A Reinvestigation of the Markup and the Business Cycle*

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Abstract

A fresh interpretation is provided of the influential finding that the markup of prices over marginal costs is counter-cyclical. Using Rotemberg and Woodford's (1991) data set we argue that the markup is best modelled as a variable that is integrated of order one. A consequence of this finding is that the markup cannot be related in the long run with business-cycle variables since these are traditionally thought of as being stationary. A distinction must therefore be made between the long- and the short-run behaviour of the markup. It is shown that the markup is negatively related to inflation in the long-run, while stationary transforms of the markup are counter-cyclical in the short-run.

Keywords: Inflation, Prices, Markup, Productivity, Business Cycles, Cointegration.

JEL Classification: C22, C32, C52, D40, E31, E32

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1. INTRODUCTION

In their influential paper, Rotemberg and Woodford (1991) (henceforth called R-W) argue that the markup of prices on marginal costs is counter-cyclical.[1] Their argument begins with a systematic theoretical analysis of the sources of the variation in the marginal cost markup over the business cycle in imperfectly competitive models. R-W then show empirically that the markup is counter-cyclical with a graphical analysis of the markup and the business cycle before estimating regressions of the markup on variables that proxy current and future demand.

The purpose of our analysis is to argue that there is a further level of complexity to the empirical modelling above that is worthy of investigation. The argument has two elements. First that the markup and inflation are best modelled as integrated series of order one. And second, that while business cycle variables certainly have an impact on the markup, inflation has an important influence on the markup in the long run.[2] Another way of expressing this idea is to make a distinction between the short- and long-run effects on the markup as captured by the business cycle variables and inflation respectively. This observation is inspired by the empirical analysis in R-W and our own work (Banerjee, Cockerell and Russell (2001), Banerjee and Russell (2001a, b) and Banerjee, Mizen and Russell (2002)) that establishes a long-run relationship between inflation and the markup.

If one grants the existence of a long-run relationship between inflation and the markup, this implies that empirical investigations of short-run relationships between the markup, productivity, inflation and the business cycle may provide spurious results unless the long-run relationship between inflation and the markup is explicitly allowed for. Our paper allows

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1 A number of authors argue in favour of a counter-cyclical markup. For example, see the staggered pricing models of Calvo (1983) and Rotemberg (1982), elasticity-of-demand models of Gali (1994), customer market models of Phelps and Winter (1970), and the implicit collusion model of Rotemberg and Woodford (1992). Alternatively, see the macroeconomic models of Layard, Nickell and Jackman (1991), Lucas (1973), Kydland and Prescott (1988), and Blanchard and Kiyotaki (1987).

2 The long run is defined in the sense of Engle and Granger (1987).

for such a long-run relationship while empirically examining the cyclical behaviour of the markup and inflation.

The paper proceeds in the next section by considering briefly theories that link inflation and the markup. Theories of the cyclical behaviour of the markup are reviewed exhaustively in Rotemberg and Woodford (1999). In section 3 we begin the empirical analysis by estimating the R-W model of the markup. For comparability we use the R-W data to highlight the interrelationship between the short-run cyclical and long run influences on the markup.[3] We show using a number of standard techniques that the specification of the empirical model in R-W is inadequate because of a missing integrated variable. In particular, this implies that the *t*-statistics cannot be interpreted using standard critical values and that the coefficient estimates may therefore demonstrate spurious significance.

We speculate that the 'missing integrated variable' is inflation (in line with the theoretical arguments presented in Banerjee, Cockerell and Russell (2001)) and re-estimate the model to incorporate a long-run relationship between inflation and the markup while retaining all the variables in the original R-W specification. This specification of the model shows that the change in the markup is counter-cyclical and that there exists a negative long-run relationship between inflation and the markup.

In order to interpret the results of this re-estimation we show graphically that a subset of the variables (the trend and levels of output and Tobin's *q*) in the long-run relationship, when combined linearly with the estimated coefficients, may act as a proxy for the level of productivity. Furthermore, on a conceptual level, the difference between the markup on marginal cost in R-W and the markup of price on average wages is due to the business cycle and this difference is presumably stationary. Consequently, the stationary component should not affect estimates of the long-run relationship between inflation and the markup. This

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3 Using the actual R-W data allows us to focus directly on the theoretical and empirical issues without recourse to explaining the coefficient estimates of the R-W specification over a longer sample. See also footnote 9.

implies that, in identifying the long-run relationship it does not matter if the markup is measured on marginal costs (as in R-W) or on average costs.[4]

To make this same point empirically we complete the empirical analysis by estimating a three variable system. The system is conditioned on the business cycle and comprises the markup on average wages (replacing the markup on marginal costs in the R-W specification), productivity (replacing the trend and levels of output and Tobin's *q* in the R-W specification) and inflation. We show that similar estimates of the long-run inflation coefficient can be obtained under this reduction. This shows that, as expected, the results are qualitatively insensitive to whether the analysis makes use of the markup on marginal costs or the markup on unit costs, conditional on allowing for integrated data and the presence of a long-run relationship between inflation and the markup. The counter-cyclicality perforce is between the business cycle and the *difference* of the markup, since the markup is an integrated variable of order one.

2. INFLATION AND THE MARKUP

Rotemberg (1983), Kuran (1986), Naish (1986), Danziger (1988) and Koniezny (1990) set out monopolistic models where inflation interacts with small 'menu' costs of price adjustment. One result of these papers is that inflation affects the *average* markup of firms for a given profit maximising markup. Bénabou and Koniezny (1994) show in an encompassing monopolistic model that the relationship between inflation and the markup depends on the skewness of the profit function and the relative size of inflation, the 'menu' costs and the discount rate. Bénabou and Koniezny also make an important distinction between the partial equilibrium analysis of these papers where firms operate independently of each other and an equilibrium analysis where inflation impacts on the profit maximising markup.[5]

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4 See Banerjee and Russell (2002) for further detail on this argument.

5 The assumption of independent firms implies that the cost and demand functions are exogenous and that the profit maximising price consistent with zero inflation or zero 'menu' costs is constant.

Bénabou (1992) proposes an equilibrium model where higher rates of inflation reduce the markup on marginal costs. Focusing on the behaviour of customers, he argues that higher inflation leads to greater customer search which in turn increases competition and leads to a lower markup.[6]

While Bénabou implicitly argues that the relationship is of a short-run nature, the argument can be extended to imply that inflation and the markup may also be negatively related in the long run.[7] If competition remains higher with higher inflation in the long-run (because of permanently higher customer search due to greater variance of price changes) then the relationship between inflation and the markup will remain in the long run. Russell (1998), Russell, Evans and Preston (2002) and Chen and Russell (2002) explicitly argue that the markup and inflation are negatively related in the long run. These papers focus on the difficulties that non-colluding imperfectly competitive firms face while coordinating price changes when information is missing. These difficulties increase with higher inflation, as prices need to be changed more frequently, by larger amounts in real terms, or some combination of the two. Given the basis of the difficulty in coordinating price changes the difficulties persist in the long run, and consequently so does the negative relationship between inflation and the markup.

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6 Simon (1999) draws on an insight in Athey, Bagwell and Sanichiro (1998) in his empirical analysis of the inflation-markup relationship. The insight is that higher inflation increases the variance of cost increases making it more difficult for firms to collude when changing prices. The reduction in collusion increases competition and the markup falls with higher inflation.

7 The empirical analysis of Bénabou (1992) and Simon (1999) assume, either implicitly or explicitly, that inflation and the markup are stationary and consequently the possibility of a long-run relationship is not explored. The partial equilibrium models mentioned above may also provide an explanation of a long-run relationship between inflation and the markup. However, a partial equilibrium analysis where the demand and cost curves are exogenous would appear to be inconsistent with the concept of the 'long-run'.

3. I(1) SYSTEM ESTIMATES OF THE MARKUP

3.1 Re-estimating the Rotemberg and Woodford Model

R-W build on Solow (1957) and Hall (1988) and derive an expression for the variation in the markup on marginal costs, μ_t, around its steady state value in terms of the business cycles in output, y_t, the capital stock, k_t, employment, h_t, and the markup of price on average wages, $(p-w)_t$. The expression can be written in the form:

$$\bar{\mu}_t = \delta_0\,\bar{y}_t + \delta_1\,\bar{k}_t - \delta_2\,\bar{h}_t + \overline{(p-w)}_t \tag{1}$$

where, δ_0, δ_1, and δ_2 are positive parameters that depend on the steady state values of the marginal cost markup, the elasticity of substitution between the two factor inputs (capital and labour) and the factor shares of capital and labour.[8] Lower case variables are in logarithms and the 'bar' on a variable indicates deviation from the trend value of the variable.

The R-W theoretical model is in terms of deviations from trend while the associated empirical modelling in the paper is in terms of the level of the marginal cost markup, μ_t. That is, the level of the marginal cost markup is constructed using (1) with levels of the variables instead of deviations from trend (i.e. ignoring the bars in equation (1)). Consequently, this measure of the marginal cost markup displays a distinct trend decline due to the trend increase in the real wage over the sample. R-W detrend the markup in their subsequent empirical analysis by use of a linear trend in their regressions in contrast with the equivalent method of prior de-trending of the component variables as in (1).

Having constructed the marginal cost markup using (1), R-W regress this variable on output and Tobin's 'q' that represent measures of current and future demand respectively. R-W present (on page 95) the following 'baseline' results from United States quarterly data for the period June 1952 to December 1988:

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[8] Equation 3.6 on page 84 of Rotemberg and Woodford (1991). Johri (2001) derives a similar expression. See also Bils and Chang (2000) and Basu (2000).

$$\mu_t = \underset{(0.5)}{0.77} + \underset{(0.0007)}{1.4\times10^{-5}}\, trend - \underset{(0.08)}{0.63}\, y_t + \underset{(0.015)}{0.058}\, q_t \qquad (2a)$$
$$R^2 = 0.983 \qquad DW = 0.16$$

$$\mu_t = -\underset{(0.6)}{0.72} - \underset{(0.0007)}{0.002}\, trend - \underset{(0.09)}{0.42}\, y_t + \underset{(0.014)}{0.035}\, q_t \qquad (2b)$$
$$\rho = 0.934 \qquad R^2 = 0.997 \qquad DW = 1.54$$

where μ_t is the constructed markup of price on marginal costs assuming a steady state marginal cost markup of 1.6 and an elasticity of capital labour substitution of 1.0, and q_t is Tobin's '*q*'. Results reported as (2b) are estimated allowing for first order serial correlation.[9] The data appendix provides details of the R-W data used in estimating (2a) and (2b). The same data are used in the empirical analysis later in the paper.

The R-W estimates implicitly assume (by including the deterministic trend variable in (2a) and (2b)) that the variables are trend stationary. However, both univariate and multivariate unit root tests indicate that the markup, output and Tobin's '*q*' variables are integrated processes.[10] It appears, therefore, that (2a) should be interpreted as a long run or static cointegrating regression as long as the residuals are stationary. The autocorrelation coefficient estimate, ρ, of 0.93 in the Cochrane-Orcutt transformed model (2b) however suggests that the residuals of (2a) contain a variable that is close to integrated.[11] The low value of the Durbin-Watson statistic in relation to R^2 also suggests the same phenomenon. It

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9 Estimating (2a) and (2b) with updated data to March 2001 we find Tobin's *q* insignificant and the coefficient on output considerably smaller (- 0.26 compared with – 0.63 in (2a)). The change in the relationship between Tobin's *q* and the markup is not surprising given the size and persistence of the stock market 'boom' in the 1990s.

10 This conclusion can be verified using PT and DF-GLS univariate unit root tests from Elliot, Rothenberg and Stock (1996) and by looking at the correlogram for each series that indicate high degrees of persistence of the correlations. Multivariate stationarity tests (*i.e.* generalisations of univariate Dickey-Fuller tests) confirm these results.

11 This is in addition to the problem that the common factor restrictions implied by the Cochrane-Orcutt transformation in (2b) are strongly violated in this data set upon testing. Correcting for first-order autocorrelation is therefore not justified.

appears, therefore, that (2a) is a spurious regression.[12] Hence, including a deterministic trend is an inappropriate device for detrending the variables of the regression and the *t*-statistics cannot be interpreted conventionally. We corroborate this finding more formally in what follows in two ways. The first uses the slightly complicated device of simulating pseudo-data series for μ_t in order to determine the sampling properties of the coefficient estimates in (2a) and (2b) by means of repeated estimation of these equations on pseudo data. The second approach is to determine the cointegrating rank of the variables in (2) by maximum likelihood methods.

3.1.1 Re-estimating the R-W Model – A Bootstrap Approach

Given the well-known unreliability of tests for unit roots, a specification test may be undertaken that does not rely on such pre-testing. We construct 100,000 pseudo-series for μ_t denoted $\tilde{\mu}_t$ using the actual data series for y_t and q_t, the estimated coefficients (including that for the autocorrelation coefficient) reported by (2b) and re-sampling from the residuals ε_t. The latter are taken to be white noise since the results are not sensitive to the value chosen for their variance. For each of these pseudo-series, regressions (2a) and (2b) are estimated for the given series y_t and q_t, and the empirical densities of the *t*-statistics for the coefficient estimates are tabulated.

For Data Generation Process 1 (labelled DGP1 in the appendix), the $\tilde{\mu}_t$ series are generated simply as a constant and trend with an AR(1) error term (with ρ varying from 0 to 1), so that the null hypothesis of no influence from y_t and q_t onto μ_t is true for this process.[13]

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12 Granger and Newbold (1974) gave $R^2 > DW$ as an informal criterion for judging a spurious regression.

13 It may be seen from the specification of the data generation process given in the appendix, that the higher the value of ρ, the more dominant is the stochastic trend component in the $\tilde{\mu}_t$ process while for smaller values of ρ, the deterministic component predominates. This is seen most easily by re-writing the process for $\tilde{\mu}_t$ in slightly more expanded form as $\tilde{\mu}_t = [c(1-\rho)+\gamma\rho]+\gamma(1-\rho)t+\rho\tilde{\mu}_{t-1}+\varepsilon_t$. We would therefore expect difficulties to arise with the use of standard critical values in cases where the error process

Models (2a) and (2b) are estimated with the generated $\tilde{\mu}_t$ and the actual y_t and q_t. Table A1 in the appendix provides the rejection frequencies (at 5 per cent significance level) of the *t*-statistics in models (2a) and (2b) if the critical values of a standard normal density are used. Thus, if the use of normal critical values were valid, one should expect a rejection frequency of the null hypothesis of insignificance of the coefficient estimates of y_t and q_t to be approximately 5 per cent. This indeed occurs when the value of ρ is 0. Significant size distortions are however evident even for values of ρ of 0.5 while if the value of the autocorrelation coefficient is 0.93 as in (2b), the rejection frequency is in fact close to 70 per cent.[14]

Two further points may be noted. Following on from above, the empirical quantiles of the *t*-densities (that may be computed as a by-product of this bootstrapping exercise) show that for the particular sample size the critical value that gives the correct size of 5 per cent is 6.0 for the coefficient of output, y_t, and 3.5 for the coefficient of Tobin's *q*. Therefore, both the coefficient estimates in (2b) reported above are in fact *insignificant* at 5 per cent if judged by the size adjusted critical values. This is not to suggest that y_t and q_t do not have an influence on the markup but simply to state that (2a) and (2b) are inadequate empirical descriptions of the data. Second, generating data with more stationary values of the autocorrelation coefficient lead to rejection frequencies closer to the nominal levels, so that the over-rejections can be attributed to the persistence in the residuals.

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is persistent, *i.e.* ρ is large. This observation is borne out fully in the results of the experiments described below.

14 This finding replicates the results reported in Banerjee, Dolado, Galbraith and Hendry (1993) in their simulations on spurious regressions. Note that as expected the null hypothesis of $\rho = 0$ is rejected with frequency approximately equal to 5 per cent when ρ is zero in the data generation process and is rejected 100 per cent of the time for non-zero values of ρ in the data generation process. The rejection frequency of the constant is surprisingly low. This however, is a consequence of the asymptotic behaviour of the estimator of the constant that is driven to zero at rate $1/T$ in a spurious regression (see Phillips (1986)). The rejection frequency for the trend coefficient is also small because of the low value for this parameter in the data generation process.

Table A2(i) and A2(ii) in the appendix report the results from repeating the above exercise where the data generation process and the model are given by (2b), labelled DGP2. The trend is omitted from the data generation process given that the actual variables in the data are already highly trending but is included in the model to mimic the effect of deterministically detrending variables that are trending stochastically. Rejection frequencies for test statistics computed as deviations from zero (Table A2(i)) and as deviations from their true values (Table A2(ii)) confirm the over-rejection phenomenon. The estimation of (2b) - which involves a Cochrane-Orcutt transform - is not justified, given the violation of the common factor restrictions on the data. Nevertheless, such a transform does ameliorate the over-rejection phenomenon to a certain extent (with the gain diminishing as the residuals become more integrated), although the sizes of the tests are still too large.

Thus based only on the data, and free from all pre-testing, we see that proper specification of the markup equation requires consideration of a missing integrated variable. The spurious-regression-like behaviour of the models above is generated entirely by the actual data, so that the markup inherits the integration properties of the right hand variables and must be related not only to y_t and q_t but also to a variable that, in accordance with theoretical models presented in Section 2, we shall take to be inflation. We show below that estimating a system with inflation incorporated, leads to well-specified and interpretable results that encompass the R-W findings and lead to interesting observations on the behaviour of the markup in the short- and long-run.

3.1.2 Re-estimating the R-W Model by Maximum Likelihood

In the terminology developed by Engle and Granger (1987), the presence of a near-integrated variable in the residuals in equation (2b) may be taken to imply that the markup, output and Tobin's q are not cointegrated. That is, the cointegrating rank, r, of this trivariate system is zero. We therefore close this section by re-estimating equation (2a) using I(1) system techniques developed by Johansen (1988). In order to mimic equation (2a) the system comprises the markup, output, Tobin's q and a trend in the cointegrating space. The constant is unrestricted. Table 1 indicates that the variables are not cointegrated since the trace statistic of $r = 0$ versus $r = 3$ lies well below the 90 per cent asymptotic critical value.[15] The result of no cointegration is entirely consistent with the findings reported above.

Table 1: Testing for the Number of Cointegrating Vectors
Estimated Values of Q(r)

$H_0 : r =$	Eigenvalues	*Q(r)*
0	0.0952	29.80 {39.08}
1	0.0770	15.20 {22.95}
2	0.0237	3.50 {10.56}

Notes: Statistics are computed with three lags of the core endogenous variables. The sample is from December 1952 to March 1989. *Q(r)* is the likelihood ratio statistic for determining the rank, *r,* in the I(1) analysis. 90 percent critical values shown in curly brackets { } are from Table 15.4 of Johansen (1995).

3.2 Introducing Inflation into the Rotemberg and Woodford Model

The results so far on the R-W empirical model indicate that it is more appropriate to model the markup as an integrated variable than a trend stationary one and that it appears the model is missing an integrated component. As argued above, there may exist a long-run relationship between the markup and inflation, implying that the missing integrated variable in the R-W framework may be inflation. To provide a unified framework, we proceed to estimate a four variable system comprising the R-W marginal cost markup, real output, Tobin's *q and* inflation with a trend in the cointegrating space. The estimated system is conditioned on the business cycle measured as linearly de-trended logarithm of the hours of employment.[16] The results are reported in Table 2 and we can now conclude in favour of one cointegrating vector.

Table 2: I(1) System Estimates

[15] This result is supported by the eigenvalues of the companion matrix where we find two roots on the unit circle and the third close to unity with a value of 0.8801.

[16] Private sector employment data is from the R-W dataset.

Markup on Marginal Costs, Output, Tobin's *q* and Inflation

Cointegrating Vector: $\underset{}{\mu} - \underset{[0.207]}{0.710\,y} + \underset{[0.024]}{0.095\,q} + \underset{[1.341]}{7.646\,\Delta p} + \underset{[0.002]}{0.010\,t}$

Equation $\Rightarrow$	$\Delta\mu$	Δy	Δq	$\Delta^2 p$
Error Correction Term	- 0.089 (- 3.5)	0.017 (0.6)	- 0.080 (- 2.3)	- 0.080 (-6.5)
Business Cycle	- 0.104 (- 2.6)	- 0.059 (- 1.4)	- 0.847 (- 2.9)	- 0.090 (- 4.7)
R^2	0.64	0.52	0.39	0.59

$LM(1): \chi^2_{16} = 18.35$, $p-value = 0.30$; $LM(4): \chi^2_{16} = 19.07$, $p-value = 0.26$;
$D-H : \chi^2_8 = 8.91$, $p-value = 0.35$

Notes: Data sample is December 1952 to March 1989. Standard errors reported in [ ] brackets and *t* statistics are in ( ) brackets.

The system is estimated with three lags of the endogenous variables. Predetermined variables are the business cycle lagged one quarter and a series of 'spike' dummies for December 1952, March and December 1953, September 1959, March 1971, September 1974, June 1978, June 1980, December 1982 and March 1989.

Trace statistics for 0, 1, 2 and 3 cointegrating vectors are 88.53 {58.96}, 35.85 {39.08}, 17.12 {22.95} and 3.66 {10.56} respectively where numbers in { } are the relevant 90% critical values. Inference concerning the number of cointegrating vectors is maintained if the system is estimated without pre-determined variables. Moduli of the first five roots of the companion matrix are 1.0, 1.0, 1.0, 0.6461, 0.6461.

$LM(1)$ and $LM(4)$ are Lagrange multiplier tests for first and fourth order autocorrelation and D-H is Doornik-Hansen normality test.

The results in Table 2 raise three interesting issues. First, the finding in R-W that the markup is related to output and Tobin's q is re-established. However, given that we argue that the data are difference stationary and not trend stationary, the level of output, y , and Tobin's q in the cointegrating vector can no longer be interpreted as business cycle measures of current and future demand. Moreover, as the markup is an integrated process it is now no longer appropriate to consider a relationship between the business cycle and the level of the markup unless the business cycle is also an integrated process. As the business cycle variable considered here is stationary, the only possible relationship is that between the change in the markup and the business cycle.

These observations are of course conditioned on the variables behaving like integrated processes for the time-period under consideration. In episodes where this is not the case, it may be possible to recover a R-W type specification in its entirety, that is, without the need to incorporate inflation and the markup would be influenced mainly by the business cycle. This is in contrast to the analysis above where the business cycle acts as a stationary perturbation around a long-run relationship between inflation and the markup.

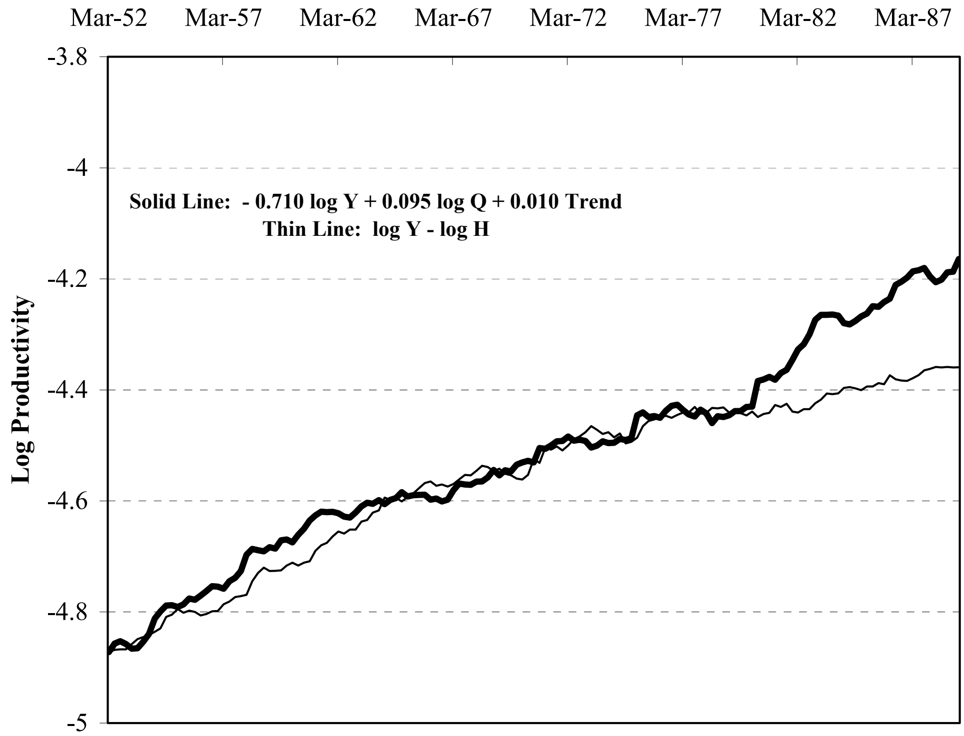
Second, the question arises of how to interpret the trend and level of output and Tobin's *q* in the long-run relationship. The long-run relationship is similar to the long-run relationship between inflation and the markup estimated in our earlier work that is based on the markup of price on average wages, productivity and inflation. This differs from the long-run specification above in that it includes the levels of output and Tobin's *q* and the trend but excludes the level of productivity. One may therefore suspect, that y and q combine with the trend in the long-run relationship above to 'proxy' productivity that is missing from the standard inflation markup long-run relationship. The thick line in Graph 1 shows this estimated 'proxy' as measured by output, Tobin's *q* and the trend variable weighted by the estimated coefficients from the cointegrating vector. The thin line shows average productivity, $(y-k)_t$. We see that up until the last few years there is a high correspondence between these two measures of productivity indicating our suspicion may be correct.[17]

Finally, the cointegrating vector establishes a negative long-run relationship between the markup on marginal costs and inflation.

[17] The divergence in the 1980s is due to the boom in the stock market that increases Tobin's *q*.

Graph 1

PRODUCTIVITY



3.3 Estimating a Unit Cost Markup Model

Measures of the markup on marginal costs used in the R-W empricical work diverge from the markup of price on average wages due to the impact of the business cycle on marginal productivity. Assuming the business cycle influence is short-run in nature, this implies the markup on marginal costs will be integrated since the markup series that it is derived from is integrated. These short-run influences should be identified not in the long-run cointegrating relationship but by a short-run business cycle term. Furthermore, these short-run influences should not affect the long-run estimates.

An alternative way to proceed, therefore, is to estimate a three variable system comprising the markup of price on average wages, $p-w$, productivity, $y-h$, and inflation, conditioned on the business cycle. This specification substitutes the 'proxy' productivity terms of y_t, q_t and the time trend with actual productivity, $y-h$, and the markup on marginal costs with the markup of price on average wages, $p-w$.

The estimated long-run relationship is of the form:

$$mu \equiv (p-w)+(y-h)=\omega-\lambda\,\Delta p \tag{3}$$

where mu is the markup of price on unit labour costs net of the cost of inflation, ω is the 'gross' markup, Δp is price inflation, and λ is a positive parameter and termed the inflation cost coefficient.[18] Short-run deviations from the long-run relationship are due to shocks and the business cycle. Estimating (3) allows us to identify separately the impact of the business cycle on the markup on average wages, productivity and inflation. Imposing linear homogeneity on the model also allows us to estimate the long-run relationship between inflation and the markup on unit labour costs.[19]

In estimating the long-run relationship the usual tension between precision and stability is encountered. More observations lead to greater precision in estimating the long-run relationship but simultaneously increase the likelihood of breaks in the series due to non-modelled influences such as changes in competition and changes in data reporting regimes. The initial estimates indicate that there is a level shift in the relationship between inflation and the markup in June 1968.[20] The estimation proceeds with the inclusion of a shift dummy for the period December 1952 to March 1968. Four spike dummies are also included to capture the somewhat erratic short-run dynamics of the price, wage and productivity data.

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18 The long-run relationship is derived and considered in some detail in Banerjee *et al.* (2001). The form of the long-run relationship is a generalisation of de Brouwer and Ericsson (1998).

19 The term $p-w+y-h$ is equivalent to the markup on unit labour costs and the inverse of labour's income share. Linear homogeneity implies a unit coefficient on productivity and implies, all else equal, that a change in costs is fully reflected in higher prices in the long run leaving the markup unchanged. In this model, 'all else equal' includes no change in the rate of inflation in the long run.

20 A similar break in the long-run relationship is evident in the United States estimates reported in Banerjee and Russell (2001a, 2001b).

3.3.1 The System Results

Table 3 reports the trace statistics of the estimated system showing acceptance of the hypothesis of one cointegrating vector.[21] The normalised cointegrating vector is reported in Table 4 with linear homogeneity imposed. We see that the inflation cost coefficient, λ, is significant and positive indicating a negative long-run relationship between the markup on unit labour costs and inflation. Importantly, the markup and productivity are found not to cointegrate if inflation is not included in the cointegrating space and this is shown in the lower panel of Table 3.

The dynamics of the estimated system are reported in Table 5. We see that the change in productivity is counter-cyclical at the 5 per cent level of significance. Furthermore, as the change in the markup shows no significant cyclical behaviour this implies that the change in the markup on unit costs is counter-cyclical. Consistent with the standard view that inflation accelerates in booms and decelerates in recessions, the business cycle has a significant positive impact on the change in inflation.

[21] Before estimating the I(1) system the series were tested for the presence of unit roots using PT and DF-GLS univariate tests from Elliot *et al.* (1996). It is found that the markup, productivity and inflation are best characterised as I(1) processes and the business cycle is clearly I(0). The system analysis that follows supports the conclusions of the univariate tests.

Table 3: Testing for the Number of Cointegrating Vectors

Three Variable System: Markup on Average Wage, Productivity and Inflation

Null Hypothesis H_0: r =	Eigenvalues	Estimated Trace Statistic $Q(r)$
0	0.1633	30.58 {26.70}
1	0.0306	4.55 {13.31}
2	0.0001	0.01 {2.71}

Two Variable System: Markup on Average Wage and Productivity

Null Hypothesis H_0: r =	Eigenvalues	Estimated Trace Statistic $Q(r)$
0	0.0687	6.38 {13.31}
1	0.0051	0.34 {2.71}

Notes: Statistics are computed with 4 lags of the core variables. The sample is December 1952 to March 1989 with 146 observations and 127 degrees of freedom.

The shaded cell indicates acceptance at the 10 per cent level of significance. Critical values shown in curly brackets { } are from Table 15.3 of Johansen (1995).

Table 4: Normalised Cointegrating Vector

	$(p-w)_t$	$(y-h)_t$	Δp_t
Unrestricted	1	1.056	2.873 (0.655)
Linear Homogeneity Imposed	1	1	3.006 (0.653)

Notes: Likelihood ratio tests (a) linear homogeneity is accepted $\chi_1^2 = 2.02$, p-value = 0.15; (b) test of coefficient on inflation is zero is rejected, $\chi_1^2 = 19.25$, p-value = 0.00. Standard error reported in brackets.

4. INTERPRETING THE RESULTS AND CONCLUSION

The general empirical results of R-W *inter alia* that the markup on marginal costs is counter-cyclical is reinterpreted as a short-run relationship between changes in the markup and the business cycle as shown in Tables 2 and 5. The system results in Table 5 identify the source of the cyclical variation in the markup on unit costs as due to the counter-cyclical behaviour of the change in productivity and not due to any cyclical behaviour of the change in the markup of price on average wages.

Furthermore, the finding of a long-run negative relationship between inflation and the markup in Banerjee *et al.* (2001), Banerjee and Russell (2001a, 2001b), and Banerjee *et al.* (2002) is re-established and it is shown that the markup and productivity are not cointegrated unless inflation is included in the long-run relationship. The long-run relationship between inflation and the markup is likely to be important in an economic sense with a 1 percentage point increase in annual inflation leading to a 0.75 of a percentage point fall in the markup in the long-run. This estimate is numerically close to the value of 0.5 in Banerjee and Russell (2001a) using quarterly aggregate US national accounts data for the period December 1961 to June 1997 and 0.65 in Banerjee and Russell (2001b) using annual US industrial sector data for the period 1947 to 1997. This robustness of the estimate of the inflation cost coefficient to diverse data sources and time-periods serves to link our analysis to existing traditional studies of markup behaviour.

The paper suggests that business cycle analyses of the markup should carefully take into account the integration properties of the data. A further implication is that earlier studies of the negative relationship between inflation and the markup may be due to the long-run link between inflation and the markup.

Table 5: Dynamics of the I(1) System
Markup on Average Wage, Productivity and Inflation
December 1952 – March 1989

Dependent Variable ⇒	Lag ⇓	*Markup* $\Delta(p-w)$	*Productivity* $\Delta(y-h)$	*Inflation* $\Delta^2 p$
<u>*Loading Matrix*</u>				
Error Correction Term	1	0.050 (1.9)	- 0.076 (- 2.0)	- 0.052 (- 2.3)
<u>*Short-run Matrices*</u>				
Δ Log P/W	1	0.076 (0.7)	- 0.054 (- 0.4)	- 0.001 (- 0.0)
Δ Log P/W	2	- 0.178 (- 1.7)	0.137 (0.9)	- 0.206 (- 2.3)
Δ Log P/W	3	- 0.075 (- 0.8)	- 0.015 (- 0.1)	- 0.023 (- 0.3)
Δ Log Productivity	1	0.154 (2.5)	- 0.135 (- 1.6)	0.045 (0.9)
Δ Log Productivity	2	0.005 (0.1)	0.133 (1.6)	0.057 (1.1)
Δ Log Productivity	3	0.118 (2.0)	- 0.256 (- 3.2)	0.125 (2.5)
Δ Inflation	1	- 0.151 (- 1.3)	0.148 (0.9)	- 0.564 (- 5.6)
Δ Inflation	2	- 0.055 (- 0.5)	0.157 (1.0)	- 0.208 (- 2.1)
Δ Inflation	3	- 0.044 (- 0.5)	0.057 (0.5)	- 0.095 (- 1.3)
<u>*Predetermined Variables*</u>				
Constant		- 0.026 (- 2.1)	0.037 (2.2)	0.022 (2.1)
Step Dummy		- 0.007 (- 4.3)	0.005 (2.4)	0.002 (1.3)
Business Cycle	1	0.012 (0.8)	- 0.064 (- 3.2)	0.045 (3.6)

Notes: Reported in brackets are *t*-statistics. $ECM_t \equiv (p-w)_t + (y-h)_t + 3.006\Delta p_t$

Step dummies: December 1952 to March 1968. Spike dummies: December 1953, March 1965, September 1970, and September 1974.

<u>*System Diagnostics for the Restricted Model*</u>

Tests for Serial Correlation

Ljung-Box (36) χ^2 (294) = 318.94, p-value = 0.15

LM(1) χ^2 (9) = 9.54, p-value = 0.39

LM(4) χ^2 (9) = 11.13, p-value = 0.27

Test for Normality: Doornik-Hansen Test for normality: χ^2 (6) = 4.94, p-value = 0.55

5. DATA APPENDIX

The data are from Rotemberg and Woodford (1991) where further details can be found. The data are for the United States, for the period June 1952 to December 1988, and seasonally adjusted. The variables are in natural logarithms.

Sources and Details of the Data

Variable	Source[(a)]	Details[(b)]
Prices	BEA	Private sector gross national product (GNP) implicit price deflator. Measured as the ratio of current price to constant price value added GNP for the private sector.
Wages	BEA BLS	Private sector average wage rate. Measured by dividing total labour compensation divided by hours of non-agricultural employment for the private sector.
Output	BEA	Private sector constant price GNP.
Employment	BLS	Hours of non-agricultural private sector hours of employment. Measure is total hours in non-agricultural payrolls less hours employed by the government.
Business cycle	BLS	The business cycle is the residuals of employment regressed on a constant and trend.
Tobin's q		The data comes directly from Blanchard, Rhee and Summers (1990).
Productivity		Average productivity. Measured as the logarithm of output less the logarithm of employment.

(a) Mnemonics: BEA – National Income and Product Accounts tables published by the Bureau of Economic Analysis. BLS – Establishment survey, Bureau of Labor Statistics.

(b) The private sector is defined as the total for the variable less the contribution of federal, state and local governments.

6. APPENDIX – BOOTSTRAP DETAILS

$Data\ Generation\ \mathrm{Pr}ocess\ 1$

$\widetilde{\mu}_t = -0.72 - 0.002t + u_t$

$u_t = \rho\, u_{t-1} + \varepsilon_t$

$\varepsilon_t \sim N(0, 1)$

$u_0 \sim N(0, (1-\rho^2)^{-1})$

$Model\ A\ (static\ regression)$

$\widetilde{\mu}_t = c + \gamma t + \beta_1 y_t + \beta_2 q_t + \omega_t$

$100{,}000\ simulated\ series\ for\ y_t\ from\ DGP\ 1\ above$

$Actual\ y_t\ and\ q_t\ series\ used$

$T = 149$

$Model\ B\ (Dynamic\ OLS\ or\ Cochrane-Orcut\ (C-O))$

$\widetilde{\mu}_t = c + \gamma t + \beta_1 y_t + \beta_2 q_t + u_t$

$u_t = \rho\, u_{t-1} + \varepsilon_t$

$100{,}000\ simulated\ series\ for\ y_t\ from\ DGP\ 1\ above$

$Actual\ y_t\ and\ q_t\ series\ used$

$T = 149$

Table A1: Rejection Frequencies of *t*-statistics at Normal 5% critical value for 100,000 bootstrapped regressions

	$t_{c=0}$	$t_{\gamma=0}$	$t_{\beta_1} = 0$	$t_{\beta_2} = 0$	$t_{\rho=0}$
$\rho = 0$					
Static	0.0549	0.0656	0.0523	0.0524	
C-O	0.0644	0.0765	0.0620	0.0612	0.0581
$\rho = 0.50$					
Static	0.2512	0.2611	0.2487	0.2597	
C-O	0.1205	0.1276	0.1181	0.1260	0.9997
$\rho = 0.80$					
Static	0.4841	0.4885	0.4839	0.5171	
C-O	0.3003	0.3058	0.2985	0.3408	1.0000
$\rho = 0.934$					
Static	0.6677	0.6640	0.6680	0.6925	
C-O	0.6051	0.6007	0.6050	0.6417	1.0000

$Data\ Generation\ \Pr ocess\ 2$

$\widetilde{\mu}_t = -0.72 - 0.42 y_t + 0.035 q_t + u_t$

$u_t = \rho\, u_{t-1} + \varepsilon_t$

$\varepsilon_t \sim N(0,1)$

$u_0 \sim N(0,(1-\rho^2)^{-1})$

$Model\ A\ (static\ regression)$

$\widetilde{\mu}_t = c + \gamma t + \beta_1 y_t + \beta_2 q_t + \omega_t$

$100{,}000\ simulated\ series\ for\ y_t\ from\ DGP\ 2\ above$

$Actual\ y_t\ and\ q_t\ series\ used$

$T = 149$

$Model\ B\ (Dynamic\ OLS\ or\ Cochrane-Orcutt\ Estimation)$

$\widetilde{\mu}_t = c + \gamma t + \beta_1 y_t + \beta_2 q_t + u_t$

$u_t = \rho\, u_{t-1} + \varepsilon_t$

$100{,}000\ simulated\ series\ for\ y_t\ from\ DGP\ 2\ above$

$Actual\ y_t\ and\ q_t\ series\ used$

$T = 149$

Table A2(i): Rejection Frequencies of *t*-statistics (from zero) at Normal 5% critical value for 100,000 bootstrapped regressions

	$t_{c=0}$	$t_{\gamma=0}$	$t_{\beta_1} = 0$	$t_{\beta_2} = 0$	$t_{\rho=0}$
ρ=0					
Static	0.0549	0.0523	0.0910	0.0667	
C-O	0.0644	0.0620	0.1029	0.0778	0.0581
ρ=0.50					
Static	0.2166	0.2201	0.2078	0.2722	
C-O	0.1206	0.1233	0.1205	0.1533	0.9996
ρ=0.80					
Static	0.4242	0.4532	0.4033	0.6330	
C-O	0.3148	0.3541	0.2982	0.5759	0.9999
ρ=0.934					
Static	0.6018	0.5979	0.6048	0.7467	
C-O	0.5424	0.5417	0.5449	0.7155	0.9999

Table A2(ii): Rejection Frequencies of *t*-statistics (from true values) at Normal 5% critical value for 100,000 bootstrapped regressions

	$t_{c=-0.72}$	$t_{\gamma=0}$	$t_{\beta_1=-0.42} = 0$	$t_{\beta_2=0.035}$	$t_{\rho=\rho_o}$
ρ=0					
Static	0.0524	0.0523	0.0523	0.0524	
C-O	0.0619	0.0620	0.0620	0.0612	0.0581
ρ=0.50					
Static	0.2103	0.2207	0.2106	0.3198	
C-O	0.1160	0.1223	0.1163	0.1881	0.0433
ρ=0.80					
Static	0.4193	0.4532	0.4210	0.6584	
C-O	0.3101	0.3542	0.3154	0.6061	0.2734
ρ=0.934					
Static	0.6029	0.5979	0.6012	0.7543	
C-O	0.5431	0.5417	0.5416	0.7253	0.0675

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