Manufacturing price determination in OECD countries; markups, demand and uncertainty in a dynamic heterogeneous panel.

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Abstract

Manufacturing price markup equations are estimated for 15 OECD countries using annual data. The specification is based on a CES production function where the markup depends on demand, competitors' prices and uncertainty. We test for cointegration using the Pedroni panel tests and a panel version of the Johansen test, and find evidence for unique cointegrating vectors. Estimation of the long-run parameters is performed with a pooled mean group method, leaving the short run heterogeneous dynamics unconstrained. Tests for homogeneity of the long-run parameters do not reject the hypothesis. Markups are pro-cyclical and rise with both competitors' prices and uncertainty.

Keywords: Pricing behaviour, markups, panel tests for order of integration, panel cointegration, dynamic heterogeneous panels, pooled mean group estimation.

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2.0

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1 Introduction

The world economy has returned to a low inflation regime. Price inflation in Europe and much of the rest of the developed world is now close to zero. The demand management controversies of the past are now largely consigned to history, and the attention of policymakers is now almost exclusively fixed on inflation and price determination. At the same time, macroeconomists are increasingly focusing on the price determination process.[1] In Europe there is now a common monetary policy among the members of EMU.The gain from this institutional change is that inflation credibility is enhanced. However, the size of the costs following inflationary shocks will depend on the degree of heterogeneity in the real side of the member economies. Differences in real supply side behaviour will surface in both wage and price setting. Similarly, if the world is to move back to widespread fixed exchange rates, it is vital to know whether different economies have similar price responses to shocks. This all makes it worthwhile, and arguably vital, to establish what international regularities in price setting behaviour exist. Yet, surprisingly, there is very little single country analysis of the price markup equation, and even less cross-country work.[2] This paper attempts to address this imbalance by examining the markup over marginal costs in a small panel of OECD countries. Moreover, in a recent paper in this Journal, Neiss (2001) uses unit labour costs to proxy international variation in the markup, and equivalently competitiveness, in a model of the inflation bias. Other authors (eg, Gali et al (2001)) have recently reinterpreted the Phillips curve as a dynamic pricing equation, where marginal costs are proxied by unit labour costs. Yet this relies upon the maintained hypothesis that production technology is Cobb Douglas. Relaxing this assumption allows a more general specification of the marginal cost. In this paper, we are able to reject the Cobb Douglas assumption, and this has implications for the work referred to above.

Use of a panel of this type raises some important methodological issues.[3] There has been an increasing interest in the use of panel techniques in macroeconomics, rather than the micro and labour based data sets in which panel methods have traditionally been used. Country panels tend to have dimensions in T and N of roughly equal orders.[4] As static models are rarely adequate for typical time series, dynamic models are usually appropriate. The small T problems with dynamic panels[5] are not relevant here as the fixed-effects problem from the initial conditions declines rapidly as T rises. But instead, there are profound problems that result from heterogeneity in the model parameters that emerge as soon as a lagged dependent variable is introduced.

[1]See the references cited in Martin (1997).

[2]This disregards the large literature on purchasing power parity, which does not attempt, except indirectly, to estimate the markup over costs. For the UK, the main papers are Martin (1997), Smith (2000) and Price (1991) and (1992). Alogoskoufis et al (1990) uses a panel of OECD countries, focussing on the role of competitors' prices.

[3]See Hall and Urga (1998) for a recent survey and analysis of some panel estimation issues that are relevant to the case we consider here, and the November 1999 Special Issue of the **Oxford Bulletin of Economics and Statistics**.

[4]In the current paper which uses annual data, T and N are 27 and 15 respectively.

[5]Arellano and Bond (1991).

This problem was forcefully addressed by Pesaran and Smith (1995). Unlike in static models, estimates are inconsistent even in large samples. Happily, in typical data sets T is sufficiently large to allow individual country estimation. Pesaran and Smith observe that while it is implausible that the dynamic specification is common to all countries, it is at least conceivable that the long-run parameters of the model may be common. We can then exploit the cross-sectional dimension to gain more precise estimates of these average long-run parameters. They then propose estimation by either averaging the individual country estimates, or (Pesaran, Shin and Smith (1999)) by pooling the long-run parameters, if the data allows.

With non-stationary data, as in our case, all this is conditional on the existence of a unique long-run cointegrating relationship. Panel tests of this hypothesis were not available in 1995, but since then a range of residual based tests have emerged.[6] Residual based tests in single countries are well known to have non-standard distributions and low power. The advantage of panel tests is that the distributions tend to asymptotic normal as the cross-section dimension rises, and power tends to increase. But there is a further, often neglected, issue to consider. An aspect of the analysis of non-stationary data is to establish the number of cointegrating relationships, r. If this exceeds one, it becomes essential to tackle this in estimation. Larsson, Lyhagen and Löthgren (2001) have developed a simple panel test for the number of cointegrating vectors. We apply this test in the current paper, as well as single equation residual based tests.[7]

We estimate a model of pricing behaviour for the manufacturing sector in 15 OECD countries.[8] After conducting single country and panel tests for order of integration (Im, Pesaran and Shin (1997)) we then test for the existence and number of cointegrating vectors using the panel approaches of Pedroni (2000) and the panel version of the Johansen test. We then estimate the long-run parameters using a pooled mean group (PMG) estimator. We test for and find equality of the long-run coefficients using a Hausman test. We reject both PPP and Cobb Douglas technology, and find the markup is positively related to demand, import prices and uncertainty.

In the next section we set out the simple theory underlying the estimates. Section 3 spells out the econometric methodology. After summarising the data in Section 4, we test for cointegration before moving on to the dynamic panel estimation in Section 6. Section 7 concludes.

[6]Very little of this rapidly expanding literature has yet been published; Pedroni (1999) is an exception, and other papers in the Oxford Bulletin special edition. McCoskey and Kao (2001) provide some Monte Carlo comparisions of the different tests available.

[7]However, see Banerjee, Marcellino and Osbat (2000) for a cautionary view of these and other similar tests, all of which assume there are no cross country relations between the variables and that the cointegrating rank r is common. If these conditions are true, they show the panel Johansen test we apply has good properties. However, if they are false, results can be misleading. But the examples they give are ones where cross unit relations are extremely likely, focussing on interest rates. In our case there may be less of a problem.

[8]Countries and variables are listed and described in the Appendix.

2 Theory

In our work we look at the manufactured sector, so we can ignore the non-traded sector.[9] It is standard[10] to assume firms are monopolistically competitive. They face a demand curve for output of the following form:

$$Q^d = D(Z, P/P^*) \tag{1}$$

where $P/P*$ is the price relative to competitors' and Z is an index of demand. Production is determined by a CRS production function

$$Q^s = F(NA, K) \tag{2}$$

where N is employment, A labour augmenting technical progress and K is capital.[11] The first order condition can be written as

$$P = \frac{W/AF_1}{1 - 1/\epsilon(Z, P/P^*)} \tag{3}$$

where W/AF_1 is marginal cost and $1 - 1/\epsilon(Z, P/P^*)$ determines the markup.[12]

To make (3) an estimable function we need to specify a functional form. Part of this will follow from a particular production function, and part from more informal considerations. It is commonly assumed the elasticity depends negatively on capacity utilisation as an index of demand. This is not totally uncontroversial. There are arguments suggesting the markup is countercyclical. Bils (1989) or more recently Ireland (1998) argue that firms use booms to attract new customers; Rotemberg and Saloner (1986) and (eg) Rotemberg and Woodford (1995) argue that collusive behaviour is less likely in booms, although their argument is restricted to exceptional price wars. Chevalier and Scharfstein (1996) have a model in which capital market imperfections lead to countercyclical pricing. Rotemberg and Woodford (1991) provide evidence of countercyclical markups for the US. Note that a positive coefficient on capacity utilisation may also indicate rising marginal cost.[13] The markup may also be positively related to competitors' prices. This follows if demand becomes more inelastic as competitors prices rise: the theoretical rationale for this is discussed in Bulow *et al* (1985). Under perfect competition PPP holds and the markup equation collapses either to $p = MC$ or $p = p^*$. Finally, the markup may be related to the level of demand uncertainty as in Price (1991). Recall that the markup is interpretable as a supply equation. In general, factor demand and supply will depend on uncertainty.

[9]See Martin (1997) for a comprehensive discussion of price determination in a small open economy, the UK.

[10]For example, Smith (2000).

[11]We can easily extend this to other inputs.

[12]There are ways of introducing other factor prices into the markup equation, essentially by solving for the cost function. Then we need to allow for time varying shares of all factors.

[13]Bils' (1987) paper relies largely on estimated countercyclicality in marginal costs, following inflexible employment levels, to provide evidence for countercyclical markups. The overall evidence, discussed in Layard, Nickell and Jackman (pp 339-340, 1991), is mixed.

There has long been an argument that uncertainty *increases* factor demands, thus reducing the markup: see Hartman (1976) for one of the earliest papers. This follows from the possibility that the expected marginal return on factors may be increasing in uncertainty for many production structures. More recently, the theory of investment has tended to suggest the opposite may be true, and uncertainty will reduce investment. Uncertainty increases the 'returns to waiting' before doing an investment.[14] But even there, ambiguity remains. However, the evidence, which is mainly related to investment, suggests more uncertainty will reduce factor demand.[15] Thus we expect uncertainty to reduce supply, implying higher profit margins. We measure uncertainty by the conditional variance of output estimated from GARCH processes.

Some form of constant elasticity of substitution (CES) production technology is often assumed. In that case, using lower case letters to indicate logs, one version of (3) is

$$p = \alpha_0 + w - (1/\sigma)(q-n) - (\sigma-1)/\sigma\, a_t + \alpha_1(p-p^*) + \alpha_2\rho + \alpha_3 var \qquad (4)$$

where σ is the elasticity of substitution, P^* is the competitors' price, ρ is an index of demand and var is the level of uncertainty. Notice that the coefficient on 'technical progress' a_t may be positive or negative, depending on whether σ is less than or greater than one. Note also that we can test the CES specification as there are within-equation restrictions. This raises the question of how wedded we are to the structure. We take the view that statistical rejection of the CES structure is interesting information, but should not invalidate the results. However, it does make interpretation of some parameters less straightforward. In particular, under CES the coefficients on wages and productivity can be interpreted purely in terms of σ. The corollary is that if we reject CES we cannot make statements about the elasticity of substitution. But this does not mean that any results are uninformative. The unrestricted version of this equation is

$$p = \beta_0 + \beta_1 w + \beta_2(q-n) + \beta_3 t + \beta_4(p-p^*) + \beta_5\rho + \beta_6 var \qquad (5)$$

where a_t is modelled in practice as a time trend, so that $\beta_3 = -(1-\beta_2)\theta$ where θ is a positive unknown and unidentified scale factor. Thus the CES specification is testable by the hypothesis that $\beta_3 > (<) 0$ if $\beta_2 < (>) -1$. Given the level of aggregation, P^* is the price of imports. We will measure ρ by an index of capacity utilisation; specifically, deviations from a Hodrick Prescott filtered trend of manufacturing output. A test of short-run PPP - which amounts to assuming perfect competition - is $\beta_4 = \beta_5 = 0$. However, it should be clear that PPP is generally considered to be a long-run phenomenon. Moreover, as labour costs and productivity are endogenous the system within which (4) is embedded may exhibit PPP even if the individual equation does not appear to exhibit it. Static homogeneity, which follows from the first order condition, is an essential property of the price relationship.[16] It is sometimes argued that dynamic homogeneity is important too, so that the price level is unaffected by the rate of

[14] This is known as the 'real-options' approach to irreversible investment under uncertainty. See Dixit and Pindyck (1994) for an exemplary treatment.

[15] Driver and Moreton (1992) and Price (1996); see also Price (1994) for evidence on employment.

[16] With commonly assumed Cobb-Douglas technology σ is unity so $w-(1/\sigma)(q-n)$ may be replaced by unit labour costs. In that case $p = \beta_0 + \beta_1(w-q+n) + \beta_4(p-p^*) + \beta_5\rho + \beta_6 var$ where static

inflation. But this is not obvious. While static homogeneity follows from fundamental properties of cost and demand functions, this is not the case for dynamic homogeneity, which amounts to assuming super-neutrality in a monetary model. Indeed, there are models (and some evidence) that suggest inflation does affect the markup: see Banerjee, Cockerell and Russell (2001). In any case, our main focus in this paper is on the pooled long-run estimates. Given these arguments together with the small sample used to estimate the dynamics, we did not attempt to impose or test the implied dynamic restrictions.

We interpret these as long-run relationships. In estimation, we use an ECM specification to capture short-run behaviour consistent with this long-run. Thus, as is usually the case in related empirical work, we do not formally derive the dynamics. However, it is possible to do so; for example, in Price (1992) firms set prices according to expected future costs and demand. A particular intertemporal quadratic loss function then implies restrictions on the coefficients on future forcing variables.

3 Econometric methodology

3.1 Dynamic heterogeneous panels

The data set we are examining covers 15 countries ($N = 15$) over 27 years.[17] The data also has complex dynamics and is characterised by strong trends and non-stationarity. Such data sets[18] raise special problems in estimation. Pesaran and Smith (1995) show that, unlike in static models, pooled dynamic heterogeneous models generate estimates that are inconsistent even in large samples. In the type of data set we are considering T is sufficiently large to allow individual country estimation. Nevertheless, we may still be able to exploit the cross-section dimension of the data to some extent. Pesaran, Shin and Smith (1999) (PSS) observe that while it is implausible that the dynamic specification is the same in all countries, it is at least conceivable that the long-run parameters of the model may be common. They propose estimation by either averaging the individual country estimates, or by pooling the long-run parameters, if the data allows, and estimating the model as a system.[19] They refer to this as the

homogeneity implies $\beta_1 = 1$. The Cobb Douglas case is strongly rejected in the results below. This is not surprising, given the wealth of macroeconometric evidence rejecting unit elasticity of substitution. What is more surprising, is that most empirical price markup equations use unit labour costs as the cost variable.

[17] Although we have more data for manufacturing output, needed to calculate a filtered trend.

[18] Termed 'data fields' by Danny Quah, eg, Quah (1993).

[19] Baltagi and Griffin (1997) argue that the efficiency gains of pooling the data outweigh the losses from the bias induced by heterogeneity. They support this argument in two ways. Firstly, they informally assess the plausibility of the estimates they obtain for a model of gasoline demand using different methods. This is hard to evaluate as it relies upon a judgement about what is 'plausible'. Monte Carlo simulations would make the comparision clearer. Secondly, they compare forecast performance. However, this is a weak test to apply to the averaging technique, which is designed only to estimate long-run parameters and not the short-run dynamics. Baltagi and Griffin do not consider the next method to be discussed, the PMG.

pooled mean group estimator, or PMG. It combines the efficiency of pooled estimation while avoiding the inconsistency problem flowing from pooling heterogeneous dynamic relationships. It is this latter method we apply.

The unrestricted specification for the system of ARDL equations for $t = 1, 2, \ldots, T$ and $i = 1, 2, \ldots, N$ is

$$y_{it} = \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + \sum_{j=1}^{q} \delta'_{ij} x_{i,t-j} + \mu_i + \varepsilon_{it} \tag{6}$$

where $x_{i,t-j}$ is the $(k \times 1)$ vector of explanatory variables for group i and μ_i are the fixed effects. In principle the panel can be unbalanced and p and q may vary across countries. (6) can be reparametrised as a VECM system.

$$\Delta y_{it} = \theta_i (y_{i,t-1} - \beta'_i x_{i,t-1}) + \sum_{j=1}^{p-1} \gamma_{ij} \Delta y_{i,t-j} + \sum_{j=1}^{q-1} \gamma'_{ij} \Delta x_{i,t-j} + \mu_i + \varepsilon_{it} \tag{7}$$

where the β_i are the long-run parameters and θ_i are the equilibrium (or error) correction parameters. The pooled mean group restriction is that the elements of β are common across countries.

$$\Delta y_{it} = \theta_i (y_{i,t-1} - \beta' x_{i,t-1}) + \sum_{j=1}^{p-1} \gamma_{ij} \Delta y_{i,t-j} + \sum_{j=1}^{q-1} \gamma'_{ij} \Delta x_{i,t-j} + \mu_i + \varepsilon_{it} \tag{8}$$

Estimation could proceed by iterated least squares, imposing and testing the cross-country restrictions on β. However, this will be inefficient as it ignores the contemporaneous residual covariances. A natural estimator is Zellner's SUR method,[20] which is a form of feasible GLS. SUR estimation is only possible if N is smaller than T. Thus PSS suggest a maximum likelihood estimator. In our case SUR is feasible, but we use the efficient PSS method.[21] They argue that in panels omitted group specific factors or measurement errors are likely to severely bias the country estimates. It is a commonplace in empirical panel studies to report a failure of the 'poolability' tests based on the group parameter restrictions.[22] So PSS propose a Hausman test. This is based on the result that an estimate of the mean long-run parameters in the model can be derived from the average (mean group) of the country regressions. This is consistent even under heterogeneity. However, if the parameters are in fact homogeneous, the mean and the individual parameters coincide and the PMG estimates are more efficient. Thus we can form the test statistic

$$H = \hat{q}'[var(\hat{q})]^{-1}\hat{q} \sim \chi^2_k$$

where $\hat{q}$ is a $(k \times 1)$ vector of the difference between the mean group and PMG estimates and $var(\hat{q})$ is the corresponding covariance matrix. Under the null that

[20]Zellner (1962).

[21]Implemented in a GAUSS program made available at Hashem Pesaran's website: we are grateful to PSS for making this available.

[22]For example, Baltagi and Griffin (1997, p 308) state that although the poolability test is massively failed (F(102,396) = 10.99; critical value about 1.3), 'like most researchers we proceed to estimate pooled models.'

the two estimators are consistent but one is efficient, $var(\hat{q})$ is easily calculated as the difference between the covariance matrices for the two underlying parameter vectors. If the poolability assumption is invalid then the PMG estimates are no longer consistent and the test rejects.

3.2 Panel cointegration

The other issue here is the existence of a long-run equilibrium. With non-stationary data, as in our case, estimation is conditional on the existence of a unique long-run cointegrating relationship. In Pesaran, Shin and Smith this is assumed, rather than tested.

3.2.1 Residual based tests

There is now a range of panel residual based tests available.[23] Residual based tests in single countries are well known to have non-standard distributions and low power. In panels the distributions tend to asymptotic normal as the cross-section dimension rises, and power usually increases.[24] The Pedroni tests which we use[25] allow for heterogeneity among the panel members. All are based on the residuals from the (most general) regressions

$$y_{it} = \alpha_i + \delta_i t + \beta_i x_{it} + e_{it}. \tag{9}$$

Pedroni constructs seven tests, four of which are based on pooling along the 'within-dimension' and three the 'between-dimension'. The former effectively pool the autoregressive coefficient in the residual based test and the latter take the average, allowing more heterogeneity. Pedroni refers to the within statistics as panel cointegration statistics, and to the between statistics as group mean panel cointegration statistics, a natural terminology given our discussion above. The panel tests are the panel v-statistic (a variance bounds test), the panel ρ-statistic (analagous to the Phillips Perron ρ test), and nonparametric and parametric panel t-statistics (or more accurately, ADF statistics). The group tests are the group ρ-statistic and the two group t-statistics.

3.2.2 Multiple cointegrating vectors

Another important issue not addressed by the residual based tests, however, is the number of cointegrating relationships, r. If this exceeds one, it becomes essential

[23] See for example Pedroni (1999), McCoskey and Kao (2001).

[24] Nevertheless, the tests are powerful against an arguably uninteresting alternative. They examine the null that there are unit roots in all of the N groups. Rejection does not, therefore, imply that all the series are I(0), simply that they are not all I(1). We are grateful to Ron Smith for making this point clear to us. Karlsson and Löthgren (2000) make the same point in their paper which examines the power of various panel unit root tests against various alternatives.

[25] We are grateful to Peter Pedroni for making his RATS procedures available to us.

to tackle this in estimation. In single time series, this is normally tested with the Johansen procedure. It is only recently that a panel test for the existence of multiple cointegrating vectors has been developed (Larsson, Lyhagen and Löthgren (2001)). Larrsson *et al* examine the mean of the standard Johansen LR (trace) tests from a heterogeneous panel.[26] They show analytically that the first two moments of the statistic exist and perform Monte Carlo experiments to derive small sample properties. They define their test statistic as

$$\Upsilon_{\overline{LR}}(H(r)|H(p)) = \frac{\sqrt{N}(\overline{LR}_{NT}(H(r)|H(p)) - E(Z_k))}{\sqrt{var(Z_k)}} \tag{10}$$

where $E(Z_k)$ and $var(Z_k)$ are the mean and variance of the asymptotic trace statistic, Z_k, and where

$$\overline{LR}_{NT}(H(r)|H(p)) = \frac{1}{N}\sum_{1=1}^{N} LR_{iT}(H(r)|H(p)) \tag{11}$$

is the simple average of the standard Johansen trace statistic for the hypothesis of reduced rank from (full rank) p to r.[27] The hypothesis tested is

$$H_0 : rank(\Pi_i) = r_i \leq p \qquad \text{for all } i = 1, \ldots, N$$

against the alternative

$$H_1 : rank(\Pi_i) = p \qquad \text{for all } i = 1, \ldots, N.$$

This pair of hypotheses is the panel analogue of the single time series trace statistic. $E(Z_k)$ and $var(Z_k)$ may be calculated by simulation (Johansen (1995)). The properties of the statistic are asymptotic. In our sample, we have a relatively short T. Reimers (1992) has suggested a small sample adjustment to the test statistics (namely, $(T - pk)/T$) where p is the number of variables in the model and k is the lag length used). There is no theoretical justification for the adjustment, but as Johansen (1995) observes, the approximation to the limit distribution appears to be better with the correction. Thus we use this adjustment in the empirical work.

4 Data

A data set of this size is hard to summarise. The data in the specification of the long-run relationship are the logs of manufacturing producer prices ($p = ln(P)$), labour

[26] The model is restrictive because it assumes there are no covariances between equations errors, and that there are no influences across countries. In particular, the α matric of loadings is assumed to be diagonal. These assumptions are relaxed in Larsson and Lyhagen (1999). However, in our case the dimensions of the problem make estimation of either the non-diagonal α or covariance matrices impossible.

[27] Results are derived for the case of an unrestricted constant and no deterministic trends in or out of the cointegrating space. Gerdtham and Löthgren (2000) conjecture the results also hold for the case where there is a linear trend in the cointegrating space, the one applicable to our case. The moments are calculated by simulation, and we are grateful to Mickael Löthgren for making his GAUSS procedure available to us.

costs (w), productivity ($q - n$ where $q = ln(Q)$, $n = ln(N)$, Q is output and N employment) and the ratio of domestic to foreign prices ($p - p^*$ where $p^* = lnP^*$ and P^* is manufacturing import prices). All variables are described in the Appendix. Inspection of the data suggest that the series are non-stationary. Individual ADF tests reported in Table 1 generally suggest the series are I(1) for prices and wages. However, the evidence for productivity and relative prices is less clear cut. As with tests for cointegration, use of a panel increases the power of tests for cointegration and we therefore apply the Im, Pesaran and Shin (1997) test for cointegration. These results are also given in Table 2, and provide strong evidence that the series are I(1).

There are also two constructed series which enter as I(0) variables. Firstly, we used deviations from a Hodrick-Prescott trend[28] as our demand indicator. The Hodrick-Prescott filter is well known to suffer from statistical deficiencies, of which possibly the worst is the 'end sample' problem, whereby the filtered trend tends to fix on the actual values. Thus we used a longer data series to generate the HP trend than in estimation of the model. Secondly, we model the conditional variance of output growth with individual GARCH(1,1) processes estimated for each country where the mean is estimated as a random walk.[29]

5 Tests for cointegration

Given the results reported above, we maintain the hypothesis that all the explanatory series are I(1) with the exception of demand and uncertainty.

5.1 Pedroni tests

We estimated (5), repeated here for the reader's convenience:

$$p = \beta_0 + \beta_1 w + \beta_2 (q - n) + \beta_3 t + \beta_4 (p - p^*) + \beta_5 \rho + \beta_6 var.$$

As ρ and var are I(0) by construction they are not included in the tests. The results are given in Table 2.

The (one sided) test statistics are distributed asymptotic standard normal. The critical value for the panel v-statistic is positive, while the others are negative. The properties of the tests have not been investigated in detail for the multivariate case, but we have a prior belief that heterogeneity in the autoregressive process is likely, so we prefer to be guided by the group statistics.[30] We report results for the unrestricted case and also when we impose homogeneity; we concentrate on the latter results. In the panel

[28]Using $\lambda = 7$ as we have annual data. This value is the annual equivalent to the 1600 recommended by Hodrick and Prescott for quarterly data. The rationale is that it allows cycles with longer than business cycle frequencies to pass through.

[29]An AR(1) process produces similar results.

[30]In the next section we report that the estimated ecm coefficients differ widely, so this was an appropriate assumption.

Table 1
Unit root tests on Δy

	Price p	Labour costs w	Productivity $q-n$	Relative price $p-p^*$
Individual ADF tests (lag=1)				
Austria	-3.047	-3.134	-2.351	-2.888
Belgium	-2.33	-3.85	-2.631	-2.053
Canada	-3.182	-3.196	-2.745	-2.552
Denmark	-2.757	-2.749	-2.001	-0.187
France	-2.777	-2.802	-2.19	-2.303
Germany	-2.582	-3.119	-1.709	-1.29
Ireland	-4.024	-3.541	-2.856	-2.285
Italy	-3.301	-3.323	-1.299	-2.735
Japan	-3.611	-3.689	-2.151	-1.374
Netherlands	-2.771	-2.905	-1.27	-1.731
Sweden	-2.819	-1.936	-2.057	-2.078
Spain	-2.803	-3.832	-2.174	-3.327
Switzerland	-2.614	-1.097	-0.249	-0.547
UK	-2.77	-3.737	-0.349	-1.927
US	-3.055	-2.699	-3.247	-2.893
Panel tests				
t-bar	-2.963	-3.041	-1.952	-2.011
Panel test	-7.663	-7.954	-2.263	-2.574
Individual ADF tests (lag=2)				
Austria	-4.354	-3.762	-3.097	-2.656
Belgium	-2.444	-3.753	-4.489	-2.213
Canada	-4.154	-3.5	-2.433	-2.246
Denmark	-3.288	-5.968	-2.3	-0.305
France	-3.412	-5.129	-3.199	-1.798
Germany	-3.832	-2.275	-1.558	-2.072
Ireland	-4.164	-4.311	-2.205	-1.994
Italy	-4.738	-5.053	-1.353	-2.922
Japan	-5.357	-5.475	-1.678	-0.897
Netherlands	-3.333	-3.417	-1.346	-1.143
Sweden	-3.566	-3.622	-2.289	-2.329
Spain	-3.103	-2.211	-1.298	-2.634
Switzerland	-3.1	-0.741	-0.259	-0.209
UK	-4.462	-3.625	-0.349	-1.74
US	-4.617	-2.679	-3.451	-2.67
Panel tests				
t-bar	-3.862	-3.701	-2.087	-1.855
Panel test	-12.455	-11.630	-3.316	-2.123

Individual tests: 5% critical values for 1 and 2 lags are -2.895 and -2.990 respectively.
Panel tests: 5% critical values is -1.65.

Table 2
Pedroni panel cointegration tests

	Non-homogeneous	Homogeneous
panel v	-0.51379	0.55013
panel ρ	1.59449	0.67276
panel PP t	-0.47479	-1.48322
panel t	-2.28033	-3.05892
group ρ	2.12882	1.32954
group PP t	-1.07620	-1.83090
group t	-6.82999	-8.16296

N= 15 , T= 27

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statistics, only the ADF statistic supports cointegration. In the group statistics, there is weak evidence from the Phillips Perron test and strong evidence from the ADF test. Thus we consider these results to give limited support for the hypothesis that cointegration exists given the CES specification imposing homogeneity.

5.2 Panel Johansen tests

The alternative test we use is the panel version of the Johansen test described above, which also enables us to to test for the number of cointegrating vectors. Table 3 gives Johansen trace test statistics for each country in our sample, maintaining the CES function and homogeneity (which reduces the maximum rank by one). We report both the uncorrected and Reimers-corrected test statistics for all the possible ranks. The uncorrected test statistics reject $r = 0$ for all but five countries (or all but two at 10%); the panel statistic strongly rejects $r = 0$. Moreover, for six countries we reject $r = 1$ at 5%, and the panel test statistic confirms this. When we use the small sample correction the individual statistics now reject $r = 0$ for six countries, but reject $r = 1$ for only three. The panel statistics now clearly indicate that we can reject the hypothesis of no cointegration and cannot reject the hypothesis that there is a single cointegrating vector. We interpret this as strong evidence for the existence of unique cointegrating vectors among this group. Given that the data is certainly non stationary, the fact that we are confident no more than one cointegrating vector exists allows us to apply the dynamic panel methodology reported below. Remember that we have not imposed a homogeneous cointegrating vector among the countries. We test for this in the next section when we estimate the vectors.

Table 3
Johansen Likelihood Ratio (Trace) statistics

	Likelihood Ratio Test Statistics			Reimers adjusted: (T-pk)/T		
	$r=0$	$r=1$	$r=2$	$r=0$	$r=1$	$r=2$
Austria	40.41	12.27	3.36	35.56	10.80	2.96
Belgium	69.74	32.36	8.27	61.37	28.48	7.27
Canada	95.29	29.92	2.79	83.85	26.33	2.46
Denmark	41.05	16.63	4.47	36.12	14.63	3.93
France	77.29	34.81	7.76	68.01	30.63	6.82
Germany	36.25	11.69	2.76	31.90	10.29	2.43
Ireland	37.33	10.26	2.76	32.85	9.03	2.43
Italy	41.91	17.48	4.66	36.88	15.39	4.10
Japan	43.07	26.28	10.79	37.91	23.13	9.50
Netherlands	47.78	13.53	3.05	42.05	11.91	2.68
Sweden	40.55	11.84	1.56	35.68	10.42	1.37
Spain	58.87	28.17	6.24	51.81	24.79	5.49
Switzerland	46.81	10.03	1.25	41.19	8.83	1.10
UK	50.61	23.53	5.25	44.54	20.71	4.62
US	51.74	19.76	3.18	45.53	17.38	2.79
5% crit. value	42.44	25.32	12.25	42.44	25.32	12.25
panel stat.	12.18	2.62	-2.07	8.66	0.81	-2.72

I(1) variables included are p, $p - p^*$, w and $q - n$ with homogeneity imposed. VAR based on one lag including ρ and var as I(0) regressors.
5% critical values are standard Johansen values with deterministic trend in the cointegrating space.
Panel statistic distributed asymptotic standard normal.

6 Estimation of a heterogeneous dynamic panel

In this section we report the results of estimating the parameter of our markup equation (5).

Table 4 reports the unrestricted estimates. Recall they are based on ecm specifications. The lag structure is determined empirically by the SBC, and is allowed to vary between countries. The completely unrestricted (unpooled) individual country results are of no particular interest and are not reported. It is worth noting, however, that as Pesaran, Shin and Smith (1999) observe, in common with most similar studies, we 'find differences in coefficients which are not only statistically significant, but economically implausible.' They go on to observe that the pooled long-run coefficients are more sensible, and this turns out to be the case here. They argue that omitted group specific factors or measurement errors severely bias the country estimates.

It is legitimate to ignore the unrestricted estimates because the Hausman tests do not reject poolability of the long-run parameters, either individually or jointly. This means that the efficient estimates of the common long-run parameters are given by the PMG method (although we also report the mean group results). These are the key results. All the coefficients are very well determined. The coefficient on labour costs has a particularly low standard error and, while numerically greater, is insignificantly different from one. Thus we accept homogeneity. The productivity coefficient is correctly signed but implies an elasticity of substitution larger than is normally assumed; it is around 2. This implies the trend term should be negative, but it is significantly positive, which rejects the CES structure. As argued above, this makes it more difficult to interpret particular coefficients but should not lead to rejection of other results. The coefficient on relative prices is in line with our prior beliefs (higher foreign prices allow margins to increase). A rise in demand raises margins.[31] A rise in uncertainty lowers margins, as we expected. All the error correction coefficients (not reported) lie between -0.1 and -1. Of the 15, three are insignificant. However, the group mean estimate is -0.546 and is highly significant. This is more evidence for cointegration.[32] We report the inefficient mean group estimates. They differ markedly from the PMG estimates but are also much worse determined, reflecting the inefficiency of the mean group method for this dataset. We cannot reject the individual hypotheses that the mean group estimates take their PMG values.

As homogeneity cannot be rejected we reestimated the equations imposing this restriction. Table 5 gives the results. Pooling cannot be rejected for individual coefficients, but is only just accepted jointly. The coefficient on the variance term becomes much smaller and insignificant. The other parameters are much the same, although the coefficient on demand rises. The mean group estimates now differ drastically (especially for the variance), although nothing is significant. As the variance results appear somewhat unrobust, we drop the term. If we impose homogeneity poolability is accepted

[31] These last two results could be interpreted to mean that we have made a small contribution to the PPP debate; PPP is rejected.

[32] The widely dispersed error correction estimates support our judgement regarding the use of the Pedroni group tests.

Table 4
Unrestricted markup equation.

Dependent variable p_{it}.

	Pooled MGE Estimates			MGE Estimates				
	Coef.	St. Er.	t-ratio	Coef.	St. Er.	t-ratio	h-test	p-val
w	1.028	0.016	63.195	0.904	0.271	3.339	0.21	0.65
$q-n$	-0.536	0.060	-8.928	0.244	1.263	0.193	0.38	0.54
$p-p^{*}$	-0.400	0.047	-8.521	0.827	0.785	1.053	2.45	0.12
ρ	0.725	0.096	7.554	0.513	0.471	1.089	0.21	0.65
var	6.696	2.038	3.286	76.254	39.074	1.952	0.18	0.67
Joint Hausman test							4.55	0.47
θ	-0.546	0.089	-6.171	-0.828	0.161	-5.127		
constant	-0.007	0.001	-5.061	-0.019	0.016	-1.213		
trend	0.107	0.029	3.705	0.328	0.440	0.747		

θ mean group ecm coefficient.

---

and the well determined parameters are extremely close to those reported in Table 4. Thus although there is clearly a question surrounding the role of uncertainty, the results regarding productivity, homogeneity, competitors' prices and demand appear robust. These conclusions could not have been reached by examining the individual country estimates alone. Fianlly, Table 7 gives the results for the key parameters of a simple fixed effects regression. The parameters are plausible, but numerically different from the PMG results.

7 Conclusions

The availability of country panels where the data can be described as 'data fields' with roughly equivalent N and T make it possible to exploit information from the cross-section about time series relationships. However, where dynamic models are called for, which will be the rule with typical non-stationary macroeconomic time series, standard pooled models are not simply inefficient but may also be highly inconsistent. This is also true for dynamic models with stationary regressors. Nevertheless, there may be a case for estimating long-run parameters in a pooled framework, as suggested by Pesaran and Smith (1995) and Pesaran, Shin and Smith (1999). This restricted poolability can be tested for by a Hausman type test. We examine a sample of 15 OECD countries for which we have consistent data. In our data set, individual and panel tests for order of integration reveal that the core variables are non-stationary. Thus for estimation to be valid, the data must also satisfy tests for the unique existence of long-run relationships (which may nevertheless be heterogeneous between countries). We test for cointegration using Pedroni's residual based panel unit root tests, and a panel version of the Johansen trace statistic. There is indeed evidence for

Table 5
Restricted markup equation: homogeneity.

Dependent variable p_{it}; coefficient on w_{it} restricted to unity.

	Pooled MGE Estimates			MGE Estimates				
	Coef.	St. Er.	t-ratio	Coef.	St. Er.	t-ratio	h-test	p-val
$q-n$	-0.505	0.082	-6.132	-6.536	6.195	-1.055	0.95	0.33
$p-p^*$	-0.374	0.064	-5.858	-15.827	16.408	-0.965	0.89	0.35
ρ	0.917	0.137	6.689	-26.753	26.566	-1.007	1.08	0.30
var	0.044	4.628	0.010	-750.275	614.408	-1.221	1.49	0.22
Joint Hausman test							8.87	0.06
θ	-0.396	0.080	-4.951	-0.519	0.105	-4.939		
constant	-0.003	0.001	-4.348	-0.009	0.004	-2.460		
trend	0.080	0.022	3.629	0.222	0.071	3.124		

Table 6
Restricted markup equation: homogeneity; excluding VAR.

Dependent variable p_{it}; coefficient on w_{it} restricted to unity.

	Pooled MGE Estimates			MGE Estimates				
	Coef.	St. Er.	t-ratio	Coef.	St. Er.	t-ratio	h-test	p-val
$q-n$	-0.597	0.063	-9.519	-0.181	0.806	-0.224	0.27	0.60
$p-p^*$	-0.387	0.046	-8.432	-0.327	0.562	-0.581	0.01	0.91
ρ	0.735	0.108	6.805	-0.088	0.644	-0.137	1.68	0.20
Joint Hausman test							3.16	0.37
θ	-0.443	0.089	-4.988	-0.517	0.114	-4.555		
constant	-0.002	0.001	-3.623	-0.004	0.003	-1.310		
trend	0.068	0.023	3.003	0.112	0.067	1.658		

Table 7
Static fixed effects estimates of the unrestricted markup equation.

Dependent variable p_{it}.

		FE		Robust	
	Coef.	St. Er.	t-ratio	St. Er.	t-ratio
w	0.9781	0.0184	53.2406	0.0546	17.9235
$q-n$	-0.4848	0.0444	-10.9122	0.1259	-3.8506
$p-p^*$	-0.1962	0.0469	-4.1810	0.1295	-1.5146
ρ	0.5489	0.1417	3.8744	0.1402	3.9153
var	11.3933	3.2343	3.5227	3.1331	3.6364

the existence of unique cointegrating relationships in the panel members. Moreover, the Hausman test cannot reject pooling of the long-run parameters. The pooled estimates accord with our expectations and are well determined. It appears that for these countries long-run price formation can be modelled as a variable markup over costs driven by productivity, labour costs and technical progress. The markup also depends on competitors' prices, deviations from trend output and uncertainty. These conclusions could not have been reached by examining the individual country estimates alone. Nor could they have been reached by conventional panel estimation. Reliance on a static panel would be misleading. There are many applications to cross-country data where the same methodology can fruitfully be applied.

Appendix: Data sources

The countries in the sample were Austria (AU), Belgium (BE), Canada (CN), Denmark (DK), France (FR), Germany (GE), Ireland (IR), Italy (IT), Japan (JP), Netherlands (NL), Sweden (SD), Spain (SP), Switzerland (SW), the United Kingdom (UK) and the United States (US). All data were annual over the period 1970 to 1996, with the exception of manufacturing output, which was 1966 to 1997. The longer period was essential to avoid the 'end-point' bias in the Hodrick-Prescott filter.

Unit labour costs (ULC = WN/Q) (1990=100) - series for all countries apart from Australia, Ireland, Spain and Switzerland from BLS. For the US, data from 1977-96 from BLS. Data for 1970-76 constructed using manufacturing output from OECD National Accounts and labour compensation from the National Income and Product Accounts. For Australia, data for 1981-95 from OECDHS. Data for 1970-80 and 1996 constructed using information on earnings, output and employment in manufacturing from OECDHS and MEI. For Ireland, data constructed using structed using data on manufacturing employment and earnings and industrial production from OECDHS. For Spain, data from 1981-95 from OECDHS. Data for 1970-80 interpolated in line with whole economy unit labour costs constructed from compensation and GDP in OECD National Accounts. For Switzerland, data from OECDHS, various issues.

Manufacturing Import Prices (P^*) (1990=100) - from EO unless otherwise stated. For the UK, prices constructed using the value and volume of imported manufactures from the National Accounts. For Switzerland, data from IFS country page (line no. 76). For Ireland, data for aggregate merchandise imports. German data for 1996 estimated using growth rate of merchandise import prices in OECD Economic Outlook, June 1998.

Producer Prices (P) (1990=100) - data for US, Canada, Japan, Germany, France, UK, Italy and Spain from NiGEM. Data for Denmark, Ireland, Netherlands, Sweden and Switzerland from IFS. Data for Australia from OECDHS. For Belgium, data for 1980-96 from IFS, data for 1970-79 from OECDHS. Definitions of series may differ across countries, with some having wholesale price data.

Manufacturing Employment (N) (1990=100) - data from BLS unless otherwise stated. For Australia, Ireland, Spain and Switzerland data from OECDHS. Adjustments made to Australian series in 1989 and Switzerland series in 1985 to allow for breaks in definition.

Manufacturing Output (Q) (constant prices, 1990=100) - data from BLS unless otherwise stated. For US, data for 1966-76 from OECD National Accounts. For Australia, data from OECDHS. For Ireland, Spain and Switzerland we use industrial production from IFS.

BLS - International Comparisons of Manufacturing Productivity and Unit Labour Cost Trends, Bureau of Labor Statistics. EO - OECD Economic Outlook, historical data diskettes. IFS - International Financial Statistics Yearbook, International Monetary Fund. MEI - Main Economic Indicators, OECD. NiGEM - National Institute Global

Econometric Model database. Further details available on request. OECDHS - OECD Historical Statistics 1960-95.

References

Alogoskoufis, G, Martin, C and Pittis, N (1990) 'Pricing and product market structure in open economies: an empirical test' **CEPR Discussion Paper** No 486.

Arellano, M and S Bond (1991) 'Some tests of specification for panel data: Monte Carlo evidence and an application to employment equations' **Review of Economic Studies 58** pp 277-320.

Baltagi, B H and J M Griffin (1997) 'Pooled estimators vs. their heterogeneous counterparts in the context of dynamic demand for gasoline' **Journal of Econometrics 77** pp 303 -327.

Banerjee, A, M Marcellino and C Osbat (2000) 'Some cautions on the use of panel methods for integrated series of macro-economic data' EUI.

Bannerjee, A, L Cockerell and B Russell (1998) 'An I(2) analysis of inflation and the markup' **Journal of Applied Econometrics 16** pp 221-240.

Bannerjee, A, J Dolado, D F Hendry and G W Smith (1986) 'Explaining equilibrium relationships in econometrics through static models: some Monte Carlo evidence' **Oxford Bulletin of Economics and Statistics 48** pp 253-277.

Bils, M (1987) 'The cyclical behavior of marginal cost and price' **American Economic Review 77** pp 838-855.

Bils, M (1989) 'Pricing in a customer market' **Quarterly Journal of Economics 104** 699-718.

Breshanam, T (1989) 'Empirical studies of industries with market power' Chapter '17 in **Handbook of Industrial Organisation Volume II** eds R Schmalensee and R Willig (Amsterdam: North Holland).

Bulow, J, J D Geanakoplos and P D Klemperer (1985) 'Multimarket oligopoly: strategic substitutes and complements' **Journal of Political Economy 93** pp 488-511.

Chevalier, J A and D S Scharfstein (1996) 'Capital-market imperfections and countercyclical markups: theory and evidence' **American Economic Review 86** 703-725.

Dixit, A and R S Pindyck (1994) **Investment Under Uncertainty** (Princeton University Press).

Driver, C and D Moreton (1992) **Investment, Expectations and Uncertainty** (Basil Blackwell, Oxford).

Gali, J, M Gertler and D López-Salido (2001) 'European inflation dynamics' **European Economic Review 45** pp 1237-1270.

Gerdtham, U-G and M Löthgren (2000) ' On Stationarity and Cointegration of international health expenditure and GDP' **Journal of Health Economics 19** pp 461-475.

Hall, S and G Urga (1997) 'Recent developments in the econometrics of panel data and the implications of non-stationarity' LBS CEF Discussion Paper DP 06-97.

Hartman, R (1972) 'Factor demand with output price uncertainty', **American Economic Review 66** pp 675-681.

Im, K S, M H Pesaran and Y Shin (1997) 'Testing for unit roots in heterogeneous panels' Cambridge University.

Ireland, P (1998) 'Customer flows, countercyclical markups and the output effects of technology shocks' **Journal of Macroeconomics 20** 649-664.

Johansen, S (1995) **Likelihood-Based Inference in Cointegrated Vector Auto-Regressive Models** OUP (Oxford).

Larsson, R, J Lyhagen and M Löthgren (2001) 'Likelihood-based cointegration tests in heterogeneous panels' **Econometrics-Journal 4** pp 109-42.

Larsson, R and J Lyhagen (1999) 'Likelihood-based inference in multivariate panel cointegration models' Stockholm School of Economics.

Karlsson, S and M Löthgren (2000) 'On the power and interpretation of panel unit root tests' **Economics Letters 66** pp 249-255.

Martin, C (1997) 'Price formation in an open economy: theory and evidence for the United Kingdom, 1951-1991' **Economic Journal 107** pp 1391-1404.

McCoskey, S and Kao, C (2001) 'A Monte Carlo comparison of tests for cointegration in panel data' **Journal of Propagations in Probability and Statistics 1** pp 165-198.

Neiss, K (2001) 'The markup and inflation: evidence in OECD countries' **Canadian Journal of Economics 34** pp 570-587.

Pain, N and K Wakelin (1998) 'Export performance and the role of foreign direct investment' **Manchester School 66** Supplement pp 62-88.

Pedroni, P (1999) 'Critical values for cointegration tests in heterogeneous panels with multiple regressors' **Oxford Bulletin of Economics and Statistics 61** pp 653-670.

Pesaran, M H, Y Shin and R Smith (1999) 'Pooled mean group estimation of dynamic heterogeneous panels' **JASA 94** pp 621-634.

Pesaran, M H and R Smith (1995) 'Estimation of long-run relationships from dynamic heterogeneous panels' **Journal of Econometrics 68** pp 79-113.

Pesaran, M H, R Smith and K S Im (1996) 'Dynamic linear models for heterogenous panels' in L Mátyás and P Sevestre **The Econometrics of Panel Data** Kluwer Academic (London).

Price, S (1991) 'Costs, Prices and Profitability in UK Manufacturing' **Applied Economics 23** pp 839-850.

Price, S (1992) 'Forward Looking Price Setting in UK Manufacturing' **Economic**

Journal 102 pp 497-505.

Price, S (1994) 'Aggregate uncertainty, forward looking behaviour and the demand for manufacturing labour in the UK' **Oxford Bulletin of Economics and Statistics 56** pp 267-283,

Price, S (1996) 'Aggregate uncertainty, investment and asymmetric adjustment in the UK manufacturing sector' **Applied Economics 28** pp 1369-1379.

Quah, D (1993) 'Empirical cross-section dynamics in economic growth' **European Economic Review 37** pp 426-434.

Reimers, H-E (1992) 'Comparisons of tests for multivariate cointegration' **Statistical Papers 33** pp 335-359.

Rotemberg, J J and G Saloner (1986) 'A supergame-theoretic model of price wars during booms' **American Economic Review 76** pp 390-407.

Rotemberg, J J and M Woodford (1991) 'Markups and the business cycle' in **NBER Macroeconomics Annual** (NBER) eds O J Blanchard and S Fischer pp 63-128.

Rotemberg, J J and M Woodford (1995) 'Dynamic general equilibrium models with imperfectly competitive product markets' in **Frontiers of Business Cycle Research** (Princeton University Press) ed. T F Cooley pp 243-93.

Smith, P N (2000) 'Output price determination and the business cycle' **Economic Modelling 17** pp 49-69.

Zellner, A (1962) 'An efficient method of estimating seemingly unrelated regressions and tests of aggregation bias' **Journal of the American Statistical Association 57** pp 348-368.