

Boug, Pål

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Pål Boug

The Demand for Labour and the Lucas Critique Evidence from Norwegian Manufacturing

Abstract:

This paper uses neoclassical theory as a foundation for modelling labour demand in Norwegian manufacturing. Applying the Johansen (1988,1991) methodology, we obtain a single cointegrating vector between employment, production, relative factor prices, total factor productivity and the stock of real capital. Normalised on employment, the estimated long run elasticities are 1.37 (production), -0.32 (relative factor prices), -0.57 (total factor productivity) and -1.00 (the stock of real capital). Next, we develop a conditional labour demand model that exhibits parameter constancy. In addition to equilibrium correction effects, we find contemporaneous effects of production and relative factor prices. We cannot reject super exogeneity to be present in our labour demand equation. Hence, the evidence on labour demand in Norwegian manufacturing does not lend support to the Lucas critique.

Keywords: Labour demand, cointegration, conditioning, equilibrium correction model, parameter constancy, exogeneity, Lucas critique.

JEL classification: C22, C32, E13, J23.

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Address: Pål Boug, Statistics Norway, Research Department. E-mail: Pal.Boug@ssb.no

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1. Introduction

The structure of the demand for labour is central to many key policy questions. In particular, fluctuations in unemployment are determined largely by factors explaining the demand for labour. Similarly, the effects of any policy that changes relative factor prices depend on the structure of labour demand. Thus, to predict impacts of changes in payroll taxes or employment subsidies, it is crucial to have reliable estimates of underlying parameters. In response, many international studies have been devoted to the estimation of labour demand schedules [see Hamermesh (1986, 1993) for surveys].¹ It is however surprising that the recent surge of research has been quite disinclined to employ well-established advances in time series econometrics. Consequently, previous labour demand studies may be subject to important limitations that are, and should be, a matter of future empirical investigations.

Firstly, labour demand studies often use restrictive dynamic specifications, and fail to test properly for residual misspecification, weak exogeneity and parameter stability. The standard practice is to estimate labour demand models using Engle and Granger's (1987) two-step method, or to estimate conditional models within a single equation analysis. It is well-known that the former approach may lead to biased long run estimates and cointegration tests with low power [cf. Kremers *et al.* (1992)]. Conditional analysis on its own hand is only valid if the conditioning variables are weakly exogenous for all parameters of interest [cf. Engle *et al.* (1983), Hendry and Neale (1988) and Johansen (1992)]. The Johansen (1988, 1991) multivariate cointegration method effectively takes these econometric issues into account. This method is, however, rarely utilised in labour demand studies. Exceptions include the contributions by Risager (1993), Engsted and Haldrup (1994) and Chiarini (1998).

Secondly, and relatedly, the relevance of the Lucas (1976) critique has received deficient attention in labour demand studies [cf. Ericsson and Irons (1995)]. Lucas (1976) argues that regression models may by construction be misspecified, and thus behave poorly in the event of policy regime shifts. For instance, if agents base their decisions on forward-looking expectations and empirical models fail to account for that, those models may mispredict when expectations change. Unstable regression models do not however, preclude the possibility of an underlying constant behavioural function. That is, non-constancies in empirical models may stem from misspecification rather than shifts in the underlying function [cf. e.g. Hendry and Ericsson (1991)]. Hence, when marginal processes are subject to shifts,

¹ Early contributions include Brechling (1965), Ball and St Cyr (1966), Feldstein (1967), Craine (1973) and Sargent (1978). Since then, the applied literature on labour demand is extended in various directions, see for instance Meese (1980), Nickell and Andrews (1983), Symons and Layard (1984), Symons (1985), Jenkinson (1986), Clark *et al.* (1988), Flaig and Steiner (1989) and Barrell *et al.* (1996).

valid conditioning is crucial for parameter constancy. Consequently, it is essential for policy simulations with labour demand equations to examine the legitimacy of conditioning and to investigate general dynamic specifications. Following the exogeneity literature [see e.g. Engle *et al.* (1983)], the Lucas critique is an example of the absence of super exogeneity. Since super exogeneity is a testable property, it is possible to confirm or refute the applicability of the Lucas critique in practice [see Hendry (1988) and Engle and Hendry (1993)]. Nevertheless, most labour demand studies assume that the estimated models are invariant to policy interventions and changes in expectations. Some studies have indeed estimated labour demand schedules assuming rational expectations [see e.g. Sargent (1978), Kennan (1979), Palm and Pfann (1990) and Engsted and Haldrup (1994)]. These studies do not however, check whether competing models suffer from the Lucas critique, a possibility that should be tested rather than assumed.

Previous labour demand studies on Norwegian data may also be subject to the limitations discussed above. For instance, Cappelen *et al.* (1992) estimates a conditional labour demand equation for the mainland economy without testing for weak exogeneity and parameter constancy. Likewise, Bowitz and Cappelen (1994) and Boug (1999) estimate conditional labour demand models for various industries, but fail to test properly for the applicability of the Lucas critique. The contribution by Moene and Nymoen (1991) may to some extent be viewed as an exception. In fact, this study concludes that unemployment is super exogenous in an estimated labour demand model for the manufacturing and construction industries. Compelling evidence against the Lucas critique is however lacking, as super exogeneity of the other conditioning variables that enter the model is not established.

The purpose of the present paper is to conduct a labour demand study that overcomes these limitations. We first outline a conventional model of neoclassical producer behaviour. Applying the Johansen's (1988, 1991) method to Norwegian manufacturing data over the period 1978-96, we then obtain a significant cointegrating vector between employment, production, relative factor prices, total factor productivity and the stock of capital. Tests of weak exogeneity suggest that both labour demand and the stock of capital correct deviations from the estimated cointegrating vector. Normalising a restricted cointegrating vector on employment yields long run elasticities of 1.37 (production), -0.32 (relative factor prices), -0.57 (total factor productivity) and -1.00 (the stock of capital). Next, having established this long run relationship, we develop a conditional labour demand model that passes several tests for residual misspecification, weak exogeneity and parameter stability. Additional to equilibrium correction effects, we find positive effects of growth in production and negative influences of growth in relative prices. No significant short run effects of the stock of capital are detected. Finally, we focus on the relevance of the Lucas critique. To that end, we employ various tests for super exogeneity as proposed by Engle and Hendry (1993). Although our approach is similar in spirit to that in Moene and Nymoen (1991), it differs regarding econometric specification and details in the

testing procedure.² In particular, we pay special attention to the modelling of marginal processes, and we apply variable addition tests for economic variables entering these processes. We cannot reject the hypothesis that production and relative factor prices are super exogenous for the parameters in our labour demand equation. Hence, the evidence in this paper suggests that the Lucas critique may not be relevant to the modelling of labour demand in Norwegian manufacturing.

2. The Economic Framework

Our theoretical setting is that of neoclassical producer behaviour. We assume Norwegian manufacturers to act as cost-minimising firms for given output demand and factor prices. Additionally, we assume the capital stock to be predetermined in order to avoid the well-known problem of measuring the cost of capital.³ The variable factors of production are thus assumed to include labour and material inputs. In line with related studies, we also make the simplifying assumption that labour and the stock of capital both are homogenous goods [see e.g. Jenkinson (1986), Flaig and Steiner (1989) and Barrell *et al.* (1996)].

One may argue that labour demand models derived from cost-minimisation are most suitable in situations where firms are rationed in the output market. A previous study on Norwegian data supports the assumption of cost-minimisation behaviour for mainland industries [see Cappelen *et al.* (1992)]. Furthermore, the possibility that Norwegian firms are constrained in the output market may be justified by business survey data. According to these data, a significant proportion of manufacturing firms reports shortage of sales and orders to be the most prominent cause for production limitations. It seems thus natural to assume production to a large extent to be demand driven. Doing so, we implicitly abstract from profit maximisation, which would imply specifying the pricing behaviour in the output market.

Applying Shepard's Lemma⁴ to the dual cost function associated with variable factors of production, we obtain a *conditional* labour demand function, which generally reads as

$$(1) \quad L_t = f(X_t, K_t, W_t/PH_t),$$

where L_t denotes labour demand (usually expressed in numbers of employees or hours worked), X_t some activity variable expressed in real terms (usually a production or an output measure), K_t the real

² Our theoretical framework and the questions raised in this paper are also quite different from those in the study mentioned.

³ The exogeneity properties of activity, factor prices and the stock of capital are investigated below.

stock of capital, W_t the wage rate and PH_t the price of material inputs. Following standard theory, factor price homogeneity is imposed in (1), so that factor price effects are captured using a relative price ratio defined as W_t/PH_t . The subscript t denotes time. In order to get a tractable model in the empirical analysis, we assume that (1) can be approximated by the following log-linear equation:

$$(2) \quad l_t = \alpha_0 + \alpha_1 \cdot x_t + \alpha_2 \cdot k_t + \alpha_3 \cdot (w-ph)_t + \alpha_4 \cdot \tau + \varepsilon_t,$$

where lower case letters indicate logs. Noticeably, (2) is augmented with a deterministic time trend, τ , as a proxy for total factor productivity, its econometric implementation being discussed below.⁵ In addition, the equation is extended with an error term, ε_t . We note that the elasticities of labour demand with respect to x_t , k_t and $(w-ph)_t$ are given by α_1 , α_2 and α_3 , respectively. The sign of the elasticities is hypothesised as follows: $\alpha_1 > 0$, $\alpha_2 < 0$ and $\alpha_3 < 0$, meaning that labour demand should increase with output and decrease with the stock of capital and the relative factor price ratio, *ceteris paribus*. It also follows from (2) that the inverse of α_1 is the scale elasticity with respect to the variable factors of production, while $|\alpha_2|/\alpha_1$ is the scale elasticity with respect to the stock of capital. Thus, the total scale elasticity is given by the formula $(1+|\alpha_2|)/\alpha_1$. This elasticity is allowed to be different from unity in the estimations of (2).

As it stands, we may interpret (2) as a *static* equilibrium describing a firm's labour demand in the long run. Hence, it will serve as the starting point for the cointegration analysis in Section 4. It is however likely that labour demand adjusts gradually to changes in the explanatory variables, so that (2) may not hold in the short run. In the literature on labour demand, adjustment costs are usually invoked to rationalise short run deviations from long run equilibrium [see Nickell (1986) for a survey]. An empirical representation of (2) therefore ought to be specified dynamically. A related issue concerns expectations about the conditioning variables. That is, a dynamic version of (2) may arise because manufacturers act on expectations of, say, x_t and $(w-ph)_t$.⁶ Such expectations can be purely data-based or may be formed according to model-based expectations.⁷ And, as discussed by Hendry and Neale (1988), weak exogeneity of x_t and $(w-ph)_t$ is precluded if agents act on expectation models of those variables. Even if expectations are not an issue, both x_t and $(w-ph)_t$ may well be endogenous by

⁴ Shepard's Lemma states that the optimal level of labour inputs under cost-minimisation is given by the derivative of the dual cost function with respect to its price, that is the wage rate.

⁵ This approximation to factor productivity is standard practice in the literature, cf. Hamermesh (1986, 1993) and Barrell *et al.* (1996) among others. Examples of Norwegian studies using a trend variable to proxy factor productivity include Cappelen *et al.* (1992), Bowitz and Cappelen (1994) and Boug (1999).

⁶ Dynamic labour demand schedules derived from expectation theory are documented in Sargent (1978) and Meese (1980) among others.

⁷ See e.g. Favero and Hendry (1992) for a discussion of data-based and model-based expectations.

nature. For instance, output and labour demand could be simultaneously determined owing to profit maximisation behaviour and monopolistic competition. Similarly, wage costs may be correlated with ϵ_t in (2) for reasons such as insiders who get hold of a portion of the profit and the use of overtime in the production. Alternatively, wages may be endogenously determined in the labour market. These potential caveats are addressed below.

A final point to make before we continue with data considerations concerns aggregation. Since we are interested in labour demand for the entire manufacturing sector in Norway, we have to assume that (2) remains valid at such a level of aggregation. This assumption may be questionable in our case. If the various industries constituting the aggregate have different production structures, or these structures have changed substantially during the estimation period, then the differences between the industries may lead to instability in the estimated parameters of (2). The stability properties of (2) and its dynamic counterpart are discussed at great length below.

3. The Data

The empirical analysis is conducted using quarterly, seasonally unadjusted data that spans the period 1978(1)-1996(4).⁸ Allowing for lags and transformations, the sample period used for estimations is 1979(1)-1996(4) unless otherwise noted. Generally speaking, quantity variables are measured in fixed prices and indices or implicit deflators measure prices with 1993 as the base year. Detailed data definitions and sources are provided in an appendix.

Employment is the central variable of this study. Choosing an appropriate measure for it is not unambiguous, as labour inputs may be altered by either employing more workers for a given number of hours or by using overtime for a given number of employees. We define employment as hours worked (L) rather than the number of employees in manufacturing. Our understanding that hours worked is a better proxy for labour inputs than employees motivates this choice.⁹

To ensure comparability with related studies, we employ gross production (X) as a proxy for manufacturing output. In modelling the marginal process for X , we use domestic absorption (ABS) and the rate of unemployment (U) as indicators of demand pressures in the Norwegian economy. These variables are easily observable, and may thus be used by Norwegian manufacturers to decide current

⁸ The data is taken from the Quarterly National Accounts (QNA), published by Statistics Norway unless otherwise noted. Due to a revision of the QNA, we were prevented from extending the series by data before 1978.

⁹ Similar labour demand measures have been employed by e.g. Barrell *et al.* (1996), while others have used number of employees, see e.g. Flaig and Steiner (1989).

production level and to form expectations about future domestic market conditions. Similarly, we employ an indicator of the activity level abroad (MII) that may be a helpful candidate in explaining gross production aimed for exports. This indicator is defined as a volume index for total commodity exports demand faced by Norwegian manufacturing. Formally, we postulate a general long run equilibrium relationship for the marginal process for X_t that reads as

$$(3) \quad X_t = f(ABS_t, MII_t, U_t).$$

We use average wage costs (inclusive of pay roll taxes) per hour worked (W) rather than wage costs per hour paid. The former seems to be a better proxy for the price of labour inputs as wage costs per hour paid are higher than the wage costs per hour worked due to holidays, paid leave and sick leave. The price of material inputs (PH) is represented by its implicit deflator. We draw heavily on results in Nymoen (1989), Johansen (1995) and Bowitz and Cappelen (1997) in modelling the marginal process for $(W/PH)_t$. Summarising briefly, we model the long run level of relative factor price according to the following general function:

$$(4) \quad W_t/PH_t = f(Y_t, P_t, IP_t, U_t),$$

where Y_t denotes average value added labour productivity, P_t the implicit deflator of factor income, IP_t the implicit deflator of imports and, as previously noted, U_t the rate of unemployment. Inspired by the above mentioned studies, we include short run effects from average normal working time (H) and the official consumer price index (CPI) in the dynamic counterpart of (4). In addition, the dynamic $(W/PH)_t$ -equation is expanded with dummy variables for the effects of wage freeze regulations and central wage settlements that were undertaken during the estimation period [cf. Bowitz and Cappelen (1997)]. Specifically, these variables are $D8889$, $D90(3)$ and DS where $D8889$ represents the wage freeze regulation in 1988-89, $D90(3)$ the catch up effects in 1990(3) of the relatively large centrally negotiated wage settlement in 1990(2) and DS the general effects of biannual central settlements that took place during the period 1978-96. A detailed description of the restrictions on these dummy variables is given in the appendix.

Finally, the series for the stock of capital (K) is an aggregate containing buildings, machinery and transport equipment. Bowitz and Cappelen (1994) and Boug (1999) employ a similar capital measure. Alternatively, the aggregate could be split into separate categories such as buildings (KB) and machinery and transport equipment (KM). One may then, in accordance with Cappelen *et al.* (1992),

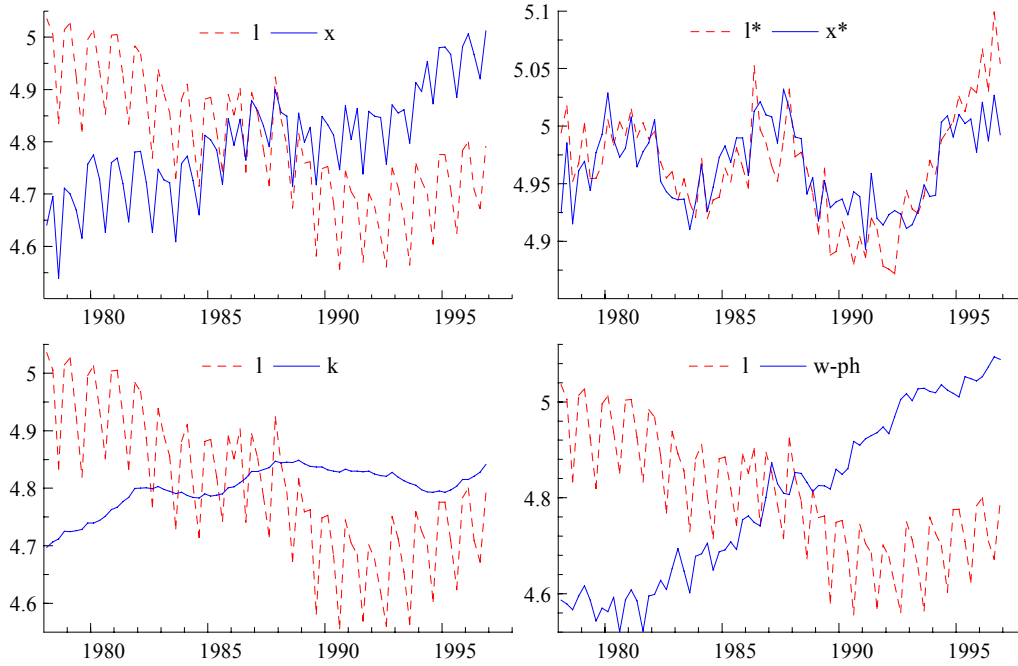
allow for the possibility that labour demand is complementary with buildings, but substitutes with machinery and transport equipment.¹⁰ We considered this hypothesis using a VAR-model in l_t , x_t , kb_t , km_t and $(w-ph)_t$. The applicability of the Johansen (1988, 1991) method was ambiguous, however, as it turned out difficult to obtain a VAR with satisfactory diagnostics. One reason may be that the dimensionality of the VAR was too large relative to the degrees of freedom. Confronted with this problem, we decided to handle a more tractable lower-dimensional VAR by aggregating the two categories of capital. We realise, however, that this aggregation may pose problems in the modelling of labour demand, especially if the two categories of capital have developed differently over the estimation period. Time series (not reported) show that both categories of capital have an upward trend, but machinery and transport equipment seem to vary somewhat more around the trend than buildings. Hopefully, this relatively small discrepancy between the movements in the two capital series will cause negligible ambiguities in the empirical analysis.

Most of the series exhibit seasonality. Specifically, the l_t -series is generally lowest in the second and third quarter and highest in the first and fourth quarter due to vacations and public holidays. To account for seasonality, we use three centred seasonal dummies, labelled S_{1t} , S_{2t} and S_{3t} . Figure 1 displays l pairwise with x , k and $(w-ph)$ over the period 1978(1)-1996(4).¹¹ To highlight correlation in the data with regards to employment and production, the figure also contains a plot of l^* together with x^* . These series represent l and x adjusted for deterministic seasonality and trend, respectively. Several aspects can be seen from the figure. First, we observe that l has a downward trend, while x has an upward trend.

¹⁰ This hypothesis is indeed supported numerically, but not statistically, by conventional significance levels in Cappelen *et al.* (1992).

¹¹ The scale of x , k and $(w-ph)$ are adjusted to match that of l in Figure 1.

Figure 1. Time series plots of (l, x) , (l^*, x^*) , (l, k) and $(l, w-ph)$



A closer look at the movements in the series in terms of l^* and x^* reveals a strong and positive correlation over the business cycle, consistent with the $\alpha_l > 0$ hypothesis. For instance, the severe decline in l during the 1987-93 recession coincides well with the fall in x . Likewise, both increased production and employment accompany the subsequent period of recovery in the Norwegian economy. We also observe that k shows a slight upward trend that may in part account for the general fall in l . Similarly, the wage costs increase relative to the price of material inputs throughout the period. Hence, l and $(w-ph)$ are negatively correlated, and may reflect long run substitution from labour to material inputs, consistent with the hypothesis that $\alpha_3 < 0$. Although both k and $(w-ph)$ seem to be important explanatory variables, it is clear that the overall fall in l cannot be entirely explained by these two series. Hence, using a linear trend to proxy the observed underlying productivity growth in l appears reasonable. We pursue this in the succeeding section.

4. Integration and Cointegration

Cointegration analysis helps clarify the empirical counterpart (or lack thereof) of the theoretical relationship stated in (2). Before applying the Johansen (1988, 1991) method, it is useful to establish the orders of integration for the variables considered. According to Augmented Dickey-Fuller (ADF) tests, we cannot reject the null hypothesis of $l \sim I(1)$, $x \sim I(1)$ and $(w-ph) \sim I(1)$. Some ambiguity arises

over the series for capital, which appears to be $I(2)$ if inferences are made on the ADF statistics alone. However, the estimated roots for Δk_t are numerically far from unity, suggesting that $\Delta k_t \sim I(0)$.¹² Thus, we conduct the Johansen analysis under the assumption that our selected variables are $I(1)$ processes.

The cointegration analysis commenced from a 5th order VAR in $[l_t, x_t, k_t, (w - ph)_t]$. This model was augmented with an unrestricted intercept and the three mentioned centred seasonal dummies. We also include a linear trend that enters the cointegration space, thereby precluding quadratic trend.¹³ The system was validly simplified to a third order VAR. We note however that the preferred VAR includes two unrestricted impulse dummies that account for outliers in the l_t -equation in 1986(2) and 1996(3). One may argue that this is unfortunate since the information set was chosen to explain l_t . An alternative approach to deal with the non-normality problem is to enlarge the VAR by relevant variables that explain the outliers. However, we are again confronted with the problem of increased dimensionality and the risk of losing degrees of freedom. We also emphasise that the added dummies in the preferred VAR are not important for parameter constancy to be present in the equations for x_t , k_t and $(w - ph)_t$. Moreover, the conclusions from the cointegration analysis are not significantly affected by these variables.

Residual misspecification tests for the third order VAR are reported in Table 1. The results are statistically acceptable, although there is indication of non-normality in the x_t -equation. As pointed out by Johansen and Juselius (1992), this is probably less important for the Johansen analysis, particularly if x_t is weakly exogenous for the long run parameters. That hypothesis is confirmed below.

For our VAR to be considered as a valid starting point of the cointegration analysis, it should also contain reasonably constant parameters. Recursive estimates (with ± 2 standard errors) and sequences of break-point Chow tests (scaled by their 1% critical values) are displayed in Figure 2. We conclude that the system is constant over the sample. The next step is thus to investigate the cointegration properties between the selected variables by means of our preferred system.

¹² It is worth pointing out that the conventional ADF-tests used here only give rough guidance to the properties of integration. In fact, such tests may have low power, as they do not take into account seasonal integration [cf. Hylleberg *et al.* (1990)]. For our series for the stock of capital, seasonal integration may be an important source of non-stationarity, as the data seems to display some seasonality but probably not altogether deterministic seasonality. The results of the ADF-tests are not reported, but are provided upon request.

¹³ Doornik *et al.* (1998) also strongly recommends commencing the analysis with a linear trend restricted to the cointegration space and an unrestricted constant to ensure asymptotic similarity to the nuisance parameters of these effects. As previously noted, the trend proxies for total factor productivity.

Table 1. Residual misspecification tests¹

Equation	Statistics						
	AR_{1-5} $F(5, 49)$	$ARCH_{1-4}$ $F(4, 46)$	$NORM$ $\chi^2(2)$	HET_1 $F(26, 27)$	AR^V_{1-5} $F(80, 124)$	$NORM^V$ $\chi^2(2)$	HET^V_1 $F(260, 194)$
l	0.366	0.834	1.577	0.548			
x	0.320	1.201	8.817*	0.856			
k	0.960	0.554	2.433	1.258			
$(w-ph)$	1.115	0.142	2.301	0.746			
VAR					1.333	10.938	0.578

¹ AR_{1-5} is Harvey's (1981) test for 5th order residual autocorrelation; $ARCH_{1-4}$ is the Engle (1982) test for 4th order autoregressive conditional heteroscedasticity in the residuals; $NORM$ is the normality test described in Doornik and Hansen (1994) and HET_1 is a test for residual heteroscedasticity due to White (1980). Similar tests for the entire VAR is denoted by ^V [see Doornik and Hendry (1997)]. $F(\cdot)$ and $\chi^2(\cdot)$ represent the null distributions of F and χ^2 , with degrees of freedom shown in parenthesis. Asterisk * denotes rejection at the 5% significance level.

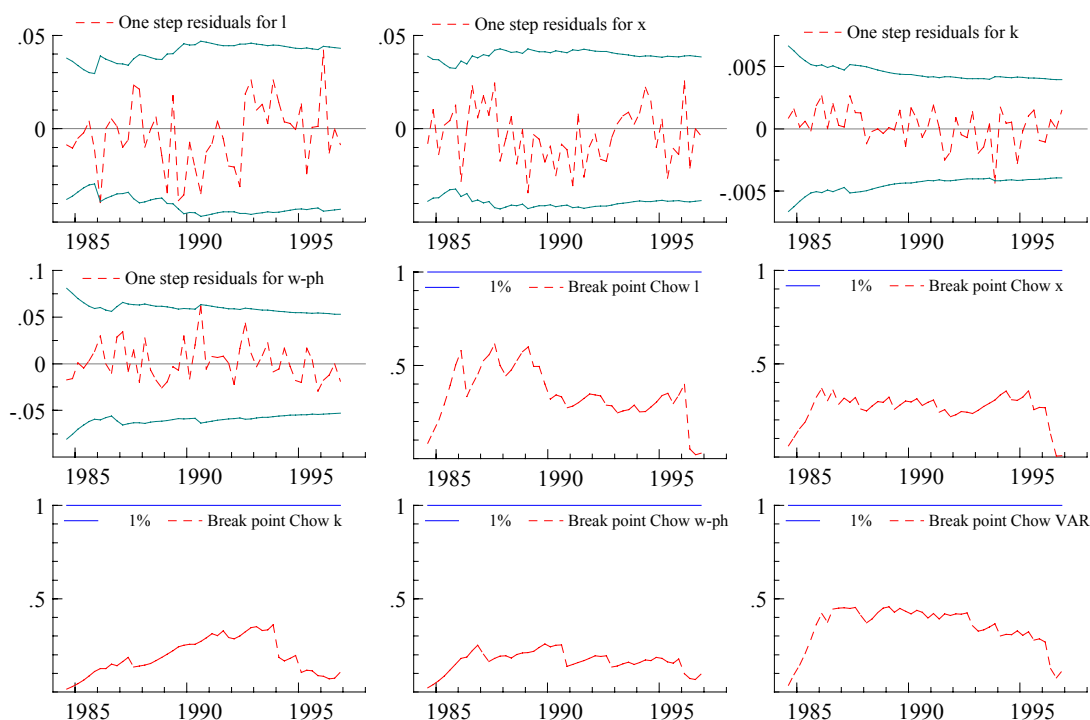
Figure 2. Recursive test statistics for the preferred VAR

Table 2 contains the results from applying Johansen's method to the VAR.¹⁴ The maximum eigenvalue and trace statistics, both with and without a degrees-of-freedom-adjustment, reject the null of no cointegration at the 1% level, but the null of at most one cointegration vector is not rejected by any of

¹⁴ Noticeably, the analysis is performed under the assumption that the presence of the impulse dummies does not affect the asymptotic critical values of the cointegration tests.

the statistics. Recursive estimation of the eigenvalues (not reported) supports our conclusion that there is a single cointegrating vector between l , x , k and $(w-ph)$. Normalising on l yields the *unrestricted* estimate of the vector, which reads as

$$(5) \quad l = \hat{\alpha}_0 + 1.349 \cdot x - 1.129 \cdot k - 0.355 \cdot (w - ph) - 0.0052 \cdot \tau, \\ (0.164) \quad (0.156) \quad (0.137) \quad (0.0014)$$

where standard errors are in parentheses. The corresponding vector of adjustment coefficients is given by

$$(6) \quad \hat{\alpha}' = (-0.399, -0.059, -0.050, -0.049).$$

The estimates in (5) and (6) suggest that the cointegrating vector can be interpreted as a long run labour demand equation consistent with the neoclassical model in Section 2. To be specific, the estimated adjustment coefficients indicate that l_t is strongly error correcting in the case of a disequilibrium in (5), while there is weak error correction, if at all, in the equations for x_t , k_t and $(w-ph)_t$. Also, the elasticities in (5) have their expected signs, and they are all highly significant.

Table 2. Johansen's cointegration tests¹

Information set: $[l, x, k, (w-ph)]$						
Eigenvalues: 0.486, 0.259, 0.138, 0.079						
Hypothesis	Statistics					
	λ_{max}	λ^a_{max}	95%	λ_{trace}	λ^a_{trace}	95%
$r=0$	48.63**	40.63**	31.5	87.40**	73.03**	63.0
$r \leq 1$	21.93	18.33	25.5	38.77	32.40	42.4
$r \leq 2$	10.80	9.03	19.0	16.84	14.07	25.3
$r \leq 3$	6.04	5.05	12.3	6.04	5.05	12.3

¹ r denotes rank, or the number of cointegrating vectors. The λ_{max} and λ_{trace} statistics are the maximum eigenvalue and trace statistics, whereas λ^a_{max} and λ^a_{trace} are the corresponding statistics with a degrees-of-freedom-adjustment. The 95% quantiles are taken from Table 2* in Osterwald-Lenum (1992). Asterisk ** denotes rejection of the null hypothesis at the 1% significance level.

Formal tests of this interpretation are given in Table 3, which reports likelihood ratio statistics for tests of weak exogeneity and structural hypotheses on the long run parameters [see Johansen and Juselius (1990, 1992)]. The statistics are asymptotically distributed as $\chi^2(\cdot)$ with degrees of freedom given in parenthesis. Firstly, weak exogeneity of employment and capital for the long run parameters is strongly rejected, implying that the cointegrating vector enters both the l_t -equation and the k_t -

equation.¹⁵ It seems however safe to assume that gross production and the factor price ratio are weakly exogenous for the long run parameters, both individually and jointly. Secondly, the exclusion tests show that each of l , x , k , $(w-ph)$ and τ enters significantly in the cointegrating vector. Thirdly, the coefficient on k in (5) is close to unity, and the restriction of long run homogeneity between l and k indeed cannot be rejected.

Table 3. Tests of long run weak exogeneity and structural hypotheses¹

Weak exogeneity tests ² :		Variable				
Statistics		l	x	k	$w-ph$	$(x, w-ph)$
$\chi^2(\cdot)$		8.087**	0.311	24.276**	0.120	0.538
p -value		(0.004)	(0.577)	(0.000)	(0.729)	(0.764)
Exclusion tests:		Variable				
Statistics		l	x	k	$w-ph$	τ
$\chi^2(1)$		22.649**	22.928**	23.888**	4.837*	7.131**
p -value		(0.000)	(0.000)	(0.000)	(0.027)	(0.008)
Tests of homogeneity between l and k :						
Statistics		Unconditionally		Conditional on $(x, w-ph)$ being weakly exogenous ³		
$\chi^2(1)$		0.445		0.370		
p -value		(0.505)		(0.543)		

¹ The test statistics are calculated under the assumption that $r=1$, and are asymptotically distributed as $\chi^2(\cdot)$ under the null, with degrees of freedom in parenthesis. Asterisks * and ** denote rejection at the 5% and 1% significance levels. ² The weak exogeneity tests for individual variables are distributed as $\chi^2(1)$, and as $\chi^2(2)$ for the joint test of $(x, w-ph)$. ³ The conditional homogeneity test statistics parallel the statistics in Johansen and Juselius (1990).

Finally, imposing the homogeneity restriction and weak exogeneity of x_t and $(w-ph)_t$, we obtain

$\chi^2(3)=0.908$ (with a p -value of 0.823) and the following *restricted* estimate of the cointegrating vector (normalised on l):

$$(7) \quad l = \alpha_0 + 1.368 \cdot x - k - 0.316 \cdot (w - ph) - 0.0057 \cdot \tau ,$$

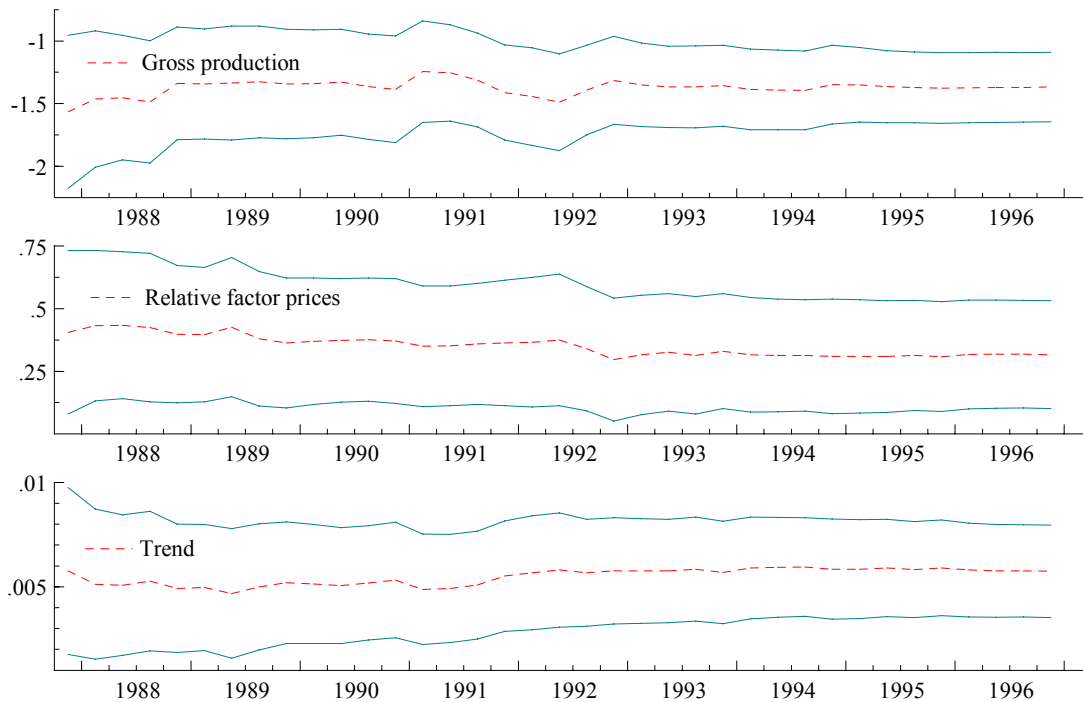
(0.155) (0.119) (0.0012)

where standard errors are in parentheses. The estimates in (7) are virtually unchanged from the unrestricted ones in (5). The similarity of coefficient estimates further points to the validity of the imposed restrictions in (7). A sequence of $\chi^2(3)$ test statistics (not shown) also confirms the validity of

¹⁵ This finding motivates the use of the Johansen procedure, so as to obtain efficient estimates of the cointegrating vector.

the restrictions at a good margin for any sample size between 1986 and 1996. Additionally, as Figure 3 shows, the recursively estimated parameters of x , $(w-ph)$ and τ in (7) are acceptably constant.¹⁶

Figure 3. Recursive estimates of parameters in (7) $\pm 2 \cdot \sigma_t$



The long run output elasticity entails that l increases by 1.37% if x increases permanently by 1%, *ceteris paribus*. We note that the implied scale elasticity with respect to variable factors of production is estimated to 0.73. Owing to the homogeneity restriction, we observe the same scale elasticity with respect to the stock of capital. It thus follows that the total scale elasticity sums to 1.46, its economic interpretation being increasing returns to scale in Norwegian manufacturing. Interestingly, in a related study Flaig and Steiner (1989) find the scale elasticity to be around 1.5 for German manufacturing. Also, in a related but somewhat different model, Harvey *et al.* (1986) estimates a similar magnitude of the total scale elasticity for UK manufacturing. Conversely, Barrell *et al.* (1996) gains support for a hypothesis of constant returns to scale in a study of aggregate labour demand in UK, France and Germany. Our findings are also in agreement with previous Norwegian studies.¹⁷ The estimated long run own price elasticity in (7) means that l decreases by somewhat more than 0.3% if wage costs

¹⁶ From an economic point of view, it is interesting to notice the constancy of the estimated coefficient of τ . We may interpret this finding as a support to the claim that factor productivity has not changed significantly over the sample. It thus seems reasonable to proxy underlying productivity by a *deterministic* trend rather than a stochastic one [cf. Harvey *et al.* (1986)].

¹⁷ See Cappelen *et al.* (1992), Bowitz and Cappelen (1994) and Boug (1999).

relative to the price of material inputs rise permanently by 1%, *ceteris paribus*. This finding is consistent with the best guess in the literature for the long run constant-output labour demand elasticity [cf. Hamermesh (1991)]. As regards the trend effects, the negative coefficient of τ indicates that technical progress has contributed to the observed decline in employment. That is, the estimated semielasticity implies that employment on average falls by 0.57% ($-0.0057 \cdot 100$) each quarter and by 2.28% ($-0.0057 \cdot 400$) each year, *ceteris paribus*. Flaig and Steiner (1989) obtain a yearly trend effect of 2.4% for German manufacturing. In contrast, albeit using a stochastic trend, Harvey *et al.* (1986) estimates the annual productivity effect to be 2.6% for UK manufacturing.

The adjustment coefficient associated with (7) is estimated to -0.346 . Economically, the estimate means that around 35% of disequilibrium in (7) is corrected within one quarter. The short run dynamics in l is not however fully described by this coefficient. Given its importance for policy, the short run adjustment of l is investigated in the next section.

5. A Parsimonious Labour Demand Model

We now turn to dynamic modelling of labour demand in Norwegian manufacturing. For this purpose, we formulate a dynamic labour demand equation, using deviations from (7) as an equilibrium correction mechanism (*EqCM*). Such a model is of great interest from a policy perspective as it enables us to study the dynamic adjustment of employment to changes in gross production, the capital stock and the factor price ratio. Applying the same information set as in the cointegration analysis leads to the following general equilibrium correction model:¹⁸

$$(8) \quad \Delta l_t = \beta_0 + \sum_{i=1}^3 \beta_{1i} S_{it} + \sum_{i=1}^4 \beta_{2i} \Delta l_{t-i} + \sum_{i=0}^4 \beta_{3i} \Delta x_{t-i} + \sum_{i=0}^4 \beta_{4i} \Delta k_{t-i} + \sum_{i=0}^4 \beta_{5i} \Delta(w-ph)_{t-i} \\ + \delta[l_{t-1} - 1.368 \cdot x_{t-1} + k_{t-1} + 0.316 \cdot (w-ph)_{t-1} + 0.0057 \cdot \tau] + e_t,$$

where Δ is the difference operator, the expression in $[\cdot]$ is the *EqCM* and e_t is the error term, assumed to be white noise. Our modelling strategy in the search for a parsimonious representation of (8), is that of the *general to specific approach* advocated by Davidson *et al.* (1978). However, we have to address the issue of whether single equation analysis or system analysis should be applied in our case. As it stands, (8) includes the contemporaneous variable Δk_t , and the conditioning of this variable is subject to a potential caveat if (8) is to be estimated within a single equation framework. We pursue the

¹⁸ The unrestricted impulse dummies from the VAR, i.e. *D86(2)* and *D96(3)*, are suppressed for ease of exposition.

modelling device made by Boswijk and Urbain (1997) in this manner, a device that reads as follows when translated to our particular case: "If there is error correction in the marginal process generating k_t , but the orthogonality condition, i.e. $COV(\Delta k_t, e_t) = 0$, is maintained, then an efficient estimator of the cointegrating vector may be obtained from the Johansen's (1988, 1991) procedure. Conditional upon this estimate, the conditional model in (8) may then be estimated using least squares (*OLS*)."

The analysis in the preceding section yielded an efficient estimate of the cointegrating vector. It remains to establish the exogeneity status of k_t as regards the short run parameters. Preliminary estimations and subsequent simplifications of (8) were thus conducted with instruments for Δk_t .¹⁹ This variable and its lags were however far from being significant in any of the regressions considered. So were Δk_t and its lags when (8) was simplified using least squares. The parsimonious labour demand model in (11) therefore omits short run effects of the stock of capital, and it is validly derived within a single equation analysis.

The conditioning of the two other contemporaneous variables, Δx_t and $\Delta(w-ph)_t$, may also pose caveats. Although, the results in Table 3 clearly support weak exogeneity of these variables for the cointegrating vector, they need not be so for the short run parameters in (11) [cf. Urbain (1992) and Boswijk and Urbain (1997)]. To account for potential simultaneity, we apply Wu-Hausman tests for orthogonality of Δx_t and $\Delta(w-ph)_t$ and the errors in the parsimonious labour demand model.

Digressing briefly, these tests are based on the modelled marginal processes of Δx_t and $\Delta(w-ph)_t$ reported in (9) and (10). Seasonal dummies are omitted for convenience. Both equations were simplified from general specifications. Estimated standard errors are in parentheses, and T , R^2 , σ and DW are the number of observations, the squared multiple correlation coefficient, the residual standard error and the Durbin-Watson statistic, respectively. In addition to AR_{1-5} , $ARCH_{1-4}$, $NORM$ and HET_1 defined in Table 1, we report HET_2 and $RESET$. The former tests whether the squared residuals depend on the levels, squares and cross products of the regressors [cf. White (1980)], while the latter tests for functional form misspecification [cf. Ramsey (1969)]. None of the diagnostics is significant at the 1% level. Both models also appear reasonably constant.²⁰

¹⁹ Various combinations of instruments were used for Δk_t . The instruments were: Δk_{t-i} , Δx_{t-i} , Δl_{t-i} , Δabs_{t-i} , Δmii_{t-i} and Δu_{t-i} with $i=1,2,3,4$. See the appendix for details.

²⁰ Recursive test statistics for (9) and (10) are not reported, but are available upon request.

$$\begin{aligned}
(9) \quad \Delta x_t = & 3.547 - 0.219 \cdot \Delta x_{t-1} + 0.359 \cdot \Delta abs_{t-1} \\
& (0.844) (0.070) \quad (0.058) \\
& + 0.649 \cdot \Delta mii_t - 0.066 \cdot \Delta_2 u_{t-3} \\
& (0.092) \quad (0.016) \\
& - 0.376 \cdot (x - 0.369 \cdot mii + 0.047 \cdot u)_{t-1} \\
& (0.089)
\end{aligned}$$

Method: *OLS* $T=71[1979(2)-1996(4)]$ $R^2=0.948$ $\sigma=1.917\%$ $DW=2.24$
 $AR_{1-5}:F(5, 59)=0.882$ $ARCH_{1-4}:F(4, 56)=0.234$ $NORM:\chi^2(2)=3.437$
 $HET_1:F(11, 52)=0.778$ $HET_2:F(26, 37)=0.968$ $RESET:F(1, 63)=0.007$

$$\begin{aligned}
(10) \quad \Delta(w-ph)_t = & -0.058 - 0.308 \cdot \Delta_3(w-ph)_{t-1} - 1.176 \cdot \Delta h_t \\
& (0.019) (0.072) \quad (0.419) \\
& - 0.063 \cdot \Delta u_{t-3} + 0.645 \cdot \Delta cpi_{t-3} - 0.158 \cdot [(w-ph) - y - p + ip]_{t-1} \\
& (0.022) \quad (0.284) \quad (0.043) \\
& - 0.029 \cdot D8889 + 0.069 \cdot D90(3) + 0.028 \cdot DS \\
& (0.011) \quad (0.022) \quad (0.011)
\end{aligned}$$

Method: *OLS* $T=72[1979(1)-1996(4)]$ $R^2=0.626$ $\sigma=2.097\%$ $DW=2.03$
 $AR_{1-5}:F(5, 57)=0.256$ $ARCH_{1-4}:F(4, 54)=0.544$ $NORM:\chi^2(2)=1.761$
 $HET_1:F(15, 46)=0.757$ $HET_2:F(36, 25)=0.726$ $RESET:F(1, 61)=0.021$.

The Δx_t model includes short run effects of domestic absorption (*abs*), the rate of unemployment (*u*) and foreign demand (*mii*). According to the estimates, increases in domestic and foreign demand pressure induce manufacturing production to rise. No significant long run effects from domestic absorption were detected. Around 38% of deviations from equilibrium in production, measured by the expression in (\cdot), is corrected within one quarter. In agreement with Nymoen (1989), Johansen (1995) and Bowitz and Cappelen (1997),²¹ the $\Delta(w-ph)_t$ model contains short run effects of average normal working time (*h*), consumer prices (*cpi*) and dummies for wage freeze regulations (*D8889*) and wage settlements [*D90(3)* and *DS*]. The former effect in particular is consistent with the negotiated compensation schemes for shorter hours in Norwegian manufacturing. Turning to the long run estimates, we observe that the cointegrating vector, which enters (10) significantly at the 1% level,

²¹ Not surprisingly, (10) differs somewhat from those in the mentioned studies. The main cause of the difference may be that we model the factor price ratio (W/PH), rather than wages alone. Besides, Nymoen (1989) estimates a wage model using fourth differences, whereas Johansen (1995) uses annual data.

consists of the wage costs share ($w-y-p$), the price of material inputs (ph) and the import price (ip).²² Interestingly, and unlike the mentioned studies, no long run wage responsiveness to unemployment is found regardless of specifying $f(U)$ as $\log(U)$ or U^{-2} , see e.g. Johansen (1995). This finding does however coincide with the results in Rødseth and Holden (1990).

We are now ready to apply the announced Wu-Hausman tests conditional on the weak exogeneity of Δx_t and $\Delta(w-ph)_t$ for the cointegration parameters. Adding each of the derived estimates of Δx_t and $\Delta(w-ph)_t$ to (11) yielded p -values of 0.872 and 0.553. Following Urbain (1992), we also tested whether the $EqCM$ from (11) is significant if added to (9) and (10). The $EqCM$ was however far from being significant, with p -values of 0.507 (Δx_t) and 0.398 [$\Delta(w-ph)_t$]. We thus conclude that gross production and the factor price ratio are weakly exogenous for all the parameters of interest in (11), parameters of which thereby are consistently estimated by *OLS*. Having established these results, we now focus on the parsimonious labour demand model in (11).

$$\begin{aligned}
 (11) \quad \Delta l_t = & 1.076 - 0.181 \cdot \Delta_3 l_{t-1} + 0.670 \cdot \Delta x_t - 0.226 \cdot \Delta(w-ph)_t \\
 & (0.223) (0.064) \quad (0.075) \quad (0.079) \\
 & - 0.331 \cdot [l_{t-1} - 1.368 \cdot x_{t-1} + k_{t-1} + 0.316 \cdot (w-ph)_{t-1} + 0.0057 \cdot \tau] \\
 & (0.068) \\
 & - 0.056 \cdot S_{1t} - 0.095 \cdot (S_{2t} + S_{3t}) + 0.069 \cdot D86(2) + 0.053 \cdot D96(3) \\
 & (0.013) \quad (0.016) \quad (0.020) \quad (0.019)
 \end{aligned}$$

Method: *OLS* $T=72[1979(1)-1996(4)]$ $R^2=0.972$ $\sigma=1.881\%$ $DW=1.89$
 $AR_{1-5}:F(5, 58)=1.019$ $ARCH_{1-4}:F(4, 55)=1.021$ $NORM:\chi^2(2)=2.296$
 $HET_1:F(12, 50)=0.952$ $HET_2:F(26, 36)=0.762$ $RESET:F(1, 62)=2.590$.

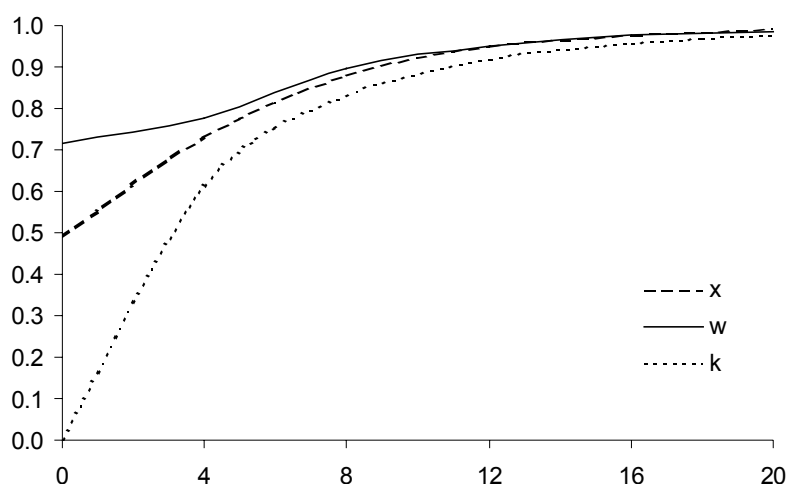
As before, estimated standard errors are in parentheses. The diagnostic tests are those previously defined. Our preferred model passes the misspecification tests at a good margin. According to the R^2 and the σ statistics, the fit of the model is satisfactory. A plot (not reported) also shows that (11) tracks the movements in Δl_t reasonably well. Moreover, the economic variables entering (11) are all highly significant. For instance, the $EqCM$ appears in the model with a t -value of -4.87 , hence adding force to the results obtained in Section 4. Besides, the adjustment coefficient of -0.33 is virtually identical to the adjustment coefficient related to (7). As regards gross production and the factor price ratio, the estimated impact elasticities of 0.67 and -0.23 are considerably smaller than their long run

²² The cointegrating vector in (10) and its imposed homogeneity restrictions commenced from experimenting with long run relationships for W and PH , each of which was hypothesised to be $W=P \cdot Y \cdot f(U)$ and $PH=P^0 \cdot IP^{(1-\phi)}$. Together, the long run factor price ratio was hypothesised to equal $W/PH=(P/IP)^{(1-\phi)} \cdot Y \cdot f(U)$. Noticeably, and in line with Johansen (1995) findings, we assume wedge effects to be unimportant in explaining the wage costs. No wedge effects may also be argued from theory [cf. e.g. Layard *et al.* (1991 chapter 2)].

counterparts. Hence, substantial smoothing of employment with respect to changes in x_t and $(w-ph)_t$, seems to be the case in Norwegian manufacturing. Next, equation (11) includes strongly and significant lagged effects of the third quarter growth in the dependent variable ($\Delta_3 l_{t-1}$), with a coefficient of -0.18 . Lastly, the dummies capture the effects of seasonality and outliers. Lags of Δx_t and $\Delta(w-ph)_t$ were however insignificant, and so omitted from (11).

To illustrate the dynamics implied by (11) in more detail, we calculated the adjustment of employment to partial and permanent shifts in gross production, the stock of capital and the factor price ratio. Figure 4 plots the simulated response of l , measured as standardised interim multipliers, to a 1% increase in each of x , k and $(w-ph)$. We observe that employment adjusts gradually to changes in these variables. Somewhat more than 75% of the total adjustment in l following an increase in $(w-ph)$ is accomplished within one year, whereas around 90% of the long run response is completed after two years. Much the same may be said of the adjustment of employment to increases in gross production. These interim multipliers may indicate presence of adjustment costs that slow down the adjustment of employment in Norwegian manufacturing [cf. Nickell (1986)]. Moreover, the apparent slow adjustment of l to shifts in x and $(w-ph)$ may reflect the institutional systems characterising the Norwegian labour market. That is, stringent overtime and redundancy regulations may make it difficult for manufacturers to adjust employment to changes in factors that determine labour demand. Barrell *et al.* (1996) argues that the relatively slow speed of adjustment of employment in France may be explained by such legal and institutional factors. The response of l to changes in k is even more gradual, with only 60% and 80% of the long run adjustment completed after one and two years, respectively. This finding is consistent with the notion that capital is fixed in the short run due to frictional and installation costs (cf. Tobin's q theory of capital investment).

Figure 4. Dynamics implied by (11). Standardised interim multipliers



6. Tests of the Lucas Critique

We use the concept of super exogeneity to examine the extent to which the labour demand model (11) suffers from the Lucas critique. Following Engle *et al.* (1983), Δx_t and $\Delta(w-ph)_t$ are viewed as super exogenous in our model if (i) they are weakly exogenous, and (ii) the parameters in (11) are invariant to changes in the two marginal distributions of x_t and $(w-ph)_t$. These conditions fail if manufacturers act on expectation models of x_t and $(w-ph)_t$. If so, and if such expectation models change sufficiently, then the Lucas critique applies to (11) [cf. Hendry (1988)]. Even if expectations are not an issue, (11) may still "break down", as its parameters need not be invariant to all conceivable interventions [cf. Favero and Hendry (1992)].

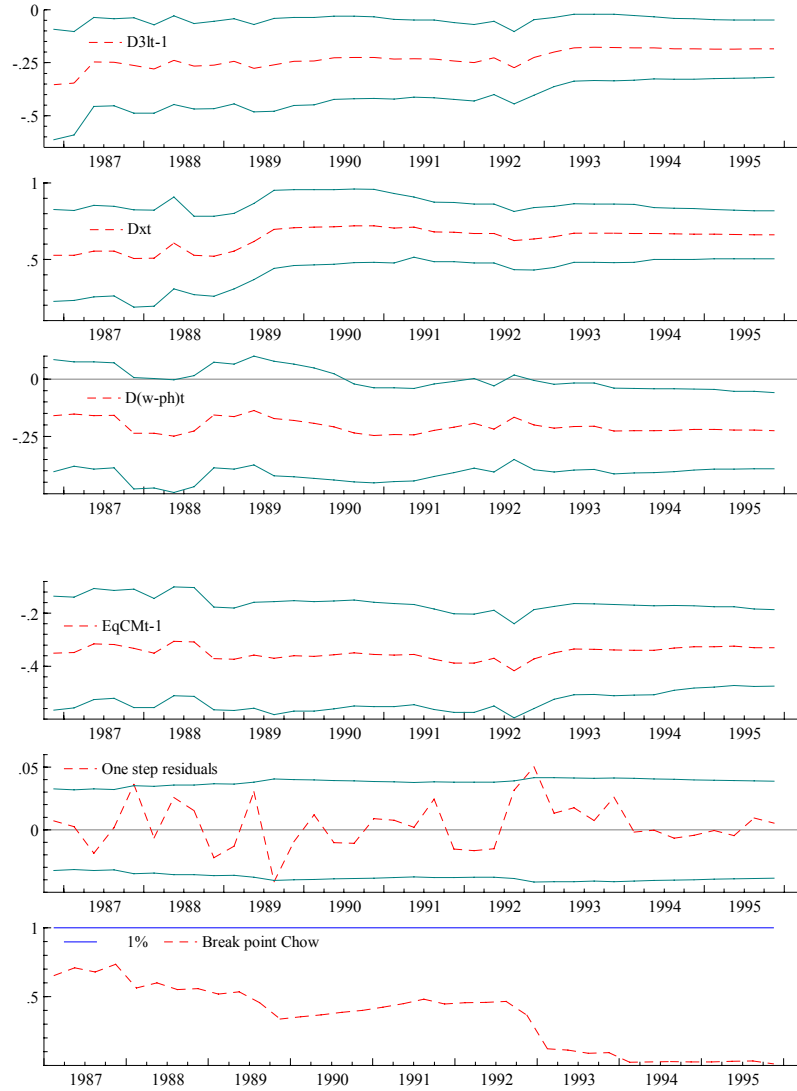
We have already found that Δx_t and $\Delta(w-ph)_t$ seem to be weakly exogenous for the parameters of interest. Our next step is thus to test the second condition for super exogeneity. To that end, we may follow the testing procedure described in Hendry (1988) and Engle and Hendry (1993). That is, to test the null hypothesis of invariance of the parameters in our labour demand model with respect to changes in marginal processes governing x_t and $(w-ph)_t$. Such tests are ostensibly promising, given the major shocks to Norwegian manufacturers in the eighties and nineties. However, applying the test approach to (11) will lack power since the marginal processes appear to be reasonably stable over the sample.²³ Hence, we ought to use other testing procedures in this respect. First, we evaluate whether (11) is invariant to interventions that occurred over the 1978-96 period by means of recursive methods. Then, we apply variable addition tests to (11) based on marginal models for Δx_t and $\Delta(w-ph)_t$ in order to examine the possible role of expectations [cf. Engle and Hendry (1993, pp. 127-131)].

Figure 5 shows recursive test statistics and recursively estimated coefficients of four central variables entering equation (11), namely those of Δs_{t-1} , Δx_t , $\Delta(w-ph)_t$ and $EqCM$. Neither the one-step residuals (with ± 2 standard errors) nor the sequence of break point Chow tests (scaled by their 1% critical values) indicate non-constancies in the model. Similarly, the coefficients of the economic variables show a high degree of stability over the sample. We also recall that the long run coefficients of the variables constituting the $EqCM$ are found to be empirically constant. Based on these results, we

²³ Both the x_t model and the $(w-ph)_t$ model of the VAR in Section 4 and simplifications of those models are constant without intervention dummies [also without $D86(2)$ and $D96(3)$]. This is also the case for the marginal models reported in (9) and (10) in Section 5, and in (12) and (13) below.

conclude that our labour demand model demonstrates invariance to interventions that took place over the 1978-96 period.²⁴

Figure 5. Recursive estimates and test statistics of (11)



Turning to the examination of the role of expectations, we first note that the Wu-Hausman tests in Section 5 do not support the claim that manufacturers act on expectations models of Δx_t and $\Delta(w-ph)_t$.

²⁴ Our claim that (11) exhibits invariance to all conceivable interventions over the sample is made with a proviso. One may interpret the impulse dummies, $D86(2)$ and $D96(3)$, as variables picking up effects of radical events in the economy that reject invariance [cf. Favero and Hendry (1992)]. In particular, it is likely that $D86(2)$ captures the effects on employment of the 12% devaluation of the Norwegian krone in May 1986. That is, manufacturers may have anticipated higher production, and thereby increased employment, following the devaluation. If such expectations indeed were present, then the Lucas critique applies to (11).

That said, those tests do not preclude agents from acting on model-based expectations of Δx_{t+1} and $\Delta(w-ph)_{t+1}$. We considered this possibility by testing the significance of determinants of Δx_{t+1} and $\Delta(w-ph)_{t+1}$ in (11). Significance of any of these determinants will provide evidence in favour of the Lucas critique. We are however prevented from using the already constructed marginal models as their Δx_{t+1} and $\Delta(w-ph)_{t+1}$ values include leaded variables that would be endogenous in (11). Our next task is thus to develop new marginal models based on those in (9) and (10). The chosen models for Δx_t and $\Delta(w-ph)_t$ are reported in (12) and (13) (seasonal dummies are omitted for convenience).

$$(12) \quad \Delta x_t = 3.120 - 0.574 \cdot \Delta x_{t-1} + 0.410 \cdot \Delta mii_{t-1} \\ (1.053) \quad (0.126) \quad (0.135) \\ - 0.056 \cdot \Delta_2 u_{t-3} - 0.330 \cdot (x - 0.369 \cdot mii + 0.047u)_{t-2} \\ (0.021) \quad (0.112)$$

Method: *OLS* $T=71[1979(2)-1996(4)]$ $R^2=0.917$ $\sigma=2.410\%$ $DW=2.37$
 $AR_{1-5}:F(5, 59)=0.921$ $ARCH_{1-4}:F(4, 56)=0.724$ $NORM:\chi^2(2)=0.039$
 $HET_1:F(10, 53)=1.074$ $HET_2:F(24, 39)=0.575$ $RESET:F(1, 63)=0.178$

$$(13) \quad \Delta(w-ph)_t = -0.038 - 0.345 \cdot \Delta_3(w-ph)_{t-1} - 0.068 \cdot \Delta u_{t-3} \\ (0.019) \quad (0.075) \quad (0.026) \\ + 0.567 \cdot \Delta cpi_{t-3} - 0.114 \cdot [(w-ph) - y - p + ip]_{t-2} \\ (0.299) \quad (0.044) \\ - 0.022 \cdot D8889 + 0.069 \cdot D90(3) + 0.024 \cdot DS \\ (0.012) \quad (0.025) \quad (0.011)$$

Method: *OLS* $T=72[1979(1)-1996(4)]$ $R^2=0.558$ $\sigma=2.280\%$ $DW=2.06$
 $AR_{1-5}:F(5, 57)=0.286$ $ARCH_{1-4}:F(4, 54)=0.375$ $NORM:\chi^2(2)=5.298$
 $HET_1:F(14, 47)=0.603$ $HET_2:F(35, 26)=0.500$ $RESET:F(1, 61)=0.637$.

Equation (12) and (13) include nine regressors that were not used in the construction of (11). Table 4 reports test results from adding the one period lead of these regressors, either individually or jointly, to our labour demand model. It is seen that none of the regressors proved to be significant. Similarly, adding the one period lead of all nine regressors jointly to (11) yields $F(9, 54)=0.20$ and a p -value of 0.993. In other words, the added terms are far from being significant. This conclusion is not altered if the seasonal dummies enter unrestrictedly in (11).

Table 4. Tests of model-based expectations of Δx_{t+1} and $\Delta(w-ph)_{t+1}$

Single variable tests of significance in (11):	One period lead of regressors in (12) ¹			One period lead of regressors in (13) ¹					
Statistics	Δmii_t	$\Delta_2 u_{t-2}$	ecm_{t-1}	Δu_{t-2}	Δcpi_{t-2}	ecm_{t-1}	$D8889$	$D90(3)$	DS
<i>t</i> -value	0.625	0.299	1.071	-0.039	-0.181	-0.340	0.106	-0.453	-0.672
<i>p</i> -value	0.534	0.766	0.289	0.969	0.857	0.735	0.916	0.652	0.504
Joint tests of significance in (11):	One period lead of regressors in (12)			One period lead of regressors in (13)					
Statistics	Δmii_t	$\Delta_2 u_{t-2}$	ecm_{t-1}	$\Delta u_{t-2}, \Delta cpi_{t-2}, ecm_{t-1}, D8889, D90(3), DS$					
<i>F</i> (·)-value ²	0.444			0.106					
<i>p</i> -value	0.723			0.995					

¹ The *ecm* term equals $(x-0.369 \cdot mii+0.047 \cdot u)$ and $[(w-ph)-y-p+ip]$. ² The *F*-statistics are distributed as *F*(3,60) and *F*(6,57).

Hence, the results do not support the hypothesis that Norwegian manufacturers act on model-based expectations of Δx_{t+1} and $\Delta(w-ph)_{t+1}$. Hendry and Ericsson (1991) have proposed an explanation for the absence of expectation models. Manufacturers may form their expectations of gross production and relative factor prices using data-based predictors rather than formal expectation models. Such behaviour may prove rational when information is costly to collect and process. In that case, our results need not imply that manufacturers are not forward-looking. That is, if data-based predictors are important, then (11) has a forward-looking interpretation. Nevertheless, we do not find significant evidence in favour of the Lucas critique.

7. Conclusions

This paper has studied the determinants of labour demand in Norwegian manufacturing during the period 1978-96. We first applied Johansen's multivariate cointegration procedure as a basis for modelling labour demand. It was shown that employment, production, relative factor prices, total factor productivity and the stock of real capital enter significantly in a single cointegrating vector. Normalised on employment, the long run elasticities were found to be 1.37 (production), -1.00 (the stock of real capital), -0.32 (relative factor prices) and -0.57 (total factor productivity). The former two elasticities imply increasing returns to scale in production. Noticeably, the estimated cointegrating vector entered significantly not only the labour demand equation, but also the equation for capital in the VAR. Consequently, previous studies on Norwegian data may have failed in obtaining efficient estimates as *long run* labour demand elasticities typically are estimated within a single equation framework.

Next, we developed a conditional labour demand model taking the estimated cointegrating vector as given. In addition to a highly significant equilibrium correction term, this model contained significant

effects of the growth in production and relative factor prices, both of which were economically important. The thrust of the *dynamic* modelling was grounded in that no significant short run effects of the stock of capital were found, findings that allowed the application of single equation analysis rather than a system one [cf. Boswijk and Urbain (1997)]. The estimated dynamics imply that employment adjusts gradually to partial and permanent shifts in production, relative factor prices and the stock of capital. Presence of adjustment costs and the institutional structure characterising the Norwegian labour market may explain this implication of our model.

Finally, we investigated the stability properties of the estimated labour demand model, and we tested whether the Lucas critique is relevant in our case. Recursive methods proved the specified model to be reasonably stable. Based on this finding, we concluded that the labour demand model is invariant to interventions that occurred during the period 1978-96. Furthermore, variable addition tests revealed no support to the claim that manufacturers act on model-based expectations of production and relative factor prices. At the outset, this is perhaps rather surprising in a world with adjustment costs and changing market conditions. One possibility is that manufacturers, whilst forward-looking, employ data-based predictors to forecast the evolution of production and relative factor prices. Nevertheless, the evidence in this paper suggests that the Lucas critique may not be relevant to the modelling of labour demand in Norwegian manufacturing.

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Appendix

Data Definitions and Sources

<i>L</i>	Hours worked by wage earners in manufacturing, expressed in 1000. Source: Statistics Norway.
<i>X</i>	Gross production in manufacturing at fixed 1993 prices. Source: Statistics Norway.
<i>ABS</i>	Domestic absorption at fixed 1993 prices, defined as $ABS=C+J+G$, where <i>C</i> is private consumption, <i>J</i> is private gross investments in fixed capital and <i>G</i> is government consumption expenditure. Source: Statistics Norway.
<i>U</i>	The rate of unemployment. Source: Statistics Norway.
<i>MII</i>	Weighted volume index for total commodity exports demand faced by Norwegian manufacturing (1993=1). Gross production in each manufacturing industry is used as weights. Source: OECD and Lindquist (1993).
<i>W</i>	Average wage costs (inclusive of pay roll taxes) per hour worked in manufacturing, expressed in nominal terms. Source: Statistics Norway.
<i>PH</i>	Implicit price deflator of material inputs in manufacturing (1993=1). Source: Statistics Norway.
<i>Y</i>	Average value added labour productivity in manufacturing, computed as value added at fixed 1993 prices per hour worked. Source: Statistics Norway.
<i>P</i>	Factor income deflator in manufacturing (1993=1). Source: Statistics Norway.
<i>IP</i>	Implicit deflator of imports of manufacturers (1993=1). Source: Statistics Norway.
<i>H</i>	Average normal working time (hours) per quarter in manufacturing. Source: Statistics Norway.
<i>CPI</i>	The official consumer price index (1993=1). Source: Statistics Norway.
<i>D8889</i>	Dummy variable for the wage freeze in 1988-89. Equals 1 in 1988(2)-1989(1), zero otherwise, cf. Bowitz and Cappelen (1997).
<i>D90(3)</i>	Dummy variable for catch up effects of the centrally negotiated wage settlement in 1990(2). Equals 1 in 1990(3), zero otherwise, cf. Bowitz and Cappelen (1997).
<i>DS</i>	Dummy variable for general effects of biannual settlements that were undertaken in Norwegian manufacturing during the period 1978-96. Equals 0.5 in quarters of the first year of central settlements and -0.5 in quarters belonging to the second year, cf. Bowitz and Cappelen (1997).
<i>K</i>	The stock of capital in manufacturing at 1993 prices, defined as $K=KB+KM$, where <i>KB</i> is buildings and <i>KM</i> is machinery and transport equipment. Source: Statistics Norway.
<i>S_{it}</i>	Centred seasonal dummy for quarter <i>i</i> , equals 0.75 in quarter <i>i</i> , -0.25 otherwise, <i>i</i> =1,2,3.
<i>D86(2)</i>	Dummy variable used to account for an outlier in the equation for <i>L</i> in the VAR. Equals 1 in 1986(2), zero otherwise.
<i>D96(3)</i>	Dummy variable used to account for an outlier in the equation for <i>L</i> in the VAR. Equals 1 in 1996(3), zero otherwise.