

Using two methods, the econometric analysis will assess the relationship between unionization and income inequality. At first, it will be calculated by a fixed effects panel; subsequently, a cointegration model of panel data will be estimated in order to take into account a possible endogeneity and the possibility that the relationship between unionization and income inequality is not homogeneous across countries, as Herzer (2001) affirms.

3.2 Data

Regarding to the data used in this investigation, to measure the inequality level we use the Estimated Household Income Inequality (EHII) database, developed and presented by the Inequality Project of the University of Texas. The EHII index has the advantage of being comparable between countries and across time. It is measured in the format of the Gini index (on a scale from 0 to 100) and it is estimated through a regression between the Gini coefficients appearing in the Deininger-Squire database and the inequality measures of Theil that appears in the PICU-KINGDOM database, using the predicted values as estimates of the Gini coefficient. A clear limitation of the EHII index is due to the fact that it results from an estimation, it could have some kind of bias that affects the results.

The level of unionization will be approximated by the level of union density. This variable is obtained from the Institutional Characteristics of Trade Unions, Wage Setting, State Intervention and Social Pacts (ICTWSS) database compiled by Jelle Visser of Amsterdam Institute for Advanced Labour Studies of the University of Amsterdam. The rate of union density is defined as the net participation of those workers who are affiliated to a union as a proportion of the total workers employed.

In the case of Chile, the ICTWSS database does not report any data between that period defined between 1978 and 1986. Therefore, to standardize the series and include them in the study we uses the information reported in Island, Tarud and Jorquera (1978). This paper presents data about the unionization rate in the 1970s, missing data was estimated using a linear interpolation.

From the ICTWSS database we also get data about the collective bargaining level. This is defined as a categorical variable that reflects the predominant level in which collective bargaining is performed.
under the labor laws in force in each year. It takes the following values: (5) if collective bargaining is carried out predominantly at national or industrial level and there are legally binding rules negotiated at central level that must be respected at the time to negotiate to lower levels; (4) if collective bargaining is carried out predominantly at an intermediate level or it is alternating between central and industrial level; (3) if collective bargaining is carried out predominantly at industrial or sectoral level; (2) if collective bargaining is predominantly at sectoral or enterprise level; and finally, (1) if collective bargaining takes place predominantly at local or business level.

Beside Chile, the selected countries were Australia, South Korea, United States, Finland, Holland, Italy, Japan, Norway, Sweden and the United Kingdom. As cointegration methodology requires completely balanced panels, in the cases that there was no data were available, these were estimated. The result is a balanced panel consisting of 396 observations for 11 countries.

3.3 Fixed-effect panel model

The first model used in this study consists in a fixed effect’s panel, where we seek to analyze the impact of the unionization rate over income dispersion from an inter-temporal perspective. The dependent variable is the Gini Index, following the specifications formulated by Atkinson and Brandolini (2009) and Durán (2011).

In general, a panel model is specified as follows:

\[ Y_{i,t} = \beta_1 X_{i,t} + \gamma Z_i + u_i + \varepsilon_{i,t} \]

Where \( X_{i,t} \) is the vector of explanatory variables that have variations for both individuals and time, \( Z_i \) corresponds to the vector that contains variables that only vary at an individual level and not over time, \( u_i \) corresponds to a vector of individual level effects, and \( \varepsilon_{i,t} \) is the vector of errors of the model.

In a first approach we use a fixed effects regression because it allows to control or isolate the effect of omitted variables, under the assumption that they do not vary over time but they do vary between individuals. The model assumes that individual level effects are not random and are correlated with the independent variables of the model and, therefore, are presented as parameters that need to be
estimated, resulting in the following expression of the model:

\[ Y_{i,t} = u_i + \beta_1 X_{1i,t} + \beta_2 X_{2i,t} + \varepsilon_{i,t} \]

Particularly, for this investigation, the specification would be the following:

\[ Inequality_{i,t} = \alpha_i + \beta_1 DS_{i,t} + \beta_2 X_{2i,t} + \varepsilon_{i,t} \]

Where:

- \( Inequality_{i,t} \) = Estimated inequality coefficient for country \( i \) at period \( t \) on Ginin format.
- \( DS_{i,t} \) = Union density for country \( i \) at period \( t \).
- \( X_{i,t} \) = Vector of control variables for country \( i \) at period \( t \). In this case, it details the level at which collective bargaining is developed and the unemployment rate.

### 3.4 Controlling for endogeneity using Dynamic Ordinary Least Squares

#### 3.4.1 Panel data cointegration

When we proceed to estimate the previous model, we do not take into account the fact that the estimated parameter and the error term might be correlated, a phenomenon called endogeneity. This would mean that it is not possible to rule out that the impact measured by the fixed effects model is in fact the effect of inequality over unionization instead of the other way around.

One of the methodologies found in the literature to correct possible endogeneity is to take advantage of cointegration within panel data (Pedroni, 2007; Herzer and Vollmer, 2012; Kao et. al, 1999; Herzer, 2014; Herzer and Nunnekamp, 2012; Alfonso and Tovar, 2012). Cointegrated series are series that are non-stationary (meaning that they can be approximated by an stochastic process instead of a deterministic one), but the lineal combination of them is stationary. Cointegration leads to a long term relationship between variables, that is, as they grow in time, the error between them does not grow (Montero, 2013). This produces an extremely consistent estimator for the independent variable’s coefficient when using Ordinary Least Squares (OLS) as long as the variables are cointegrated.
To test the non-stationarity of union density and income inequality, we will be using a unit root test, since OLS requires us particularly that both series are cointegrated in first order I(1). Specifically, we will use the Im, Pesaran and Shin (2003) test that contrasts the null hypothesis that all panels have a unit root, against the alternative that at least one panel is stationary. This test is based on the regression of the Augmented Dickey-Fuller test, in the following way (Hoang and Mcnown, 2006):

\[ y_{i,t} = \alpha_i + \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} \theta_{i,j} \Delta y_{i,t-j} + \epsilon_{i,t} \]

Where \( \rho_i \) is the first order autoregressive parameter for country \( i \), where the t-statistic is obtained to test \( \rho = 1 \). The Im, Pesaran and Shin statistic, consists basically in the average of each of the sections of the cross country data that conform the panel:

\[ \bar{t}_{N,T} = \frac{1}{N} \sum_{i=1}^{N} t_{i,T} \]

The super-consistency of the estimators implies that they converge to their true value when the sample size increases more quickly than an ordinary OLS estimator, allowing the model to have more precise estimations than the alternatives. In addition, this estimators are robust in terms of endogeneity, since they maintain their qualities in the presence of correlation between the error term and the independent variables (Stock, 1987).

To test for cointegration we will use the method presented by Westerlund (2007), describing four tests where the null hypothesis is that the error correction term in a conditional error correction model is equal to zero, so that if the null hypothesis is rejected, the variables are cointegrated (Baltagi, 2008; Ruiz-Fuensanta, 2010). The statistical tests proposed by Westerlund are the following:

a) \( G_t = \frac{1}{N} \sum_{i=1}^{N} \frac{\hat{\alpha}_i}{SD(\hat{\alpha})} \)

b) \( G_\alpha = \frac{1}{N} \sum_{i=1}^{N} \frac{T\hat{\alpha}}{\hat{\alpha}_i} \)

c) \( P_t = \frac{\hat{\alpha}}{SD(\hat{\alpha})} \)

d) \( P_\alpha = T\hat{\alpha} \)
3.4.2 Dynamic ordinary least squares

As indicated above, when there is cointegration, the OLS estimator can be defined as “super-consistent”. However, its distribution does not tend to be standard because of the presence of finite sample bias. This can be caused both by the endogeneity of the independent variables or by the autocorrelation of the error term.

To solve this problem, the usual approach is Dynamic Ordinary Least Squares (DOLS), proposed by Stock and Watson in 1993 (Barcenillas et al., 2009; Olivera-Chaves et al., 2010; Herzog, 2014; among others). Broadly speaking, the DOLS model includes differences on the regressors to correct the endogeneity problem, and includes leads and lags of these differences to correct the residual autocorrelation.

Following the model suggested by Herzer (2014), we considered an equation that includes only the variables of unionization and inequality in the distribution of income, so the model takes the following form:

\[
\text{Inequality}_t = \alpha_i + \beta_i \text{DS}_{i,t} + \sum_{j=-q}^{p} d_{i,j} \Delta \text{DS}_{i,t} + \epsilon_{i,t}
\]

Where:
- \( \text{Inequality}_{i,t} \) = Estimated inequality coefficient for country \( i \) at period \( t \) on Ginin format.
- \( \text{DS}_{i,t} \) = Union density for country \( i \) at period \( t \).
- \( d_{i,j} \) = Operator of leads and lags.
- \( p \) = Lag length.
- \( q \) = Advancement length.

A weakness of the method of panel cointegration is that while it ensures a long-term causal relationship between two variables, it does not tell us anything about the direction of the causality. To determine the direction of the short-term causality between unionization rate and the indicator of inequality, an error correction model is estimated using the two-stage method proposed by Engel and Granger (1987). According to this method, the first step is to obtain the estimated residuals from the DOLS long-term equation to use them as a term of error correction (TEC) as follows:

\[
t\text{cee}_{i,t} = \text{Inequality}_{i,t} - [\hat{\alpha}_i + \hat{\beta}_i \text{DS}_{i,t}]
\]

11
The second step of this method is to estimate the dynamic error correction model defined by the following equations:

\[
\Delta \text{Inequality}_{i,t} = \alpha_{1,i} + \sum_{j=1}^{k_i} \beta_{1,1,i,j} \Delta \text{Inequality}_{i,t-j} + \sum_{j=1}^{k_i} \beta_{1,2,i,j} \Delta \text{DS}_{i,t-j} + \gamma_{1,i} \text{tce}_{i,t-1} + \mu_{1,i,t}
\]

\[
\Delta \text{DS}_{i,t} = \alpha_{2,i} + \sum_{j=1}^{k_i} \beta_{2,1,i,j} \Delta \text{DS}_{i,t-j} + \sum_{j=1}^{k_i} \beta_{2,2,i,j} \Delta \text{Inequality}_{i,t-j} + \gamma_{2,i} \text{tce}_{i,t-1} + \mu_{2,i,t}
\]

Where \(\gamma_{1,i}\) and \(\gamma_{2,i}\) are the coefficients that capture how inequality and union density adjust to any deviation of the long term equilibrium. So, to infer a causal long-term relationship, these coefficients must be negative and significant.

Regarding short-term causality, this can be inferred from the coefficients of the differences variables. In particular, if from the first equation we obtain that the \(\beta_{1,2}\) parameter is significant, it allows us to reject the hypothesis that in the short-term trade union density does not cause inequality. In the same way, if from the second equation we obtain that \(\beta_{2,2}\) is significant, we can reject the hypothesis that inequality does not cause union density in the short-term.

Following Ruiz-Fuensanta (2010), these equations will be estimated by the Pooled Mean Group (PMG) estimator proposed by Pesaran, Shin and Smith (1999). This method requires that the long-term effects must be equal for all the individuals in the sample, but relaxes this restriction in the case of short-term effects.

3.5 Data and descriptive statistics

This section presents descriptive statistics of the variables used in the study. The first thing to be considered is that one of the biggest limitations for the realization of this study was the low availability of panel data, both for unionization and income inequality. Considering this, the country selection was based mainly on availability of comparable data.

Table 1 shows descriptive statistics of the EHHI Gini Index for the 11 countries included in our sample, between the years 1978 and 2005. As we can see, Chile is the country with highest levels of inequality, followed by Japan and South Korea, while Sweden has the lowest level, followed by Finland and Norway.
<table>
<thead>
<tr>
<th>Country</th>
<th>Mean</th>
<th>Stand. Deviation</th>
<th>Max</th>
<th>Min</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>35.1</td>
<td>1.6</td>
<td>37.0</td>
<td>31.6</td>
</tr>
<tr>
<td>Chile</td>
<td>47.8</td>
<td>1.5</td>
<td>50.3</td>
<td>44.8</td>
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<tr>
<td>South Korea</td>
<td>38.1</td>
<td>1.1</td>
<td>40.0</td>
<td>36.6</td>
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<tr>
<td>United States</td>
<td>37.9</td>
<td>1.0</td>
<td>40.1</td>
<td>35.9</td>
</tr>
<tr>
<td>Finland</td>
<td>32.3</td>
<td>1.1</td>
<td>34.4</td>
<td>30.3</td>
</tr>
<tr>
<td>Holland</td>
<td>35.2</td>
<td>1.2</td>
<td>37.6</td>
<td>32.8</td>
</tr>
<tr>
<td>Italy</td>
<td>36.7</td>
<td>1.0</td>
<td>38.3</td>
<td>34.9</td>
</tr>
<tr>
<td>Japan</td>
<td>38.1</td>
<td>3.0</td>
<td>43.2</td>
<td>35.4</td>
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<tr>
<td>Norway</td>
<td>34.1</td>
<td>1.4</td>
<td>36.6</td>
<td>31.9</td>
</tr>
<tr>
<td>United Kingdom</td>
<td>34.5</td>
<td>2.1</td>
<td>37.2</td>
<td>29.5</td>
</tr>
<tr>
<td>Sweden</td>
<td>28.3</td>
<td>1.9</td>
<td>30.5</td>
<td>23.1</td>
</tr>
<tr>
<td>Overall</td>
<td>36.2</td>
<td>4.9</td>
<td>50.3</td>
<td>23.1</td>
</tr>
</tbody>
</table>

Table 1: Descriptive statistics of EHII Gini coefficient (1978-2005).

Figure 1 shows the dynamics of inequality in the sample countries during the same period. It follows from the figure that all countries, with the exception of South Korea and Sweden, increased their level of inequality, being the United Kingdom, Japan and Korea the ones that increased the most.

![Figure 1](image-url)

Table 2 shows how trade union density behaves in the sample countries in the same time frame available for the EHII Gini Index, being the Nordic countries the ones with higher rates of unionization, with United States, South Korea and Chile on the opposite end, having the lowest rates.
<table>
<thead>
<tr>
<th>Country</th>
<th>Mean</th>
<th>Stand. Deviation</th>
<th>Max</th>
<th>Min</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>39.7</td>
<td>9.8</td>
<td>49.8</td>
<td>24.4</td>
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<tr>
<td>Chile</td>
<td>17.3</td>
<td>4.1</td>
<td>27.7</td>
<td>13.1</td>
</tr>
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<td>South Korea</td>
<td>13.5</td>
<td>2.4</td>
<td>18.6</td>
<td>9.9</td>
</tr>
<tr>
<td>United States</td>
<td>16.0</td>
<td>3.6</td>
<td>25.0</td>
<td>11.8</td>
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<td>Finland</td>
<td>73.5</td>
<td>4.4</td>
<td>80.7</td>
<td>66.9</td>
</tr>
<tr>
<td>Holland</td>
<td>26.4</td>
<td>4.6</td>
<td>37.0</td>
<td>21.3</td>
</tr>
<tr>
<td>Italy</td>
<td>39.9</td>
<td>5.3</td>
<td>50.4</td>
<td>33.6</td>
</tr>
<tr>
<td>Japan</td>
<td>25.5</td>
<td>4.1</td>
<td>32.6</td>
<td>18.8</td>
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<td>Norway</td>
<td>56.5</td>
<td>1.6</td>
<td>58.5</td>
<td>53.9</td>
</tr>
<tr>
<td>United Kingdom</td>
<td>39.4</td>
<td>8.6</td>
<td>51.9</td>
<td>27.1</td>
</tr>
<tr>
<td>Sweden</td>
<td>81.0</td>
<td>3.4</td>
<td>87.4</td>
<td>75.6</td>
</tr>
<tr>
<td>Overall</td>
<td>39.0</td>
<td>22.5</td>
<td>87.4</td>
<td>9.9</td>
</tr>
</tbody>
</table>

Table 2: Descriptive statistics of union density (1978-2005).

In figure 2, we see the dynamics of trade union density. In this case, the majority of the countries in the sample saw a reduction in this variable. Both the cases of Australia and United Kingdom are particularly interesting: both have an important union tradition, having rates over 40% in the beginning of the 1970s, but saw this rates drastically decrease by 2005. Japan also had a significant reduction in its trade union density, while the Nordic countries maintained their rates relatively steady for the most part.

**Figure 2**

At first glance it is clear that, at least within this sample, there is a general correlation between rate of unionization and income inequality: countries that have a higher rate of the first have lower income inequality. Figures 3 and 4 are aligned with the previous reflection. The first figure pairs the averages of union density and inequality, showing clearly the trend indicated.

**Figure 3**

In figure 4 we show a country by country analysis, where it can be observed that for most countries the trend of unionization and the one for income inequality have opposite slopes, which would reflect a
negative relationship between unionization and inequality throughout time. Only for the Nordic countries and South Korea the relationship is not so clear. This could be due to the fact that effects of unionization on inequality are not homogeneous among the countries that are in the sample, as posed by Herzer (2014).

Figure 4

4 Results

4.1 Fixed-effects panel

To estimate the panel model that relates the union density with inequality, we use a fixed effects model. An advantage that this methodology provides is that it allow us to control for omitted variables, as long as these variables do not vary across time, only varying between units of observation: countries. To control for these effects we will use the unemployment rate for the year in question and a dummy variable indicating whether the collective bargaining is done at a level above firms. The robust correction in Stata is used to correct heteroskedasticity in the errors.

<table>
<thead>
<tr>
<th>Independent Variable</th>
<th>Fixed-effect estimator</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Union density</td>
<td>-0.149</td>
<td>0.034</td>
</tr>
<tr>
<td>Unemployment</td>
<td>0.256</td>
<td>0.035</td>
</tr>
<tr>
<td>Higher Level of negotiation to company</td>
<td>-0.755</td>
<td>0.220</td>
</tr>
<tr>
<td>Constant</td>
<td>40.390</td>
<td>0.000</td>
</tr>
</tbody>
</table>

Table 3: Summary of results of the fixed-effects panel model.

The previous table shows the results of the estimation with fixed effects. This tells us that on average, union density has a negative effect on inequality, in particular, that an increase of 1 base point in union density will lead to a reduction of 0.15 base points in inequality (99\% of confidence). The results also indicate the presence of a positive and significant relationship between the rate of unemployment and inequality to (99\% of confidence), specifically, that an increase of 1 base point on the unemployment rate diminishes in 0.25 base points income inequality. Finally, the dummy variable indicates that if collective bargaining is carried out at a level greater than the firm it results in a negative relationship
with inequality: the higher the level of collective bargaining, the lower the level of inequality. However, these results are not significant.

These results must be considered with caution, since this model does not take into account the possible endogeneity. It is entirely plausible that both union density and unemployment are simultaneously determined with inequality. This can lead to a bias in the fixed effects estimators. To correct this issue we can use co-integrated panel estimation, developed in the following section.

4.2 Model of dynamic ordinary least squares (DOLS)

4.2.1 Test for cointegration

To prove that there is co-integration between union density and inequality as variables, the first requirements that both series must meet are that both must be non-stationary and have a unit root. As we indicated in the methodological section, we use the Im, Pesaran and Shin (2003) statistical, taking the average of each of the cross country sections that conform the panel.

<table>
<thead>
<tr>
<th></th>
<th>Level</th>
<th>First Differences</th>
</tr>
</thead>
<tbody>
<tr>
<td>Inequality</td>
<td>Union density</td>
<td>Inequality</td>
</tr>
<tr>
<td>IPS with statistical trend</td>
<td>1.76</td>
<td>-0.86</td>
</tr>
<tr>
<td>P-value</td>
<td>0.96</td>
<td>0.19</td>
</tr>
</tbody>
</table>

Table 4: Summary of results of the unit root test of Im-Pesaran-Shin.

Table 4 shows us the results of the Im-Pesaran-Shin test for the original variables and the first differences. The selection of the optimum lag is based on the Akaike Information Criterion (AIC). When the analysis is done by levels, there is not enough statistical evidence to reject the null hypothesis: that the series are non-stationary. Furthermore, when we apply the test to the first difference of the variables, the null hypothesis is strongly rejected, confirming that the variables are cointegrated at the first order, and therefore, the first requirement is met.

Once we reject the hypothesis of non-stationarity, the next step is to prove that this series are co-integrated and therefore, have a relationship in the long run. To do this, we use the method proposed
by Westerlund (2007), since this method takes into account the possibility of heterogeneity between the different countries that conform the panel. The results of the test and the p-values are presented in the following table.

<table>
<thead>
<tr>
<th></th>
<th>Value</th>
<th>Z-value</th>
<th>P-value</th>
<th>P-value robust</th>
</tr>
</thead>
<tbody>
<tr>
<td>$G_t$</td>
<td>-5.42</td>
<td>-12.64</td>
<td>0.00</td>
<td>0.43</td>
</tr>
<tr>
<td>$G_{\alpha}$</td>
<td>-6.45</td>
<td>2.72</td>
<td>1.00</td>
<td>0.04</td>
</tr>
<tr>
<td>$P_t$</td>
<td>-6.34</td>
<td>0.78</td>
<td>0.78</td>
<td>0.01</td>
</tr>
<tr>
<td>$P_{\alpha}$</td>
<td>-10.60</td>
<td>-0.91</td>
<td>0.18</td>
<td>0.09</td>
</tr>
</tbody>
</table>

Table 5: Summary of results of the unit root test of Im-Pesaran-Shin.

It needs to be pointed out that for both the number of lags and the future values the selection criterion used was AIC. To obtain the robust p-value we used a bootstrap of a 100 iterations. Analyzing these p-values, to control for the possible transversal dependency of the data, we find that all but one statistical significantly reject the null hypothesis of no co-integration between the inequality indicator and trade union density. This allows us with a high degree of certainty reject the null hypothesis (Ruiz-Fuensanta, 2010), which implies that there is a long-term relationship between the variables. Therefore, by meeting the necessary requirements, we can proceed with the estimation based on co-integrated panels.

4.2.2 DOLS: General model and causality

To estimate the coefficient that quantifies the long-term relationship between union density and inequality, the model used is dynamic ordinary least squares. To test the robustness of the model, we propose several different equations, by varying the amount of future variables and lags. The results are presented in the following table.
In particular, we have that in average and for the countries in the sample, an increase by one percentage point in the rate of union density will result in a decrease of approximately 0.155 percentage points in the index of inequality. These results have a confidence level of 95%, so controlling for the endogeneity of the variables, the result holds: there is a relationship in the long-term that implies that higher rates of unionization is related to less inequality.

Despite the fact that the results of the cointegration tests and the DOLS model indicate that there is a long-term relationship between the rate of union density and the level of inequality, these tests do not provide us information to determine the causality of this relationship. To investigate the causality we proceed with a correction of errors model, following the two-stage procedure proposed by Engle and Granger (1987), as we describe in the methodological section. The correction error value is estimated from the two lag and one future variable DOLS model. Table 6 presents a summary of the results obtained for this model, estimated through the Pooled Mean Group (PMG) estimator, proposed by Pesaran et al. (1999), using AIC to select the optimal number of lags.

Table 6: Summary of results of the DOLS model.

| Lags | Advances | Coefficient | D.E. | Z   | P>|Z|   | 95% Confidence Interval |
|------|----------|-------------|------|-----|-------|------------------------|
| 1    | 1        | -0.159      | 0.043| -3.660 | 0.000 | -0.244 -0.074          |
| 2    | 1        | -0.158      | 0.045| -3.500 | 0.000 | -0.247 -0.070          |
| 2    | 2        | -0.156      | 0.047| -3.300 | 0.001 | -0.248 -0.063          |
| 3    | 2        | -0.154      | 0.049| -3.110 | 0.002 | -0.250 -0.057          |
| 3    | 3        | -0.151      | 0.052| -2.920 | 0.003 | -0.252 -0.050          |

Table 7: Summary of results of causality test.
The coefficients associated the error correction terms of the DOLS model are significant at the 1% in the two equations. This result indicates that on average, in the long term there is a two-way causality, which is consistent with the arguments raised by Herzer (2014). Sleekly, one could say that the bidirectional transfer implies that changes in inequality are a cause and an effect of union density. However, the results do not provide evidence in favor of short term causality in any way, since the coefficients associated with changes in union density and inequality as an independent variable are not significant.

4.2.3 Effects by country

When analyzing the individual effects, we find that these are heterogeneous among the selected countries in the study, just as Herzer (2014) enunciates. Table 7 shows the results of the coefficients calculated by dynamic ordinary least squares.

| Country       | Coefficient | D.E. | Z    | P>|Z| | 95% Confidence Interval |
|---------------|-------------|------|------|------|------------------------|
| Australia     | -0.07       | 0.06 | -1.05| 0.30 | -0.19 0.06             |
| Chile         | -0.22       | 0.08 | -2.61| 0.01 | -0.39 -0.05            |
| South Korea   | -0.15       | 0.16 | -0.95| 0.34 | -0.47 0.16             |
| The US        | -0.33       | 0.04 | -7.66| 0.00 | -0.41 -0.24            |
| Finland       | 0.20        | 0.06 | 3.15 | 0.00 | 0.07 0.32              |
| Holland       | -0.21       | 0.03 | -6.41| 0.00 | -0.28 -0.15            |
| Italy         | 0.01        | 0.09 | 0.07 | 0.94 | -0.18 0.19             |
| Japan         | -0.76       | 0.21 | -3.66| 0.00 | -1.17 -0.35            |
| Norway        | -0.39       | 0.21 | -1.89| 0.06 | -0.80 0.01             |
| United Kingdom| -0.18       | 0.04 | -4.18| 0.00 | -0.27 -0.10            |
| Sweden        | 0.46        | 0.08 | 5.62 | 0.00 | 0.30 0.63              |

Table 8: Descriptive statistics of EHII Gini coefficient (1978-2005).

While this results should be taken with caution, due to the limitations in the sample, they show that most countries present results as the ones we expected: raising unionization diminishes inequality. The countries that have the higher effects are Norway, United States and Chile, respectively. We have to
consider that, while Australia and South Korea have the right direction in their coefficients, their results are not significant. On an opposite direction, Sweden and Finland present the contrary relationship, a fact that is consistent with Herzer (2014). For Italy we find no relationship between unionization and inequality.

Herzer argues that a possible explanation for this heterogeneity in the results is the fact that in the countries that he finds a positive coefficient, trade union density and the range of collective bargaining are very high, resulting in possible rigidness in the labor market, pushing wages up, generating unemployment and therefore inequality. While this possibility cannot be ruled out, it’s fundamental to notice that this analysis doesn’t take into account the fact that the countries that present this positive correlation have consistently lower levels of inequality, compared to the rest of the studied countries.

As a consequence of the previous analysis, we could argue that the marginal effect that institutions in the labor market have over inequality might be decreasing, resulting in eventual negative effects in the long run for countries that have already reached a minimum level of inequality induced by unions. Since in this sample most of the countries have lower levels of inequality, raising trade union density would raise income equality.

5 Conclusions and policy implications

The main objective of this study is to determine the effect of unionization on inequality of income. Although the results should be viewed with caution due to the relatively short panel, these indicate that on average, a higher rate of unionization implies a lower level of income inequality. These results are robust when controlling for the possible endogeneity in the variables unionization rate and coefficient of inequality. Also we find that in the long run causality runs in both directions so you could say that in the long term, changes in inequality are the cause and effect of union density. The results also show that there is heterogeneity in the effects when analyzing individual countries, particularly countries with low inequality present a contrary to that expected. A contrary explanation to those presented in the literature is that the marginal effects of labor market institutions on the decline in inequality levels are decreasing, which could have negative effects once it has reached a level minimum of inequality induced by unions. Under this premise, for countries that have a low rate of unionization and a high level of inequality, there is enough space to increase unionization to reduce inequality without actually affect
the level of employment.

Analyzing the case of Chile and if we start from the basis that as a society we are averse to inequality, the results presented in this research motivate to discuss some aspects of the union sector institutions. Remember that the legal framework of trade union action is delimited by what was termed as the “Labor Plan” of 1979 which must be contextualized as part of a series of reforms neoliberal that took place during the military dictatorship which strongly restricted the trade union and the development of means of pressure in collective bargaining. The results suggest progress in incorporating workers unions, for which a policy of automatic enrollment could be useful.

Similarly further progress is needed in research with both unions and various other labor market institutions on levels of income inequality in each country.

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