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Publication date:
2000

Document Version
Publisher's PDF, also known as Version of record

The Relationship between the Markup and Inflation in the G7 plus One Economies*

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Abstract

An I(2) analysis of inflation and the markup is undertaken for the G7 economies and Australia. We find that the levels of prices and costs are best described as I(2) processes and that except for Japan a linear combination of the log levels of prices and costs cointegrate to the markup that is integrated of order 1. It is also shown that the markup in each case cointegrates with inflation and that higher inflation is associated with a lower markup in the long-run.

Keywords: Inflation, Wages, Prices, Markup, I(2), Polynomial Cointegration.
JEL Classification: C22, C32, C52, E24, E31

Presented at the Econometric Society 8th World Congress

Seattle, USA, August 11 – 16 2000

This paper is forthcoming in the *Review of Economics and Statistics*

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1 INTRODUCTION

The proposition examined in this paper is that there exists a long-run relationship in the sense proposed by Engle and Granger (1987) where the markup decreases as inflation increases and *vice versa*.[1] This paper estimates this relationship using data from the G7 economies and Australia. A central feature of our analysis is that the level of prices and costs may be taken to be integrated of order 2, denoted I(2), for the purposes of modelling. In other words, both the differences of prices and costs and their levels that comprise the markup display persistent behaviour over the samples investigated. This requires us to make use of recently developed techniques for the estimation of I(2) processes developed by Johansen (1995a, b) *inter alia*.

Bénabou (1992) argues within a price-taking model that higher inflation leads to greater competition and therefore a lower markup. In contrast, Russell, Evans and Preston (1997), Chen and Russell (1998) and Simon (1999) focus on the difficulties that price-setting firms face when adjusting prices in an inflationary environment where there is missing information. In this case the lower markup with higher inflation is interpreted as the higher cost of overcoming the missing information with higher inflation. Importantly, Russell *et al.* and Chen and Russell argue that information remains missing in the steady state and that the relationship between rates of steady state inflation and the markup will also remain in the steady state.[2]

1 The logarithm of the markup, mu, is defined as $mu \equiv p - \sum_{i=1}^{n} \psi_i c_i$ where p and the c_i's are the logarithms of prices and the costs of production respectively, and $\sum_{i=1}^{n} \psi_i = 1$. If the latter condition is not satisfied then the relationship between prices and costs cannot be termed the markup.

2 The steady state is defined as all nominal variables growing at the same constant rate.

Banerjee, Cockerell and Russell (1998) investigate the proposition using Australian inflation data and find strong empirical support of the proposition. An important question is whether the findings in Banerjee *et al.* are in some way peculiar to the Australian data. The 'peculiarity' of the data may be due to the nature of the shocks encountered over the sample examined, the behaviour of the Australian monetary authorities or the structure of the economy. Alternatively, the findings may be applicable to developed western economies in general when inflation is non-stationary. To this end we proceed to examine the proposition for the G7 economies and Australia.

The empirical investigation proceeds in two stages. First we estimate an I(2) system for each economy of the core variables of interest, namely prices and costs. Except for Japan, we find that a polynomially cointegrating relationship is present between the level of the markup and the changes in the core variables.[3] Having obtained an estimate from the I(2) analysis of the long-run relationship between the markup and general inflation of the core variables, we proceed to estimate an I(1) system in order to obtain the direct relationship between price inflation alone and the markup. The estimated I(1) system is a particular and full reduction of the I(2) system and corroborates the findings in the I(2) system.

While differences emerge between the economies, the finding of polynomial cointegration for the G7 economies and Australia is remarkably robust. The only exception is Japan where the levels of prices and costs cointegrate to an I(1) variable but it cannot be interpreted as the markup. Therefore, it appears that except for Japan the proposition that there exists a

[3] Polynomial cointegration occurs when the cointegrated levels of the data cointegrate with the differences in the levels. In our case the I(2) levels of prices and costs cointegrate to the markup which is I(1) and the markup then cointegrates with inflation which is also I(1). For a detailed discussion concerning polynomial cointegration see Johansen (1995b).

negative long-run relationship between inflation and the markup is consistent with the data in the G7 economies as well as in Australia.

2 AN IMPERFECT COMPETITION MARKUP MODEL OF PRICES

We propose estimating an imperfect competition markup equation in the Layard / Nickell tradition for the eight economies.[4] It is assumed that in the long-run firms desire a constant markup, q, of prices, p, on unit costs net of the cost of inflation. Short-run deviations in the markup are due to the business cycle and non-modelled shocks. For an open economy the main inputs are labour and imports and we can write the inflation cost long-run markup equation as:[5]

$$mu = p - \delta\, ulc - (1-\delta)\, pm = q - \lambda \Delta p \qquad (1)$$

where ulc and pm are unit labour costs and unit import prices respectively and δ and λ are positive parameters. Lower case variables are in logarithms and Δ represents the change in the variable.

When the inflation cost coefficient, λ, is zero, inflation imposes no costs on the firm in the long-run and the long-run markup equation collapses to the standard Layard / Nickell model.

[4] For the standard Layard / Nickell model see Layard, Nickell and Jackman (1991) or Carlin and Soskice (1990). For a detailed discussion of empirical models relating the markup with inflation see Cockerell and Russell (1995) and Banerjee *et al.* (1998).

[5] Banerjee *et al.* (1998) derives equation (1) and considers in some detail issues concerning the integration properties of the data. The form of the long-run price equation is a generalisation of that estimated in de Brouwer and Ericsson (1998). Two other papers estimating markup models of inflation are Richards and Stevens (1987) and Franz and Gordon (1993).

In the more general case when $\lambda > 0$ inflation imposes costs on the firm in terms of a lower markup net of the cost of inflation.[6] This is given by $q - \lambda \Delta p$.

The coefficients δ and $1-\delta$ in (1) are the long-run price elasticities with respect to unit labour costs and import prices respectively. Linear homogeneity is imposed as the coefficients sum to one so that q represents the markup of prices on costs. Linear homogeneity suggests that all else equal an increase in costs is fully reflected in higher prices in the long-run leaving the markup unchanged.

2.1 The I(2) System

The I(2) system analysis is an extension of the now standard I(1) system analysis. For a detailed theoretical outline of the I(2) analysis see Haldrup (1998), Johansen (1995a, b) and Paruolo (1996). Alternatively, for a brief 'penetrable' survey of the I(2) theory in relation to the model estimated here see Banerjee *et al.* (1998). Other empirical applications of the I(2) theory can be found in Engsted and Haldrup (1998) and Juselius (1998).

For illustration, suppose the long-run price equation can be written as a second order vector autoregression of the core variables, x_t, of dimension $n \times 1$:

$$x_t = \Pi_1 x_{t-1} + \Pi_2 x_{t-2} + \Phi D_t + \mu + \varepsilon_t \qquad (2)$$

6 The long-run price equation (1) cannot be strictly true as it implies that the markup approaches zero as inflation tends to an infinite rate. Russell (1998) overcomes this problem by specifying the cost of inflation in the form; $\lambda[\Delta p/(\Delta p + \phi)]$ where ϕ is trend productivity. Consequently, as inflation tends to an infinite rate the cost of inflation approaches λ. It is assumed that the proposed log-linear model of inflation costs is a fair approximation of the 'true' relationship over the small range of inflation experienced by the economies examined.

where μ is a vector of unrestricted constant terms and D_t is a vector of predetermined variables that are assumed not to enter the cointegration space and on which the empirical analysis is conditioned. The lower case variables are in logs and in our case $n = 3$ and the core variables, x_t, are the price level, unit labour costs and import prices. It is assumed that the variable ε_t is a $n-$ dimensional Gaussian vector of errors.

The I(2) analysis provides us with the orthogonal decomposition into the I(0), I(1) and I(2) relationships of the data with dimensions, r, s and $n-r-s$ respectively. Furthermore, the number of polynomially cointegrating vectors is equal to the number of I(2) trends, $n-r-s$.

2.2 The Data

The data are quarterly, seasonally adjusted and taken from the June 1997 OECD Data Compendium.[7] The length of the data sample for each economy is the maximum possible from that source given the series involved. West German data is used for Germany to avoid data problems associated with the reunification with East Germany.

Except for the United States the price index is the private consumption implicit price deflator at 'factor cost'.[8] Unit labour costs are calculated as total labour compensation divided by constant price GDP. Import prices is the implicit price deflator for the imports of goods and services.

7 See the data appendix for further details.

8 The private consumption implicit price deflator at 'factor cost' is calculated as: $P = P_{MP}/(1+tax)$ where P_{MP} is the consumption implicit price deflator at market prices and tax is the proportion of indirect tax less subsidies in nominal GDP. While the 'factor cost' adjustment is theoretically necessary in practice it has little impact on the results.

The consumption deflator at factor cost was initially used for the United States but gave conflicting results. While the I(2) analysis indicated that the level of prices and costs were best described as I(2) statistical processes, there were a number of indicators to suggest that these series did not cointegrate to the markup. As the 'no markup' result is not useful in investigating the proposition, the GDP implicit price deflator at factor cost was used.[9]

The predetermined variables are the log change in the unemployment rate and a number of spike intervention dummies to capture the sometimes erratic short-run wage and price behaviour of firms and labour.[10] This is especially the case during the OPEC oil price shocks and large shifts in exchange rates and tax regimes. A step dummy is introduced for the period leading up to March 1968 for the United States, March 1975 for France, and March 1970 for Canada. These capture a level shift in the markup that is observable in the data and can be interpreted as reflecting a shift in the competitive environment in these economies. Further details of the pre-determined variables are available in Appendix B.

The log change in the unemployment rate represents the business cycle in the model. An alternative specification of the empirical model would be to include the level of unemployment in the cointegrating space as an endogenous or exogenous variable. However, it is not clear what the economic relationship between the markup, inflation and the level of unemployment would be in the long-run. There is some indication that the relationship may be highly non-linear and may differ substantially among economies. Furthermore, such an inclusion would alter the interpretation of this variable from that of an indicator of the business cycle. It was therefore decided to allow for the effects of the business cycle by

9 The failure to estimate the markup using the consumption deflator may be because the unit labour cost variable is for the whole economy and a poor proxy for unit labour costs associated with consumption expenditures for the United States.

conditioning on a stationary pre-determined variable given by the log change in the unemployment rate and its lags. The data appendix describes in more detail the data and its sources.

The integration properties of the data were investigated using PT and DF-GLS univariate unit root tests from Elliot, Rothenberg and Stock (1996).[11] Prices are clearly I(2) except for Japan and West Germany which are marginally I(2). Similarly unit labour costs are mostly I(2) or marginally I(2). One exception is Australia where it appears that unit labour costs may be I(1). The tests also indicate that import prices may be I(1) for many of the economies. However, univariate tests of the logarithm of the *ratios* of prices to unit labour costs and prices to import prices show clear acceptance of the hypothesis that they are I(1) which can occur only if *all* the core variables are I(2), given that prices are I(2). Consequently we proceed under the assumption that the core variables are I(2). This assumption is supported by the I(2) and I(1) systems analysis below where the results are consistent only with the assumption that the core variables are I(2). Finally, the log of the unemployment rate is found to be best described as an I(1) variable.

2.3 The I(2) System Results

Table 1 shows the results of the joint trace tests for determining r and s for the eight economies. In the case of the United States, Japan, Germany, France and the United Kingdom the hypothesis of $r=1$, $n-r-s=1$ is accepted and our findings are corroborated by looking

10 Three lags of the unemployment variable are initially incorporated with insignificant terms subsequently excluded.

11 These results are available on request from the authors.

at the roots of the companion matrix (see Appendix B).[12] The results therefore show that the levels of prices and costs in each of these economies contain an I(2) trend. Moreover, since $r=1$ there is only one cointegrating vector and hence it is of the polynomially cointegrating type.

For the remaining economies, Italy, Canada and Australia, there is a marginal rejection of $r=1$, $n-r-s=1$. However we choose to accept this null hypothesis since the critical values on which inference is based are asymptotic and have been computed under the assumption that there are no pre-determined variables, including dummies, in the system. Not only would taking account of pre-determined variables raise the critical values (thereby leading to *acceptance* of the maintained hypothesis), the evidence from the roots of the companion matrix for these economies are unambiguously in favour of our hypothesis.[13] The subsequent I(1) system analysis in the next section confirms these results.

Imposing $r=1$ and $n-r-s=1$ on each system imposes a polynomial cointegrating vector on the analysis in each case. Table 2 reports the normalised cointegrating vectors with linear homogeneity imposed for each economy. Except for Japan the hypothesis of linear

[12] The 90 % and 95 % critical values for the case of no pre-determined variables are taken from Paruolo (1996) and are reported in the table below. The 95 % critical values are in italics. Other critical values are available in tables compiled by Rahbek, JØrgensen and Kongsted (1998) and Johansen (1995b).

Critical Values for the Joint Trace Test *Q(s, r)*

n-r	r				
3	0	66.96 *70.87*	47.96 *51.35*	35.64 *38.82*	26.70 *29.38*
2	1		33.15 *36.12*	20.19 *22.60*	13.31 *15.34*
1	2			11.11 *12.93*	2.71 *3.84*
n-r-s		3	2	1	0

[13] The moduli of the first four roots are 1.0, 1.0, 1.0, 0.7144 for Italy, 1.0, 1.0, 0.9881, 0.8161 for Canada and 1.0, 1.0, 0.9417, 0.6533 for Australia under the assumption of $r=1$. A finding of $n-r-1=0$ would therefore not be consistent with the third root of close to unity for these economies if $r=1$ is maintained.

homogeneity is accepted and, therefore, the levels of prices and costs cointegrate to the markup in the polynomially cointegrating vector.

For Japan, Germany, France and Canada import prices enter the markup with an insignificant coefficient. The analysis is therefore re-estimated excluding import prices and the results of the joint trace tests for the two variable systems are reported in Table 1 and again support the hypothesis that $r=1$ and $n-r-s=1$. Reported in Table 2 are the normalised cointegrating vectors. The results now hold as before for Germany, France and Canada but the estimated coefficients for Japan are not interpretable as the markup since the test for linear homogeneity continues to be rejected strongly.

Since the steady state is defined by the condition $\Delta p=\Delta ulc=\Delta pm$ we see in Table 2 that for the economies where the markup is defined, the sum of the coefficients on the difference terms is negative. This implies that there is a negative relationship between general inflation and the markup in the long-run.

3 ESTIMATING THE I(1) SYSTEM

The I(2) analysis provides estimates of polynomial cointegration between a linear combination of the markup and the differences in the core variables. In an economic sense it is necessary for $\Delta p = \Delta ulc = \Delta pm$ in the very long-run. However, the method of summing the coefficients on the difference terms provides only an approximate estimate of the relationship between inflation and the markup, given that the variables may grow at different rates over the finite samples. Furthermore, the theoretical models of Russell *et al.* (1997) and Chen and Russell (1998) posit a long-run relationship between the markup and steady state price inflation alone.

Having established polynomial cointegration in the I(2) analysis, a particular reduction to I(1) space helps us establish the relationship of primary concern to us, namely; between price inflation and the markup. In order to implement this reduction we make use of the result that the decomposition into the I(0), I(1) and I(2) directions is an orthogonal one.

In particular, the vectors β_1' and β_2' lie in the space orthogonal to β_3'. Thus if $\beta_3' \equiv (1, a, b)$, then a basis for the space orthogonal to β_3' is given by the matrix $H = \begin{pmatrix} 1 & 1 \\ -1/a & 0 \\ 0 & -1/b \end{pmatrix}$.

Therefore $\begin{pmatrix} H' x_t \\ f' \Delta x_t \end{pmatrix}$, where f is any 3×1 vector that satisfies the restriction that $f' \beta_3 \neq 0$, provides the transformation to I(1) which keeps all the cointegrating and polynomially cointegrating information. Hence if we take f to be $(1,0,0)'$, then the trivariate system given by $\begin{pmatrix} \Delta p_t \\ mulc_t \\ rer_t \end{pmatrix} = \begin{pmatrix} \Delta p_t \\ p_t - 1/a \, ulc_t \\ p_t - 1/b \, pm_t \end{pmatrix}$ is a valid full reduction and under linear homogeneity $a = b = 1$.[14] Furthermore we can retrieve the implicit markup of prices on unit costs from this I(1) system by rearranging the estimated long-run or cointegrating relationship.[15]

Tests of the number of cointegrating vectors in the I(1) system $(\Delta p_t, mulc_t, rer_t)'$ show that except for the United States the hypothesis of one cointegrating vector is accepted.[16] For the United States there is a marginal rejection of the hypothesis although the eigenvalues of the

14 Hans Christian Kongsted suggested this transformation in Banerjee *et al.* (1998).

15 The markup of prices on import prices might be loosely referred to as the 'real exchange rate' due to its similarity with the relative price of traded and non-traded goods as used by Swan (1963) as a measure of the real exchange rate in his classic article.

16 Appendix C reports the results of the I(1) analysis in more detail.

companion matrix strongly support the finding of 1 cointegrating vector. Given also the argument in Section 2.3 that the critical values are likely to be affected by the presence of dummy variables we proceed on the basis of one cointegrating vector for all the economies.

Table 3 reports the adjustment coefficients and the error correction terms for each economy. We see that the ECM appears strongly in each of the 'markup' equations and, except for Italy, is insignificant in the 'real exchange rate' equations. We see also that the adjustment coefficient in the 'Markup Equation' is on average three times that in the 'Inflation Equation'. This suggests that when these economies are shocked away from the long-run relationship, adjustment back to equilibrium is more through changes in the markup, via the goods and labour markets, than by changes in the rate of inflation through actions of the monetary authorities.

Table 4 reports the implicit long-run price elasticities with respect to costs from the I(1) analysis and the equivalent estimates from the I(2) analysis. Also shown are the estimated inflation cost coefficients, λ, from the I(1) and I(2) analyses.[17] The long-run impact of a one percentage point increase in annual steady state inflation on the markup is shown in the final column and range between 0.3 percent for the United States and 2 percent for Italy. It appears likely, therefore, that the long-run relationship between inflation and the markup is important in an economic sense.

4 CONCLUSION

One explanation of the negative long-run relationship in the data is that the 1970s were a period when supply shocks from the energy and labour markets were very prevalent. The low markup, therefore, simply reflects the lags in price adjustment following the shocks. The

17 The latter are an approximation calculated by assuming $\Delta p = \Delta ulc = \Delta pm$ for each economy in Table 1.

adjustment appears to be very slow for economies with little or no price controls. In most cases the relatively low markups persist for around 10 years following the shocks and the markup does not fully recover until the economy again experiences low inflation.

Graph 1 presents the long-run relationship, LR, for the United States and the United Kingdom from the I(1) analysis along with the realisations of the markup and inflation for five distinct inflationary periods indicated by different symbols.[18] If the 'supply shocks' argument is correct then different mean levels of inflation would not affect the behaviour of the markup. Consequently, realisations of the markup and inflation from different periods of inflation would be distributed evenly along the entire curve in Graph 1. This however is not the case.

It may be seen clearly from Graph 1 that if the data were subdivided into periods of inflation with different means, the associated mean levels of the markup are different. For example, for both the United States and the United Kingdom the early 1960s are shown as crosses on Graph 1 and we see that the markup is high during a period of low inflation. The late 1960s and early 1970s are shown as squares and was a period of slightly higher inflation and a slightly lower markup. We can follow the relationship through each inflationary period until the observations return to hover around low inflation and a high markup for the period following the early 1990s recession.

If the actual observations are followed individually (and not by periods as in the graph) a loose negative short-run relationship between inflation and the markup may sometimes be observed in the data. However, any short-run relationship is confined to different sections of

[18] Similar graphs can be constructed for the other economies but for brevity only the United States and the United Kingdom is shown here. Appendix D reports scatter graphs of inflation and the estimated markup for each economy along with the long-run relationship, LR, for each economy.

the long-run curve depending on the general rate of inflation. Thus while short-run mechanisms are almost certainly reflected in some of the data the relationship is strongly driven by the general rate of inflation.

The ability to separate actual observations of inflation and the markup into distinct period with higher inflation associated with a lower markup and *vice versa*, is further confirmation that inflation is a non-stationary process.

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Table 1: The 'Joint Procedure' for Estimating r and s

Estimated Values of $Q(s, r) = Q(s|r) + Q(r)$

United States

n-r	r				
3	0	156.87	91.41	40.15	36.95
2	1		78.70	13.32	8.37
1	2			23.98	1.33
n-r-s		3	2	1	0

Japan

n-r	r				
3	0	112.50	79.90	52.24	46.10
2	1		41.75	13.40	12.11
1	2			5.24	2.54
n-r-s		3	2	1	0

Germany

n-r	r				
3	0	102.83	62.40	33.80	31.82
2	1		56.40	21.65	15.79
1	2			24.29	3.95
n-r-s		3	2	1	0

France

n-r	r				
3	0	140.76	92.47	61.31	60.33
2	1		64.03	21.36	20.81
1	2			2.80	1.79
n-r-s		3	2	1	0

Italy

n-r	r				
3	0	118.53	88.08	64.70	60.13
2	1		46.25	24.07	21.73
1	2			21.35	3.47
n-r-s		3	2	1	0

United Kingdom

n-r	r				
3	0	172.64	97.53	56.72	54.77
2	1		78.87	9.04	6.34
1	2			9.89	0.75
n-r-s		3	2	1	0

Canada

n-r	r				
3	0	121.73	72.90	51.85	49.36
2	1		44.33	23.08	22.33
1	2			4.83	2.43
n-r-s		3	2	1	0

Australia

n-r	r				
3	0	171.41	111.78	70.76	55.43
2	1		86.23	26.93	15.02
1	2			20.89	4.53
n-r-s		3	2	1	0

Prices and Unit Labour Costs Only

Japan

n-r	r			
2	0	65.54	34.84	30.34
1	1		4.30	3.61
n-r-s		2	1	0

Germany

n-r	r			
2	0	43.96	20.05	18.48
1	1		6.91	1.83
n-r-s		2	1	0

France

n-r	r			
2	0	62.54	33.69	32.61
1	1		5.54	4.47
n-r-s		2	1	0

Canada

n-r	r			
2	0	71.67	29.67	26.96
1	1		5.58	4.96
n-r-s		2	1	0

Notes: Statistics are computed with 4 lags of the core variables. See Appendix B for details of the predetermined variables on which the analysis is conditioned. $Q(s|r)$ is the likelihood ratio statistic for determining *s* conditional on *r*. $Q(r)$ is the likelihood ratio statistic for determining *r* in the I(1) analysis. Critical values are given in Paruolo (1996) as shown in footnote 12.

Table 2: Cointegrating Vectors of the I(2) System Analysis

Sample Periods	US 61:4-97:2	Japan 66:1-96:1		Germany 71:1-94:4		France 71:4-97:1	
Levels							
Prices	1	1	1	1	1	1	1
Unit Labour Costs: δ	- 0.937	- 0.767	-1	- 1.279	-1	- 1.030	1
Import Prices: $1-\delta$	- 0.063	- 0.233		0.279		0.030	
'Standard Errors' for *ulc* & *pm*	0.012	0.073		0.096		0.030	
Differences							
Δ Prices	- 0.357	0.718	- 0.243	- 0.607	- 1.839	- 0.687	- 1.378
Δ Unit Labour Costs	- 0.334	1.027	- 0.243	- 0.809	- 1.839	- 0.695	- 1.378
Δ Import Prices	- 0.699	- 0.301		- 1.534		- 0.953	
Sum of the Coefficients							
Differences of P, ULC, & PM	- 1.390	1.444	- 0.486	- 2.95	- 3.678	- 2.333	- 2.756
Test and Diagnostics							
Linear Homogeneity	0.35 [0.55]	23.58 [0.00]	23.11 [0.00]	0.01 [0.93]	2.52 [0.11]	0.23 [0.63]	0.47 [0.49]
Weight on Imports: $1-\delta=0$	9.76 [0.00]	0.40 [0.53]		2.26 [0.13]		0.43 [0.51]	
LM(1)	15.41 [0.08]	10.87 [0.28]	3.08 [0.55]	14.05 [0.12]	0.76 [0.94]	13.48 [0.14]	2.34 [0.67]
LM(4)	6.93 [0.64]	3.96 [0.91]	3.80 [0.43]	31.81 [0.00]	10.65 [0.03]	8.48 [0.49]	6.22 [0.18]
D-H(N)	5.60 [0.47]	27.10 [0.00]	10.63 [0.03]	4.19 [0.65]	5.85 [0.21]	7.49 [0.28]	2.55 [0.64]

Sample Periods	Italy 72:1-97:1	UK 61:4-97:1	Canada 62:1-97:1		Australia 67:1-97:1
Levels					
Prices	1	1	1	1	1
Unit Labour Costs: δ	- 0.717	- 0.877	- 0.922	- 1	- 0.785
Import Prices: $1-\delta$	- 0.283	- 0.123	- 0.078		- 0.215
'Standard Errors' for *ulc* & *pm*	0.064	0.024	0.038		0.051
Differences					
Δ Prices	- 2.735	- 0.690	- 1.591	- 2.219	-1.600
Δ Unit Labour Costs	- 2.840	- 0.658	- 1.572	- 2.219	- 1.364
Δ Import Prices	- 2.468	- 0.915	- 1.817		- 2.463
Sum of the Coefficients					
Differences of P, ULC, & PM	- 8.043	- 2.263	- 4.980	- 4.538	- 5.427
Test and Diagnostics					
Linear Homogeneity	7.27 [0.01]	6.49 [0.01]	1.11 [0.29]	1.23 [0.27]	4.22 [0.04]
Weight on Imports: $1-\delta=0$	10.48 [0.00]	6.13 [0.01]	2.43 [0.12]		14.75 [0.00]
LM(1)	6.19 [0.72]	16.94 [0.05]	16.98 [0.05]	4.40 [0.40]	20.51 [0.02]
LM(4)	16.15 [0.06]	10.33 [0.32]	13.33 [0.15]	4.34 [0.36]	11.73 [0.23]
D-H(N)	3.87 [0.69]	7.32 [0.29]	3.98 [0.68]	7.41 [0.12]	4.77 [0.57]

Notes: Figures reported in [ ] are probability values. LM(1) and LM(4) are Lagrange multiplier tests of autocorrelation of order 1 and 4 respectively. D-H(N) are Doornik-Hansen test for normal errors. Reported as tests of linear homogeneity and zero weight on coefficient are likelihood ratio tests distributed as χ_1^2.

Table 3: I(1) System Adjustment Coefficients and Error Correction Terms

Dependent Variable	*'Markup' Equation* $\Delta mulc$	*'Real Exchange Rate' Equation* Δrer	*Inflation Equation* $\Delta^2 p$	*Error Correction Term*
United States	- 0.298 (- 5.7)	- 0.182 (- 1.2)	- 0.061 (- 2.0)	$mulc_t + 0.059\, rer_t + 1.960\,\Delta p_t$
Germany	- 0.116 (- 4.7)		- 0.017 (- 1.4)	$mulc_t + 4.748\,\Delta p_t$
France	- 0.194 (- 4.9)		- 0.092 (- 3.7)	$mulc_t + 2.672\,\Delta p_t$
Italy	- 0.039 (- 2.7)	- 0.079 (- 2.3)	- 0.030 (- 5.1)	$mulc_t + 0.459\, rer_t + 11.926\,\Delta p_t$
United Kingdom	- 0.278 (- 6.4)	0.009 (0.1)	- 0.080 (- 3.2)	$mulc_t + 0.139\, rer_t + 2.874\,\Delta p_t$
Canada	- 0.085 (- 3.0)		- 0.068 (- 4.6)	$mulc_t + 4.318\,\Delta p_t$
Australia	- 0.189 (- 4.0)	0.125 (1.5)	- 0.041 (- 2.0)	$mulc_t + 0.166\, rer_t + 6.276\,\Delta p_t$

Notes: Reported in brackets are *t*-statistics.

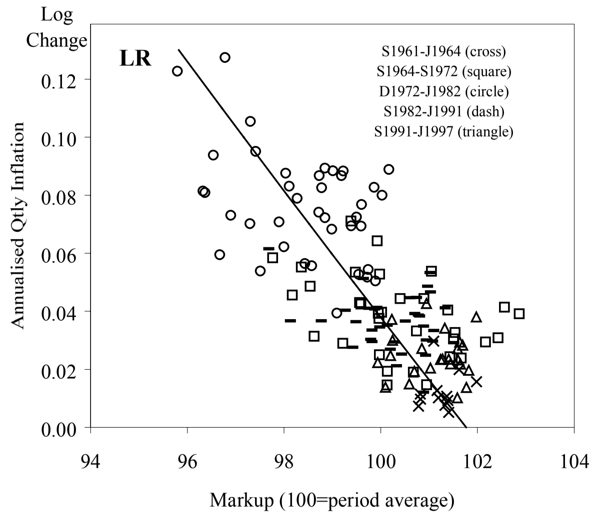
Table 4: I(1) and I(2) Estimates of the Markup and the Inflation Cost Coefficient λ

	Analysis	*Prices*	*Unit Labour Costs*	*Import Prices*	*Inflation Cost Coefficient* λ	*Long-run Effect on the Markup of a 1 Percentage Point Increase in* Δp *
United States	I(1)	1	- 0.944	- 0.056	- 1.851	0.5
	I(2)	1	- 0.937	- 0.063	- 1.390	0.3
Germany	I(1)	1	- 1		- 4.748	1.2
	I(2)	1	- 1		- 3.678	0.9
France	I(1)	1	- 1		- 2.672	0.7
	I(2)	1	- 1		- 2.756	0.7
Italy	I(1)	1	- 0.685	- 0.315	- 8.174	2.0
	I(2)	1	- 0.717	- 0.283	- 8.043	2.0
United Kingdom	I(1)	1	- 0.878	- 0.122	- 2.523	0.6
	I(2)	1	- 0.877	- 0.123	- 2.263	0.6
Canada	I(1)	1	- 1		- 4.318	1.1
	I(2)	1	- 1		- 4.538	1.1
Australia	I(1)	1	- 0.858	- 0.142	- 5.383	1.3
	I(2)	1	- 0.785	- 0.215	- 5.427	1.4

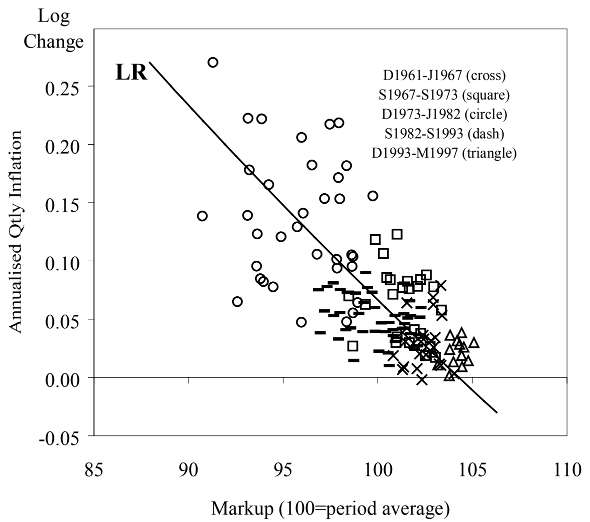
* A percentage point increase in annual inflation is equivalent to an increase in Δp of 0.25 per quarter.

Graph 1: Periods of Inflation and the Markup

UNITED STATES
September 1961 - June 1997



UNITED KINGDOM
December 1961 - March 1997



APPENDIX A: DATA SOURCES AND TRANSFORMATIONS

The data are quarterly and drawn from the June 1997 *OECD Statistical Compendium*. The table below reports the identification codes of the series used in the estimation of the models.

Data Codes for the OECD Statistical Compendium

Series	*United States*	*Japan*	*Germany*	*France*
Current Price GDP	421008SC	461008SC	131008SC	141008SC
Constant Price GDP	421108SR	461108SR	131108SR	141108SR
Indirect Taxes less Subsidies	421304SC	461304OC*	131304OC*	141304SC
Private Consumption Deflator	421201SK	461201SP	131201SP	141201SP
Total Labour Compensation	421301SC	461301OC*	131301OC*	141301SC
Standardised Unemployment Rate	4242889J	464286A3	134280A2	144286A3(2)
Imports of Goods and Services Deflator	421205SK	461205SP	Derived(1)	141205SP

Series	*Italy*	*United Kingdom*	*Canada*	*Australia*
Current Price GDP	Series 29(5)	261008SC	441008SC	541008SC
Constant Price GDP	Series 29(5)	261108SL	441108SL	541108S1
Indirect Taxes less Subsidies	Series 28(5)	261304SC	441304SC	541304SC
Private Consumption Deflator	161201SP	261201SP	141201SP	541201S2
Total Labour Compensation	161301SM	261301SC	141301SC	541301SC
Standardised Unemployment Rate	164286A3	UKOCSUN%E(3)	144286A3	544286A3(4)
Imports of Goods and Services Deflator	161205SP	261205SP	141205SP	541205S2

* Not seasonally adjusted.

(1) Derived from 131006SC and 131106SR (current price and constant price imports of goods and services respectively).

(2) Prior to March 1982 use 144295A3.

(3) Prior to March 1975 use UKOCUNE%E plus 0.954839.

(4) Prior to March 1978 use 544295A3.

(5) Italian data from www.bbs.istat and *Conti economici nazionali trimestroli* 70.1-97.4 (03/98). Constant price data from C3VAGKD, current price data from C3VAGLD.

Notes: The following transformations of the data were performed.

(a) Unit labour costs = total labour compensation divided by constant price gross domestic product (GDP).

(b) The private consumption implicit price deflator at 'factor cost' is calculated as: $P = P_{MP}/(1+tax)$ where P_{MP} is the consumption implicit price deflator at market prices and tax is the proportion of indirect tax less subsidies in current price GDP.

(c) Total labour compensation and indirect taxes less subsidies for Japan and Germany were seasonally adjusted by exponential smoothing using ESMOOTH in RATS.

APPENDIX B: ESTIMATING THE I(2) SYSTEM

B.1 The Predetermined Variables

United States: 2 lags of the first difference of the log of the unemployment rate, a step dummy up to and including March 1968 and not restricted in the cointegrating space and dummies for: June 1972, June 1973, March 1974, March 1982, and March 1991.

Japan: For $n=3$ and $n=2$. 3 lags of the first difference of the log of the unemployment rate and dummies for: March 1974, March 1975, June 1975.

Germany: For $n=3$. 3 lags of the first difference of the log of the unemployment rate, and dummies for: March 1974, June 1974, September 1974, December 1974, June 1979, September 1986, March 1993. For $p=2$. 1 lag of first difference of the log of the unemployment rate, and dummies for: December 1973, December 1974, June 1980, September 1986, and March 1993.

France: For $n=3$ and $n=2$. 2 lags of the first difference of the log of the unemployment rate, a step dummy up to and including March 1975 and not restricted in the cointegrating space and dummies for: March 1974, December 1977, and September 1982.

Italy: 2 lags of the first difference of the log of the unemployment rate, and dummies for: September 1972, March 1976, June 1976, December 1979, December 1984, December 1994.

United Kingdom: 2 lags of the first difference of the log of the unemployment rate, and dummies for: March 1974, March 1975, December 1975, March 1978, September 1979, and September 1980.

Canada: For $n=3$ and $n=2$. 3 lags of the first difference of the log of the unemployment rate, a step dummy up to and including March 1970 not restricted in the cointegrating space and dummies for: September 1974, December 1976, December 1990, December 1991.

Australia: 3 lags of the first difference of the log of the unemployment rate, and dummies for: June 1973, September 1973, June 1974, September 1974, December 1975, March 1977, March 1982, September 1982, June 1985, and September 1986.

B.2 Roots of the Companion Matrix

Modulus of the Roots of the Companion Matrix
(First 5 Values Reported, $r=1$ and Linear Homogeneity Not Imposed)

	n	1	2	3	4	5
United States	**3**	1.0000	1.0000	0.9006	0.7704	0.6092
Japan	**3**	1.0000	1.0000	0.9833	0.6800	0.6800
	2	1.0000	0.9871	0.6070	0.6070	0.5378
Germany	**3**	1.0538	1.0000	1.0000	0.7864	0.7864
	2	1.0000	0.8590	0.7910	0.7910	0.5295
France	**3**	1.0000	1.0000	0.9936	0.6797	0.6797
	2	1.0064	1.0000	0.6966	0.5650	0.5650
Italy	**3**	1.0071	1.0000	1.0000	0.7144	0.7144
United Kingdom	**3**	1.0000	1.0000	0.9502	0.6839	0.6839
Canada	**3**	1.0000	1.0000	0.9881	0.8161	0.6943
	2	1.0000	0.9836	0.7834	0.5403	0.5403
Australia	**3**	1.0000	1.0000	0.9417	0.6533	0.4837

APPENDIX C: RESULTS OF THE I(1) ANALYSIS

Table C1: Testing for the Number of Cointegrating Vectors
Estimated Values of Q(r)

United States

$H_0: r =$	**Eigenvalues**	***Q(r)***
0	0.1982	47.65 {26.70}
1	0.0805	15.83 {13.31}
2	0.0257	3.75 {2.71}

United Kingdom

$H_0: r =$	**Eigenvalues**	***Q(r)***
0	0.2474	44.81 {26.70}
1	0.0287	4.44 {13.31}
2	0.0021	0.30 {2.71}

Germany

$H_0: r =$	**Eigenvalues**	***Q(r)***
0	0.1833	21.24 {13.31}
1	0.0142	1.40 {2.71}

France

$H_0: r =$	**Eigenvalues**	***Q(r)***
0	0.2203	25.47 {13.31}
1	0.0009	0.09 {2.71}

Italy

$H_0: r =$	**Eigenvalues**	***Q(r)***
0	0.3472	47.76 {26.70}
1	0.0454	4.69 {13.31}
2	0.0000	0.00 {2.71}

Australia

$H_0: r =$	**Eigenvalues**	***Q(r)***
0	0.1647	32.55 {26.70}
1	0.0728	10.77 {13.31}
2	0.0333	1.62 {2.71}

Canada

$H_0: r =$	**Eigenvalues**	***Q(r)***
0	0.1474	22.65 {13.31}
1	0.0011	0.16 {2.71}

Notes: Statistics are computed with 4 lags of the core variables. *Q(r)* is the likelihood ratio statistic for determining *r* in the I(1) analysis. 90 percent critical values shown in curly brackets { } are from Table 15.3 of Johansen (1995b).

Table C2: Modulus of the Roots of the Companion Matrix
(First 5 Values Reported, $r = 1$ imposed)

	n	1	2	*3*	4	5
United States	**3**	1.0000	1.0000	0.6981	0.6981	0.6622
Germany	**2**	1.0000	0.7560	0.7560	0.6673	0.6673
France	**2**	1.0000	0.6718	0.5052	0.4896	0.4896
Italy	**3**	1.0000	1.0000	0.7152	0.7152	0.7089
United Kingdom	**3**	1.0000	1.0000	0.6622	0.6622	0.6132
Canada	**2**	1.0000	0.7885	0.5696	0.5696	0.5696
Australia	**3**	1.0000	1.0000	0.7961	0.7878	0.7878

Table C3: I(1) System Analysis: The United States
September 1961 – June 1997

Dependent Variable	*'Markup' Equation* $\Delta mulc$	*'Real Exchange Rate' Equation* Δrer	*'Inflation' Equation* $\Delta^2 p$
<u>*Loading Matrix* α</u>			
Error Correction Term	- 0.298 (- 5.7)	- 0.182 (- 1.2)	- 0.061 (- 2.2)
Constant	- 1.553 (- 5.7)	- 0.948 (- 1.2)	- 0.316 (- 2.2)
R^2	0.39	0.64	0.52

Notes: Number of observations: 144. Lags in the core variables = 4. Reported in brackets are *t*-statistics. The ECM is calculated: $ECM_t = mulc_t + 0.059\, rer_t + 1.960\, \Delta p_t$. Implicit markup: $mu_t = p_t - 0.944\, ulc_t - 0.056\, pm_t$. Predetermined variables are 1 lag of log unemployment, a step dummy up to June 1968 not in the cointegrating space and dummies for: June 1972, March 1974, June 1978, March 1982, and March 1991.

Tests for Serial Correlation

LM(1) χ^2 (9) = 12.59, prob-value = 0.18 LM(4) χ^2 (9) = 4.52, prob-value = 0.87

Test for Normality

Doornik-Hansen Test for normality: χ^2 (6) = 10.15, prob-value = 0.12

Table C3: I(1) System Analysis: Germany
September 1970 – December 1994

Dependent Variable	*'Markup' Equation* $\Delta mulc$	*'Inflation' Equation* $\Delta^2 p$
<u>*Loading Matrix*</u> α		
Error Correction Term	- 0.116	- 0.017
	(- 4.7)	(- 1.4)
Constant	- 0.572	- 0.084
	(- 4.7)	(- 1.4)
R^2	0.49	0.41

Notes: Number of observations: 98. Lags in the core variables = 4. Reported in brackets are *t*-statistics. The ECM is calculated: $ECM_t = mulc_t + 4.748\,\Delta p_t$. Markup: $mulc_t = p_t - ulc_t$. Predetermined variables are 1 lag of log unemployment and dummies for: December 1973, December 1974, June 1980, September 1986 and March 1993.

Tests for Serial Correlation

LM(1) χ^2 (4) = 1.65, prob-value = 0.80 LM(4) χ^2(4) = 7.70, prob-value = 0.10

Test for Normality

Doornik-Hansen Test for normality: χ^2(4) = 6.99, prob-value = 0.14

Table C3: I(1) System Analysis: France
December 1971 – March 1997

Dependent Variable	*'Markup' Equation* $\Delta mulc$	*'Inflation' Equation* $\Delta^2 p$
<u>*Loading Matrix*</u> α		
Error Correction Term	- 0.194	- 0.092
	(- 4.9)	(- 3.7)
Constant	- 0.864	- 0.407
	(- 4.9)	(- 3.7)
R^2	0.49	0.71

Notes: Number of observations: 102. Lags in the core variables = 4. Reported in brackets are *t*-statistics. The ECM is calculated: $ECM_t = mulc_t + 2.672\,\Delta p_t$. Markup: $mulc_t = p_t - ulc_t$. Predetermined variables are 2 lags of log unemployment, a step dummy up to June 1975 not in the cointegrating space, and dummies for: March 1973, March 1974, December 1977, September 1979, and September 1982.

Tests for Serial Correlation

LM(1) χ^2 (4) = 0.93, prob-value = 0.92 LM(4) χ^2(4) = 5.76, prob-value = 0.22

Test for Normality

Doornik-Hansen Test for normality: χ^2(4) = 1.24, prob-value = 0.87

Table C3: I(1) System Analysis: Italy
March 1972 – March 1997

Dependent Variable	*'Markup' Equation* $\Delta mulc$	*'Real Exchange Rate' Equation* Δrer	*'Inflation' Equation* $\Delta^2 p$
<u>*Loading Matrix*</u> α			
Error Correction Term	- 0.039	- 0.079	- 0.030
	(- 2.7)	(- 2.3)	(- 5.1)
Constant	- 0.193	- 0.386	- 0.150
	(- 2.7)	(- 2.3)	(- 5.1)
R^2	0.40	0.56	0.60

Notes: Number of observations: 101. Lags in the core variables = 4. Reported in brackets are *t*-statistics. The ECM is calculated: $ECM_t = mulc_t + 0.459\,rer_t + 11.926\,\Delta p_t$. Implicit markup: $mu_t = p_t - 0.685\,ulc_t - 0.315\,pm_t$. Predetermined variables are 3 lags of log unemployment and dummies for: September 1972, March 1976, June 1976, December 1979, December 1984, and December 1992.

Tests for Serial Correlation

LM(1) χ^2 (9) = 5.53, prob-value = 0.79 LM(4) χ^2(9) = 12.08, prob-value = 0.21

Test for Normality

Doornik-Hansen Test for normality: χ^2(6) = 1.56, prob-value = 0.96

Table C3: I(1) System Analysis: The United Kingdom
December 1961 – March 1997

Dependent Variable	*'Markup' Equation* $\Delta mulc$	*'Real Exchange Rate' Equation* Δrer	*'Inflation' Equation* $\Delta^2 p$
<u>*Loading Matrix*</u> α			
Error Correction Term	- 0.278	0.009	- 0.080
	(- 6.4)	(0.1)	(- 3.2)
Constant	- 1.479	0.047	- 0..424
	(- 6.4)	(0.1)	(- 3.2)
R^2	0.39	0.37	0.70

Notes: Number of observations: 142. Lags in the core variables = 4. Reported in brackets are *t*-statistics. The ECM is calculated: $ECM_t = mulc_t + 0.139\,rer_t + 2.874\,\Delta p_t$. Implicit markup: $mu_t = p_t - 0.878\,ulc_t - 0.122\,pm_t$. Predetermined variables are 2 lags of log unemployment and dummies for: March 1974, March 1975, December 1975, March 1978, September 1979, and September 1980.

Tests for Serial Correlation

LM(1) χ^2 (9) = 14.66, prob-value = 0.10 LM(4) χ^2(9) = 9.80, prob-value = 0.37

Test for Normality

Doornik-Hansen Test for normality: χ^2(6) = 8.92, prob-value = 0.18

Table C3: I(1) System Analysis: Canada
March 1962 – March 1997

Dependent Variable	*'Markup' Equation* $\Delta mulc$	*'Inflation' Equation* $\Delta^2 p$
<u>*Loading Matrix*</u> α		
Error Correction Term	- 0.085 (- 3.0)	- 0.068 (- 4.6)
Constant	- 0.426 (- 3.0)	- 0.342 (- 4.6)
R^2	0.35	0.57

Notes: Number of observations: 141. Lags in the core variables = 4. Reported in brackets are *t*-statistics. The ECM is calculated: $ECM_t = mulc_t + 4.318\,\Delta p_t$. Markup: $mu_t = p_t - ulc_t$. Predetermined variables are 3 lags of log unemployment, a step dummy up to March 1970 and dummies for: September 1974, December 1976, December 1990, December 1991.

Tests for Serial Correlation

LM(1) χ^2 (9) = 2.78, prob-value = 0.60   LM(4) χ^2(9) = 2.00, prob-value = 0.74

Test for Normality

Doornik-Hansen Test for normality: χ^2(6) = 12.74, prob-value = 0.01

Table C3: I(1) System Analysis: Australia
March 1967 – March 1997

Dependent Variable	*'Markup' Equation* $\Delta mulc$	*'Real Exchange Rate' Equation* Δrer	*'Inflation' Equation* $\Delta^2 p$
<u>*Loading Matrix*</u> α			
Error Correction Term	- 0.189 (- 4.0)	0.125 (1.5)	- 0.041 (- 2.0)
Constant	- 1.001 (- 4.0)	0.665 (1.5)	- 0.215 (- 2.0)
R^2	0.40	0.40	0.52

Notes: Number of observations: 121. Lags in the core variables = 4. Reported in brackets are *t*-statistics. The ECM is calculated: $ECM_t = mulc_t + 0.166\,rer_t + 6.276\,\Delta p_t$. Implicit markup: $mu_t = p_t - 0.858\,ulc_t - 0.142\,pm_t$. Predetermined variables are 3 lags of log unemployment and dummies for: June 1974, March 1982, June 1985, and September 1986.

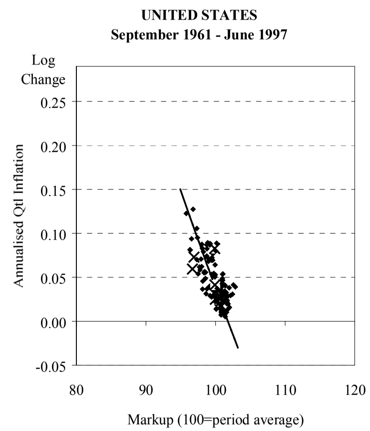
Tests for Serial Correlation

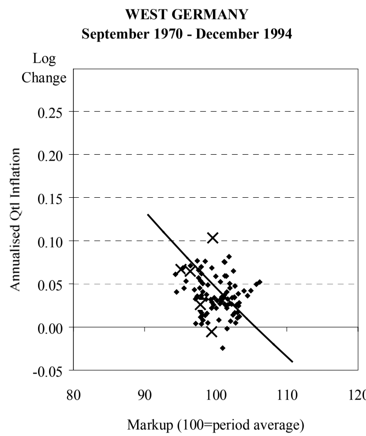
LM(1) χ^2 (9) = 15.63, prob-value = 0.08   LM(4) χ^2(9) = 7.47, prob-value = 0.59

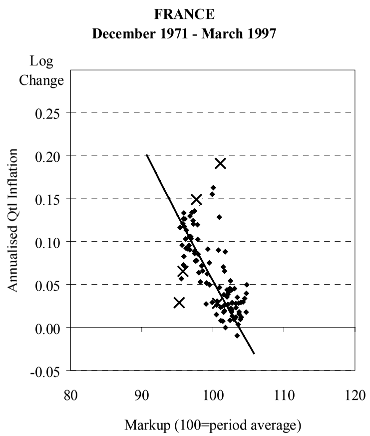
Test for Normality

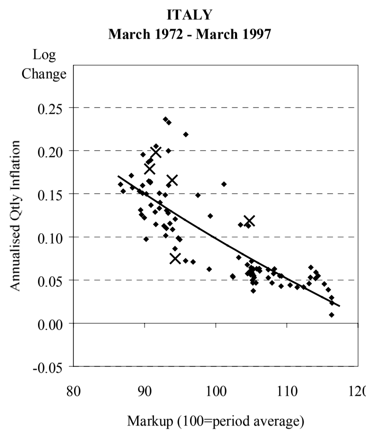
Doornik-Hansen Test for normality: χ^2(6) = 9.75, prob-value = 0.14

APPENDIX D: I(1) ANALYSIS OF THE MARKUP AND INFLATION



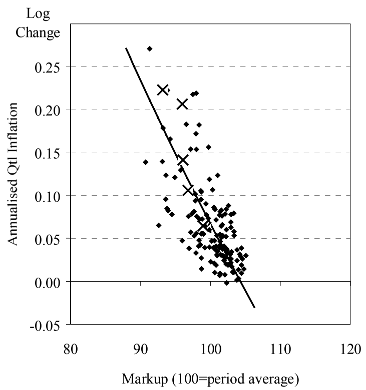




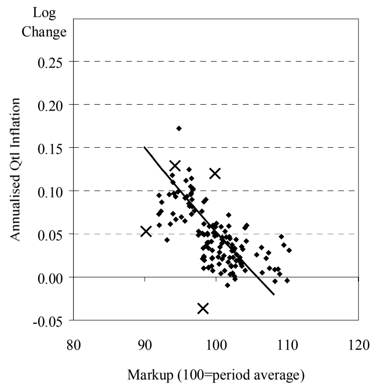


The solid line shows the estimated cointegrating relationship from the I(1) analysis between the markup and price inflation assuming the change in unemployment, spike dummies and the differences of the core variables and their lags are zero. Shown as dots are the realisations of quarterly inflation and the estimated markup from the I(1) analysis. The step dummies for the United States, France and Canada 'adjust' the markup for the respective periods. The crosses indicate the observations that correspond to the spike dummies.

UNITED KINGDOM
December 1961 - March 1997



CANADA
March 1962 - March 1997



AUSTRALIA
March 1967 - March 1997