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TESTING FOR A TIME-VARYING PRICE-COST MARKUP IN THE EURO AREA INFLATION PROCESS

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Testing for a time-varying price-cost markup in the Euro area inflation process*

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Abstract

Empirical models of inflation often incorporate equilibrium correction effects based upon levels of prices and input costs. Such models assume that the steady-state price-cost markup is constant, but recent research suggests that this may not be true for the Euro area economy, which has undergone major structural reforms over the last 25 years. We allow for permanent shifts in the markup factor through estimating an inflation equation that includes a time-varying intercept. The model suggests that a reduction in the markup contributed to disinflation in the Euro area during the period 1981-2000.

Keywords: *inflation, price-cost markup, cointegration, time-varying intercept, dynamic modelling.*

JEL classification: *C22,C32,E31*

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1 Introduction

The empirical analysis of inflation dynamics often makes use of equilibrium correction models. These models treat the target price level as a markup on some combination of input costs, and fluctuations in inflation are then interpreted as partial adjustment of the price level towards that target, i.e. inflation is an equilibrium correcting process. This treatment of price adjustment is central to the modern literature on inflation, see de Brouwer and Ericsson (1998), Aron and Muellbauer (2000), Banerjee and Russell (2001), Bank of England (1999), and Hendry (2001). In the empirical analysis of the Euro area economy, markup models of inflation have been used in the Area Wide Model (AWM) developed by Fagan et al. (2001), see Jansen (2004) for a comparison of the AWM and some competing models of Euro area inflation.

An important (but rarely challenged) assumption in this literature is that the percentage markup of prices over costs targeted by firms is fixed. A number of recent studies suggest that this is unlikely to be true in the case of the Euro area economy. The reforms implemented as part of the single market programme may have reduced the price-setting power of firms through stimulating product market competition, see Blanchard and Philippon (2003), whilst theories of pricing decisions in the presence of search costs suggest that markups could have fallen because of an increase in macroeconomic stability, see Benabou (1992) and Ball and Romer (2003). A reduction in inflation that occurs because of a decline in the price-cost markup is not identified in a linear equilibrium correction model. This not only means that the importance of such a channel cannot be assessed, but also that the estimation of the model may be biased.

One response to the problems posed by time variation in the price-cost markup is to generalise the equilibrium correction equation to include a time-varying intercept that is estimated jointly with the coefficients of the model. If the estimation does not detect any time variation in the intercept then there is no evidence that breaks in the price-cost markup have affected inflation, and the linear specification is adequate. However, if the time-varying intercept turns out to be statistically significant then it follows that the equilibrium price level has changed independently of input costs, and that the target markup factor has shifted. The behaviour of the estimated intercept term would then cast light on the timing and magnitude of these shifts, and a comparison of this model with one that imposes a constant intercept would indicate the implications, in terms of parameter estimation, of erroneously assuming that the long-run markup factor is fixed. We follow this approach in modelling Euro area inflation over the period $1981q1 - 2000q4$. The time-varying intercept that we obtain is statistically significant, and its evolution is consistent with the hypothesis that a reduction in the equilibrium markup factor has restricted inflation, particularly during the first half of the $1980s$.

The remainder of the paper expands upon these points and is structured as follows. In section 2 we discuss markup models of the price level and consider their relevance to the Euro area economy of the past 25 years. Section 3 presents the data used for the empirical analysis. Section 4 reports a model for the Euro area price level, and two separate models for inflation, one that assumes a fixed long-run markup factor, and one that is consistent with permanent shifts in the markup. The results are used to evaluate the importance of different channels for inflation adjustment, and to assess the implications of wrongly assuming a constant equilibrium markup factor. Finally, section 5 rounds off the paper with a summary of the main arguments.

2 Markup models of inflation

The standard approach to modelling inflation treats the equilibrium price level as a fixed markup on input costs, see, for instance, de Brouwer and Ericsson (1998). A plausible log-linear representation of the equilibrium price level in the Euro area is the following:

$$p_t = \log \Psi + \kappa ulc_t + \gamma pcom_t + \lambda tax_t + \varphi trend \tag{1}$$

where all variables are in natural log form, and p represents the price level, ulc is the level of unit labour costs (wages and salaries paid per unit of production), $pcom$ is the domestic currency price of oil and other raw materials, tax is the percentage tax wedge and $trend$ is a time trend. Each of the time series referred to is in index form and the base period is common across series. This expression for the price level can be derived as the profit maximising outcome when the underlying production technology is Cobb-Douglas. We follow Hendry (2001) in adding a time trend to the model in order to control for elements of total costs that may have been omitted from the information set, e.g. the compensation of the self-employed, or imports that take the form of finished consumer goods rather than raw materials.[1]

A dynamic version of (1) can be reparameterised to give the equilibrium correction model in (2), see Hendry and Doornik (2001) for further details. In equation (2) it is assumed that the price level is linearly homogeneous in unit labour costs and raw material prices, i.e. $\kappa + \gamma = 1$, and that there is complete pass-through from the tax wedge to the price level, i.e. $\lambda = 1$, so that the equilibrium correction term can be written as the sum of the two relative prices, $(ulc - p + tax + \varphi trend + \log \Psi_1)$ and $(pcom - p + tax + \varphi trend + \log \Psi_2)$, where $\log \Psi = \log \Psi_1 + \log \Psi_2$. Also, we allow for short-run variation in the percentage price-cost markup by conditioning on the log ratio of actual GDP to permanent, or trend, GDP, a variable that we denote gap. It is important to note that this term reverts towards an equilibrium value of zero, and therefore cannot induce changes in the markup factor in the long run.[2] In equation (2) the parameters (κ^*, γ^*) are equal to (κ, γ) in (1) multiplied by α, the speed at which the price level converges on its equilibrium value. The inflation equation contains a large number of parameters. In section 4 we describe

[1] The Euro area import price index is not included in (1) because it is an aggregate of national import price indices and therefore includes intra-area trade in addition to trade with countries outside the Euro area. Note that the commodity price index will account for an important component of Euro area imports from the rest of the world, which means that equation (1) should still represent a valid model for the price level.

[2] In the New Keynesian literature the output gap and real marginal unit labour costs are viewed as *alternative* control variables, see e.g. Galí and Gertler (1999). The idea is that real marginal unit labour costs behave cyclically, e.g. because overtime hours have to be paid at a premium rate during economic expansions, such that there is no need to control for the output gap. Equation (2) includes real average unit labour costs, which do not exhibit the same cyclicality as real marginal unit labour costs. Consequently, there is still a role for the output gap in this equation.

the general-to-specific modelling strategy that is used to reduce the equation to a more parsimonious form.

$$\begin{aligned}\Delta p_t \quad = \quad & \sum_{q=0}^{3} \varsigma_q gap_{t-q} + \sum_{i=1}^{3} \theta_i \Delta p_{t-i} + \sum_{j=0}^{3} \kappa_j \Delta ulc_{t-j} + \sum_{l=0}^{3} \gamma_l \Delta pcom_{t-l} \\ & \sum_{p=0}^{3} \lambda_p \Delta tax_{t-p} + \kappa^* (ulc - p + tax + \varphi trend + \log \Psi_1)_{t-1} \\ & + \gamma^* (pcom - p + tax + \varphi trend + \log \Psi_2)_{t-1}\end{aligned} \qquad (2)$$

The equilibrium correction model has a clear time series interpretation. The levels of price and cost variables are considered to be integrated of order one, I(1), and cointegrate such that the linear combination of them included in (2) is I(0). This term, along with the output gap and log differences in prices and costs, which are also stationary, explain the inflation rate, which is I(0) under the assumption that the price level is I(1). In section 4 of the paper, we show that these assertions concerning integration and cointegration in the data are consistent with the outcomes of formal econometric tests.

The model in (1) and (2) assumes that the markup factor that firms target in equilibrium, Ψ, is constant, which implies that all changes in input prices relative to consumer prices must induce inflation adjustment. A number of recent papers suggest that this assumption may be inappropriate, particularly in the case of the Euro area economy in the period since 1980. Rogoff (2003) suggests that greater product market competition, driven by increased international trade, privatisation of state controlled industries and deregulation of goods markets, has decreased the equilibrium price-cost markup. Such an argument seems particularly relevant to the Euro area over the past 25 years, which has seen the marketisation of several southern European economies in preparation for accession to the European Community, and the introduction of more rigorous competition policy as part of European Union law. Empirical support for this view is presented in Blanchard and Philippon (2003), who show that the average levels of barriers to entrepreneurship and international trade, as measured by the OECD, have fallen over the last 20 years, particularly during the 1990s. Blanchard and Phillipon contend that structural change of this sort has decreased the share of profits in total output, and seek to explain European unemployment evolutions in terms of this trend. In terms of the framework presented in this paper, the mechanism emphasised by Blanchard and Phillipon implies a smaller Ψ parameter and will therefore affect inflation.

A related body of literature concentrates on theories of price-setting and the markup at the micro level. Benabou (1992) shows that in (S, s) pricing models, lower inflation is associated with lower price dispersion, a result that is confirmed empirically by Nath (2004). If consumer search costs are relatively high so that in equilibrium agents do not always purchase from the cheapest supplier, a reduction in price dispersion will ensure that fewer consumers end up paying a price that is far in excess of the lowest price offered in the market. This implies that the average price-cost markup will decrease. Similar results are derived by Ball and Romer (2003) using a model in which buyers and sellers form long-term relationships, so that current prices provide signals concerning the direction of future prices. As inflation and price dispersion fall, current prices become more informative for future prices and consumers are therefore better able to identify the cheapest supplier, i.e. demand curves are more price elastic. This in turn restricts the size of the price-cost markup that can be sustained in equilibrium. In empirical studies, the propensity for inflation to force down steady-state profit margins has been demonstrated by Chirinko and Fazzari (2000), using U.S. data. Gwin and Taylor (2004) find evidence for the same relationship using data for a different set of U.S. industries, though the correlation is only apparent amongst industries in which search costs are relatively high. These papers suggest that the reduction in Euro area inflation secured through the channels identified in equation (2), e.g. slower growth in unit labour costs caused by labour market reforms or a tough monetary policy, may be self-reinforcing in the sense that they lead to a reduction in the Ψ parameter in (2), which then leads to further reductions in inflation.[3] It is also important to note that the inception of the European Monetary System (EMS) in 1979 may have restricted price volatility, both within and across Euro area countries, and that this may have reinforced the downward pressure on equilibrium markups via the mechanisms emphasised by Benabou (1992) and Ball and Romer (2003).

A permanent reduction in the price-cost markup implies that (2) is not a valid model for inflation. To see this, note that a decrease in the target price-cost markup should cause the two relative price terms to fall (through decreases in Ψ_1 and Ψ_2) so that inflation decreases for some time and brings the price level into line with its new target level. If these two parameters are held fixed, however, then the reduction in inflation will be left unexplained, which will manifest itself either as departures from white noise in the residuals, or as biases in the coefficients fitted to the model.

It follows that some generalisation of the linear inflation equation is necessary. One option would be to construct a measure of the aggregate profit share and then condition on that variable in the empirical analysis. However, obtaining the relevant data for all Euro area countries is likely to prove a particularly demanding exercise. Furthermore, it is not clear how one might extract the underlying target markup factor from the observed price-cost markup (or profit share), which will be dominated by relative price disequilibrium. An alternative approach would be to condition on dummy variables that allow for a step shift in the relationship between inflation and the markup. However, the precise timing of the shifts is not known and can be difficult to test for using Chow-type tests if the sample size is small. Further, if changes in competitive forces diffuse through the economy at a gradual pace then shifts in the relationship between inflation and the markup are not best modelled in this way. Instead, we adopt an approach related to that in Aron and Muellbauer (2000), which involves augmenting (2) with a time-varying intercept, or local level, which we denote $\psi\mu_t$.[4]

The equilibrium correction model extended to include a local level term, and

[3] A reduction in the markup factor can induce permanent reductions in the price level but not the inflation rate. However, if adjustment towards the new target price level is sluggish then even transitory changes in the inflation rate may appear persistent. Furthermore, if structural change is a gradual process, or if its effects filter through to price-setting very gradually, then it may appear that inflation changes permanently.

[4] In the remainder of the paper the terms 'time-varying intercept' and 'local level' are used interchangeably.

also the error term that was not included in previous specifications, is as follows:

$$
\begin{aligned}
\Delta p_t \quad = \quad & \sum_{q=0}^{3} \varsigma_q gap_{t-q} + \sum_{i=1}^{3} \theta_i \Delta p_{t-i} + \sum_{j=0}^{3} \kappa_j \Delta ulc_{t-j} \\
& + \sum_{l=0}^{3} \gamma_l \Delta pcom_{t-l} + \sum_{p=0}^{3} \lambda_p \Delta tax_{t-p} \\
& + \alpha * \kappa (ulc - p + tax + \varphi trend)_{t-1} \\
& + \alpha * \gamma (pcom - p + tax + \varphi trend)_{t-1} + \psi \mu_t + u_t
\end{aligned} \tag{3}
$$

$$
\begin{aligned}
& (\kappa + \gamma = 1) \\
& u_t \sim (0, \sigma_u^2)
\end{aligned}
$$

$$
\mu_t = \mu_{t-1} + \varepsilon_t, \quad \varepsilon_t \sim (0, \sigma_\varepsilon^2) \tag{4}
$$

The local level component of the *measurement equation* in (3), μ_t, is constructed as an I(1) process using the STAMP package of Koopman et al. (2000). Its evolution is determined by the variance of the shocks that enter the *transition equation* in (4). The local level measures the variation in inflation that cannot be explained by deviations of the price level from a long-run target based upon a constant markup of prices over costs, or by the output gap and other dynamic adjustment terms. It therefore captures the movements in inflation attributable to time variation in the markup factor. If structural changes have not contributed to price adjustment in the Euro area, and the price-cost markup has in fact been constant over the past quarter of a century, then σ_ε^2 will be zero and μ_t will be a constant, such that (3) is equivalent to the linear equilibrium correction model. If, however, there has been time variation in the price-cost markup, then μ_t will capture movements in inflation arising from it. The model does not require any priors as to when breaks occur, and as such is a suitable technique for modelling inflation in an economic area that has been subject to structural change.

The estimation of the model in (3) and (4) proceeds in two steps. First, it is postulated that the two error terms, u_t and ε_t, are normally and independently distributed, so that maximum likelihood estimates of σ_u^2 and σ_ε^2 can be computed using numerical optimisation techniques. The coefficients of the measurement equation are then calculated using the Kalman filter, which performs the same function in the estimation of models in SSF as does least squares in the estimation of a standard regression model.

The fact that the local level term is constructed as an I(1) process raises some questions concerning the cointegrating properties of the model for inflation. As noted above, the usual presumption is that the price level is I(1) and cointegrates with I(1) input cost measures, such that the price-cost markup is I(0). This term then explains the stationary variation in Δp_t, i.e. the model is balanced. The inclusion of the local level term complicates the picture through making the right-handside of (3) non-stationary. As the inflation rate on the left-handside of the equation is I(0), the equation is unbalanced. Our interpretation of this relationship is as follows:

Inflation and the markup fluctuate in I(0) space, but can be subject to deterministic shifts that take effect gradually, e.g. following a break in the price-cost markup. In finite samples the effects of these slowly evolving changes can be modelled by a low variance I(1) process - to see this point note that in the limiting case in which σ_ε^2 is set to zero, the local level term is simply a constant and is therefore obviously stationary. Thus, while (3) cannot represent the actual data generating process when price inflation is stationary, it is a plausible model in finite samples.[5]

3 The data

The variables used in the empirical analysis are all measured at the quarterly frequency and expressed in seasonally adjusted form. Each variable is constructed as a weighted average of the corresponding national series, with the weight accorded to each country set equal to its share in constant 1995 Euro area GDP, measured at market prices using purchasing power parity (PPP) exchange rates, see Fagan et al. (2001) for further details.

The price level, p_t, is measured by the consumer expenditure deflator and ulc_t is defined as whole economy wages and salaries paid per unit of GDP at factor cost (where wages and salaries include payroll taxes but exclude the compensation of the self-employed). The variable $pcom_t$ is based upon the HWWA (Hamburgerisches Welt-Wirtschaftliches-Archiv) world commodity price index, which measures the US$ prices of 29 important raw materials at 36 market places around the world (the price of crude oil is the dominant component in the index and receives a weight of 55%). This index is then multiplied by the €-US$ exchange rate in order to give a measure of the cost of crude oil and raw material imports to the Euro area. The variable tax is the ratio of GDP at market prices to GDP at factor cost. The output gap, gap_t, measures the deviation of the natural log of GDP from the natural log of trend GDP, where the latter is obtained using the Hodrick-Prescott filter with the smoothing parameter set to 1600.

In Figure 1 we plot p, ulc, $pcom$ and tax and some relevant linear combinations of those variables (recall that all variables are log transforms of the original series). We also graph the first difference of the log consumer expenditure deflator, Δp_t, which measures the quarterly inflation rate, and the output gap variable, gap_t.

[5] A second logical possibility, which we find less plausible *a priori,* is that the price-cost markup is I(1), in which case the linear equilibrium correction model in (2) is certainly unbalanced, in that the model would entail regressing an I(0) inflation rate on I(1) relative prices. If the local level term cointegrates with relative prices so that we obtain a stationary righthandside to equation (3), a balanced regression obtains (this point has also been made, albeit in a slightly different context, by Naug (2000)). In such instances the interpretation of the local level term is that it measures the I(1) contribution to inflation arising from permanent shifts in the price-cost markup.

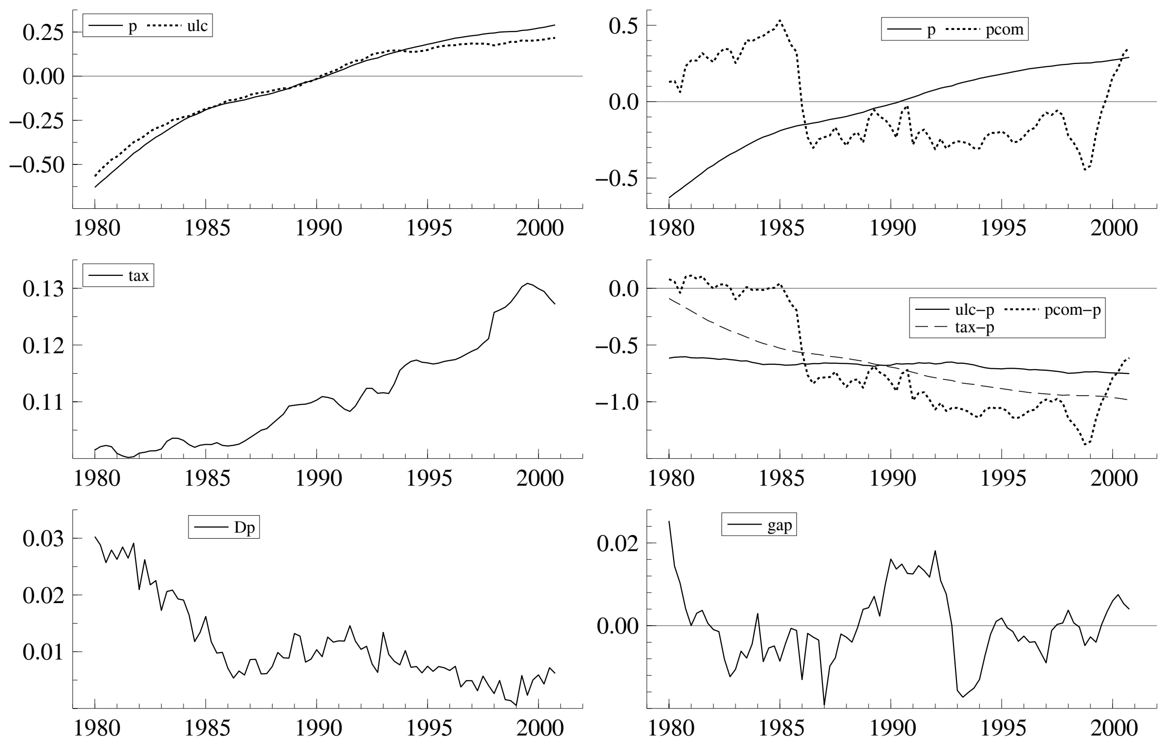

Figure 1: The data series used in the empirical analysis.

4 Empirical results

The empirical strategy that we follow can be divided into three parts. In the first part we make the standard assumption that the target markup factor is fixed and then estimate the long-run relationship between prices and costs conditional upon that assumption. In the second part of the analysis we embed this relationship in a dynamic conditional model for inflation and test down to a parsimonious form. In the third part of the analysis this parsimonious specification is generalised to include a time-varying intercept. Such a model allows for the possibility that there are systematic movements in inflation that occur independently of the disequilibrium between prices and input costs, even after controlling for the output gap. These are the movements that we attribute to shifts in the equilibrium markup factor. A comparison of the two single equation models casts some light on the importance of structural change in setting Euro area inflation.

4.1 Modelling the long run

There are two key reasons for estimating the long-run relationship between the price level and input costs. First, it is important to demonstrate that price and cost indices cointegrate from I(1) to I(0), thereby confirming the time series interpretation of those variables suggested in section 2.[6] Second, we need to check that the price level is linearly homogeneous in unit labour costs and raw materials costs, and that the tax wedge feeds into the price level with a unit coefficient. Recall that these two assumptions are necessary for the parameterisations of the dynamic inflation models given in (2) and (3). If the assumptions fail then the profit markup is not identified from the available price and cost measures, and the unobserved component estimated in (3) will not measure the variation in inflation arising from shifts in the equilibrium markup factor.

In order to do this we estimate a vector autoregression (VAR) for the vector $[p\ ulc\ pcom]'$.[7] Following the discussion in section 2, each line of the VAR also includes the non-modelled variables tax_{t-1} and gap_{t-1}, and a linear time trend. We also include a dummy variable set to unity in 1992(3) and 1992(4), and to zero otherwise. This dummy, denoted $i92q3q4$, is intended to control for any effects of the global recession that are not captured by the output gap variable, and also the events in foreign exchange markets that led to the break-up of the ERM. The VAR contains three lags in each of the endogenous variables and is estimated in the isomorphic Vector Equilibrium Correction Mechanism (VEqCM) form using data for 1980(4) to 2000(4). The estimation procedure is the Johansen (1988) cointegration technique.[8] The tax wedge is treated as I(1) and is therefore allowed to enter the cointegrating space that defines the long-run solution for the price level, while the output gap is treated as I(0) and does not enter the cointegration space.

The results obtained for the unrestricted VEqCM are used to determine the cointegrating properties of the model. The *Trace*-statistic for the hypothesis that the number of cointegrating vectors, r, is equal to zero is 62.55, while that for r less than or equal to one is 33.19 and that for r less than or equal to two is 12.78 (degrees of freedom corrected). The 5% critical values quoted in Harbo et al. (1998) for the case in which one non-modelled variable enters the cointegration space are 49.6, 30.5 and 15.2 respectively. Formally, this evidence implies that there are two cointegrating vectors in the model. However, the evidence for the second of those two relations is very marginal, as the test statistic of 33.19 is only slightly larger than the critical value of 30.5. Furthermore, the critical values used here are only likely to be approximate because of (i) the small sample size; (ii) the inclusion of *gap* outside the cointegration space, which implies that the distribution of the *Trace*-statistic is characterised by nuisance parameters that do not affect the distribution simulated by Harbo et al. (1998), see Rahbek and Mosconi (1999), and (iii) the

[6] We rule out the possibility that the price and cost variables are I(2). In defence of this assumption, we note that Hendry (2001) argues that variables which appear to be I(2) are actually I(1) processes that are subject to breaks in the mean.

[7] We also considered a VAR with four endogenous variables, the level of import prices being the fourth variable. However, the results did not yield a plausible coefficient on the import price term (see Bowdler and Jansen (2004) for further details). We suspect that this is due to an important measurement error outlined previously, namely that the import price series includes the average price of goods and services traded *between* Euro area countries in addition to imports from outside the area. The import price term was therefore dropped from the analysis. Note that the raw material price index, *pcom*, will control for many types of imports to the Euro area.

[8] All of the empirical results in Section 4.1 were obtained using *PcFiml 9.3* — see Doornik and Hendry (1997)

inclusion of a dummy variable in the unrestricted part of the model, which implies further nuisance parameters in the distribution for the *Trace*-statistic. In view of these considerations, we conclude that there exists a unique I(1) to I(0) cointegrating vector amongst the endogenous variables.

Table 1 provides various estimates of the cointegrating vector, each normalised so that there is a unit coefficient on the price level. Panel 1 refers to the just identified case in which we do not impose theory-driven restrictions. The tax wedge is positively signed and one cannot reject the hypothesis there is full pass-through to consumer prices - the imposition of this restriction (see panel 2) yields a test statistic of 0.02, which generates a p-value of 0.90 using a $\chi^2(1)$ distribution. In panel 3 the linear homogeneity of p in ulc and $pcom$ is imposed and it is clear from the test results quoted beneath panel 3 that such a restriction is compatible with the data. The relative weights in the long run equation are 0.96 and 0.04, the latter being small but clearly significant. The time trend enters the long-run equation for the price level in panel 3 with a coefficient that implies autonomous annual growth in the price level equal to 0.39%. As suggested previously, this can be interpreted as the result of some form of measurement error such as the omission of the self-employed from the calculation for unit labour costs. It is important to note that if the VAR is estimated without the time trend then the p-value for a test of the homogeneity restriction (after first imposing the unit tax wedge effect) is .03, indicating that homogeneity fails when the system does not condition on a trend. This illustrates the need for the trend term in (1).

Table 1: Estimates of the price level equation derived from the cointegrated VAR model.

Panel 1: The just identified price level equation

$$p_t = \underset{(0.06)}{0.90} ulc_t + \underset{(0.01)}{0.04} pcom_t + \underset{(1.19)}{0.83}\, tax_{t-1} + \underset{(0.0009)}{0.0014}\, trend_t$$

Panel 2: Full effect of indirect taxation

$$p_t - tax_{t-1} = \underset{(0.03)}{0.90} ulc_t + \underset{(0.01)}{0.03} pcom_t + \underset{(0.0002)}{0.0013}\, trend_t$$

$$\chi^2(1) = 0.02[0.90]$$

Panel 3: As Panel 2 and homogeneity

$$p_t - tax_{t-1} = \underset{(-)}{0.96} ulc_t + \underset{(0.01)}{0.04} pcom_t + \underset{(0.00014)}{0.00098}\, trend_t$$

$$\chi^2(2) = 1.37[0.50],\ \chi^2(1) = 1.35[0.24]$$

The sample is 1980(4) to 2000(4), 81 observations.

System mis-specification tests for the underlying VAR:

AR_v *1-5* $F(45, 140)$	1.31[0.12]
$Normality_v$ $\chi^2(6)$	14.67[0.02]**
$Heteroscedasticity_v$ $F(132, 217)$	0.70[0.99]

References:

AR-test (Godfrey (1978) and Doornik (1996)),

Normality test (Doornik and Hansen (1994)), and

Heteroscedasticity test (White (1980) and Doornik (1996)).

The residuals from the estimated cointegrating vector incorporating price-cost homogeneity and a unit coefficient on the tax wedge measure the markup of prices over costs at each point in time. Some minor non-stationarities occur in this series. Indeed, ADF(4) tests[9] for the presence of a unit root do not reject the null hypothesis, the test statistic being -2.63 and the 5% critical value -2.90 (a high level of persistence in the residuals is mirrored in the relatively large error autocorrelation statistic for the underlying VAR). This finding is at odds with the outcome of the Johansen analysis, and no doubt partly reflects the low power of the ADF procedure, see Gregory et al. (2004) for some evidence on the contradictory results that can arise from single equation and multi-equation tests for cointegration. Nevertheless, it constitutes some evidence for the view that differences between the levels of prices and costs dissipate only very slowly. This could be because the estimated VAR is affected by permanent shifts in the target markup factor. Any such shifts are clearly not large enough to induce residual autocorrelation in the VAR, but some attempt to quantify their size and timing could prove interesting.

4.2 A linear equilibrium correction model for inflation

The next stage in the empirical analysis entails formulating a conditional model for inflation, maintaining the assumption that the equilibrium markup factor is fixed. Tests for the weak exogeneity of unit labour costs and commodity prices within the VAR, i.e. tests of the hypotheses that the loadings on the cointegrating vector in the unit labour cost and commodity price equations are zero, yield the outcomes $\chi^2(1) = 2.11[0.15]$ and $\chi^2(1) = 0.25[0.62]$ respectively, and the joint hypothesis yields the outcome $\chi^2(2) = 2.14[0.34]$. Non-rejection of these hypotheses implies that switching from a system analysis to a single equation model for inflation can be done without any loss of efficiency, and we now pursue that option. Specifically, we estimate the linear inflation equation in (2) that assumes a constant markup factor. This model is then tested down to a parsimonious form using a general-to-specific modelling strategy that entails deleting the least significant term, re-estimating the model, and then repeating the procedure until each of the variables included in the regression is individually significant.

The parsimonious inflation equation is reported below, along with a series of residual diagnostic tests.[10] The references for these tests are given in Table 1, except for the *ARCH 1-4 F*-test, which is a test for autoregressive conditional heteroscedasticity, see Engle (1982), and the *RESET*-test, which is the specification test due to Ramsey (1969).

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[9]The ADF specification chosen was that with the largest number of lags and a significant coefficient on the final lag (lag orders above 12 were not considered).

[10]All of the empirical results in Section 4.2 were obtained using *PcGive10* - see Hendry and Doornik (2001).

$$
\begin{aligned}
\Delta p_t \quad = \quad & \underset{(0.01)}{0.04} + \underset{(0.07)}{0.35}\ \Delta p_{t-2} + \underset{(0.002)}{0.01}\ \Delta pcom_t + \underset{(0.027)}{0.05}\ gap_{t-1} \\
& + \underset{(0.04)}{0.13}\ \Delta ulc_t + \underset{(0.25)}{0.57}\ \Delta tax_{t-1} - \underset{(0.001)}{0.006}\ i92q3q4 \\
& + \underset{(0.02)}{0.13}\ (ulc - p + tax + \varphi trend)_{t-1} \\
& + \underset{(0.001)}{0.007}\ (pcom - p + tax + \varphi trend)_{t-1}
\end{aligned} \tag{5}
$$

$\sigma = 0.001630$ $\qquad$ $\varphi = 0.00098$

AR 1-5 $F(5, 66)$	0.94[0.46]
ARCH 1-4 $F(4, 63)$	1.04[0.39]
Normality $\chi^2(2)$	0.41[0.81]
Heteroscedasticity $F(15, 55)$	0.69[0.78]
RESET $F(1, 70)$	4.75[0.03*]

The properties of the parsimonious model are somewhat mixed. The diagnostic tests indicate that there is no evidence of heteroscedasticity or non-normality in the residuals. However, the RESET test, which evaluates whether or not squares and higher powers of the explanatory variables are correlated with the model residuals, indicates an inadequate functional form, and the autoregressive term appears to play an important role in the model. Once again, these findings could be explained by the fact that the target markup of prices over costs has not been constant over the sample period. We now investigate this possibility through estimating the local level equation described in section 2.

4.3 A local level equation for the inflation process

We apply the local level estimation technique to the tested down linear equilibrium correction model rather than the unrestricted specification in (3). The validity of the restrictions imposed in obtaining the parsimonious model are then checked by adding back in a distributed lag in each of the variables, one at a time, and testing down once more through repeated local level estimation. We opt for this strategy rather than repeat the entire general-to-specific modelling exercise because we find that highly parametrised models tend to be associated with implausibly large variation in the local level term. The strategy that we follow, namely developing the best fitting linear model and then studying the properties of a non-linear version of that restricted specification, is consistent with the approach to non-linear modelling advocated by Teräsvirta (1998).

The results are reported in Table 2.[11] Beneath each estimated equation we quote the Box-Ljung test for lack of residual serial correlation up to order 4 and the Doornik-Hansen test for residual normality, see Koopman et al. (2000). These tests refer to the *auxiliary residuals*, which are a smoothed version of the error processes

[11] All of the empirical results in section 4.3 are obtained using *STAMP 6* — see Koopman et al. (2000).

u_t and ε_t in equations (3) and (4). Further details on the interpretation of auxiliary residuals can be found in Koopman et al. (2000). The equation standard error (σ) is the square root of the variance of the model residuals (the u_t series in equation (3)), and is therefore comparable with the standard error reported for the linear inflation equation. To the right in Figure 2 we plot the time-varying intercept, which is the term $\psi\mu_t$ from equation (3) multiplied by the reciprocal of one minus the estimated autoregressive parameter in Table 2. This scaling factor is used in order to illustrate the steady-state contribution of the time-varying intercept to the inflation rate. To the left in Figure 2 the stretched time-varying intercept is plotted alongside inflation (a constant has been subtracted from the time-varying intercept in order to ensure that it has the same mean as the inflation series).

Table 2: Local Level estimates of the parsimonious inflation equation.

$$\Delta p_t = \underset{(0.01)}{0.04}LL_t + \underset{(0.08)}{0.17}\Delta p_{t-2} + \underset{(0.04)}{0.10}\,\Delta ulc_t + \underset{(0.002)}{0.01}\,\Delta pcom_t$$
$$+\underset{(0.24)}{0.49}\Delta tax_{t-1} + \underset{(0.04)}{0.09}gap_{t-1} + \underset{(0.02)}{0.14}(ulc - p + tax + \varphi trend)_{t-1}$$
$$+\underset{(.002)}{0.008}(pcom - p + tax + \varphi trend)_{t-1} - \underset{(.001)}{0.006}i92q3q4$$
$$\sigma = 0.139\%$$
$$\varphi = 0.00098$$
$$q\text{-}ratio^{*)} = 0.14$$

AR 1-4 statistic = 3.00, 95% *critical value is* 3.84 ($\chi^2(2)$)

Normality $\chi^2(2)$ 1.50[0.47]

The sample is 1981(1) to 2000(4), 80 observations.

*) The *q-ratio* is defined as the ratio $\sigma_\varepsilon^2/\sigma_u^2$, where σ_ε^2 is the residual variance of (4) and σ_u^2 is the residual variance of (3)

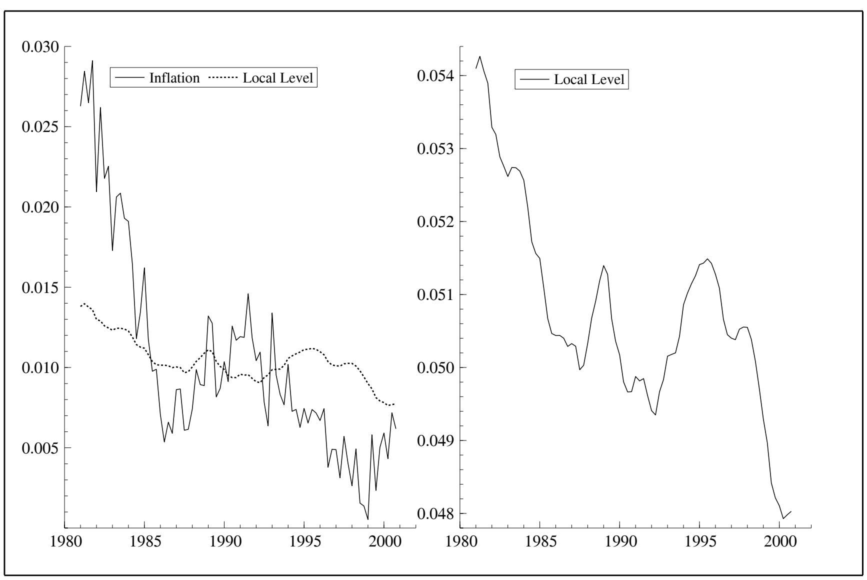

Figure 2: The first graph shows the local level term plotted against inflation and the second graph is a close-up of the local level term. In each case the local level has been scaled by the reciprocal of one minus the autoregressive parameter in Table 2..

The addition of the local level term to the model reduces the regression standard error by approximately one sixth. The t-ratio for the coefficient multiplying the local level in the second model in Table 2 is 3.4. Given that the variables are not in zero mean form, this partly reflects the significance of the intercept. When the regression is estimated with all variables in zero mean form, the absolute t-ratio is 2.34 and the corresponding p-value is 0.02. The local level term therefore makes a contribution to inflation that is significant at the conventional level.

The autoregressive coefficient, which was estimated to be 0.35 in the linear model, is only 0.165 in the local level model. One interpretation of this parameter shift, which is more than twice the standard error for the autoregressive term estimated for the linear and local level models, is that in the linear model the autoregressive coefficient is biased upwards because of the exclusion of the time-varying intercept. To the extent that lagged inflation proxies the inflation reducing effect of a fall in the equilibrium price-cost markup, the relatively large autoregressive coefficient in the linear model may explain why that specification does not fail more of the residual diagnostic tests. One implication of this is that the application of a linear equilibrium correction model to an economy that has been subject to structural change may give the impression that inflation has been more dependent on its own lags than has actually been the case.

The parsimonious local level model in Table 2 yields a q-ratio of 0.14, which means that the variance of the shocks impacting the local level term is equal to 14% of the variance of the residuals for the estimated model. An inspection of the time series plots for the local level term in Figure 2 casts some light on the role that it plays in explaining inflation. Its main feature is a downward trend over the first half of the 1980s, which accounts for approximately one fifth of the eight percentage points reduction in annual Euro area inflation during the period 1980−85. This may in fact underestimate the importance of the time-varying markup factor if some of its effects are captured by the weighted time trend derived from the VAR model. However, when we allow the trend to enter the local level model unrestrictedly (rather than with a pre-determined coefficient as part of the relative price terms), the results do not change in a meaningful way. In any case, the reduction in inflation (and hence the price level) that occurred independently of the costs of production during the early 1980s is certainly important in an economic sense. Over the remainder of the sample period the local level term behaves somewhat cyclically, suggesting few permanent changes in the markup factor. There is some evidence of a sustained decline in the markup factor during the late 1990s, but one would need to consider the way in which the local level evolved over an extended sample in order to give a firm interpretation to this movement.

The behaviour of the local level term is important because it suggests that the substantial reduction in Euro area inflation that occurred in the 1980s cannot be entirely explained in terms of well known disinflation strategies such as a deceleration of nominal wages, or an appreciation of the currency. Additionally, the price level fell independently of cost and demand conditions, and we attribute this part of the disinflation to firms reducing their claims on output through choosing a smaller percentage markup of prices over costs. Such a change in price-setting behaviour cannot be attributed, with certainty, to a single set of economic events. However, an argument that we find plausible is that a reduction in price dispersion, arising from both the disinflation of the early 1980s and the inception of the new European Monetary System (EMS) in 1979, led to stronger price competition in product markets and hence a reduction in the profit share.

The evidence for a reduction in Euro area inflation caused by a decrease in the equilibrium price-cost markup contrasts with the evidence presented in Bowman (2003), who notes that actual markups in OECD countries do not appear to have fallen during the period since 1980. There are two reasons for the apparently conflicting evidence. First, the markups considered in the latter study refer to prices relative to unit labour costs, not prices relative to unit labour costs, commodity prices and the tax wedge. Second, Bowman focuses on the actual markup of prices over costs, not the underlying equilibrium markup. If macroeconomic policy causes a large reduction in, say, unit labour costs, then the actual price-cost markup may rise even if the equilibrium, or target, markup factor is decreasing (this logic suggests that equilibrium markups can only be observed at times of constant inflation, during which price-cost disequilibrium is minimal). The results presented in Table 2 indicate some evidence of a reduction in inflation after controlling for disequilibrium in prices, indicating that there may be more support for theories of negative drift in the markup than is suggested by simple plots of prices relative to costs.

The local level model indicates that the price level is linearly homogeneous in input costs only after controlling for the effects of the time-varying intercept. This is at odds with the results from the linear VAR model, which indicated that a linear homogeneity restriction for prices and input costs could not be rejected. One interpretation of this finding is that the coefficients on input costs in the long-run equation for the price level sum to less than one, but are sufficiently close to one that linear homogeneity cannot be ruled out. This observation raises an important point: While the reduction in inflation resulting from a decrease in the target price-cost markup is statistically significant, it is not important enough to completely undermine standard linear analyses of prices and inflation.

If the local level term does account for shifts in the target markup of prices over costs, it should help to control for some of the local non-stationarities that occur in the markup when it is derived from the VAR model under the assumption that the long-run markup factor is fixed. A regression of the markup calculated from panel 3 in Table 1 on the estimated local level term should yield a residual series that is more obviously stationary than the unadjusted markup. An ADF(4) test based upon these residuals rejects the unit root hypothesis at the 5% level of significance. Recall that we were unable to reject the hypothesis that the unadjusted markup factor contained a unit root.[12] The reasons for the change in the test outcome can be seen in Figure 3, which plots autocorrelegrams for the equilibrium correction factor before and after it has been regressed on the local level. There is clearly less persistence in the equilibrium correction factor after controlling for the local level term.

[12] Of course, non-rejection of the unit root hypothesis most likely reflects the low power of the ADF procedure. Our point here is simply that the markup is more obviously stationary after controlling for the local level.

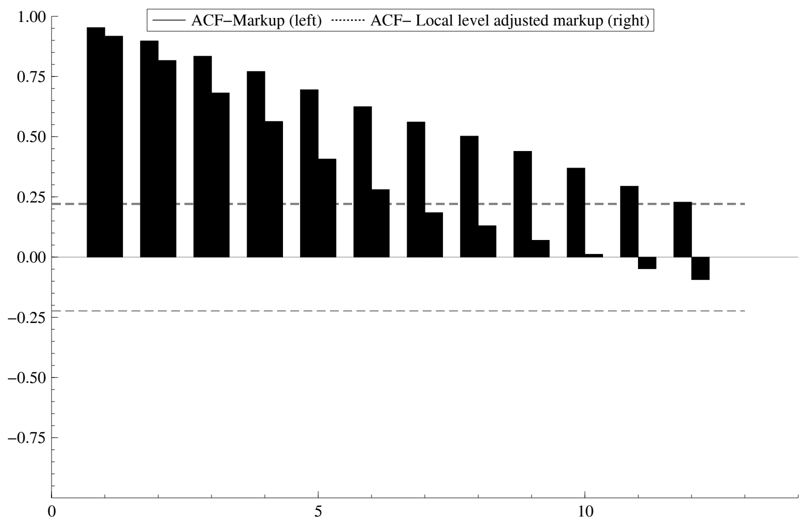

Figure 3: The autocorrelograms for the price-cost markup (left) and the residuals from a regression of the price-cost markup on the local level term (right).

5 Summary

This paper has drawn attention to the fact that standard equilibrium correction models of the price level assume that the markup of prices over costs is constant in the long-run. This assumption is arguably less attractive in the case of the Euro area economy because changes in product market competition are likely to have decreased the markup factor. Consequently, linear equilibrium correction models may omit an important channel for Euro area inflation adjustment and thus provide a poorer fit to the data. One solution to this problem is to include a time-varying intercept, or local level, in the equilibrium correction model. In the case of the Euro area inflation equation, such a term accounts for the negative drift in inflation that a linear model is unable to explain. As a result the extended model provides a better fit to the data. Although the structure of the inflation equation does not change dramatically compared to the linear case, the approach highlights the importance of controlling for the effects of structural change when modelling inflation.

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