BANK OF FINLAND DISCUSSION PAPERS

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Economics Department

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Wage Determination in the Long Run, Real Wage Resistance and Unemployment:

Multivariate Analysis of Cointegrating Relations in 10 OECD Economies

SUOMEN PANKIN KESKUSTELUALOITTEITA • FINLANDS BANKS DISKUSSIONSUNDERLAG

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ISBN 951-686-451-1
ISSN 0785-3572

Suomen Pankin monistuskeskus
Helsinki 1995

Summary*

Over the past twenty years or so, unemployment has been increasing in most OECD economies. In the same period, there has been a considerable increase in the wedge between the real cost to the employer of hiring a worker and the net real wage received by the worker. The present study examines whether changes in the wedge (including various tax rates) may have generated long-lasting effects on real labour costs. Behaviour which generates this kind of outcome is called "real wage resistance". If there is real wage resistance, higher taxes lead to higher unemployment. If this outcome persists in the long run, the primary problem related to the functioning of the wage setting mechanism is not necessarily the speed of adjustment but rather the equilibrium in which adjustment terminates.

The countries examined are the United States, Japan, Germany, France, the United Kingdom, Italy, Canada, Australia, Sweden and Finland. The study covers the private business sector and the estimation method applied is the procedure proposed in Johansen & Juselius (1990) for estimation of long-run relationships.

Signs of real wage resistance were discovered in all the economies examined although it differs in degree between countries. Although this can to some extent be related to characteristics of wage setting mechanisms in the countries concerned, there is not a simple story to tell. The outcome depends both on labour market characteristics and on the actual development of the wedge.

As far as the effect of actual changes in the wedge on actual changes in unemployment are concerned, the most unfavourable case is that found in France: not only considerable degree of real wage resistance but also a large rise in the wedge was detected. Between the mid-70's and early 90's, the impact of real wage resistance on the unemployment rate is also important in Australia, Canada, Finland and Sweden. In the latter two countries in particular this is due to recent increases in the wedge.

When the contribution of taxes is separated it becomes clear that, in Canada unfavourable impacts of real wage resistance are not related to taxation. In Japan, taxes are not a primary cause either whereas in France, Italy and Sweden taxes have played a major role. In Germany, appreciation of the exchange rate has "created room" for consumption taxes to rise without the harmful effect on consumer prices which would have generated wage claims. In Finland, large tax increases have taken place in recent years. In the other countries, the effect of taxes on labour cost is in the range of ± 2 per cent.

For four countries (Germany, France, Canada and Finland), effects of three alternative revenue-neutral shifts in taxation were simulated. A shift from income tax to payroll tax has no long-run impact but in the short run (which lasts for five years or more) the effect on (un)employment is unfavourable. A shift from taxes on income to taxes on consumption has a favourable impact on (un)employment, not only in the short run but also in the long run. Given current budget constraints, a precondition for this policy option would seem to be that the government allows the real value of such non-wage incomes as pensions, unemployment benefits and social transfers to fall as a result of the increase in the consumption tax.

* This paper has been prepared at the OECD as part of *the OECD Jobs Study*. Its methods and results have been highlighted in a more compact report "Real Wage Resistance and Unemployment: Multivariate Analysis of Cointegrating Relationships in 10 OECD Economies" (Tyrväinen, 1995a).

Tiivistelmä**

Viimeisten kahdenkymmenen vuoden aikana työttömyys on lisääntynyt useimmissa OECD-maissa. Samaan aikaan on kasvanut "kiila", joka on työnantajan maksaman reaalisen työvoimakustannuksen ja työntekijän reaalisen, käteen jäävän (netto)palkan välillä. Tässä tutkimuksessa arvioidaan onko kiilan — johon sisältyy mm. erilaisia veroja — muutoksilla ollut pitkäaikaisia vaikutuksia työvoimakustannuksiin. Jos näin olisi, korkeammat verot olisivat lisänneet työttömyyttä. Jos työttömyyden nousu näyttää jäävän pysyväksi, ensisijainen palkanmuodostusmekanismiin liittyvä ongelma ei ehkä olekaan sopeutumisen hitaus. Ongelma voikin olla sen tasapainon luonne, jossa palkkasopeutuminen pysähtyy. Tällaisessa tasapainossa korkeakaan työttömyys ei hillitse palkkojen nousua.

Tutkimuksen kohteena ovat Yhdysvallat, Japani, Saksa, Ranska, Englanti, Italia, Kanada, Australia, Ruotsi ja Suomi. Tarkastelussa on yksityinen yrityssektori. Estimoinnit on suoritettu ns. Johansenin menetelmällä, joka on erityisen sopiva pitkän aikavälin vaikutusten analysointiin.

Kaikissa tutkituissa maissa näkyi merkkejä siitä, että kiilan muutokset ovat vaikuttaneet työvoimakustannuksiin — joskin vaikutuksen voimakkuus on erilaista. Vaikka maittaiset erot näyttävät ainakin jossakin määrin heijastavan palkanmuodostusmekanismin erilaisuutta, liiallista yksinkertaistusta on syytä varoa. Toteutuneet vaikutukset riippuvat sekä työmarkkinoiden erityispiirteistä että kiilan (ts. myös verotuksen) todellisesta kehityksestä.

Kun arvioidaan kiilan toteutuneiden muutosten vaikutusta työttömyyteen, Ranskan tapaus näyttää kaikkein epäsuotuisimmalta. Ranskassa a) kiilan muutoksen palkkavaikutus on merkittävä ja b) kiila on kasvanut paljon. 1970-luvun puolivälin ja 1990-luvun alun välillä kiilan kasvun vaikutus työttömyyteen on merkittävä myös Australiassa, Kanadassa, Suomessa ja Ruotsissa. Erityisesti kahdessa viimeksi mainitussa tämä tulos kertoo viime vuosien kehityksestä.

Kun tutkitaan erikseen nimenomaan verotuksen roolia työttömyyden kasvussa, se ei ole ollut kovin suuri Kanadassa. Myöskään Japanissa se ei ole ollut olennainen, kun taas Ranskassa, Italiassa ja Ruotsissa verojen rooli on ollut keskeinen. Suomessa verotus on kiristynyt merkittävästi viime vuosina. Saksassa valuuttakurssin vahvistuminen on "luonut" kulutusveroille nousutilaa. Kun tuontihintojen aleneminen kompensoi verotuksen kiristymisen hintavaikutukset, palkkapaineita ei synny. Muissa maissa verotuksen vaikutus työvoimakustannuksiin on ollut ± 2 prosenttia.

Neljän maan (Saksa, Ranska, Kanada ja Suomi) osalta tutkimuksessa arvioidaan sellaisten verorakenteen muutosten vaikutuksia, jotka jättävät valtion verotulot ennalleen. Tuloverotuksen keventäminen työnantajain sotu-maksuja korottamalla ei vaikuta pitkän aikavälin tulemiin, mutta lyhyellä aikavälillä (joka kestää noin 5 vuotta) työllisyys heikkenee. Tuloverotuksen alentaminen arvonlisäverotusta nostamalla on työllisyyden kannalta myönteistä sekä pitkällä että lyhyellä aikavälillä. Ehtona on kuitenkin se, että ei-palkkatulojen (eläkkeet, työttömyyskorvaukset ja sosiaaliavustukset) reaaliarvon annetaan alentua kulutusverotuksen kiristyessä.

Suomen osalta tutkimuksen tärkeimmät tulokset ovat seuraavat. Verotuksen kiristyminen on heikentänyt työllisyyttä, koska verot ovat johtaneet työvoimakustannusten kallistumiseen. Noin puolet verojen noususta on jäänyt työvoimakustannuksiin ja vastaavasti noin puolet verotuksen kiristymisestä on näkynyt vähennyksenä palkansaajien käteen jäävässä reaalipalkassa.

Koska verotuksen kiristyminen on ollut erityisen voimakasta viime vuosina ja koska sopeutumisviiveet ovat melko pitkät, verotuksen haittavaikutuksia lienee vielä tulevina vuosina lisäämässä palkkapaineita ja heikentämässä työllisyysnäkymiä.

Tutkimuksen mukaan siirtyminen yhdestä työntekoon kohdistuvasta veromuodosta toiseen ei ole pitkän aikavälin ratkaisu. Jos halutaan vähentää verotuksen haitallisia vaikutuksia työttömyyteen, tulee alentaa työntekoon kohdistuvan verotuksen kokonaistasoa.

** Tämä tutkimus on tehty osana OECD:n Työllisyystutkimusta. Sen menetelmiä ja tuloksia esitellään lyhyemmin raportissa "Real Wage Resistance and Unemployment: Multivariate Analysis of Cointegrating Relationships in 10 OECD Economies", *The OECD Jobs Study: Working Paper Series.*

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1 Introduction[1]

Over the past twenty years or so, unemployment has risen in most OECD countries. In the same period, many countries have experienced an increase in the "wedge" between the real labour cost paid by the firm and the real take-home pay received by the employee. If wages do not fully absorb changes in wedge factors (including various tax rates), real wage resistance exists. For the present study, the key question is how the real labour cost responds to changes in elements of the wedge because "if there is such a positive response, then the wedge will influence the equilibrium unemployment" (Layard, Nickell & Jackman, 1991, p. 210).

The paper is organized as follows. Section 2 introduces the economic model and evaluates shifts in equilibrium unemployment. The concept of real wage resistance is discussed as well as earlier evidence. Section 3 defines an unrestricted VAR model derived from the theoretical considerations and discusses the maximum likelihood method proposed in Johansen & Juselius (1990) for estimation of long-run relationships in multivariate systems. The final operational model is arrived at by reducing the full model by conditioning, i.e. by considering part of the variables as weakly exogenous whenever a test allows it. In section 4, the identifying restrictions of the structural hypotheses are specified and tested in the partial model. In section 5 the effect of taxes on real labour costs via real wage resistance is calculated using the estimated elasticities in conjunction with the actual tax data. Section 6 presents simulations in which effects of actual changes in wedge factors are examined using three-equation models specified for each country in Appendix 2. Section 7 summarizes the paper.

2 The Theoretical Model

In a world in which perfect competition prevails, wages adjust to whatever level needed to clear the labour market. Accordingly, real labour costs deviate only temporarily from the level of the labour productivity and all changes in unemployment can only be considered as variations around a "natural" rate of unemployment. This should find its confirmation in the long-run stationarity of unemployment rates. As indicated by Layard, Nickell & Jackman (1991, below: LNJ), in a perspective of one hundred years or so, this seems like a plausible description of the history of unemployment.

In the literature, it has been conventional to present models in which higher unemployment leads to wage moderation. In Phillips curve models, the relation is typically between the (real) wage *change* and the unemployment rate. This specification makes sense as a description of the short-run interaction in a

[1] This paper has benefitted from useful comments and suggestions by Sven Blondal, David Grubb, Jorgen Elmeskov, Steve Englander, David Hendry, Mark Pearson, Stephan Thurman and Dave Turner. Helpful discussions with Katarina Juselius, Søren Johansen and Antti Ripatti also are gratefully acknowledged. Needless to say, the usual disclaimer applies.

world in which unemployment is stationary, i.e. when it fluctuates around a mean and returns to the mean value often. However, if the unemployment rate is non-stationary, a Phillips curve type of relationship is not necessarily consistent with the time-series properties of the data.

Over the past twenty years or so, persistent shifts in the ("natural") unemployment rate seem to have taken place in almost all industrialized countries[2]. This challenges standard models in many respects and leads us to search for an explanation which derives from imperfect competition in labour markets.

Wage setting is considered in Figure 1. On the vertical axis is the real labour cost which we here simply define as the wage, W. On the horizontal axis is the labour force, L, or alternatively employment, N. Unemployment is the difference between L and N, i.e. u = L–N.

With a given labour demand schedule, LDS, a shift in the wage setting schedule from WSS_0 to WSS_1 induces a change in the equilibrium relation (W^*,u^*) from point A to point B. The rise in equilibrium wage is $W^*_1 - W^*_0$ and the rise in equilibrium unemployment is $u^*_1 - u^*_0$. The new equilibrium prevails when all interaction between wages, employment and unemployment has taken place. Below we examine whether persistent shifts in the wage setting schedules could be due to real wage resistance[3].

To consider an imperfectly competitive labour market, we apply a bargaining model common from literature. No distinction is made between bargaining at plant, firm or industry level and country wide negotiation. The decisive feature is the imperfect competition embodied in collective contracts.

There are n identical firms which have a production function Q = AF(N) with one input, labour (N). A is a productivity index. Imperfect competition prevails in the product market. The firm maximizes profits which are defined as the difference between sales revenues and production costs:

$$\pi = \hat{P}[\tilde{Z}\,AF(N)]\,AF(N) - W(1+s)N \tag{i}$$

where $\hat{Q}^d = \hat{P}^{-1}(P)\tilde{Z}^{-1} \equiv D(P)Z$ is a downward sloping demand curve of the separable form introduced by Nickell (1978). Here, $Z = \tilde{Z}^{-1}$ 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 represents the competitors' producer price, W the nominal wage, s the employers' social security contribution rate. The output of the firm, $\hat{Q}$, is considered endogenous.

Employers bargain with representatives of workers. The welfare of the latter depends on the after tax real wage of employed and the (real) unemployment benefit received by the unemployed, $V = V(W(1-\tau)/P_c,N,B)$ where P_c represents the consumer price, τ the income tax, and B the

[2] A stationarity test included in the program for Cointegration Analysis of Time Series (CATS) in RATS rejects stationarity of the unemployment rate in all countries concerned (in estimation periods applied). Katarina Juselius has kindly supplied this software for estimation.

[3] Manning (1992b) defines a model which allows for multiple equilibria of unemployment. His empirical model — which also includes wedge factors — suggests that the British economy may have moved from a "low unemployment equilibrium" to a "high unemployment equilibrium" although no economic fundamentals changed.

replacement ratio (unemployment benefit in relation to the relevant wage level). As far as the partial derivatives are concerned, we assume that $V_1', V_2', V_3' > 0$ and $V_1'', V_2'', V_3'' < 0$ respectively. This general specification covers most of the common preference functions.

The widely used bargaining models differ as regards the factors which are assumed to be bargained over. In the "right-to-manage" model, wages are bargained over and the profit maximizing firm sets employment unilaterally.

Let us specify the game as a standard Nash solution of a cooperative game after Binmore et al. (1986):

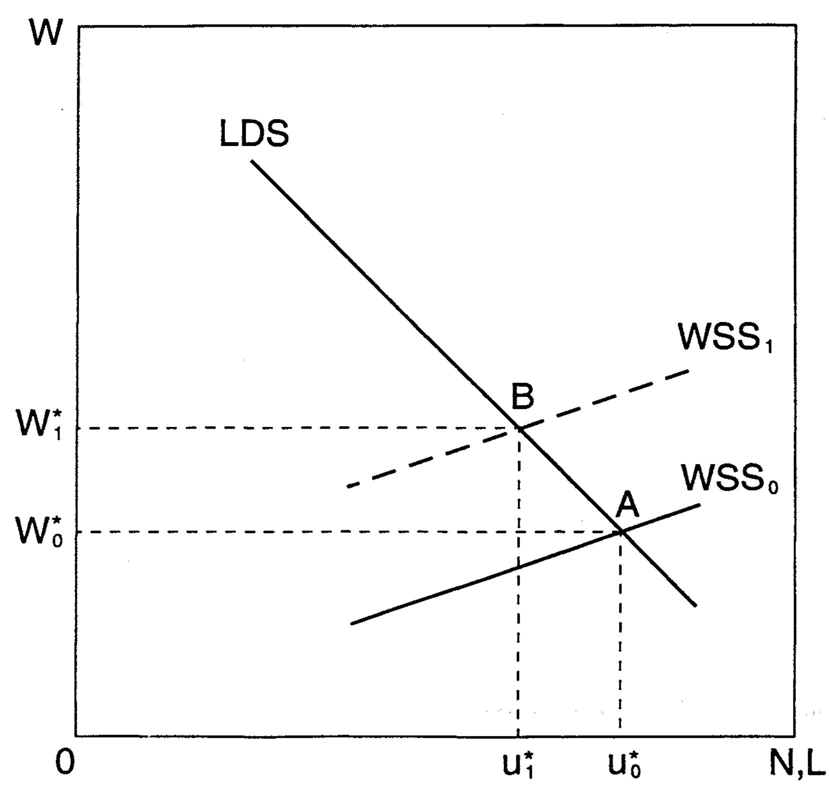
$$\max_{W}(V - V_0)^{\theta}(\pi - \pi_0)^{1-\theta} \qquad \text{s.t.} \quad N(.) = \arg\max_{N} \pi \tag{ii}$$

where θ refers to the bargaining power of the employees, $0 < \theta < 1$. If θ is either zero or unity, the wage level is not subject to bargaining. If θ is zero, the firm defines the wage level unilaterally. If it is unity, the wage is set by the union.

Bargaining power is an unobservable variable which probably depends positively on the unionization rate and negatively on the unemployment rate. Data on union density are, unfortunately, not available as time series of sufficient length. So, we assumed simply that employees' bargaining power is weaker when the unemployment rate, u, is higher (when all other factors are given) and *vice versa*.

Figure 1.

Wage setting, demand for labour and equilibrium unemployment



V_0 is the fall-back utility of the workers in the event an agreement is not reached. The alternative income in this case could be the unemployment benefit, UB, or a strike allowance, SA. 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.

If the trade-offs incorporated in (ii) represent long-run targets of the social partners — a plausible presumption — it is natural to consider the solution as an equilibrium relationship which refers to the long run. The resulting model for the equilibrium (real) wage consists of variables influencing profits, on the one hand, and on the other, the utility of the employees. In addition, a role is played by determinants of the fall-back utilities of the parties. Finally, bargaining power matters. In its most general form the wage setting schedule is

$$W^* = W(\underset{+}{P}, \underset{-}{s}, \underset{+}{\tau}, \underset{+}{\frac{P_c}{P}}, \underset{-}{u}, \underset{+}{Z}, \underset{+}{A}, \underset{+}{UB}, \underset{+}{SA}). \qquad \text{(iii)}$$

All signs in (iii) are according to evaluations in Tyrväinen (1995b). Although we have stressed above the bargaining aspects in modelling, discrimination between bargaining models and other models is not straightforward. For instance, market clearing models can be specified so that they produce schedules which are very much like those in this paper. McKee et al. (1986), e.g., derive a role for the wedge in a set-up in which labour supply depends on taxes. However, when the wedge variables enter as determinants of union behaviour we think that persistent effects could be more probable than when they enter as determinants of labour supply of individuals. On the other hand, there are studies (e.g. Calmfors & Driffil, 1988) which seem to suggest that if wage setting is sufficiently centralized real wage resistance would not necessarily be strong either.

In the empirical part of the paper, an unrestricted VAR model is first estimated. In this set-up, significant presence of each tax variable is tested. By including tax variables into the theoretical model we simply allow the significance of these variables to be tested — nothing more. We believe that this is a more appropriate way to proceed than to exclude certain variables *a priori.*

Series included in (iii) tend to be non-stationary over the observation period. This leads to well-known problems if standard estimation methods are used. As a result of introduction of the concept of cointegration, Engle & Granger (1987) proposed a two-step method for estimation of the long-run relationships between non-stationary variables. In the present study, we use the maximum likelihood procedure introduced in Johansen (1991b) for estimation of multivariate systems. As the two-step method only picks one potential candidate for the relevant long-run relationship with no consideration of the others, the Johansen method allows the vector space to be examined in a more thorough manner, i.e. it allows

- to make (in the estimation period) an explicit distinction between the time-invariant relationships and unstable relationships;
- to analyse simultaneously several cointegrating vectors;

- to make a distinction between long-run relationships and short-run dynamics and to estimate all related parameters simultaneously;
- to test hypotheses and discuss identification in a straightforward manner (see Johansen & Juselius, 1992, 1994b).

If data support the existence of a time-invariant long-run relation like (iii), a cointegrating vector has been discovered. This vector acts as an attractor[4] which incorporates an equilibrium relation between the wage level, unemployment and the rest of the variables. The decisive property of an attractor implies that if the wage is on it, there is no incentive for the wage to change. A shift to the new equilibrium B in Figure 1 represents an unfavourable shift in the attractor. Because B is an equilibrium, unemployment which exceeds an earlier record does not generate wage adjustment.

A subset of the variables in (iii) sum up to "WEDGE" which consists of taxes with an additional contribution coming from relative import prices influencing the price wedge, P_c/P. Let variable X summarize the rest of the variables in (iii) including the productivity variable. If long-run homogeneity between wages and prices is assumed to hold, the relation of interest looks like

$$(W/P)^* = W(\underset{-}{u}, \underset{+}{\mathrm{WEDGE}}, X) \tag{iv}$$

In the context of the Johansen method, attempts have been made to avoid all *a priori* structures which would bias the estimation in either finding or rejecting wedge effects in (iv). *If* the data indicate that both the WEDGE and the unemployment rate enter a cointegrating vector like (iv), equilibrium unemployment will be influenced by (exogenous) changes in the wedge. Furthermore, *if* an increase in the WEDGE takes place, then the equilibrium level of the (real) wage is higher for any given level of unemployment. *If* both unemployment and wages are endogenous, it is natural to expect that in the new equilibrium both the real wage and the unemployment rate are higher — for all levels of other variables including productivity (as in Figure 1 above).

If the actual real wage is off the attractor, pressure to correct the deviation emerges. Therefore, a cointegrating relation like (iv) in (log) levels defines the error-correction part in the dynamic error-correction equation in (log) differences. In the full model, the estimation defines for each (endogenous) variable a difference equation which contains all long-run relationships present in the system. In so far as wages are considered — and allowing two lags in

[4] Let us consider two non-stationary variables x and y such that y = Ax. A acts as an attractor if there is some mechanism such that if y departs from Ax there will be a tendency to get back near to it. Because of uncertainties, rigidities, contracts etc., the mechanism may not immediately bring the points exactly to the attractor. "If the economy lies on A, a shock will take it away. If there is an extended period with no exogenous shocks, the economy will definitely go to the line an remain there. Because of this property, the line A can be thought of as an 'equilibrium', of the centre of gravity type" (Engle & Granger, 1991, p.2) .

levels which seems to be appropriate in all countries examined here — the wage equation is as follows[5]

$$
\begin{aligned}
\Delta\log(W/P)_t = {} & \gamma_{0,Q/N}\Delta\log(Q/N)_t + \gamma_{0,\tau_a}\Delta\log((1-\tau_a)_t + \gamma_{0,\tau_m}\Delta\log(1-\tau_m)_t \\
& + \gamma_{0,s}\Delta\log(1+s)_t + \gamma_{0,P_c/P}\Delta\log(P_c/P)_t + \gamma_{0,u}\Delta u_t \\
& + \gamma_{1,W/P}\Delta\log(W/P)_{t-1} + \gamma_{1,Q/N}\Delta\log(Q/N)_{t-1} + \gamma_{1,\tau_a}\Delta\log(1-\tau_a)_{t-1} \\
& + \gamma_{1,\tau_m}\Delta\log(1-\tau_m)_{t-1} + \gamma_{1,s}\Delta\log(1+s)_{t-1} + \gamma_{1,P_c/P}\Delta\log(P_c/P)_{t-1} \\
& + \gamma_{1,u}\,\Delta u_{t-1} + \text{possible constant} + \text{possible dummies} \\
& + \alpha_{W/P}\,[\beta_0\log(W/P)_{t-1} + \beta_{Q/N}\log(Q/N)_{t-1} + \beta_{\tau_a}\log((1-\tau_a)_{t-1} \\
& \qquad + \beta_{\tau_m}\log(1-\tau_m)_{t-1} + \beta_s\log(1+s)_{t-1} + \beta_{P_c/P}\log(P_c/P)_{t-1} + \beta_u(u_{t-1})]
\end{aligned}
\tag{v}
$$

In the present study we are particularly interested in the long-run coefficients β_i which are in the last two rows of (v). A significant constant term in the short-run part generates a trend to the level relationship (see Johansen, 1991c). It should be noted that in (v) wage dynamics is influenced by the unemployment rate both in levels and in differences, by labour productivity both in levels and in differences etc. However, the long-run convergence is towards the attractor defined by the β-coefficients.

A coefficient of special interest is $\alpha_{W/P}$ which reports the share of the equilibrium error which is corrected in the first period. The α-coefficient is often considered as a crude measure of speed of adjustment.

If the lagged dependent variable enters significantly the dynamic part of the equation, it may influence importantly the adjustment speed. A significant presence of differenced shock variables have similar impacts. Furthermore, point estimates of $\alpha_{W/P}$ should be considered cautiously because of two reasons. First, the short-run part of our model will remain more or less in an unrestricted VAR format and, therefore, much less parsimonious than the long-run part. Second, part of the OECD series have been readily seasonally adjusted and in some series — tax series in particular — the time disaggregation is more or less *ad hoc*.

In addition, when there are more than one cointegrating relationships — as there usually are — they all enter the difference equation above. The system becomes even more complicated than the one in (v) and several cointegrating vectors may influence estimation of the α-coefficient we are interested in. On the other hand, if one of the $\alpha_{W/P}$'s is considerably large in comparison to the others and, in addition, it is related to the long-run relation considered as a wage setting schedule, then the case is probably not problematic.

[5] A description of the underlying statistical model (in vector notation) is in Appendix 1. Equation (v) describes a case in which there is only one cointegrating vector in the system. If there are more cointegrating vectors, their equilibrium errors also enter equation (v) and additional α-coefficients are estimated. Furthermore, in (v) all variables except real wages are weakly exogenous. If one of the other variables is endogenous, its current difference does not enter (v) and, instead, an additional difference equation is estimated (simultaneously with the present one) with the other endogenous variable on the left-hand-side.

Discussion about $\alpha_{W/P}$-coefficients should not overlook these complications. The estimation methods available did not give us a straightforward means to consider this matter more explicitly.

In the Johansen estimations, the role of dummy variables differs importantly from their role in standard regressions. The dummies enter the short-run part of the model but not the long-run vectors (see Appendix 1). Use of *economically meaningful* dummies has been advocated because sudden shifts in variables (e.g. due to oil price shocks or tax reforms) create when differenced outliers which may make estimation of the short-run coefficients in (v) potentially arbitrary. As this also concerns the α's, problems could be generated on inference about conditioning, i.e. on the decision whether to consider part of the variables as weakly exogenous (see footnote 23 below). Of course, dummies should be allowed to enter only if formal tests related to residual analysis indicate that they are necessary.

2.1 Rigidities versus real wage resistance: The evidence

McKee, Visser & Saunders (1986) estimate the size of the "tax wedge[6]" in various countries. In 1983, the average tax wedge (at the level of an "average production worker) was 30–40 per cent in the USA, Canada and Australia. In Japan it was somewhat lower while in Germany and the UK it was somewhat higher. In Finland[7], the wedge was estimated to be slightly below and in France and Italy slightly above the 50 per cent level. In Sweden, the tax wedge exceeded the 60 per cent level, implying that the real after-tax wage which the worker receives is less than 40 per cent of the effective labour cost.

McKee et al. (1986, p. 53) argue that "a simple, but incorrect, comparison of the no-tax and tax models alone might suggest that the tax wedge ... is a measure of what labour 'pays'... Workers may not, in the end, 'pay' the ... taxes to the extent that pre-tax wages may rise to compensate for the taxes — so that the tax 'burden' is shifted to the owners of capital." Finally, they state that "the interest in tax wedges is *not* that these can tell us anything directly about the economic consequences of taxation, but rather they provide the necessary basic

[6] The "wedge" is the difference between the (gross) real labour cost paid by the employer and the after-tax real wage received by the employee. The "tax wedge" defines the contribution of taxes to this difference.

[7] As far as the employers' social security contribution rate is concerned, several OECD studies have used and presented misleading figures for Finland. McKee et al. (1986), e.g., state that this rate was 5 per cent in 1983 although the effective rate was more than four times higher. The confusion arises because Finland has a programme which has been considered as privately run even though the schemes are funded by mandatory contributions and are similar in all other ways to publicly run systems elsewhere. In the revised National Accounts of 1993, these schemes have been redefined to be part of the public sector. Inclusion of the effective rates would shift Finland into the group of countries with the highest tax wedge, similar to that in Sweden.

input for making such assessment." The aim of this paper is to carry out such analysis[8].

In *The OECD Jobs Study: Evidence and Explanations* (Vol. II, p. 247), earlier evidence on real wage resistance is reviewed. Many studies discover long-run effects of taxes on labour costs. A cross-country analysis by Symons & Robertson (1990) indicates, however, that in the long run the wedge is fully borne by labour[9]. This is in spite of considerable "short-run" effects which are long-lasting: on average, for 16 OECD countries, a 1 per cent rise in the wedge induces an immediate rise in labour costs of 1/2 per cent, and nearly half of this effect remains after 5 years. Given the further lags in the system this implies that a change in the wedge can have a significant impact on employment for at least a decade. LNJ (pp. 210—211) refer to these long lags found by Symons et al. and suggest that researchers who have considered the effects as "permanent" may have had difficulties in discriminating between permanent and temporary effects.

So, the most one can say is that there is plenty of evidence that taxes have very long-lasting effects on product wages, and hence on the equilibrium of the economy, operating via real wage resistance.

On the other hand, the distinction between the long run and the short run (or equilibrium and adjustment) has been adequately addressed in very little of the research carried out in the 1970's or 1980's. Methods which can be supposed to perform better in this respect are fairly new. The Johansen procedure allows us not only to distinguish between the long-run equilibrium and short-run dynamics. It also allows us to avoid problems related to "spurious regressions" between trended variables and to judge (indirectly) whether structural breaks had "first-order" impacts on the relationships of interest. Of course, the inference only concerns the data set and the observation periods which are available. The "very-very long term", which is not tractable by the data, remains beyond inference. This limitation, however, concerns all empirical studies.

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[8] In a recent OECD study, Turner, Richardson & Rauffet (1993) suggest that the inability of wage growth to adjust to a slowdown in productivity growth is the primary factor behind high unemployment in major economies. We are inclined to consider our study as an attempt to search for an explanation for this long-lasting deviation.

[9] When standard estimation methods are used, the evidence is heavily influenced by the short-run structure of the information set. The "long-run" is typically solved simply by shifting the lagged dependent variable to the left-hand side.

3 The Empirical Vector Autoregressive Model

In comparison with the sample size which is available, the number of variables in the theoretical model is such that the risk of overparametrization cannot be overlooked. VAR models share the property of all other models that the estimations become potentially vulnerable if the number of variables grows "too large". As the sample size is the most important problem for our study, we have searched for solutions which enable the dimension of the system to be reduced.

The most important compromise which has been made is to impose homogeneity between wages and prices *a priori*. Of course, we first tested the plausibility of this restriction which binds together wages and prices in one variable, the real wage W/P. Earlier evidence in a different context (see Tyrväinen, 1995b) indicates that it is preferable to impose this restriction as part of the estimation because allowing short-run deviations from the homogeneity conjecture improves the overall fit of a wage relation. However, since the deficiency is probably of secondary order, we gave priority to the reduction of the dimension of the full model.

Most (seasonally adjusted) semiannual series come from the OECD Analytical Data Base (ADB). Because of earlier evidence about a profound difference in the public sector wage behaviour in comparison with the rest of the economy, we emphasize the private sector only. A description of the data can be found in Appendix 2 below.

The operational counterparts of the variables are as follows. The wage series is the wage paid per wage earner. Employment is measured by the number of employed persons because data on working hours were not available. The producer price, P, is the aggregate value added deflator of private sector firms.

Our conjecture was that the employee side gathers information about inflation by monitoring consumer price index, CPI, which is published with a short lag, instead of private consumption deflator, PCP, which is published with a much longer lag. However, since the relationship between these two price measures varies much between countries[10] we considered both alternatives.

Productivity is measured by the output-employment ratio, Q/N, the growth of which is presumably the driving force behind the long-run growth in real wages. Since imperfect competition was assumed to prevail in the product market, relation (iii) includes a demand shift factor, Z. It is well-known from other contexts that there is no straightforward operationalization of this variable. Since in the long run demand and output presumably grow conjointly, one could operationalize Z as the value added, Q. It would, however, be difficult to distinguish the independent effect of Q from the effect working through Q/N. In order to keep the dimension of the model under control and to avoid problems related to multicollinearity, only the latter enters our empirical model.

[10] In Germany, France, Finland and Australia, CPI and PCP have moved more or less hand-in-hand in the long run. – In the USA, CPI has risen 5 percentage points more than PCP within the observation period and there are fluctuations in CPI/PCP which have lasted several years in a row. In Japan as well as in the UK, CPI/PCP has risen around 6 percentage points. The biggest rise was found in Canada, 10 percentage points. In Italy and Sweden, CPI has risen almost 10 per cent less than PCP from early 1970's to early 1990's.

Theoretical models stress the role of unemployment benefits, UB, (or replacement ratios) as factors which define the reservation wage or the position of workers who lose their jobs. New data on replacement ratios was collected for *The OECD Jobs Study*. However, the first experiments revealed that, in most countries, there was so little variation in these ratios over our estimation period that no significant impacts could be seen. Therefore, we omitted UB at the out-set. Strike allowances, SA, are usually defined strike-by-strike and may change in the course of each dispute. Because of lack of time series, SA was accordingly left out.

The majority of the empirical literature characterizes the income tax system with one parameter only, either the average tax rate, τ_a, or the marginal tax rate, τ_m. The analysis in Tyrväinen (1995b) as well as in Lockwood & Manning (1993), however, indicates that this may be insufficient since both tend to matter and have separate roles.

Jackobsson (1978) proposes a progressivity index τ_p which links the average and the marginal income tax rate as

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

When τ_p is subtracted from unity and a logarithm is taken, one gets

$$\begin{aligned} \log(1-\tau_p) &= \log\left(1 - \frac{\tau_m - \tau_a}{1-\tau_a}\right) \\ &= \log(1-\tau_m) - \log(1-\tau_a). \end{aligned} \tag{vii}$$

Below, we include both τ_a and τ_m and expect that $W^*_{\tau_a} \geq 0$ and $W^*_{\tau_m} \leq 0$.[11] When reporting the results, use of the Jackobsson index (vii) will be made.

Operationalization has been conducted by our special interest in studying real wage resistance, i.e. the impact of wedge variables in wage setting, and our 7-dimensional log-linear unrestricted VAR-model contains

1) real wages, W/P, where W is the wage paid per employee and P is the value added deflator,
2) labour productivity, Q/N,
3) the average income tax rate, $1-\tau_a$,
4) the marginal income tax rate, $1-\tau_m$,
5) the employers' social security contributions rate (including both voluntary and statutory contributions), 1+s,

[11] As will be seen, in some cases the structure of the information set is such that some tax variables enter significantly whereas some others do not. As far as income taxation is concerned, this may happen if either τ_a or τ_m is flat over the estimation period. On the other hand, in some countries τ_a and τ_m seem to cointegrate. In these cases, the result of insignificancy may be due to time series properties of certain series and it should not necessarily be considered as an analytical device. This stresses the importance of careful examination of the data.

6) the price wedge, i.e. consumer price relative to the producer price P_c/P, which contains the effect of consumption taxes,
7) unemployment as measured by a) the unemployment rate, u, b) log(u) or c) log of number of unemployed persons.

It should be recognized that shifts in (world market) prices of raw-materials (including energy) and in exchange rates have been reflected in the two deflators which enter the model as well as in the price wedge. The observation period is 1972S1—1992S2 except that for the UK, Italy and Japan the data are complete only up to 1991S2, for Sweden and Finland to 1990S2 and for Australia to 1990S1.

The income tax data used in this study differ from those used in earlier studies. Turner et al. (1993) and Symons et al. (1990), e.g., approximate income taxes with a relation of all taxes paid by households to all pre-tax incomes of households. As indicated by McKee et al. (1986), this is not without problems. Our data are derived for an "average production worker"[12] with a dependent spouse and two children from the OECD publication "The tax/benefit position of production workers". For an average worker with similar status but with a working spouse, data have been recently constructed at the OECD. Income tax series include employees' social security contributions. Consumption tax rates used in the calculations are new OECD estimates.

In order to examine the properties of the series in full VAR-models, stationarity tests and exclusion tests[13] were carried out. None of the variables seems to be generally non-relevant and could, hence, be excluded *a priori* (exclusion-test). Stationarity of the series is generally rejected (stationarity-test).[14] As far as the lag length is considered, misspecification tests indicate that we do not lose anything by restricting it to 2.[15]

At the outset, a 7-dimensional model was estimated for each country.[16] In this model, a joint test[17] is performed which defines a) the cointegration rank,

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[12] For a discussion of the concept of "average production worker" and problems related, see McKee et al. (1986).

[13] For the test procedures, see Juselius & Hargreaves (1992).

[14] It has been shown recently that if there is I(2)-ness in the model, this test needs reinterpretation (see Johansen, 1995).

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

[16] Besides this procedure we have estimated — as a "control solution" for Finland — a similar model using the quarterly data of the BOF4 model of the Bank of Finland which was also estimated in Tyrväinen (1995b). This model concists of same variables as the ones for the other countries with one exception due to the fact that the data file of BOF4 also includes producer prices. Since we use producer prices as the real wage deflator, the real price of imported energy and raw-materials will also enter the "control model". Inclusion of real prices of imports of energy and other raw-materials, P_m/P, increases the dimension of the unrestricted VAR-model to 8. In Table A4, we report results of these estimations as well.

[17] For the test procedures, see Johansen (1992), Johansen & Juselius (1990) and Tyrväinen (1995b). The asymptotic critical values tabulated in Johansen & Juselius (1990) were used. Test results can be found in Appendix 3 below.

r, which specifies the number of linearly independent stationary relations between the levels of the variables and b) the presence of a linear trend[18]. Cointegrating relations are the eigenvectors corresponding to the r largest eigenvalues in the system[19]. Trace tests are in Table A2 in Appendix 3[20].

In a 7-dimensional model (n = 7) with three long-run relations (r = 3), there are four common trends (c = n − r = 4). In many countries the process seems to be I(2) in one or two direction(s) ($c_2 = c - c_1 = 1$) implying that there are one or two common I(2) trend(s) which drive(s) the system[21]. So, a three-dimensional cointegration space would be stationary in two or one direction(s) and non-stationary in one or two direction(s) such that differenced I(2) variables are needed to obtain stationary. This is an example of multicointegration.

Residual analysis showed that at the outset all Gaussian assumptions were not always satisfied. In order to reduce the problem we introduced three dummies to each country model. The dummies refer to discrete shifts in the price of energy in 1973, 1979 and 1986. Their significance was tested for each country separately. In many cases we could drop one or two of the dummies as insignificant without an effect on residuals. In some countries, major tax reforms had to be accounted for[22]. As stressed in Section 2 above as well as in Appendix 1, dummies only enter the dynamic part of models and leave the long-run relationships unaffected.

All variables are endogenous at the outset. Since the parameters of interest are the long run parameters, β, we examined whether some of the variables

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[18] In the set-up applied, the test regarding the presence of a linear trend is a test about significance of a constant term in a difference equation like (v) above (see Johansen, 1991c).

[19] 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). 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 the assumption concerning the existence of the linear trend (yes or no). The distribution has broader tails if the trend is absent.

[20] 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. For example, for France $H_0^*, ..., H_2^*$ and $H_0, ..., H_2$ are rejected as well as H_3^*. However, $Q_3 < CV_{95\%}$. So we conclude that r = 3 and reject the hypothesis of no linear trend.

[21] The evidence is given in Table A3 in Appendix 3. As shown by Johansen (1991d), inference in the presence of I(2)-ness can be conducted using the tables prepared for the analysis of cointegration of I(1) variables. Investigation of the so called $\beta_\perp^2$-vector resulting from the system (see Johansen, 1992a) reveals that I(2)-ness is in many cases mainly to be found in the complicated properties of the tax series. In Germany, France, Italy and Australia, there is no sign of I(2)-ness in the information set.

[22] In cases where tax reforms have made the residuals "wild", a dummy was introduced and its significance tested. A full set of dummies is available from the author on request.

could be considered weakly exogenous. Of course, if there are, for example, three cointegrating vectors in a VAR-model, we cannot reduce the set of system variables to a smaller number. Weak exogeneity indicates that a variable does not react to a disequilibrium in any of the cointegrating vectors. This can be tested in the full model.[23] One can also evaluate qualitatively whether some of the variables are exogenously determined (tax rates, e.g.). The time series properties may also indicate whether the data results from an endogenous data generating process.

In all cases, tests indicate that one or more of the variables could be considered weakly exogenous. Conditioning varies from country to country according to test results (for the test, see Juselius & Hargreaves, 1992). When a test was at the limit, we chose the smaller number of endogenous variables but checked whether the choice influences the rest of the inference. Results of residual analysis of the partial models were encouraging[24] (see Appendix 3 below).

4 Structural Restrictions which Identify a Long-run Wage Setting Schedule

In order to identify a long-run wage setting schedule in the multivariate vector space, identifying restrictions can be defined. Theoretical considerations indicate that a log linear relation

$$\begin{aligned}\log W/P = {} & \beta_{Q/N}\cdot\log(Q/N) + \beta_{\tau_a}\cdot\log(1-\tau_a) + \beta_{\tau_m}\cdot\log(1-\tau_m) \\ & + \beta_s\cdot\log(1+s) + \beta_{P_c/P}\cdot\log(P_c/P) + \beta_u\cdot u\end{aligned} \tag{viii}$$

should be considered as a wage setting schedule only if $1 \geq \beta_{Q/N} \geq 0$, $\beta_{\tau_a} \leq 0$, $\beta_{\tau_m} \geq 0$, $\beta_s \leq 0$, $\beta_{P_c/P} \geq 0$, and $\beta_u \leq 0$.

As far as real wage resistance is emphasized, a wedge-restriction like

[23] The hypothesis related to weak exogeneity is that for selected equations, the α_i's are zeros. The test statistic is similar to that described in footnote 28 below. It has been shown recently that if there is I(2)-ness in the model, the results of exogeneity tests included in CATS must be considered cautiously (see Paruolo & Rahbek, 1995). This is because an extra term related to the I(2)-ness enters the test statistic. If the test rejects weak exogeneity this term does not influence the inference. If the hypothesis of weak exogeneity is accepted, the extra term may or may not influence the inference. So far, appropriate test producers to take account of this have not been available.

[24] Eitrheim (1991) shows that the parameter estimates of the long-run relation are not sensitive for misspecification in other respects than the one generating autocorrelation. In broad terms, Cawing & La (1993) seem to reach a similar conclusion. Autocorrelation is not a problem in our models.

$$-\beta_{\tau_a} = \beta_{P_c/P} = 1 + \beta_s = \beta_{WEDGE} \tag{ix}$$

is tested. If employees have full dominance in wage setting, β_{WEDGE} equals unity. If the firms dominate, β_{WEDGE} equals zero. In the former case taxes fall fully on the firm, in the latter they fall fully on the worker.

To consider wedge-restrictions which are seemingly identical to (ix) — when imposed on (viii) — Table 1 introduces the H_0-hypotheses in some well-established estimating equations. The *a priori* structure can be seen by investigation of the β_i-coefficients. For example, the H_0-hypothesis in (II) is that taxes are fully borne by unions. In equation (III), according to H_0 taxes are fully borne by firms. (I) and (IV) are combinations which imply asymmetric incidence.

Model (II) has been used by, for instance, Calmfors & Forslund (1990) and Calmfors & Nymoen (1990) with the *a priori* restriction $|\beta'_s| = |\beta'_\tau| = |\beta'_{P_c/P}|$. Model (III) has been used, among others, by Eriksson et al (1990) and by Pencavel & Holmlund (1988) with the *a priori* restriction $|\beta''_s| = |\beta''_\tau| = |\beta''_{P_c/P}|$. Model (IV) has been estimated by Rødseth & Holden (1990) with restriction $\beta'''_s = \beta'''_\tau = \beta'''_{P_c/P} = 0$.

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

Coefficient	Independent variables	Dependent variable (I) W	(II) $\left(\frac{W(1+s)}{P}\right)$	(III) $\left(\frac{W(1-\tau)}{P(1+v)}\right)$	(IV) $\left(\frac{W}{P(1+v)}\right)$
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: $\beta'''_v = 0 \Leftrightarrow \beta_v = 1$

The fact that seemingly similar restrictions have very different implications in different specifications[25] is of profound importance for the inference. This is even more crucial if the elasticities related to taxes are not precisely defined as often happens. In the present paper, we apply a procedure and a testing strategy which should help avoid implicit *a priori* structures which could bias the test in one direction or another.

We expect one well-specified relation, i.e. wage setting schedule to show up. However, as the tests indicate that there are other long-run relations in the

[25] This has seldomly been stressed in the literature with Calmfors & Nymoen (1990) being one of the few exceptions.

data set, we have additional vectors to consider[26]. This implies also that the relation we are interested in can be a linear combination of several vectors. Here, the identifying restrictions become vital. In what follows, short-term dynamics is determined freely but on the long-run part we impose restrictions which identify a long-run wage relation. *All* competing hypotheses discussed above will be tested and no *a priori* structure is imposed on β_i's which relate to real wage resistance. Structural restrictions are tested in partial models and the structure which is best in accordance with the data defines the preferred relation.

Structural restrictions are tested in partial models and the structure which is best in accordance with the data defines the preferred relation marked with a star, *, in Table A4[27].

In the present context, we occasionally reconsider earlier inference on cointegrating rank, r. This is because of the small sample problems discussed above. Choosing a "too high" r implies that the tests imposed are "too loose". On the other hand, if the correct choice is, for example, r = 4 but we choose r = 3, the tests are excessively stringent and the resulting p-values are definitely the low limits of the appropriate ones. Whether the "last" eigenvector contains relevant information about the long-run relations of interest can also be evaluated by comparing the parameter estimates discovered including and excluding this vector.

Each vector of β's is linked to a vector of α's with at least one element, α_i, different from zero. The α's are the weights with which the cointegrating relation enters a dynamic equation and they embody the error correcting structure in the system.

A restriction on β-coefficients is data consistent if the eigenvalues related to the restricted estimation do not differ "too much" from the unrestricted estimation. Each restriction is always compared to the original unrestricted estimation and all r eigenvalues contribute the test statistic[28] which follows the χ^2-distribution with degrees of freedom indicated in Table A4.

[26] These vectors may mimic various processes. As pointed out by Johansen & Juselius (1994), in macroeconomic behaviour a role is often played by various types of agents with disparate goals (demanders versus suppliers, producers versus consumers, employers versus employees etc.) interacting in such a way that equilibrium is restored once it has been violated. This complicates evaluations because one may also pick up vectors describing either one of the sides or a mixture of both of the sides influencing the variables concerned. In the present context, the "left-over" vectors could describe determination of productivity, prices, employment and/or unemployment and complicated structures between various tax rates may also play a role. A closer look at the variables in our unrestricted VAR makes it clear that none of these potential relationships could be well specified. These "semirelations" may also be mixtures of two or more competing but misspecified relationships. Hence, one should not put too much emphasis on overinterpretation of the "left-over" vectors (see also the discussion in Tyrväinen, 1995c).

[27] In Table A4, the results are in vector notation. So, all elasticities with respect to the real wage have the opposite sign compared with that usually seen in standard regressions where one of the variables (= the real wage) has been removed to the left-hand side.

[28] The test statistic concerned is $T\sum_{1}^{r}\ln((1-\hat{\lambda}_i)/(1-\tilde{\lambda}_i))$ 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).

The necessary condition for unique identification is that the minimum number of restrictions is one less than the number of cointegrating vectors, r−1. It is two whenever we conclude that r = 3. If r = 4, the minimum number of restrictions is three.

We started by testing the restriction $\beta_{W/P} = -\beta_{Q/N}$ which imposes long-run homogeneity between labour productivity and real wages. The other restriction imposed at the outset was the wedge restriction (ix) augmented with (vii) which allows the effect of τ_m. If (ix) passed the test, we continued by testing whether β_{WEDGE} differs significantly both from zero and from unity. When the test indicated that $0<\beta_{WEDGE}<1$, we tested whether coefficients of all components differ significantly from zero and unity.[29]

If rejected, the β_{WEDGE}-restriction was relaxed. It was then tested whether any of the coefficients could separately be restricted to zero. In some cases we found a coefficient which differs significantly from zero and has a value which is close to unity. In these cases we also tested whether the deviation from unity is significant. This procedure was continued until a parsimonious description of the long-run relationship was reached in which only significant elasticities enter.

As can be seen from the p-values in Table A4, the identifying restrictions generally pass the LR test at a relatively high significance level. As indicated above, a hypothesis is usually rejected if the p-value is less than .05. In the restricted vectors, the α-coefficients are larger in magnitude than in the non-restricted vectors which indicates success in search for an attractor[30].

Homogeneity conjecture between the real wage and labour productivity pass the test in all cases although the adjustment lags seem to be of considerable length.

[29] In the estimation procedure applied, standard t-statistics cannot be used to evaluate significance of the long-run coefficients. Instead, the likelihood ratio test allows us to examine whether (in an acceptable structure) a long-run coefficient which is close to zero (or unity) is actually zero (or unity). For example, if in the preferred relation for Japan, β_u is restricted to zero, the difference between the original test statistic (3.20) and the resulting test statistic (11.56) follows $\chi^2(1)$-distribution which gives 3.84 as the critical value on 5 per cent significance level. So unemployment variable can definitely not be omitted from the wage relationship in Japan (because $11.56 - 3.20 = 8.36 > 3.84$).

[30] Please recall 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. — Whether we correctly interpret a long-run relation can also be considered by investigation of the weights, $\alpha_i\beta_i$. The preferred long-run vectors enter the short-run equations for wages, ΔW, with a considerable weight. The rest of the weights are usually considerably smaller in magnitude. This surely does not indicate that the interpretation suggested for these long-run relations would be arbitrary. Finally, "left-over" relations do not usually influence wage equations as the related α-coefficients are close to zero in many cases. Hence, the fact that we cannot give a plausible interpretation of the "left-over" relations hardly undermines the credibility of identification of the relation of interest.

5 Estimation Results

Table 2 summarizes the long-run elasticities related to real wage resistance which derive from Table A4. So, all structures in Table 2 have been tested.

In eight out of the ten countries, the β_{WEDGE} -restriction imposing one single coefficient on all relevant wedge variables is in accordance with the data. In the USA and Sweden this was rejected[31]. As far as Sweden is concerned, we argue below that the rejection is due to particular time-series properties of the tax data in the observation period. We can also reject the hypothesis that the numerical value of the β_{WEDGE}-coefficient is the same in countries where restriction (ix) passed the LR test.

In four countries, a separate impact of marginal income tax rate could be found which indicates that steeper progessivity has tended to moderate wage claims. In (vi) and (vii) above, the Jackobsson progressivity index τ_p was defined. In our results, wage elasticity with respect to τ_p is .6 in Italy, in Japan and Finland the elasticity is .5 and in Canada it is .2.

Whether our interpretation about a long-run relation is correct can also be considered by investigation of the weights, $\alpha_i\beta_i$[32]. Our preferred long-run vectors enter the short-run equations for real wages, $\Delta(W/P)$, with a considerable weight. The rest of the weights are usually clearly smaller in magnitude. This surely does not indicate that the interpretation suggested for these long-run relations would be arbitrary. Finally, "left-over" relations do usually not influence wage equations as the related α-coefficients are close to zero in many cases[33]. Hence, the fact that we cannot give a plausible interpretation to the "left-over" relations hardly undermines the credibility of identification of the relation of interest.

Because interpretation of Table A4 is perhaps not without problems, we also discuss the results country by country. When we make use of the Jackobsson index (ix) and write the wage relationships so that the real labour cost – i.e. the real product wage $W_{pr} = W(1+s)/P$ – is on the left-hand side, we arrive from Table A4 to the elasticities reported first country by country below and summerized thereafter in Table 2.

Generally, we find evidence of a more considerable amount of real wage resistance than many earlier researchers. We believe that this is because the method applied is more suitable for identification of long-run relationships.

The cointegrating wage relations for the ten countries are as follows.

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[31] At face value, this rejection could allow long-run "free lunch" policies. As will be indicated below, caution is required in this respect.

[32] The strength of error correction can be evaluated by multiplying each α_i with β_i. In Table A4, the disequilibrium error of the long-run relation (1) for Canada enters the difference equation of wages with weight (–.254)*(1.000) = –.254 which indicates rapid error correction. In the productivity equation, the same relation (1) enters with weight (.032)*(–1.000) = –.032.

[33] For example, when the preferred structure for the USA (with r = 2) is considered we have one "left-over" cointegrating vectors. For it we have $\beta_1\alpha_1$ = +.03. In Canada, in addition to the preferred vector (with r = 3) we have two "left-over" cointegrating vectors. For the first of them we have $\beta_1\alpha_1$ = –.024 and for the second $\beta_1\alpha_1$ = +.01.

Germany:

$$\log(W_{pr}) = 1.00\cdot\log\left(\frac{(CPI/P)(1+s)}{1-\tau_a}\right) \qquad \text{(x)}$$
$$+1.00\cdot\log(Q/N) - .17\cdot\log(u)$$

In Germany, the Trace test suggests that the cointegration rank is four and that there is a linear trend in the data set (see Appendix 3, Table A2 for Germany). Under this conjecture, the structure above passes the LR-test with a p-value[34] of .68 when income tax data for two-earner families (an average production worker with a working spouse) are used. For one earner families, the p-value is .13 (see relations (1) and (2) for Germany in Table A4).

In Germany, the degree of real wage resistance appears to be very high which contradicts some earlier results (see Turner et al., 1993, e.g.). However, the magnitude of the effect of real wage resistance on wages can only be evaluated when actual changes in the wedge are taken into account as we do in section 6.1 below.

Around 15 per cent of a disequilibrium is corrected within the first half-year. The unemployment rate variable (in logarithmic form) is highly significant. A dynamic wage equation which incorporates the long-run relation (x) explains 79 per cent of the semester-to-semester variation in real wages.

Canada:

$$\log(W_{pr}) = 0.80\cdot\log\left(\frac{(PCP/P)(1+s)}{1-\tau_a}\right) + .20\cdot(\tau_a - \tau_m) \qquad \text{(xi)}$$
$$+1.00\cdot\log(Q/N) - .02\cdot u$$

We conclude that r = 3 and that a linear trend is present in the data (see Appendix 3, Table A2 for Canada). Under this conjecture, the p-value related to the preferred structure is .06. The preferred structure indicates almost full wage compensation and allows a separate role for progressivity (marginal income tax rates increased considerably particularly in the late 1980's). In fact, the hypothesis that the wage response is unity cannot be rejected. In fact, the p-value related to that structure is .30 (see relation (2) in Table A4) or .14 depending on which tax rates are applied. The former result relates to a model for an average wage earner with a dependent spouse. In the latter case the spouse is working. The structure is qualitatively identical in both cases. Despite these facts, we have chosen a relation which indicates partial wage compensation as the preferred one. This is because even in this model the wage response exceeds that found in some earlier studies.

[34] The concept of the "p-value" refers to the significance level. Usually, a hypothesis is rejected if the p-value related to the test is smaller than .05. In statistics, the concept of p-value is related to the "type II error" which indicates acceptance of H_0-hypothesis when it is in fact false.

In Canada, the Consumer Price Index, CPI, has risen 10 per cent more than the deflator of the private consumption, PCP, within our estimation period. A model in which the price wedge is defined as PCP/P gives a somewhat better overall explanation of the data generating process.

A tightening in income taxation which takes the form of an equi-proportional rise in average and marginal tax rates, leads to a considerable wage compensation. However, if the progressivity steepens, wage compensation diminishes[35].

Wages were not found to respond strongly to an increase in employers' social security contributions. So, much of a rise in employers' social security taxes remains as a higher labour cost. Part of the explanation could be in the generous unemployment benefit system because of which economic losses facing workers if they become unemployed are smaller than in many other countries.

An increase in the price wedge leads to a high degree of adjustment of nominal wages. COLA clauses are probably only part of the explanation.

The coefficient of the unemployment variable is highly significant and it's magnitude is considerable. Around 15 per cent of a disequilibrium in the wage relation is corrected within the first semester. The dynamic equation incorporating the long-run relation (xi) explains 85 per cent of the short-run variation in real wages.

Japan:

$$\log(W_{pr}) = .50\cdot\log\left(\frac{(PCP/P)(1+s)}{1-\tau_a}\right) + .50\cdot(\tau_a - \tau_m) + 1.00\cdot\log(Q/N) - .05\cdot u \qquad \text{(xii)}$$

At face value, the Trace test indicates that the cointegration rank is five (see Appendix 3, Table A2 for Japan). However — given the complexity of the model — the test statistic indicating rejection of $r = 4$ is not much above the critical level and, in addition, examination of the residual of the fifth eigenvector reveals that the cointegration property in this relationship cannot be particularly strong. If we drop out the fifth eigenvector, we can condition on three tax variables, i.e. average and marginal income tax rates as well as employers' social security contributions. We prefer a model with less parameters to be estimated and assume that $r = 4$ and condition on three tax variables. This choice was facilitated by the result of Cheung et Lai (1993) which indicates that with small samples the asymptotic tests are potentially biased towards finding cointegration slightly "too often".

Under $r = 4$, the relation above passes the LR-test with a p-value of .20. The preferred structure indicates partial wage compensation and allows a role for the progressivity. Japan is one of the countries with a systematically increasing gap between CPI and PCP. A specification of the price wedge with

[35] This kind of effect is more probable in countries with an important influence of trade unions. Please note also that in Canada, marginal income tax rates increased considerably particularly in the late 1980's.

the latter was preferred. The speed of error correction in Japan is high. Almost half of a disequilibrium is corrected within the first half year.

Due to conditioning, the magnitude of the eigenvalue related to the fourth cointegrating vector diminished from .46 in the full model to .27 in the partial model. So, it could be more appropriate to assume that there are three cointegrating vectors in the system. As can been from relation (2) for Japan in Table A4, under r = 3, a relation which is numerically by and large identical to the one above has a p-value of .07. So, our findings are not sensitive for the choice between three or four cointegrating vectors.

A dynamic equation containing (xii) explains 71 per cent of the short run variation in real wages.

Finland:

$$\log(W_{pr}) = .5\cdot\log\left(\frac{(CPI/P)(1+s)}{1-\tau_a}\right) + .5\cdot(\tau_a-\tau_m) + 1.00\cdot\log(Q/N) - .31\cdot\log(u) \qquad \text{(xiii)}$$

The preferred relation (xiii) above stems from a set-up in which — as in the models for the USA and Italy below — homogeneity between real labour cost and labour productivity — easily accepted by data — was imposed *a priori* In this structure, we find two cointegrating vectors and no linear trend in the data set (see Appendix 3, Table A2 for Finland).

As far as the control model[36] is concerned, the Trace test indicates that r = 4. However, having conditioned on three of the variables the eigenvalue related to the fourth eigenvector had diminished considerably. Therefore, the appropriate choice could be r = 3. We tested the relevant restrictions both under hypothesis r = 4 and r = 3. As can been seen from relations (3) and (4) in Table A4, the structure is identical in both cases. Under r = 4 the p-value is .94 and under r = 3 it is .74.

Elasticities related to taxes are qualitatively similar in relation (1) and in the "control model" which uses somewhat different data and a different (quarterly) time disaggregation. The "original" data accept with a high p-value a predefined structure which is between in (1) and (4). This model indicates a considerable role of the unemployment variable as well as of the error correcting forces. So, relation (2) was chosen as the preferred one.

The preferred structure passes the test with a p-value of .88. The error correction coefficient has the value of –.30. Half of a change in employers' social security taxes as well as in the price wedge is shifted to higher labour costs. The same holds for an increase in income taxes which leaves the progressivity index intact. However, if progressivity is tightened, wage push

[36] The structure and the data in the "control model" for Finland is somewhat different from the other countries. Because we use data about producer prices — which is not available in ADB — instead of the value added deflators, we also include real energy-prices, P_m/P. Because P_m/P had not a significant long-run effect on wages, we wrote relations (3) and (4) in a form seemingly identical to the other ones. Of course, $\Delta(P_m/P)$ enters difference equations (i.e. vectors of α-coefficients in Table A4). In order to avoid further complication in reporting, these coefficients — which are .075 and .048 in (3) and (4) — were left out from the Table.

related to the income tax is reduced by around 50 per cent of the magnitude of the change in the progressivity index.

In the preferred relation, data rejects omission of the unemployment variable. The relevant change in the LR-test statistic is 4.79 which clearly exceeds the critical value of 3.84. The dynamic equation incorporating (xiii) explains 77 per cent of the short-run variation in real wages.

Australia:

$$\log(W_{pr}) = .50 \cdot \log\left(\frac{(PCP/P)(1+s)}{1-\tau_a}\right) + 1.00 \cdot \log(Q/N) - .04 \cdot u \qquad \text{(xiv)}$$

The result comes from an estimation in which the cointegration rank is three and a linear trend enters the model. As Table A2 in Appendix 3 indicates, the choice of r was not straightforward. At face value, the Trace test indicates that r = 4 and that a linear trend is probably absent. On the other hand, the test statistic only slightly exceeds the critical values implying r = 3 with or without a linear trend. Therefore, we considered four cases: r = 4 with and without and r = 3 with and without a linear trend. Because the sample size was even smaller than in other countries, we strongly preferred a smaller rank particularly as this did not influence much the results. Exclusion of the linear trend would leave the preferred structure unchanged but the error correction property disappears. Therefore, despite all the problems related, we preferred relationship (2) in Table A4.

In estimations, one could not have both τ_a and τ_m in the wage relation[37] since they tended to play a similar role. This appears to reflect the fact that in Australia marginal income tax rates have been very high and the average rates have risen permanently because more and more wage earners have been moving upwards in the wage scales (see various issues of OECD Country Studies on Australia). As a matter of fact, a test reveals that τ_a and τ_m move so closely hand-in-hand that they are actually cointegrated. As they tended to cancel out each others' significance in the estimation, we excluded the latter *a priori.*

According to (xiv), in Australia wages have absorbed half the changes in employers' social security taxes in the long run. In the short-run the effect on real labour costs is, however, full. Changes in income taxes as well as in the price wedge have been half compensated. The structure in (xiv) has a p-value of .60.

The speed of error correction is fairly fast (α is almost .4). A dynamic equation incorporating (xiv) explains 62 per cent of the semester-to-semester variation in real wages.

[37] In Australia, income tax data for two-earner households generated slightly better behaving residuals than the data for one-earner families. Therefore, results according to the former have been reported although they were almost identical in both cases.

France:

$$\log(W_{pr}) = .42\cdot\log\left(\frac{(CPI/P)(1+s)}{1+\tau_a}\right) + 1.00\cdot\log(Q/N) - .29\cdot\log(u) \qquad \text{(xv)}$$

The Trace test suggests that r = 3 and that a linear trend is in the data set (see Appendix 3, Table A2 for France). The preferred structure found under this conjecture is very similar to that of the countries above. Income tax effects could, however, be distinguished only after the marginal tax variable was excluded *a priori*. This is because $\tau_a-\tau_m$ (which is a difference of two non-stationary variables) has been almost flat in France: most of the time it has fluctuated between 5.2 and 5.8 percentage points, with only two observations outside this range. As in the case of Australia, a formal test indicates that the two income tax variables have been cointegrated in the observation period.

The preferred structure passes the LR-test with a p-value of .42. The coefficient of the wedge term is .4 and this estimate differs significantly from .5. (The test statistic indicating this has the value 5.98–.65 = 5.33 which exceeds considerably the critical value of 3.84; see relations (1) and (2) for France in Table A4).

The strength of error correcting forces is similar to that in many other countries. Around 30 per cent of a disequilibrium is corrected in the first half-year. The dynamic equation incorporating (xv) explains 76 per cent of the semester-to-semester variation in real wages.

Italy:

$$\log(W_{pr}) = .40\cdot\left(\frac{(PCP/P)(1+s)}{1-\tau_a}\right) + .60\cdot(\tau_a-\tau_m) + 1.00\cdot\log(Q/N) - .33\cdot\log(u) \qquad \text{(xvi)}$$

In a model with — as in the USA and Finland — long-run homogeneity between productivity and real wages imposed *a priori*, cointegration rank is three (see Appendix 3, Table A2 for Italy). Under this conjecture, the structure above has a p-value of .44. The error correction coefficient has the value of .15 and the dynamic equation which incorporates (xvi) explains 81 per cent of the short-run variation in real wages.

The result above indicates partial response of real labour costs to wedge factors with a strong effect of income tax progressivity[38].

In estimations, PCP performed better than CPI. Therefore, the former enters the price wedge in the reported relations.

[38] In broad terms, this seems to be in accordance with the result in Padoa-Schioppa(1992).

Sweden:

$$\log(W_{pr}) = 1.00 \cdot \log(CPI/P) + 1.00 \cdot \log(Q/N) - .23 \cdot \log(u) \qquad \text{(xvii)}$$

We conclude that r = 3 and that a linear trend is in the information set. The p-value related to the preferred relation is .21. In this structure, changes in the price wedge have been fully compensated by changes in wages. Since wage indexation was not widely used in the observation period, the reason is presumably in union behaviour in wage setting. Wage contracts as well as the wage drift have been driven by consumer price developments.

Changes in employers' social security taxes have been fully absorbed by wages. A look at the dynamic part of the model reveals that most of the absorption takes place already in the very short term.

Income taxes were not found to have major effects on wage determination. This seems to reflect the fact that the variation in the average rate has been small except at the end of the estimation period. The LR-test for omission of τ_m is on the limit: the test statistic is 3.95 against a critical value of 3.84. Since the sign of the coefficient is against our *a priori* expectations we excluded it from the preferred structure.

Finally, the unemployment variable plays a definitely significant role in the long-run wage relation and its coefficient is fairly large. When unemployment rate is in a logarithmic form, residuals of the full model behave better than otherwise.

Error correction is fairly slow (α is around .1). The dynamic equation which contains long-run restrictions incorporated in (xvii) explains 93 per cent of the short-run variation in the real wage.

USA:

$$\log(W_{pr}) = 1.00 \cdot \log\left(\frac{1}{1-\tau_a}\right) + 1.00 \cdot \log(Q/N) - .01 \cdot u \qquad \text{(xviii)}$$

There are two sets of results in Table A4. Relations (1) and (2) stem from a set-up with seven variables and a cointegration rank of three. If we impose homogeneity between real wage and productivity *a priori* — in (1) and (2) it is imposed as part of the estimation — the cointegration rank drops to two. Relationships (3) and (4) stem from the latter specification with six variables.

The preferred relation (3) in Table A4 passes the LR-test with a p-value as high as .64. If one restricts the coefficient of τ_a to zero, the LR test statistic jumps from 1.68 to 16.00 and the p-value drops to zero[39]. Therefore, wage

[39] One should note that here we do not test acceptance of an additional restriction in which case the critical value of the test statistic would be 3.84. In the present context, all four competing structures implement an identical number of restrictions. The hypotheses are 1) $\beta_a = -\beta_v = -1$, 2) $\beta_a = 0$, $\beta_v = 1$, 3) $\beta_a = -1$, $\beta_v = 0$, and finally 4) $\beta_a = \beta_v = 0$. So, the inference can not be based on a standard LR test but one must evaluate otherwise which of the competing structures is best in accordance with the data.

effect of income taxes can definitely not be rejected in the present set-up[40]. Partial compensation is also rejected (vector (2) outperforms vector (1) for the USA in Table A4).

This result contradicts most existing beliefs. It indicates that in the USA, a rise in average income tax rate has been fully compensated in pre-tax wages. So far, the only explanation we could think of for this finding — in addition to local bargaining or bargaining about after-tax wages — refers to efficiency wage considerations: if employees' effort depends on the purchasing power offered by the firm, employers should care about real take-home pay. So, employers might be willing to compensate for changes in income taxes when — as in the USA — tax shifts take place discreetly with common knowledge about tax reforms causing changes in living standards.

The level of progressivity — measured by the Jackobsson-index — has been fairly stable in the USA and its effect could not be distinguished.

An increase in employers' social security contributions leads to downward adjustments in wages. Since the response is one-to-one, no long-term effect on real labour costs remains. In the short run, employers' social security tax influences real labour costs by its full amount. Simulations indicate that it may take as long as ten years for this effect to die out fully. In the second and the third year after the shock, the unfavourable impact on (un)employment is of considerable order. Although it vanishes in the long-run, a rise in unemployment can still be seen after five years of adjustment (see Figure 5).

Evidence related to the price wedge (incorporating indirect taxes) is mixed. Relations (1) and (2) indicate that an increase in the price wedge leads to an upward adjustment of nominal wages with the full amount. Because of COLA clauses we did not find this surprising. On the other hand, these relations (1) and (2) lack a significant unemployment variable.

When homogeneity between the real wage and productivity was imposed *a priori*, the results changed in three important respects. First, cointegration rank dropped from three to two as could be expected. Second, unemployment variable became significant — although the magnitude of the coefficient is tiny[41]. Third, the long-run response to the price wedge disappeared although a short run response of .8 remains.

The error correction feature in relation (3) is clear: 20 per cent of a disequilibrium is corrected in the first half-year. This so although the speed of adjustment is hampered by the presence of staggered contracts. The dynamic equation incorporating the long-run relation (xviii) explains 67 per cent of the short-run, semester-to-semester variation in real wages.

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[40] In our average income tax data there are — in addition to a trendwise increase — two shifts. An upwards shift takes place in the first two years of the 1980's. After the middle of the decade a (partly) compensating shift downwards can be seen. — Because the standard variable used in many other studies (= direct taxes of households relative to their pre-tax disposable income) is — with some variation — almost flat over our observation period, it is not surprising that a significant effect on a non-stationary dependent variable is not found when it is used.

[41] As wages are non-stationary and the unemployment rate, u, is close-to-stationary, the small coefficient is quite understandable. If u were uniformly an I(0) variable, it would hardly be significant in the long-run relationship.

Relation (4) is identical to (3) except that the CPI has been replaced by the private consumption deflator. It is easy to see that this has no impact on the preferred structure although the p-value has jumped to .97.

United Kingdom:

$$\log(W_{pr}) = .25\cdot\log\left(\frac{(PCP/P)(1+s)}{1-\tau_a}\right) \qquad \text{(ixx)}$$
$$+1.00\cdot\log(Q/N) - .10\cdot\log(u)$$

The Trace test suggests that the cointegration rank is five and that a linear trend is present in the data (see Appendix 3, Table A2 for the UK). Since the hypotheses r = 4 is not rejected with a particularly large margin[42] we consider also whether the choice of r influences the outcome.

The preferred structure in (ixx) passes the LR test with a p-value of .27. In the UK, CPI and PCP have tended to deviate systematically. Over the estimation period, the former has increased 6 per cent more than the latter. Specification including PCP gave a better overall explanation of the data generating process.

There are relatively modest signs of long-run real wage resistance. However, a look at the dynamic part of the estimation reveals that, in the short-run, higher employers' social security contributions add to labour costs by their full amount. A short-run response to income tax is considerable as well. Simulations indicating the adjustment are in Appendix 4.

The coefficient of the unemployment rate is fairly small but highly significant. Around two thirds of an equilibrium error is corrected within the first half-year[43]. Because the second "left-over" vector enters the dynamic part of the wage equation with a considerable loading, one cannot rule out bias in the estimation of the error correction coefficient. The dynamic equation which incorporates the long-run relationship (ixx) explains 76 per cent of the semester-to-semester variation in real wages.

The structure remains the same if cointegration rank is reduced to four. The p-value is .09 and the explanatory power of the dynamic equation is .70. Relation (2) for UK in Table A4 shows that the structure passes also if we choose a rank of three. Now the explanatory power has diminished further (to .55). So, although the bulk of the information related to wage setting is in the first three eigenvectors, the fourth and the fifth also incorporate some relevant pieces of information.

As far as the elasticities in Table 2 are concerned, one should recall that they do not predict the final impact on wages of a rise in taxes. This is because higher wages lead to higher unemployment. And higher unemployment leads to

[42] Please recall the result of Cheung & Lai (1993) which indicates that with small samples the asymptotic tests are potentially slightly biased towards finding cointegration "too often".

[43] Manning (1992b) finds an error correction coefficient which varies between .68 and .74 in various wage equations (which include wedge terms) although the framework, the estimation method as well as the data differ from ours in many important respects

wage adjustment. Accordingly, simulations discussed in Section 7 below (and in Appendix 4) show that in each country both employers and employees bear part of the increase in the tax burden. This is so also in Germany where the driving elasticity in the wage schedule is unity in Table 2.

In Table 3, countries have been characterized as ones with high, intermediate or low real wage resistance although these kind of rankings are always somewhat arbitrary. The criteria come from the β-coefficients in Table 2. If the sum $\beta_s + \beta_{P_c/P} + \beta_{\tau_a}$ is above two, we consider it as high degree of real wage resistance. When the sum is more than one but less than two, it is an intermediate case. If the sum is one or somewhat less, real wage resistance is low. If it were zero, there would be no sign of long-run real wage resistance.

At face value, the degree of real wage resistance has been particularly high in Germany and Canada. In the USA, Sweden and the UK, it has been fairly weak. Japan, Finland, Australia, Italy and France are intermediate cases.

Table 3 also reports the $\alpha_{W/P}$-coefficients of the preferred structures as well as the R^2's of difference equations like (v) above. For the reasons discussed above, α-coefficients should be emphasized qualitatively and too much should not be made of small differences between countries.

Wage response to a change of 1 percentage point in the unemployment rate is in the second column in Table 3. The sensitivity of wages with respect to unemployment is very similar in Germany, France and Italy. It is slightly higher in Australia and slightly lower in Canada. Our estimates seem in general to be well in the range of elasticities in studies referred to e.g. in LNJ, although in some countries our estimates are perhaps slightly higher than the "consensus estimates".

In Japan and in the two Nordic countries, the sensitivity to unemployment is particularly high. The incredible elasticity found for Sweden reflects the following facts. Between 1975 and 1983, real wages declined by almost 20 per cent[44] at the same time as the unemployment rate rose from around 1 1/4 per cent to 3 per cent. From 1985 to 1990 real wages grew more than 20 per cent at the same time as the unemployment rate fell from 3 to around 1 1/2 per cent.

When the matter is evaluated with the future in mind, caution is needed. This is because — as Lindbeck (1993, p.81) argues — in Sweden and Finland "the apparent sensitivity of the product wage to changes in unemployment rate ... was exhibited in the context of recurring discretionary devaluations rather than spontaneous market-induced product wage reductions by way of nominal wage moderation. It is open to doubt whether the fall in the product wage in these countries in connection with rather modest increases in unemployment can really be interpreted as high spontaneous sensitivity of the product wage to the unemployment rate".

The tiny coefficient found for the UK is in accordance with the view that wages are not particularly sensitive to unemployment in Great Britain. The interpretation of the small coefficient found for the USA is exactly the opposite. Wages seem to adjust quickly enough to allow the unemployment rate to

[44] Hourly wages in Swedish manufacturing industry declined much less than average wages in the private business sector. The difference appears at least partly to be due to implementation of part time employment schemes which may have played a more important role in other sectors than in manufacturing industry. For manufacturing industry, the unemployment elasticity of wages would therefore show up as considerably smaller than the estimates reported here.

fluctuate around a (more or less) stable mean and to cross this mean frequently. Still, these fluctuations seem to be sufficiently slow to lead to a rejection of stationarity in a formal test. In this light, it would be a mistake to consider the small coefficient as a sign of weak wage response. Rather, wages seem to adjust to shocks in third variables in such a flexible manner that unemployment is near to drop out from the long-run relationship although its short-run impact on wages is quite strong (see the simulation model for the USA in Appendix 4, Section 3).

Table 2 **The long-run response of real labour costs to changes in wedge factors**

	Employers' social security contributions	Consumption tax as part of the price wedge	Income tax[1]
Germany	1.0	1.0	1.0
Canada	0.8	0.8	0.8
Japan	0.5	0.5	0.5
Finland	0.5	0.5	0.5
Australia	0.5	0.5	0.5
France	0.4	0.4	0.4
Italy	0.4	0.4	0.4
Sweden	0.0	1.0	0
USA	0.0	0.0	1.0
UK	0.25	0.25	0.25

[1] As far the income taxation is concerned, the elasticities in this Table incorporate the implicit assumption that marginal and average rates move conjointly. This assumption is of importance as far as Japan, Canada, Finland and Italy are concerned.

Table 3 **Summary of 1) real wage resistance, 2) the response of the wage level to a 1 percentage point increase in the unemployment rate, 3) the share (= $\alpha_{W/P}$) of the deviation from the equilibrium which is corrected within the first half-year, and 4) the explanatory power of the wage equation**

	Degree of real wage resistance	Wage response to a one percentage point change in the unemployment rate	$\alpha_{W/P}$ (= share of an equilibrium error corrected in the first half-year)	R^2 of the dynamic equation incorporating the preferred long-run properties
Germany	high	– 3[1]	–.15	.79
Canada	high	– 2	–.16	.85
Japan	intermediate	– 5	–.43	.71
Finland	intermediate	– 6[1]	–.30	.77
Australia	intermediate	– 4	–.38	.62
France	intermediate	– 3 1/2[1]	–.34	.76
Italy	intermediate	– 3 1/2[1]	–.15	.81
Sweden	low	– 10[1]	–.08	.93
USA	low	– 1	–.20	.67
UK	low	– 1[1]	–.70	.76

[1] In the regression concerned, the unemployment rate (u) is in logarithmic form. The effect reported here has been evaluated at the mean unemployment rate of the estimation period.

6 Real Wage Resistance and Labour Costs

The estimated elasticities only tell part of the story since developments in the wedge have differed considerably between countries. Changes in tax structures may also have generated offsetting tendencies. When the elasticities discovered are analyzed in conjunction with actual data on the wedge, we obtain a more accurate picture of the impacts on wages.

In Table 4 and Figure 2, we have simply combined the estimated long-run elasticities with annual data on the wedge. Because all adjustment lags and feedbacks have been ignored, the results should be considered as "a first" look at the effect of real wage resistance on real labour costs. The effect seems to be particularly large in France, Canada and Finland. It is only slightly smaller in Sweden, Italy, Japan and Australia. In the rest of the countries, the effect is considerably smaller or even negative.

Of course, unemployment can be generated by factors which are not covered by our model. If a level relationship between the real wage and unemployment is found without any impact due to real wage resistance, the primary reason for the level shift in (equilibrium) unemployment is beyond our model. The UK could be an example.

If a shift in unemployment (which is consistent with each level of real wages) is generated by a rise in the wedge but persists after a reduction in the wedge, then "hysteresis" due to factors beyond our model is at work. Signs of this can be seen in Germany although the effects of reunification make strong conclusions difficult.

6.1 The contribution of taxation

There are two major components in the price wedge: the consumption tax and (a certain part of) the relative prices of imports (see e.g. LNJ, p. 210). When Table 5 was calculated, the price wedge was replaced by the consumption tax rate.

The contribution of non-tax element of the wedge varies dramatically between countries (see also Figure 2). The gap between the change in the price wedge and in the consumption tax is largest in Canada, Australia and Japan: 10–12 percentage points from 1975 to 1991. In the USA, France and Finland, it was only slightly less. In Italy, the UK and Sweden the rise in the price wedge was below the contribution of the consumption tax.

Germany is the only country – perhaps because appreciation of the exchange rate has moderated import price increases – where the price wedge has fallen (by 2 percentage points) despite an increase in the level of the consumption tax (by 3 1/2 percentage points). In a way, the exchange rate has created room for taxes to rise by offsetting the unfavourable price effects.

In France and Italy, taxes have played a first-order role. In Finland, the tax wedge has been growing almost systematically over the past ten years. In Sweden, the recent rise is considerable. In Canada, Japan, Australia, the USA and the UK, the overall increase in real labour costs due to a response to taxation is in the range of around ± 2 per cent.

Table 4

Increases in real labour costs due to real wage resistance in 10 OECD economies, percentage points

	From 1975 to early 80's	From early 80's to mid-80's	From mid-80's to 1991/92	From 1975 to 1991/92
Germany	0	+ 5 1/2	− 7	− 1 1/2
Canada[1]	− 1/2	+ 4 1/2	+ 5	+ 9
Japan[1]	+ 2	+ 3	+ 2	+ 7
Finland[1]	− 1	+ 3 1/2	+ 5	+ 7 1/2
Australia	+ 2 1/2	+ 3	+ 1/2	+ 6
France	+ 4	+ 4	+ 1	+ 9
Italy[1]	− 1	+ 5	+ 1/2	+ 4 1/2
Sweden	+ 2 1/2	− 1/2	+ 3	+ 5
USA	+ 5	− 2	− 1	+ 2
UK	+ 2 1/2	− 1/2	− 2 1/2	− 1/2

[1] The regressions for this country include a separate effect of a progressivity index, which — when increasing — tends to moderate wage claims generated by higher average income tax rates.

Table 5

Increases in real labour costs due to real wage resistance *with respect to taxes*[1] in 10 OECD economies, percentage points

	From 1975 to early 80's	From early 80's to mid-80's	From mid-80's to 1991/92	From 1975 to 1991/92
Germany	0	+ 3	+ 1	+ 4
Canada[2]	0	0	− 1/2	− 1/2
Japan[2]	− 1	+ 1	+ 2 1/2	+ 2 1/2
Finland[2]	− 2 1/2	+ 1 1/2	+ 4 1/2	+ 3 1/2
Australia	+ 1 1/2	+ 2	− 2	+ 1 1/2
France	+ 2	+ 2 1/2	+ 1	+ 5 1/2
Italy[2]	− 1	+ 5	+ 1	+ 5
Sweden	+ 2	+ 2 1/2	+ 2	+ 6 1/2
USA	+ 5	− 2	− 1	+ 2
UK	+ 2 1/2	− 1	− 1	+ 1/ 2

[1] The numbers summarize the joint effect of employers' social security contributions, income taxes and consumption taxes.

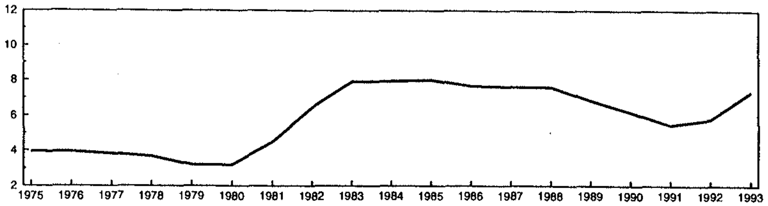
[2] The figures for the country concerned include the separate effects of marginal income tax rates, which — when increasing — tend to moderate wage claims generated by higher average income tax rates.

Figure 2.

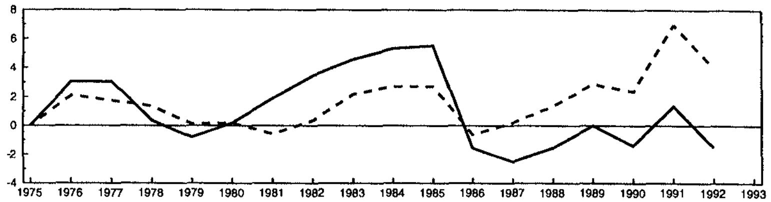
Unemployment and real wage resistance: A first look

GERMANY

Unemployment rate, per cent.

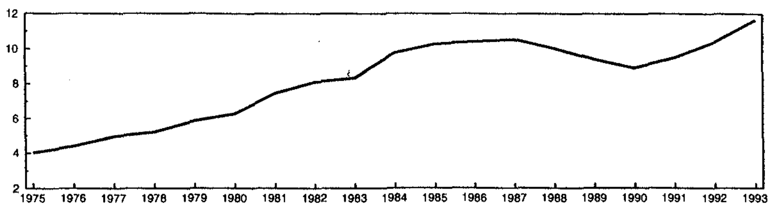


An estimate of the impact effect of real wage resistance on the level of real labour cost (——) and the contribution of taxation (- - -) when all lags and feedbacks are ignored. Percentage points with 1975 as the base year.



FRANCE

Unemployment rate, per cent.



An estimate of the impact effect of real wage resistance on the level of real labour cost (——) and the contribution of taxation (- - -) when all lags and feedbacks are ignored. Percentage points with 1975 as the base year.

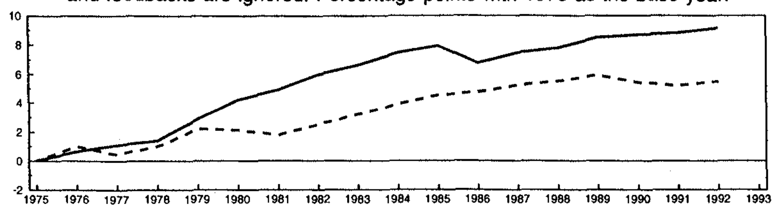
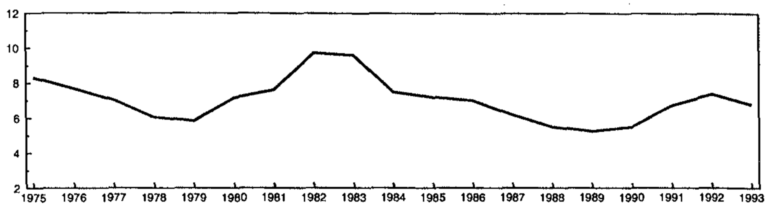


Figure 2. (cont.)

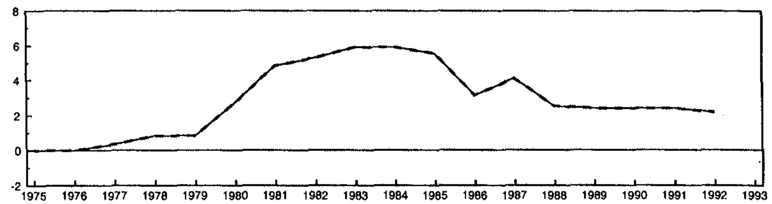
Unemployment and real wage resistance A first look

UNITED STATES

Unemployment rate, per cent.

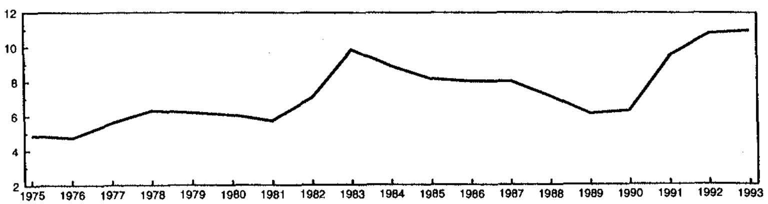


An estimate of the impact effect of real wage resistance on the level of real labour cost (——) and the contribution of taxation (- - -) when all lags and feedbacks are ignored. Percentage points with 1975 as the base year.



AUSTRALIA

Unemployment rate, per cent.



An estimate of the impact effect of real wage resistance on the level of real labour cost (——) and the contribution of taxation (- - -) when all lags and feedbacks are ignored. Percentage points with 1975 as the base year.

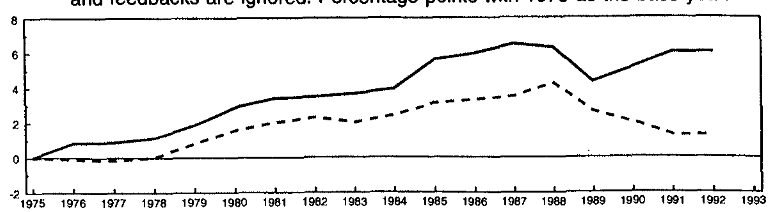
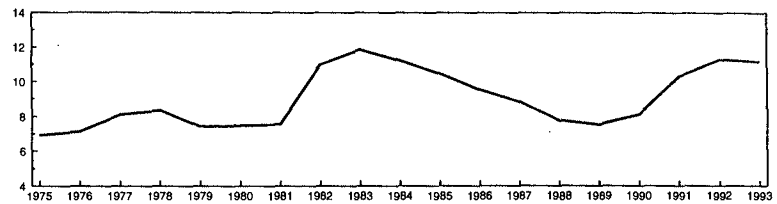


Figure 2. (cont.)

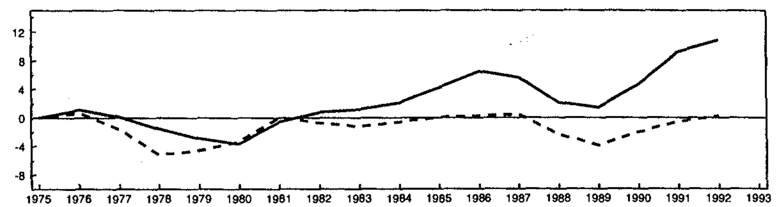
Unemployment and real wage resistance
A first look

CANADA

Unemployment rate, per cent.

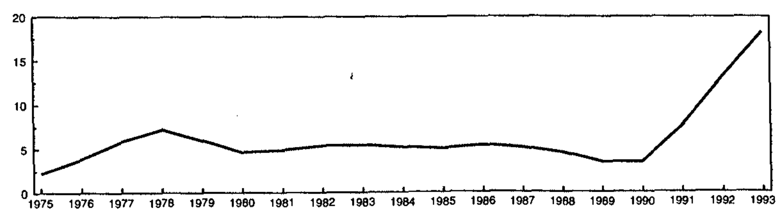


An estimate of the impact effect of real wage resistance on the level of real labour cost (——) and the contribution of taxation (- - -) when all lags and feedbacks are ignored. Percentage points with 1975 as the base year.



FINLAND

Unemployment rate, per cent.



An estimate of the impact effect of real wage resistance on the level of real labour cost (——) and the contribution of taxation (- - -) when all lags and feedbacks are ignored. Percentage points with 1975 as the base year.

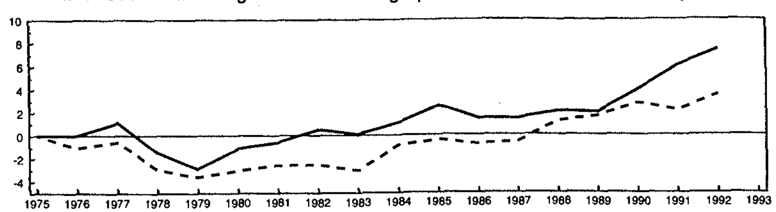
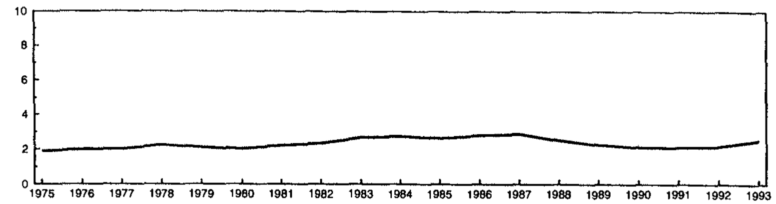


Figure 2. (cont.)

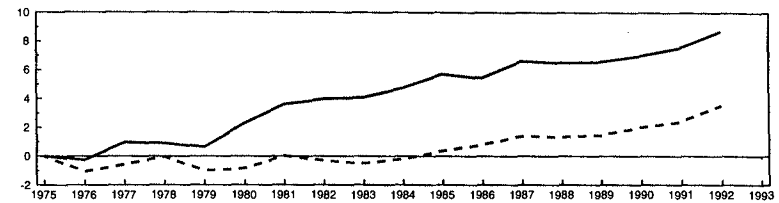
Unemployment and real wage resistance
A first look

JAPAN

Unemployment rate, per cent.

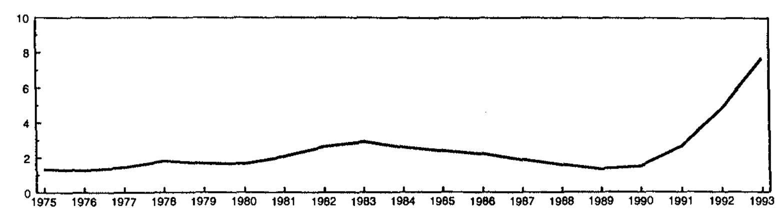


An estimate of the impact effect of real wage resistance on the level of real labour cost (——) and the contribution of taxation (- - -) when all lags and feedbacks are ignored. Percentage points with 1975 as the base year.



SWEDEN

Unemployment rate, per cent.



An estimate of the impact effect of real wage resistance on the level of real labour cost (——) and the contribution of taxation (- - -) when all lags and feedbacks are ignored. Percentage points with 1975 as the base year.

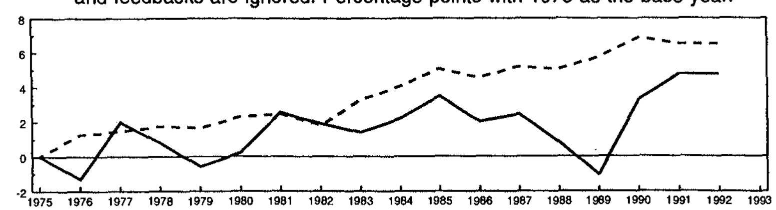
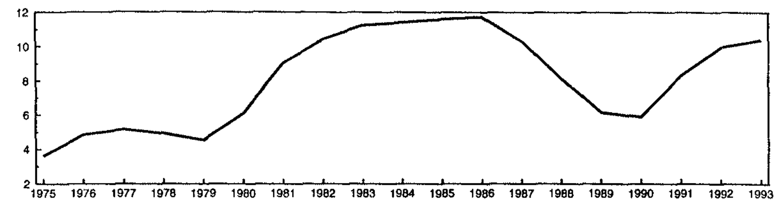


Figure 2. (cont.)

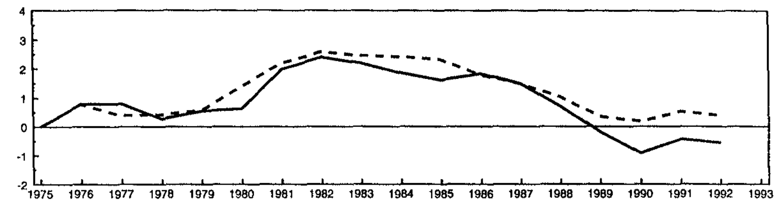
Unemployment and real wage resistance
A first look

UNITED KINGDOM

Unemployment rate, per cent.

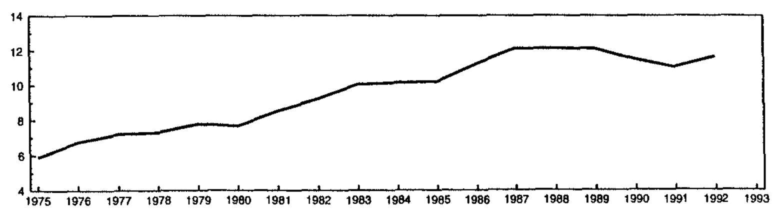


An estimate of the impact effect of real wage resistance on the level of real labour cost (——) and the contribution of taxation (- - -) when all lags and feedbacks are ignored. Percentage points with 1975 as the base year.

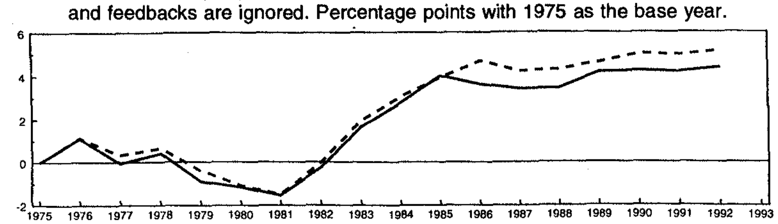


ITALY

Unemployment rate, per cent.



An estimate of the impact effect of real wage resistance on the level of real labour cost (——) and the contribution of taxation (- - -) when all lags and feedbacks are ignored. Percentage points with 1975 as the base year.



7 Some simulations

In this section we report the results of simulations[45] with models consisting of wage and employment equations as well as of an equation defining the response of unemployment to a change in employment. These three-equation models can be found in Appendix 4. The shocks correspond to changes in wedge factors which took place in the mid-70's[46] and the latest observation is in the early 90's. Because the simulation model contains just the key relationships in the labour market, and takes the rest of the economy as given, the estimates should be considered as qualitative[47].

The key results are in Table 6 which gives estimates not only for the full period but for some subperiods as well. This is because, as Figure 2 shows, in some countries there have been subtrends in the wedge which have presumably generated offsetting impacts on unemployment. Of course, the point estimates for subperiods should be treated even more cautiously because adjustment lags create overlapping.

In Figure 3, we illustrate for four countries the effects of shocks on 1) the real labour cost, 2) real take home pay, 3) employment, and 4) the unemployment rate.

As far as the effect of real wage resistance on real labour costs is concerned, results in Table 6 differ from the *ceteris paribus* estimates in Table 4 because the feedback effects via employment and unemployment have now been taken into account.

In the present section we do not distinguish explicitly between taxes and other contributing factors. This would not have added anything to the conclusions in Section 6 above.

[45] The simulations have been carried out using the "MAQUETTE" simulation software developed and kindly supplied by Dave Turner.

[46] The reason for choosing 1975 as the base year of the simulations is as follows. Developments in 1973—1975 were heavily dominated by the first oil shock, OPEC I, which took place in the latter half of 1973. A considerable part of its immediate effects seems to have taken place in the first two years after the shock: the number of unemployed people in the OECD area rose from 12—13 millions in 1972 to 14—15 millions in 1975 with some stabilization after that (OECD, 1994a). When 1975 is the base year, the effective changes in the wedge start to cumulate from the second half of 1975 onwards. This choice allows us to avoid an overwhelming dominance of OPEC I in discussions and to keep the main emphasis in an intra-OECD dimension.

[47] In a model for the full economy, a key adjustment channel runs from competitiveness to output via exports as well as via business sector confidence. In our model, allowing this would lower the long-run impact of tax shocks on labour costs because of weaker productivity and probably higher unemployment. Of course, in a more realistic model in which everything is endogenous, there are plenty of other adjustment channels as well.

Figure 3

Simulation results: France.

A simulation in which all elements of the wedge change simultaneously by the percentage they actually changed between 1975 and 1992.

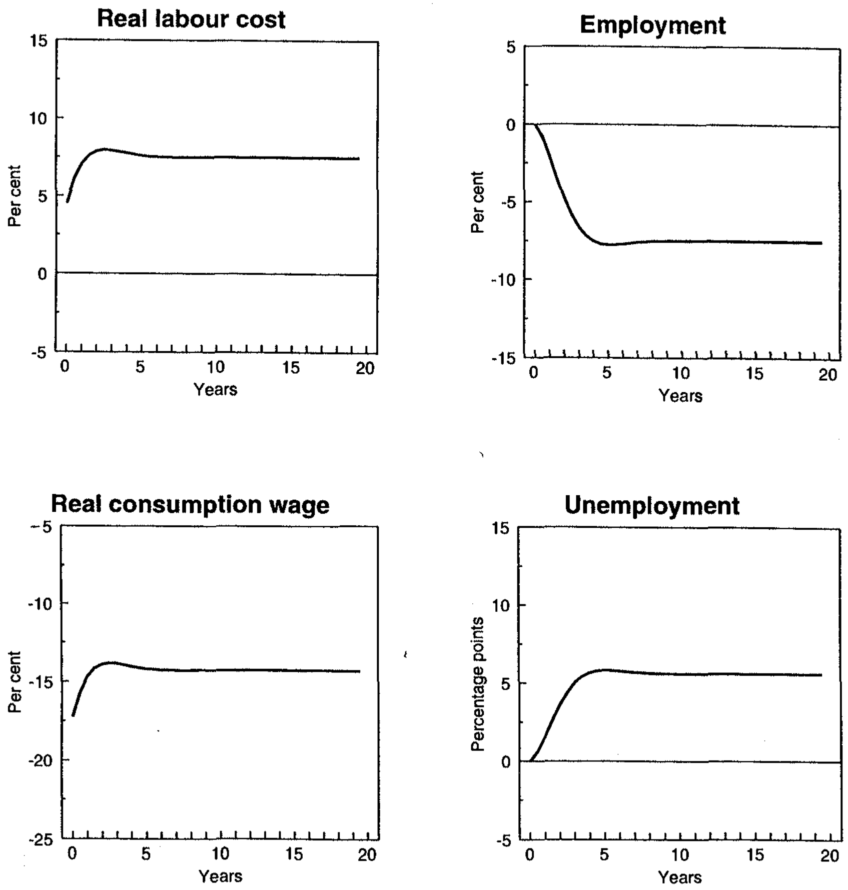


Figure 3 (cont.)

Simulation results: Canada.

A simulation in which all elements of the wedge change simultaneously by the percentage they actually changed between 1975 and 1992.

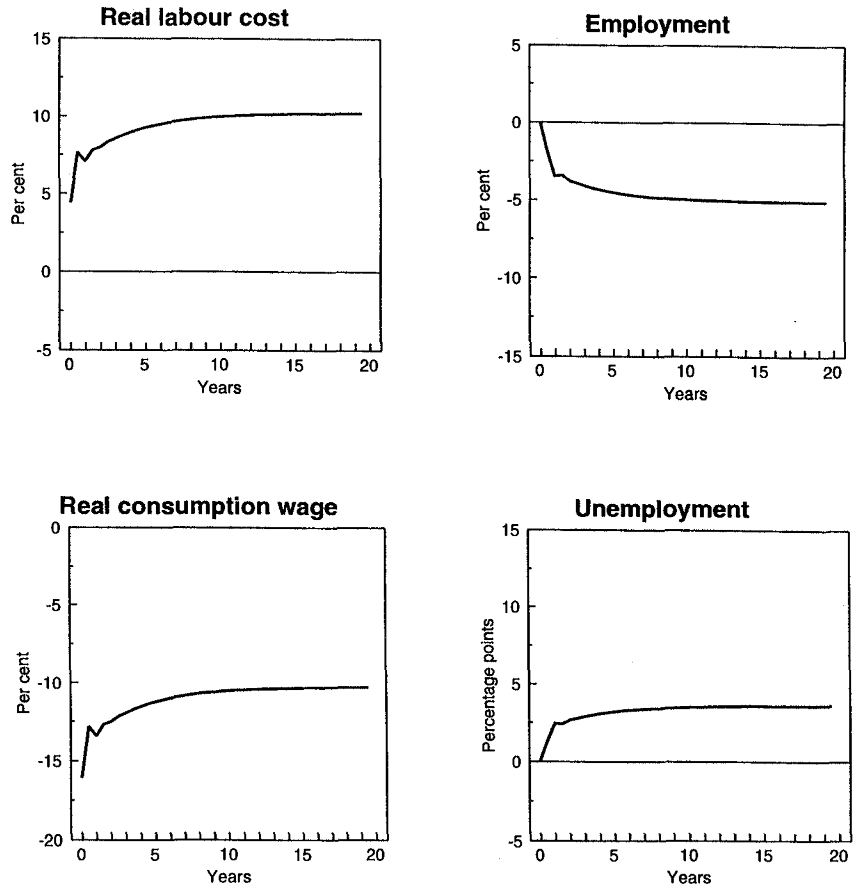


Figure 3 (cont.)

Simulation results: United States.

A simulation in which all elements of the wedge change simultaneously by the percentage they actually changed between 1975 and 1992.

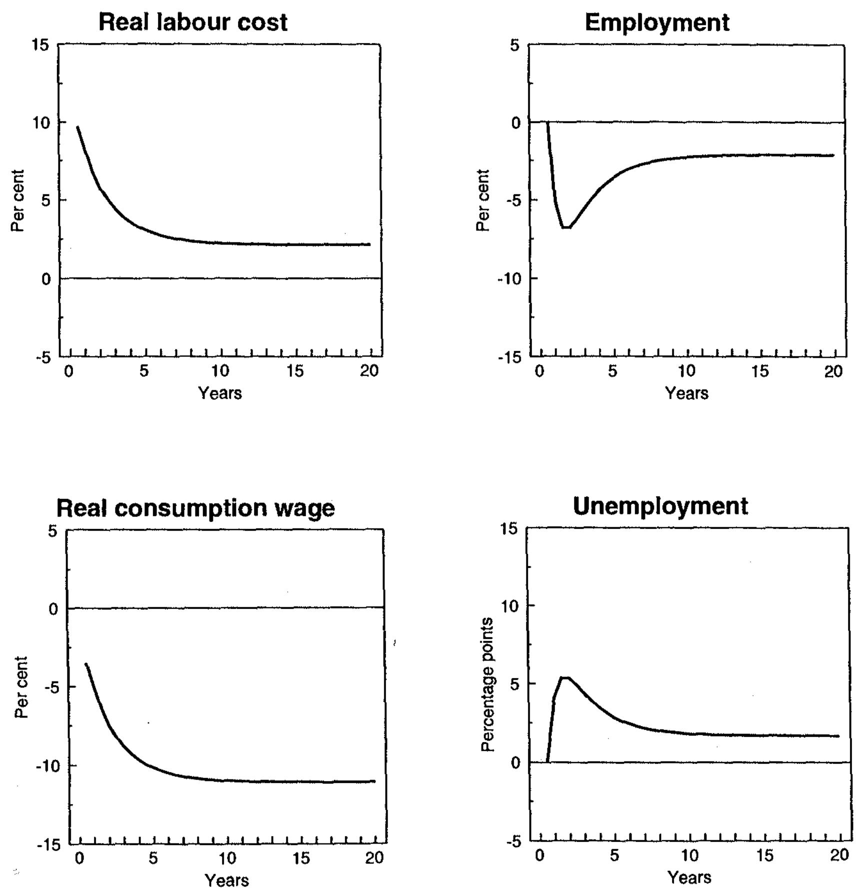


Figure 3 (cont.)

Simulation results: United States.

A simulation in which all elements of the wedge change simultaneously by the percentage they actually changed between

a) 1975 and 1983 (——),
b) 1983 and 1992 (- - -).

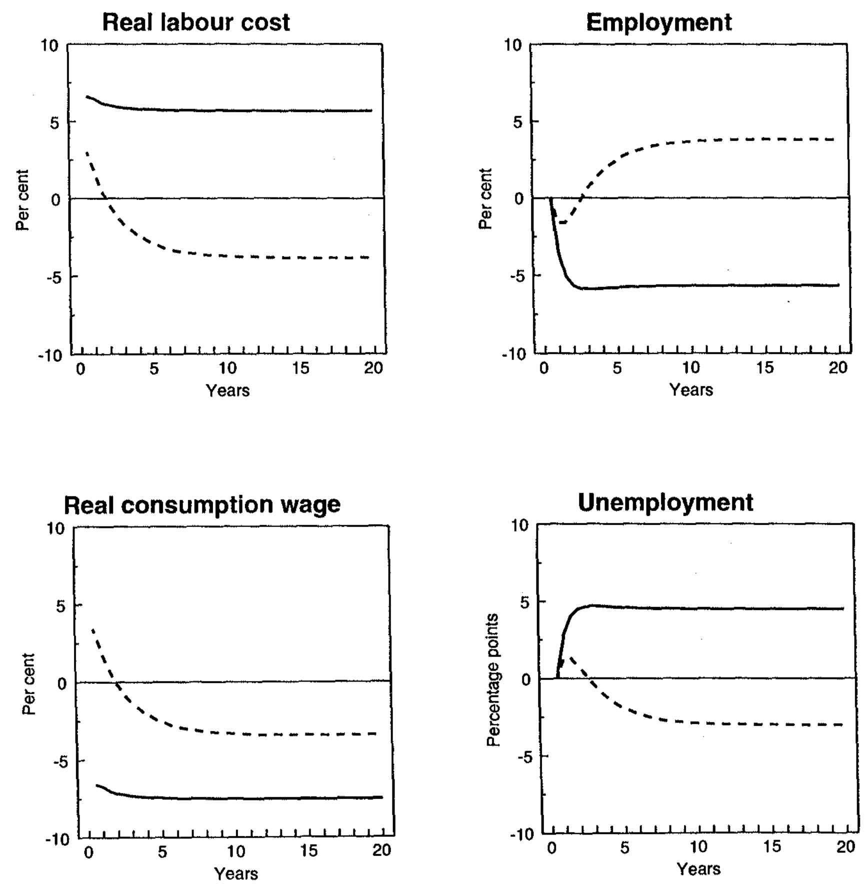


Figure 3 (cont.)

Simulation results: Germany.

A simulation in which all elements of the wedge change simultaneously by the percentage they actually changed between 1975 and 1992.

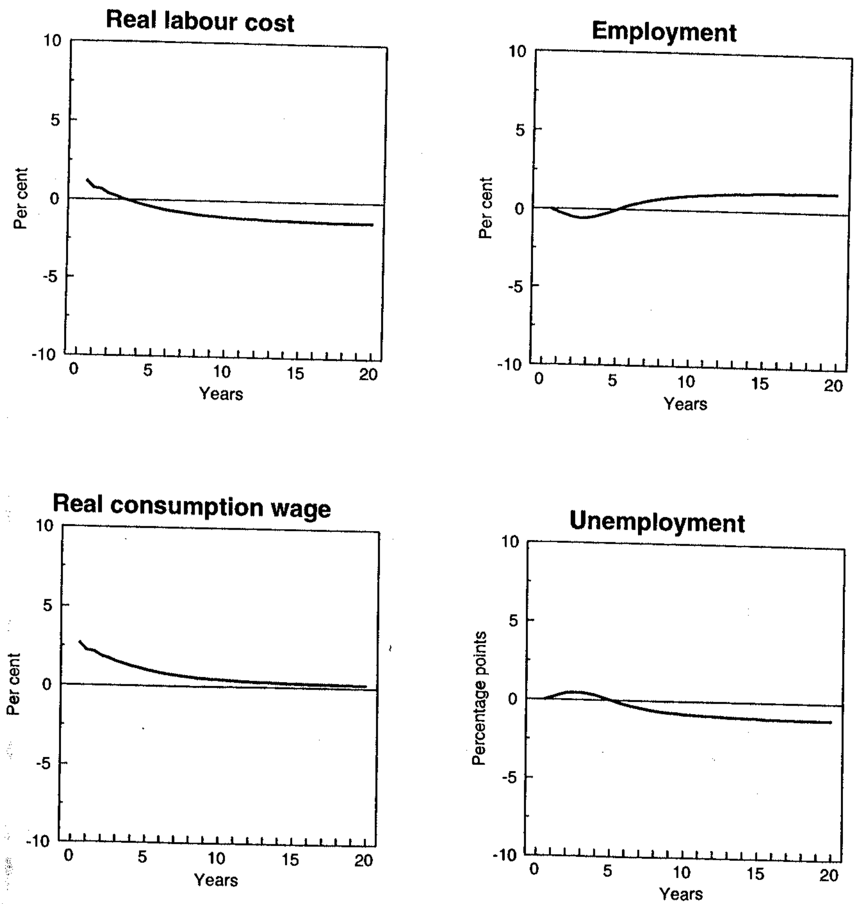
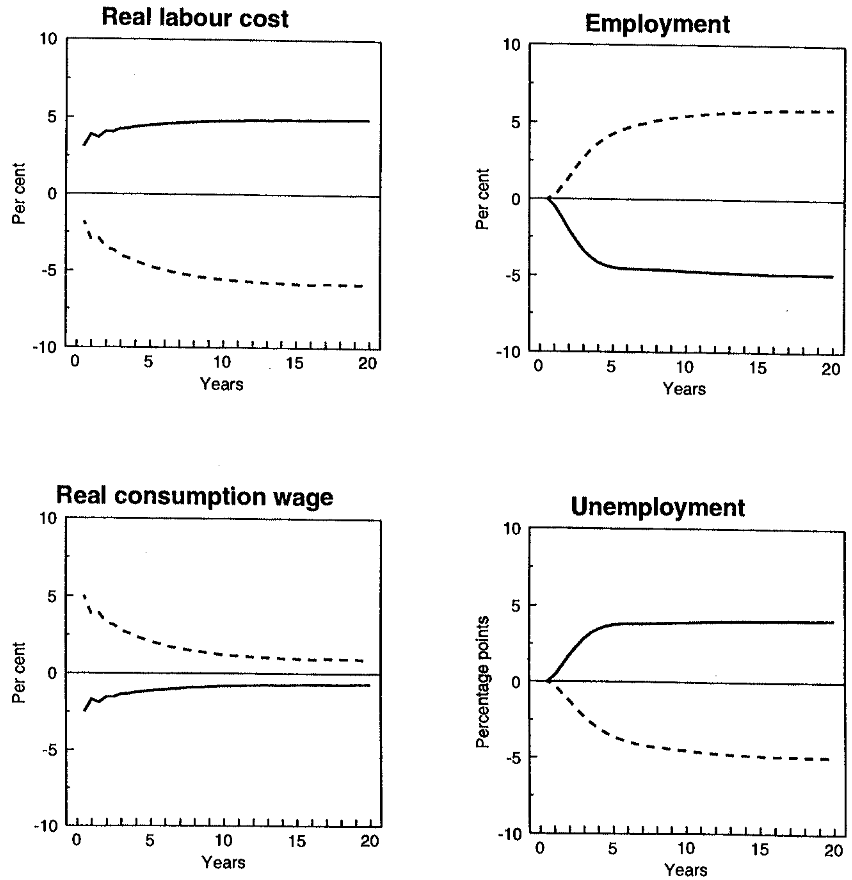


Figure 3 (cont.)

Simulation results: Germany.

A simulation in which all elements of the wedge change simultaneously by the percentage they actually changed between

a) 1975 and 1985 (——),
b) 1985 and 1992 (- - -).



A rise in the wedge reduces the size of the "cake" which is being bargained over[48]. For a given level of disposable income of the economy (or nominal GDP) the reduction in the "cake" is (b)–(a) with (a) and (b) as defined in Table 6. In France, the rise in the wedge has reduced the "cake" by more than 20 per cent since 1975. In Japan and Canada the amount is only slightly smaller. In the USA, Italy, Finland and Australia the reduction is 12–14 per cent. In Sweden and Italy, the increase in the wedge was considerably smaller. In Germany and the UK, the wedge was smaller in the early 90's than in the mid-70's.

The division of the burden can also be calculated. The share of the reduction in the "cake" which is borne by employees (in the form of lower real after-tax wages) is (b)/((b)–(a)). This share is lowest in Germany (around 20 %) and highest in the USA (85 %). In Italy the share is 70 per cent. In the rest of the countries it is around 60 per cent with two exceptions. In Canada half of the increase in the wedge and in Sweden 40 per cent has been absorbed by a reduction in real after-tax wages.

Table 6 **The simulated effect of a shock in which all wedge factors change simultaneously by the amount by which they actually changed in 1975–1991/92 on a) the real labour cost, b) real take home pay, c) employment, and d) the unemployment rate, percentage points**

	(a) Real labour cost	(b) Real take home pay	(c) Employment	(d) Unemployment rate
Germany	**– 1 1/4**	**+ 1/4**	**+ 1 1/4**	**– 1**
Pre –85	+ 5 1/4	– 3/4	– 5 1/4	+ 4 1/2
Post –85	– 6 1/2	+ 1	+ 6 1/2	– 5 1/2
Canada	**+ 10**	**– 10**	**– 5**	**+ 3 1/2**
France	**+ 8**	**– 13 1/2**	**– 8**	**+ 6**
Finland	**+ 6**	**– 8**	**– 6**	**+ 4**
Australia	**+ 5 1/2**	**– 7**	**– 5 1/2**	**+ 4**
USA	**+ 2**	**– 11**	**– 2**	**+ 1 1/2**
Pre –83	+ 6	– 7 1/2	– 6	+ 4 1/2
Post –83	– 4	– 3 1/2	+ 4	– 3
Sweden	**+ 4**	**– 3**	**– 3 1/2**	**+ 2 1/4**
Pre –83	+ 1	– 9 1/2	– 1	+ 3/4
1983–89	– 2	– 3	+ 2	– 1 1/2
Post –89	+ 5	+ 10	– 4	+ 3
Japan	**+ 7**	**– 12**	**– 6**	**+ 1 1/2**
UK	**– 1**	**+ 1 1/2**	**+ 1**	**– 3/4**
Pre –83	+ 2	– 7 1/2	– 2	+ 1 1/4
Post –83	– 3	+ 9	+ 3	– 2
Italy	**+ 4**	**– 10**	**– 2**	**+ 1 1/2**

[48] Evaluation of the fact that higher taxes tend to go hand-in-hand with an increase in public services is beyond the scope of the present paper.

In countries where the Cobb-Douglas function was found to be an appropriate description of the technology, (c) = –(a) holds because the wage elasticity of demand for labour is unity. In Canada, Italy, Japan and Sweden, the Cobb-Douglas was rejected but a CES function passed. With CES, the wage elasticity is below unity and therefore (c)<(a) in Table 6. This difference is particularly important in Canada and Italy.

Finally, the right-hand side column reports changes in the unemployment rate when feedback effects have been taken into account. As indicated in Appendix 4, in all countries labour supply — or perhaps public sector employment — responds to changes in employment. Because of this (d) < –(c) in all countries. This difference is particularly large in Japan.

Although our model is not meant to be a full explanation for changes in unemployment — far from it — it is still interesting to evaluate how its predictions fit with the actual history of unemployment.

In the ***USA***, the estimated effect of real wage resistance on the unemployment rate is + 4 1/2 percentage points from 1975 to 1983. The actual change in the unemployment rate from the level of 4 1/2 — 5 1/2 per cent prevailing in the early 1970's to 9 1/2 per cent in 1982—83 is of a similar order. Afterwards a reduction of 3 percentage points due to a decrease in the wedge should have taken place. In real life, the unemployment rate in 1992 was around 2 1/2 percentage points below the level of 1982—83 and around 1 1/2 percentage points above the level of the early 1970's. This is in accordance with the prediction in column (d).

In ***France***, the estimated effect of real wage resistance on the unemployment rate is 6 per cent for the full period. The actual change from 4 per cent in 1975 to around 10 per cent in 1992 is of the same order. In ***Canada*** the predicted impact of 3 1/2 percentage points corresponds to the observed rise of 4 percentage points from 7 to 11 per cent. Because of the considerable increase in the wedge since 1989 — and because of adjustment lags — some unfavourable effects on unemployment may still be in the pipeline in Canada.

For ***Australia***, the estimated effect on the unemployment rate is 4 per cent. In real life, the unemployment rate has risen slightly more, from 5 per cent to 10—11 per cent. In ***Finland*** we find an estimated wedge effect of 4 per cent, most of which dates from the final years of the observation period. In 1991, the unemployment rate in Finland was 7 1/2 per cent, which is around 5 percentage points above the level of the early 1970's and 3—4 percentage points above the level of the late 1980's. In 1992—93, the unemployment rate rocketed and was around 20 per cent at the end of 1993. This can be linked to the dramatic fall in real GDP, which amounted to 13—14 per cent in 1991—93[49]. However, one can hardly avoid the conclusion that real wage resistance has contributed to the recent rise in unemployment. Because of adjustment lags, it is probable that the unfavourable effects have not fully materialized yet.

In ***Japan***, the estimated effect of real wage resistance on the unemployment rate is 1 1/2 percentage points for the full period. This estimate is low in comparison with the large effect found on real labour costs and

[49] The fall in real GDP can be attributed to the collapse of trade with the former Soviet Union, to the excessive (foreign) indebtness of domestic agents and to the banking crisis, which has been exacerbated by the first two factors.

reflects the strong response of labour supply (see Turner et al, 1993, and Elmeskov and Pichelman, 1993). In real life, the unemployment rate doubled from around 1 1/2 per cent prevailing in the early 1970's to almost 3 per cent in 1986—87. The explanation for the decline in unemployment afterwards is beyond our model. More recently, continuing weak demand has been producing clear signs of rising underemployment either in the form of labour hoarding or withdrawal, especially of women, from the labour force (see OECD, 1994a).

In ***Sweden***, real wage resistance was estimated to have increased the unemployment rate by around 3/4 percentage points from the early 1970's to 1983 and the actual rise is around 1 percentage point. For 1983—89, the model predicts a decline of 1 1/2 percentage points, which is in line with the actual decline from 3 1/2 per cent in 1983 to 1 1/2 per cent in 1989. After that, a sharp increase in unemployment is predicted which fits facts. It is probable that all the effects due to real wage resistance have not yet worked their way through.

One of the most interesting cases concerns ***Germany*** where the wedge increased considerably up to 1985 and declined thereafter. The estimated impact on unemployment is + 4 1/2 percentage points in the first period. The actual change was some 5 1/2 percentage points (from 1 — 1 1/2 per cent prevailing in the first half of the 1970's to around 7 per cent in the mid-80's). Thereafter, the decline in the wedge exceeds the earlier rise with an effect on the unemployment rate which should be –5 percentage points. The actual decline has been around 3 percentage points. The reason for the smaller-than-predicted decline in unemployment is beyond our model. Insider power may have generated obstacles to wage moderation. Reunification of Germany has also played a role.

In ***Italy***, the unemployment rate has risen by around four percentage points, from 7—8 per cent in the latter half of the 1970's to the level of 11—12 per cent which has prevailed in most years since the mid-80's. The estimated impact of wedge factors is +1 1/2 percentage points. Most of the relevant rise in the wedge took place between 1981 and 1985[50].

In the ***UK***, the unemployment rate rose from the level of 3—4 per cent prevailing in the early 1970's to around 12 per cent in the early 1980's. According to our estimates, real wage resistance contributed one percentage point by. Afterwards an opposite effect of real wage resistance of the order of –2 percentage points relates to the actual decline in unemployment by –3 1/2 percentage points (from 12 1/2 per cent in 1983 to 8 3/4 per cent in 1991). The

[50] Italy is among the countries that was worst hit by the recession which began in the late 70's. GDP growth virtually stagnated for three years in 1981—83. In general, a complication related to the interpretation of the Italian results is that most of the rise in the unemployment rate is due to an increase in female and youth unemployment. The difference between male and total unemployment rates has increased steadily almost throughout. Youth unemployment accounted for 50—60 per cent of total unemployment in the 1980's. Male unemployment as a percentage of male labour force was 4.0 per cent on average in 1968—73, 4.2 per cent in 1974-79 and 6.6 per cent in 1980—89.

explanation for the rise in unemployment in 1975–83 is beyond our model as too is its persistence[51].

Replication of Figure 1 allows us to discuss these results in terms of shifts in the wage setting schedule[52].

In Figure 4, the upper part could describe France, Finland, Australia, Canada, Japan and Italy. In these countries the dominating change has been an upward shift in the wage setting schedule from WSS_0 to WSS_1. The resulting change in equilibrium unemployment is $u^*_1 - u^*_0$. The magnitude of the change is indicated in column (d) in Table 6.

The lower part of Figure 4 describes the USA, Germany, Sweden and the UK. In Germany we first find a shift from WSS_0 to WSS_1 and then back to WSS_0 or even slightly below it. The resulting change in equilibrium unemployment should be from u^*_0 to u^*_1 and then back to u^*_0. As indicated above, the latter part of the process is still on its way in Germany.

In the USA, we first see a shift from WSS_0 to WSS_1 and then to WSS_2, which is above the original position. The resulting change in equilibrium unemployment should be from u^*_0 to u^*_1 and then to u^*_2. As indicated above, our estimates indicate an increase of 4 1/2 per cent in the equilibrium unemployment rate in the pre-1983 period and a reduction of 3 percentage points thereafter. The resulting level should exceed the original equilibrium by 1 1/2 percentage points. This is well in accordance with basic trends in the actual unemployment rate in the USA.

In Sweden, we first predict a shift from WSS_0 to WSS_2. Then a shift back to WSS_0 (or below it) should follow. Finally, a strong upward shift to WSS_1 should take place. The first two shifts are estimated to be small. The actual variation in unemployment in Sweden as well as its magnitude seem to fit well with our predictions. This also applies to the recent rise in the unemployment rate[53].

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[51] The United Kingdom was particularly severely hit by the recession which started at the end of the 70's. Aggregate GDP in 1981 was 3 1/2 per cent below the level of 1979. The decline in manufacturing output from 1978 to 1982 was around 15 per cent. This must be the core of an explanation for the increase in unemployment in these years.

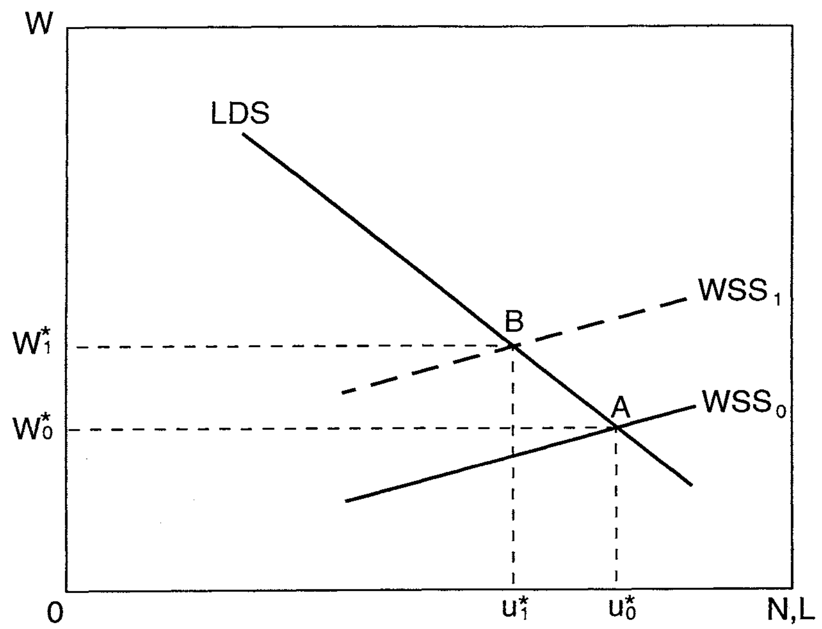
[52] It should be recalled that when the FIML estimation method proposed by Johansen (1991) is used, a (wage) relationship is considered as a long-run cointegrating vector only if it is time-invariant. When the test "approves" the existence of such relationships, it implies that the role of structural breaks, institutional changes etc. is of secondary order.

[53] To be frank, this point is complicated by the fact that the real labour cost has not risen in recent years in a manner predicted by the increasing wedge. On the other hand, this may underline the possibility that some unfavourable effects are still in the pipeline.

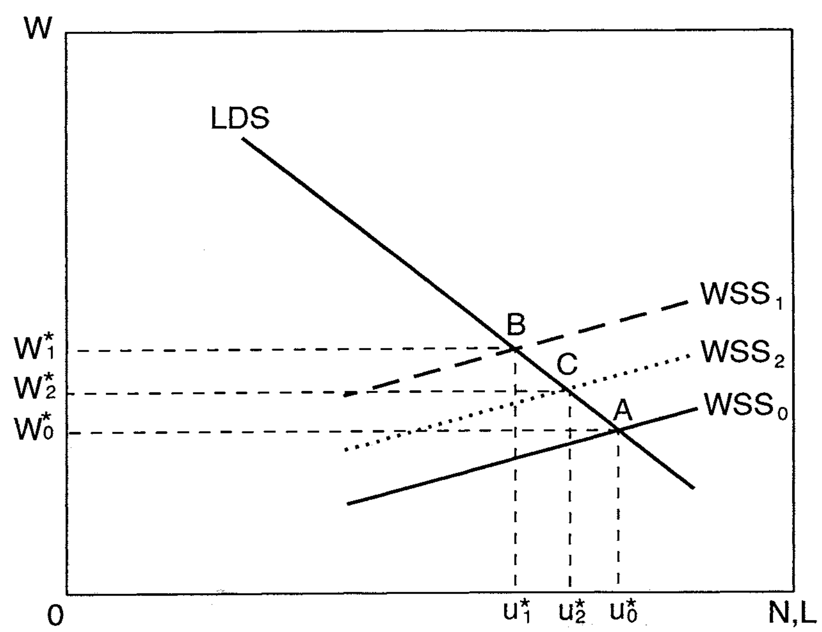
Figure 4.

Wage setting, demand for labour and equilibrium unemployment revisited

A. A stylized model for France, Canada, Japan, Finland, Australia and Italy.



B. A stylized model for the USA, Germany, UK and Sweden.



7.1 An excursion: A rise in employers' social security contributions

Let us now consider the impact of a rise in employers' social security contributions more thoroughly. Figure 5 presents simulated effects of a rise in the relevant tax rate in the USA, Canada and France. In Canada, the long-run effect on labour cost was estimated to be .8. In France, the estimate is .4 . Although the long-run elasticity was found to be zero in the USA, it takes ten years for the effect to die out fully[54]. A marked effect on (un)employment can still be observed after five years of adjustment[55].

In all countries, in the first, second and third year the unemployment rate exceeds the control solution. In France, adjustment is slower but its path is similar to that in the USA. The divergence between the three countries becomes visible in the longer run. So, in countries with highly different long-run effects the short-run effects may have much in common.

In comparison with these countries, Sweden is at the other extreme in that not only is the long-run effect on real labour cost nil but the short-run effect as well. We believe that this is because (in central bargaining) wage setters have taken account of changes in indirect labour costs[56].

7.2 An excursion: Revenue-neutral shifts in the tax mix

As discussion on revenue-neutral reforms in the tax mix is going on in many countries, we present simulations related to three alternative policy programs for four countries (Germany, France, Canada and Finland). Figure 6 examines a shift from income tax to employers' social security tax. Because the tax base is identical (i.e. the wage bill) in both cases, the shock has been specified as a cut of 1 percentage point in the former with an identical rise in the latter. The impact on (un)employment is nil in the long run but the short-run effect, which lasts for five years or more, is unfavourable.

Figure 7 describes a shift from employers' social security tax to consumption taxes. The tax base related to the latter is much wider than the wage bill. In 1991, wage income was 56 per cent of total income in Germany, 52 per cent in France, 64 per cent in Canada and 61 per cent in Finland. Let us

[54] To give an example, let us consider a hypothetical increase of 7 percentage points in employers' social security contributions. In our model, it leads to an immediate increase of 7 per cent in the real labour cost in the USA. Three years later around half of this effect is still there. After the shock, employment starts to decline with some lag. In the second and the third years after the shock, unemployment is 3 – 3 1/2 percentage points higher than in the control solution. Although unemployment declines thereafter, a considerable effect can still be observed five years after the shock occurred.

[55] As far as the simulated adjustment path of the unemployment rate in the USA is concerned, the shape of the curve in Figure 5 is almost identical to the one reported in Turner et al.(1993, Figure 6) although adjustment is slightly slower in Turner et al.

[56] In recent years, centralized wage setting has lost ground in Sweden. Therefore, extrapolation of this reasoning may be inappropriate where future behaviour is concerned.

assume a reduction of one percentage point in the social security tax. To cover the revenue loss, the government raises consumption taxes by .56 per cent in Germany, by .52 per cent in France, by .64 per cent in Canada and by .61 per cent in Finland. The impact on (un)employment is favourable both in the long run and in the short run.

Figure 8 reports simulations in which both income tax and employers' social security tax have been cut by half a percentage point. So, these cuts sum to unity. Because the rise in consumption tax is less than unity, real labour costs fall and real take-home pay increases. The effect on (un)employment is favourable. In Germany and France, overshooting lasts for five years or so. In the long run, the employment effect is biggest in Germany whereas the other countries seem to converge towards a similar impact.

Of course, the favourable impacts predicted are only valid on the assumption that the government allows the purchasing power of pensions, unemployment benefits and social transfers to fall as a consequence of the rise in the consumption tax.

Finally, we once more stress that in estimation we have put much more weight on analysis of long-run relationships than on analysis of short-run dynamics. Accordingly, as far as short-run adjustment is concerned, the message of policy simulations is qualitative rather than quantitative.

Figure 5

Simulation results: a rise in employers' social security tax

The simulated effect of a one percentage point rise in the employers' social security contribution rate.

United States: ——
Canada: - - - -
France: ....

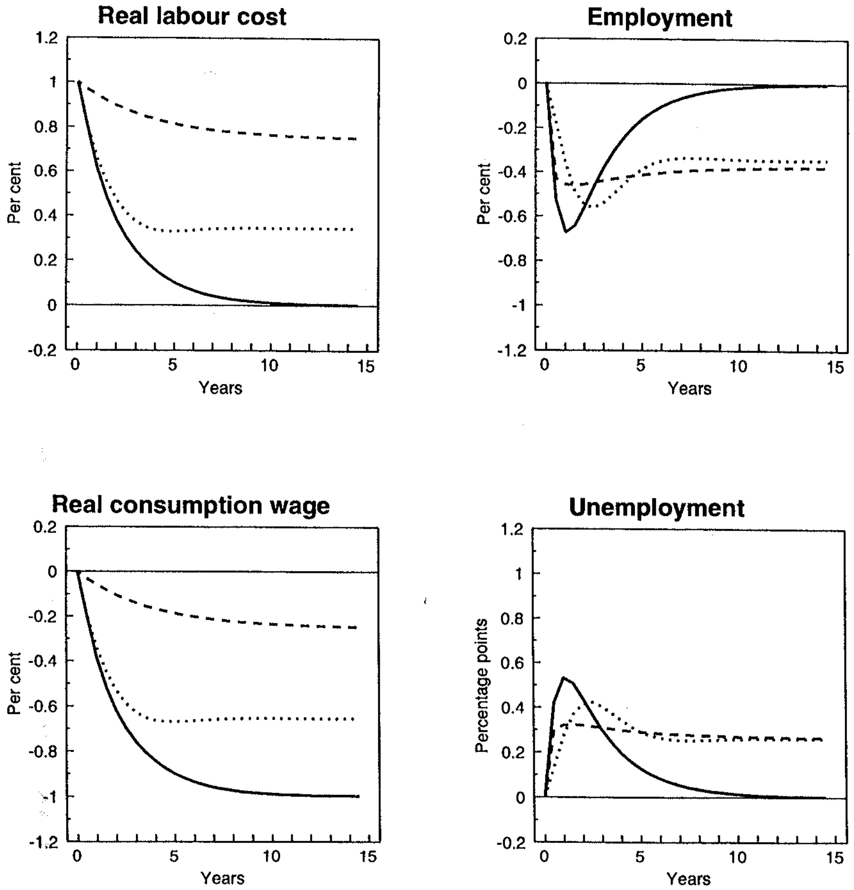


Figure 6

Simulation results: a revenue neutral shift from income tax to employers' social security tax

The simulated effect of a simultaneous one percentage point cut in the income tax rate and one percentage point rise in the employers' social security contribution rate.

Germany: ——
Finland: - - - -
France: ......
Canada: —·—·—

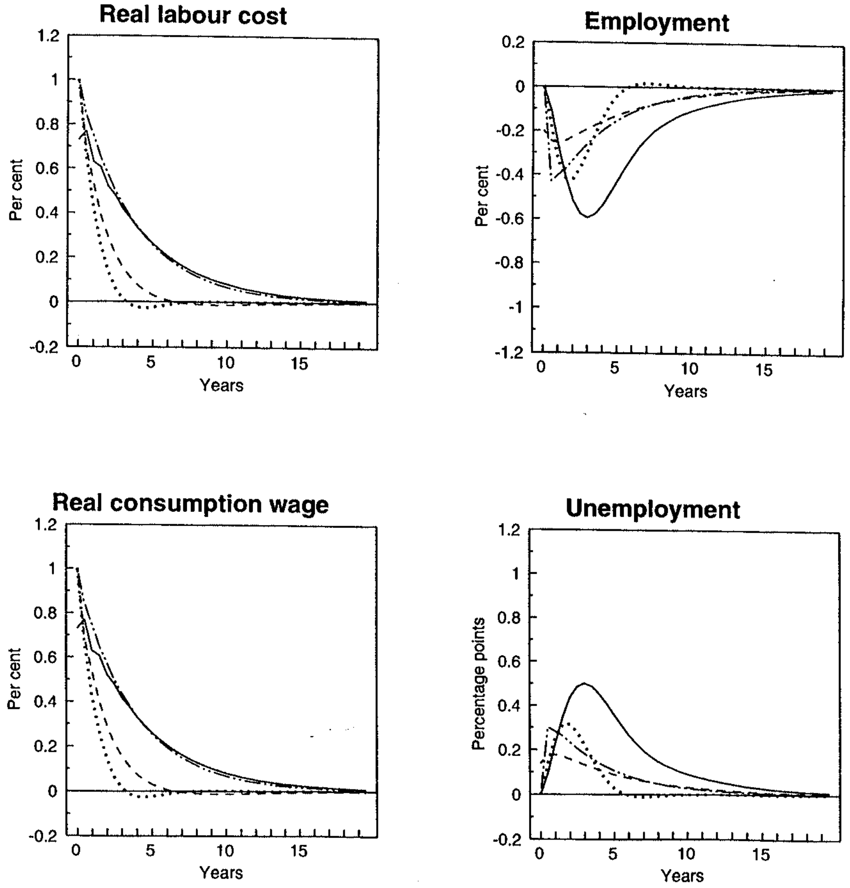


Figure 7

Simulation results: a revenue neutral shift from employers' social security tax to consumption tax

The simulated effect of a simultaneous one percentage point cut in the employers' social security contribution rate and a revenue-neutral rise in the consumption tax.

Germany: ——— France: ·······
Finland: - - - - Canada: -··-··-

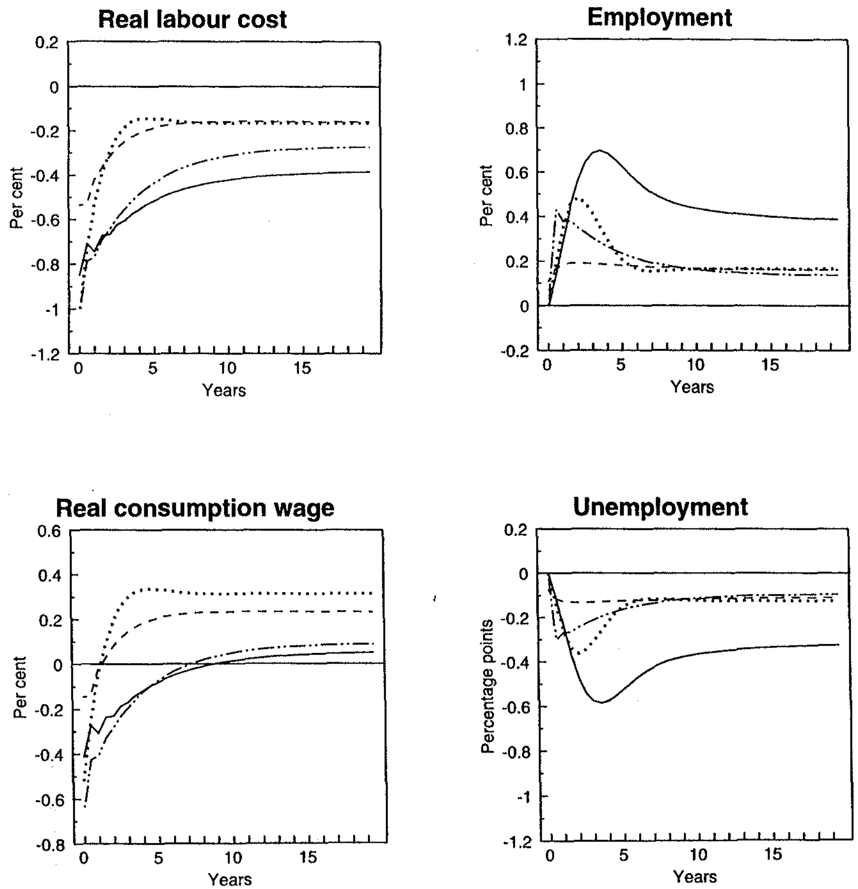
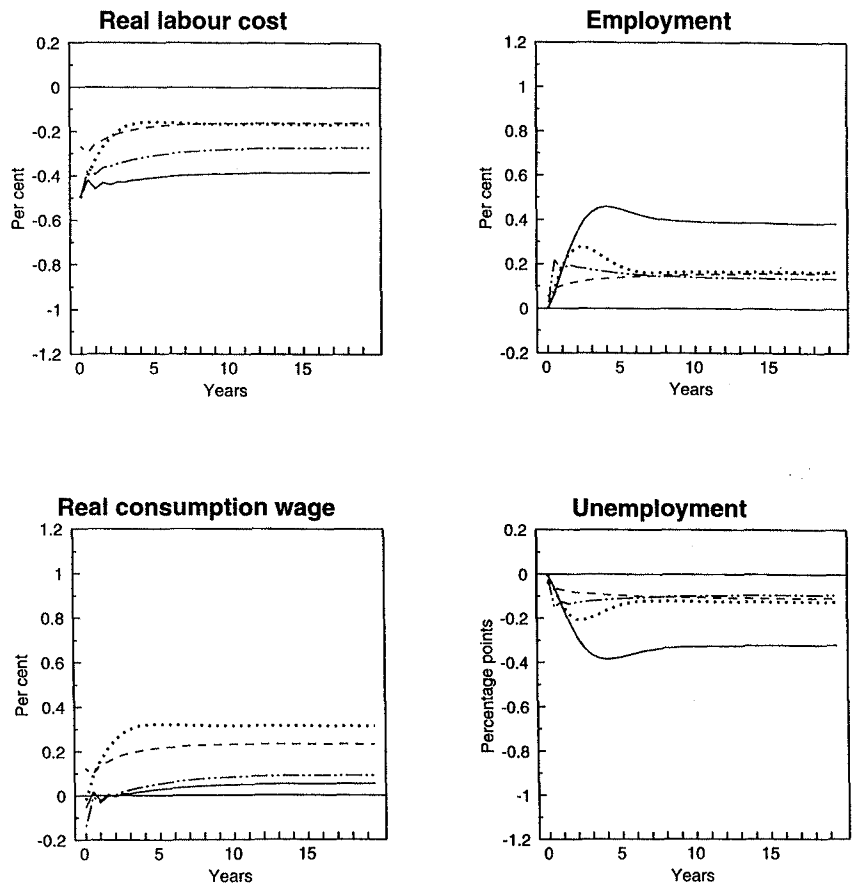


Figure 8

Simulation results: a revenue-neutral shift from income taxes to consumption taxes

The simulated effect of a half-a-percentage point cut in both the income tax rate and the employers' social security contribution rate and a simultaneous revenue-neutral rise in the consumption tax rate.

Germany: ——— France: ·······
Finland: - - - - Canada: -··-··-



8 Summary

We have examined whether changes in the wedge (including tax rates) may have generated effects on real labour costs which have resulted in unfavourable (un)employment outcomes in the countries concerned. Behaviour which generates such results has been called "real wage resistance". The countries examined are the United States, Japan, Germany, France, the United Kingdom, Italy, Canada, Australia, Sweden and Finland and the study covers wage earners in the private business sector. The main interest is in the long run, i.e. in persistency of the effects, although adjustment is also analyzed by means of dynamic simulations.

The estimation method is the FIML-procedure proposed by Johansen (1991b). A 7-dimensional vector space is defined which includes the real wage and labour productivity, a measure of unemployment and four wedge variables. Whenever accepted by the data, the full Vector Autoregressive model is reduced by means of conditioning, i.e. by considering part of the variables as weakly exogenous. Structural restrictions which identify a time-invariant long-run wage setting schedule are specified, imposed and tested. No structure has been imposed *a priori.* Generally, identifying restrictions pass at fairly high significance levels and generate coefficient estimates with plausible economic interpretation.

The results can be summarized as follows:

a) in all countries, the conjecture of long-run homogeneity between real wages and productivity passes the tests.

b) Signs of real wage resistance can be found in all economies although its degree differs between countries. These differences can at least partly be related to the characteristics of the wage setting patterns in the countries concerned.

c) In eight out of the ten countries, a hypothesis of identical long-run response to changes in all relevant tax rates passed the test. This is in accordance with the conjecture of (long-run) irrelevance of *de jure* incidence. In the short run, however, there are important differences. A cut in the employers' social security tax combined with an offsetting rise in other tax rates tends to ease unemployment in the short run although the beneficial effect vanishes as the adjustment proceeds.

d) As far as adjustment lags are concerned, simulations related to shifts in wedge variables gave two results of interest. First, differences in the speed of adjustment are considerable and in some countries lags are fairly long. Second, long lags are in some cases primarily due to slow response of wages to changes in the wedge and in some other cases due to slow response of employment to higher wages.

e) Generally, higher wages have also in the long run compensated at least a fraction of upward changes in the price wedge (which contains the effects of consumption taxes). In the USA, the long-run effect is nil despite a strong short-run response. In Canada, Germany and Sweden one cannot reject the hypothesis of full long-run wage compensation.

f) In all countries, wages have responded to changes in income tax rates. A separate effect related to the progressivity of the tax system could be distinguished in Canada, Japan, Italy and Finland. In these countries, evidence indicates that steeper progressivity has tended to reduce wage claims.

g) In the USA and Sweden, wages have fully absorbed changes in employers' social security contributions. Although no long-run effect on real labour cost remains, a substantial temporary effect lasting several years was found in the USA. Higher social security taxes induced higher labour costs in Germany and Canada and to a lesser extent in Japan, Finland, Australia, France, Italy and the UK.

h) In all countries, the data definitely reject omission of an unemployment variable from the wage setting schedule. Thus, there appears to be a level relationship between real wages and unemployment. In European countries, the unemployment rate expressed in a logarithmic form accorded somewhat better with the structure of the information set whereas the unemployment rate as such was the more appropriate measure in the USA, Canada, Australia and Japan. Relevant elasticities are not sensitive to the choice of unemployment measure. In addition to the unemployment rate and log(unemployment rate) we experimented with the log of the number of unemployed persons.

i) When the estimates of long-run elasticities are analyzed in conjunction with actual data from the mid-1970's to early 1990's, real wage resistance seems to have had a particularly large effect on real labour costs in France and Canada. It is also high in Australia, Finland and Japan, and the impact is only slightly smaller in Italy and Sweden.

In the USA, Germany and the UK, the effect for the full period was small or negative. However, even in these economies there were subperiods during which real wage resistance seems to have generated deviations of real labour costs from labour productivity. This is particularly important for the analysis of unemployment in the USA and Germany.

j) According to the simulations, the impact of real wage resistance on unemployment differs considerably between countries. In the USA and Germany, real wage resistance may have contributed strongly to the increase in unemployment in the late 1970's and the early 1980's. The decline in unemployment thereafter also fits with predictions. In the USA the magnitude of the decline is in accordance with our estimate whereas in Germany the reduction over the past few years is smaller than predicted. Presumably, this is because of factors beyond our model which have influenced German labour markets over the past ten years or so.

Over the full estimation period, the impact of real wage resistance on the unemployment rate is particularly unfavourable in France. It is also important in Australia, Canada, Finland and Sweden. In the latter two, the increase in the wedge has been particularly large in recent years. Because of adjustment lags, this may imply that some unfavourable effects are still in the pipeline in Finland and Sweden.

In Canada, the unfavourable impact of real wage resistance is not due to taxation. Taxes have not been a primary factor in Japan either whereas in France, Italy and Sweden taxes have played a major role. In Finland, large tax rises have taken place in recent years. In Germany, the exchange rate appreciation has "created room" for consumption taxes to rise without any harmful effects due to real wage resistance. In other countries, the contribution of taxation to changes in the real labour cost is in the range of ± 2 per cent.

Although countries with better and worse unemployment records each have some features in common, there are differences as well. The outcome appears to depend both on labour market characteristics and on actual developments in the wedge factors. The most unhappy combination is that in France and Canada, where we found a considerable degree of real wage resistance and a big rise in the wedge. Although the feedback from unemployment to wages is clear in both countries, real wage resistance has dominated.

In the USA, the degree of real wage resistance is fairly low and adjustment is quick. In addition, the rise in the wedge over the full observation period is small. Therefore, it should not be surprising that there are less signs of a trendwise increase in unemployment than in most other countries. In Japan, we found a considerable amount of real wage resistance, a large rise in the wedge and an adjustment profile which is not particularly rapid. The success of Japan in keeping the unemployment rate low is attributable to the lower response of demand for labour and, particularly, to the strong response of labour supply to changes in the demand for labour. The strong feedback effect from unemployment to wages is also a contributory factor. As indicated in OECD(1994), more recently the Japanese economy has shown signs of rising underemployment either in the form of labour hoarding or withdrawal, especially of women, from the labour force.

k) According to simulations carried out for four countries (Germany, France, Canada and Finland), revenue-neutral shifts from income taxes to employers' social security taxes have no long-run impact but in the short run (which lasts for five years or more) the effect on (un)employment is unfavourable. A shift from taxes on income to taxes on consumption has a favourable impact on (un)employment not only in the short run but in the long run as well. However, a precondition for this policy option is that the government allows the real value of such non-wage incomes as pensions, unemployment benefits and social transfers to fall as a result of the price increases generated by higher consumption taxes.

Appendix 1 The statistical model

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

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

where the X is a vector of stochastic variables, D is a vector of deterministic or non-normal variables (dummies),[57] μ contains constant terms and ε the Gaussian residuals.

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 $(n-r) \times c_1$ 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 (xx) 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 contain 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).

[57] 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

$$\begin{aligned} f(X_t;\psi D_t) &= f(X_t|X_{t-1},D_t;\psi) \\ &= f(\hat{X}_t|\bar{X}_t,X_{t-1},D_t;\psi_1)\cdot f(\bar{X}_t|X_{t-1},D_t;\psi_2)\,, \end{aligned} \tag{xxi}$$

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 in 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 lose any information concerning the long-run relations if part of the data is considered as weakly exogenous.

Appendix 2 Data and sources

The theoretical model was specified in the main text. To make this model operational, following series were collected from the Analytical Database (ADB) of the OECD for each of the countries:

1) Wage rate, private sector, WR,
2) Private dependent employment, EEP,
3) Wages and salaries, WAGE,
4) Employers' contributions to social security, TRPBTH,
5) Direct taxes, TYH,
6) Deflator for private consumption, PCP,
7) Consumer price index, CPI,
8) Deflator for business sector value added, PGDPB,
9) Business sector value added, volume, GDPBV,
10) Number on unemployed persons, UN,
11) Unemployment rate, UNR.

As far income tax rates were considered, the basic source of information was the OECD-publication TAX/BENEFIT POSITION OF AN AVERAGE PRODUCTION WORKER. There we found average tax rates for *a married couple with (two) children with a principle with an average earning and a non-working spouse*. The data is annual for 1974—91 with missing values for 1975, 1977, 1980 and 1982. The missing data points were calculated as an arithmetic mean of the previous and of the proceeding observation. Whenever the difference between these two observations was large, additional information was sought from the OECD Country Studies as well as from other sources. This was how we attempted to evaluate in which year the big shifts were likely to have taken place.

Furthermore, from an early OECD issue TAX/BENEFIT POSITION OF SELECTED INCOME GROUPS we found data on average tax rates for *a single male average production worker* for 1972—74. Relative changes in these rates were used when the series discussed above were approximated for 1972—73. As a result, we have annual series on average tax rates for 1972—92. All figures include own social security contributions.

New data on marginal income tax rates were produced for *the OECD Job Study*. These annual observations, however, only cover five years: 1978, 1981, 1985, 1989, 1992 and for some countries 1991. Missing observations were calculated using the difference between average rate (τ_a) and the marginal rate (τ_m), i.e. $\tau_m-\tau_a$, as a reference. For 1972—76, full information on this difference is in the TAX/BENEFIT POSITION OF SELECTED INCOME GROUPS. For the rest of the period, we started by assuming that this difference changes smoothly over time. This assumption was relaxed whenever the OECD Country Studies gave alternative information.

New data produced for *the OECD Jobs Study* on average and marginal income tax rates contain information on various types of wage earners. There are several characterizations depending on wage earner's family status and earnings level relative to the average earning. In the present study, we also experimented with data on *a married couple with (two) children with a prin-*

ciple with an average earning and a spouse with an earning which is 66.66 percent of the average earning. As far as the marginal income tax is concerned, we considered the rate faced by the principle as the relevant one.

In some calculations as well as in simulations, we used new data produced for *the OECD Jobs Study* on the consumption tax rates.

We also experimented with new data constructed for *the OECD Jobs Study* on replacement ratios (for the first six months of unemployment). Data on average wage earners as well as .666*average wage earners with a dependent and a working spouse were used. These series consist of observations for every second year and gaps were filled with methods similar to those discussed above.

The resulting income tax and replacement rate series are annual series which were transformed to semiannual series by two different methods. First, technical extrapolation procedures were used. Second, the series were allowed to proceed in a stepwise manner. Both types of series were experimented with. Usually, the former worked better.

Appendix 3 Germany

Table A1 **Residual analysis**

A. The full model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	1.510	1.435	-0.099	0.567	0.600
$\Delta\log(Q/N)$	1.378	4.068	-0.539	1.163	4.189
$\Delta\log(1-\tau_a)$	2.031	0.352	-0.139	-0.379	0.369
$\Delta\log(1-\tau_m)$	0.993	0.051	0.258	-0.100	0.460
$\Delta\log(1+s)$	1.010	0.137	-0.186	0.363	0.450
$\Delta\log(CPI/P)$	1.282	1.033	-0.095	-0.205	0.130
$\Delta\log(u)$	0.788	0.817	0.574	0.826	3.331

B. The partial model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	1.919	5.025	-0.193	0.252	0.356
$\Delta\log(Q/N)$	1.416	2.793	-0.085	1.262	2.701
$\Delta\log(1-\tau_m)$	1.264	0.011	0.321	2.700	12.838**
$\Delta\log(1+s)$	0.989	0.202	-0.406	0.684	1.878
$\Delta\log(u)$	0.869	0.059	0.836	0.504	5.078

* indicates significance at the 5 per cent and ** at the 1 per cent level. 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 Jarque & Bera test statistic for normality of the residuals (derived from columns three and four) which is distributed $\chi^2(2)$.

Table A2 **A joint Trace test for the cointegration rank, r, and the existence of a linear trend, H_r^* versus H_r.**

	No linear trend, H_r^*			Linear trend, H_r		
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	.802	251	132	.802	228	124
1	.754	187	102	.754	163	94
2	.685	130	76	.665	107	68
3	.634	84	53	.559	64	47
4	.448	44	35	.356	30	30
5	.350	20	20	.267	13	15
6	.073	3	9	.017	.7	4

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

Table A3 **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	3	112	35
1	2	61	20
2	1	23	9

* In a 7-dimensional system with r = 4, the number of common trends is c = 7-4 = 3.

Appendix 3 (cont.) Canada

Table A1 **Residual analysis**

A. The full model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
Δlog(W/P)	1.486	1.202	0.222	1.112	2.391
Δlog(Q/N)	1.951	0.046	0.389	-0.315	1.171
$\Delta\log(1-\tau_a)$	0.864	2.519	0.163	0.877	1.458
$\Delta\log(1-\tau_m)$	2.225	3.326	-0.152	0.222	0.237
Δlog(1+s)	1.378	3.053	-0.061	0.926	1.456
Δlog(PCP/P)	1.313	0.001	-0.223	-0.344	0.528
Δ(u)	1.053	3.086	0.168	0.295	0.334

B. The partial model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
Δlog(W/P)	1.234	0.460	0.161	0.359	0.388
Δlog(Q/N)	1.678	2.701	-0.103	0.815	1.176
$\Delta\log(1-\tau_a)$	0.769	2.729	0.178	0.851	1.417
Δlog(PCP/P)	1.167	0.380	-0.099	-0.590	0.646
Δ(u)	0.439	2.097	0.607	1.092	4.443

* indicates significance at the 5 per cent and ** at the 1 per cent level. 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 Jarque & Bera test statistic for normality of the residuals (derived from columns three and four) which is distributed $\chi^2(2)$.

Table A2 **A joint Trace test for the cointegration rank, r, and the existence of a linear trend, H_r^* versus H_r.**

	No linear trend, H_r^*			Linear trend, H_r		
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	.912	289	132	.899	252	124
1	.825	192	102	.794	160	94
2	.770	122	76	.711	97	68
3	.481	63	53	.373	47	47
4	.363	37	35	.285	28	30
5	.276	19	20	.272	15	15
6	.140	6	9	.057	2	4

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

Table A3 **A Trace test for I(2) common trends***

Number of		Test statistic, Q_{3,c_1}	Critical value, $CV_{95\%}$
I(1) common trends, c_1	I(2) common trends, $c_2 = c - c_1$		
0	4	116	53
1	3	68	35
2	2	29	20
3	1	8	9

* In a 7-dimensional system with r = 3, the number of common trends is c = 7−3 = 4.

Appendix 3 (cont.) Japan

Table A1 **Residual analysis**

A. The full model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	1.120	5.827	0.104	0.337	0.248
$\Delta\log(Q/N)$	1.891	0.053	0.105	-0.708	0.863
$\Delta\log(1-\tau_a)$	1.238	4.692	0.053	0.052	0.022
$\Delta\log(1-\tau_m)$	1.957	5.893	-0.362	-0.672	1.547
$\Delta\log(1+s)$	2.544	0.529	0.533	0.026	1.803
$\Delta\log(PCP/P)$	0.755	2.765	0.140	0.508	0.533
$\Delta(u)$	1.428	0.537	1.344	3.503	30.881**

B. The partial model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	1.366	10.536	0.135	0.381	0.345
$\Delta\log(Q/N)$	1.538	1.198	0.159	-0.579	0.691
$\Delta\log(PCP/P)$	1.384	1.171	0.392	0.448	1.291
$\Delta(u)$	1.423	0.461	0.652	1.750	7.538

* indicates significance at the 5 per cent and ** at the 1 per cent level. 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 Jarque & Bera test statistic for normality of the residuals (derived from columns three and four) which is distributed $\chi^2(2)$.

Table A2 **A joint Trace test for the cointegration rank, r, and the existence of a linear trend, H_r^* versus H_r.**

	No linear trend, H_r^*			Linear trend, H_r		
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	.799	219	132	.782	187	124
1	.754	158	102	.621	129	94
2	.617	105	76	.571	93	68
3	.509	68	53	.491	60	47
4	.386	41	35	.386	35	30
5	.341	23	20	.341	15	15
6	.165	7	9	.013	.5	4

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

Table A3 **A Trace test for I(2) common trends***
under r = 4 (the upper panel) and r = 5 (the lower panel).

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	3	64	35
1	2	24	20
2	1	.8	9

Number of		Test statistic, Q_{5,c_1}	Critical value, $CV_{95\%}$
I(1) common trends, c_1	I(2) common trends, $c_2 = c-c_1$		
0	2	24	20
1	1	.3	9

* In a 7-dimensional system with r=4, the number of common trends is c = 7–4 = 3 and with r = 5 it is c = 7–5 = 2.

Appendix 3 (cont.) Finland

Table A1 **Residual analysis**

A. The full model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log[(W/P)-(Q/N)]$	1.403	2.058	0.386	-0.716	1.660
$\Delta\log(1-\tau_a)$	0.827	0.013	-0.331	3.133	15.379**
$\Delta\log(1-\tau_m)$	0.599	0.112	0.247	2.342	8.593*
$\Delta\log(1+s)$	2.062	10.877**	0.018	-0.301	0.137
$\Delta\log(CPI/P)$	0.927	0.972	-0.250	-0.240	0.460
$\Delta\log(u)$	0.761	1.291	-0.450	0.636	1.823

B. The partial model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log[(W/P)-(Q/N)]$	1.470	0.738	0.225	-0.577	0.802
$\Delta\log(u)$	0.182	0.846	−0.086	0.641	0.060

C. The full "controll model"

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	1.064	2.420	-0.136	-0.242	0.454
$\Delta\log(Q/N)$	0.984	2.473	-0.122	0.551	1.237
$\Delta\log(1-\tau_a)$	2.311	0.200	1.165	0.631	42.212**
$\Delta\log(1-\tau_m)$	2.306	1.773	0.940	1.744	22.464**
$\Delta\log(1+s)$	1.075	0.312	0.825	3.129	42.763**
$\Delta\log(PC/P)$	3.875	4.085	0.019	-0.183	0.119
$\Delta\log(P_m/P)$	0.729	0.611	0.775	1.688	17.947**
$\Delta\log(u)$	1.730	4.350	0.002	1.355	6.269*

D. The partial "controll model"

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	1.298	6.298*	-0.271	-0.044	1.014
$\Delta\log(1-\tau_a)$	1.638	3.977	-0.393	4.587	73.984**
$\Delta\log(1-\tau_m)$	1.524	11.048**	−0.859	10.060	355.867**
$\Delta\log(P_m/P)$	0.511	0.545	0.940	2.518	33.737**

* indicates significance at the 5 per cent and ** at the 1 per cent level. 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 Jarque & Bera test statistic for normality of the residuals (derived from columns three and four) which is distributed $\chi^2(2)$.

Table A2 **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	.768	138	102	.767	134	94
1	.614	86	76	.607	82	68
2	.514	52	53	.500	48	47
3	.363	26	35	.363	23	30
4	.185	93	20	.162	7	15
5	.053	2	9	.018	.7	4

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

Table A3 **A Trace test for I(2) common trends***

Number of		Test statistic, Q_{2,c_1}	Critical value, $CV_{95\%}$
I(1) common trends, c_1	I(2) common trends, $c_2 = c-c_1$		
0	4	97	53
1	3	43	35
2	2	16	20
3	1	7	9

* In a 6-dimensional system with r = 2, the number of common trends is c = 6−2 = 4.

Appendix 3 (cont.) Australia

Table A1 **Residual analysis**

A. The full model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	0.667	1.225	0.114	0.110	0.093
$\Delta\log(Q/N)$	0.508	0.188	0.358	0.312	0.888
$\Delta\log(1-\tau_a)$	1.478	2.745	0.232	0.279	0.429
$\Delta\log(1+s)$	1.495	3.121	-0.569	2.333	9.826**
$\Delta\log(PCP/P)$	1.354	0.851	-0.155	-0.837	1.161
$\Delta(u)$	1.468	3.778	-0.394	0.345	1.079

B. The partial model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	0.819	1.891	0.223	-0.062	0.295
$\Delta\log(Q/N)$	1.073	0.223	-0.020	-0.436	0.280
$\Delta\log(1-\tau_a)$	1.809	3.668	0.644	0.038	2.423
$\Delta\log(PCP/P)$	1.613	0.420	−0.353	-0.753	1.555

* indicates significance at the 5 per cent and ** at the 1 per cent level. 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 Jarque & Bera test statistic for normality of the residuals (derived from columns three and four) which is distributed $\chi^2(2)$.

Table A2 **A joint Trace test for the cointegration rank, r, and the existence of a linear trend, H_r^* versus H_r.**

	No linear trend, H_r^*			Linear trend, H_r		
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	.853	194	102	.853	178	94
1	.787	127	76	.787	111	68
2	.647	73	53	.488	57	47
3	.456	37	35	.446	33	30
4	.301	15	20	.251	12	15
5	.075	3	9	.073	3	4

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

Table A3 **A Trace test for I(2) common trends***

Number of		Test statistic, Q_{3,c_1}	Critical value, $CV_{95\%}$
I(1) common trends, c_1	I(2) common trends, $c_2 = c-c_1$		
0	3	63	35
1	2	36	20
2	1	13	9

* In a 6-dimensional system with r = 3, the number of common trends is c = 6−3 = 3.

Appendix 3 (cont.) France

Table A1 **Residual analysis**

A. The full model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	1.010	0.295	-0.128	-0.628	0.768
$\Delta\log(Q/N)$	1.258	1.011	-0.311	0.859	1.877
$\Delta\log(1-\tau_a)$	2.147	4.378	0.809	1.794	9.733**
$\Delta\log(1+s)$	0.801	1.556	0.299	-0.236	0.689
$\Delta\log(CPI/P)$	0.490	4.691	-0.034	-0.396	0.269
$\Delta\log(u)$	0.340	6.095	0.166	-0.463	0.541

B. The partial model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	1.070	11.567**	0.016	-1.075	1.926
$\Delta\log(1-\tau_a)$	2.735	4.909	1.095	2.884	21.858**
$\Delta\log(CPI/P)$	0.573	3.728	0.041	-0.322	0.184

* indicates significance at the 5 per cent and ** at the 1 per cent level. 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 Jarque & Bera test statistic for normality of the residuals (derived from columns three and four) which is distributed $\chi^2(2)$.

Table A2 **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	.702	164	102	.701	144	94
1	.652	115	76	.594	96	68
2	.562	73	53	.523	60	47
3	.374	40	35	.367	30	30
4	.361	21	20	.194	12	15
5	.081	3	9	.080	3	4

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

Table A3 **A Trace test for I(2) common trends***

Number of		Test statistic, Q_{3,c_1}	Critical value, $CV_{95\%}$
I(1) common trends, c_1	I(2) common trends, $c_2 = c-c_1$		
0	3	74	35
1	2	30	20
2	1	13	9

* In a 6-dimensional system with r = 3, the number of common trends is c = 6–3 = 3.

Appendix 3 (cont.) Sweden

Table A1 **Residual analysis**

A. The full model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	0.668	1.400	-0.379	-0.027	0.863
$\Delta\log(Q/N)$	1.238	0.591	-0.075	0.996	1.522
$\Delta\log(1-\tau_a)$	2.241	4.057	0.139	0.082	0.126
$\Delta\log(1-\tau_m)$	0.983	0.327	-0.077	-0.722	0.817
$\Delta\log(1+s)$	0.877	0.242	0.415	-0.532	1.457
$\Delta\log(CPI/P)$	2.166	3.938	-0.124	0.427	0.365
$\Delta\log(u)$	1.541	4.366	-0.064	-0.826	1.048

B. The partial model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	1.367	3.282	0.148	-0.102	0.147
$\Delta\log(Q/N)$	1.821	0 .317	-0.007	-0.220	0.073
$\Delta\log(1-\tau_a)$	2.969	3.548	0.369	-0.301	0.953
$\Delta\log(1-\tau_m)$	0.973	0.517	0.150	-0.157	0.173
$\Delta\log(CPI/P)$	1.683	1.509	0.060	-0.421	0.287

* indicates significance at the 5 per cent and ** at the 1 per cent level. 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 Jarque & Bera test statistic for normality of the residuals (derived from columns three and four) which is distributed $\chi^2(2)$.

Table A2 **A joint Trace test for the cointegration rank, r, and the existence of a linear trend, H_r^* versus H_r.**

	No linear trend, H_r^*			Linear trend, H_r		
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	.888	217	132	.885	188	124
1	.730	138	102	.690	110	94
2	.632	91	76	.572	69	68
3	.477	55	53	.458	37	47
4	.386	32	35	.214	15	30
5	.214	15	20	.167	7	15
6	.151	6	9	.006	.2	4

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

Table A3 **A Trace test for I(2) common trends***

Number of		Test statistic, Q_{3,c_1}	Critical value, $CV_{95\%}$
I(1) common trends, c_1	I(2) common trends, $c_2 = c-c_1$		
0	4	110	53
1	3	57	35
2	2	14	20
3	1	4	9

* In a 7-dimensional system with r = 3, the number of common trends is c = 7–3 = 4.

Appendix 3 (cont.) United States

Table A1 **Residual analysis**

A. The full model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log[(W/P) - (Q/N)]$	0.448	1.229	-0.074	-0.660	0.762
$\Delta\log(1-\tau_a)$	2.373	13.688**	-0.365	1.300	3.708
$\Delta\log(1-\tau_m)$	2.245	0.073	-1.094	2.667	19.837**
$\Delta\log(1+s)$	1.718	0.522	-0.152	-0.015	0.155
$\Delta\log(CPI/P)$	0.842	3.048	0.079	-0.137	0.073
$\Delta(u)$	0.821	0.063	1.030	3.952	33.109**

B. The partial model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log[(W/P)-(Q/N)]$	1.256	0.563	-0.240	-0.167	0.429
$\Delta\log(1-\tau_m)$	3.541	0.089	-1.280	2.835	24.322**
$\Delta\log(1+s)$	1.062	0.489	-0.160	0.621	0.812

* indicates significance at the 5 per cent and ** at the 1 per cent level. 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 Jarque & Bera test statistic for normality of the residuals (derived from columns three and four) which is distributed $\chi^2(2)$.

Table A2 **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	.877	176	102	.865	157	94
1	.614	92	76	.607	77	68
2	.473	54	53	.420	40	47
3	.331	29	35	.320	19	30
4	.223	12	20	.059	3	15
5	.058	2	9	.016	.6	4

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

Table A3 **A Trace test for I(2) common trends***

Number of		Test statistic, Q_{2,c_1}	Critical value, $CV_{95\%}$
I(1) common trends, c_1	I(2) common trends, $c_2 = c-c_1$		
0	4	96	53
1	3	45	35
2	2	19	20
3	1	2	9

* In a 6-dimensional system with r = 2, the number of common trends is c = 6–2 = 4.

Appendix 3 (cont.) Italy

Table A1 **Residual analysis**

A. The full model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log[(W/P)-(Q/N)]$	1.094	0. 535	0.614	0.233	2.474
$\Delta\log(1-\tau_a)$	1.562	0.155	0.221	-0.257	0.414
$\Delta\log(1-\tau_m)$	1.669	11.849**	0.240	0.309	0.515
$\Delta\log(1+s)$	3.893	1.820	-0.285	0.257	0.618
$\Delta\log(PCP/P)$	0.888	2.561	0.787	1.099	5.837
$\Delta\log(u)$	1.086	1.815	0.210	0.030	0.280

B. The partial model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log[(W/P)-(Q/N)]$	1.253	0.802	0.654	-0.013	2.711
$\Delta\log(1-\tau_a)$	1.118	23.155**	0.867	1.102	6.680
$\Delta\log(1-\tau_m)$	0.716	2.680	0.566	0.808	3.066
$\Delta\log(1+s)$	3.893	2.525	0.002	0.075	0.009
$\Delta\log(u)$	0.749	0.311	-0.417	3.545	20.999**

* indicates significance at the 5 per cent and ** at the 1 per cent level. 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 Jarque & Bera test statistic for normality of the residuals (derived from columns three and four) which is distributed $\chi^2(2)$.

Table A2 **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	.911	219	102	.872	201	94
1	.717	127	76	.714	123	68
2	.593	80	53	.593	75	47
3	.558	45	35	.520	28	30
4	.227	14	20	.225	10	15
5	.114	5	9	.090	4	4

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

Table A3 **A Trace test for I(2) common trends***

Number of		Test statistic, Q_{3,c_1}	Critical value, $CV_{95\%}$
I(1) common trends, c_1	I(2) common trends, $c_2 = c-c_1$		
0	3	85	35
1	2	39	20
2	1	13	9

* In a 6-dimensional system with r = 3, the number of common trends is c = 6–3 = 3.

Appendix 3 (cont.) United Kingdom

Table A1 **Residual analysis**

A. The full model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	0.906	0.747	-0.419	0.539	1.572
$\Delta\log(Q/N)$	1.621	2.183	0.084	1.633	4.269
$\Delta\log(1-\tau_a)$	0.559	4.806	0.355	1.065	2.597
$\Delta\log(1-\tau_m)$	1.068	0.010	0.306	-0.916	1.920
$\Delta\log(1+s)$	1.181	0.500	-0.087	0.138	0.078
$\Delta\log(PCP/P)$	0.792	0.237	-0.476	-0.491	1.816
$\Delta\log(u)$	0.610	0.283	0.041	-0.009	0.011

B. The partial model

	B-P.Q(10)/8	ARCH(2)	SKEW.	EX.KURT.	J-B.NORM.
$\Delta\log(W/P)$	0.906	0.794	-0.419	0.550	1.591
$\Delta\log(Q/N)$	1.529	2.281	-0.086	2.245	8.030*
$\Delta\log(1-\tau_a)$	0.749	5.522	0.278	1.509	4.094
$\Delta\log(1-\tau_m)$	1.141	0.073	0.306	-0.791	1.582
$\Delta\log(1+s)$	1.289	0.213	0.181	0.260	0.315
$\Delta\log(PCP/P)$	0.826	0.586	-0.524	-0.252	1.842

* indicates significance at the 5 per cent and ** at the 1 per cent level. 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 Jarque & Bera test statistic for normality of the residuals (derived from columns three and four) which is distributed $\chi^2(2)$.

Table A2 **A joint Trace test for the cointegration rank, r, and the existence of a linear trend, H^*_r versus H_r.**

	No linear trend, H^*_r			Linear trend, H_r		
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	.895	286	132	.885	242	124
1	.780	201	102	.729	160	94
2	.722	143	76	.690	111	68
3	.663	95	53	.576	66	47
4	.515	53	35	.405	33	30
5	.306	26	20	.297	13	15
6	.267	12	9	.015	6	4

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

Table A3 **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	3	66	35
1	2	24	20
2	1	6	9

* In a 7-dimensional system with r = 4, the number of common trends is c = 7–4 = 3.

Table A4 **Cointegrating relations with characteristics of a wage setting relation under cointegration rank (=r) as indicated in the head of the Table[a]**

		USA				CAN		JAP	
		r = 3	r = 3	r = 2	r = 2	r = 3	r = 3	r = 4	r = 3
		(1)	(2)	(3)*	(4)	(1)*	(2)	(1)*	(2)
coeff.	variables	long run coefficients β_i							
$\beta_{W/P}$	W/P	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
$\beta_{Q/N}$	Q/N	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000
β_{τ_a}	$1-\tau_a$	.714	1.000	1.000	1.000	1.000	1.300	1.000	1.000
β_{τ_m}	$1-\tau_m$	.000	.000	.000	.000	−.198	−.300	−.500	−.500
β_s	1+s	1.000	1.000	1.000	1.000	.198	.000	.500	.500
$\beta_{Pc/P}$	P_c/P	−.714	−1.000	.000	.000	−.802	−1.000	−.500	−.500
β_u	u	.000	.000	.007	.007	.019	.020	.051	.039
		adjustment coefficients α_i							
$\alpha_{W/P}$	$\Delta W/P$	−.124	−.092	−.199	−.143	−.157	−.253	−.431	−.498
$\alpha_{Q/N}$	$\Delta Q/N$	..	..	..	..	.066	.032	−.399	−.408
α_{τ_a}	$\Delta\tau_a$	−.044	−.041	..	..	−.077	−.100	..	..
$\alpha_{\tau m}$	$\Delta\tau_m$	−.397	−.324	−.370	−.508	..	..	..	..
α_s	Δs	−.012	−.011	.018	.025	..	..	..	..
$\alpha_{Pc/P}$	$\Delta P_c/P$	..	..	..	..	.245	.147	.244	.239
α_u	Δu	−.083	−.061	..	..	−2.894	−6.608	1.942	2.294
number of restrictions imposed on relation i, (=η_i)		6	6	4	4	4	4	4	4
characterization of restrictions imposed		$\beta_{W/P} = -\beta_{Q/N} = \beta_s$, $\beta_{\tau_a} = -\beta_{Pc/P}$, $\beta_{\tau_m} = \beta_u = 0$	$\beta_{W/P} = -\beta_{Q/N} = \beta_{\tau_a} = \beta_s = -\beta_{Pc/P}$, $\beta_{\tau_m} = \beta_u = 0$	$\beta_{W/P} = -\beta_{Q/N} = \beta_{\tau_a} = \beta_s$, $\beta_{\tau_m} = \beta_{Pc/P} = 0$	$\beta_{W/P} = -\beta_{Q/N} = \beta_{\tau_a} = \beta_s$, $\beta_{\tau_m} = \beta_{Pc/P} = 0$	$\beta_{W/P} = -\beta_{Q/N} = \beta_{\tau_a}$, $\beta_{\tau_a} + \beta_{\tau_m} = -\beta_{Pc/P} = 1-\beta_s$	$\beta_{W/P} = -\beta_{Q/N} = \beta_{\tau_a} + \beta_{\tau_m} = -\beta_{Pc/P} = 1-\beta_s$	$\beta_{W/P} = -\beta_{Q/N} = \beta_{\tau_a}$, $\beta_{\tau_m} = 1-\beta_s = -\beta_{Pc/P} = .5$	$\beta_{W/P} = -\beta_{Q/N} = \beta_{\tau_a}$, $\beta_{\tau_m} = 1-\beta_s = -\beta_{Pc/P} = .5$
LM test statistic, χ^2 ($\Sigma(\eta_i-(r-1))$)		1.45	1.26	1.68	.13	5.62	2.40	3.20	7.16
critical value, $CV_{5\%}$ ($\Sigma(\eta_i-(r-1))$)		7.81	9.49	7.81	7.81	5.99	5.99	5.99	7.81
p-value		.69	.87	.64	.97	.06	.30	.20	.07
eigenvalues, λ_i		.621	.614	.640	.639	.734	.767	.573	.577

[a] All variables are in logs, those considered as weakly exogenous have been underlined, and r = number of cointegrating vectors, r_1 = number of cointegrat ng vectors to be restricted, r_2 = number of non-restricted cointegrating vectors = $r-r_1$. In so far as the α-coefficients are concerned, when a variable has been concerned as a weakly exogenous one, there is ".." in the place reserved for the α-coefficient.
For all countries in this table, u refers to the unemployment rate. For the USA, the unrestircted VAR behind relations (1) and (2) is identical to that applied for other countries. In relations (3) and (4), the homogeneity constraint — which was approved to hold in relations (1) and (2) — between real wages and labour productivity was imposed *a priori*. In relation (4), CPI has been replaced by the deflator for private consumption.

Table A4 (cont.) **Cointegrating relations with characteristics of a wage setting relation under cointegration rank (= r) as indicated in the head of the Table**[a]

		GER		FRA		ITA		UKM	
		r = 4	r = 4	r = 3	r = 3	r = 3	r = 3	r = 5	r = 3
		(1)*	(2)	(1)*	(2)	(1)	(2)*	(1)*	(2)
coeff.	variables	long run coefficients β_i							
$\beta_{W/P}$	W/P	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
$\beta_{Q/N}$	Q/N	−1.000	−1.000	$\underline{-1.000}$	$\underline{-1.000}$	−1.000	−1.000	−1.000	−1.000
β_{τ_a}	$1-\tau_a$	$\underline{1.000}$	$\underline{1.000}$	.421	.500	1.000	1.000	.250	.250
β_{τ_m}	$1-\tau_m$	.000	.000	..	..	−.611	−.600	.000	.000
β_s	1+s	.000	.000	$\underline{.579}$	$\underline{.500}$	.611	.600	.750	.750
$\beta_{Pc/P}$	P_c/P	$\underline{-1.000}$	$\underline{-1.000}$	−.421	−.500	$\underline{-.389}$	$\underline{-.400}$	−.250	−.250
β_u	u	.168	.262	$\underline{.289}$	$\underline{13.659}$	.332	.330	$\underline{.098}$	$\underline{.097}$
		adjustment coefficients α_i							
$\alpha_{W/P}$	$\Delta W/P$	−.149	−.146	−.344	.005	−.147	−.150	−.702	−.668
$\alpha_{Q/N}$	$\Delta Q/N$	.018	.068	..	..	..	..	.173	−.236
α_{τ_a}	$\Delta\tau_a$	..	..	−.093	−.002	−.042	−.043	.087	.159
$\alpha_{\tau m}$	$\Delta\tau_m$	.083	−.073	..	..	.065	.060	−.130	−.005
α_s	Δs	−.010	.004	..	..	−.146	.144	−.186	−.180
$\alpha_{Pc/P}$	$\Delta P_c/P$	..	..	−.210	−.007	..	..	−.152	−.201
α_u	Δu	−.181	.872	..	..	.302	.315	..	..
number of restrictions imposed on relation i, $(=\eta_i)$		5	5	3	4	3	4	5	5
characterization of restrictions imposed		$\beta_{W/P} = -\beta_{Q/N} = \beta_{\tau_a} = -\beta_{Pc/P} = 1-\beta_s$, $\beta_{\tau_m} = 0$	$\beta_{W/P} = -\beta_{Q/N} = \beta_{\tau_a} = -\beta_{Pc/P}$, $= 1-\beta_s$, $\beta_{\tau_m} = 0$	$\beta_{W/P} = -\beta_{Q/N}$, $-\beta_{Pc/P} = \beta_{\tau_a} = 1-\beta_s$	$\beta_{W/P} = -\beta_{Q/N}$, $-\beta_{Pc/P} = \beta_{\tau_a} = 1-\beta_s = .5$	$\beta_{W/P} = -\beta_{Q/N} = \beta_{\tau_a}$, $\beta_{\tau_a} + \beta_{\tau_m} = 1-\beta_s = -\beta_{Pc/P}$	$\beta_{W/P} = -\beta_{Q/N} = \beta_{\tau_a}$, $\beta_{\tau_a} + \beta_{\tau_m} = 1-\beta_s = -\beta_{Pc/P} = .4$	$\beta_{W/P} = -\beta_{Q/N}$, $1-\beta_s = \beta_{\tau_a} = -\beta_{Pc/P} = .25$, $\beta_{\tau_m} = 0$	$\beta_{W/P} = -\beta_{Q/N}$, $1-\beta_s = \beta_{\tau_a} = -\beta_{Pc/P} = .25$, $\beta_{\tau_m} = 0$
LM test statistic, $\chi^2\ (\Sigma(\eta_i-(r-1))$		.78	4.04	.65	5.98	1.65	1.65	1.19	7.17
critical value, $CV_{5\%}\ (\Sigma(\eta_i-(r-1))$		5.99	5.99	3.84	5.99	3.84	5.99	3.84	7.81
p-value		.68	.13	.42	.05	.20	.44	.27	.07
eigenvalues, λ_i		.489	.358	.570	.419	.506	.506	.787	.786

[a] All variables are in logs, those considered as weakly exogenous have been underlined, and r = number of cointegrating vectors, r_1 = number of cointegrating vectors to be restricted, r_2 = number of non-restricted cointegrating vectors = $r-r_1$. In so far as the α-coefficients are concerned, when a variable has been concerned as a weakly exogenous one, there is ".." in the place reserved for the α-coefficient.
For all countries in this Table u refers to log of the unemployment rate.

Table A4 (cont.) **Cointegrating relations with characteristics of a wage setting relation under cointegration rank (= r) as indicated in the head of the Table[a]**

		AUS		SWE		FIN		$FIN^{control}$	
		r = 3	r = 3	r = 3	r = 3	r = 2	r = 2	r = 4	r = 3
		(1)	(2)*	(1)	(2)*	(1)	(2)*	(3)	(4)
coeff.	variables	long run coefficients β_i							
$\beta_{W/P}$	W/P	1.000	1.000	1.000	1.000	1.000	1.000	1.000	1.000
$\beta_{Q/N}$	Q/N	−1.000	−1.000	−1.000	−1.000	−1.000	−1.000	$\underline{-1.000}$	$\underline{-1.000}$
β_{τ_a}	$1-\tau_a$	.398	.500	.000	.000	$\underline{1.000}$	$\underline{1.000}$	1.000	1.000
β_{τ_m}	$1-\tau_m$	..	..	.107	.000	$\underline{-.614}$	$\underline{-.500}$	−.443	−.442
β_s	1+s	$\underline{.602}$	$\underline{.500}$	$\underline{1.000}$	$\underline{1.000}$	$\underline{.614}$	$\underline{.500}$	$\underline{.443}$	$\underline{.442}$
$\beta_{Pc/P}$	P_c/P	−.398	−.500	−1.000	−1.000	$\underline{-.386}$	$\underline{-.500}$	−.557	−.558
β_u	u	$\underline{.033}$	$\underline{.035}$	$\underline{.234}$	$\underline{.256}$	.332	.314	$\underline{.000}$	$\underline{.000}$
		adjustment coefficients α_i							
$\alpha_{W/P}$	$\Delta W/P$	−.380	−.377	−.161	−.082	−.297	−.303	−.919	−.887
$\alpha_{Q/N}$	$\Delta Q/N$	.040	.034	.227	.150	..	..	..	..
α_{τ_a}	$\Delta\tau_a$	−.061	−.066	−.016	−.018	..	..	−.378	−.403
$\alpha_{\tau m}$	$\Delta\tau_m$	..	..	.126	.174	..	..	−.505	−.617
α_s	Δs	..	..	..	..	..	..	..	..
$\alpha_{Pc/P}$	$\Delta P_c/P$	.356	.350	.300	.310	..	..	.508	.710
α_u	Δu	.354	..	..	..	1.265	1.275	..	..
number of restrictions imposed on relation i, $(=\eta_i)$		4	4	5	5	3	4	6	6
characterization of restrictions imposed		$\beta_{W/P} =$ $-\beta_{Q/N}$, $1-\beta_s$ $=-\beta_{Pc/P}$ $=\beta_{\tau_a}$	$\beta_{W/P} =$ $-\beta_{Q/N}$, $1-\beta_s =$ $-\beta_{Pc/P}=$ $\beta_{\tau_a}= .5$	$\beta_{W/P} =$ $-\beta_{Q/N}$ $= \beta_s$ $=-\beta_{Pc/P}$, $\beta_{\tau_a}=0$	$\beta_{W/P}=$ $-\beta_{Q/N}$ $=\beta_s$ $=-\beta_{Pc/P}$, $\beta_{\tau_a}=$ $-\beta_{\tau_m}$ $=0$	$\beta_{W/P}=$ $\beta_{Q/N}=$ β_{τ_a}, $\beta_{\tau_a}+$ $\beta_{\tau_m}=$ $1-\beta_s$ $=-\beta_{Pc/P}$	$\beta_{W/P}=$ $\beta_{Q/N}= \beta_{\tau_a}$, $\beta_{\tau_a}+$ $\beta_{\tau_m}=$ $1-\beta_s$ $=-\beta_{Pc/P}$ $=.5$	$\beta_{W/P}=$ $\beta_{Q/N}= \beta_{\tau_a}$, $\beta_{\tau_a}+$ $\beta_{\tau_m}=$ $1-\beta_s$ $=-\beta_{Pc/P}$, $\beta_u=0$	$\beta_{W/P}=$ $-\beta_{Q/N}$ $=\beta_{\tau_a}$, $\beta_{\tau_a}+$ $\beta_{\tau_m}=$ $1-\beta_s$ $=-\beta_{Pc/P}$, $\beta_u=0$
LM test statistic, $\chi^2\ (\Sigma(\eta_i-(r-1))$		.96	1.02	.60	4.55	.58	.68	.42	1.97
critical value, $CV_{5\%}\ (\Sigma(\eta_i-(r-1))$		3.84	5.99	5.99	7.81	5.99	7.81	7.81	9.49
p-value		.33	.60	.74	.21	.75	.88	.94	.74
eigenvalues, λ_i		.539	.539	.589	.544	.444	.454	.309	.309

[a] All variables are in logs, those considered as weakly exogenous have been underlined, and r = number of cointegrating vectors, r_1 = number of cointegrating vectors to be restricted, r_2 = number of non-restricted cointegrating vectors = $r-r_1$. In so far as the α-coefficients are concerned, when a variable has been concerned as a weakly exogenous one, there is ".." in the place reserved for the α-coefficient.
For AUS, u refers to the unemployment rate and for SWE and FIN to the logarithm of the unemployment rate.

Appendix 4 The Simulation Model

We have derived and estimated above wage relationships. In order to design a simulation model which could be used in evaluations about related effects on employment and unemployment we need to specify relationships which define 1) demand for labour and 2) responses of unemployment to changes in employment. That is the purpose of this appendix.

1 Modelling Labour Demand

In section 3 of the main text, we specified a wage setting relation in a model with only one production factor, labour. In order to "show" that this simplification does not make a big difference as far as empirical application is concerned, we now define the relationship acting as the optimization restraint in the Nash maximand (ii) in the main text in a standard set-up used by Wallis et al., 1984, e.g. Let us derive a relationship in which employment depends on real labour cost and output by writing the production function as

$$Q = F(N, K, t) \tag{xxii}$$

where Q is output, N is employment, K is capital stock and t is time trend which measures the technical progress. The marginal productivity condition for labour is

$$W(1+s)/P = F_N(W(1+s)/P, Q, t)$$

where W is the nominal wage, s is the payroll tax rate and P is the value added deflator.

Since capital stock is often considered as a matter which is particularly poorly measured it is usually eliminated. Using (xxii) for this purpose yields

$$N^d = N^d((W(1+s)/P, Q, t) \tag{xxiii}$$

with $N_W^d < 0$ and $N_Q^d > 0$. The general from (xxiii) hides certain restrictions. If constant elasticity of substitution (CES) is assumed a general log linear version of (xxiii) is obtained, namely

$$\log N = [(1+b/v)/(1+b)] \log Q - (1+b)^{-1} \log(W(1+s)/P) \tag{xxiv}$$

where $(1 + b)^{-1}$ is the elasticity on substitution and v returns to scale. When b tends to zero, CES tends to Cobb-Douglas. An important feature of (xxiv) is the presence of output, which is an endogenous decision variable under profit maximization. In longer-run estimations, prices for aggregate sectors can be considered as endogenous a well.

So, in the long run, we expect to find a (log linear) relation like

$$\log N = \beta_Q \log Q + \beta_W \log(W(1+s)) + \beta_P \log(P) \tag{xxv}$$

with $\beta_Q > 0$ and $-\beta_W = \beta_P > 0$. A CES technology is in concern if $\beta_Q = 1$ and $0 < -\beta_W = \beta_P \leq 1$. If $\beta_Q = -\beta_W = \beta_P = 1$, the technology is Cobb-Douglas (see Wallis et al, 1984, e.g.).

Our empirical unrestricted log-linear VAR-model contains the following four variables for the private sector:

1) employment, N, is the number of employed persons[57],
2) output, Q, is the value added in real terms,
3) the wage variable is the labour cost, W(1+s), where W is wage paid per an employee and s refers to employers' social security contributions,
4) output price, P, is the value added deflator.

The seasonally adjusted semiannual series come from the analytical data base (ADB) of the OECD. Because of the availability of the relevant series, estimation period varies from country to country (see Table A5).

All variables (N, Q, W, P) will be considered as endogenous at the outset. One (or more) of the variables will be then considered as weakly endogenous only if the relevant test indicates that this would be appropriate (see Appendix 1, and for the test, Juselius & Hargreaves, 1992).

If the data approves the existence of a long-run relation like (xxv), it defines a time-invariant attractor which acts as an equilibrium relation between employment and its determinants. If employment is off of the attractor, a pressure to correct this deviation emerges. Therefore, to a cointegrating relation like (xxv) in (log) levels, one can always relate a dynamic error correction equation in (log) differences. Applying the lag length which seems to be appropriate i.e. k = 2, the employment equation is as follows[58]:

$$\begin{aligned}\Delta\log N_t = {} & \gamma_{0,Q/N}\Delta\log Q_t + \gamma_{0,w}\Delta\log((W(1+s)_t + \gamma_{0,P}\Delta\log P_t \\ & + \gamma_{1,N}\Delta\log N_{t-1} + \gamma_{1,Q}\Delta\log Q_{t-1} + \gamma_{1,W}\Delta\log((W(1+s))_{t-1} \\ & + \gamma_{1,P}\Delta\log P_{t-1} + \text{possible constant} + \text{possible dummies} \\ & + \alpha_N\,[\beta_0 \log N_{t-1} + \beta_Q \log Q_{t-1} + \beta_W \log W_{t-1} + \beta_P \log P_{t-1}]\end{aligned} \tag{xxvi}$$

In the present study we are particularly interested in the long-run coefficients β_i which are on the last row of (xxvi). The other coefficient of special interest is the error correction coefficient α_N which reports the share of the equilibrium error which is corrected in the coming semester. In a difference equation like (xxvi), a significant constant term takes the role of the trend present in the level relationship (see the discussion in the main text).

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[57] The OECD data base does not include data on hours worked.

[58] The discussion in footnote 5 above also concerns the present equation.

The significant presence of the time trend will be tested simultaneously with the choice of the cointegration rank.

As far as the preliminary analysis of the data is considered, a few remarks are worth making. First, with few exceptions the appropriate choice for cointegration rank, r, is 2[59]. This indicates that there are two cointegrating vectors in the system. When three appeared to be the appropriate choice, we examined the residuals of the second and third vectors in order to emphasize whether the test result should be challenged. Second, with few exceptions the data appear not to contain linear trends. Third, as indicated above, dummies are often needed to reduce risk of bias which could be due to outliers in the dynamic counterpart of equation (x). In the present case, the dummies are usually related to oil shocks[60]. In general, in models including the relevant dummies, no signs of misspecification could be found in residual analysis[61]. Particularly, none of the estimations suffered from autocorrelation which would have been the most severe problem for estimation of the long-run coefficients (see Eitrheim, 1991, and Cheung et al., 1993). Signs of excess kurtosis were found in some cases but it seems not to be a severe problem for estimation in the context of the Johansen procedure. Finally, in six cases the data do not allow us to consider any of the four variables as weakly exogenous.

There are two countries in which the standard procedure worked less well. These countries are Sweden and the UK. In the former, a shift in the relation between wage level and the level of the employment/output ratio appeared to take place in the middle of the 1970's. In order to take account of this, a dummy which takes value of 1 during 1963S1—1975S2 had to be included into the model. In the UK, a similar shift appears to have taken place in the late eighties which is a familiar observation from other contexts as well. One explanation has been that at this point of time, a considerable number of low-skilled workers were obliged to leave unemployment and enter employment (even at a lower wage) which induced a negative effect on average productivity in the private business sector.

For all countries, restriction $\beta_N = \beta_Q = 1$ seems to be well in accordance with the data. In addition, the restriction $\beta_W = \beta_P$ was never rejected by the LR-test. Table A6 reports the long-run elasticities incorporated in the preferred relationships. The related α_N-coefficients are indicated as well. Finally, the Table reports the p-values related to the structure imposed and R^2's of employment equations like (xxvi) explaining changes in employment from one semester to the next. The α-coefficients which may potentially suffer from problems of imprecise estimation discussed above, are in parentheses although similarity of the estimates between countries with similar characteristics reduces fears in this respect.

Restrictions with an economic interpretation pass at a fairly high significance level. In six of the countries concerned (USA, AUS, FRA, GER,

[59] Test tables indicating this are available from the author upon request.

[60] A full set of dummies is available from the author upon request. It should be recognized that the role of dummies in the present set-up differs radically from the role they have in traditional estimation procedures. For a discussion of this issue, see the main text.

[61] Tables indicating this are available from the author upon request.

UK and FIN), elasticies in accordance with Cobb-Douglas technology were discovered. In an additional two countries (JAP and SWE), the deviation of the wage elasticity from unity is fairly small and the specification test does not reject the hypothesis that this elasticity is actually unity. In the remaining two economies (CAN and ITA), a general CES function seems to be a more appropriate description of the technology.

Table A5 **Summary of the preliminary analysis of the unrestricted VAR model for 10 OECD economies.**

	Estimation period	Number of cointegrating relations	Linear trend or not	Number of dummies	Weakly exogenous variables
USA	72S1–92S2	2	not	1	none
Canada	70S1–92S2	2	not	0	Q
Australia	69S2–90S1	2	not	2	none
France	70S1–92S2	2	yes	1	W(1+s), Q
Germany	60S1–92S2	3	not	2	none
Japan	68S1–91S2	3	not	1	none
Italy	60S1–91S2	3	yes	0	none
UK	61S1–91S2	2	not	1	P
Finland	60S1–90S2	1	not	1	W(1+s)
Sweden	63S1–90S2	2	yes	2	Q

Table A6 **Summary of the long-run demand for labour schedules for 10 OECD economies and the share (= α_N) of the deviation from the equilibrium which is corrected within the first half-year.**

	Employment elasticity with respect to output	Employment elasticity with respect to the real labour cost	α_N	p-value related to the preferred structure	R^2 of the difference equation for $\Delta \log N$ with preferred long-run properties
USA	1	–1	–.53	.19	.65
Canada	1	–.5	(–.58)	.14	.88
Australia	1	–1	–.25	.28	.71
France	1	–1	(–.18)	.16	.81
Germany	1	–1	–.15	.23	.71
Japan	1	–.8	(–.15)	.56	.51
Italy	1	–.5	(–.05)	.63	.43
UK	1	–1	–.05	.42	.62
Finland	1	–1	–.05	.48	.40
Sweden	1	–.9	(–.05)	.91	.66

2 Response of Unemployment

In order to close the simulation model, we need to consider the relation between employment and unemployment. According to various documents, in most OECD countries the response of unemployment to changes in employment is far from full. With all variables in logarithmic form, we can write $\Delta u = \rho \, \Delta N^d$, with $0<\rho<1$. In the present study, we do not estimate ρ. Instead, we make conjectures about its magnitude in light of two recent OECD studies.

Turner, Richardson & Rauffet (1993) report estimates for the USA, Japan and Germany. These numbers will be used as such below. Elmeskov & Pichelmann (1993) report estimates for all ten countries examined here. In some countries the estimates differ so much from their estimates for comparable countries that we preferred to do some adjustment. This concerns Australia, Italy, Sweden and to a lesser extent Canada. The figures used in simulations can be seen in the right-hand-side column of Table A7.

In Sweden, the choice of ρ is somewhat problematic. The analysis in Elmeskov & Pichelmann (1993) appears to indicate that a considerable change in ρ has taken place within the observation period. For the earlier part, the "low" value found by Elmeskov & Pichelmann seems to be an appropriate one whereas in the later years a considerably higher value would probably be a more correct choice. Because we are more interested in pressures generated in recent years, we have chosen to work with the higher elasticity reported in Table A7.

Table A7 **Estimates related to the response of unemployment to changes in employment.**

	Elmeskov &Pichelmann (1993)	Turner, Richardson & Rauffet (1993)	Values used in the present study
USA	–.65	–.79	–.79
Japan	–.19	–.28	–.28
Germany	–.52	–.84	–.84
France	–.75	n.a.	–.75
UK	–.67	n.a.	–.67
Italy	–.48	n.a.	–.70
Canada	–.59	n.a.	–.70
Australia	–.50	n.a.	–.70
Sweden	–.43	n.a.	–.70
Finland	–.69	n.a.	–.69

3 Simulation models

For each country, the simulation model consists of three equations. The first two equations incorporate the long-run wage setting and demand for labour schedules estimated above. These relationships which generate the long-run properties of the system are in the square brackets at the end of each equation. Deviations from long-run relationships generate error correction the strength of which depends on the magnitude of the coefficient before the square brackets. The right-hand-side variables in differences derive from the short-run part of equations (v) and (xxvi). Difference terms enter the simulation equations when they differ significantly from zero. Inference on significance derives from the t-ratios.

In all simulation models the unemployment rate is in logarithmic form. This should be kept in mind when wage equations below are compared with those in Table A4. Transformations required were evaluated at the average unemployment rate of the observation period.

GERMANY

Wage equation:

$$\begin{aligned}\Delta\log(W/P)t = & - .55\Delta\log(W/P)_{t-1} + .24\Delta\log(Q/N)_{t-1} + .27\Delta\log(1-\tau_a)_t \\ & + .26\Delta\log(CPI/P)_t + .29\Delta\log(CPI/P)_{t-1} - .15\,[\log(W/P)_{t-1} \\ & - \log(Q/N)_{t-1} - 1.00\log(1-\tau_a)_{t-1} - 1.00\log(CPI/P)_{t-1} + .17 \\ \log(u_{t-1})]\end{aligned}$$

Labour demand equation:

$$\begin{aligned}\Delta\log(N_t) = & .55\,\Delta\log(N_{t-1}) - .15\,[\log(N_{t-1}) + 1.00\log(W(1+s))_{t-1} \\ & - 1.00\log(P_{t-1}) - \log(Q_{t-1})]\end{aligned}$$

The unemployment rate "definition": $\log(u_t) = -.84 \log(N_t)$

CANADA

Wage equation:

$$\begin{aligned}\Delta\log(W/P)_t = & - .33\,\Delta\log(W/P)_{t-1} + .54\,\Delta\log(Q/N)_{t-1} + .51\,\Delta\log(PCP/P)_{t-1} \\ & - .16\,[\log(W/P)_{t-1} - \log(Q/N)_{t-1} - 1.00\log(1-\tau_a)_{t-1} \\ & + .20\log(1-\tau_m)_{t-1} + .20\log(1+s)_{t-1} - .80\log(PCP/P)_{t-1} \\ & + .18\log(u_{t-1})]\end{aligned}$$

Labour demand equation:

$$\begin{aligned}\Delta\log(N_t) = & .62\,\Delta\log(Q_t) + .14\,\Delta\log(Q_{t-1}) - .14\,\Delta\log(W(1+s))_{t-1} \\ & + .19\,\Delta\log(P_{t-1}) - .58\,[\log(N_{t-1}) + .5\log(W(1+s))_{t-1} - .5\log(P_{t-1}) \\ & - \log(Q_{t-1})]\end{aligned}$$

The unemployment rate "definition": $\log(u_t) = -.7 \log(N_t)$

JAPAN

Wage equation:

$$\begin{aligned}\Delta\log(W/P)_t = & - .50\ \Delta\log(W/P)_{t-1} \\ & - .43\ [\log(W/P)_{t-1} - \log(Q/N)_{t-1} - 1.00\ \log(1-\tau_a)_{t-1} \\ & + .50\ \log(1-\tau_m)_{t-1} + .50\ \log(1+s)_{t-1} - .50\ \log(PCP/P)_{t-1} \\ & + .14\ \log(u_{t-1})]\end{aligned}$$

Labour demand equation:

$$\Delta\log(N_t) = -.15\ [\log(N_{t-1}) + .8\ \log(W(1+s))_{t-1} - .8\ \log(P_{t-1}) - \log(Q_{t-1})]$$

The unemployment rate "definition": $\log(u_t) = -.28\ \log(N_t)$

FINLAND

Wage equation:

$$\begin{aligned}\Delta\log(W/P)_t = & .76\ \Delta\log(CPI/P)_t - .23\ \Delta\log(W/P)_{t-1} + .23\ \Delta\log(Q/N)_{t-1} \\ & -.30\ [\log(W/P)_{t-1} - \log(Q/N)_{t-1} - 1.00\ \log(1-\tau_a)_{t-1} + .50\ \log(1-\tau_m)_{t-1} \\ & + .50\ \log(1+s)_{t-1} - .50\ \log(CPI/P)_{t-1} + .31\ \log(u_{t-1})]\end{aligned}$$

Labour demand equation:

$$\begin{aligned}\Delta\log(N_t) = & -.2\ \Delta\log(W(1+s))_t + .35\ \Delta\log(N_{t-1}) \\ & -.05\ [\log(N_t) + 1.00\ \log(W(1+s)_{t-1} - 1.00\ \log(P_{t-1}) - \log(Q_{t-1})]\end{aligned}$$

The unemployment rate "definition": $\log(u_t) = -.69\ \log(N_t)$

FRANCE

Wage equation:

$$\begin{aligned}\Delta\log(W/P)_t = & - .34[\log(W/P)_{t-1} - \log(Q/N)_{t-1} - .42\log(1-\tau_a)_{t-1} + .58\log(1+s)_{t-1} \\ & - .42\log(CPI/P)_{t-1} + .29\log(u_{t-1})]\end{aligned}$$

Labour demand equation:

$$\begin{aligned}\Delta\log(N_t) = & .4\Delta\log(N_{t-1}) + .36\ \Delta\log(Q_{t-1}) - .18[\log(N_{t-1}) + 1.00\ \log(W(1+s)_{t-1} \\ & - 1.00\ \log(P_{t-1}) - \log(Q_{t-1})]\end{aligned}$$

The unemployment rate "definition": $\log(u_t) = -.75\ \log(N_t)$

ITALY

Wage equation:

$$\Delta\log(W/P)t = -.28\ \Delta\log(W/P)_{t-1} - .23\ \Delta\log(u)_{t-1} - .21\ [\log(W/P)_{t-1} - \log(Q/N)_{t-1} - 1.00\ \log(1-\tau_a)_{t-1} + .60\ \log(1-\tau_m)_{t-1} + .60\ \log(1-s)_{t-1} - .40\ \log(PCP/P)_{t-1} + .33\ \log(u_{t-1})]$$

Labour demand equation:

$$\Delta\log(N_t) = .3\Delta\log(N_{t-1}) - .05\ [\log(N_{t-1}) + .50\ \log(W(1+s))_{t-1} - .50\ \log(P_{t-1}) - \log(Q_{t-1})]$$

The unemployment rate "definition": $\log(u_t) = -.7\ \log(N_t)$

AUSTRALIA

Wage equation:

$$\Delta\log(W/P)_t = -.02\Delta\log(u_{t-1}) - .38[\log(W/P)_{t-1} - \log(Q/N)_{t-1} - .50\log(1-\tau_a)_{t-1} + .50\log(1+s)_{t-1} - .50\log(PCP/P)_{t-1} + .23\log(u_{t-1})]$$

Labour demand equation:

$$\Delta\log(N_t) = -.2\ \Delta\log(W(1+s))_{t-1} - .25\ [\log(N_{t-1}) + 1.00\ \log(W(1+s))_{t-1} - 1.00\ \log(P_{t-1}) - \log(Q_{t-1})]$$

The unemployment rate "definition": $\log(u_t) = -.7\ \log(N_t)$

USA

Wage equation:

$$\Delta\log(W/P)_t = .82\ \Delta\log(CPI/P)_t - .05\ \Delta\log(u_{t-1}) - .2\ [\log(W/P)_{t-1} - \log(Q/N)_{t-1} + 1.00\ \log(1+s)_{t-1} - 1.00\ \log(1-\tau_a)_{t-1} + .05\ \log(u_{t-1})]$$

Labour demand equation:

$$\Delta\log(N_t) = .38\ \Delta\log(Q_{t-1}) - .53\ [\log(N_{t-1}) + 1.00\ \log(W(1+s))_{t-1} - 1.00\ \log(P_{t-1}) - \log(Q_{t-1})]$$

The unemployment rate "definition": $\log(u_t) = -.79\ \log(N_t)$

SWEDEN

Wage equation:

$$\Delta\log(W/P)_t = -1.00\ \Delta\log(1+s)_t - .08\ [\log(W/P)_{t-1} - \log(Q/N)_{t-1} + 1.00\ \log(1+s)_{t-1} - 1.00\ \log(CPI/P)_{t-1} + .23\ \log(u_{t-1})]$$

Labour demand equation:

$$\Delta\log(N_t) = -.05\ [\log(N_{t-1}) + .9\ \log(W(1+s))_{t-1} - .9\ \log(P_{t-1}) - \log(Q_{t-1})]$$

The unemployment rate "definition": $\log(u_t) = -.7\ \log(N_t)$

UK

Wage equation:

$$\Delta\log(W/P)_t = -\ .7\ [\log(W/P)_{t-1} - \log(Q/N)_{t-1} - .25\log(1-\tau_a)_{t-1} + .75\log(1+s)_{t-1} - .25\log(PCP/P)_{t-1} + .10\log(u_{t-1})]$$

Labour demand equation:

$$\Delta\log(N_t) = -\ .14\ \Delta\log(W(1+s)_{t-1} + .17\ \Delta\log(Q/N)_{t-1} + .51\ \Delta\log(N_{t-1}) - .05\ [\log(N_{t-1}) + 1.00\ \log(W(1+s)_{t-1} - 1.00\ \log(P_{t-1}) - \log(Q_{t-1})]$$

The unemployment rate "definition": $\log(u_t) = -.67\ \log(N_t)$

3.1 Some simulation results

Simulations have been carried out using the "MAQUETTE" simulation software developed and kindly supplied by Dave Turner. Figure A1 presents a selection of results which clarify the diagnostics of the models defined above. A more comprehensive set is available from the author upon request.

Figure A1

Simulation results: Germany

The simulated effect of a simultaneous one percentage point rise in
1) the average income tax rate, 3) the employers' social security contribution rate,
2) the marginal income tax rate, 4) the indirect tax rate.

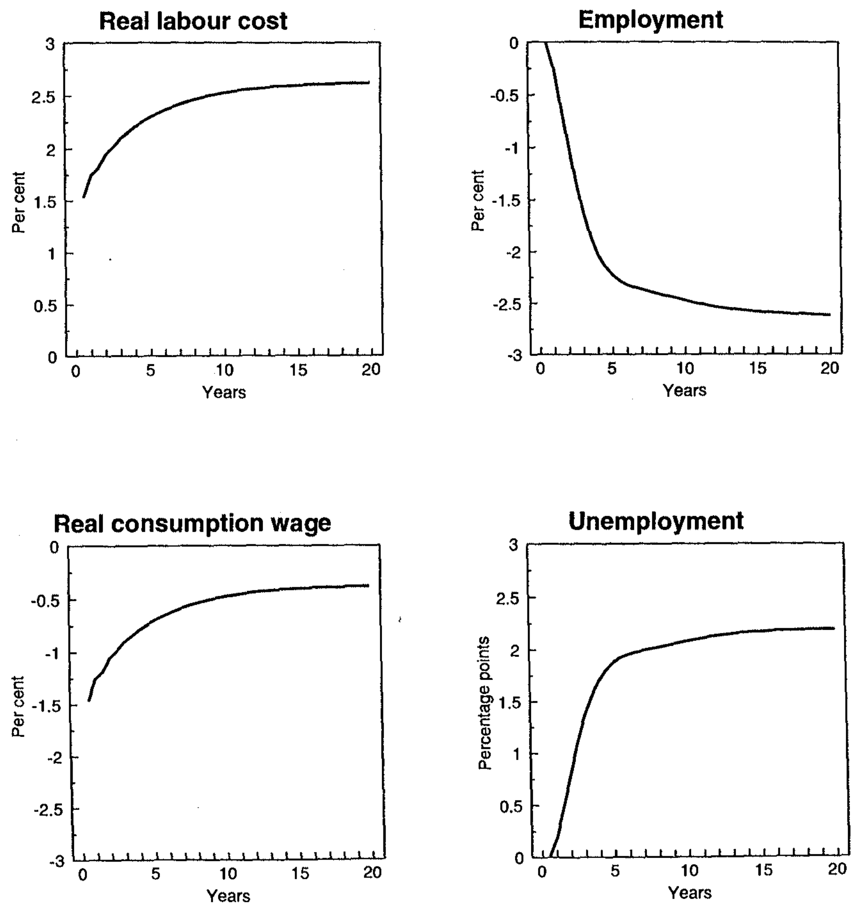


Figure A1 (cont.)

Simulation results: Canada

The simulated effect of a simultaneous one percentage point rise in
1) the average income tax rate, 3) the employers' social security contribution rate,
2) the marginal income tax rate, 4) the indirect tax rate.

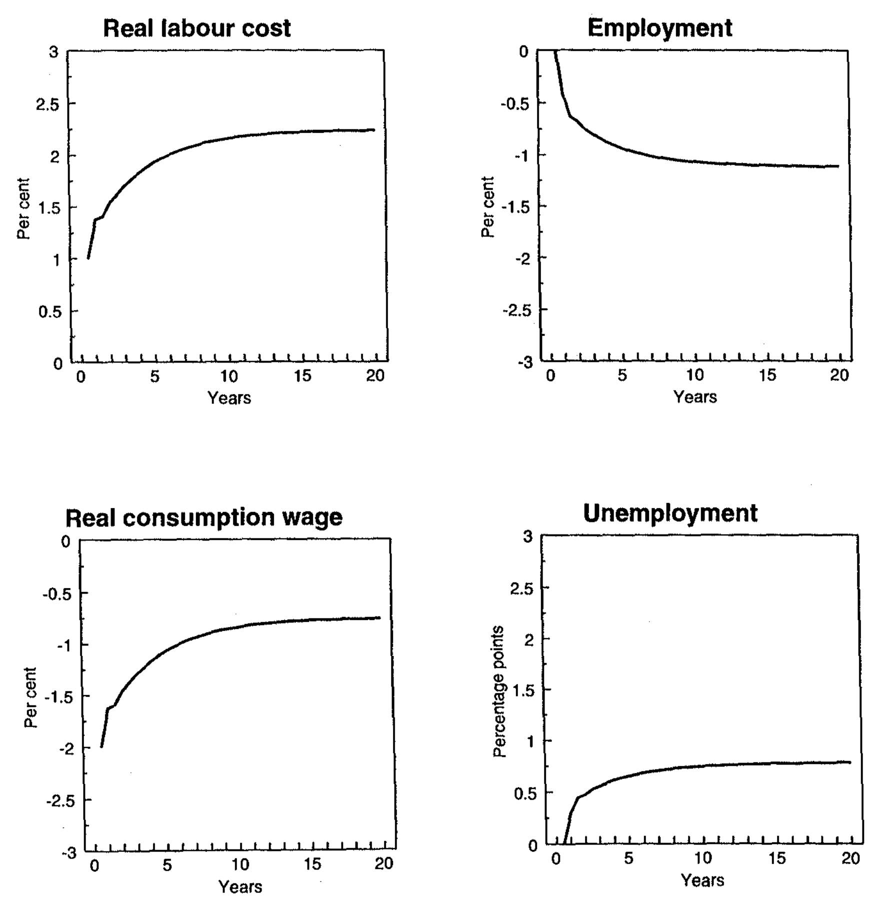


Figure A1 (cont.)

Simulation results: France

The simulated effect of a simultaneous one percentage point rise in
1) the average income tax rate, 3) the employers' social security contribution rate,
2) the marginal income tax rate, 4) the indirect tax rate.

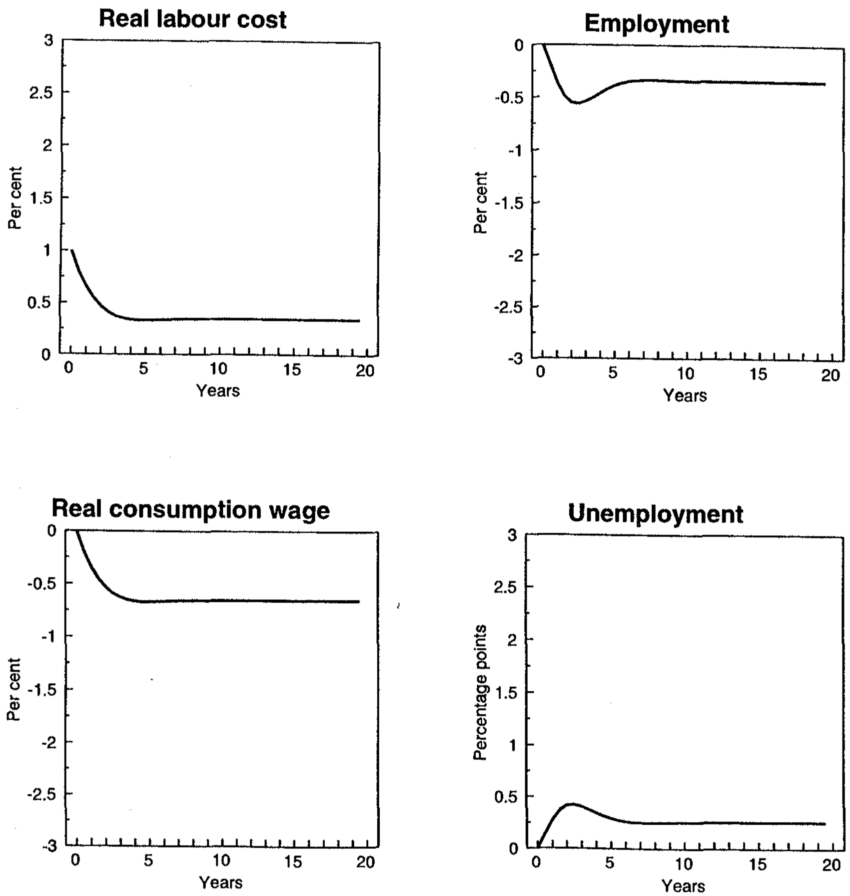


Figure A1 (cont.)

Simulation results: Australia

The simulated effect of a simultaneous one percentage point rise in
1) the average income tax rate, 3) the employers' social security contribution rate,
2) the marginal income tax rate, 4) the indirect tax rate.

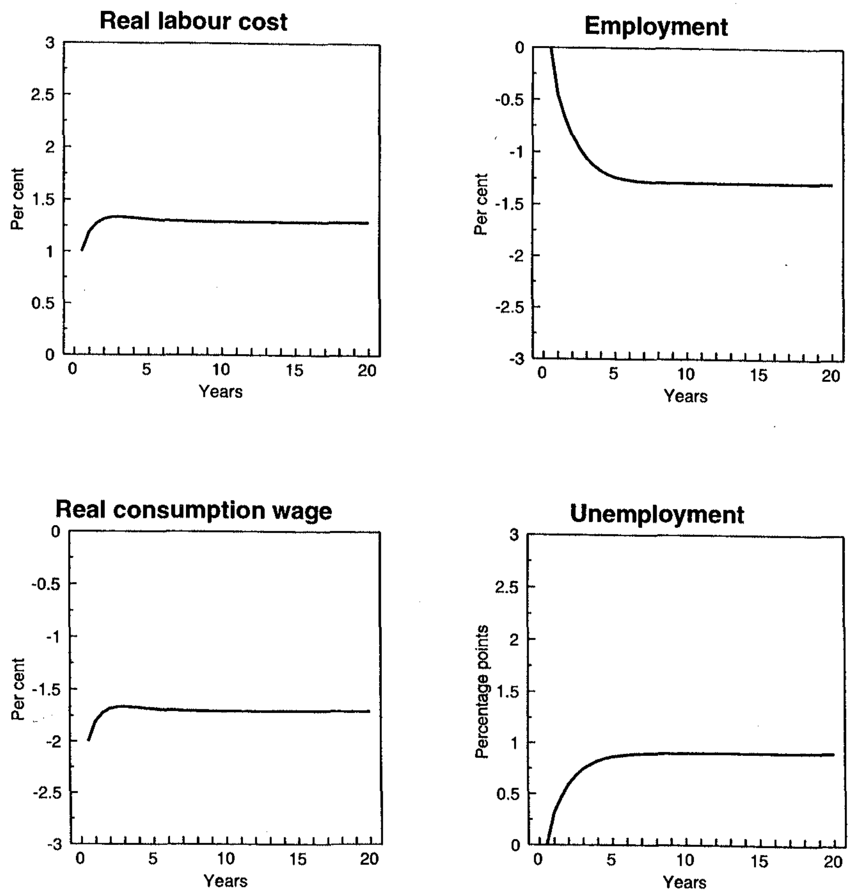


Figure A1 (cont.)

Simulation results: Japan

The simulated effect of a simultaneous one percentage point rise in
1) the average income tax rate, 3) the employers' social security contribution rate,
2) the marginal income tax rate, 4) the indirect tax rate.

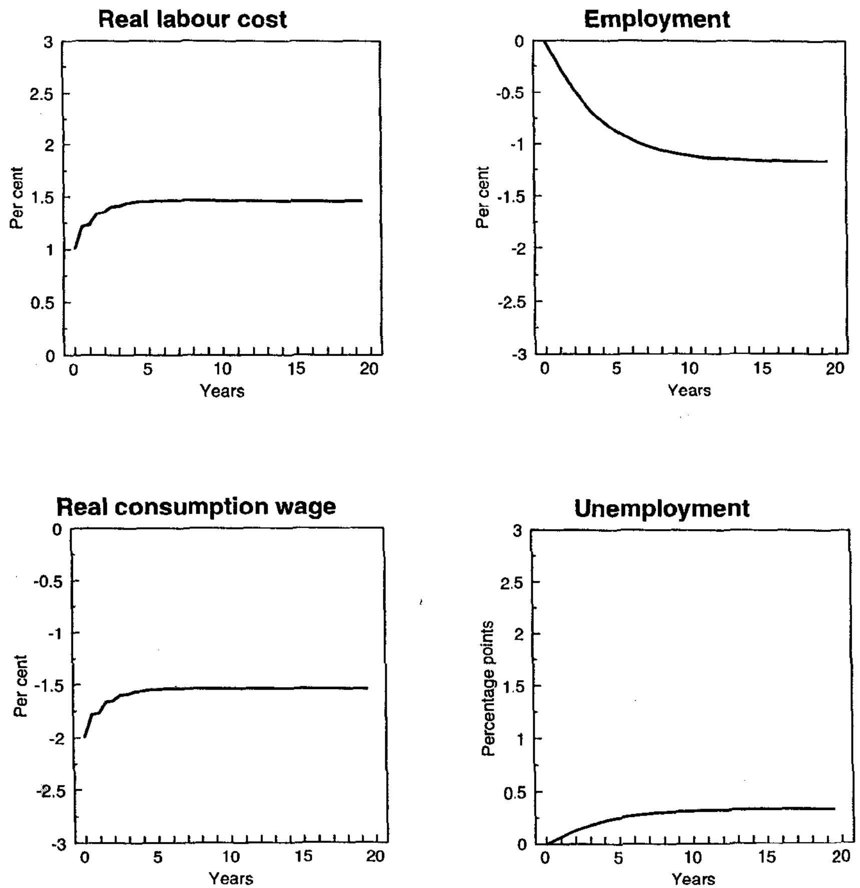


Figure A1 (cont.)

Simulation results: Finland

The simulated effect of a simultaneous one percentage point rise in
1) the average income tax rate, 3) the employers' social security contribution rate,
2) the marginal income tax rate, 4) the indirect tax rate.

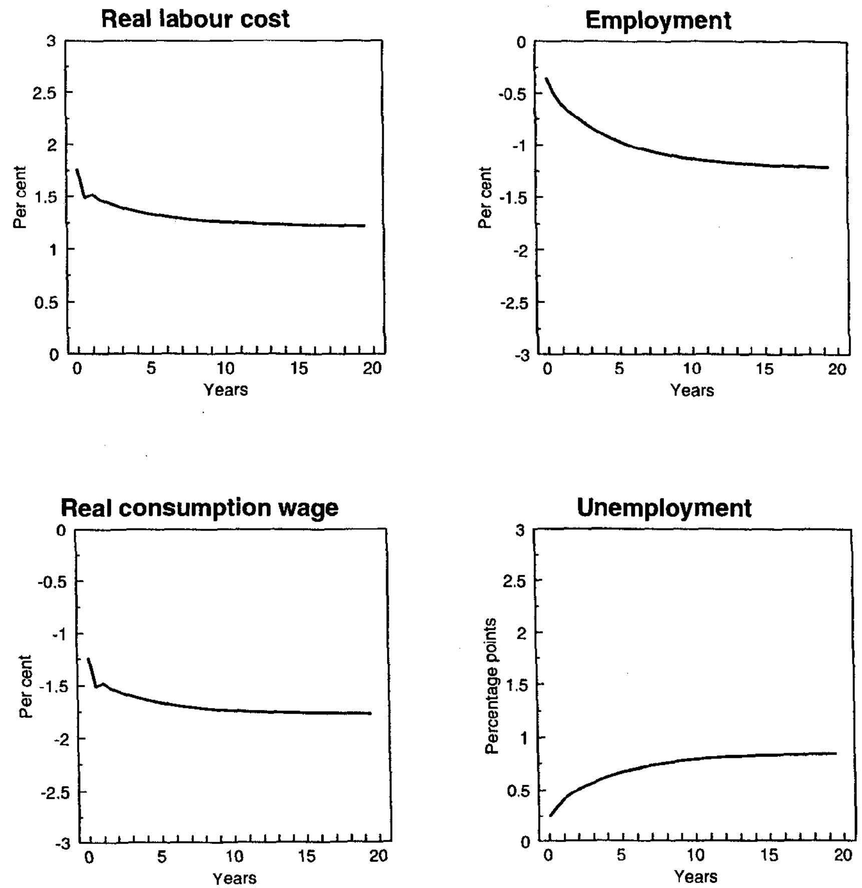


Figure A1 (cont.)

Simulation results: United States

The simulated effect of a simultaneous one percentage point rise in
1) the average income tax rate, 3) the employers' social security contribution rate,
2) the marginal income tax rate, 4) the indirect tax rate.

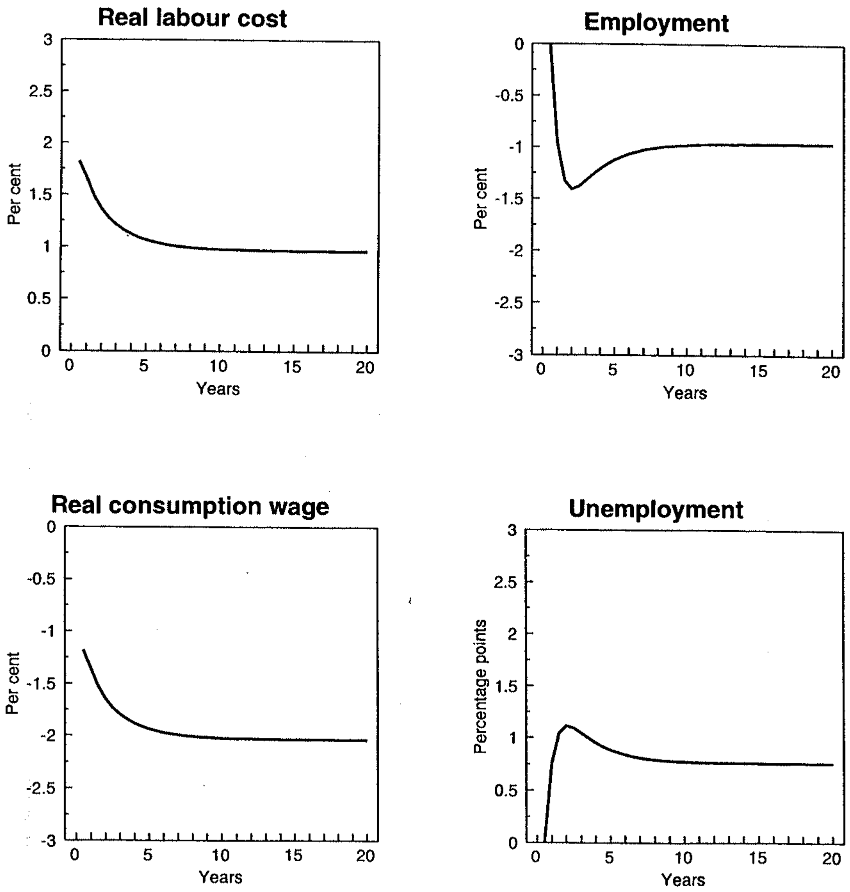


Figure A1 (cont.)

Simulation results: Sweden

The simulated effect of a simultaneous one percentage point rise in
1) the average income tax rate, 3) the employers' social security contribution rate,
2) the marginal income tax rate, 4) the indirect tax rate.

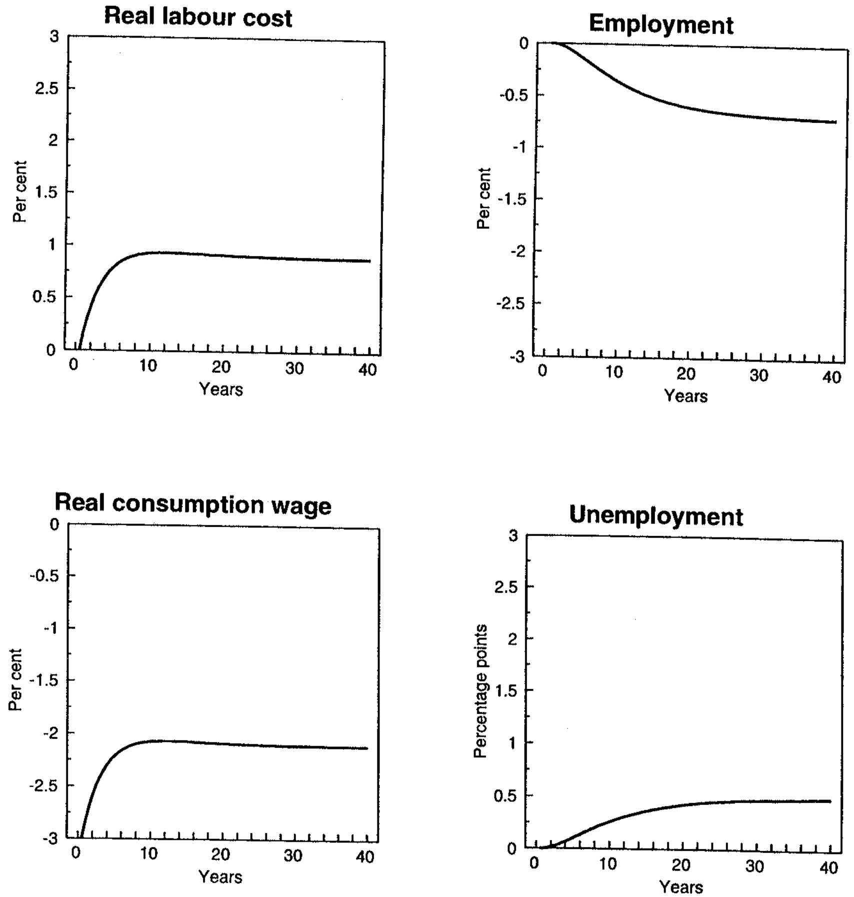


Figure A1 (cont.)

Simulation results: United Kingdom

The simulated effect of a simultaneous one percentage point rise in
1) the average income tax rate, 3) the employers' social security contribution rate,
2) the marginal income tax rate, 4) the indirect tax rate.

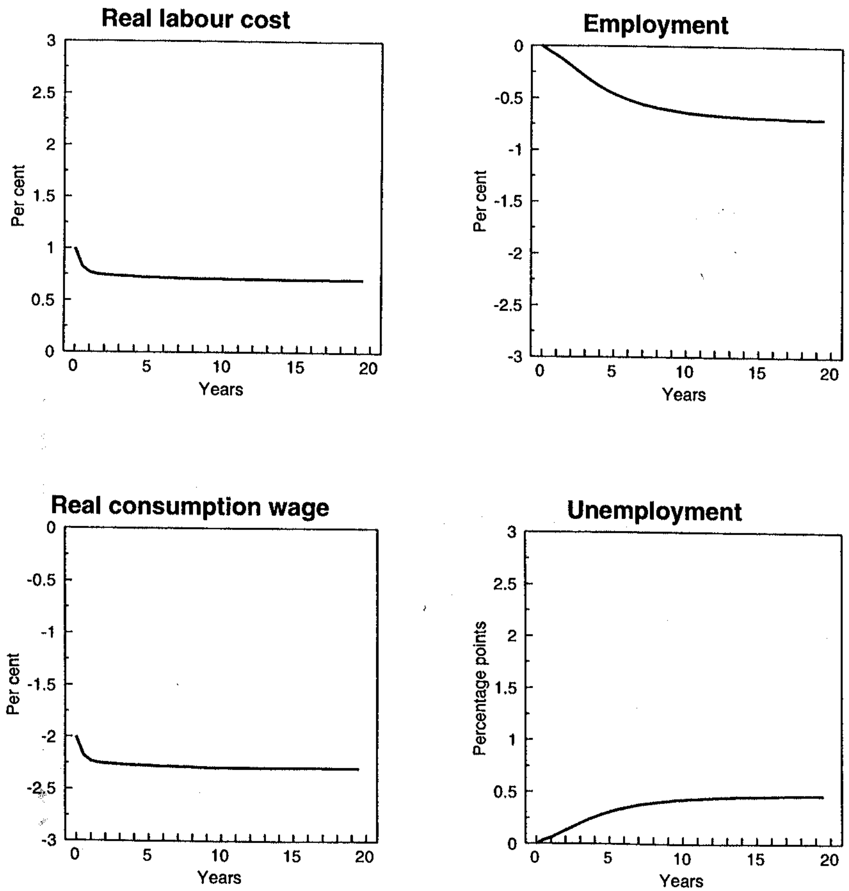
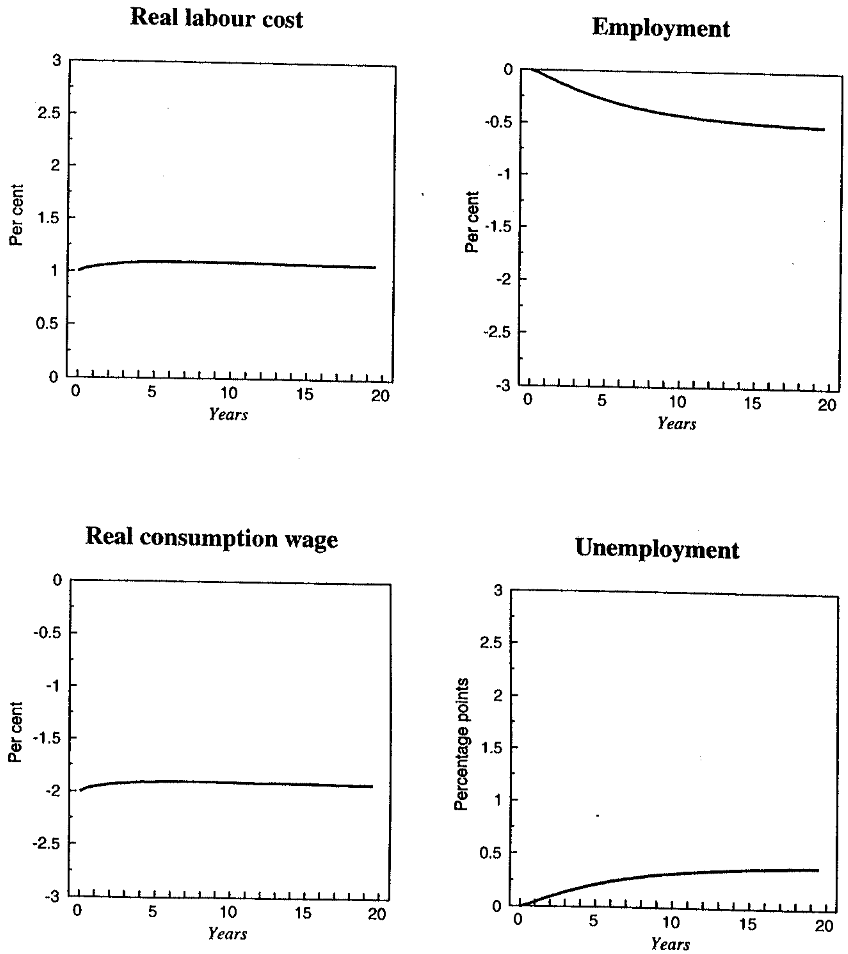


Figure A1 (cont.)

Simulation results: Italy

The simulated effect of a simultaneous one percentage point rise in
1) the average income tax rate, 3) the employers' social security contribution rate,
2) the marginal income tax rate, 4) the indirect tax rate.



3.2 Summary of the properties of the simulation models

In Table A8, the most important results characterizing the simulation properties have been summarized. The results describe the long-run response to increases in all relevant wedge factors by 1 percentage point[62]. In Table A8 (b)—(a) equals minus three which is the effect of the shock on the "cake" the division of which is defined by the Nash maximand in (ii) in the main text.

In Germany and Canada, most of the increase in the wedge leads to higher real labour costs. In Japan, France, Australia and Finland the burden is fairly evenly shared between employees and employers. In the USA, Sweden, Italy and the UK, most of the burden falls on wages.

Evaluation of the third column reveals the impact of the wage elasticity found in the demand for labour relationship. In countries with Cobb-Douglas technology, (c) = – (a). In countries in which we found a CES function as the most appropriate description of the technology, (c) < – (a). This difference is particularly important in Canada and Italy. Finally, the right hand side column reports changes in the unemployment rate when all feedback effects on wages have been taken into account. The difference between (d) and – (c) is determined by the magnitude of ρ (see Table A7 above) which is particularly low in Japan.

Table A9 reports some statistics related to the speed of adjustment. The numbers indicate the time (in years) which it takes before 50 and 80 per cent of the impact on unemployment of a shock on wedge has passed through. Although these estimates should be considered cautiously, it is straightforward to conclude that differences between countries seem to be considerable. Unemployment adjusts to the shock particularly quickly in the USA. Adjustment is fairly rapid in Australia as well. In most of the countries, it takes 4—6 years to conclude 80 per cent of the adjustment. The process is particularly sluggish in Sweden reflecting the small error correction coefficients in both wage and employment equations.

Tables A10—A12 reveal the source of the slowness in adjustment. Table A10 gives a rough approximation of the wage adjustment which follows an exogenous shift in the wedge. Table A11 gives a rough approximation of the wage adjustment following an exogenous shift in the wedge. Table A11 reports simulations with the employment equation only when an exogenous permanent increase by 1 per cent is introduced to the wage level. Table A12 reports simulations with the wage equation only when an exogenous shift of 1 percentage point in the unemployment rate takes place. The reaction of the demand for labour seems to be slowest in Sweden, Italy and the UK although it is slow also in Finland and Japan. The response of wages to higher unemployment has been slowest in Germany and Canada and perhaps in Sweden and Japan.

[62] As far as the countries in which progressivity matters are concerned, it should be noticed that an identical rise in average and marginal tax rates leave the progressivity index unchanged.

Table A8 **The simulated effect of a simultaneous rise of one percentage point in a) employers' social security contribution, b) average income tax rate, c) marginal income tax rate, and d) price wedge, percentage points**

	(a) Real labour cost	(b) Real take home pay	(c) Employment	(d) Unemp- loyment
Germany	+ 2 1/2	− 1/2	− 2 1/2	+ 2 1/4
Canada	+ 2 1/4	− 3/4	− 1	+ 3/4
France	+ 1 1/4	− 1 3/4	− 1 1/4	+ 1
Finland	+ 1 1/4	− 1 3/4	− 1 1/4	+ 1
Australia	+ 1 1/4	− 1 3/4	− 1 1/4	+ 1
USA	+ 1	− 2	− 1	+ 3/4
Sweden	+ 1	− 2	− 3/4	+ 1/2
Japan	+ 1 1/2	− 1 1/2	− 1	+ 1/3
UK	+ 1/2	− 2 1/2	− 1/2	+ 1/2
Italy	+ 1	− 2	− 1/2	+ 1/2

Table A9 **The time it takes for 50 per cent (= median lag) and 80 per cent of the response in unemployment to take place after a permanent shift in the wedge factors, years**

	50 per cent	80 per cent
USA	1[1]	2[1]
Japan	3 1/2	5 1/2
Germany	2 1/2	5
France	2	4
UK	4	6 1/2
Italy	5	9
Canada	2	5
Australia	1 1/2	3
Finland	1 1/2	5
Sweden	9	15

[1] In the USA, a quick overshooting increase in unemployment takes place in the beginning of the post-shock period (see Figures A1 and A2).

Table A10 **Diagnostics of the wage equation[1]: (a) The long-run effect on wage level of an exogenous, permanent increase in a wedge factor[2] by 1 percentage point and the time, in years, which it takes for (b) 50 per cent (= median lag) and (c) 80 per cent of the effect to pass through.**

	(a) The effect on real labour cost level, percentage points	(b) The time for 50 per cent of the passthrough, years	(c) The time for 80 per cent of the passthrough, years
USA	+ 1	2–3	4–5
Japan	+ 1 1/2	2	3
Germany	+ 3	2–3	7–8
France	+ 1 1/4	2	3
UK	+ 3/4	1	2
Italy	+ 1	3	7
Canada	+ 2 1/2	3	5
Australia	+ 1 1/2	2	3
Finland	+ 1 1/2	2	4
Sweden	+ 1	3	6

[1] In these simulations all feedback effects through employment and unemployment have been ignored.

[2] Please note that reactions differ from one element of the wedge to an other. For example, the direction of the response to an increase in employers' social security rates is usually opposite to that related to other wedge factors. In most countries, employers' social security contributions add fully labour costs in the first event and the impact reduces over time if the long-run response is below unity. The response to other tax variables is often nill at the outset and the push effect starts to influence over time. Any numbers which aggregate these kind of opposite processes must be considered cautiosly.

Table A11 **Diagnostics of the employment equation[1]: (a) The long-run effect on employment of an exogenous, permanent 1 percentage point increase in the real wage level and the time, in years, which it takes for (b) 50 per cent (= median lag) and (c) 80 per cent of the effect to pass through**

	(a) Effect on the employment level, percentage points	(b) Time for 50 per cent of the effect to pass through, years	(c) Time for 80 per cent of the effect to pass through, years
USA	– 1.0	1	1 1/2
Japan	– 0.8	3	5
Germany	– 1.0	2	3
France	– 1.0	2	3
UK	– 1.0	4 1/2	7
Italy	– 0.5	5	9
Canada	– 0.5	1	1 1/2
Australia	– 1.0	1 1/2	3
Finland	– 1.0	3	7 1/2
Sweden	– 0.9	7 1/2	20

Table A12 **Diagnostics of the wage equation[1]: (a) The long-run effect on the wage level of an exogenous, permanent 1 percentage point increase in the unemployment rate and the time, in years, which it takes for (b) 50 per cent (= median lag) and (c) 80 per cent of the effect to pass through**

	(a) Effect on the wage level, percentage points	(b) Time for 50 per cent of the effect to pass through, years	(c) Time for 80 per cent of the effect to pass through, years
USA	– 1	1	1
Japan	– 5	3	5 1/2
Germany	– 3	4	6
France	– 3 1/2	1 1/2	3
UK	– 1	1	1 1/2
Italy	– 3 1/2	2	4
Canada	– 2	3 1/2	5 1/2
Australia	– 4	1 1/2	2 1/2
Finland	– 6	2	3 1/2
Sweden	– 10	2 1/2	9 1/2

[1] In these simulations feedback effects through other equations have been ignored.

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ISSN 0785-3572

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