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Wage Setting, Taxes and Demand for Labour: Multivariate Analysis of the Cointegrating Relations

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Abstract

As an outcome of modelling wage setting and demand for labour in a bargaining framework, a 10-dimensional vector space is defined. This overparametrized system is reduced by means of conditioning. Structural restrictions which identify the long-run relations of interest are specified. The restrictions which characterize wage setting and labour demand schedules are imposed and tested first separately and then jointly.

Generally, the restrictions which satisfy the condition for formal identification pass the tests on a fairly high significance level. The plausibility of resulting cointegrating relations applies not only to the signs but to the magnitudes of the coefficients as well. This concerns the economic identification. Preliminary evidence indicates that the preferred relations satisfy conditions for empirical identification as well.

The relations show up almost identically in partial and joint analysis. They appear not to be sensitive for the choice between r=4 or r=3 which is the factor defining the number of cointegrating relations in the system. Finally, the relations are hardly influenced by alternative conjectures concerning the endogeneity of various tax rates.

According to the results, in the private sector in Finland neither of the bargaining parties – employers or employees – has gained full dominance in the wage setting process. However, the coefficient estimates indicate that the adjustment in the long run has more closely reflected union goals. An equiproportional increase in average and marginal income tax rates has been shifted to higher pre tax wage level with an elasticity of around two thirds. Interestingly, the adjustment coefficient related to other taxes is considerably smaller, around one third or slightly less. Stronger unions have pushed up the wage level. Higher real prices of raw-materials have reduced wages. The driving force of the real wage growth is the productivity growth. As far as demand for labour is concerned, the negative impact of real labour cost is clear-cut. The additional output effect is an indication of imperfect competition in the product market.

Contents

1 Introduction[1]

The introduction of the concept of cointegration has provided an interesting avenue for investigation of the long-run relations between non-stationary variables. For estimation, Granger & Engle (1987) proposed a two-step method which was applied in Tyrväinen (1988, 1992a) to investigate determination of wages and employment in Finland. More recently, Sören Johansen (1991b) has proposed a maximum likelihood procedure for estimation of the multivariate systems. As the two-step method only picks up one of the potential candidates for the relevant long-run relations with no consideration of the others, the Johansen method is a considerable step forward as it allows us to evaluate the vector space in a more thorough manner. Furthermore, testing of hypotheses and discussion on identification is more straightforward in this framework (see Johansen & Juselius, 1992a,b).

In the present paper, wage setting and demand for labour in Finland is investigated applying the Johansen procedure. Furthermore, we wish to put considerable emphasis on issues related to identification of the relations of interest in the three contexts relevant in econometric work (see Johansen & Juselius, 1992b). That is, we are interested in 1) formal identification, which is related to the statistical model, 2) empirical identification, which is related to the actual estimated parameter values, and last but not least, 3) economic identification, which is related to the economic interpretability of the estimated coefficients of a formally and empirically identified model.

The paper is organized as follows. Section 2 introduces the economic model. Section 3 discusses the statistical model applied. In section 4 an unrestricted VAR-model deriving from the theoretical considerations is specified. The final operational model is arrived at by means of reducing this overparametrized model by conditioning. In section 5, the identifying conditions of the structural hypotheses are specified. In section 6, these restrictions are tested in terms of the partial model. Section 7 discusses the results concerning tax effects on wages in light of earlier evidence. Section 8 summarizes the paper.

[1] This paper has benefitted from useful comments of Søren Johansen, Katarina Juselius, David Hendry and participants of the Workshop on Multivariate Cointegration financed by the Joint Commitee of the Nordic Social Science Research Councils. Helpful discussions with Antti Ripatti and Erkki Koskela are also gratefully acknowledged. Needless to say, the usual disclaimer applies.

2 The economic model

There are n identical firms which have a production function F(N,m,K) with three inputs, labour (N), raw materials (m) and capital ($\underline{K}$) which is considered as predetermined. Furthermore, F(.) = $e^{\lambda t}Q(N,m,\underline{K})$ with the (steady) technical progress embodied in t. Imperfect competition is assumed to prevail in the product market. The firm maximizes profits which are defined as the difference between sales revenues and production costs:

$$\pi = \hat{P}[ZF(N,m,\underline{K})]F(N,m,\underline{K}) - W(1+s)N - P_m m, \qquad \text{(i)}$$

where $\hat{Q}^d = \hat{P}^{-1}(P)Z^{-1} \equiv D(P)Z$ is a downward sloping demand curve of the separable form introduced by Nickell (1978, p. 21). Z is a parameter describing the position of the demand curve faced by the firm and $\hat{P}$ is the (endogenous) producer price of the firm, P = competitors' producer price, W = nominal consumer wage, s = payroll tax, P_m = price of raw materials (incl. energy). The output of the firm, $\hat{Q}$, is considered as endogenous. According to the marginal product condition, optimal use of an input is determined by the relative price. If the firm uses raw materials optimally, the demand for labour schedule has the following standard form[2]

$$N^d = N^d\left(\underset{-}{\frac{W(1+s)}{P}}, \underset{+}{Z}, \underset{(-)}{P_m}, \underset{?}{\underline{K}}, \underset{?}{t}\right). \qquad \text{(ii)}$$

In an organized labour market the firm bargains with a union. The welfare of the union depends on the after tax real wage of its employed members and the (real) unemployment benefit received by the unemployed members, $U = U(W(1-\tau)/P_c, N, B)$ where P_c = consumer prices, τ = income taxes, and B = replacement ratio (unemployment benefit in real terms). As far as the partial derivatives are concerned, we assume that $U_1', U_2', U_3' > 0$ and $U_1'', U_2'', U_3'' < 0$ respectively. This general specification covers most of the common preference functions.

The widely used union models differ as regards the factors assumed to be bargained over. In the "right-to-manage" model preferred here, wages are bargained over and the profit maximizing firm sets employment unilaterally. The game is specified as a standard Nash solution of a cooperative game after Binmore et al. (1986):

$$\max_W (U - U_0)^{\theta}(\pi - \pi_0)^{1-\theta} \qquad \text{s.t.} \quad N(.) = \arg\max_N \pi \qquad \text{(iii)}$$

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[2] All signs in this section are according to Tyrväinen (1992a,b). The ones which do not necessarily hold generally are in parentheses.

where θ refers to the bargaining power of unions, $0 < \theta < 1$. U_0 is the fall-back utility of the union in the event an agreement is not reached. In Finland, the relevant alternative for a contract is a strike not only in economy-wide but in local negotiations as well. So, U_0 is assumed to depend on strike allowances, $U_0 = U_0(A)$. π_0 is the fall-back profit which reflects fixed costs during a production stoppage. When π_0 is deducted from the "under-contract" profits, fixed costs cancel out. For simplicity, fixed costs were already omitted from (i) above. The model defined in (iii) gives the monopoly union model and the efficient bargaining model as special cases.

The model for equilibrium (real) wage consists of variables influencing profits, on the one hand, and the utility of the union, on the other hand. In addition, a role is played by determinants of the fall-back utilities of the parties. Finally, the relative bargaining power matters. In its most general form, the wage setting schedule is

$$W^* = W(\underset{+}{P}, \underset{-}{s}, \underset{+}{\tau}, \underset{+}{\frac{P_c}{P}}, \underset{+}{\theta}, \underset{(-)}{P_m}, \underset{+}{Z}, \underset{+}{B}, \underset{+}{A}, \underset{+}{\underline{K}}, \underset{+}{t}). \tag{iv}$$

Indirect tax, v (value added tax, e.g.), influences as part of the price wedge, P_c/P.

Discrimination between bargaining models and other models is not straightforward. For instance, market clearing models can be specified so that they produce wage and employment schedules which are very much like those above. The role of bargaining power is, however, the distinguishing feature of bargaining models.

When the model is solved under the assumption of perfect competition on the product market, the demand shift variable drops out (see Tyrväinen, 1988). Hence, a significant presence of a variable describing aggregate economic activity in the labour demand schedule would give support to the hypothesis of monopolistic competition. "It is this channel that distinguishes this model from the competitive special case, and consequently the search for the significant presence for aggregate demand variables is an important aspect of their empirical implementation" (Andrews, 1988, p. 29).

3 The statistical model

The statistical analysis is carried out in terms of an n-dimensional vector autoregressive model

$$\Delta X_t = \Gamma \Delta X_{t-1} + \Pi X_{t-1} + \psi D_t + \mu + \varepsilon_t, \tag{v}$$

where the X's are stochastic variables, the D's are deterministic or non-normal variables (dummies),[3] μ is the constant term and ε is a Gaussian residual.

The case of particular interest is when Π is neither of rank zero nor full rank, $0 < r < n$. Now the hypothesis of cointegration indicates that we can write

$$\Pi = \alpha\beta'$$

where α (the adjustment coefficients) and β (the cointegration relations) are n x r matrices, and

$$\alpha'_{\perp}(I - \Gamma)\beta_{\perp} = \rho\zeta',$$

where $\alpha'_{\perp}$ is orthogonal to α and $\beta'_{\perp}$ is orthogonal to β ($\alpha'_{\perp}\alpha = 0$, $\beta'_{\perp}\beta = 0$) and ρ and ζ are (n–r) x c_1 matrices. If c_1 = n–r, we have an I(1) model; c_1 indicates the number of I(1) common trends. If, however, c_1 < n–r the model is I(2); c_2 = n–r–c_1 indicates the number of I(2) trends.

If (v) is I(1), the constant term can be partitioned into

$$\mu = \alpha\beta_0 + \alpha_{\perp}\gamma_0$$

where β_0 (which is r x 1) represents the intercept in the cointegration relations and γ_0 (which is (n–r) x 1) is a vector of linear trend slopes in the data. If $\alpha_{\perp}\gamma_0$ is zero, the data contains no linear trends (see Johansen, 1991b). This is a testable hypothesis. Since technical progress can be approximated by a linear trend, we expect the data to contain linear trends.

In empirical work, it is often advantageous to partition X_t into

$$X_t = \begin{bmatrix} \hat{X}_t \\ \bar{X}_t \end{bmatrix}$$

where the $\hat{X}$'s are the variables to be modelled in the system of equations and $\bar{X}$'s are the potentially weakly exogenous variables (see Engle, Hendry & Richard, 1983).

[3] To account for various extraordinary effects one almost always has to condition on dummies.

Let the parameters of interest be β (= the long-run parameters). We can partition the likelihood function into the conditional distribution and marginal distribution

$$f(X_t;\psi D_t) = f(X_t|X_{t-1}, D_t;\psi)$$
$$= f(\hat{X}_t|\bar{X}_t, X_{t-1}, D_t;\psi_1)\cdot f(\bar{X}_t|X_{t-1}, D_t;\psi_2). \tag{vi}$$

where ψ_1 and ψ_2 are functions of ψ. If β only depends on the parameters of the conditional model, $\beta = f(\psi_1)$, and ψ_1 and ψ_2 are variation free, then the inference is fully efficient about β from the conditional model. Johansen (1992b) shows that a sufficient and necessary condition is that $\alpha_{\bar{X}} = 0$, i.e. that the $\bar{X}$'s are weakly exogenous. This is a testable hypotheses and – if satisfied – it ensures that we do not loose any information concerning the long-run relations if part of the data is considered as weakly exogenous.

4 The empirical vector autoregressive model

4.1 The full model

In order to reach reasonably operationalized empirical model, a further note on the tax matters is worth to make. The majority of empirical literature characterizes the income tax system – summarized above with τ – with one parameter only, either the average tax rate, τ_a, or the marginal tax rate, τ_m. A more thorough analysis of the tax incidence in union models in Tyrväinen (1992b), however, shows that this may be insufficient. Since both matter and have separate role, they should be both included. In accordance with Tyrväinen (1992b), we expect that $W^*_{\tau_a} > 0$ and $W^*_{\tau_m} < 0$.

The evaluations above leave us with too many variables to deal with. Because a VAR model becomes vulnerable when the number of variables grows, the dimension of the system must not be "too large". Our choices are conducted by our special interest in studying the impact of taxes on wage setting and demand for labour. Therefore, such variables of great interest are reluctantly left out as capital stock and the related user cost, unemployment benefits and strike allowances. Earlier evidence indicates that the deficiency is not considerable except as far as the capital stock is concerned. Tyrväinen (1991), however, shows that wage equations for Finland remain by and large intact when the capital-labour ratio (K/N) is replaced by a productivity measure (Q/N). Notably, output, Q, enters our system as a proxy for the demand shift factor Z (for discussion, see Tyrväinen, 1992a) and N is in the model as well. Of course, the coherence of the resulting set-up will be scrutinized using common misspecification tests. As it stands, our system appears to be the largest analysed with the procedure concerned.

To sum up, our 10-dimensional VAR-model contains the following variables

1) nominal wages, W,
2) producer prices, P,
3) employment, N,
4) output volume, Q, which enters as a proxy for the demand shift factor, Z (for discussion, see Tyrväinen, 1992a),
5) income tax rate, average, $1-\tau_a$,
6) income tax rate, marginal, $1-\tau_m$,
7) rate of employers' social security contributions, 1+s,
8) rate of indirect tax, v, which is proxied by the price wedge, i.e. consumer price relative to the producer price, CPI/P,
9) (import) price of raw-materials (incl. energy), P_m,
10) unionization rate, UNION, which is a proxy for the bargaining power of the unions.

The observation period is 1970Q1–1990Q4 and majority of the (seasonally adjusted) series come from the data-base of the quarterly model of the Bank of Finland (BOF4). Some of the data are only available covering the aggregate economy, namely variables 5), 6), 9), and 10) above. In light of this, the analysis of the economy as a whole would be a straightforward choice. Earlier analysis, however, indicates that there is a profound difference in behaviour of the public sector wage and employment series as compared to the rest of the economy. This leads us to exclude public sector and to put the emphasis in analysis of the private sector.

According to tests none of the variables is non-relevant and could, hence, be excluded (exclusion-test). In addition, none of them is stationary (stationary-test).

4.2 The preliminary analysis of the model

The estimations are carried out with the maximum likelihood procedure proposed by Johansen (1991b).[4] To start, we test jointly the cointegration rank, r, and the existence of a linear trend. The rank defines the number of linearly independent stationary relations between the levels of the variables. The cointegrating relations are estimated as the eigenvectors corresponding to the r largest eigenvalues in the system with n variables. The magnitude of an eigenvalue, λ_i, indicates how strongly the cointegrating relation is correlated with the stationary part of the process. The test for a specific value of r concerns the hypothesis that $\lambda_{r+1} = ... = \lambda_n = 0$, whereas $\lambda_1, ..., \lambda_r > 0$ (see Johansen, 1992a).[5] As our earlier studies (Tyrväinen, 1991, 1992a) indicate that there may be I(2) series in our data, it must be taken into account. This is especially so, because a closer look at the residuals of the cointegrating relations also confirms the presence of I(2)-ness.

The 10-dimensional model is estimated and the results of the Trace test are in Table A1. Hypothesis for $r = \underline{r}$ is rejected if $H_0, ..., H_{\underline{r}-1}$ are rejected and further,

$$Q^*_{\underline{r}} > CV^*_{95\%} \text{ and } Q_{\underline{r}} > CV_{95\%}$$

where the superscript, *, derives from a system with no linear trend (i.e. $\alpha_\perp \gamma_0 = 0$ above). In Table A1, $H^*_0, ..., H^*_3$ and $H_0, ..., H_3$ are rejected as well as H^*_4. However, $Q_4 < CV_{95\%}$. So, the hypothesis of no linear trend is rejected. There appears to be four long run relations (r=4) and six common trends (c=n−r=6).

[4] Katarina Juselius has kindly supplied me the program for Cointegration Analysis of Time Series (CATS) in RATS.

[5] The likelihood ratio test statistic of the hypothesis of r cointegrating vectors in n-dimensional system is given by the so-called trace statistic, $Q_r = -T\sum_{r+1}^{n} \ln(1-\hat{\lambda}_i)$ where T is the number of observations. The distribution of the test statistic, which is a non-standard Dickey–Fuller type (involving a multivariate Brownian motion), has been tabulated for the asymptotic case in Johansen & Juselius (1990). The distribution depends on which assumption concerning the existence of the linear trend (yes or no) is maintained. The distribution has broader tails if the trend is absent.

According to the test statistics in Table A2 we conclude that the process is I(2) in one direction ($c_2=c-c_1=1$) indicating that there is one common I(2) trend which drives the system. So, the four-dimensional cointegration space is found to be stationary in three directions and non-stationary in one direction such that the differenced I(2) variable is needed to obtain stationarity. This is an example of multicointegration.

Investigation of the so called $\beta_{\perp}^{2}$-vector resulting from the system above (see Johansen, 1992a) reveals that I(2)-ness is mainly to be found in W, P and P_m. In addition, the coefficients of the first two in $\beta_{\perp}^{2}$ are approximately equal in absolute value. Hence, it is likely that the common I(2) trend could be eliminated by specifying the system in real terms.

An experiment with a 9-dimensional system comprising W/P and P_m/P indicates that r is reduced from 4 to 3 and, in addition, there is no I(2) common trend left. A profound deficiency of the "real model" is that price homogeneity is imposed to hold not only in the long run but in the short run as well which contradicts empirical evidence. To allow for temporary short-run deviation, we prefer to work with the nominal model. The question related to I(2)-ness will be reconsidered below.

The residual analysis indicates that some of the series do not seem to satisfy the Gaussian assumptions (see Table A3.A).

4.3 The partial model

As compared to the sample size, the number of variables is such that it creates the risk of the model to be overparametrized. Since the parameters are unrestricted in the model, the total number of estimated parameters increases very fast with increasing values of n and k (= lag length). Particularly in small samples overparametrization often leads to undesirable statistical properties. Let us first consider the choice of k. Residual misspecification tests indicate that we do not loose anything by restricting the lag length to k=2.[6]

Since the parameters of interest are the long run parameters, β, we proceed by considering weak exogeneity, i.e. whether some of the variables do not react to a disequilibrium in the long-run relations. There are several ways to evaluate the matter. First, weak exogeneity can be tested in terms of the full model although the deviations from the Gaussian assumptions obviously reduce the credibility of the test.[7] Second, we can evaluate qualitatively whether some of the variables are exogenously determined by nature (tax rates, e.g.). Third, the time series properties may give an indication of whether the data results from an endogenous data generating process.

Having indicated that there are probably four cointegrating vectors in our VAR-model, we cannot reduce the set of system variables to a smaller number. Wages, prices and employment are are with no doubt among the endogenous variables. Import prices of raw-materials are obviously exogenous. The

[6] This indicates that there is one lagged difference term in each equation.

[7] The hypothesis is that for selected equations, the α_i's are zeros. The test statistic is similar to that described in footnote 15 below.

quarterly time disaggregation of the union membership has been maintained by technical routines from annual observations. Hence, we are not willing to put too much weight on this quarter-to-quarter path. The discussion on taxes is less straightforward although there is a great deal of arbitrariness in the quarterly time paths here as well. To begin, since payroll tax rates are flat rates which are announced in advance it is natural to consider them as weakly exogenous. The price wedge contains both an endogenous and a weakly exogenous element. The latter is in the form of the sales tax which is a stable rate with a few stepwise shifts within the observation period. We expect the latter to dominate and start by considering the term as weakly exogenous. As far as income taxes are concerned, the conjecture of weak exogeneity is even more vague. One could argue that the union optimizes by taking the marginal income tax rate as given and strains after the optimal combination between the after tax wage and employment. In this case the marginal tax would be (weakly) exogenous and the average tax endogenous. Again there is, however, such an amount of arbitrariness in the quarterly time path of income tax series that we prefer to consider them as weakly exogenous at the outset. This issue will, however, be reconsidered at the end of the paper. At the moment, we prefer to consider output as the fourth system variable. This is so although tests in the present context as also those discussed in Tyrväinen (1992a) indicate that it could be considered as weakly exogenous as well.

So, we proceed by examining a model with four system variables,

W, P, N, Q,

and six weakly exogenous variables[8]

$1-\tau_a, 1-\tau_m, 1+s, CPI/P, P_m$, UNION.

The residual analysis of the partial model indicates no special problems (Table A3.B, see also Figure 1). Importantly, the simulations in Eitrheim (1991) indicate that the parameter estimates of the long-run relation are not sensitive for misspecification in other respects than the one generating autocorrelation.

[8] The present version of CATS can handle at most six weakly exogenous variables. This adds a technical restraint for the size of the model as far as introduction of additional weakly exogenous variables is concerned.

As can be seen in Table A3.B, autocorrelation appears not to be a severe problem in our partial model.[9]

[9] Discrete shifts in the data in level form may when differenced generate outliers which distort the estimation of the long-run parameters of interest. To avoid this we have included 6 quarterly dummies to the short run part of the model. These are 1) $D\tau_1$ and 2) $D\tau_2$ which are due to discrete changes in income tax schedules which became effective in 1976Q1 and 1989Q2 respectively, 3) Ds which is due to a discrete reduction in the employers social security contributions in 1990Q1, 4) DM is a dummy for the three week strike in the metal and engineering industry in 1971Q1, 5) DOPECI refers to the first oil shock in 1973Q3. Finally, the sixth dummy captures an important feature related to the institution of centralized wage setting. In Finland, wage agreements covering almost entire economy become effective simultaneously. This generates peaks to the differenced wage data in the contract quarters. Account of this has been taken by introducing a multiplicative, centered dummy, DCONT. In estimations, each of the six dummies enter at least one of the short-run equations significantly.

5 Specifying structural hypotheses

The model in section 2 was designed for analysis of wage setting and the demand for labour. To proceed, we need to define how to identify the relevant relations in the vector space. The problem is solved by imposing structure implied by theoretical considerations on the long-run relations. Below, these restrictions are made explicit.

First, let us define a log linear relation

$$\begin{aligned}\log W = {} & \beta_P \cdot \log P + \beta_N \cdot \log N + \beta_Q \cdot \log Q \\ & + \beta_{\tau_a} \cdot \log(1-\tau_a) + \beta_{\tau_m} \cdot \log(1-\tau_m) + \beta_s \cdot \log(1+s) + \beta_v \cdot \log(CPI/P) \\ & + \beta_{P_m} \cdot \log(P_m) + \beta_U \cdot \log(UNION)\end{aligned} \qquad \text{(vii)}$$

In so far as this relation is to be considered as a wage setting schedule (see (iv) above), the signs should be $\beta_P \geq 0$, $\beta_Q \geq 0$, $\beta_{\tau_a} \leq 0$, $\beta_{\tau_m} \geq 0$, $\beta_s \leq 0$, $\beta_v \geq 0$, $\beta_{P_m} \leq 0$ probably, and $\beta_U \geq 0$. Setting $\beta_N = -\beta_Q \leq 0$ is the restriction which defines the productivity term (Q/N) discussed above.

One can think of two (contradicting) extreme hypotheses about wage setting. In the first, unions are definitely dominating. According to the second, the opposite is true and the firms dominate. These hypotheses also indicate how price homogeneity is to be defined in the wage relation (vii).

If the unions dominate the process, price homogeneity is defined in terms of the target variable of the union, the level of the after tax real wage. This implies $\beta_P = \beta_v = -\beta_{\tau_a} = 1$ in (vii). If the firms dominate, the price homogeneity is defined in terms of the real labour cost. This implies $\beta_P = -\beta_s = 1$, and $\beta_v = 0$. If it is the real raw-material price which matters then instead of imposing $\beta_P = 1$ as we did above, the relevant restriction is $\beta_P + \beta_{Pm} = 1$ in (vii). Note also, that – if accepted – this restriction probably removes the I(2)-ness from the long-run relation.

If we take the discussion on raw-material prices into account and augment the hypothesis of the dominating unions by assuming that the tax burden is fully born by the firms, this implies $\beta_P + \beta_{P_m} = \beta_v = -\beta_{\tau_a} = 1$, $\beta_s = 0$ and $\beta_U > 0$. Under the same conditions, if the model with dominating firms is augmented to imply that the tax burden is fully born by unions, we have $\beta_P + \beta_{P_m} = -\beta_s = 1$, and $\beta_v = \beta_{\tau_a} = \beta_{\tau_m} = \beta_U = 0$.

In addition to those discussed above, several hypotheses concerning tax influence on wages have been applied in the literature. For instance, the so called wedge restriction which implies that $W^*_\tau = W^*_v = -W^*_s$ is widely used. In the relevant studies, no difference has usually been made between proportional and marginal income tax rates.

In empirical applications, the choice of the dependent wage term implies considerable amount of implicit structure to the model. This has seldom been

Table 1. **H_0-hypothesis about tax incidence in some wage equations**

		Dependent variable			
		(A)	(B)	(C)	(D)
Coefficient	Independent variables	W	$\left(\frac{W(1+s)}{P}\right)$	$\left(\frac{W(1-\tau)}{P(1+v)}\right)$	$\left(\frac{W}{P(1+v)}\right)$
β_P	P	H_0: $\beta_P = 0$	H_0: $\beta'_P = 0 \Leftrightarrow \beta_P = 1$	H_0: $\beta''_P = 0 \Leftrightarrow \beta_P = 1$	H_0: $\beta'''_P = 0 \Leftrightarrow \beta_P = 1$
β_s	(1+s)	H_0: $\beta_s = 0$	H_0: $\beta'_s = 0 \Leftrightarrow \beta_s = -1$	H_0: $\beta''_s = 0 \Leftrightarrow \beta_s = 0$	H_0: $\beta'''_s = 0 \Leftrightarrow \beta_s = 0$
β_τ	(1−τ)	H_0: $\beta_\tau = 0$	H_0: $\beta'_\tau = 0 \Leftrightarrow \beta_\tau = 0$	H_0: $\beta''_\tau = 0 \Leftrightarrow \beta_\tau = -1$	H_0: $\beta'''_\tau = 0 \Leftrightarrow \beta_\tau = 0$
β_v	(1+v)	H_0: $\beta_v = 0$	H_0: $\beta'_v = 0 \Leftrightarrow \beta_v = 0$	H_0: $\beta''_v = 0 \Leftrightarrow \beta_v = 1$	H_0: $\beta'''_v = 0 \Leftrightarrow \beta_v = 1$

discussed in the literature[10] despite the profound importance for empirical work and testing of behavioural hypotheses. To clarify the issue, Table 1 introduces the H_0-hypotheses in some of the commonly applied estimating equations. The a priori structure can be analysed by investigation of the β_i-coefficients. For instance, a closer look at equation (B) reveals that according to H_0-hypothesis tax incidence is fully borne by unions. In equation (C), according to H_0 tax incidence is fully borne by the firms.

Model (B) has been used by, for instance, Calmfors & Forslund (1990) and Calmfors & Nymoen (1990) with the a priori wedge restriction, $|\beta'_s| = |\beta'_\tau| = |\beta'_v|$. Model (C) has been used, among others, by Eriksson et al (1990) and by Pencavel & Holmlund (1988) with the a priori wedge restriction $|\beta''_s| = |\beta''_\tau| = |\beta''_v|$. Model (D) has been estimated by Rødseth & Holden (1990) with restrictions $\beta'''_s = \beta'''_\tau = \beta'''_v = 0$ imposed. As is obvious from the Table, the restrictions which may look similar at the first sight have very much different implications in different specifications. This may be of profound importance particularly if the elasticities related to taxes are not very precisely defined as it often appears. The procedure applied in this paper gives an opportunity to distinguish between the alternative hypothesis in a very flexible manner.

Finally, Padoa Schioppa (1991) suggests that the income tax system can be adequately described with the progressivity index proposed by Jackobsson (1978):

$$\tau_p = \frac{\tau_m - \tau_a}{1 - \tau_a}.$$

Taking logarithms and subtracting τ_p from unity, gives

$$\log(1-\tau_p) = \log\left(1 - \frac{\tau_m - \tau_a}{1-\tau_a}\right)$$
$$= \log(1-\tau_m) - \log(1-\tau_a).$$

So, τ_p just imposes a restriction for the coefficients of the two parameters describing the tax rates to be of opposite sign and equal in absolute value. In what follows, we take this as one of the testable restrictions.[11]

The discussion on the demand for labour schedule is more straightforward. To see this, let us write

[10] Calmfors & Nymoen (1990) is one of the few exceptions.

[11] All the three characteristics, τ_a, τ_m and τ_p, of the income tax system can be seen in Figure 1.

$$\log N = \beta_W \cdot \log W + \beta_P \cdot \log P + \beta_Q \cdot \log Q + \beta_{\tau_a} \cdot \log(1-\tau_a) + \beta_{\tau_m} \cdot \log(1-\tau_m) + \beta_s \cdot \log(1+s) + \beta_v \cdot \log(CPI/P) + \beta_{P_m} \cdot \log(P_m) + \beta_U \cdot \log(UNION) \qquad \text{(viii)}$$

If this relation is to be considered as a labour demand schedule, we expect that $\beta_W \le 0$, $\beta_P \ge 0$, $\beta_Q \ge 0$, $\beta_{\tau_a} = 0$, $\beta_{\tau_m} = 0$, $\beta_s \le 0$, $\beta_v = 0$ and $\beta_{P_m} \le 0$ probably. The impacts of the income tax as well as of the indirect tax derive from their wage effect. If the union has a direct influence on employment (for a given wage), then $\beta_U \neq 0$. The restriction $\beta_W = -\beta_P = -\beta_Q = \beta_s > 0$ implies that it is the real labour cost for a produced unit which matters. Again, if it is the relative raw-material price which matters, we write $\beta_w = -\beta_P - \beta_{P_m} = -\beta_Q = \beta_s$. If $\beta_Q > \beta_W = -\beta_P - \beta_{P_m} = \beta_s$, the product market demand has an additional effect which is in accordance with the hypothesis of imperfect competition on the product market.

We concluded above that there are probably four cointegrating vectors in our data space. Two of them have been specified so far. Further evaluations are less straightforward. The reason is as follows. As pointed by Johansen & Juselius (1992b), in macroeconomic behaviour a role is often played by at least two types of agents with disparate goals (demanders versus suppliers, producers versus consumers etc.) interacting in such a way that equilibrium is restored once it has been violated. Therefore, one may also pick up vectors describing either the demand side or the supply side of the variables concerned.

In the present context, we could for instance discover a relation describing the supply of output, i.e. the production function. Since the capital stock is not included in the model any relation can, of course, only be a poor description of the appropriate technology. On the other hand, the demand side may show up because the data set covers the largest component of houshold disposable income, Y, i.e. the after tax real income of the wage earners:

$$Y \sim \frac{W(1-\tau_a)N}{P(CPI/P)}. \qquad \text{(ix)}$$

The model, however, lacks export demand and government demand. Therefore we can only hope to discover a misspecified demand relation suffering from problems related to omitted variables. However, if the demand for output schedule shows up we expect that $\beta_Y > 1$. The reason is as follows. Since the government demand has grown considerably more than the houshold demand, the level of Q has risen more than Y. As a consequence, the misspecified output demand relation should produce an elasticity above unity. Strikingly, in so far as the demand effect is concerned, the wage level has a positive and the deflator negative impact on output. Looking from the cost side which emphasises profitability, the effect to expect is supposedly the opposite.

A simple price setting schedule links producer prices to the unit labour costs and prices of other inputs. If, in addition, we allow shifting backwards of indirect taxes, we end up with the relation

$$P = P\left(\frac{W(1+s)N}{Q},\ P_m,\ CPI/P\right). \tag{x}$$

The first two terms capture the constant mark-up feature in the price setting. The third term indicates whether producer prices adjust when consumer prices rise due to an increase in the value added tax, for instance. As this kind of price adjustment is a reaction to the effect of taxes on the real product demand, this term captures a demand side effect. Of course, product demand may also influence mark-ups directly which could generate a positive effect of Q instead of the negative effect apparent in (x)

To sum up, we expect the two well-specified relations, i.e. wage setting schedule and demand for labour condition, to show up. However, as the tests indicate that there are other long-run relations in the data set, we have additional vectors to investigate. We expect them to mimic 1) the supply of output and 2) the demand for output and 3) the constant mark-up pricing rule and 4) the price setting conducted by the product market demand conditions. The resulting "semirelations" may also be mixtures of two or more competing but misspecified relations. Hence, one should not put too much emphasis on the overinterpretation of the "left over" vectors.[12]

Our strategy below is as follows. Because the amount of possible joint hypotheses is very large, we start with partial analysis. As far as wage setting is concerned, we first test the two extreme structures indicating that one of the two parties – unions or employers – dominate. This is to investigate which is a better "initial" approximation of the process concerned. Whether deviations from the extreme structure are better in accordance with the data is then evaluated. Having found the preferred partially estimated relations, we test whether they are jointly accepted by the data. In this exercise, we take the numerical values of all the elasticities from the partially estimated vectors. These vectors are, in other words, considered as known ex ante. This is not to test whether the partially estimated vectors are identical to those which would result from a joint estimation. This is to see whether the partially discovered vectors are "too far" from the "appropriate" ones or not.

The final exercise is to estimate and test the relevant structure jointly for wage setting and the demand for labour. For the joint analysis, the preferred structure resulting from the partial analysis is the point of departure.

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[12] Some restrictions for these "semirelations" can be defined. In the constant mark-up pricing model $\beta_W = \beta_N = -\beta_Q = \beta_s > 0$ captures the unit labour cost restriction. If, additionally, homogeneity holds between unit labour costs and producer prices, we have $\beta_W = -\beta_P = \beta_N = -\beta_Q = \beta_s = 1$. The restriction $\beta_W = -\beta_P = \beta_N = -\beta_Q = \beta_{\tau_a} = -\beta_v$ generates the proxy for the houshold demand factor. This restriction is supposed to hold in the output demand relation but also in the price setting function allowing mark-up to vary with demand. In the relation concerned, real wages have positive output effect. In the output supply relation we expect that the amount of input (labour) is the driving force. A Cobb-Douglas technology would imply $\beta_N = 1$. A (permanent) shift in the relative price of energy, e.g., may additionally influence output, however. If income taxes and unions influence (cost) price setting only through wages, then $\beta_{\tau_a} = \beta_{\tau_m} = \beta_U = 0$ in the relevant vector. Because the present model only allows us to test these restrictions in the context of misspecified "semirelations", any tests as also related parameter estimates will be biased.

6 Testing structural hypotheses

6.1 Partial analysis

Table A4 introduces the four non-restricted cointegrating vectors. Interestingly, λ_4 is fairly small. This could indicate that instead of four cointegrating vectors the appropriate number would be three. Therefore, we evaluate the model not only under the assumption r=4 but under r=3 as well. Choosing a "too high" r implies that the tests imposed are "too loose". Consequently, if the correct choice is r=4 but we choose r=3, the tests are excessively stringent and the resulting p-values are definitively the low limits of the appropriate ones. Whether the fourth vector contains relevant information about the long-run relations of interest can also be evaluated by comparing the parameter estimates discovered including and excluding the fourth vector.

The β_i-coefficients indicate the long-run relationships embodied in the eigenvectors. To each cointegrating vector there is a corresponding vector of α's with at least one element, α_i, different from zero. These elements are the weights with which the cointegration relation enters the equation concerned. Hence, the α_i-coefficients embody the error correcting behaviour in the system.

In Table A4, some of the desirable properties can be detected. Additionally, the error correction property is quite clear in these freely estimated vectors.[13] However, the relation we are looking for can be a linear combination of the freely estimated vectors. So, mapping of the processes of interest requires more discipline to the analysis. This is where the theoretical considerations and the discussion on identifying restrictions are indispensable.

We proceed by imposing restrictions describing long-run properties which the vectors of interest are expected to fulfil. Short-term adjustment (to the changes of the process and to the long-run steady-states) is determined freely. However, as far as relevant cointegrating relations with behavioural interpretation will be detected, we expect the error correcting property to show up even more clearly.[14]

A restriction on β-coefficients is data consistent if the resulting eigenvalues do not change significantly as compared to the unrestricted estimation. Accordingly, each restricted vector is always compared to the original

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[13] This can be evaluated by multiplying each α_i with β_i. In the wage equation the disequilibrium error of relation (3) enters with weight .008·(−52.3) = −.40 which indicates rapid error correcting behaviour. In the employment equation, relation (2) enters with weight −.20. Finally, error correction is showing up in the price equation with respect to the first relation.

[14] The importance of this notion dates back to the statement of Granger (1986, p. 217) on the special relation between cointegration and error correction: "Not only must cointegrated variables obey such a model but the reverse is also true; data generated by an error-correction model ... must be cointegrated." Of course, this result is technical by nature, and the interpretability of the coefficients indicate the economic plausibility of the outcome.

unrestricted vector and all r eigenvalues contribute the test statistic[15] which follows the χ^2-distribution with degrees of freedom indicated in Table 2 which reports the results of the partial analysis.

6.1.1 Wage setting schedule

The necessary condition for unique identification is that the minimum number of restrictions is one less than the number of cointegrating vectors, r−1. In the present context this is three as we start by assuming that r=4. First, we evaluate the two extreme hypotheses concerning the dominance of the bargaining parties and the tax incidence augmented with restriction ($\beta_N = -\beta_Q$) which makes productivity the driving force of real wages. Under r=4 both of the contradicting hypotheses passed the test.[16] The hypothesis of union dominance (with taxes fully born by firms) generates a p-value of .34 whereas the opposite hypothesis has a p-value of .07. At face value, this could indicate that the former is better in accordance with the data. Under r=3 both extremes are, however, rejected. Therefore, we expect an intermediate case to outperform the extreme hypotheses. To proceed, we assumed that the restriction suggested by Padoa Schioppa (1991), $\beta_{\tau_a} = -\beta_{\tau_m}$, holds and that the coefficients of payroll tax and the price wedge are equal in absolute value, $\beta_s = -\beta_v$. These assumptions passed although on a fairly low significance level. As the general structure of the resulting vector was implausible, we preferred to relax the τ_P-restriction and, instead, to test whether the absolute value of the coefficients of ($1-\tau_m$), (1+s), (CPI/P) are equal in absolute value ($-\beta_{\tau_m} = \beta_s = -\beta_v$). The restriction was accepted but now the elasticity β_{τ_a} was above unity which can be considered as implausibly large. As can be seen from columns (1) & (2) in Table 2, β_{τ_a} can easily be restricted to have the value of unity. Here, the results appear not to be sensitive for the choice between r=4 or r=3.

The likelihood ratio test can also be used to evaluate a further issue of special importance for empirical identification. This is whether (in an acceptable structure) any of the coefficients which are close to zero are actually zeros. If in equation (2), for instance β_U is restricted to zero, the difference between the original test statistic (6.54) and the resulting (not reported) test statistic (12.70) follows $\chi^2(1)$ which gives 3.8 as the critical value on 5 per cent significance level. This test indicates that union density can not be omitted under r=3. The same concerns raw-material prices. Under r=4, these rejections do not emerge.

Finally, columns (3) & (4) indicate that the long run homogeneity between the real wage and productivity is rejected both under r=3 and under r=4 (the

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[15] The test statistic concerned is $T\sum_{1}^{r}\ln((1-\tilde{\lambda}_i)/(1-\hat{\lambda}_i))$ where $\hat{\lambda}_i(\tilde{\lambda}_i)$ is calculated without (with) the restrictions on β. Permanently, the H_0-hypothesis is that the restriction imposed is accurate (for details, see Johansen & Juselius, 1990).

[16] The fact that competing hypotheses pass the test may indicate that the surface of the likelihood function is so flat that it can not distinguish between hypotheses on the standard 5 % significance level. It may also be that we are testing a relatively small number of restrictions as compared to the number of parameters. As it happens, we choose to adjust the significance level upwards. This is in accordance with prefering a hypothesis which generates a higher p-value.

Table 2. **Partially identified cointegrating relations, under r = 4 and r = 3***

		wage setting schedule				labour demand schedule				output "semi-schedule"	price setting "semi-schedule"		
		(1) $r = 4$	(2) $r = 3$	(3) $r = 4$	(4) $r = 3$	(5) $r = 4$	(6) $r = 3$	(7) $r = 4$	(8) $r = 3$	(9) $r = 4$	(10) $r = 4$	(11) $r = 4$	(12) $r = 4$
coeff.	variables	long run coefficients β_i											
β_W	W	1.000	1.000	1.000	1.000	.237	.213	.210	.188	-1.361	-1.000	-1.000	-1.000
β_P	P	-1.139	-1.150	-1.054	-1.083	-.257	-.228	-.210	-.188	1.361	1.000	1.000	1.000
β_N	N	.721	.719	1.000	1.000	1.000	1.000	1.000	1.000	-1.000	-1.000	-1.000	-1.000
β_Q	Q	-.721	-.719	-1.000	-1.000	-.320	-.307	-.294	-.283	1.000	1.000	1.000	1.000
β_{τ_a}	$1-\tau_a$	1.000	1.000	1.000	1.000	.000	.000	.000	.000	-1.361	.000	.000	.000
β_{τ_m}	$1-\tau_m$	-.383	-.355	-.547	-.377	.000	.000	.000	.000	.000	.000	.000	.000
β_s	1+s	.383	.355	.547	.377	.237	.213	.210	.188	.000	-1.000	-1.000	-1.000
β_v	CPI/P	-.383	-.355	-.547	-.377	.000	.000	.000	.000	1.361	1.000	.000	.650
β_{P_m}	P_m	.139	.150	.054	.083	.019	.015	.000	.000	.000	-.030	-.043	-.034
β_U	UNION	-.300	-.317	-.008	.027	.000	.000	.000	.000	.000	.000	.000	.000
		adjustment coefficients α_i											
α_W	ΔW	-.368	-.431	-.381	-.164	-.410	-.601	-.367	-.569	.451	.283	.332	.305
α_P	ΔP	.167	.183	.150	.179	.313	.279	.309	.273	-.101	-.143	-.150	-.152
α_N	ΔN	-.033	-.027	-.046	-.045	-.253	-.227	-.261	-.231	.019	.072	.091	.081
α_Q	ΔQ	.423	.146	.371	.000	-.018	-.134	-.044	-.134	-.253	-.158	-.112	-.153
number of restrictions imposed (η)		5	5	6	6	6	6	7	7	8	8	8	8
characterization of restrictions imposed		$\beta_W = -\beta_P-\beta_{P_m} = \beta_{\tau_a}$, $\beta_N = -\beta_Q$, $\beta_{\tau_m} = -\beta_s = \beta_v$	$\beta_W = -\beta_P-\beta_{P_m} = \beta_{\tau_a}$, $\beta_N = -\beta_Q$, $\beta_{\tau_m} = -\beta_s = \beta_v$	$\beta_W = -\beta_P-\beta_{P_m} = \beta_N = -\beta_Q = -\beta_{\tau_a}$, $\beta_{\tau_m} = -\beta_s = \beta_v$	$\beta_W = -\beta_P-\beta_{P_m} = \beta_N = -\beta_Q = -\beta_{\tau_a}$, $\beta_{\tau_m} = -\beta_s = \beta_v$	$\beta_W = -\beta_P-\beta_{P_m} = \beta_s$, $\beta_{\tau_a} = \beta_{\tau_m} = \beta_v = \beta_U = 0$	$\beta_W = -\beta_P-\beta_{P_m} = \beta_s$, $\beta_{\tau_a} = \beta_{\tau_m} = \beta_v = \beta_U = 0$	$\beta_W = -\beta_P = \beta_s$, $\beta_{\tau_a} = \beta_{\tau_m} = \beta_v = \beta_{P_m} = \beta_U = 0$	$\beta_W = -\beta_P = \beta_s$, $\beta_{\tau_a} = \beta_{\tau_m} = \beta_v = \beta_{P_m} = \beta_U = 0$	$\beta_W = -\beta_P = \beta_{\tau_a} = -\beta_v$, $\beta_N = -\beta_Q$, $\beta_{\tau_m} = \beta_s = \beta_{P_m} = \beta_U = 0$	$\beta_W = -\beta_P = \beta_N = -\beta_Q = \beta_s = -\beta_v$, $\beta_{\tau_a} = \beta_{\tau_m} = \beta_U = 0$	$\beta_W = -\beta_P = \beta_N = -\beta_Q = \beta_s$, $\beta_{\tau_a} = \beta_{\tau_m} = \beta_v = \beta_U = 0$	$\beta_W = -\beta_P = \beta_N = -\beta_Q = \beta_s$, $\beta_{\tau_a} = \beta_{\tau_m} = \beta_U = 0$, $\beta_v = .65$
LM test statistic, $\chi^2(\eta-r_2)r_1$		3.89	6.54	7.93	15.34	3.12	6.47	4.99	8.08	5.22	7.15	7.65	6.91
critical value, $CV_{5\%}((\eta-r_2)r_1)$		5.99	7.82	7.82	9.49	7.82	9.49	9.49	11.07	11.07	11.07	11.07	11.07
p-value		.14	.09	.05	.00	.37	.17	.29	.15	.39	.21	.18	.23
eigenvalues, λ_i		.309	.311	.250	.246	.418	.432	.394	.410	.245	.246	.182	.221

*All variables are in logs, and r = number of cointegrating vectors, r_1 = number of cointegrating vectors to be restricted, r_2 = number of non-restricted cointegrating vectors = $r-r_1$

test is the one above). In fact, as the trend is supposed to capture the contribution of technical progress, a coefficient below unity for Q/N is what we expect to discover. The raw-material prices enter in the form of the relative price, P_m/P, which – when combined with the price homogeneity – probably removes the I(2)-ness as argued above.

6.1.2 Demand for labour

The demand for labour schedule we are looking for was defined above. The sole tax factor which enters (ii) directly is the payroll tax. Other taxes influence indirectly through wages. So, we expect that $\beta_{\tau_s} = \beta_{\tau_m} = \beta_v = 0$. A hypothesis that (in the long run) the firms operate on the labour demand curve imposes components of real labour cost to have equal effect, i.e., $\beta_W = -\beta_P = \beta_s$. The restrictions above as well as omission of union effects on employment are easily accepted as can been seen in columns (5) and (6). So, the demand for labour is guided by real labour cost and the level of activity. Columns (7) and (8) indicate that the effect of (real) raw-material prices differs not significantly from zero. The test discussed above gives the evidence despite the drop of the p-values. As a final experiment, we restricted the output elasticity to be equal to the labour cost effect. Under r=4 this restriction passes the test and generates a unit labour cost elasticity of .49. However, the p-value of .19 is considerably lower than the one related to vector (7). Under r=3 the restriction concerned is definitively rejected. This indicates that there is such a significant demand effect on employment which is in accordance with the hypothesis of imperfect competition on product market.

As a reference, Nickell & Symons (1990, p.1) conclude that as far as U.S. employment equations are concerned, "the present situation is one of total confusion" because negative, positive and zero wage elasticities have been discovered due to inappropriate specifications. In Finland as well the issue has been somewhat unsettled. In this perspective, our results are highly encouraging.

It is also interesting to evaluate the results in light of the recent contribution of Hamermesh (1991). Having surveyed a wide selection of studies concerning various countries, he argues that "a 'best guess' for the long-run constant-output labour-demand elasticity based on this literature is –0.30" (p. 5). Our results are fairly close to that. On the other hand, "with a value of labor's share in developed economies of around 0.70, this 'best guess' value ... is roughly consistent with Cobb-Douglas production". Finally, "labour and energy are p-(= price)substitutes, with a very small cross-price elasticity" (p. 6). So, the fact that we could not distinguish an impact for raw-material prices appears not to be in strong contradition with findings elsewhere.

6.1.3 Output / price setting "semischedules"

The various effects which may show up in the "left over" vectors were discussed in section 4. Under r=4 we discover some interesting results. First, relation (9) in Table 2 resembles a demand for output schedule. On the other hand, the hypothesis for the Cobb-Douglas technology, $\beta_Q = -\beta_N$ is also accepted. Hence, the vector concerned appears to capture both demand side and supply side elements. The relation is in accordance with Hamermesh (1991, p. 5) summarizing that "... the estimates based on both micro and aggregate data suggests the Cobb-Douglas function is a satisfactory way of describing aggregate production..."

As far as price setting is concerned, prices appear to be driven by unit labour costs. The results indicate that β_v is not precisely defined as the data accepts two contradicting hypotheses. According to the first, producer prices do adjust fully to an increase in sales tax ($\beta_v = 1$). According to the second, there is no adjustment at all ($\beta_v = 0$). This lead us to search by iteration that specific value of the adjustment coefficient which is most easily accepted by the data taken the rest of the structure as given. This is analogous to adjusting in a stepwise manner upwards the standard significance level of 5 %, ceteris paribus. The procedure concerned gives an elasticity in the range of $\beta_v = .6–.7$. A vector indicating this is well in accordance with the data and the p-value has risen to .23. This fragile evidence indicates that around 2/3 of a higher sales tax will be shifted backwards to lower producer prices.

The results above derive from tests carried out under the assumption r=4. Under r=3, all structures which mimic output schedule or price setting schedule are rejected. This indicates that for the "semirelations", the information cointained in the fourth vector is of crucial importance.

6.1.4 Summary of the partial analysis

We now test whether the data accepts the simultaneous coexistence of the partially identified vectors in Table 2. The results are highly encouraging. When the co-existence of the wage and employment relations characterized by restrictions in relations (1)&(7) is tested under the assumption r=4, the p-value is as high as .62. When we add the output schedule (9) the test is passed with a p-value .09. Next, we replaced the output schedule with the pricing rule (12). Now the three vectors were accepted with a p-value of .08. The coexistence of all the four vectors concerned is definitely rejected. Having indicated in section 4 that the composition of our data set does not allow us to identify a well-specified output schedule or pricing rule, this is as could be expected. Under r=3 the coexistence of wage and employment schedules (2) and (8) is accepted with a p-value .39.

Introduction of restrictions has led to a strengthening in the α-coefficients as compared to the non-restricted estimates. In the resulting structure the error correction property shows up very clearly. This is an indication of success in search for the fundamental long-run processes.

Of course, one can ask whether we correctly interpret the context of the various long-run relations and especially, are the vectors different enough that we can distinguish between the schedules suggested? The test concerning the formal identification will be discussed below (see Table A5). However, investigation of the weights in Table 2 also give insight on the matter of interest. Relation (2) which is considered as a wage setting schedule enters the equation for short-run adjustment of wages, ΔW, with a weight $\alpha_W = -.43$

indicating rapid error correction. Vector (8) considered as a labour demand schedule enters the dynamic equation, ΔN, with a high weight as well ($\alpha_N \sim -.23$). The rest of the weights, $\alpha_i \beta_i$, are all considerably smaller. As it happens, each of the two well-specified long-run relations with a special interpretation strongly contributes the short-run adjustment of the factor concerned. This surely does not indicate that interpretations suggested for these long-run relations would be arbitrary.

The partial nature of the analysis above reduces the credibility of the inference. Therefore, for better evidence we carry out the identification tests simultaneously for the two schedules of special interest, i.e. wage setting and the demand for labour.

6.2 Joint hypotheses

The final exercise is to estimate relations characterized with the relevant identifying restrictions jointly. As the wage setting and the demand for labour schedules appeared not to be sensitive for the choice between r=4 or r=3 in partial analysis, we prefer to take the risk of making the test excessively stringent by choosing r=3 in joint analysis. Afterwards, we investigate the data consistency of the resulting relations under r=4 as well.

Table 3 gives the results of the joint identification. There are five pairs of relations the first of which has its counterparts in relations (2) combined with either (6) or (8) in Table 2. In each pair, 5 restrictions have been imposed on the wage setting schedule and 6 or 7 restrictions on the labour demand curve. The joint hypothesis (A) with separate effect of real import prices on employment passes the test on a significance level of 11 per cent. Since β_{P_m} does not, however, differ significantly from zero in employment relation, hypothesis (B) is strictly preferable to (A) as far as empirical identification is concerned. The other pairs, (C)–(E), have been designed to evaluate whether the results are sensitive for the decision to consider tax rates as weakly exogenous. As discussed above, especially average income tax rate could well be endogenously generated in a process in which the union optimizes net wages taking as given the marginal tax scales defined by the authorities ex ante. In (C), the average income tax rate has been endogenized. Restrictions which are identical to those in (A) pass the test with a p-value of .19 and coefficients hardly change. In addition, β_{P_m} does not differ significantly from zero in the demand for labour schedule. So, earlier conclusions are valid also when the average income tax rate is considered as endogenous. In (E) not only average income tax but the marginal tax as well as the price wedge have been endogenized. The only considerable change concerns the magnitude of the wage elasticity of the "subwedge"-term. In addition, in this structure restricting β_{P_m} to zero in the labour demand relation is definitely rejected. The high p-value in (E) may be an artifact since as compared to (B) we have taken a considerable step towards our original overparametrized, full model. This is why we prefer the relations identified in (B) which have been shown to be robust also for considering τ_a as endogenously determined as in (D). For empirical identification it is of special interest that the LM test rejects restrictions of β_{P_m} and/or β_U to zero in all wage relations.

Table 3. **Jointly identified cointegrating relations, under r = 3, with weakly exogenous taxes in (A) and (B) and endogenous taxes in (C) – (E)***

		(A)		(B)		(C)		(D)		(E)	
		wage setting schedule	**labour demand schedule**	**wage setting schedule**	**labour demand schedule**	**wage setting schedule**	**labour demand schedule**	**wage setting schedule**	**labour demand schedule**	**wage setting schedule**	**labour demand schedule**
		(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
coeff.	variables	long run coefficients β_i									
β_W	W	1.000	.237	1.000	.199	1.000	.228	1.000	.181	1.000	.255
β_P	P	-1.147	-.257	-1.144	-.199	-1.147	-.247	-1.146	-.181	-1.168	-.281
β_N	N	.731	1.000	.714	1.000	.724	1.000	.706	1.000	.775	1.000
β_Q	Q	-.731	-.328	.714	-.293	-.724	-.319	-.706	-.277	-.775	-.349
β_{τ_a}	$1-\tau_a$	1.000	.000	1.000	.000	1.000	.000	1.000	.000	1.000	.000
β_{τ_m}	$1-\tau_m$	-.340	.000	-.346	.000	-.332	.000	-.345	.000	-.150	.000
β_s	1+s	.340	.237	.346	.199	.332	.228	.345	.181	.150	.255
β_v	CPI/P	-.340	.000	-.346	.000	-.332	.000	-.345	.000	-.150	.000
β_{P_m}	P_m	.147	.020	.144	.000	.147	.020	.146	.000	.168	.026
β_U	UNION	-.294	.000	-.320	.000	-.292	.000	-.322	.000	-.184	.000
		adjustment coefficients α_i									
α_W	ΔW	-.485	.068	-.492	.065	-.600	.291	-.582	.273	-.393	.073
α_P	ΔP	.053	.214	.090	.168	.198	.061	.222	.039	.782	-.405
α_N	ΔN	-.145	-.393	-.115	-.357	.150	-.336	.122	-.314	.163	-.304
α_Q	ΔQ	.453	-.735	.406	-.662	.479	-.623	.432	-.559	.514	-.495
α_{τ_a}	$\Delta\tau_a$	-	-	-	-	-.101	.194	-.086	.175	-.256	.466
α_{τ_m}	$\Delta\tau_m$	-	-	-	-	-	-	-	-	-.231	.489
α_v	Δ(CPI/P)	-	-	-	-	-	-	-	-	-.684	.520
number of restrictions imposed on relation i, (η_i)		5	6	5	7	5	6	5	7	5	6
characterization of restrictions imposed		$\beta_W = -\beta_P-\beta_{P_m} = \beta_{\tau_a}$, $\beta_N = -\beta_Q$, $\beta_{\tau_m} = -\beta_s = \beta_v$	$\beta_W = -\beta_P - \beta_{P_m} = \beta_s$, $\beta_{\tau_a} = \beta_{\tau_m} = \beta_v = \beta_U = 0$	$\beta_W = -\beta_P-\beta_{P_m} = \beta_{\tau_a}$, $\beta_N = -\beta_Q$, $\beta_{\tau_m} = -\beta_s = \beta_v$	$\beta_W = -\beta_P = \beta_s$, $\beta_{P_m} = \beta_{\tau_a} = \beta_{\tau_m} = -\beta_v = \beta_U = 0$	$\beta_W = -\beta_P - \beta_{P_m} = \beta_{\tau_a}$, $\beta_N = -\beta_Q$, $\beta_{\tau_m} = -\beta_s = \beta_v$	$\beta_W = -\beta_P - \beta_{P_m} = \beta_s$, $\beta_{\tau_a} = \beta_{\tau_m} = \beta_v = \beta_U = 0$	$\beta_W = -\beta_P - \beta_{P_m} = \beta_{\tau_a}$, $\beta_N = -\beta_Q$, $\beta_{\tau_m} = -\beta_s = \beta_v$	$\beta_W = -\beta_P = \beta_s$, $\beta_{P_m} = \beta_{\tau_a} = \beta_{\tau_m} = \beta_v = \beta_U = 0$	$\beta_W = -\beta_P - \beta_{P_m} = \beta_{\tau_a}$, $\beta_N = -\beta_Q$, $\beta_{\tau_m} = -\beta_s = \beta_v$	$\beta_W = -\beta_P - \beta_{P_m} = \beta_s$, $\beta_{\tau_a} = \beta_{\tau_m} = \beta_v = \beta_U = 0$
LM test statistic, $\chi^2\,(\Sigma(\eta_i-(r-1))$		11.69		14.18		10.00		12.84		7.01	
critical value, $CV_{5\%}\,(\Sigma(\eta_i-(r-1))$		14.07		15.51		14.07		15.51		14.07	
p-value		.11		.08		.19		.12		.43	
eigenvalues, λ_i		.511	.411	.511	.411	.531	.490	.531	.490	.583	.518

*In (C) and (D) the average income tax rate, $1-\tau_a$, has been considered as endogenous. In (E) both the average and the marginal income tax rates, $1-\tau_a$ and $1-\tau_m$, as well as the price wedge, CPI/P, have been endogenized. All variables are in logs, and r = number of cointegrating vectors, r_1 = number of cointegrating vectors to be restricted, r_2 = number of non-restricted cointegrating vectors = $r-r_1$

It is noteworthy, that the third ("leftover") relation does not influence wage setting or demand for labour equations introduced in the Table. The relevant α-coefficient is close to zero (~ .02) in both equations. Hence, the fact that we cannot give a plausible interpretation to the third relation hardly reduces the credibility of identification of the relations of interest. Under r=4, the co-existence of the pairs of relations which are numerically identical to those reported as pairs (A), (B), (C), (D) and (E) in Table 3, pass the test with a p-value of .83, .65, .82, .51 and .29 correspondingly. Again, the first four which are almost identical show up as the strongest candidates.

As a summary, the preferred cointegrating relations resulting from the joint identification can be written as follows:

$$\log W = 1.00\cdot\log P + .71\cdot\log(Q/N) - 1.00\cdot\log(1-\tau_a) + .35\cdot\log\left(\frac{(1-\tau_m)(CPI/P)}{(1+s)}\right) - .14\cdot\log(P_m/P) + .32\cdot\log(UNION)$$

$$\log N = -.20\cdot\log\left(\frac{W(1+s)}{P}\right) + .29\cdot\log(Q)$$

These jointly identified wage and employment relations are almost identical to the partially identified relations reported in Table 2. Finally, the formal identification of the structure present in the preferred relations is to be checked. This is consistent with testing that no linear combination of β_j's with $i\neq j$ can produce a vector that "looks like" β_i. The rank condition for formal identification is defined in Johansen & Juselius (1992b). In the present context it is satisfied as far as the wage setting and demand for labour schedules of interest are concerned. The result which shows that there are several overidentifying restrictions can be seen in Table A5 in Appendix.

7 Wages and taxes: Discussion

One of the interesting novelties in the present study consideres the opposite effects discovered for the average and marginal income tax rates. The signs are in accordance with evaluations in Tyrväinen (1992b). As the elasticity related to the former is unity and the opposite effect to the latter is around .35, the net impact of an equiproportional increase in average and marginal income tax rates on wages is positive and approximately .65.

Table 4 contains the wage elasticities introduced in some other studies concerning Finland. There are four notions of special interest. First, the parameter estimates vary a lot. Second, the real outlier is still Calmfors & Nymoen (1990). Third, Eriksson et al. (1990) discover a zero effect of the average income tax and a positive effect of the marginal income tax rate. In their estimations which suffer from small number of observations, the effects of the two tax factors which behave quite similarly (see Figure 1) appear to be mixed. This conjecture is supported by the fact that in Tyrväinen (1992a) with only the marginal rate entering, this factor captures the role of the average rate which plays the dominant role in the present estimations. Fourth, the rest of the elasticities discovered in Tyrväinen (1992a) applying the Granger & Engle procedure (with OLS) are quite close to those discovered in the present context.

Interestingly, Padoa Schioppa (1990) discovers for Italy a positive unitary elasticity of the average tax rate and a considerable negative effect of progressivity which is very close to our results. Manning (1992) reports the effects of +.3 and –.3 respectively after having imposed them to be equal in absolute value according to the τ_p-restriction discussed above.

As can be seen from Figure 1, the average income tax has risen around 10 per centage points between 1965 and 1990 in Finland. The marginal rate increased by around 15 per centage points. The coefficients discovered in the present study indicate that this would have generated a compensating increase in the pretax nominal real wage level of close to 5 per cent.

Table 4. **The Wage effects of Taxes derived from some studies.***

	Payroll tax	Price wedge (incl. VAT)	Income tax, average	Income tax, marginal
This study	–.3	.3	1.0	–.3
H–H–K	–.4	.6	.6	..
B–L–N	–.7	.2	.2	..
P	–.7	.9	1.0	..
E–S–V	–1	.9	0	.3
C–N	–1	0	0	..

* The studies concerned are H–H–K: Holm, Honkapohja & Koskela (1991), E–S–V: Eriksson, Suvanto & Vartia (1990), P: Pehkonen (1990), C–N: Calmfors & Nymoen (1990), B–L–N: Bean, Layard & Nickell (1986).

8 Conclusions

As an outcome of modelling wage setting and demand for labour in a bargaining framework, a 10-dimensional vector space was defined. Structural restrictions which identify the long-run relations of interest were specified and tested first in partial and then in joint analysis.

Generally, the restrictions which satisfy the condition for formal identification pass the tests on a fairly high significance level. The plausibility of resulting cointegrating relations applies not only to the signs but to the magnitudes of the coefficients as well. This concerns the economic identification. Testing empirical identification is not straightforward for the time being. However, the preliminary results applying the LM test indicate that the preferred relations also satisfy this condition.

The relations show up almost identically in partial and joint analysis. They appear not to be sensitive for the choice between r=4 or r=3 which is the factor defining the number of cointegrating relations in the system. Finally, the relations are hardly influenced by alternative conjectures concerning the endogeneity of various tax rates. Therefore, we consider these long-run relations as robust.

According to the results, in the private sector in Finland neither of the bargaining parties – employers or employees – has gained full dominance in the wage setting process. However, the coefficient estimates indicate that the adjustment in the long run has more closely reflected union goals. An equiproportional increase in average and marginal income tax rates has been shifted to a higher pretax wage level with an elasticity of around two thirds. Interestingly, the adjustment coefficient related to other taxes is considerably smaller, around one third or slightly less. Stronger unions have pushed up the wage level. The significant role of the union density variable is in accordance with bargaining models. Higher real prices of raw-materials have reduced wages. The driving force of the real wage growth is the productivity growth. As far as demand for labour is concerned, the negative impact of real labour cost is clear-cut. The additional output effect is an indication of imperfect competition in the product market. The (negative) effect of higher raw-material prices is insignificant.

Table A1. **A joint Trace test for the cointegration rank, r, and the existence of a linear trend, H_r^* versus H_r.**

r	No linear trend, H_r^*			Linear trend, H_r		
	Eigen-value, λ_i	Test statistic, Q_r^*	Critical value, $CV_{95\%}$	Eigen-value, λ_i	Test statistic, Q_r	Critical value, $CV_{95\%}$
0	.686	397	244	.628	351	233
1	.555	302	203	.555	270	193
2	.536	236	166	.526	203	156
3	.444	174	132	.443	142	124
4	.399	126	102	.316	93	94
5	.316	84	76	.285	63	68
6	.285	53	53	.188	35	47
7	.174	25	35	.174	18	30
8	.108	10	20	.028	3	15
9	.003	.2	9	.003	.2	4

The critical values are from Johansen & Juselius (1990, Tables A1 and A3) and Osterwald-Lenum (1990).

Table A2. **A Trace test for I(2) common trends***

Number of		Test statistic, Q_{4,c_1}	Critical value, $CV_{95\%}$
I(1) common trends, c_1	I(2) common trends, $c_2 = c-c_1$		
0	6	426	102
1	5	248	76
2	4	162	53
3	3	92	35
4	2	30	20
5	1	8	9

* In a 10-dimensional system with r=4, the number of common trends is c = 10–4 = 6.

Table A3. **Residual analysis.**

A. The full model

Equation	B-P.Q(20)/18 (1)	ARCH(2) (2)	SKEW. (3)	EX.KURT. (4)	J-B.NORM. (5)
ΔW	1.030	1.783	-.367	-.112	1.887
ΔP	1.666	1.185	.112	.458	.889
ΔQ	.919	.034	-.248	.466	1.579
ΔN	1.117	4.703	.015	.012	.004
$\Delta(1-\tau_a)$	1.642	9.508	.189	2.104	15.612
$\Delta(1-\tau_m)$	1.517	18.161	-.129	6.191	131.188
$\Delta(1+s)$	1.083	.638	.783	5.061	95.887
$\Delta(CPI/P)$	1.603	.943	-.142	.155	.357
ΔP_m	.707	.167	1.217	4.999	105.616
$\Delta UNION$	1.330	4.701	-.317	.621	2.690

B. The partial model

Equation	B-P.Q(20)/18 (1)	ARCH(2) (2)	SKEW. (3)	EX.KURT. (4)	J-B.NORM. (5)
ΔW	1.252	3.997	-.338	-.132	1.621
ΔP	1.036	1.848	.306	.117	1.327
ΔN	.826	4.610	.048	-.446	.709
ΔQ	.919	1.516	.080	-.095	.119

All variables are in logs. The first column introduces a scaled Box-Pierce autocorrelation test statistic. The second column is the test statistic for conditional heteroscedasticity which follows the χ^2-distribution with degrees of freedom indicated in parenthesis. Third column measures skewness and the fourth measures excess kurtosis-3 (these two statistics are zeros with the normal distribution). Finally, the fifth column introduces a test statistic for normality of the residuals (derived from columns three and four) which is distributed $\chi^2(2)$.

Table A4. **Unrestricted cointegrating relations**

		(1)	(2)	(3)	(4)
coeff.	variables	long run coefficients β_i			
β_W	W	40.340	-20.704	-52.284	16.496
β_P	P	-66.575	-9.973	61.414	-19.257
β_N	N	68.195	-99.963	24.934	-41.835
β_Q	Q	-8.002	59.477	19.693	-8.841
β_{τ_a}	$1-\tau_a$	91.366	-11.813	-170.146	-149.500
β_{τ_m}	$1-\tau_m$	-32.449	-10.599	48.410	47.316
β_s	1+s	73.878	-56.979	-4.876	-139.612
β_v	CPI/P	-42.120	4.973	50.275	-1.709
β_{P_m}	P_m	10.951	10.598	-5.005	5.893
β_U	UNION	-6.729	20.070	11.995	11.168
		adjustment coefficients α_i			
α_W	ΔW	-.001	.004	.008	-.004
α_P	ΔP	.004	.000	.001	.000
α_N	ΔN	.000	.002	-.001	.000
α_Q	ΔQ	.000	.001	-.005	-.002
eigenvalues, λ_i		.511	.411	.335	.133

All variables are in logs.

Table A5. **Verification of the rank condition for formal identification in the long-run structures (A)-(E) in Table 3.***

(A) = (C) = (E)

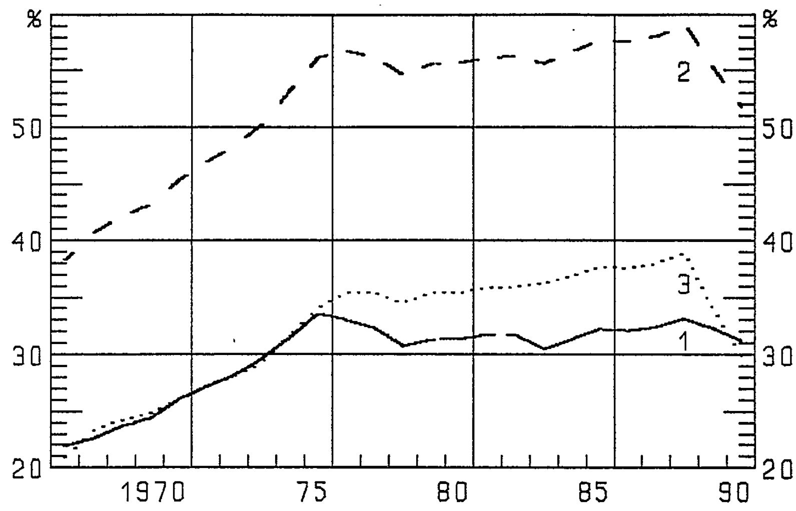
	$H_{i.j}$		$H_{i.jk}$
1.2:	2	1.23:	5
1.3:	5		
2.1:	4	2.13:	6
2.3:	7		

(B) = (D)

	$H_{i.j}$		$H_{i.jk}$
1.2:	2	1.23:	5
1.3:	5		
2.1:	4	2.13:	6
2.3:	7		

*For the method, see Johansen & Juselius (1992b). Formal identification requires that the value of each $H_{i,j}$ is greater than 1 and each $H_{i,jk}$ exceeds 2. The statistics reported have been kindly calculated by Katarina Juselius.

Figure 1. **Income tax rates in Finland, per cent**



1 Average tax rate, τ_a
2 Marginal tax rate, τ_m
3 Progressivity index, $\tau_P = (\tau_m - \tau_a)/(1 - \tau_a)$

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