Can Power Resources Theory Explain Variations in Labor’s Share of National Income?

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Abstract

This paper shows that labor’s share of national income – the part that does not go to capital incomes – tends to increase when union density increases, and fall when unemployment increases, in a sample of 16 rich countries 1960-2007. The paper also explores heterogeneity between countries and over time. It is shown that the effect of union density on labor’s share is not uniformly positive: in a few corporatist countries the effect is actually weakly negative. It is also shown that the effect has varied over time: from positive in the 1960s to negative in the 1980s, and then again weakly positive in the 1990s and early 2000s. Two other power resources variables tested, frequency of work conflict and unemployment insurance generosity however have no significant estimated effects on labor’s share.

Keywords: power resources theory, labor’ share, functional income distribution
I. Introduction

Power resources theory (PRT), developed in the 1980s by sociologist Walter Korpi, has been and is quite successful in explaining variations in welfare state generosity and income inequality between rich countries (Bradley et al, 2003; Brady, 2009; Volscho and Kelly, 2011). PRT holds that there is a distributive conflict in capitalist societies between capital owners and wage earners, as well as conflicts between different categories of wage earners depending on their amounts of human capital (Shalev and Korpi, 1980: 31-32; Korpi, 2006:172-173). The approach stresses the importance of collective organisation of wage earners in shaping welfare states and income distribution: more organisation in (left-wing) political parties and trade unions is associated with more generous welfare states and less income inequality. So far, it has been used most in comparative studies of welfare states, but also in studies of income distribution in terms of the gini coefficient or the distance between for example high income earners (the 90th percentile) and low (10th percentile) or medium income earners (50th percentile). However, there is one research question that PRT would seem to fit perfectly but hasn’t been much used for: the distribution of income between labour incomes and capital incomes, the so-called ‘functional income distribution’.

The doyen of PRT Walter Korpi has discussed the decreasing labor share in rich countries in the perspective of PRT (Korpi, 2002), but only in passing and descriptively, without econometric analysis. The perspective has been used to great effect by Tali Kristal in studies of labor’s share in Israel (Kristal, 2007) and in 16 rich countries (Kristal, 2010). Kristal has shown the importance of unionisation and strike activity as well as welfare state generosity for determining labor’s share. With a PRT perspective it’s a natural assumption that greater unionisation should be associated with a larger labor’s share, and that more generous welfare states should as well, since in theory a more generous welfare state strengthens labour’s bargaining position vis-à-vis capital.

This paper uses a dataset with sixteen rich countries in the years 1960-2007 to investigate whether variables chosen from a power resources perspective can in regression analysis explain the variations in labor’s share. It has already been shown by Kristal (2010) as well as other researchers that union density and some versions of welfare state generosity (Stockhammer, 2009) are positively associated with labor’s share. Therefore, this paper goes one step further by focusing on two types of heterogenous effects: heterogenous between countries, and heterogenous over time. We already know – and I will show it again in section III of this paper – that overall there is a positive association between for example union density and labor’s share in the most researched period, i.e. rich countries since 1960 or 1970. But which sub-effects drive this overall result? In other words, does unionisation really matter equally for labor’s share in each country, and in the same way over time? To investigate the first type of
question, heterogenous effects between countries, I use cross-validation as well as a random coefficients model (Beck, 2001; Beck and Katz, 2007). I show that in three countries – Canada, Norway and Sweden – the positive association that overall is to be found in the dataset between union density and labor’s share is not to be found. I hypothesise that this is due to “competitive corporatism”. To investigate the second type of question, heterogenous effects over time, I employ rolling regression (Baum, 2004), allowing coefficients to vary over time. I show that the effect of union density on labor’s share was at its most positive in the 1960s, then switched to negative in the 1980s and then again became weakly positive in the 1990s and 2000-2007.

The paper is structured as follows. Next section discusses why and how power resources variables should have effects on the functional income distribution, and which variables should be included. I present the paper’s hypotheses. The following section describes the dataset used, with 16 rich countries 1960-2007. Section IV describes the empirical analyses and the results. Section V concludes.

II. Power resources theory and labor’s share

Capitalist societies contain a power asymmetry between the working class and the capital owners, but in a democratic polity with capitalism – what Korpi calls “welfare capitalism” – workers can organise to even out power relations (Korpi, 2002: 369). Workers have more egalitarian preferences regarding income and wealth distribution than capitalists have, and so there is a distributive conflict between them (cf. Korpi, 2006: 172-3). In power resources research typically two arenas of working class activism are studied: politics and the labour market (cf. Hancké, 2012: 12). Through political activity and trade union activity, workers can shape the income and wealth distribution in a more egalitarian way. A large literature shows that this is indeed the case (Pontusson et al, 2002; Bradley et al, 2003; Kenworthy and Pontusson, 2005; Brady 2009).

Since power resources theory explicitly postulates a conflict of interest between workers and capital owners, it is evident that it the theory has implications for functional income distribution: from a PRT perspective, in countries with stronger unions, stronger left parties and more generous welfare states labor’s share of national income should be larger than in countries with weak unions and left parties and small welfare states. However, only recently has PRT been applied to functional income distribution in any thorough way. Korpi (2002: 405-8) has clarified that the profit share since 1960 in most rich countries has seen a U-shape: first a decrease with a profit squeeze in the 1970s, and thereafter a rapid increase in the late 1980s and 1990s. Korpi
interprets this development as resulting from power shifts between workers and capital owners and political shifts between the Left and Right. However, he did not examine the issue econometrically. I will here discuss three types of PRT variables in the context of labor’s share: union density, unemployment insurance generosity and other types of social spending, and unemployment levels. The discussion will lead to formulated hypotheses to test in section IV of the paper.

Union density

PRT claims that workers more organised in trade unions will have more influence on society including its income distribution. Therefore it’s an easy assumption to make that higher union density – meaning, a larger proportion of workers organised in unions – should be associated with a higher share for labor in national income. However, the logic is not that straightforward. A large literature on social pacts and wage moderation directs our attention to the issue that strong unions might actually consciously agree to wage moderation in exchange for job creation, welfare state expansion or some other desirable good (cf. Arpaia and Pichelmann, 2008: 32; Woldendorp, 2011). The economists Calmfors and Driffill in the 1980s (Calmfors and Driffill, 1988) claimed that wage restraint and non-inflationary wage policy will be easiest to achieve in countries with either very weak and decentralized unions, or with very strong and centralized ditto. From this I get two partially contradictory hypotheses on union density and labor’s share:

- Hypothesis 1a. Union density will be positively correlated with labor’s share
- Hypothesis 1b. The effect of union density will be weaker, and possibly negative, in countries with high density and high corporatism

There is already some evidence that can be interpreted in the direction of hypothesis 1b. Stockhammer (2009: 42-3) includes a dummy for Ghent countries – in which unions manage unemployment insurance so workers have strong incentives to join the union – and gets a positive effect of union density on labor’s share in non-Ghent countries, but no effect in Ghent countries. Stockhammer interprets this as that higher union density doesn’t mean as much for bargaining power in Ghent as in non-Ghent countries, since the average union member in a Ghent country is likely to be less ideologically motivated and militant than the average member in a non-Ghent country. However, I think that it could be the corporatist wage moderation side of high union density that explains the weaker effect on labor’s share in Ghent countries, rather than the Ghent system itself. Recently Hancké (2012) has pointed out that the fall in labor’s shares actually has been particularly severe in countries with strong unions. Hancké claims that this is because unions in these countries have been met by strong independent inflation-averse central banks that punish wage increases seen as inflationary. Because of this, unions in these countries internalise expectations on monetary policy in their wage policies, and strong unions do not lead to higher wage shares. There is a lot of merit to
Hancké’s argument, although I do not accept his conclusion that increased union density never is associated with higher wage shares; I argue, and show in section IV, that it depends on the country and the period. Hancké has data for 1980-1999; with a longer period, 1960-2005, Tali Kristal (2010) has shown that higher union density is indeed associated with larger shares of national income for labor. I extend Kristal’s study by considering differing effects between countries and over time.

There is at least one other hypothesis regarding union density to test. There has been much discussion on union decline; union density has fallen in all rich countries and unions’ influence as well. This is varyingly explained with globalisation strengthening employers’ bargaining power vis-à-vis workers, a turn to the right in politics that lends less attention to union demands today than before, and other factors. Baccaro (2008) has explored the possibility that trade unions might be less redistributive today than they used to be. Baccaro and Howell (2011) in a paper on the neoliberalisation of industrial relations also have explained how institutions – like unions and coordinated wage bargaining – might be converted so that institutions that on the surface look the same, have in function changed significantly. Baccaro and Howell (2011: 9) claim that coordinated wage bargaining once played a role for “political correction of market inequalities”, but nowadays is more likely to produce wage increases systematically below productivity growth, to foster competitiveness. Thus hypothesis 1c:

- Hypothesis 1c. The effect of union density will decline over time; in the 1990s and 2000s unions will not have as positive an effect on labor’s share as in the “golden years” in the 1960s and 1970s.

**Strike frequency**

Another indicator of workers’ mobilisation and militancy is strike frequency, typically measured as number of work days per a certain number of employees lost in a country in a year due to strikes. Of course the reason workers go on strike is to improve their employment conditions including wages, so there is a reason to believe that more strikes should be correlated with increases in labor’s share. However, it is not that simple; as Olin Wright (1984) and others have pointed out, it is also quite possible that workers will go on strike especially when their bargaining position is weak and they fail to reach their aims by other measures such as negotiations. Then strike frequency might actually be negatively correlated with changes in labor’s share, even though both movements in the variables would be caused by a third variable (bargaining power, unmeasured) rather than them causing each other’s movements. Still, the hypothesis is worth formulating:

- Hypothesis 2. Strike frequency should be positively correlated with changes in labor’s share
Unemployment insurance generosity and social spending

As stated above, the main focus of PRT when it comes to working class organisation and power is one the one hand unions, and on the other hand Left parties. Left parties are assumed to increase income redistribution and welfare generosity, benefitting workers. Two PRT variables that can be included in quantitative analysis of determinants of labor’s share are then proportion of cabinet seats held by Left parties, and government social expenditure, typically as percentage of GDP. However, I am a bit uncomfortable with these variables: they are quite far removed from labor’s share; what is the assumed smoking gun here? Regarding Left cabinet seats, there is obviously another variable inbetween the cabinet seats and labor’s share, and that variable is policies that the Left government enact. So, if it is possible to operationalise those policies, I think that it’s a preferable way of modeling the theorised relationship between Left government and functional income distribution. Social expenditure of GDP is a policy variable just like that. However, this variable has problems of its own as an indicator of worker-friendly policy (Korpi and Palme, 2003: 426, 432; Scruggs and Allan, 2006: 56). It fluctuates automatically with business cycles, without any political effort, with higher unemployment raising social expenditure and negative GDP growth increasing it since the measure is a percentage of GDP. A variable that does not vary with the economy like that, and is a more direct measure of government intent, is generosity of social insurance systems. This is measured as replacement levels: if a person becomes sick or unemployed, how large a proportion of his or her previous wage does s/he get from the social insurance system? Unemployment insurance obviously offers workers a “social wage” that sets a floor for real wages (cf. Korpi and Palme, 2003: 432). This variable is then theoretically a good measure of welfare policy’s effect on the functional income distribution: more generous unemployment insurance should strengthen workers’ bargaining positions.

Hypothesis 3. Unemployment insurance generosity will have a positive effect on labor’s share

As robustness checks, I will also use Left government and social expenditure as percentage of GDP as independent variables, but not in my main models.

Unemployment

The unemployment level (locally, nationally, globally – cf. Glyn, 2006) affects the bargaining relation between employer and employee. An employee in a situation with higher unemployment has a narrower set of options and must accept worse conditions. Thus, higher unemployment should have a negative effect on labor’s share, especially on the changes in the share.

Hypothesis 4. The unemployment level will have a negative effect on labor’s share
The macroeconomic hypothesis here, that has been presented before by for example Glyn (1997), also has a microeconomic foundation in the so-called wage curve. The wage curve is the negative correlation between local unemployment and local wage increases, a correlation described by Blanchflower and Oswald (2005) as “an empirical law of economics”. Wages increase slower when unemployment is higher since competition between workers about the jobs is tougher. Unemployment levels should therefore have a negative effect on the change in labor’s share rather than on the levels; the model specifications in this paper focus on changes (cf. De Boef and Keele, 2008: 188; Podesta, 2006).

III. Data

This paper studies 16 rich countries: Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Ireland, Italy, Japan, the Netherlands, Norway, Sweden, the United Kingdom and the United States.

From the preceding section follows that four independent variables are necessary: union density, strike frequency, unemployment insurance generosity and unemployment level. All these variables are included in the dataset in one form or other. Union density is fetched from the Amsterdam-based dataset ICTWSS, administered by Jelle Visser (2009). It has no missing values for the 16 countries and 47 years that I am working with here. Strike frequency is more difficult to measure: the measure that I have located is actually work days lost due to conflict days per 100,000 workers. This includes lockouts as well as strikes. The measure comes from the Comparative Political Dataset administered by Klaus Armingeon and colleagues (see Armingeon et al, 2009). This variable has clear problems: for example the conflict frequency in France is understated since the data only includes legal conflicts while the majority of French strikes have been wildcat strikes. Unemployment insurance generosity data access is also problematic. The OECD has a dataset called “The OECD summary measure of benefit entitlements, 1961-2007”, but it does not seem trustworthy, at least not for older figures. For Sweden in 1961 the OECD dataset states that the unemployment insurance replacement rate for an average production worker was 5.4 percent of his or her previous wage, which can be compared with that the Swedish-produced dataset SCIP, produced by people more familiar with the Swedish situation, states 63.5 percent for the average production worker in Sweden in 1961. Since the OECD dataset can not be trusted, instead I use data calculated by Lyle Scruggs of the University of Connecticut. His unemployment insurance replacement rate measure also focuses on the replacement rate for a person with an average production workers’ wage. This has data for most countries from 1971 to 2002 or 2003. One exception is the UK where the variable goes
back to 1963. In total the replacement rate variable has 525 country-years (see Scruggs and Allan, 2006 for discussion on the data).

<table>
<thead>
<tr>
<th>Table 1. Description of dataset</th>
<th>Source</th>
<th>N</th>
<th>Mean</th>
<th>Minimum</th>
<th>Maximum</th>
<th>Standard deviation</th>
</tr>
</thead>
<tbody>
<tr>
<td>Labor’s share</td>
<td>AMECO</td>
<td>768</td>
<td>69.21</td>
<td>49.23</td>
<td>81.44</td>
<td>5.10</td>
</tr>
<tr>
<td>Union density</td>
<td>ICTWSS</td>
<td>768</td>
<td>42.97</td>
<td>7.62</td>
<td>83.86</td>
<td>18.60</td>
</tr>
<tr>
<td>Strike frequency</td>
<td>CPD</td>
<td>754</td>
<td>16.50</td>
<td>0.00</td>
<td>197.90</td>
<td>26.91</td>
</tr>
<tr>
<td>UI replacement rate</td>
<td>Scruggs</td>
<td>525</td>
<td>61.15</td>
<td>6.70</td>
<td>94.30</td>
<td>16.01</td>
</tr>
<tr>
<td>Unemployment</td>
<td>OECD</td>
<td>768</td>
<td>5.56</td>
<td>0.49</td>
<td>17.15</td>
<td>3.50</td>
</tr>
<tr>
<td>Openness</td>
<td>PWT</td>
<td>758</td>
<td>60.82</td>
<td>9.27</td>
<td>184.30</td>
<td>32.30</td>
</tr>
<tr>
<td>GDP growth per capita</td>
<td>PWT</td>
<td>757</td>
<td>2.76</td>
<td>-8.94</td>
<td>13.09</td>
<td>2.69</td>
</tr>
<tr>
<td>Logged productivity per worker</td>
<td>PWT 7.1</td>
<td>758</td>
<td>10.77</td>
<td>9.36</td>
<td>11.49</td>
<td>0.33</td>
</tr>
</tbody>
</table>

The unemployment rate has been fetched from the OECD. No missing values.

The regressions in this paper also includes control variables. The first control variable is the country’s trade openness, measured as exports+imports as share of GDP. The idea is that in a globalised world the bargaining power of workers in rich countries has been eroded by increasing competition for investment with workers in poor countries. The measure comes from Penn World Tables and the only missing values are Germany 1960-1969. GDP growth per capita is included in the difference regressions because labor’s share tends to fall in years with high growth, and conversely rise in years with slow growth. This is because wages increase at a steadier pace than productivity growth, so in years with unexpectedly high productivity growth wages don’t rise as much as productivity, but on the other hand wages tend to increase faster than productivity in years with slow growth (cf. Kristal, 2007: 7). Productivity per worker – GDP per worker – is included to control for decreases in labor’s share in national income due to increasing capital intensity in production (Blanchard, 1997). The measure comes from Penn World Table and is PPP Converted GDP Chain per worker at 2005 constant prices.

The relationships, in terms of correlation, between the variables are stated in table 2 below.
Now that the dataset is in place, the empirical investigation can begin.

### IV. Empirical investigation

The main focus of this paper is variation in effects of power resources variables on labor’s share between countries and over time. However, to be able to discuss these two issues, first it should be established that there is an overall association between the variables to begin with. Thus, this section begins with basic models that do not allow for varying coefficients between countries and over time. With these models I show that union density is positively associated and unemployment levels negatively associated with labor’s share. However, for unemployment insurance generosity and strike frequency I do not get significant results. After establishing these main results, this section goes on with random coefficients models and rolling regression, allowing effects of the independent variables varying first between countries and then over time.
**Basic models**

The de facto standard model in comparative political economy is a fixed effects model with a lagged dependent variable included among the right hand side variables (Plümper et al, 2005; Wilson and Butler, 2007).

\[ Y_{it} = \rho Y_{it-1} + \beta X_{it} + e_i + \epsilon_{it} \]

Country fixed effects are used since the units in comparative political economy – countries – are highly heterogenous entities where we do not have measurements for everything that we want to control for. It is then handy to be able to use a country dummy variable to control for non-observed heterogeneity (Beck, 2001: 284; Halaby, 2004: 522-3; Plümper et al, 2005: 330-4; Adolph et al, 2005; Wilson and Butler, 2007: 106). With country fixed effects the estimated coefficients are just for explaining within-country variation, so the estimator is also called a within estimator. In comparative political economy the focus used to be on explaining between-country variation, but the interest in explaining variation within countries over time has increased as the time series have grown longer (cf. Pontusson, 2007: 326, 330). The lagged dependent variable is de facto standard since the dependent variables in comparative political economy tend to be highly autoregressive. When a lagged dependent variable is included among the independent variables the \( \beta \)'s for the other variables must be interpreted differently compared to a usual static model (cf. Shalev, 2007: 285). Since the lagged dependent variable (LDV) in itself is determined by previous values of the independent variables, the LDV soaks up effects that are really by the independent variables. The \( \beta \) coefficient for the independent variables should be interpreted as the immediate and short-run effect, and the long-run effects is given by \( \beta / (1 - \rho) \) (Williams and Whitten, 2011: 577).

I also estimate first differences specifications. Their form is:

\[ \Delta Y_{it} = \alpha + \beta \Delta X_{it} + \epsilon_{it} \]

Hence, the change in labor’s share is determined by an intercept, effects of changes in the independent variables, and errors. With a differences specification, we only get information about the short-run effects of the independent variables. I still use differences specification, since I believe that the variables strike frequency and GDP growth will have short-run effects on changes in labor’s share but no long-run effect, on the level of labor’s share. The FD specification also gets rid of nonstationarity problems that can lead to spurious results (Kittel and Winner, 2005: 278).

In table 3 below we see the result for four basic specifications. There are within specifications with and without unemployment insurance generosity – as I have explained in section III with unemployment insurance the dataset only covers 1971-2003 and so I estimate specifications with and without that variable separately. And there are first differences specifications with and without unemployment insurance. Within parantheses are panel-corrected standard errors.
Table 3. Basic models. Estimated with OLS, panel-corrected standard errors within parentheses

<table>
<thead>
<tr>
<th></th>
<th>Without unemployment insurance</th>
<th>With unemployment insurance</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Model 1. Within</td>
<td>Model 2. First differences</td>
</tr>
<tr>
<td>Intercept</td>
<td>0.18 *** (0.07)</td>
<td>0.18 *** (0.08)</td>
</tr>
<tr>
<td>Lagged dependent variable</td>
<td>0.91 *** (0.02)</td>
<td>0.82 *** (0.03)</td>
</tr>
<tr>
<td>Union density</td>
<td>0.02 ** (0.01)</td>
<td>0.12 *** (0.04)</td>
</tr>
<tr>
<td>Strike frequency</td>
<td>0.00 (0.00)</td>
<td>0.00 (0.00)</td>
</tr>
<tr>
<td>Unemployment rate</td>
<td>-0.12 *** (0.02)</td>
<td>-0.31 *** (0.06)</td>
</tr>
<tr>
<td>Unemployment insurance generosity</td>
<td>-0.01 (0.01)</td>
<td>0.01 (0.01)</td>
</tr>
<tr>
<td>Openness</td>
<td>-0.02 *** (0.01)</td>
<td>-0.07 *** (0.01)</td>
</tr>
<tr>
<td>GDP growth</td>
<td>-0.18 *** (0.02)</td>
<td>-0.18 *** (0.02)</td>
</tr>
<tr>
<td>Logged productivity per worker</td>
<td>0.11 (0.33)</td>
<td>-0.11 *** (0.02)</td>
</tr>
<tr>
<td>Adj R²</td>
<td>0.88</td>
<td>0.31</td>
</tr>
<tr>
<td>N</td>
<td>743; n=16, t=38-47</td>
<td>743; n=16, t=37-48</td>
</tr>
</tbody>
</table>

*** significant at 99 % level; ** significant at 95 % level; * significant at 90 % level

As we can see, union density as expected has a positive and significant effect in all four specifications. Unemployment as expected has a negative and significant effect. Likewise, labor’s share as expected decreases in years with high GDP growth, and as a level is lower with higher capital intensity/productivity with worker, at least in the specification with unemployment insurance included. The coefficients of strike frequency aren’t significant in either of the first differences models. This is however not so surprising considering the data problems and that strikes might actually be most common in years with weak bargaining positions for workers.

The graph below shows effects of the independent variables in model 1, with 95 percent confidence intervals. The confidence interval for productivity is huge; the effect is not significant. The confidence intervals for the other three variables on the other hand all stay clear of the 0 line; trade openness and unemployment have significant negative effects and union density a significant positive effect.
Robustness checks/alternative specifications

In section II I discussed public social expenditure as share of GDP and Left party participation in government as two potential political power resources variables. I dismissed them both in favour of unemployment insurance generosity as a better measure of political support for workers’ bargaining power. However, in the models above unemployment insurance generosity did not achieve any significance. In this section I re-estimate models 1 and 2 from table 3 above, adding either social expenditure or left government. The social expenditure variable comes from the OECD and has data for all 16 countries 1980-2007, except for Austria and Norway where the years 1981-84 and 1986-89 are missing. It should be noted that when using this variable the sample is just above half the size of the sample used in the original models, and that this changes the meaning of the coefficients: now we are not looking for example at the effect of union density all the way from 1960 to 2007, but just in the 1980s, 90s and 00s. The left government variable comes from Comparative Political Dataset I and has information for all the years 1960-2007 for all 16 countries. It measures the percentage of cabinet seats held by Left parties during the year, weighted by number of days. The final alternative specification is one with a slightly different dependent variable. Several economists have remarked that labor’s share is a imperfect measure for workers’ share of national income, since the labor’s share measure includes salaries and renumerations for corporate officers, CEOs and the like, who are in a class schedule actually closer to being capitalists than workers, even though they receive labour income (cf. Krueger, 1999; Buchele and Christiansen, 2003; Kristal, 2010). One way of correcting for this is
to take labor’s share and deduct the incomes of the 1 percent who earn the most from this share. The assumption is that this top percentile encompasses the CEOs and lawyers whose class interests are closer to capital owners than to workers, and that after deducting their incomes from labor’s share we get something approximating a “workers’ share”. I do this adjustment using top percentile income data from Atkinson, Piketty and Saez’ Top Incomes Database. This has data for 13 of my 16 countries, with a varying amount of years; the model using this dependent variable becomes seriously unbalanced, with t varying from 11 to 47 for the 13 countries.

The results for the three alternative specifications – including social expenditure in a within and a first differences model, the same with left government, and finally a model with workers’ share as dependent variable – are shown in table 4 below.

Table 4. Robustness checks. Panel-corrected standard errors within parantheses

<table>
<thead>
<tr>
<th></th>
<th>Within with social expenditure</th>
<th>Within with left government</th>
<th>FD with social expenditure</th>
<th>FD with left government</th>
<th>“Workers’ share” as DV. Within</th>
</tr>
</thead>
<tbody>
<tr>
<td>Intercept</td>
<td>0.01 (0.01)</td>
<td>0.18 *** (0.07)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Lagged dependent variable</td>
<td>0.80 *** (0.03)</td>
<td>0.91 *** (0.02)</td>
<td></td>
<td></td>
<td>0.88 *** (0.03)</td>
</tr>
<tr>
<td>Union density</td>
<td>0.03 (0.02)</td>
<td>0.02 ** (0.01)</td>
<td>0.10 ** (0.05)</td>
<td>0.12 *** (0.04)</td>
<td>0.04 *** (0.02)</td>
</tr>
<tr>
<td>Social expenditure</td>
<td>0.03 (0.05)</td>
<td></td>
<td>0.36 *** (0.09)</td>
<td></td>
<td></td>
</tr>
<tr>
<td>Left government</td>
<td></td>
<td>0.00 (0.00)</td>
<td></td>
<td>-0.00 (0.00)</td>
<td></td>
</tr>
<tr>
<td>Strike frequency</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Unemployment rate</td>
<td>-0.20 *** (0.05)</td>
<td>-0.12 *** (0.02)</td>
<td>-0.56 *** (0.08)</td>
<td>-0.31 *** (0.06)</td>
<td>-0.13 *** (0.03)</td>
</tr>
<tr>
<td>Openness</td>
<td>-0.04 *** (0.01)</td>
<td>-0.02 *** (0.01)</td>
<td>-0.07 *** (0.01)</td>
<td>-0.07 *** (0.01)</td>
<td>-0.03 *** (0.01)</td>
</tr>
<tr>
<td>Logged productivity</td>
<td>-1.86 * (1.01)</td>
<td>0.08 (0.32)</td>
<td>-17.19 *** (3.19)</td>
<td>-11.13 *** (2.18)</td>
<td>-0.44 (0.56)</td>
</tr>
<tr>
<td>Adj R²</td>
<td>0.91</td>
<td>0.91</td>
<td>0.42</td>
<td>0.31</td>
<td>0.85</td>
</tr>
<tr>
<td>N</td>
<td>434; n=16, t=20-28</td>
<td>743; n=16, t=38-47</td>
<td>421; n=16, t=20-28</td>
<td>743; n=16, t=37-48</td>
<td>496; n=13, t=11-47</td>
</tr>
</tbody>
</table>

*** significant at 99 % level; ** significant at 95 % level; * significant at 90 % level
Overall the results hold up: union density has a positive and significant effect in four of these five models, and unemployment rate and openness have negative and significant effects in all five. The important exception to note here is that union density loses its significance in the within model including social expenditure among the covariates. Is this a sign that I have omitted variable bias in my main models, union density there unjustly absorbing effects really from the not included variable social expenditure? No, that’s not the case. First, note that social expenditure is not significant in the model. Two, note that since social expenditure data is only available after 1980. And as I will show below in the section on varying effects over time, the positive effect of union density on labor’s share is weaker (or non-existent) in the 1980s than in especially the 1960s. Therefore the results are different for a model using the years 1980-2007 than for one with 1960-2007.

Varying effects between countries

In comparative political economy studies using regression analysis, too little attention has been paid to heterogeneity between the units being studied. The panel data methods used in CPE are borrowed from economics where typically the units are anonymous individuals that in themselves aren’t interesting; what the economists are interested in are overall patterns. However, in CPE this is not the case. The units are typically, as in this study, countries, with all of their specific histories and ways of functioning. Therefore, more attention should be paid to unit heterogeneity; as the sociologist Michael Shalev (2007: 264; cf. Western, 1998) states in an article critical of the use of regression analysis in CPE, “most producers and consumers of comparative political economy are intrinsically interested in specific cases. Why not cater to this interest by keeping our cases visible?”. One reason that CPE hasn’t lived up to this is that the regression techniques for varying slopes etc are not very developed and practitioners have been uncertain about their meaning. The leading political science methodologist Nathaniel Beck (2001: 286) stated ten years ago that it was uncertain how well random coefficient models would work with larger T as in CPE. Beck and Katz (2007: 184) more recently pointed out that random coefficient models are under-used in CPE and actually work better than practitioners may think. I estimate two random coefficient models – where separate coefficients for the independent variables are estimated for each unit, in this case country – below to investigate heterogeneity in union density’s effects on labor’s share.

First, however, it makes sense to also use a more “low-tech” (Beck, 2001: 291) technique to look at country heterogeneity: jackknifing, i.e. estimating the base models repeatedly (in my case 16 times) while excluding one country at a time. This is a robustness check, in the sense that we can see if an outlier country causes the coefficient of some independent variable to be overstated (cf. Kittel and Winner, 2005: 285). But it’s also interesting from a heterogeneity perspective: with which exclusion do the
coefficients inflate and deflate? In table 5 we see the coefficients for model 1 and model 2 from table 3, estimated 16 countries, each time excluding a country from the sample.

<table>
<thead>
<tr>
<th>Table 5. Jackknifing</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
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<tr>
<td></td>
</tr>
<tr>
<td>Lagged dependent variable</td>
</tr>
<tr>
<td>Union density</td>
</tr>
<tr>
<td>Strike frequency</td>
</tr>
<tr>
<td>Unemployment rate</td>
</tr>
<tr>
<td>Openness</td>
</tr>
<tr>
<td>GDP growth</td>
</tr>
<tr>
<td>Logged prod. per worker</td>
</tr>
</tbody>
</table>

The results are very robust to this test; no coefficients switch signs or even change very much. We see that the coefficient for union density is strongest when Finland (in the within model) or Norway or Sweden (in the first differences model) are excluded; this is a hint that in these Nordic countries with high unionisation, the effects of union density on labor’s share are relatively weak.

Next step is to estimate random coefficients models\textsuperscript{viii}. Their form is: \[ Y_{it} = \rho Y_{it-1} + \beta_iX_{it} + c_i + \varepsilon_{it} \]

As can be seen, I still use a lagged dependent variable on the right hand side, as well as country-specific intercepts. What’s new is the i-subscript to the betas: that is, the coefficients for the independent variables are now allowed to vary between countries. I am using the RCMs as diagnostics rather than as the main analysis, so I only report the coefficients for union density, the power resources variable that I have got consistently significant results for. In table 6 we see country-specific coefficients for union density for the 16 countries, both within-models, the first without and the second with
unemployment insurance as an independent variable (so the first includes years 1960-2007 and the second 1971-2003).

Table 6. Country-specific coefficients for union density, from random coefficient models

<table>
<thead>
<tr>
<th>Country</th>
<th>Model 1 as RCM</th>
<th>Model 3 as RCM</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>0.23</td>
<td>-0.32</td>
</tr>
<tr>
<td>Austria</td>
<td>0.23</td>
<td>0.36</td>
</tr>
<tr>
<td>Belgium</td>
<td>0.24</td>
<td>0.22</td>
</tr>
<tr>
<td>Canada</td>
<td>-0.03</td>
<td>-0.18</td>
</tr>
<tr>
<td>Denmark</td>
<td>0.13</td>
<td>0.26</td>
</tr>
<tr>
<td>Finland</td>
<td>0.11</td>
<td>0.45</td>
</tr>
<tr>
<td>France</td>
<td>0.24</td>
<td>0.80</td>
</tr>
<tr>
<td>Germany</td>
<td>0.01</td>
<td>0.01</td>
</tr>
<tr>
<td>Ireland</td>
<td>0.23</td>
<td>-0.02</td>
</tr>
<tr>
<td>Italy</td>
<td>0.17</td>
<td>0.22</td>
</tr>
<tr>
<td>Japan</td>
<td>0.34</td>
<td>0.76</td>
</tr>
<tr>
<td>Netherlands</td>
<td>0.14</td>
<td>0.17</td>
</tr>
<tr>
<td>Norway</td>
<td>-0.43</td>
<td>-0.34</td>
</tr>
<tr>
<td>Sweden</td>
<td>-0.05</td>
<td>-0.16</td>
</tr>
<tr>
<td>United Kingdom</td>
<td>0.10</td>
<td>0.17</td>
</tr>
<tr>
<td>United States</td>
<td>0.03</td>
<td>0.12</td>
</tr>
</tbody>
</table>

In the first model, three countries actually have negative coefficients for union density: more union members are not associated with larger labor shares. These three countries are Canada, Norway and Sweden. Norway and Sweden, two Nordic countries with high union density, are usual suspects. In the second model, including unemployment insurance generosity among the covariates, the same three countries plus Australia and Ireland get negative coefficients for union density. Australia is rather unexpected but for Ireland it can be said that even though the country’s union density is rather low, it is still a corporatist country (cf. Woldendorp, 2011; Hancké, 2012: 3).

Is the effect of union density on labor’s share overall less positive in countries with high union density? This is, recall the discussion in section II, hypothesis 1b of this paper. In the figure below I have plotted the coefficients from the first random coefficient model above against the mean union density for the country in the entire period.
Indeed, overall the relationship is negative but the difference is not statistically significant, which is not surprising given that N here is 16). I interpret this as tentative support for hypothesis 1b.

Beck and Katz (2007) point out that just because it’s possible to model TSCS models with full flexibility – allowing country-specific coefficients for all the independent variables – we don’t have to use all this flexibility. Rather, if we only except one independent variable to have different effects between countries – like union density here – it’s more parsimonious with a model where only that coefficient is allowed to vary. In table 7 below are results for two linear mixed models, fitted by maximum likelihood. It should be noted that we have here moved from working with ordinary least squares to maximum likelihood and that the country-specific intercepts are now “random” rather than “fixed” effects.
Table 7. Linear mixed models fitted by maximum likelihood with random effects of union density and country-specific intercepts. Standard errors in parantheses

<table>
<thead>
<tr>
<th></th>
<th>Same covariates as model 1</th>
<th>With unemployment insurance</th>
</tr>
</thead>
<tbody>
<tr>
<td>Lagged dependent variable</td>
<td>0.94 (0.01)</td>
<td>0.89 (0.01)</td>
</tr>
<tr>
<td>Union density</td>
<td>0.01 (AUS), 0.02 (AUT), 0.06 (BEL), 0.01 (CAN), 0.01 (DK), 0.00 (FIN), -0.00 (FR), 0.00 (GER), 0.05 (IRL), 0.01 (ITA), 0.04 (JP), 0.01 (NL), -0.03 (NO), 0.00 (SWE), 0.02 (UK), 0.02 (USA)</td>
<td>0.03 (AUS), 0.04 (AUT), 0.11 (BEL), 0.01 (CAN), 0.01 (DK), -0.00 (FIN), 0.01 (FR), -0.01 (GER), 0.12 (IRL), 0.02 (ITA), 0.11 (JP), -0.00 (NL), -0.06 (NO), 0.00 (SWE), 0.04 (UK), 0.04 (USA)</td>
</tr>
<tr>
<td>Unemployment insurance generosity</td>
<td>-0.00 (0.01)</td>
<td>-0.00 (0.01)</td>
</tr>
<tr>
<td>Unemployment rate</td>
<td>-0.14 (0.02)</td>
<td>-0.17 (0.02)</td>
</tr>
<tr>
<td>Trade openness</td>
<td>-0.01 (0.00)</td>
<td>-0.02 (0.00)</td>
</tr>
<tr>
<td>Logged productivity</td>
<td>0.48 (0.08)</td>
<td>0.89 (0.12)</td>
</tr>
<tr>
<td>Number of observations</td>
<td>N=743</td>
<td>N=525</td>
</tr>
<tr>
<td>Goodness-of-fit</td>
<td>AIC=2604, BIC=2641</td>
<td>AIC=1915, BIC=1953</td>
</tr>
</tbody>
</table>

In the first linear mixed model, the zero or negative coefficients for union density are for Finland, France, Germany, Norway and Sweden. In the linear mixed model including unemployment insurance, the countries are Finland, Germany, the Netherlands, Norway and Sweden. The difference, then, is that France is in the first group and the Netherlands in the second; the other four countries are the same. Recall that in the random coefficients model – where all coefficients were allowed to vary, not only union density like in these ones – the countries with negative or zero effects for union density were Canada, Norway and Sweden in both specifications, plus Australia and Ireland in the specification including unemployment insurance. Norway and Sweden are the two countries that persistently turn out zero or negative effects.

**Varying effects over time**

In a study with panel data (or more specifically, time-series-cross-section data; cf. Beck, 2001), the effects of the independent variables on the dependent variable may not only vary between units but also over time. This kind of variation is also highly interesting in comparative political economy; the context for for example a trade union to do their stuff is obviously very different over time in a sample that stretches from the middle of “the golden age of capitalism”, 1960, to the early 21st century just before the financial crisis, 2007. Plümper et al. (2005: 345-6) put it bluntly: “it is a heroic assumption that parameter values remain constant over time”. To investigate varying effects over time,
estimate a so-called rolling regression version of the basic model (model 1 from table 3).

Rolling regression (or moving windows regression) means estimating a regression which moves through the sample’s time dimension from the beginning to the end, in each step considering a specified number of years, and calculating separate coefficients for the independent variables for each step. In this way, we can find out whether the effects of the independent variables vary over time (cf. Pontusson, 2007: 328; Kwon and Pontusson, 2010). I have estimated model 1 from table 3 above as a rolling regression. The coefficients produced are country-specific but these coefficients, specific both to period – I have used a moving window of seven years – and country, are very unstable and estimated with very large errors; Zanin and Marra (2012: 93) state that “This methodology suffers from several well-known problems, which include the problematic window size choice that typically results in unstable parameter estimates.”. Therefore I have averaged the sixteen coefficients (one per country) for each window and this average is shown in the figure below.

The line begins in 1967 since my sample begins in 1960 and I have set the window width to seven years; at each point the estimated coefficient is for the seven year period ending this very year. So the point estimate for 1967 is a coefficient for 1960-67, and so on. We see that mainly the coefficient moves trendless around 0; recall that the total coefficient – hiding all this heterogeneity – for model 1 in table 3 is 0.02. Hypothesis 1c
of this paper is that the effect of union density should have decreased over time, but this does not fit with the evidence in the graph. There is one fairly long period where the average coefficient consistently stays below zero, but this is the eighties, or 1981-1990 to be more precise. According to my model, the time in my sample where union density did the least to increase labor’s share of national income was in the 1980s; not later, in the 1990s or 2000s as I expected. (Note, though, that the confidence interval for the RCM estimates is very wide and actually at every point includes 0.) The average coefficients noted per decade are 0.41 in the 1960s, 0.00 in the 1970s, -0.63 in the 1980s, 0.35 in the 1990s and 0.30 in 2000-2007.

V. Conclusions

The hypotheses of the paper were:

- Hypothesis 1a. Union density will be positively correlated with labor’s share
- Hypothesis 1b. The effect of union density will be weaker, and possibly negative, in countries with high density and high corporatism
- Hypothesis 1c. The effect of union density will decline over time; in the 1990s and 2000s unions will not have as positive an effect on labor’s share as in the “golden years” in the 1960s and 1970s.
- Hypothesis 2. Strike frequency should be positively correlated with changes in labor’s share
- Hypothesis 3. Unemployment insurance generosity will have a positive effect on labor’s share
- Hypothesis 4. The unemployment level will have a negative effect on labor’s share

Hypothesis 1a has been confirmed. Union density has a quite robustly positive and significant effect in the model specifications tried in this paper. However, this is not a novel finding; Kristal (2010) has shown the same thing. Hypotheses 1b and 1c, on varying effects of union density between countries and over time, are more original. I would say that hypothesis 1b has got tentative support: the jackknife analysis as well as the random coefficient models used have shown that overall the effect of union density is the strongest when Nordic countries (with high union density) are excluded, and that the estimated effects in at least two of the Nordic countries, Sweden and Norway, actually are negative. I have not found support for hypothesis 1c; surprisingly, the rolling regression shows that the effects of union density on labor’s share in my specification are most negative in the 1980s but then come back to being positive in the 1990s and 2000s. Hypothesis 2 has not found support; strike frequency has not become significant in the regressions. However this is not very surprising considering both the theoretical uncertainty about this variable (see discussion in section II) and the
measurement issues (see section III). Hypothesis 3 has not found support, surprisingly. Estimated effects of unemployment insurance have not been significant in a single specification tried in this paper. Two alternative operationalisations of political effort in favor of workers, public social expenditure as percentage of GDP and left government, were mere “successful” in getting the expected positive and significant effects on labor’s share, but I would still say that the theoretical foundation for including these variables is weaker than for unemployment insurance generosity. Hypothesis 4 has been confirmed: in all specifications tried in this paper the estimated coefficients for unemployment level are negative and significant, much in line with a “industrial reserve army” argument where greater unemployment creates pressure on workers’ wages.
References

Data

AMECO. http://ec.europa.eu/economy_finance/db_indicators/ameco/index_en.htm
Penn World Table 7.1. http://pwt.econ.upenn.edu/

Literature


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i For economic-theoretical discussion on why workers’ organisation should be able to have an effect on functional income distribution, see Blanchard (1997) and Harrison (2002). The basic idea is that if competition is not perfect – and in real markets it seldom is – than there will be “rents” in firms, and the distribution of these rents depend on relative bargaining power of employers and employees. Relative bargaining power is affected by union density, unemployment levels and the other variables used here.

ii Hancké’s stress on monetary policy is important; the effects of monetary policy on income inequality is overall an under-researched problem (cf. Coibion et al., 2012).

iii More or less the “usual suspects” in comparative political economy. Spain is excluded since I do not have union density data for that country until 1981. New Zealand is excluded because I do not have labor’s share data for that country until 1986.

iv I used the Hausman test to test for use of random effects vs fixed effects. It dismissed random effects.

v A precondition for this kind of explanation to work is that there is enough variation in the variables within countries over time. In many cases in comparative political economy we are interested in the effects of institutional variables that change very slowly, and then there’s a problem with using within estimators. In my case however the period is long enough and the variables change enough so that this shouldn’t be a problem. For example, the dependent variable labor’s share has a within standard deviation of 4.26 in the sample and only 2.9 between sd. Union density on the other hand has more variation between than within countries: the between standard deviation is 17.56 and the within is 7.5.

vi I have inspected the autocorrelation functions and partial autocorrelation functions for labor’s share and the country-wise series shows a pattern conforming to an AR(1) data generating process: a “damped sine wave or exponential decay” in the autocorrelation function and a ”single positive spike at lag 1” in the partial autocorrelation function (cf. Cowpertwait and Metcalfe, 2009: 79-81).

vii The use of “corrected” or “robust” standard errors has caught some flak recently by eminent methodologists (King and Roberts, 2012; Beck, 2012) who argue that these corrections are really just sweeping problems under the rug instead of modelling the complexities of the data. However, using them is still standard and shouldn’t at least do harm. I have used Beck and Katz-corrected standard errors from the R package plm.

viii Estimated by the command pvcm from the R package plm. For description see Croissant and Millo (2008: 19-21).


x For discussion of random versus fixed effects intercepts see Bartels (2008), Bell and Jones (2012) and Clark and Linzer (2012).
I have used the Stata command rollreg, created by Christopher Baum. The procedure is described in Baum (2004).