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Unemployment in a Small Open Economy: Finland and New Zealand

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Abstract

Unemployment is now the key issue for economic policy in the OECD and Europe in particular. By examining data from the period 1962–1996 for two highly different small open OECD economies, Finland and New Zealand, in a VEC model this paper seeks to cast light on three questions: the degree to which unemployment has been the result of slow adjustment to large external shocks; the degree to which differences in labour market structures can lead to different responses to shocks; the importance of the exchange rate and the external sector in resolving the problem. The approach uses a fairly general model of the labour market that includes wages, unemployment, the capital stock and the terms of trade. It uses cointegration analysis to establish long-run relationships among the four variables. In the case of Finland we find that the short-run response of unemployment to shocks (to the long-run relationship) is large relative to the response of real wage and the terms of trade. In New Zealand on the other hand both real wages and the terms of trade, in particular, adjust more rapidly. As a result the burden of short-run adjustment in the New Zealand economy appears to fall more heavily on (relative) prices. Since the unemployment rate in both countries displays hysteresis, these results suggest that relative price adjustment in the New Zealand economy is more effective in preventing adverse aggregate shocks from becoming adverse unemployment shock.

Keywords: unemployment, open economy, structural change, labour market

Työttömyys pienessä avotaloudessa: Suomi ja Uusi-Seelanti

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Tutkimusosasto

Tiivistelmä

Tutkimuksessa tarkastellaan kahden pienen avotalouden, Suomen ja Uuden-Seelannin, työttömyyteen vaikuttavia tekijöitä viimeisen yli 30 vuoden aikana. Tutkimuksella pyritään hakemaan lisävalaistusta kolmeen kysymykseen: missä määrin työttömyyden kasvu johtuu hitaasta sopeutumisesta talouteen kohdistuviin suuriin häiriöihin, missä määrin työmarkkinoiden rakenteelliset erot synnyttävät erilaisuutta työmarkkinoiden vasteessa häiriöihin ja lopuksi, mikä on valuuttakurssin ja avoimen sektorin merkitys työttömyysongelman ratkaisussa. Tutkimuksen teoreettiset tarkastelut perustuvat suhteellisen tavanomaiseen palkat, työttömyyden, pääomakannan ja vaihtosuhteen käsittävään avotalouden työmarkkinamalliin. Yhteisintegroituvuusanalyysin avulla pyritään saamaan selville, löytyykö maiden havaintoaineistoista tukea mallin muuttujien pitkän aikavälin riippuvuuksille. Tulosten perusteella Suomen kokonaistaloudellinen sopeutuminen poikkeaa Uuden-Seelannin talouden käyttäytymisestä sikäli, että Suomessa työttömyys reagoi lyhyellä aikavälillä suhteellisen voimakkaasti häiriöihin, jotka kohdistuvat työttömyyden, reaalipalkan, vaihtosuhteen ja pääomakannan väliseen pitkän aikavälin riippuvuuteen. Tasapainottomuudet ilmenevät siten Suomessa vahvasti juuri työttömyydessä. Uudessa-Seelannissa vaihtosuhteen ja vähäisemmässä määrin reaalipalkan muutokset ovat vastaavassa asemassa. Kummankin maan työttömyydessä ilmenee kuitenkin hystereesia eli voimakasta jälkivaikutusta; jos työttömyys yllättävästi muuttuu, ovat nämä muutokset suhteellisen pysyviä. Tulokset näyttävätkin näin ollen viittaavan siihen, että suhteellisten hintojen – kuten reaalipalkkojen ja vaihtosuhteen – muutokset vaimentavat tehokkaammin kokonaistaloudellisten häiriöiden kielteisiä työttömyysvaikutuksia Uudessa-Seelannissa.

Asiasanat: hidas sopeutuminen, avotalous, yhteisintegroituvuus, virheenkorjaus

Contents

1 Introduction

1.1 The problem to be addressed

While price stability is still a key focus of economic policy, unemployment has taken over in OECD as the major problem urgently seeking a solution. This paper seeks to make a contribution to the understanding of the causes of unemployment and its successful reduction. It explores the example of two rather different small open economies, Finland and New Zealand, as their contrasts shed light on three key issues:

– the degree to which unemployment has been the result of slow adjustment to large external shocks
– the degree to which differences in labour market structures can lead to different responses to shocks
– the importance of the exchange rate and the external sector in resolving the problem

The policy prescriptions for trying to handle unemployment are very different if the problem is largely that adjustment to shocks takes an unacceptably long time rather than if it is that structural change in the economy means that unemployment will continue at a higher level. In the first case the effort needs to go into improving the operation of the labour market itself, while in the second it is the structure affecting the demand and supply of labour that needs to change.

The two countries have experienced step and trend increases in unemployment. Both New Zealand and Finland have been subject to major shocks, which if not permanent are certainly enduring – the restriction of traditional agricultural markets in the case of New Zealand and the collapse of the former Soviet Union in the case of Finland. They have also become aware of increasing difficulties (non-sustainable trends) with underlying aspects of their institutional structures. Finland, to a lesser extent than Sweden, has found that there are limits to the 'Nordic model' of the extensive social benefit system and increasingly important role of the public sector. New Zealand found that the process of increasing regulation and insulation from consequences of external distortions was not sustainable.

The historical profiles of unemployment are fairly similar. Both New Zealand and Finland have seen unemployment rise markedly (from 4 % in the mid 1980s to 11 % in 1992 in New Zealand and from 5 % in the mid-1980s to 18 % in Finland by the end of 1993) and seen it fall away fairly rapidly since (to 6 % in New Zealand in 1995 and to 13 % in Finland in 1996). In both countries even the starting values for unemployment in the mid-1980s were thought to be unacceptably high compared with post-war history and urgent action was thought to be needed. The desired process of reduction in unemployment is by no means over. Nor are the responses to the shocks fully played out yet. Our evidence is thus partial but so it is in all the rest of the OECD countries where unemployment rose on this sort of scale.

There are, however, important differences between the two countries that make them a good illustration of the range of possible responses. In New Zealand the response has been wide-reaching, involving the rapid reduction of external barriers to trade, the liberalisation of the domestic economy and the introduction of substantial change into the operation of the labour market itself. The response in Finland has been much more muted, although smaller moves in all three directions have occurred particularly with membership of the EU. As our analysis shows the behaviour of the two labour markets has also been different. The role of unemployment, in particular, in the adjustment to aggregate shocks appears to be different in two economies; whereas unemployment in both economies displays extreme persistence in the sense of a unit root in the generating process – hysteresis – the burden of adjustment to aggregate shocks is more biased towards unemployment in Finland.

As in so many instances it is difficult to isolate these various effects. The nature of these adjustment processes to shocks has been heavily obscured by the business cycle, as the period of observation is short. Both countries have seen a major economic cycle and a period of fiscal reform, which has exacerbated unemployment in the short run but contributed to its reduction in the longer term. However, while New Zealand has returned to a budget surplus, a sustainable long run fiscal position and has debt to GDP ratio of 30 % and falling, Finland is inside the Maastricht criteria of a 3 % deficit to GDP ratio and a 60% debt to GDP ratio but has longer-run problems still to resolve.

Nevertheless, one feature stands out in both of the adjustment processes, namely the role of the real exchange rate in having a rapid favourable impact on competitiveness (without much inflation) in the turnround and up phases of the cycle and a major impact on bringing down inflation (without a decline in activity in New Zealand in the opposite phase, until the impact of the Asian crisis came through). The UK and Sweden have had similar experiences following their 1992 devaluations. Finland will enter Stage 3 of EMU in the first wave in 1999 and the apparent stability of the nominal exchange rate with respect to the other euro countries already anticipated Finland’s entry prior to the actual decision in the beginning of May 1998. Hence while New Zealand’s nominal exchange rate has already moved downwards substantially as the economy moves towards the recovery phase of the cycle when, it is hoped, unemployment will be able to resume its downward momentum, Finland will no longer have this mechanism in euro markets.

1.2 Our approach

We can only hope to shed light on a process which is incomplete and we do so largely by disentangling key features of the data relating to unemployment in the two countries and its evolution over time. In the next section we develop a very simple model of the characteristics of the labour market that builds on the existing literature. This model enables us to distinguish progressive influences and permanent and temporary shocks to the labour market but its principal focus is on the dynamics of the interactions of the labour market so that we can establish the extent of the ‘persistence’ of unemployment in response to those shocks. Section 3 then explains the statistical approach based on cointegration analysis, which is used to estimate the model and explores the outcomes for Finland and New Zealand in turn. The threads of the two examples are drawn together in Section 4, which sets out our conclusions for the three key issues of persistence, differences in structure and the role of external competitiveness.

2 The model and its context

Unemployment has, over the years, gained in importance among the set of socio-economic problems faced in the OECD, particularly in Europe, and is currently high on agenda of the candidate countries for the third stage of Economic and Monetary Union. In one sense, unemployment has become a ‘problem’ in the region as a whole only since the mid-1970s or early 1980s. Although the unemployment rate also increased in Europe prior to 1975, the increase was relatively modest (from around 3 % in 1960 to less than 4 % in 1973–1974). However in both Finland and New Zealand it was considerably lower than the average before the mid 1970s, so figures as ‘low’ as 3 % were regarded as being unacceptably high.

With the two oil-price shocks in the 1970's the picture has changed. The second crisis, in particular, marked a difference between the behaviour of the European and other OECD economies. Whereas the OECD average unemployment rate peaked at approximately 8 % in 1982–1983, it continued to rise in Europe and exceeded 10 % in 1986–1987 before starting to fall. In the 1990's unemployment rate has once again risen in the OECD area in general and in Europe in particular, where it was around 10.5 % in mid-1998 compared with the OECD average of 7.2 %. Hence, the unemployment gap between OECD-Europe and the rest of the OECD has not come down since it emerged in 1982–1983, but has actually risen in recent years. The time series pattern of the European unemployment rate then clearly suggests that there is something like ‘the European Unemployment Problem’ notion used by Ljunqvist and Sargent (1998) to head their recent contribution. New Zealand has followed a middle course, neither showing the stability of employment at lower levels in the US and Japan nor the more persistent higher levels of Europe.

Various theories have been proposed to explain the rise in European unemployment. Blanchard and Summers (1986) and Lindbeck and Snower (1988) impute the outcome to insider-outsider conflicts between employed and unemployed workers that arose in the highly unionized economies in Europe. Bentolila and Bertola (1990) study the idea that excessive European hiring and firing costs contributed to higher unemployment. Blanchard (1998) emphasizes capital accumulation and factor prices when he revisits the European unemployment problem. There has been wage moderation in Europe since the mid-1980s but it has not led to a decrease in unemployment, because there has been an adverse shift in labour demand by firms. Blanchard argues that either a shift in the distribution of rents away from workers or a technological bias against labour explains this shift. Malinvaud (1994) focuses on the effects of capital shortage, which, according to his argument, was caused in Europe by high real wages in the 1970's and high real interest rates in 1980's. This hypothesis of a capital shortage can, quite naturally, also be the outcome of various other factors. The underlying idea is simply that some exogenous factor(s) reduce the profitability of production, e.g. by inducing a persistent fall in (anticipated future) output prices, which, in turn, tend to reduce incentives to invest and the need for capital. It is prima facie likely that New Zealand and Finland will have quite a lot in common in this regard up till the second half of the 1980s.

An interesting feature of most of these explanations is, as noted by Ljunqvist and Sargent (ibid. p. 6), that they assign the problem to the *demand* for labour, making the decisions either of employers or of unionized employed workers sustain a high unemployment rate.[1] Ljunqvist and Sargent (ibid.), on the other hand, focus on the effects of the welfare state on the supply of labour and hope to contribute a sense of how the welfare state adversely affects the dynamic responses to economic shocks and to increasing turbulence in the economic environment. The starting point of their analysis is the well-known claim that high-income taxation and generous welfare benefits distort workers' labour supply decisions.[2] This factor forms an important difference between Finland and New Zealand.

Unfortunately, there is no general agreement among economists regarding the best approach to modelling unemployment in Europe. Saying this we do not want to convey the impression that there has been no theoretical progress in the attempts to identify the determinants of (the natural rate of) unemployment[3] nor that the various theories proposed in the literature are not broadly consistent the European unemployment numbers. Rather, as also argued by Nickell (1998), there seems to be a lack of a satisfactory *empirical* explanation of the time series pattern of unemployment in Europe, or OECD for that matter.[4]

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[1]The growth oriented approaches to unemployment build on the Schumpeterian (1942) notion of 'creative destruction' and emphasize how growth and technological progress exert a continuous pressure on the economy to reallocate its resources from contracting sectors to expanding ones. This on going process of reallocation of the economy's factors of production is often called 'churning' (Caballero and Hammour, 1998c). The transaction between capital and labour – when these are combined to form new production units – suffers from an 'appropriability' problem, whenever investments exhibit some degree of *specificity* w.r.t. labour (Caballero and Hammour, 1986, 1998b), which in turn gives rise to quasi rents and insider power to workers. Unemployment in this context acts as an equilibrium response to the economic system that restrains the bargaining position of the insiders and preserves the profitability of investments (Caballero and Hammour, 1986, p. 818; on the macroeconomic effects of specificity, see Caballero and Hammour, 1998a).

[2]The assertion of Layard, Nickell and Jackman (1991, p. 62) that "unconditional payment of benefits for an indefinite period is clearly a major cause of high European unemployment" should also be mentioned, as it is born out by the analysis of Ljunqvist and Sargent. Ljunqvist and Sargent (1998) formulate a general equilibrium search model where workers' skills depreciate during unemployment spells, and unemployment benefits are determined by workers' past earnings. Simulations of the model clearly bring out the sensitivity of the equilibrium unemployment rate to the amount of skills lost by lay-offs. Their analysis attributes the persistently higher unemployment from 1980's to the increased turbulence, as measured by a special 'turbulence' parameter, in the economic environment, while also explaining how lower unemployment rates in the 1950's, the 1960's and the early 1970's were sustainable under more tranquil economic conditions.

[3] Indeed, Blanchard and Katz (1996) even go as far as to argue that the theoretical progress over the last 30 years or so has actually produced a framework for the analysis of the determinants of the natural rate.

[4]Ljunqvist and Sargent (ibid. p. 6) also point to the possibility that one reason for the past lack of emphasis on workers' distorted incentives as an explanation for high European unemployment must be the scarce empirical support for the idea. They argue that many empirical studies have failed to find any cross-country correlation between unemployment benefits and aggregate unemployment, and this has given rise to the conclusion that generous entitlement programs are not to be blamed for high unemployment rates. It appears, however, that unemployment rates do respond to benefit entitlements, but only with a considerable lag of 5 to 10 years, and in some cases 10 to 20 years. Why are the lags so long then? Ljunqvist and Sargent argue that lags are purely coincidental, and the real explanation for persistently higher European unemployment from the 1980's is to be found in a changed economic environment.

Quite recently, fresh ideas on modelling the determinants of the interaction between (lagged) labour market dynamics and (time-varying) natural rate, in particular, have suggested approaches that have high empirical content. One of the most interesting ones of these, proposed by Karanassou and Snower (1997b),[5] explores the implications of the interaction of persistent effects of temporary shocks (or of unemployment persistence) and delayed effects of permanent shocks (or of imperfect unemployment responsiveness) on long-term unemployment and analyzes these effects by deriving the long-term implications of the "chain reaction" theory of unemployment. The idea is simply that inherently lagged dynamics in the labour market interact with growing exogenous variables (population growth, capital stock etc.) to produce long-term effects from labour market lags, in addition to the typical short-term effects.

Our objectives in this paper are more modest than in the recent literature on long-term unemployment, as exemplified e.g. by Karanassou and Snower (ibid.). We attempt to model long-term unemployment determination with a vector time series of relatively low-dimension consisting of unemployment and a number of potentially important factors correlated with unemployment as a (possibly) cointegrated system. Cointegrating vector(s) then give us estimates of the long-term (static) relationship between the system variables. In particular, we test for the exogeneity of the other variables w.r.t. the parameters of the cointegration relationship. This is the first step in trying to identify an unemployment equation from the system. Granger's representation theorem (Granger and Engle, 1991) implies that a cointegrated system can equivalently be represented in an Error Correction Form, which summarizes the short-run dynamics of the changes of the system. It is well known that cointegration implies causality among the variables in the system. Testing for non-causality of unemployment w.r.t. to other variables in the system can finally be used to check whether single-equation approach to modelling unemployment is fully efficient.

2.1 A macroeconomic model of the labour market in a small open economy

The analytical framework, which underpins the structure of the estimated cointegration vectors in section 3, follows the mainstream of the macro-oriented labour economics in consisting of a labour and supply function as well as a wage setting equation. We take the Jacobson et al. (1996) model and modify it to reflect the openness of the economy.

Production possibilities are described by the (log of the) Cobb-Douglas functional form

$$y_t = \beta k_t + \alpha l_t + \mu_t \qquad (1)$$

where y, k and l denote, respectively, output, capital and labour, and μ is a stochastic shock to technology or production possibilities (i.e. a productivity shock), whose generating process may contain a unit root. Employment is assumed to be given by

[5] See also Karanassou and Snower (1997a, 1998).

$$l_t = \gamma_0 + \gamma_1 k_t - \gamma_2 \nu_t + \zeta_t \qquad (2)$$

where $\nu = \ln(W/P)$ is the real product wage, W the nominal wage, P the output price and ζ an employment shock. If the firm is maximizing profits, then $\gamma_1 = \beta\gamma_2$ and $\gamma_2 = 1/(1-\alpha)$ and $\zeta_1 = \gamma_2\mu$. Capital is determined by

$$k_t = \rho_0 + \rho_1 l_t - \rho_2 r_t + \zeta_t \qquad (3)$$

where r is the (log of the) rental price of capital, R, or, simply, the real interest rate. In what follows we assume that the real interest rate is constant[6] and hence subsumed by the constant ρ_0. Note that employment and capital stock share the same shock ζ, which essentially derives from the underlying shocks to technology, μ. For this reason ζ will be simply refered to as 'a technological shock'. Under profit maximization $\rho_1 = \alpha/(1-\beta)$.

Labour supply is given by

$$l^s_t = \theta w_t + \xi_t \qquad (4)$$

where $w = \ln(W/\Pi)$ is the consumption real wage and Π is the consumer price index with the share of domestic consumer goods in the index denoted by ϕ.[7] Non-modelled factors affecting labour supply decisions by households[8] are represented by the stochastic shift variable ξ, which is generated by a process that potentially contains a unit root. Furthermore, we assume that the wage setting relation is

$$w_t = -\sigma_1 u_t + \sigma_2 q_t + \omega_{w,t} \qquad (5)$$

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[6]The real interest rate has in fact varied over the period and has affected other variables in the model, particularly unemployment, as also emphasized by Phelps and Zoega (1998) inter alia. However, we could not disentangle those effects from the data. In any case reliable interest rate data may be hard to come by, since a large part of the sample comes from a period of financial (rate) regulation in the two countries.

[7]Such a labour supply function can be derived from an underlying (static) constrained utility maximization problem, where the utility function is separable in leisure, H, and aggregate consumption, C, and 'quasi-linear' in form, i.e. $u(C, 1 - H) = C - \kappa(1 - H)^\rho$ and where the budget constraint can be written as $C + wH = w$ (time endowment normalized to 1), where w denotes the real consumption wage, W/Π. Aggregate consumption, on the other hand, corresponds to the Cobb-Douglas utility aggregator in the constrained utility maximizing decision on the allocation of consumption across domestic and foreign goods, so that Π can be thought of as the cost minimizing price of a unit of aggregate consumption and can be written as $P^\phi(P^*S)^{1-\phi}$, where P* and S denote the foreign price level and nominal exchange rate, respectively.

[8]These non-modelled factors could conceivably include preference shocks (between consumption and leisure), changes in the labour flow across borders (net immigration), changes in the birth and mortality rates, reservation wages and social safety nets (unemployment and other benefits) etc.

where u and q denote, respectively, the log of unemployment U and the terms of trade Q.[9] This particular form of the wage setting equation differs slightly from the standard one in the literature (e.g. Jacobson et al. eq. (4) p. 4). Typically, the real product wage is expressed in terms of unemployment and productivity, the former affecting real wages negatively and the latter positively. Above, real consumption wage depends on the terms of trade movements and negatively on unemployment. This particular form of the wage setting relation basically results from the fact the difference between real product and real consumption wage in an open economy is related to the terms of trade. For a Cobb-Douglas type consumer price index (see fn. 7), we have $\nu - w = -(1-\phi)q$. Since labour supply decisions depend on the consumption real wage and labour demand decisions on the real product wage, changes in the terms of trade will affect the state of the labour market.[10] In any case, equation (5) is particularly useful in the sense that it emphasizes the terms of trade as a potential source of aggregate fluctuations affecting the economy. Historically Finland, for example, has been subject to rather large terms of trade swings, particularly at the business cycle frequencies, which also contributes to explaining why the terms of trade is included in the data set for the empirical analysis.[11]

As for the non-modelled factors affecting wage formation, i.e. shocks to wage formation, $\omega_{w,t}$, it is assumed that it too is generated by a process that potentially contains a unit root.[12] Hence, the wage, employment and labour supply shock and the shock to the capital stock can, without loss of generality w.r.t. to the long-run properties of the model, be represented simply as an AR(1) process of the form

$$V_t = \psi V_{t-1} + \varepsilon_{V,t} \tag{6}$$

where $0 < \psi \leq 1$, (i.e. (6) can be a pure random walk).[13]

The main reason for wanting flexibility in the stochastic structure of the shocks is that the cointegration implications of the model depend on the number of unit roots in the exogenous stochastic processes. For example, if shocks to

[9]Terms of trade usually refers to the price of domestic exports relative to imports in domestic currency. Hence, it is the relative price associated with tradable goods. The real exchange rate, on the other hand, is often defined as the ratio of domestic and foreign price levels (in domestic currency). In the theoretical context of a single good small open economy, these two coincide. This is the main reason for us to use 'terms of trade' in a loose sense in theoretical context of the main text.

[10]The literature on wage indexation (Gray, 1976; Fischer, 1977; Karni, 1983), especially in open economies (Turnovsky, 1983; Aizenmann and Frenkel, 1985a, b) under optimal wage indexing (Devereux, 1988; Vilmunen, 1992) provides a very useful theoretical background in this context.

[11]Blanchard and Katz (1996) argue that research development has resulted in a tractable framework for the (macroeconomic) analysis of unemployment problems, and that the set of exogenous determinants affecting long-run unemployment has to be adjusted to reflect the context of application. The idea here is similar in spirit; the emphasis here is not only on the nature of shocks as such, but also on the propagation mechanism underlying unemployment determination in small open economies.

[12]These non-modelled factors could include (institutional etc.) parameters related to wage bargaining as well as various restrictions, such as minimum wage laws, on wage formation. Also, shock to real interest rates could affect the stochastic behaviour of shocks to wage formation.

[13] We could generalize equation (6) by letting $|\psi| \leq 1$ and retaining the property that $\varepsilon_{V,t}$ is stationary. In this more general case, V_t is stationary, but not necessarily an AR(1) process, when $|\psi| < 1$. On the other hand, when $|\psi| = 1$, V_t is integrated of order 1, I(1), but not necessarily a pure random walk.

wage formation are generated by a random walk and $u_t \sim I(1)$ and $q_t \sim I(1)$, then the trivariate system $X_t = (w_t, u_t, q_t)$ cannot be cointegrated (i.e. linear combinations $\beta' X_t$ in the present context cannot be I(0)), whereas stationary, mean reverting shocks to wage formation imply cointegration among these variables. This implication of the wage shock is emphasised also by Jacobson et al. (1996 p. 6).

Finally, fluctuations in the terms of trade evolve according to

$$q_t = -\delta_1 k_t - \delta_2 l_t + \omega_{q,t} \tag{7}$$

where $\delta_i > 0$. (7) can be derived by combining an IS-schedule linking competitiveness to aggregate demand (at the constant real interest rate)[14] with an aggregate supply behaviour implied by (1)–(3), i.e. the terms of trade balances aggregate demand and supply. Under this interpretation, the terms of trade shocks ω_q is a combination of aggregate demand and supply shocks (or shocks to the production technology). Thus it is possible that ω_q is generated by a process containing a unit root, so that ω_q can be represented as $\omega_{q,t} = \psi_q \omega_{q,t-1} + \varepsilon_{q,t}$, $0 < \psi_q \leq 1$. An alternative interpretation relies on mark-up pricing under exogenous terms of trade shocks.

From equations (2) and (4) we can obtain an expression for (the rate of) unemployment[15] (the constant has been ignored)

$$u_t = l_t^s - l_t = (\theta + \gamma_2)\, w_t - \gamma_2 (1 - \phi)\, q_t - \gamma_1 k_t + \xi_t - \zeta_t . \tag{8}$$

The demand for capital function (3), on the other hand, can be written in an alternative form as (once again ignoring the constant)

$$\begin{aligned} k_t &= -\rho_1 [l_t^s - l_t] + \rho_1 l_t^s + \zeta_t \\ &= -\rho_1 u_t + \rho_1 \theta\, w_t + \zeta_t + \rho_1 \xi_t \end{aligned} \tag{9}$$

which may prove useful when interpreting possible cointegration relationship; in particular, the stochastic properties of the linear combination $k_t + \rho_1 u_t - \rho_1 \theta w_t$ will depend on the stochastic properties of the linear combination, $\zeta_t + \rho_1 \xi_t$, of shocks to labour demand and supply.

[14] I.e. the underlying IS-schedule is $y = -aq + u$ for some positive a, where u denotes exogenous IS or aggregate demand shocks, and includes fiscal policy shocks, exogenous shocks to consumption and investment etc. Implicitly we are abstracting from interest rate determination and simply take the real interest rate is as parametrically given.

[15] Note that according to (9), long-run unemployment need not be independent of shocks to labour demand and supply. Lindbeck (1993) argues that a realistic macroeconomic theory should have long-run unemployment independent of productivity and labour supply shocks. This is very similar in spirit to the identification scheme used by Blanchard and Quah (1987) to identify aggregate demand and supply shocks, and is similar to what Blanchard and Katz (1996, p. 9) argue. Above, independence will prevail essentially if $\xi_t - \zeta_t$ is stationary, which, in the case of unit roots in labour demand and supply shocks, boils down to ξ_t and ζ_t sharing a common trend. Productivity, or a similar particular form of technological progress, would probably be the most plausible interpretation of this common trend, since shocks to productivity would affect not only labour demand, but also labour supply (see e.g. Blanchard and Katz, 1996, pp. 9–10).

Now, we can use the preceding model to motivate our cointegration analysis in the next section. Without solving the model explicitly (see Jacobson et al. 1996), we know that the vector of the four endogenous variables of interest, X = (u, w, q, k)' can be solved[16] as a linear function of the underlying vector of shocks $F = (\zeta, \omega_w, \xi, \omega_q)' = (V, \omega_q)'$, i.e.

$$X_t = A\,F_t \tag{10}$$

where A is (4 x 4)-matrix summarizing the impact effects of the shocks.[17]

Equation (10) is a particularly useful representation of the solution, since it enables us to control for the number of cointegrating relations more explicitly (as the mirror image of the number of stochastic trends, i.e. independent unit root processes). Clearly, the system in (10) consisting of the four endogenous variables may be driven by up to four independent unit root processes, in which case there exists no cointegrating relations among the variables (u, w, q, k) (Stock and Watson, 1991). If, on the other hand, one of the processes in F_t is stationary, then there is one cointegrating relation among the variables (u, w, q, k), i.e. there are three stochastic trends. At the other extreme, if only one of the exogenous driving processes, i.e. shocks to the labour supply, is a unit root process, then there are three cointegrating relations among the variables (u, w, q, k).

One particularly interesting case occurs when technological and labour supply shocks are both generated by stationary, mean reverting processes, i.e. there are at least two cointegrating relations among the variables in (u, w, q, k), the unemployment equation (8) gives (one of the) cointegrating relations:

$$u_t - (\theta + \gamma_2)\,w_t + \gamma_2(1 - \pi)\,q_t + \gamma_1 k_t = \xi_t - \zeta_t \sim I(0) \tag{11}$$

An equation like (11) is an example of a relationship between the variables in (u, w, q, k) that we hypothesise exist among the observed time series counterparts of these variables in the two countries concerned. As the logic of the preceding model suggests, however, (11) need not be the only cointegrating relation between the variables; interestingly, two pairs of variables in (11) could be cointegrated, so that (11) in effect represents a linear combination of these two stationary variables or cointegrating pairs.

[16]There will be a unique solution, provided the determinant of the "structural" matrix linking the four endogenous variables together in the model structure is non-zero, see Jacobson et al. (1996) for an analogous condition (p. 5, where they denote the determinant by ψ).
[17]Note that a vector of constant is missing from the solution (10), since we have abstracted from the relevant constants in developing the theory.

3 Empirical analysis of the Finnish and New Zealand data

3.1 Econometric preliminaries: cointegration analysis

In our statistical analysis we are mainly be interested in the long-run comovement of unemployment, real (consumption) wages, capital stock and terms of trade. The main reason for focusing on this particular variable set is the desire to combine fluctuations or shocks predominantly at business cycle frequencies (terms of trade movements) with shocks to potential output (evolution of the capital stock) to get a quantitative estimate of their (joint) contribution to the rise in the observed unemployment rate in these two small open economies. Both of these economies are cyclically sensitive and in both of them major changes have taken place, suggesting that the longer-run growth performance of the economies may have been affected. In New Zealand measures to increase external competition or measures to improve the operation of the labour market have been taken, while in Finland major external shocks, the collapse of the trade with the former Soviet Union, impinging on the economy have occurred. Furthermore, Finnish growth policy in the past has resulted in inefficiency through over-investment,[18] which has certainly contributed to explaining the observed decline of the capital stock in the 1990's after the crisis set in.

We use cointegration analysis to model the possible long-run relationship between these variables. To this end, consider the four-dimensional vector time series $X_t = (u_t, w_t, q_t, k_t)$, where u, w, q and k denote, respectively, log of the unemployment rate, real consumption wage, terms of trade and net (business) capital stock. A statistical model for the vector time-series X_t, the component series of which are assumed to be I(1)-processes, is provided by an unrestricted vector autoregression of length k,VAR(k)

$$X_t = \Pi(L)\, X_{t-1} + \varepsilon_t \tag{12}$$

where

$$\Pi(L) = \Pi_1 + \Pi_2 L + ... + \Pi_k L^{k-1},\ \varepsilon_t \sim NID_4(0, \Omega)$$

for fixed initial values of X_{-k+1}, ..., X_0. In (12) L is the lag-operator $L^j X_t = X_{t-j}$. The hypothesis of interest is the number of possible cointegration relations among the component series of the vector time series X_t. We follow Johansen's (1988, 1991)[19] FIML approach to the statistical inference in cointegrated systems[20] here. The procedure involves estimating a VAR(k) by maximum likelihood and starts

[18]See Pohjola (1996).

[19]See also Johansen and Juselius (1990, 1992). Johansen's book (1995) provides a unifying reference.

[20]Full efficiency of the Johansens's procedure may well enhance its attractiveness in the approach to the statistical analysis of cointegrated systems. However, to our minds the fact that it provides a unified framework for inference in such systems is perhaps a more compelling reason to follow the procedure. Inder (1993) provides a comparison of some of the different methods of estimating long-run relationships in economics.

by transforming the VAR(k) in (12) into its equivalent vector error correction form (VECM, for short):[21]

$$\Delta X_t = \Pi X_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \varepsilon_t \tag{13}$$

where

$$\Pi = \sum_{i=1}^{k} \Pi_i - I_4, \ \Gamma_i = -\sum_{j=i+1}^{k} \Pi_j.$$

The number of cointegration relations depends on the rank, r, of the long-run Π matrix in the known way; if X_t is not a difference (r = 0) or level (r = 4) stationary VAR(k), then there are exactly r, 0 < r < 4, cointegration relations among the four variables. This is equivalent to saying that the series can be jointly characterized by 4–r common or stochastic trends,[22] i.e. the vector representation of these series has 4–r unit roots.

The hypothesis that there are r cointegration relations in the system,[23] H(r) for short, can thus be stated formally as a reduced rank condition on the matrix Π,[24]

$$\Pi = \alpha \beta' \tag{14}$$

where α and β are (4xr) matrices of full column rank, called the (factor) loading matrix and the matrix of cointegrating vectors, respectively. Since the hypotheses form a sequence of nested hypotheses (a lower dimensional space is embedded in a higher one), the idea is to test for a increasing number of cointegration vectors in the system, starting from the hypothesis that r = 0.

Specific hypotheses about the long-run cointegrating vectors β,[25] which can be formalized as linear restrictions on β, can be formulated and tested within the Johansen framework.[26] Analogously, we can impose and test restrictions on the

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[21]A cointegrated system can, under mild regularity conditions, be equivalently represented as a VECM. This equivalence is the result of the Granger's representation theorem (see e.g. Granger and Engle, 1991, or Johansen, 1995).

[22]Johansen (1995, p. 41) emphasizes the relationship between the error correction and the common trends model: they are complementary in the sense that the two approaches are mathematically equivalent. They could, of course, appeal to different types of intuition. Stock and Watson (1991) propose tests that can alternatively be viewed as tests of the number of common trends, linearly independent cointegrating vectors or autoregressive unit roots of the vector process. See also Warne (1990a,b, 1993).

[23]Only the dimension of the cointegration space will be estimated by the Johansen's procedure. The set of vectors that generate this space will not be identified by the procedure, i.e. the set is not unique, since a non-singular transformation of the set of cointegration vectors will do as well and will amount to representing the cointegration space in a different co-ordinate system.

[24]Anderson (1984) is the relevant reference for the theory underlying reduced rank regressions.

[25]The estimates of the long-run parameters $\hat{\beta}$ are asymptotically mixed Gaussian (see e.g. Johansen, 1995, ch. 15), which has implications in the context of the discussion on standard confidence intervals built upon the cointegrating vectors later in the text.

[26]We could, in particular, test for the stationarity of any of component series by testing whether the unit vector (0, ..., 1, ..., 0)' (1 in the i^{th} place) belongs to the cointegration space.

loading matrix α.[27] Perhaps the most interesting of the hypotheses concerning the loading matrix relates to the concept of weak exogeneity, which, if it holds for a subset of the variables, means that inference about cointegration (including loadings) in the partial system not involving weakly exogenous variables is valid.[28] Weak exogeneity can be a very useful property of a system, especially when it is easier to model the conditional model of the endogenous variables given the exogenous variables satisfactorily and the marginal distribution of the exogenous variables shows irregular behaviour, which is difficult to model using a VAR.

Johansen has proposed two test statistics for testing for the number of cointegrating relations and general linear hypotheses, the maximum eigenvalue and trace tests respectively. The latter is based on the sum of the p–r smallest canonical correlations (or eigenvalues) between (standardized) residuals from projecting changes in X_t and the level of X_{t-1} on lagged changes in X_t. The asymptotic distribution theory is non-standard, and tabulated distributions have to be consulted.[29] The (asymptotic) χ^2 distribution theory can be, however, invoked when testing linear restrictions on the cointegrating vectors. Analogously, restrictions on the loading matrix can be tested comparing the estimates of the eigenvalues in the cases when α is restricted and unrestricted. Inference in these tests relies on a χ^2 distribution.[30]

3.2 Estimation results for Finland

Semi-annual data for Finland are used covering the sample period 1960.1–1996.2. In general, it is difficult to choose the appropriate periodicity for the data. Quarterly data tend to have a high level of noise and short-run dynamics can obscure the more fundamental analysis we are concerned with here. Semi-annual data, however, are much more stable. Since both annual and semi-annual data provide similar outcomes we report the semi-annual results to minimize any loss of information. The variables included in the empirical analysis are, as explained earlier, the logarithm of the rate of aggregate unemployment (u), real consumption wage (the ratio of average earnings to the CPI, 1990 = 100, w), terms of trade (ratio of export prices to import prices, 1990 = 100, q) and net capital stock in the business sector (k).

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[27]The matrix α gives (factor) loadings "along the stationary dimension" of the cointegrated system. Hence, loadings "along the non-stationary" dimension, denoted by $\alpha_\perp$, i.e. coefficients of the common trends belong to the orthogonal complement of the subspace generated by the columns of α, i.e. $\alpha'\alpha_\perp = 0$. Cointegrating vectors and common trends are obtained from the solution to the same eigenvalue problem, the former corresponding to the r largest and the latter to the p–r smallest eigenvalues, so they are estimated in a dual manner in the Johansen's cointegrated VAR context (see Johansen, 1995, section 8.3).

[28]For a more thorough analysis of an 'I(1) model', see Johansen (1995, especially ch. 5); ch. 8, on the other hand, deals with weak exogeneity and valid conditional inference.

[29]See e.g. Johansen (1995, ch. 15).

[30]Johansen (1995, ch. 8).

3.2.1 Unit root testing on the Finnish data

In conducting the unit root test, lags up to 4 (i.e. two years) were included. The maximum of four lags appeared to be sufficient to filter out most of the residual autocorrelation present in the residuals after fitting an AR-process to the time series. Appendix 1 contains summary tables from the unit root testing. Table 1.1 summarizes the results of the Augmented Dickey-Fuller tests for Finland.

Even though the formal test suggests a unit root in the (log) level of the variables in the system, the unemployment rate as well as the capital stock appear borderline cases, but in different directions. Whereas the log of the unemployment rate comes close to being stationary, the log of the capital stock appears to come very close to being an I(2) process! In particular the null of a unit root in the DGP for the rate of unemployment cannot be rejected when more lags are included in the ADF test. This may be a reflection of the reduction of power in unit root testing when the number of lags is increased.[31] On the other hand, the sum of the estimated AR-coefficients, the beta Y_1 column, in the case of the log of the capital stock appears to be large in comparison to the corresponding ones of other variables in the system, and even exceeding one, once lags are added.[32] Further

[31]The argument that unemployment rate cannot have a unit root, because it is bounded by 0 and 1 – so that random labour market shocks would drive the unemployment rate to 0 or one with the passage of time (see e.g. Karanassou and Snower, 1997a, p. 4 fn. 11) – needs qualification. First of all, unemployment rate cannot be an *unrestricted* (linear) random walk or Brownian motion because of the bounds. But it can be *regulated* Brownian motion (or even a Brownian bridge). Hence, there can be a unit root in the unemployment rate, but its fluctuations are constrained by barriers, most plausibly by reflecting barriers, since absorbing barriers would imply that the unemployment rate stays at a particular level, once it reaches that level. These barriers can e.g. reflect the internal workings of the economy itself or they can result from policy regulation. The existence of such barriers raises the possibility of a non-linear relationship between the unemployment rate and its fundamental determinants. Alternatively, there could be non-linearities in the process generating observed unemployment – e.g. regime shifts – so that observations on the unemployment rate look favourable to a unit root, in the context of unit root testing. At a general level, however, the unemployment rate is not much different from many other economic series, in the sense that these series cannot strictly speaking be modelled as (symmetric) unrestricted Brownian motion at least because of the existence of non-negativity constraints, which are typically ignored in the unit root tests of these series. Consumption, output, prices etc. are subject to non-negativity constraints and cannot thus be, strictly speaking, modelled as (symmetric) unrestricted Brownian motion. In these circumstances 'unit root econometrics' is implicitly assumed to give a reasonable basis for statistical inference in the context of (statistical) modelling these series. A similar line of reasoning is applied in the present context.

[32]The notation used in PcGive 9.0 for Windows (1996) by Hendry and Doornik may cause confusion at this particular point; p. 212 informs the reader that the beta Y_1 is 'the coefficient on the lagged level: β' while the test equation is given in equation (16.7) p. 211: $\Delta x_t = \alpha + \mu t + \beta x_{t-1} + \Sigma\gamma_i\Delta x_{t-i} + u_t$. Our Table 1.1 is representative of the output produced by PcGive; for example β equals 0.847 in a DF-test (number of lags equals 1) for a unit root in the unemployment rate, which seems to imply that the estimated AR(1)-coefficient, $\hat{\rho} = 1 + \hat{\beta} = 1 + 0.847 = 1.847$! Clearly, then, beta Y_1 in the PcGive output table from a unit root test is not the estimated β in the above ADF test equation. Rather, beta Y_1 equals the sum of the estimated AR-coefficients from the equation, as argued in the main text. An estimate of the β is obtained by subtracting one from beta Y_1, $\hat{\beta}$ = beta Y_1 – 1.

evidence is provided by Table 1.2, which tests for a unit root in the first difference of the series (i.e. growth rates).[33]

At face value, Table 1.2 seems to suggest that there is also a unit root in the growth rate of real wages and the capital stock. However, the β-coefficient on the relevant lagged difference for the growth rate of real wages is 0.48 and 0.42 at lags 3 and 4 respectively and the ADF-test suggests a unit root in the process generating growth in real wages. For the capital stock, on the other hand, the β-coefficients are higher, but still well bounded above by 0.9. This suggests that there is a substantial amount of autocorrelation in the growth rate of the capital stock or that the growth rate of the capital stock series is relatively 'smooth'.[34]

Further specification tests[35] suggest the following observations; (i) general as well as ARCH-type heteroscedasticity is present in the residuals from the ADF-test equation, which cannot be removed by the usual procedure of adding in further lagged differences. (ii) graphical inspection indicates that there are two large residuals around the year 1990, which give rise to deviations from normality in the form of thick tails and skewness to the left. These deviations from the ideal conditions underlying the ADF-tests tend to reduce the efficiency of these tests in finite samples.[36] Overall, then, the decision to reject the hypothesis of unit root nonstationary growth rates of the variables in our system should be viewed with caution, especially because of the test results for the growth rate of the capital stock. With this in mind, we will continue to the cointegration analysis under the assumption that the vector time series $X_t = (u_t, w_t, q_t, k_t)$ is generated by an I(1) vector process.

3.2.2 Cointegration analysis of the Finnish data

Johansens's procedure was followed to test formally for the dimension of the cointegration space as well as to run weak exogeneity tests on some of the variables, most notably the capital stock, w.r.t. the parameters of the cointegration relationship among four variables in our system. Appendix 2A gives the relevant

[33]The unit root test for the terms of trade is in line with (the extensive literature on) testing the validity of PPP (purchasing power parity) in the sense that the test here suggests that deviations from PPP are nonstationary (see Rogoff, 1995, for an excellent survey on PPP and long-run real exchange rates).

[34]The shape of the estimated spectrum for the growth rate of the capital stock also confirms that there is substantial autocorrelation in the series. The estimated spectrum for the growth of the real wages, on the other hand, is U-shaped, where the minimum occurs approximately at frequency 3/4, implying that cycles shorter than 1 1/3 years make a considerable contribution to the variance of the growth rate of real wages.

[35]Available from the authors upon request.

[36]We will return to the possible I(2) of the capital stock in the context of the cointegration analysis, where the growth rate of the capital stock is also used instead of the log-level. As far as the other diagnostics are concerned, we checked the outcome from the ADF test by including two impulse dummies in the test equation to mitigate the effects of the outliers around 1990. The t-statistics, t_β, did generally rise at various lags, even to the extent that at lags 3 and 4, the null of a unit root was rejected at the 5 % significance level. The 10 % critical value for the ADF test (from McKinnon, 1991), on the other hand, is around –3.1619, which tends to lend support for the rejection of a unit root in the growth rate of the capital stock, but does not, unfortunately, fully sustain the conclusion in the main text that the growth rate of the capital stock is (trend) stationary.

summary tables from the (unrestricted) cointegration analysis.[37] According to Table 2.1b, both the maximal and trace test, in the uncorrected form in particular, seem to suggest that the cointegration rank be 1, i.e. there is evidence of one cointegration relationship between the unemployment rate, real consumption wage, capital stock and terms of trade in Finland. The unrestricted estimates of the β-matrix as well as of the loading matrix (vector) α corresponding to the proposed cointegration vector – maximal eigenvalue or canonical correlation – (emboldened) are Table 2.1c.

Given that the cointegration dimension is one, the estimated cointegration vector appears to be reasonable from an economic point of view. In particular, if we could conclude that it is an unemployment equation, the signs of the individual coefficients are consistent with theory, although they are perhaps a little too large (in absolute value).

Overall, then, Table 2.1c seems to suggest that the following linear combination of the unemployment rate, real consumption wage, capital stock as well as the terms of trade in the Finnish data, normalized by the unemployment rate,

$$ECM_t = \hat{\beta}' X_t = u_t - 3.23\, w_t + 1.53\, k_t + 1.63\, q_t \tag{15}$$

(where the hat signifies 'estimated'), is stationary, i.e. I(0). The corresponding standardized (factor) loadings in Table 2.1d, which interpret the effect of the disequilibrium error corrected for lagged differences,[38] indicate stable error correction dynamics.

The structure of the estimated loading matrix has an interesting structure. In the present context we can write the VECM as

$$\begin{aligned}\Delta X_t^{FI} = \begin{pmatrix} \Delta u_t \\ \Delta w_t \\ \Delta k_t \\ \Delta q_t \end{pmatrix} &= \begin{pmatrix} -0.341 \\ -0.011 \\ -0.001 \\ -0.027 \end{pmatrix} (u_{t-1} - 3.322\, w_{t-1} + 1.526\, k_{t-1} + 1.626\, q_{t-1}) \\ &+ \sum_{j=1}^{3} \hat{\Gamma}_j \Delta X_{t-j}^{FI} + \hat{\varepsilon}_t^{FI}\end{aligned} \tag{16}$$

[37] A battery of specification tests were run on the ECM representation underlying the cointegration analysis. The ECM of the VAR(4) captures most of the observed variation in the growth rates of variables (i.e. Δx_t); no residual autocorrelation is left in the equations; a small ARCH-effect at lag 2 is present in the residuals for the growth rate of capital, but no more general form of heteroscedasticity can be detected from the residuals. The null of normal residuals, however, is rejected in the case of the growth rate of unemployment and capital stock. It appears from the graphs of the residuals that this stems mainly from skewness. The sample distribution of unemployment residuals appears to be slightly positively skewed due to the sharp increase in the observed unemployment rate in 1990–1991. Residuals from the growth rate of the capital stock appear to be a mirror image of those of the unemployment rate; i.e. their sample distribution is skewed to the left.

[38] And hence involves all the parameters of the model.

Table 2.1d seems to suggest that the factor loading on the capital stock is zero, and very small on real wages and perhaps also on the terms of trade. We tested the hypothesis of zero loading on the capital stock using Johansen's likelihood ratio test.[39] The numerical value of the test statistic, which is asymptotically $\chi^2(3)$ (see Johansen, 1995, Th. 8.2, p. 126), is 0.4257 with p-value 0.9349. Hence, weak exogeneity of the capital stock w.r.t. to the long-run parameters cannot be rejected, and the adjustment matrix can be written as $\hat{\alpha} =$ (–0.34, –0.01, 0, –0.03)´ in accordance with the order of variables (u, w, k, q). The corresponding cointegration vector, β_α, say, is (1.008, –3.178, 1.383, 1.543) or, in a standardized form, (1.000, –3.153, 1.372, 1.531) with (asymptotic standard errors 0.173, 1.323, 0.820 and 0.787 respectively). Hence, restricting the capital stock to be weakly exogenous slightly reduces the coefficients of the cointegrating vector ('long-run elasticities').[40]

We also tested for the weak exogeneity of the capital stock and real wage for the long-run parameters, because it is an interesting hypothesis in itself in the sense that if not rejected it implies that real wages are not error correcting. Unemployment, on the other hand, will respond to all sources of exogenous shocks directly and through endogenous adjustment.[41] This hypothesis cannot be rejected (p-value is 0.605), and the corresponding unrestricted β_α vector is now (1.027, –3.121, 1.246, 1.591).[42] Once again, a slight reduction in the 'long-run elasticities' occurs. Finally, we tested for the weak exogeneity of the capital stock, real wages and terms of trade for the parameters of the cointegrating relations. The p-value drops drastically, to 0.059, so that formally this hypothesis is a borderline case. The estimated unrestricted β_α vector is now (1.003, –2.864, 1.154, 1.166).[43] It should, however, be noted that though formally a borderline case, the sharp drop in the p-value in the last exogeneity test is perhaps best interpreted as a warning to the modeller; the numerical test results should not be taken at face value and interpreted too rigidly to avoid running the risk of accepting too easily. In order to check for this possibility, we tested for the weak

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[39]All the subsequent tests are conditional on r = 1, i.e. that the cointegration rank is one.

[40]Note that the asymptotic standard errors of the components of the β_α-vector are relatively large. For the capital stock and terms of trade components, the $\beta_\alpha \pm$ 2S.E, where S.E. is the estimated standard error of the component, i.e. 'approximate 95 % confidence intervals', include zero, although the precision in the recursive estimates increases sharply as the sample size increases recursively. These confidence intervals may, however, give misleading results of the significance of a component, because of the usual problem that the t-values do not adequately reflect the correlation between the coefficient estimates. The more appropriate LR-tests for the significance of each of the components indicate that the null hypotheses of a zero cointegrating component can be rejected decisively for u, w and k. The p-values for the terms of trade q, on the other hand, range from 0.1 to 0.15, depending on the exact test. Hence, the evidence for the inclusion of the terms of trade in the cointegration vector is weaker.

[41]This set up would correspond to our theoretical model with $\rho_1 = \sigma_1 = 0$. Under this assumption, we can immediately see that wages do not respond to labour demand or technology and labour supply shocks, only to shocks to wage formation and terms of trade. Since the capital stock is driven by technology shocks and terms of trade by its own autonomous component and shocks to wage formation, the bulk of the adjustment falls on unemployment in the sense that it is affected by all sources of shocks.

[42]Which gives (1.000, –3.039, 1.213, 1.549) in standardized form and the asymptotic standard errors are 0.1934, 1.2944, 0.7975 and 0.7717. Note that β_α – 2S.E is only slightly above the zero for the terms of trade component. See fn. 40 for comments on the interpretation of these confidence intervals.

[43]The standardized or reduced form vector is (1.000, –2.856, 1.151, 1.163) and the asymptotic S.E's are 0.2107, 1.4103, 0.8689 and 0.8409.

exogeneity of the terms of trade alone w.r.t. the long-run parameters and the test result has a low p-value of 0.065. Once again, this is formally a borderline case, but strongly suggesting that the terms of trade is, in the end, not weakly exogenous.

Finally, we re-estimated the system with the (log of the) capital stock replaced by the growth rate of the capital stock (i.e. log-difference of the capital stock). We wanted to see, whether there was any effect on the results from taking the capital stock as an I(2)-process (see fn. 36). Test results do indicate that the cointegration rank is one – the uncorrected maximum eigenvalue statistic is 29.12 and the trace statistic is 47.36 – and the unrestricted estimate of the cointegration vector is $\hat{\beta} = (1, \; -1.23, \; 0.45, \; 0.72)'$, and the associated estimate of the loading matrix is $\hat{\alpha} = (-0.328, \; -0.006, \; 0.183, \; -0.032)'$. We can see immediately that the unrestricted estimates of the components of the cointegration vector are now much smaller. The original estimates are, as argued earlier, probably a little too large (in absolute terms), so the present ones appear to imply a move in the right direction. Weak exogeneity tests suggest that the growth rate of the capital stock is weakly exogenous – the p-value is 0.229 – and there is now stronger evidence also of the weak exogeneity of real wages; the p-value is 0.329 for the weak exogeneity of the growth rate of the capital stock and real wages jointly. Adding the terms of trade to list results, once again, in the p-value dropping sharply to p = 0.1. Finally, these results did not change much, when we added two impulse dummies – corresponding to 1989.1 and 1990.1 – to control for the observed ARCH-type behaviour of the growth rate of the capital stock observed also earlier in the context of the unit root tests (see fn. 37). Evidence in favour of the weak exogeneity of the capital stock and real wages (and perhaps also of the terms of trade) is slightly stronger in this case, however.

3.3 Estimation results for New Zealand

Semi-annual data were not available for New Zealand, so that the empirical analysis is conducted using annual data over the period 1960–1995. This implies that there are fewer observations available than in the Finnish case, even though the time span of the sample is the same. Also, because of time aggregation, caution must be exercised when making comparisons with results from the Finnish data. Various issues of International Financial Statistics were used to construct the terms of trade.[44] As in the case of Finland, let $X_t = (u_t, w_t, q_t, k_t) = (LRUE_t, LRWAGE_t, LTOT_t, LCSTOCK_t)$ denote the vector time series of the logarithm of the aggregate unemployment rate, real consumption wage, the terms of trade and the (aggregate) capital stock.[45] Before reporting the results from the cointegrating analysis, the Dickey-Fuller tests for number of unit roots in the series is given are reviewed. Tables 1.3 and 1.4 in Appendix 1 contain the relevant numbers and statistics from the ADF-tests. In the case of the unemployment rate, real wage and terms of trade, the ADF test equation is AR(2), while for the capital stock it is an AR(4),[46] possibly with a constant and a time trend included (see Tables 1.3 and 1.4).

[44]The ratio of export and import price deflators.
[45]Aggregate instead of business sector capital stock is used for New Zealand.
[46]Similar results were also obtained from under an AR(2) process for the capital stock.

3.3.1 Unit root testing

According to unit root testing on the levels of the variables in Table 1.3, there also seem to be uncertainties about the existence of unit roots in the generating processes, particularly for the unemployment rate and the terms of trade, in the case of New Zealand. Again one should perhaps notice the relatively small β-coefficients on the lagged level of the unemployment rate and terms of trade. Under the assumed AR(p) process for the relevant DGP, β equals the sum of the AR-coefficients. Accordingly, the implied estimate for the sum of the AR-coefficients ranges from 0.38 to 0.57 and from 0.44 to 0.53, respectively, in the case of unemployment and terms of trade. It could, of course, be argued that one should not put too much weight on the size of the estimated β-coefficient as such when making a decision on the (lack of) unit roots, the low values of these coefficients in the present context at least warns of the possibility that the low power of the ADF-tests in small samples may be responsible for the difficulty of rejecting the null of a unit root in these series. The residual variance is still particularly high in the case of the unemployment rate.

As for possible unit roots in the first differences, or growth rates of the series (Table 1.4), the test results once again indicate a unit root in the growth rate of the capital stock as well as of the real consumption wage. The estimated β-coefficients are again relatively low, ranging from 0.39 to 0.52 and 0.05 to 0.39, respectively, for the growth rate of the real wage and the capital stock, militating against the unit root in the growth rates. From the shape of the estimated spectra, on the other hand, we can infer that a nontrivial amount of positive autocorrelation is present in these growth rates, particularly in the capital stock.[47] So, small sample problems associated with the ADF-tests are very acute here. Since for the rest of the empirical analysis we require uniform time series properties of the variables in the data set, we will assume that the vector time series $X_t = (u_t, w_t, q_t, k_t)'$ is I(1). Given this assumption, we continue to the cointegration analysis.

3.3.2 Cointegration analysis

Appendix 2B reports the results from the cointegration analysis of the data from New Zealand. The unrestricted estimates of β-vectors as well as the estimate of the loading matrix α corresponding to maximal canonical correlation (eigenvalue) are reported in Table 2.2c and d.

Table 2.2a gives evidence in favour of one cointegration relationship between the unemployment rate, real wage, terms of trade and capital stock in the data from New Zealand.[48] The ordering of the variables in the estimated cointegration vector Table 2.2c is different from the one used for Finland. The reason is, to

[47]Even though the shape of the estimated spectrum for the growth rate of the capital stock displays the ‘typical spectral shape’ (Granger and Hatanaka, 1964) of many economic series, it is not as steep near the zero frequency as the estimated spectrum for the growth rate of the capital stock in the Finnish data.

[48]The null of no cointegration (r = 0) is, according to Table 2.2a, is rejected at 5 % significance level even after the correction for small sample bias is taken into account.

anticipate the results, that the data seem to suggest that the terms of trade, in particular, is error correcting, i.e. changes in the terms of the trade constitute an important short-run dynamic channel whereby the New Zealand economy adjust to shocks. Table 2.2a indicates that the particular linear combination

$$ECM_t = \hat{\beta}' X_t = q_t + 0.185 u_t + 0.742 w_t - 0.49 k_t \tag{17}$$

is I(0). The loading matrix, on the other hand, Table 2.2b, suggests a sizable effect of the disequilibrium error on the terms of trade and, in particular, on unemployment. A unit shock to the long-run relation gives rise to a 0.6 and 1.9 percentage point short-run response, corrected for lagged differences, of the terms of trade and unemployment rate respectively. The unemployment response, in particular, is perhaps too large and small sample bias may be one factor affecting the size of the adjustment coefficient and make the inference about the burden of short-run adjustment based on the estimated coefficient uncertain. Weak exogeneity tests of the variable w.r.t. the long-run parameters suggested that this is a distinct possibility.

We performed a sequence of weak exogeneity tests to see how far we can go restricting the loading matrix.[49] This would also give us further information about the possible nature of the cointegrating relation. First of all, we tested for the weak exogeneity of the capital stock for the parameters of the long-run relation; the p-value for the test of a zero factor loading on the capital stock is 0.9987 (i.e. $\chi^2(3) = 0.0294$). The corresponding cointegrating vector is β_α = (1.337, 0.175, 0.892, –0.381)′ with asymptotic standard errors 0.1973, 0.0466, 0.2075 and 0.1872.[50] The normalized vector is thus (1.000, 0.131, 0.667, –0.285)′. Hence, there is evidence that the capital stock is weakly exogenous to the long-run parameters.

We further tested, in sequence, whether in addition to the capital stock, real wages and the unemployment rate, are also exogenous to the parameters of the long-run relation. None of these hypotheses could be rejected at conventional significance level. It should, however, be noted that the p-value falls sharply in the last test; it is 0.1853 and the corresponding cointegrating vector is β_α = (1.278, 0.180, 0.833, –0.442) with asymptotic standard errors 0.1851, 0.0437, 0.1946 and 0.1756.[51] Again, then, the dramatic drop in the p-value should perhaps best be interpreted as a warning signal to the modeller. One should not take the test result at face value and accept weak exogeneity too easily. We followed a similar procedure as in the Finnish case, and tested for the weak exogeneity of the unemployment rate w.r.t. the long-run parameters without and with weak exogeneity of the capital stock; in the former case, the p-value of the test is 0.2162 and in the latter case it is 0.3833. Hence, we tempted to conclude in favour of at least weak exogeneity of the capital stock. Further testing indicated that the factor loadings for the unemployment rate could be reduced, even considerably, and that of the real wage increased, with a restricted loadings matrix (–1.087, –0.131, –0.131, 0)′ and the corresponding standardized cointegrating vector of (1.000, 0.138, 0.727, –0.367). In this context we could reject the hypothesis that the terms of trade is weakly exogenous to long-run parameters decisively, as the p-value is as low as 0.0007.

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[49]These tests are conditional on r = 1.
[50]The estimated long-run parameters are mixed Gaussian.
[51]The normalized vector is (1.000, 0.140, 0.651, 0.346)’.

Finally, we tested for the significance of each of the components in the cointegrating vector, i.e. we tested whether $\beta_i = 0$, for i = 1, 2, 3, 4, in which case the corresponding variable does not belong to the stationary linear combination of the variables. All these hypotheses are decisively rejected, with p-values ranging from 0.0000 (real wage) to 0.0005 (rate of unemployment).[52] In this sense we can say that the data strongly suggests cointegration among all of the four variables and not just in the subset of these variables.[53]

4 Summary and discussion

For Finland the data seem to suggest a model, where the capital stock is taken as (weakly) exogenous to the long-run parameters, i.e. the capital stock is not error correcting w.r.t. shocks to the cointegration relationship, which can be regarded as an unemployment relationship. At conventional significance levels at least, formal tests indicate that we could also take real wages and terms of trade as (weakly) exogenous. Since we are primarily interested how the burden of adjustment to shocks to the unemployment relationship is distributed across unemployment, real wages and terms of trade in like Finland and New Zealand, we will focus on the relevant error correction representations of the cointegrated system of variables. In the Finnish case, we have from equation (16)

$$\Delta X_t^{FI} = \begin{pmatrix} \Delta u_t \\ \Delta w_t \\ \Delta k_t \\ \Delta q_t \end{pmatrix} = \begin{pmatrix} -0.34 \\ -0.01 \\ 0 \\ -0.03 \end{pmatrix} (u_{t-1} - 3.32\, w_{t-1} + 1.53\, k_{t-1} + 1.63\, q_{t-1}) + \sum_{j=1}^{3} \hat{\Gamma}_j \Delta X_{t-j}^{FI} + \hat{\varepsilon}_t^{FI} \qquad (16')$$

[52] Approximate (asymptotic) 95 % confidence intervals $\beta_i \pm 2$S.E. in the case of New Zealand are in line with these (more appropriate) tests for the significance of the 'cointegrating coefficients' as they do not include zero. Compare this with case of Finland, where the confidence intervals suggest that a coefficient is insignificantly different from zero, where the more appropriate (LR) test rejects the hypothesis of a zero coefficient. The reason for these conflicting outcomes is explained in fn. 40.

[53] As in the case of Finland, a battery of specification tests on the VAR(2) model was performed. According to the results, the VAR(2) appears to provide a reasonable statistical model for the vector time series. There are no signs of autocorrelation in the vector of estimated residuals, nor of heteroscedasticity. Evidence against normal errors occurs, however, in the residuals from the terms of trade, unemployment and capital stock equations. The source of this non-normality appears to be skewness to the right, i.e. sizable increases in the terms of trade, rate of unemployment and the capital stock. These large deviations show up also in the graphs of the residuals. Also, from the graphs of the estimated spectral densities, one can see negative autocorrelation in the residuals of the terms of trade and unemployment equations, and positive autocorrelation in those of the real wage and capital stock equations, although the estimated autocorrelation is not significant at 5 % (residuals from the capital stock equation come close to have significant autocorrelation at lags 1–2).

where a zero factor loading has been imposed on the capital stock, k.

(16') suggests that the bulk of the short-run adjustment to a shock to the unemployment relation falls on unemployment. The speed of adjustment of unemployment is also relatively fast, whereas real wages and perhaps also terms of trade adjust only sluggishly; unemployment will adjust at a speed of 0.34 percentage points per 6 months to a unit shock to the long-run unemployment relationship (e.g. a permanent fall in the capital stock), while the adjustment speed of real wages is as low as 0.01 percentage points per 6 months. In the light of evidence from other studies of the flexibility of wages to unemployment, this point estimate may be towards the low end of the range of estimates obtained.[54]

The finding that real wages are only mildly error correcting may be associated with real wage rigidity which, together with slightly stronger error correction on the terms of trade, in turn could be related to the wage formation under centralized wage bargaining and fixed exchange rates with fairly regular devaluations that characterised Finland over much of the period. It may be extremely difficult to achieve real wage adjustments, in the face of adverse shocks to employment, through nominal wage cuts in a ‘corporatist economy’ like Finland.[55] Devaluations of the domestic currency in bad times - trying to counteract a fall in the price level – contribute to making nominal wage cuts unnecessary.[56]

This distribution of the burden of adjustment has important policy implications, because not only does unemployment display substantial hysteresis, which tends to make shocks to unemployment highly persistent, but the low (endogenous) response of wages and the terms of trade to labour market shocks provides a very weak cushion against unemployment increases in the presence of adverse shocks. The low response of real wages contributes to making unemployment increasingly exposed to shocks; i.e. increases the likelihood of

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[54]See e.g. Parjanne (1997). With annual data, real wages appear to fall by 0.1 percentage points after a one percentage point increase in unemployment.

[55]Finland, along with other Nordic countries, has often been considered as an archetype of *social corporatism*, i.e. an economic system whose labour market is characterized by two basic features: i) centralized wage bargaining and ii) formal or informal involvement of the government economic and social policies in the process (Pekkarinen et al. 1992, p. 2; see also Vartiainen 1995). In Finland the wage bargaining process has typically been a two- or three- tier process, with a centralized agreement, an agreement signed by industrial unions and the corresponding employers’ associations and adjustments at the plant or firm level agreed by the workers’ and employers’ representatives. In the successive tiers after the centralized agreement, there is strong bias towards positive wage drift, which has not always been fully anticipated at the central level. Unions tend to improve upon the central agreement, and plant level applications usually involve (options for) positive adjustments to union level wage rates (see also Vartiainen 1995, pp. 4–6). Holden (1991) has formally proven – using Nash bargaining theory – that the existence of collective agreements changes the threat points of wage bargains at the lower level of bargaining in a way that generates positive wage drift. Furthermore this multi-tier scheme tends to work better in an environment with some inflation, since when inflation is low, some nominal wages may need to be cut to achieve a given aggregate outcome.

[56]The implications here is, thus, that real wage flexibility in Finland, if there has been any, has been at least partly the result of monetary and exchange rate policy. There is an empirical content in this claim – more corporatist economies have ‘softer’ exchange rate policies – and it certainly qualifies the interpretation of the Calmfors – Driffil ‘smile’ (Calmfors and Driffil 1988; see also Pohjola 1992). Tabellini’s analysis of discretionary monetary policy equilibria under centralized wage setting could probably provide a formal setting for the analysis of this claim (Tabellini 1988, pp. 105–106). It could be argued that Austria is a clear exception but Austria is also an exception in the sense that it has not performed so well in terms of (changes in) employment. Although its record seems to be good in terms of (changes in) unemployment (see e.g. Pohjola 1992, pp. 51–52, graphs 3.3–3.6).

poor unemployment performance of the economy in the presence of adverse shocks. This, in turn, contributes to the possibility of sharp increases and subsequently low convergence of the unemployment rate. Since the factor loadings depend on a variety of institutional and structural features of the labour market,[57] this implies policy efforts should be directed to those reforms that, in addition to measures that potentially reduce hysteresis in unemployment, redistribute the burden of short-run adjustment away from unemployment. This involves, inter alia, increasing the unemployment responsiveness of real wages in the economy.[58]

In the case of New Zealand, on the other hand, the evidence may be more difficult to interpret from the point of view of labour market adjustment, because uncertainty in the parameter estimates, in particular of the factor loadings, appears to be larger than for Finland. It is also more difficult because the evidence in the data in favour of a long-term unemployment equation seems to be much weaker than for Finland. Furthermore, comparison with the result from the Finnish data is complicated by the fact that the New Zealand data are annual.

Although there is strong evidence in favour of cointegration among the terms of trade, real consumption wages, rate of unemployment and capital stock in the data from New Zealand, with the hypothesis of cointegration only among the subset of variables decisively rejected by the data, results from weak exogeneity tests indicate that we have essentially estimated a long-term 'terms of trade equation' of the form $\beta' X_t = q_t + 0.14u_t + 0.73w_t - 0.37k_t$ with the associated loading matrix $\alpha = (-1.09, -0.13, -0.13, 0)'$ from the data. Hence, the data appear to suggest that the error correction form of the cointegrated system for New Zealand is

$$\Delta X_t^{NZ} = \begin{pmatrix} \Delta q_t \\ \Delta u_t \\ \Delta w_t \\ \Delta k_t \end{pmatrix} = \begin{pmatrix} -1.09 \\ -0.13 \\ -0.13 \\ 0 \end{pmatrix} (q_{t-1} + 0.14\, u_{t-1} + 0.14\, w_{t-1} - 0.37\, k_{t-1}) + \sum_{j=1}^{1} \hat{\Gamma}_j \Delta X_{t-j}^{NZ} + \hat{\varepsilon}_t^{NZ} \tag{18}$$

Formally, the sample may just be too small and, hence, the estimated factor loadings too imprecise for us to be able to infer the correct VECM. However, some observations are warranted on the basis of the estimated VECM in (18). First of all, as noted earlier, the unemployment rate in New Zealand has swung sharply during the last 10–15 years: from 4 % in 1984 to 11 % in 1992 while it is currently around 7 %. Hence, as these figures indicate labour market adjustment

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[57]In the context of a different model, Nickell (1997, p. 2) also emphasizes the dependence of the model's parameters on the institutional features of the labour market.

[58]Yet another possibility for the non-responsiveness of real wages to unemployment variations may involve aspects from human capital development during unemployment spells. The average ability of unemployed workers fall during unemployment, which, among other things, tends to increase the mismatch between vacancies and unemployed workers. Hence, anything that reduces the effectiveness of the long-term unemployed as fillers of vacancies, such as long periods receiving benefit, will tend to lower the responsiveness (Nickell, 1997, p. 2).

through unemployment can be sizable; whether it is 1.9 percentage points p.a. for a unit shock to the long-run equilibrium, as the unrestricted loading estimates indicate (Table 2.2d), is another matter. Formal tests indicate that, under r = 1 and weakly exogenous capital stock, one can reduce the factor loading of the (growth of the) unemployment rate by as much as 1.5 percentage points without essentially no reduction in the test statistic.

Second, the numerical estimates of the factor loadings in (18) (and in the unrestricted case) indicate larger factor loading on real wages than in Finland. Hence, real wages appear to be more strongly error correcting in New Zealand than in Finland. This means that the rate of convergence of real wages to the long-run equilibrium is more rapid in New Zealand than in Finland.[59]

Third, terms of trade movements appear to be an important adjustment channel to shocks to the long-run equilibrium in the New Zealand economy. As noted in Section 3, formal tests decisively rejected the hypothesis that the terms of trade is weakly exogenous to the parameters of the long-run equilibrium. In the context of a sticky price model, like the one in Section 2, nominal exchange rate movements are perhaps the most important single source of terms of trade movements in an open economy, and this ‘sticky price logic’ may actually explain the signs of the estimated coefficients in the cointegrating vector, with the capital stock capturing important supply side effects on the nominal value of the New Zealand currency.[60] Furthermore, the rate of convergence of the terms of trade to the long-run equilibrium is rapid and certainly faster than in Finland.

We thus see two very different labour markets but both seem to have coherence over the period as a whole. As noted by Chapple, Harris and Silverstone (1996) New Zealand adjusted quite flexibly even before the reform programme of the last fifteen years. Finland on the other hand did not adjust readily even when it had a floating exchange rate. It will therefore need new mechanisms if it is to respond more flexibly under Stage 3 of EMU. However, experience of the last year suggests that it has already been possible to have a faster rate of economic growth consistent with price stability as the new market pressures from membership of the euro area are anticipated. Thus, while it may be too early to identify structural breaks, they may become obvious with time. Centralised bargaining may permit more flexible wage adjustment in monetary union. The lesson from both countries may turn out to be that institutional change outside the labour market can nevertheless have a clear effect on the way the labour market adjusts to external shocks.

[59]One factor that may sustain higher wage responsiveness in New Zealand is the benefit reform of 1991, which, through reduction of benefits, appears to have increased incentives to work and, hence labour market participation (Maloney 1997).

[60]Hansen and Hutchison (1997) find support for such an emphasis on ‘real’ determinants of nominal exchange rates in New Zealand. Their model of the real side of the economy is different from ours.

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Appendix 1

Unit root tests for the log of the unemployment rate u, log of the real consumption wage w, log of the capital stock k and log of the terms of trade q.

A) Finland

Table 1.1 **ADF unit root tests for the levels u, w, k and q, 1962.1–1996.2**

Asymptotic critical values: 5 % = –3.476, 1 % = –4.097;
Constant and Trend included

Variable	Lag	t-adf	Beta (t–1)	Sigma	t–Δ(lag)	t-prob	F-prob
u	4	–2.586	0.855	0.127	–0.581	0.563	
	3	–3.099	0.842	0.126	–0.859	0.394	0.563
	2	–3.962 *	0.821	0.126	1.287	0.203	0.589
	1	–3.760 *	0.847	0.127	6.922	0.000	0.446
w	4	–2.693	0.911	0.014	1.732	0.088	
	3	–2.441	0.919	0.014	–0.259	0.796	0.088
	2	–2.552	0.917	0.014	3.015	0.004	0.224
	1	–2.078	0.929	0.015	0.151	0.881	0.010
q	4	–2.207	0.866	0.030	0.815	0.418	
	3	–2.071	0.879	0.030	1.689	0.096	0.418
	2	–1.705	0.902	0.031	–1.247	0.217	0.182
	1	–2.146	0.881	0.031	1.528	0.132	0.177
k	6	1.119	1.008	0.003	–1.180	0.243	
	5	0.918	1.006	0.003	–0.329	0.743	0.243
	4	0.874	1.006	0.003	–1.102	0.275	0.478
	3	0.584	1.004	0.003	–2.494	0.015	0.446
	2	–0.131	0.999	0.003	–0.185	0.854	0.078
	1	–0.209	0.998	0.003	11.910	0.000	0.130

Note: The test equation is $x_t = \alpha + \mu t + \rho_1 x_{t-1} + \rho_2 x_{t-2} + \ldots + \rho_p x_{t-p}$ or $\Delta x_t = \alpha + \mu t + \beta x_{t-1} + \Sigma\gamma_i \Delta x_{t-i} + \varepsilon_t$, where we sum from 1 to p–1 and where $\beta = \rho_1 + \rho_2 + \ldots \rho_p - 1$; t–adf = t-value on the lagged level, t_β; Beta (t–1) = sum of the estimated AR-coefficient, $\hat{\rho}_1 + \hat{\rho}_2 + \ldots + \hat{\rho}_p$; sigma = standard error of regression; t–Δ(lag) = t–value of the longest lag, $t_{\gamma j}$; t-prob = significance of the longest lag: $1 - P(|\tau| \leq |t_{\gamma j}|)$; F-prob = significance level of the F-test on the lags dropped up to that point; * significant at 5 %.

Table 1.2 **ADF unit root tests for the first differences Δu, Δw, Δk and Δq 1962.2–1996.2**

Asymptotic critical values: 5 % = **–2.904**, 1 % = **–3.528**;
Constant included (and Trend for CSTOCK)

Variable	Lag	t-adf	Beta (t–1)	Sigma	t–Δ(lag)	t-prob	F-prob
Δu	4	–4.643**	0.227	0.132	0.821	0.415	
	3	–4.980**	0.299	0.132	1.687	0.097	0.415
	2	–4.726**	0.421	0.134	2.447	0.017	0.182
	1	–3.916**	0.553	0.134	0.444	0.659	0.029
Δw	4	–2.825	0.423	0.015	0.837	0.406	
	3	–2.696	0.478	0.015	–1.715	0.091	0.406
	2	–3.752**	0.334	0.015	0.387	0.700	0.172
	1	–4.026**	0.365	0.015	–2.954	0.044	0.296
Δq	4	–3.165*	0.165	0.031	–0.345	0.731	
	3	–3.697**	0.126	0.031	–0.247	0.806	0.731
	2	–4.368*	0.098	0.031	–1,250	0.216	0.915
	1	–6.559**	–0.063	0.031	1.775	0.081	0.640
Δk	6	–2.452	0.781	0.003	–0.260	0.796	
	5	–2.870	0.771	0.003	0.985	0.329	0.796
Crit. Value;	4	–2.701	0.802	0.003	0.135	0.893	0.603
5 % = **–3.478**	3	–2.903	0.806	0.003	0.884	0.380	0.792
1 % = **–4.101**	2	–2.773	0.826	0.003	2.403	0.019	0.773
	1	–2.175	0.863	0.003	0.247	0.806	0.214

Note: See Table 1.1

B) New Zealand

Table 1.3 **ADF unit root tests for the levels u, w, k and q, 1962–1995**

Asymptotic critical values; 5 % = **–3.556**, 1 % = **–4.271**;
Constant and Trend included

Variable	Lag	t-adf	Beta Y(t–1)	Sigma	t–ΔY(lag)	t-prob	F-prob
u	2	–1.962	0.572	0.536	–1.411	0.169	
	1	–3.448	0.387	0.546	2.033	0.052	0.169
w	2	–1.669	0.902	0.028	0.758	0.455	
	1	–1.564	0.914	0.028	2.433	0.022	0.455
q	2	–2.410	0.531	0.084	2.385	0.447	
	1	–3.577*	0.443	0.084	–0.772	0.024	0.447
k	4	–2.133	0.952	0.004	–0.518	0.609	
	3	–2.167	0.952	0.004	–0.602	0.553	0.609
	2	–2.194	0.952	0.004	–2.286	0.031	0.736
	1	–2.514	0.942	0.004	3.693	0.001	0.164

Note: See Table 1.1

Table 1.4 **ADF unit root tests for the first differences Δu, Δw, Δk and Δq 1963–1995**

Asymptotic critical values: 5 % = –2.959, 1 % = –3.657;
Constant included (and Trend of CSTOCK)

Variable	Lag	t-adf	Beta (t–1)	Sigma	t–Δ(lag)	t-prob	F-prob
Δu	2	–3.939**	–0.430	0.577	0.065	0.948	
	1	–6.221**	–0.411	0.567	2.939	0.007	0.948
Δw	2	–2.843	0.393	0.029	1.420	0.167	
	1	–2.430	0.522	0.030	–0.579	0.567	0.167
Δq	2	–4.021**	–0.294	0.091	0.110	0.913	
	1	–5.787**	–0.269	0.090	2.532	0.017	0.913
Δk	4	–2.571	0.053	0.004	0.930	0.362	
Crit. Value;	3	–2.492	0.248	0.004	0.447	0.659	0.362
5 % = **–3.567**	2	–2.665	0.316	0.004	0.545	0.590	0.595
1 % = **–4.295**	1	–2.884	0.394	0.004	2.554	0.017	0.720

Note: See Table 1.1

Appendix 2

Cointegration analysis of the system consisting of the log of the unemployment rate u, log of the real consumption wage w, log of the capital stock k and log of the terms of trade q

A) Finland

Table 2.1 **Cointegration rank of the system 1962.2–1996.2**

(a) Eigenvalues

Eigenvalue	Loglik. for rank	
	1096.50	0
0.3636	1112.09	1
0.1613	1118.16	2
0.1014	1121.85	3
0.0028	1121.94	4

(b) Test statistics

H0: rank=r	Max λ	T–nm	CV 95 %	Trace	T–nm	CV 95 %
R = 0	31.18	27.57	27.1	50.89	45.00	47.2
R ≤ 1	*	*	21.0	*	17.42	29.7
R ≤ 2	12.14	10.73	14.1	19.71	6.69	15.4
R ≤ 3	7.38	6.52	3.8	7.57	0.17	3.8
	0.19	0.17		0.19		

Notes: Max λ = maximal eigenvalue test for the rank. T–4k corrects the tests for small sample bias, i.e. uses T–4k instead of T. CV 95 % = 95 % Critical value. Trace = trace test for the rank. T–nm corrects the test for small sample bias, where n = dimension of the VAR and m = log length of the VAR.

(c) Standardized β (eigenvectors)

u	w	k	q
1.000	**–3.322**	**1.526**	**1.626**
–0.032	1.000	–0.754	–0.878
0.470	–9.990	1.000	18.899
–0.158	–7.625	4.723	1.000

(d) Standardized α-coefficients

u	–0.341
w	–0.011
k	–0.001
q	–0.027

B) New Zealand

Table 2.2 **Cointegration rank of the system 1960–1995**

(a) Eigenvalues

Eigenvalue	Loglik. for rank	
	408.564	0
0.6851	427.632	1
0.4319	436.961	2
0.2107	440.865	3
0.0093	441.019	4

(b) Test statistics

H0: rank=r	Max λ	T–4k	CV 95 %	Trace	T–nm	CV 95 %
R = 0	31.18 *	28.89 *	27.1	64.91	49.17*	47.2
R ≤ 1	18.66	14.14	21.0		20.28	29.7
R ≤ 2	7.81	5.91	14.1	**	6.15	15.4
R ≤ 3	0.31	0.23	3.8	26.77	0.23	3.8
				8.12		
				0.31		

Notes: Max λ = maximal eigenvalue test for the rank. T–4k corrects the tests for small sample bias, i.e. uses T–4k instead of T. CV 95 % = 95 % Critical value. Trace = trace test for the rank. T–nm corrects the test for small sample bias. * = significant at 5 %. ** = significant at 1 %.

(c) Standardized β (eigenvectors)

q	u	w	k
1.000	**0.185**	**0.745**	**–0.490**
–2.224	1.000	1.077	–2.929
–0.222	–0.090	1.000	1.963
0.716	0.005	–2.216	1.000

(d) Standardized α-coefficients

u	–0.596
w	–1.880
k	–0.059
q	–0.001