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I. INTRODUCTION

Although the empirical issues involved in the Scandinavian model of inflation have received a fair amount of attention, the first papers that consolidated the foundations of the theory were published before the unit roots and cointegration revolution. For example, Calmfors (1977) has tested the main assumptions and propositions of the model. However, since no attention is paid to the temporal properties of the variables under consideration, the standard estimation techniques used by him were likely inappropriate.

In this paper the Swedish manufacturing wage equation is estimated using the Granger-Engle two step estimation procedure. That is, the variables involved are first tested to determine their level of integratedness and the obtained information is then used to search for stationary linear combination of individually nonstationary variables. This search for cointegration can then be thought of as a pre-test to avoid 'spurious regression' results (Granger, 1986). Finally, using the Granger representation theorem according to which for a set of variables that is cointegrated there always exists an error-correction formulation, a dynamic model including the 'equilibrium error' term will be estimated.

The error-correction formulations of the wage formation in Nordic countries have risen into prominence since the 1980s. For example, Nymoen (1989) uses it to estimate the wage equation for Norwegian manufacturing, augmenting its model with wedge terms to account for the effect of payroll and income taxes. Nymoen and Calmfors (1990) use error-correction model to analyse wage formation in four Scandinavian countries. The hallmark of their work is heavy emphasis on the institutional aspects of their rigidly regulated and centralised labour markets. This paper uses the most basic approach, partly because the data that were used in the two papers alluded to beforehand were not available to us.

It should be mentioned at the outset that the paper is a report of failure to find empirical evidence for the Scandinavian model, at least in the format that the latter is presented below. Although once the cointegrating regression is properly augmented, we cannot reject the null hypothesis of cointegration between wages, import prices and productivity - which is the main conjecture of the model - the error-correction term based on this relationship turns out to be highly insignificant in the dynamic model. Finally, we end up with the model that apart from the seasonal and oil shock dummies has a lagged change in import prices as the only explanatory variable! Whether this result is due to misspecification, poor luck, inappropriateness of the theory, or some combination of them, remains at this stage an open-ended question.

We proceed as follows. Firstly, we expose the main assumptions and implications of the Scandinavian model (Section II). Secondly, we try to search for connections between the Scandinavian and Phillips curve theories of inflation. Following Nymoen (1989), it will be argued that the error-correction model encompasses both the Scandinavian and Phillips curve type models (Section III). Thirdly, since the Scandinavian model predicts that the wage share is stationary, we examine the time series properties of the wage share data. In particular, we inspect whether the observed nonstationarity of the wage share might have been caused by the structural break produced by the oil price shock. We also discuss the temporal properties of several other time series that are relevant to wage modelling (Section IV). Fourthly, by trying to find a cointegrating vector relating four variables, we analyse the long-run behaviour of the system (Section V). Fifthly, we formulate a dynamic wage equation for Swedish manufacturing wages that embraces the established cointegration results (Section VI).

II. SCANDINAVIAN MODEL OF INFLATION

The Scandinavian model combines the essential elements of the 'structural' hypothesis of inflation (*eg* Baumol, 1967) with a special transmission mechanism of price changes from the world market to a small open economy. 'Smallness' in this context is defined by the assumption that the country is a price-taker in the world market. By emphasising the 'structural' aspects and the direct price links as primary channels for the inflation transmission, the Scandinavian model is in sharp contrast with the demand-oriented Keynesian and monetarist hypotheses.

For more formal exposition, let us define the following variables: P = the domestic currency price of the tradable commodity; P_W = world market price in the foreign currency; P_N = price of the nontradable good; P_C = composite price index; E = exchange rate (the price of the foreign currency in domestic currency); W and W_N = wages in the tradable and nontradable sectors; Q and Q_N = respective productivities; and α = weight of the tradable sector in total output. The basic assumptions of the Scandinavian model can now be written as follows (Calmfors, 1977; Lindbeck, 1979)[1]

[1] A dot above a lowercase variable denotes the relative rate of change of the variable.

Perfect arbitrage for a homogenous aggregate tradable commodity

$$\dot{p} = \dot{p}_{\mathrm{W}} + \dot{e} \qquad (1)$$

Constant factor income shares in tradable sector

$$\dot{w} = \dot{p} + \dot{q} \qquad (2)$$

Homogenous labour market

$$\dot{w}_{\mathrm{N}} = \dot{w} \qquad (3)$$

Constant factor income shares in nontradable sector

$$\dot{p}_{\mathrm{N}} = \dot{w}_{\mathrm{N}} - \dot{q}_{\mathrm{N}} \qquad (4)$$

Price index with constant weights, expressing sector shares of output

$$\dot{p}_{\mathrm{C}} = \alpha\,\dot{p} + (1-\alpha)\,\dot{p}_{\mathrm{N}} \qquad (5)$$

Combining the equations (1) through (3) yields the expression for the wage growth in the sheltered sector

$$\dot{w}_{\mathrm{N}} = \dot{p}_{\mathrm{W}} + \dot{e} + \dot{q} \qquad (6)$$

which together with the constant labour share assumption (4) means that the prices of nontradables grow at the rate of

$$p_{\mathrm{N}} = \dot{p}_{\mathrm{W}} + \dot{e} + \dot{q} - \dot{q}_{\mathrm{N}} = \dot{p} + \dot{q} - \dot{q}_{\mathrm{N}} \qquad (7)$$

Substituting (7) into the expression of the aggregate price index given by (5) results in the main message of the Scandinavian model

$$\dot{p}_{\mathrm{C}} = \dot{p} + (1-\alpha)\left(\dot{q} - \dot{q}_{\mathrm{N}}\right) \qquad (8)$$

All variables on the RHS are exogenous. Therefore, the domestic rate of inflation in a small open economy is fully explained both by the world rate of inflation and the difference in between the rates of increase of labour productivity in two sectors weighted by the share of the nontradables sector. The last term of (8) can be interpreted as 'structural' inflation, whereas the first term expresses the direct price link from the world market to domestic economy. The hypothesised causal links are demonstrated graphically in Figure 1, where the variables enclosed in the rectangles are exogenous while those in the ellipses are explained by the model.

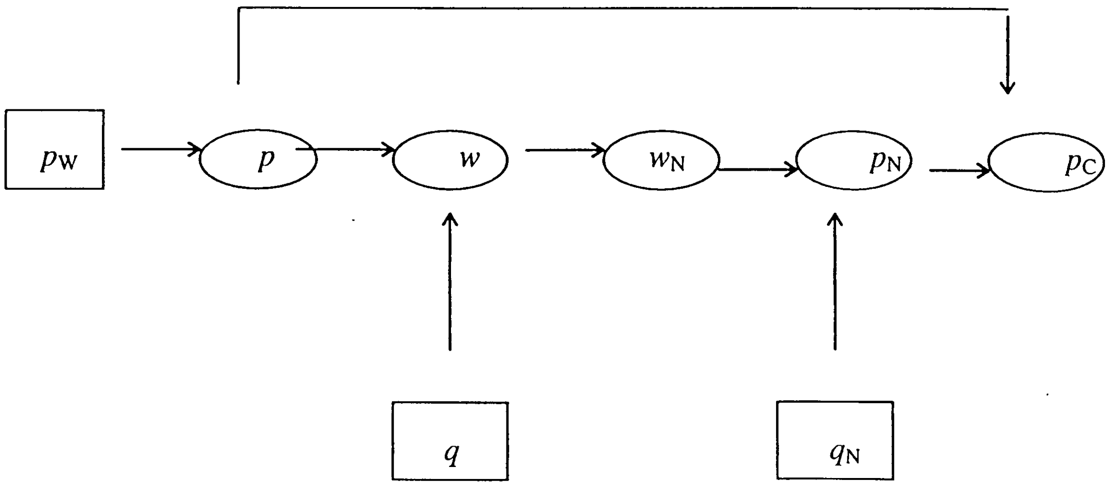

Figure 1. Diagrammatic representation of the Scandinavian model of inflation (Frisch, 1977).

Thus according to the model, a course through time of the national wage level is ultimately set by the world inflation, by the chosen exchange rates, and by productivity increases of the exposed industries. In the literature such a time path is usually referred to as the 'main course' of wages (*eg*, Aukrust, 1977). In what follows we concentrate our attention exclusively to the manufacturing as the representative of the exposed sector, which according to the model acts as a wage leader and for which, differently from the sheltered sector, data are readily available.

III. INCORPORATING THE DEMAND-SIDE INTO THE SCANDINAVIAN MODEL

One of the testable implications of the presented model is that wage-shares in the exposed industries are stationary stochastic variables. Or viewed from the opposite angle, in the long-run, wages in the exposed sectors adjust as to leave profits with a long-run normal share (π).

Retaining the definitions of variables given on the page 2, the main-course theory of wages can be expressed as[2]

$$w_t = \ln(1-\pi_t) + p_t + q_t + v_t \qquad ($$

where v_t is white noise. The composite variable $p_t + q_t (= mc_t)$ is often called the main-course variable. In general, mc_t is regarded to be strongly exogenous, *ie* it is affected neither by the current nor lagged values of wages (Kennedy, 1992).

[2] From here onwards, lowercase letters denote logs unless otherwise specified.

The main-course hypothesis as represented by the (9) is clearly a supply-side model, and although it may well be an adequate theory over a reasonably long time horizon, it is difficult not agree with Branson and Myhrman who argue that 'explaining the rate of inflation from the supply or cost side alone is equivalent to assuming that it is only the supply blade of the Marshallian supply-and-demand scissors that does the cutting' (1976, p. 17).

To allow wages to respond to the changing demand conditions and to cope with the apparent nonstationarity of most of the relevant time series, most empirical models of manufacturing wages on the Scandinavian data have added a term to measure labour market pressure and used differenced data

$$\Delta w_t = \alpha_0 + \alpha_1 u_t + \beta\,(L)(\Delta p_t + \Delta q_t) + \varepsilon_t \qquad (10)$$

where $\alpha_1 < 0$, u is the log of the unemployment rate, $\beta(L)$ a polynominal in the lag operator, and the notation $\beta(1) = 1$ is used to denote the sum of the coefficients, *ie* $\beta(1) = 1 - \beta_1 - \beta_2 - \ldots - \beta_i = 1$

In order to examine the long-run implications of (10) we make the simplifying assumption that the main-course variable grows along a deterministic trend, *ie* $E(\Delta w_t) = (\Delta p_t + \Delta q_t) = \Delta mc_t = g$, where g is the deterministic trend. Taking expectations of (10) gives

$$E(\Delta w_t) = g = \alpha_0 + \alpha_1 E(u_t) + E[\beta\,(L)(\Delta p_t + \Delta q_t)]$$

Substituting $\Delta p_t + \Delta q_t$ with Δmc_t and expanding $\beta(L)$ yields

$$g = \alpha_0 + \alpha_1 E(u_t) + E[(1 - \beta_1 L - \beta_2 L^2 - \ldots - \beta_i L^i)\Delta mc_t]$$

$$g = \alpha_0 + \alpha_1 E(u_t) + E(\Delta mc_t) - \beta_1 E(\Delta mc_{t-1}) - \beta_2 E(\Delta mc_{t-2}) - \ldots - \beta_i E(\Delta mc_{t-i}) \qquad (11)$$

Since, by assumption, the main-course variable grows along the deterministic trend, it follows that $E(\Delta mc_{t-i}) = g$ for every i, and we can rewrite (11) as

$$g = \alpha_0 + \alpha_1 E(u_t) + g(1 - \beta_1 - \beta_2 - \ldots - \beta_i)$$

but since according to (10) $\beta(1) = 1$, we have

$$g = \alpha_0 + \alpha_1 E(u_t) + g \qquad (12)$$

From (12) it follows that the Phillips curve model is consistent with the main-course if $E(u) = -\alpha_0/\alpha_1 \equiv u^N$, where u^N is interpreted as a natural rate of unemployment. Thus, if unemployment is at its natural rate, then

$$E\left[w_t - mc_t \middle| u^{\mathrm{N}}\right] = \ln(1-\pi) \tag{13}$$

ie, the expected wage-share develops as predicted by the main-course theory at every point in time. The major problem with the equation (10) is that although it does take account the neglected demand side of the main-course theory, it does not encompass the stationarity hypothesis of the latter.

Nymoen (1989) has shown, however, that a wage equation in the error-correction form, encompasses both the stationarity hypothesis of the main-course and the Phillips curve model. Assuming that u is stationary and weakly exogenous, Nymoen considers the following model.

$$\Delta w_t = a_0 + a_1 u_t + b(L)(\Delta p_t + \Delta q_t) - \gamma\left(w_{t-1} - p_{t-1} - q_{t-1}\right) + e_t \tag{14}$$

where $a_1 \leq 0$ and $b(1) \neq 1$. Taking expectation of (14) yields

$$E(\Delta w_t) = g = a_0 + a_1 E(u_t) + E[b(L)(\Delta p_t + \Delta q_t)] - \gamma\, E(w_{t-1} - p_{t-1} - q_{t-1})$$

In the steady state the last term is clearly equal to ln(1 - π). Thus we have

$$g = a_0 + a_1 \bar{u} + (1 - b_1 - b_2 - \ldots - b_i)g - \gamma \ln(1-\pi)$$

Solving for the wage-share

$$\ln(1-\pi) = \frac{1}{\gamma}\left[a_0 + a_1\bar{u} + (b(1) - 1)g\right] < 0 \tag{15}$$

The equation (14) gives us thus an error-correction representation of wages, corresponding to a situation where wages are cointegrated with the main-course variable. Clearly the Phillips curve formulation (10) is now encompassed by (14). That is, the error-correction model is capable of explaining the results produced by the Phillips curve equation (Gilbert, 1986), for the latter corresponds to the special case of (14) when $\gamma = 0$, *ie* there is no cointegration between the main-course and wages, or in other words, v_t in (9) is $I(1)$.

The above interpretation is crucially dependent on the assumption that both unemployment and wage-share are stationary. If, however, unemployment happens to be $I(1)$, it is either an irrelevant explanatory variable, or it is cointegrated with the wage-share, in which case the wage share itself is $I(1)$. In the latter case (14) explains the nonstationarity of the wage-share with nonstationarity in unemployment. The next logical step is thus to analyse the temporal properties of the relevant time series.

IV. UNIT ROOTS AND STRUCTURAL BREAKS

Before proceeding to examine the sets of variables for cointegration, it is necessary to establish the temporal properties of the individual series. Information about the integratedness limits the set of possible cointegrating relationships, for only the variables that are integrated of the same order can be cointegrated (Hall, 1986). Table 1 reports the *DF* and *ADF*(1) statistics for the levels and first differences for the five series.

TABLE 1

Time Series Properties. DF and ADF(4) Statistics 1972Q-1993Q3 T=86. Critical Values are obtained from Microfit 3.0 (Pesaran and Pesaran, 1991)

	DF		*ADF(4)*	
Variable	*Without trend*	*With trend*	*Without trend*	*With trend*
w	-2.112	-1.214	-2.883	-1.333
u	1.593	0.906	-1.961	-2.183
p	-3.697	-1.323	-3.094	-1.918
p_C	-4.080	0.422	-3.645	-0.856
q^*	-0.548	-1.893	0.774	-4.852
l	-2.670	-2.021	-2.530	-2.540
Δw	-10.322	-10.778	-2.115	-2.835
Δu	-6.840	-7.396	-1.463	-1.845
Δp	-6.268	-6.833	-3.508	-4.647
Δp_C	-7.108	-8.588	-2.336	-4.606
Δq^*	-4.295	-4.224	-2.651	-2.541
Δl	-7.455	-7.598	-3.451	-3.465
Critical values	-2.895	-3.462	-2.897	-3.465

Definitions (see Appendix for details).

w	Hourly earnings in manufacturing
u	Unemployment rate
p	Import prices
p_C	Consumer prices
q^*	Smoothed productivity
$l = [\ln(1 - \pi)] = w - p - q$	Labour share

In the view of the alternative possible interpretations of the error-correction model given by (14), our primary interest lies in establishing the dynamic properties of the u and l series. Both *DF* and *ADF* statistics in Table 1 show that the log of the unemployment rate is clearly not stationary. Whether it is integrated of order one remains ambiguous, *DF* statistics shows that u is $I(1)$, whereas *ADF* suggests that even Δu is nonstationary. As regards the labour share (l), both tests show that it is $I(1)$. This means that wages cannot be cointegrated with the main course variable and that we have to consider the alternative specification of the error-correction model by adding unemployment to the set of cointegrating variables.

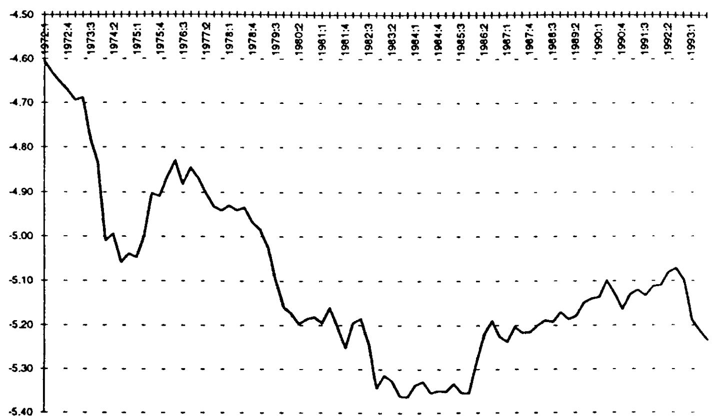

Figure 2. The log of the labour share $[(w - p - q^*)_t]$ in Sweden over 1972Q1-1993Q3.

A glance at the Figure 2 verifies that the labour share has indeed been far from being stable over the time period under consideration. On the other hand, it shows that the conspicuous falls have taken place in the middle of 1970s and at the beginning of 1980s, and are therefore most probably caused by the oil shocks. Another feature to be mentioned is that whereas after the first oil price hike around 1974, the labour share showed subsequently signs of rather brisk recovery, the second shock seems to have pushed it to a permanently lower level.

Thus the graph suggests that there might have been a structural break at the beginning of 1980s. It is well established, however, that unit root tests that do not allow for the possibility of one or more structural breaks under the null and alternative hypotheses are biased towards the nonrejection of a unit root (*eg* Enders, 1995; Holden and Perman, 1994). Looking at the graph, although always helpful, is not enough, for seemingly similar patterns may have been caused either by an one-time jump in the level of a unit root process or by an one-time change in the intercept of a trend stationary process. One possible way to test for unit root in the presence of a structural break involves splitting the sample into two parts and using Dickey-Fuller tests on each part. But in general, it is preferable to have a single test based on the full sample in order to avoid the problems caused by the diminished degrees of freedom.

Perron (1989) has developed a formal procedure to test for unit roots in the presence of a structural change at time period $t = \tau + 1$. He compares two models

$$y_t = \alpha_1 + \beta_1 t + \mu_1 D_L + \mu_2 D_T + \varepsilon_t$$ (1

where t is the time trend, D_L represents a level dummy variable such that $D_L = 1$ if $t > \tau$ and zero otherwise; and $D_T = t - \tau$ for $t > \tau$ and zero otherwise. Alternatively he considers

$$y_t = \alpha_1 + \rho\, y_{t-1} + \mu_2 D_T + \phi\, D_P + \varepsilon_t \tag{17}$$

where D_P is a pulse dummy variable such that $D_P = 1$ if $t = \tau + 1$ and zero otherwise. The equation (17) is the unit root null hypothesis whereas the equation (16) is the 'trend stationary about a breaking trend' alternative. Perron proposes the estimating equation that is obtained by pooling explanatory variables in (16) and (17) and adding lagged values of Δy_t, as in the augmented Dickey-Fuller framework. The equation to be estimated is therefore

$$y_t = \alpha + \rho\, y_{t-1} + \beta\, t + \gamma\, D_T + \phi\, D_P + \theta\, D_L + \sum_{i=1}^{i=k} c_i \Delta y_{t-1} + \varepsilon_t \tag{18}$$

Under the null hypothesis we have the restrictions $\rho = 1$ and $\gamma = \beta = 0$, whereas under the alternative hypothesis of a 'trend stationary' process we expect ρ to be less than one; β, γ, and θ to be non-zero and ϕ to be close to zero. The equation (18) was estimated using the data for the labour share over the period of 1972Q1-1993Q3. Since we are primarily interested whether after allowing for the structural break in the third quarter of 1979[3], l is still $I(1)$, we do not report full results of the regression here. The point estimate of ρ turned out to be 0.86 and its standard error 0.048, which implies that the relevant t-ratio is -2.92. The critical value is obtained from Table VI.B on page 1377 in Perron (1989) and it is dependent on the relative position of the break in the sample (λ). In our case $\lambda = 30/87 \cong 0.3$, and the critical value at a 5 per cent level is therefore -4.17. This implies that even after allowing for the structural break, we cannot reject the null hypothesis that the labour share is nonstationary. Together with our findings about the temporal properties of the unemployment series, this implies that a set of cointegrating variables is likely to include the unemployment rate.

But before we start to look for the cointegrating regressions, we have to establish the order of integration for the rest of the series. In almost every case, there is some contradicting evidence. For example, $ADF(4)$ statistics seems to suggest that nominal wages and productivity are integrated of higher order than one, and using DF and $ADF(4)$ without trend as a guidance does not allow to reject the hypothesis that both consumer and import prices are

[3] The break date $t = \tau + 1$ may be treated as known, or it may be determined endogenously. Perron (1994) suggests two alternative methods for choosing the breakpoints. We have treated the break date as known *a priori*, which is not perhaps theoretically very rigorous. For although the alleged structural break is in our opinion clearly related to the oil shock, there is some leeway in deciding exactly what quarter to choose for the date of the break. Thus the third quarter of 1979 is somewhat arbitrary.

$I(0)$. These problems are inherent in applying the *DF* and *ADF* tests. Although the introduction of lagged variables removes the serial correlation if the latter happened to be a problem, it may at the same time cause multicollinearity, if autocorrelation was not there in the first place. The reported tests can thus give only a rough guidance to the temporal propertied and may indeed, for the reasons mentioned beforehand, be misleading. The ultimate decision is thus necessarily the matter of judgement. Since there is no hard evidence to the contrary, we are therefore in a position to make the assumption that suits us most. Namely, in what follows all six variables in Table 1 will assumed to be $I(1)$.

V. COINTEGRATION AND THE TWO-STAGE PROCEDURE

In the last section we rejected the hypothesis that the $l_t = \mathbf{a}'\mathbf{x}_t$, $\mathbf{x}_t = (w_{t,}, p_t, q_t^*)'$ is stationary for a cointegrating vector $\mathbf{a} = (1, -1, -1)'$ whose value was suggested by the main-course theory. We also found that a set of cointegrating variables is likely to include the unemployment rate. Thus we are looking for a new cointegrating vector $\mathbf{b}$, such that $z_t = \mathbf{b}'\mathbf{m}_t$, is $I(0)$, where $\mathbf{m}_t = (w_{t,}, p_t, q_t^*, u_t)'$. Therefore, differently from the previous case, the theoretical model of the system dynamics does not suggest now a particular value for the cointegrating vector $\mathbf{b}$, and we have to first estimate it by *OLS*.

In the previous case when the cointegrating vector was fully specified, it was appropriate to use conventional unit root tests in checking for cointegration. Now it is different, because by using *OLS* to estimate the cointegrating vector we are essentially choosing between all possible vectors, thereby encountering the type of distributional problems associated with order statistics and multiple comparisons. In the present context this means that it is difficult to reject the null hypothesis that there are no stationary linear combinations when the observed data are used to estimate the most stationary-looking linear combinations before testing for cointegration (Dickey, *et al*, 1994). The practical upshot of all this is that the critical values for the Dickey-Fuller statistics for a unit root in the residuals from the cointegrating regression differ from those for a unit root in the variables involved in that regression.

The alternative to the Dickey-Fuller statistics is to use the Durbin-Watson statistic for the cointegrating regression known as the cointegrating regression *DW* statistics or simply *CRDW*. The rationale for using this statistic is provided by Sargan and Bhargava (1983) who show that the *CRDW* statistic has a probability limit of zero under the null hypothesis of non cointegration, which is therefore rejected for large values of the *CRDW* statistic. The problem with the CRDW statistic is that its critical values are not sufficiently constant across the various experiments, and therefore it has been suggested that the augmented Dickey-Fuller test is preferable (Holden and Perman, 1994).

Based on the discussion in Section III, a potential cointegrating regression for earnings per hour is:

$$w_t = \alpha_0 + \alpha_1 p_t + \alpha_2 q_t^* + \alpha_3 u_t + V_t \tag{19}$$

The results for the tests of stationarity of the residuals are shown below.

TABLE 2

Testing for the existence of the long-run equilibrium

Dependent variable: w_t	
Regressor	Parameter estimate
Constant	-5.13
p_t	0.24
q_t	1.87
u_t	-0.06
R^2	0.99
DF	-2.22
ADF(4)	-4.32
Time period	1972Q1-1993Q3
Number of observations	87

The critical value for the *DF* and *ADF* statistic is -4.23 (obtained from *Microfit 3.0*)

Neither standard errors nor t-ratios are given, for in the presence of nonstationarity the distributions of t-tests diverge so that for the conventional significance tests critical values will not be correct even asymptotically (Hendry, 1986). However, if cointegration holds, the parameter estimates will be superconsistent, in the sense that they will converge to their true value at a faster rate than normal *OLS* estimates (Jenkinson, 1986). Again after comparing the *DF* and *ADF* statistics with their critical values, we are left with uncertainty. While *DF* statistic is well below its critical value, *ADF* shows stationarity. Since we are dealing with quarterly data, it is probably more appropriate to use *ADF* statistic and to infer that the residuals are stationary. However, the point estimates of the coefficients are not what we would have hoped them to be. Although since the labour share is nonstationary, it is not entirely surprising to find that the coefficients of p and q^* are not equal to one, the long-run implications of the calculated coefficients are nevertheless staggering if not implausible. For the productivity elasticity of nominal wages well above one means that *ceteris paribus* in the long run labour gets all the factor income, whereas the small value of price elasticity implies the converse outcome.

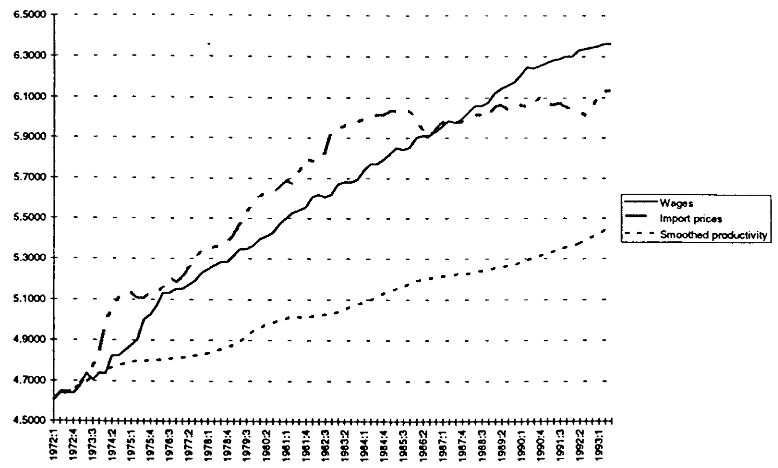

Figure 3. The logs of wages, productivity and import prices in Sweden over 1972Q1-1993Q3.

Figure 3 corroborates in a sense the findings in Table 2. Nominal wages have indeed grown well in excess of productivity over the last twenty years, whereas import prices have for the most of the period outpaced wages. This is in turn consistent with what we saw in Figure 2, where two sharp decreases in the labour share were caused by the two oil shocks. It may be that the point estimates given in Table 2 were to a certain extent influenced by multicollinearity that is bound to be there since we are dealing with trended variables. However, as Figure 3 demonstrated, there is some justification to the values of the coefficients of the cointegrating regression. This, however, prompts the question of whether the equation (19) really represents the long-run equilibrium relationship. For example, our findings are in sharp contrast with both Nymoen (1989) who has used the data of Norwegian manufacturing and Hall (1986) who has worked with the UK aggregate wage data. Nymoen estimated the following equation:

$$wc_t = \beta_0 + \beta_1 mc_t + \beta_2 \left(t_1 + t_2 + p_c - q\right)_t + \varepsilon_t \qquad (2$$

where wc_t is average wage costs per hour, mc_t the main course variable, t_1 and t_2 employees' and employers' tax rates, and p_C consumer prices, so that the last term of (20) is the so-called wedge between the product and consumption real wages.[4] Nymoen found the coefficient of the main-course variable to be equal to 1.02, which is consistent with the Scandinavian theory. Hall who assumed a closed economy estimated the following equation

$$w_t = \gamma_0 + \gamma_1 p_{c,t} + \gamma_2 q_t^* + \gamma_3 u_t + \gamma_4 h_t + \upsilon_t \qquad (21)$$

where h_t are average hours. According to Hall the coefficient of consumer prices[5] was 1.02 and that of smoothed productivity 0.93. Thus in the light of Hall's and Nymoen's work, our results in Table 2 seem to be rather implausible. Nevertheless at this stage we make a tentative conclusion that the four variables in (19) are cointegrated, although the first symptoms that something is to go wrong are already there.

Before to proceed further we test whether the exclusion of any of the variables would affect the stationarity of the residuals in (19). The idea is that if after dropping any of the variables cointegration could still be retained, it would suggest that this variable may be redundant in the cointegrating regression. Thus each of the three variables (p, q^* and u) was dropped, one at the time, and cointegration tests were carried out. Table 3 reports the results.

TABLE 3

Exclusion Test for the Three Variables

	Excluded Variable		
	p	q^*	u
DF	-1.41	-1.14	-1.70
ADF(4)	-3.65	-1.72	-3.71
R^2	0.98	0.91	0.99

In all cases the test statistics are considerably lower than in equation (19) and the exclusion of the productivity variable seems to have had a particularly strong effect. Therefore in order to estimate a valid error-correction model we must include the full cointegrating vector as in (19) in the level part of the model.

There is one more difficulty, however, to be solved before moving on to the dynamic specification. We have reached a tentative conclusion that equation (19) is a valid cointegrating regression involving four variables. But in cases where more than two time-series are being considered, there can be more than one stable linear combination. If cointegrating vectors are thought of as constraints that an economic system imposes on the movement of the variables in the long-run, the more cointegrating vectors there are, the 'more stable' the system (Dickey, *et al*, 1994). In the present context it means that instead of

[4] The similar approach was tried for Sweden, but no cointegrating relationship was found.

normalising w, *ie* the first element of the cointegrating vector **b** to be unity we could have as well normalised with respect to any other of the three variables. Given the properties of *OLS*, however, the resulting equilibrium relationship implied by the regression would be identical to (19) only in the limiting case of R^2 being equal to one (Hamilton, 1994). It is therefore interesting to see how different the equilibrium relationships using other dependent variables would be. Table 4 shows the various inversions of equation (19), we have rearranged the regressions so that w is always on the LHS for ease of comparison.

TABLE 4

The Effects on the Equilibrium Relationship of Changing the Dependent Variable

	Coefficients						
Dependent variable	*Constant*	p	q^*	u	R^2	*DF*	*ADF(4)*
w	-5.13	0.24	1.87	-0.06	0.9871	-2.22	-4.32
p	2.92	0.83	-0.77	0.02	0.9125	-2.02	-2.57
q^*	5.87	-0.11	2.17	0.01	0.9831	-1.80	-3.97
u	6.57	-0.09	-2.40	-0.52	0.2155	-0.66	-3.01

The critical value for the *DF* and *ADF* statistics is -4.23 (obtained from *Microfit 3.0*)

As expected, different inversions do produce considerably different estimates of the equilibrium parameters. However, the only inversion that gives a significant *ADF* value is the one using wages as a dependent variable. This equation also produces the highest R^2. The latter is important because although the estimates of the cointegrating regression are superconsistent, they are nevertheless subject to a finite sample bias. It has been argued (*eg* Hall, 1986) that the bias seems to be related to the overall goodness of fit of the regression, and so the regression with the highest R^2 should be subject to the smallest bias. Thus the dynamic modelling will be continued on the basis of the equation normalised on w which was only one producing stable residuals and giving the best fit.

[5] We also experimented with different prices in the cointegrating regression (from producer prices to export prices, and in what is essentially a departure from the Scandinavian model, consumer prices). However, the only regression that yielded stable residuals was the one that used import prices.

VI. DYNAMIC ECONOMETRIC MODELLING

After having found an appropriate specification of the cointegrating relationship, we can now proceed to the second stage of the estimation procedure. Granger and Engle have shown that if a set of variables is cointegrated then there always exists an error-correction formulation of the dynamic model and *vice versa*. It must be emphasised, however, that compared with the tests for cointegration, formulating and estimating a full dynamic model is considerably more intricate. Although we have already resorted quite a few times to our discretion, the specification of the dynamic adjustment process amplifies its scope even further even if broadly agreed rules are being followed.

In what follows the 'broadly agreed rules' are those of D. Hendry's 'general to specific' methodology, where one starts with a very general hypothesis and then narrows it down by looking for simplifications that are acceptable on the data. In moving from the general to simple, the investigator therefore has to confine his attention to specifications that are in the words of Gilbert (1986) *F*-acceptable. But as there may be alternative *F*-acceptable simplifications of the general representation, the role of discretion can at best be mitigated but not ruled out entirely.

We start with the following general specification

$$w_t = \sum_{j=0}^{4}\alpha_j \Delta w_{t-j-1} + \sum_{j=0}^{5}\beta_j \Delta u_{t-j} + \sum_{j=0}^{5}\gamma_j \Delta p_{c,t-j} + \sum_{j=0}^{5}\delta_j \Delta p_{t-j} + \sum_{j=0}^{5}\phi_j \Delta q^*_{t-j} + ec_{t-1} + \mathbf{h}'\mathbf{a}_t + \varepsilon_t \qquad (22)$$

where ec is the error-correction term, or the residuals from equation (19) and $\mathbf{a}_t$ is a vector consisting of the constant term, three seasonals and a oil shock dummy.[6] In including a distributed lag in consumer price inflation $\sum\gamma_j\Delta p_{c,t-j}$ we have followed the path taken by Nymoen (1989), and it is probably worth mentioning that not all models of manufacturing wages based on the Scandinavian theory contain this variable. The presence of consumer prices suggests that in the short-run the wage inflation in the exposed sector may be influenced by the developments in the nontradables sector which are then captured in this particular variable. The implicit - and not very innocuous - assumption here is that p_c is at least weakly exogenous.

Thus equation (22) was estimated using *OLS* for the period of 1972Q1-1990Q3 which leaves 12 observations out of sample for forecasting. Given the fact that we have 69

[6] The specific definition of the oil shock dummy is one example of exercised discretion. In various different regression specifications the third quarter of 1975 distinguished itself as an influential outlier. Given the

observations to estimate no less than 34 parameters, there are no doubt serious collinearity problems to be encountered.[7] In the view of the number of degrees of freedom this drawback should not definitely be downplayed. On the other hand, as Davidson *et al* (1978) have pointed out that collinearity problems are likely to occur in conjunction with omitted variables problems. They show that if the n initially excluded variables are important in determining the regressand, then adding them may help to resolve what appears to be a collinearity problem between the m originally included variables. The estimation results of the general model are given in Table 5.

The reported results do suggest that multicollinearity may be a problem, for out of 34 estimated parameters only two - Δw_{t-1} and oil dummy (D_t) appear to be significant. On the other hand, the general model meets all the requirements of the diagnostic statistics (for the F and χ^2-tests the p-values are given in the parentheses). The fourth column of the Table 5 reports the results of the variable exclusion tests for the respective groups of variables. As it turns out, only the $\sum\delta_j\Delta p_{t-j}$ group is significant at the 5 per cent level.

centralised nature of the Swedish wage bargaining procedure, it is plausible to conclude that shock fed through system in this particular period. Thus, oil dummy is equal to 1 in 1975Q3 and 0 otherwise.

[7] In fact different lag structures and lengths (but never higher than five) were tried, including those that allowed for fourth-differences and first-differences of the fourth-differenced data, but they were either not accepted on the grounds of diagnostic statistics or produced essentially the same results as equation (22).

TABLE 5

The General Dynamic Model

Dependent variable: Δw_t			
Variable	*Coefficient*	*t-value*	*F-test*
Constant	0.009	0.82	
Δw_{t-1}	-0.283	-2.25	
Δw_{t-2}	-0.011	-0.08	
Δw_{t-3}	0.113	0.78	$F_{6,34} = 1.64(0.18)$
Δw_{t-4}	-0.141	-1.02	
Δw_{t-5}	-0.120	-0.79	
Δu_t	-0.006	-0.21	
Δu_{t-1}	-0.024	-0.69	
Δu_{t-2}	0.008	0.26	$F_{6,34} = 0.55(0.77)$
Δu_{t-3}	0.017	0.56	
Δu_{t-4}	-0.022	-0.72	
Δu_{t-5}	-0.037	-1.24	
$\Delta p_{c,t}$	0.102	0.33	
$\Delta p_{c,t-1}$	0.402	1.32	
$\Delta p_{c,t-2}$	0.219	0.66	$F_{6,34} = 0.75(0.61)$
$\Delta p_{c,t-3}$	0.271	0.97	
$\Delta p_{c,t-4}$	0.110	0.37	
$\Delta p_{c,t-5}$	0.012	0.04	
Δp_t	-0.124	-1.43	
Δp_{t-1}	0.219	1.89	
Δp_{t-2}	-0.096	-0.93	$F_{6,34} = 2.42(0.05)$
Δp_{t-3}	-0.079	-0.73	
Δp_{t-4}	-0.161	-1.79	
Δp_{t-5}	0.088	1.08	
Δq^*_t	0.316	0.61	
Δq^*_{t-1}	-0.648	-1.31	
Δq^*_{t-2}	-0.221	-0.43	$F_{6,34} = 0.81(0.57)$
Δq^*_{t-3}	0.491	0.93	
Δq^*_{t-4}	-0.588	-1.14	
Δq^*_{t-5}	0.069	0.13	
ec_{t-1}	-0.054	-0.66	
D_t	0.094	3.89	
$S_{1,t}$	0.015	1.64	
$S_{2,t}$	0.017	1.59	$F_{6,34} = 1.87(0.15)$
$S_{3,t}$	-0.006	-0.67	

Time period: 1973Q3-1990Q3; $T = 69; k = 34$; $R^2 = 0.77$; $R^2_{adj} = 0.54$; Serial correlation: $F_{4,30} = 1.93(0.13)$; Functional form: $F_{1,33} = 1.30(0.26)$; Normality: $\chi^2_2 = 0.97(0.62)$; Heteroskedasticity: $F_{1,67} = 0.24(0.63)$; Predictive failure: $F_{12,34} = 0.61(0.82)$.

The next step was to move from the model given in Table 5 towards a more parsimonious equation. First we eliminated the groups of variables with the lowest values of the F-statistic. Each time checking not only the F-values calculated from the new and simpler regressions, but also making sure that all the diagnostics criteria were satisfied. Thus the exclusion of the $\sum\beta_j\Delta u_{t-j}$ and $\sum\gamma_j\Delta p_{c,t-j}$ groups was acceptable both in terms of their low F-values and the diagnostics statistics. Actually most of the diagnostic statistics improved after the elimination of these two sets of variables. The removal of the $\sum\phi_j\Delta q^*_{t-j}$ group, although acceptable in terms of the F-statistic, gave rise to the serial correlation. Therefore, after the elimination of the $\sum\beta_j\Delta u_{t-j}$ and $\sum\gamma_j\Delta p_{c,t-j}$ groups, the new sets of variables to be excluded were combined from the different sets of variables, the criterion for the exclusion being a low value of the t-statistic. As before, no more than five variables were removed at one time.

It was at this stage when the error-correction term dropped out from the equation, giving the second serious signal that the exercise was going to haywire. But since its t-value was only -0.47 after the number of regressors had shrunk to 18, there was little alternative to its exclusion.[8] The fact the error-correction term was after all insignificant casts obviously serious doubt to our previous finding that wages, unemployment, productivity and import prices were cointegrated. We must probably recall here that the DF-statistic suggested nonstationarity, but we nevertheless reached the opposite conclusion because ADF statistic allowed it, if only marginally. Therefore the finding of cointegration may have been an example of wishful thinking in the first place that was now rejected by the data. Also we must recall the discussion of the plausibility of the long-run equilibrium relationship implied by the calculated coefficients of the cointegrating equation.

Unfortunate as it from our perspective may be, at the end of the day, working at the five percent significance level, it was impossible not to drop all the explanatory variables except lagged changes in import prices, and both the seasonal and oil shock dummies. We cannot probably emphasise strongly enough that other the lag structures and lengths were experimented in the general equation. However, given the general framework of equation (22), they all tended to simplify down to one version. This speaks in a sense for the route independence, but probably only if we constrain our attention to this particular set of explanatory variables and take the implied long-run equilibrium equation at its face value. Table 6 reports the final specification.

[8] As it has been stressed already several times, different routes were tried, but the error-correction term never proved significant in the final equation.

TABLE 6

The Final Model

Dependent variable: Δw_t		
Variable	*Coefficient*	*t-value*
Constant	0.016	4.51
Δp_{t-1}	0.157	2.97
D_t	0.103	6.91
$S_{1,t}$	0.013	2.65
$S_{2,t}$	0.012	2.45
$S_{3,t}$	-0.017	-3.63

Time period: 1972Q3-1990Q3; $T = 73$; $k = 6$; $R^2 = 0.59$; $R^2_{adj} = 0.55$; $DW = 2.55$; Serial correlation: $F_{4,63} = 2.28(0.07)$; Functional form: $F_{1,66} = 3.60(0.06)$; Normality: $\chi^2_2 = 2.13(0.34)$; Heteroskedasticity: $F_{1,71} = 0.27(0.61)$; Predictive failure: $F_{12,67} = 1.06(0.40)$; Parameter stability (Chow): $F_{6,73} = 1.93(0.09)$.

Obviously after going through all the tests for cointegration it is a bit disappointing - to say the least - to end up with such a simple equation with only lagged import prices as a nondummy explanatory variable. But since this final specification was arrived at from rather many various starting points, we decided to present these results. Although the *p*-values of the various diagnostic statistics are in some cases (*eg* serial correlation and functional form) quite low, it does pass all the tests at the 5 per cent significance level. From the viewpoint of the Scandinavian model, the only comforting fact to be found in Table 6 is that import prices do seem to affect wages in exposed sector more than any other variable initially included in the general model. Although the point estimate of the import prices coefficient is rather low - approximately of the similar magnitude as that in the cointegrating regression - its *t*-value is highly significant. The fact that productivity did not affect the short-run wage formation may have to do with the way we constructed the q^*_t variable - smoothing probably eliminated most of the short-term dynamics. The lack of consumer prices from the final specification may also be interpreted as a demonstrating the prevalence of foreign factors over the domestic ones. It is somewhat paradoxical however that the unemployment rate that entered the cointegrating equation did not have any short-term effects. This is contrary to what the natural-rate-type stories have led us to believe.

The fitted values are shown together with the actual values in Figure 4. The equation fits actual data better in 1980s than in inflationary seventies. In this respect the unfortunate counterexamples are 1981 and 1989 where the model does not perform very well. The track record for seventies is much less impressive which may be blamed upon the oil price shocks (the same may apply with respect to 1981). The figure also shows that Swedish wage inflation is very much influenced by seasonal factors and that therefore quite a lot of explanatory work in our final model is actually done by the seasonal dummies.

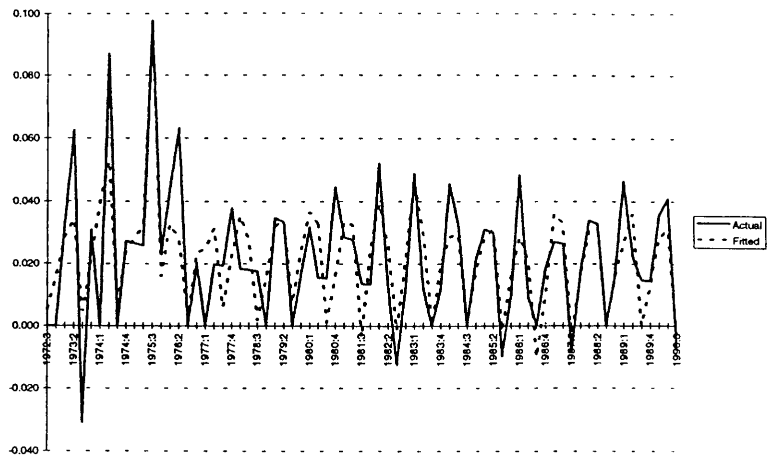

Figure 4. Fit of the model in Table 6. Estimation period 1972Q3-1990Q3

In order to test for the parameter stability, we re-estimated it for the period of 1972Q1-1980Q3. The results of the re-estimation are given in Table 7. As compared with the results in Table 6, serial correlation and functional form tests have improved quite remarkably and all other tests pass also at the 5 per cent level. As far as the parameter estimates are concerned, the picture is less satisfactory, for two variables - constant and the first quarter seasonal dummy - have ceased to be significant. Which is suggestive of parameter instability. However, the Chow tests do not show either parameter instability or predictive failure over the 1980Q4-1993Q3 forecast horizon. The fact that all the signs of the variables have remained unchanged and that the short-run elasticity of the lagged import prices is approximately of the same magnitude as before, indicates the stability of the model as well.

TABLE 7

Testing for Parameter Stability. Re-estimation of the Final model for the Period of 1972Q3-1980Q3

Dependent variable: Δw_t		
Variable	*Coefficient*	*t-value*
Constant	0.013	1.82
Δp_{t-1}	0.232	2.32
D_t	0.110	5.58
$S_{1,t}$	0.006	0.64
$S_{2,t}$	0.019	2.13
$S_{3,t}$	-0.019	-2.13

T = 33; k = 6; R^2 = 0.61; R^2_{adj} = 0.54; DW = 2.46; Serial correlation: $F_{4,23}$ = 1.14(0.36); Functional form: $F_{1,26}$ = 2.21(0.15); Normality: χ^2_2 = 0.56(0.76); Heteroskedasticity: $F_{1,31}$ = 0.05(0.83); Predictive failure: $F_{52,27}$ = 0.52(0.98); Parameter stability (Chow): $F_{6,73}$.= 1.49(0.19).

We have also calculated sequences of the predictive failure tests over increasing and decreasing forecasting periods. Figure 5 presents the results of the increasing horizon predictive failure tests. The estimation period is again 1972Q3-1980Q3, the beginning of the forecasting period is thus 1980Q4 and its end varies from 1981Q3 to 1993Q3. The length of the period increases each time by for quarters. As it can be seen from the figure, our final model does not present any evidence of predictive failure as the forecast period is gradually increased. None of the calculated F-values is significant even at the 35% level.

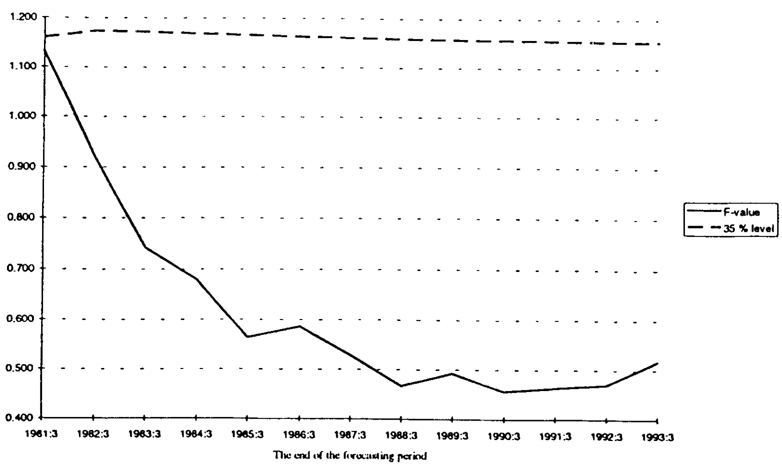

Figure 5. Predictive failure tests with an increasing forecasting horizon of the final model. (The estimation period is 1972Q3-1980Q3, the critical values are obtained from *Excel 5.0a*)

Figure 6 demonstrates the predictive stability of our final model using the same kind of test with a decreasing forecasting horizon. In this case it is the end of the estimation period that increases from 1980Q3 to 1992Q3 and pushes thereby each time back the beginning of the forecasting horizon. Again there are no indications of the predictive failure, as all the *F*-values are insignificant even at the 15 per cent level. Thus given the route of research that we have taken and the evidence that we have obtained through it, we have to conclude that the model given in Tables 6 and 7 is the best we are going to get provided that we do fundamentally change the underpinnings of our analysis. However unsatisfactory the model may be, it did show a considerable degree of route independence, provided that we stayed within the overall analytical framework of equation (22) which as most would agree, is an adequate representation of the Scandinavian model.

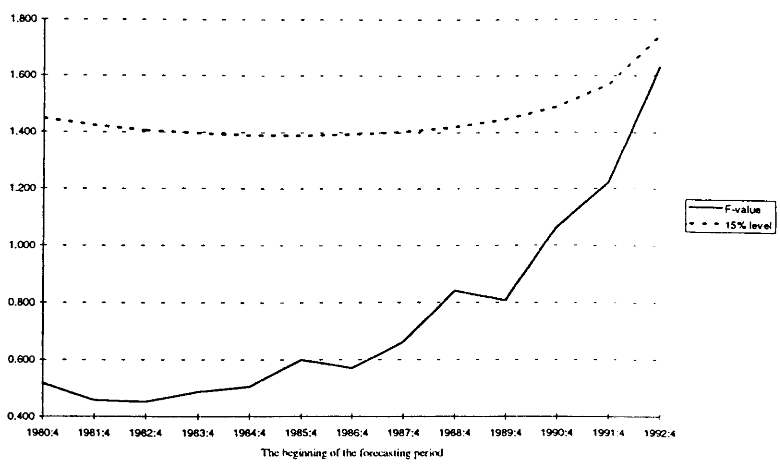

Figure 6. Predictive failure tests with a decreasing forecasting horizon of the final model. (The end of the estimation period increases from 1980Q3 to 1992Q3, the critical values are obtained from *Excel 5.0a*)

Gilbert (1986) has recommended that after the parameter constancy tests have been carried out, the model should be re-estimated using the entire sample in order to increase efficiency. This is done in Table 8 which is the final result of our analysis.

TABLE 8

The Final Model. Re-estimation over the Entire Sample

Dependent variable: Δw_t		
Variable	*Coefficient*	*t-value*
Constant	0.015	4.70
Δp_{t-1}	0.151	3.05
D_t	0.102	6.84
$S_{1,t}$	0.010	2.25
$S_{2,t}$	0.011	2.46
$S_{3,t}$	-0.016	-3.56

Time period: 1972Q3-1993Q3; $T = 85$; $k = 6$; $R^2 = 0.54$; $R^2_{\text{adj}} = 0.51$; DW = 2.35; Serial correlation: $F_{4,75} = 1.24(0.30)$; Functional form: $F_{1,78} = 3.98$ (0.05); Normality: $\chi^2_2 = 1.53(0.46)$; Heteroskedasticity: $F_{1,31} = 0.61(0.44)$

Since the parameter signs have essentially remained the same, there is no need to comment on them any more. As regards diagnostic statistics, the only serious problem has to do with the functional form. But since its F-value is still insignificant at the 5 per cent level, we can probably continue to assume the correct specification, although with considerable scepticism. The other fact to be mentioned is that the value of R^2_{adj} which in all specifications thus far has been of the magnitude of 0.54-0.55, has fallen to 0.51. The noticeable decline in the goodness of fit may in this case have to do with the effects of 1992 devaluation (which might have also affected the already poor functional form statistic).

VII. CONCLUSIONS

This paper tested the main proposition of the Scandinavian model of inflation: in the long run wages in the exposed sectors of an open economy are determined by world market prices and productivity. Statistically this implies that wage shares in the exposed industries must be stationary. Integration and cointegration tests for Swedish manufacturing data suggest that the unemployment rate should be included in the cointegrating equation, so that nonstationarity of the wage share will be explained with nonstationarity in unemployment. However, the parameter estimates of the cointegrating regression prompted us to question the plausibility of the implied long-run behaviour.

Next we developed a dynamic model that is based on the earlier estimated cointegration results. The apprehensions that were first raised by the coefficients of the cointegrating regression, were justified when we ran into a contradiction. Namely, the term that showed the divergence from the long-run equilibrium implied by the cointegrating regression, was insignificant in the dynamic specification. In fact the only nondummy explanatory variable that survived all exclusion and simplification tests was an one period lagged change in the import prices. This speaks for Scandinavian model in a sense that import prices are definitely important in determining the exposed sector's wage increases. On the other hand, the final equation does seem to be too simple to be true. The further reasons for fearing misspecification were given by the high values of the functional form statistics, which in general though were not significant at the 5 per cent level. At the same time, the various parameter stability tests did not give any signs of structural breaks even at the significance levels well beyond 5 per cent. Thus, it is hard to be too dismissive about our final specification.

In all, however, it is difficult to end this paper otherwise than at a highly sceptical note. One of the reason for this is that it is hard to imagine that the wage modelling in Sweden can be meaningfully done without any reference to the institutional framework of highly centralised wage bargaining and the effects of the generous system of unemployment benefits and the labour market programmes. It may be that the complete ignoring of these factors - which occurred mainly because of the lack of data - caused the final model to be what it is, *ie* rather unsatisfactory.

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APPENDIX: DATA DEFINITONS

OECD refers to the *OECD: Main Economic Indicators* and *IMF* refers to *IMF: International Financial Statistics.* Both *OECD* and *IMF* are CD-ROM databases at the Cambridge University's Faculty of Economics.

W is index of hourly earnings in manufacturing. *OECD:* 60431502.

U is standardised unemployment rate. *OECD:* 60428614.

P_C is consumer price index net of indirect taxes. *OECD:* 60446202

P is index of unit value of imports. *IMF:* 144 75.

Y is index of industrial production for manufacturing. *OECD:* 60204502.

H is index of monthly hours worked. *OECD:* 60428802.

Q is output per man-hour, defined as $Q = Y/H$.

Q^* is a smoothed version of Q defined as

$$Q^* = \frac{1}{8}\sum_{i=0}^{7} Q_{t-i}$$

All indices have been rebased so that 1972Q1=100.

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