INSTITUT FÜR VOLKSWIRTSCHAFTSLEHRE

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UNIVERSITÄT AUGSBURG

Volkswirtschaftliche Diskussionsreihe

STABILITY AND DYNAMIC PROPERTIES OF LABOUR DEMAND IN WEST-GERMAN MANUFACTURING

von

Gebhard Flaig und Viktor Steiner

Beitrag Nr. 29

Universität Augsburg
Memminger Straße 14
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Zusammenfassung

In dieser Arbeit wird die Stabilität und die dynamischen Eigenschaften der Arbeitsnachfragefunktion des Verarbeitenden Gewerbes untersucht. Erklärende Variable sind Output, ein Index der Faktorproduktivität und relative Faktorpreise. Die Schätzung der dynamischen Arbeitsnachfragefunktion erfolgt mit Hilfe des Kointegrations-Ansatzes. Die Stabilität der Regressionsparameter wird mit dem CHOW- und CUSUM-Test überprüft. Eine dynamische Simulation zeigt, daß eine Arbeitsnachfragefunktion, in der alle Variablen nur in ersten Differenzen verwendet werden, die tatsächliche Entwicklung nicht wiedergeben kann. Abschließend werden die ökonomischen Implikationen des Modells mit Hilfe dynamischer Multiplikatoren illustriert.

Summary

This paper studies the stability and dynamic properties of labour demand in West-German manufacturing with output, an index of total factor productivity and relative factor prices as independent variables. The model is estimated using the cointegration-techniques. Parameter stability is tested by Chow- and CUSUM-test. A dynamic simulation reveals that a labour demand function with all variables in differenced form is not able to track the actual employment behaviour. Finally, the economic implications of the model are illustrated by dynamic multipliers.

Sachs 1987). But these authors assume profit maximization which implies that the firms are never rationed in the output market. This seems to us a too strong assumption.

In this paper we therefore set up a simple neoclassical labour demand model derived from cost minimization, incorporating relative factor prices as well as the demand for manufacturing output. We estimate our model with quarterly data using the cointegration techniques recently proposed by Engle and Granger (1987). In a recent study for the U.K. Jenkinson (1986) used this technique to study labour demand in the manufacturing sector and failed to 'discover' (a specific form of) a neoclassical labour demand function. Although our approach is similar in spirit, it differs regarding econometric specification and details in the estimation procedure. Furthermore, the results of our investigation are quite different from those in the study mentioned.

Applying a two–step procedure to our model, we first estimate the long–run static labour demand function as derived from the assumed economic model. The adjustment of labour demand to its equilibrium level is considered in the second step of the estimation procedure, where we estimate the corresponding error correction model. We then compare these estimates with those obtained from an unrestricted error correction model and from an adjustment equation without levels information, respectively. We conclude that the adjustment model corresponding to the cointegrating regression produces the most reliable estimates. Using our preferred specification, we finally show how employment is adjusted to variations in output and relative prices.

II. ECONOMETRIC SPECIFICATION

Our theoretical setting is a cost–minimizing firm facing exogenously given factor prices and output demand. Total production costs are a function of the mentioned exogenous variables and an index of total factor productivity. The labour demand function can be derived by applying Shephard's Lemma: The optimal level of labour input is given by the derivative of the cost function with respect to the wage rate (see e.g. Varian 1978, p. 32). In order to get a tractable model, we assume that the labour demand function can be approximated by the following loglinear equation:

$$(1) \qquad E = a_1 \cdot \lambda + a_2\,(pim/w) + a_3\,(uc/w) + a_4\,Y\,.$$

The variables labour input E, output Y, the normalized price of intermediate input (pim/w) and normalized user costs of capital (uc/w) are all in logs. Since the input prices pim and uc are normalized by the wage rate w the labour demand equation is homogeneous of degree zero in all factor prices. λ is an unobservable index of total factor productivity, whose econometric implementation is discussed below.

Equation (1) is a simple neoclassical labour demand function with all inputs variable. We may therefore interpret it as an equilibrium relationship describing a firm's labour demand 'in the long run'. Given the output level, an increase in the wage rate relative to the price of any of the two other factors should reduce the demand for labour, and vice versa. A higher level of demand for goods will, other things equal, increase the demand for labour.

As we are interested in employment behaviour for the whole manufacturing sector, we have to assume that equation (1) remains valid at this level of aggregation. Furthermore, to interpret it as a labour demand equation, we make the realistic assumption that employment is determined by the demand side of the labour market.

As it stands, equation (1) is an equilibrium relationship which need not necessarily hold in the short run. In the literature on short run employment functions adjustment costs are usually invoked to rationalize short run deviations from long run equilibrium. It can be shown that under certain assumtions about adjustment costs optimizing firms will adjust gradually to their desired employment level (for a survey see Nickell 1986).

In econometric work the error correction model (ECM) has become a very popular specification to model dynamic adjustment processes. Following the work of Davidson et al. (1978), one would have an error correction term in the adjustment equation being composed of the variables of the model in levels. As the error correction term seems to have a natural interpretation as the amount the system is out of equilibrium, the ECM specification implies that a certain proportion of the disequilibrium from one period is corrected in the next.

Contrary to the traditional approach where all level information was differenced out in the adjustment equation, the ECM can be solved to obtain the equilibrium (long run) values of the model. A decisive advantage of the ECM over the traditional approach was seen in the possibility that it provided a dynamic specification ensuring a long run solution which refers to some kind of equilibrium relationship as derived from economic theory. Referring to our employment model, the ECM specification promises the possibility to model dynamic adjustment of employment under the restriction that our equilibrium labour demand model in equation (1) holds in the long run.

Error correction modelling of adjustment processes can theoretically be rationalized as the outcome of dynamic optimization in the presence of quadratic adjustment costs if the target variable has a unit root. In this case both the target and the control variable are integrated variables of order one, but a linear combination of the two variables is stationary (see Nickell 1985). The two variables are said to be cointegrated and – as Engle and Granger (1987) have shown – the adjustment process can be described by an error-correction model.

In this case a dynamic model fullfilling the restriction imposed by some equilibrium relationship may be estimated in two steps. In the first step the parameters of the levels equation can be estimated consistently by an ordinary least squares regression. In the second step the residuals from this regression can be used as an error correcting term in the dynamic adjustment equation, thus incorporating levels information without introducing nonstationarity. As this two-step procedure seems quite appealing, we use it to estimate our employment model set up above.

III. ESTIMATION

Our empirical work is based on the employment function in equation (1) which we estimate for the manufacturing sector of the W.-German economy for the period 1964:1 to 1986:4 using quarterly seasonally adjusted time series. The exact definitions of the variables and references to data sources are supplied in an appendix. As we are concerned with the whole manufacturing sector, only the price of imported raw materials seems relevant here.

In a first step we tested whether all variables in our model are integrated of the same order. As many macroeconomic variables seem to be integrated of order one (see Nelson and Plosser 1982, Wasserfallen 1986), we have tested for unit root in the autoregression. To test the null against the alternative that the series in the model are stationary around a deterministic trend, we estimated equations of the following form (see Dickey and Fuller 1981):

$$\Delta x_t = \text{constant} + \alpha\,\text{trend} + \beta x_{t-1} + \sum_{j=1}^{k} \gamma_j \Delta x_{t-j} + e_t \tag{2}$$

where x_t is the log level of the series and Δx_t is its first difference. We may conclude that x is integrated of order one if Δx is stationary, implying $\alpha = \beta = 0$.

The number of lags (k) in (2) was determined on the basis that the residuals in the equation are empirically white noise. The usual procedure to test for stationarity in the differenced series is to run an ordinary least squares regression on (2) to obtain a 't–value' for β. In most applications a time trend is neglected and the null is rejected if a negative and significant value turns up. This test statistic is not t–distributed under the null–hypothesis, but significance tests are possible using the critical values tabulated in Dickey and Fuller (1981, Table III, p. 1062). We resisted, however, to exclude a priori the possibility of $\alpha \neq 0$, so we tested the joint hypothesis $\alpha = \beta = 0$ and compared the test statistics obtained from an ordinary F–test with the values given in Dickey and Fuller (Table VI, p. 1063). The results of these tests are summarized in the following table.

TABLE 1

Testing for unit root in time–series

	t–stastic		F–statistic	number		
Variable	$\alpha = 0$	$\beta = 0$	$\alpha=\beta=0$	of lags	DW	$\chi(12)$
E	–2.32	–1.97	2.70	5	1.86	4.32
Y	1.07	–1.62	2.17	6	2.04	4.26
pim/w	–0.17	–2.06	2.81	1	1.89	10.27
uc/w	–1.51	–1.27	1.79	2	2.0	14.23

Period of estimation: 1962 Q4–1986:Q4

Note: χ (12) is the Box–Pierce statistic for serial correlation in the residuals based on 12 lags; the optimal lag length for each variable was chosen based on the AKAIKE criterion.

Comparing the test statistics in Table 1 with the corresponding critical values given by Dickey and Fuller for a sample size of N = 100, we could not reject the null hypothesis on the basis of either test. Our next step in the estimation procedure therefore was to test if the variables of our model are cointegrated.

At this point a discussion of the role of the productivity index λ is in order (for discussion of this problem see also Harvey et al. 1986). If λ is a deterministic function of time plus a stationary random element, it is obvious that the variables E, Y, (pim/w) and (uc/w) can (after detrending) be cointegrated. If λ behaves as a random walk (with drift), those variables are generally not cointegrated. A noticeable exeption arises from the "Common Trends Model" developed by King et al. (1987). In this case the observed variables E, Y, (pim/w) and (uc/w) all depend on the same technology factor. A priori this seems to be a quite realistic assumption in the present context, but ultimately the question whether the variables are cointegrated is an empirical one.

As cointegration refers to non–deterministic time series (Granger 1986, p. 216), we first removed the deterministic part of each variable by subtracting a linear time trend from the (log) levels, then we adjusted it for its mean. In what follows, we therefore work with seasonally adjusted and detrended time–series.

To test for cointegration one simply estimates equation (1) by ordinary least squares and tests whether the residuals of this regression are stationary. The tests usually employed are the cointegrating Durbin–Watson test (CRDW) and the ADF–test, discussed above (see Engle and Granger 1987 who also provide a number of alternative tests).

The parameters of a set of cointegrating variables can be estimated consistently despite the complete omission of the underlying dynamics. Theoretically, it can even be shown that the estimated parameters converge to the 'true' parameters of the model at a faster rate than ordinary least squares estimates (see Stock 1987).

Leaving aside the question of uniqueness of the cointegrating parameters for a moment, we now present the results of the cointegrating regression which we interpret as our long run employment function.

TABLE 2
Cointegrating Regression Employment Function
Dependent Vairable: Employment

Regressor	Coefficient
pim/w	–0.028
uc/w	0.129
Y	0.612
Summary statistics	
R^2(adj.)	0.86
s.e.e.	0.019
CRDW	0.81
ADF(4)	–3.92

Estimation period: 1963:4–1968:4

First we turn to the CRDW and the ADF statistics to see whether the variables of our model are in fact cointegrated. The CRDW tests against a random walk with drift and accepts the null of non–cointegration if the test statistic lies below the critical value. We do not know the exact critical values of the CRDW– and ADF–test for our four–variable model, but for the three–variable case they are reported in Hall (1986) as obtained by Granger from a simulation study. At the five percent level these critical values are given by 0.367 for the CRDW and – 3.13 for the ADF–test. Given this qualification, the value of either test statistic allows us to reject the hypothesis that our employment model represents a non–cointegrating vector of the relevant variables.

The estimated coefficients in Table 2 show the 'long run' – elasticities of employment with respect to relative factor prices and output, respectively. As the errors in this regression are strongly correlated, t–values are biased and hence are not reported. Not surprisingly, the output variable is the dominating force in employment determination. The long–run output elasticity is 0.61 indicating increasing returns to scale. Interestingly, in a related but somewhat different model Harvey et al. (1986) found almost the same value for U.K. manufacturing. Relative factor prices have only a modest effect. The coefficient of the normalized price of imported raw materials has a negative sign, indicating complementarity between labour and raw material, although to a rather small extent. Capital and labour seem to be substitutes, as indicated by the relatively strong positive effect normalized user costs have on employment. Finally, the 'long–run' own–price elasticity of employment turns out to be around 0.10. This is a rather low estimate compared with other time–series studies in this field (see e.g. Hamermesh 1986, Table 8.2). Note, however, that our estimates have been obtained for a given output level.

The value of 0.86 for the adjusted R^2 is rather high considering the detrended nature of the variables in the model. It must be noted, however, that it lies below the value of 0.95 sometimes claimed to be required for the desired properties of a cointegrating regression (i.e. rapid convergence of estimated to true parameter values, see e.g. Hendry 1986, p. 206) to be valid.

As cointegration of the variables of our model seems to obtain, we may use the residuals from the cointegrating regression to construct the corresponding error correction model, see Engle and Granger (1987) for discussion. In its general form this is given by

$$(3)\ \Delta E_t = \alpha_o\, RES_{t-1} + \sum_{j=1}^{k} \alpha_j \Delta E_{t-j} + \sum_{j=0}^{k} \beta_j \Delta Y_{t-j} + \sum_{j=0}^{k} \gamma_j \Delta\left(\frac{pim}{w}\right)_{t-j} + \sum_{j=0}^{k} \delta_j \Delta\left(\frac{uc}{w}\right) + \epsilon_t$$

where $RES_{t-1} = E_{t-1} + 0.028\,\left(\frac{pim}{w}\right)_{t-1} - 0{,}129\,\left(\frac{uc}{w}\right)_{t-1} - 0{,}612Y_{t-1}$.

The model in (3) differs from the conventional ECM specification in that the level term is restricted by the parameter vector obtained from the cointegrating regression in (1).

Before we present the results of various specifications of our dynamic employment model, we must briefly return to the problem mentioned above concerning the lack of uniqueness in the cointegrating relationship. We estimated the different inversions of equation (1) and tested each relationship for cointegration. It turned out that except for the equation with (pim/w) as the dependent variable, all inversions of equation (1) seem to be cointegrated.

Hence, to account for the existence of three cointegrating relationships in our model, equation (3) should have been extended to the corresponding multivariate representation. Although this is theoretically possible (see Engle and Granger 1987), for empirical applications the two–step procedure seems to get rather involved and therefore a specific inversion of the cointegrating regression is usually assumed on a priori grounds.

To see whether the cointegrating relationship in (1) can actually be treated as employment function, we added to equation (3) the error terms obtained from regressions on the different inversions of (1). It turned out that only the error from the employment function was statistically significant. In our opinion, this may warrant our interpretation of (3) as a dynamic employment model, although we realize that our model remains somewhat unsatisfactory in this respect.

Given this qualification, we now present the results for various specifications of our dynamic employment model in the following table.

TABLE 3:
Dynamic Employment Model, Dependent Variable: ΔE_t
Parameter estimates (t-statistic)

Regressor	Model 1		Model Specification Model 2		Model 3	
RES_{t-1}	-0.086	(2.2)	-----		----	
E_{t-1}	---		-0.070	(1.6)	---	
Y_{t-1}	---		0.030	(0.9)	---	
$(\frac{pim}{w})_{t-1}$	---		-0.003	(1.3)	---	
$(\frac{uc}{w})_{t-1}$	---		0.004	(0.4)	---	
ΔE_{t-1}	0.623	(5.6)	0.589	(4.8)	0.617	(5.4)
ΔE_{t-2}	-0.104	(0.9)	-0.093	(0.8)	-0.151	(1.3)
ΔE_{t-3}	-0.185	(1.7)	-0.189	(1.7)	-0.196	(1.8)
ΔE_{t-4}	0.696	(6.3)	0.685	(6.1)	0.701	(6.2)
ΔE_{t-5}	-0.337	(3.4)	-0.314	(3.0)	-0.408	(4.2)
ΔY_t	0.095	(4.7)	0.091	(4.3)	0.093	(4.5)
ΔY_{t-1}	0.074	(2.5)	0.091	(2.5)	0.115	(4.9)
ΔY_{t-2}	0.051	(1.8)	0.066	(1.8)	0.082	(3.1)
ΔY_{t-3}	0.018	(0.7)	0.027	(0.8)	0.038	(1.4)
ΔY_{t-4}	-0.044	(1.7)	-0.038	(1.4)	-0.032	(1.3)
ΔY_{t-5}	-0.045	(2.1)	-0.043	(1.9)	-0.037	(1.7)
$\Delta(\frac{pim}{w})_t$	-0.009	(1.2)	-0.009	(1.0)	-0.008	(1.1)
$\Delta(\frac{pim}{w})_{t-1}$	0.003	(0.3)	0.003	(0.3)	0.002	(0.2)
$\Delta(\frac{pim}{w})_{t-2}$	0.005	(-)	0.002	(0.2)	-0.001	(0.1)
$\Delta(\frac{pim}{w})_{t-3}$	0.012	(1.2)	-0.011	(1.0)	-0.013	(1.3)
$\Delta(\frac{pim}{w})_{t-4}$	0.008	(0.8)	0.009	(0.9)	0.009	(0.9)
$\Delta(\frac{pim}{w})_{t-5}$	0.008	(0.9)	0.010	(1.0)	0.010	(1.0)
$\Delta(\frac{uc}{w})_t$	0.006	(0.5)	0.003	(0.2)	0.002	(0.2)
$\Delta(\frac{uc}{w})_{t-1}$	-0.003	(0.3)	-0.0003	(0.01)	0.0005	(0.05)
$\Delta(\frac{uc}{w})_{t-2}$	-0.015	(1.5)	-0.013	(1.2)	-0.015	(1.4)
$\Delta(\frac{uc}{w})_{t-3}$	-0.002	(0.3)	-0.002	(0.2)	-0.003	(0.3)
$\Delta(\frac{uc}{w})_{t-4}$	-0.001	(0.1)	-0.0003	(0.01)	-0.001	(0.1)
$\Delta(\frac{uc}{w})_{t-5}$	-0.003	(0.3)	-0.002	(0.2)	-0.003	(0.3)
Summary statistics						
R^2(adj.)	0.85		0.84		0.84	
s.e.e.	0.0034		0.0035		0.0035	
DW	1.99		1.99		2.02	
Q(12)	8.5		8.2		13.3	

Period of estimation: 1964:1-1986:4

The models in Table 3 differ in the specification of the error correcting–variable, i.e. the form level information is contained in the dynamic adjustment equation. In Model 1 this variable is represented by the residuals obtained from the cointegrating regression, in Model 2 it is replaced by the lagged levels of the variables in the model, whereas no level information is included in Model 3. The coefficients of the lagged levels in Model 2 are all statistically insignificant, a F–test on these variables yielded a test statistic of 1.46 with (4,65) degrees of freedom. The insignificance of the level terms in Model 2 should not be too surprising in face of the severe multicollinearity in this regression. As dropping the level terms from the regression is known to be a completely inadequate way to deal with multicollinearity (see Hendry 1986, p. 207), we decided to leave them in the model.

Although the specifications in Table 3 may seem excessively over–parameterized, for the following reasons we decided not to exclude differenced regressors at insignificant lags from the model. First, on the basis of various test statistics five lags for employment and output were necessary to account for the autocorrelation in the residuals. This may indicate that our seasonal–adjustment procedure (see Appendix) did not completely remove seasonality from these series. Secondly, although a F–test on all price terms in the adjustment model did not show a significant effect of relative prices on employment (F(12,68) = 0.86), dropping prices from our model resulted in a relatively poor forecasting performance of this more 'parsimonious' model (see below). Thirdly, dropping individual regressors at insignificant lags may distort the autocorrelation structure in the data.

The estimated coefficients of the differenced variables are surprisingly similar in the different specifications of the dynamic employment model. In addition, regarding the summary statistics there is hardly a difference in the overall performance between these specifications. As we show below, the similarity of the dynamic structure of our alternative specifications does not remain, however, if the model is scrutinized more closely for structural stabiltiy and predictive performance.

Testing for structural stability

In view of two oil price shocks which presumably had a profound effect on the manufacturing sector of the national economy, it seems essential to test for structural stability of our dynamic labour demand model. We used the Chow–test for structural change in 1973:4 and 1979:4, respectively, as well as the Cusum– and the Cusum of Squares–test over the whole period. As the latter is known to reject the null of no structural change more often, we only report results for the Cusum of Squares–test.[1] The test–statistics for the Chow–test are given in Table 4, the graphs for the Cusum of Squares–statistic and the corresponding critical values are plotted in figure 2.

TABLE 4
Testing for stability of regression coefficients
Chow–test

Structural break in	Model 1 F(24,44)	Model 2 F(27,38)	Model 3 F(23,46)
1973:4	1,51 (0,12)	1,66 (0,08)	1,31 (0,21)
1979:4	0,74 (0.78)	0,65 (0.88)	0,86 (0.65)

Significance levels in parentheses.

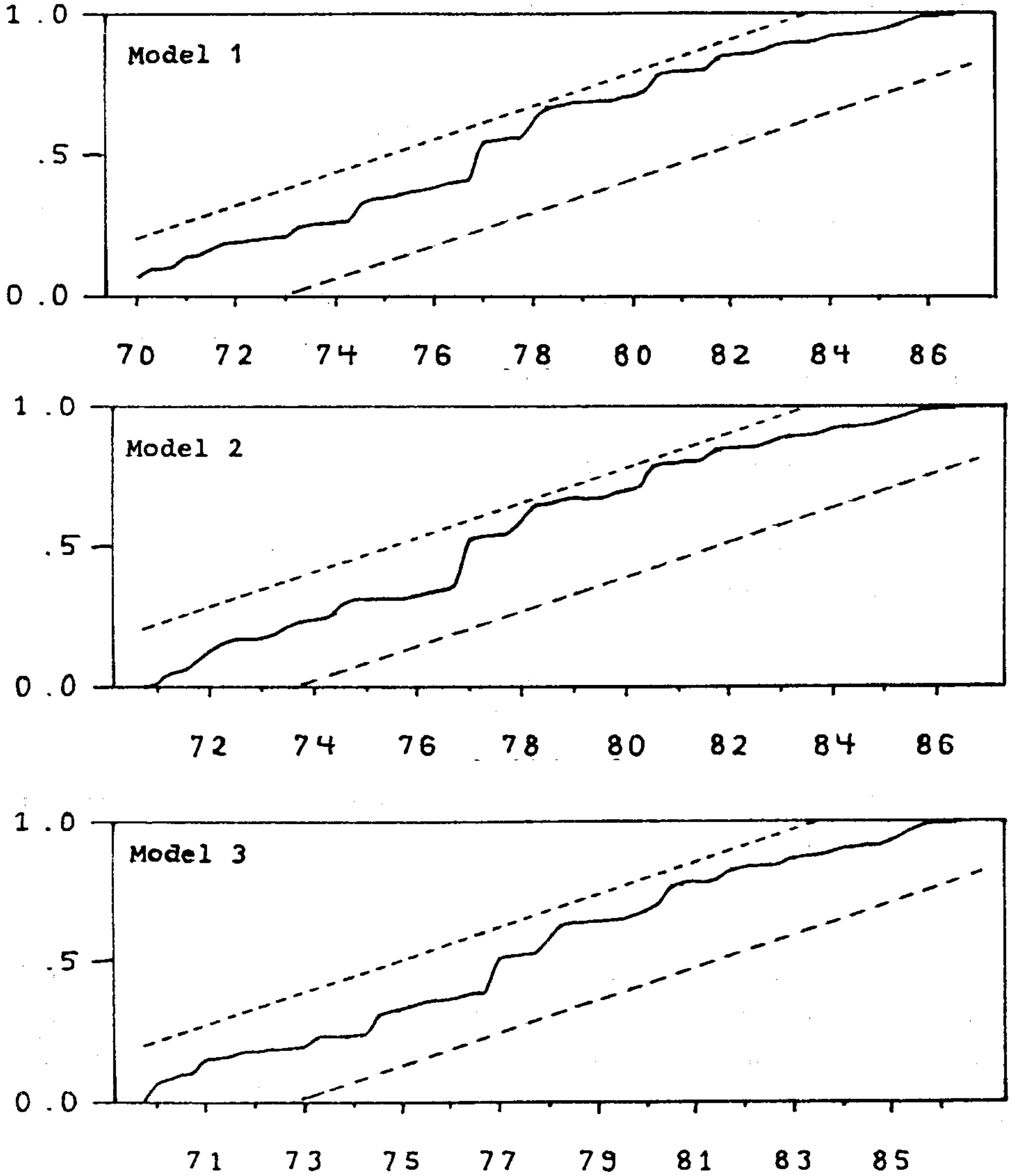
Comparing the test–statistics in Table 4 with the relevant critical values of the F–distribution shows no indication of parameter instability in either specification.

That this result is not just due to the selection of the specific reference periods, is revealed by the plots for the Cusum of Squares–test in Figure 2.

[1] For discussion of these tests see e.g. Johnston (1984), Krämer and Sonnberger (1986); for calculation of the recursive residuals required to construct the test–statistic of the Cusum and Cusum of Squares–tests we used the Kalman filter–procedure which is implemented in RATS.

FIGURE 2

Testing for stability of regression coefficients: Cusum of Squares–test



Note: The critical values for the test statistic were calculated for $\alpha = 10$ % using the formula and tabulated values given in Johnston (1984, Table B–8, p. 560).

Although the test statistics for Model 1 and Model 2 come very close to cross the upper critical line we may accept the hypothesis of parameter stability at a high level of significance. Surprisingly, Model 3 comes out best in this respect.

Testing for predictive performance

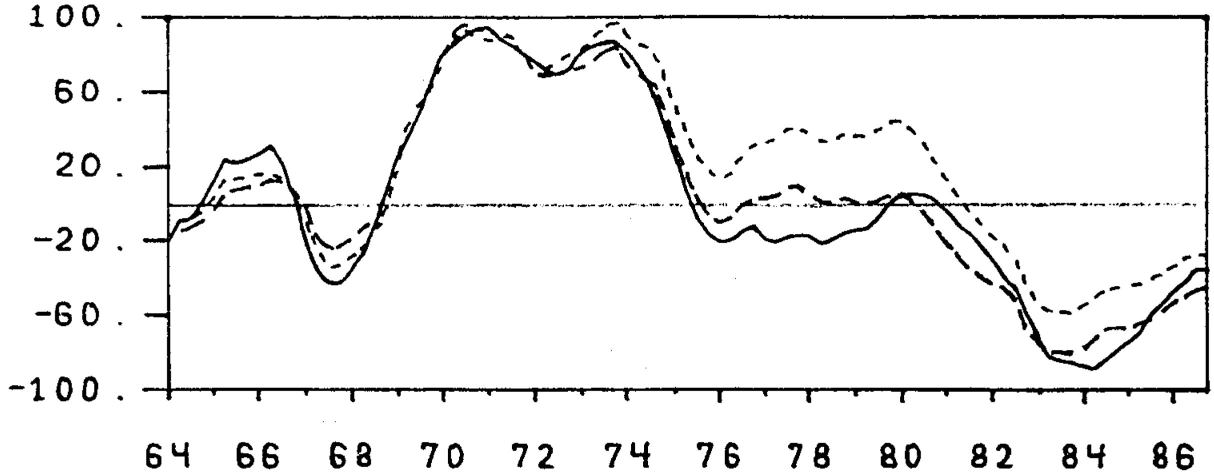
We performed dynamic simulation over the whole estimation period and ex post forecasting for the period 1982:1 – 1986:4 with the alternative specifications of our employment model. This certainly is a demanding test for the validity of a model. We must note, however, due to the structure of the model only employment and the cointegrating error are endogeneously determined by the dynamic simulation.

The results of the simulations are illustrated in Figure 3, where in the upper part we compare the behaviour of our employment series with the estimates obtained from Model 1 and Model 3, respectively. In the lower part of Figure 3, the simulation results for Model 1 and Model 2 are plotted against the employment series.

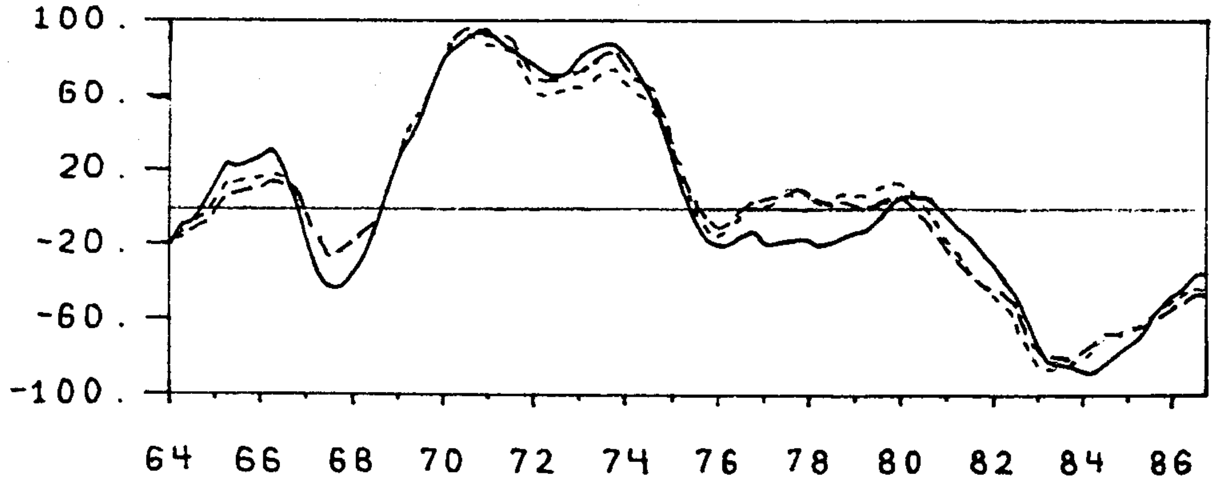
Apart from the second half of the 70s, Model 1 as well as Model 2 seem to perform quite satisfactory. Compared to these two specifications which incorporate levels information, the performance of Model 3 is distinctly worse. As can be seen from Figure 3(a) forecasts based on Model 3 drift apart after 1974 and need quite a long time to be corrected compared with Model 1. This seems to be caused by the exclusion of any level information in Model 3.

Referring to the lower part of Figure 3, the simulation results for Model 1 and Model 2 overall show no distinctly different pattern. In order to allow for a more precise comparison of these models' simulation performance, we calculated their root mean squared errors (r.m.s.e.) which are tabulated in the first line of Table 5. In the second line of this table we report the r.m.s.e obtained from ex post forecasts of our three specifications for the period 1982:1 – 1986:4, using parameter estimates up to 1981:4.

FIGURE 3
Predictive Performance of Dynamic Labour Demand Model
Dynamic Simulations 1964:1 – 1986:4



(a)
Model 1 (— —), Model 3 (- - -) and actual employment behaviour (——)



(b)
Model 1 (— —), Model 2 (- - -) and actual employment behaviour (——)

Note: The figures show detrended and seasonally adjusted series (Numbers of employees in thousands)

TABLE 5

Dynamic simulation and ex post forecasting

r.m.s.e.	Model Specification Model 1	Model 2	Model 3
dynamic simulation 1964:1 – 1986:4	0.0123	0.0123	0.0259
ex–post forcasting 1982:1 – 1986:4	0.0091	0.0089	0.0106

There is virtually no difference in forcasting performance between Model 1 und Model 2, whereas Model 3 performs relatively poor, as expected. It might be interesting to report that the exclusion of relative factor prices from the model, as possibly suggested by the lack of significance of the estimated parameters in Table 3, would have resulted in a very poor forcasting performance of our model. If we drop the price terms in Model 1 and calculate the r.m.s.e. for the dynamic simulation and the ex–post forcast, we obtain values of 0.0772 and 0.0955, respectively.

Long–run implications of the dynamic employment model

In evaluating a dynamic adjustment model, it is essential to take into consideration its long run implications. Missing any level information, Model 3 does not assure that there is an economically meaningful long run solution of the model. As shown above, on the basis of the dynamic employment model, it is hardly possible to discriminate statistically between the two specifications with level information.

It is worth mentioning, however, that they do have quite different long run solutions. Normalizing the estimated parameters of the levels in Model 2 for a unit coefficient on employment, the following long run employment function is implied by this specification:

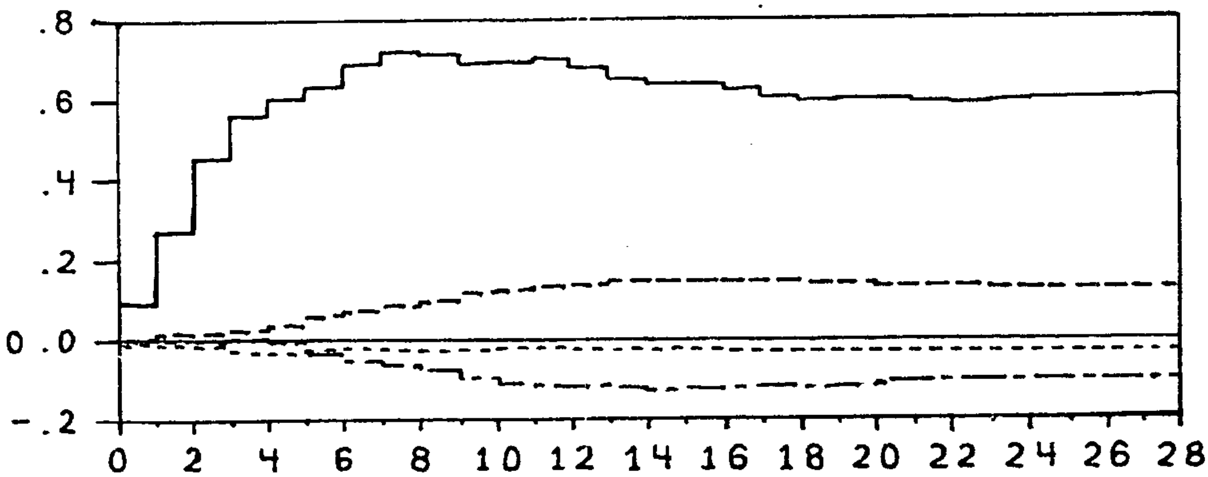
$$E = 0.429\,Y - 0.04\,(pim/w) + 0.057\,(uc/w). \quad (4)$$

Comparing the estimated coefficients to those with the cointegrating regression reveals marked differences between Model 1 and Model 2 with respect to the underlying equilibrium relationship. Due to the multicollinearity problem referred to above, the estimated parameters in equation (4) are not precisely determined. On this account we a priori consider the cointegrating parameters the more reliable estimates.

To illustrate the dynamic behaviour of our model, we finally calculate the adjustment process of the employment variable following a permanent increase in output and (input) prices, respectively. We performed this exercise with Model 1, so the long–run elasticities are, of course, given by the estimated parameters of the cointegrating regression.

In Figure 4 the simulated behaviour of our employment variable for an increase in the value of an exogeneous variable by 1 percent is plotted. We have already commented on the long run elasticities of our preferred model, so some remarks on the dynamics of the model may suffice here.

FIGURE 4
Dynamic adjustment of labour demand
to changes in exogenous variables



Response to output (——)
Response to user costs of capital (— —)
Response to price of raw materials (- - -)
Response to wage rate (—·—)

An increase in output achieves its largest impact on employment after seven quarters, then declines gradually and obtains its equilibrium level after about 18 quarters. Contrary to this adjustment path, an increase in the user costs of capital relativ to the wage rate shows its largest impact on employment after approximately four years and remains roughly constant thereafter. This long lag length does not seem implausible regarding the limited substitution possibilities for the existing capital stock. Much the same may be said of the adjustment of employment to an increase in the wage rate, other things equal. Finally, employment responds only moderately to an increase in the price of imported raw materials, the adjustment process being completed after roughly five quarters.

IV. CONCLUSION

In this paper we have applied the concept of cointegration to estimate a neoclassical labour demand model for the manufacturing sector of the W.–German economy. We showed that all variables of the model are cointegrated which allowed us to take advantage of the two–step procedure of Engle and Granger and to include level information in the error correction part of the model. In addition we estimated an unrestricted error correction model and an adjustment equation not including level information. In terms of predictive accuracy the latter specification performed relatively poor, although the usual test statistics do not indicate misspecification of the model. In contrast we could statistically not distinguish between the restricted and unrestricted specification of the error correction model, although their long run implications are somewhat different.

We argued, however, that the estimated cointegrating parameters of the employment function may be the more reliable estimates and illustrated the economic implications of our results by simulating employment behaviour following a change in relative factor prices and output, respectively. Not surprisingly, changes in output is the dominant force in determining employment behaviour, relative prices playing only a minor role. Most of the decline of employment in the manufacturing sector may therefore be attributed to a reduction in output. We realize, however, that our results should be interpreted with care, because of the partial nature of the model specification and the ambiguity of the two–step procedure in the case of more than two variables.

DATA APPENDIX

Quarterly data for the manufacturing sector of the W.–German economy all in logs, are used for the period 1960:1–1986:4. As we lose observations in constructing the user cost variable and in applying our seasonal adjustment procedure, our estimation period is 1964:1–1986:4.

(1)	E:		Employment (all employees); Source: OECD, Main Economic Indicators. Due to a change in the sample of reporting firms, there is a break in the original employment series. We adjusted the series by multiplying the data from 1960:1 to 1969:4 by the factor 1.0341. This factor was computed using Table 5.7 in Statistisches Bundesamt, Lange Reihen zur Wirtschaftsentwicklung 1984.
(2)	Y:		Index of industrial production (1980 = 100); Source: OECD, Main Economic Indicators. To account for an outlier in the series due to a strike in 1984:2, we used a strike dummy which takes the value 1 in 1984:2 and –1 in 1984:3.
(3)	pim:		Price index of imported raw materials and intermediate goods (1980 = 100); Source: Statistisches Bundesamt, Fachserie 17, Reihe 8.
(4)	w:		Hourly wage rates (1980 = 100); Source OECD, Main Economic Indicators.
(5)	uc:		User costs of capital = $q(r + d - wq^*)$;
		q	Price index of investment goods (equipment) (1980 = 100); Source: Deutsches Institut für Wirtschaftsforschung, Vierteljährliche Volkswirtschaftliche Gesamtrechnung, Sozialprodukt und Einkommenskreislauf.
		r	Interest rate (Current yield for long–term bonds); Source: Deutsche Bundesbank, Monatsberichte.
		d	Depreciation rate. d was interpolated from yearly date using a quadratic time trend. Source: Statistisches Bundesamt, Fachserie 18.
		wq^*	Expected inflation rate of investment goods. wq^* was constructed as follows: We estimated a VAR–system with 4 lags for the inflation rates of output goods and of investment goods. wq^* was constructed as the mean of the predicted values for q for the next two years in order to remove some erratic movements of the user costs series.

The original data for E, Y, w and q were seasonally unadjusted. We computed seasonally adjusted values of growth rates as the residuals from regressions on a constant and three seasonal dummies. Seasonally adjusted level values were obtained by accumulation of the adjusted growth rates.

REFERENCES

Artus, J. (1984). 'The Disequilibrium Real Wage Rate Hypothesis. An Empirical Evaluation', *IMF Staff Papers,* Vol. 31, pp. 249–302.

Burda, M. and Sachs, J. (1987). 'Institutional Aspects of High Unemployment in the Federal Republic of Germany', *NBER Working Papers No. 2241.*

Davidson, J., Hendy, D., Srba, F., Yeo, S. (1978). 'Econometric Modelling of the Aggregate Time–Series Relationship Between Consumers' Expenditure and Income in the United Kingdom', *Economic Journal,* Vol. 88, pp. 661–692.

Dickey, D. and Fuller, W. (1981). 'Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root', *Econometrica,* Vol. 49, pp. 1057–81.

Engle, R., Granger, C. (1987). 'Co–Integration and Error Correction: Representation, Estimation and Testing', *Econometrica,* Vol. 55, pp. 251–276.

Granger, C. (1986). 'Developments in the Study of Cointegrated Economic Variables'. *Oxford Bulletin of Economics and Statistics,* Vol. 48, pp. 213–28.

Hall, S. (1986). 'An Application of the Granger & Engle Two–Step Estimation Procedure to United Kingdom Aggregate Wage Data', *Oxford Bulletin of Economics and Statistics,* Vol. 48, pp. 229–39.

Hamermesh, D. (1986). 'The Demand for Labor in the Long Run', in Ashenfelter, O. and Layard, R. (eds.), *Handbook of Labor Economics,* Amsterdam etc., North Holland.

Harvey, A., Henry, S., Peters, S., Wren–Lewis, S. (1986). 'Stochastic Trends in Dynamic Regression Models: An Application to the Employment–Output Equation', *Economic Journal,* 96, pp. 975–85.

Hendry, D. (1986). 'Econometric Modelling with Cointegrated Variables: An Overview', *Oxford Bulletin of Economics and Statistics,* Vol. 48, pp. 201–12.

Jenkinson, T. (1986). 'Testing Neo–Classical Theories for Labour Demand: An Application of Cointegrating Techniques', *Oxford Bulletin of Economics and Statistics,* Vol. 48, pp. 241–251.

Johnston, J. (1984). 'Econometric Methods', Auckland etc., McGraw–Hill.

King, R., Plosser, Ch., Stock, J., Watson, M. (1987). 'Stochastic Trends and Economic Fluctuations', *NBER Working Papers No. 2229.*

Krämer, W. and Sonnberger, H. (1986). The Linear Regression Model under Test, Heidelberg, Physica.

Nelson, Ch., Plosser, Ch. (1982). 'Trends and Random Walks in Macroeconomic Time Series. Some Evidence and Implications'. *Journal of Monetary Economics,* Vol. 10, pp. 139–62.

Nickell, S. (1985). 'Error–Correction, Partial Adjustment and All That: An Expository Note', *Oxford Bulletin of Economics and Statistics,* Vol. 47, pp. 110–29.

Nickell, S. (1986). 'Dynamic Models of Labour Demand', in Ashenfelter, O., and Layard, R. (eds.), Handbook of Labor Economics, Amsterdam etc., North Holland.

Sachs, J. (1983). 'Real Wages and Unemployment in the OECD Countries', *Brookings Papers on Economic Activity*, Vol. 14, pp. 255–89.

Stock, J. (1987). 'Asymptotic Properties of Least Squares Estimators of Cointegrating Vectors', *Econometrica*, Vol. 55, pp. 1035–56.

Varian, H. (1978). 'Microeconomic Analysis', New York, Norton.

Wasserfallen, W. (1986). 'Non-stationarities in Macro-economic Time Series–Further Evidence and Implications', *Canadian Journal of Economics*, Vol. 19, pp. 498–510.

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