Inflation and Measures of the Markup

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Abstract

Theoretical models of the markup-inflation relationship focus on the markup of price on marginal costs in contrast with empirical models that typically focus on the markup on unit costs. Using nearly 50 years of quarterly United States data we identify a negative long-run relationship between inflation and the markup of price on unit costs on the one hand and with the markup on marginal costs on the other. We find that the impact of inflation on the marginal cost markup is larger than the impact on the unit cost markup.

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

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

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1. INTRODUCTION

In recent years there has been mounting empirical evidence of a negative relationship between inflation and the markup. For example, the error correction term in the models of inflation estimated by Richards and Stevens (1987), Franz and Gordon (1993), Cockerell and Russell (1995), and de Brouwer and Ericsson (1998) may be interpreted as the markup that is negatively related with inflation. Further evidence includes work by Bénabou (1992), Simon (1999) and Batini, Jackson and Nickell (2000). All these models assume that inflation and the markup are stationary.

In contrast, Banerjee, Cockerell and Russell (2001), Banerjee and Russell (2000, 2001a, 2001b) and Banerjee, Mizen and Russell (2002) argue that the variables should be treated as integrated and identify a negative long-run relationship between inflation and the markup.[1] Characterising the relationship as the long run between two integrated variables allows us to consider two types of empirical relationships. The first is a short-run relationship between the stationary components in both series that may, or may not, involve the business cycle. The second is the long-run relationship between the integrated components in both inflation and the markup.

Standard theoretical explanations of the negative relationship 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 measure, most empirical work has estimated the relationship using the markup on unit or average costs since these can be measured directly. Furthermore, one may argue that the empirical relationship between inflation and the markup is dependent on the way the markup is measured.

The important issue for this paper is that the markup of price on marginal costs diverges from the markup of price on unit costs over the business cycle and might imply relationships of

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

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

different magnitude in the long run with inflation.[3] Both relationships are of interest depending on the economic phenomenon we wish to explore. If we are concerned about the impact of inflation on fixed capital formation then, given imperfect capital markets and the predilection of firms to fund investment through retained earnings, the relationship with the unit cost markup is more relevant.[4] Alternatively, if our interest were more in the employment consequences of persistently higher inflation then the impact on the marginal cost markup would be more informative.

Two broad issues that follow from the inconsistency between the theoretical and empirical measures of the markup 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 examine this issue, we estimate two cointegrated systems, with the markup measured on unit costs in one system and on marginal costs in the other. Using quarterly United States data for the period June 1953 to March 2000 we re-establish the negative long-run relationship between the unit cost markup and inflation identified in earlier work. We also identify a negative long-run relationship between inflation and the marginal cost markup.

Second, we show that the marginal cost markup can be thought as equivalent to the unit cost markup plus an adjustment related to the business cycle. The adjustment reflects differences over the business cycle between marginal productivity and average productivity as well as differences between the marginal markup of price on wages and the average markup. If the adjustment is a stationary variable it should not influence the estimate of the long-run relationship suggesting that the long-run estimates computed from the two systems should not be significantly different. Conversely, if the adjustment is non-stationary then the long-run estimates may differ.

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[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.

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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). Using this approach we show that any difference in the estimated long-run relationship using the two measures of the markup is due to the omitted dynamics associated with the business cycle. The cointegrated systems are estimated in Section 3. We show that the results are qualitatively insensitive to whether the empirical analysis is in terms 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.

2. 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 to 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:

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

where F_H is the marginal product of labour, and RW_t is the real wage defined as W_t/P_t where 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, $\overline{\mu}_t$:[5]

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

where $\mu*$ 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.[6] Equation (3) represents the direct and indirect effects of the business cycle on the markup of price on

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

[6] 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).

marginal costs. The direct effect is through the markup of price on average wages, $\overline{p-w}$. The remaining influences, through $\overline{y}$, $\overline{k}$ and $\overline{h}$ represents 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.[7] Explanations of the finding of a counter-cyclical markup usually focus on increasing marginal costs due to firms introducing less productive inputs into the production process as the economy expands or on increased wage and cost pressures with higher demand.[8] In summary, given the counter-cyclical nature of the markup of price on wages, $p-w$, in (3) this 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. 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 that productivity is actual productivity instead of trend productivity which is the implicit assumption of (3). In this case the marginal cost

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[7] 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.

[8] 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.

markup, $mcmu_t$, can be thought of as an 'adjustment' to the unit cost markup, $(p - w)_t + (y - h)_t$, such that:

$$mcmu_t = \overline{a}_t + (p - w)_t + (y - h)_t \qquad (4)$$

where $\overline{a}_t = \dfrac{e - \mu * S_K}{e - e\,\mu * S_K}\,\overline{y}_t + \dfrac{(1-e)\mu * S_K}{e - e\,\mu * S_K}\,\overline{k}_t - \dfrac{\mu * S_H}{1 - \mu * S_K}\,\overline{h}_t$ is the marginal cost adjustment to unit costs due to the indirect effects of the business cycle.[9]

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, $\overline{a}_t$. If $\overline{y}$, $\overline{k}$ and $\overline{h}$ are stationary then the adjustment, $\overline{a}_t$, will also be stationary and the statistical properties of the marginal cost markup and the unit cost markup will be the same. This implies the long-run relationship will be the same irrespective of whether the markup is defined on marginal or unit costs. Alternatively, if $\overline{y}$, $\overline{k}$ or $\overline{h}$ 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-linearities, 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) that variations in the marginal cost markup can be thought of as a function of deviations in the unit cost markup subject to the effects of the business cycle on output, capital and hours of work is likely to be maintained.

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[9]   We can write (3) in the following form: $\overline{\mu}_t = \delta_1\,\overline{y}_t + \delta_2\,\overline{k}_t + \delta_3\,\overline{h}_t + \big[(p - w)_t + \delta_4\,trend + \delta_5\big]$. If we interpret the trend in the markup 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).

3.   THE RELATIONSHIP BETWEEN THE MARKUP AND INFLATION

We now turn to estimating the long-run cointegrating relationship between inflation and the markup. We proceed by estimating a three variable cointegrating system using now 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) where, along with Banerjee and Russell (2001), 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:

$$mu + prod = q - \lambda \Delta p \qquad\qquad (5)$$

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 'inflation cost' parameter, and Δ represents the change in the variable. 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.[10] Alternatively, if, mu_t, is defined as $(p - w)_t + \bar{a}_t$ where $\bar{a}_t$ is defined as in equation (4) then $mu + prod$ is the markup of prices on marginal costs. Consequently, the inflation cost parameter, λ, represents the impact, or cost, of inflation in terms of either a lower unit or marginal cost markup in the long run depending on the measure of the markup we use.

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[10]   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.

3.1 The data

The cointegrated systems are estimated with quarterly United States data for the period June 1953 to March 2000. The markup and inflation are derived from national accounts data. Prices, wages and output are measured on a 'private sector' basis excluding the contribution of federal, state, and local governments. The national accounts data are from the National Income and Product Accounts tables published by the Bureau of Economic Analysis. 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 is 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 to represent 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 so not suitable as a predetermined variable while national accounts measures of the business cycle suffer 'errors in measurement' problems when used in association with national accounts price data.[11] To avoid these difficulties the business cycle is represented by de-trended natural logarithm of non-agricultural private sector hours of employment.

The systems were 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.

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[11] 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.

3.2 The Marginal Cost 'Adjustment' Factor

We follow Rotemberg and Woodford (1991) and ignore the capital stock component of (3) and choose their 'baseline' values of $e = 1$ and $\mu^* = 1.6$.[12] The output and employment components of $\bar{a}_t$ 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 endogenous break in the series.[13] For our sample, labour's share of income, S_H, is 0.659 with the level of the marginal cost adjustment:

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

The level of the marginal cost adjustment is shown in the top panel of Graph 1. Using Perron (1997), we find a break in the level and trend of the marginal cost adjustment, a_t, at June 1974.[14] Once the breaks are accounted for, unit root tests unambiguously indicate the de-trended marginal cost adjustment series, $\bar{a}_t$, is stationary. The de-trended marginal cost adjustment series is shown in the lower panel of Graph 1. Using this series we calculate the marginal cost markup, $\bar{a}_t + (p - w)_t + (y - h)_t$ 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

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[12] Sensitivity of the results to the choice of steady state markup is considered following the estimation.

[13] This approach allows for co-breaking in the output and labour input series that would be of importance if the series were de-trended individually.

[14] 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.

and DF-GLS (Elliot, Rothenberg and Stock (1996)) univariate unit root tests. [15] All three unit root tests indicate that the markup, $(p - w)_t$, the marginal cost markup, $(p - w)_t + \bar{a}_t$, 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, $\bar{a}_t$, is stationary. The results from the system analysis and the system unit root tests are consistent with these findings.

3.3 Results from Estimating the Markup and Inflation Systems

The trace statistics that test 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.[16] The diagnostics of the systems reported in the notes to Tables 1 and 2 show that both systems have 'well behaved' diagnostics, but those of the marginal cost markup system perform considerably better.

The normalised long-run coefficients with 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 latter is defined on marginal or unit costs. We note that the estimate of the inflation cost coefficient in the unit cost markup system is 2.677 implying a 1 percentage point increase in inflation is associated with approximately 0.67 of a percentage point decline in the markup of price on unit costs in the long run.[17] This estimate is very similar to the annual estimates for the United States reported in Banerjee and Russell (2001a, b) of 0.62 using annual 'private sector' gross domestic product data and 0.46

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[15] Results available from the authors on request.

[16] 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.

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

using aggregate gross domestic product data. In contrast, the estimate of the inflation cost coefficient in the marginal cost markup system is 6.602 which is around 2 ½ times the estimate from the unit cost markup system.

It appears that the two estimates of the inflation cost coefficient are significantly different.[18] The difference can be explained in two ways. First, it is well known that even in medium to large samples, specifying the dynamics properly is an issue of some importance. 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.[19] Therefore, the different estimates may simply be due to the inclusion of the marginal cost adjustment providing better dynamics that improve the estimate of the inflation cost coefficient in finite samples. The improved diagnostics of the marginal cost markup system compared with the unit cost markup system lends some support to this argument.

The second explanation is that the de-trended marginal cost adjustment series may be non-stationary with a changing mean even though the evidence from unit root tests is against integration.[20] The lower panel of Graph 1 suggests that the de-trended marginal cost adjustment series may have a downward shift in mean in the mid 1960s and then a reversal of that shift at the start of the 1980s. These shifts may then be negatively correlated with movements in the mean rate of inflation for the same periods and this may affect the long-run relationship and explain the larger inflation cost coefficient in the marginal cost markup system.

To investigate this proposition we adjust for the shifts in mean in the de-trended marginal cost adjustment, a_t, and re-estimate the marginal cost markup system using this de-trended

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[18]   The two estimated coefficients are more than two standard deviations from each other although the confidence intervals from the two estimates overlap. Formal testing of the proposition is complicated as the coefficients are from two different models.

[19]   See Banerjee, *et al.* (1986).

[20]   The low power of unit root tests to reject the null of unit roots in the presence of shifts in mean is well documented. See Hendry and Neale (1987).

variable.[21] The results are reported in Table 3 along with the trace statistics and critical values. We again accept one negative long-run relationship between inflation and the markup on marginal costs and the inflation coefficient is now similar to that estimated in the unit cost system.

4.   INTERPRETING THE RESULTS AND CONCLUSION

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, μ^*) may be considerably lower than 1.6. Therefore, to determine the sensitivity of the estimates from the marginal cost markup system we repeat the system analysis with the marginal cost adjustment, $\bar{a}_t$, calculated assuming a range of values between 1.0 and 1.5 for the steady state markup, μ^* . The results are qualitatively the same as those reported above in the sense that we continue to identify a negative long-run relationship between inflation and the markup. As the steady state markup is increased in increments of 0.1 from 1.0 to 1.5 the estimate of the inflation coefficient increases steadily from 2.035 to 5.883 for the system without the mean of $\bar{a}_t$ adjusted (i.e. the system estimated in Table 2). Adjusting for the shift in mean in the marginal cost adjustment series, $\bar{a}_t$, results in the inflation cost coefficient being less sensitive to the choice of steady state markup and increasing from 2.318 to 3.052 as the steady state markup increases from 1.0 to 1.5 (i.e. the system reported in Table 3).

It appears, therefore, that the long-run estimates are sensitive not only to the marginal cost adjustment but also to the value of the steady state markup. Removal of the shift in mean in the marginal cost adjustment series reduces the estimated inflation cost coefficient and is indicative of the source of the reduction. The slowdown in productivity following the first OPEC oil price shock leads to a downward shift in the mean of the marginal cost adjustment as the ratio of output to labour input changes. This shift in mean is accentuated by the choice of steady state markup. Consequently, a high value for the steady-state markup amplifies the

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[21] We also de-trended the marginal cost adjustment, a_t, with a second order polynomial trend. However, $\bar{a}_t$ continued to display a shift in mean and the estimated inflation cost coefficient was very similar to those reported for the marginal cost markup system in Table 2 when a linear trend was used.

shift in mean of the marginal cost adjustment induced by the slowdown in productivity. Removing the shift in mean eliminates the source of this amplification and leads to the marginal cost adjustment having less effect on the inflation cost coefficient.

In light of the above arguments, we need to ask whether we are justified in an economic sense in adjusting the marginal cost adjustment series for the mean shift. If the slowdown in productivity has no persistent effect on the marginal cost markup, this suggests that the marginal cost adjustment should be stationary and removing the shift in mean is justified. In this case the appropriate estimate of the inflation cost coefficient for the marginal cost markup system (from Table 3) is - 3.255 assuming a steady state markup of 1.6. Alternatively, if the slowdown in productivity is, in part, a response to the fall in the unit cost markup (in turn partly due to higher inflation) and the effect of the slowdown in the growth in productivity is persistent, it is not justified to account for the shift in means. In this latter case the marginal cost adjustment is non-stationary (in particular stationary with shifting means) and the estimate of the inflation cost coefficient is considerably larger with a value of – 6.602 assuming a steady state markup of 1.6.

Finally, the analysis indicates that the empirical identification of a negative long-run relationship between inflation and the markup is not dependent on how the markup is measured. However, taking account of better short-run dynamics by adjusting the unit cost definition of the markup for business cycle effects leads to a larger value of the trade-off in which we are interested, and additionally, alters the value of the short-run relationship.

5.   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 $\bar{a}_t$		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 break in constant and 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.
De-meaned Marginal cost Adjustment		MCA regressed on a constant, step dummy June 1965 to March 1980, step dummy June 1980 to March 2000, spike dummy June 1965, spike dummy June 1980.

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

6. REFERENCES

Banerjee, A., J. Dolado, D.F. Hendry and G.W. Smith (1986). Exploring Equilibrium Relationships in Econometrics through Static Models: Some Monte Carlo Evidence, *Oxford Bulletin of Economics and Statistics*, vol. 48, pp. 253-77.

Banerjee, A, L. Cockerell and B. Russell, (2001). An I(2) Analysis of Inflation and the Markup, *Journal of Applied Econometrics*, Sargan Special Issue, vol. 16, pp. 221 -240.

Banerjee, A., J. Dolado, J.W. Galbraith and D.F. Hendry, (1993). *Cointegration, Error Correction, and the Econometric Analysis of Non-Stationary Data*, Oxford University Press, Oxford United Kingdom.

Banerjee, A., P. Mizen and B. Russell (2002). The Long-Run Relationships among Relative Price Variabiltiy, Inflation and the Markup, European University Institute Working Papers, Eco No. 2002/1.

Banerjee, and B. Russell (2000). The Markup and the Business Cycle Reconsidered, European University Institute Working Papers, Eco No. 2000/21.

Banerjee, A, and B. Russell, (2001a). The Relationship between the Markup and Inflation in the G7 Economies and Australia, *Review of Economics and Statistics*, vol. 83, No. 2, May, pp. 377-87.

Banerjee, A, and B. Russell, (2001b). Industry Structure and the Dynamics of Price Adjustment, *Applied Economics*, May, vol. 33, pp. 1889-1901.

Barsky, R. and G. Salon (1986). Real Wages over the Business Cycle, NBER Working Paper 2888, March.

Basu, S. (2000). Understanding how Price Responds to Costs and Production: A Comment, Carnegie-Rochester Conference Series on Public Policy, vol. 52, pp. 79-85.

Batini, N., B. Jackson and S. Nickell (2000). Inflation Dynamics and the Labour Share in the UK, Bank of England External MPC Unit Discussion Paper, no. 2, November.

Bénabou, R. (1992). Inflation and Markups: Theories and Evidence from the Retail Trade Sector, *European Economic Review,* 36(3), pp. 566-574.

Bénabou, R. and J.D. Konieczny (1994). On Inflation and Output with Costly Price Changes: A Simple Unifying Result, *American Economic Review*, March, 84(1), pp. 290-7.

Bils, M. (1987). The Cyclical Behavior of Marginal Cost and Price, *American Economic Review*, vol. 77, December, pp. 838-57.

Bils, M. and J.A. Kahn (1996). What Inventory Behaviour tells us about Business Cycles, Rochester Center for Economic Research Working Paper, no. 428.

Bils, M. and Y. Chang (2000). Understanding how Price Responds to Costs and Production, Carnegie-Rochester Conference Series on Public Policy, vol. 52, pp. 33-77.

Blanchard, O.J. and N. Kiyotaki (1987). Monopolistic Competition and the Effects of Aggregate Demand, *American Economic Review*, 77, September, pp. 647-66.

Calvo, G. (1983). Staggered Prices in a Utility-Maximising Framework, *Journal of Monetary Economics*, 12, pp. 383-98.

Cockerell, L. and B. Russell (1995). Australian Wage and Price Inflation: 1971-1994, Reserve Bank of Australia Discussion Paper, no. 9509.

Danziger, L. (1988). Costs of Price Adjustment and the Welfare Economics of Inflation and Disinflation, *American Economic Review*, September, 78(4), pp. 633-46.

Dunlop, J.T. (1938). The Movement of Real and Money Wage Rates, *Economic Journal*, vol. 18, pp. 413-34.

de Brouwer, G and N. R. Ericsson (1998). Modelling Inflation in Australia, *Journal of Business & Economic Statistics,* vol. 16, no. 4, pp. 433-49.

Elliott, G., T.J. Rothenberg and J.H. Stock (1996). 'Efficient Tests for an Autoregressive Unit Root', *Econometrica*, vol. 64, no. 4, July, pp. 813-36.

Engle, R.F. and C.W.J. Granger (1987). 'Co-integration and Error Correction: Representation, Estimation, and Testing', *Econometrica*, vol. 55, pp. 251-76.

Franz, W. and R.J. Gordon (1993). German and American Wage and Price Dynamics: Differences and Common Themes, *European Economic Review*, vol. 37, pp. 719-62.

Gali, J. (1994). Monopolistic Competition, Business Cycles and the Composition of Aggregate Demand, *Journal of Economic Theory*, 63, pp. 73-96.

Hall, R.E. (1988). The Relation Between Price and Marginal Cost in US Industry, *Journal of Political Economy*, vol. 96, no. 5, pp. 92-147.

Hendry, D.F., and A.J. Neale (1987). Monte Carlo Experimentation using PC-NAIVE, in T. Fomby and G. Rhodes (eds.), *Advances in Econometrics*, Basil Blackwell, Oxford.

Johansen, S. (1988). Statistical Analysis of Cointegration Vectors. *Journal of Economic Dynamics and Control*, 12, pp. 231-54,

Johansen, S., (1995). *Likelihood-Based Inference in Cointegrated Vector Autoregressive Models*, Oxford University Press, Oxford.

Johri, A. (2001). Markups and the Seasonal Cycle, *Journal of Macroeconomics*, vol. 23, no. 3, pp. 367-95.

Konieczny, J.D. (1990). Inflation, Output and Labour Productivity when Prices are Changed Infrequently, *Economica*, May, 57(226), pp. 201-18.

Kuran, T. (1986). Price Adjustment Costs, Anticipated Inflation, and Output, *Quarterly Journal of Economics*, December, 71(5), pp. 1020-7.

Kydland, F.E., and E.C. Prescott (1988). Cyclical Movements of the Labor Input and its Real Wage, Working Paper 413, Research Department, Federal Reserve Bank of Minneapolis.

Kwiatowski, D., P.C.B. Phillips, P. Schmidt and Y. Shin (1992). Testing the Null Hypothesis of Stationarity Against the Alternative of a Unit Root: How Sure are we that Economic Time Series have a Unit Root?, *Journal of Econometrics*, vol. 54, pp. 159-78.

Layard, R., S. J. Nickell and R. Jackman (1991). *Unemployment Macro-economic Performance and the Labour Market*, Oxford University Press, Oxford.

Lucas, R.E.J. (1973). Some International Evidence on Output-Inflation Tradeoffs". *American Economic Review,* 1973, vol. 63(3), pp. 326-34.

Martins, J.O., S. Scarpetta and D. Pilat (1996). Mark-up Ratios in Manufacturing Industries: Estimates for 14 OECD Countries, OECD Economics Department Working Papers, no. 162.

Myers, S.C. and N. Majluf (1984). Corporate Financing and Investment Decisions When Firms Have Information Investors Do Not Have, *Journal of Financial Economics*, vol. 13, pp. 187-221.

Naish, H.F. (1986). Price Adjustment Costs and the Output-Inflation Trade-off, *Economica*, May, 53(210), pp. 219-30.

Perron, P. (1997). Further Evidence on Breaking Trend Functions in Macroeconomic Variables, *Journal of Econometrics*, vol. 80, no. 2, pp. 355-85.

Pigou, A.C. (1927). *Industrial Fluctuations*, London, Macmillan.

Phelps, E.S. and S.G. Winter (1970). Optimal Price Policy under Atomistic Competition, in E. Phelps ed., *Microeconomic Foundations of Employment and Inflation Theory*, W.W. Norton and Co, New York.

Richards, T. and G. Stevens (1987). Estimating the Inflationary Effects of Depreciation, Reserve Bank of Australia Discussion Paper, no. 8713.

Rotemberg, J.J. (1982). Sticky Prices in the United States, *Journal of Political Economy*, 90, pp. 1187-211.

Rotemberg, J.J. (1983). Aggregate Consequences of Fixed Costs of Price Adjustment, *American Economic Review*, June, 73(3), pp.219-30.

Rotemberg, J.J. and M. Woodford (1991). Markups and the Business Cycle, NBER Macroeconomics Annual, 63.

Rotemberg, J.J. and M. Woodford (1999). The Cyclical Behaviour of Prices and Costs, in J.B. Taylor and M. Woodford (eds.), *Handbook of Macroeconomics*, North Holland Press, Amsterdam.
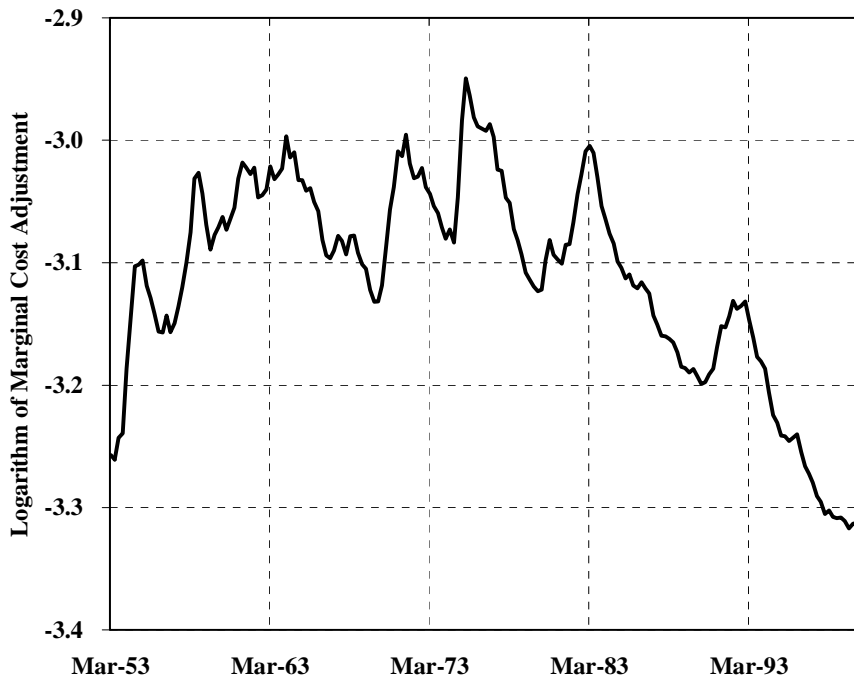
Simon, J. (1999). Markups and Inflation, Department of Economics, Mimeo, MIT.

Solow, R.M. (1957). Technical change and the Aggregate Production Function, *Review of Economics and Statistics*, vol. 39, August, pp. 312-20.
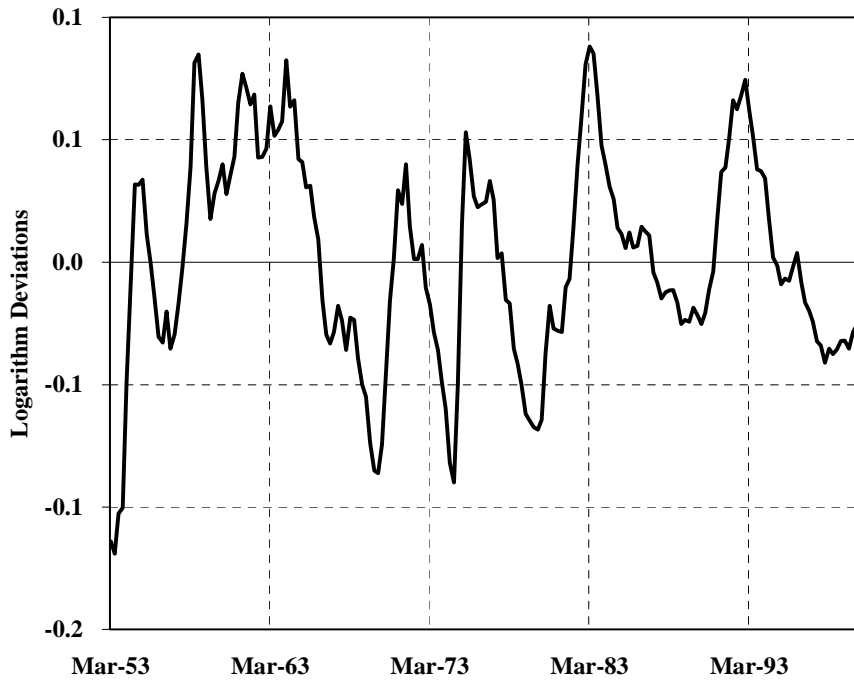
Tarshis, T. (1939). Changes in Real and Money Wage Rates, *Economic Journal*, vol. 19, pp. 150-4.
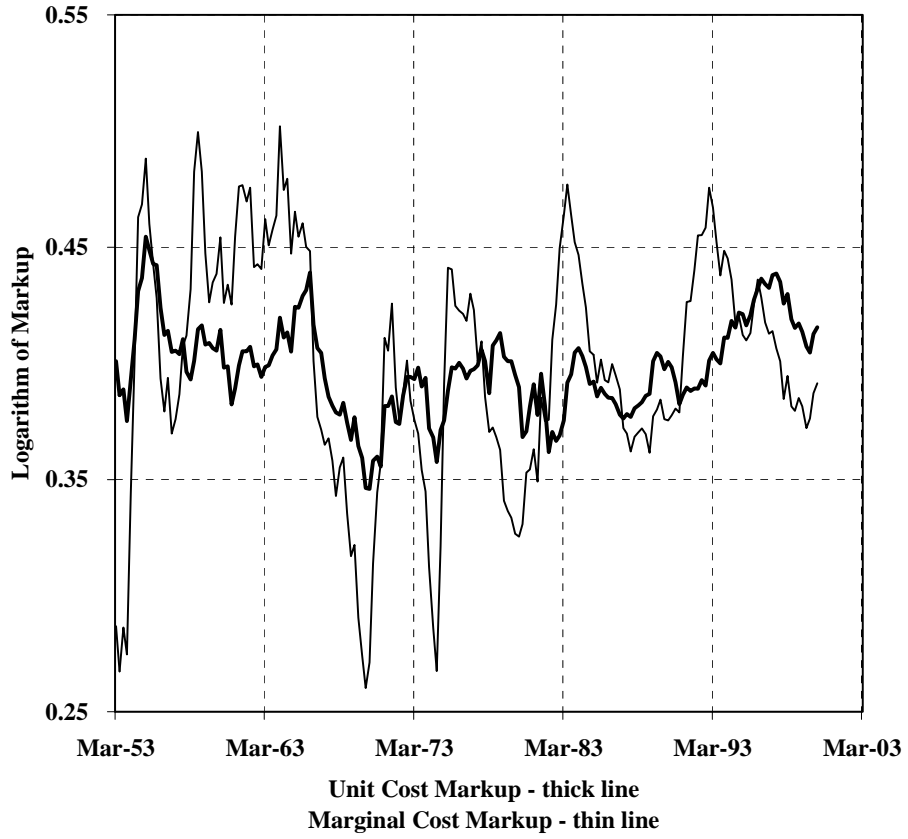
Graph 1: Marginal Cost Adjustment

Level of the Marginal Cost Adjustment

Detrended Marginal Cost Adjustment

Graph 2: Unit Cost and Marginal Cost Markups

Unit Cost Markup - thick line
Marginal Cost Markup - thin line

Table 1: Normalised Cointegrating Vectors

'Unit Cost Markup' System

	$p - w$	$y - h$	Δp
Unrestricted	1	1.019	2.677
	(0.019)	(0.019)	(0.735)
Linear Homogeneity Imposed	1	1	2.526
			(0.700)
Adjustment Coefficients	- 0.006	- 0.121	- 0.006
	[- 0.3]	[- 3.9]	[- 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) $\chi^2 (393) = 442.59$, p-value = 0.04

LM(1) $\chi^2 (9) = 8.49$, p-value = 0.49

LM(4) $\chi^2 (9) = 16.12$, p-value = 0.06

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

Table 2: Normalised Cointegrating Vectors

'Marginal Cost Markup' System

	$\bar{a}+(p-w)$	$y-h$	Δp
Unrestricted	1	1.044	6.333
	(0.024)	(0.023)	(1.048)
Linear Homogeneity Imposed	1	1	6.602
			(0.956)
Adjustment Coefficients	- 0.094	- 0.091	- 0.009
	[- 4.9]	[-5.7]	[- 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) $\qquad$ $\chi^2(393) = 418.39$, p-value = 0.18

LM(1) $\qquad$ $\chi^2(9) = 13.29$, p-value = 0.15

LM(4) $\qquad$ $\chi^2(9) = 14.08$, p-value = 0.12

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

Table 3: Normalised Cointegrating Vectors

'Marginal Cost Markup' System

Adjusting for the Shift in Mean in the Marginal Cost Adjustment

	$\bar{a} + (p-w)$	$y - h$	Δp
Unrestricted	1 (0.017)	1.017 (0.017)	3.222 (0.763)
Linear Homogeneity Imposed	1	1	3.255 (0.712)
Adjustment Coefficients	- 0.157 [- 6.8]	- 0.091 [-4.7]	- 0.001 [- 0.2]

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:

$$ECM_t \equiv (p - w)_t + (y - h)_t + 3.222 \Delta p_t.$$

Likelihood ratio tests (a) linear homogeneity is accepted $\chi_1^2 = 0.86$, p-value = 0.35; and (b) test of coefficient on inflation is zero is rejected, $\chi_1^2 = 12.37$, p-value = 0.00.

Predetermined Variables: Spike dummies – September 1965, June 1974, March 1975 and June 1980. Step dummies for June and September 1965, June and September 1980. Business cycle variable lagged one period.

The level of the marginal cost adjustment, $\bar{a}_t$, was demeaned by regressing $\bar{a}_t$ on a constant, two step dummies for June 1965 to March 1980 and June 1980 to March 2000, and spike dummies for June 1965 and June 1980.

Testing for the Number of Cointegrating Vectors

Estimated trace statistic for the null hypothesis $H_0 : r = 0$ is 56.84 {26.70}, $H_0 : r = 1$ is 10.84 {13.31}, and $H_0 : r = 2$ is 0.83 {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 168 degrees of freedom.

System Diagnostics for the Model with Linear Homogeneity Imposed

(a) *Tests for Serial Correlation*

Ljung-Box (47) $\chi^2 (393) = 421.30$, p-value = 0.16

LM(1) $\chi^2 (9) = 10.15$, p-value = 0.34

LM(4) $\chi^2 (9) = 16.89$, p-value = 0.05

(b) *Test for Normality:* Doornik-Hansen Test for normality: $\chi^2 (6) = 10.06$, p-value = 0.12