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Increased Price Markup from Union Coordination – OECD Panel Evidence

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Abstract Although coordination of wage bargaining probably affects entry barriers and competition in product markets, research on price determination has typically not considered such factors. We argue that the price markup depends on wage setting institutions and present empirical evidence in form of estimated price equations in a panel of 15 OECD countries. The estimates show that consumer prices may be as much as 21 percent higher in coordinated compared to uncoordinated countries, solely due to the effect of coordination on the markup. Since other studies find that coordination has a dampening effect on wages, this may explain why there is no clear effect of coordination on unemployment.

JEL C23, E31, J51
Keywords Imperfect competition model; price markup; labor market institutions; unemployment; panel data model

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1 Introduction

Labor market characteristics like unemployment benefits, employment protection, union power, union coverage and union coordination have received considerable attention in the economic and political literature, since they offer a potential explanation for the relative low unemployment rate in the US and the persistently high unemployment in many European countries. Hence, there has been a substantial amount of empirical research aiming at identifying the effect of such factors on the unemployment level. See for example Layard et al. (1991), Addison and Grosso (1996), Bleaney (1996), Nickell (1997), OECD (1997, 1999), Siebert (1997), Scarpetta (1996, 1998), Blanchard and Portugal (1998), Elmeskov et al. (1998), Nickell and Layard (1999), Blanchard and Wolfers (2000), Daveri and Tabellini (2000), Bertola et al. (2000, 2002), Nickell et al. (2001), Baker et al. (2002), Chen et al. (2002), IMF (2003), Belot and van Ours (2004) and Nickell et al. (2005).

One of the main results in these studies is that countries with a high degree of coordination of wage bargaining are associated with the best macroeconomic performance and the lowest unemployment when controlling for other factors. However, the evidence is inconclusive when it comes to identifying the effect of coordination at an intermediate level relative to a low level. Some of the studies find a hump-shaped relationship, i.e. that coordination at an intermediate level produces the highest unemployment, yet others find a monotonically decreasing relationship between coordination and unemployment. Hence, it is commonly concluded that the effect of coordination on unemployment is ambiguous.

The underlying assumption in the studies listed above is typically that coordination only affects wage setting, and that the effect of coordination on nominal wages is unclear as well. However, since all the studies only consider the effect of coordination in wage bargaining on unemployment, they cannot tell us anything about the causal relationship between the two variables. The ambiguity should instead lead us to investigate weather coordination has a separate effect in the determination of prices as well.

Coordination in wage bargaining may be characterized both as coordination between unions and employer confederations. There may be at least three theoretical justifications for including coordination as a separate explanatory variable in the price equation. First, as coordination between unions and employer confederations increases, firms realize that their competitors suffer from the same rise in wages. Hence, pass-through to prices may tend to be faster and more comprehensive, increasing the average level of the markup. Second, since wage bargaining implies profit sharing (see Layard et al. 1991, Ch. 2), both unions and employers may wish to create entry barriers for new competitors. As coordination increases the unions and employer confederations may be more successful in creating these entry barriers, especially if they have political power. There may be several other motivations for the unions to fight new entrants as well. For example, the unions may want to protect their existing members against potential joblessness (insiders vs. outsiders), or they want to protect themselves in case they suspect that new firms primarily want to hire non-unionized workers. Moreover, unions may fear that new firms will challenge several of the rights the unions have accomplished throughout the years of bargaining.

Third, other types of entry costs in a highly unionized labor market may exist as well, i.e. the cost of negotiating rigorous tariff agreements. In completely coordinated regimes unions face a dilemma; protecting their "rights" and fight new entrants or enhance employment growth and welcome new entrants. However, the sustainability of the regime requires that only new entrants that are "union-friendly" will be welcomed.

Price formation has been subject to some research, although, interestingly, there seems to be no consensus about which variables determine prices at the aggregate level.[1] Nevertheless, there has been little or no research linking institutional factors such as trade unions and the role of coordination of wage bargaining explicitly to price determination. This paper aims to correct for this deficiency in the literature.

There has been some work studying the role of institutions in wage setting for prices implicitly in reduced-form-equations. However, this literature assumes that the role of institutions enters through their effect on nominal wages. Sen and Dutt (1995) set up a theoretical model where bargaining power affects wages and where the wage level influences the markup. Furthermore, Bowdler and Nunziata (2005) conduct an empirical analysis where they replace unit labour cost in a price equation with wage determining factors. Layard et al. (1991) consider the issues of wage bargaining and trade unions in some detail, but without linking coordination

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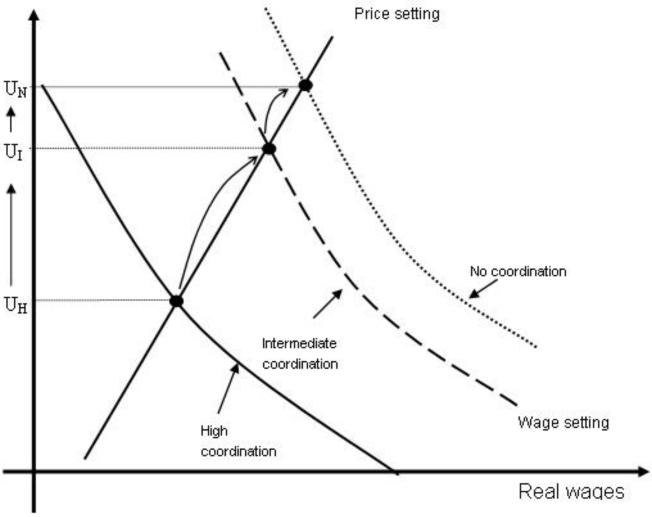
[1] See inter alia, Price (1991), Martin (1997) and Ashworth and Byrne (2003).

of wage determination and price determination. These studies have given valuable insight to the concepts of corporatism, trade unions and nominal inertia plus its effects on unemployment.

The existence of involuntary unemployment as a steady state phenomenon is probably best understood within the framework of wage and price curves put forward by e.g. Layard et al. (1991). Figure 1 gives an illustration of the equilibrium relationship between real wages and unemployment in this model framework.

Figure 1: Equilibrium in the Labor Market, the Effect of Coordination of Wage Bargaining for the Wage Curve



The two solid curves show the anticipated level of real wages in wage setting and price setting, respectively, under the assumption that wage setting is highly coordinated. The wage curve is decreasing in real wages and unemployment because unions, which can determine or at lest influence on nominal wages, must make a trade-off between higher real wages and lower employment or lower real wages and higher employment. In the theoretical literature, the price curve is often assumed to increase in real wages and unemployment since the markup over unit labor cost tends to be low, and hence real wages high, when demand (and employment) is weak. Equilibrium unemployment is U_H, indicating a real wage level consistent with steady state in both the labor market and the product market.

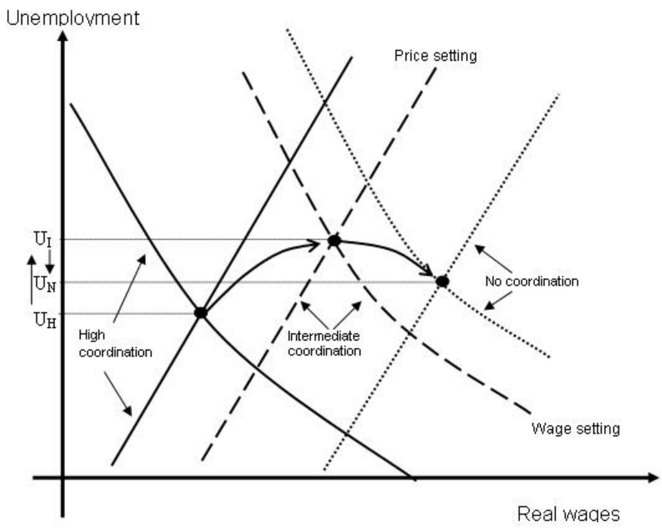
Moving now to more uncoordinated wage setting regimes, the figure illustrate the common assumption that coordination only affects wage setting. The dashed and dotted wage setting curves indicate shifts in the level of coordination to an intermediate and low level, respectively. The shifts imply a higher real wage as wage setting becomes de-coordinated. As long as we keep the price setting curve independent of coordination, we see that a lower degree of coordination will monotonically increase unemployment as well, to U_I when coordination is at an intermediate level and U_N when there is no coordination.

If we, as we do in Figure 2, also assume that the price-setting depends on the degree of coordination in wage setting consistent with the hypothesis in this article, the effect on unemployment may be ambiguous. As we move from a highly coordinated regime to an intermediately and further to an uncoordinated regime, real wages increases monotonically. However, while unemployment increases from U_H when wage setting is highly coordinated to U_I at an intermediate level of coordination, unemployment is reduced to U_N when wage-setting is de-coordinated further.

Figure 2 indicates that real wages may increase monotonically with de-coordination, while the relationship between coordination and unemployment is unambiguous. Nunziata (2005)—which to our knowledge is the only published article where the effect of coordination is estimated in a multi-country wage bargaining mode—tested the humped shaped relationship between coordination and real labor cost on a panel of 20 OECD countries. He could not reject a monotonically decreasing relationship between real labor cost and the degree of coordination.[2]

[2] Podrecca (2004) analysed the importance of labour market institutions in wage equations for 20 OECD countries as well. However, she did not test for a separate effect of coordination alone, but whether union coordination made a difference on the wage responsiveness to changes in unemployment. She found no significant effect.

Figure 2: Equilibrium in the Labor Market, the Effect of Coordination of Wage Bargaining for the Wage Curve and the Price Curve



Moving from an uncoordinated wage setting to an intermediately coordinated wage setting may leave workers worse off, if real wages decrease and unemployment increase. This may seem paradoxical and raises the question why workers should want to unionize, but is indeed the essence of the issue at hand; as strong, rational unions/workers fail to incorporate the adverse macroeconomic effects of their increased wage claims at intermediate coordination. The workers would be even worse off if workers in other industries joined industry-wide unions and they did not. As coordination increases from an intermediate to a high level, firms are perceivably better off due to higher price markup and lower product real wages. Workers must necessarily also be better off, even though they experience a decrease in real wages, if not they would not coordinate. Hence, the benefits of increase in employment outweigh the adverse effects in real wages.[3]

[3] There might be other advantages for workers in coordinated regimes as well, such that equality and worker-friendly social reforms.

The empirical evidence presented in this paper is founded on panel data for 15 OECD countries observed from the 1960s to 2000. We use an index developed by Kenworthy (2001) as an indicator of the level of coordination of wage bargaining. The main finding is that coordination significantly increases the level of consumer prices. An increase in the wage coordination index from the lowest level (1) to the highest level (5) will induce a long-run price level increase of 21 percent according to the estimates in our baseline model.

The remainder of the paper develops these points and is structured as follows. First, in Section 2, we set up an imperfect competition model for price determination, where the markup depends on coordination of wage bargaining and the relationship between import prices and prices on domestically produced goods and services. Then, after discussing the econometric methodology and data in Sections 3 and 4 respectively, we estimate the model in Section 5. Conclusions are drawn in Section 6.

2 Literature and Economic Theory

Price determination is essential to the understanding of complex issues of unemployment and inflation. The increasing attention of policymakers regarding inflation targets in monetary policy has necessitated further research in this field. Nevertheless, there is surprisingly little literature on the role of labor market institutions on price setting behavior.

In the price determination literature there are two central theories - purchasing power parity theory (PPP or the law of one price) and the pricing-to-market or the markup theory (see e.g. Dornbusch 1987; Krugman 1987; Froot and Klemperer 1989). The purchasing power theory argues that domestic prices are determined in the long run by world market prices, thus it emphasizes the importance of import prices.[4] Although the professional opinion concerning the validity of PPP seems to have shifted several times in the post-war period, the main conclusion from the recent literature appears to be that PPP may be viewed as a valid long-run international parity condition, and that mean reversion in real exchange rates exhibit significant nonlinearities (Taylor, Peel and Sarno (2001)). The markup

[4] See Dornbusch (1992) for further details.

theory developed by Michal Kalecki states that producers set prices as a markup over average variable costs[5], partly as an insurance against variability in input prices and partly to earn greater profits. This latter theory has become a widely accepted approach to price determination, and has, besides being a standard assumption in many macroeconomic models[6], also been subject to considerable empirical testing.[7] In, inter alia, Isard (1977), Frenkel (1978), Richardson (1978), McKinnon (1979), Bruce and Purvis (1985), Giovannini (1988) the determinants of the price level have been analyzed thoroughly. As the theory of price determination is central in the conduct of macroeconomic policy (see e.g. Blanchard and Kiyotaki 1987, Ball et al. 1988, Layard et al. 1991, Dixon and Rankin 1994), the need for consensus about the determinants for the aggregate price level is obvious. However, the lack of consensus is apparent.[8] Both the PPP and the markup theories are utilized in the literature, and even combinations of the two. Moreover, most empirical studies draw on average or other easily obtainable cost measure as a substitute for the hard-to-quantify marginal cost.

Martin (1997) advocates, through his theoretical and empirical assessment of the UK economy, that reality is somewhat in the middle of the two central price theories. He argues that interactions between domestic and foreign agents cannot be neglected when it comes to theories of price formation, thus both domestic costs and import prices are important in determining the domestic price level. Ashworth and Byrne (2003) and Asteriou et al. (2002) specify and estimate price equations of OECD countries. The estimations in both papers put a roughly 50-50 percent weight on unit labor costs and import prices in the price equation, respectively, which corresponds relatively well to the estimates in single-country studies as well (see inter alia Boug et al., 2006, and Bårdsen et al., 1998).

We utilize the markup theory in this paper; by including an index for the coordination of wage determination, we attempt to investigate whether there is an omitted variable bias in the standard price determination model. To our knowledge

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[5] See Kalecki (1943).

[6] See for example Weintraub (1958), Okun (1981), Dutt (1990), Taylor (1991) or Dornbusch and Fischer (1994).

[7] See for example Scherer and Ross (1990).

[8] For surveys see e.g. Gordon (1981, 1990), Domberger and Smith (1982), Carlton (1989), Hay and Morris (1991), Layard et al. (1991).

there has been no attempt to study the role of labor market institutions for the price setting behavior. However, some studies aim at describing their effect on prices in reduced-form-equations, i.e. inserting for wage determining factors in price-equations.

Sen and Dutt (1995) set up a theoretical model where bargaining power affects wages and where the wage level influences the markup. Their idea was closely related to ours, since they believed bargaining power would influence the markup. However, we argue that the level of coordination has a separate effect on the markup, and not only through the wage level. Therefore we control for unit labour cost in the empirical model below. Furthermore, Bowdler and Nunziata (2005) conduct an empirical analysis where they replace unit labour cost in a price-equation with wage determining factors. They do find effects of wage setting institutions, but the interpretation is still different from the findings in the present paper. Layard et al. (1991) consider the issues of wage bargaining and trade unions in some detail, but without linking coordination of wage determination and price determination. Nevertheless, these studies have given valuable insight to the concepts of corporatism, trade unions and nominal inertia plus its effects on unemployment.

2.1 The Theoretical Foundation of the Estimated Equation

A formal derivation of the price equation is now presented. The aim of this section is to show how coordination in wage bargaining may enter the price equation and thereby provide a starting point for an econometric specification of the price equation with coordination as a separate explanatory variable. First, the standard consumer price relation (1) is put forward. *CP* is consumer prices, *DP* is prices on domestically produced goods and services and *PI* is price of imports. The subscript *i* denotes country.

$$CP_i = DP_i^{\beta} \cdot PI_i^{1-\beta} \tag{1}$$

The weight on domestic prices is assumed to be a constant β . Further, according to the markup theory, domestic prices are set as a markup over marginal costs. Unit labor costs (*ULC*) are used as a proxy for marginal costs[9], such that:

$$DP_i = (1+m_i)\cdot ULC_i \; . \tag{2}$$

Layard et al. (1991, p. 338) show that the markup, (1 + m_i), depends on the elasticity of demand with respect to own price, i.e. that (1 + m_i) depends on product-market competitiveness. In aggregated price-equations the markup is usually assumed to be either a constant or, in an open economy setting, dependent only on the relationship between import prices and prices on domestically produced goods and services:

$$(1+m_i) = e^{\tilde{m}_{0i}}\left(\frac{PI_i}{DP_i}\right)^{m_{1i}} , \tag{3}$$

As argued in above, we wish to explore whether coordination of wage bargaining influence competition in the product market. Since $\tilde{m}_{0i}$ reflects the degree of competition, we therefore need to specify the markup as a function of the degree of coordination, as justified previously. Exactly how coordination affects the markup (e.g. linearly or non-linearly) is without prior knowledge an empirical question. Here, we approximate this theoretical link by defining $\tilde{m}_{0i} = m_{0i} + m_2 CO_i$, where CO_i is an index for the degree of coordination of wage bargaining and m_2 is the effect of an increase in the index on the markup. Hence, we may rewrite (3) as:

$$(1+m_i) = e^{m_{0i}+m_2\cdot CO_i}\left(\frac{PI_i}{DP_i}\right)^{m_1} , \tag{4}$$

where we have assumed that $m_{1i} = m_1$ and $m_{2i} = m_2$, for all *i*.

Inserting the expression for $(1+m_i)$ given by (4) in (2) and rearranging we obtain:

9 With a Cobb–Douglas production technology unit labour costs are proportional to marginal costs.

$$DP_i = \left(e^{(m_{0i} + m_2 \cdot CO_i) \cdot \frac{1}{1+m_1}} \right) \cdot \left(PI_i^{\frac{m_1}{1+m_1}} \right) \cdot \left(ULC_i^{\frac{1}{1+m_1}} \right) \tag{5}$$

Inserting the equation for domestic prices (5) into the consumer price relation (1) yields:

$$CP_i = \left(e^{\cdot \frac{\beta(m_{0i} + m_2 \cdot CO_i)}{1+m_1}} \right) \cdot \left(PI_i^{\left(1-\beta+\frac{\beta \cdot m_1}{1+m_1}\right)} \right) \cdot \left(ULC_i^{\left(\frac{\beta}{1+m_1}\right)} \right) \tag{6}$$

By taking logs of equation (6) we obtain (henceforth lower case letters denote natural logarithms):

$$cp_i = \alpha \cdot m_{0i} + \delta \cdot CO_i + \alpha \cdot ulc_i + (1-\alpha) \cdot pi_i \, , \tag{7}$$

where $\alpha = \dfrac{\beta}{1+m_1}$ and $\delta = \alpha \cdot m_2$.

Hence, the level of consumer prices depends on unit labor costs (*ulc*), the price of imports (*pi*) and the degree of coordination of wage bargaining (*CO*). Equation (7) may serve as a starting point for an econometric specification. However, when estimation the relationship between the variables one must take into account the typical dynamic properties of macro data. Therefore, (7) will serve as the long run relationship between the variables when we specify the dynamic price equation in next section.

2.2 An Econometric Specification of Price Setting

We estimate the price setting using an error correction model (see Sargan 1980), where parameters related to the short-term price growth and the long-run price level given in equation (7) are estimated simultaneously. Hence, the following shows consistency between the error correction model employed and the economic theory derived in the previous section.

The error correction model used in the analysis is given by:

$$\Delta cp_{it} = \tau_i + \gamma_t - \beta_{1i}(cp_i - ulc_i)_{t-1} - \beta_{2i}(ulc_i - pi_i)_{t-1} + \beta_{3i}CO_{it-1} \tag{8}$$
$$+ \beta_{4i}\Delta pi_{it} + \beta_{5i}\Delta ulc_{it} + \beta_{6i}\Delta cp_{it-1} + \beta_{7i}z_{it} + \varepsilon_{it},$$

where τ_i is a fixed country specific effect, γ_t is a specific time dummy and ε_{it} is an error term. The β's are non-negative parameters and z_{it} is a vector containing all other variables and dummies. Δ denotes the first difference of a variable.

The link to equation (7) is further illustrated by calculating the steady-state solution for (8). Let steady state be defined as:

$$\Delta ulc_{it} = \omega_i,$$
$$\Delta pi_{it} = \rho_i,$$
$$\Delta\omega_{it} = \Delta\rho_{it} = \Delta^2 cp_{it} = \Delta z_{it} = \Delta CO_{it} = \varepsilon_{it} = 0,$$

i.e. that unit labor costs, import prices and consumer prices grow according to constant rates, and that z and coordination remain constant. For simplicity, suppose also that $\gamma_t = 0$ in steady state. We can then express the steady state level of consumption prices in the following way:

$$cp_i - [\alpha_i \cdot ulc_i + (1-\alpha_i)\cdot pi_i] = \gamma_{0i} + \delta_i \cdot CO_i \tag{9}$$

where

$$\alpha_i = 1 - \frac{\beta_{2i}}{\beta_{1i}},$$
$$\gamma_{0i} = \frac{1}{\beta_{1i}}[\tau_i + \beta_{7i}z_i + (\beta_{5i} + \alpha_i\beta_{6i} - \alpha_i)\omega_i + (\beta_{4i} + (1-\alpha_i)\beta_{6i} - (1-\alpha_i))\rho_i], \text{ and}$$
$$\delta_i = \frac{\beta_{3i}}{\beta_{1i}}.$$

Equation (9) is the long-run solution for the estimated price equation, and it is the empirical counterpart to equation (7). The detailed empirical results and interpretations are presented in Section 5.

3 Data Appreciation

The bulk of the data in the paper is retrieved from the OECD databases[10]. Our original data set consists of 20 OECD-countries. However, we did not have coordination data for Portugal and Spain, and for Switzerland, New Zealand and Germany the data had less data points than the number of parameters to estimate. Therefore, we have been able to estimate price equations for only 15 OECD-countries. The complete data set is documented in Appendix A.

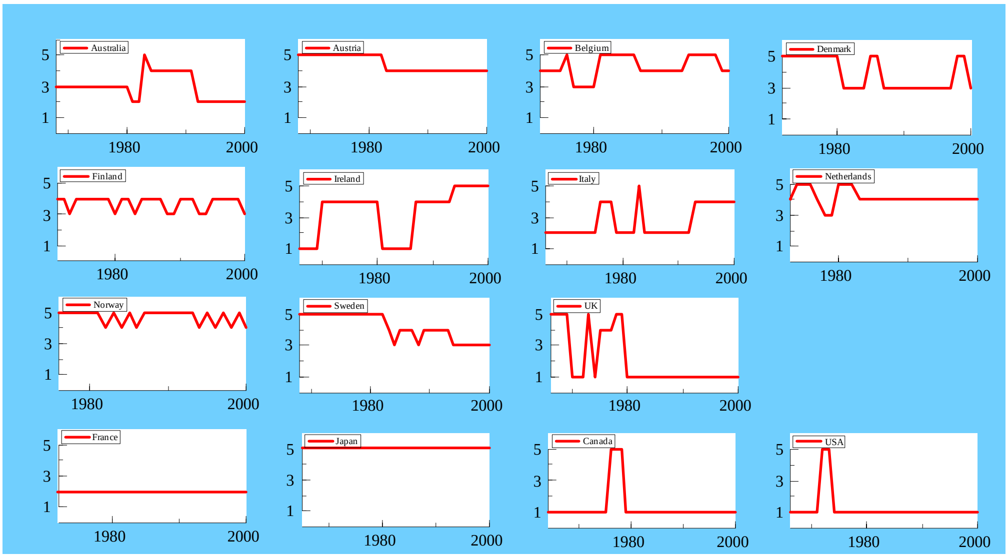
The import prices are constructed by taking the ratio of the value and the volume of imported goods and services. While in almost all previous papers predominantly data for the manufacturing sector has been used, we use the OECD unit labor cost index that covers the whole economy. As our macroeconomic data is from the OECD databases, this should aid consistency in terms of weights, definitions and comparability across OECD member countries for the series used in this analysis.

The coordination variable (*CO)* is retrieved from Professor Lane Kenworthy's dataset (Kenworthy 2001), and constructed as an index from 1 to 5. The index comprises both coordination between both labor unions and employer confederations. The index draws on a variety of much cited references from the wage-setting literature, and is elaborated on in Appendix A. The Thatcher regime in the United Kingdom in the 1980s is an illustrative case, which entailed a strong decentralization of the wage-setting process, bringing the score for the UK from an intermittent high level down to 1 from 1980 onwards. Figure 3 shows the coordination scores for each country. Four countries have little or no variation in the coordination scores. The four countries are France and Japan, which have a constant *CO* variable, and Canada and the US, which have coordination scores equal to 1 in all years except during the price and wage legislation in the 1970's. In the analysis *CO* is held constant also in Canada and the US since changes in consumption prices obviously would be influenced by the price controls and not necessarily the change in wage setting. Hence, in our panel regression with fixed effects these four countries do not contribute to identify the effect of coordination. Obviously, it is always desirable to acquire better and more robust data, and economic and political indicators such as the coordination index utilized in this

[10] OECD *Economic Outlook* and *Main Economic Indicator* (MEI) databases.

paper are inherently subjective and as such subject to measurement errors (Kenworthy 2003). Nevertheless, it serves well as an indicator of the degree of coordination in the various economies, and is in our view a richer and more extensive index than the previous attempts to measure coordination.[11]

Figure 3: Coordination Scores, 1-5



Source: Kenworthy (2001).

4 Econometric Issues

The strength of the evidence of an econometric model relies heavily on the error terms being independently and identically distributed. There are several potential sources of misspecification that have to be examined. In this section we address these issues.

[11] See e.g. Soskice (1990), Iversen (1999), Traxler, Blashke, and Kittel (2001), Golden, Wallerstein, and Lange (1997), Ferner and Hyman (1998).

4.1 Unit Root and Cointegration

Macroeconomic time series are rarely stationary and frequently characterized by trends. The variables in a balanced error correction model with a stationary variable on the left hand side must either be stationary or cointegrated, and we must therefore determine the order of integration of the variables in our model. Generally, any process that has a single unit root is said to be integrated of order one, that is $I(1)$, implying that the first difference of the process is stationary.

We have performed four different panel unit root tests; The Levin–Lin–Chu test (Levin et al. 2002), the Im–Pesaran–Shin test (Im et al. 2003), the Fisher ADF test and the Fisher PP test (Maddala and Wu 1999; Choi 2001). The Levin–Lin–Chu test assumes common unit root processes; the others assume individual root processes. The results are reported in Appendix B.

It is important to keep in mind that the unit root tests have low power. Nevertheless, the unit root tests generally support the stationarity assumptions when it comes to the growth rates entering the dynamic part of equation (8). However, the tests also reveal that the variables in levels, which are entering the long run part of (8), are non-stationary. Hence, we need to test their cointegrating properties.

Pedroni (1999) suggests a suite of seven tests designed to test the null hypothesis of no cointegration in dynamic panels with multiple regressors and a rank equal to 1. The first four tests are based on the within panel estimator (see Hsiao, 1986). The last three tests are labeled Group Mean Panel Tests by Pedroni, and are calculated by pooling along the between dimension. The tests allow for heterogeneity of the long-run coefficients and autoregressive parameters under both the null and the alternative. The tests are presented in Appendix B.

The test results are inconclusive. However, the Pedroni tests also have low power. Since the tests nevertheless give some support to the assumption that the variables of the long run part of (7) are in fact cointegrated, we choose to go ahead assuming that they in fact are cointegrated. This is a common strategy. After all, there are sound theoretical arguments in favor of a long run relationship between the variables, and special events in the data period may have prevented us from detecting the cointegrating relationships. A general to specific model reduction of the dynamic part of the model as we do may give more precise estimation of the cointegrating properties. We will return to the question in the empirical results section below.

4.2 Nickell-Bias

Nickell (1981) shows that OLS estimation may be inconsistent when applied to models that include fixed effects and a lagged dependent variable. The bias is of the order $1/T$, where T is the number of observations along the time dimension of the panel. The panel data set used in this paper is an unbalanced dataset, and the time dimension varies from 21 to 37 when lags of variables are included.[12] Hence, it is likely that the 'Nickell bias' will be very small. Moreover, Judson and Owen (1999) largely confirm this and show that OLS estimation of dynamic fixed effects models perform well for $T \geq 30$, i.e. with a T dimension similar to ours. Even when T = 20, the fixed effects estimator was almost as good as the alternatives (GMM and Anderson–Hsiao).

4.3 Poolability

The pooled panel data regression is valid only under the assumption that the slope coefficients are homogeneous across countries. As shown by Pesaran and Smith (1995), if homogeneous coefficients are falsely imposed, the pooled estimator is inconsistent even if T approaches infinity. The test statistics of all homogeneity restrictions in our pooled model is: $\chi^2(105) = 301.26$ [0.00]. Hence, the test clearly rejects the null of homogeneous coefficients. However, as pointed out by Baltagi (1995, Ch. 4) the pooled model can yield more efficient estimates at the expense of bias, and one must therefore balance the two concerns.

We have chosen to assume homogeneous coefficients. Our main objective in this analysis is to investigate the role of coordination in wage bargaining for price setting. Therefore, existing price equations constitute the most important benchmark in this process. We believe that the effect of coordination is best assessed if our price model otherwise reasonably well replicate the findings of others. The estimation results presented in the next section show that the pooled model is in fact in line with other empirically specified price equations, both in

12 See the appendix for the regression period for each country.

single country studies and when using a panel of countries. This indicates that the estimator of the pooled model probably has very little bias.

4.4 Non-Spherical Errors

The OLS estimator assumes spherical errors. Consequently, we must test the assumption of homoskedasticity and error independence in the panel regressions. We consider three cases of non-spherical errors, namely serially correlated errors, contemporaneous correlations and panel heteroskedasticity.

Because of the dynamic nature of our model specification the test for first order auto regressive errors in Table 1 indicates serially uncorrelated errors. However, there will most likely be contemporaneous correlations in a macro panel like ours. Therefore, we have included time dummies in order to correct for such cross-country dependence. In addition, we have included the price of crude oil as a proxy for economic shocks in the global economy. Furthermore, in the column labeled M2 in Table 1 we also present a cross-section SUR (PCSE) estimator. This estimator allows for unrestricted and unconditional correlation between the contemporaneous residuals.

We have also performed a panel homoskedasticity test using a likelihood ratio test between the log-likelihood value of the fixed effects specification with complete parameter heterogeneity ($l_{restricted}$), where the residual variances are restricted to be equal, and the sum of the log-likelihood values from the separate (unrestricted) models ($l_{unrestricted}$), i.e.:

$$\chi^2(14) : -2(l_{restricted} - l_{unrestricted}) = 59{,}5\,[0.00]**,$$

The p-value is given in square brackets. The null of homoskedasticity across countries is clearly rejected, so we have used estimation techniques that are robust to panel heteroscedasticity.

In addition, we have used the estimated residuals of the single country analysis to correct for the cross-country heteroskedasticity prior to the panel estimation. The correction is done by multiplying all variables for each country with the respective estimated residual variance. In Figure 4 we show the density distributions of the residuals for each country before and after this correction. The

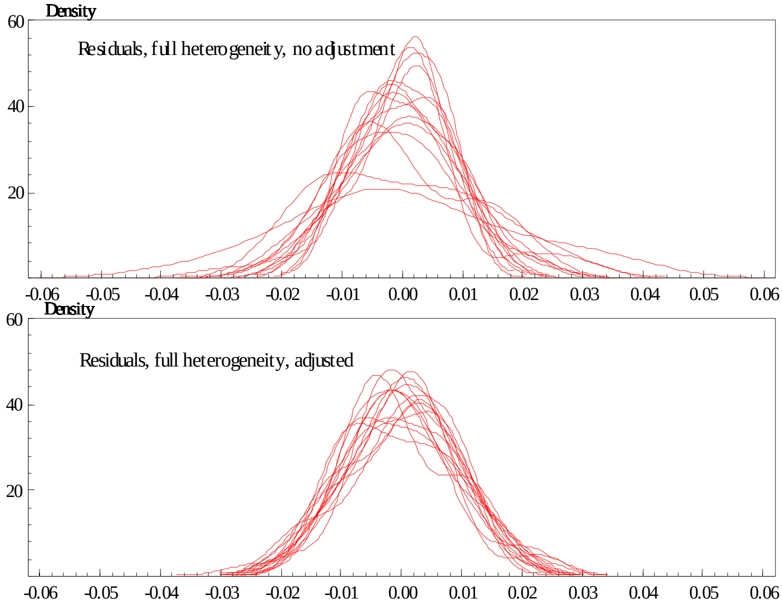
Table 1: Panel Estimation Results of the Determinants of Consumer Prices in 15 OECD Countries, 1964(–80)– 2000 (Dependent Variable is Δcp)

	M1	M2	M3	M4
Constant	0.122 (4.5)**	0.133 (3.8)**	0.103 (3.9)**	0.107 (4.1)**
$(cp\text{-}ulc)_{t-1}$	-0.039 (-3.2)**	-0.041 (-2.7)**	-0.039 (-3.2)**	-0.038 (-2.8)**
$(ulc\text{-}pi)_{t-1}$	-0.018 (-4.0)**	-0.014 (-3.5)**	-0.018 (-4.0)**	-0.019 (-4.2)**
Δcp_{t-1}	0.271 (7.9)**	0.260 (6.8)**	0.272 (7.9)**	0.291 (5.3)**
Δcp_{t-2}	0.091 (3.0)**	0.086 (2.1)**	0.093 (3.1)**	0.090 (3.0)**
Δulc	0.387 (18.1)**	0.411 (17.7)**	0.386 (17.9)**	0.358 (6.0)**
Δpi	0.028 (2.2)*	0.018 (1.1)	0.027 (2.1)*	0.031 (1.5)
Δur	-0.025 (-3.3)**	-0.030 (-3.6)**	-0.025 (-3.2)**	-0.026 (-5.3)**
Δpo	0.021 (3.6)**	0.021 (1.7)	0.022 (3.6)**	0.021 (2.5)*
Δe_{t-1}	0.034 (4.6)**	0.042 (4.6)**	0.034 (4.6)**	0.033 (6.1)**
CO_{t-1}	0.0020 (3.2)**	0.0025 (3.6)**		0.0019 (4.1)**
$DCO{=}2_{t-1}$			0.0061 (1.3)	
$DCO{=}3_{t-1}$			0.0083 (1.9)	
$DCO{=}4_{t-1}$			0.0100 (2.4)*	
$DCO{=}5_{t-1}$			0.0115 (2.7)*	
Std. error in %	0.90	0.94	0.90	0.90
$N_{AR\text{-}1}$	-0.286 [0.78]	0.026 [0.56]	-0.364 [0.72]	-0.708 [0.48]
Sargan test $\chi^2(7)$=	na	na	na	8.049 [0.33]

Note: The columns marked M1 and M3 show estimation using OLS with heteroskedasticity-corrected (robust) standard errors (see Section 4). Cross-section SUR (PCSE) estimation on unadjusted data is presented in column M2. M3 is identical to M1 except that there are dummy variables instead of the coordination variable CO. Estimation using IV on heteroskedasticity-corrected series and treating Δpi and Δulc as endogenous variables is shown in column M4. All exogenous variables, dummies and Δpi_{t-1}, Δpi_{t-2}, Δulc_{t-1}, Δe, Δpo_{t-1}, EP, EP_{t-1}, BRR and BRR_{t-1} are instruments relative to this regression. *t*-values are given in parentheses and p-values are given in square brackets. * and ** denote significance at the 5% and 1% level, respectively. $N_{\text{AR-1}}$ has a standard normal distribution under the null of no 1. order autoregressive errors. The Sargan test (Sargan, 1964) is χ^2 distributed under the null of valid instruments, and the degrees of freedom are given in the parenthesis.

correction envelopes reasonably well distributed residuals. The SUR (PCSE) estimation in M2 in Table 1 is conducted on unadjusted data because this estimator is robust to panel heteroskedasticity.

Figure 4: Density Distribution of the Residuals before (Top) and after (Bottom) Correcting for Heteroskedasticity



4.5 Identification

Unit labor costs (*ulc*) and the price of imports (*pi*) are explanatory variables in our consumer price equation (8). As the consumer price variable itself might influence unit labor costs, the *ulc* variable could be treated as an endogenous variable in the estimation of (8). Similarly, according to the pricing to market theory importers may set their prices dependent on the consumer prices in the country they are operating in. This may entail the use of instrumental variables to correct for potential simultaneity problems.

However, the Durbin–Wu–Hausman test indicates that both unit labor costs and price of imports are exogenous variables. The test has two steps; first we estimated an equation for Δulc using its own lags and an index of the degree of employment protection (*EP)*, level and lagged, as instruments (lags of Δcpi were found to be insignificant). An equation for Δpi was estimated using Δpi_{t-2}, $\Delta\, e_t$ and Δpo_t as instruments. Second, we included the estimated residuals from the two above equations as explanatory variables in the equation for Δcp separately. The residuals from the Δulc equation had a *t*-value equal to 0.52 (t-probability: 0.60), and the residual from the Δpi equation returned a *t*-value equal to 0.18 and a *t*-probability of 0.86. Hence the residuals had no significant effect on Δcp indicating that Δulc and Δpi may be treated as exogenous variables. A Sargan-test fails to reject the validity of the instruments employed in these endogeneity tests. In M4 in Table 1 we nevertheless report an IV estimator where we treat Δulc and Δpi as endogenous variables.

5 Empirical Results

The sample comprises 468 observations, but 14 of them are accounted for by including impulse dummies to control for large residuals (*t*-values between 4.6-25.7) due to special events in the sample period. The estimation results are as summarized in Table 1.

Model M1 is estimated using OLS with heteroskedasticity-corrected (robust) standard errors (see section 4). In model M2 we present a cross-section SUR (PCSE) estimation. While the model in M1 corrects for cross-sectional heteroskedasticity, the SUR (PCSE) estimation in M2 also corrects for contemporaneous correlations between countries. Since the standard errors of the coefficients are only moderately changed, the results from this estimation suggest that the problems with contemporaneous correlations in model M1 are rather small.

In M3 we exclude the *CO* variable and include four dummies instead. DCO=2 is 1 when CO=2 and zero otherwise, DCO=3 is 1 when CO=3 and zero otherwise, and so on. When CO=1 all dummies are zero. Hence, the corresponding coefficients measure the estimated effect on consumer prices by moving from CO=1 to another level. While the effect of increased coordination is restricted to be the same regardless of the initial level of coordination in M1, the separate effects are estimated freely in M3.

As explained in the previous section, Δpi and Δulc were found to be exogenous variables using the Durbin–Wu–Hausman test. Therefore, in models M1–M3 we have treated both Δulc and Δpi as exogenous variables. In model M4 we nevertheless show the results of IV estimation treating Δulc and Δpi as endogenous variables. All exogenous variables, dummies and Δpi_{t-1}, Δpi_{t-2}, Δulc_{t-1}, Δe, Δpo_{t-1}, *EP* (an index of the degree of employment protection), EP_{t-1}, *BRR* (the benefit replacement ratio) and BRR_{t-1} are instruments in this regression. The Sargan test for the model specification M4 supports the validity of the instruments. Moreover, treating Δulc and Δpi as endogenous variables does not change the results noticeably.

In the estimations, the price of crude oil is included both as a level and differenced variable. The price of oil may act as a proxy for the price of intermediate goods. This is relatively established in empirically estimated price equations. It is however not clear how it should be theoretically implemented in the analysis in Section 2. It is nevertheless found to have an insignificant effect on prices in the long run, and therefore excluded from the long-run relationship. The exchange rate between the USD and the local currency (*E*) is included as an explanatory variable in the short-run part of the models. We believe that changes in the exchange rate against the USD and changes in the world market price of oil may correct for some common shocks to the countries in the panel. In addition we have included time dummies.

The estimation results demonstrate that changes in consumer prices in the short run are determined by changes in the price of imports (insignificant in M2 and M4), unit labor costs, the unemployment rate, the price of oil (insignificant in M2) and the lagged exchange rate. Higher unit labor costs, price of imports and price of oil, and a depreciation of the exchange rate are empirically estimated to increase inflation. A short-run effect of the unemployment rate indicates evidence in favor of a pro-cyclical markup.

The long-run solution of M1 is shown in Table 2. The results are in line with the common findings in the literature; in the long run there is a roughly 50–50 percent weight on unit labor costs and import prices, respectively[13]. This means that an increase in either the unit labor costs or the price of imports of one percent increases the level of consumer prices by approximately 0.50 percent. In the table c_i is a country-specific constant.

[13] See inter alia, Ashworth and Byrne (2003), Asteriou et al. (2002), Boug et al. (2006) and Bårdsen et al. (1998).

Table 2: Long-Run Solution of Model M1 in Table 1

LONG-RUN SOLUTION: M1
$cp = c_i + 0.53\, ulc + 0.47\, pi + 0.053\, CO$

More interestingly for the focus of this paper is the long-run coefficient of 0.053 on *CO*. This implies that a movement from, say, 2 to 3 on the coordination score, will increase the level of consumer prices by 5.3 percent. Correspondingly, a complete decentralization of wage bargaining, i.e. a movement from 5 to 1 on the index, is thus supposed to decrease the level of consumer prices by approximately 21 percent in the long run. Nevertheless, the effect of a change in coordination is slow. The loading coefficient indicates an increase in the price level following a unit increase in the coordination index of only 0.20 percent the following year. Half of the deviation from the long run solution is corrected for after approximately 11.5 years. As can be seen from M3, the restriction in M1 of a monotonically increasing effect of increased coordination on consumer prices is clearly not rejected, formally: $\chi^2(4) = 0.76$ which corresponds to a p-value equal to 0.94. However, coordination has a slightly stronger estimated impact on consumer prices in M3 as compared to M1.

As a robustness test we have also estimated the equations without the impulse dummies. The estimated coefficients in the dynamic part of the model were generally very little affected. The coefficients in the long run part of the model were changed somewhat more, but not significantly. Most importantly, the estimated effect of coordination on consumer prices was still significant at a 5 percent level. The estimated residual standard error increased by approximately 0.2 percentage points when excluding the dummies.

The diagnostic tests at the bottom of Table 1 are also of interest when considering the econometric validity of the model. First, Sargan's test statistic is insignificant for M4, indicating that the price equation is a valid model. Second, there is no significant residual autocorrelation in neither of the estimations. In sum, the results suggests that inflation equilibrium corrects with respect to an open economy long-run price equation. Hence, our interpretation of the cointegration tests in Appendix B is supported by the results for the dynamic econometric models in Table 1.

On the whole, these results are quite remarkable, and suggest that the low level of real wages in countries with a coordinated wage bargaining system at least partly stems from higher prices. Seen together with empirical wage-equations, the results indicate that there are two separate effects of coordination on unemployment; first, coordination increases the (price) markup over wages, which lowers demand and increase unemployment, second, coordination lowers the wage markup over prices, which raises demand and decrease unemployment. This suggests that increased coordination of wage bargaining has a clear negative effect on real wages and an ambiguous effect on unemployment, consistent with the empirical evidence cited in the introduction.

The results may potentially have policy implications, since there might be further macroeconomic benefits in coordinated regimes. A strong focus on increasing consumer power, anti trust legislations, fighting monopoly tendencies and reducing entry barriers, may in fact increase real wages and reduce unemployment further in countries with a high degree of coordination in the wage determination.

On the other hand, looking at the results, it seems that highly coordinated regimes have emerged as a consequence of an income transfer from consumers to firm profits. Because of profit sharing in wage determination, the income transfer is in reality from consumers to both capitalists and wage earners. The increased profits from "taxing" consumers in coordinated regimes may in fact be necessary for the sustainability of such a social contract. One might say then, that reduced competition in the product market is a price we should be willing to pay for having a high degree of coordination in wage determination.

6 Conclusions

The objective of this paper was to investigate the hypothesis that coordination of wage bargaining increases the level of consumer prices since union coordination may influence entry barriers and competition in product markets. The empirically specified price equations estimated in this paper are based on panel data for 15 OECD countries observed from the 1960s to 2000. The main finding is a significant long-run coefficient (0.053) of the coordination variable *CO*. A movement from a completely coordinated regime to a fully uncoordinated regime will according to these empirical results decrease the long-run price level by 21 percent.

Furthermore, the results indicate that the level of consumer prices in the short run is determined by the price of imports, unit labor costs, the unemployment rate, the price of oil and the exchange rate. In the long run, the price level is affected positively by unit labor costs and the price of imports. The long-run coefficients of unit labor costs and the price of imports reconcile very well with the common findings in the literature, i.e. that the two variables are approximately of equal importance for the consumer prices.

The findings in this paper may offer an explanation for why empirical research on the impact of coordination of wage bargaining on unemployment has been inconclusive. While increased coordination reduces wage claims, and hence lowers unemployment, it also increases the price markup, which raises unemployment. However, it is worth noting that the empirical evidence referred to in the introduction suggests that the best macroeconomic performance is associated with a very high level of coordination of wage bargaining. If so, the moderating effect coordination at this level has on wages outweighs the adverse effect it has on prices.

The price effect from union coordination can serve as an explanation for the sustainability of the coordinated regimes in many countries. A common characteristic of countries with a coordinated regime is that the employer organizations are encouraging the system, at least implicitly, and in the light of these findings this is obvious. As coordination moderates the pay claims *and* increases prices, the firms in these coordinated regimes may consider themselves better off due to decreased real wages. Workers, on the other hand, experience this decrease in real wages but still gain through increased employment. At the same time, there are most certainly institutional variations in the labor markets across the panel that can lead to different national effects of a hypothetical change in the degree of coordination. Exploring these aspects is beyond the scope of this paper.

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Appendix A – Data

The data consist of annual time series from 1960 for some countries and up to 2000 for all, for a selection of variables for the 15 OECD countries indicated in the table below. Some of the variables do not exist for the whole period, and similarly some countries' variables are not available. The longest time series is for the period 1964-2000, and the shortest is for the period 1980-2000. Consequently, we possess an unbalanced panel data set.

Most of the data (except for the *CO* variable) used in this paper is retrieved from or constructed by using the Organisation for Economic Co-operation and Development (OECD) Economic Outlook and Main Economic Indicators (MEI) Databases.[14]

List of Countries in the Data Sample

COUNTRY	REGRESSION PERIOD
Australia	1968–2000
Austria	1968–2000
Belgium	1972–2000
Canada	1964–2000
Denmark	1972–2000
Finland	1971–2000
France	1972–2000
Ireland	1980–2000
Italy	1966–2000
Japan	1965–2000
Netherlands	1973–2000
Norway	1976–2000
Sweden	1968–2000
UK	1966–2000
USA	1966–2000

[14] By using Xvision Fame 8.0.2, a programme licensed by SunGard Data Management Solutions.

Description of Variables

CP – Consumer Prices

The *CP* variable is constructed by using a Purchasing Power Parity index (*PPP*) and multiplying it with the consumer price index for USA (CPI_{US}), which is 1 in 2000. The *PPP* variable is a price relative which measures the number of units of each country's currency that are needed in the country to purchase the same quantity of an individual good or service as 1 USD will purchase in the US, i.e. $PPP_i = P_i/P_{US}$, where P_i and P_{US} are the price levels in country i and in US, respectively. The calculation gives us:

$$CP_i = PPP_i \cdot CPI_{US} = \frac{P_i}{P_{US}} \cdot \frac{P_{US}}{P_{US\_2000}} = \frac{P_i}{P_{US\_2000}}$$

The denominator (P_{US_2000}) is the price level in the US for year 2000 and is simply a constant, which just adds to the constant in the regression.

CO – Wage Setting Coordination Scores

These data are retrieved from Professor Lane Kenworthy's dataset (http://www.u.arizona.edu/~lkenwor/). The data draw on a variety of sources, including Soskice (1990), Iversen (1999), Traxler et al. (2001), the Golden–Lange–Wallerstein indexes of wage centralization (Golden et al. 1997), Ferner and Hyman (1998), the monthly European Industrial Relations Review, and the European Industrial Relations Observatory website (http://www.eiro.eurofound.ie).

The coordination scores consist of an index from 1–5[15]:

1 = Fragmented wage bargaining, confined largely to individual firms or plants (Canada, Ireland 1960–69 and 1981–1987, New Zealand since 1988, United Kingdom since 1980, United States)

[15] Kenworthy, L. *Wage-setting coordination scores*. http://www.u.arizona.edu/~lkenwor/WageCoorScores.pdf

2 = Mixed industry- and firm-level bargaining, with little or no pattern-setting and relatively weak elements of government coordination such as setting of basic pay rate or wage indexation (Australia since 1992, France, Italy in most years)

3 = Industry-level bargaining with somewhat irregular and uncertain pattern-setting and only moderate union concentration (Denmark in most years since 1981, Finland in a few years, Sweden since 1994)

Government wage arbitration (Australia prior to 1981, New Zealand prior to 1988)

4 = Centralized bargaining by peak confederation(s) or government imposition of a wage schedule/freeze, without a peace obligation (Belgium and Finland in most years, Ireland 1970–1980 and 1987–1993)

Informal centralization of industry- and firm-level bargaining by peak associations (Italy since 1993, Netherlands since 1983, Norway in some years, Switzerland) Extensive, regularized pattern-setting coupled with a high degree of union concentration (Germany, Austria since 1983)

5 = Centralized bargaining by peak confederation(s) or government imposition of a wage schedule/freeze, with a peace obligation (Denmark 1960–1980, Ireland since 1994, Norway in some years, Sweden 1960–1982)

Informal centralization of industry-level bargaining by a powerful, monopolistic union confederation (Austria prior to 1983)

Extensive, regularized pattern-setting and highly synchronized bargaining coupled with coordination of bargaining by influential large firms (Japan)

PI – Price of Imports

The ratio of import value in local currency and import volume is used as a proxy for the price of imports.

PO – Price of Oil

The world dated price of Brent crude oil in USD per barrel multiplied with the average annual exchange rate between the local currency and USD, so that the variable proxies the price of oil in local currency.

UR – Rate of Unemployment

The OECD standardized unemployment rate gives the number of unemployed persons as a percentage of the civilian labour force.

ULC – Unit Labor Costs

ULC is an index of unit labour costs (2000=100) provided by the OECD.

E – Exchange Rate

E is the exchange rate given by USD ($)/Local currency; average spot rates.

EP – Employer Protection

The data comprise an index of the degree of employer protection. Data provided by Dr. Luca Nunziata, Nuffield College, University of Oxford, UK, see Nunziata (2005). The series are extended with the 1995 value to year 2000.

BRR – Benefit Replacement Ratio

The data comprise an index of benefit replacement ratio. Data provided by Dr. Luca Nunziata, Nuffield College, University of Oxford, UK, see Nunziata (2005). The series are extended with the 1995 value to year 2000.

Appendix B – Stationarity and Cointegration

We have performed four different panel unit root tests; The Levin–Lin–Chu test (Levin et al. 2002), the Im–Pesaran–Shin test (Im et al. 2003), the Fisher ADF test and the Fisher PP test (Maddala and Wu 1999; Choi 2001). The Levin–Lin–Chu test assumes common unit root processes; the others assume individual root processes. The results are reported in Table B1.

Table B1: Panel Unit Root Tests

Panel Unit Root Test on (country-specific effects and linear trends):					
Null: Unit root	*cp*	*ulc*	*pi*	*ur*	*e*
Levin-Lin-Chu, t-stat.	**0.53 (0.70)**	**2.11 (0.98)**	**3.53 (1.00)**	**1.71 (0.96)**	**-1.15 (0.12)**
Im-Pesaran-Shin, W-stat.	**2.37 (0.99)**	**6.04 (1.00)**	**6.19 (1.00)**	**1.67 (0.95)**	**-2.52 (0.01)****
ADF - Fisher, χ^2 - stat.	**18.3 (0.95)**	**7.59 (1.00)**	**3.01 (1.00)**	**28.1 (0.56)**	**43.9 (0.03)***
PP - Fisher, χ^2 - stat.	**2.45 (1.00)**	**4.76 (1.00)**	**2.22 (1.00)**	**10.6 (1.00)**	**18.6 (0.91)**
Null: Unit root	*Δ cp*	*Δ ulc*	*Δ pi*	*Δ ur*	*Δ e*
Levin-Lin-Chu, t-stat.	**-4.70 (0.00)****	**-6.48 (0.00)****	**-10.9 (0.00)****	**-0.29 (0.39)**	**-13.5 (0.00)****
Im-Pesaran-Shin, W-stat.	**-1.50 (0.07)**	**-3.89 (0.00)****	**-8.21 (0.00)****	**-6.77 (0.00)****	**-10.0 (0.00)****
ADF - Fisher, χ^2 - stat.	**36.3 (0.20)**	**61.3 (0.00)****	**116.5 (0.00)****	**106.6 (0.00)****	**138.4 (0.00)****
PP - Fisher, χ^2 - stat.	**29.9 (0.47)**	**48.9 (0.02)***	**113.0 (0.00)****	**46.3 (0.03)***	**127.8 (0.00)****
Null: Unit root	*Δ^2 cp*	*Δ^2 ulc*	*Δ^2 pi*	*Δ^2 ur*	*Δ^2 e*
Levin-Lin-Chu, t-stat.	**-21.2 (0.00)****	**-18.1 (0.00)****	**-18.3 (0.00)****	**-7.74 (0.00)****	**-22.24 (0.00)****
Im-Pesaran-Shin, W-stat.	**-20.4 (0.00)****	**-16.8 (0.00)****	**-21.4 (0.00)****	**-13.5 (0.00)****	**-23.2 (0.00)****
ADF - Fisher, χ^2 - stat.	**335.1 (0.00)****	**267.5 (0.00)****	**360.8 (0.00)****	**206.8 (0.00)****	**387.2 (0.00)****
PP - Fisher, χ^2 - stat.	**1043 (0.00)****	**830.8 (0.00)****	**2496 (0.00)****	**288.5 (0.00)****	**3385 (0.00)****

Note: The Levin–Lin–Chu test assumes common unit root processes (see Levin, Lin and Chu, 2002). The Im–Pesaran–Shin test (Im, Pesaran and Shin, 2003), the Fisher ADF test and the Fisher PP test (Maddala and Wu, 1999, and Choi, 2001) assume individual root processes. P-values are given in parentheses, * and ** denote significance at the 5% and 1% level, respectively. Δ and Δ^2 denote that the variable is in first and second difference, respectively.

The null hypothesis of non-stationarity is not rejected for any of the variables that enter the long run part of the model, i.e. log of consumption prices (*cp*), log of unit labor cost (*ulc*) and log of import prices (*pi*).[16] While the log of exchange rates between USD and each country's local currency (*e*) show signs of being stationary, the unemployment rates are clearly not. The same tests on the first differences are for the most part rejected. However, according to the Im–Pesaran–Shin test and the two Fisher tests, the null of unit root for the inflation rates are not rejected. The Levin–Lin–Chu test rejects the null with a significance level below 1

[16] As for the coordination indices these are sometimes constant for long periods, and as such are not applicable to unit root testing. However, as they are indices and therefore bounded processes, they will not introduce spurious non-stationarity to the model.

percent, while the test statistics of the Im–Pesaran–Shin test has a p-value equal to 0.07. In addition to the variables in Table 1, we use the price of crude oil measured in dollars, and ADF-tests reveal that this variable is *I(1).* Keeping in mind that these tests have low power, on most parts the unit root tests support the stationarity assumptions when it comes to the growth rates entering the dynamic part of equation (8). Furthermore, the tests also reveal that the variables in levels, which are entering the long run part of (8), are non-stationary.

Pedroni (1999) suggests a suite of seven tests designed to test the null hypothesis of no cointegration in dynamic panels with multiple regressors and a rank equal to 1. The first four tests are based on the within panel estimator (see Hsiao, 1986). The last three tests are labeled Group Mean Panel Tests by Pedroni, and are calculated by pooling along the between dimension. The tests allow for heterogeneity of the long-run coefficients and autoregressive parameters under both the null and the alternative.

While macro panels typically exhibit cross-sectional dependence, the panel unit root tests and the Pedroni panel data cointegration tests all assume cross-country independence. As shown in Banerjee et al. (2004) and Banerjee and Carrion-i-Silvestre (2004) using Monte Carlo simulations, falsely assuming cross-sectional independence may cause severe size distortions. We have computed the Pedroni tests using a RATS code written by Peter Pedroni, where inclusion of time dummies in order to control for this type of cross-sectional dependence is optional.[17] The test statistics of the seven tests, both with and without time dummies, are shown in Table B2 in the same order as in Pedroni (1999). As can be seen from the table, the null of no cointegration is clearly rejected when time dummies are not included, while the tests are inconclusive when time dummies are included.

[17] We are indebted to Professor Pedroni for providing us with the latest version of the code.

Table B 2: Panel Cointegration Tests

Pedroni (1999) Panel cointegration tests

Test number	**1**	**2**	**3**	**4**	**5**	**6**	**7**
Heterogeneous intercepts included							
Pedroni stats.	-1.9[0.06]*	2.5[0.01]**	2.3[0.02]**	2.1[0.04]**	3.8[0.00]***	3.5[0.00]***	3.4[0.00]***
Heterogeneous intercepts and time dummies included							
Pedroni stats.	2.1[0.04]**	-0.7[0.46]	-1.7[0.08]*	-1.7[0.09]*	-0.1[0.92]	-1.9[0.06]*	-1.8[0.07]*

Note: Tests 1-4 are based on the within panel estimator (see Hsiao, 1986). Tests 5-7 use the between dimension, see Pedroni (1999). The test are performed using Pedroni's RATS code (Pedroni, 2006). P-values are given in parentheses. *, ** and *** denote significance at the 10%, 5% and 1% level, respectively.

The Pedroni tests also have low power, especially when 39 time dummies (as is the time dimension of our data set) are included. The tests nevertheless give some support to the assumption that the variables of the long run part of (7) are in fact cointegrated.