Inflation, relative price variability and the markup: Evidence from the United States and the United Kingdom

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Abstract

This paper links two separate empirical literatures on inflation, relative price variability (\textit{RPV}) and the price–cost markup. Measures constructed from annual and quarterly data for the US and UK are used to examine the nature of the relationships among these variables. The results show that two long-run relationships can be identified from the data: a negative relationship between inflation and the markup and a positive relationship between inflation and \textit{RPV}. We eliminate the possibility that inflation is a proxy for \textit{RPV} in the inflation–markup relationship and show systems estimates to be superior to single-equation estimates.

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1. Introduction


Two broad sets of explanations of the negative relationship are offered in the literature. The first set is the ‘demand side’ arguments that focus on the role that relative price dispersion has on the level of the markup. In this case, higher inflation leads to higher price dispersion which in turn reduces the markup. The second set is the ‘supply side’ arguments that consider how inflation interacts with the way firms adjust prices and, in the process, how inflation affects the markup.

It is intriguing that in the empirical work examining the veracity of the demand-side explanations the models are estimated using inflation and the markup rather than price dispersion and the markup. Consequently, inflation in these papers is only a proxy for relative price dispersion. If the demand-side arguments and this modelling approach are valid, we should be able to replace inflation with a measure of price dispersion to yield a similar relationship.

This paper attempts to distinguish between these two broad explanations of the inflation–markup relationship by considering the relationship between inflation, price dispersion and the markup simultaneously. In doing so we draw together two separate literatures; one on inflation and price dispersion (as measured by relative price variability, \( RPV \)) and the other on the markup and inflation.\(^1\) While previous studies were able to establish, independently, the consensus that there is a negative relationship between inflation and the markup and a positive relationship between inflation and \( RPV \) they were not able to do so in a single unified modelling approach.\(^2,3\)

Our framework allows us to consider five important questions concerning the relationships between inflation, \( RPV \) and the markup to which the earlier literature could only provide partial answers. We ask, first, ‘What relationships exist between inflation, the markup and \( RPV \)?’ Second, ‘Are the relationships between the variables short- or long-run phenomena?’ An answer to this question takes us to the heart of the issue of the appropriate econometric treatment of the variables as stationary or cointegrated variables.\(^4\) Third, to distinguish between the two broad sets of explanations we ask, ‘Does the system involve two separate relationships, between inflation and the markup and between inflation and \( RPV \), or a single relationship among all three variables?’ If inflation acts as a proxy for \( RPV \) then we should be able to replace inflation with \( RPV \) in the inflation–markup relationship to yield a similar relationship. Fourth, nearly all of the existing empirical analysis of the inflation–\( RPV \) relationship assumes that single-equation estimation is appropriate. We ask ‘Is the assumption of single-equation estimation appropriate and sustainable?’

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\(^1\) The separate treatment of these relationships in the literature is surprising given that the estimated relationships are often derived from common theoretical bases.


\(^3\) The only paper of which we are aware that models inflation and \( RPV \) (but excludes the markup) in a systems framework is Mizon (1991) using quarterly United Kingdom data for the period 1965–1987.

And lastly, ‘Can we discriminate between the demand-side and supply-side arguments that potentially offer support for the relationships between these variables?’

We estimate a multivariate system comprising inflation, the markup and RPV. For comparability with much of the existing empirical literature we estimate the system with both quarterly and annual data for the United States and the United Kingdom. The following results emerge from our analysis. First, in all four data sets, we re-establish the negative relationship between the markup and inflation and a positive relationship between inflation and RPV. Second, these relationships are characterized as long-run relationships as identified from cointegration analysis. This is a central finding since it proceeds from the testable assumption that the data for inflation, markup and RPV are all integrated of order one. Third, we are able to accept the restriction that RPV does not appear in the markup—inflation relationship, and that the markup does not appear in the inflation—RPV equation. From this it follows that inflation does not act as a proxy for RPV in the first relationship. Fourth, estimating within a systems framework, we are also able to establish that single-equation estimation of the inflation—RPV relationship is valid for three of the four data sets while single equation estimation of the inflation—markup relationship is not valid in any of the data sets examined. Therefore, the inflation—RPV relationship should be investigated within a system, since this facilitates the simultaneous consideration of exogeneity, integration of the variables and inter-linkages between the cointegrating relationships. Lastly, we find evidence consistent with a supply-side argument for a relationship between inflation, RPV and the markup. All five findings establish the existing results while making a compelling case that the markup—inflation and RPV—inflation relationships should be considered in a systems context.

The economic importance of the empirical issue discussed above arises from the literature often referred to as the ‘sand in the wheels’ view, which is attributed to Lucas (1973, 1994) and Barro (1976) inter alia. It underlies the opinion that inflation is costly to welfare and is often justified by demonstrating that the dispersion of prices increases with the average level of inflation (cf. Hercowitz, 1981). However, many authors have challenged the opinion that higher inflation generates welfare losses by claiming that the countervailing influences of inflation on competition and market power need also to be considered. Therefore it is important to estimate the inflation—RPV—markup relationship carefully in order to make a proper judgement on the welfare considerations of inflation.

In the next section we briefly consider the economic literature on the inflation—markup and inflation—RPV relationships. Section 3 sets out the long-run relationships, discussing the measurement issues, data sources and the econometric framework for estimation. Section 4 reports the empirical results. Section 5 concludes.

2. Economic background

2.1. Inflation and the markup

The literature on the inflation—markup relationship can be separated into ‘demand-side’ and ‘supply-side’ explanations. On the demand-side, Bénabou (1988, 1992) and Diamond (1993) focus on the interaction of inflation with the demand for a firm’s output. The essential element in this approach is that consumers on the demand side search across firms to find the best priced goods. The consumer gains utility from consuming the good and this is offset against the cost of purchase i.e. the price. In these models higher inflation increases price dispersion and consumers have a greater incentive to engage in search behaviour to identify ‘bargains’. Like the menu cost literature, firms face

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5 Banerjee et al. (2001) and Banerjee and Russell (2001a) have argued at length the case in favour of regarding inflation and markup as $I(1)$ variables. We re-establish this finding for all the data sets examined in this paper including the RPV series.
costs of changing posted prices of goods in their inventory and therefore when there is inflation, goods
that have been kept on the stocks for some time will appear to be ‘cheap’ relative to newly produced
goods. When inflation increases, consumers engage in search to discover ‘bargain’ priced goods with
their fixed posted prices that have not kept pace with inflation. This search behaviour by consumers
leads to an increase in competition among sellers and results in a subsequent fall in the markup. This
explanation implies that the increase in RPV associated with higher inflation is the cause of the lower
markup. We return to this issue in Section 4 when we discuss whether inflation is a proxy for RPV.

The most prominent supply-side explanation is the ‘menu’ cost argument of Mankiw (1985) and
Parkin (1986) which relies on the proposition that firms face costs when adjusting posted prices i.e.
catalogue prices in response to realised demand shocks. Although small, these costs have large
welfare effects. Firms have insufficient incentives to adjust prices and thereby bring the market back
to equilibrium, but the costs to society from being out of equilibrium are substantial. The asymmetry
in the incentive structure results in prices that can be ‘stuck’ at levels that are too high.

Users of this approach include Rotemberg (1983), Kuran (1986), Naish (1986), Danziger (1988),
Konieczny (1990) and Bénabou and Konieczny (1994) who model the price setting behaviour of
firms in this way. They show that when there is inflation in the economy this has a negative impact on
the average markup. These models assume that firms follow a fixed price adjustment rule, which
may not be a sustainable assumption in the long run. For example, if these firms notice that their
average markup diverges from the profit maximising markup when inflation is stable then they could
simply adjust their price rule to eliminate the divergence.

Non-menu cost supply-side explanations may provide a long-run relationship between inflation
and the markup. These approaches do not rely on the asymmetry of the incentive structure to
motivate such a relationship, but argue that co-ordination between firms becomes more difficult as
the marketplace becomes more ‘noisy’. For example, Athey, Bagwell and Sanichiro (1998)
indirectly argue that higher inflation is associated with higher variability in input costs making it
harder for firms to collude on final goods prices. This results in greater competition and the markup is
reduced in the long run if the RPV is sustained. Further supply-side explanations are provided by
Russell, Evans, and Preston (2002), and Chen and Russell (2002). The first paper argues that
marginal costs are difficult to apportion between jointly-produced output and are not known with
certainty. Furthermore, it argues that inflation increases the uncertainty about the appropriate price of
the output. Firms faced with these uncertainties will have incentives to minimise the expected loss
from setting an incorrect markup, and will veer towards a ‘low’ markup to avoid losses when
inflation increases. The result is that the markup can deviate systematically from its ‘correct’ setting.
The second paper explores the difficulties that firms face when attempting to co-ordinate price
changes in an inflationary environment. These papers argue explicitly that there exists a negative
relationship between the markup and steady state rates of inflation that may be thought of as a long-
run relationship. In these models, the lower markup associated with higher inflation is interpreted as
the cost to firms of overcoming the missing information when setting prices.

6 Unlike the menu cost literature the costs of changing goods prices are not fixed. It is assumed that there is no cost to
changing the price of a newly produced good, but there is a cost to changing the price of an item in the inventory.
Therefore, prices can change costlessly if a good is sold and replaced by a newly produced good.
7 Bénabou and Konieczny (1994) show, in an encompassing model of the menu cost papers, that the relationship
between inflation and the markup may be either positive or negative depending on the relative size of inflation, the
‘menu’ costs, the discount rate, as well as whether the profit function is left- or right-skewed. However, under
‘reasonable’ assumptions it is likely the relationship between inflation and the markup is negative.
8 See Ball, Mankiw, and Romer (1988), Sims (1988) and Dotsey, King, and Wolman (1999).
2.2. Inflation and relative price variability

Explanations of the relationship between inflation and RPV can be separated into three categories. The first is also based on the Mankiw (1985) and Parkin (1986) ‘menu’ cost literature mentioned above. This model can generate a positive relationship between inflation and RPV if inflation pushes prices above the upper threshold price in an \((S, s)\) framework (see Sheshinski and Weiss (1977) and the majority of the empirical literature supports this view). However, when there is negative inflation, pushing prices below the lower threshold, there may no longer be a positive relationship. Furthermore, Rotemberg (1983) and Ball and Mankiw (1994) argue that price dispersion may fall with inflation if firms try to avoid the menu costs of continual adjustment or the costs of losing market share.

The second set of explanations is based on the ‘sand in the wheels’ argument of Lucas (1973). Hercowitz (1981), Lach and Tsiddon (1992) and Debelle and Lamont (1997) argue that higher inflation causes an increase in misperceptions that lead in turn to higher RPV. The issue facing buyers is to decide whether they have been offered a price by a high-priced seller or an average-priced seller (in a situation where the price rises in line with inflation). The increase in the inflation rate makes it more difficult for the buyer to distinguish between the signal and the noise, and RPV increases.

A third explanation suggests that the relationship is simply an artefact of the aggregation of the data or the time period examined. Fischer (1981), Hartman (1991) and Driffill, Mizon, and Ulph (1990) argue that shocks to food and energy prices in the 1970s simultaneously increased inflation and measures of RPV and thereby created or overstated the positive relationship between inflation and RPV. In this case there is no underlying economic reason for the relationship, only a statistical association.

3. The long-run relationships

The long-run relationships between inflation and the markup on the one hand and inflation and RPV on the other can be derived from any of the previous theories. We will determine which of the theories are consistent with the evidence by examining the nature of the relationships in the next section, but in this section we define the long-run relationships, consider the measurement issues, discuss our data sources and describe our econometric methodology.

The long-run relationship between the markup and inflation from the economic models above can be written as:

\[
\mu_t = q - \lambda \Delta p_t \tag{1}
\]

and

\[
rpv_t = \zeta_0 + \zeta_1 \Delta p_t \tag{2}
\]

In Eq. (1), \(\mu_t\) is the estimated markup of price on unit costs ‘net’ of the costs of inflation, and \(\Delta p_t\) represents inflation; \(q\) is the ‘gross’ markup, and \(\lambda\) is a positive parameter and is termed the ‘inflation cost’ coefficient. In Eq. (2), \(rpv_t\) is relative price variability and \(\zeta_0\) and \(\zeta_1\) are positive parameters. Lower-case variables are measured in natural logarithms. In our systems approach described in Section 3.3 we need not impose coefficient restrictions that result in these long-run relationships although the data support them and we describe the system with the appropriate restrictions imposed.

3.1. Measurement issues

The markup, \(\mu\), for \(k\) inputs of the production process is defined as:

\[
\mu = p - \sum_{i=1}^{k} \psi_i c_i \tag{1a}
\]
where the $c_i$ represents the logarithms of the costs of production for input $i$, and $\sum_{i=1}^{k} \psi_i = 1$. If the latter condition is satisfied then the relationship between prices and costs can be termed the markup on unit costs. The condition $\sum_{i=1}^{k} \psi_i = 1$ imposes linear homogeneity and implies that, ceteris paribus, a change in unit costs will be fully reflected in the price level in the long run leaving the markup unchanged. In our case, ceteris paribus includes no change in the rate of inflation. In a ‘standard’ macroeconomic framework $\lambda = 0$ and inflation has no effect on the markup in the long run. In the estimates reported here, $\lambda > 0$ and the markup, net of the cost of inflation, is lower with higher inflation and vice versa in the long run.

Imposing linear homogeneity, we can write the markup defined in Eq. (1a) as the weighted sum of the markup of price on unit labour costs, and the ‘real exchange rate’ where the latter is defined appropriately for the measure of prices used in the estimation. When the markup is calculated using consumer prices, we write Eq. (1) as:

$$
u_{1t} = \nu_{Lt} + \rho \nu_{Mr} = q - \lambda \Delta p_t$$

(3)

where $\nu_{Lt}=p_{Ct}-ulc_t$, ($ulc_t$ is unit labour costs) and the difference between the consumer price and unit import costs is the real exchange rate, $rer_{Mt}=p_{Ct} - pm_t$, ($pm_t$ is unit import costs). The estimated markup written in the form of Eq. (1a) with linear homogeneity imposed and derived from Eq. (3) is:

$$
u_{1t} = p_{Ct} - \frac{1}{1+\rho} ulc_t \left( 1 - \frac{1}{1+\rho} \right) pm_t$$

(4)

and may be interpreted as the markup of price on unit labour and import costs. Alternatively, if the markup in Eq. (1a) is calculated using the GDP deflator, $p_{GDP}$, the real exchange rate must instead be defined as $rer_{Xt}=p_{GDP} - px_t$ to adjust the price series for export prices, $px_t$. The estimated markup is then:

$$
u_{2t} = p_{GDP} - \frac{1}{1+\rho} ulc_t \left( 1 - \frac{1}{1+\rho} \right) px_t$$

(5)

We can interpret $\nu_{2t}$ as the markup of the price of non-traded GDP on unit labour costs. This interpretation is explained later in the Appendix.

As the next section makes clear, we use both annual and quarterly data sets to cross-check our findings. For quarterly data we use consumer prices to construct the markup using the first measure, while for annual data we rely on GDP deflators and use the second measure of the markup.


$$RPV_t = \left( \sum_{i=1}^{n} s_i (\Delta p_{it} - \Delta p_T)^2 \right)^{\frac{1}{2}}$$

(6)

where $s_i$ is the share of each component index in the total index, and $\Delta p_{it}$ is the rate of inflation of the individual component index and $\Delta p_T$ is the rate of inflation of the total price index. $RPV_t$ is therefore
the weighted average of the component inflation rates relative to the aggregate measure of annualised inflation.

Four alternative measures of $RPV$ are set out in Table 1 of Fielding and Mizen (2001). While all these measures capture the variability of prices around the average, a closer look at the relevant literature shows that the overwhelming majority of studies use the measure of $RPV$ used here. For example, Hartman (1991), Mizon (1991), Ball and Mankiw (1994), Parsley (1996) and Fielding and Mizen (2001) have examined the inflation–$RPV$ relationship using variants of our measure of $RPV$. In addition, our chosen measure of $RPV$ avoids the problem of trending relative prices in the component indexes due to different productivity growth rates between the sectors.

3.2. Data

The sources and definitions of the data are reported in Table 1. We use both quarterly and annual data for the US and the UK. This allows comparison of our results, and reduces the possibility that we will be drawing country- or data-frequency specific conclusions. It also imposes on us the discipline that the relationships between the variables for the same country at different frequencies should give a common interpretation.\(^{12}\)

We use two measures of prices in the calculation of the markup on unit labour costs, the real exchange rate, and inflation. The first is the private consumption implicit price deflator and the second is the gross domestic product implicit price deflator. Both are derived from national accounts data. Unit labour costs are measured on an aggregate economy-wide basis while imports and exports prices are the respective goods and services implicit price deflators.

Except for the quarterly United States data, $RPV$ is calculated using the component indices of the aggregate price index used for measuring inflation and the markup. The measures of $RPV$ are therefore consistent with the measures of inflation and the markup. For the quarterly United States data, results obtained from $RPV$ (calculated from the All Urban Consumer Price Index (CPI-U)) and inflation (calculated from the private consumption implicit price deflator or the Consumer Price Index) showed some instability and were not robust to changes in the lag structure of the system. Consequently, although there were numerous indications that the results were consistent with those from the other data sets, the United States quarterly inflation and the markup are calculated using national accounts gross domestic product data and $RPV$ is calculated using CPI-U data.

The short-run impact of the business cycle on inflation, the markup and $RPV$ is represented by the difference in the logarithm of the unemployment rate. The variable is stationary and avoids the ‘errors in measurement’ problem when using de-trended constant price gross domestic product in association with national accounts price data.\(^{13}\)

The annual and quarterly measures of $RPV$ are highly skewed and bounded at zero.\(^{14}\) The explosive nature of the ‘spikes’ in the data often leads to non-normal and serially correlated system

\(^{12}\) There are some obvious qualifications to this statement as a result of the different data definitions employed at each frequency.

\(^{13}\) Measurement errors in national accounts data often have a simultaneous impact on the price and output series so as to offset each other. Consequently, estimates of the relationship between price and output data would be contaminated by the presence of common measurement errors. This contamination is not likely to be serious if the price series spans a small component of the output series. However, in our case the span of the price series and output series are the same or very similar.

\(^{14}\) Vining and Elwertowski (1976) and Mizon et al. (1991) similarly find that the assumption of normality is strongly rejected for $RPV$ due to strong skewness and kurtosis. Banerjee, Mizen and Russell (2002) report graphs and descriptive statistics of the annual and quarterly measures of $RPV$. 

residuals. We identified in Section 2 that shocks to certain components of the price series used to generate inflation and \( RPV \) measures could be responsible for a positive association between the two variables, quite independently of any economic relationship. In order to control for this effect we wish to remove the largest shocks from the \( RPV \) series which are responsible for the spikes in the series. While the contemporaneous impact of these ‘spikes’ on the dependent variables can be accounted for by the introduction of dummy variables in the long-run relationships it is more difficult to account for their impact on the dynamic system. Instead we pre-filter the series to reduce the skewness of the data by regressing \( RPV \) on a series of dummy variables that coincide with observations of more than 2.5 standard deviations from the mean of the series.\(^{15}\) The dummies coincide with periods of high inflation and, therefore, removing these observations is likely to make it more difficult to find a positive relationship between inflation and \( RPV \). We use the de-spiked \( RPV \) series in place of the original, and even after removing the spikes we continue to find that our results below are consistent with the general view in the literature of a positive relationship between inflation and \( RPV \). We can then be sure that the criticisms of earlier empirical work made by Fischer (1981), Hartman (1991) and Driffill et al. (1990) that simultaneous shocks to inflation and \( RPV \) are the cause of a positive relationship between \( RPV \) and inflation do not apply here.

\(^{15}\) The initial system estimates based on \( RPV \) provide similar results to those we report below. For the quarterly United Kingdom data it was found that to reduce the skewness to ‘manageable’ levels in terms of the system diagnostics, dummy variables were necessary for observations greater than 2 standard deviations from the mean and reported in the data appendix of Banerjee et al. (2002).
Prior to estimation, the integration properties of the data were investigated using PT and DF–GLS univariate unit root tests from Elliot, Rothenberg and Stock (1996).\(^\text{16}\) We find that inflation, the markup, the real exchange rate and \(RPV\) variables may be characterized as \(I(1)\), and the logarithm of the unemployment rate is \(I(1)\). The estimates, therefore, are given by an \(I(1)\) system consisting of four core variables: the markup on unit labour costs, the ‘real exchange rate’, inflation and \(RPV\). Estimation of the system is conditioned upon the change in the logarithm of unemployment representing the business cycle and a set of spike dummies to capture the occasionally erratic behaviour of the data particularly in the turbulent mid- to late-1970s.\(^\text{17}\) To allow for the possibility of structural change in the economy that leads to shifts in the long-run relationships, a trend is included in the cointegrating space and omitted whenever insignificant. Estimates are given at both the annual and quarterly frequencies for the United States and the United Kingdom.

3.3. Econometric methodology

Taking our two log–linear Eqs. (1) and (2) we include intercepts and trends (where appropriate) to ensure that the long-run reduced form disturbances have zero means:

\[
m_{ut} = \delta_0 + \delta_1 t + \delta_2 \Delta p_t + \delta_3 rer_t + \xi_{1t} \\
r_{pt} = \zeta_0 + \zeta_1 t + \zeta_2 \Delta p_t + \zeta_{2t}
\]

the Eqs. (7) and (8) can then be written as a long-run compact form as:

\[
\xi_t = \beta' z_t - b_0 - b_1 (t)
\]

where

\[
\xi_t = (\xi_{1t}, \xi_{2t})', z_t = (m_{ut}, r_{pt}, \Delta p_t, rer_t)',
\]

\[
b_0 = (\delta_0, \zeta_0)', b_1 = (\delta_1, \zeta_1)
\]

and

\[
\beta' = \begin{pmatrix}
1 & 0 & -\delta_2 & -\delta_3 \\
0 & 1 & -\zeta_2 & 0
\end{pmatrix}
\]

In estimation we embed the two long-run relations given by Eqs. (7) and (8) in the vector error correction framework given by:

\[
\Delta z_t = a_z - \alpha \xi_{t-1} + \sum_{j=1}^{J} \Gamma_{zj} \Delta z_{t-j} + u_{zt}
\]

where \(a_z\) is a \((4 \times 1)\) matrix of intercepts, \(\alpha\) is a \((4 \times 2)\) matrix of loading coefficients, \(\Gamma_{zj}\) is a set of \((4 \times 4)\) matrices of short-run coefficients. Finally, \(u_{zt}\) is a \((4 \times 1)\) vector of disturbances assumed to be iid(0, \(\Sigma_z\)). Here we have followed Sims (1980), Pesaran, Shin and Smith (2002), Pesaran and Shin (2002) and Garratt, Lee, Pesaran, and Shin (2003) in assuming that deviations from long-run reduced forms can be approximately captured by a finite number of lags in the explanatory variables. The variables in \(z_t\) are all assumed to be first-difference stationary.

\(^{16}\) These results are available on request from the authors.

\(^{17}\) The dummies differ from those used to ‘de-spike’ the RPV series and are reported in Tables 3a–3d.
4. Empirical results

Our initial concern is with the long run properties of the system. We therefore first determine the number of cointegrating vectors using the Johansen (1988, 1995) maximum likelihood systems estimator following a methodology outlined in Banerjee et al. (1993). Table 2 establishes the existence of two cointegrating vectors in each system. In all four cases, for US and UK data at both annual and quarterly frequencies, we can reject the null that there is at least one cointegrating relationship between the variables. We cannot reject the null that there are at least two (the value of the test statistic can be compared to the critical value in brackets beneath). Further evidence of two cointegrating vectors is provided by the roots in modulus of the companion matrix and the stationarity of both cointegrating vectors. This finding that there are two cointegrating vectors in all four systems considered implies that we can be reasonably assured that there are two distinct relationships between our four variables identified in the data.

Tables 3a–3d report the two long-run relationships detected in the system. The first four columns in each table give the adjustment coefficients in the dynamic equations for each of the variables and report the feedback on each variable from the deviation from the estimated long-run relationship (given in the final column of each table).

The estimated long-run relationship is obtained by normalising each of the two cointegrating vectors on one of the variables. On theoretical grounds we interpret one of the relationships to be between the markup and inflation and the other to be between inflation and $RPV$. Accordingly we normalize on the markup on unit labour costs ($mu$) in the first row and on inflation ($\Delta p$) in the second row. Later in the paper we consider an alternative specification of the long-run relationships (see Table 4). The diagnostic tests reported in Tables 3a–3d accept the null of no serial correlation and normality of the residuals in all cases suggesting that the system is correctly specified.

The common feature of all four estimated inflation–markup relationships is a negative and significant relationship between the markup and inflation and a positive and significant relationship between inflation and $RPV$. The second result is obtained even though the outliers have been removed from the $RPV$ series, which a priori makes the finding less likely since the correction removes high $RPV$ values associated with times of high inflation. In some cases there is also a trend in the cointegrating relation, which has a small but significant coefficient. These findings confirm the most commonly reported results in the empirical literature since they deliver the expected signs of the coefficients in all cases. Thus, in answer to our first and second questions, we can confirm the finding of two long-run relationships within the system and can demonstrate that the long-run relationships have the same signs as the relationships reported previously in the separate empirical literatures.

Once annualised, the estimates from quarterly data are similar to those derived from annual data for the United States. This is encouraging since the two systems were estimated independently, and were not constrained in any way to yield similar results. The results are numerically similar to those reported in Banerjee and Russell (2001a,b) for the United States even though the samples differ and the system estimated here contains $RPV$.

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18 The roots in modulus of the companion matrix are consistent with the maintained hypothesis of two cointegrating vectors in a four variable system where we expect the first two roots to be unity and the remaining roots bounded away from unity. Both cointegrating vectors in each case visually appear stationary and this can be verified by univariate unit root tests.

19 Normalisation does not imply direction of causation because we could normalise on any of the variables in the cointegrating vector whilst maintaining the same long-run relationship between them.

20 The long-run coefficients in the inflation–markup relationship are also very similar to those reported in Banerjee and Russell (2001a,b) for the United States.
Table 2
Testing for the number of cointegrating vectors estimated values of $Q(r)$

<table>
<thead>
<tr>
<th>Annual</th>
<th>Quarterly</th>
</tr>
</thead>
<tbody>
<tr>
<td>United States</td>
<td></td>
</tr>
<tr>
<td>$H_0$: $r =$</td>
<td>$H_0$: $r =$</td>
</tr>
<tr>
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<td>0</td>
</tr>
<tr>
<td>0.7750</td>
<td>0.2945</td>
</tr>
<tr>
<td>121.65 [58.96]</td>
<td>85.93 [58.96]</td>
</tr>
<tr>
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<td>1</td>
</tr>
<tr>
<td>0.5530</td>
<td>0.1737</td>
</tr>
<tr>
<td>50.05 [39.08]</td>
<td>39.53 [39.08]</td>
</tr>
<tr>
<td>2</td>
<td>2</td>
</tr>
<tr>
<td>0.1485</td>
<td>0.0727</td>
</tr>
<tr>
<td>11.41 [22.95]</td>
<td>14.15 [22.95]</td>
</tr>
<tr>
<td>3</td>
<td>3</td>
</tr>
<tr>
<td>0.0741</td>
<td>0.0304</td>
</tr>
<tr>
<td>3.69 [10.56]</td>
<td>4.11 [10.56]</td>
</tr>
<tr>
<td>United Kingdom</td>
<td></td>
</tr>
<tr>
<td>$H_0$: $r =$</td>
<td>$H_0$: $r =$</td>
</tr>
<tr>
<td>0</td>
<td>0</td>
</tr>
<tr>
<td>0.5247</td>
<td>0.2292</td>
</tr>
<tr>
<td>77.80 [58.96]</td>
<td>72.07 [43.84]</td>
</tr>
<tr>
<td>1</td>
<td>1</td>
</tr>
<tr>
<td>0.4234</td>
<td>0.1755</td>
</tr>
<tr>
<td>41.35 [39.08]</td>
<td>33.54 [26.70]</td>
</tr>
<tr>
<td>2</td>
<td>2</td>
</tr>
<tr>
<td>0.2532</td>
<td>0.0321</td>
</tr>
<tr>
<td>14.37 [22.95]</td>
<td>4.98 [13.31]</td>
</tr>
<tr>
<td>3</td>
<td>3</td>
</tr>
<tr>
<td>0.0013</td>
<td>0.0010</td>
</tr>
<tr>
<td>0.06 [10.56]</td>
<td>0.15 [2.71]</td>
</tr>
</tbody>
</table>

Notes: Reported are the test statistics of the parsimonious models reported in Tables 3a–3d. $Q(r)$ is the likelihood ratio statistic for determining the number of cointegrating vectors, $r$, in the I(1) analysis. 90% critical values shown in curly brackets {} are from Tables 15.3 and 15.4 of Johansen (1995).

Table 3a
Cointegrating vectors and adjustment coefficients

<table>
<thead>
<tr>
<th>United States Annual 1950 to 1997</th>
</tr>
</thead>
<tbody>
<tr>
<td>Markup equation</td>
</tr>
<tr>
<td>$\Delta \mu$</td>
</tr>
<tr>
<td>-0.240</td>
</tr>
<tr>
<td>(-4.0)</td>
</tr>
<tr>
<td>1</td>
</tr>
<tr>
<td>$\Delta_{rer}X$</td>
</tr>
<tr>
<td>-0.097</td>
</tr>
<tr>
<td>(-1.0)</td>
</tr>
<tr>
<td>2</td>
</tr>
<tr>
<td>$\Delta p$</td>
</tr>
<tr>
<td>-0.454</td>
</tr>
<tr>
<td>(-8.1)</td>
</tr>
<tr>
<td>3</td>
</tr>
<tr>
<td>$\Delta_{rpv}$</td>
</tr>
<tr>
<td>1.641</td>
</tr>
<tr>
<td>(0.4)</td>
</tr>
<tr>
<td>1</td>
</tr>
<tr>
<td>$0.10\mu + 0.319_{rer} + 1.483_{}\Delta p - 0.001927$</td>
</tr>
<tr>
<td>(0.171)</td>
</tr>
<tr>
<td>{0.063}</td>
</tr>
<tr>
<td>{0.188}</td>
</tr>
<tr>
<td>{0.00032}</td>
</tr>
<tr>
<td>(-0.1)</td>
</tr>
<tr>
<td>(-1.2)</td>
</tr>
<tr>
<td>(-1.3)</td>
</tr>
<tr>
<td>(9.4)</td>
</tr>
<tr>
<td>2</td>
</tr>
<tr>
<td>$0.037_{rpv}$</td>
</tr>
<tr>
<td>(0.195)</td>
</tr>
<tr>
<td>{0.003}</td>
</tr>
</tbody>
</table>


Implicit relationships Implicit long-run markup–inflation relationship: $\mu_2 + 1.124\Delta p$ where the implicit markup is $\mu_2 = [p - 0.242_{rpv}] - 0.758ulc$.

Tests for serial correlation
- LM(1) $\chi^2(16)=10.76$, $p$-value=0.82
- LM(4) $\chi^2(16)=11.50$, $p$-value=0.78

Test for normality
- Doornik–Hansen test for normality: $\chi^2(8)=11.29$, $p$-value=0.19

Likelihood ratio test of cointegrating vector restrictions
- Estimated coefficients for $RPV$ in cointegrating vector 1 and for the markup, ‘real exchange rate’ and trend in cointegrating vector 2 accepted $\chi^2(2)=1.61$, $p$-value=0.45.
- Reported in () are $t$-statistics and in {} are standard errors of the estimate. Normalized cointegrating vectors reported after imposing 2 vectors on the cointegration space.

The annual estimation began with two lags of the core variables, the contemporaneous and first lag of the unemployment variable and a trend in the cointegrating space. The quarterly estimation began in a similar fashion except with four lags of the core variables and four lags of the unemployment variable. The parsimonious form of the model was sought with the trend variable and the longest lags of the core variables and the unemployment variable eliminated if insignificant. Spike dummies were introduced for periods with residuals greater than 3 standard errors from zero.
There are however some minor divergences evident between annual and quarterly systems for the United Kingdom. For example, for quarterly data, the annualised ‘inflation cost’ coefficient on inflation is 0.904, which contrasts with 0.299 derived from annual data. The sensitivity of the United Kingdom results to sampling frequency may be explained by the different price indexes used for the two different frequencies and the presence of a trend in the proportion of ‘non-traded GDP’ in the economy (in the annual data estimates). This point may also explain the presence of the significant trend terms in the United States cointegrating vectors.

The third question we addressed in the introduction is whether inflation acts as a proxy for the RPV measure. This amounts to asking the question whether RPV could replace inflation in the markup equation if all three variables appeared in the same system. We test and can accept for all the data sets the restrictions that RPV is not significant in the inflation–markup long-run relationship and the markup is not significant in the inflation–RPV long-run relationship. This identifies the inflation–markup and inflation–RPV pairs as the long-run relationships in the data. The chi-squared tests for excluding the variables from each long-run relationship are reported in Tables 3a–3d.

We note that we could alternatively have identified the cointegrating vectors by restricting the inflation term to zero in the first cointegrating vector, leaving the identifying restriction for the second cointegrating vector unchanged. Table 4 supports our argument in favour of the first identifying scheme instead of the alternative. Table 4 reports a reformulation of the systems in Tables 3a–3d without imposing identifying restrictions (see notes to Table 4). In order for the alternative identifying scheme to be appropriate, we require that mu is insignificant in Cointegrating Vector 1.

---

### Table 3b
Cointegrating vectors and adjustment coefficients

<table>
<thead>
<tr>
<th>United States Quarterly June 1968 to June 2001</th>
</tr>
</thead>
<tbody>
<tr>
<td>Markup equation</td>
</tr>
<tr>
<td>Δmu</td>
</tr>
<tr>
<td>1</td>
</tr>
<tr>
<td>(-4.3)</td>
</tr>
<tr>
<td>2</td>
</tr>
<tr>
<td>(1.8)</td>
</tr>
</tbody>
</table>

---

Core and pre-determined variables

Implicit relationships Implicit long-run markup–inflation relationship: \( mu_2 + 4.850Δp \) where the implicit markup is \( mu_2 = [p - 0.214px] - 0.786ucl \).

Tests for serial correlation
LM(1) \( \chi^2(16) = 22.05, p\text{-value}=0.14 \)
LM(4) \( \chi^2(16) = 23.89, p\text{-value}=0.09 \)

Test for normality
Doornik–Hansen test for normality: \( \chi^2(8) = 10.71, p\text{-value}=0.22 \)

Likelihood ratio test of cointegrating vector restrictions
Estimated coefficients are zero for RPV in cointegrating vector 1 and for the markup and ‘real exchange rate’ in cointegrating vector 2 accepted \( \chi^2(1) = 0.03, p\text{-value}=0.87 \).

See also the notes at the bottom of Table 3a.

There are however some minor divergences evident between annual and quarterly systems for the United Kingdom. For example, for quarterly data, the annualised ‘inflation cost’ coefficient on inflation is 0.904, which contrasts with 0.299 derived from annual data. The sensitivity of the United Kingdom results to sampling frequency may be explained by the different price indexes used for the two different frequencies and the presence of a trend in the proportion of ‘non-traded GDP’ in the economy (in the annual data estimates). This point may also explain the presence of the significant trend terms in the United States cointegrating vectors.

The third question we addressed in the introduction is whether inflation acts as a proxy for the RPV measure. This amounts to asking the question whether RPV could replace inflation in the markup equation if all three variables appeared in the same system. We test and can accept for all the data sets the restrictions that RPV is not significant in the inflation–markup long-run relationship and the markup is not significant in the inflation–RPV long-run relationship. This identifies the inflation–markup and inflation–RPV pairs as the long-run relationships in the data. The chi-squared tests for excluding the variables from each long-run relationship are reported in Tables 3a–3d.

We note that we could alternatively have identified the cointegrating vectors by restricting the inflation term to zero in the first cointegrating vector, leaving the identifying restriction for the second cointegrating vector unchanged. Table 4 supports our argument in favour of the first identifying scheme instead of the alternative. Table 4 reports a reformulation of the systems in Tables 3a–3d without imposing identifying restrictions (see notes to Table 4). In order for the alternative identifying scheme to be appropriate, we require that mu is insignificant in Cointegrating Vector 1.

---

21 We are grateful for correspondence with Neil Ericsson that drew our attention to this alternative restriction that generated the subsequent analysis.
and significant in Cointegrating Vector 2. Table 4 shows that this combination of restrictions cannot be supported for any of the four data sets. For three of the data sets we reject the restriction that $\mu$ is equal to zero in Cointegrating Vector 1 and accept the restriction that the coefficient on $\mu$ is equal to zero in Cointegrating Vector 2. For the fourth data set (annual United Kingdom) we reject the hypothesis that the coefficient on $\mu$ is zero in both cointegrating vectors. This is evidence in favour of considering the relationship between inflation and the markup to be the ‘primitive’ one, with the alternative identifying scheme being a rewriting of the first.

These results suggest that inflation cannot be replaced by $RPV$ in the inflation–markup relationship; nor can the markup be included in the inflation–$RPV$ long-run relationship. By implication, the explanations provided by Bénabou (1988, 1992), and Diamond (1993) of the inflation–markup relationship (that rely on the impact of $RPV$ on competition) do not explain our long-run inflation–markup results.\footnote{Acceptance of these restrictions does not preclude the possibility that these explanations may explain any short-run relationship between the variables through the dynamics of the system.}

Turning to the fourth question posed in the introduction, we have shown the existence of two cointegrating vectors that have the expected signs and functional form corresponding to the long-run relationships described in Section 3. However, the existence of two distinct cointegrating vectors does not necessarily imply that single-equation estimation of each relationship is valid. Such a simplification requires weak-exogeneity conditions to hold and these requirements can be investigated using our results in Tables 3a–3d by noting how the error-correction terms enter the system.

For example, our results based on annual data for the United States (Table 3a) and the annual and quarterly data for the United Kingdom (Tables 3c and 3d) shows that the second cointegrating
relationship enters significantly only into the $RPV$ equation (this is the only adjustment coefficient that is significantly different from zero on the basis of the $t$-statistics in brackets beneath the estimated coefficients). Taken together with the result that the $RPV$ variable does not enter the first cointegrating relationship (a result we confirmed in Table 4), we expect that if we examine a bivariate system consisting of inflation and $RPV$, the cointegrating relationship between inflation and $RPV$ would not enter with a significant coefficient in the inflation equation. This would establish the weak exogeneity status of inflation in this bivariate system and justifies the estimation of the relationship between $RPV$ and inflation (with $RPV$ being the dependent variable) using single-equation methods.

The quarterly United States results in Table 3b (where the second cointegrating relationship does enter with a significant coefficient in the inflation equation) is the only exception. Here inflation is unlikely to be weakly exogenous in this bivariate system and we can verify this formally.

Table 3d
Cointegrating vectors and adjustment coefficients
United Kingdom Quarterly June 1964 to March 2001

<table>
<thead>
<tr>
<th>Cointegrating vector</th>
<th>Markup equation</th>
<th>'RER' equation</th>
<th>Inflation equation</th>
<th>$RPV$ equation</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>$\Delta mu$</td>
<td>$\Delta rer_M$</td>
<td>$\Delta p$</td>
<td>$\Delta rpv$</td>
</tr>
<tr>
<td>2</td>
<td>$(-0.136, 0.020, -0.074)$</td>
<td>$(0.020, 0.059, -0.003)$</td>
<td>$(1.800, 24.035)$</td>
<td>$(1.0, 0.023)^{rpv,}$</td>
</tr>
</tbody>
</table>

Core and pre-determined variables
Three lags of the core variables: the markup on unit labour costs, the ‘real exchange rate’, inflation and $RPV$. Predetermined variables: dummies for March 1973, March and December 1975, September 1979 and June 1991 and the change in the logarithm of the unemployment rate lagged one period. Number of observations, 148.

Implicit relationships
Implicit long-run markup–inflation relationship: $\mu = 3.617 \Delta p$ where the implicit markup is $\mu = p - 0.850 ulc - 0.150 pm$.

Tests for serial correlation
LM(1) $\chi^2(16) = 21.09$, $p$-value = 0.17
LM(4) $\chi^2(16) = 11.71$, $p$-value = 0.76

Test for normality
Doornik–Hansen test for normality: $\chi^2(8) = 15.08$, $p$-value = 0.06

Likelihood ratio test of cointegrating vector restrictions
(a) Estimated coefficients are zero for $RPV$ and trend in cointegrating vector 1 and for the markup RER and trend in cointegrating vector 2 accepted, $\chi^2 = 1.63$, $p$-value = 0.65.
(b) Estimated coefficients are zero for $RPV$ in cointegrating vector 1 and for the markup and RER in cointegrating vector 2 accepted $\chi^2(1) = 1.58$, $p$-value = 0.21.

See also the notes at the bottom of Table 3a.

relationship enters significantly only into the $RPV$ equation (this is the only adjustment coefficient that is significantly different from zero on the basis of the $t$-statistics in brackets beneath the estimated coefficients). Taken together with the result that the $RPV$ variable does not enter the first cointegrating relationship (a result we confirmed in Table 4), we expect that if we examine a bivariate system consisting of inflation and $RPV$, the cointegrating relationship between inflation and $RPV$ would not enter with a significant coefficient in the inflation equation. This would establish the weak exogeneity status of inflation in this bivariate system and justifies the estimation of the relationship between $RPV$ and inflation (with $RPV$ being the dependent variable) using single-equation methods.

The quarterly United States results in Table 3b (where the second cointegrating relationship does enter with a significant coefficient in the inflation equation) is the only exception. Here inflation is unlikely to be weakly exogenous in this bivariate system and we can verify this formally.

Table 4
$\chi^2(1)$ tests of significance of markup in cointegration vectors

<table>
<thead>
<tr>
<th>Cointegrating Vector 1</th>
<th>Cointegrating Vector 2</th>
</tr>
</thead>
<tbody>
<tr>
<td>US Annual</td>
<td>13.62, $p$-value = 0.00</td>
</tr>
<tr>
<td>US Quarterly</td>
<td>5.65, $p$-value = 0.02</td>
</tr>
<tr>
<td>UK Annual</td>
<td>12.09, $p$-value = 0.00</td>
</tr>
<tr>
<td>UK Quarterly</td>
<td>17.28, $p$-value = 0.00</td>
</tr>
</tbody>
</table>

Note: The systems reported in Tables 3a–3d are reformulated as

Cointegrating Vector 1: $v_0 \mu + v_1 \Delta p + v_2 + v_3 \text{trend}$
Cointegrating Vector 2: $\omega_0 \mu + \omega_1 \Delta p + \omega_2 \\Delta rpv + \omega_3 \text{trend}$

The reported individual $\chi^2(1)$ tests are that $v_0 = 0$ in cointegrating vector 1 and $\omega_0 = 0$ in cointegrating vector 2.
Even though inflation is not weakly exogenous in any of the four variable systems that we examine, our intuition based on the results from our system estimates is correct in that we can establish for the data used in Tables 3a−3d that inflation is weakly exogenous in a bivariate system comprising inflation and RPV. Furthermore, inflation is not weakly exogenous in a bivariate system using the quarterly United States data from Table 3b as expected.23

In order to highlight the impact of the endogeneity of inflation in this inflation–RPV relationship, Table 5 reports the single-equation and system estimates of the inflation–RPV long-run relationship for all four data sets. The single equation estimates are derived from estimating the Bewley (1979) transform of the ADL(m,n) model given by:

$$rp_v = c + \vartheta \Delta p_t + \sum_{i=0}^{n-1} \alpha_i \Delta rp_{v_{t-1}} + \sum_{j=0}^{n-1} \beta_j \Delta^2 p_{t-j}$$

The Bewley transform is estimated using 2SLS with the lagged endogenous variables and current and lagged exogenous variables as instruments. A trend is included if significant. The system estimates are from Tables 3a to 3d normalized on rp_v.

<table>
<thead>
<tr>
<th>Data Set</th>
<th>Single equation estimate</th>
<th>Systems estimate</th>
</tr>
</thead>
<tbody>
<tr>
<td>United States Annual</td>
<td>30.200</td>
<td>27.036</td>
</tr>
<tr>
<td></td>
<td>{11.200}</td>
<td>{5.274}</td>
</tr>
<tr>
<td>United States Quarterly</td>
<td>98.548</td>
<td>147.152</td>
</tr>
<tr>
<td></td>
<td>{23.817}</td>
<td>{23.986}</td>
</tr>
<tr>
<td>United Kingdom Annual</td>
<td>21.849</td>
<td>24.448</td>
</tr>
<tr>
<td></td>
<td>{5.332}</td>
<td>{5.134}</td>
</tr>
<tr>
<td>United Kingdom Quarterly</td>
<td>51.795</td>
<td>44.281</td>
</tr>
<tr>
<td></td>
<td>{22.095}</td>
<td>{21.392}</td>
</tr>
</tbody>
</table>

Note: Reported in brackets {} are standard errors. The single equation estimates of the cointegrating parameter, \( \vartheta \) and its standard error are derived from estimating the Bewley (1979) transform of the ADL(m,n) model given by:

Even though inflation is not weakly exogenous in any of the four variable systems that we examine, our intuition based on the results from our system estimates is correct in that we can establish for the data used in Tables 3a−3d that inflation is weakly exogenous in a bivariate system comprising inflation and RPV. Furthermore, inflation is not weakly exogenous in a bivariate system using the quarterly United States data from Table 3b as expected.23

In order to highlight the impact of the endogeneity of inflation in this inflation–RPV relationship, Table 5 reports the single-equation and system estimates of the inflation–RPV long-run relationship for all four data sets. The single equation estimates are derived from estimating the Bewley (1979) transform of the ADL(m,n) model and the system results are taken from the results reported in Tables 3a−3d. Consistent with our finding that inflation is endogenous in the quarterly United States data, the single equation estimates differ from the system estimate by approximately 50%. For the remaining data sets, where inflation is weakly exogenous, such divergences do not appear.

Based on the results presented in Tables 3a−3d on the adjustment coefficients and the associated t-statistics we may deduce that single-equation estimation of the inflation–markup relationship is unlikely to be valid, since neither inflation nor the markup on unit labour costs are weakly exogenous in the four-variable system.

Finally we turn to the graphical illustration of our annual results.24 Graphs 1a and 1b show as solid lines the long-run relationships between inflation and the markup and between inflation and RPV and are labelled as LR(1) and LR(2) respectively.25 Shown as dots on the graphs are the actual values for inflation and RPV along with the estimated markup from the system analysis. Marked with crosses are the observations corresponding to the dummy variables in the system analysis.

We have shown that paying due regard to the integration properties of the data, the interlinkages between the relationships and the exogeneity status of the variables are all important for a proper analysis of the two interrelated themes in the literature. In particular, only a complete

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23 These results are available from the authors.
24 Equivalent graphs of the quarterly results are reported in Banerjee et al. (2002).
25 For ease of interpretation, the graphs show annual inflation as the change in the natural logarithm of the price index multiplied by 100.
systems analysis is capable of judging the validity of the simplifications adopted in much of the existing empirical literature. Our results are estimated more consistently than those of the single equation studies and permit short-term dynamic interactions that the former studies ignore.

5. Conclusions

In this paper we argue that although the empirical literatures on the relationships between inflation and the markup and between inflation and $RPV$ were developed separately they should be considered together in order to address five important questions. The questions are: ‘What relationships exist between the variables in our system?’, ‘Are they long-run or short-run relationships?’, ‘Are there two relationships (between inflation and the markup on the one hand and $RPV$ and inflation on the other) or one relationship between all three variables?’, ‘Is a system estimate more appropriate than single equation estimation?’ and ‘Can we determine whether demand- or supply-side arguments are more consistent with the empirical relationships we have identified?’
The first two questions can be answered positively, since we reconfirm the findings of a negative relationship between inflation and the markup and a positive relationship between inflation and RPV in the long run. In answer to the third question there are two relationships between the three variables; these are the inflation–markup and inflation–RPV relationships. RPV is not admissible in the inflation–markup equation, and that the markup can be excluded from the inflation–RPV relationship in the long run. We answer our fourth question by concluding that single equation estimation of the inflation–RPV and inflation–markup relationships are unlikely to be valid although it cannot be categorically ruled out.

Our conclusions are that inflation is associated with greater uncertainty as measured by an increase in RPV and a reduction in the markup in both the US and the UK data over our sample period. However, although price dispersion increases, this is not the cause of the reduction in the markup. The systems approach indicates that RPV cannot substitute for inflation, indicating that inflation is not a proxy for price dispersion. This answers our fifth question, since we find evidence that is inconsistent with the demand-side explanation for the reduction of market power when inflation increases, but is consistent with supply-side reasoning. The fact that these results can be supported using data at several frequencies for more than one country is encouraging, although more research is required to establish the generality of the result. Our paper offers an initial step towards integrating the two literatures on inflation–RPV and the inflation–markup that have previously been investigated separately.

Appendix A. Interpreting the Estimated Markup

We can interpret \( \mu_2 \) as the markup of the price of non-traded GDP on unit labour costs in the following way. The GDP deflator may be defined as the weighted average of the implicit price deflators for exports and domestically consumed GDP, such that:

\[
p_{\text{GDP}} = \omega p_x + (1-\omega) p_D
\]

where \( p_D \) is the domestically consumed GDP implicit price deflator and \( \omega \) is the share of exports in GDP. Substituting for the GDP deflator in Eq. (3) using Eq. (A1) provides the following expression:

\[
\mu_2 = [(1-\omega)p_D - (\delta - \omega)p_x] - (1-\delta)ulc
\]

If the impact of export prices on the GDP deflator is confined only to the direct impact of exports in the price index then the estimate of \( \delta \) will equal the share of exports in GDP, \( \omega \), and the estimated markup, \( \mu_2 \), collapses to \( p_D - ulc \). That is, the estimated markup is of the price of domestically consumed GDP on unit labour costs and will not be affected in the long run by changes in \( \text{rer}_X \). However, changes in \( \text{rer}_X \) will affect in the long-run the markup, \( p_D - ulc \). Alternatively, if export prices have indirect effects on the price of domestically consumed GDP then \( \delta > \omega \) and \( \text{rer}_X \) will have an impact upon the markup \( p_D - ulc \) and \( p_{\text{GDP}} - ulc \) in the long-run.\(^{26}\) In this case, if we denote by ‘non-traded GDP’ that part of GDP not subject to the direct and indirect effects of changes in export prices then the estimated markup, \( \mu_2 \), represents the markup of the price of non-traded GDP on unit labour costs.

\(^{26}\) Indirect effects of higher export prices on the price of domestically consumed GDP may operate through ‘supply’ and ‘demand’ channels. The ‘supply’ effect is due to domestic producers diverting sales overseas reducing the supply of goods and services domestically leading to an increase in the markup of \( p_D - ulc \). The ‘demand’ effect occurs if real export prices increase due to a depreciation of the exchange rate leading to a similar increase in real import prices and therefore to an increase in the price of domestically produced substitutes.
References


