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Wage coordination and unemployment dynamics in Norway and Sweden

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Wage coordination and unemployment dynamics in Norway and Sweden[#]

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May 25., 2001

Abstract

From a detailed study of the yearly wage bargaining rounds in Norway and Sweden, we construct time series of five complemetary coordination indices. Econometrics is used to evaluate the importance of the coordination indicators for our understanding of the changes in the rates of unemployment in the two countries. The results show that there is considerable similarity between the two countries, e.g., in terms of the estimated e;ects of macroeconomic variables on unemployment, The coe>cients are recursively stable over the the 1980s and 1990s. Simulation shows that the models imply numerically important e;ects of wage coordination on the rates of unemployment in both countries.

Keywords: *unemployment, wage setting, economic institutions, corporatism, centralization, decentralization, income policies, econometric modelling.*

JEL classification: *C32, J51, J52, P11, P17, P52.*

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1 Introduction

In most European countries, individual wages are a;ected by collective bargaining between employers and employees. The institutional framework in which this bargaining takes place vary strongly from country to country, in particular with respect to the degree of coordination. These variations are now typically considered to play a key role in the understanding of the European unemployment problem. A fully centralised bargaining system may produce low unemployment because the bargainers induced to take the aggregate (equilibrium) consequences of their behaviour into account, and hence avoid damaging wage-price spirals. However, a centralised system may be less responsive towards demand- and supply forces that call for changes in relative wages. As pointed out by Calmfors and Dri>ll (1988), the relationship between wage coordination and unemployment may be non-monotonic. Both extremes may produce better results than middle-of-the-road systems, as the wage aspirations of e.g., branch-level bargainers are neither checked by the competitive forces associated with completely local wage setting, nor by internalising the aggregate outcome associated with completely centralized bargaining.

In recent years, there has been a burst of research into the relationship between the degree of wage coordination on the one hand and the rate of unemployment (or other macroeconomic success indicators) on the other. Based on cross-section analyses of country-specific unemployment rates, the result that coordinated wage bargaining systems produce lower unemployment than more decentralized systems, seems to be fairly robust, see Calmfors and Dri>ll (1988), Soskice (1990); Bean (1994),Jackman et al. (1996), Scarpetta (1996),OECD (1997),OECD (1999), Nickell and Layard (1999) and Iversen (1999). But there are at least three serious problems associated with these pieces of evidence. The first is that they are based on comparisons of countries that do not only di;er in terms of their wage bargaining systems, but also in terms of other structural characteristics, as well as in their macroeconomic policy practices. And given that the few degrees of freedom available in cross-section studies makes it virtually impossible to condition on all potentially relevant factors, it is a possibility that the correlation across countries between wage coordination and unemployment is spurious. The second problem is that of measurement. The analyses referred to above are all based on the construction of crude indices that rank the countries in terms of wage coordination or centralization and also quantify the di;erences between them. It is clear that the construction of such indices involve a large element of subjective judgement, and as demonstrated by Flanagan (1999), apparently small changes in these judgements may drastically change the qualitative results. The third problem is that the cross-section analyses disregard the substantial element of variation in coordination within countries over time (although some of the studies do represent each country with two or three observations corresponding to di;erent periods).

The aim of the present paper is to take advantage of within-country variations in wage setting coordination over time in order to identify the relationship between wage coordination and unemployment. For this purpose, we have engaged in a detailed study of the yearly wage bargaining in Norway and Sweden during the past four decades. Qualitative information about each year's wage settlements has then been used to construct time series of five complementary coordination indicators describing the level of the bargaining, the degree of government participation (incomes policies), the degree of coordination on the employer- and employee sides, respectively, and the bargaining climate between the two parties. We use this comprehensive information, together with other structural and macroeconomic variables, to assess the causal relationship between wage coordination or 'corporatism' and the rate of unemployment.

The paper proceeds as follows. In section 2, we review the intellectual foundation for the type of wage-coordination endeavours that have dominated much of Norwegian and Swedish wage setting, with the aid of a simple theoretical bargaining model. Section 3 gives a brief historic overview of wage bargaining in Norway and Sweden and presents the five coordination indicator series for each country. Section 4 discusses methodology and presents the results of the econometric analyses, and section 5 concludes.

2 Theory: How does coordination a;ect unemployment?

Our main theoretical point can be made in the, by now, standard modelling framework of Layard et al. (1991). Wages are set in a decentralized bargaining process, while product prices are determined by monopolistically competing profit maximizing firms. Wage- and price setters target a (consumer or producer) real wage, and their respective wage claims depend on the rate of unemployment. Let $i = 1, 2, ..., I$ be the set of monopolistically competing firms. Output in firm i, Y_i, is proportional to labour inputs N_i, i.e., $Y_i = AN_i$. Assume that each firm faces a demand function of the form

$$(1) \qquad Y_i = \left(\frac{P_i}{P}\right)^{P@} D(P, X), \quad @ > 1,$$

where P_i is the price set by firm i, P is the aggregate producer price level and $D(.)$ denotes aggregate demand with X as (vector of) demand determining variables. The price elasticity of demand, @, depends on the degree of competition in the product market, but is assumed to be independent of the level of aggregate demand. To maximize profits, each firm equalizes marginal revenues to unit costs. This implies that the price set by firm i is a mark-up on the (productivity-adjusted) wage cost, W_i , in that firm:

$$(2) \qquad P_i = \left(1 + \frac{1}{@ \quad 1}\right) \frac{W_i}{A}.$$

The union in firm i maximizes the utility function

$$U_i = N_i \left(\frac{\mathrm{F}W_i \quad ((1 \quad u)\mathrm{F}W + uB)}{P_c} \right), \tag{3}$$

where F is the fraction of the firms' wage costs that accrues to the worker, P_c is the consumer price level, u is the rate of unemployment, W is the wage rate prevailing in other firms and B is the income allocated to unemployed workers (and/or the value of leisure). The wedge between the consumer and the producer real wage is determined by income- and payroll taxes and by import prices. We have that $\mathrm{F} = (1 \quad)_1)(1 +)_2)^{\mathrm{P}1}$ and $P_c = (1 +)_3)I(P, VP^{imp})$, where $)_1$ is the income tax, $)_2$ is the payroll tax, $)_3$is the commodity tax and $I(.)$ is a price index depending on producer prices P, and import prices P^{imp} adjusted for the exchange rate V.

The wage is determined through a generalized Nash-bargaining over wages only, i.e., $W_i = \arg\max \left(U_i^{-} <_i^{(1\mathrm{P}^{-})} \right)$,where $<_i$ is the profit of the firm. The solution to this maximization problem is

$$\mathrm{F}W_i = \frac{1}{1 \quad c} ((1 \quad u)\mathrm{F}W + uB), \tag{4}$$

where

$$c = \frac{-}{- + @ \quad 1}. \tag{5}$$

The take-home wage in each firm is a mark-up on alternative income, and the mark-up depends on relative bargaining strengths and the degree of product market competition. In a symmetric equilibrium we have that $P_i = P$ and $W_i = W$, hence the aggregate price- and wage setting equations reduce to

$$\frac{W}{P} = \frac{@ \quad 1}{@} A \tag{6}$$

$$u = \mathrm{F}W \frac{c}{\mathrm{F}W \quad B} \tag{7}$$

From (6), the real producer wage is determined solely by the price setting mechanism, i.e., it is a fraction of the labour productivity which is larger the more fierce is the product market competition. Wage setting behaviour in turn does not a;ect the equilibrium real wage at all, only the rate of unemployment, see 7.

We wish to take the argument one step further by noting that for the economy as a whole it is unreasonable to treat the nominal level of unemployment income, B, as given. Unemployment benefits are normally (mechanically or discretionary) tied to previous earnings. And even other sources of welfare, such as income support from family members, social assistance, capital income and the provision of public services are likely to be closely related to the general standard of living in the society. This implies that in the long run, it is appropriate to assume that there is a fixed replacement ratio, Q,attached to the nominal take-home pay, in that $B = \mathtt{Q}FW$. Consequently, the wage setting curve determines the equilibrium rate of unemployment uniquely:

$$u = \frac{c}{1 \quad \mathtt{Q}} \tag{8}$$

Wage setters try to set wages relative to the income derived from unemployment, but as the income derived from unemployment catches up, they are unable to do so. Nevertheless, because they are uncoordinated, they keep on trying until the unemployment rate obtains the value at which the desired relationship between the wage and the income derived from unemployment exactly matches the exogenously determined replacement ratio, Q. At this point the rate of unemployment reaches its equilibrium level. The equilibrium rate of unemployment is higher, the higher is the replacement ratio, the lower is the elasticity of labour demand and the higher is the bargaining power of the workers. It is independent of the level of productivity, as well as taxes and other wedge variables. Since the real wage is determined by price setting behaviour in the long run, the welfare of the unions is strictly higher the weaker is their own bargaining power. Each union of course pursues its own self-interest. But the collective result of this uncoordinated activity is that all lose. This conclusion forms a strong intellectual basis for collective wage restraint, and may explain why unions in Norway and Sweden have indeed embarked on co-ordinating activities that aim at halting the nominal wage growth for their own members. In the present model, such a policy would simply amount to a reduction in the bargaining power of local unions, ¯. However, even though collective wage moderation is a desirable equilibrium outcome for all agents in the economy, it may be very di>cult to accomplish in practice. Each local union has obvious incentives to defect from the co-operative strategy, particularly if wage setting is viewed as a one-shot game. In a dynamic setting, co-operative wage setting may be more sustainable, see Holden and Raaum (1991). But even if the coordination problem is solved, there may be short-term gains associated with more aggressive wage setting if the price setting mechanism operates sluggishly.

More generally, the steady state is likely to involve productivity- and hence real wage growth as well as inflation. If wages grow over time, the precise way in which unemployment benefits are indexed to wages may be of importance. In both Norway and Sweden, there is a lag in the determination of benefits, arising from the fact that these payments are indexed to previous nominal earnings. As a result, the replacement ratio is, in relation to expected earnings, lower the higher is the rate of nominal wage growth. In times of low inflation, this may not matter much. But during periods of very high inflation, the consequences may be quite severe for the unemployed workers concerned. To illustrate a potentially interesting consequence of this latter point, assume that there exists a steady state characterized by a constant yearly nominal wage growth given by g_w. Assume also that the income of the unemployed is indexed to the nominal wage level in the previous year. The steady state wage curve can then be written

$$u = \frac{c}{1 \quad \mathrm{Q}^{\mathrm{t}}}, \quad \text{where} \quad \mathrm{Q}^{\mathrm{t}} = \frac{1}{1+g_w}\mathrm{Q},$$

hence the replacement ratio is discounted with a rate that is equal to the nominal wage growth. Again, there is no trade o; between the rate of unemployment and the level of the wage rate in the long run. There is, however, a trade-o; between the rate of unemployment and the rate of nominal wage growth. The conclusion that the equilibrium rate of unemployment depends on the structural factors that a;ect wage setting (bargaining power, product market competition and everything that a;ects the living standard of the unemployed) continues to hold.

The main merit of our theoretical model is that it makes transparent why the institutional aspects of wage setting play an important role for unemployment in a steady state equilibrium situation. However, the theory makes a list of assumptions that are unlikely to hold in practice, e.g., identical firms and one type of labour input, fixed labour supply, a particular form of unions' utility function, and no adjustment lags.[1] Consequently, equation (8) may lead to a model with undesirable statistical properties when estimated on time series data. On the other hand, the possibility of a causal mechanism between wage coordination and unemployment is not limited by the specifics underlying the derivation of (8), but is of interest also for a wider class of models. Thus, it is important that judgement of the relevance of wage coordination is based on a model of the rate of unemployment that has good statistical properties. In general, this will lead to inclusion of additional explanatory variables to those already motivated by the formal theoretical model, see section 4 below. First, however, the corporative facet of wage formation has to be made operational. The construction of our new coordination indices is explained in the next section.

3 Coordination indices for Norway and Sweden

Norway and Sweden have strong traditions for coordinated wage setting. In the 1960s and 1970s, the degree of coordination was high in both countries. Starting in the 1980s, Sweden has moved towards less centralised and less coordinated wage settlements. This trend has been strengthened in the 1990s. Norway also embarked upon less coordinated wage settlements in the beginning of the 1980s, but in this

[1] For example, Kolsrud and Nymoen (1998) investigate a dynamic wage-price system that implies a steady state for inflation, the real wage and the real exchange rate for any given level of unemployment.However, if that model is modified so that unemployment is linked the real wage rate, the e;ects of increased wage pressure will be qualitatively the same as in the static model in this section.

case the decentralisation was reversed during the late 1980s. The revitalisation of coordination in Norway has continued in the 1990s. There are some traditional di;erences in bargaining patterns between Norway and Sweden that seem to have persisted. First, there is a greater number of bargainers in Sweden. Second, government intervention, or incomes policy, has been more actively used in Norway. The Norwegian system has permanent institutions for income policy, whereas in Sweden incomes policy has been used only under exceptional circumstances.

There are several dimensions of coordination in collective wage bargaining, and it is interesting to consider the e;ects of each dimension separately. Our definition of coordination hence includes a traditional indicator, the level of bargaining, as well as less traditional indicators as incomes policies from the government, coordination between labour unions and between employers' associations, and the overall bargaining climate. These five coordination indicators have been constructed after a thorough study of each year's collective bargaining in Norway and Sweden in the period from 1961 to 1999, as documented in Barkbu (2000) The operational definitions of the indicators are therefore partly adapted to the tradition of collective bargaining in these two countries, but it should be possible to construct similar indicators for other countries.

The five indicators have been given values {0, 0.5, 1} according to the degree of coordination each year, where 1 generally indicates a high level of coordination. 0 indicates a low level of coordination. The full set of indicator values is shown in table 1.

The first indicator, *I1*, reflects the level of bargaining. Both in Norway and Sweden, collective bargaining has traditionally been dominated by the strongest labour union confederation (LO) and a counterpart employers' association (NHO in Norway; SAF in Sweden). The agreements signed by these organisations are often wage-leading. Prior to each bargaining round, the two organisations decide whether bargaining will be at the peak (national) level or at the industry level, a decision which is believed to be of great importance for the outcome of that year's collective bargaining round. Thus, our operational definition of level of bargaining has been to allocate the value 1 to *I1* if bargaining between LO and NHO/SAF is at the peak or national level, and 0 if bargaining is at the industry or sector level.

The second indicator, *I2*, considers coordination in the form of incomes policies, i.e., government contributions such as tax reliefs, price subsidies etc. for instance in order to moderate wage demands or to increase employment. It is important to note that government contributions are considered as incomes policy solely when they are made conditional on the organisations' actions in the collective bargaining. *I2* is given the value 1 when the government intervenes in the bargaining and is 0 when the government does not actively intervene in order to a;ect the results of the bargaining.

Table 1: Indicators of wage formation coordination and incomes policy.

	Norway						Sweden					
	I1	I2	I3	I4	I5	average	I1	I2	I3	I4	I5	average
1961	0	0	1	1	1	0.6	1	0	1	1	1	0.8
1962	1	0	1	1	1	0.8	1	0	1	1	1	0.8
1963	1	1	1	1	1	1	1	0	1	1	1	0.8
1964	1	1	1	1	1	1	1	0	1	1	1	0.8
1965	1	1	1	1	1	1	1	0	1	1	1	0.8
1966	1	0	1	1	1	0.8	1	0	1	1	1	0.8
1967	1	0	1	1	1	0.8	1	0	1	1	1	0.8
1968	0.5	1	0.5	1	1	0.8	1	0	1	1	1	0.8
1969	1	1	1	1	1	1	1	0	1	1	1	0.8
1970	1	1	1	1	1	1	1	0	0.5	1	1	0.7
1971	1	1	1	1	1	1	1	0	0.5	1	0.5	0.6
1972	1	1	0.5	1	1	0.9	1	0	0.5	1	0.5	0.6
1973	1	1	1	1	1	1	1	1	0.5	1	1	0.9
1974	0	1	0.5	1	1	0.7	1	1	1	1	1	1
1975	1	1	1	1	1	1	1	1	0.5	1	0.5	0.8
1976	1	1	1	1	1	1	1	1	0.5	1	0.5	0.8
1977	1	1	1	1	0.5	0.9	1	0	0.5	1	0.5	0.6
1978	1	1	1	1	0.5	0.9	1	0	1	1	1	0.8
1979	1	1	1	1	0.5	0.9	1	0	1	1	1	0.8
1980	0.5	1	0.5	0.5	0	0.5	1	0.5	0.5	1	0	0.6
1981	1	0.5	0.5	0.5	0	0.5	1	0	0	1	0.5	0.5
1982	0	0	0.5	0.5	0	0.2	1	0	0	1	0.5	0.5
1983	1	0	0.5	0.5	0	0.4	0	0	0	0	0.5	0.1
1984	0	0	0	0.5	0	0.1	0	0.5	0	0	0.5	0.2
1985	1	0	0.5	0.5	0	0.4	0,5	1	0	0	0.5	0.4
1986	0	0	0	0	0	0	0,5	0.5	0	0	0.5	0.3
1987	1	0	1	1	1	0.8	0,5	0.5	0	0	0.5	0.3
1988	1	1	1	1	1	1	0	0	0.5	0	1	0.3
1989	1	1	1	1	1	1	0,5	0.5	0.5	0	1	0.5
1990	0.5	0	1	1	1	0.7	0,5	0.5	0.5	0	1	0.5
1991	1	0.5	1	1	1	0.9	1	0.5	1	1	1	0.9
1992	1	1	0.5	1	1	0.9	1	0.5	1	1	1	0.9
1993	1	1	1	1	1	1	0	0	1	0	1	0.4
1994	0	1	1	1	1	0.8	0	0	1	0	1	0.4
1995	1	1	0.5	1	1	0.9	0	0	0.5	0	0	0.1
1996	0	1	0	1	1	0.6	0	0	0.5	0	0	0.1
1997	1	1	1	1	1	1	0	0	0.5	0	0.5	0.2
1998	0	1	0	0.5	1	0.5	0	0	1	0	1	0.4
1999	1	1	1	1	1	1	0	0	1	0	1	0.4

The indicators *I3* and *I4* reflect the degree of coordination among participants on the employee- and employer side, respectively. As discussed in the theoretical section, labour unions have a common interest in moderate wage demands, but there are obvious reasons for each union, at least in the short run, to deviate from moderation. Often, there have been attempts to coordinate the demands of the di;erent labour unions. The indicator *I3* has been allocated the value 1 when there have been successful coordination e;orts, and 0 if there have been no e;orts or e;orts have been unsuccessful. We have focused on the level of coordination within LO and between LO and the other unions. Similarly, the indicator I4 has been given the value 1 when the employers' associations have not attempted to divert from the main content of other associations' agreements. We have focused on the level of coordination within NHO/SAF and between NHO/SAF and the other associations.

I5 is an indicator of the overall bargaining climate. We have focused on the climate between key labour unions and employers' associations before and during the collective bargaining. In addition, we have taken into consideration the number of strikes appearing during each round. The indicator has the value 1 if the bargaining climate was relatively favourable to an agreement , whereas it obtains the value 0 if there were great disparities between the parties.

An alternative index of corporatism is provided by Iversen (1999, Ch 3). It builds on the idea that the fewer is the number of participants at the crucial levels of the bargaining processes, the easier it is to coordinate the outcomes. Let w_{jt}be a weight (subjectively) assigned to level j bargaining (firm, industry or whole economy) in year t, and let p_{ijt} be the fraction of unionised workers at that level belonging to union i. Iversen's index is then equal to $\sum_j \sum_i w_{jt} p_{ijt}^2$. This index takes a larger value the fewer and the larger (relative to other unions) are the unions operating at the most important levels of wage bargaining. In section 4.2 and 4.3 we investigate the empirical relevance of both our own and Iversen's indices in econometric models of the rates of unemployment in the two countries.

4 The empirical relationship between wage coordination and unemployment

In this section we investigate the role of the coordination indices in explaining the rates of unemployment in Norway and Sweden. Section 4.1 first discusses how econometric methodology can help bridge the gap between our qualitative and static theory on the one hand and the dynamic and non-stationary aspects of observed unemployment rates on the other. Section 4.2 and 4.3 then reports our econometric models. Finally, simulation is used to illustrate the e;ects of hypothetical changes in collective bargaining set-up.

4.1 Methodology

The rates of unemployment in the two countries are observed over a relatively short sample period, and display heterogeneity, persistence and possibly non-stationarity caused by "regime shifts". In contrast, the formal theory in section 2 is static and partial in nature, and a direct application of say, equation (8), after suitable linearization and replacing $\mathtt{Q}$ and c by the relevant data series, is unlikely to provide useful results. A better strategy is to explicitly model the unemployment rate, seeking econometric relationships that are congruent with the available data evidence on the measured attributes.

Our starting point is a general and unrestricted model, denoted *GUM*, that has white-noise residual properties. The final model is chosen on the basis of a general to specific modelling strategy, denoted *Gets:* see Hendry (1995, Ch 8 and 14). Common concerns with the *Gets* specification procedure include the question about the meaning of "significant" t-values after repeated testing; worries that chance correlations and "over-fitting" make the final model become sample dependent or sample specific; and the issue about "path dependency" and "investigators bias", meaning that only a small number of possible simplification paths are investigated and documented. However, recent work by Hoover and Perez (1999) and Hendry and Krolzig (2000) have shown that *Gets* performs surprisingly well in controlled Monte-Carlo situation, especially against the background of the pessimistic results in Lovell (1983) whose work they extend: In most "states of nature" the simplified model is either the correct specification or comes very close to it. Also, the t-statistics of the tested-down models are well behaved.

However, the performance of *Gets* also depends on the size of the *GUM* and specifically on whether it includes the correct local data generating process. In our case, the large number of coordination indices that we wish to test, the need to include additional explanatory variables for congruency and the low (nominal) sample size T, invokes "the curse of dimensionality": Ideally, in order to accommodate all hypotheses, we would like a larger *GUM.* Thus, although we derive our results from a congruent *GUM*, there may be competing *GUMs* that are also congruent and this may induce an element of researcher's bias or "GUM specificity" in our final model. Although this problem by its very nature cannot be resolved in a single study, we seek to reduce the problem by reporting the results of neighbouring *GUM* specifications and other robustness tests.

Given that the *GUM* is well specified, the short sample $T < 40$ may be less of a problem for the applicability of *Gets* than one might think. Experience from "short-sample modelling" in e.g., Nymoen (1992) and Campos and Ericsson (1999), reminds us that what matters is not the nominal sample size but what we may think of as the information content per observation, see Campos and Ericsson (1999). Information content may be large if the data span years with large changes in the individual variables and in the correlation structure between variables. In our case, the late 1980s saw a sharp increase in unemployment in Norway. In the early 1990s an even more dramatic increases occurred in Sweden. Other major events followed the increase in unemployment: A large devaluation (Sweden), revitalization of incomes policies (Norway), and late in the sample an increased role of monetary policy stabilizing inflation.

Our modelling strategy is summarized in three steps:

1. For each country, a single equation model is first presented, in terms of a *GUM* and a final model is chosen from *Gets* modelling.

2. The model resulting from step 1, includes explanatory variables that are potentially endogenous (i.e., reacting to the rate of unemployment, contemporaneously or lagged), and that may be themselves be explainable with the aid of our coordination indices. Therefore, a system of equations model is estimated by *Fiml* and its properties investigated.

3. Based on the model in step 2, the numerical importance of the e;ects of the coordination indices is illustrated by simulation (intervention analysis).

4.2 Norway

The dependent variable is the natural logarithm of the total rate of unemployment, tu_t, defined as the sum of registered unemployed and labour programmes participants divided by the size of the labour force. We have considered to instead model the rate of "open" employment, i.e., without workers on labour market programmes, or a standardized (Labour Force survey) definition of the unemployment rate. Importantly, the substantial findings below are not dependent on the choice of definitions, but (for both countries) we prefer to present the econometric relationships that were achieved for the total rate of unemployment. The use of the logarithmic transform likewise aids congruent modelling.

In addition to the corporative indices we include the following explanatory variables:

1. The log of the replacement ratio, rpr_t.

2. The rate of GDP growth, $^{1} y_t$.

3. The log the wage-share in the manufacturing sector, wsh_t, as a measure of profitability in that sector of the economy.

4. The log of the labour market programme rate, pro_t.

The role of the replacement ratio was motivated theoretically in section 2. The other variables generalizes the model to include the other explanatory factors discussed at the end of section 2. Specifically, Rødseth (1997) and Lindbeck (1997) have emphasised the importance of aggregate demand and exchange rate policy in keeping down unemployment in Norway and Sweden, see also Calmfors and Nymoen (1990) and Rødseth and Nymoen (1999). The two first terms (¹ y_t) and (wsh_t) accommodate this view. The rationale for the third term is that labour market programmes have been used actively to counter swings in employment and to combat skill deterioration and discouraged worker e;ects. The e;ect of these programmes at the level of individual workers remain in dispute, see e.g., Dølvik et al. (1997), but they are an integral part of the determination of the measured rate of unemployment in Norway as well as in Sweden. Finally, we include the lagged rate of unemployment in order to capture the evident adjustment lags. Similar sets of explanatory variables are also used in existing econometric models on quarterly data, see e.g., Bårdsen et al. (1999).

Estimation showed that the replacement ratio failed to assert itself empirically.[2] Part of the reason is that in the sample the replacement ratio is driven by variation in the inflation rate and is correlated with the wage-share variable. Instead of throwing out the wage-share we kept that variable since it represents a channel that coordination indices may a;ect unemployment. Moreover, since Holden and Nymoen (2001) and Rødseth and Nymoen (1999) find significant e;ects of the replacement rate in a model of manufacturing wages, it is possible that the final system of equations, where the wage-share is endogenous, will be more suited to represent the e;ects of the replacement rate on the unemployment rate.

Thus, table 2 shows the results of a general model where only four economic explanatory variables are included (GDP growth, the wage-share, programmes and lagged unemployment). Because of this parsimony in the "economic" part of the equation, there are enough degrees of freedom to estimate freely the coe>cient of all the coordination indices with one lag each.

The bottom part of the table contains statistics that can be used to evaluate whether this general unrestricted model is congruent with the data: The multiple correlation coe>cient (R^2), the residual standard error (ˆ), the Durbin-Watson statistic (DW), the Fdistributed tests of residual autocorrelation (F_{AR}), autoregressive conditional heteroscedasticity (F_{ARCH}). Finally, we include the Doornik and Hansen (1994) Chi-square test of residual non-normality ($:^2_N$), see Doornik and Hendry (1999). The numbers in brackets are p-values for the respective null hypotheses, e.g., of no residual 2. order autocorrelation in the case of $F_{AR}(2,16)$,—they show

[2] A test of the hypothesis that the omission of rpr_t and $rpr_{t\ 1}$from table 2 gives $F(2,18) = 0.95932[0.4019]$: The result does not depend on whether the log-transform (i.e., rpr_t) is used or not.

Table 2: Norway: Modelling Δtu_t from 1963 to 1998. (OLS estimation results of a general model)

Variable	Coefficient	Standard Error	t-probability
Constant	5.1597	3.6452	0.1723
tu_{t-1}	-0.32532	0.10545	0.0058
Δy_t	-5.8475	1.4800	0.0008
wsh_{t-1}	0.81767	0.57023	0.1670
Δpro_t	0.53405	0.06800	0.0000
pro_{t-1}	0.22183	0.10652	0.0503
$I1_t$	-0.064850	0.064973	0.3301
$I1_{t-1}$	-0.14090	0.064295	0.0404
$I2_t$	-0.019158	0.060659	0.7554
$I2_{t-1}$	-0.037660	0.044807	0.4106
$I3_t$	-0.047620	0.080262	0.5596
$I3_{t-1}$	-0.025782	0.081161	0.7540
$I4_t$	0.16788	0.15857	0.3024
$I4_{t-1}$	0.35464	0.27058	0.2048
$I5$	0.090701	0.14272	0.5323
$I5_{t-1}$	-0.23347	0.23883	0.3400

$R^2 = 0.922$ $\hat{\sigma} = 0.090$ $DW = 2.1$
for 16 variables and 36 observations.
$F_{AR}(2,18) = 1.1177[0.3487]$,
$F_{ARCH}(1,18) = 2.898[0.1059]$
$\chi^2_N(2) = 0.54848[0.7601]$

that none of the diagnostic tests are significant.

Turning to the coefficient estimates, note that almost every variable in the economic part of the equation is significant, the least significant variable is the lagged wage-share, with a significance level of 0.17. Among the coordination indices, only $I1_{t-1}$ is significant at the 5% level, and the majority of the variables have very high significance levels. Thus, an F-test on the omission of all 10 indicator variables gives $F(10,20) = 1.09[0.42]$, so the lack of significance of the majority of the indicator variables is not due to multicollinearity as such. Omission of *all* indicator variables other than $I1_t$ and $I1_{t-1}$ yielded $F(6,22) = 0.41[0.86]$, so it seemed reasonable to keep only those two indicator variables. The lagged wage-share variable now became significant, and it was also convenient to restrict the coefficient of pro_t to be twice that of pro_{t-1} (already evident in table 2) resulting in a single term $\Delta pro_t + pro_t$. Lastly, and perhaps exposing ourselves to the danger of "overfitting", we joined up $I1_t$ (t-value 1.01) together with its significant lag (2.7), in the form of the variable $I1_t + 2 \cdot I1_{t-1}$.

The final equation is therefore

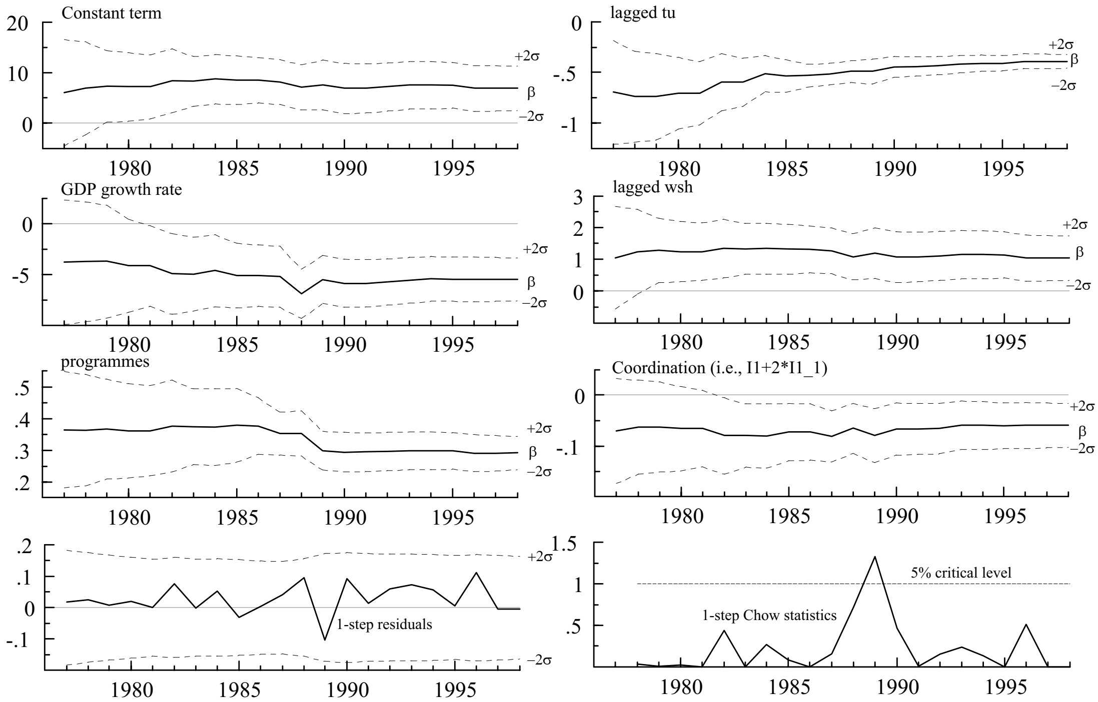

Figure 1: Recursive stability of final model of the Norwegian rate of unemployment.

$$
\begin{aligned}
(9) \qquad \Delta\widehat{tu}_t = {} & \underset{(2.1878)}{6.94} \quad \underset{(0.034)}{0.388}\, tu_{t-1} \quad \underset{(1.055)}{5.457}\, \Delta y_t + \underset{(0.349)}{1.046}\, wsh_{t-1} \\
& + \underset{(0.0259)}{0.292}\, (\Delta pro_t + pro_t) \quad \underset{(0.022)}{0.0594}\, (I1_t + 2 \cdot I1_{t-1})
\end{aligned}
$$

$$R^2 = 0.903 \quad \hat{\sigma} = 0.082 \quad DW = 2.23 \quad F_{ENC}(10, 20) = 0.47[0.89]$$

$$F_{AR}(2, 28) = 2.56[0.10] \quad F_{ARCH}(1, 28) = 1.34[0.26]$$

$$\chi^2_N(2) = 1.16[0.56] \quad F_{X2}(10, 19) = 1.30[0.30]$$

The estimated coefficients of the economic variables have barely changed from the general model, and their standard errors (reported below the coefficient estimates) have been reduced considerably.[3] In all, the equation imposes 10 restrictions on the general model in table 2. The F-test of the joint significance of the restriction is reported below the equation as $F_{ENC}(10, 20)$, showing that with 10 fewer parameters, equation (9) fits the data almost as well as the model in table 2. That the other residual statistics in equation (9) are insignificant is also evidence that the reduction is valid.[4]

[3] Note that in table 2, the estimated coefficient of Δpro_t is almost twice as large as the coefficient of pro_{t-1}. Imposing that restriction gives the variable $\Delta pro_t + pro_{t-1}$ in equation (9).

[4] The diagnostics to equation (9) also includes a test of heteroscedasticity due to squares of the regressors (F_{X2}), as implemented in PcGive 9.3.

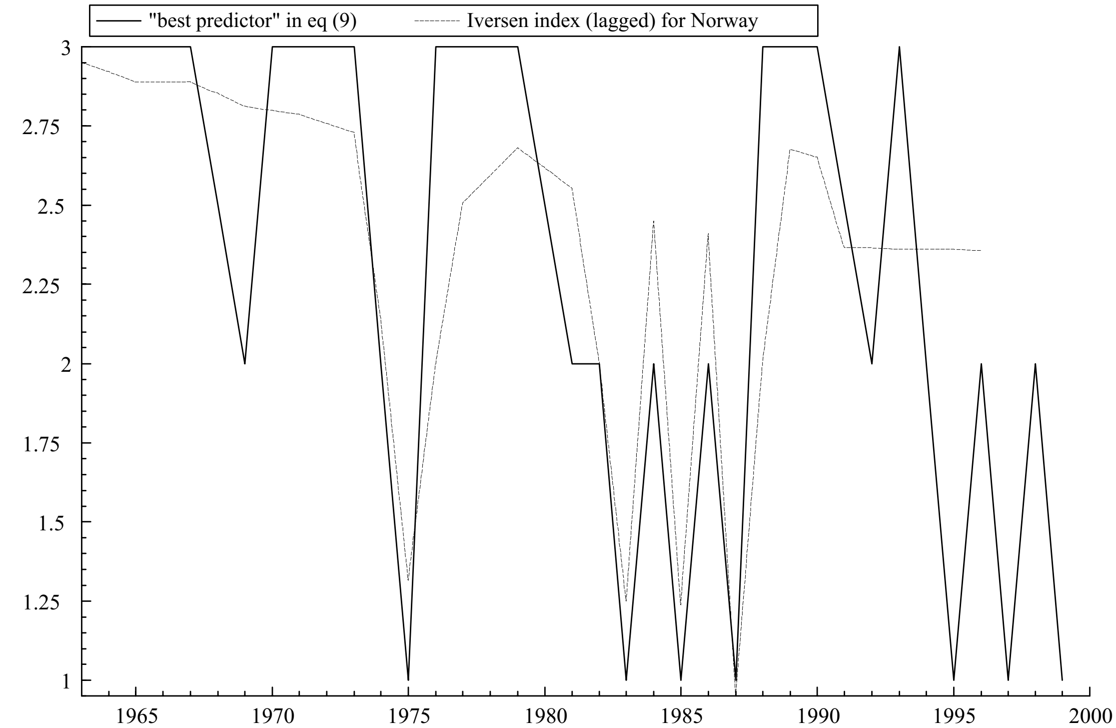

Figure 2: Comparison with Iversen (1999) index with the index used as regresson in equation (9), i.e., "best predictor". Scale adjusted for readability

Figure 1 shows the stability of equation (9) over the period 1977-98. The six first graphs show the recursively estimated elasticities in (9), with ±2 estimated coe>cient standard errors, denoted ¯ and ±2¨ in the graphs. The last two graphs show first the 1-step residuals with ±2 residual standard errors, ±2¨ in the graph, and finally the sequence of 1-step Chow statistics scaled with their 5% critical levels. All graphs show a high degree of stability.

The potential omission of the replacement ratio can be addressed again using equation (9). For the null hypothesis that both rpr_t and $rpr_{t\text{P}1}$ have zero coe>-cients the relevant statistic became $F(2, 28) = 1.3178[0.2838]$. The most significant e;ect obtained was for the di;erenced variable ¹ rpr_t, which obtained a coe>cient of 0.43013 (with "t-value" of 1.614) when added to equation (9). The estimated coe>cients in equation (9) were practically una;ected by the inclusion of ¹ rpr_t. As already noted we will consider the possibility of indirect e;ect of the replacement rate, when we model a system of equations that contains the wage-share as an endogenous variable. A variable for the coverage by the main trade confederation (LO) was always tried added to equation (9), but neither the union coverage itself nor its di;erence were statistically significant.

As already noted, there is a danger that the inclusion of the corporative variable in the final model is fortuitous. Interestingly, it is possible to get a check on that result by using the index of centralization of wage bargaining in Iversen (1999, Chs 1 and 3): Replacing $I1_t{+}2{\cdot}I1_{t\mathrm{P}1}$ in (9) with Iversen's index (lagged) yields a coe>cient of -0.38947 with a "t-value" of -3.122 over a sample that ends in 1996.[5] Moreover, Iversen's index of wage centralization is highly correlated with $I1_t + 2 \cdot I1_{t\mathrm{P}1}$, as showed in figure 2. An approximation to Iversen's index, based on regression over the period 1963-1995, is:

$$
\text{(10)} \qquad \begin{aligned}
\widehat{Iver}_t &= \underset{(0.04)}{0.32} + \underset{(0.05)}{0.18}\ I1_t + \underset{(0.06)}{0.15}\ I3_t \\
R^2 &= 0.60 \qquad DW = 1.53
\end{aligned}
$$

which suggests that in terms of our own operational definitions, Iversen's index can be expressed as a combination of the level of bargaining and the degree of worker side coordination.[6]

In sum there is evidence that the level of bargaining, and possibly also the level of worker side coordination, has a;ected Norwegian unemployment. In order to gauge the numerical significance of the level of bargaining e;ect in (9), assume that the going rate of unemployment is 10%. According to the model, a change from decentralized to centralized wage settlements then lowers the rate of unemployment to 8.8% in the course of two years. However, that illustration is too partial since both the rate of labour market programmes (pro_t) and the wage-share (wsh_t) change when the rate of unemployment is reduced. Hence, the the impact on unemployment of moving from centralization to decentralization may be both larger and smaller than the single equation results suggest.

The model in table 3 provides a wider setting to consider these issues. We started from an unrestricted 1st order vector autoregression (VAR) in the three variables $^1\, tu_t$, pro_t and wsh_t. Even though only $I1$ was found to play a role in the tu equation, there may be indirect e;ects of the other indicators, (via pro_t and wsh_t). Therefore all five indicators and their lags were included in the VAR. For the same reason the replacement ratio rpr was also included.

The estimated unemployment equation is virtually unchanged from the single equation results. For ease of interpretation we have however omitted $I1_t$, the least significant part of the composite dummy $I1_t + 2 \cdot I1_{t\mathrm{P}1}$in equation (9). The second equation, explaining the programme rate, is basically a reaction function that pegs the programme rate to the total rate of unemployment. However, it appears that a higher (manufacturing sector) wage-share lowers the programme rate, reflecting that the government is making the provision of active labour market programmes conditional on the maintenance of cost competitiveness of the manufacturing sector.

The third equation explains the wage-share by the lagged rate of unemployment and the lagged replacement rate. The significance of $rpr_{t\mathrm{P}1}$ is evidence of

[5] The estimates of the other coe>cients in (9) are practically una;ected.

[6] We note that, using $\widehat{Iver}_{t\ \ 1}$ in the place of $I1_t + 2 \cdot I1_{t\ \ 1}$ gives a coe>cient of -0.325 with a t-value of -2.245.

Table 3: Norway: FIML estimation results.

The unemployment equation

$$\widehat{{}^{1}tu_t} = \underset{(2.678)}{7.029} \quad \underset{(0.075)}{0.381}\, tu_{t\text{P}1} \quad \underset{(1.070)}{5.618}\ {}^{1}y_t$$

$$+\ \underset{(0.071)}{0.297}\ ({}^{1}pro_t + pro_t) +\ \underset{(0.420)}{1.061}\ wsh_{t\text{P}1} \quad \underset{(0.037)}{0.099}\ I1_{t\text{P}1}$$

$$\hat{\ } = 0.083$$

The programme rate equation

$$\widehat{pro}_t = \underset{(6.15)}{19.81} + \quad tu_{t\text{P}1} \quad \underset{(0.97)}{2.94}\ wsh_{t\text{P}1}$$

$$\hat{\ } = 0.307$$

The wage-share equation

$$\widehat{wsh}_t = \underset{(0.814)}{2.311} + \underset{(0.130)}{0.634}\ wsh_{t\text{P}1} \quad \underset{(0.008)}{0.014}\ tu_{t\text{P}1}$$

$$\underset{(0.012)}{0.033}\ I2_t + \underset{(0.010)}{0.015}\ (I3_t + I3_{t\text{P}1})$$

$$\underset{(0.023)}{0.054}\ {}^{1}I5_t + \underset{(0.047)}{0.120}\ rpr_{t\text{P}1} \quad \underset{(0.030)}{0.114}\ i1979_t$$

$$\hat{\ } = 0.028$$

Diagnostics

$:^2_{ENC}(35)$	=	36.4221[0.4024]
$vF_{AR}(18,68)$	=	1.8398[0.0387]
$v:^2(6)$	=	8.7506[0.1881]

The sample is 1963 to 1998, 36 observations.

an indirect e;ect of the labour insurance system on unemployment. Finally, there are four indicator variables in the wage-share equation. First, the incomes policy indicator ($I2$) which has a negative coe>cient. Second, the two year average of $I3$ (worker side coordination) has a positive e;ect on the wage-share. Third, the term ${}^{1}I5_t$ captures a short-run e;ect of a change in what we have called the bargaining climate. Finally, and outside the list of bargaining indicators, $i1979_t$ is an impulse dummy for 1979, when there was a wage and price freeze.

The statistic $:^2_{ENC}(35)$ in the *Diagnostics* part of the table shows there are 35 (overidentifying) restrictions on the underlying VAR, which are jointly insignificant. Hence, the econometric model in the table encompasses the VAR, see Hendry and Mizon (1993). The other test statistics are vector tests of 2nd order residual autocorrelation, and of residual non-normality. These statistics are explained in Doornik and Hendry (1996b). We note that the test of autocorrelation is significant at the 5% level, so there is room for improvements in terms of fit. However, the estimated model is su>ciently congruent to serve as an illustration of the numerical importance of changes in the bargaining indicator variables.

Consider first a change (from 0 to 1) in $I1$, representing a move from decentralization to centralization. Assuming initial rates of 10% and 4% for total unemployment and labour market programmes, total unemployment is predicted to be reduced to 6.4%, the programme rate rate to 2.8% and the rate of open unemployment from 6% to 3.6%, all within a time span of 4 years. A policy change to "active" incomes policy (*I2*) leads to quite di;erent predictions: Total unemployment increases to 10.4% because the lower wsh allows an increase in the programme rate, to 5%, and consequently the rate of open unemployment is reduced from 6% to 5.6%. However, in the longer run both unemployment rates are reduced in this experiment. Finally, consider a change in worker coordination ($I3$) from "low" to "high", which according to the model leads to an increased wage-share. Thus there is lowering of the programme rate, from 4% to 3.2%, and therefore total unemployment is reduced (to 9.7%), while the rate of open unemployment is increased, to 6.4%.

4.3 Sweden

The estimated coe>cients of the general model in table 4 shows that the results for the economic explanatory variables (¹ y_t, $wsh_{t\mathrm{P}1}$ and pro_t) come close to the results for Norway in table 2.[7] However, neither the economic variables nor the institutional indictor variables are of any help in explaining the raise in the rate of unemployment early in the sample—in 1964-1967 and in 1967-1976. Hence, we include two dummies, $DUM1_t$ and $DUM2_t$ for these time periods, as one way of whitening the residuals.[8]

Interestingly, the results for the indicator variables are somewhat di;erent from what we obtained for Norway. The current value of the income policy indicator $I2_t$ has a t-probability of 0.09. Moreover, both the employer coordination indicator $I4$, and the climate variable $I5_t$ also have levels of significance better than 0.5 which, following our liberal strategy for simplifying the GUM, means that we keep the variables. At first sight, the positively signed coe>cient of the level of bargaining indicator $I1_t$ is troublesome. However, note that there is no variation in this indicator between 1964 and 1982, as all wage settlements are classified as centralized. Thus, the coincidence of a sharp increase in unemployment and a (temporary) return to centralized bargaining in 1991 probably gets a high leverage for the coe>cient of $I1_t$.

Based on table 4, we have obtained the final equation for Sweden shown in equation (11). The included economic variables are the same as for Norway (compare equation (9)), and the estimated coe>cients are strikingly similar. The equation retains $I2$ (incomes policy), the lagged value of $I4$ (employer coordination) and $I5$

[7] In the case of Sweden the wage share variable is for the private sector, and ! y_t is a measure of aggregate demand, see Forslund and Kulm (2000)

[8] $DUM1$ is one in the years 1964-1967, zero elsewhere. $DUM2_t$ is one in the period 1967-1976, zero elsewhere.

Table 4: Sweden: Modelling Δtu_t from 1964 to 1997. (OLS estimation results for a general model)

Variable	Coefficient	Std. Error	t-prob
Constant	1.0639	0.61736	0.1041
tu_{t-1}	-0.18414	0.13925	0.2047
Δy_t	-6.2393	0.99710	0.0000
wsh_{t-1}	0.95619	1.0075	0.3567
Δpro_t	0.71079	0.094271	0.0000
pro_{t-1}	0.25640	0.15798	0.1241
$I1_t$	0.20389	0.12596	0.1251
$I1_{t-1}$	-0.0011157	0.13236	0.9934
$I2_t$	-0.10358	0.057187	0.0889
$I2_{t-1}$	0.0027546	0.055332	0.9609
$I3_t$	-0.021825	0.091881	0.8153
$I3_{t-1}$	0.054207	0.081781	0.5169
$I4_t$	-0.13983	0.13396	0.3121
$I4_{t-1}$	-0.17576	0.10951	0.1281
$I5$	-0.084775	0.080386	0.3073
$I5_{t-1}$	0.0078693	0.067400	0.9085
$DUM1_t$	0.24635	0.10003	0.0255
$DUM2_t$	0.40040	0.15568	0.0205
$R^2 = 0.948$ $\hat{\sigma} = 0.070$ $DW = 2.6$			
for 18 variables and 34 observations.			
$F_{AR}(2,14) = 2.7892[0.0956]$,			
$F_{ARCH}(1,14) = 0.67408[0.4254]$			
$\chi^2_N(2) = 3.6591[0.1605]$			

(climate). Taken at face value, this equation therefore indicates a richer response of unemployment to changes in coordination indicators than was the case of Norway, with a particular vigorous effect of employer coordination on the rate of unemployment.

(11)
$$\Delta \widehat{tu}_t = \underset{(0.447)}{1.342} \quad \underset{(0.057)}{0.264}\, tu_{t-1} \quad \underset{(0.833)}{6.017}\, \Delta y_t + \underset{(0.694)}{1.274}\, wsh_{t-1}$$
$$+ \underset{(0.036)}{0.334}\, (\Delta pro_t + pro_t) \quad \underset{(0.035)}{0.094}\, I2_t \quad \underset{(0.050)}{0.183}\, I4_{t-1}$$
$$\underset{(0.044)}{0.083}\, I5_t + \underset{(0.075)}{0.281}\, DUM1_t + \underset{(0.111)}{0.476}\, DUM2_t$$
$$R^2 = 0.931 \quad \hat{\sigma} = 0.065 \quad DW = 2.3 \quad F_{ENC}(8,16) = 0.64[0.73]$$
$$F_{AR}(2,22) = 0.43[0.65] \quad F_{ARCH}(1,22) = 0.00[0.99]$$
$$\chi^2_N(2) = 3.53[0.17] \quad F_{X2}(15,8) = 0.33[0.97].$$

There is considerable stability in estimated coefficients, as shown in figure 3, despite that the shocks to the labour market in Sweden around 1990 were even larger

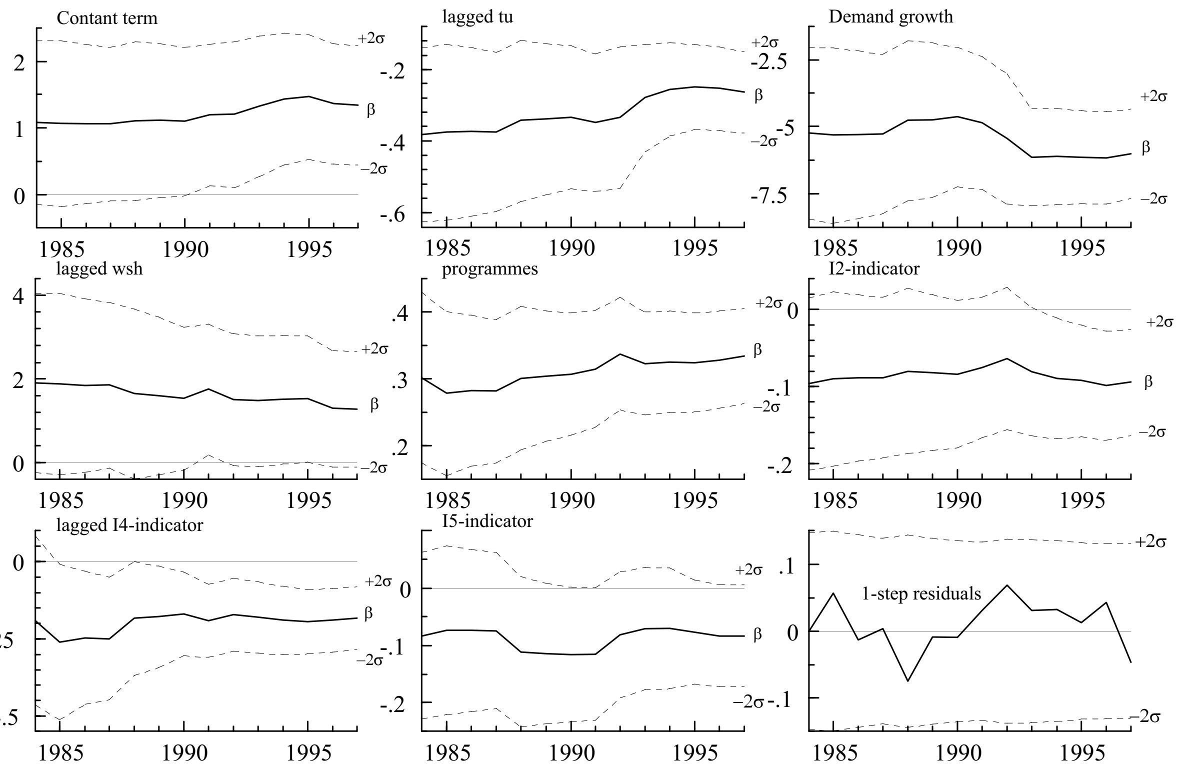

Figure 3: Recursive stability of final model of the Swedish rate of unemployment in equation (11).

than in Norway.[9]

In the same manner as for Norway, we investigated the e;ects of adding the replacement rate (and its lag) to equation (11), and with similar results: Both the current and lagged value obtained negative coe>cient estimates, and the statistic for the joint null hypothesis of both coe>cients being zero was F(2, 22) = 1.471 [0.2514].

However, introduction of Iversen's indicator for Sweden produced interesting results: It obtains a significant "t- value" of 2.4 when added to (11) and the estimated coe>cient of $I4_{t\text{P}1}$ is changed from 0.18 to 0.08, i.e., it is now in line with the coe>cients of $I2_t$ and $I5_t$. Using our set of indicator variables in an attempt to explain Iversen's index yields the following regression over the sample 1965-1998:

$$
\begin{array}{lrcl}
(12) & \widehat{Iver}_t & = & \underset{(0.02)}{0.39} + \underset{(0.04)}{0.11}\, I1_t + \underset{(0.03)}{0.09}\, I4_{t\text{P}1} \\
 & R^2 & = & 0.66 \qquad DW = 0.52.
\end{array}
$$

Although there is positive residual autocorrelation which makes it di>cult to assess

[9]In order not to glut the figure, the panel with the 1-step Chow statistics has been omitted, but none of the statistics are significant over the period 1979-1997.

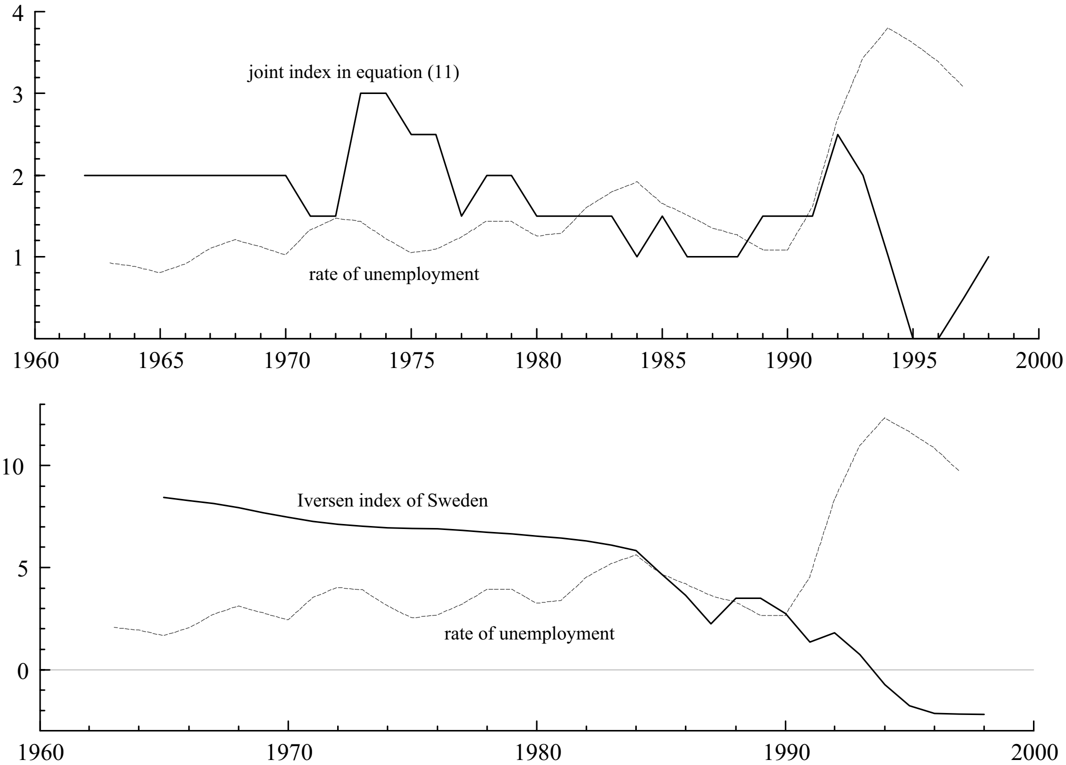

Figure 4: Sweden. Upper panel: The best predictor indicator variable $I2_t + I4_{t-1} + I5_t$ and the rate of unemployment. Lower panel: Iversen's index and the rate of unemployment. Graphs in both panels are scale adjusted.

the statistical significance of the two retained regressors, equation (12) helps explain why the numerical importance of $I4_t$ is reduced when $Iver_t$ is added to model (11). Moreover, the fact that equation (12) fails to explain $Iver_t$ very well, suggests that Iversen's index contains separate information relative to our set of indicator variables. Figure 4 compares the two indicators with the Swedish rate of unemployment in two separate panels. The visible correlation in the upper panel reflects that the index by construction is the best predictor of the rate of unemployment among the possible linear combinations of the indices included in table 4. The interpretation of Iversen index is hampered by its trend behaviour. However, it is strongly correlated with the rate of unemployement in the 1990s.

Using both Iversen's and our own set of variables yields the equation (13) for the Swedish rate of (total) unemployment.

$$
\begin{aligned}
(13) \qquad \Delta \widehat{tu}_t &= \underset{(0.310)}{1.231} - \underset{(0.045)}{0.343}\, tu_{t-1} - \underset{(0.627)}{4.680}\, \Delta y_t + \underset{(0.470)}{0.985}\, wsh_{t-1} \\
&\quad + \underset{(0.028)}{0.340}\, (\Delta pro_t + pro_t) - \underset{(0.022)}{0.080}\, (I2_t + I4_{t-1} + I5_t) \\
&\quad - \underset{(0.190)}{0.644}\, Iver_t + \underset{(0.047)}{0.227}\, DUM1_t + \underset{(0.068)}{0.453}\, DUM2_t \\
R^2 &= 0.95 \qquad \hat{\sigma} = 0.051 \quad DW = 2.1
\end{aligned}
$$

Table 5: Sweden: FIML estimation results.

The unemployment equation

$$\widehat{{}^{1}tu_t} = \underset{(0.49412)}{1.45} \quad \underset{(0.101)}{0.327}\, tu_{t\text{P}1} \quad \underset{(0.667)}{4.503}\ {}^{1}y_t$$
$$+ \underset{(0.992)}{1.335}\, wsh_{t\text{P}1} + \underset{(0.101)}{0.330}\, ({}^{1}pro_t + pro_t) \quad \underset{(0.027)}{0.087}\, (I2_t + I4_{t\text{P}1} + I5_t)$$
$$\underset{(0.224)}{0.698}\, Iver_t + \underset{(0.063)}{0.251}\, DUM1_t + \underset{(0.106)}{0.427}\, DUM2_t$$
$$\hat{} = 0.057$$

The programme rate equation

$$\widehat{pro}_t = \underset{(0.431)}{0.321} + \underset{(0.062)}{0.832}\, tu_{t\text{P}1} \quad \underset{(0.684)}{3.400}\, wsh_{t\text{P}1}$$
$$\hat{} = 0.174$$

The wage share equation

$$\widehat{wsh}_t = \underset{(0.072)}{0.587} + \underset{(0.104)}{0.269}\, wsh_{t\text{P}1} \quad \underset{(0.008)}{0.082}\, tu_{t\text{P}1}$$
$$+ \underset{(0.009)}{0.048}\, (I1_t + I3_t) \quad \underset{(0.0076)}{0.034}\, (I4_t + I5_t + I5_{t\text{P}1})$$
$$\underset{(0.011)}{0.049}\, DUM1_t \quad \underset{(0.018)}{0.107}\, DUM2_t$$
$$\hat{} = 0.013$$

Diagnostics

$:^2_{ENC}(35)$	$=$	48.34[0.07]
$vF_{AR}(18,54)$	$=$	0.95[0.53]
$v:^2(6)$	$=$	8.64[0.20]

The sample is 1964 to 1997, 34 observations.

In table 5 the model in (13) is estimated jointly with equations for the programme rate and the wage-share. In line with the results for Norway, we find that some of the indicator variables are a;ecting the rate of unemployment indirectly, via the wage-share equation. One indicator, $I4$ (union coordination), enters in both the equation for the rate of total unemployment and in the wage-share equation.

The model in table 5 can be used to illustrate the numerical importance of changes in the di;erent dimensions of Swedish corporatism. The indicators divide into two groups: A shift in the level of bargaining to centralization ($I1 = 1$)), or to strong worker coordination ($I3 = 1$) is predicted to *increase* unemployment from 10% to 16.6% (total), and from 6% to 10.2% (open), after fours years.[10] The mechanism behind this is a rise in real wages.[11] On the other hand, a shift in the incomes policy indicator ($I2_t$) from decentralization to centralization, to high firm side coordination ($I4$), or to a favourable bargaining climate ($I5$), all produce the expected

[10] In the same manner as for Norway we assume inital rates 10% and 4% for total unemployment and labour market programmes

[11] $Iver_t$ is unchanged in this eperiment even though $I1$ is a regressor in equation (12). The reason is of course that its significance may be spurious due to residual autocorrelation.

fall in unemployment: For incomes policy ($I2$), the rate of total unemployment is predicted to fall to 8.3%, and open unemployment to 3.9%. A shift to high firm side coordination ($I4 = 1$) leads to a sharp reduction in total unemployment, to 4.9%[12]. However, because the provision of programmes also is reduced, the fall in open unemployment is more moderate, from 6% to 5%. Finally, for the "climate" indicator $I5$, the results suggest a 4.8 percentage point reduction in total unemployment and a 3 percentage point fall in open unemployment.

5 Conclusion

We have presented econometric models of the Swedish and Norwegian unemployment rates that have a similar structure, even in terms of the estimated coe>cient values of individual explanatory variables. This similarity supports the view that the models capture important factors behind unemployment in the two neighbouring economies. Moreover, the models are congruent with the data properties and have coe>cients that are stable over the 1980s and 1990s. Thus, the models represent a valid framework for testing the individual and joint e;ects of the institutional developments that we have measured with a new set of coordination indices.

As discussed in section 3, there are several dimensions of coordination in collective wage bargaining. This suggest that there may be some value added in a set of indices for each country, i.e., for each facet of coordination, in addition to joint indices of the type presented in the seminal work of Iversen (1999). Our empirical results lend some support to that view. First, it appears that some but not all facets of "high coordination" are favourable to low unemployment. For example, the Swedish results indicate that periods of highly centralized wage bargaining ($I1$) and worker side coordination ($I3$) have made it di>cult to attain the goal of full employment, since firms' profitability have su;ered. On the other hand, active income policy ($I2$), high firm side coordination and a favourable climate ($I5$) between the bargaining parties, appear to have been good for keeping the rate of unemployment low. Second, it is not obvious that the di;erent facets of wage settlements have the same impact on unemployment in the two countries, since the eventual e;ects will depend on the interactions with other parts of the political economy. For example, while highly centralized settlements have increased unemployment via increased wage-share in Sweden, a high value of the $I1$ indicator is singled out as one beneficial factor for keeping unemployment low in Norway. Taken at face value, these results indicates that the label "high centralization" actually means di;erent things in the Swedish and Norwegian political and economic systems. This possibility has been discussed by e.g., Freeman (1997) and Rødseth and Nymoen (1999), and one corroborating observation is that Sweden seems to have abandoned centralization

[12] Again, $Iver_t$ is kept unchanged.

permanently in 1993, while in Norway it remains very much a part on the picture, see table 1 above.

There is a possibility that these and other substantive conclusions lack robustness because the econometric models on which they are based may be influenced by e.g., fortuitous correlation and “path dependency” in going from the general to the final model, cf. section 4.1. Typically, new data when they become available, will corroborate some of our conclusions, while others may indeed turn out to reflect “data mining”. For the time being, we venture that the estimated e;ects of changes in wage coordination are statistically more reliable for Sweden than for Norway. We also note that Iversen’s (1999) indices of wage bargaining centralization, when included into our econometric equations for Sweden and Norway , support the conclusions based on our new set of indices. Another issue that we want to return to in future research, concerns the possible endogeneity of the coordination indices themselves, possibly within a framework that pools the data set of the two countries, thus taking advantage of the similarities in structure that this paper has disclosed.

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