

# *Short-run Employment Functions for Irish Manufacturing\**

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*Abstract:* Cyclical fluctuations in employment are a major concern of policy makers. When an industry's output changes, the level of employment also changes—but usually only with a time lag. Sensible policy and economic planning by the government requires knowledge of the magnitude of the response in employment which will follow the output change and the speed with which it will take place. The extent of response and the speed of adjustment differs quite considerably from industry to industry. For total manufacturing only 25 per cent of the eventual adjustment is made in the first quarter, a further 25 per cent of the balance in the next quarter and so on. Taken industry by industry, speeds of adjustment are faster in Ireland than in the United Kingdom probably as the result of higher unemployment rates in Ireland.

THIS paper estimates short-run employment functions for the Irish manufacturing sector. The model used is a modification of the formulations employed in earlier studies. Estimates are made using quarterly data for total manufacturing and ten sub-classifications by means of the  $R$ 'th order least squares estimation procedure. The model is found to fit total manufacturing and most of the subclassifications quite well.

## *I. The Model*

In recent years there have been a number of studies of the relationship between employment and output in which desired employment is a function of output and actual employment is determined by the incorporation of a dynamic adjustment mechanism. These studies include, for the United Kingdom, Brechling (1965), Ball and St. Cyr (1966) and Scott and Smyth (1970); for the United States, Solow (1964), Soligo (1966), Fair (1969) and Smyth (1974*a* and *c*); for the German Federal Republic, O'Brien (1967); for Australia, Smyth and Ireland (1967), Norton

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and Sweeney (1971) and Smyth (1972); for New Zealand, Hazeldine and Woodfield (1971 and 1973); for Canada, Swan (1972); and for various OECD countries Brechling and O'Brien (1967) and Smyth (1974*b*). The models used have a common basis; by the introduction of assumption (ii) below most of the earlier models appear as special cases according to the value assigned to the parameter  $\gamma$ . This more general model was introduced in Smyth (1974*a*, *b* and *c*).

Our notation is as follows:

- $q$  = output,
- $e$  = number of employees,
- $e^*$  = desired number of employees,
- $h$  = hours worked per employee,
- $k$  = stock of capital,
- $x$  = rate of utilisation of capital stock,
- $a$  = shift variable indicating state of technology,
- $t$  = time period,
- exp = exponential.

Lower case letters denote actual values. Upper case letters will be used to denote natural logarithms so that

- $Q$  = logarithm of output,
- $E$  = logarithm of number of employees,
- $E^*$  = logarithm of desired number of employees.

Parameters will be described when they are introduced.

We use the Cobb-Douglas production function<sup>1</sup>

$$q_t = a_t(k_t x_t)^\beta (e_t h_t)^a \quad (1)$$

and make a number of operational assumptions.

(i) Output, capital stock and techniques of production are exogenous in the short-run.<sup>2</sup>

1. The assumption that the production function is Cobb-Douglas in form is not a crucial one—for instance, Ireland and Smyth (1970) use a constant elasticity of substitution production function. However, as the results are independent of the elasticity of substitution and as the Ireland-Smyth procedure does not yield an estimate of the elasticity we present the model in the expositionally simpler Cobb-Douglas form.

2. Relative factor prices do not appear explicitly in the model. The assumption that capital stock is exogenous for the purpose of the model and approximated by a time trend means that we are assuming that the effect of changes in relative factor prices on the optimal long-run capital stock-employment mix is a smooth one.

(ii) The rate of utilisation of capital stock and man-hours are related by

$$x_t = j_t(e_t h_t)^\gamma \quad j_t > 0, \gamma \geq 0 \quad (2)$$

This assumption is made because estimates of the rate of utilisation of capital stock are not available for Ireland and in any case the inclusion of  $k_t x_t$  in the same regression as  $e_t h_t$  or  $q_t$  leads to multicollinearity. The parameter  $\gamma$  measures the elasticity between the rate of utilisation of capital stock and man-hours worked. If  $\gamma = 0$  then the rate of utilisation of capital stock is given (at full utilisation or less) for  $x_t = j_t$ . This is implied in the work of Brechling (1965), Ball and St. Cyr (1966), Brechling and O'Brien (1967) and others. If  $0 < \gamma < 1$  the capital utilisation rate changes proportionately less than man-hours. If  $\gamma = 1$ ,  $x_t = j_t(e_t h_t)$  and capital utilisation and man-hours change proportionately: this case is considered by Smyth and Ireland (1967) and Ireland and Smyth (1970) and there is some recent support for this hypothesis for the United Kingdom—see Ireland, Briscoe and Smyth (1973). If  $\gamma > 1$  then the rate of capital stock utilisation varies proportionately more than man-hours. As total man-hours worked grows over time we expect a time trend in  $j_t$ .

Substituting for  $x_t$  from (2) in (1) gives

$$q_t = a_t(j_t k_t)^\beta (e_t h_t)^{\alpha + \beta\gamma} \quad (3)$$

(iii) Normal hours and the ratio of overtime to standard pay are either constant or vary smoothly with time so that desired employment,  $e_t^*$ , may be used instead of man-hours,  $e_t h_t$ . Any time trends in normal hours or in the ratio of overtime to standard pay are subsumed in a trend term. So too are any changes in the mix between different categories of labour with different productivities.<sup>3</sup> Denoting this trend term by  $m_t$ , we have

$$e_t h_t = m_t e_t^* \quad (4)$$

and substitution in (3) yields for the production function

$$q_t = a_t(j_t k_t)^\beta (m_t e_t^*)^{\alpha + \beta\gamma} \quad (5)$$

(iv) The growth of capital stock, technical progress and the trends in equations (2) and (4) may be represented by an exponential growth trend of the form

$$a_t(j_t k_t)^\beta (m_t)^{\alpha + \beta\gamma} = b \cdot \exp(\rho t) \quad (6)$$

Substituting for (6) and rearranging gives for the desired level of employment

$$e_t^* = [q_t / b \cdot \exp(\rho t)]^{1/(\alpha + \beta\gamma)} \quad (7)$$

3. Between clerical and manual, male and female, youth and adult, for instance.

Taking logarithms (denoted by upper case letters) we have

$$E_t^* = \frac{1}{\alpha + \beta\gamma} [Q_t - B - \rho t] \quad (8)$$

(v) The level of employment is determined by the simple adjustment process

$$E_t - E_{t-1} = \lambda(\eta E_t^* - E_{t-1}) \quad 0 < \lambda \leq 1, \eta \geq 1 \quad (9)$$

Desired employment,  $E_t^*$ , rises over time and we allow for the series not being stationary by introducing the expectancy element  $\eta$  thus assuming that firms extrapolate the desired level of employment; hence we assume that the short-run adjustment of employment is to  $\eta E_t^*$  where  $\eta \geq 1$ . The assumption  $0 < \lambda \leq 1$  indicates that a constant proportion of the logarithmic difference between the desired and actual levels of employment is eliminated in any time period.<sup>4</sup> Substituting for  $E_t^*$  from (8) into (9) and rearranging gives

$$E_t = \frac{\lambda\eta}{\alpha + \beta\gamma} [-B + Q_t - \rho t] + (1 - \lambda)E_{t-1} \quad (10)$$

We shall fit

$$E_t = a_0 + a_1 Q_t + a_2 t + a_3 E_{t-1} \quad (11)$$

where

$$a_0 = \frac{-\lambda B \eta}{\alpha + \beta\gamma} \quad (12)$$

$$a_1 = \frac{\lambda \eta}{\alpha + \beta\gamma} \quad (13)$$

$$a_2 = \frac{-\lambda \rho \eta}{\alpha + \beta\gamma} \quad (14)$$

$$a_3 = (1 - \lambda) \quad (15)$$

Hence

$$\lambda = 1 - a_3 \quad (16)$$

$$(\alpha + \beta\gamma)/\eta = (1 - a_3)/a_1 \quad (17)$$

$$\rho = -a_2/a_1 \quad (18)$$

$$B = -a_0/a_1 \quad (19)$$

4. We also attempted to fit a model in which  $\lambda$  was a function of the level of unemployment as in Smyth (1972, 1974a and b). The results were unsatisfactory and are not reported here.

It is not possible to untangle fully the estimate of  $(\alpha + \beta\gamma)/\eta$  and obtain estimates of the individual components  $\alpha$ ,  $\beta$ ,  $\gamma$  and  $\eta$ . Bearing in mind that quarterly data are used we would not expect  $\eta$  to be much greater than one—the average quarterly rate of growth of total manufacturing employment is about 1.5 per cent so that even if firms extrapolated growth as much as three quarters ahead  $\eta$  would not be above 1.05. Even if we take an arbitrary estimate of  $\eta$  we still do not obtain estimates of  $\alpha$  and  $\beta$  unless we are prepared to assume that  $\gamma$  is zero; in this case  $(1 - a_3)/a_1$  provides an estimate of  $\alpha$ , the returns to labour. As  $(1 - a_3)/a_1$  is widely interpreted in the short-run employment function literature as measuring returns to labour it follows that most earlier studies implicitly assume  $\gamma = 0$ . It is difficult to accept this interpretation of  $(1 - a_3)/a_1$  for earlier studies and for our results in this paper because of the pervasive tendency for the ratio to be markedly greater than unity which would imply that the marginal unit of labour is more productive than the average unit—see Ireland and Smyth (1970). If  $\gamma$  is put equal to one then (17) provides an estimate of  $\alpha + \beta$ , that is, of returns to scale. Such an interpretation is much more acceptable for previous studies than the returns to labour interpretation—see Ireland and Smyth (1970) and Ireland, Briscoe and Smyth (1973). But we shall see in Section 4 below that the estimated coefficients obtained are rather too high for this interpretation to hold for the Irish data.

## 2. The Data

Our study uses quarterly data covering the period 1959 (i) to 1971 (i). At the time the statistical analysis was undertaken the first quarter of 1971 was the latest quarter for which data were available. The source of our data is various issues of the *Irish Statistical Bulletin* and *Quarterly Industrial Production Inquiry*. The output series consists of an index of production with 1953 = 100. The employment series is the number of people employed in thousands. The classification of manufacturing industries is the same for employment as for output.

The industrial groups covered are the following: Food; Drink and Tobacco; Textiles; Clothing and Footwear; Wood and Furniture; Paper and Printing; Chemicals and Chemical Products; Clay Products, Glass, Cement, etc.; Metal and Engineering; Other Manufacturing Industries. Together these ten industries comprise Total Manufacturing.

The data are not seasonally adjusted so we include zero-one seasonal dummies in the regressions,  $S_1$ ,  $S_2$  and  $S_3$ , representing the first three quarters respectively.

## 3. Estimation Procedures

To estimate the short-run employment functions we use the  $R$ 'th Order Least Squares (RALS) package. RALS is a programme developed by D. F. Hendry which enables one to obtain consistent estimates for an equation with autoregressive errors and lagged endogenous variables. In addition to providing the ordinary least squares (OLS) estimates RALS also estimates for first order up to

rth order autocorrelation in the error term where  $r$  is specified. As we are dealing with quarterly data we consider the possibility of autocorrelation up to and including the fourth quarter. The principle behind the method is very simple and can best be illustrated by an example. Consider

$$Y_t = b_0 + b_1 Y_{t-1} + u_t \quad (20)$$

where

$$\left. \begin{aligned} u_t &= \xi u_{t-1} + \epsilon_t \\ E(\epsilon_t) &= 0 \\ E(\epsilon_t^2) &= \sigma^2 \\ E(\epsilon_t \epsilon_{t-1}) &= 0 \quad i \neq 0 \end{aligned} \right\} \quad (21)$$

The application of OLS to the above structural equation will yield inconsistent estimates due to the presence of autocorrelation with a lagged endogenous variable,  $Y_{t-1}$ .

Consider now the following transformation

$$Y_t - \xi Y_{t-1} = b_0(1 - \xi) + b_1 Y_{t-1} - b_2 \xi Y_{t-2} + u_t - \xi u_{t-1} \quad (22)$$

and hence

$$Y_t = b_0(1 - \xi) + (b_1 + \xi) Y_{t-1} - b_2 \xi Y_{t-2} + \epsilon_t \quad (23)$$

Notice that the error term of the transformed equation is non-autocorrelated so minimising  $\sum \epsilon_t^2$  will give consistent estimates. But note that  $b_1$  and  $\xi$  cannot be estimated directly by OLS since there is a non-linear restriction between the parameters.

The preceding example can be generalised to more variables and higher order autocorrelation—the programme handles up to a seventh order autoregressive structure. In addition likelihood ratios are constructed so that the order of autocorrelation may be tested by means of a  $\chi^2$  distribution. For large samples and autoregressive structure of low order the estimates will effectively be maximum likelihood estimates.

In the regression results presented we give the “best” obtained with the RALS programme, hence the order of the autocorrelation estimated varies with the equations. Here “best” is used in the sense that allowance is made for that degree of autocorrelation, which may be zero autocorrelation, which the programme suggested as most suitable for the data-set in question.

#### 4. Empirical Results

Table 1 presents the RALS regression estimates with standard errors in parentheses. Table 2 gives the derived estimates of  $\lambda$ ,  $(\alpha + \beta\gamma)/\eta$  and  $\rho$ .

TABLE I: *Rals regression results*

<i>Industry</i>	<i>Estimated regression coefficients and standard errors</i>										
	<i>Constant</i>	<i>Q<sub>t</sub></i>	<i>100<sub>t</sub></i>	<i>E<sub>t-1</sub></i>	<i>S<sub>1</sub></i>	<i>S<sub>2</sub></i>	<i>S<sub>3</sub></i>	<i>ε<sub>t-1</sub></i>	<i>ε<sub>t-2</sub></i>	<i>ε<sub>t-3</sub></i>	<i>ε<sub>t-4</sub></i>
Food	1.009 (0.450)	0.009 (0.062)	0.09 (0.08)	0.718 (0.103)	-0.032 (0.007)	-0.089 (0.020)	-0.029 (0.008)	—	—	—	—
Drink and tobacco	0.708 (0.347)	0.069 (0.046)	0.08 (0.04)	0.561 (0.131)	-0.043 (0.007)	-0.046 (0.006)	-0.039 (0.005)	—	—	—	—
Textiles	1.177 (91.153)	0.000 (0.079)	0.21 (0.16)	0.607 (30.611)	0.008 (0.039)	0.004 (0.300)	-0.006 (0.234)	0.610 (0.305)	—	—	—
Clothing and footwear	0.924 (0.641)	0.186 (0.051)	-0.53 (0.74)	0.412 (0.216)	-0.017 (0.008)	-0.005 (0.006)	-0.009 (0.006)	0.357 (0.278)	0.347 (0.189)	-0.005 (0.215)	—
Wood and furniture	0.313 (0.258)	0.304 (0.065)	-0.03 (0.07)	0.112 (0.189)	0.001 (0.007)	0.016 (0.008)	-0.004 (0.007)	0.490 (0.250)	-0.155 (0.217)	-0.126 (0.201)	-0.230 (0.166)
Paper and printing	0.448 (0.298)	0.095 (0.038)	-0.02 (0.05)	0.660 (0.111)	-0.005 (0.009)	-0.011 (0.009)	-0.024 (0.009)	—	—	—	—
Chemicals and chemical products	0.657 (0.253)	0.241 (0.032)	0.40 (0.15)	-0.134 (0.132)	-0.019 (0.009)	0.016 (0.010)	-0.009 (0.013)	0.605 (0.210)	0.396 (0.212)	-0.261 (0.201)	-0.169 (0.165)
Clay products, glass, cement, etc.	0.625 (0.387)	0.211 (0.041)	0.06 (0.26)	0.048 (0.193)	0.000 (0.014)	-0.009 (0.015)	-0.015 (0.013)	0.431 (0.250)	—	—	—
Metal and engineering	0.157 (0.173)	0.132 (0.034)	-0.01 (0.06)	0.748 (0.059)	0.003 (0.006)	-0.008 (0.006)	-0.004 (0.006)	—	—	—	—
Other manufacturing	0.215 (0.250)	-0.008 (0.058)	0.14 (0.13)	0.915 (0.098)	0.009 (0.009)	0.019 (0.008)	-0.003 (0.009)	—	—	—	—
Total manufacturing industries	0.711 (1.039)	0.130 (0.041)	-0.06 (0.13)	0.739 (0.198)	-0.012 (0.004)	-0.018 (0.003)	-0.015 (0.003)	0.239 (0.247)	0.380 (0.207)	0.152 (0.179)	-0.344 (0.175)

TABLE 2: *Derived parameter estimates*

	$\lambda$	$(\alpha + \beta\gamma)/\eta$	$\rho$
Food	.282	30.000	-.0957
Drink and tobacco	.439	6.403	-.0117
Textiles	.393	—	—
Clothing and footwear	.588	3.170	.0286
Wood and furniture	.888	2.917	.0010
Paper and printing	.340	3.595	.0021
Chemicals and chemical products	1.134	4.709	-.0167
Clay products, glass, cement, etc.	.952	4.520	-.0028
Metal and engineering	.252	1.908	.0008
Other manufacturing industries	.085	-10.872	.1795
Total manufacturing industries	.261	2.019	.0046

The regression coefficient for  $Q_t$  is positive and more than twice its standard error for Total Manufacturing and six sub-classifications. The estimated reaction coefficient (given by one minus the estimated coefficient of  $E_{t-1}$ ) differs from zero by more than two standard errors for Total Manufacturing and all sub-classifications except for the Other Manufacturing industries grouping. The reaction coefficient is less than unity by more than two standard errors for Total Manufacturing and all sub-classifications except Clothing and Footwear, Wood and Furniture, Chemicals and Chemical products and Clay products, Glass, Cement, etc. For these four industrial groupings we conclude that the reaction coefficient is not significantly different from unity implying complete adjustment during the unit time period, one quarter.

It would seem that the model works satisfactorily for Total Manufacturing and six sub-classifications. Because of the non-significance of the coefficient of  $Q_t$  it fails for the Food, Drink and Tobacco, Textiles and Other Manufacturing groupings. The failure of the model for the heterogeneous Other Manufacturing grouping should not surprise us. In the studies by Ball and St. Cyr (1966) and Smyth and Ireland (1967) for the United Kingdom and Australia respectively it was found that the coefficients of  $Q_t$  for composite Food, Drink and Tobacco groupings were not significant so the Irish results are similar to these. However, in both countries the grouping including Textiles yielded significant results, in contrast to the negative results obtained here for Textiles. Overall the significance of the results appears to be comparable with those obtained for the United Kingdom and Australia industry studies.

Ruling out the non-significant regressions we see that the estimated reaction coefficients vary from 0.252 for Metal and Engineering to 1.134 for Chemicals and Chemical Products and that the estimated values of  $(\alpha + \beta\gamma)/\eta$  range from 1.908 for Metal and Engineering to 4.709 for Chemicals and Chemical Products.



The values of  $(\alpha + \beta\gamma)/\eta$  are so high that we can rule out the  $\gamma = 0$  possibility and probably also the  $\gamma = 1$  possibility in favour of the hypothesis that the percentage change in the rate of capacity utilisation is greater than the percentage change in man-hours.

There is no systematic tendency for the exponential trend (measured by  $\rho$ ) to be either positive or negative. The trend is a combination of a number of trends and we have no hypothesis on its sign.

We find fourth order autocorrelation for three series, Total Manufacturing, Chemicals and Wood and Furniture, third order autocorrelation for Clothing and Footwear, first order autocorrelation for Textiles and Clay Products, Glass, Cement, etc., and no autocorrelation for the remaining five industry classifications.

There exist two earlier studies that have estimated short-run employment functions for Ireland—O'Herlihy (1966) and Brechling and O'Brien (1967)—and a comparison of our results with those obtained there is in order.

O'Herlihy's analysis is in terms of actual values, not logarithms.<sup>5</sup> He made employment a function of output and lagged values of output, time and lagged employment. His data, consisting of quarterly observations measured in deviations from seasonal means for 1954 to 1963 were for total transportable industries. O'Herlihy found a long-run elasticity of about 0.6 between employment and output which implies, in our notation, a value of  $(\alpha + \beta\gamma)/\eta$  of about 1.7. Our estimate, for total manufacturing is approximately 2. O'Herlihy's reaction coefficient (applied to actual values not logarithms) is about one-third, our estimate is approximately one-quarter. The magnitude of the time trends in the regressions cannot be compared directly but both O'Herlihy's and ours are negative.

Brechling and O'Brien (1967) included Ireland in their study of OECD countries. They used OECD data covering SIC groups 20-38 for a period extending from the second quarter of 1952 to the second quarter of 1964 and, as we did, included seasonal dummies and estimated their model with data transformed into logarithms. Their results imply an estimate of  $(\alpha + \beta\gamma)/\eta$  of about 1.5 and an adjustment coefficient of about 0.3 which are of the same order as our estimates and those of O'Herlihy. Brechling and O'Brien's results yield an estimate of  $\rho$  of .0031; our estimate is .0046.<sup>6</sup>

O'Herlihy's study differs from ours with respect to his use of actual values rather than logarithms, his inclusion of more than one output term, his seasonal adjustment procedure and his time period. Brechling and O'Brien's period is different to ours. In addition both O'Herlihy and Brechling and O'Brien use ordinary least squares while we use  $R$ th order least squares. In view of these differences it is encouraging that the various results obtained are fairly close.

Finally, it is of interest to compare our estimates with those for the United Kingdom by Ball and St. Cyr (1966). The comparison is presented in Table 3—the

5. Brechling (1965) made estimates for the UK with his variables in actual values as well as logarithms.

6. We have converted the figure given by Brechling and O'Brien as their estimates were made using logarithms to base 10 while ours are to base  $e$ .

TABLE 3: Comparison of derived parameters, Ireland and the United Kingdom

	$\lambda$		$(\alpha + \beta\gamma)/\eta$	
	Ireland	UK	Ireland	UK
Clothing and footwear	.588	.240	3.170	1.730
Wood and furniture	.888	.264	2.917	2.650
Paper and printing	.340	.119	3.595	2.166
Chemicals and allied industries	1.134	.035	4.709	.327
Clay products, glass, cement, etc.	.952	.226	4.520	1.982
Metal and engineering	.252	.193	1.908	1.422
Total manufacturing industries	.261	.180	2.019	1.165

non-significant Irish industries are excluded.<sup>7</sup> The pattern is very consistent. For all industry comparisons the estimated values of  $(\alpha + \beta\gamma)/\eta$  are larger for Ireland than for the United Kingdom so that either there are greater returns to scale in Ireland or the elasticity of the rate of utilisation of capital stock to man-hours is much larger in Ireland or a combination of these.<sup>8</sup> Also the reaction coefficients are all larger for Ireland indicating faster adjustment in Ireland. A possible explanation here, lies in the higher unemployment rates to be found in Ireland. In an inter-country comparison Hughes (1971) has shown a tendency for speeds of adjustment to be positively correlated with unemployment rates.<sup>9</sup>

##### 5. Conclusions

A model in which actual employment is linked to desired employment by a dynamic adjustment mechanism was derived from a production function. The model was fitted to quarterly data for Total Manufacturing and ten sub-classifications using the RALS procedure. For Total Manufacturing and six industry groupings the model performed well. Accordingly, we conclude that the short-run employment function approach that has proved useful in earlier studies is of value in analysing short-run employment movements in the Irish economy.

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7. The UK results are from Table I of Ball and St. Cyr (1966). The UK classification is fairly comparable to the Irish classification but not identical. The UK series used in successive rows of Table 3 are as follows: Textiles, Leather, Clothing and Fur; Timber, Furniture, etc.; Paper, Printing and Publishing; Chemical and Allied industries; Bricks, Pottery, Glass and Cement; Total of Engineering and Allied industries; Total All Manufacturing industry.

8. The expectations coefficient,  $\eta$ , may also differ between the two countries but the differences found for  $(\alpha + \beta\gamma)/\eta$  are too marked and consistent to be fully accounted for by this.

9. Note, however, that the correlation coefficient between average unemployment of an industry and its reaction coefficient over the six sub-classifications for which the model is satisfactory is only 0.31.

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