Business cycles: real facts or fallacies?*

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Abstract

We contrast two alternative approaches to the analysis of price behaviour over the business cycle: the "stylized facts" methodology and econometric models. Using the "stylized facts" apparatus on United Kingdom and Norwegian data, we find that the relationships between wages and unemployment, are unstable over even short sample periods. We design two small wage–price models for the United Kingdom and Norway that appear constant, interpretable, data coherent and that can account for the instabilities of the "stylized facts" correlations. Simulation is used to show that even when underlying behaviour is constant, we can expect to observe quite different correlations of macroeconomic variables over the business cycle, hence we encompass the result of the "stylized facts" analysis.

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Contents

1 Introduction

Conflicting views on the causes and nature of business cycle fluctuations can usually be attributed to different beliefs about the adjustment of prices and wages. For example, Kydland and Prescott (1990)—*KP* hereafter—conclude that US prices have been countercyclical since 1954 and interpret this as evidence against demand driven models of the cycle. These findings are based on *KP*'s innovative "stylized facts" methodology which is further developed in Kydland and Prescott (1991) and contrasted with the econometric "system-of-equations" approach of Koopmans (1947, 1949).

The main reason for *KP*'s dismissal of econometric models seems to be that results derived from "system-of-equation models" are model-dependent and represent biased evidence on the nature of business cycle fluctuations, e.g. the "myth" that prices behave procyclically. As an alternative, *KP* offers their own "stylized facts" methodology which involves only a minimum of assumptions, thus allowing the data to speak directly, instead of through the veil of an econometric model: *KP* apply a filtering technique known in the economics literature as the Hodrick-Prescott (*HP*) filter to the raw data in order to identify and remove the trend in the individual series. The filtered or "cyclical" series are then used to analyse business cycle fluctuations by calculation of bivariate correlations between e.g. detrended prices and *GDP* at different leads and lags. The full set of bivariate correlations constitute the "*stylized facts*".

"Stylized facts" methodology have been successful in attracting the interest of business cycle analysts. Blackburn and Ravn (1992) find results for the United Kingdom that are consistent with *KP*'s results: that prices are countercyclical and the real wage is procyclical. Other studies using the "stylized facts" approach convey a more mixed picture, see Brandner and Neusser (1992), Danthine and Girardin (1989), Correia, Neves, and Rebelo (1992), Englund, Persson, and Svensson (1992) and Schlitzer (1993). Although some of these studies find countercyclical prices, an even more conspicuous finding is that correlations change over sample periods and across countries. Thus Englund et al. (1992), using Swedish data for 128 years, find unstable correlations between wages and *GDP*-output. For a much shorter sample period and on quarterly data, Blackburn and Ravn (1992) find large fluctuations in the correlation between prices and output.[1] Building on a multi-country study, Andersen (1994, p. 19-18) reports instabilities both across time and between countries. Instabilities of this kind obviously represent a problem for business cycle analysis, since, to constitute stylized facts, the "statistical properties of cyclical fluctuations must remain broadly invariant to the passage of time".[2]

KP makes no distinction between the "system-of-equations" approach advocated by Koopmans and modern econometric models and methodology. In this paper we reconsider the econometric approach of explaining business cycle correlations, taking account of the last 10-15 years of methodological debate and subsequent developments see e.g. Granger (1990) and Banerjee, Dolado, Galbraith, and Hendry (1993, Chapter 9) for discussion and overview. One product of this debate has been a renewed interest in several of the concepts and problems that figured markedly

[1] i.e. between -0.8 and 0, see Blackburn and Ravn (1992, Figure 5).

[2] Blackburn and Ravn (1992, p. 393)

during the formative years of econometrics, see Morgan (1990, p. 262-263). These include dynamic systems, which was essential to Tinbergen's models in the 1930s but which died away under the weight of static Keynesian systems, and the concept of autonomy of relationships which was introduced by Ragnar Frisch and further discussed and refined by Trygve Haavelmo, see Frisch (1938), (1948) and Haavelmo (1944).[3]

Broadly speaking a relationship is the more autonomous the wider the range of circumstances under which the relationship is valid. Several aspects of the concept of autonomy reappear in the recent discussion of "exogeneity", "fragility" and "structure", e.g. Engle, Hendry, and Richard (1983), Pagan (1987), and Hendry (1993). Relationships with low degree of autonomy Frisch and Haavelmo called confluent relationships. Today, the distinction between autonomous and confluent relationships is drawn in several contexts, e.g. between identified cointegration relationships and spurious regressions, see Hendry and Morgan (1989) and Hendry (1993).

In this paper we demonstrate that "system-of-equations" models have much to offer in explaining business cycles generally and specifically in resolving the puzzles that have arisen from the applications of the "stylized facts" methodology. A main theme in our account is that while regime shifts and structural breaks will necessarily induce changes and fluctuations in bivariate correlations, this is not necessarily true for econometric relationships. Thus, invoking Frisch's terminology, *KP*'s bivariate correlations constitute relationships that are confluent *relative* to econometric relationships—they are less autonomous. Conversely, changing correlations constitute information that is vital for the discovery of structure in the form of stable and invariant econometric relationships, i.e. relationships with a certain degree of autonomy.

Central to our approach is the notion of *structure* as the set of invariant features of the economic mechanism, see Hendry (1993) and Morgan and Hendry (1995). Invariance under extensions of the information set (time, regime, sources of information) is a key property of structure. Conversely, discovering structure implies a certain degree of autonomy. Necessary attributes to structure are testable, e.g. stability, even if sufficient ones are not. Hence claims of structural interpretation of *models* can be tested, Ericsson and Irons (1994) collect several practical examples. However, models are simple constructs relative to the complex real systems they mimic—so we can only determine partial structure, and invariance properties are always relative.

The rest of the paper is organized as follows. Section 2 briefly reviews the use of the *HP* filter in extracting trends. The use of simple bivariate cross-correlations of detrended data to analyse business cycles is shown to yield results which are nonrobust and may actually hide existing structure. We end Section 2 by a brief account of our own methodology. Section 3 describes a theoretical wage-price model, which we find has the potential to encompass both the empirical results of *KP* —and the earlier results they reject.

In Section 4 the model derived in Section 3 is estimated both on Norwegian and United Kingdom data. The empirical models are stable and convey partial

[3] See Morgan (1990, Chapter 4.4) on Frisch's views on Tinbergen's (first) League of Nations report, and Aldrich (1989) for a history of the concept of autonomy. An excellent discussion of autonomy is also found in Johansen (1977, Chapter 4.4).

structure in the form of identified cointegrated relations. Marginal analysis of these econometric models reported in Section 5 shows that the observed cross-correlations of macroeconomic variables depend entirely on the original source of the cyclical fluctuation. We argue that, since observed business cycles in the past 50 years have arisen from many different sources, it is not surprising that it is possible to obtain apparently contradictory bivariate correlations. Furthermore, we are able to account for unstable correlations by means of stable econometric models. Section 6 offers a brief summary and conclusions.

2 Business cycle analysis

We start by giving a critical account of the "stylized facts" methodology and then sketch the ingredients in an alternative econometric approach to business cycle analysis.

2.1 The *KP*-approach

The studies cited above generate cross correlations from detrended data using the *HP* filter, which identifies a time-varying trend based on the following minimization:

$$\min_{\tau_i} \sum_{t=1}^{T} (y_t - \tau_t)^2 + \lambda \sum_{t=2}^{T-1} [(\tau_{t+1} - \tau_t) - (\tau_t - \tau_{t-1})]^2 \tag{1}$$

for $i = 1, \ldots, T$ where y_t is the series of interest, τ_t is the trend and λ is a positive parameter to be determined.[4] The first part of equation (1) encourages the trend to follow the actual series, while the second part is minimized if the trends behaves smoothly. Several potential problems with the statistical properties of the filter are discussed by King and Rebelo (1993), Harvey and Jaeger (1993) and Cogley and Nason (1995). For example Cogley and Nason (1995) show that when applied to persistent data series, the *HP* filter can generate spurious cycles. This is worrying because experience shows that many macroeconomic series appears to be well approximated by stochastic trends. Despite these reservations, the filter may be useful in analysing business cycles—see Section 4. Our point here is simply that its limitations need to be recognized and interpretations adjusted accordingly.

Computing cross-correlations, at different leads and lags, is the second step in the "stylized facts" methodology and is our main concern here. There is always a limit to the usefulness of bivariate correlation techniques—especially when dealing with two variables which are both endogenous to the macro economy. It is by no means obvious that real *GDP* should have a stable correlation coefficient with any of the variables examined by *KP*—production inputs, output and income components, monetary aggregates, the price level or real wages.

In the following we concentrate on the cyclical properties of nominal wages and prices and of real wages. Our own computations of the *GDP*-inflation correlations shows the expected instability and thus corroborate the findings in e.g. Andersen

[4] Kydland and Prescott (1990, p. 8) conclude that $\lambda = 1600$ is a "reasonable" value for US $\log GNP$.

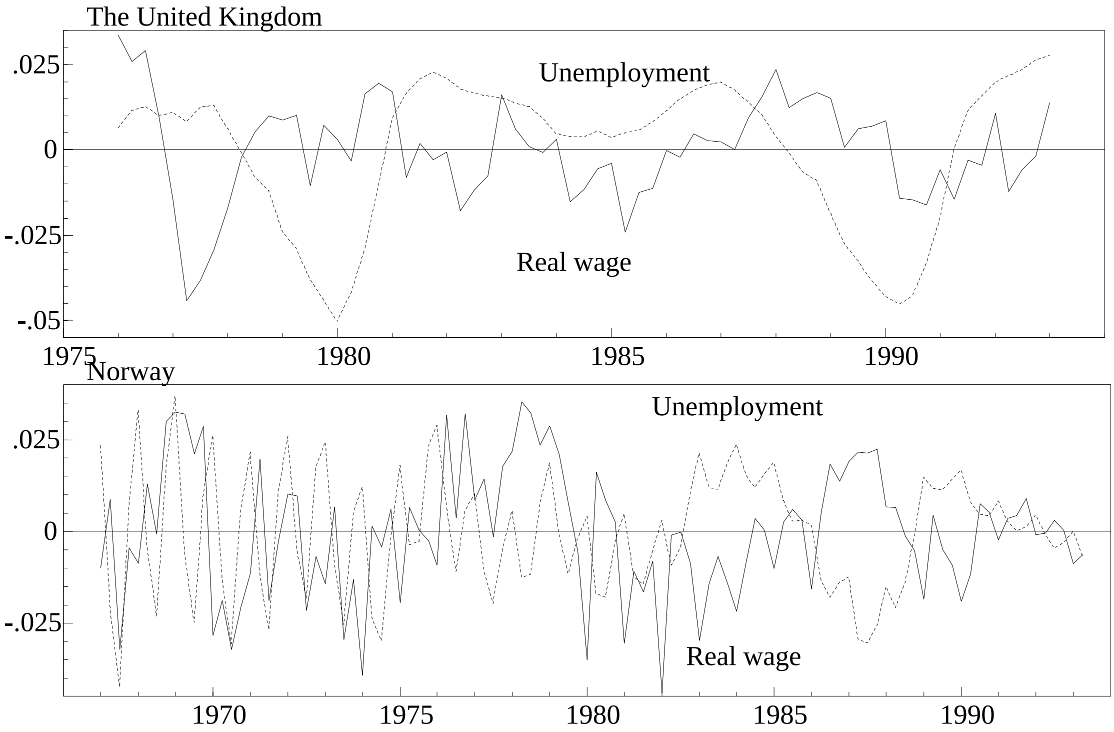

Figure 1: Deviations from Hodrick-Prescott trends for unemployment and real wage, scales adjusted.

(1994, Figure 2.2). The same is true for real wages and unemployment, which is more relevant in the light of the model of price-wage adjustment discussed below. Figure 1 shows the *HP* cycles for consumer real wages and unemployment, using quarterly data for Norway and the United Kingdom.[5] We then consider the correlation between detrended unemployment and real wages at time t and at 4-period leads and lags of real wages. We compute *rolling* correlation coefficients—using a fixed window width of 32 quarters. The result of these exercises are reported in Figure 2. The correlation coefficients are numerically unstable and they even change sign over the sample periods.

One could suspect that maybe the window width of eight years is too small to obtain robust information. However, very little improvement in the results are obtained when the correlations are computed *recursively*, as Figure 3 demonstrates. One possible interpretation of graphs like Figures 2 and 3 is that there is little structure to discover, the cyclical properties of real wages are evidently very transient and little can be learnt, at least with linear models. This however is too pessimistic, since bivariate correlations may actually hide structure. To see this, note that the correlation r_{yx}^2 can be written as

$$r_{yx}^2 = \hat{\beta}\,\hat{b} \tag{2}$$

[5]In estimating the trends we use $\lambda = 1600$.

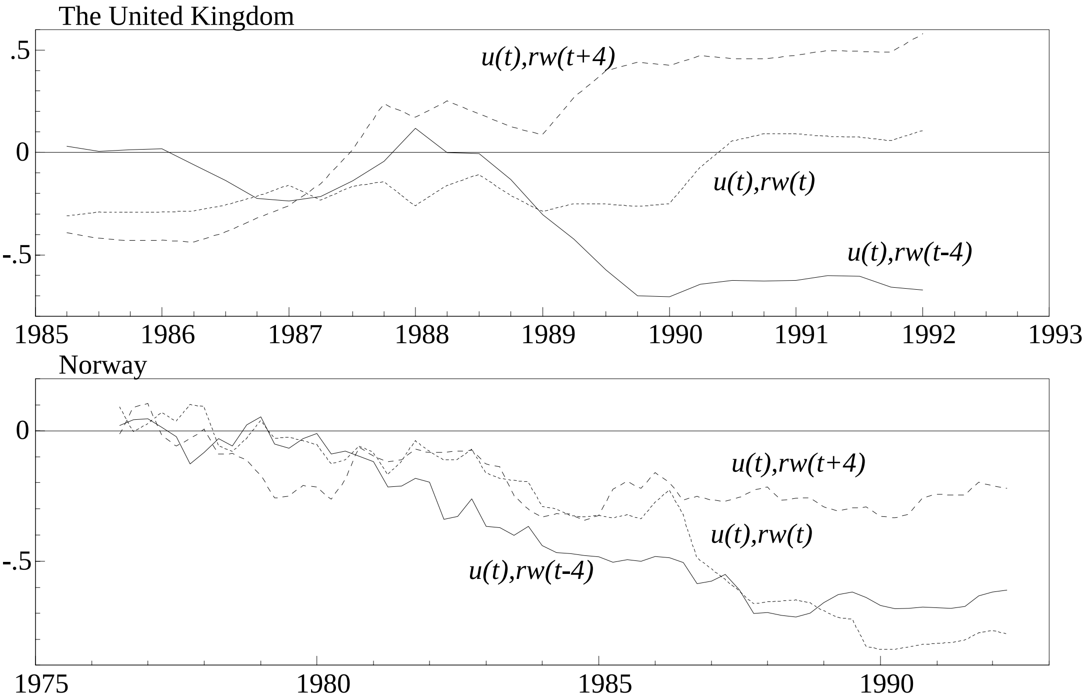

Figure 2: Rolling correlation coefficients between unemployment and the real wage, using a window width of 32 quarters.

where $\hat{\beta}$ is the OLS estimate from the regression model

$$y_t = \alpha + \beta x_t + \varepsilon_{w,t}, \tag{3}$$

and $\hat{b}$ is the estimate from the inverse regression

$$x_t = a + b y_t + \varepsilon_{y,t}. \tag{4}$$

Instability in r_{yx}^2 does not in itself imply that $\hat{\beta}$ is recursively unstable— it is possible that all instability in the correlation coefficient is matched by the inverse regression estimate $\hat{b}$. In this (possibly rare) case equation (3) is invariant to the regime shift made manifest by the instability in the correlation coefficient, and hence it can be claimed that equation (3) has a structural interpretation. Note that it is the occurrence of regime shifts that makes it possible to discover the structural nature of (3), hence they represent useful information, not a nuisance. Conversely, the correlation coefficient is non-structural by definition and therefore is unlikely to be stable even when the underlying behaviour is invariant.

In practice we will not expect to discover structure from simple bivariate regressions like equation (3) very often. But the same logic goes through for multivariate models, with r_{yx} interpreted as a partial correlation coefficient. Thus, it is changes in correlations that makes it possible to refute that an estimated relationship represents structure, i.e. the relationship breaks down. Non-rejection of invariance over the sample strengthens the claim of a structural interpretation, although any invariance property of econometric models is always relative to a set

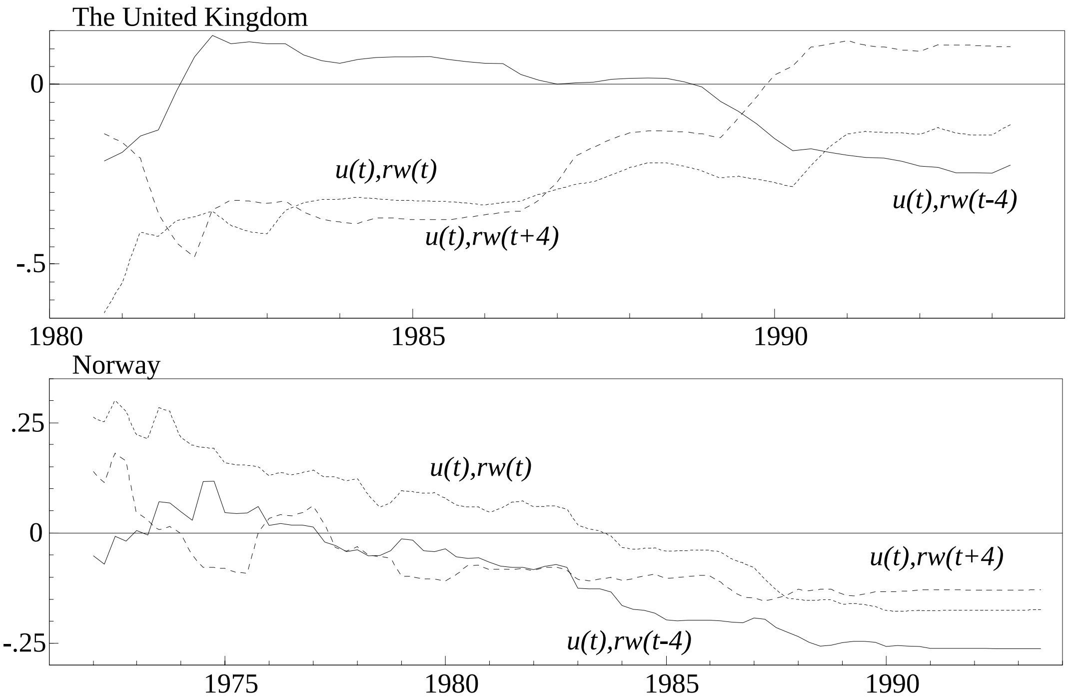

Figure 3: Recursive correlation coefficients between unemployment and real wage, using a starting period of 12 quarters.

of interventions, thus autonomy and structure are relative concepts—see Haavelmo (1944) and Hendry (1993).

2.2 A macro-econometric approach

In the rest of this paper we set out to investigate whether econometrics can help resolve the puzzles that have arisen from modern business cycle analysis and elucidate the cyclical properties of aggregate wages and prices. Before we set out it may be helpful to summarize our approach under the following three headlines.

1. Multivariate framework

We use a multivariate framework as opposed to *KP*'s bivariate framework. The outlines of the models that we estimate follow from a simple theoretical model discussed in Section 3 below. In economics we cannot carry out true laboratory experiments to isolate cause and effect. In interpreting the data we need to allow for all possible influences—hence the prevalence of multivariate equation systems. As we will show in Section 3, the observed correlations can depend entirely on the source of stochastic variation. In particular, the trend estimation using the *HP* filter smooths over different cycles and the estimated correlation coefficients represent average relationships. Hence a demand driven contraction is treated identically to a supply driven contraction, even though the signs of bivariate correlations may be reversed.

2. Cointegration analysis

Experience shows that actual data for wages and prices contain stochastic trends, in such a way that, for example, wages w_t is integrated of order 1, which we write as $w_t \sim \mathsf{I}(1)$. Postulated relationships between $\mathsf{I}(1)$ variables may prove to be spurious, but this possibility can be tested by using cointegration analysis, see Engle and Granger (1987) and Johansen (1988). In a multivariate set-up there may be more than one relationship holding between the $\mathsf{I}(1)$ variables, hence there is a problem of identifying the structural cointegration vectors, a problem that is related to Frisch's confluence analysis, see Hendry and Morgan (1989). Identification of cointegration vectors draws on subject matter economic theory and thorough testing of the relevant restrictions on the cointegration space, see Johansen and Juselius (1992). Inclusion of empirically valid long-run cointegration relationships in the model of cyclical price behaviour is important for theory consistency and it identifies a separate propagation mechanism: equilibrium correction with respect to underlying equilibrium relationships. Finally, inclusion of identified cointegrating relationships enhances structure in that they remain invariant to expansions of the information set, see e.g. Hendry (1993) and Banerjee et al. (1993, Chapter 9.2).

3. Simultaneous equation model

Embedding identified cointegration vectors in dynamic adjustment equations brings the analysis into $\mathsf{I}(0)$-space and ensures economic interpretability and econometric stability of the steady state of the system. The importance of this stage thus lies in the possibility that the simultaneous equation model can parsimoniously encompass the unrestricted reduced form, see Hendry and Mizon (1993) and Bårdsen and Fisher (1995).

3 A theoretical model of wage-price adjustment

In this section we set out a simple model of wage and price interactions which might account for the correlation results reported in KP. We use a wage-price system for an open economy, based on theories of imperfect competition in goods and labour markets. The model is adapted from Kolsrud and Nymoen (1996) who discuss its dynamic properties in more detail. The variables used are nominal wages W, retail prices P, productivity PR, the import price deflator PI, the unemployment rate U, the employers tax-rate $T1$, the income tax-rate $T2$, the tax rate on the retail price basket $T3$, and the deviation gap of output from trend. Lower case letters denote logs, except that (for convenience) $ti = ln\,(1 + Ti)$, $i = 1, 3$, and $t2 = \ln(1 - T2)$.

Theories of wage-formation, see e.g. Carlin and Soskice (1990) and Lindbeck (1993), states that the bargained nominal wage level depends on

- firm-side variables, e.g. productivity, producer prices and the payroll-tax rate
- factors affecting workers' take home pay, e.g. consumer prices and the income tax-rate
- labour-market pressure
- institutional features, e.g. the existence of centralized wage-bargaining institutions and the degree of mismatch.

These items are represented by the following text-book wage equation:

$$w_t^* = \alpha_1 pp_t + (1-\alpha_1)\, p_t + \theta_1 pr_t - \beta_1 t1_t - \kappa_1 t2_t - \zeta_1 u_t \tag{5}$$
$$0 \le \alpha_1 \le 1, \quad \beta_1, \theta_1, \kappa_1, \zeta_1 \ge 0.$$

We have in mind a log-linear relationship, hence w_t^* denotes the logarithm of the predicted wage level. The variables pp_t, pr_t, and $t1_t$ represent the firm-side variables. Institutional and structural features are reflected in the coefficients of (5). Changes in the impact of institutions on wage setting can therefore be tested by looking at the empirical stability of (5) over the sample period. To simplify the exposition, we will in the following set $\theta_1 = 1$ and $\kappa_1 = 0$. Both restrictions are tested and accepted for the empirical models in Section 4.

In an economy with imperfect competition firms set their equilibrium prices to reflect a constant mark-up m_2 over normal unit-labour costs:

$$pp_t^* = m_2 + w_t - pr_t + t1_t. \tag{6}$$

Equation (6) is kept deliberately simple. For example, the pricing to market hypothesis suggest that producer prices depend on competing prices in addition to unit labour costs, cf. Naug and Nymoen (1996). However, in the following sections we focus on the modelling of consumer price behaviour, and in that equation a separate pricing to market coefficient would not be identifiable.

Wage growth is a function of expected price growth, Δp_{t+1}^e and Δpp_{t+1}^e and last period's deviation from the desired wage level (i.e. equation (5)):

$$\Delta w_t = c_1 + \gamma_{1,1} \Delta p_{t+1}^e + \gamma_{1,2} \Delta pp_{t+1}^e - \delta_1 (w - w^*)_{t-1} + \epsilon_{1t}, \tag{7}$$
$$0 \le \gamma_{1,1} + \gamma_{1,2} \le 1, \quad \delta_1 \ge 0.$$

Superscript e denotes "expectation" and ϵ_{1t} is an error term. As already noted, integratedness of the price and wage series is an important assumption underlying our work. Consequently, Δw_t, Δp_{t+1}^e and Δpp_{t+1}^e are all assumed to be stationary $\mathsf{I}(0)$, and furthermore (7) can only become a balanced equation if $(w - w^*)_{t-1} \sim \mathsf{I}(0)$, i.e. the wage level predicted by theory is *cointegrated* with the actual wage level, see e.g. Banerjee et al. (1993) for an exposition. Conversely, the equilibrium-correction term $\delta_1 (w - w^*)_{t-1}$, with $\delta_1 > 0$, is implied by cointegration. Hence, in the empirical sections we first test that (5) is indeed a valid cointegrating equation before we proceed to the dynamic equation (7).

In the short run (i.e. the capital stock is fixed) the marginal cost curve is upward sloping and hence increase in output exerts a positive pressure on prices. The variable gap_t represents the deviation of log output from trend. In addition product price inflation is governed by expected wage growth in excess of productivity growth $\Delta (w - pr)_{t+1}^e$ as well as by corrections from last period's deviation from the equilibrium price:

$$\Delta pp_t = c_2 + \gamma_2 \Delta (w - pr)_{t+1}^e - \delta_2 (pp - pp^*)_{t-1} + \zeta_2 gap_{t-1} + \epsilon_{2t}, \tag{8}$$
$$0 \le \gamma_2 \le 1, \quad \delta_2, \zeta_2 \ge 0.$$

Theoretical models of price adjustment under imperfect information lead expressions similar to (8). For example, if the information set-up is such that agents only get to know the optimal full-information price with a lag, then it is rational to adjust current prices with respect to the difference between the optimal and actual period $t-1$ price, see Andersen (1994, Chapter 6.3).

The model is closed in two steps. First, the import price in domestic currency pi_t is treated as exogenous in the consumer price index equation, together with the indirect tax-rate $t3_t$:

$$p_t = \eta\, pp_t + (1-\eta)\, pi_t + t3_t, \;\; 0 < \eta < 1. \tag{9}$$

Second, expectations develop according to

$$\begin{aligned} \Delta pp^e_{t+1} &= \Delta pp_t, \\ \Delta p^e_{t+1} &= \Delta p_t, \\ \Delta (w - pr)^e_{t+1} &= \Delta (w-pr)_t \,. \end{aligned} \tag{10}$$

One interpretation of (10) is that forward-looking aspects of price and wage adjustment arise from data-based predictors rather than model-based (e.g. rational) predictors; cf. Campos and Ericsson (1988), For an integrated process x_t which is $\mathsf{I}(d)$, a simple forecast is obtained from $\Delta^{d+1} x_{t+1} \approx 0$. For $d = 1$, that implies $\Delta x^e_{t+1} = \Delta x_t$ or that $x^e_{t+1} = x_t + \Delta x_t$. x^e_{t+1} is unbiased if Δx_t is $AR(q)$ with symmetrical error process and could be 'rational' if information is costly, see Hendry and Ericsson (1991) for a discussion. However, if agents do act on model based expectations (10) is obviously wrong and the Lucas critique applies. Ultimately, it is an empirical question whether structural models must build on theory of dynamic optimizing behaviour and model consistent rational expectations.

The theoretical model (5)–(10) is easily condensed to a simultaneous equations model for wages and consumer prices:

$$\begin{aligned} (11)\quad \Delta w_t &= -\delta_1 \left[(w - p - pr) + (a_1 - \alpha_1)(pi - p) + \zeta_1 u + \beta_1 t1 + a_1 t3 \right]_{t-1} \\ &\quad + c_1 + \left(\gamma_{1,1} + g_{1,2}\right) \Delta p_t - \left(g_{1,2} - \gamma_{1,2}\right) \Delta pi_t - g_{1,2} \Delta t3_t + v_{wt} \\ (12)\quad \Delta p_t &= -\delta_2 \left[(1-\eta)(p - pi) - \eta (w - p - pr + t1) - t3 \right]_{t-1} + \zeta_2 gap_{t-1} \\ &\quad + \eta (c_2 + \delta_2 m_2) + g_2 \Delta (w - pr)_t + (1-\eta) \Delta pi_t + \Delta t3_t + v_{pt} \end{aligned}$$

where $a_1 = \alpha_1/\eta$, $g_{1,2} = \gamma_{1,2}/\eta$, $g_2 = \gamma_2 \eta$, $v_{wt} = \epsilon_{1t}$ and $v_{pt} = \eta \epsilon_{2t}$.

The first equation, (11), is a wage equation in the equilibrium-correction form introduced by Sargan (1964). We have derived it with reference to theories of wage bargaining. The equation also embodies earlier "conflict" theories of inflation.[6] Lastly, (11) represents a generalization of conventional Phillips-curves, which entail that the only wage equilibrating mechanism is the rate of unemployment, where there is a unique natural rate of unemployment which only depends on the parameters of the wage and price equations. The second equation, (12), is a cost-mark-up price equation, where the mark-up is a positive function of output fluctuations.

The main relevance of the model (11) and (12) is that it implies that correlations between output and inflation, and between wages and unemployment, are

[6] See Nymoen (1990, Chapter 1).

non-structural and will depend on where shocks arise; and on the propagation mechanisms embodied in the model dynamics, see Frisch (1933).

As a reminder of the importance of the origin of shocks, consider first the situation where markets are imperfectly competitive, prices exceed marginal costs, so a positive *demand shock* might be expected to induce increased supply even at the initial price level. Hence, productivity will pick up and this may induce an initial tendency for prices to be countercyclical. However, if the change in output then induces an increase in employment, unemployment will start to fall and wages will pick up, eventually prices will also rise and inflation becomes procyclical. Hence, in this scenario, both nominal and real wages are procyclical, but out of phase. However, these responses can be reversed if there is a strong effect of excess demand in price formation, e.g. if the demand shock occurs in a "heated" situation where marginal costs are rising very sharply. Excess demand will push up prices and wages, we will observe a *positive*, procyclical, correlation between inflation and output. Depending on the wage pressure stemming from reduced unemployment, real wages can go either way.

Suppose now that we see a *supply side shock* which increases wage costs—c_1 goes up. Real wages rise, so employment and output fall. Prices will then catch up to restore profit margins and eventually the system will revert either to the starting point, or to a new natural rate depending on the permanence of the shock—assuming a non-inflationary policy response. We will have observed a *negative*, countercyclical, correlation between inflation and output.

A first point to make about propagation mechanisms, is that an inherent property of dynamic systems is that the short-run response of an endogenous variable to a shock is different from the long-run response, determined by the steady-state solution of the system. To find the steady state in our case, we follow Kolsrud and Nymoen (1996) and solve for the *producer* real wage $w_{p,t} = w_t - pp_t - pr_t$ and the real-exchange rate $pi_{p,t} = pi_t - pp_t$. Kolsrud and Nymoen (1996) show that the producer real wage $w_{p,t}$ is negatively related to unemployment in the medium run. However, if $0 \leq \alpha_1 < 1$, the steady-state producer real wage is shown to be independent of the level of unemployment. In other words, the predicted short-run responsiveness of producer real wages is larger than long-run responsiveness. The same is true for inflation: a negative short-run effect from a rise in unemployment follows directly from mark-up pricing. However, in steady state we have by definition $\Delta pi_p = \Delta pi - \Delta pp = 0$ and thus $\Delta p = \Delta pi$, saying that inflation is unresponsive to unemployment and the output-gap in the long-run, despite being positively related to domestic demand pressure in the short-and medium run. As a final example, anticipating the analysis of the empirical models below, consider the steady-state solution for *consumer* real wages[7]

[7] i.e. for fixed growth rates of productivity (Δpr) and import prices (Δpi), fixed tax-rates t_1 and t_3, constant rate of unemployment and $gap = 0$.

$$
\begin{aligned}
(w-p-pr) \;&=\; \left(\frac{\dfrac{c_1}{\delta_1}+\alpha_1\left(\dfrac{c_2}{\delta_2}+m_2\right)}{1-\alpha_1}\right)-\frac{1}{\left(1-\alpha_1\right)\delta_1}\Delta pr \\
&\quad -\left(\frac{\dfrac{1-\gamma_{1,1}-\gamma_{1,2}}{\delta_1}+\dfrac{\left(1-\gamma_2\right)\alpha_1}{\delta_2}}{1-\alpha_1}\right)\Delta pi + \qquad (13)\\
&\quad -\frac{\zeta_1}{1-\alpha_1}u+\frac{\alpha_1-\beta_1}{1-\alpha_1}t1.
\end{aligned}
$$

The degree of long-run real wage responsiveness to unemployment is

$$\frac{\zeta_1}{1-\alpha_1}.$$

Depending on the structural parameters this magnitude can be bigger or smaller than the degree of short-run responsiveness $\delta_1\zeta_1$. The latter case is referred to as "wage hysteresis", following Nickell (1987), and is seen to be encompassed by our model of wage-price dynamics.

Within this framework we can conclude that there is no unique correlation between output fluctuations and price inflation or between unemployment and real wages. Both the source of shocks and the persistence of the shocks matters. This result is hardly new (see for example, Andrews, Bell, Fisher, Wallis, and Whitley (1985)) and could be stated for any (pair wise) correlation between price and quantity.

In the following we investigate the idea that simple and aggregate systems of equations can encompass any "fact" based on filtering and cross-correlations. In order to demonstrate this in practice, the two next sections present estimation and simulation results based on Norwegian and United Kingdom data.

4 Econometric wage and price models for the United Kingdom and Norway

In this section we summarize the results from applying the macro-econometric approach set out in Section 2.2 to the modelling of open-economy wages and prices. In a separate paper—Bårdsen, Fisher, and Nymoen (1994)—we analyse the wage-price process in Norway and the United Kingdom. We rely heavily upon the econometric results from that paper to illustrate our points.

4.1 The United Kingdom

In the data set for the United Kingdom the wage variable, w, is average actual earnings. The price variable, p, is the retail price index, excluding mortgage interest payments and the Community Charge. The non-modelled variables consist of employers' taxes, $t1$, indirect taxes, $t3$, mainland productivity, pr, import prices, pi, the unemployment rate, u, and a measure of the output gap, gap, approximated by mainland *GDP* cycles estimated by the Hodrick-Prescott filter. Finally two dummies are included to take account of income policy events—see appendix A for precise details.

The estimated equilibrium of the model—by means of the Johansen procedure—is given as:

$$
\begin{aligned}
(14) \qquad i) \quad w &= p + pr - t1 - 0.065u + constant \\
ii) \quad p &= 0.89(w + t1 - pr) + 0.11pi + 0.6t3 + constant.
\end{aligned}
$$

Wages are homogenous in consumer prices. The elasticity of productivity is also unity. According to (14), all changes in employers taxes ($t1$) are reflected in wages in the long-run. The estimated wage moderating effect of unemployment seems reasonable, compared to what others have found on U.K. manufacturing data. For example Rowlatt (1987, equation (20)) implies a long-run elasticity of -0.08, while in Nickell (1987, equation (22)) the estimate is -0.104 (i.e. for a fixed long-term unemployment proportion). The appearance of the rate of unemployment, u_t, in (14 i) is consistent with $u_t \sim \mathsf{I}(1)$ and cointegrating with wages, prices and productivity. On the other hand, if w_t, p_t and pr_t cointegrate, u_t can nevertheless be in the levels part of the wage-model, as a separate $\mathsf{I}(0)$ variable, which is also consistent with (14 i).

According to equation (14 ii), prices are partly adjusted for changes in the indirect tax-rate. Furthermore, prices are set as a weighted average of mark-up over labour costs and import prices. Compared to the corresponding Norwegian estimates in (15) below, it is interesting to note how the size—and therefore openness—of the economies is reflected in this weighted average.

The estimated dynamic model for the United Kingdom is given in Table 1,while Figure 4 shows one-step residuals and residual standard errors—i.e. $\left\{y_t - \hat{\beta}_t x_t\right\}$ and $\{\pm 2\hat{\sigma}_t\}$ in a standard notation—for each equation. The third panel records Sargan's (1964) test of overidentifying restrictions against its 5% critical value, for every sample size {1976(3)–1984(3),1976(3)–1984(4), ...1976(3)–1993(1). Following Hendry and Mizon (1993) the interpretation of an insignificant Sargan-statistic is that the model in Table 1 encompasses the underlying *VAR*-system. Finally, the fourth panel shows the "break point" Chow-statistic for the sequence {1984(3)–1993(1), 1984(4)–1993(1),...1992(4)–1993(1), 1993(1)} for the model in Table 1.

The model in Table 1 is estimated by *FIML*. The model equilibrium is incorporated into the dynamic model as equilibrium-correction terms to represent the adjustment to disequilibria and their importance is clearly shown. In addition to the equilibrium-correction term, wages are driven by growth in consumer prices over the last two periods and by productivity gains. With a elasticity estimate of 0.66 and a standard error of 0.039, short-run homogeneity is clearly rejected. The negatively estimated coefficient for the change in the indirect tax-rate ($\Delta t3_t$) is surprising at first sight, but inspection of (11) shows that the result is in accordance with theory: Consumer prices respond when the tax-rate is increased, cf. the price equation (12), which in turn is passed on to wages. Hence, the net effect of a discretionary change in the indirect tax-rate on wages is estimated to be effectively zero in the short-run and positive in the intermediate and long-run. The effect of an increase in the payroll-tax rate is to reduce earnings, both in the short and long-run.

Table 1: The model for the United Kingdom.

The wage equation

$$\widehat{\Delta w_t} = \underset{(0.075)}{0.187}\,\Delta w_{t-1} + \underset{(0.039)}{0.332}\,(\Delta_2 p_t + \Delta pr_t) - \underset{(0.100)}{0.341}\,\Delta^2 t1_t$$
$$- \underset{(0.064)}{0.162}\,\Delta_2 t3_t - \underset{(0.023)}{0.156}\,(w_{t-2} - p_{t-2} - pr_{t-1} + t1_{t-2} + 0.065 u_{t-1})$$
$$+ \underset{(0.071)}{0.494} + \underset{(0.003)}{0.013}\,BONUS_t + \underset{(0.001)}{0.003}\,IP4_t$$
$$\hat{\sigma} = 0.45\%$$

The price equation

$$\widehat{\Delta p_t} = \underset{(0.149)}{0.963}\,\Delta w_t - \underset{(0.118)}{0.395}\,\Delta pr_t + \underset{(0.059)}{0.153}\,\Delta\,(p + pr)_{t-1} - \underset{(0.019)}{0.044}\,\Delta u_{t-1} + \underset{(0.092)}{0.536}\,\Delta t3_t$$
$$- \underset{(0.047)}{0.480}\left[p_{t-1} - 0.89\,(w + t1 - pr)_{t-2} - 0.11 pi_{t-2} - 0.6 t3_{t-1}\right]$$
$$+ \underset{(0.099)}{0.238}\,gap_{t-1} - \underset{(0.131)}{1.330} - \underset{(0.005)}{0.019}\,BONUS_t - \underset{(0.001)}{0.005}\,IP4_t$$
$$\hat{\sigma} = 0.71\%$$

Diagnostic tests

Overidentification $\chi^2(16)$	=	24.38[0.08]
$AR\ 1-5\ F(20,94)$	=	0.97[0.50]
Normality $\chi^2(4)$	=	3.50[0.48]
Heteroscedasticity $F(84,81)$	=	0.63[0.98]

Notes

The sample is 1976(3) to 1993(1), 67 observations.
Estimation is by FIML. Standard errors are in parantheses below the estimates.
The symbol $\hat{\sigma}$ denotes the estimated percentage residual standard error.
The p-values of the diagnostic tests are in brackets.

According to the second equation in Table 1, prices respond sharply (by 0.96%) to a one percentage change in wage costs. Hence short-run homogeneity is likely to hold for prices. In addition to wage increases and equilibrium-correction behaviour, price inflation is seen to depend on the output gap, as captured by the variable *gap*.

Finally, note that the two income policy-dummies, $BONUS$ and $IP4$ are significant in both equations, albeit with different signs. Their impact in the first equation is evidence of wage-raising incomes policies, and their reversed signs in the price equation indicate that these effects were not completely anticipated by price-setters.

The diagnostics reported at the bottom of Table 1 give evidence of a well determined model. In fact, Figure 4 shows that the model encompasses the underlying system for almost every sample size. Hence the model constitutes a valid parsimonious model of the underlying reduced form system. The diagnostics are vector tests and hence multivariate tests on the residuals of the model—see Doornik and Hendry (1994) for details.

4.2 Norway

In the Norwegian data set, the wage variable, w, is average hourly wages in the mainland economy, i.e. excluding the North-Sea oil producing sector and international shipping. The productivity variable, pr, is defined accordingly. The price

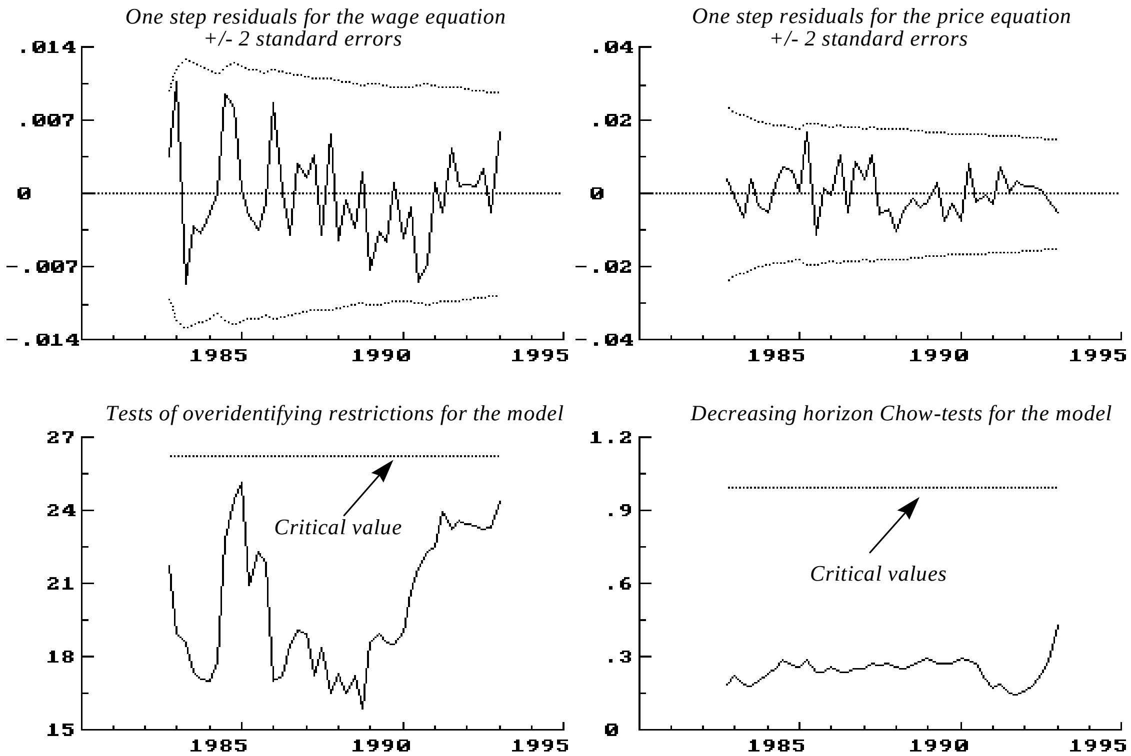

Figure 4: Recursive stability tests for the United Kingdom model. 5% critical values for tests of overidentification and for Chow-tests.

index, p is measured by the official consumer price index. pi again denotes import prices. The unemployment variable, u, is defined as a "total" unemployment rate, i.e. including labour market programmes.

The other non-modelled variables are normal working time, h_t, and the indirect tax-rate, $t3$. The change in the length of the working day, Δh_t captures wage compensation for reductions in the length of the working day—see Nymoen (1989b). Incomes policies and direct price controls, have been in operation on several occasions during the sample period. Two intervention variables, WD and PW, and one impulse dummy, $d80q2$, are used to capture the impact of these policies. Finally, $d70q1$, is a VAT dummy. We use seasonally unadjusted quarterly data, so the estimated models includes three seasonal dummies (S_i, $i = 1, 2, 3$).

Cointegration analysis yielded the two following estimated long-run relationships :

$$(15) \qquad \begin{aligned} i) \quad w &= p + pr - 0.08u + constant \\ ii) \quad p &= t3 + 0.65(w + t1 - pr) + 0.35pi + constant \end{aligned}$$

which are remarkably close to the corresponding results for the United Kingdom. The import price index is attributed a larger weight in the Norwegian price equation, which is reasonable. It is interesting that the estimated unemployment elasticity is only a marginally larger (in absolute value) than in the United Kingdom model.

Table 2: The estimated model for Norway.

The wage equation

$$\widehat{\Delta w_t} = \Delta p_t - \underset{(0.027)}{0.104}\,\Delta_2 (w-p-pr)_{t-1} - \underset{(0.035)}{0.104}\,\Delta pi_t - \underset{(0.190)}{0.728}\,\Delta (t1_t + t1_{t-2})$$
$$- \underset{(0.085)}{0.496}\,\Delta t3_{t-2} - \underset{(0.102)}{0.294}\,\Delta h_t - \underset{0.0193)}{0.131}\left[(w-p-pr)_{t-2} + 0.08 u_{t-1}\right]$$
$$- \underset{(0.016)}{0.106} + \underset{(0.009)}{0.027}\,D80q2_t - \underset{(0.004)}{0.021}\,WD_t - \underset{(0.010)}{0.046}\,D70q1_t + \underset{(0.003)}{0.021}\,PD_{1t}$$
$$- \underset{(0.005)}{0.022}\,S_{1t} + \underset{(0.008)}{0.041}\,S_{2t} + \underset{(0.005)}{0.009}\,S_{3t}$$
$$\hat{\sigma} = 0.92\%$$

The price equation

$$\widehat{\Delta p_t} = \underset{(0.015)}{0.115}\,\Delta\left[w_t + (w - 0.5pr)_{t-1}\right]$$
$$- \underset{(0.008)}{0.090}\,[p_{t-3} - 0.65\,(w_{t-3} + t1_{t-3} - pr_{t-2}) - 0.35 pi_{t-1} - t3_{t-2}]$$
$$+ \underset{(0.016)}{0.066}\,gap_{t-1} + \underset{(0.001)}{0.013} + \underset{(0.004)}{0.050}\,D70q1_t - \underset{(0.002)}{0.013}\,PD_{1t}$$
$$- \underset{(0.002)}{0.001}\,S_{1t} - \underset{(0.002)}{0.004}\,S_{2t} - \underset{(0.002)}{0.005}\,S_{3t}$$
$$\hat{\sigma} = 0.39\%$$

Diagnostic tests

Overidentification $\chi^2(41)$	=	36.31[0.68]
AR $1-5$ $F(20, 162)$	=	0.97[0.51]
Normality $\chi^2(4)$	=	3.78[0.44]
Heteroscedasticity $F(168, 99)$	=	0.84[0.85]

Notes

The sample is 1967(4) to 1993(2), 103 observations.
Estimation is by FIML. Standard errors are in parantheses below the estimates.
The symbol $\hat{\sigma}$ denotes the estimated percentage residual standard error.
The p-values of the diagnostic tests are in brackets.

This finding contradicts the widely held view that Nordic wages are extremely responsive to changes in unemployment, see e.g. the reported elasticities in Layard, Nickell, and Jackman (1991, Ch. 9, Table 2). Our estimates correspond much better with Nymoen (1990, Table 4.1), who estimates a long-run elasticity of -0.13 on a similar aggregate data-set. Johansen (1995) finds the same long-run wage responsiveness to unemployment as we do, albeit for annual observations of manufacturing wages. Earlier models of manufacturing wages have obtained somewhat stronger wage responsiveness, e.g. in Nymoen (1989a, equation (16)) the estimated elasticity is -0.21 and Calmfors and Nymoen (1990, Table 7) find -0.17 on annual data.

The estimated dynamic model is shown in Table 2. There is considerable within sample constancy, as Figure 5 shows. This suggest invariance to such "shocks" as the credit market liberalization that occurred in the early 1980s, the 50% reduction in oil-prices in 1986 and the changes in exchange-rate regime in the late 1980s.

The *wage equation* in Table 2 imposes short-run homogeneity with respect to consumer prices. At first glance this suggests real wage rigidity. However, closer inspection of the equation shows that this is not the case in general: The wage equation includes an indirect tax-rate with a negative sign, a negative coefficient of the *VAT* dummy ($D70q1_t$) and (ceteris paribus) positive effects of price controls

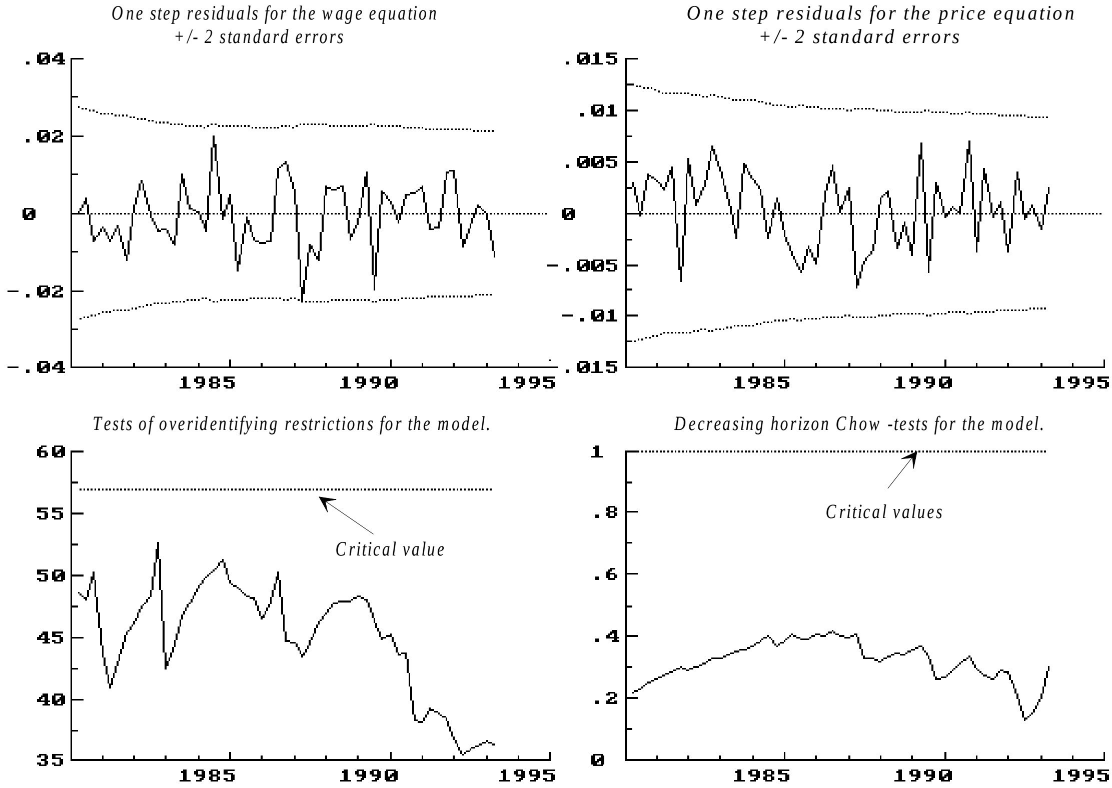

Figure 5: Recursive stability tests for the Norwegian model.

(PD_t). Hence discretionary policies have clearly succeeded in affecting consumer-real wage growth over the sample period. However, without such policies, Norwegian aggregate wages *do* seem to react very quickly to "normal" or expected consumer price increases. Import price growth is likely to be the most important "unexpected" part of price inflation, so given the unit coefficient on Δp_t, it is not surprising that, Δpi_t, is attributed a negative estimated coefficient. The equilibrium-correction term is highly significant, with a "t-value" of -6.8. Finally, the change in normal working-time, Δh_t, enters as expected in the wage equation, i.e. with a negative coefficient.

>From the second part of Table 2 we find that, in addition to equilibrium correction and the dummies representing incomes policy, price inflation is significantly influenced by wage growth and excess demand. The overall speed of adjustment seems to be slower for the Norwegian consumer prices than for U.K. retail prices.

5 Simulating the "stylized facts"

The estimated wage-price models described in the previous section are stable, data coherent models estimated by *FIML*. The models can be thought of as the structural forms of restricted open *VAR* systems. The non-modelled variables are unemployment, productivity, import prices and tax-rates on employment and consumer prices. In this section the relevance of the estimated structure (propagation) and the importance of the non-modelled variables (impulses) is illustrated with the aid

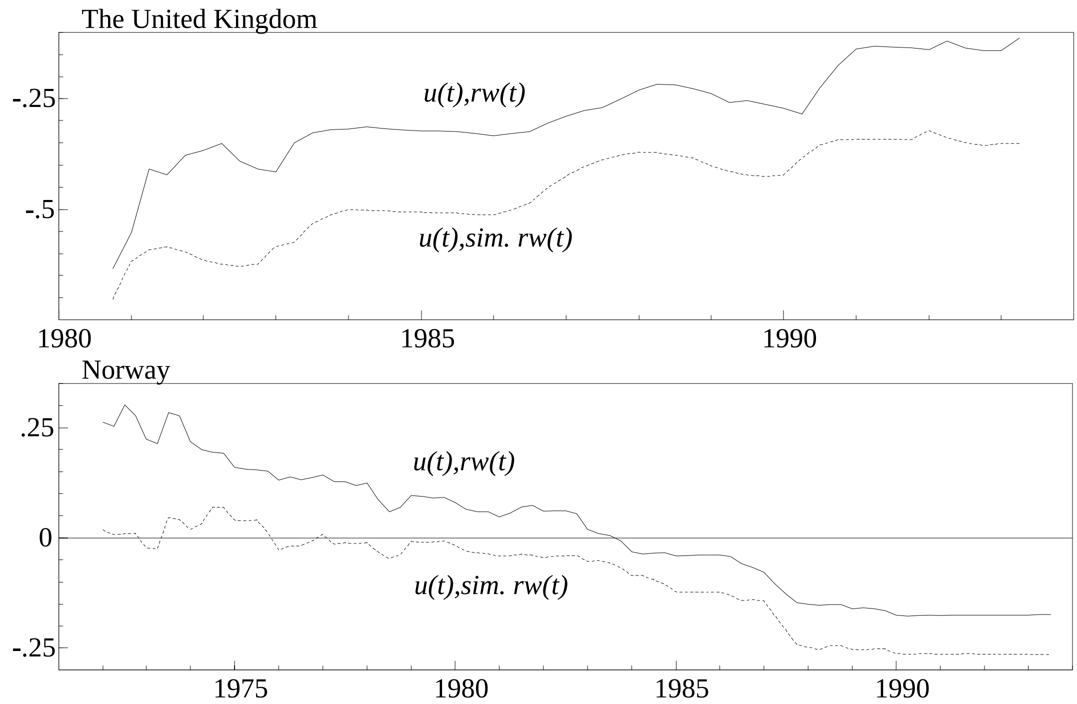

Figure 6: Recursive correlation coefficients between unemployment and real wage, and between unemployment and simulated real wage, using a starting period of 12 quarters.

of simulation exercises. First, Section 5.1 demonstrates the models' ability to match both sample paths and moments. Against that background, Section 5.2 shows how different paths of the non-modelled variables can generate different correlations in the data.

5.1 Matching trends and moments

An econometric model of I(1) variables, which is data coherent, must reproduce the trends in the modelled variables: nominal prices and wages. Given the homogeneity restrictions, the models must also track the trend in the implied real-wage, $w_t - p_t$. Hence, dynamic simulation of the two models produces a "good fit" for the real-wage in the two countries, reflecting the numerical, as well as statistical, significance of the non-modelled variables in Tables 1 and 2.

It is less obvious that econometric models will necessarily reproduce those features of the data that were not taken into account during the modelling process. Hence we attempt to reproduce the non-constancy of the bivariate correlations using the dynamic simulation results for real-wages. This is demonstrated in Figure 6 which shows how the models reproduce the recursive correlation coefficients. The instability of the bivariate correlations are just as evident with the simulated data from the two constant parameter models, thus illustrating the non-structural nature of "stylized facts".

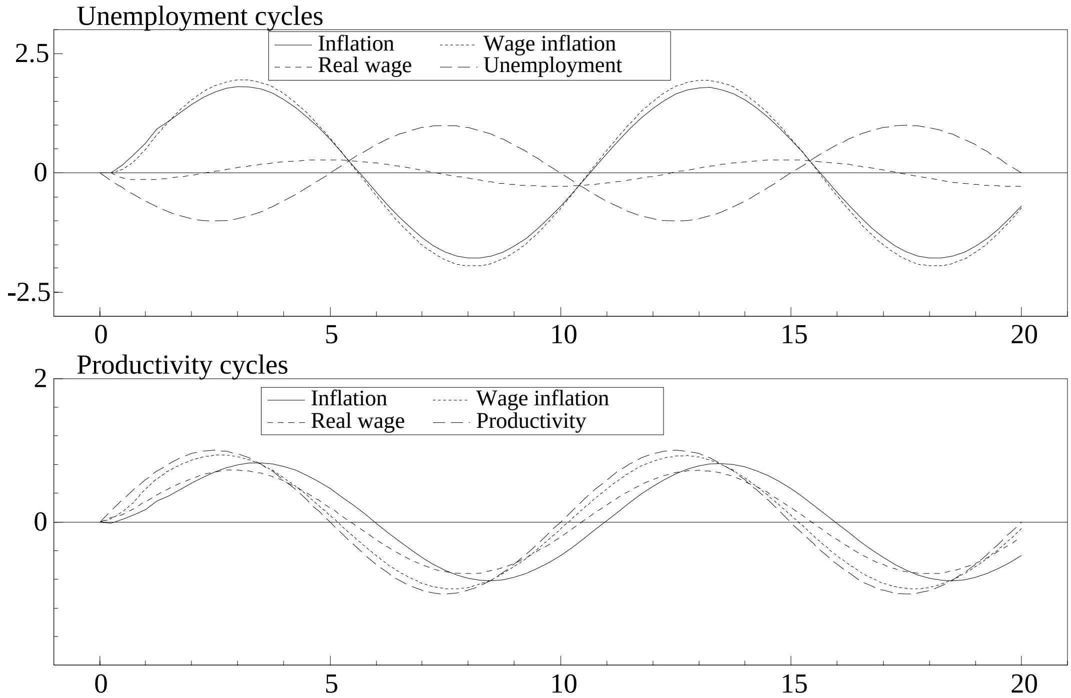

Figure 7: Unemployment and productivity 10 year cycles in the United Kingdom model.

5.2 Matching correlations with the cycle

The high contribution from the non-modelled variables to the close match between the predicted and observed correlations, makes it interesting to examine the partial dynamic response of earnings, prices and the real wage to shocks in these "exogenous" variables. The non-modelled variables are by no means independent of each other and this makes the partial responses something of a thought experiment rather than a realistic counterfactual simulation. Nevertheless the results prove to be a useful diagnostic in understanding why certain cross-correlations are observed.

We have undertaken a variety of simulations and report just two. Both can be thought of as resulting from a cycle in output taking the form of a sine wave with a 1% amplitude, with the period of the cycle (frequency) to be varied:

1. A cosine wave in the log of the unemployment rate with amplitude of 10% —hence a 1% amplitude in the level of the rate around a base value of 10% unemployment—for the United Kingdom and with an amplitude of 5% for Norway. This implicitly assumes that the output cycle is totally reflected in employment with no change in recorded productivity.
2. A sine wave in recorded productivity with an amplitude of 1%. This implicitly assumes that employment and hence unemployment, remain constant.

Ideally we would use a closed macroeconomic model to isolate the independent sources of economic fluctuations—changes in policy variables, external events, demographic and technology trends etc. However, this would serve to cloud the issue and

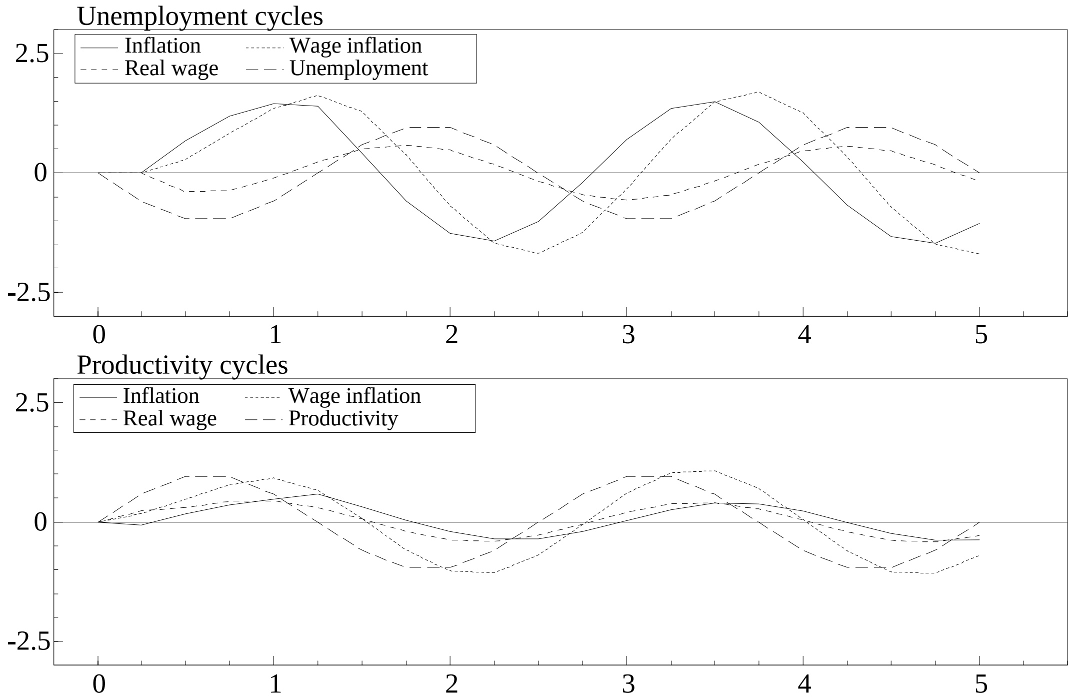

Figure 8: Unemployment and productivity 2.5 year cycles in the United Kingdom model.

we concentrate on the transmission of real shocks to prices, which covers some of the major debating points when examining business cycle correlations. To complete this assumption and to preserve homogeneity in nominal variables we assume that import prices remain fixed to their foreign currency values (p^f) and that the exchange rate (ex) adjusts instantaneously to changes in domestic prices. This allows us to substitute out the import price for domestic prices plus a real exchange rate term (rex). Effectively, we are introducing a third equation ($pi_t = p_t^f - ex_t \equiv p_t + rex_t$).

The wavelength of the shock is variable. The current cycle in most western countries will have a period not far short of ten years and so this is our base case. But some cycles are shorter and so we have experimented with the case of four cycles in ten years as well.

5.2.1 The United Kingdom

Taking the United Kingdom 10 year cycle in unemployment first, unemployment is negatively correlated with price and wage inflation. The real wage is relatively constant and is out of phase (see Figure 7). When the input cycle is productivity then both wage and price inflation are procyclical but weaker, whilst the real wage response is procyclical and significantly larger.

Hence the correlation of cycles in activity with cycles in the real wage depends entirely on whether fluctuations in output lead to fluctuations in employment or measured productivity.

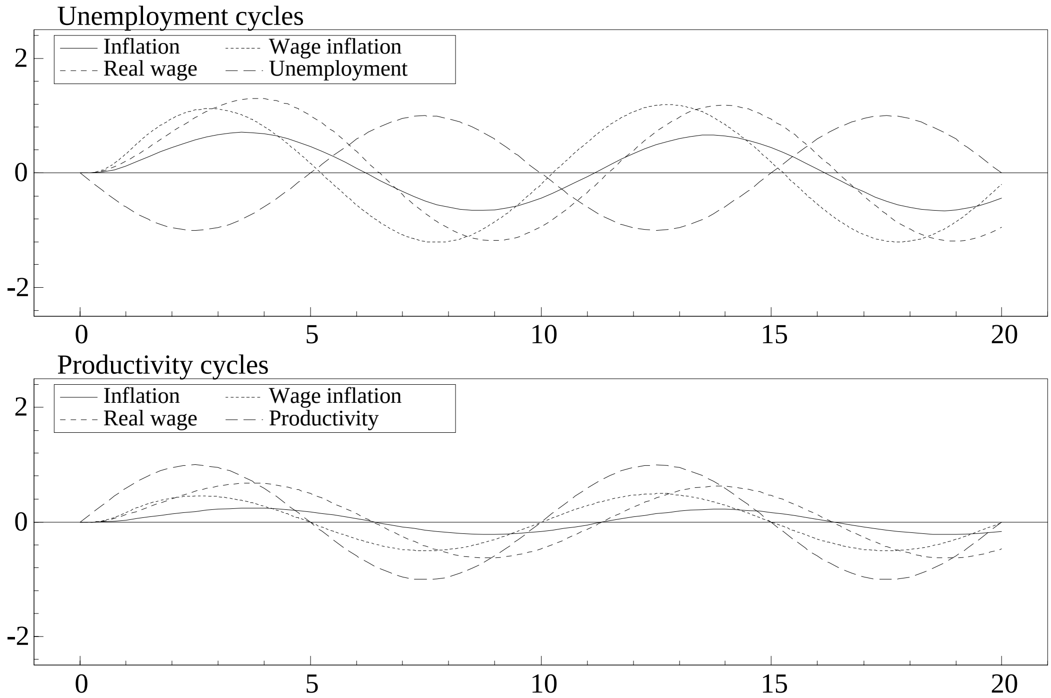

Figure 9: Unemployment and productivity 10 year cycles in the Norwegian model.

The results of a shorter input cycle are depicted in Figure 8. Taking the unemployment cycle first: price and wage inflation are now so much out of phase that they appear to lead unemployment. Real wages are still countercyclical, but out of phase compared to the pattern for the long input-cycle in Figure 7. As regards the productivity cycle the real wage response is now practically zero.

These results are sufficient to demonstrate that the correlations of endogenous variables are sensitive to the nature of the shock—including the periodicity. Further experiments with other shocks—such as the real exchange rate or an autonomous shock to wages demonstrate that a wider range of correlations is possible. We also note in passing that the results depend slightly on whether the input cycle is started up from cold or the continuation of an existing cycle.

5.2.2 Norway

Having demonstrated the sensitivity of partial correlations within one country, the main interest in examining a second country is the fact that we have similar structures yielding different results. The inflation responses in Norway are smaller—despite very similar long-run elasticities (by constraining the real exchange rate to be constant)[8] and a large *real* wage response. The dynamics of the response are also much more lagged.

[8]The smaller shock compensates for the non-linearity in the association of output with the unemployment rate. But this does not explain the totality of the smaller inflation response.

Comparing Figures 7 and 9 we find that inflation response is more vigorous for the United Kingdom, while real wages response is stronger and procyclical in Norway. Note that this emerges despite the fact that the estimated long-run wage response is estimated to be quite similar in both systems. This highlights the importance of considering *both* wage and price formation when predicting real wage behaviour over the cycle. The more muted responsiveness in U.K. real wages is explained by our estimation results. In terms of the structural parameters in equations (11) and (12), the estimation results tell us that price equilibrium correction, δ_2, and the simultaneous pass-through of wages on prices, g_2, are much higher in the United Kingdom than in Norway. Conversely, expected inflation has a bigger impact, $(\gamma_{1,1} + g_{1,2})$, on Norwegian wages than on UK wages. As a result, U.K. nominal wage-claims are almost instantaneously cancelled by price adjustment. Against the background of a more open economy and of the institutional set-up in Norway, e.g. more centralization and coordination, this explanation has some credibility.

In summary, these diagnostic simulations confirm that, even when underlying behaviour is invariant, we can expect to observe quite different correlations of macroeconomic variables over the business cycle depending on the nature and source of the shocks which disturb the system.

6 Discussion and summary

A set of "stylized facts" in the form of univariate and bivariate moments summarizes salient properties of marginal and bivariate distributions that are of interest to business-cycle analysts. A primary use of the correlations is the calibration of *RBC* models. They represent key "facts" that the theoretical model of interest should be able to account for.

Calibration give rise to a range of methodological and technical issues, e.g. the absence of measures of sampling uncertainty and of statistical evaluation criteria, and on several fronts progress appears to be made, see e.g. Canoe, Finn, and Pagan (1994), Eichenbaum (1995) and Wickens (1995) for discussion. So far, less attention has been paid to the validity of the assumption that the correlations are stable over time. Therefore, a new kind of problem is posed by the empirical instability of the correlations, which has now been demonstrated for several of the correlations used for the evaluation of *RBC* models, see e.g. the results for price/wage-output correlations in Blackburn and Ravn (1992), Englund et al. (1992) and Andersen (1994, p. 19-18) , and our own results for Norway and the U.K. in Section 2 above. An immediate concern is that the reliance on the correlations per se in *RBC* analysis is apt to generate fragile results. Indeed, as pointed out by Hendry (1995), the current practice of matching subsets of moments that are inherently non-constant induces a sample-dependence problem for the very models that originated from a desire to avoid exactly this difficulty.

We have concentrated on the correlations (r_{yx}) between real wages and the unemployment-cycle and have demonstrated that econometric modelling can determine (partial) structure and resolve the puzzle of unstable r_{yx}'s. A simple theory model of wages and prices that allows imperfect competition elucidates why there is no reason to expect a consistent pattern of bivariate correlations. Each cycle is likely to reflect different shocks to the economic system—both the sign and magnitude of the correlations will depend on the implicit shocks. The text-book suggestions of procyclical price variations originate from the analysis of demand-led business cycles. Some of the recent cycles in many countries appear to have had major contributions from external or supply side shocks. It is therefore hardly surprising that observations from different samples yield different bivariate correlations.

The real issue of interest is to identify which shocks cause which correlations and how these combine to explain the historical record. There is more than one approach possible to resolve this question. We have chosen the route of specifying and estimating structural econometric models, based on the encompassing of *VAR*-systems. We have focused on the response of nominal variables to real shocks, hence we have constructed partial wage and price systems for the United Kingdom and Norway.

Our econometric estimates yield stable, data coherent models which we show are able to explain the unstable observed correlations in the data sample. For several of the key parameters, the estimates for the United Kingdom and Norway are surprisingly similar in appearance. Simulation results from the estimated models reveal that the observed correlations between the cycle, inflation and the real wage will depend entirely on the nature of the real shock—we concentrate on whether the output shock affects productivity or unemployment. The correlation also depends on the frequency of the shock and varies across countries due to the different dynamic adjustment processes.

These results demonstrate that, even when the underlying behaviour is invariant, we cannot expect to observe consistent pair-wise correlations either in macroeconomic time series or in cross-country comparisons. It is nevertheless possible to capture such complicated movements in the data by the use of "systems-of - equation" models—ideally in a closed model of the macro economic system.

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Appendix

A Data definitions

Small letters denote the natural logarithms of the variables.

A.1 United Kingdom

All variables are seasonally adjusted except for the RPI. ONS is the Office for National Statistics (formerly the CSO).

W Wages: index of actual quarterly average earnings: ONS identifier DNAB.

P Retail Price Index: ONS identifier: CHAW.

PR Non-north sea productivity: ONS identifiers: CKJL/DYDC.

PI Price deflator for expenditure on imported goods and services: ONS identifiers: DJBC/DJDJ.

U Unemployment rate, registered number of unemployed: ONS identifier: BCJE.

gap Output gap defined as log of non-north sea GDP (NNO) deviations from trend, where the trend is estimated by the Hodrick-Prescott filter using $\lambda = 1600$. ONS identifier for NNO: CKJL.

$T1$ Employers tax rate: $(GTAY + AIIR + TSET)/(34.203 * DNAB * (BCAD + BCAH)/1000)$ which are ONS identifiers apart from $TSET$ which is receipts of selective employment tax. $GTAY$ and $AIIR$ give employers' National Insurance and other contributions and the last term gives the aggregate employed salary bill.

$T3$ Tax rate on the retail price index basket, excl. mortgage interest payments. This tax rate is derived from the Bank of England's spreadsheet for construction of RPIY (see Beaton and Fisher (1995)).

$BONUS$ Impulse dummy for retimed bonus payments brought forward one month before a General Election. Takes the value 1 in 1992:1 and -1 in 1992:2.

$IP4$ Incomes policy dummy. Zero after 1979:2: source Whitley (1986).

A.2 Norway

W Nominal mainland hourly wages. Source: Quarterly National Accounts (QNA).

P Consumer price index. 1991=1. Source: Norges Bank's database of economic time series.

PR Mainland economy value added per man hour at factor costs. Mill. 1991 NOK. Source: QNA.

PI Deflator of total imports. 1991=1. Source: QNA.

gap Output gap defined as mainland GDP deviations from trend, where the trend is estimated by the HP-filter using $\lambda = 1600$. Mill. 1991 NOK. Source: QNA.

$T1$ Employers tax rate. Source: RIMINI-database. RIMINI is the Central Bank of Norway's quarterly econometric model.

U Rate of unemployment. Registered unemployed plus persons on labour market programmes as a percentage of the "labour force". The labour force is calculated as employed wage earners plus unemployment. Source: QNA and RIMINI-database.

$T3$ Indirect tax rate. Source: RIMINI-database.

H Normal working hours per week.

$D80q2$ Dummy for lift of wage-freeze. 1 in 1980.2, zero otherwise.

WD Composite dummy for wage freeze: 1 in 1979.1, 1979.2,1988.2 and 1988.3.

$D70q1$ VAT dummy. 1 in 1970.1, zero otherwise.

PD Composite dummy for introduction and lift of direct price regulations. 1 in 1971.1, 1971.2,1976.4,1979.1. -1 in 1975.1,1980.1,1981.1,1982.1. Zero otherwise.