Inflation Adjustment in the Open Economy: An I(2) Analysis of UK Prices*

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Abstract

We analyse a cointegrated VAR comprising UK data on consumer prices, unit labour costs, import prices and real consumption growth. The nominal variables, treated as I(2) here, form a linearly homogeneous relation, suggesting a transformation of the system to one comprising inflation and relative prices. This is then estimated in I(1) space. An impulse response analysis using the results suggests that higher real import prices reduce real wages, such that the impact of an external shock on domestic inflation is moderated. This explains why the depreciation of sterling in 1992 left inflation unchanged. In contrast, high real import prices in 1974 increased inflation because wage accommodation effects were absent.

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JEL Classification: C32, C51, C53, E31, F0.

1 Introduction

Empirical models of inflation typically assign a central role to import prices in explaining inflation fluctuations. For example, in the conditional analysis of Australian inflation due to de Brouwer and Ericsson (1998) 'error corrections' of consumer prices with respect to import prices are a key determinant of inflation. In the case of the United Kingdom the import price index has been used in a range of inflation forecasting models employed by the Bank of England, see Bank of England (1999).

In recent years, however, several anomalies have arisen concerning the relationship between import price shocks and inflation. For instance, after the UK exited the European Exchange Rate Mechanism (ERM) in 1992 the ratio of import prices to consumer prices increased by more

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than 10%, but there was no subsequent increase in inflation. Such an episode stands in stark contrast to the British experience of the 1970s, during which time large increases in import prices induced bouts of high inflation.

In this paper we demonstrate that the key to understanding the impact of external shocks on the rate of inflation lies in a joint analysis of consumer prices, import prices and unit labour costs. We analyse these variables, in addition to a cyclical indicator, consumption growth, in a vector autoregression (VAR). We allow the three nominal variables to be integrated of second order, I(2), see Johansen (1992, 1996), and analyse the resulting model using the novel Maximum Likelihood (ML) estimation algorithm of Johansen (1997). Within the system we find one common I(2) trend and two stationary long-run relations. We demonstrate that the common I(2) trend affects each of the nominal variables proportionately, and that consumer prices can therefore be written as a linearly homogeneous function of unit labour costs and import prices. This permits a transformation of the system to one comprising real unit labour costs (equivalently, productivity adjusted real wages), real import prices, consumer price inflation and the growth rate of real consumption. The first three of these variables are integrated of order one, I(1), and may therefore be analysed alongside the consumption growth term using the Vector Error Correction Mechanism (VECM) described in Johansen (1996). The results of this analysis indicate that the first long-run relation links the inflation rate to real unit labour costs and real import prices, and the second relation links consumption growth to its constant steady-state value.

We then describe how impulse response analysis on the VECM can be used to interpret the effects of two types of permanent shock impacting the system, a domestic shock and a foreign shock. A key finding is that increases in real import prices induced by the foreign shock are principally corrected through a downward adjustment of productivity adjusted real wages, such that the pass through to inflation is small, and in some cases insignificant. This real wage accommodation has theoretical foundations in the competing claims models proposed by Layard, Nickell and Jackman (1991), and explains why the post-ERM depreciation of sterling had a relatively benign impact on UK inflation. Next, we show that a version of our results computed for the post-1974 subsample suggests that real wage accommodation was absent at the time of the first oil price shock in 1974, a finding that we attribute to the effects of Phase III of the Heath administration's incomes policies. This increases the partial derivative of the inflation rate with respect to real import prices and thereby provides an explanation of the 'great inflation' of the mid-1970s. Finally, we use our results to assess recent conjectures concerning the future direction of UK inflation. We argue that the upturn in UK inflation predicted to take place once the real import price index returns to its historical average may be less serious than some commentators suggest, for a reduction in real wages will accommodate any adverse trend in import prices and thereby ensure automatic stabilisation of the inflation rate.

The remainder of the paper expands on these points and has the following structure. Section 2 illustrates how a markup model of the price level can be interpreted in light of the time series properties of the variables. Section 3 presents quarterly data on those variables spanning more than three decades. Section 4 provides details of the econometric tools that we employ in our analysis and Section 5 presents results from both the I(2) and I(1) analyses. Section 6 presents an impulse response analysis and discusses its implications for understanding the mechanisms behind open economy inflation adjustment. Finally, Section 7 concludes.

2 Theoretical Framework and Time Series Interpretation

The empirical analysis of the relationship between consumer prices, P_t, unit labour costs, U_t, and import prices, M_t, has a basis in markup models of the price level, see, for example, de Brouwer and Ericsson (1998). In such models the price level is a markup over total unit costs, which we take to be a combination of unit labour costs and import prices. Assuming that the price level is a linearly homogeneous function of input costs (a proposition that we test in Section 5) a partial adjustment model for the price level can be written as follows:

$$\Delta p_t = \omega_0 \Delta p_{t-1} - \omega_1 \cdot (p_{t-1} - \gamma \cdot u_{t-1} - (1-\gamma) \cdot m_{t-1} - \ln \mu), \tag{1}$$

where lower case letters denote logarithms of variables, $\mu - 1$ measures the markup factor, the elasticities of P_t with respect to U_t and M_t are γ and $(1-\gamma)$ respectively and the conditions $\omega_0 > 0$, $\omega_1 > 0$ imply partial adjustment of the price level towards its steady-state value.[1] The presence of the first order autoregressive term in (1) indicates that we consider an eclectic theory of price adjustment, encompassing such factors as adaptive expectations amongst agents, or aggregation over heterogeneous sectors of the economy.

The dynamic relationship in (1) constitutes a simple theoretical starting point for the empirical analysis. The exact interpretation of the dynamics, and the time series properties of different linear combinations of the variables, depends on the order of integration of the different series. If the price variables, $(p_t : u_t : m_t)'$ are integrated of first order, I(1), as is often assumed in econometric analyses, then they could potentially cointegrate to stationarity through the I(1) to I(0) cointegrating relation

$$p_{t-1} - \gamma \cdot u_{t-1} - (1-\gamma) \cdot m_{t-1} \sim I(0), \tag{2}$$

see *inter alia* de Brouwer and Ericsson (1998). In this case (1) describes a simple error correction mechanism with some additional dynamics. If we find the price variables to be I(1) in the empirical analysis, (2) is a likely candidate for a long-run relation.

To motivate an analysis of I(2) variables, subtract Δp_{t-1} from both sides of (1) to obtain

$$\Delta^2 p_t = (\omega_0 - 1) \cdot \Delta p_{t-1} - \omega_1 \cdot (p_{t-1} - \gamma \cdot u_{t-1} - (1-\gamma) \cdot m_{t-1} - \ln \mu). \tag{3}$$

If $(p_t : u_t : m_t)'$ is an I(2) process, the relation (3) allows for an alternative analysis based on a dynamic steady state relation. The price levels may still cointegrate, in this case from I(2) to I(1), in general, such that the linearly homogeneous left hand side of (2) is an I(1) process, as opposed to an I(0) process. This first level of cointegration allows for a transformation from

[1] A partial adjustment process could also imply that further lags in the inflation rate and other cost terms enter (1). We allow for this in the empirical analysis that follows through analysing lags of up to order two. At this stage we restrict the adjustment equation to a parsimonious form as the objective is simply to determine the set of variables that will be useful in modelling inflation, and to illustrate the interpretation of the markup approach.

nominal prices to the real variables, namely the real unit labour cost[2], $(u-p)_t$, and the real import price, $(m-p)_t$. A second layer of cointegration could be the polynomially cointegrating relation

$$\Delta p_{t-1} - \theta \cdot (\gamma \cdot (u_{t-1} - p_{t-1}) + (1-\gamma) \cdot (m_{t-1} - p_{t-1})) \sim I(0), \tag{4}$$

which links the I(1) inflation rate to the I(1) relation between unit labour costs and import prices. The relation (4) can be thought of as a kind of error correction, where the inflation rate, Δp_t, corrects deviations from the I(2) relation given by the left hand side of (1). Deviations from (4) could now be stationary and the stationary second order difference $\Delta^2 p_t$ could error correct to this relation as in (3).

In the empirical analysis we allow the markup to fluctuate over the short-term through business cycle effects. We use the log-linear approximation of a time-varying markup:

$$\ln \mu_t = \theta_0 + \theta_1 z_t, \tag{5}$$

where θ_0 denotes the log of the autonomous component of the steady-state markup and z_t measures the cyclical position of the economy. In this paper we choose the growth rate of real consumer expenditure, Δc_t (c_t being the log of real consumer spending), as the cyclical indicator. Such a variable captures periods of rapid expansion and contraction in the consumer sector and therefore represents an appropriate measure of the cyclical pressures affecting consumer price inflation. Inserting the time varying markup (5) in the dynamic relation (1) modifies the candidates for the stationary long-run relation, (2) or (4), since a consumption growth term would then be included in each of them. If consumption growth is stationary, however, implying that the level of consumption is I(1), then a relationship including Δc_t is not identified by the requirement of stationarity alone, as linear combinations of stationary variables are themselves stationary. But we can still model the effects of the stationary variable within the system through adopting a unit vector as a second long-run relation.

Banerjee, Cockerell and Russell (2001) also consider a time varying markup, where $\ln \mu_t$ explicitly depends on the inflation rate. They argue that firms have imperfect information on market prices and face a comparatively large loss if prices are set too high, e.g. due to the presence of a kinked demand curve. That will cause firms to act cautiously and choose relatively low markups. Taking inflation as a measure of uncertainty, high levels of inflation will be accompanied by a relatively low markup, inducing a link between the markup and inflation and hence between inflation and the price levels. This suggests an alternative interpretation of the polynomially cointegrating relation (4), namely one in terms of the impact of relative prices on the time varying markup and hence inflation.

In the presentation above the focus was solely on the dynamics of the inflation rate and the time series interpretations of such a relationship. It may be the case that wage-setting practices lead unit labour costs to respond to import and consumer price indices, see Layard, Nickell and Jackman (1991), making unit labor costs endogenous to the parameters in the long-run

[2]The real unit labour costs referred to are equivalent to the productivity adjusted real wage facing consumers. However, due to the fact that firms are both producers and retailers in this analysis, and also the fact that we do not model such things as the tax wedge, real unit labour costs are also equivalent to the productivity adjusted real wage facing producers.

relation. To take account of this, the empirical analysis allows for causation in all directions in (4) through estimating a system of equations for the vector of variables $X_t = (p_t : u_t : m_t : \Delta c_t)'$ and the lagged relationships between them.

3 The Data

In the empirical analysis we study X_t for the effective sample $t = 1969 : 1 - 2000 : 4$. We use the natural log of the implicit deflator for household consumption in measuring p_t, the log of average unit labour costs for u_t, the log of the implicit deflator for imports of goods and services for m_t, and the log of real household consumption for c_t, see Appendix A for further details.

The data and some relevant linear combinations are presented in Figure 1. Graph (A) depicts the logs of consumer prices, p_t, unit labour costs, u_t, and import prices, m_t. Over the sample period the total increases in consumer prices and unit labor costs have been quite similar, while the increase in import prices has been somewhat smaller. These differences are reflected in Graph (C), in which real import prices, $m_t - p_t$, are more obviously negatively trended than real unit labour costs, $u_t - p_t$. A further interesting feature of the series plotted in Graph (C) is that for much of the past thirty years real unit labour costs and real import prices have been negatively correlated. For example, real unit labour costs declined rapidly following a large increase in real import prices after sterling exited the ERM in 1992. Similarly, the appreciation of sterling that decreased real import prices during 1996-2000 has been accompanied by upward drift in real unit labour costs. Such co-movements suggest that real wages accommodate the impact of shocks to real import prices. In contrast, real import prices and real unit labour costs appear to correlate positively following the first oil price shock of 1974, suggesting that real wage accommodation effects did not operate at that time. A central theme of the empirical results that we report in Sections 5 and 6 is that real wage accommodation effects have been effective over most of the past three decades, except at the time of the first oil price shock.

The final variable in our empirical analysis is the growth rate of real consumption, Δc_t. This measure of cyclical conditions is more appropriate than an alternative based on GDP, for the inflation rate that we study is derived from the deflator for total consumer expenditure, and is therefore more likely to be a function of fluctuations in the consumer sector than in the aggregate economy.[3]

A priori and after inspection of Figure 1, the most likely scenario in terms of cointegration is that in which there is one I(2) trend affecting $(p_t : u_t : m_t)'$ proportionately. In this case the polynomially cointegrating relation (4) between price inflation, real unit labour costs and real import prices is a candidate for a stationary relation. From graph (D) consumption growth looks stationary suggesting a unit vector as a second stationary relation.

[3] The late 1990s provide a good example of how the choice of cyclical indicator can be important: GDP growth during that period suggested that the economy was expanding at its trend rate, but that masked strong demand pressures in the consumer sector that were offset at the aggregate level by a manufacturing recession, as exporters struggled to cope with the effects of a high sterling exchange rate.

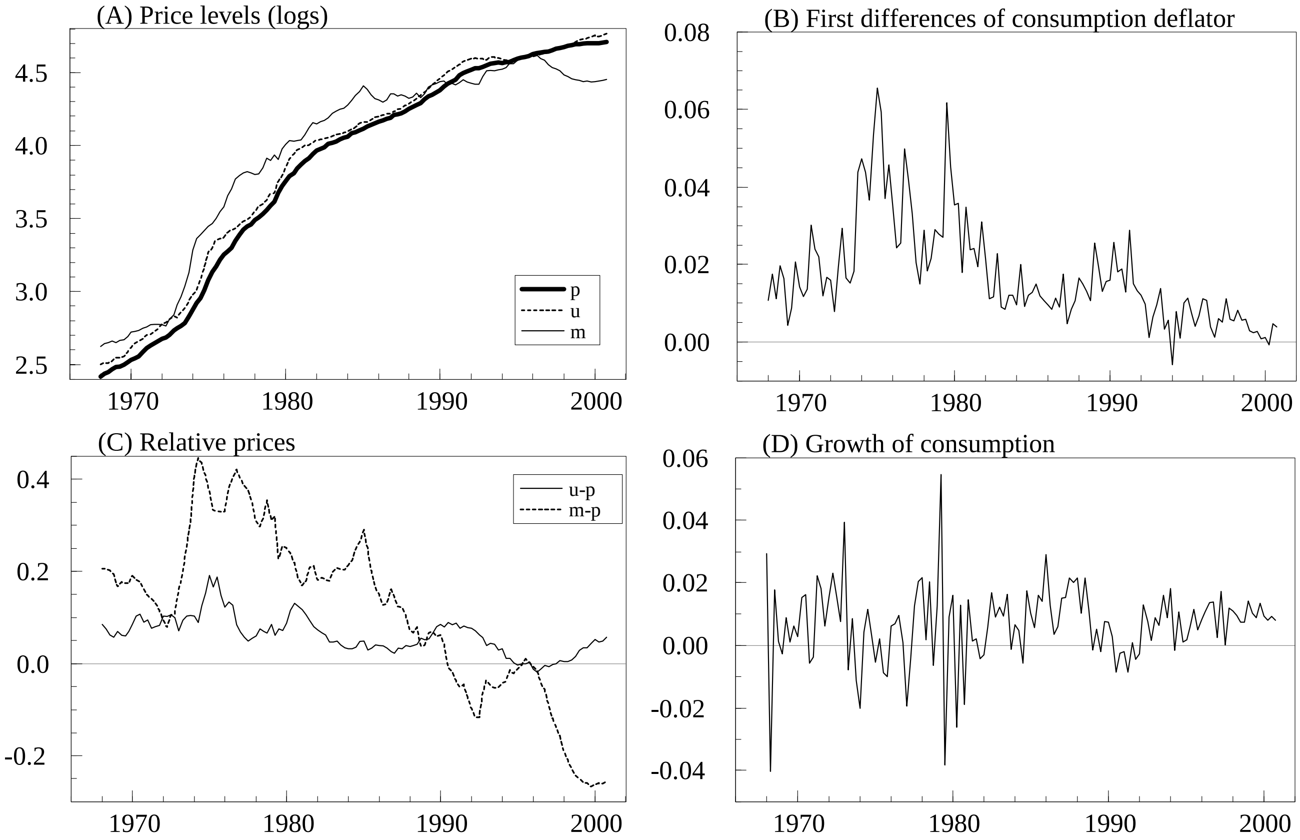

Figure 1: Data and certain linear combinations.

4 The Econometric Approach

The empirical analysis follows the general-to-specific approach of Hendry and Mizon (1993) applied to a VAR incorporating I(1) and I(2) cointegration restrictions, see Johansen (1992, 1996). The basic statistical model for analysing the $p-$dimensional data vector X_t, $t = 1, ..., T$, is a VAR of order k, which can be parametrised as:

$$\Delta^2 X_t = \Pi X_{t-1} - \Gamma \Delta X_{t-1} + \sum_{i=1}^{k-2} \Psi_i \Delta^2 X_{t-i} + \mu_0 + \mu_1 t + \phi D_t + \epsilon_t. \tag{6}$$

The innovations are assumed to be identically and independently Gaussian, $\epsilon_t \sim N(0, \Omega)$, and the initial values, $X_{-k+1}, ..., X_0$, are taken to be fixed. The k matrices of autoregressive parameters, $\Pi, \Gamma, \Psi_1, ..., \Psi_{k-2}$, are each of dimension $p \times p$ and the deterministic specification is given by a constant, μ_0, a linear drift term, $\mu_1 t$, and a set of dummy variables, D_t, with coefficients, ϕ. We further discuss the deterministic specification of the model later in this section.

The cointegrated I(2) model, denoted $H_{r,s}$, is characterised by r stationary directions, s directions integrated of first order, and $p - r - s$ directions integrated of second order. This is a submodel of (6) defined by the reduced rank restrictions[4]

$$\Pi = \alpha\beta' \quad \text{and} \quad \alpha'_\perp \Gamma \beta_\perp = \xi\eta', \tag{7}$$

[4]An additional full rank condition is necessary to exclude integration of higher order.

where α and β are matrices of dimension $p \times r$, ξ and η are matrices of dimension $(p-r) \times s$, and $\alpha_\perp$ and $\beta_\perp$ are the orthogonal complements to α and β respectively, see Johansen (1992, 1996) for additional details. To characterise directions with different degrees of persistence we define the matrices

$$\beta_1 = \overline{\beta}_\perp \eta, \quad \beta_2 = \beta_\perp \eta_\perp \quad \text{and} \quad \delta = \overline{\alpha}' \Gamma \overline{\beta}_2,$$

where for a matrix β we define $\overline{\beta} = \beta \left(\beta'\beta\right)^{-1}$. Using this notation $\beta_2' X_t$ are the $p-r-s$ dimensional non-cointegrating linear combinations of the data. The $r+s$ directions $(\beta : \beta_1)' X_t$ are integrated of less than second order and thereby cointegrate. They can be further divided into s directions, $\beta_1' X_t$, that remain I(1), and r relations that cointegrate to stationarity with the first differences of the I(2) directions through the following polynomially cointegrating relations:

$$S_t = \beta' X_t - \delta \beta_2' \Delta X_t.$$

If $r > p-r-s$ then $\delta_\perp' S_t = \delta_\perp' \beta' X_t$ will be linear combinations of the levels alone, allowing for direct cointegration from I(2) to I(0).

Previous applications of the I(2) model have employed the two-step estimator of Johansen (1995), see *inter alia* Juselius (1998), Diamandis, Georgoutsos and Kouretas (2000), Banerjee at al. (2001) and Nielsen (2002a) for examples. The two-step estimator is basically a sequential application of the reduced rank regression employed in the analysis of VAR models featuring I(1) processes, and is not the ML estimator for the I(2) model. In this paper we rely on the novel ML estimator proposed by Johansen (1997). This technique is based on the parametrisation

$$\begin{aligned} \Delta^2 X_t &= \alpha \left(\rho'\tau' X_{t-1} + \psi' \Delta X_{t-1}\right) + \Omega \alpha_\perp \left(\alpha_\perp' \Omega \alpha_\perp\right)^{-1} \kappa' \tau' \Delta X_{t-1} \\ &\quad + \sum_{i=1}^{k-2} \Psi_i \Delta^2 X_{t-i} + \mu_0 + \mu_1 t + \phi D_t + \epsilon_t, \end{aligned} \tag{8}$$

where $\tau = (\beta : \beta_1)$, $\rho = (I : 0)'$, $\psi' = -\left(\alpha'\Omega^{-1}\alpha\right)^{-1} \alpha'\Omega^{-1}\Gamma$, and $\kappa' = -\left(\alpha_\perp' \Gamma \overline{\beta} : \xi\right)$. There does not exist a closed form solution for the ML estimator in this context, but the estimates can be obtained through an iterative procedure, see Johansen (1997).

The inclusion of deterministic variables in the I(2) model is far from straightforward, as unrestricted deterministic variables in (6) or (8) will accumulate in the solutions for the levels of the variables, e.g. an unrestricted constant will produce a quadratic trend in the I(2) directions. In this paper we use the deterministic specification suggested by Rahbek, Kongsted and Jørgensen (1999), which allows for a linear trend in all directions, including the stationary polynomially cointegrating relations, S_t. This model entails the following restrictions on (8):

$$\begin{aligned} \mu_1 &= \alpha \rho' \tau_0' \\ \mu_0 &= \alpha \psi_0' + \Omega \alpha_\perp \left(\alpha_\perp' \Omega \alpha_\perp\right)^{-1} \kappa' \tau_0', \end{aligned}$$

where τ_0 and ψ_0 are free parameter matrices of dimension $1 \times (r+s)$ and $1 \times r$ respectively. In the empirical analysis the linear trends provide an alternative to stochastic trends in describing the non-stationarity in the data, and can generally be interpreted as the effect of some form of measurement error, e.g. the use of different methods of quality adjustment in constructing consumer and import price deflators. Some directions spanned by the data may feature a zero

Table 1: Lag length determination

k	log-likelihood	Information criteria SW	HQ	AIC	Likelihood Ratio test $k\,\vert 5$	$k\,\vert 4$	$k\,\vert 3$	$k\,\vert 2$
5	2440.36	-33.8851	-35.3667	-36.3806				
4	2432.73	-34.3724	-35.6423	-36.5114	.75			
3	2423.66	-34.8371	-35.8954	-36.6196	.74	.55		
2	2410.63	-35.2401	-36.0867	-36.6661	.47	.27	.15	
1	2346.65	-34.8469	-35.4818	-35.9164	.00	.00	.00	.00

Note: SW, HQ and AIC are the values of the Schwarz, Hannan-Quinn, and Akaike information criteria respectively. The Likelihood Ratio tests $k\,|m$ are the $F-$transforms of the tests for the last $m-k$ lags being insignificant. The reported figures are the corresponding $p-$values.

coefficient on the linear trend, a property of the model that can be tested during the analysis. Importantly, the actual coefficients on the linear trends do not affect the asymptotic properties of estimation and hypothesis testing carried out in other parts of the analysis, see Nielsen and Rahbek (2000) for a discussion of similarity issues in cointegration models.

5 Empirical Analysis

We first set up a congruent statistical model for the data $X_t = (p_t : u_t : m_t : \Delta c_t)'$, where $t = 1969 : 1 - 2000 : 4$, giving 128 effective observations.[5] To take account of special events in the data we include six intervention dummies,[6] namely

$$D_t = (D73q1 : d74q1 : D75q1 : d75q3 : D79q2 : D80q1)'_t.$$

The dummies are balanced, i.e. $\sum_{t=1}^{T} D_t = 0$, ensuring that they do not produce changes in the slopes of the linear trends in the I(2) directions. The dummies $Dxxqi$ are of the form $(0, ..., 0, 1, -1, 0, ..., 0)$ where the value 1 occurs at $19xx : i$, whereas the dummies $dxxqi$ are defined as $(0, ..., 0, 1, -.5, -.5, 0, ..., 0)$. $D73q1$ handles the blips in UK consumer spending due to the fiscal expansion undertaken by the Heath administration, and also the effects of decimilisation of the currency at the start of that year. $d74q1$ controls for fluctuations in both import prices and consumer prices arising at the time of the first oil price shock, while $D75q1$ and $d75q3$ control for the uneven nature of earnings growth in the UK during the period in which the Wilson-Callaghan 'social contract' applied to labour market bargaining. Finally, $D79q2$ and $D80q1$ control for, respectively, the second oil price shock and the effects of increases in Value Added Tax under the first Thatcher administration.

We next determine an appropriate lag length, k, for the unrestricted VAR. Table 1 reports log-likelihood values and various information criteria for different values of k, as well as LR tests for successive lag deletions, see Lütkepohl (1991). Both the LR tests and the information criteria point towards two lags, and we set $k = 2$ in the analysis that follows.

[5]The empirical analysis was carried out using a set of procedures programmed in Ox 3.0, see Doornik (2001).

[6]The location and type of the dummies was identified by looking at the maximised values of the likelihood function. The results of the analysis are not dependent on the particular form of the dummies, and, by and large, identical results can be obtained for a model with no dummies.

Table 2: Tests for mis-specification of the unrestricted VAR(2)

	AR(1)	AR(1-4)	ARCH(4)	Normality
p_t	.02 [.90]	.98 [.42]	.82 [.51]	3.29 [.19]
u_t	.20 [.65]	2.03 [.09]	3.19 [.02]	1.11 [.58]
m_t	.60 [.44]	1.21 [.31]	1.72 [.15]	4.47 [.11]
Δc_t	.00 [.97]	.89 [.48]	.66 [.62]	2.21 [.33]
Multivariate tests:	1.36 [.16]	1.58 [.01]	...	11.96 [.15]

Note: Figures in square brackets are $p-$values. AR(1) are the $F-$tests for first order autocorrelation and are distributed as F(1,111) and F(16,321) for the single equation and vector tests respectively. AR(1-4) are tests for up to fourth order autocorrelation and are distributed as F(4,108) and F(64,366) respectively. ARCH (4) tests for ARCH effects up to the fourth order and is distributed as F(4,104). The last column reports the Jarque-Bera asymptotic tests for normality, which are distributed as $\chi^2(2)$ and $\chi^2(8)$ respectively.

The results of a battery of mis-specification tests for the unrestricted VAR(2) are reported in Table 2. These are generally satisfactory. There is a tendency for ARCH in the errors in the equation for unit labour costs, u_t, but this is due to a greater frequency of large shocks in the first part of the sample and is difficult to remedy within the VAR framework. The multivariate test for no residual autocorrelation up to the fourth order also results in a rejection, and, again, this seems to be due to some persistent shocks affecting the series for unit labour costs. There is no evidence of residual autocorrelation in the individual equations and steps that might remedy the problem, for example including additional lags in the model, do not affect the main conclusions of the analysis. We therefore proceed with the assumption that the VAR(2) is a suitable framework within which to analyse the long-run properties of X_t.

5.1 I(2) Analysis

The rank indices (r, s) defining the cointegrating properties of the I(2) model, $H_{r,s,}$, can be determined via an iterative application of Likelihood Ratio (LR) tests calculated from the maximised log-likelihood values due to different models. Table 3 reports LR tests for each of the restricted models, $H_{r,s}$, against the unrestricted stationary VAR, $H_{4,0}$, as well as the corresponding $p-$values, which are in square brackets.[7]

The columns of Table 3 contain models with the same number of I(2) trends but different numbers of I(1) trends; the last column sets $p - r - s = 0$ and therefore comprises the I(1) models. The usual strategy is to first test the most restricted model, $H_{0,0}$, then $H_{0,1}$, and so on, row-wise, rejecting a model only if each of the more restricted models have also been rejected, see Johansen (1996) for details. All models with no stationary relations, $r = 0$, are safely rejected, and the same is the case for models with $r = 1$. The model $H_{2,1}$ generates a test statistic of 23.44 and a $p-$value of .52. This model features $r = 2$ stationary relations, $s = 1$ I(1) trend, and $p - r - s = 1$ I(2) trend. One of the stationary relations is directly cointegrating from I(2) to I(0) and one is a polynomially cointegrating relation involving first differences. This is

[7] The distributions of the LR tests are non-standard and depend on the precise deterministic specification, see Rahbek et al. (1999). In this paper we use the approximation of Doornik (1998) based on $\Gamma-$distributions.

Table 3: Test for the rank indices of the I(2) model

r	LR tests				
0	535.98	251.85	168.98	126.54	117.51
	[.00]	[.00]	[.00]	[.00]	[.00]
1		160.11	102.31	70.84	65.72
		[.00]	[.00]	[.00]	[.00]
2			54.35	23.44	19.80
			[.01]	[.52]	[.24]
3				10.77	7.28
				[.58]	[.33]
$p-r-s$	4	3	2	1	0

Note: Likelihood Ratio tests for the rank indices (r, s) based on the Maximum Likelihood estimator of the I(2) model. Figures in square brackets are asymptotic $p-$values based on the approximate critical values derived from $\Gamma-$distributions by Doornik (1998).

potentially consistent with the relationships set out in the theoretical framework, in which the polynomially cointegrating relation is given by (4), and the second cointegrating relation is the unit vector $(0:0:0:1)'$, which implies stationarity of the series for consumption growth.

The results of the tests for rank determination are confirmed by the roots of the characteristic polynomial of the autoregressive model. The eight eigenvalues of the companion matrix, which are the inverses of the roots of the characteristic polynomial, are as follows:

$$1.000,\ 1.000,\ 1.000,\ .595,\ .427,\ .405,\ .299 \pm .157i.$$

Note that three of the eigenvalues of the companion matrix are restricted to unity given that the model comprises a unit root from the I(1) trend and a double unit root from the I(2) trend. The unrestricted eigenvalues are all far from the unit circle, reflecting the clear-cut rank determination apparent from Table 3.

A Nominal-to-Real Transformation of the System to I(1). If the nominal variables of the system are found to be first-order homogeneous, it follows that relative magnitudes such as the real wage are invariant to the values taken by nominal aggregates, ruling out such phenomena as 'money illusion'. Thus, a test for first-order homogeneity of the nominal variables also constitutes a check on whether or not the model is consistent with Neo-Classical economic theory. Furthermore, and of particular relevance to the analysis reported in this paper, first-order homogeneity of the nominal variables permits a transformation of the relatively complicated I(2) model to one expressed in I(1) space, which can then be analysed using standard techniques, see Kongsted (1998).

Homogeneity of the data vector $X_t = (p_t : u_t : m_t : \Delta c_t)'$ implies that the loadings applied to the I(2) trend in the evolution of the nominal variables are proportional, i.e. $sp(b) = sp(\beta_2)$, where $b = (1:1:1:0)'$. The estimate of the loadings applied to the I(2) trend is given by

$$\begin{array}{rcccccccl} & p_t & & u_t & & m_t & & \Delta c_t & \\ \widehat{\beta}_2' = (& 1.000 & : & 0.974 & : & 1.309 & : & 0.003 &) \\ b' = (& 1 & : & 1 & : & 1 & : & 0 &) \end{array}$$

which is not too far from the theoretical vector, although the loading applied to the import price is somewhat larger than that applied to consumer prices and unit labour costs. This difference is probably a small sample phenomenon; as the sample extends into the future we would expect the total growth in import prices to converge on that for the other I(2) variables in the system, and for the loading coefficient to get closer to unity.

Since β_2 is orthogonal to $\tau = (\beta : \beta_1)$, homogeneity entails the restriction $b'\tau = 0$ and under the null the LR test statistic for this hypothesis is asymptotically distributed $\chi^2(r+s)$, see Johansen (2000). Using the ML estimation algorithm, we obtain a LR test statistic of 7.38, which corresponds to a $p-$value of .06 in a $\chi^2(3)$ distribution. Our estimated model is therefore consistent with the hypothesis that consumer prices are a linearly homogeneous function of unit labour costs and import prices. Applying the homogeneity restriction we can construct the transformed I(1) data set $Y_t = (X_t'B : \Delta X_t'v)'$, where $B = b_\perp$ and $|b'v| \neq 0$, see Kongsted and Nielsen (2002). In the present paper we choose $Y_t = (u_t - p_t : m_t - p_t : \Delta p_t : \Delta c_t)'$, to obtain a measure of real wages, $u_t - p_t$, real import prices, $m_t - p_t$, and consumer price inflation, Δp_t, as well as consumption growth, Δc_t. Note that $u_t - p_t$ can be interpreted as the share of labour in GDP, a variable frequently used in the analysis of inflation, see Batini, Jackson and Nickell (2000).[8] The analysis of the transformed variables can then be carried out within the well-known I(1) framework

$$\Delta Y_t = \widetilde{\alpha}\widetilde{\beta}^{*\prime}\begin{pmatrix} Y_{t-1} \\ t \end{pmatrix} + \widetilde{\Gamma}_1 \Delta Y_{t-1} + \widetilde{\mu}_0 + \widetilde{\phi} D_t + \epsilon_t. \tag{9}$$

The cointegration rank, r, determined in the I(2) analysis carries over to (9), and the polynomially cointegrating relations from the I(2) model, S_t, are embedded in the new system as I(1) cointegrating relations. This model transformation simplifies hypothesis testing relating to the long-run structure, see Kongsted and Nielsen (2002) for a full discussion.

5.2 Identifying the Long-Run Structure within the I(1) Model

In the I(1) model the space spanned by the columns in $\widetilde{\beta}$ is identified but individual coefficients are not, and in order to obtain an exactly identified structure we have to impose two restrictions on the cointegrating parameters. Table 4 reports, under the heading $\mathcal{H}_0$, the coefficients of two exactly identified long-run relationships, together with $t-$values based on asymptotic standard errors. The first long-run relation is obtained through normalising the vector with respect to consumption growth and then placing a zero coefficient on the inflation term, consistent with one of the long-run relations being directly cointegrating. The second long-run relation resembles (4). The loading coefficients of this exactly identified structure, $\widetilde{\alpha}$, clearly suggest that the real import price is weakly exogenous for the long-run structure of the model.

In the next version of the model, $\mathcal{H}_1$, we test the hypothesis that the first cointegrating relation is a unit vector. The hypothesis implies three over-identifying restrictions on $sp\left(\widetilde{\beta}^*\right)$ and produces a LR test statistic of 9.25, which corresponds to an asymptotic $p-$value of .03 when

[8] Strictly speaking, $u_t - p_t$ constitutes the labour share if p_t represents the GDP deflator, not the consumer expenditure deflator. However, the differences between these two indices have been negligible in the UK over the sample period, and $u_t - p_t$ is therefore very close to the labour share.

Table 4: Identification of the long-run structure

	$\mathcal{H}_0$				$\mathcal{H}_1$				$\mathcal{H}_2$			
	$\widetilde{\beta}^*$		$\widetilde{\alpha}$		$\widetilde{\beta}^*$		$\widetilde{\alpha}$		$\widetilde{\beta}^*$		$\widetilde{\alpha}$	
$u_t - p_t$	.1229	−.2275	.0383	.6172	0	−.2193	.1813	.6586	0	−.2077	.1940	.6547
	(4.01)	(−7.47)	(.27)	(4.32)	...	(−7.36)	(1.55)	(4.87)	...	(−7.10)	(1.65)	(4.88)
$m_t - p_t$	.0070	−.0567	.0850	.1542	0	−.0562	.0048	.1237	0	−.0541	0	0
	(.81)	(−6.65)	(.29)	(.51)	...	(−6.73)	(.02)	(.43)	...	(−6.59)	...	...
Δp_t	0	1	.1891	−.4676	0	1	.1781	−.4953	0	1	.1675	−.4882
	...	...	(2.06)	(−4.94)	...	...	(2.30)	(−5.53)	...	...	(2.16)	(−5.85)
Δc_t	1	0	−.7459	−.0825	1	0	−.5177	.0655	1	0	−.5357	0
	...	...	(−6.71)	(−.72)	...	...	(−5.20)	(.57)	...	...	(−5.67)	...
t	.0001	−.0002			0	−.0002			0	−.0002		
	(2.13)	(−3.94)			...	(−3.90)			...	(−3.58)		
Log-likelihood value	2396.67378				2392.04660				2391.85038			
(a) Asymp. test	...				9.25 [.026]				9.65 [.140]			
(b) Bartlett cor. test	...				6.74 [.081]				7.51 [.276]			
(c) Adj. Bartlett cor.	...				6.05 [.109]				7.27 [.297]			
Asymp. distribution	...				$\chi^2(3)$				$\chi^2(6)$			

Note: $t-$values based on asymptotic standard errors in parentheses. (a) is the uncorrected asymptotic test statistic. (b) is the simulated Bartlett corrected test statistic, calculated as the uncorrected test multiplied by the ratio of the mean of its asymptotic distribution to the simulated mean of the test statistics under the null, where 5000 Monte Carlo replications are used in the simulation. (c) is a bias adjusted test that uses a further correction for the effect of the downward bias in the autoregressive coefficients, see Appendix B for further details.

using a $\chi^2(3)$ distribution. However, LR tests pertaining to cointegrating coefficients are often found to be over-sized in simulations, see *inter alia* Li and Maddala (1997), Jacobson, Vredin and Warne (1998) and Gredenhoff and Jacobson (2001). Further, in Appendix B we report on a simulation exercise that shows that the test for $\mathcal{H}_1$ at a 5% nominal level produces an actual size of around 15% for a sample of 128 observations. Consequently we employ a simulated Bartlett correction which entails adjusting the test statistic by the ratio of its simulated mean under the null to the expectation of its asymptotic distribution. Applying this correction, and also a bias adjusted simulated Bartlett correction which further corrects for the small sample bias affecting the estimated autoregressive coefficients, brings the actual size of the test down to 5% and 4% respectively, see Appendix B.

Constructing the corrected test statistics for the reduction of $\mathcal{H}_0$ to $\mathcal{H}_1$ yields 6.7 and 6.0 respectively; the corresponding $p-$values, based on a $\chi^2(3)$, are .08 and .11. Thus, the hypothesis that consumption growth is a stationary process cannot be rejected. This is the result that one would expect from a visual inspection of Figure 1 (D).

We next impose the restrictions that real import prices are weakly exogenous, and that consumption growth does not react to disequilibrium in the pricing relation, leading to the structure, $\mathcal{H}_2$. The test statistic for $\mathcal{H}_2$ against $\mathcal{H}_0$ is 9.65 and follows a $\chi^2(6)$ distribution. The marginal restrictions embodied in $\mathcal{H}_2$ compared to $\mathcal{H}_1$ are thus easily accepted, and the total structure is accepted with an asymptotic $p-$value of .14 and $p-$values based on the simulated Bartlett corrected tests of approximately .30.

Under the structure $\mathcal{H}_2$ the long-run inflation relation can be written as

$$\Delta p_t = -.2618 \cdot (p_t - .7934 \cdot u_t - .2066 \cdot m_t - .0008 \cdot trend) \tag{10}$$

indicating an import share of just above 20%, a plausible estimate for this sample period. The relationship can be interpreted as a two-step error correction. In the first step the I(1) inflation rate error corrects with respect to the homogeneous long-run pricing relation comprising the I(2) variables. In the second step $\Delta^2 p_t$ error corrects in response to disequilibrium in (10). The adjustment coefficient of $-.49$ in the second step error correction is relatively high, indicating that half of any disequilibrium is removed in each quarter.

One of the fundamental results illustrated by the structure in $\mathcal{H}_2$ is that productivity adjusted real wages error correct with respect to disequilibrium in the long-run relation between inflation, real unit labour costs and real import prices. As we emphasise in Section 6, this implies that shifts in real import prices elicit adjustment of both the inflation rate and the productivity adjusted real wage. Further, the error correction parameter defining real wage responses is of opposite sign to that in the inflation equation, indicating that real wages tend to accommodate higher real import prices and thereby limit their impact on inflation.

Lagged consumption growth exerts a positive effect on both the rate of change of inflation and the rate of change of the real wage, suggesting that erosion of spare capacity in the consumer goods sector tends to accelerate price and wage adjustment. However, the asymptotic standard deviations of the estimated parameters are relatively large, reflecting difficulties in identifying the exact channels through which capacity shortages raise prices. An additional restriction that removes the capacity effect operating via real wages can be imposed on $\mathcal{H}_2$. This raises the direct capacity effect on inflation from .17 to a statistically significant .23. Still, we take $\mathcal{H}_2$ as the preferred model in the analysis that follows and therefore continue to allow for indirect capacity effects operating through real wages.

6 A Structural VAR Interpretation

In order to cast further light on the macroeconomic effects of increases in real import prices we construct an interpretable moving average representation of the preferred model, $\mathcal{H}_2$, and perform an impulse response analysis. We employ the methodology of King, Plosser, Stock and Watson (1991), Melander, Vredin and Warne (1992) and Warne (1993), see also Nielsen (2002c) for the exact implementation.

The cointegrated VAR in (9) has a solution, in terms of innovations and deterministic variables, given by the so-called Granger representation

$$Y_t = C \sum_{i=1}^{t} \epsilon_i + C^* (L) \epsilon_t + f(t), \tag{11}$$

where $f(t)$ comprises the deterministic terms in the model, namely a constant, a linear trend and the dummy variables. The matrix $C = \widetilde{\beta}_\perp \left(\widetilde{\alpha}'_\perp \left(I - \sum_{i=1}^{k-1} \widetilde{\Gamma}_i \right) \widetilde{\beta}_\perp \right)^{-1} \widetilde{\alpha}'_\perp$ defines the long-run impacts of the different innovations and $C^*(L) = \sum_{i=0}^{\infty} C_i^* L^i$ is a convergent matrix polynomial in the lag operator, L. The long run impact matrix, C, has reduced rank, indicating that only $p - r$ linear combinations of the innovations have permanent effects. Furthermore, the innovations, ϵ_t, are in general correlated, which makes it difficult to give the shocks a structural interpretation.

In order to obtain a structural interpretation of the results we require a representation of the form

$$Y_t = \Upsilon \sum_{i=1}^{t} \varphi_i + R^* (L) \begin{pmatrix} \varphi_t \\ \psi_t \end{pmatrix} + f(t). \tag{12}$$

In (12) the innovations $v_t = \left(\varphi_t' : \psi_t'\right)'$ are decomposed into $p - r$ innovations with permanent effects, φ_t, and r innovations with transitory effects, ψ_t, and we assume orthogonality, $E\left(v_t v_t'\right) = I_4$. The long-run impact matrix, Υ, is of dimension $p \times (p - r)$ and in order to exactly identify the driving trends separately, we have to impose *a priori* $\frac{1}{2}(p-r)(p-r-1)$ restrictions corresponding to the zeros in the lower diagonal $(p - r) \times (p - r)$ matrix, see Warne (1993). The exactly identified structure (12) can be obtained from a rotation of (11), i.e. $v_t = F\epsilon_t$, $(\Upsilon : 0_{p\times r}) = CF^{-1}$ and $R^*(L) = C^*(L) F^{-1}$ for some $p \times p$ rotation matrix F, see Warne (1993) or Nielsen (2002c) for the construction of F.

In the present application we wish to identify the two driving trends as a foreign I(1) price trend and a domestic I(1) price trend, which entails the imposition of a single identifying restriction. We assume that the domestic trend does not exert a long-run impact on real import prices, implying no pricing-to-market in the long run. Using this identifying assumption we obtain the following long-run impact matrix, $\widehat{\Upsilon}$, which is based on a one standard deviation shock to each of the innovations underpinning the driving trends. The numbers in parentheses are 90% confidence intervals based on 5000 bootstrap replications.

$$\begin{pmatrix} u_t - p_t \\ \\ m_t - p_t \\ \\ \Delta p_t \\ \\ \Delta c_t \end{pmatrix} = \begin{pmatrix} -.5360 & 1.3633 \\ (-1.0951; -.0239) & (.9724; 1.7088) \\ 3.9376 & .0000 \\ (2.9028; 5.0424) & (...) \\ .1015 & .2832 \\ (-.0161; .2130) & (.1812; .3760) \\ .0000 & .0000 \\ (...) & (...) \end{pmatrix} \begin{pmatrix} \text{Foreign driving trend} \\ \text{Domestic driving trend} \end{pmatrix} + ...$$

The dynamic effects of one standard deviation shocks to the I(1) trends are given in Figure 2.

The domestic shock exerts a positive and significant long-run impact on both inflation and real wages, suggesting that this process may describe the effect of shifts in labour supply, e.g. the increased costs to producers arising from trade union attempts to secure additional employee compensation, which the UK was particularly affected by during the 1970s.

As we would expect, a shock to the foreign trend exerts a positive effect on real import prices, but it also induces a statistically significant reduction in real unit labour costs. This reflects our earlier finding that real wages error correct with respect to disequilibrium in the pricing relation, and provides an explanation for the lack of any significant upturn in UK inflation following the post-ERM devaluation of sterling. Specifically, the increase in real import prices in 1992 did not fuel large increases in the rate of inflation because the resulting supply pressures were offset by reductions in real labour costs implemented in response to the import price shock. Such 'automatic stabilisation' of the inflation rate is evident in the third graph in the first column of Figure 2, which shows that while inflation increases in response to a shock to the foreign driving

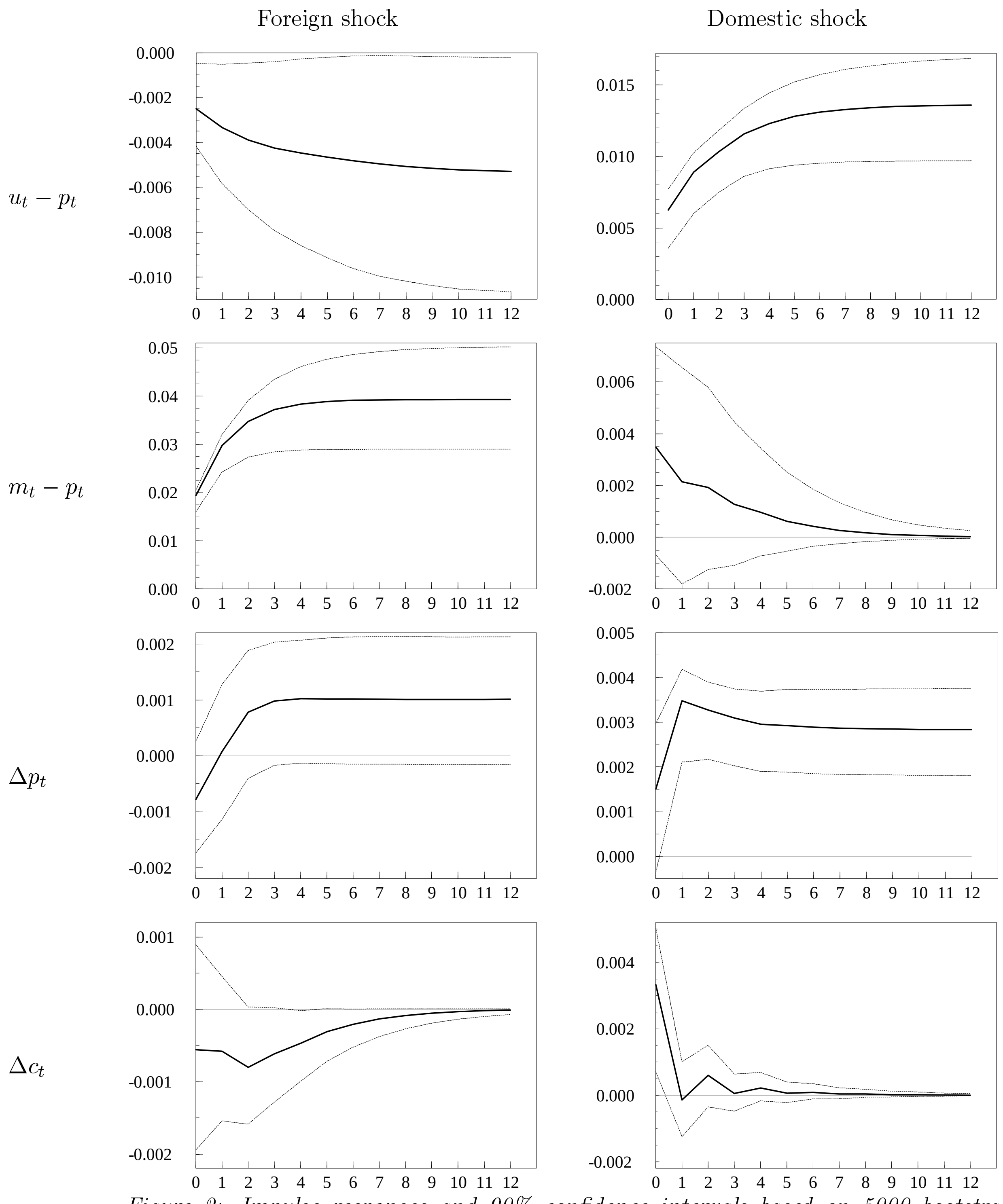

Figure 2: Impulse responses and 90% confidence intervals based on 5000 bootstrap replications. Horizontal axis is time in quarters.

trend, the effect is marginally insignificant. Real wage accommodation effects of this sort have a theoretical basis in wage bargaining models, the central idea being that labour unions reduce real wage claims following adverse shocks to the terms of trade in order to restrict the extent of job losses, see Layard, Nickell and Jackman (1991) for further details.

The role of real wage accommodation in the import price transmission mechanism is especially relevant to the current debate on the future direction of UK inflation. Batini, Jackson and Nickell (2000) estimate a conditional model of UK inflation which confirms the roles of the labour share and the import share in explaining inflation. They then note that at the end of the 1990s the inflation rate was being held down by low real import prices, while the labour share was high by historical standards, injecting inflation into the system. Batini, Jackson and Nickell conclude that any tendency for the real import price index to revert to its historical average will lead to an upturn in inflation. Such an analysis is conditional upon a very crucial assumption, which the authors acknowledge:

> " ... if the relative price of imports reverts to its pre-1995 level, a high level of the [labour] share as the one currently observed may eventually drive up inflation, other things being equal." (Batini et al., 2000, p. 35).

Our results demonstrate that the 'other things equal' assumption underpinning the Batini-Jackson-Nickell analysis is unlikely to hold. The labour share has been shown to be endogenous with respect to disequilibrium in the long-run relation for inflation and relative prices, undermining the results of any comparative static exercise featuring real import prices and inflation. In the case of the UK in the late 1990s, it is precisely because real import prices have hit such a low level that the labour share has been allowed to rise. Any future weakening of sterling that leads to an increase in the real import price index is likely to reverse this trend, as firms squeeze the labour share in order to pay overseas suppliers. This suggests that the impact of higher real import prices on inflation will be very modest.

Analysing the Robustness of the Long-Run Responses. We now assess the robustness of the impulse responses to using a shorter sample period for the underlying estimation. An inspection of Figure 1 suggests that real import prices were an important determinant of inflation during the first half of the $1970s$, but not during the $1990s$. We therefore explore the possibility that inflation adjustment in the UK was governed by a different model at the time of the first great oil price shock of 1974. Obviously it is not possible to address this question through estimating a model for the period up to 1974, as the sample would then comprise just 24 observations. Instead, we analyse the sub-sample $t = 1975 : 1 - 2000 : 4$ and seek to draw inferences concerning the structure of the model in the early $1970s$ through comparing our new results with those obtained for the full sample.

The long-run responses to the permanent shocks over the shorter sample period are given below, together with 90% confidence bands (recall that the ordering of the variables is real unit labour costs, real import prices, inflation, consumption growth).

$$\widehat{\Upsilon}_{\text{Post 1974}} = \begin{pmatrix} -.8553 & .97466 \\ (-1.3426; -.3712) & (.6462; 1.2600) \\ 2.9880 & .0000 \\ (2.1959; 3.7942) & (...) \\ .0509 & .33049 \\ (-.0820; .1828) & (.2250; .4241) \\ .0000 & .0000 \\ (...) & (...) \end{pmatrix}$$

The key result from the sub-sample analysis is that real wage accommodation has been much stronger post 1974, so strong, in fact, that the response of inflation to a one standard deviation shock to the foreign trend is clearly insignificant for that period. This suggests that real wage accommodation was not a feature of price adjustment in the UK at the time of the first oil price shock of 1974, but has been an important factor in the determination of inflation since then.

We attribute this finding to the effects of a system of wage indexation that was in place in 1974 as part of Phase III of the incomes policy of the Heath administration. This regime guaranteed that wage increases would exactly compensate for price inflation above an annual rate of 7%, see Greenaway and Shaw (1988). Consequently, when consumer prices accelerated following a quadrupling of oil prices in 1974, unit labour costs were dragged along at the same rate, preventing real wage accommodation. The absence of real wage accommodation effects constitutes a potential explanation for the bout of high inflation experienced during the mid 1970s, and hence for the fact that the long-run response of inflation to the external shock is larger over the full period than the sub-period. More generally, it illustrates the central role played by real wage accommodation in the transmission of real import price shocks to inflation. Specifically, when real wage accommodation is present fluctuations in real import prices exert a benign effect on inflation (as observed during the mid 1990s), whereas when real wage accommodation is absent real import prices exert a large effect on inflation (as in the 1970s).

7 Summary

This paper has analysed UK data, on consumer prices, unit labour costs, import prices and real consumer spending, within a VAR in which the variables were permitted to be integrated of second order, I(2). Results obtained using the novel ML estimation algorithm due to Johansen (1997) indicated that a single I(2) trend drives the nominal variables, and that there exist two stationary long-run relations in the system. The system satisfied the conditions necessary for a transformation of the variables to I(1) space, and the long-run relations were recovered within the I(1) system.

The first long-run relation links the inflation rate to real unit labour costs and real import prices, whilst the second links the rate of consumption growth to its constant steady-state value. A crucial finding was that real unit labour costs error correct with respect to disequilibrium in the long-run relation between inflation and the relative price measures, suggesting that increases in real import prices may be accommodated through reductions in real wages. As a result of this, a unit innovation to the trend driving real import prices induces an increase in inflation that is actually marginally insignificant. Such a finding is consistent with the British inflation experience following the depreciation of sterling in 1992. On the other hand, results estimated for the post−1974 subsample indicated that real wage accommodation did not operate during the first half of the 1970s, implying that the full burden of adjustment following the first oil price shock was borne by the inflation rate.

These results suggest that the evolution of real wages is a key factor in understanding the transmission of real import price shocks to inflation, and should therefore be taken into account when trying to predict the response of inflation to changes in the macroeconomic environment. In Section 6 we discussed one example of how such reasoning can turn out to be crucial: The Batini-Jackson-Nickell conjecture that UK inflation will increase when the real import price index returns to its historical average may not be borne out in practice, for it is likely that real unit labour costs will adjust in order to moderate inflationary pressures arising from overseas.

Appendix A: Data

The data source is the United Kingdom Office for National Statistics (ONS).

Our measure of consumer prices is the implicit deflator for household final consumption, obtained as the ratio (RPQM/NPSP) using the ONS series codes. Import prices are measured as the implicit deflator for total imports of goods and services, which is defined as (IKBI/IKBL) using the ONS codes. The unit labour cost series is an average measure for the whole economy and is calculated as the ratio of wages and salaries to real gross domestic product, which is (ROYJ/AMBI) using the ONS codes. Real consumption expenditure is defined as the volume index, NPSP, used in the construction of the consumer expenditure deflator.

All of the series obtained from the ONS are seasonally adjusted, and the price series that we construct are indexed such that $1995 = 100$.

Appendix B: Small Sample Properties of LR Tests

As a solution to the small sample size distortion affecting tests relating to the long-run structure, we apply a simulated Bartlett correction proposed in Rocke (1989) in the context of seemingly unrelated regression and in Nielsen (2002b) in the context of tests on cointegrating coefficients. The idea is to estimate the mean of the test statistic under the null for a given sample size, and then use the ratio of that statistic to the mean of the asymptotic distribution as a Bartlett type correction factor that can be applied to the test statistic in order to ensure that its expectation gets closer to that of the asymptotic distribution. To estimate the mean of the test statistic a Monte Carlo simulation is used, with the model estimated under the null from the actual data acting as the data generating process (DGP). The simulated Bartlett correction is closely related to the Bootstrap methodology, see *inter alia* Gredenhoff and Jacobson (2001) for an application of the Bootstrap to tests on cointegrating coefficients and Nielsen (2002b) for a comparison of the two approaches.

The expectation of the corrected test statistic may still be different to that of the asymptotic distribution because the correction is derived from estimated rather than true parameters. Further, due to a downward bias in the estimated autoregressive coefficients, the implied difference in the test statistics tends to be systematic rather than random. To take account of this small sample estimation bias a second level Monte Carlo simulation can be used, see Nielsen (2002b).

To illustrate the properties of the different test procedures we consider the hypothesis imposed under $\mathcal{H}_1$ in Table 4, i.e. stationarity of consumption growth. We use the model estimated from the actual data under $\mathcal{H}_1$ as a DGP and perform a small Monte Carlo simulation. In each of 5000 replications we generate T observations through using historical starting values and drawing pseudo-random Gaussian innovations from distributions assumed to have a covariance structure identical to that estimated from the data. We then test the hypothesis $\mathcal{H}_1$ on each generated data set using the different test procedures, and calculate the actual size of a test as the proportion of the 5000 replications which lead to a rejection of the hypothesis at the nominal 5% level when a $\chi^2(3)$ is the reference distribution.

The rejection frequencies in the Monte Carlo simulations for the uncorrected as well as the two simulated Bartlett corrected tests are reported in Figure A.1 for different numbers of

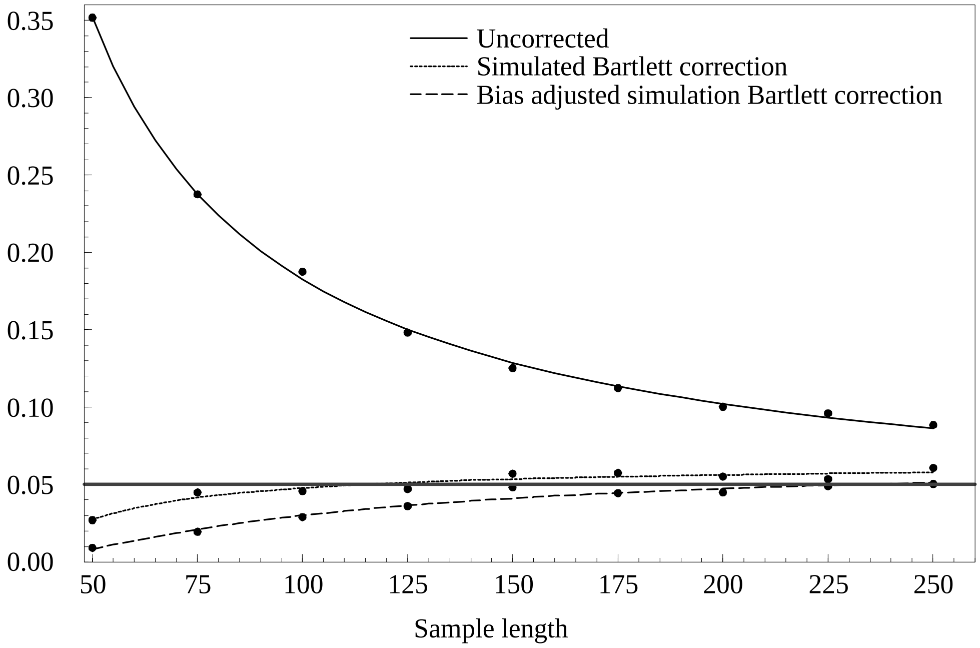

Figure A.1: Actual sizes of uncorrected and corrected LR tests for restrictions on β^*. The solid circles are results of the Monte Carlo simulation and the lines are simple response surface regressions. Based on 5000 Monte Carlo replications and 200 pseudo samples in each replication to estimate the mean.

observations. For very small samples the actual size of the uncorrected test is very large, but it converges to the nominal size as the number of observations increases. For the actual number of observations used for estimation in this paper, $T = 128$, the actual size of the uncorrected test is around 15%. Both of the simulated Bartlett corrected tests have actual sizes closer to the nominal size of 5%, and the rate of convergence to that nominal size is faster. In very small samples the correction factor is too big for the present DGP and the corrected tests are under-sized. For 128 observations the actual sizes in the simulations are 5% and 4% respectively. Finally, we note that simulations in Nielsen (2002b) indicate that these test corrections can be implemented without any substantial reduction in (size corrected) power.

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