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Nordic wage formation and unemployment seven years later

By
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List of the last 10 Memoranda

No 28	By Hilde Bojer: Equivalence Scales and Intra-household Distribution. 12 p.
No 01	By Gabriela Mundaca and Jon Strand: Speculative attacks in the exchange market with a band policy: A sequential game analysis. 45 p.
No 02	By Pedro P.Barros and Tore Nilssen: Industrial Policy and Firm Heterogeneity. 18 p.
No 03	By Steinar Strøm: The Economics of Screening Programs. 8 p.
No 04	By Kai Leitemo and Øistein Røisland: Choosing a Monetary Policy Regime: Effects on the Traded and Non-Traded Sectors. 39 p.
No 05	By Eivind Bjøntegård: The Composite Mean Regression as a Tool in Production Studies. 9 p.
No 06	By Tone Ognedal: Should the Standard of Evidence be reduced for White Collar Crime? 27 p.
No 07	By Knut Røed and Tao Zhang: What Hides Behind the Rate of Unemployment? Micro Evidence from Norway. 39 p.
No 08	By Geir B. Asheim, Wolfgang Buchholz and Bertil Tungodden: Justifying Sustainability. 24 p.
No 09	By Asbjørn Rødseth and Ragnar Nymoen: Nordic wage formation and unemployment seven years later. 53 p.

Nordic wage formation and unemployment seven years later*

Asbjørn Rødseth[†] and Ragnar Nymoen[‡]

April 28, 1999

Abstract

The paper performs a reinvestigation of Nordic wage formation along the lines of Calmfors and Nymoen (1990) with data that include the recent period of high unemployment and low inflation. Among the main findings are: 1)There has been no shift in Nordic wage formation since 1987. 2)The wage formation in Denmark, Finland, Norway and Sweden is fairly similar, even though the unemployment histories differ. 3)Real wage flexibility is not particularly high in the Nordic countries. 4)No strong conclusions can be drawn about the effects of active labour market programs, but an increase in the supply of such programs seems to increase the equilibrium *total* unemployment rate. 5)Only in Norway is there evidence that incomes policy has been effective. The paper gives some, but limited, support to those who attribute the good employment performance of Norway and Sweden until the late 1980s to aggregate demand factors and devaluations.

*The paper draws heavily on joint work with Oddbjørn Raaum and Fredrik Wulfsberg. We have benefitted from comments from Anders Forslund, Kalle Moene, Steinar Holden and from seminar participants at the Universities of Stockholm and Uppsala and at FIEF in Stockholm. We are also grateful to Torben Andersen, Dominique Anxo, Tor Eriksson, Anders Forslund, Anders Harkman, Kari Hämalainen, Jens Holst Jensen, Oddbjørn Raaum and Niels Westergård-Nielsen for supplying data, and to Snorre Evjen, Colin Forthun, Espen Erlandsen and Tore Langset for research assistance. Financial support was received from the Nordic Council of Ministers. Much of Rødseth's work on the paper was done during visits to the Institute of International Economic Studies at the University of Stockholm and to Stanford University.

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1 Introduction

In the late 1980s there was a sharp increase in unemployment in Norway. In the early 1990s even more dramatic increases in unemployment followed in Sweden and Finland (Figure 1). Old explanations for the previous low unemployment in these countries - centralised bargaining and an active labour market policy - have been called into question, and a new explanation - progressive income taxes (Lockwood and Manning (1993)) - have been advanced. Lindbeck (1997) and Rødseth (1997) have emphasised the importance of aggregate demand and exchange rate policy in keeping down unemployment in Sweden and Norway respectively. These developments merit a new study of Nordic wage formation along the lines of Calmfors & Nymoen (1990), especially since there has been a debate about what Europe could learn from the earlier low unemployment in Sweden (and Norway)[1]. Our sample consists of annual observations from the 1960s to 1994. The sample period in Calmfors & Nymoen (1990) is thus extended by seven new and turbulent years.

Has there been a structural change in Nordic wage formation, or do earlier findings, based on samples with low unemployment, still hold? This is a main question in the present paper. Its relevance is enhanced by other major events which followed the increase in unemployment: large devaluations (Sweden and Finland), a productivity 'miracle' (Finish manufacturing), active incomes policies (Norway). This provides a strong testing ground for wage equations.

Unlike its Nordic neighbours, Denmark has had high unemployment since the late 1970s. However, the increase in the unemployment rate from 1987 to 1993 was greater there than in Norway. In Finland unemployment increased in the second half of the 1970s, as in most of Europe, but Finland avoided the second round of increased unemployment in the early 1980s. An underlying theme in the present study is the question of heterogeneity versus common features of Nordic wage formation. May the different unemployment

[1] See Calmfors (1993) for further references and an assessment of the debate

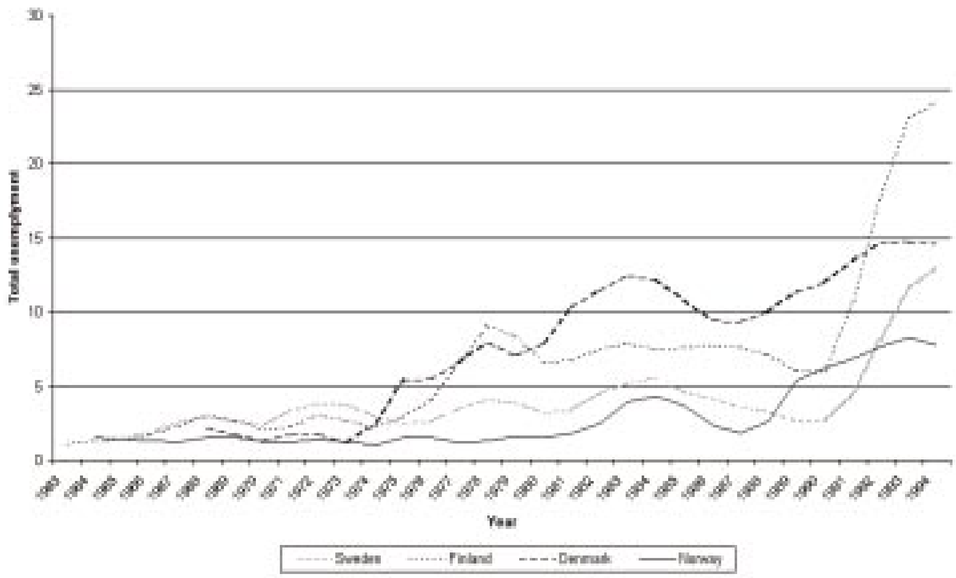

Figure 1: Total unemployment rates (including participants in labour market programs)

histories of e.g. Denmark and Norway be explained by structural differences in wage formation?

The main emphasis in the present paper is on the stability of the wage equations, on the degree of long run real wage flexibility, on the effects of active labour market policy and on possible shifts in the equilibrium unemployment rate[2]. The estimated wage equations are deliberately traditional, as we wanted to connect to the discussion in Calmfors & Nymoen (1990), Layard, Nickell and Jackman (1991) and the country studies in Calmfors (1990). We also discuss some methodological problems within this tradition. As in most earlier studies, we focus on manufacturing wages.

In the next section we review the economic theory which is used to interpret the estimated wage equations. In Section 3 we present the data and discuss some econometric issues. The estimation results are presented in Section 4, while Section 5 concludes.

[2] A separate paper, Evjen and Nymoen (1997), focuses on the effect of incomes policy in Norway and on the appropriate functional form of the wage curve. Another separate paper, Nymoen and Rødseth (1996), looks in more detail at the effects of taxes on wages.

2 Theory

Imagine a small open economy. There are a large number of identical firms. All of them produce for the world market and are price takers. All have production functions which are homogeneous of degree one in labour and capital. Each firm and its local union first bargain for the wage rate. The firm then decides how much labour to employ.

The firm's objective is to maximise profits. In the short run its capital stock is given, and the profit function of a typical firm is then

$$\Pi = \pi(\omega)K \tag{1}$$

where Π is real profits, π profits per unit of capital, ω the real wage, and K the capital stock in the firm.

The union's objective is to maximise the utility of an employed worker, $\nu(\omega)$, where $\nu' > 0$ and $\nu'' \leq 0$. We apply the asymmetric Nash-bargaining solution. This means that the outcome of the wage bargain is the real wage which maximises the Nash-product

$$[\nu(\omega) - \nu_0(\Omega, U)]^{\beta}[\pi(\omega)K]^{1-\beta}$$

ν_0 is a fall-back or reference utility which depends positively on the average real wage in the economy, Ω, and negatively on the economy-wide unemployment rate, U. $\beta \in [0,1]$ measures the relative bargaining strength of the unions.

The first order condition for maximisation of the Nash-product can be written

$$\frac{\nu'(\omega)\omega}{\nu(\omega) - \nu_0(\Omega, U)} = \frac{1-\beta}{\beta}\frac{n(\omega)\omega}{\pi(\omega)} \tag{2}$$

In the derivation we have used that the firm's demand for labour is $N = n(\omega)K = -\pi'(\omega)K$. The first order condition says that at the chosen real wage the relative gain to the union of a wage increase should be proportional to the relative loss to the firm, the factor of proportionality being determined by relative bargaining strength.

Since all firms are identical, we can solve for the aggregate wage level by setting $\Omega = \omega$ in (2). Provided that $\nu_{0\Omega}$ is not too large, this yields a downward sloping relationship between the real wage and unemployment, $\omega = g(U)$, which is called the wage (setting) curve.[3]

For later reference it is useful to define the l.h.s. of (2) when $\Omega = \omega$ as

$$\mu = \frac{\nu'(\omega)\omega}{\nu(\omega) - \nu_0(\omega, U)}$$

Anything which raises μ for a given level of ω will raise ω and shift the wage curve upwards. Thus, when the reference utility goes up, this raises the wage level.

The reference utility ν_0 has two different interpretations depending on how the Nash bargaining solution is justified in game theory. In one approach ν_0 is the outside opportunity of the workers; i.e. the expected utility if they leave the firm. In a more recent non-cooperative approach, where the parties bargain over a shrinking cake, ν_0 is the utility during a conflict. In either case the value of the alternative is related to the wage level. This is obvious if the alternative is to work somewhere else, or to receive unemployment benefits that are indexed to wages. But the value of time for leisure or household work must also be higher when the real wage is high. At least this should be the case when an eventual unemployment spell is short.

One outside opportunity is to take part in a labour market program. Participants are usually assumed to obtain higher welfare than the openly unemployed. This seems to be confirmed by the facts that program participation is usually voluntary and that there is often a queue to get into programs. The increased welfare could have at least three reasons: 1)higher financial remuneration for program participants; 2)programs give training and experience which raises expected future income from work; 3)participating in programs may be more enjoyable than being idle at home. In particular, program participation may give an income when the right to unemployment benefits has expired or the participant did not qualify for benefits in the first place. Thus, an increased supply

[3] For a given level of Ω, the second-order condition ensures that an increase in U reduces ω at each firm. However, there is a second round effect as lower wages in all firms leads to a lower Ω. For this process to converge at a lower real wage, $\nu_{0\Omega}$ must not be too large. That $\nu_{0\Omega} < \nu_\omega$ is sufficient, but not necessary.

of program places is expected to increase the reference utility and real wages[4].

The positive welfare effect of programs is for a given set of rules for income support for the openly unemployed. In the Nordic context the policy choice has often been between income support with or without the condition that the recipient participates in active programs. A high supply of programs has been a tool for making income support conditional. If the time spent in programs is costly, the reference utility may be lower when income support is conditional.

We may interpret U as the total unemployment rate, including both program participants and openly unemployed. For a given total unemployment rate, it is not just the welfare effect of programs which can shift the wage curve. For the presently employed, the reference utility depends on how effectively they compete with the unemployed for work. Training and experience gained in programs can make the unemployed more effective competitors. The fact that they are occupied and receive income support can make them less effective. In addition there are well known problems with measuring unemployment. A high supply of labour market programs is an incentive to register at job centres also for those who do not qualify for benefits. An increased propensity to register will raise the total unemployment rate without necessarily increasing the competition for jobs.

If we take the view that only 'internal' alternatives matter, it is less obvious that the replacement rate or the program share (or even the unemployment rate) should be in the wage equation. In the Nordic countries, it is not possible to receive unemployment benefits or participate in labour market programmes when one are on strike or in a lock-out. However, outside alternatives may still be brought in by assuming that unions care about the utility gain $\nu(\omega) - \nu_0$ instead of about $\nu(\omega)$.

So far we have concentrated on the union side of the equation. By a simple substitu-

[4] See Calmfors (1994) and Calmfors and Lang (1995) for an overview of effects of labour market programs.

tion we can simplify the producer side. By definition

$$\frac{n(\omega)\omega}{\pi(\omega)} = \frac{S}{1-S} = \frac{\omega/a}{1-\omega/a} \tag{3}$$

where S is the wage share and a is average labour productivity (output divided by employment). Thus the only remaining variables on the firm side are ω and a. In particular the effects of labour market programs on productivity are subsumed in a.

Both productivity and the real wage have historically had a strong positive trend. Unemployment have had no such trend over this century. This conjunction of facts seem to require that

$$\frac{\omega}{a} = g(U) \tag{4}$$

Otherwise deterministic trends in the real wage and in productivity would imply a trend also in unemployment. In order for the wage curve to have this form, μ must obviously be independent of ω. This independence assumption is maintained throughout the analysis. A sufficient condition for it to hold is that ν and ν_0 both have the same constant elasticities with respect to ω.

One consequence of the independence assumption is that if we drive a wedge between producer and consumer real wages, e.g. by taxing labour income or consumption proportionally, this has no effect on the wage curve. If producer and consumer real wages differ by a factor κ, the argument in μ should really be $\kappa\omega$. However, if the independence assumption holds, μ is also independent of κ.

A progressive income tax is different. A high marginal tax rate reduces the gain from a wage increase in the numerator of (2), see Hersoug (1984). However, a progressive income tax can also raise the gain from being employed, the denominator in (2). This is because a more progressive tax system raises the income of those who are employed only part of the year relative to those who are employed the whole year (Rødseth (1999)).

The above model can be extended in various directions without fundamentally changing the nature of the wage curve. If the traded goods sector is a wage leader, as is often assumed in the Nordic countries, adding a non-traded goods sector does not change anything. This is relevant since our data are for manufacturing wages. One can also introduce monopolistic competition in the traded goods industry. The producer side of the bargain can still be represented by the wage share. Union preferences can be extended to include employment. Suppose the union wants to maximise not $\nu - \nu_0$ but $N(\nu - \nu_0)$ (the sum of the utility gains of all workers). Then the equation for the wage curve can be written as

$$\mu + \varepsilon = \frac{1-\beta}{\beta}\frac{\omega/a}{1-\omega/a} \tag{5}$$

where ε is the wage elasticity of employment. The concern for employment fosters wage moderation. How the wage curve is altered depends on the relationship between ε and ω. In general $\varepsilon = -\sigma/(1-\omega/a)$, where σ is the elasticity of substitution between labour and capital. Thus, if the elasticity of substitution is constant, the wage curve has the same form as before (proportional effect of a).

Sometimes labour market programs are set up inside firms in order to prevent lay-offs there. Such programs can reduce the perceived wage elasticity of labour demand. They may then shift the wage curve upwards. However, these programs are usually a small share of active labour market programs.

In the Nordic countries it is common to have bargaining at both the central and the local level. The first and most difficult question to ask about central bargaining is what the preference function looks like. One suggestion is to use the same preferences as for the local union. However, one may object that the reference utility is not the same at the central as at the local level. What is 'outside' to a local union may be 'inside' to a central union, and industrial action may have different costs at the two levels. Since our main interest here is which variables to include in the wage equation and what restrictions to place on their coefficients, not the merits of local versus central bargaining, we shall leave the question at this level. Thus, we assume that the preferences of the central union have the same form as previously assumed for the local union, although the content of the reference utility need not be the same.

Given this type of preferences, a central union should take account of that it affects ν_0 through both Ω and U, and possibly also through tax rates. The effect of labour market programs become more ambiguous, because a fully centralised union will know that its members also pay the costs of the programs, see Rødseth (1995).

The wider setting

In Figure 2 the wage curve is drawn as a function, g, of total employment, N, on the assumption that unemployment is a decreasing function of N. In equilibrium N is determined by the intersection of the wage curve and the aggregate labour demand curve

$$N = n(\omega)K \quad where \;\; n' < 0 \tag{6}$$

as illustrated by point A in Figure 2. (For simplicity we use the symbols as if the number of firms was just one. Multiplying with the number of firms in appropriate places does not change anything). The rate of return on real capital is $\pi(\omega)$. Over time the capital stock will grow or decline depending on whether $\pi(\omega)$ is higher or lower than the rate of return in international financial markets, ρ_*. This means that the long run equilibrium is characterised by

$$\pi(\omega) = \rho_* \tag{7}$$

which alone determines the equilibrium real wage $\bar{\omega}$. The long run labour demand curve is horizontal at $\bar{\omega}$ in Figure 2. The long run equilibrium is where the horizontal labour demand curve interects with the wage curve. The latter is independent of the capital stock, and remains the same in both the short and the long run. Thus, point A in the figure is both a both a short run and a long run equilibrium. In the long run unions and wage bargaining have no influence on the real wage. Instead, the wage equation determines employment, and, thus, implicitly unemployment[5].

[5] The way the long run equilibrium is determined here is different from in Layard *et al.* (1991), where the rate of return requirement is replaced by a requirement of balanced trade. However, in the long run national savings must adapt to the external borrowing constraint somehow. This need not place any restriction on the real wage.

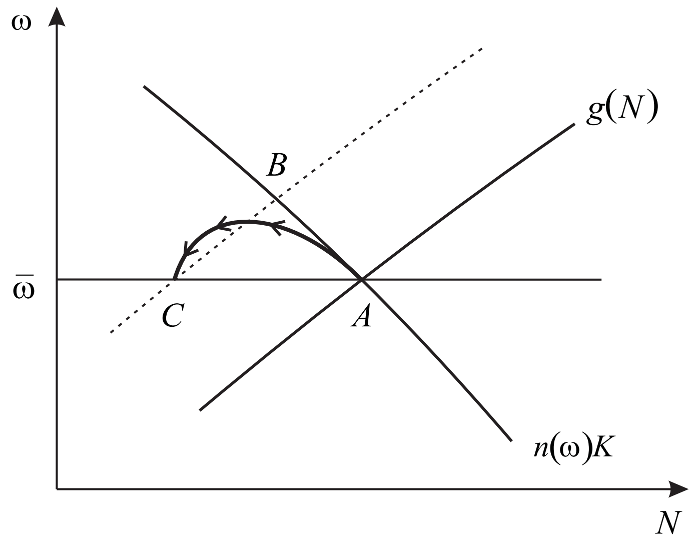

Figure 2: Union wage determination

The dynamics following a shift in the wage curve to the dotted line are illustrated informally in the figure. From the starting point A there is a rapid movement towards the short run equilibrium B, where profits are too low for the capital stock to be sustained. Slow adjustments of the capital stock means that it takes a long time to move from B to C. As the movement goes on, the short run labour demand curve shifts to the left, meaning that the short run equilibrium moves along the wage curve towards C.

In the simplest model there is no room for demand management by the government. However, if we allow for some government employment, which is not determined by profit maximization, or for a non-traded industry, the basic picture is still the same as in Figure 2. Note then that if the government by some form of demand management were to keep total employment constant at a level above the equilibrium A, the real wage would increase, but only until the economy hits the wage curve. It would not increase without limits, and the loss of profitability would be limited. Furthermore, since we are looking at a real model here, the inflation rate at this point could be anything. Nothing says that inflation must accelerate or that there must be a continuous loss of competitiveness even if unemployment is kept below its equilibrium level. Unemployment below the long run equilibrium level cannot be sustained forever. It may be maintained for quite a while, though, if the decline of the capital stock in the traded goods industry is slow and the expansion of government demand is done with a balanced budget.

The wage equation

Above we derived the wage curve from a simple bargaining model. In practice the economy will not be on the wage curve continuously. One reason is that wage contracts last for fixed periods, another may be that the reference utility lags behind. There will be some short run dynamics around the wage curve as illustrated informally in Figure 2. Our estimated wage equation has a fairly standard form:

$$\begin{aligned}\Delta wc_t &= \beta_0 + \beta_1 \Delta cpi_t + \beta_2 \Delta p_t + \beta_3 \Delta pr_t \\ &\quad + \beta_4 (wc_{t-1} - p_{t-1} - pr_{t-1}) + \beta_5 u_{t-1} + \beta_6 z_{t-1} + \beta_7 \Delta z_t + \varepsilon_t \end{aligned} \tag{8}$$

where wc is hourly wage cost, cpi the consumer price index, p the producer price index, pr average labour productivity, u the total unemployment rate, z a vector of other explanatory variables, and ε a stochastic error term. All variables are in logs. The β's are unknown parameters to be estimated, and the subscript t is for time period. Note that $wc - p - pr$ is the same as the log of the wage share. More lags may be included in the equation. The vector z will contain variables such as the replacement rate, active labour market programs, and possibly taxes and other wedge variables. The error term ε is assumed to be independently normally distributed with conditional expectation zero and constant variance.

We get an empirical counterpart of the wage curve by setting all rates of change in (8) equal to zero, which yields

$$wc_t - p_t - pr_t = -\frac{\beta_0}{\beta_4} - \frac{\beta_5}{\beta_4} u_t - \frac{\beta_6}{\beta_4} z_t \tag{9}$$

If $\beta_4 < 0$, (8) is an equilibrium correction model for the wage rate. This means that when all explanatory variables are constant, the wage rate will approach the level determined by the wage curve (9).

Economic theory is less of a guide when it comes to the short-run dynamics (the Δ-terms in equation (8)). Generally one would tend to include different lags of the variables that are arguments in the wage curve. In addition it seems natural to include the changes in so-called wedge variables (the difference between consumer and producer prices and tax rates). Even if the independence assumption holds, these variables may have short run effects.

The wage curve (9) is often seen as the long run equilibrium of the wage equation (8). However, even in the long run prices and productivity may grow. If they do so, the dynamic specification (8) means that the economy may never actually get to the wage curve. Suppose that technical progress is purely labour augmenting, that the international rate of return is constant, and that capital goods are traded. As before, investment responds to the difference between the rate of return on capital and the international rate of return. Then the economy will approach a steady state where the wage share is constant and wage costs grow at the rate

$$\Delta wc = \Delta p + \Delta pr$$

From this and the wage equation, assuming $\Delta z = 0$, we get a long run equilibrium condition

$$\begin{aligned} wc - p - pr \quad &= \quad -\frac{\beta_0}{\beta_4} - \frac{\beta_5}{\beta_4}u - \frac{\beta_6}{\beta_4}z + \frac{1-\beta_1-\beta_2}{\beta_4}\Delta cpi \\ &\quad -\frac{1-\beta_2}{\beta_4}(\Delta cpi - \Delta p - \Delta pr) + \frac{\beta_2-\beta_3}{\beta_4}\Delta pr \qquad (10) \end{aligned}$$

This steady state condition can be simplified considerably by two common assumptions:

- *Dynamic homogeneity*, $\beta_1 + \beta_2 = 1$, which means that the long run equilibrium is independent of the rate of inflation.

- *The scope restriction*, $\beta_2 = \beta_3$. If the rate of return on capital is going to be constant, $\Delta p + \Delta pr$ is the scope for wage increases. If $\beta_2 = \beta_3$, only the scope, not its individual components, matter for wage growth.

Both assumptions are tested in our empirical study. Even if the two assumptions, are granted the term $\Delta cpi - \Delta p - \Delta pr$ remains, and affects the long-run equilibrium. In order to determine the long run equilibrium we need, in addition to (10) and the condition that capital shall get the international rate of return, also an equation determining $\Delta cpi - \Delta p - \Delta pr$. A simple alternative is the Scandinavian model of inflation, which implies that in the long run $\Delta cpi - \Delta p - \Delta pr = -\Delta \bar{pr}$, where $\Delta \bar{pr}$ is the average rate of productivity growth in the economy[6]. Given that $\beta_2 < 1$, this means that fast productivity growth reduces wage pressure and unemployment. A more detailed model might recognise that also trends in the terms of trade and in commodity taxes may affect the equilibrium. That the long run equilibrium depends on productivity growth is an unaknowledged feature of many estimated wage equations.[7].

The wage equation (8) can be related to some older theories of wage formation. An augmented Phillips-curve emerges when $\beta_4 = 0$. The Scandinavian model of inflation assumed that the wage share in the traded goods industries would tend towards a constant in the long run. A strong version of this theory emerges if $\beta_4 < 0$ and $\beta_5 = \beta_6 = 0$. However, even if the condition that $\beta_5 = \beta_6 = 0$ is not met, the wage share in long-run equilibrium is constant as long as ρ_* is constant, as we have seen above.

There is a strong presumption that a permanent increase in the general level of inflation should not change equilibrium unemployment; i.e. that dynamic homogenity should hold. (Long-run price homogeneity is already imposed in (8)). The presumption has been questioned for periods with low inflation (see e.g. Holden (1994)). We do not

[6] The Scandinavian model of inflation assumes that wage growth in the non-traded sector follows that in the traded goods sector, and that the wage share in the non-traded sector is constant. Then

$$\Delta cpi = \alpha \Delta p + (1-\alpha)(\Delta wc + \Delta pr_n)$$

where α is the share of traded goods in the consumer price index, and Δpr_n is productivity growth in the non-traded goods industry. If we use that in the long run equilibrium $\Delta wc = \Delta p + \Delta pr$, then

$$\Delta cpi - \Delta p - \Delta pr = -\alpha \Delta pr - (1-\alpha)\Delta pr_n = -\Delta \bar{pr}$$

[7] If the dynamics around the wage curve are due to expectational errors and misperceptions only, then it seems difficult to defend that $\Delta cpi - \Delta p - \Delta pr$ should have any long run effect. Technically the only way to avoid this and still allow $\Delta cpi - \Delta p - \Delta pr$ to have a short run effect, is to include lagged values of $\Delta cpi - \Delta p - \Delta pr$ in the equation and constrain the coefficients for the different lags to sum to zero. However, it could be that in times of fast productivity growth the reference utility actually lags behind, and that this reduces wage pressures.

impose dynamic homogeneity a priori. Potentially, testing the hypothesis of dynamic homogeneity involves specifying expectations for consumer prices and product prices, as homogeneity might hold for expectations, but fail for realised values. A standard approach is to assume that expected inflation is an unbiased predictor of actual inflation taking account of all available information. This means that expected inflation can be replaced by actual inflation, and the discrepancy between the two subsumed in the random error term. As one have learned from the discussion of 'the Peso-problem' in exchange rate economics, this procedure may be misleading if the sample window is short and there is a small probability of some events with large effects on inflation. Big devaluations, changes in monetary policy regimes, and wage and price freezes may be cases in point. Given the usual sample sizes, there may be large efficiency gains from better modelling of the expectational error, even if expectations are essentially unbiased predictors[8]. However, here we follow the more common procedure of including the contemporaneous inflation rates together with lagged inflation rates. The latter allows, in an ad hoc manner, for expectations which lag behind actual developments, or alternatively for wage contracts which actually depend on lagged inflation as in some indexing and bonus schemes. This procedure may be satisfactory if we are interested in e.g. the effects of labour market policy, but hardly if the interest is in the costs of an abrupt change in the monetary regime.

3 Data and some econometric considerations

Some main features of the data

We have seen that the unemployment history of the Nordic countries shows some marked dissimilarities. Given this, the overall pattern of wage growth was surprisingly similar (Figure 3). Nominal wage growth reached a peak in the mid 1970s in all countries, then fell to the level of the late 1960s. There was a new, lower peak, at the end of the last boom, before it fell to the lowest level of the observation period. Consumer price inflation

[8] Rødseth and Holden (1990) gives an example

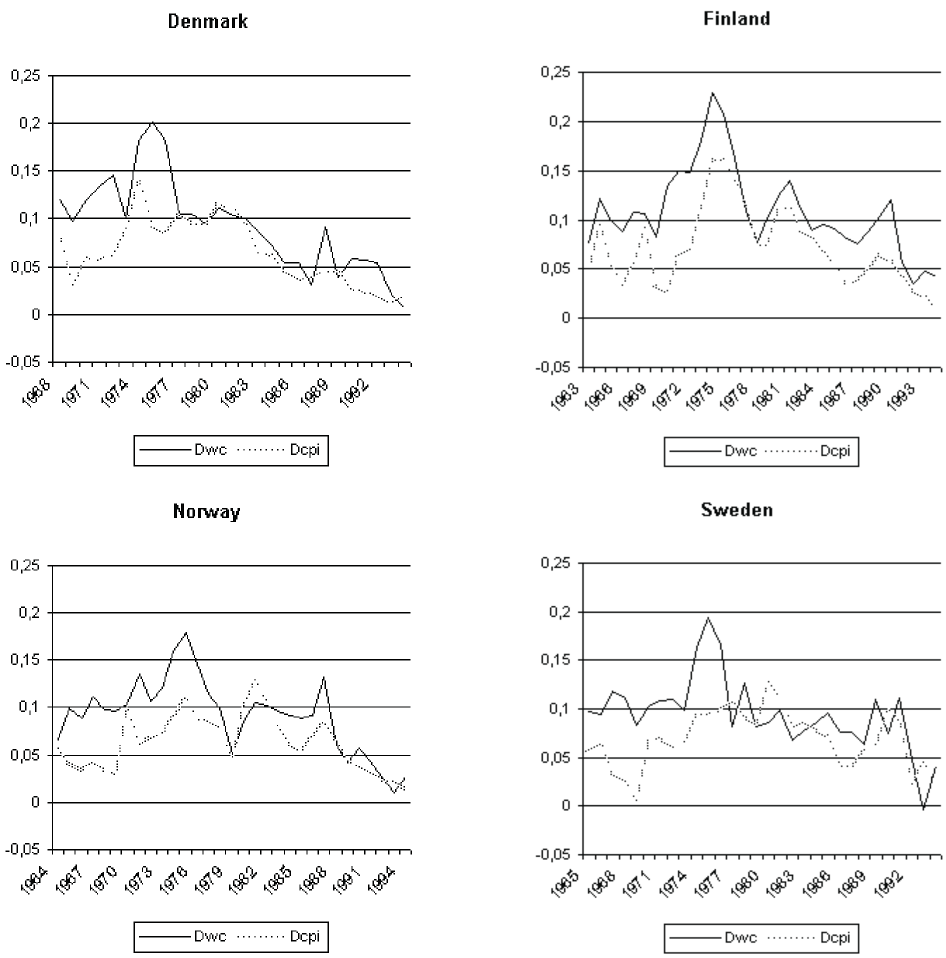

Figure 3: Growth in wage costs and consumer prices

shows the same general pattern, but with an extra peak after OPEC II. Since nominal wage and price inflation have the same peaks and troughs, it will not be surprising to find dynamic homogeneity. However, in no instance has the average nominal wage (net of pay-roll tax) been reduced. The difference between the curves shows that real wage growth was rapid until about 1977. Real wages then stagnated and even fell. Real wage growth resumed during the boom in the mid or late 1980s, in Finland somewhat earlier.

Wage shares show no clear trends, which means that the growth in producer real wages has been roughly in line with productivity (Figure 4). However, there has been relatively long swings in wage shares. In Sweden and Norway the second half of the 1970s stands out with high wage shares. Denmark seems to have experienced a particularly

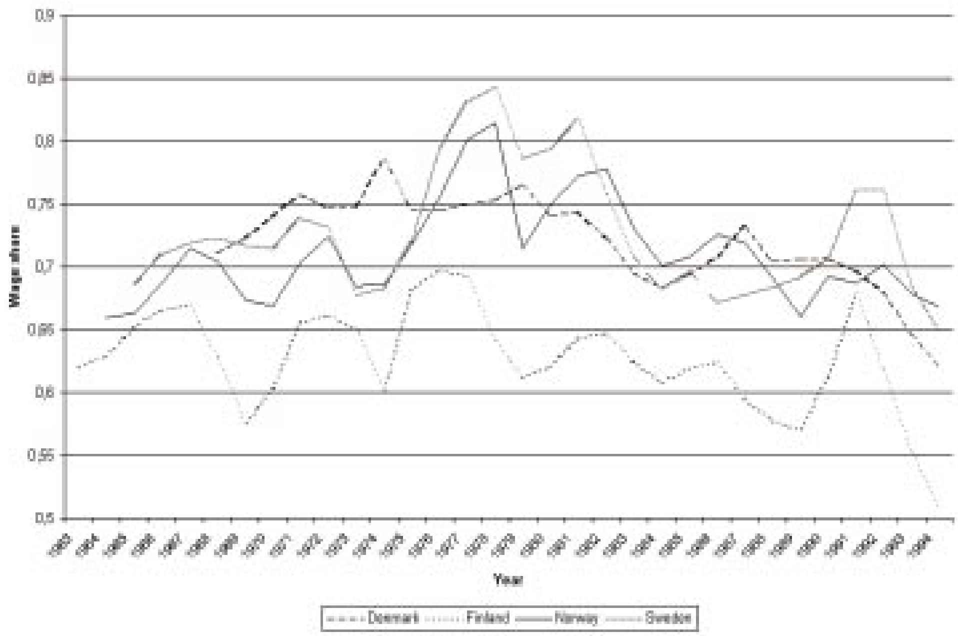

Figure 4: Wage shares in gross product, manufacturing sector

long swing, and the stationarity of the wage share there can be questioned.

Most data are from standard sources[9]. However, the data on unemployment, labour market programs and replacement rates need some comments.

The methods for computing replacement rates for insured workers differ from country to country. The rates were supplied to us by individual researchers. Summary measures of replacement rates and their time paths are strongly influenced by the assumptions made about the length of unemployment spells and about the wage level before unemployment[10].

A special data set with detailed information on the number of participants in labour market programs has been created. The number of participants include those in training and in job creation schemes. Programs aimed specifically at the disabled are not included. The distinctions between job schemes and regular employment and between training

[9] Data sources are documented in a separate report, Evjen and Langset (1997). Compared to Calmfors & Nymoen (1990) there are some data revisions for the overlapping part of the sample, probably due to the use of different editions of the national accounts. The different starting points for the samples are decided by breaks in the national accounts.

[10] OECD has recently provided a new data set on replacement rates which clearly illustrates this. OECD's summary measures seem to have unreasonably high weights on long unemployment spells. The new data arrived too late for the pressent study, and they are bi-annual.

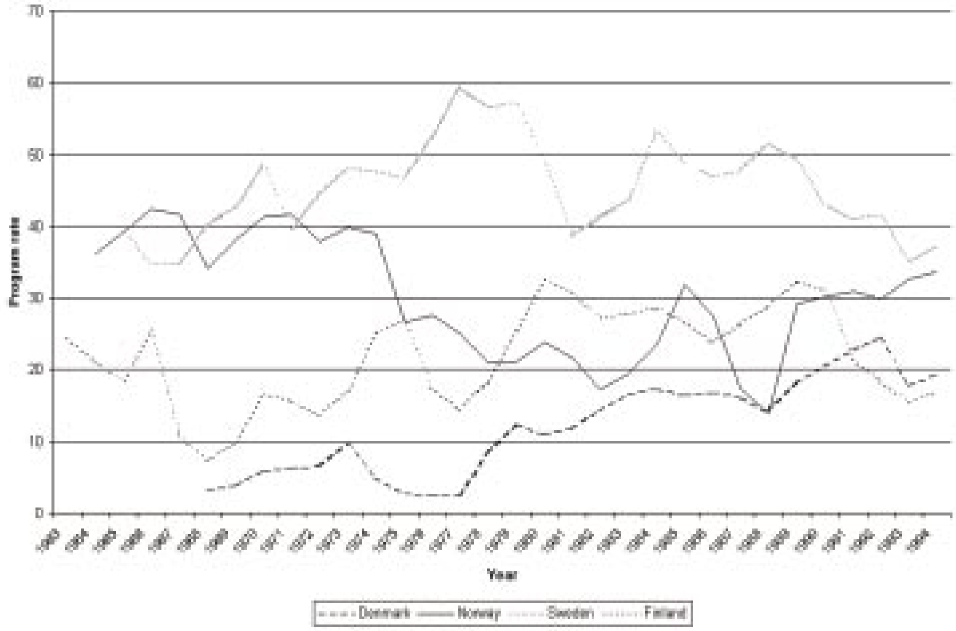

Figure 5: Program shares (numbers in active labour market programs relative to the sum of program participants and openly unemployed)

programs and regular education are difficult both conceptually and in practice. The main criterion which has been used, also in our study, is administrative. The supposition is that programs run by the labour market authorities differ from regular jobs and regular education in content, in selection of applicants and in renumeration of participants. However, borderline cases do arise. Finland in the 1960s labeled as labour market programs a number of public works that would have been classified as regular investment projects in other countries. We chose not to include them. Some training programs in Denmark and Sweden accept participants who have a job and who take a leave. On the whole, we believe that our series are fairly consistent over time and between countries.

The volume of labour market programs has increased with the level of unemployment, but the shares of the unemployed who participated in programs have fluctuated widely (Figure 5). In Norway the share was higher in the 1960s than later. In Denmark a rising share of the unemployed have taken part in labour market programs, but the program share still remains lower there than in Sweden.

It is important to avoid double counting of the unemployed. This has forced us to use national unemployment figures even where standardised unemployment rates were available[11]. In some countries it has also forced us to use the numbers registered at employment offices instead of labour force surveys. We use official figures for the labour force. These are not always consistent in the way they treat participants on labour markets programs. Unfortunately there seems to be no way to correct for this in the earlier part of the sample. Probably the resulting errors are small.

Integration and cointegration

Equation (8) is suitable for discussing and tackling important issues related to integration and exogeneity. Nearly all contemporaneous modelling of wages and prices take the view that stochastic trends are an integral part of the apparent non-stationarity of these series. Hence, they are integrated series and have to be differenced before they can be treated as stationary variables that can be modelled by conventional time series methods. Typically, wages or prices are seen as being integrated of order 1 or 2, which is written $I(1)$ and $I(2)$ respectively. In this paper we assume that all variables are $I(1)$, at most. Hence we assume that any non-stationarity lingering in e.g. wage growth Δwc_t can be modelled by deterministic terms, e.g. dummy variables. Given the $I(1)$ interpretation, the left hand side of the equation is $I(0)$, but on the right-hand side we need a "mapping" from $I(1)$ to $I(0)$ in order to achieve a balanced equation and to stay clear of problems of spurious regression.

If we assume that the rate of unemployment is $I(0)$, it is seen that the Phillips curve model achieves the transformation from $I(1)$ to $I(0)$ by using differenced variables along with only one stationary level variable, i.e. the rate of unemployment itself. Likewise, the strong version Scandinavian model of inflation produces an equation that is balanced as long as the wage share is $I(0)$. One might express the same by saying that both

[11] Since standardised unemployment rates for Denmark became available, they have usually been around two percentage points lower than the national figures, while for the three other countries the national figures have more often been below the standardised rates.

the Phillips curve and the Scandinavian model assumes that there is one a priori known cointegrating vector, respectively the rate of unemployment and the wage share. Of course, including both the wage share and the rate of unemployment causes no problems provided that they both are $I(0)$, hence (8) generalises on earlier specifications.

The need to test for cointegration nevertheless arises in practical situations. First, even though variables in rate form (the wage-share and the rate of unemployment) must conceptually be $I(0)$ (since their variance must be bounded), actual time series of these variables may be highly autocorrelated over the given sample window. In this case, small sample inference becomes more reliable if one treats the variables as $I(1)$ and test whether they cointegrate using the appropriate distribution theory.

A simple way to test for cointegration in single equation models is provided by Kremers, Ericsson and Dolado (1992) In our context, if we assume that $u_t \sim I(0)$, the hypothesis that $(wc_t - p_t - pr_t) \sim I(1)$ can be tested by comparing the $t-$value of $\hat{\beta}_{4,OLS}$ with the usual Dickey Fuller table. More generally, we may assume that both $u_t \sim I(1)$ and $(wc_t - p_t - pr_t) \sim I(1)$. The hypothesis about no-cointegration can then be tested by comparing the $t-$value of $\hat{\beta}_{4,OLS}$ with the appropriate critical value in MacKinnon (1991). This procedure extends to more than two $I(1)$ variables, e.g. contained in the $\beta_6 z_t$ term, and is the test for cointegration reported below. A drawback with this methodology is that it rules out discovery of multiple cointegrating relationships. Another caveat is that the testing procedure is only approximately valid if some of the right-hand side variables are also error-correcting with respect to the estimated cointegrating relationship.

For prices, wages and productivity the Dickey-Fuller statistics reported in Table 7 in the Appendix show a pattern consistent with these variables being $I(1)$; i.e., their differences are stationary. However, the tendency to mean reversion in nominal wage and price inflation has been slow and not statistically significant. The tendency to mean reversion in the unemployment rate and the program share is also extremely slow and usually not statistically significant. This underscores that the period of observation is short relative to the swings we observe in inflation and unemployment. There is some tendency to mean reversion of the wage share, but non-stationarity is rejected only for Finland. However, the real test of whether there is a tendency to equilibrium correction in the wage share comes in the next section.

Identification

As shown in Johansen and Juselius (1994), identification in co-integrated models can be discussed in two steps: first identification of the long-run structure and then identification of the short-run dynamics. At each step standard identification criteria apply.

A formal discussion of identification of the long run relationship requires a complete macro model, and is left for another occasion. All variables in the wage curve (the wage share, the total unemployment rate, the program share and the replacement rate) are potentially endogenous. Assume provisionally, as in most previous discussions, that the replacement rate and the program share are exogenous. The wage curve is then one equation between two endogenous variables. There are more than enough exogenous variables excluded from the wage equation for the necessary condition for identification to be fulfilled. Since we are looking at small open economies, foreign interest rates, foreign prices and foreign output surely are exogenous. Any worry about identification must then be because of another equation, or a linear combination of two or more equations, contain the wage share and the unemployment rate, but none of the exogenous variables that are not in the wage curve. It is hard to see how this can come about when one remembers that the wage share is for manufacturing only, while the unemployment rate is for the whole economy. Thus, the worry about identification of *aggregate* wage equations in a *closed economy* expressed in Manning (1993) seems exaggerated in our context.

Furthermore, one should remember the recursive structure of the long run equilibrium. The wage share in manufacturing is determined by technology, demand elasticities, and the required international rate of return. Then the wage curve determines unemployment. This means that the international rate of return should identify the wage equation.

The more difficult identification problem is related to the possible endogeneity of labour market policy. Identification would then seem to require that we can observe at least one exogenous variable which influences the program share without being an argument in the wage equation. The political colour of the government may be one, but this suffers from a lack of variation (and given the ties between the social democratic parties and the unions, we cannot be sure that the colour of government should not also be an argument in the wage equations).

As for identification of the short run dynamics, a sufficient condition is that deviations from equilibrium in one of the cointegrating equations leads to adjustments in just one endogenous variable. For example, deviations from the wage curve must lead to adjustments in the wage rate only, and disequilibria in other markets must not affect the wage rate directely. This seems to conform well with how we think of price dynamics in economic theory.

Functional form

The functional form chosen for the program variable is different from that in earlier papers. We use

$$lmp = \ln(1 - LMP), \tag{11}$$

where

$$LMP = \frac{\#\text{ on programs}}{\#\text{ on programs } + \#\text{ in open unemployment}}.$$

i.e. the program share. lmp is the log of the share of open unemployment in total unemployment. This is combined in the equation with u which is ln(total rate of unemployment). This allows us to test two hypothesises

a. Only total unemployment matters, which is the hypothesis that the coefficient in front of lmp is zero.

b. Only open unemployment matters, which is the hypothesis that the coefficients in front of lmp and u are equal, since $u + lmp$ is the log of the rate of open unemployment.

Previous studies use functional forms that makes it straightforward to test a or b, but not both. With our functional form the elasticity of wage increases with respect to the program share increases with the program share itself. This seems reasonable if at low levels programs are targeted at special problem groups (e.g. long term unemployed).

General to specific

A popular econometric practice is the general to specific methodology. One starts from a fairly long list of potential explanatory variables and allows a flexible lag structure. Then variables and lags are successively eliminated from the equation by imposing parameter restrictions which are not rejected by the data until one arrives at an equation which is parsimonious in the number of parameters and statistically well specified. Usually the variables or lags with the lowest t-values are eliminated first.

A main issue in the present paper is the degree of similarity among the Nordic countries. If the general to specific procedure is carried out for each country separately, the resulting equations almost certainly exaggerate the differences. For example, a variable which has the same estimated coefficient in the general model may drop out for one country and remain in the final equation for another just because of a small difference in t-values in some intermediate step.

One answer may be to conduct the reduction process under the constraint that the same variables are retained for all countries. However, if all variables and all lags which are significant for one country are kept in the equations, one easily ends up with an overparametrised model. There are reasons to expect some idiosyncrasies, especially in the short run dynamics. Bargaining in the different countries take place at different times of the year and the contract length has varied. Thus, one has to compromise.

When idiosyncrasies are detected, one possible conclusion is that wage formation actually differs. However, since the countries we study have similar institutions, idiosyncrasies may also be taken as an indication that the equation is misspecified or the correlation spurious. The interpretation should depend on whatever supporting evidence there is from other sources, and on a close examination of the features of the data that produced the different results.

4 Empirical results

We started with the same list of variables and lags for all countries. The list included

- Δcpi_t, Δp_t and Δpr_t (growth in consumer and producer prices and in productivity), all with current value and one lag.
- Δpt_t, (pt_t = the log of one plus the rate of pay-roll tax), and Δh_t (h_t = the log of normal working hours per year) without lag.
- Δu_t, u_{t-1}, Δlmp_t and lmp_{t-1} (u and lmp being respectively the *total* unemployment rate and the share of open unemployment in total unemployment, both in logs).
- $s_{t-1} = wc_{t-1} - p_{t-1} - pr_{t-1}$ (the wage share) and rpr_{t-1} (the after tax replacement rate) with one lag.
- Income policy dummies for Norway.

As usual, it is necessary to reduce the number of parameters to get a well specified model. We decided a priori on a fairly long list of variables which should remain in the equations irrespective of the outcome of any tests. The list includes the wage share and the labour market variables which are of particular interest to us (i.e. $(wc-p-pr)_{-1}$, lmp_{-1}, Δlmp, u, Δu, and rpr_{-1}). This means that all levels variables are kept in the equation.

For the remaining variables, which all relate to short-run dynamics, we carried out the following reduction procedure:

- Test for dynamic homogeneity. As this was not rejected at the five per cent level, dynamic homogeneity was assumed[12]

- Test whether Δp and Δpr can be combined in a scope variable $(\Delta p + \Delta pr)$. As this was not rejected for any of the countries, we proceeded on this assumption.

- Test for zero and equality restrictions on the remaining coefficients. Generally restrictions were imposed when they were within a band of plus / minus one standard deviation. For ease of comparison it was eventually decided to keep Δpt in the equation in all cases.

The final results are shown in Table 1 and discussed at length throughout this section[13]. After some general comments on the estimated equations, we shall in turn discuss stability, short run dynamics, real wage flexibility, labour market policy and incomes policy. We show how certain features of the data have created the results, and use this as the basis for a critical evaluation. At the end of the present section we present another set of estimated equations, where we have taken account of some of the criticism and allowed more flexibility in finding the lag structure which suits each country.

On first sight the equations appear to be well specified for all countries. Table 1 reports the usual diagnostic statistics for autoregressive errors, heteroscedasticity and non-normality. None are significant. The standard errors of the regressions range from 1 per cent for Norway to 1.9 per cent for Denmark and Sweden. The implied prediction intervals are on the high side if wage equations are to be useful for predicting inflation.

[12] There is some doubt about dynamic homogeneity for Finland. Originally the sample for Finland was from 1962 to 1994, and dynamic homogeneity passed without problems. However, 1962 was the year of a major social security reform in Finland with an extreme increase in the pay-roll tax. Tests indicated a structural break in 1962. We decided to leave out 1962. This mainly affects the coefficient for the pay-roll tax. Unfortunately when 1962 is left out, dynamic homogeneity is marginally rejected at the 5 per cent level when OLS-estimates are used, marginally not rejcted when IV-estimates are used. OLS was used throughout the reduction procedure for the other countries. The problem with dynamic homogeneity for Finland seems to be related to events in the 1960s, not to the period of high unemployment and low inflation in the 1990s.

[13] OLS-estimates of the same equation are provided in Table 8 in the Appendix. The differences are relatively small. One could argue that we should also have included the change in the average rate of income tax. However, when this variable was added, it was in most cases insignificant and the estimates did not change much. When a measure of tax progressivity was added to the Norwegian equation, it was also insignificant,and the point estimate did not have the expected sign. Tax issues are dealt with in Nymoen & Rødseth (1996) and Rødseth (1999).

Table 1: Wage equations, IV-estimates. Dependent variable Δwc.

Variable	Denmark	Finland[a]	Norway	Sweden
Δcpi^b	0.658 (0.290)**	1.000	0.643 (0.070)***	0.469 (0.172)**
$\Delta scope^c$	0.342 (0.290)		0.357 (0.070)***	0.531 (0.172)***
Δpt	0.274 (0.411)	0.771 (0.481)	0.439 (0.339)	0.985 (0.292)***
Δh	-0.627 (0.516)	-0.751 (0.608)	-0.231 (0.210)	
rpr_{-1}	0.092 (0.085)	0.055 (0.031)*	0.022 (0.030)	0.000 (0.041)
Δu	-0.017 (0.018)	-0.026 (0.024)	-0.025 (0.012)*	0.003 (0.031)
u_{-1}	-0.032 (0.010)***	-0.012 (0.004)***	-0.020 (0.005)***	-0.046 (0.012)***
Δlmp	-0.148 (0.158)	-0.038 (0.080)	0.117 (0.052)*	-0.001 (0.044)
lmp_{-1}	-0.089 (0.108)	-0.035 (0.053)	-0.049 (0.047)	-0.070 (0.054)
s_{-1}	-0.200 (0.167)	-0.169 (0.058)***	-0.180 (0.070)**	-0.284 (0.075)***
$Constant$	-0.060 (0.041)	0.027 (0.029)	-0.055 (0.019)***	-0.191 (0.059)***
$i1979$			-0.044 (0.013)***	
$i1988$			-0.051 (0.013)***	
$i1989$			-0.005 (0.017)	
Sample	1968-94	1963-94	1964-94	1965-94
SER	0.0188	0.0146	0.0098	0.0189
AR1-2	F(2,15) = 0.29	F(2,21) = 3.36*	F(2,16) = 0.05	F(2,19) = 0.10
ARCH 1	F(1,15) = 0.14	F(1,21) = 0.17	F(1,16) = 0.06	F(1,19) = 0.11
Normality	$Chi^2(2) = 1.44$	$Chi^2(2) = 0.15$	$Chi^2(2) = 1.06$	$Chi^2(2) = 0.75$
Validity[d]	$Chi^2(5) = 7.84$		$Chi^2(7) = 11.57$	$Chi^2(6) = 10.56$
t_{ECM}	-1.66	-2.91*	-2.54	-3.92***

Standard deviations in parenthesis. * indicates significant difference from zero at 10 per cent level, ** at 5 per cent level, and *** at 1 per cent level (two-sided tests for coefficients, one-sided tests for diagnostic statistics including t_{ECM}). Additional instruments used were Δwc_{t-1}, Δp_{t-1}, Δpr_{t-1}, Δcpi_{t-1}, Δimp (the log rate of change of the import price index), Δimp_{t-1}, Δat_1 (Δat being the log rate of change of one minus the marginal tax rate, not available for Denmark) and for Norway only Δvat (the change in the log of one plus the rate of VAT).

[a] OLS-estimates, with $\Delta wc - \Delta cpi$ on the l.h.s. since the coefficient for $\Delta scope$ is set equal to zero.

[b] No lag for Denmark and Finland, lagged one year for Norway and Sweden. The coefficient for Δcpi and $\Delta scope$ are constrained to add to one for all countries, and the coefficient for Δcpi is set equal to one for Finland.

[c] Average over two years for Norway. No lags for the others.

[d] Test of instrument validity. Tests the overidentifying restrictions.

Table 2: Estimated long-run elasticities.

	Denmark	Finland	Norway	Sweden
Of the real wage with respect to				
u	-0.157 (0.139)	-0.071 (0.050)	-0.113 (0.040)	-0.163 (0.059)
lmp	-0.444 (0.751)	-0.206 (0.338)	-0.271 (0.333)	-0.248 (0.157)
rpr	0.457 (0.345)	0.327 (0.202)	0.125 (0.176)	0.002 (0.157)
$\Delta\bar{pr}^a$	-3.287 (2.838)	-5.929 (2.038)	-3.569 (1.528)	-1.653 (0.650)
Of unemployment with respect to				
lmp	-2.822 (2.839)	-2.884 (4.611)	-2.407 (2.374)	-1.516 (1.127)
rpr	2.905 (2.678)	4.591 (3.402)	1.109 (1.389)	0.010 (0.885)
$\Delta\bar{pr}^b$	-20.876 (13.75)	-83.209 (25.70)	-31.675 (8.972)	-10.125 (3.190)
$s(= w - p - pr)$	-6.352 (5.621)	-14.033 (6.318)	-8.874 (3.178)	-6.125 (2.231)

Standard deviations in parenthesis.

[a] Semi-elasticity

[b] Semi-elasticity

This is a price we pay for imposing a common structure.

Equilibrium correction is tested by looking at the t-values of the coefficient on the lagged wage share, as explained in Section 3. Equilibrium correction is strongly indicated for Sweden, and at a lower level of significance for Finland. For Norway we are also close to rejecting the null of no equilibrium correction. Given the different behaviour of the wage share in Denmark, it should not surprise that the tendency to equilibrium correction is weaker there[14].

The overall impression from Table 1 is the great similarity between the four countries. Most coefficient estimates are within two standard deviations of each other and have the same sign. Departures from similarity are in the short run dynamics and in the inclusion of income policy dummies for Norway. None of the long-run coefficients reported in Table 2 differ by more than two standard errors.

Stability

Part of the motivation for this paper is to see whether the structure of wage formation in the Nordic countries has changed after 1987. Figure 6 plots the one-step residuals

[14] Other studies which find a tendency to the same type of equilibrium correction are for Finland Saikkonen and Teräsvirta (1985), for Norway Johansen (1995) and Nymoen (1989), for Sweden Forslund and Risager (1994)

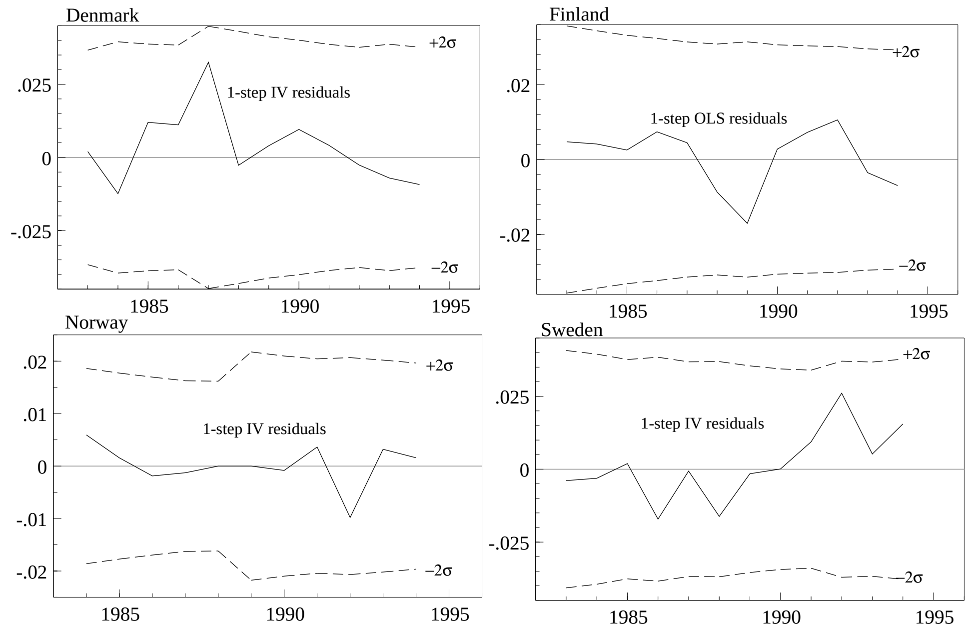

Figure 6: Recursive residuals

($y_t - \hat{\beta}_t' x_t$ in the usual notation) and the corresponding equation standard errors, $\pm 2\sigma$. The plots show no obvious signs of structural breaks. This impression is confirmed by recursive plots of the estimated coefficients (available from the authors). In no cases do the coefficients move outside the confidence intervals estimated on earlier data[15]. There is thus no indication of a structural break after 1987, which was the final year in the sample of Calmfors & Nymoen (1990). As we argue when we come to incomes policy, the 1998-dummy for Norway should not be seen as a structural break.

However, one should note from Figure 6 the relatively large residuals for Denmark in 1987 and Sweden in 1992. To these we shall return. The turbulence in 1987 helps considerably in pinning down several of the estimates for Denmark. The period after 1987 has contributed to narrow some confidence intervals in all countries, especially for the coefficient in front of the total unemployment rate. Still, even after seven turbulent years confidence interals remain large.

If we compare the wage equations estimated here to those in Calmfors & Nymoen

[15] The impression of stability is reinforced by the Hansen tests for parameter stability reported in Table 8 in the Appendix, which in no case indicate parameter instability. Nor does Chow tests indicate any structural breaks after 1987.

(1990), they have a number of properties in common. They display the same type of equilibrium correction, the slopes of the wage curves are not significantly different, and, except for Denmark, the adjustment speeds are roughly the same. The replacement rate came out as significant in the old Finnish equation, and insignificant in the others, as now. Dynamic homogeneity is accepted in all cases. We get somewhat different short run dynamics, but this can largely be explained by a different reduction procedure, by the data revisions for Denmark and Finland, and by the inclusion of the import price index as an extra variable in Calmfors & Nymoen (1990). The different reduction strategies also impinge on the estimated wage curves.

One visible difference with important consequences is that unlike Calmfors & Nymoen (1990) we do not include union density in the Danish equation. The new equation for Denmark is more similar to the equations for the other countries. In particular the estimated adjustment speed is brought down to a much lower level (from 0.66 to 0.20). The reason we now get a well-specified equilibrium correction model for Denmark without the union density variable is that we start the sample later (1968 instead of 1962). This was done because of difficulties in bridging Danish national accounts data around this time.

We have also reestimated the old equations on the new data set with the same observation period as before, and then tested the stability of the resulting equations from 1987 to 1993/94 with traditional Chow tests. There was a serious miss for Sweden in 1992 when the rate of unemployment rose sharply, but otherwise the old equations passes the stability test[16]. It should be admitted though that this stability test was rather weak, as the old equations got considerably higher standard deviations when they were reestimated on the new data.

[16] 1992 is an outlier also with the new equation. The miss was even greater with the old equation because it had a somewhat higher effect of the unemployment rate.

Short run dynamics

Δcpi enters without lag for Denmark and Finland, with a lag of one year for Norway and Sweden. Thus, consumer prices seem to affect wages more quickly in Denmark and Finland than in Norway and Sweden. This does not appear to be an arbitrary result of the reduction procedure. One explanation may be that Denmark and Finland were the only countries which had automatic index clauses for large sections of the estimation period. However, studies on quarterly data have found within year effects of consumer prices also in Norway and Sweden, so we should not exaggerate the difference. Another marked difference is the absence of a scope effect in the short run in Finland. Such effects were included in Calmfors & Nymoen (1990) and Nymoen (1992) with data up to 1987, and this may be a cause for concern. Movements in producer prices and productivity were rather dramatic in Finland after 1987, partly as a result of the breakdown of the Soviet Union. Confidence in the estimated wage equations would have increased if scope effects had remained significant in the Finnish equation.

The change in normal working hours, Δh, was included in the wage equations because of the lumpiness of hours reductions. Actually the lumpiness is most pronounced for Norway, where hours reductions were concentrated on just four years, but the variable seems to be more important for Denmark and Finland. If both leisure and consumption are normal goods, one would expect hours reductions and real wage growth to go hand in hand. There is thus a danger that the coefficients for Δh are spurious and reflect omitted variables rather than a causal effect (in spite of that hours reductions often came through 'exogenous' legislation)[17].

Increases in VAT and in the general sales tax have been subtracted from Δcpi in the Norwegian equation. These tax rates were changed only four times during the sample

[17] An accidental discovery is that the fit of the Swedish equation can be improved substantially if we include the *lagged* change in *weekly* hours. However, since the reductions in weekly hours resulted from legislation, and usually were known well in advance, it is not clear why there should be such a lag. The Danish equation can be similarly improved by allowing two lags in hours, but then the point estimate becomes unbelievably high.

period. By far the largest change was due to a comprehensive tax reform which took place on Janury 1 1970. Since consumer prices affect wages with a lag of one year, a 1971-dummy gives virtually the same result. The 5.6 per cent increase in indirect taxes from 1969 to 1970 was compensated by reductions in income tax. There is independent historical evidence that further compensation for the increase VAT-induced change in consumer prices was not demanded in central wage bargainging, and that expected inflation for 1971, was not directly affected. Without any correction the Norwegian equation appears seriously misspecified[18].

Increases in pay-roll taxes consistently seem to raise wage costs in the short run, although the standard deviations of the estimated coefficients are high. Some of this may be due to a lack of variation in pay-roll taxes, but our impression is also that years with major social security reforms (1962 for Finland, which we deleted, 1967 and 1971 for Norway) create difficulties. The sharp result for Sweden may be due to the fact that the social security system was well established before our sample started. We conjecture that a close attention to years with major social reforms and tax reforms may help to devise wage equations with better fit and better predictive ability.

Real wage flexibility and adjustment speeds.

Real wage flexibility, according to one definition, is the same as the long run elasticity of the real wage with respect to unemployment. The estimates reported in the second line of Table 2 are not far from the proverbial typical value of 0.1 (see Blanchflower and Oswald (1994)). As in earlier studies, Sweden is a bit on the high side, Finland a bit on the low side, and the differences are high enough to be of economic significance.

[18] Interestingly, this shows up as an extreme residual for 1967. Researchers who did not correct Δcpi have resorted to a 1967 dummy. With this alternative solution Δpt drops out (the unconstrained point estimate is slightly negative). These results can be reconciled if one knows that there were major reforms of social security in Norway in 1967 and 1971. These years have the highest increases in pay-roll taxes in our sample. The pay-roll tax was otherwise fairly constant. We can explain the high increases in wage costs in 1967 and 1971 with increases in the pay roll tax, or we can explain the high wage increases in 1967 with a dummy and the high wage increases in 1971 with the increase in indirect taxation the year before. The second explanation actually works better in terms of goodness of fit as the SER is reduced to 0.0057. We still prefer the first explanation because it fits the historical record and accords with the finding of positive effects from pay-roll taxes in other Nordic countries.

Considering the standard deviations, however, it is impossible to claim that real wage flexibility is greater in one country than in another. In Calmfors & Nymoen (1990) the long run elasticities of the real wage with respect to employment were in the same range; -0.13 for Denmark, -0.03 for Finland, -0.17 for Norway and -0.22 for Sweden (calculated at 5 per cent total unemployment).

Layard *et al.* (1991, Ch.9) reported that Finland, Norway and Sweden had particularly high real wage flexibility. This is not confirmed here. Our estimated real wage flexibilities are lower than theirs except for very low unemployment rates (below 1,5 per cent in Sweden, even lower in Finland and Norway)[19]. At total unemployment rates around six per cent the real wage flexibility in Denmark, Norway and Sweden seems to be near the median of the countries studied in Layard *et al.* (1991), and in Finland even lower.

Sweden has the highest coefficient for the total unemployment rate. A look at the recursive plot in Figure 7 shows that the estimate comes to rest just after the large devaluations in 1981-82, but the confidence interval remains high. Then, around 1991 when unemployment increases the point estimate starts to decline and the confidence bounds get narrower. The shift is not statistically significant. However, one suspects that the two episodes with high unemployment in Sweden, the first around 1984 and the second after 1990, are crucial in determining the coefficient for u, and they seem to point in slightly different directions.

In Figure 8 we have drawn the wage curves which are embedded in Table 1. Recall that the wage curve is an equilibrium relationship between the real wage and unemployment, and that we can derive it from the wage equation by setting all rates of change equal to zero. According to (9) the relationship is of the form $wc - p - pr = \alpha - (\beta_u/\beta_s)u$, where α is a constant which depends on the program share and on the replacement rate. Since it is difficult to compare the absolute level, all curves have been drawn through the same point: five per cent unemployment and a real wage equal to 1. We have used levels, not

[19] The comparison is to the coefficient $\bar{\gamma}_1$ in Table 2 on p. 406 of Layard *et al.* (1991). Actually their estimated real wage flexibility for Norway was quite low, as is clear from the appendix to their Chapter 9. In the main text they use an average of their own estimate and one by Aligoskofis and Manning.

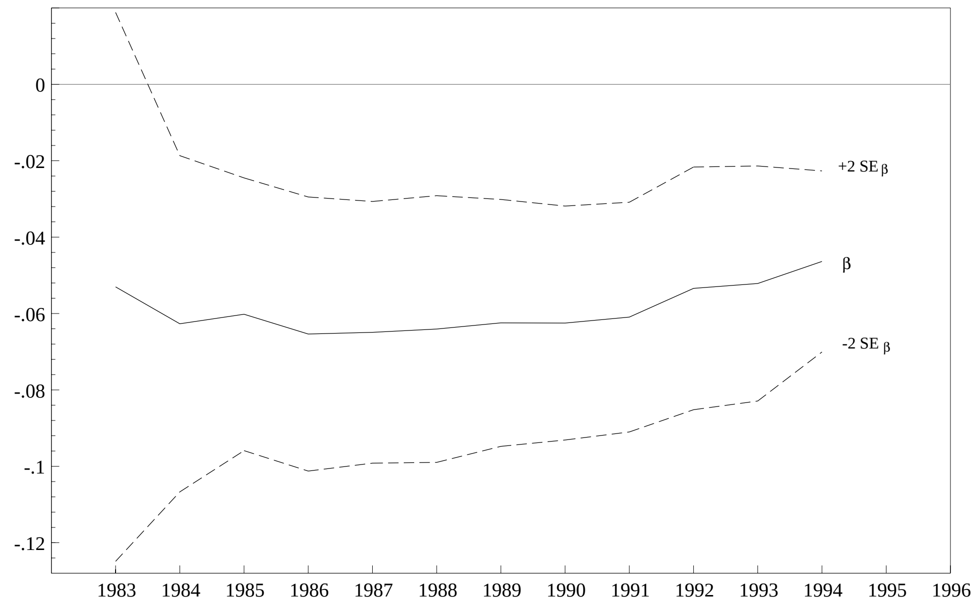

Figure 7: Recursive estimates of the coefficient for u in Sweden

logs on the axises. In Section 2 we emphasised that even if employment was kept below the equilibrium level, the wage curve would limit wage increases. Figure 8 can tell us by how much. We do not know what the equilibrium unemployment rate is, but take an 8 per cent total unemployment rate as an arbitrary example. If the total unemployment rate is instead kept at 4 per cent, the wage curve shows that the resulting long run increase in the real producer wage is in the range from 5 (Finland) to 12 (Sweden) per cent. This appears to be significant in terms of international cost competitiveness. However, it is hardly high enough to invalidate explanations of low unemployment based on aggregate demand. The crucial question is, of course, how quickly an increase in the level of the real wage of e.g. around twelve per cent destroys the traded goods industries.

The coefficient in front of the lagged wage share in Table 1 is the speed of adjustment towards the wage curve. The point estimates are in the range (0.15,0.3), indicating a halving time of two to three years. This says something about the potential for exchange rate policy. Most of the real wage effects of a devaluation will disappear within a period of four to six years (the time it takes for 75 per cent of the effect to disappear if we disregard

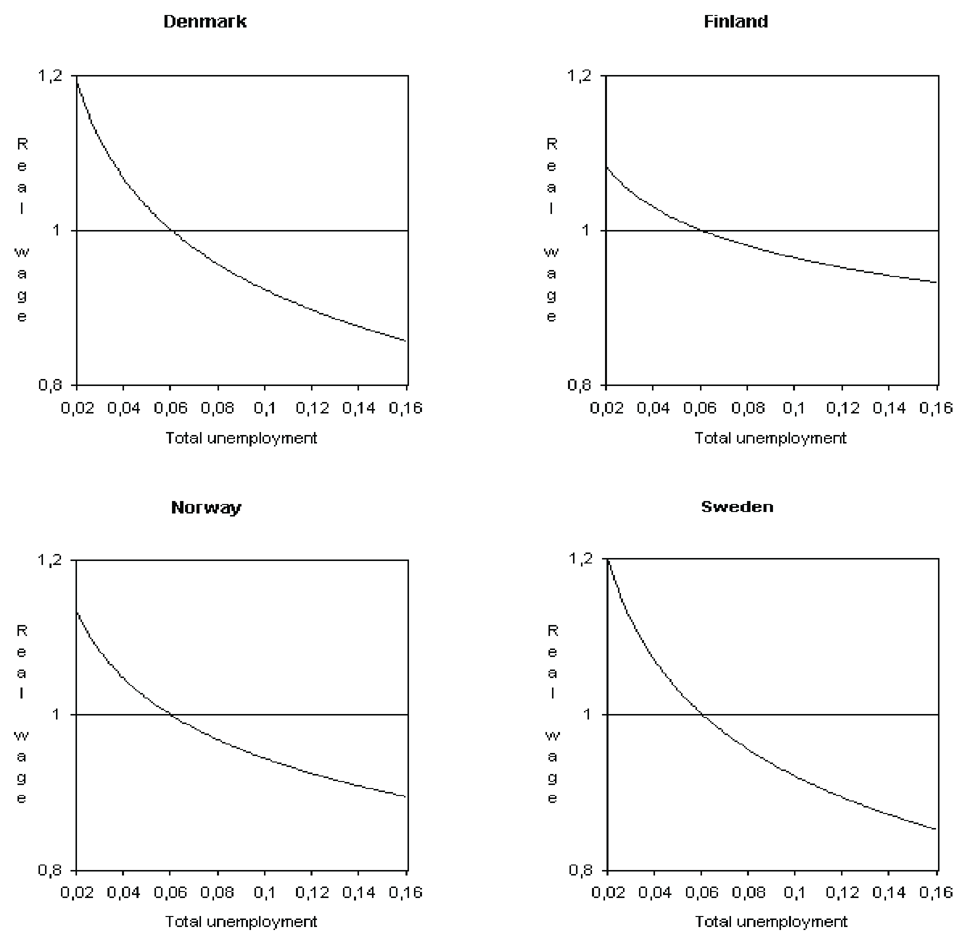

Figure 8: Estimated wage curves

the short run dynamics). However, depending on how the devaluation is transmitted to Δcpi and Δp, the effects may disappear much faster. A thorough investigation of this lies outside the scope of the present study. It would seem that Norway and Sweden, with their slower response of wages to prices, are in a somewhat better position than Denmark and Finland to benefit from devaluations.

The adjustment lags also applies to the effects of low unemployment on the real wage. Thus, if unemployment is kept at four per cent rather than eight per cent, it may take four to six years before the real wage is close to the wage curve. A large devaluation may then give four to six new years. However, if this is repeated too often, the short run dynamics may change.

The question of hysteresis in unemployment has attracted much attention in recent years. From Table 1 there is no doubt that a high level of unemployment reduces wage increases. Thus, there does seem to be a unique equilibrium unemployment rate for each country; i.e. there is not complete hysteresis. However, the change in the unemployment rate is marginally significant for Norway, and has the expected sign also for Denmark and Finland.

Labour market policy

The first column of Table 3 shows t-statistics for the hypothesis that only total unemployment, not its composition, matter for wage formation. The statistics are just the t-values for the coefficients for labour market programs in Table 1. With one exception they are all negative, but not significantly different from zero. The exception is a significantly positive value for Norway in the short run. This would seem to indicate that in the short run labour market programs have a dampening effect on wages in Norway.

The second column presents t-statistics for the hypothesis that the coefficients for lmp and u are equal, meaning that only open unemployment reduces wage pressure. Here a negative sign means that for a given open unemployment rate an increase in program participation, and thus in total unemployment, increases wage pressure. Clearly we

Table 3: T-tests for the effect on wages of numbers in active labour market programs

	Same as unemployed[a]	Same as employed[b]
Short run		
Denmark	-0,93	-0,79
Finland	-0,47	0,12
Norway	2,26	2,42
Sweden	-0,01	-0,07
Long run		
Denmark	-0,82	-0,57
Finland	-0,65	-0,42
Norway	-1,04	-0,61
Sweden	-1,30	-0,46

[a] H_0 is that β_{lmp} respectively β_{lmp} is equal to zero, negative values indicates that numbers in programs lead to more wage pressure than numbers in employment. The degrees of freedom are: D: 17, D: 23, N: 18, S: 21.

[b] H_0 is that $\beta_{lmp} = \beta_u$ respectively that $\beta_{lmp} = \beta_{Dlmp}$, negative values indicates that numbers in programs leads to more wage pressure than numbers in unemployment.

should expect positive signs. However, with two exceptions they are negative. Again only the coefficient for the short run in Norway is significantly positive.

How should one react to these findings? Take the long run first. For the individual countries there are no significant results. However, there is a consistent pattern. When $\beta_{lmp} = 0$, and the error terms are independent between countries, the likelihood of getting four t-values lower than -0.65 (the highest value in the first column) is less than one per cent. Taken together the results thus seem to indicate that, for a given total unemployment rate, a larger share on programs increases wage pressures. In fact all point estimates are higher for β_{lmp} than for β_u. That is what the negative signs in the second column tells. By the same logic we applied to the first column, we are almost forced to the conclusion that workers on programs contribute more to wage pressure than employed workers. However, the degree of significance is lower in this case, as the lower t-values indicate. The similarity in results comes in spite of that lmp has followed very different paths from country to country[20].

[20] Calmfors & Nymoen (1990) found that labour market programs increased wage pressures in Denmark, Finland and Sweden, but not in Norway. The classical study which found wage-increasing effects of labour market programs relative to open unemployment in Sweden is Calmfors and Forslund (1991).

Still we believe that these results should not be regarded as the final words on the long run effects of labour market programs. Our experience is that wide confidence intervals often go together with a lack of robustness to small changes in the specification. Table 6 gives one example where the sign changes for Finland. Micro studies also tend to indicate that labour market programs do not increase wage pressures (Forslund (1994), Edin, Holmlund and Östros (1995), Raaum and Wulfsberg (1997)). Consider the estimated long run elasticities with respect to lmp and u in Table 2. If the size of the labour force were unaffected by programs, they would indicate that one worker on programs "increases wage pressure" by the same amount as 1.5 to 3 employed workers. The welfare effect effect can hardly be of this magnitude. If the elasticity is really so big as the point estimates suggest, the registration effect must be a primary suspect. It is often said that programs are used to 'hide' unemployment, but when there is a registration effect it could be the opposite: Programs make unemployment more visible. The observed total unemployment rate is then a distorted measure of the actual tightness of the labour market. Since the measurement error is correlated with the program share, this can explain the high estimtated elasticities.

Note that the estimated coefficients in the wage equation are not sufficient to draw any conclusions about the effects of labour market programs on *regular employment*. The latter is the combined effect of changes in the total unemployment rate and in total labour supply. If programs increase measured labour supply sufficiently, employment may go up in the long run equilibrium even if the total unemployment rate increases. This is conceivable if the registration effect is strong. Consider also that in the long run the stock of former program participants will be several times as high as the number of participants at a point in time.

The results on the short run effects of labour market programs are less clearcut. Apparently there is an extremely strong and statistically significant negative effect on wage growth in Norway. However, look at Figure 9. The series for Δlmp is dominated

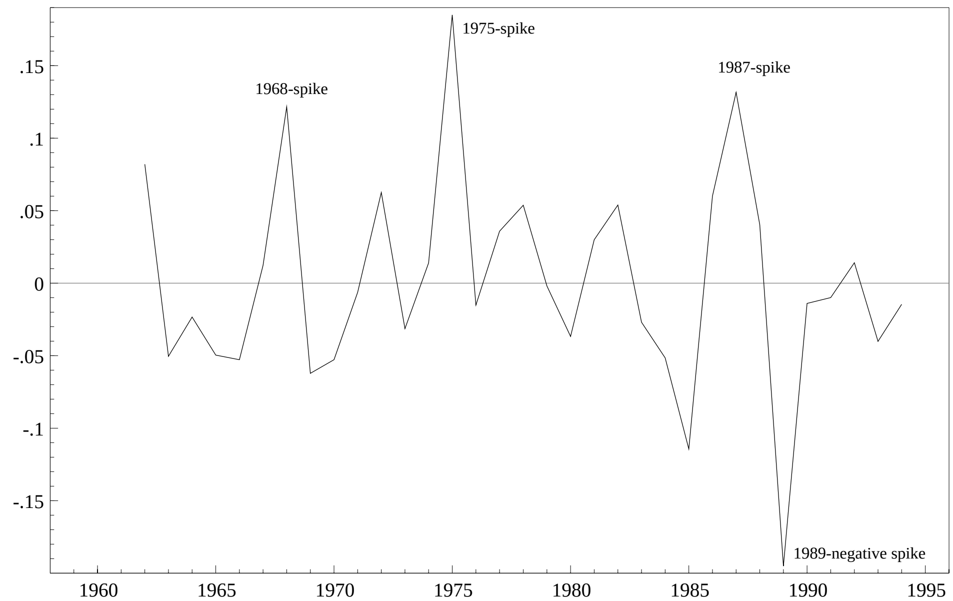

Figure 9: Δlmp for Norway

by three positive spikes and one or two negative spikes. The most markedly negative spike is in 1989, for which we have a dummy in the wage equation, so it does not count. The contrast between the three positive spikes and the remaining observations largely determines the estimate. The three positive spikes, which are years with large reductions in the program share, were in 1968, 1975 and 1987. This is not arbitrary. The economy boomed in 1967, 1974 and 1986. In the following years there were sharp upturns in unemployment. At the same time there were fairly high wage increases, especially in 1975 (the highest in the sample) and in 1987 (the highest in the later half of the sample). The reason for the positive spikes in Δlmp is obvious. Because the change in the cycle was sudden and rapid, the authorities did not manage to step up labour market programs at the same pace as unemployment rose. However, the positive Δlmp can hardly explain the rapid wage increases in the same year. We know that almost all wage increases for manual workers in 1987 were the result of a contract concluded in May 1986.

An alternative hypothesis is that we underestimate the pressure for wage increases at the height of booms. Our wage equation may not be sufficiently non-linear in unemployment. However, Nymoen & Rødseth (1996) shows that the effect does not disappear even if we include an inverse quadratic of the unemployment rate. The series for u in Figure 1 gives a hint. At each successive boom the total unemployment rate has bottomed out at a higher level. If we use vacancy rates as an alternative measure of the state of the labour market, the three years seem less different. The unemployment rate may have been a poor indicator of labour market pressures at the height of strong booms.

For the other countries we also see some of the same tendency to spikes in Δlmp just after the end of booms, weakest perhaps in Sweden. The connection to high wage increases is not so clear, however. One reason may be that some of the turning points are at higher levels of unemployment.

The coefficients for rpr_{-1} in Table 1 are positive for all countries, although only just so for Sweden[21]. It is marginally significant for Finland. Including the contemporaneous instead of the lagged value for Finland gives an even larger and strongly significant coefficient, as in Calmfors & Nymoen (1990). If labour market programs tend to raise wages through welfare effects, one should expect that high replacement rates would also increase wages. The pattern is consistent with this, but the evidence is not strong.

A closer look at what is behind the Finnish result is revealing. In Figure 10 the two curves for real wage increases and the replacement rate follow each other quite closely until 1990. Especially important is the peak in rpr around 1970, which was accompanied by a peak in real wage increases. No definite conclusion can be drawn from this graph. The replacement rate comes into the equation not because a high replacement rate explains the period with high unemployment, but because it explains high wage increases at a time when unemployment was low. An obvious alternative explanation of the high wage increases is the low level of unemployment. The estimated

[21] In the Nordic countries unemployment benfits are more or less proportional to wages up to a ceiling. For Sweden, unlike in the other countires, the series we have got for rpr is based on the ratio of the ceiling; i.e. it is equal to the ratio between the benefits at the ceiling and the average wage in manufacturing. The presumption behind the calculation is that the ceiling to a large extent has been effective. Because the adjustment of the ceiling has often lagged in periods with rapid wage increases, there has been a tendency to a negative correlation between rpr and wage increases which is absent in the other countries. This may explain the different result for Sweden.

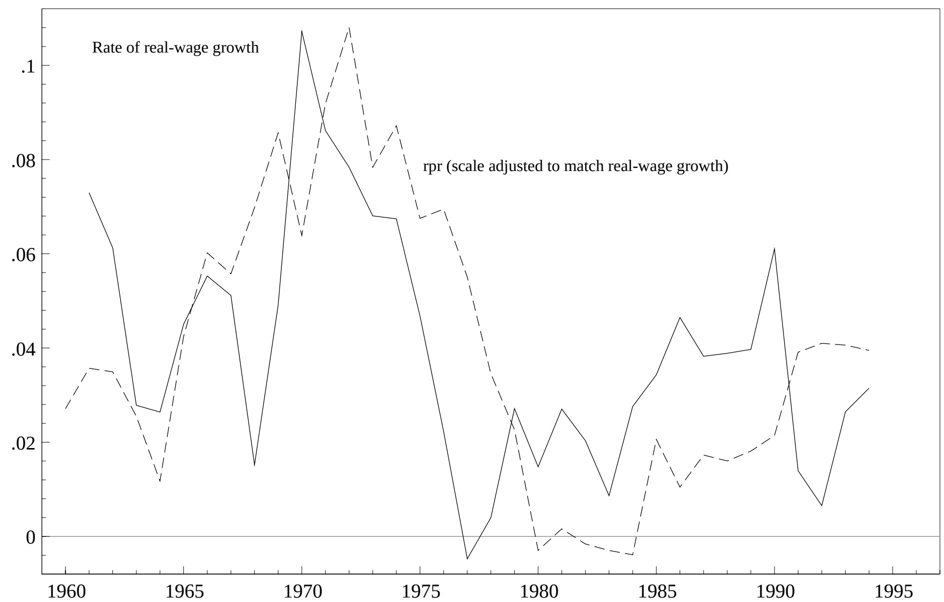

Figure 10: Δwc and rpr for Finland

effect of unemployment in Finland is unusually low. If we shorten the sample by dropping the early years, the coefficient for rpr is reduced and becomes statistically insignificant, while the coefficient for u increases towards the same level as in Norway and Denmark. This happens in spite of the high correlation depicted in Figure 10. Finland could be even more similar to the other Nordic countries than it first appears.

Incomes policy

Incomes policy has a long tradition in the Nordic countries (see Calmfors (1990)). In all of them some form of incomes policy was also a response to the macroeconomic events of the late 1980s and early 1990s.

Incomes policy has been pursued most vigorously in Norway, where special laws have been used to forbid wage increases (from September 1978 to December 1979) and to extend to the whole labour market wage increases agreed on by the main union and the main employers confederation (1988 and 1989). In Table 1 the 1979 and 1988 dummies are highly significant. The 1989 dummy has been found significant in a number of other studies[22]. Its value here is reduced because of the inclusion of Δlmp_{-1}. Other studies also find significant negative effects of the 1978-79 wage freeze on wage increases in 1980. When we do not, it is because the low inflation in 1979 is sufficient to explain the relatively low wage increases from 1979 to 1980. However, everyone knew that past inflation was a poor indicator of expected inflation in 1980. Thus, an alternative interpretation where incomes policy worked in 1980 is possible.

The success of incomes policy in Norway can be seen as a particular form of real wage flexibility. The estimated long run elasticities in Table 2 then underrate the overall wage flexibility. The effects of incomes policy fades out with a halving time of about three years.

In the other countries there has not been the same strong compulsion. There is also less agreement that income policy had any effect. In Sweden a three year wage contract conducted at the central level, "Rehnberg-avtalet" after the name of the mediator, was supposed to keep wage increases moderate through 1991-93. However, when we include dummies for these years in the equation from Table 1, the coefficients we get are (standard deviations in parenthesis): 1991: 0.008 (0.026), 1992: 0.039 (0.026), 1993: 0.011 (0.031). None are statistically significant, nor is a dummy equal to one for all three years significant. If anything, the Rehnberg-agreement seems to have raised wage increases. One should not jump to strong conclusions though. The positive dummies for 1991-93 are directly related to the relatively high estimate of the short-run effect of u in Sweden which is behind the tendency to underprediction of wage increases in the early 1990s[23].

In 1983 the Swedish Employers Confederation stopped bargaining with its counterpart LO at the central level. Bargaining at the industry level continued, though. Unlike Forslund & Risager (1994) we find that this may have affected wages. A dummy for 1983-90 added to the equation in Table 1 yields a coefficient of -0.027 (0.011), significant

[22] See e.g. Bowitz and Cappelen (1996) and Johansen (1995)

[23] We have also looked at earlier episodes of incomes policy in Sweden without finding significant effects. These include 1983-85 (the post devaluation period, coefficient -0.008 (0.013)), and 1974-76 (coefficient 0.025 (0.017).

at the 5 per cent level. The period with more decentralised bargaining was interrupted by the Rehnberg-agreement, but continued from 1994 onwards. A dummy for 1983-90 and 1994 gives a coefficient of -0.026 (0.013), still significant. This seems to indicate that decentralisation actually lead to more moderate wage settlements. Again one should not jump to conclusions. The period that really stands out as exceptional in the Swedish experience is 1975-81 with its consistently high wage share. A dummy for these years is highly significant (coefficient 0.045 (0.017)) and knocks out the 1983-90 dummy completely. The period after 1982 then does not appear to deviate from the period with centralised bargaining before 1975. In 1975-81 Sweden used massive subsidies and within-firm labour market programs to prevent lay-offs. If unions care about employment, this policy of guaranteeing jobs at the firm level may have greatly increased wage pressures.

Denmark attempted incomes policy when the country was still in a boom in 1987. The result was not as intended, as one can already see from Figure 6. A 1987-dummy in the equation from Table 1 yields a coefficient of 0.053 (0.025)[24].

Shifts in equilibrium unemployment

Changes in the level of unemployment that persist over time are nowadays usually explained as changes in equilibrium unemployment. Unfortunately we cannot calculate equilibrium unemployment since we do not know the equilibrium labour share. However, we can calculate how much the equilibrium wage curve has shifted over time. In our model such shifts will be due to changes in the replacement rate, in the supply of labour market programs or in the average rate of productivity growth[25]. The latter is used as

[24] Following Forslund & Risager (1994) we have also tried a dummy for 1983-85 (representing a shift in macro policy). Unlike their finding, this comes out small and statistically insignificant (Coefficient -0.011 (0.014)).

[25] From (10), and given the assumptions made, the equation for the wage curve is

$$wc - p - pr = -\frac{\beta_0}{\beta_s} - \frac{\beta_u}{\beta_s}u - \frac{\beta_{\Delta cpi}}{\beta_s}(\Delta cpi - \Delta p - \Delta pr) - \frac{\beta_{lmp}}{\beta_s}lmp - \frac{\beta_{rpr}}{\beta_s}rpr$$

while equilibrium unemployment is

$$u = -\frac{\beta_0}{\beta_u} - \frac{\beta_s}{\beta_u}s_* - \frac{\beta_{\Delta cpi}}{\beta_u}(\Delta cpi - \Delta p - \Delta pr) - \frac{\beta_{lmp}}{\beta_u}lmp - \frac{\beta_{rpr}}{\beta_u}rpr$$

where s_* is the level of $wc - p - pr$ which corresponds to the reguired rate of return on captial.

Table 4: Shift in wage curve since 1965-73. Per cent.

	In direction of ω	Due to[a] lmp	rpr	$\Delta\bar{pr}$
Denmark				
1974-80	14.5	0.6	7.9	6.0
1981-87	16.8	5.6	5.5	5.7
1988-94	12.6	7.8	2.3	2.6
Finland				
1974-80	18.6	2.4	-4.4	20.6
1981-87	8.0	4.1	-11.2	15.1
1988-94	9.0	3.4	-7.1	12.6
Norway				
1974-80	0.6	-4.3	2.9	1.9
1981-87	1.5	-6.2	3.3	4.4
1988-94	3.3	-5.2	3.8	4.6
Sweden				
1974-80	9.6	5.6	0.0	3.9
1981-87	3.9	2.0	0.0	1.9
1988-94	2.2	0.6	0.0	1.6

[a] Percentage shift in the wage curve as caused by the change in the average value of each explanatory variable. Measured in the direction of ω.

an approximation to the long run trend in $\Delta cpi - \Delta p - \Delta pr$[26]. Table 4 shows the shifts in the wage curve between major periods. Note that it is the average positions of the wage curve (i.e. given the period averages of *lmp*, *rpr* etc.) which are computed. The shift is measured in the vertical direction in per cent of the real wage. Because of the chosen functional form, the size of the shift is the same everywhere along the curve.

Since the base period, the estimated wage curve has shifted upwards everywhere, but more in Denmark and Finland than in Norway and Sweden. Everywhere reduced productivity growth gave a main contribution to the positive shift[27]. Reduced productivity growth had a particularly strong effect in Finland because of the absence of a scope effect in the short-run dynamics there. However, reductions in *rpr* from the late 1970s served to dampen the shift. In Denmark the shift was reinforced by increased 'generosity' towards the unemployed, first in the form of better benefits, then in the form of labour market programs. Norway also raised the replacement rate, but, more importantly, reduced the

[26] The period averages of $\Delta cpi - \Delta p - \Delta pr$ cannot be used as they are influenced by the deviation between actual and equilibrium wage increases.

[27] In contrast, Bean (1994) in his survey reports that "one invariably finds the effect of the productivity slowdown to be unimportant".

program share. In Sweden a high program share reinforced the upward shift in the wage curve in the late 1970s, but the shift there has subsided since then.

An important caveat is that the periods we look at may be too short to discern long-run changes in the rate of productivity growth. After OPEC I The fall in the rate of productivity growth was stronger in Denmark and Finland than in Norway and Sweden. (The opposite was the case for manufacturing, though). The poor productivity performance may have been a result of, rather than a cause for, the greater increase in unemployment rates.

Table 5 shows a tentative calculation of equilibrium unemployment rates. It is based on a rather bold assumption, namely that the equilibrium wage share has always been equal to the average wage share over the observation period[28] . Still it gives a picture of how, according to our estimates, $\Delta\bar{pr}$, rpr and lmp have moved the equilibrium unemployment rate over time. Indeed we find that the equilibrium unemployment rates have increased more after 1974 in Denmark and Finland than in Norway and Sweden. However, except for an odd case in Finland, actual unemployment varies much more than estimated equilibrium unemployment. The large effects in Finland are related to the relatively low coefficient for u_{t-1} in the Finnish equation, and the statistical uncertainty is great. In Norway and Sweden shifts in equilibrium unemployment seem to contribute almost nothing to the higher unemployment towards the end of the period. Systematic trends in international terms of trade or in indirect taxes may also have made $\Delta\bar{pr}$ a poor indicator of the trend in the rate of change of the wedge.

We find similar levels of equilibrium unemployment in all four countries before 1974, in spite of substantial differences in productivity growth, compensation rates and labour market programs.

The elasticity of unemployment with respect to the wage rate along the wage curve,

[28] Over the period 1966-94 manufacturing employment declined by 34-35 percent in Denmark, Finland and Norway, by 39 percent in Sweden. This would seem to indicate that on average profitability was at a similar level in all four countries, although perhaps somewhat lower in Sweden. If that was the case, the estimated equilibrium unemployment rates are too low for Sweden.

Table 5: Actual and equilibrium unemployment rates.

	Denmark		Finland		Norway		Sweden	
	Actual	Equil.	Actual	Equil.	Actual	Equil.	Actual	Equil.
1965-73	1.4	2.6	2.3	2.0	1.4	2.3	2.5	2.6
1974-80	5.8	6.1	5.1	21.7	1.3	2.4	3.3	4.6
1981-87	10.8	7.0	7.5	5.9	2.7	2.6	4.3	3.3
1988-94	12.8	5.5	11.5	6.7	4.9	3.1	4.4	3.0
1990-94			14.5	10.6			4.9	2.7

Geometric averages. Equilibrium rates are calculated based the coefficient estimates in Table 1.

reported in the lower half of Table 2, shows the potential of high real interest rates to increase equilibrium unemployment. In the long run the elasticity of the wage share with respect to the real interest rate is $(1-\alpha)(1-\sigma)$ where α is the wage share and σ is the elasticity of substitution. If $\alpha = 0.7$ and $\sigma = 0.5$, the elasticity of the wage share with respect to the real interest rate is 0.15. This should be multiplied with the elasticities of unemployment with respect to the wage share from Table 2 to get the effects on equilibrium unemployment. For Norway this gives an elasticity of 1.3. If the required rate of return went up from say 6 to 8 per cent (an increase of one third) this would mean that the equilibrium unemployment rate increased by close to one half (somewhat less for Denmark and Sweden, somewhat more for Finland). This example is no proof, but it shows that high international real interest rates in the 1980s may have raised equilibrium unemployment markedly.

Modified wage equations

Some modified wage equations are shown in Table 6. They are an attempt to address some of the problems pointed out in our discussion of the results in Table 1. Another purpose is to give an impression of the robustness of the results. Δlmp and Δh are dropped (except Δh for Norway) because of their possible endogeneity. For Sweden Δu has been dropped because of lack of significance, and a dummy for the period with more decentralised bargaining has been inserted. Based on earlier experience we have also allowed longer lags for the wage share in Denmark and Finland. It helps considerably in reducing the standard error for Finland.

Table 6: Alternative wage equations, IV estimates. Dependent variable Δwc.

Variable	Denmark	Finland[a]	Norway	Sweden
Δcpi	0.866 (0.371)**	1.000	0.667 (0.077)***	0.582 (0.120)***
$\Delta scope$	0.134 (0.371)		0.333 (0.077)***	0.418 (0.120)***
Δpt	0.568 (0.459)	0.514 (0.381)	0.414 (0.376)	0.741 (0.293)**
Δh			-0.500 (0.191)**	
rpr_{-1}	0.057 (0.104)	0.059 (0.027)**	0.066 (0.026)**	0.019 (0.036)
Δu	-0.021 (0.020)	-0.054 (0.013)***	-0.012 (0.012)	
u_{-1}	-0.034 (0.012)**	-0.009 (0.003)**	-0.023 (0.005)***	-0.033 (0.013)**
lmp_{-1}	-0.026 (0.127)	0.039 (0.043)	-0.123 (0.038)***	-0.060 (0.030)*
s_{-1}			-0.136 (0.075)*	-0.360 (0.077)***
Δs_{-1}	0.076 (0.230)	-0.073 (0.060)		
s_{-2}	-0.275 (0.241)	-0.259 (0.060)***		
Constant	-0.060 (0.056)	0.070 (0.023)***	-0.080 (0.017)***	-0.119 (0.059)*
$i1979$			-0.048 (0.014)***	
$i1988$			-0.043 (0.014)***	
$i1989$			-0.024 (0.017)	
$iDecent$				-0.022 (0.015)
Sample	1968-94	1963-94	1964-94	1965-94
SER	0.0221	0.0128	0.0109	0.0166
AR1-2	F(2,16) = 0.17	F(2,22) = 1.56	F(2,17) = 2.02	F(2,20) = 0.41
ARCH 1	F(1,16) = 0.03	F(1,22) = 0.00	F(1,17) = 0.05	F(1,20) = 0.96
Normality	$Chi^2(2) = 0.03$	$Chi^2(2) = 2.98$	$Chi^2(2) = 1.54$	$Chi^2(2) = 0.04$
Validity	$Chi^2(4) = 0.46$		$Chi^2(7) = 15.27**$	$Chi^2(6) = 10.87$
t_{ECM}	-1.73	-4.33***	-1.80	-5.33***

See legend and footnotes to Table 1

[a] Ordinary least squares.

In general the results are not much changed from Table 1. The most marked change is the increased statistical significance of rpr_{-1} and lmp_{-1} in some cases. Note, however, the sign reversal for lmp_{-1} for Finland. Thus, the overall conclusion about the effect of labour market programs is not necessarily strengthened. For Finland Δu now comes out as highly significant, as in several earlier studies, while the same variable loses its significance for Norway. The 1989-dummy for Norway has a negative coefficient (- 2.5 per cent) indicating that incomes policy worked also this year, although the coefficient is not significant. For Sweden the point estimate of the coefficient for u_{-1} is reduced, that for s_{-1} increased in absolute value. This makes the point estimate of Sweden's real wage flexibility 0.09, and gives further support to the view that Sweden does not necessarily have higher real wage flexibility than other countries. The dummy for more decentralization in Sweden after 1982 is no longer significant.

5 Conclusions

The main findings are: 1)There has been no shift in Nordic wage formation since 1987. 2) The wage equations for Denmark, Finland, Norway and Sweden are fairly similar. Except for some differences in the short run dynamics, practically all coefficients are within two standard deviations of another. 3) Real wage flexibility is not particularly high in the Nordic countries. 4)No strong conclusions can be drawn about the effects of active labour market programs, but the weight of the evidence is in favour of that an increase in the supply of such programs increases the equilibrium total unemployment rate. In fact it is impossible to reject that the equilibrium rate of open unemployment is independent of the supply of labour market programs. This would mean that in the long run one cannot 'hide' unemployment by an active labour market policy. Any cost benefit analysis of such programs must take account of labour supply and productivity effects which are not estimated.

According to our very crude estimates all equilibrium unemployment rates were between two and three percent before 1974. Since then they have increased everywhere, but the increase in Norway and Sweden was slight. However, if we compare Denmark to Norway and Sweden, the actual unemployment rates have differed much more than the equilibrium rates. Within each of the three countries the actual unemployment rates have also increased more than the equilibrium rates. The decline in productivity growth is a common factor behind the increased equilibrium unemployment rate. In Denmark this was supplemented by a more generous treatment of the unemployed. The estimated equilibrium unemployment rate for Finland varies so much that it is hard to give much credence to them. In fact one contribution of the present paper is to show how difficult it is to give good estimates of how the equilibrium unemployment rate evolves over time.

The relatively good employment performance of Norway and Sweden until the late 1980s is partly explained by a low increase in equilibrium unemployment rates. However, since the actual unemployment rate was low compared to the equilibrium unemployment rate there is also some room for explanations based on aggregate demand factors, devaluations and - in the case of Norway - wage laws. There is a wage curve, not a unique equilibrium unemployment rate in the short run. As explained, it may then be possible to stay away from the long run equilibrium for considerable periods of time, and the estimates indicate that this has indeed happened to different degrees in different countries. The speed of equilibrium correction when there are deviations from the wage curve is slow enough to allow devaluations to have effects for several years.

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Appendix

Table 7: Means, standard deviations and Dickey-Fuller-tests (t_{D-F}).

	Denmark			Finland			Norway			Sweden		
	Means	St.dev.	t_{D-F}	Means	St.dev.	t_{D-F}	Means	St.dev.	t_{D-F}	Means	St.dev.	t_{D-F}
wc	1.1638	0.7905	0.25694	1.0924	1.1547	-1.0321	0.9411	0.9492	0.13172	0.9366	0.9036	-0.15623
Δwc	0.0918	0.0488	-3.7158*	0.1089	0.0433	-2.0166	0.0907	0.0386	-2.8042	0.0969	0.0362	-3.3605
cpi	0.8574	0.6125	-0.45963	0.7928	0.7889	-1.3797	0.6868	0.6849	-1.2338	0.7367	0.6815	1.8727
Δcpi	0.0658	0.0339	2.0439	0.0704	0.0386	-1.9246	0.0624	0.0297	-2.0569	0.0686	0.0295	-2.0745
p	0.7662	0.5566	0.91618	0.7230	0.7288	-0.42250	0.6538	0.6519	-1.3570	0.6702	0.6132	-1.2610
Δp	0.0611	0.0285	-3.5288	0.0593	0.0565	-2.9223	0.0584	0.0376	-3.4559	0.0580	0.0344	-2.5968
pr	0.4250	0.2733	-1.5990	0.3324	0.4687	1.2353	0.2291	0.2949	-2.1057	0.2493	0.2980	-2.6895
Δpr	0.0367	0.0358	-5.3561**	0.0541	0.0331	-2.9953	0.0329	0.0262	-6.4205**	0.0397	0.0312	-3.2212
Δpt	-0.0001	0.0103	-4.8718**	0.0057	0.0104	-6.6171**	0.0021	0.0075	-4.9413**	0.0084	0.0140	-4.3646**
Δh	-0.0090	0.0122	-5.1339**	-0.0058	0.0064	-2.4730	-0.0060	0.0117	-4.8537**	-0.0046	0.0096	-6.7855**
u	-2.8219	0.8746	-15657	-3.0882	0.8652	-4.1143*	-3.8081	0.6569	-2.9831	-3.3715	0.4138	-4.6417
Δu	0.0944	0.2656	-5.1980**	0.1117	0.2265	-5.3657**	0.0534	0.2385	-5.2192**	0.0598	0.2256	-4.8455**
lmp	-0.1287	0.0769	-2.4842	-0.2592	0.0950	-2.5202	-0.3678	0.1177	-1.7415	-0.6078	0.1305	-1.9260
Δlmp	-0.0062	0.0301	-4.8186**	0.0090	0.0691	-4.7808**	-0.0011	0.0714	-5.0398**	-0.0029	0.0968	-5.4599**
s	-0.0274	0.0511	-2.3095	0.0370	0.0670	−4.0517*	0.0583	0.0548	-2.2347	0.0170	0.0660	-2.8049
rpr	-0.3905	0.0659	-2.9512	-0.6120	0.1517	-1.6165	-0.4000	0.1441	-1.3327	-0.2701	0.1268	-4.5531

Notes: Samples: Denmark 1969-1994, Finland 1963-1994, Norway 1965-1994, Sweden 1966-1994. For Sweden, rpr is assumed contant from 1993 to 1994. Constant and trend included. One lag is included if the t-value is significant at the 5% level. The t_{D-F} critical value is -3.6.

Table 8: Wage equations, OLS estimates. Dependent variable Δwc.

Variable	Denmark	Finland	Norway	Sweden
Δcpi	0.517 (0.144)***	1.000	0.629 (0.064)***	0.614 (0.102)***
$\Delta scope$	0.483 (0.144)***		0.371 (0.064)***	0.386 (0.102)***
Δpt	0.114 (0.412)	0.771 (0.481)	0.439 (0.339)	0.975 (0.272)***
Δh	-0.549 (0.398)	-0.751 (0.608)	-0.223 (0.209)	
rpr_{-1}	0.131 (0.090)	0.055 (0.031)*	0.023 (0.030)	-0.011 (0.033)
Δu	-0.006 (0.019)	-0.026 (0.024)	-0.024 (0.012)*	-0.016 (0.023)
u_{-1}	-0.029 (0.008)***	-0.012 (0.004)***	-0.020 (0.005)***	-0.041 (0.010)***
Δlmp	-0.191 (0.160)	-0.038 (0.080)	0.118 (0.051)**	-0.007 (0.041)
lmp_{-1}	-0.035 (0.106)	-0.035 (0.053)	-0.048 (0.047)	-0.046 (0.044)
s_{-1}	-0.381 (0.230)	-0.169 (0.058)***	-0.177 (0.070)**	-0.268 (0.068)***
$Constant$	-0.037 (0.041)	0.027 (0.029)	-0.055 (0.019)***	-0.156 (0.045)***
$i1979$			-0.045 (0.012)***	
$i1988$			-0.051 (0.013)***	
$i1989$			-0.005 (0.017)	
Sample	1968-94	1963-94	1964-94	1965-94
SER	0.0186	0.0146	0.0098	0.0177
$AR1-2$	$F(2,14)=0.55$	$F(2,21)=3.36$	$F(2,16)=0.08$	$F(2,20)=0.73$
$ARCH1$	$F(1,14)=0.00$	$F(1,21)=0.17$	$F(1,16)=0.07$	$F(1,20)=0.67$
Normality	$Chi^2(2)=0.92$	$Chi^2(2)=0.15$	$Chi^2(2)=0.89$	$Chi^2(2)=0.57$
Stability[a]	2.14	1.87		1.48
t_{ECM}	-1.66	-2.91*	-2.54	-3.92***

See legend and footnotes to Table 1.

[a]Hansen's test for joint parameter stability.