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Price Determination in Ireland: Effects of Changes in Exchange Rates and Exchange Rate Regimes*

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Abstract: This paper finds that the wholesale price of Irish manufacturing output is closely tied to foreign prices in the long run; domestic wage costs generally do not exert a significant influence. In the short run changes in foreign currency prices affect Irish prices more rapidly than do changes in exchange rates. There is evidence of a change in pricing behaviour on Ireland's entry into the EMS. German output prices are now as important as UK prices in determining Irish prices.

I INTRODUCTION

It is well established by now that short-run purchasing power parity does not hold between major countries (see, for example, Frenkel, 1981 or Isard, 1977). This means that one cannot define an empirical counterpart to the single world price of simple small open economy (SOE) models. Several approaches can be adopted to dealing with this problem of modelling the influence of foreign prices on the domestic prices of a small country. When Ireland's exchange rate was fixed at parity with sterling the dominant approach adopted in empirical work was the use of prices in the UK, Ireland's largest trade partner, to proxy world prices. The alternative, of using a trade-weighted basket of the prices of Ireland's major trade partners, has also been given considerable attention.

In this paper we reassess these results, allowing for differences in the short-run response to exchange rates and foreign prices, and for possible changes

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in the relative importance of UK and German prices associated with Ireland's membership of the European Monetary System (EMS). We use quarterly data on the wholesale prices of the output of manufacturing industry. Wholesale prices do not include VAT or retail distribution costs, so they provide more direct evidence on the influence of foreign prices than consumer prices, which also include a much larger non-traded component. We allow for the possibility of differential response to foreign currency price changes and exchange rate changes. We investigate possible changes in the relative importance of different countries, or other changes in the influence of foreign prices arising from or associated with Ireland's EMS membership. Specifically, we examine econometric tests of the following questions:

(1) Are Irish prices proportional to foreign prices in Irish currency in the long run?
(2) Do domestic unit wage costs exert at least a short-run influence on prices, as a mark-up model of pricing would suggest?
(3) Do Irish prices react similarly in the short run to changes in foreign currency prices and to exchange rate changes, or is there a differential response?
(4) Is there evidence of an asymmetric response of Irish prices to increases and decreases in foreign prices?
(5) Which specification of foreign price variables performs best: UK prices alone, a trade-weighted index, or the prices of several major competitors?
(6) Has entry to the EMS led to a change in the relative influence of different countries, or to any other changes in the relationship between foreign and domestic prices?

The remainder of the paper is stuctured as follows. A brief review of previous work on price determination in Ireland is given in Section II. Section III describes the basic model and Section IV gives results on the six main questions outlined above, based on analysis of wholesale price indices for manufacturing industry treated as an aggregate. Section V deals briefly with the results of applying the basic model to the data on the output prices of individual industrial sectors. Section VI draws together the main conclusions.

II REVIEW OF PREVIOUS WORK

In the 1970s a number of studies of price determination in Ireland, using data from the sterling link period, found that Irish prices closely followed prices in the UK (Geary, 1976a and b, Geary and McCarthy, 1976, and Bradley, 1977). The evidence presented by Bradley indicated that consumer prices in Ireland followed movements in prices in the UK within 3 to 5 months.

The conclusions from this work, as summarised by Honohan and Flynn (1986), were that "Irish price inflation was determined by, and more or less equal to, UK inflation." (p. 175).

This evidence led to a widespread expectation that EMS membership would rapidly result in the Irish rate of inflation being determined in a similar fashion by German inflation. The experience of EMS membership has been rather different. Whereas, in the sterling link period, there was no uncertainty concerning the future value of the sterling exchange rate, the relationship between the Irish pound and the Deutschmark has not remained fixed over time and firms have been faced with uncertainty about future rates of exchange between the Irish pound and other EMS currencies. The uncertainty concerning the rate of exchange between the Irish pound and sterling, which has remained outside the EMS, has been even greater. The pattern of Irish inflation since joining the EMS cannot be explained by the simple SOE model of price determination: it is no longer closely tied to the UK rates, and it has not mirrored German rates. At first sight, this suggested a change in the foreign influence on Irish prices. However, Honohan and Flynn's analysis of the 1972-1984 period showed that a more flexible dynamic specification of the influence of foreign prices, using a trade-weighted index rather than simply UK prices, could yield an adequate model for both the sterling link and EMS sub-periods.

Neither Honohan and Flynn's trade-weighted approach, nor the earlier work based on UK prices, examined the question of whether the relative importance of different countries' prices in determining Irish prices has changed over time, in particular in response to Ireland's membership of the European Monetary System (EMS). The use of trade weights, while valuable for many purposes, implies that the relative importance changes if and only if trade patterns change, thus simply imposing an answer to the question. The use of UK prices does not even allow this limited amount of flexibility.

FitzGerald (1983) explored the issue using data disaggregating foreign price influence by country. However, that analysis was limited by the small number of post-EMS observations. The approach adopted there, and explored in more detail here, was to regress Irish output prices against the prices of Ireland's major competitors, notably the UK, Germany and the USA, allowing the relative importance of different countries to be more directly determined from the data. *A priori* one would expect considerable difference in behaviour between domestic Irish firms and foreign-owned firms, and also between different sectors. In order to examine this possibility, we repeat the analysis for selected sectors of manufacturing industry (as did FitzGerald), some of which are dominated by multinationals, and others by domestic firms.[1]

1. However, there are some limitations on the available data, which will be specified later.

FitzGerald also found a stronger response to sterling price changes in the period 1979-81 than to sterling exchange rate changes, although attempts to disaggregate the exchange rate and foreign currency price changes for other countries gave less satisfactory results. This analysis strongly suggested further investigation of differential effects of exchange rate and foreign currency price changes, a topic neglected in other analyses of Irish price determination.

Leddin's (1988) more recent analysis has investigated some of the issues addressed here, also using wholesale price index data for manufacturing industry. However, Leddin's specification imposed some crucial restrictions[2] which our tests indicate are rejected by the data. His conclusion, that short-run purchasing power parity does not hold, is supported by our more extensive tests, but our results show that a long-run version of purchasing power parity (not tested by Leddin) may hold.

III BASIC MODEL AND THE DATA

We begin with two alternative general models which express the long-run determination of the level of prices in Ireland; these two models represent the polar cases of the pure SOE model and the closed economy alternative where prices are determined as a mark-up on input costs.

$$P^*_{IR} = f_1(P_i R_i; i = 1, n) \tag{1}$$

$$P^*_{IR} = f_2(W_j; j = 1, m) \tag{2}$$

where

P^*_{IR} = the long-run desired wholesale price for Irish manufacturing output
P_i = the wholesale price for manufacturing of country i
R_i = the rate of exchange between Ireland and country i
W_j = the price index for input factor j.[3]

If prices are purely externally determined, as in Equation (1), then, assuming that purchasing power parity holds in the long run, Irish prices will be some weighted average of the prices of a possible range of other countries, all expressed in Irish pound terms. If, on the other hand, they are purely determined as a mark-up on domestic costs, then Equation (2) is appropriate. In examining the behaviour of output prices we test for whether either of these two long-run models of price determination on their own or, alternatively,

2. These include, for example, the imposition of identical coefficients on foreign currency prices and exchange rates.

3. Unit wage costs are used as a proxy for domestic factor input costs.

some mixture of the two is appropriate. This is done by combining Equations (1) and (2) into the composite Equation (3).

$$P^*_{IR} = f_3(P_iR_i, W_j; i = 1, n \text{ and } j = 1, m) \tag{3}$$

Taking Equation (3) as the underlying long-term relationship determining Irish output prices, we allow for differential short-run effects from the prices and exchange rates of the UK, Germany (GR) and the US, and from the price of domestic factor inputs. This is implemented using an error-correction type term which allows a long-run influence from prices in each of these countries and, possibly, from domestic costs; the only restriction imposed at this stage was that of long-run homogeneity, tested later. The structure of the model is, therefore, similar to that estimated on trade-weighted data by Honohan and Flynn (1986). The model can be written as follows where L stands for the lag operator, P_{IR} for the Irish price, P_j for the price of country j, R_j for the bilateral exchange rate with respect to country j, W for unit wage costs, and α, β_{ij}, γ_{ij}, ϵ_i, η_j and λ are all parameters to be estimated:

$$\begin{aligned}\Delta\log(P_{IR}) = {} & \sum_{i=0}^{3} \sum_{j=1}^{c} \beta_{ij}L^i \{\Delta\log(P_j)\} + \sum_{i=0}^{3} \sum_{j=1}^{c} \gamma_{ij}L^i \{\Delta\log(R_j)\} \\ & + \sum_{i=0}^{3} \epsilon_i L^i \{\Delta\log(W)\} + \lambda L^4 \Bigg\{ \sum_{j=1}^{c} \eta_j \log(P_j \times R_j) \\ & + \left(1 - \sum_{j=1}^{c} \eta_j\right) \log(W) - \log(P_{IR}) + \frac{\alpha}{\lambda} \Bigg\}\end{aligned} \tag{4}$$

The β_{ij}, γ_{ij}, and ϵ_i terms give the short-run responses to changes in foreign currency prices and exchange rates in the three countries and to domestic unit wage costs, while the last term is an error-correction mechanism (ECM) by which Irish prices are brought into a long-run equilibrium ratio with a weighted average[4] of the Irish currency prices of the three countries and labour costs. The constant $\frac{\alpha}{\lambda}$ in the ECM term allows for this equilibrium ratio to be some constant other than unity; in the estimated equations we report the coefficient for α. This model allows for deviations from purchasing power parity in the short run. Homogeneity in the long run is tested in a later section by adding an unrestricted term in $L^4(\log(P_{IR}))$. Appropriate seasonal dummies were included in all the equations.

4. where the η_j are weights.

The data on the wholesale price of the different categories of manufacturing industry output, including total output, are taken from the CSO databank EOLAS. The data on unit wage costs for each sector are derived from a similar source. The data on exchange rates are partly taken from the EOLAS databank and partly from the Department of Finance databank of series from the OECD *Main Economic Indicators* publication. The output prices for each of the industrial sectors in the other countries are taken from Department of Finance databank of series from the OECD *Quarterly Indicators of Industrial Activity* publication. The exchange rate data for the three main countries (the USA, the UK, and West Germany) are period averages whereas the exchange rate data for certain other countries tried in the disaggregated equations are averages of end monthly values.[5] The availability of the wholesale price indices basically determined the data period as the first quarter of 1975 to the third quarter of 1987. The estimation period, unless otherwise indicated, is the first quarter of 1976 to the third quarter of 1987.

IV RESULTS FOR TOTAL MANUFACTURING INDUSTRY

4.1 *Preliminary Tests of Co-integration*

Before estimating the model (Equation (4)) we check that the time series properties of the data satisfy the constraints which Engle and Granger (1987) have shown are implied by the model. The basic procedure suggested is in two steps. The first step is to test that the individual series share a common order of integration (i.e., are each stationary after differencing the same number of times); the second step is to test that the series move together in the long run (i.e., are "co-integrated"[6]).

Since the models ultimately used in this section involve just UK and German prices as explanatory variables, they imply a constant long-run relationship between Irish prices and a weighted average of UK and German prices, on a common currency basis. Therefore, we examine first of all the order of integration for Irish, UK and German prices, all in Irish pound terms. We test the hypothesis that the series is a random walk, using the test statistics derived by Dickey and Fuller (1979). The hypothesis that the series is a random walk without drift (so that the series is "integrated of order 1") is tested by Φ_2, while Φ_3 tests the hypothesis that the first differenced series is a random walk with drift.

The test statistics for the full period, and for the post-EMS period are shown in Table 1.

5. The data available in the databank.

6. See Thom, this *Review*, or Engle and Granger for the precise definitions of these terms, and more detail on the procedures.

Table 1: *Augmented Dickey-Fuller Test Statistics for Hypothesis that Irish Pound Price Series are Integrated of Order 1*

	Full Period (45 obs.)		*EMS Period (33 obs.)*	
	Φ_2	Φ_3	Φ_2	Φ_3
$\text{Log}(P_{IR})$	2.45	2.55	1.40	1.34
$\text{Log}(P_{UK} \times R_{UK})$	2.78	2.58	4.47	5.55
$\text{Log}(P_{GR} \times R_{GR})$	3.72	0.92	3.98	3.16
Critical values from Dickey and Fuller (1981):				
	50 obs.		*25 obs.*	
5%	5.13	6.73	5.68	7.24
10%	4.31	5.61	4.67	5.91

These test statistics do not reject the hypotheses that each of the Irish pound price series is integrated of order 1, even at the 10 per cent level. However, the results may be sensitive to the periodicity of the data, or perhaps to the exact sub-period chosen, as the results of Thom elsewhere in this *Review* might suggest. The statistics also appear to be quite sensitive to certain choices which must be made within the test procedure, an issue to which we now turn.

The augmented Dickey-Fuller regression from which these statistics are derived involve, for the log (P_{IR}) series, lagged terms in $\Delta\log(P_{IR})$. Granger (1986) indicates that the order of the autoregressive terms should be large enough that the residuals from the augmented Dickey-Fuller regression are empirical white noise. The best procedure for establishing the order of this autoregression has yet to be established. The procedure adopted here was to start with four lags, because the quarterly price data are known to involve seasonal factors, and add further lags if the F-form of the Lagrange Multiplier test indicates autocorrelation of the residuals at the 5 per cent level. In all cases where additional lags were introduced, they had highly significant t-statistics, and the F-test no longer indicated the presence of autocorrelation. If the order of the autoregression was instead set and maintained at four, the hypothesis that the Irish pound price of German goods is integrated of order 1 would be rejected for the full period, though not for the post-EMS period.

Given that the Irish, German and UK price series (converted to Irish pounds) may be regarded as integrated of a common order, we wish to test the hypothesis that they share a common trend. Most tests of this type of hypothesis (such as those synthesised in Engle and Granger, 1987) have concentrated on bivariate relationships, using a simple "co-integrating regression" which regresses one variable on the other plus a constant term. Some test statistics for higher order cases have been produced by Engle and Yoo (1987), using a multivariate co-integrating regression.

We examine first of all whether Irish prices share a common trend or long-run equilibrium relationship with either UK or German prices (in Irish pound terms), and then whether there is a common long-run equilibrium relationship between Irish prices and some weighted average of UK and German prices. The test statistic is based on a regression of the first differences of the residuals from the co-integrating regression on autoregressive terms, together with the lagged value of the levels of these residuals (again, see Thom, this *Review*, or Engle and Granger, 1987, for details). The choice of the number of auto-regressive terms to be included in this regression raises issues parallel to those discussed above, and the sensitivity of the results is similar. The hypothesis of co-integration is rejected if the calculated t-statistic for the lagged level of the residuals does not exceed the critical values. The distribution of this Augmented Dickey-Fuller (ADF) test statistic is not in fact given by the t-distribution, but critical values are tabulated by Engle and Yoo (1987). Table 2 gives the results for the full period, and the post-EMS period.

Table 2: *Tests of Co-integration between* $log(P_{IR})$, $log(P_{UK} \times R_{UK})$, *and* $log(P_{GR} \times R_{GR})$

	Full Period (44 obs.) ADF	*EMS Period (27 obs.) ADF*
P_{IR} and $P_{UK} \times R_{UK}$	1.82	1.29
P_{IR} and $P_{GR} \times R_{GR}$	1.47	3.16
Critical values[7]	*50 obs.*	*25 obs.*
5%	3.29	NA
10%	2.90	NA
P_{IR}, $P_{UK} \times R_{UK}$, and $P_{GR} \times R_{GR}$	3.42	3.50
Critical values	*50 obs.*	*25 obs.*
5%	3.75	NA
10%	3.36	NA

7. The critical values are taken from Engle and Yoo (1987).

The results for the full period illustrate the possibility that Irish prices may not be co-integrated with either UK prices or German prices, but could be co-integrated with a weighted average of the two: the only statistic which is significantly different from zero (thereby indicating co-integration) is that based on the hypothesis of a long-run equilibrium relationship between Irish prices and *both* UK and German prices. While critical values are not yet available for the number of observations corresponding to the post-EMS period, the results suggest a stronger role for German prices, but would still allow for a long-run link to some weighted average of UK and German prices.

Given that there is support from this co-integration analysis for the existence of a long-run equilibrium relationship between Irish prices and a weighted average of UK and German prices, we proceed to the estimation of Equation (4).[8]

4.2 *Basic Model: Dynamic Specification and Stability Tests*

Equation (4) was applied to the price of output of total manufacturing. The unit wage cost variable proved insignificant. This suggests that the long-run relationship determining Irish prices is closer to the pure SOE model, described by Equation (1), than to the mixed model underlying Equation (3).[9]

Data for prices and exchange rates of three different countries, the UK, Germany and the USA, were tried in the equation. However, in estimating Equation (4), it was apparent that neither US prices nor the dollar exchange rate played any significant role in the determination of wholesale prices at an aggregate level. This is somewhat surprising, given the fact that price-setting in dollars has become increasingly important, especially in the office machines and data processing equipment and pharmaceuticals sectors, which have grown rapidly over the period to form a significant part of manufacturing industry. However, there are inherent difficulties in measuring the prices of these sectors. In particular, the problems posed by rapid quality change in the computing industry have led to the exclusion of its prices from the wholesale price index in Ireland, so that the lack of US influence may reflect the exclusion from the index of the sector where prices are most obviously set in US dollars.

Three lags on the change in German and UK prices and exchange rates were initially included, together with an error correction term lagged four periods and seasonal dummies S_k. These initial estimates are reported as Equation (5).[10]

8. We do not perform similar co-integration analysis of the other variables in that equation (US prices, wages) since the simpler tests used in the next section indicate that they should be discarded.

9. As Baker (1985) has pointed out, the rapid growth of high technology sectors of the economy in the 1980s poses problems for the use of an aggregate wage cost variable covering all of manufacturing industry. As discussed in the next section, when examined on a disaggregated basis, wage costs proved to be a significant factor determining prices in the long run in two sectors, i.e., electrical engineering and paper and paper products.

10. In terms of Equation (4) $\lambda = \theta_{GR} + \theta_{UK}$ and $\eta_{GR} = \frac{\theta_{GR}}{\lambda}$ and $\eta_{UK} = \frac{\theta_{UK}}{\lambda}$.

$$\Delta\log(P_{IR}) = \sum_{i=0}^{3} \beta_{iGR} L^i\{\Delta\log(P_{GR})\} + \sum_{i=0}^{3} \gamma_{iGR} L^i\{\Delta\log(R_{GR})\}$$
$$+ \sum_{i=0}^{3} \beta_{iUK} L^i\{\Delta\log(P_{UK})\} + \sum_{i=0}^{3} \gamma_{iUK} L^i\{\Delta\log(R_{UK})\}$$
$$+ \theta_{GR} L^4 \log\left(\frac{P_{GR} \times R_{GR}}{P_{IR}}\right) + \theta_{UK} L^4 \log\left(\frac{P_{UK} \times R_{UK}}{P_{IR}}\right)$$
$$+ \alpha + \sum_{i=0}^{2} \zeta_i S_i \quad (5)$$

The fit of the relationship is very good, and most coefficients have significant values of the expected (i.e., positive) sign; the three coefficients with negative signs are not significantly different from zero. The long-run coefficients are particularly well-defined.

Equation 5 Results[11]

Coefficient:	i=0	i=1	i=2	i=3
α	1.379 (4.7)			
β_{iGR}	0.575 (2.1)	0.092 (0.3)	-0.282 (-0.8)	0.122 (0.4)
γ_{iGR}	0.165 (4.6)	0.171 (4.0)	0.096 (2.3)	0.084 (1.8)
β_{iUK}	0.169 (1.1)	-0.117 (-0.7)	0.193 (1.1)	-0.063 (-0.5)
γ_{iUK}	0.048 (1.2)	0.131 (3.4)	0.079 (2.1)	0.089 (2.6)
θ_{GR}	0.135 (4.3)			
θ_{UK}	0.096 (4.0)			
ζ_i	0.018 (5.3)	0.012 (2.8)	0.004 (1.1)	

$R^2 = .940$	SER = .00534	$\bar{R}^2 = .887$
Mean = .0229	S.D. = .0171	Chow F[12,35] = 1.2
AR F[4,20] = .906	DFFITS = 2.92	Het F[1,44] = .684

11. t-statistics are given in parentheses under the coefficients.

The diagnostic statistics are generally satisfactory. The F-versions of the Lagrange Multiplier tests do not indicate that the errors are autocorrelated or heteroscedastic. The relationship is relatively robust to the dropping of the last 12 observations, as indicated by the Chow test: more detailed analysis of possible structural breaks related to the EMS regime is taken up later in this section. The DFFITS statistic is above the critical level suggested by Krasker, Kuh and Welsch (1983) indicating that at least one observation is exerting significant leverage on the results. However, when estimated using bounded influence regression, the results were little changed (Krasker, Kuh and Welsch, 1983).

The strength of the influence from Germany, and the weakness of the UK influence are interesting. The coefficients on the ECM terms indicate that the long-run weight on German prices in Irish pound terms is 0.6 while that on the UK is 0.4. On the basis of trade shares one might have expected the UK influence to be greater.[12] The difference between the estimated short-run and long-run effects in this and other equations accounts for the autocorrelation found by Leddin (1988) in a specification which did not allow for these differential dynamics.

We want to investigate both the dynamic specification of this relationship, and its stability across the change of exchange rate regime. To make our results independent of the order in which we undertook these tasks, we have performed F-tests for a series of nested models. They involve simplifications of the dynamic structure and alternative changes associated with EMS entry.

In Figure 1 the dynamic structure of Equation (5), with three lags on all short-run coefficients, is labelled A; the simplification (suggested by the t-statistics) which drops the three lags on the price terms is labelled B; a corresponding simplification, dropping the lags on the exchange rate terms is labelled Z; while the final simplification, dropping the third lag on the German exchange rate, is labelled C. A specificiation which allows both long- and short-run relationships to change after EMS entry is labelled 1;[13] a specification which allows only the long-run relationship to change, as shown in Equation (6) below, is labelled 2; and the specification which imposes the same relationship over the full period (as in Equation (5)) is labelled 3. Thus we have a 4×3 matrix of possible specifications, with Equation (5) represented by A3. Figure 1 below indicates the specifications preferred by the F-test at the 5 per cent level: the calculated statistics are given in Appendix 1.

12. This issue is discussed later in Section 4.6.

13. This implies that only the post-EMS period can be estimated, since there are insufficient degrees of freedom to estimate it for the pre-EMS period.

Figure 1: *Summary of Results of Nested Tests*
(Significance level: 5%)

	(1) Short- & Long-run Relationships Changed by EMS		*(2) Long-run Relationship Changed by EMS*		*(3) Short- & Long-run Relationships Unchanged by EMS*
Z. Drop lags on R	Z1	⇒	Z2	⇐	Z3
	⇓		⇓		⇓
A. 3 lags on P and R	A1	⇒	A2	⇐	A3
	⇓		⇓		⇓
B. Drop lags on P	B1	⇒	B2	⇒	B3
C. Drop 3rd lag	⇓		⇓		⇑
on R_{GR}	C1	⇒	C2	⇐	C3

These F-tests contain several features of interest. The dropping of the lags on the price terms is never rejected; corresponding tests on dropping the lags on the exchange rate terms are always strongly rejected. This is strong evidence for an asymmetry between the effects of foreign price changes and exchange rate changes on Irish wholesale prices: the exchange rate effects are more spread out over time, while the price effects are concentrated in the current period. Further testing on this issue is carried out in Section 4.4 below.

Tests for a structural break after entry to the EMS are accepted for all but one specification at the 5 per cent level, and for all at the 10 per cent level. The hypothesis that this structural break is concentrated on the long-run coefficients cannot be rejected at the 5 per cent level for any of the specifications considered. This strongly suggests that there was a change in the determination of Irish wholesale prices subsequent to EMS entry. The fact that the short-run coefficients do not change significantly may reflect the difficulty in identifying the exact short-run relationships, whereas the longer-run relationships differ more clearly.

No matter what the starting point in the matrix of specifications of Figure 1, following the results of the nested tests leads to either specification C2 or B3. This ambiguity can be resolved by moving to a 10 per cent significance level for the tests. The only change in terms of Figure 1 is that the arrow from B2 to B3 is reversed. This means that no matter what the starting point,[14] the arrows now point eventually to specification C2, which is reported below as Equation (6).

14. For example, the most general specification A1, as suggested by the "general to simple" strategy set out by Hendry (1983).

In the preferred specification, Equation (6), the long-run relationship is allowed to change on joining the EMS. This is implemented by multiplying a dummy[15] by the error correction terms for the UK and Germany.

$$
\begin{aligned}
\Delta\log(P_{IR}) = {} & \sum_{i=0}^{3} \beta_{iGR} L^i\{\Delta\log(P_{GR})\} + \sum_{i=0}^{3} \gamma_{iGR} L^i\{\Delta\log(R_{GR})\} \\
& + \sum_{i=0}^{3} \beta_{iUK} L^i\{\Delta\log(P_{UK})\} + \sum_{i=0}^{3} \gamma_{iUK} L^i\{\Delta\log(R_{UK})\} \\
& + \theta_{GR} L^4 \log\left(\frac{P_{GR} \times R_{GR}}{P_{IR}}\right) + \theta_{UK} L^4 \log\left(\frac{P_{UK} \times R_{UK}}{P_{IR}}\right) \\
& + \delta_{GR} L^4 \log\left(\frac{P_{GR} \times R_{GR}}{P_{IR}}\right) \times D_{EMS} \\
& + \delta_{UK} L^4 \log\left(\frac{P_{UK} \times R_{UK}}{P_{IR}}\right) \times D_{EMS} + \alpha + \sum_{i=0}^{2} \zeta_i S_i \qquad (6)
\end{aligned}
$$

While the t-statistic on the coefficients of the two "structural break" variables are individually significant only at the 10 per cent level, the F-test of the joint hypothesis that both are equal to zero yields a value of 5.59, which is significant at the 1 per cent level. As shown in Table 3, these results imply that the long-run weights on German and UK prices changed on joining the EMS as the importance of UK output prices in determining Irish output prices was greatly reduced and there was a corresponding increase in the importance of German output prices.

A break in this estimated relationship, involving fixed weights on each country's price indices has two possible explanations. If the underlying world price determining Irish output prices were a trade-weighted average of prices in our main trading partners the significance of the dummies could merely reflect a change in trade weights, i.e., the higher weight on German prices in the later period could be due solely to its increasing importance in Irish trade. This explanation can, however, be discounted, because of the evidence from the trade-weighted regressions below.

The alternative explanation, supported by these results, is that there was a change in price setting behaviour after Irish entry into the EMS. This suggests that the relative importance of German prices as against UK prices in Irish firms' pricing decisions may have been altered for more institutional reasons, perhaps associated with a reversal in the relative stability of the Deutschmark and Sterling nominal exchange rates under the new regime.

15. The dummy is zero before EMS entry, and unity thereafter.

Equation 6 Results[16]

Coefficient:	i=0	i=1	i=2	i=3
α	1.050 (3.7)			
β_{iGR}	0.500 (2.8)			
γ_{iGR}	0.118 (3.6)	0.114 (3.0)	0.061 (1.8)	
β_{iUK}	0.270 (2.5)			
γ_{iUK}	0.067 (2.5)	0.135 (5.1)	0.081 (3.0)	0.081 (3.3)
θ_{GR}	0.024 (0.6)			
θ_{UK}	0.167 (2.9)			
δ_{GR}	0.061 (1.5)			
δ_{UK}	-0.07 (-1.4)			
ζ_i	0.016 (5.8)	0.011 (4.1)	0.003 (1.3)	

R^2 = .942	SER = .00490	$\bar{R}^2$ = .910
AR F[4,25] = 1.36	DFFITS = 1.99	Het F[1,44] = 1.12

Table 3: *Weights on German and UK Prices*

	Germany	*UK*
Pre-EMS	0.13	0.87
Post-EMS	0.47	0.53

16. t-statistics are given in parentheses under the coefficients.

4.3 *Tests of Long-run Homogeneity*

The tests of co-integration in Section 4.1 have indicated support for a long-run equilibrium relationship between Irish prices and a weighted average of the Irish pound prices of German and UK manufactures. This is a necessary, but insufficient, condition for long-run homogeneity to hold. The restriction of long-run homogeneity, or relative purchasing power parity, has been imposed in the models of the previous section: the long-run solution of these models involves a constant ratio between Irish prices and a (geometric) weighted average of UK and German prices in Irish pound terms. Given the nature of price indices it is not possible to draw any conclusions concerning absolute purchasing power parity from these results.

A stronger test of the homogeneity restriction can be derived by including as an additional variable the fourth lag of the level of Irish prices: this allows the long-run relationship to be a non-constant ratio. If the restriction is rejected by this test it would suggest the omission of some important explanatory variable. While the coefficient on this additional variable is significantly different from zero in the regressions covering the full period, it is not significantly different from zero in 3 of the 4 regressions for the post-EMS sub-period.

Thus, the evidence on long-run homogeneity is somewhat mixed. It should be remembered that the price series used are for total manufacturing industry. Differences in the weights attached to different components will lead to divergences from relative PPP if the price trends for these different components diverge.

4.4 *Tests of Short-term PPP*

We have already seen one set of F-tests which indicates that the response to exchange rate changes is slower than that to prices. Further tests confirm this as a strong result. Formulations which impose an identical response in each quarter by Irish prices to changes in foreign currency prices and exchange rates are rejected against those which allow differential effects. This result holds very generally across countries, sectors and sub-periods. For example, the F-test for imposing an identical response restriction on Equation (5) yields a value of 3.65 as against a 1 per cent critical value $F(8,2)$ of 3.41. These tests confirm the result of Leddin (1988) that purchasing power parity, expressed on a period by period basis, can be rejected: they also indicate, however, that Leddin's imposition of identical coefficients on the foreign price and exchange rate terms is rejected by the data.

However, if the identity of response to foreign price and exchange rate changes is assumed to apply within an annual context, by restricting the sum of the adjustment coefficients on prices and the exchange rates to be equal, the resulting restrictions are not rejected by the data. The F-test for imposing this restriction on Equation (5) yields a value of 1.28 as against a critical value $F(2,23)$ of 3.42.

The response to German price changes in DM terms tends to be larger (and better defined) than the response to exchange rate changes. In the case of the UK the opposite is the case with the exchange rate terms being generally better defined than the price terms. The short-term response to exchange rate changes tends to be somewhat slower than to foreign currency prices.

4.5 *Tests of Asymmetric Response to Changes in Foreign Prices*

Baker *et al.* (1986) suggested that domestic prices would respond faster to a rise in the price of foreign currency than to a fall. The possibility of a stronger response to exchange rate induced increases in competitors' prices (in Irish pound terms) than to decreases was tested as follows. First, a dummy variable was constructed taking the value 1 for an increase in the UK exchange rate and zero otherwise. An additional variable was formed by multiplying the short-term exchange rate changes by this dummy variable, thus giving a variable equal to the exchange rate change if this was positive, but zero otherwise.[17] A positive significant coefficient on this variable would indicate a more rapid response to positive exchange rate changes. The results show that the coefficients on the dummy variables, while all bearing the expected positive sign, implying a faster response to a depreciation of the Irish pound, are jointly not significant from zero. However, due to the small number of observations for an appreciation of the Irish pound against Sterling, this test is not conclusive.

4.6 *Non-nested Tests against Trade-weighted Foreign Price Variables*

Trade weights for the UK, Germany and the US move fairly steadily from 0.75, 0.13, 0.12, respectively, to 0.55, 0.20, 0.25 over the period 1975 to 1987. We have already seen that the implied weight on German prices in the aggregate level regressions is much greater than even this end-period weight, and the implied weight on the UK is correspondingly lower. However, the fact that the trade weights move over the period to reflect Germany's increasing trade share could give a trade-weighted approach some advantages over the fixed weights of the regression approach. Simple trade weights were used to create weighted averages of UK, German and US prices and exchange rates: the results reported below do not suggest that more sophisticated weighting systems would lead to different conclusions. Non-nested tests provide the appropriate framework for choosing between the trade-weighted and regression-weighted approaches.

17. This was the approach adopted by Goldstein (1974) in testing for an asymmetric response of wages to price increases and decreases.

The results indicate that the regression-weighted approach encompasses the results of the trade-weighted approach. The trade-weighted regressions are rejected against a joint model incorporating separate information on the UK and Germany (the test statistic of 4.72 exceeding the critical $F(18,15)$ at the 1 per cent level of 3.43), while the regression-weighted approach is not rejected against a joint model including trade-weighted information (the test statistic of 1.31 being below the 5 per cent critical $F(9,15)$ of 2.59). The reported statistics are based on the inclusion of three lags of each price and exchange rate variable, as in Equation (5), but similar results were obtained with more restricted specifications.

Specifications which use UK data alone to proxy foreign prices are also strongly rejected against specifications including German price data. Furthermore, the specifications involving the UK alone yield unstable and counterintuitive results, including rejection of long-run homogeneity, and a sign on the error correction term which would lead to Irish prices diverging from those in the UK.

There would seem to be two main reasons why the influence of German prices exceeds their trade weight. The first is that German prices may to some extent proxy prices for France, the Netherlands and other European countries not included in the analysis. It is doubtful, however, if this is the whole story. An alternative, mentioned earlier, is that German prices exert an influence out of proportion to their trade weight, possibly because German firms are regarded as price leaders; further investigation would be needed to establish whether or not this is the case.

V REVIEW OF DISAGGREGATED RESULTS

We also applied the model described above in Equation (4) to the price data for the output of a range of non-food sectors of manufacturing industry.[18] In each case data for prices and exchange rates in the UK, Germany and the US, were tried together with wage rates. In the case of certain sectors price and exchange rate data for some other possibly relevant countries were tried. The results of these regressions are summarised in Table 4.[19] The same dynamic specification was used in each case with no lags on the price change variables and 3 lags on the exchange rate changes. In all but one case a specification including 3 lags on the change in prices was rejected at the 5 per cent level in favour of this more restricted specification used here.

18. See Callan and FitzGerald (1988) for full details of these results.

19. While seasonal dummies were included in each regression, they are omitted from the table to save space.

While the $\bar{R}^2$ statistics reported for these disaggregated equations are typically lower than that for the aggregate equation, they are all highly significant, and compare favourably with other time series analyses of dependent variables differenced in logarithms. At such a disaggregated level the increasing importance of special factors, not captured by the regressors, may account for the somewhat lower $\bar{R}^2$ statistics reported.

The 11 equations for the different sectors of manufacturing cover the bulk of the non-food output of that sector (over 60 per cent of net output). One major omission is the electronics sector which is also omitted from the index covering total manufacturing. As can be seen from the table the results are reasonably consistent with the results obtained from the aggregate equation.

In eight sectors at least one of the θ coefficients are significant indicating a significant long-term relationship between Irish prices and the explanatory variables. Domestic costs, proxied by wage costs, are of only limited importance in determining prices; they are only significant in two sectors in the long run.[20] Thus, for most sectors the SOE model is appropriate. As with the results presented in Section 4, German prices generally have a higher weight in determining Irish prices than do UK prices. The results confirm the slower speed of adjustment of Irish prices to changes in exchange rates than to changes in foreign currency prices. Finally, these results conformed to the results in Section 4 in ascribing little or no weight to other countries' prices[21] in explaining Irish prices.

20. The domestic wage cost data used were those appropriate to the relevant sector.

21. For example, US prices.

Table 4: *Results for Individual Industrial Sectors*

Sector	*1*	*2*	*3*	*4*	*5*	*6*	*7*[22]	*8*	*9*	*10*	*11*
α	0.788	1.510	1.139	1.101	1.171	2.003	0.567	0.404	−0.056	0.665	0.493
	(3.6)	(3.6)	(1.4)	(4.8)	(4.0)	(3.4)	(1.9)	(2.3)	(0.3)	(3.4)	(1.4)
β_{0GR}		0.261	0.117	0.014	1.005	−0.007	0.176			0.270	0.744
		(0.7)	(0.1)	(0.0)	(3.9)	(0.0)	(0.9)			(1.8)	(5.4)
γ_{0GR}		0.223	−0.010	0.074	0.107	0.371	−0.004			−0.038	0.031
		(3.4)	(0.1)	(1.5)	(2.1)	(2.2)	(0.2)			(0.3)	(0.3)
γ_{1GR}		0.069	0.140	0.171	0.233	0.511	−0.020			0.281	0.415
		(1.0)	(1.9)	(3.2)	(4.1)	(2.9)	(0.0)			(2.1)	(3.9)
γ_{2GR}		0.002	0.037	0.160	0.081	0.173	0.077			−0.308	0.216
		(0.0)	(0.4)	(2.9)	(1.5)	(1.0)	(2.1)			(2.4)	(2.0)
γ_{3GR}		0.134	−0.107	0.015	0.162	0.200	0.032			0.188	0.081
		(2.3)	(1.3)	(0.3)	(3.1)	(1.2)	(1.0)			(1.4)	(0.8)
β_{0UK}	0.436	0.616	0.339	0.293	−0.088	0.183	−0.010	0.853	0.463		
	(2.8)	(2.7)	(2.1)	(1.4)	(0.4)	(0.3)	(0.0)	(3.9)	(3.8)		
γ_{0UK}	0.018	0.081	0.076	−0.063	0.109	0.106	0.090	0.155	0.019		
	(0.2)	(1.7)	(1.6)	(1.6)	(2.8)	(0.8)	(2.4)	(3.1)	(0.4)		
γ_{1UK}	0.232	0.146	0.099	0.054	−0.043	0.026	0.026	0.190	0.105		
	(3.1)	(3.2)	(2.5)	(1.3)	(1.0)	(0.2)	(0.7)	(3.8)	(2.1)		
γ_{2UK}	0.086	0.079	0.012	0.023	0.033	0.124	0.075	−0.020	0.096		
	(1.0)	(1.7)	(0.2)	(0.5)	(0.8)	(0.9)	(2.0)	(0.3)	(1.8)		
γ_{3UK}	0.072	0.111	−0.022	0.168	0.040	0.026	0.062	0.026	0.091		
	(0.9)	(2.4)	(0.5)	(4.3)	(1.0)	(0.2)	(1.6)	(0.5)	(1.9)		
ϵ_0	0.037		0.015					−0.086	0.072		
	(1.3)		(0.7)					(2.1)	(2.7)		
ϵ_1	0.012		−0.002					−0.068	0.112		
	(0.4)		(0.1)					(1.3)	(3.6)		
ϵ_2	0.024		0.041					0.124	0.087		
	(0.7)		(1.5)					(2.0)	(2.4)		
ϵ_3	0.076		0.040					0.021	−0.067		
	(2.2)		(1.5)					(0.4)	(1.4)		
θ_{GR}	0.131	0.126	0.114	0.099	0.126	0.346	0.055			0.097	0.076
	(2.6)	(3.1)	(1.5)	(3.7)	(4.2)	(3.1)	(1.8)			(3.3)	(1.4)
θ_{UK}		0.133	0.050	0.084	0.068	−0.049	0.000	0.086	−0.067		
		(3.3)	(0.7)	(3.8)	(2.4)	(1.0)	(0.0)	(1.6)	(1.4)		
θ_W	0.042		0.051					−0.014	0.079		
	(1.4)		(2.4)					(0.3)	(2.5)		
$\bar{R}^2$	0.647	0.664	0.706	0.685	0.756	0.393	0.453	0.694	0.785	0.292	0.472

Sectors:
1. Chemicals
2. Metal Products
3. Electrical Engineering
4. Motor Vehicles and Parts
5. Textiles
6. Leather and Footwear
7. Clothing
8. Timber and Furniture
9. Paper and Paper Products
10. Rubber and Rubber Products
11. Processing of Plastics

22. For this sector (clothing) Spanish prices replace German prices.

VI CONCLUSIONS

The results presented in this paper show that the SOE model of price determination is appropriate to Irish manufacturing output. Domestic wage costs play only a minor role in price determination, while foreign price influences are dominant. In this respect, the results confirm the consensus of earlier research on this topic; but the results on the nature and timing of the foreign influence are a good deal more novel.

Short-run homogeneity is rejected by the data. While long-run homogeneity is generally rejected for the full data sample, it is accepted for the post-EMS period.

Generally the effects of changes in exchange rates on prices are much slower to materialise than are the effects of changes in foreign currency prices. This difference in behaviour can be explained by the fact that firms are reluctant to change their output prices very frequently. As a result, they have to set future prices based on their expectations of the Irish pound price of foreign competitors. When foreign competitors raise their prices in foreign currency terms there is little likelihood that they will subsequently reduce them, so domestic producers will be quick to follow suit. However, if foreign prices in Irish pound terms change because of exchange rate changes, past experience suggests, especially in the case of the Sterling-Irish pound exchange rate, that there is quite a possibility that the change may be reversed in the immediate future. As a result, firms are likely to be slower to change their expectations about future values of the exchange rate as a result of a single quarter's figures.[23]

There is no evidence of an asymmetric response to appreciation and depreciation *vis-à-vis* Sterling. However, the number of observations available for the post-EMS period are still too few to properly test this hypothesis. In any event, this hypothesis may be more appropriately applied to a model of the formation of domestic consumer prices, as suggested by Baker *et al.*, 1986.

Germany has a stronger influence in determining Irish prices than the trade weights suggest. In the post-EMS period its role in output price determination is roughly equal to that of the UK. The pattern of response of Irish prices to foreign prices was significantly different from that suggested by an approach based on trade weights.

Because of a limited number of consistent observations for the pre-EMS period it is difficult to test for a change in pattern of behaviour following the break in the link with sterling consequent on EMS membership. However, the evidence from the aggregate price series does indicate a change in pricing

23. See the case of Irish Distillers in the Report of the Restrictive Practices Commission, 1988.

behaviour in 1979 with an increase in the importance of German prices and a fall in the importance of UK prices.[24] It is difficult to identify a change in short-run pricing behaviour. The differential response to exchange rate changes and foreign currency price changes has also taken on added importance now that the Sterling-Irish pound rate is flexible.

These results have a number of implications for public policy. In so far as the pricing behaviour of Irish firms reflects their views as to who their competitors are, these results suggest a change in the traditional trade-weighted approach to measuring competitiveness. On the basis of these results, EMS countries, proxied by Germany, should have a much bigger weight in competitiveness measures than would be suggested by traditional trade weights.

The slower response of domestic prices to changes in the exchange rate than to changes in foreign currency prices has implications for the Irish economy on the completion of the EC market in 1992. If the UK remains out of the EMS uncertainty about the bilateral Sterling-Irish pound exchange rate will pose problems in price determination. Without any customs or fiscal barriers the slow adjustment of prices to exchange rate changes could give rise to large temporary trade flows. The alternative is that either prices in Ireland change more rapidly in the face of changes in the Sterling-Irish pound exchange rate or else the variability of UK prices in Sterling terms increases.

24. This is confirmed for certain of the disaggregated series examined in Callan and FitzGerald, 1988.

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Appendix 1: *Calculated Values of the Test Statistics Underlying Figure 1.*

Specifications 1, 2 and 3 and dynamics A, B, C and Z are as described in Figure 1 and the related text.

Z1	F(11,17)=1.05	Z2	F(2,28)=8.04**	Z3
F(6,11)=7.29**		F(6,22)=8.05**		F(6,24)=8.12**
A1	F(11,11)=1.61	A2	F(2,22)=7.27**	A3
F(6,11)=2.26		F(6,22)=1.87		F(6.24)=0.50
B1	F(11,17)=1.18	B2	F(2,28)=3.31+	B3
F(1,17)=0.50		F(1,28)=1.65		F(1,30)=5.60*
C1	F(11,18)=1.33	C2	F(2,29)=5.59**	C3

A test statistic which is significant at 1 per cent(**), 5 per cent(*) and 10 per cent(+) indicates that the less general specification (i.e., the one with more restrictions and, hence, fewer estimated parameters) is rejected against the more general specification. For example, since Z is less general than A, the test statistic F(6,11)=7.29 indicates that Z1 is rejected against A1. Specification A is more general than B which is more general than C while dynamics 1 are more general than 2 which are more general than 3.

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