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Mikael Juselius
Swedish School of Economics and Business Administration
P.O.Box 479
00101 Helsinki, Finland

Distributor:

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The relationship between wages and the capital-labor ratio: Finnish evidence on the production function

Mikael Juselius*†
Swedish School of Economics & BA, Department of Economics

Abstract

In this paper, a positive empirical relationship between wages and the capital-labor share is established using Finnish manufacturing data. This relationship is consistent with a modeling approach for the Finnish economy that assumes a CES production function and imperfections in both product and labor markets. The popular Cobb-Douglas production function is inconsistent with the observed relationship. Moreover, the estimations are consistent with an elasticity of substitution above one. The results are further strengthened by a positive relationship between unemployment and the capital-labor share through its effect on the wage rate. The estimations also provide insights into the processes determining the output-labor ratio, capital-labor ratio and investments.

Key Words: Cointegration, Production Function, VAR-Model, Labor Markets.

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*Address: University of Helsinki, Department of Economics, P.O. Box 17, FIN - 00014 Helsinki. E-mail: mikael.juselius@helsinki.fi.

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1 Introduction

Lately, there has been a renewed interest in the functional form of the aggregate production function. This development has partly been due to the empirical observation that factor shares have not remained constant (see for example Bentolila and Saint-Paul (2003)) casting doubts on the validity of the popular Cobb-Douglas specification[1]. More importantly, the revival of growth theory during the last decade (see Quah (1996) and Solow (1994)) makes the issue topical since different functional specifications can lead to different and richer theoretical results. For instance, assuming a more general CES (Constant Elasticity of Substitution) technology allows for endogenous growth (see for instance Jones and Manuelli (1990)) for some parameter values and multiple stable equilibrium configurations with respect to the labor share for others (see Azariadis (1996)). Also, in a recent paper Kauppi et al. (2004) establish a relationship between wages and the capital-labor share by assuming a CES production function coupled with imperfections in both labor and product markets. This relationship indirectly determines equilibrium unemployment through its effect on wage formation.

Most CES based results hinges on the parameter value of the elasticity of substitution[2]. Attempts at estimating this elasticity for the CES specification have been the focus of some previous empirical studies. The seminal contribution in this respect is Arrow et al. (1961), who derive the CES function and suggest one way of estimating the elasticity of substitution under the additional assumption of perfect competition. A more recent contribution is provided by Duffy and Papageorgiou (2000), who uses panel data on 82 countries over 28 years to estimate a CES specification. They find that a CES function with an elasticity of substitution above one fits the data quite well for the most developed countries in their sample (including Finland). They do not, however, address the time series problems of the data[3]. A Finnish study is provided by Ripatti and Vilmunen (2001), who estimates a CES specification directly using non-linear cointegration techniques. They find an elasticity of substitution well below one, but are forced to make some strong assumptions on the processes of the unobservables in order to obtain their results[4]. Furthermore, their results are derived under the assumption of perfectly competitive capital and labor markets (although they allow for monopolistic competition in the product market).

In this paper I suggest an indirect way of making empirical inference on the

[1]See Mankiw et al. (1992) for an argument in favor of the Cobb-Douglas specification.

[2]The Cobb-Douglas functional form is a special case of the CES function corresponding to the case where the elasticity of substitution is equal to one, as is well known. Whether the elasticity of substitution is above or below unity is of special theoretical interest.

[3]Another problem with using a panel of countries is that it requires strong assumptions on the similarities of the countries in sample.

[4]As a consequence they implicitly restrict the number of common trends and hence the number of cointegration vectors.

form of the production function by utilizing the theoretical relationship between wages and the capital-labor share derived by Kauppi et al. (2004)[5]. They show that the nature of this relationship depends crucially on the elasticity of substitution and the degree of product market competition (captured by the elasticity of substitution between products). Furthermore, it vanishes in the Cobb-Douglas case, i.e in the case when the elasticity of substitution is one. In particular, the relationship between wages and the capital-labor share is positive when the elasticity of substitution is above one and there is sufficient product market competition. A negative relationship on the other hand arises when the elasticity of substitution is below one. The relationship is ambiguous when the elasticity of substitution is above one and the degree of product market competition is low. Thus, if it is possible to establish a long-run[6] relationship between wages and the capital-labor share empirically it would be consistent with the CES production function (and inconsistent with a Cobb-Douglas technology). Moreover, if this relationship is positive it would be consistent with an elasticity of substitution above one.

The statistical workhorse utilized in this paper is the cointegrated VAR (Vector Auto-Regression) model (see Johansen (1995)). This model is particularly well suited since it does not require assumptions on causality and it allows for non-stationarity in the data and hence for distinguishing between long-run economic relationships and (i.e. cointegration) and short-run adjustments (see Engle and Granger (1987)). The advantage of the present approach is that it allows one to model the main properties of the data[7], thus avoiding some of the more unrealistic assumptions, while still retaining its relevance to economic theory. Furthermore, the estimates provide insights into the processes that determine the variables at hand, which is of interest in itself.

The cointegrated VAR model is estimated on quarterly Finnish manufacturing data (using an information set that captures the main features of the model in Kauppi et al. (2004)) over the years 1980:1-2001:4. I find a strong positive long-run relationship between manufacturing wages and the capital-labor ratio, determining the change in wages. This relationship also enter the determination of unemployment in a positive way. These results have two implications. First, the Cobb-Douglas production function specification seems to be inappropriate for the case of Finland. As an alternative, the CES specification coupled with

[5]Bentolila and Saint-Paul (2003) uses similar indirect arguments to make inference on the parameter value of the elasticity of substitution based on a fitted equation for the labor share.

[6]The relationship in Kauppi et al. (2004) is derived under product market clearing and should be viewed as describing the "long-run" in some sense. Hence, the empirical relationship should have the same property.

[7]For example, the time series properties of the data can be handled in a convincing way. Furthermore, economic theory usually describe long-run (equilibrium) behavior but is silent about the short-run. However, in empirical applications short-run behavior also needs to be modeled in order to make valid inference on the long-run.

imperfections in both product and labor markets is more consistent with the observed behavior of the data. Second, the positivity of this relationship is broadly consistent with an elasticity of substitution above one. The fact that the relationship enter unemployment positively strengthens this conclusion. Estimates of the processes determining the changes in the output-labor share, capital-labor share, investments and inflation are also obtained as byproducts of the analysis.

The paper is organized in the following manner. In section 2 some properties of the CES aggregate production function and the most relevant results from Kauppi et al. (2004) are discussed. Section 3 provides a short summary of the main properties of the cointegrated VAR model. Section 4 is concerned with the statistical analysis of the data. The empirical results are then discussed in section 5 and section 6 finally concludes.

2 The CES production function and the relationship between wages and the capital-labor share

The theoretical relationship between wages and the capital-labor share, as established in Kauppi et al. (2004), and its relation to the empirical problem at hand is discussed in this section. This relationship is dependent on the assumption of a CES production function (and indeed on imperfections in both product and labor markets) and, in particular, on the value of the elasticity of substitution. This key feature of the model is utilized in this paper to make empirical inference on the functional form of the production function. We begin by discussing the CES production function and then turn to the results from Kauppi et al. (2004).

2.1 The CES production function

The concept of an aggregate production function is often a useful (theoretical) point of departure. Although this concept is somewhat problematic[8], most authors assume that the economy's production technology is represented by some (two factor) production function like

$$Y = AF(K,\, L) \tag{1}$$

where Y is output, K is capital, L is labor and A captures technological progress. The function F is usually assumed to be twice continuously differentiable and to satisfy INADA conditions. Since its introduction by Solow (1956), the undoubtedly most popular choice has been the Cobb-Douglas production function

$$Y = AK^{\beta}L^{1-\beta} \tag{2}$$

[8]The literature on aggregation suggests that the assumption of an aggregate production function might be hard to justify. For a nice overview see Felipe and Fisher (2003).

where $0 < \beta < 1$. The popularity of the Cobb-Douglas specification is probably mostly due to its analytical convenience rather than its realistic predictions[9]. A considerably more general alternative is provided by the CES production function

$$Y = \gamma \left[\beta K^{\frac{\sigma-1}{\sigma}} + (1-\beta) L^{\frac{\sigma-1}{\sigma}} \right]^{\frac{\sigma}{\sigma-1}} \tag{3}$$

where σ is the elasticity of substitution and γ plays a similar role as A in (1) and β captures the functional distribution of income. (3) embeds the popular Cobb-Douglas form as a special case when $\sigma = 1$. As noted in the introduction this functional form allows for much richer theoretical results, for example in endogenous growth contexts. Naturally, many other functional forms would do as production functions as well. However, as the theoretical literature has predominately focused the Cobb-Douglas or the CES function the same focus is maintained in this paper.

2.2 The relationship between wages and the capital-labor share

Kauppi et al. (2004) analyze equilibrium wage formation, unemployment and investments under the assumptions of product market imperfections a lá Dixit and Stiglitz (1977) and a 'right-to-manage' union bargaining approach to labor market imperfections. They generalize previous contributions by Blanchard and Giavazzi (2003) and Spector (2004) by assuming a CES production function which also allow them to study the role of investments. Under these assumptions they establish the following relationship between wages and the capital-labor share (see Kauppi et al. (2004))

$$\frac{\partial w^N}{\partial k} \left\{ \begin{array}{c} < \\ = \\ > \\ ? \end{array} \right\} 0 \; as \left\{ \begin{array}{c} \sigma < 1 \\ \sigma = 1 \\ 1 < \sigma < s \\ 1 < s < \sigma \end{array} \right\} \tag{4}$$

where w^N is the Nash bargaining solution of the wage rate, k is the capital-labor share (defined as K/L), σ is the elasticity of substitution from the CES production function and $s > 1$ is the elasticity of substitution between products in a CES-utility function. Thus, s can be viewed as a measure of the degree of product market competition. As can be seen from (4), the relationship between wages and the capital-labor share is negative if the the elasticity of substitution is less than one (i.e. gross complementary). In the Cobb-Douglas case, where

[9] The Cobb-Douglas functional form is usually motivated by the Kaldor's "stylized facts", for instance that factor shares have been largely constant (in fact, the Cobb-Douglas production function implies *absolutely* constant factor shares). However, this is clearly not the case for many economies as demonstrated by Bentolila and Saint-Paul (2003).

$\sigma = 1$ there is no relation between wages and the capital-labor share. If on the other hand the elasticity of substitution is above one (gross substitutability) but less than the elasticity of product demand, the relationship is positive. In the last case where $1 < s < \sigma$ the sign is ambiguous and depends on the relative bargaining power of the labor union (see Kauppi et al. (2004)). Through its effect on wages, the effects from the capital-labor share on unemployment are similar to that of (4).

This framework provides a realistic setting in which to make inferences on the functional form of the production function for the Finnish economy. The Finnish economy is indisputably best characterized by labor and product market imperfections (see for example Nickell (1997)). The empirical problem is thus to establish a long-run relationship between wages and the capital-labor share (the lack of such a relationship would lend support to a Cobb-Douglas specification). The nature of such a relationship enables us to make predictions on the region of the parameter values.

3 The statistical model

The baseline statistical model is the p-dimensional cointegrated VAR-model with k lags

$$\Delta X_t = \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Pi X_{t-1} + \mu + \Psi D_t + \varepsilon_t \tag{5}$$

where the vector process X_t is assumed to be $I(1)$, μ is a vector of constants, D_t consist of the other deterministic components and $\varepsilon_t \sim N_p(o, \Sigma)$. If there exists cointegration $\Pi = \alpha\beta'$ and $\mu = \alpha\beta_0 + \alpha_\perp\gamma_0$, where α and β are two $(p \times r)$ matrixes such that $r < p$, $\alpha_\perp$ and $\beta_\perp$ are two $p \times (p - r)$ matrixes such that $\alpha' \alpha_\perp = 0 = \beta' \beta_\perp$, then the moving average representation of the model is given by

$$X_t = B + C(\sum_{i=1}^{t} (\varepsilon_i + \Psi D_i) + \mu t) + C^*(L)(\varepsilon_t + \Psi D_t) \tag{6}$$

where B depends on the initial values and $C = \beta_\perp(\alpha_\perp' (I - \sum_{i=1}^{k-1} \Gamma_i)\beta_\perp)^{-1}\alpha_\perp$, and the polynomial $C^*(z)$ is convergent for some $z < 1 + \delta$ for some $\delta > 0$, Johansen (1995). The main properties of the model are investigated in Johansen (1995) and will not be repeated here[10]. However, for the sake of clarity some of the most crucial results are reproduced.

3.1 Linear restrictions on the β-vectors

An important part of the analysis consists of testing linear restrictions on the β-vectors with the aim of identifying a structure of empirically relevant relations.

[10]See Hendry and Juselius (1999, 2000) for a helpful overview.

The restrictions can either be applied to all the cointegrating relations simultaneously or to the individual vectors separately. Given $\Pi = \alpha\beta^{'}$ Johansen and Juselius (1990) show that the same restriction on all the vectors can be tested by formulating the alternative hypothesis

$$\mathcal{H}_2 : \Pi = \alpha\varphi' H' \tag{7}$$

that is $\beta = H\varphi$ where H is a $(p \times s)$ with $r \leq s \leq p$ matrix defining linear restrictions on β. Restrictions on the individual vectors are formulated as

$$\mathcal{H}_3 : \begin{array}{rcl} \Pi & = & \alpha\beta' \\ \beta & = & \{H_1\varphi_1, \ldots, H_r\varphi_r\} \end{array} \tag{8}$$

where H_i is $p \times s_i$, defining linear restrictions on the individual vectors. Johansen (1995) derives the LR-tests for testing the hypothesis in the form (7) and (8).

3.2 Linear restrictions on the α-matrix and weak exogeneity

The α-vectors can also be restricted in a similar fashion as in the last section. Of special interest is the case where one or several rows in α consist of zeros. A variable with a zero row in α is not affected by the long-run relationships and is hence treated as weakly exogenous. In this case one estimates

$$\Delta X_{1,t} = A\Delta X_{2,t} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \widetilde{\alpha}\beta' X_{t-1} + \mu + \psi D_t + \varepsilon_t \tag{9}$$

where $X_{1,t}$ consist of the endogenous variables while $X_{2,t}$ consist of the weakly exogenous variables and $X_t = \{X_{1,t}, X_{2,t}\}$. The dimension is now $p - h$, where h is the number of exogenous variables, Hendry and Juselius (2000).

4 Empirical analysis

The data used in the analysis is Finnish quarterly manufacturing data spanning over the years 1980:1-2001:4 (88 observations). In accordance with Kauppi et al. (2004), the process determining Finnish manufacturing production, wages and unemployment is assumed to be described by the information set $I_t = \{y_t, k_t, i_t, rw_t, u_t, \Delta p_t, \pi_t, r_t\}$ where y_t is (the log of) the output labor share ($y_t = log(Y_t/L_t)$ where Y_t is real manufacturing output and L_t is hours worked in manufacturing), k_t is the capital labor share ($k_t = log(K_t/L_t)$ where K_t is the real capital stock), $i_t = \Delta log(K_t)$ is approximately investments[11], rw_t is real

[11] i_t is included in the information set because the transforms y_t and k_t correspond to a homogeneous relation with one restriction on the (logs of the) variables Y_t, K_t and L_t. Hence, the difference of one of the variables (with the natural choice of $\Delta log(K_t)$) must be included in in the information set, Juselius (1999). Δp_t is included by similar reasoning.

manufacturing wages ($\log(W_t/P_t)$, where W_t is a index of manufacturing wages and P_t is the consumer price index), u_t is the (log of) manufacturing unemployment rate, $\Delta p_t = \Delta log(P_t)$ is inflation, $\pi_t = log(P_t/P_t^p)$ is the spread between (the logs of) consumer prices and manufacturing producer prices and r_t is the yield on government 10 year bonds. The constructed variable π_t is used as an approximation to the tax wedge and, to some extent, product market imperfections[12]. The analysis of the long-run relations was conducted using *CATS* in *RATS* while the short-run structure was estimated using *PC-Fiml*.

4.1 Dynamic long-run relations

The long-run properties (i.e. cointegration relations) of the data were investigated by estimating model (5) with $X_t = [y_t, k_t, i_t, rw_t, u_t, \Delta p_t, \pi_t, r_t]$, $k = 2$ and a linear trend restricted in the cointegration space. The reduced rank hypothesis (trace test, see Johansen (1995)) indicated that the rank should be set at 4. However, initial tests for weak exclusion indicated that r_t could be excluded[13] (with LR test value of 8.53 against $\chi^2_{0.95}(4) = 9.49$). The model (5) was hence re-estimated with $X_t = [y_t, k_t, i_t, rw_t, u_t, \Delta p_t, \pi_t]$, $k = 2$ and a linear trend in the cointegration space. The reduced rank hypothesis is reported in table 1 and indicated that $r = 3$ (thus excluding r_t allows for one cointegrating relation less).

$H_o : r$	Trace	Trace(90)	λ_i
0	220.11	141.31	0.56
1	150.19	110.00	0.50
2	90.75	82.68	0.37
3	**51.64**	58.96	0.26
4	**25.92**	39.08	0.15

Table 1: The reduced rank hypothesis.

Given $r = 3$, the variables were tested for stationarity, weak exclusion and weak exogeneity. Stationarity and weak exclusion were rejected in all variables while, π_t, the tax wedge was found to be weakly exogenous (with LR value 5.05 against $\chi^2_{0.95}(3) = 7.81$) which is not surprising[14].

[12]For example, the variable is also likely to capture changes in the mark-up on prices over marginal costs. In the reminder of this paper, π, will be loosely referred to as the 'tax wedge'.

[13]In fact, r_t was only important in the equation determining y_t. Keeping it in the analysis does not alter the interesting results reported below in any significant way, apart from inducing more constant beta parameters (see below). The advantages of excluding r_t at this stage is that it allows for simpler interpretations of the cointegrating vectors.

[14]The interest rate, r, which was excluded from the model was also found to be weakly exogenous if allowed to remain in the analysis.

Moving to a partial system, model (9) was finally estimated with $X_{1,t} = [y_t,\, k_t,\, i_t,\, rw_t,\, u_t,\, \Delta p_t]$, $X_{2,t} = [\pi_t]$, $k = 2$ and a linear trend in the cointegration space. Although the trace test is not appropriate for a partial system, the roots of the companion matrix supported the previous choice of rank ($r = 3$). Thus the rank was set at 3. The normalized cointegrating vectors with weights corresponding to this choice of rank are reported in table 2. The estimated vectors in

Table 2: The unrestricted estimates of the $\beta-$ and $\alpha-$vectors given $r = 3$.

	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$	$\hat{\alpha}_1$	$\hat{\alpha}_2$	$\hat{\alpha}_3$
y	1	1.42	0.03	-0.64	-0.07	-0.80
k	0.03	1	0.02	-0.68	-0.01	0.42
i	2.77	83.16	1	0.01	-0.01	-0.09
rw	0.35	-2.90	0.26	0.07	0.01	-0.26
u	-0.01	0.50	-0.01	-1.37	-0.15	3.20
Δp	-4.58	35.4	-1.19	0.06	0.01	0.20
π	0.02	1.83	0.23	–	–	–
$trend$	-0.02	-0.01	-0.01	–	–	–

table 2 span the cointegration space, within which linear restrictions of the form (8) can be tested. Naturally, the null hypothesis is that a particular relation is in the cointegration space. Table 3 reports the results of testing three particular hypotheses on the cointegration space (the relation in $\mathcal{H}_3$ is just identifying so no testing is involved. However since it is used in the subsequent analysis it is reported here for convenience). In table 3 c_{ij} are positive constants correspond-

Table 3: Structural hypothesis on the estimated β-vectors.

Hypothesis	y	k	i	rw	u	Δp	π	$trend$	LR, $\chi^2(df)$	p-value
$\mathcal{H}_1$	$-c_{11}$	$-c_{12}$	0	1	0	0	c_{13}	0	4.44(2)	**0.11**
$\mathcal{H}_2$	1	0	0	c_{21}	$-c_{22}$	$-c_{23}$	0	$-c_{23}$	0.00(1)	**0.99**
$\mathcal{H}_3$	0	c_{31}	1	0	c_{32}	c_{33}	c_{34}	$-c_{35}$	–	–
$\mathcal{H}_4$	1	0	0	$-c_{41}$	0	0	0	$-c_{42}$	10.71(3)	0.01

ing to the $j : th$ unrestricted estimate of the parameter in the $i : th$ relation (the signs are reported for clarity and stem form the estimates). Each vector is normalized on variables that are error correcting[15] in that relation and is significant (based on the t-values of the α-vectors in table 2. See also table 6 and

[15]A variable is error correcting if the linear combination implied by a cointegrating vector

equation (11)). The relation tested in the first hypothesis, $\mathcal{H}_1$, describes an error correcting mechanism for manufacturing wages. The stationarity of this relation cannot be rejected (p-value 0.11, i.e. the relation belongs to the cointegration space). This relation turns out to be important for making inferences about the functional form of the production function and the magnitude of its parameters. Note that k enters this relation with a negative sign which implies a positive relation between Δw and k (see equation 4 in table 6). The relation in the second hypothesis, $\mathcal{H}_2$, is error correcting in the output labor share. The natural interpretation of this relation is that of business cycle effects on the output labor share. For instance higher wages would *ceteris paribus* lower y while higher inflation leads to higher y. Clearly $\mathcal{H}_2$ cannot be rejected. The third relation describes similar effects for investments where it is error correcting (see table 6 equation 3). This relation is just identifying so no testing is involved. Finally the relation in the last hypothesis, $\mathcal{H}_4$, is of interest since the c_{41} corresponds to a direct estimate of the elasticity of substitution, σ, under the additional assumption of perfect competition[16] (which is hardly the case for Finland). The null hypothesis that this relation is in the cointegration space is rejected at 5% significance. The estimated cointegration vectors are reported in table 4 and figure 1 shows the centered linear combinations of the variables implied by the vectors, labeled *ecm*1, *ecm*2 and *ecm*3. Note that the vectors in table 4 span an

Table 4: Identified β-vectors.

	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$
y	-0.58	1	0
k	-0.02	0	0.01
i	0	0	1
rw	1	0.39	0
u	0	-0.03	0.01
Δp	0	-6.47	0.4
π	0.55	0	0.04
$trend$	0	-0.02	-0.00

identified system. Finally, parameter constancy of the vectors were tested by a recursive test described in Hansen and Johansen (1993). The results are reported

contains the variable with a positive (negative) sign and enters the equation of the variable with a negative (positive) sign. For example, if $y - \beta x$ is a linear combination and $\Delta y_t = -\alpha(y - \beta x)_{t-1}$ then y is error correcting in $y - \beta x$.

[16] Arrow et al. (1961) show that b in the regression $y_t = a + bw_t + \epsilon_t$ provides an estimate of the elasticity of substitution when labor and product markets are competitive. Taking the time series properties of the data into account, this would correspond to testing the stationarity of the relation in $\mathcal{H}_4$.

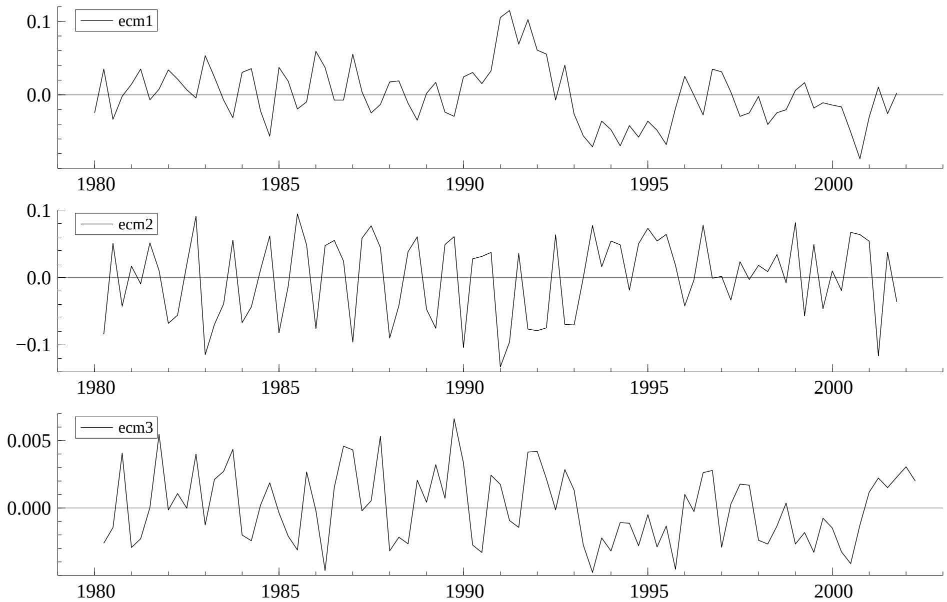

Figure 1: Centered and normalized cointegration vectors.

in figure 2 and indicate a possible structural break in the beginning of the 1990's.

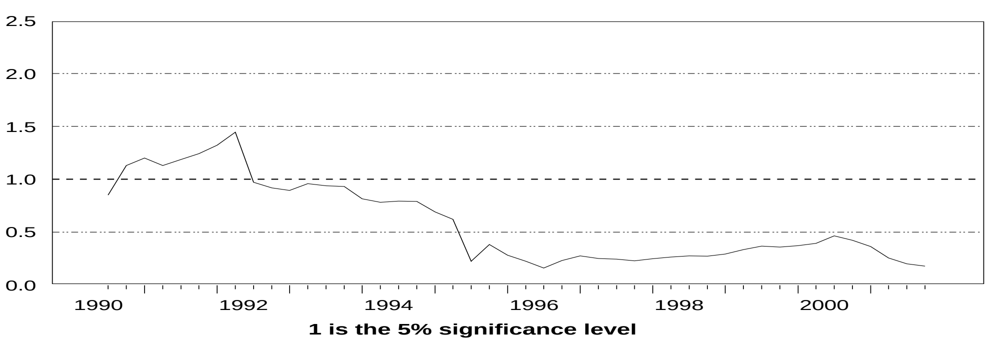

Figure 2: Test of beta constancy over the period 1985:3-2001:4. The dashed line at 1 indicates the 5% significance level.

However, if r is included in the information set the problem disappears yielding constant parameters. As noted earlier the main results are not altered by the inclusion of r.

4.2 The short-run structure of the model

The short run structure of the model was estimated in PC-fiml with the long-run relationships given in the previous section. Labeling the centered and normalized linear combinations of the variables implied by $\hat{\beta}_1$, $\hat{\beta}_2$ and $\hat{\beta}_3$ as $ecm1$, $ecm2$ and $ecm3$ respectively, the short-run structure was estimated by

$$\Delta X_{1,t} = \Gamma_1 \Delta X_t + \Gamma_2 \Delta X_{t-1} + \alpha ECM_{t-1} + \mu + \Psi D_t + \epsilon_t \tag{10}$$

where Γ_1 is a (6×1) vector, Γ_2 is a (6×7) matrix and α is a (6×3) matrix, ECM_t is a column vector with $ecm_{i,t}$ as elements. Initial testing of zero restriction on the parameters of the complete system (by F-tests) indicated that u_{t-1}, Δp_{t-1} and π_{t-1} could be excluded (p-values 0.38, 0.07 and 0.91 respectively). The system was then re-estimated with these changes. The results from misspecification tests on the estimated residuals from the equations are reported in table 5 (the numbers are p-values). By and large, the model seems to fit the data quite well, although there might be some small problems with autocorrelation in the y_t, k_t equations[17]. The rejection of normality in the unemployment series is due to one large outlier (this could be modeled by a dummy variable, but it adds little to the analysis).

Table 5: Misspecification tests on the estimated residuals from each equation. The tests are described in Doornik and Hendry (2001). The normality test is derived under the null of normality. The null in the ARCH test is no conditional heteroscedasticity and the null in the AR 1-5 test is no autocorrelation in the first 5 lags. Bold values indicate rejection at 5% significance.

Equ.	"normality χ^2"	ARCH	AR 1-5	correlation(obs/est)
Δy	0.80	0.59	**0.01**	0.80
Δk	0.50	0.25	**0.03**	0.96
Δi	0.36	0.12	0.86	0.90
Δrw	0.12	0.86	0.28	0.76
Δu	**0.02**	0.75	0.34	0.74
$\Delta^2 p$	0.99	0.48	0.98	0.86

The result from imposing zero restrictions on the individual equations are reported in table 6. Of special interest are the equations determining the changes in the real wages and unemployment (equations 4 & 5 in table 6) as noted in section 2. These equations are

$$\Delta w_t = -0.1\Delta y_{t-1} - 1.42\Delta i_{t-1} - 0.19\Delta w_{t-1} - 0.13(w - 0.58y - 0.02k + 0.55\pi)_{t-1} + \phi_1 D_t + e_t \tag{11}$$

[17]Also this problem disappears when r_t is included in the information set.

Table 6: The short run structure of the model. Bold values of the coefficients to the deterministic components (lowest part of the table) indicate significance at the 5% level.

	Equ1 Δy_t	Equ2 Δk_t	Equ3 Δi	Equ4 Δrw_t	Equ5 Δu_t	Equ6 $\Delta^2 p_t$
Δy_{t-1}	–	–	–	-0.10	–	–
			–	(-4.38)		
Δk_{t-1}	–	-0.19	0.02	–	–	–
		(-2.45)	(3.69)			
Δi_{t-1}	–	–	–	-1.42	–	–
				(-4.08)		
Δrw_{t-1}	–	1.09	-0.09	-0.19	–	–
		(3.89)	(-4.30)	(-2.27)		
$\Delta \pi_{t-1}$	–		-0.04	–	–	–
			(-4.38)			
$ecm1_{t-1}$	–	–	–	-0.13	1.24	0.06
				(-4.53)	(3.34)	(4.05)
$ecm2_{t-1}$	-0.70	-0.51	-0.01	–	–	0.09
	(-9.50)	(-6.97)	(-2.53)			(9.36)
$ecm3_{t-1}$	-7.22	–	-0.67	–	–	–
	(-5.39)		(-7.59)			
$Constant$	**0.014**	**0.02**	0.0004	**0.008**	0.008	-0.000
$Season_t$	-0.002	**0.12**	-0.002	-0.000	**-0.11**	**0.006**
$Season_{t-1}$	**-0.012**	**0.09**	**-0.003**	-0.008	**-0.37**	**0.007**
$Season_{t-2}$	**-0.007**	**0.23**	0.0002	**-0.014**	**-0.18**	**-0.004**

and

$$\Delta u_t = 1.24(w - 0.58y - 0.02k + 0.55\pi)_{t-1} + \phi_1 D_t + e_t \tag{12}$$

where the product of the cointegration vector and the variable vector, $ecm1$ is written out in full. The implications of these equations are discussed in the next section.

5 Interpretation of the results

Equation (11) describes the change in real manufacturing wages due to short-run effects (the lagged differenced terms) and imbalances in the long-run relationship (the first cointegration vector). The short-run effects are not that interesting and will not be discussed further here. The first thing to note is that real manufacturing wages are error correcting in $ecm1$ so that the other variables are (loosely) determining the wages in that relation[18]. Since the capital-labor share enter in $ecm1$, the long-run relationship between the real wages and the capital-labor share *is empirically established*. This relationship is consistent with $\sigma \neq 1$ in (3) and therefore inconsistent with the standard Cobb-Douglas production function. Furthermore, the coefficients in equation (11) imply that this relationship is positive. To see this consider a small positive increase in the capital-labor share ($\Delta k_t > 0$) assuming initial equilibrium. This would make $ecm1$ negative by the negative coefficient of k in the first cointegration relationship and hence imply a *positive* change in the real wages (by the negative coefficient to $ecm1$ in equation (11). Thus, in the long-run we would have something similar to

$$\frac{\Delta w}{\Delta k} > 0$$

which in terms of the model by Kauppi et al. (2004) (see equation (4)) either corresponds to the case where $1 < \sigma < s$ or the case where $1 < s < \sigma$. Regardless of which of these are valid, the result is consistent with $\sigma > 1$.

It is also apparent from (11) that the capital-labor share is not the only long-term influence on the wage rate since both the output-labor share (y) and the tax wedge (π) enter $ecm1$ as well. The effect from a change in y is similar to that of k which seems natural. The effect from a change in π on wages is negative by similar reasoning. Thus, an increase in the tax wedge lowers wages and increases unemployment (see equation (12)).

Additional support of the findings above is found in equation (12) that determines the change in unemployment. As noted in section 2 equilibrium unemployment is determined in part by the capital-labor share, due to the effect on unemployment from the wage rate. This is obviously the case in equation (12) due to the positive coefficient on $ecm1$. Since $ecm1$ is normalized on the

[18]Strictly speaking, since w is endogenous this is not correct.

wage rate, a positive change the capital-labor share leads to positive changes in unemployment as wages adjust upwards. However, note that there is also a direct negative response to unemployment from an increase in the capital-labor share since it enters $ecm1$ with a negative sign[19]. Hence, in the long-run we have something similar to

$$\frac{\Delta u}{\Delta k} > 0$$

which is again consistent with $1 < \sigma < s$ or $1 < s < \sigma$ in the paper by Kauppi et al. (2004). The long-run effects from y and π on unemployment are exactly the opposite to those above (note that an increase in π leads to higher unemployment).

Finally, table 6 provides additional information on the processes determining the (changes of) output-labor share (y), the capital-labor share (k), investments (i) and inflation (Δp). The output-labor share (equation 1) is determined by $ecm2$ and $ecm3$, where $ecm2$ is error correcting. The effects from $ecm2$ are intuitive and seem to capture business cycle effects. For instance, assuming initial equilibrium, a positive change in the wage rate reduces the output-labor share (implying that output falls faster than employment). Similarly, positive changes in either inflation or unemployment increase the output-labor share by $ecm2$. However, the effects from unemployment and inflation are ambiguous since these variables also enter $ecm3$ with opposite signs to that of $ecm2$. The capital-labor ratio (equation 2) is in the long-run solely determined by $ecm2$. The effects are similar to those on the output-labor share. For instance, an increase in unemployment will lead to an increased capital-labor share as labor is reduced. The long-run effects determining investments (equation 3) are captured by both $ecm2$ and $ecm3$, where $ecm3$ is error correcting. Again the effects from $ecm2$ are similar to those in equations 1 and 2. The more interesting (and stronger) effects come from $ecm3$. A positive change in the capital-labor share, unemployment, inflation or the tax wedge leads to reduced investments. Finally, inflation is determined by $ecm1$ and $ecm2$, where $ecm2$ is error correcting. Thus, a positive change in the output-labor share, the real wages or the tax wedge leads to increased inflation while a positive change in the capital-labor share or unemployment decreases inflation.

6 Conclusions

In this paper a long-run relationship between real manufacturing wages and the capital-labor share was established by estimating a cointegrated VAR-model on Finnish manufacturing data. The existence of such a relationship has implications on the underlying functional form of the production function as demonstrated in

[19] Unemployment is not error correcting in $ecm1$, so this effect is best viewed as a short-run effect.

Kauppi et al. (2004). In particular, a positive relationship between wages and the capital-labor share was found. This result is consistent with a modeling approach for the Finnish economy that uses a CES production technology with an elasticity of substitution above one coupled with imperfections on both product and labor markets. A positive relationship between unemployment and the capital-labor share was also found, strengthening these conclusions further. Estimates of the processes determining the output-labor share, the capital-labor share, investments and inflation were also obtained. Compared to approaches that attempt to derive direct estimates of 'deep parameters', the advantage of the present approach is that it allows for more data oriented modeling. Furthermore, the estimated processes provide valuable information and are of interest in their own right.

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