Inflation and Measures of the Markup

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Abstract

Theoretical models of the relationship between inflation and markup focus on the markup of price on marginal costs in contrast with empirical models that typically concentrate on the markup on unit costs. Using nearly 50 years of quarterly United States data we identify a strong negative long-run relationship between inflation and both measures of the markup. We derive the theoretical link between the two measures and empirically verify our prediction that the two inflation cost coefficients should not differ from each other significantly. We conclude that the long-run trade-off between inflation and markup does not depend on the particular measure of the markup used.

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1. Introduction

Empirical evidence in favour of a negative relationship between inflation and the markup has grown in recent years, including papers by Richards and Stevens (1987), Bénabou (1992), Franz and Gordon (1993), Cockerell and Russell (1995), de Brouwer and Ericsson (1998), Simon (1999) and Batini, Jackson and Nickell (2000). All the estimation undertaken in these papers has assumed that inflation and the markup are stationary variables.

In contrast, Banerjee, Cockerell and Russell (2001), Banerjee and Russell (2000, 2001a, 2001b) and Banerjee, Mizen and Russell (2002) have argued that these variables should be treated as integrated and have identified a negative *long-run* relationship between inflation and the markup.[1] In addition, the existence of a short-run relationship between the stationary components in both series, possibly involving the business cycle, has also been demonstrated.

Persistent shifts in the rate of inflation appear to be associated with persistent changes in the markup and *vice versa*. Econometric estimates of the long-run relationship indicate that shocks to either inflation or the markup, and their subsequent impact on each other, persist for many years. Furthermore, a negative long-run relationship between inflation and the markup implies that there is a positive relationship between the real wage and inflation for a given level of productivity. Therefore, if we believe that, *ceteris paribus*, employment, investment and the capital stock are related to the real wage, we must conclude that inflation has lasting effects on employment and investment and is likely to be important for the conduct of economic policy.

Standard theoretical explanations of the negative relationship between inflation and the markup in the literature focus on the impact of inflation on the markup of price on marginal costs.[2] However, since marginal costs are difficult to quantify, most empirical work estimates the relationship using the markup on unit or average costs since these can be measured directly.

The marginal cost and unit cost markups diverge from each other over the business cycle, suggesting the existence of relationships of different magnitudes with inflation in the long run.[3] Both relationships are of interest in their own right, depending upon the economic phenomenon we wish to explore. If we are concerned about the impact of inflation on fixed capital formation, the relationship with the unit cost markup is more relevant, given imperfect capital markets and the predilection of firms to fund investment through retained earnings.[4] Alternatively, if our interest were more in the employment consequences of persistently higher inflation, the impact on the marginal cost markup would be more informative. This choice then becomes a matter of primary importance if the strength of the estimated relationship between inflation and the markup is found to be dependent on the way the markup is measured.

Two broad issues are investigated in this paper. First, can we continue to identify a long-run relationship between the markup and inflation when the markup is measured as the markup of price on marginal costs? To answer this question, we construct an index of the marginal cost markup in terms of the unit cost markup in the tradition of Hall (1988) and extended by Rotemberg and Woodford (1991). Two cointegrated systems are next estimated with quarterly United States data for the period June 1953 to March 2000, using the two alternative measures of the markup. The negative long-run relationship between inflation and the unit cost markup identified in earlier work is re-established. We also identify a negative long-run relationship between inflation and the marginal cost markup. The results are reported in Section 4.

The second issue we investigate is whether or not the estimate of the long-run relationship between inflation and the marginal cost markup differs from that obtained from using the unit cost measure of the markup. We show that the marginal cost markup may be thought of as being equivalent to the sum of the unit cost markup and an adjustment related to the business cycle. This adjustment reflects differences over the business cycle between marginal productivity and average productivity as well as differences between the marginal and average markup of price on wages. If the adjustment is a stationary variable, as one would expect given its relation to the business cycle, the two estimates of the long-run relationship should not differ substantially from each other. The converse would be true if the adjustment were non-stationary.

To presage our results, the system using the unit cost measure of the markup estimates that an increase of 1 percentage point in the rate of inflation leads to a decrease of 0.67 of a percentage point in the unit cost markup in the long run. This contrasts with the estimate obtained from using the marginal cost measure of the markup where a 1 percentage point increase in inflation leads to a 1.65 percentage points decrease in the marginal cost markup.

This difference between the estimates may be attributed to the improved dynamics introduced by the inclusion of the effects of the business cycle in the marginal cost markup although formal statistical testing accepts the null hypothesis of the equality of the two estimates. Our baseline results are derived for a so-called ‘high’ value for the steady state markup of 1.6, consistent with Rotemberg and Woodford (1991). We show that assuming lower, and some might argue more realistic, values for the steady state markup leads to more congruent estimates between the long-run inflation-markup relationship using the two different measures of the markup.

2. Brief Explanations of the Inflation-Markup Relationship

The idea that a nominal variable such as inflation should have a persistent effect on a real variable like the markup derives from considering price-setting imperfectly competitive models. Within this framework, there are three broad approaches to considering the impact of inflation on the markup. The first is in the 'menu' cost tradition of Mankiw (1985) and Parkin (1986), which assumes that the cost and demand functions are not directly influenced by the rate of inflation. Papers by Rotemberg (1983), Kuran (1986), Naish (1986), Danziger (1988), Konieczny (1990), and Bénabou and Konieczny (1994) model the price behaviour of imperfectly competitive firms in the presence of small 'menu' costs and conclude that inflation has a negative impact on the average markup while the profit maximising markup is unaffected. However, these approaches to modelling pricing behaviour are unlikely to explain an empirical long-run relationship between inflation and the markup since the features, such as the stable pricing rule, that underpin the result are likely to disappear in the steady state or long run.[5]

The second broad approach, either explicitly or implicitly, argues that the demand functions are influenced by inflation. Bénabou (1988, 1992) and Diamond (1993) argue that higher inflation leads to greater search in customer markets, which increases competition and reduces the markup. These explanations will generate a steady state relationship between inflation and the markup under the assumption that higher inflation permanently increases search and competition. If it does not, then the relationship will only persist in the short run.

The third approach is an optimising, price-setting 'behavioural' model which argues explicitly that the inflation-markup relationship will exist in the steady state. Russell, Evans, and Preston (2002) argue that firms are uncertain about the profit maximising markup and believe that they face an asymmetric loss function. The asymmetry reflects the belief of firms that setting too 'high' a markup relative to its profit maximising value, costs the firm more in lost profits than if they set too 'low' a markup. These 'costs' are due to the presence of a kinked demand curve or increasing returns to scale. Consequently, in an uncertain environment, firms set a markup below the profit maximising level. Furthermore, if uncertainty increases with inflation then firms set an even lower markup relative to their profit maximising levels.

The explanation offered in Russell *et al.* (2002) suggests that if the uncertainty persists in the steady state, so will the negative relationship.[6] In a perfectly competitive price-taking world, firms are able to predict the profit maximising markup in the steady state, so that they can set the profit maximising level of output with certainty. However, with price-setting firms it is unlikely that the uncertainty will disappear if it is caused by the firms' inability to coordinate the adjustment in prices.[7] Furthermore, it is likely that uncertainty will increase with inflation as the frequency and / or size of the price changes must increase.

3. Measuring the Marginal Cost Markup

Hall (1988) builds on work by Solow (1957) to provide a model for estimating a constant marginal cost markup. Rotemberg and Woodford (1991) extend the model to allow for a time varying marginal cost markup. Under certain assumptions concerning the production function and market structure, Rotemberg and Woodford provide an expression that can be interpreted as the relationship between the variation in the markup of the price on the average wage and the marginal cost markup over the business cycle. This expression can then be used to provide a measure of the adjustment that is necessary for the measured unit cost markup to provide an estimate of the marginal cost markup.

Rotemberg and Woodford (1991) assume an imperfectly competitive goods market, increasing returns to scale and a production function, F :

$$y_t^i = F\left[K_t^i, z_t\left(H_t^i - \hat{H}_t\right)\right] \tag{1}$$

where y_t^i, K_t^i, and H_t^i represent output, capital input and labour input at time t for firm i. Technology at time t is represented by z_t allowing firms to be more productive during some periods and $\hat{H}_t$ is the fixed cost of overhead labour. The latter introduces decreasing average costs and provides for price greater than marginal cost in the model. With imperfectly competitive goods markets and competitive labour and capital markets the marginal cost markup, $MCMU_t$, is:

$$MCMU_t = \frac{F_H\left[K_t, z_t\left(H_t - \hat{H}_t\right)\right]}{RW_t} \tag{2}$$

where F_H is the marginal product of labour, RW_t is the real wage defined as W_t/P_t and W_t and P_t are the average wage rate and average price respectively. The marginal cost markup, $MCMU_t$, cannot be measured using (2) as z_t and $\hat{H}_t$ cannot be measured directly. Rotemberg and Woodford overcome this problem by considering a log linear approximation of the production function around the steady state growth path where the firm's labour input,

H_t, and overhead labour, $\hat{H}_t$, grow at the same rate. They provide the following expression for log deviations in the marginal cost markup from its steady state value, $\bar{\mu}_t$:[8]

$$\bar{\mu}_t = \frac{e-\mu^* S_K}{e-e\mu^* S_K}\bar{y}_t + \frac{(1-e)\mu^* S_K}{e-e\mu^* S_K}\bar{k}_t - \frac{\mu^* S_H}{1-\mu^* S_K}\bar{h}_t + \overline{(p-w)}_t \tag{3}$$

where μ^* is the steady state value of the marginal cost markup, e represents the elasticity of substitution between the two factor inputs (capital and labour) and S_K and S_H are the factor shares of capital and labour. Lower case variables are in natural logarithms and the 'bar' on a variable indicates the log deviation from the trend value of the variable.[9] Equation (3) represents the direct and indirect effects of the business cycle on the markup of price on marginal costs. The direct effect is through the markup of price on average wages, $\overline{p-w}$. The remaining influences, through $\bar{y}$, $\bar{k}$ and $\bar{h}$, represent the indirect effects of the business cycle on the marginal cost markup through the impact on marginal productivity.

At least since Dunlop (1938) and Tarshis (1939) it has been generally acknowledged that the real wage is pro-cyclical and the markup of the price on the average wage, $p-w$, which is the inverse of the real wage, is therefore counter-cyclical.[10] Explanations of counter-cyclical markups usually focus on why firms do not increase prices by enough to match the increasing nominal marginal costs when firms introduce less productive inputs into the production process as the economy expands. Alternative explanations focus on the increased wage and cost pressures with higher demand.[11] In summary, given the counter-cyclical nature of the markup of price on wages, $p-w$, (3) implies that, all else equal, the marginal cost markup will also be counter-cyclical. Furthermore, the indirect effects in (3) may cause the marginal cost markup to be either more or less counter-cyclical than the markup of prices on average wages.

Three points should be noted concerning (3). First, the expression does not identify the numerical value of the marginal cost markup but identifies the variation in the marginal cost markup in terms of the variation in output, capital stock, labour input and the markup of average prices on average wages. In the system estimation that follows, this is not a problem as only an index number of the marginal cost markup is required. Second, we can derive an index of the marginal cost markup in terms of the unit cost markup if we assume in (3) above, that productivity is actual productivity instead of trend productivity. In this case the marginal cost markup, $mcmu_t$, can be thought of as an 'adjustment' to the unit cost markup, $(p-w)_t + (y-h)_t$, such that:

$$\begin{aligned} mcmu_t &= \bar{a}_t + (p-w)_t + (y-h)_t \\ &= \bar{a}_t + ucmu_t \end{aligned} \tag{4}$$

where $\bar{a}_t = \frac{e-\mu^* S_K}{e-e\mu^* S_K}\bar{y}_t + \frac{(1-e)\mu^* S_K}{e-e\mu^* S_K}\bar{k}_t - \frac{\mu^* S_H}{1-\mu^* S_K}\bar{h}_t$ is the marginal cost adjustment to the unit cost markup, $ucmu_t$, due to the indirect effects of the business cycle.[12]

Third, equations (3) and (4) suggest that the statistical properties of the marginal cost markup depend on those of the unit cost markup, $(p-w)_t + (y-h)_t$, and the marginal cost adjustment, $\bar{a}_t$. If $\bar{y}$, $\bar{k}$ and $\bar{h}$ are stationary then the adjustment, $\bar{a}_t$, will also be stationary. The statistical properties of the marginal cost markup and the unit cost markup will therefore be the same by virtue of their sharing a common I(1) trend. This in turn implies that the long-run relationship will be the same irrespective of whether the markup is defined on marginal or unit costs. Alternatively, if $\bar{y}$, $\bar{k}$ or $\bar{h}$ are non-stationary then the marginal cost adjustment will also be non-stationary and the long-run relationship between inflation and the marginal cost markup may differ from the relationship with the unit cost markup.

Finally, the Rotemberg and Woodford model highlights both the complexity of the modelling problem and the range of simplifying assumptions that are necessary to arrive at equation (3). A more complicated model would introduce non-linearity, interaction between the parameters and variables, and imperfectly competitive factor markets, but such a model would quickly become intractable. However, the basic result embodied in (3) is likely to be maintained, namely that variations in the marginal cost markup can be thought of as being a function of deviations in the unit cost markup subject to the effects of the business cycle on output, capital and hours of work.

4. The Relationship between the Markup and Inflation

We now turn to the estimation of the long-run cointegrating relationship between inflation and the markup. We proceed by estimating a three variable cointegrating system using standard I(1) techniques developed by Johansen (1988, 1995). The core integrated variables are the markup, productivity and inflation and the estimation is conditioned on a predetermined business cycle variable and spike dummies to capture the sometimes erratic behaviour of the price, wage and productivity data that occurred during the period but especially in the turbulent 1970s. Two systems are estimated, the first with the markup measured on marginal costs and the second measured on unit costs.

The form of the long-run relationship follows Banerjee, Cockerell and Russell (2001) and

Banerjee and Russell (2001a) where further details concerning the modelling of inflation and the markup allowing for non-stationarity in the series can be found. The long-run relationship may be written as:

$$mu + prod = q - \lambda \Delta p \tag{5}$$

where mu is the markup, $prod$ is average productivity measured as $y - h$, q is the 'gross' markup, p is the price level, λ is a positive parameter termed the 'inflation cost coefficient', and Δ represents the change in the variable.[13] If the markup, mu_t, is defined as, $(p - w)_t$, where prices and wages are measured as their average values then $mu + prod$ is the markup of prices on unit costs.[14] Alternatively, if, mu_t, is defined as $(p - w)_t + \overline{a}_t$ where $\overline{a}_t$ is defined as in equation (4) then $mu + prod$ is the markup of prices on marginal costs.[15] Consequently, the inflation cost coefficient, λ, represents the impact, or cost, of inflation in terms of either a lower unit cost or marginal cost markup in the long run, depending on the measure of the markup used.

4.1 The Data

The cointegrated systems are estimated with quarterly United States data for the period June 1953 to March 2000. The markup and inflation data are derived from the National Income and Product Accounts tables published by the Bureau of Economic Analysis. Prices, wages and output are measured on a 'private sector' basis excluding the contribution of federal, state, and local governments. Labour input is measured as non-agricultural private hours of employment from the establishment survey published by the Bureau of Labor Statistics. The price level is the gross domestic product at factor cost implicit price deflator, wages consist of total labour compensation divided by labour input, and output is constant price gross domestic product. Further details concerning the data are provided in the data appendix.

A number of measures of the business cycle present themselves as candidates for the short-run impact on the estimated system. These measures include variables based on the unemployment rate, hours of employment and national accounts measures of constant price output. The unemployment rate appears to be an integrated variable and is not suitable for use as a predetermined variable for the business cycle. National accounts measures of the business cycle on the other hand suffer from 'errors in measurement' problems when used in association with national accounts price data.[16] To avoid these difficulties the business cycle is represented by de-trended natural logarithm of non-agricultural private sector hours of employment.

In summary, the systems are estimated with four lags in the core integrated variables (markup, productivity and inflation) and four lags of the business cycle variable. Lags of the business cycle were excluded on the basis of a '5 per cent' *t* criterion. Spike dummies are included for periods where residuals were greater than 3 standard errors.

4.2 The Marginal Cost 'Adjustment Factor' $\overline{a}_t$

We follow Rotemberg and Woodford (1991) and choose their 'baseline' values of $e = 1$ and $\mu^* = 1.6$. Sensitivity of the results to the choice of steady state markup is considered below. The output and employment components of a_t appear to have broken trends at the time of the first OPEC oil price shock. Therefore, we calculate the *level* of a_t using the levels of output and employment and then de-trend a_t allowing for the possibility of an exogenous break in the series.[17] For our sample, labour's share of income, S_H, is 0.659 with the level of the marginal cost adjustment given by:

$$a_t = y_t - 2.3207\, h_t \tag{5}$$

Using Perron (1997), we find a break in the level and trend of the marginal cost adjustment, a_t, at June 1974.[18] Once the break is accounted for, unit root tests unambiguously indicate the de-trended marginal cost adjustment series, $\overline{a}_t$, is stationary. The de-trended marginal cost adjustment series is shown in Graph 1. Using this series we calculate the marginal cost markup, $\overline{a}_t + (p - w)_t + (y - h)_t$ and this is shown as the thin line in Graph 2. For comparison, the unit cost markup, $(p - w)_t + (y - h)_t$, is shown as the thick line on the same graph.

The integration properties of the data used in the system estimation were investigated using augmented Dickey-Fuller, KPSS (Kwiatowski, Phillips, Schmidt and Shin (1992)) and PT and DF-GLS (Elliot, Rothenberg and Stock (1996)) univariate unit root tests.[19] All three unit root tests indicate that the markup, $(p - w)_t$, the marginal cost markup, $(p - w)_t + \overline{a}_t$, productivity, $(y - h)_t$, and inflation, Δp_t, are best described as integrated I(1) variables while hours of employment is trend stationary and the marginal cost adjustment, $\overline{a}_t$, is stationary. The results from the system analysis and the system unit root tests are consistent with these findings.

4.3 Results from Estimating the Markup and Inflation Systems

The trace statistics for the number of cointegrating vectors in the two systems are reported in the notes to Tables 1 and 2 and show acceptance of the hypothesis of one cointegrating vector. Further evidence for accepting the hypothesis of one cointegrating vector can be found from the companion matrix.[20] The notes to Tables 1 and 2 also report that both systems have 'well behaved' diagnostics, but those of the marginal cost markup system perform considerably better.

The normalised long-run coefficients with linear homogeneity imposed are also reported in Tables 1 and 2 for the two systems. We see that the inflation cost coefficient, λ, is significant and positive for both systems indicating a negative long-run relationship between inflation and the markup irrespective of whether the markup is defined on marginal or unit costs. Note that the estimate of the inflation cost coefficient in the unit cost markup system is 2.526 implying a 1 percentage point increase in annual inflation is associated with approximately 0.63 of a percentage point decline in the markup of price on unit costs in the long run.[21] This estimate is very similar to the annual estimates for the United States reported in Banerjee and Russell (2001a, 2001b) of 0.62 using annual 'private sector' gross domestic product data and 0.46 using aggregate gross domestic product data.[22] In contrast, the estimate of the inflation cost coefficient in the marginal cost markup system is 6.602 and that is around 2 ½ times the estimate from the unit cost markup system.

A formal test of the equality of the two estimates of the inflation cost coefficient can be undertaken by testing for cointegration in the three-variable system given by:

$$(mcmu,\ ucmu,\ \Delta p)' \tag{6}$$

where $mcmu_t$ is the marginal cost markup and $ucmu_t$ is the unit cost markup as defined in (4). Therefore, if and only if, $mcmu_t$ and $ucmu_t$ bear the same relation to inflation (i.e. $mcmu_t - \lambda_{MC}\,\Delta p_t$ and $ucmu_t - \lambda_{UC}\,\Delta p_t$ are both I(0) and $\lambda_{MC} = \lambda_{UC} \neq 0$), the vector $(1, -1, 0)'$ is a cointegrating vector in the system given by (6). Hence, if $\lambda_{MC} \neq \lambda_{UC}$, a cointegrating vector for (6) will be given by $(1, -1, \beta)'$ where $\beta \equiv \lambda_{UC} - \lambda_{MC}$ is non-zero and measures the difference in the magnitudes of the two inflation cost coefficients.

To test for equality of the inflation cost coefficients we therefore estimate the system given by (6), test for cointegration, extract the cointegrating vector given by $(1, b_1, b_2)'$ (after normalising on the first variable) and test the restrictions that $b_1 = -1$ and $b_2 = 0$. This may also be regarded as a systems test for whether the adjustment factor, $\overline{a}_t$, is stationary.[23]

The outcome of this procedure is reported in Table 3 where we find we are able to accept the null of one cointegrating vector between $mcmu$, $ucmu$, and Δp. We are also able to easily accept the restrictions that $b_1 = -1$ and $b_2 = 0$ with a p-value of 0.77. This not only confirms our finding reported earlier of the stationarity of the adjustment factor series, $\overline{a}_t$, but also implies that the differences in magnitude of the inflation cost coefficient are not significant in the long-run relationship and are caused by the short-run dynamics of the processes over the sample period.[24]

We next turn to a sensitivity analysis of our estimates of the long-run coefficients to the choice of the value of the steady state markup. Martins, Scarpetta and Pilat

(1996) argue that the average marginal cost markup (which may be thought of as a proxy for the steady state markup, $\mu*$) may be considerably lower than the 1.6 assumed here. Repeating the system analysis with the marginal cost adjustment, $\overline{a}_t$, calculated by assuming a range of values, between 1.0 and 1.6 for the steady state markup, $\mu*$, we see that the estimates of the inflation cost coefficient move closer to each other. Starting with a value of 1.6 and decrementing in steps of 0.1 to a value of 1.0, we find that the estimate of the inflation cost coefficient decreases steadily from 6.602 to 2.035 so that our conclusions concerning the similarity of the estimated inflation cost coefficients in the two systems are unaltered.[25]

The finding of long-run inflation-markup relationships is of course conditioned on inflation and the markup behaving like integrated processes for the time-period under consideration. In episodes where inflation is stationary, it would not be possible to investigate whether there was a long-run relationship (in the Engle and Granger sense) between inflation and the markup even if such a relationship existed during times when inflation behaved as an integrated variable.

5. Conclusion

This paper considers the question of whether or not it matters if the markup is measured on marginal or unit costs when estimating the long-run relationship between inflation and the markup. To examine this question we began by calculating the marginal cost markup following Rotemberg and Woodford (1991) and assume, as they do, a 'steady state' markup of 1.6. Using United States quarterly data we estimate two inflation markup systems with the markup measured on unit costs in one and on marginal costs in the other. Our results reestablish the long-run relationship between inflation and the unit cost markup identified in earlier work but also show the existence of a long-run relationship between inflation and the marginal cost markup.

We find that the estimate of the 'strength' of the long-run relationship is larger numerically in the marginal cost markup system but formal testing of the equality of the inflation cost coefficients derived from the two alternative measures accepts the hypothesis that they are not significantly different from each other. In any case, the numerical difference between the estimates is reduced when a lower steady state markup is chosen when calculating the marginal cost markup.

Finally, do we care how we measure the markup and does it matter? In answer to the first part of the question, the answer in our view is 'no'. This is because we conclude that the long-run estimates derived from either measure are essentially the same. The answer to the second part is also straightforward, since the estimates suggest that a 4 percentage point increase in the general rate of inflation (as occurred in the 1970s in the United States) is associated with between a 3 percent increase in the real wage relative to the level of productivity (from the unit cost markup estimate) and an 8 percent increase (from the marginal cost markup estimate). Movements of this magnitude would be expected to have a significant impact on functioning of the United States economy.

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6. Data Appendix

United States data are seasonally adjusted for the period June 1952 to March 2000. Natural logarithms are taken of all variables before estimation proceeds.

Sources and Details of the Data

Variable	Source[(a)]	Details
Price	BEA	Private sector gross domestic product (GDP) implicit price deflator at factor cost. Measured as current price GDP (value added) less value added of federal, state and local government less indirect taxes plus subsidies divided by constant price GDP
Wages	BEA BLS	Private sector average wage rate. Measured by dividing total labour compensation less government labour compensation divided by labour input.
Output	BEA	Private sector constant price GDP. Measured as chained 1996 dollars of value added GDP less value added of federal, state and local government.
Labour input	BLS	Hours of non-agricultural private hours of employment. Measured from June 1953 to March 1964 by 'private hours' of employment from Rotemberg and Woodford (1991). This measure is total hours in non-agricultural payrolls less hours employed by the government. From March 1964 labour input is quarterly average of monthly data measured as 'total private index of aggregate weekly hours' (EES00500040) taken from Table B1 'Employees on non-farm payrolls by industry'. The two series are very similar from March 1964 to the end of the Rotemberg and Woodford data in March 1989. The two series are 'back-spliced' in March 1964.
Business cycle	BLS	Measured as de-trended natural logarithm of labour input. No break in the trend or level of the series was identified using the Perron (1997) unit root test. The business cycle is the residuals of the logarithm of labour input regressed on a constant and trend.
Marginal Cost Adjustment $\overline{a}_t$		The level of the marginal cost adjustment (MCA) was calculated as $a_t = y_t - 2.3207\ h_t$. Perron (1997) unit root test identifies a shift in the constant and break in trend in June 1974. MCA de-trended by regressing a_t on a constant, dummy for June 1974 to March 2000, trend, and a short-trend June 1974 to March 2000.

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

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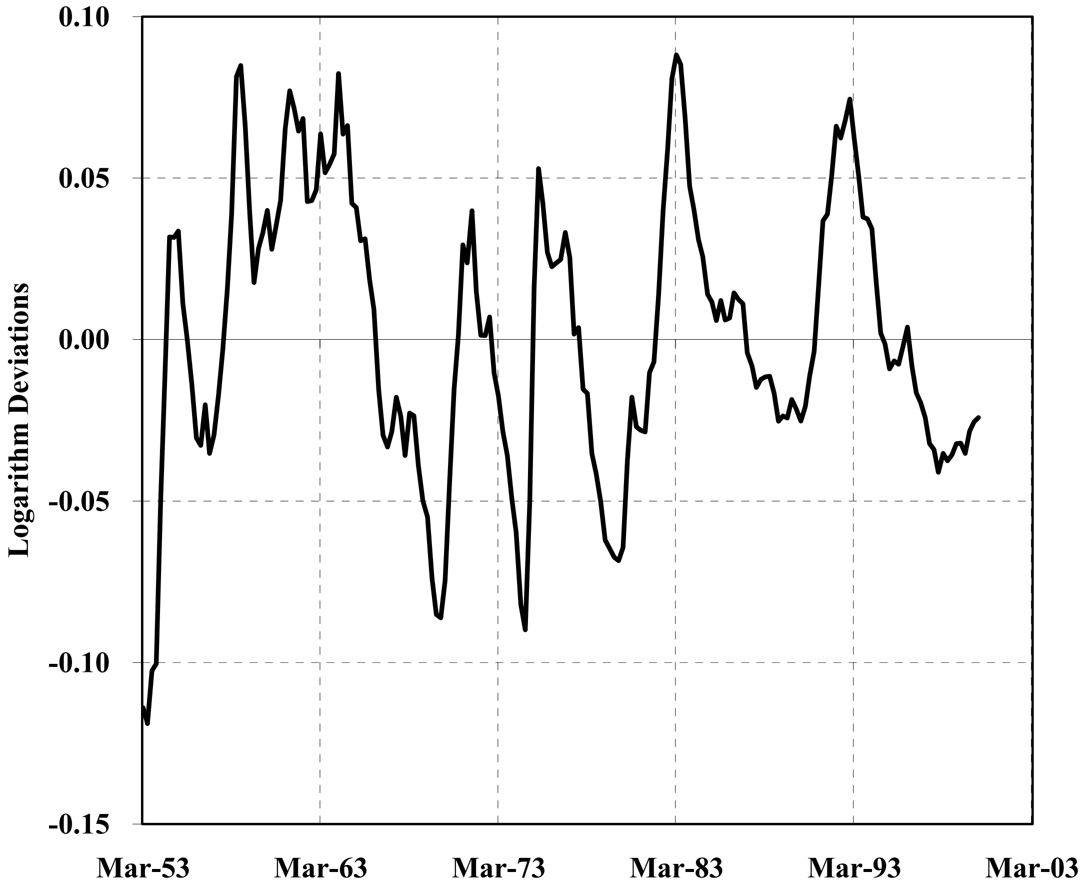
Graph 1: Detrended Marginal Cost Adjustment Factor $\overline{a}_t$



Graph 2: Unit Cost and Marginal Cost Markups

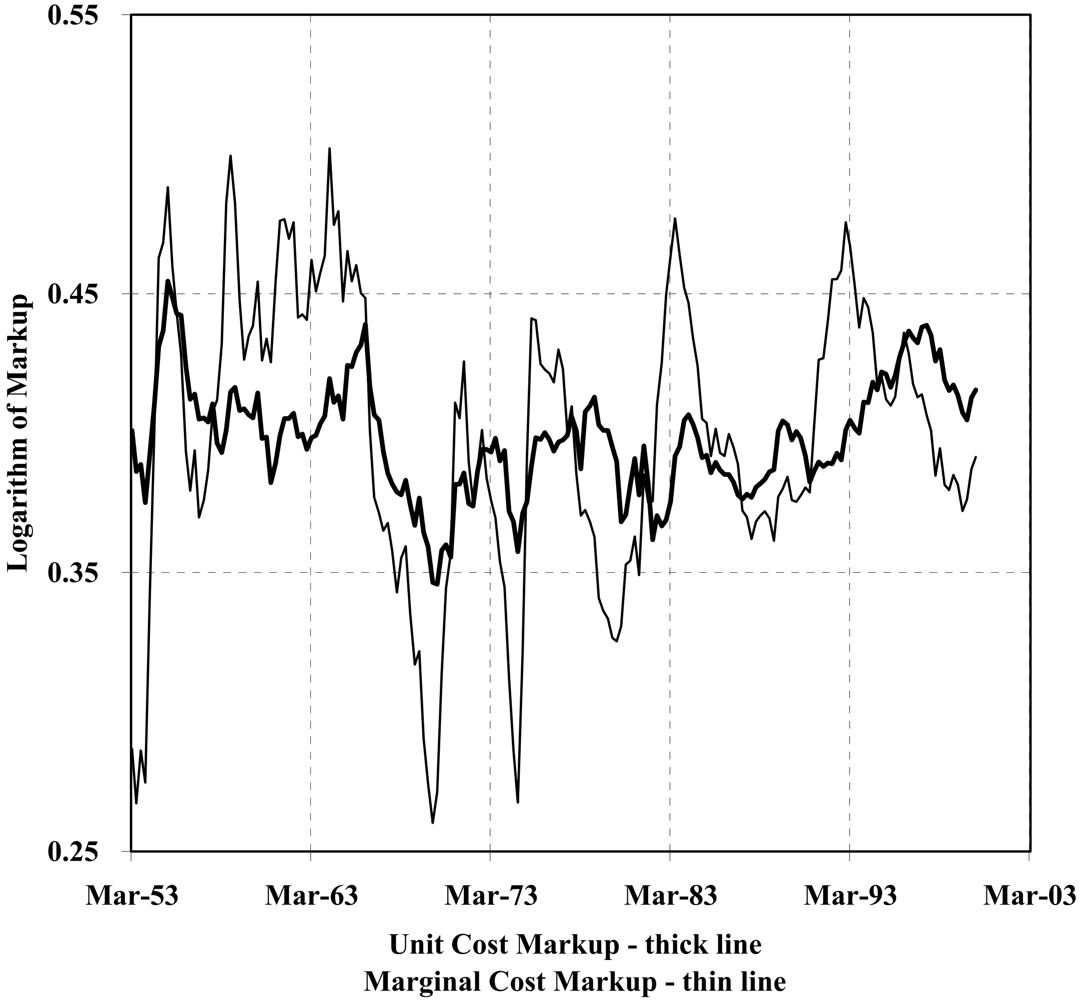


Table 1: Normalised Cointegrating Vectors

'Unit Cost Markup' System

	$p-w$	$y-h$	Δp
Unrestricted	1 (0.019)	1.019 (0.019)	2.677 (0.735)
Linear Homogeneity Imposed	1	1	2.526 (0.700)
Adjustment Coefficients	- 0.006 [- 0.3]	- 0.121 [- 3.9]	- 0.006 [- 0.4]

Standard errors reported as ( ), *t*-statistics reported as [ ]. The adjustment coefficients are the values with which the long-run enters each equation of the system with linear homogeneity imposed. This implies the long-run relationship, or dynamic error correction term, is: $ECM_t \equiv (p-w)_t + (y-h)_t + 2.526\Delta p_t$.

Likelihood ratio tests (a) linear homogeneity is accepted $\chi_1^2 = 0.39$, p-value = 0.53; (b) test of coefficient on inflation is zero is rejected, $\chi_1^2 = 6.2$, p-value = 0.01, and (c) exclusion of a trend in the cointegrating space $\chi_1^2 = 10.08$, p-value = 0.01.

Predetermined Variables: Spike dummies – June and September 1954, March 1971, and June 1975.

Testing for the Number of Cointegrating Vectors

Estimated trace statistic for the null hypothesis $H_0 : r = 0$ is 31.06 {26.70}, $H_0 : r = 1$ is 13.04 {13.31}, and $H_0 : r = 2$ is 1.33 {2.71}. Numbers in { } are the relevant 90 per cent critical values from Table 15.3 of Johansen (1995). Statistics computed with 4 lags of the core variables. The sample is June 1953 to March 2000 and has 188 observations with 171 degrees of freedom.

System Diagnostics for the Model with Linear Homogeneity Imposed

(a) *Tests for Serial Correlation*

Ljung-Box (47) χ^2 (393) = 442.59, p-value = 0.04

LM(1) χ^2 (9) = 8.49, p-value = 0.49

LM(4) χ^2 (9) = 16.12, p-value = 0.06

(b) *Test for Normality:* Doornik-Hansen Test for normality: χ^2 (6) = 4.31, p-value = 0.63

Table 2: Normalised Cointegrating Vectors

'Marginal Cost Markup' System

	$\bar{a}+(p-w)$	$y-h$	Δp
Unrestricted	1 (0.024)	1.044 (0.023)	6.333 (1.048)
Linear Homogeneity Imposed	1	1	6.602 (0.956)
Adjustment Coefficients	- 0.094 [- 4.9]	- 0.091 [-5.7]	- 0.009 [- 1.3]

Standard errors reported as ( ), *t*-statistics reported as [ ]. The adjustment coefficients are the values with which the long-run enters each equation of the system with linear homogeneity imposed. This implies the long-run relationship, or dynamic error correction term, is: $ECM_t \equiv (p-w)_t + (y-h)_t + 6.333\Delta p_t$.

Likelihood ratio tests (a) linear homogeneity is accepted $\chi_1^2 = 3.30$, p-value = 0.07; (b) test of coefficient on inflation is zero is rejected, $\chi_1^2 = 20.30$, p-value = 0.00, and (c) exclusion of a trend in the cointegrating space is accepted $\chi_1^2 = 2.55$, p-value = 0.11.

Predetermined Variables: Spike dummies – June, September and December 1974 and March 1975. Business cycle variable lagged one period.

Testing for the Number of Cointegrating Vectors

Estimated trace statistic for the null hypothesis $H_0 : r = 0$ is 45.14 {26.70}, $H_0 : r = 1$ is 5.81 {13.31}, and $H_0 : r = 2$ is 0.55 {2.71}. Numbers in { } are the relevant 90 per cent critical values from Table 15.3 of Johansen (1995). Statistics computed with 4 lags of the core variables. The sample is June 1953 to March 2000 and has 188 observations with 170 degrees of freedom.

System Diagnostics for the Model with Linear Homogeneity Imposed

(a) *Tests for Serial Correlation*

Ljung-Box (47) χ^2 (393) = 418.39, p-value = 0.18

LM(1) χ^2 (9) = 13.29, p-value = 0.15

LM(4) χ^2 (9) = 14.08, p-value = 0.12

(b) *Test for Normality:* Doornik-Hansen Test for normality: χ^2 (6) = 3.85, p-value = 0.70

Table 3: Testing for the Equality of the Inflation Cost Coefficients

Normalised Cointegrating Vectors

	mcmu	*ucmu*	Δp
Unrestricted	1 (0.361)	- 0.680 (0.154)	- 0.059 (1.094)
Restricted*	1	- 1	0

Standard errors reported as ( ).

* Likelihood ratio test of the restriction $\chi_1^2 = 0.53$, p-value = 0.77.

Predetermined Variables: Spike dummies – June 1954, September 1954, March 1971, June 1975.

Testing for the Number of Cointegrating Vectors

Estimated trace statistic for the null hypothesis $H_0 : r = 0$ is 32.00 {26.70}, $H_0 : r = 1$ is 13.26 {13.31}, and $H_0 : r = 2$ is 2.75 {2.71}. Numbers in { } are the relevant 90 per cent critical values from Table 15.3 of Johansen (1995). Statistics computed with 4 lags of the core variables. The sample is June 1953 to March 2000 and has 188 observations with 171 degrees of freedom.

Footnotes

[1] See Banerjee, Dolado, Galbraith and Hendry (1993) and Johansen (1995) for a description of the long run in the sense of Engle and Granger (1987).

[2] For example see Rotemberg (1983), Kuran (1986), Naish (1986), Danziger (1988) and Konieczny (1990), Bénabou and Konieczny (1994) and Bénabou (1992).

[3] Reasons for the divergence between the two measures of the markup include convex adjustment costs, firms insuring workers against fluctuations in their real wage, the introduction of lower quality workers and capital into the production process as output expands, and overtime hours being more expensive than straight-time hours. Rotemberg and Woodford (1991, 1999) survey these reasons at length.

[4] See for example Myers and Majluf (1984) for their theory of the pecking order of finance.

[5] Steady state defined as all nominal variables increasing at the same constant rate.

[6] Russell (1998) and Chen and Russell (2002) model price-setting firms when information is missing and also argue that the negative relationship between inflation and the markup will persist in the steady state.

[7] Eckstein and Fromm (1968), Chatterjee and Cooper (1989), Blinder (1990) and Ball and Romer (1991) argue price setting firms find it difficult to coordinate price changes.

[8] Equation 3.6 on page 84 of Rotemberg and Woodford (1991). Johri (2001) provides a more straightforward exposition of (3) with $e = 1$.

[9] Bils and Chang (2000) arrive at a similar expression to (3) assuming a CES production function with returns to scale $1+\eta$ and a capital labour rate of substitution of σ, such that

$$\dot{\mu}_t = \frac{1}{\sigma}\left(\frac{S_K}{S_H + S_K}\right)\left(\frac{\dot{h}}{k}\right)_t - \eta\left(\frac{S_K}{S_H + S_K}\right)\dot{k}_t - \eta\left(\frac{S_H}{S_H + S_K}\right)\dot{h}_t$$

and where dots on the variables indicate rate of change (in contrast with deviations from trend as in (3)). This expression differs from (3) in that there is no direct demand effect on the markup. See also Basu (2000).

10 Pigou (1927) argues the cyclical nature of the real wage depends on the particular business cycle examined. More recently Bils 1987, Kydland and Prescott (1988), Barsky and Solon (1986) and Bils and Kahn (1996) provide evidence of counter-cyclical markups.

11 Calvo (1983), Rotemberg (1982), Gali (1994), Phelps and Winter (1970), and Rotemberg and Woodford (1991) provide some explanations for why firms may accept a lower markup with higher output. The macroeconomic models of Layard, Nickell and Jackman (1991), Lucas (1973), Kydland and Prescott (1988), and Blanchard and Kiyotaki (1987) also imply a counter-cyclical markup. Rotemberg and Woodford (1991, 1999) consider the issue of a counter-cyclical markup extensively while Johri (2001) provides a survey of models of markup variation in response to fluctuations in demand.

12 We can write (3) in the following form: $\overline{\mu}_t = \delta_1 \overline{y}_t + \delta_2 \overline{k}_t + \delta_3 \overline{h}_t + \left[(p - w)_t + \delta_4\, trend + \delta_5\right]$. If we interpret the trend in the markup of average prices on average wages as due to the persistent increases in average productivity then we can replace $\delta_4\, trend + \delta_5$ with $y - h$ and arrive at (4).

13 The assumption of a linear relationship between inflation and the markup in (5) cannot strictly be true since the markup approaches zero as inflation tends to infinity. However, we assume that over the smaller range of values of inflation considered by us, the log linear relationship is a good approximation.

[14] A unit coefficient on productivity imposes linear homogeneity on the markup where a change in costs will, all else equal, lead to an equivalent change in prices leaving the markup unchanged in the long-run. In this model, 'all else equal' includes no change in the rate of inflation in the long run.

[15] The term $mu + prod$ in (5) represents either the unit cost markup or the marginal cost markup depending on how mu is defined.

[16] Measurement errors in national accounts data often have a simultaneous impact on the price and output series so as to offset each other. Consequently, estimates of the relationship between price and output data would be contaminated by the presence of common measurement errors and this contamination is likely to be serious when the span of the price and output series are the same or very similar as in our case.

[17] This approach adjusts for co-breaking in the output and labour input series that would not be captured if the series were de-trended individually.

[18] The augmented Perron (1997) unit root test allows for the presence of an endogenous one-time change in the level and slope of the trend function. The test identifies a break in the trend and constant in June 1974 with a unit root test statistic of - 5.3 compared with a 95 per cent critical value of - 3.13 indicating the null of a unit root is rejected.

[19] Results available from the authors on request.

[20] The companion matrix in both systems is consistent with the maintained hypothesis of one cointegrating vector in a trivariate system of I(1) variables where we expect two roots at unity and the other bounded away from unity.

[21] The data are quarterly and so the inflation cost coefficient is divided by four to calculate the 'annual' inflation cost coefficient.

[22] The results are of a similar order of magnitude to those reported in Banerjee *et al.* (2001), and Banerjee and Russell (2001a) using non United States data. The annualised inflation cost coefficients using I(1) estimation techniques for Australia, Canada, France, Germany, Italy and the United Kingdom, were 1.3, 1.1, 0.7, 1.2, 2.0, and 0.6 respectively. Similarly, the 'implicit' annualised inflation cost coefficient from equation (1) on page 438 of de Brouwer and Ericsson (1998) for Australia is 2.8.

[23] Since this is an important issue, we investigated the time series properties of the adjustment factor, $\overline{a}_t$, using recursive estimation. We find that for a range of values of the steady state markup between 1.0 and 1.6, there is no evidence of breaks in mean or trend for this series. These results are available from us upon request.

[24] Modelling stationary dynamic adjustments to the markup reduce biases in the estimates although asymptotically, leaving out stationary terms should not affect the long-run estimate. See Banerjee *et al.* (1986).

[25] The values of the estimated inflation cost coefficient, corresponding to the steady state markups of 1.5, 1.4, 1.3, 1.2 and 1.1, are 5.853, 5.140, 4.382, 3.547 and 2.717 respectively.