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MEASUREMENT OF INFLATION: AN ALTERNATIVE APPROACH

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The stochastic approach to index numbers has attracted renewed attention in recent times (e.g., Clements and Izan, 1981 and 1987; Diewert, 1995; Giles and McCann, 1994; and Selvanathan and Rao, 1994). One of the attractions of this approach is that it provides standard errors for the index numbers. This paper reviews the stochastic approach and extends the existing work by presenting an alternative approach to measure the rate of inflation. This approach has been demonstrated using consumption expenditure data for three countries, Australia, the United Kingdom (UK) and the United States (US).

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I. Introduction

A familiar problem in index number theory is to measure the overall rate of inflation given the two sets of prices of n goods $p_{1t}, p_{2t}, \dots p_{nt}$, and $p_{1t-1}, p_{2t-1}, \dots p_{nt-1}$ in two periods t and t-1. An estimator of the rate of inflation and its standard error are in practical use for wage negotiations, wage indexation and so on. While the utility-based *functional approach* to index number theory (see Diewert 1981, for a survey) provides an estimator of the rate of inflation, the alternative approach, the *stochastic approach*, provides in addition, its standard error as well. Under the

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stochastic approach, the proportionate change in each individual price is taken to be equal to the underlying rate of inflation plus other components, which are random and non-random (see Clements and Izan 1987, and Selvanathan and Rao 1994). If we have *n* prices, then the rate of inflation can be estimated by taking some form of average of the *n* price changes. The stochastic approach can be viewed as a signal extraction problem. To illustrate, consider the simplest case whereby each of the *n* proportionate price changes is the sum of the underlying rate of inflation and an independent random component. Here each observed price change is a reading on the rate of inflation 'contaminated' by the random term. The averaging of the price changes serves to eliminate as much as possible of the contamination and leaves an estimate of the underlying signal, the rate of inflation.

Although the stochastic approach is less well known than the functional approach, it has a long history going back to Edgeworth (see Frisch 1936 for references). In addition, this approach has recently attracted renewed attention (Clements and Izan 1981, 1987; Crompton 2000; Diewert 1995; Giles and McCann 1994; Selvanathan and Rao 1994; and Selvanathan and Selvanathan 2004). The attraction of the stochastic approach is that it provides standard errors for the price indices. These standard errors increase with the degree of relative price variability. This agrees with the intuitive notion that when the individual prices move very disproportionately, the overall rate of inflation is less well defined. The availability of the standard errors allows us to construct confidence intervals for the true rate of inflation. These interval estimates can be used in a number of practical situations (e.g., wage negotiations).

The discussions in this paper deal exclusively with the measurement of price inflation. However, the methodology could equally well be applied to quantities and be used for the measurement of real income, total factor productivity, and so on. The framework could also be used to test the purchasing power parity hypothesis (see Miller 1984) and to extend the analysis of Divisia monetary aggregates (Barnett 1981, Section 7.11).

Using the stochastic approach, Clements and Izan (1987) derive an estimator for the rate of inflation in terms of Divisia indices and its standard error by decomposing the proportionate price changes into various components. This paper, whilst also using the stochastic approach framework proposes an alternative way to derive an estimator of the overall rate of inflation and its standard error. This alternative way uses a regression model that involves expenditures rather than price changes (as in Clements and Izan) to derive an estimator of the overall rate of inflation in terms of the well-known Laspeyres price index and its standard error.

The results show that (i) the estimator of the rate of inflation is related to the well-known Laspeyres and Paasche index numbers and is approximately equal to the Divisia estimator of Clements and Izan (1987) and (ii) the standard error of the estimator is linked to the variability of the relative prices. When there are larger changes in relative prices, the standard error will be higher.

Our estimator and its standard error have three major attractions. The first is that they are simple to evaluate and have clear economic interpretation like Clements and Izan's Divisia counterparts. The second, they require only base-period budget shares for evaluation whereas the Divisia counterparts require both the current (usually unknown) and base-period budget shares which will cause some practical difficulties when evaluating the current rate of inflation. The third is that our estimator and standard error are approximately equal to the Divisia counterparts when the change in prices and budget shares are small.

The organisation of this paper is as follows. In Section II, we present the Divisia results of Clements and Izan (1981, 1987). In Section III we present an alternative approach. In the following section, we present an empirical application. Finally in Section V, we present the concluding comments.

II. Divisia estimator of inflation

Let p_{it} be the price of commodity i (i=1,...,n) in period t (t=1,…,T) and Dp_{it} = log p_{it} - log p_{it-1} be the price log-change. For each period, let each price log-change be made up of a systematic part α_t and a zero-mean random component u_{it}; that is,

$$Dp_{it} = \alpha_t + u_{it}, \quad i=1,...,n;\ t=1,\dots,T. \tag{1}$$

As $E[Dp_{it}] = \alpha_t$, we interpret α_t as the common trend in all prices. The random term u_{it} is assumed to be independent over commodities and have a common variance σ_t^2; that is,

$$\mathrm{Cov}[u_{it}, u_{jt}] = \sigma_t^2 \delta_{ij}, \quad i,j=1,...,n, \tag{2}$$

where δ_{ij} is the Kronecker delta.

From (1) we can see that $u_{it} = Dp_{it} - \alpha_t$ is the change in the i^{th} price deflated by the common trend in all prices; i.e., u_{it} is the change in the i^{th} relative price. Hence (2) is interpreted as saying that the relative prices are independent and have a common variance. Furthermore, the assumption that $E[u_{it}] = 0$ means that all relative price changes have an expected value of zero. Under these assumptions, the best linear unbiased estimator of α_t is

$$\hat{\alpha}_t = \frac{1}{n}\sum_{i=1}^{n} Dp_{it},$$

which is just the unweighted average of the n price log-changes. Also we have

$$\text{var}\ \hat{\alpha}_t = \frac{1}{n}\ \sigma_t^2. \tag{3}$$

The variance σ_{t}^2 can be estimated unbiasedly by

$$\hat{\sigma}_{\mathrm{t}}^2 = \frac{1}{n-1}\sum_{i=1}^{n} (Dp_{it} - \hat{\alpha}_t)^2. \tag{4}$$

From (3) and (4) we see that when there is substantial variation in relative prices, the sampling variance of $\hat{\alpha}_t$ will be higher. This agrees with the intuitive notion that the meaning of the overall rate of inflation becomes less well defined when there are large changes in relative prices.

Given the above interpretations, assumption (2), together with $E[u_{it}] = 0$, is obviously very stringent. We now extend the model to relax these assumptions. We continue to take the relative price changes as having expectation zero and being independent, but we now replace (2) with

$$\text{Cov}[u_{it}, u_{jt}] = \frac{\lambda_t^2}{\overline{w}_{it}}\ \delta_{ij},\ i,j=1,...,n, \tag{5}$$

where λ_t^2 is a constant with respect to commodities; and $\overline{w}_{it}$ is the arithmetic average of the budget share of i during the two periods t and t-1. That is, $\overline{w}_{it} = ½(w_{it} + w_{it-1})$. Under this assumption, the variance of the change in the relative price of i is inversely proportional to $\overline{w}_{it}$. This means that the variability of a relative price falls as the commodity becomes more important in the consumer's budget.

We write (1) in vector form as

$$\boldsymbol{Dp}_{\mathbf{t}} = \alpha_t\, \iota + \boldsymbol{u}_t, \qquad t=1,\dots,T, \tag{6}$$

where $\boldsymbol{Dp_t} = [Dp_{it}]$; $\iota = [1 \ldots 1]'$; and $\boldsymbol{u}_t = [u_{it}]$. Under (5), the $n \times n$ covariance matrix of $\boldsymbol{u}_t$ is

$$\text{Var}\ \boldsymbol{u}_t = \lambda_t^2 \overline{W}_t^{-1}, \tag{7}$$

where $\overline{\boldsymbol{W}}_t = \text{diag}\,[\overline{w}_{1t} \ldots \overline{w}_{nt}]$. Application of GLS to (6) under (7) gives

$$\tilde{\alpha}_t = (\iota' \overline{\boldsymbol{W}}_t \iota)^{-1} \iota' \overline{\boldsymbol{W}}_t \boldsymbol{Dp_t}$$

Since $\iota' \overline{\boldsymbol{W}}_t \iota = \sum_{i=1}^n \overline{w}_{it} = 1$ and $\iota' \overline{\boldsymbol{W}}_t \boldsymbol{Dp}_t = \sum_{i=1}^n \overline{w}_{it} Dp_{it}$, this simplifies to

$$\tilde{\alpha}_t = \sum_{i=1}^n \overline{w}_{it} Dp_{it}. \tag{8}$$

This expression is identical to the finite-change form of the well-known Divisia price index (Theil 1975/76).

The sampling variance of $\tilde{\alpha}_t$ is $\lambda_t^2 (\iota' \overline{\boldsymbol{W}}_t \iota)^{-1} = \lambda_t^2$; that is,

$$\text{Var}\ \tilde{\alpha}_t = \lambda_t^2. \tag{9}$$

This variance can be estimated unbiasedly by

$$\frac{1}{n-1} (\boldsymbol{Dp_t} - \tilde{\alpha}_t \iota)' \overline{\boldsymbol{W}}_t (\boldsymbol{Dp_t} - \tilde{\alpha}_t \iota) = \frac{1}{n-1} \sum_{i=1}^n \overline{w}_{it} (Dp_{it} - \tilde{\alpha}_t)^2,$$

so that

$$\tilde{\lambda}_t^2 = \frac{1}{n-1} \sum_{i=1}^n \overline{w}_{it} (Dp_{it} - \tilde{\alpha}_t)^2. \tag{10}$$

We write (9) as

$$\text{Var}\ \tilde{\alpha}_t = \frac{1}{n-1} \Pi_t, \tag{11}$$

where $\Pi_t = \sum_{i=1}^n \overline{w}_{it} [Dp_{it} - \tilde{\alpha}_t]^2$ is the finite-change form of the Divisia variance of relative price changes (Theil 1975/76). This Π measures the degree to which prices move disproportionately; $\Pi = 0$ only if all prices change proportionately, that is, if there are no changes in relative prices. From (10) and (11) again we see that the sampling variance of the estimator of inflation will be higher the larger the relative price movements.

III. An alternative approach

Let $p_{io}q_{io}$ be expenditure on commodity i (i=1,...,n) in the base period o and let $p_{it}q_{io}$ be the base-period consumption of i (q_{io}) valued at the current period prices (p_{it}). Consider a regression of ($p_{it}q_{io}$ - $p_{io}q_{io}$) on $p_{io}q_{io}$:

$$(p_{it}q_{io} - p_{io}q_{io}) = \gamma_t p_{io}q_{io} + \varepsilon_{it}, \qquad i = 1, \ldots, n, \tag{12}$$

where γ_t is a constant with respect to commodities; and ε_{it} is a disturbance term. We assume

$$\mathrm{E}[\varepsilon_{it}] = 0, \quad \mathrm{cov}[\varepsilon_{it}, \varepsilon_{jt}] = \sigma_t^2 p_{io}q_{io}\delta_{ij}\,, \tag{13}$$

where δ_{ij} is the Kronecker delta. In the next section, we shall empirically test the validity of the error structure (13). In model (12), γ_t is interpreted as the common trend in all prices or the overall rate of inflation over the periods o and t. To see this, we divide both sides of (12) by $p_{io}q_{io}$ and use (13) to give

$$\gamma_t = \mathrm{E}\left[\frac{p_{it}}{p_{io}} - 1\right].$$

We divide both sides of (12) by $\sqrt{p_{io}\, q_{io}}$ to give

$$y_{it} = \gamma_t x_{io} + \eta_{it}\,, \quad i=1,\ldots,n, \tag{14}$$

where $y_{it} = (p_{it} - p_{io})\sqrt{q_{io}/p_{io}}$; $x_{io} = \sqrt{p_{io}q_{io}}$ and $\eta_{it} = \varepsilon_{it} \,/\, \sqrt{p_{io}q_{io}}$. It follows from (13) that E[η_{it}] = 0, cov[η_{it},η_{jt}] = $\sigma_t^2\delta_{ij}$. Thus we can now apply least squares to (14) to get the BLUE of γ_t,

$$\hat{\gamma}_t = \left[\frac{\Sigma_{i=1}^n x_{io} y_{io}}{\Sigma_{i=1}^n x_{io}^2}\right] = \left[\sum_{i=1}^{n} w_{io} \frac{p_{it}}{p_{io}}\right] - 1\,, \tag{15}$$

where $w_{io} = p_{io}q_{io}/M_o$ is the budget share of i in period o; and $M_o = \sum_{i=1}^{n} p_{io}q_{io}$ is total expenditure in period o. Obviously the first right-hand term of the above equation is the well-known *Laspeyres price index* which is a weighted average of the n price ratios, the weights being the base-period budget shares. Thus equation (15) shows that the rate of inflation estimator $\hat{\gamma}_t$ differs from the Laspeyres price index by unity. This result is very attractive as it expresses the estimator of the rate of inflation in terms of the commonly used Laspeyres price index and it requires only the base-period budget shares for its evaluation.

Now we shall show that the estimator of γ_t given by (15) is approximately equal to the Divisia estimator of α_t given in (8). To show this we write (15) in the form

$$\hat{\gamma}_t = \sum_{i=1}^{n} w_{io}\left[\frac{p_{it}}{p_{io}} - 1\right], \tag{16}$$

where we have used $\sum_{i=1}^{n} w_{io} = 1.$

If the periods o and t are not very much apart (for example, period o is t-1 or t-2), then we could expect $x = (p_{it}/p_{io}) - 1$ to be small. As $ln\,|1 + x\,| \cong x$ for small x, we have $x \cong \log\,(p_{it}/p_{io}) = Dp_{it}^{o}$, the log-change in the i^{th} price from period o to t. Therefore, we can now write (16) in the form

$$\hat{\gamma}_t \approx \sum_{i=1}^{n} \overline{w}_{it}^{o} Dp_{it}^{o}$$

where we have replaced w_{io} by $\overline{w}_{it}^{o}$, which is the arithmetic average of w_{io} and w_{it}. Clearly, the above is exactly the same form as that of the Divisia estimator given in (8) when period $o = t$-l.

The variance of $\hat{\gamma}_t$ is given by

$$\text{Var } \hat{\gamma}_t = \frac{\sigma_t^2}{\sum_{i=1}^{n} x_{io}^2} = \frac{\sigma_t^2}{M_o}. \tag{17}$$

The parameter σ_t^2 can be estimated unbiasedly by

$$\hat{\sigma}_t^2 = \frac{1}{n-1}\sum_{i=1}^{n}(y_{it} - \hat{\gamma}_t x_{io})^2.$$

By substituting the above results together with the values of x_{io} and y_{it} in (17) we obtain

$$\text{Var } \hat{\gamma}_t = \frac{1}{n-1}\sum_{i=1}^{n} w_{io}\left[\left[\frac{p_{it}}{p_{io}}-1\right]-\hat{\gamma}_t\right]^2. \tag{18}$$

Equation (18) shows that the variance of $\hat{\gamma}_t$ increases with the degree of relative price variability. This agrees with the intuitive notion that when the individual prices move very disproportionately, the overall rate of inflation is less well defined.

Using the same arguments in the paragraph above equation (17), we can easily show that

$$\text{Var } \hat{\gamma}_t \approx \frac{1}{n-1}\Pi^o_t,$$

where

$$\Pi^o_t = \sum_{i=1}^{n} \overline{w}^o_{it}[Dp^o_{it} - \hat{\gamma}_t]^2$$

is the Divisia price variance which measures the degree to which prices move disproportionately. The above approximate variance is exactly the same as that of the variance given in (11) when period $o = t$-l.

Now we derive the link between $\hat{\gamma}_t$ and the Paasche index. Let $p_{it}q_{it}$ be the expenditure on good i in the current period t and $p_{io}q_{it}$ be the current-period consumption, q_{it} valued at base-year prices. Consider a regression $(p_{it}q_{it} - p_{io}q_{it})$ on $p_{io}q_{it}$:

$$(p_{it}q_{it} - p_{io}q_{it}) = \gamma^*_t p_{io}q_{it} + \varepsilon^*_{it}, \quad i=1, \ldots, n, \tag{19}$$

with

$$\text{E}[\varepsilon^*_{it}] = 0, \qquad \text{cov}[\varepsilon^*_{it}, \varepsilon^*_{jt}] = \sigma^{*2}_t p_{io}q_{it}\delta_{ij}. \tag{20}$$

Equations (19)-(20) are the same as (12)-(13) except that base-period consumption q_{io} in the former set of equations is replaced with current-period consumption q_{it}.

As before, by dividing both sides of (19) by $\sqrt{p_{io} q_{it}}$, it can be easily shown that the BLUE of γ_t^* is

$$\hat{\gamma}_t^* = \left[\sum_{i=1}^{n} w_{it}^o \frac{p_{it}}{p_{io}} \right] - 1, \tag{21}$$

and its variance is given by

$$\text{Var } \hat{\gamma}_t^* = \frac{1}{n-1} \sum_{i=1}^{n} w_{it}^o \left[\left[\frac{p_{it}}{p_{io}} - 1 \right] - \hat{\gamma}_t^* \right]^2, \tag{22}$$

where $w_{it}^o = p_{io} q_{it} / M_t^o$; and $M_t^o = \sum_{i=1}^{n} p_{io} q_{it}$. The expenditure M_t^o is current-period consumption valued at base-period prices, summed over all n goods; and w_{it}^o is the share of commodity i in M_t^o with $\sum_{i=1}^{n} w_{it}^o = 1$. Obviously the first right-hand term of equation (21) is the well-known *Paasche price index*. Thus equation (21) shows that the estimator of the rate of inflation also differs from the Paasche index by unity. As before, we can show that the estimator $\hat{\gamma}_t^*$ and its variance given by equations (21) and (22) are approximately equal to the Divisia estimator and its variance given by equations (8) and (11).

IV. An illustrative application

Now we present an illustrative application of the results derived in Section III by using price and expenditure annual data covering the period 1963-1996 for the three countries, Australia, the UK and the US. These data are obtained from Selvanathan and Selvanathan (2003), which originated from the *Yearbook of National Account Statistics* (United Nations: New York, various issues) and *National Accounts of OECD Countries* (OECD: Paris, various issues). The number of commodity classifications are 9 for Australia and the US and 8 for the UK. The nine commodity groups are food, clothing, housing, durables, medical, transport, education, recreation and miscellaneous. For the UK, education and recreation are combined into one group. For comparison of the performance of our alternative estimator with Divisia estimator, we consider period *o* as *t*-1. Table 1 presents the results based on the data for the three countries. For each country, the first column presents the official published rate of inflation (obtained from the DX database); the next 2 columns present the Divisia based estimate of inflation (based on equation 8), and the corresponding standard error (based on equation 11); the last 2 columns present the alternative approach based estimate of inflation (based on equation 16), and the corresponding standard error (based on equation 18). For example, as can be seen from columns 2-6 of the table, the Divisia estimate of inflation for Australia in 1966 is 2.94 with standard error 0.40 and is close to the corresponding alternative approach estimate for the rate of inflation of 2.97 with standard error 0.41 and to the corresponding official inflation estimate of 3.00. In general, the two sets of estimates and the respective standard errors of the rate of inflation presented in columns 3-4 and 5-6 for Australia; in columns 8-9 and 10-11 for the UK and in columns 13-14 and 15-16 for the US are reasonably close to each other and are also similar to the corresponding official published rate of inflation figure presented in columns 2, 7 and 12 supporting model (12). The last row of Table 1 presents the sample averages and the corresponding standard errors.

We now test the specification that the errors in (12) satisfy the error covariance structure (13), using the Park (1966) procedure. That is, we test the particular type of heteroscadasticity we assumed in (13), which provides the basis for our estimated results presented in Table 1. The Park procedure involves regressing the logarithm of the squared residuals from (12) on the logarithm of the regressor and testing if the slope parameter is unity. The Park test statistic follows a t-distribution with *n*-2 (*n*=9 for Australia and the US and *n*=8 for the UK) degrees of freedom. Table 2 presents the value of the Park test statistic and the corresponding *p*-values. As can be seen, in almost all years, the data from the three countries support the form of heteroscedasticity assumed in (13) and hence support the underpinning model (12).

Figure 1 presents a scatter plot of the estimate of inflation versus the corresponding standard error for the three countries; the solid line is the LS regression line. As can be seen, the standard error mostly increases along with increasing inflation. In other words, the higher the rate of inflation, the more difficult it is to measure precisely.

Figure 2 plots against time the estimated inflation and the 95% confidence band constructed using the normal distribution as (inflation $\pm 1.96 \times$ standard error). It is worth noting the jump in inflation and increase in the width of the confidence bands in mid 1970’s and in 1980 in the three countries.

Table 1. Inflation: Official rate, Divisia estimates and alternative approach estimates with standard errors

	Australia					UK					US				
Year	Official data	Divisia		Alternative Approach		Official data	Divisia		Alternative Approach		Oficial data	Divisia		Alternative Approach	
	Inflation	Inflation	SE	Inflation	SE	Inflation	Inflation	SE	Inflation	SE	Inflation	Inflation	SE	Inflation	SE
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
1963	0.50	0.46	1.04	0.48	1.05	2.10	1.75	1.01	1.79	1.03	1.20	1.42	0.18	1.43	0.18
1964	2.40	3.13	0.58	3.20	0.60	3.20	3.53	0.49	3.59	0.51	1.30	1.47	0.24	1.48	0.24
1965	4.00	3.31	0.56	3.36	0.58	4.80	5.07	0.52	5.20	0.55	1.70	1.63	0.22	1.65	0.22
1966	3.00	2.94	0.40	2.97	0.41	3.90	3.83	0.45	3.90	0.47	3.00	2.73	0.49	2.77	0.50
1967	3.20	3.49	0.49	3.55	0.51	2.40	2.54	0.40	2.57	0.41	2.80	2.48	0.53	2.52	0.55
1968	2.70	3.17	0.53	3.22	0.55	4.70	4.40	0.51	4.49	0.53	4.20	3.99	0.37	4.08	0.39
1969	2.90	3.13	0.88	3.22	0.90	5.50	5.48	0.29	5.63	0.30	5.40	4.41	0.39	4.51	0.40
1970	3.90	5.44	0.83	5.59	0.87	6.40	5.76	0.48	5.93	0.50	5.90	4.46	0.26	4.56	0.28
1971	6.10	6.47	0.78	6.69	0.85	9.40	8.22	0.36	8.57	0.39	4.30	4.28	0.44	4.36	0.46
1972	5.90	7.45	1.85	7.89	2.17	7.10	5.79	0.84	5.97	0.90	3.30	3.40	0.43	3.48	0.44
1973	9.50	8.88	1.55	9.42	1.67	9.20	7.09	2.90	7.40	3.00	6.20	7.85	1.99	8.23	2.13
1974	15.10	16.22	1.11	17.64	1.30	15.90	15.50	0.83	16.78	0.97	11.00	9.84	0.75	10.34	0.82
1975	15.10	14.31	0.89	15.43	1.02	24.20	21.28	1.13	23.73	1.38	9.10	7.74	0.58	8.05	0.64
1976	13.50	11.05	0.77	11.69	0.87	16.50	14.29	0.72	15.36	0.83	5.70	5.75	0.77	5.93	0.81
1977	12.30	9.21	0.55	9.63	0.60	15.90	15.19	7.10	16.54	8.06	6.50	6.41	0.70	6.61	0.75
1978	7.90	8.78	0.85	9.20	0.94	8.20	9.15	0.74	9.58	0.81	7.60	6.77	0.61	7.03	0.65
1979	9.10	9.91	0.58	10.42	0.63	13.50	12.70	0.82	13.54	0.93	11.30	8.74	0.77	9.16	0.84

Table 1 (cointinued). Inflation: Official rate, Divisia estimates and alternative approach estimates with standard errors

	Australia					UK					US				
Year	Official data	Divisia		Alternative Approach		Official data	Divisia		Alternative Approach		Oficial data	Divisia		Alternative Approach	
	Inflation	Inflation	SE	Inflation	SE	Inflation	Inflation	SE	Inflation	SE	Inflation	Inflation	SE	Inflation	SE
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
1980	10.10	9.24	0.44	9.69	0.48	18.00	15.61	1.16	16.88	1.35	13.50	10.40	1.15	11.00	1.28
1981	9.70	9.21	0.67	9.64	0.74	11.90	10.93	1.46	11.52	1.63	10.30	8.89	0.72	9.31	0.78
1982	11.10	10.65	0.86	11.22	0.96	8.60	8.57	1.06	8.96	1.15	6.20	6.02	0.83	6.21	0.88
1983	10.10	7.56	0.58	7.87	0.62	4.60	5.45	0.40	5.61	0.42	3.20	3.86	1.14	3.92	1.19
1984	4.00	7.37	2.93	7.65	3.10	5.00	4.82	0.66	4.93	0.70	4.30	3.93	0.55	4.02	0.57
1985	6.70	5.91	0.79	6.10	0.84	6.10	5.30	0.24	5.43	0.25	3.60	3.59	0.47	3.66	0.48
1986	9.10	8.35	0.29	8.73	0.31	3.40	3.80	0.67	3.87	0.69	1.90	2.48	0.82	2.51	0.83
1987	8.50	7.70	0.51	8.01	0.56	4.10	4.37	0.54	4.46	0.56	3.70	3.86	0.26	3.94	0.27
1988	7.20	7.28	0.60	7.57	0.64	4.90	5.07	0.46	5.19	0.49	4.00	4.12	0.56	4.21	0.58
1989	7.60	6.23	0.46	6.43	0.49	7.80	5.72	0.41	5.90	0.44	4.80	4.54	0.57	4.65	0.60
1990	7.30	4.92	0.40	5.04	0.42	9.50	5.38	1.09	5.53	1.14	5.40	5.16	0.58	5.30	0.61
1991	3.20	2.53	0.37	2.56	0.38	5.90	6.08	1.02	6.27	1.07	4.20	3.97	0.35	4.02	0.36
1992	1.00	1.85	0.34	1.89	0.35	3.70	5.81	1.09	6.02	1.16	3.00	3.04	0.49	3.09	0.51
1993	1.80	1.49	0.34	1.49	0.35	1.60	3.21	0.49	3.27	0.50	3.00	2.71	0.53	2.76	0.55
1994	1.90	1.25	0.41	1.26	0.42	2.50	2.70	0.69	2.74	0.71	2.60	2.33	0.42	2.38	0.43
1995	4.60	2.64	0.51	2.68	0.53	3.40	2.81	0.53	2.87	0.55	2.80	2.23	0.50	2.26	0.51
1996	2.60	1.65	0.42	1.67	0.43	2.40	2.66	0.37	2.70	0.37	2.90	2.06	0.50	2.11	0.51
Mean	6.58	6.27	0.90	6.56	0.97	7.54	7.05	1.50	7.43	1.68	5.00	4.61	0.68	4.75	0.72

Table 2. The Park test statistic values and their p-values

Year	Australia		UK		US	
	Test statistic	p-value	Test statistic	p-value	Test statistic	p-value
1963	0.56	0.59	-0.39	0.71	-0.29	0.78
1964	2.46	0.04 *	1.33	0.23	1.25	0.25
1965	1.61	0.15	-1.38	0.22	0.21	0.84
1966	0.22	0.83	4.42	0.00 *,**	1.39	0.21
1967	-0.67	0.53	2.98	0.02 *	-0.30	0.77
1968	2.25	0.06	1.56	0.17	0.99	0.35
1969	-0.63	0.55	0.70	0.51	1.69	0.13
1970	0.40	0.70	0.21	0.84	0.12	0.90
1971	0.21	0.84	1.31	0.24	3.69	0.01 *,**
1972	-1.21	0.26	1.02	0.35	0.29	1.00
1973	0.76	0.47	0.01	0.99	0.86	0.42
1974	1.40	0.20	2.28	0.06	0.08	0.94
1975	5.46	0.00 *,**	0.23	0.83	0.10	0.93
1976	4.79	0.00 *,**	1.92	0.10	1.73	0.13
1977	1.47	0.18	2.54	0.04 *	0.34	0.74
1978	-0.45	0.66	3.89	0.01 *,**	0.14	0.89
1979	0.52	0.62	0.61	0.56	-0.50	0.63
1980	-1.33	0.22	1.38	0.22	1.10	0.31
1981	0.60	0.57	0.73	0.49	0.90	0.40
1982	1.10	0.31	0.91	0.40	3.98	0.01 *,**
1983	-0.44	0.68	0.93	0.39	0.44	0.67
1984	1.01	0.36	5.93	0.00 *,**	-0.27	0.79
1985	2.44	0.05 *	1.04	0.34	1.96	0.09
1986	-0.79	0.46	1.65	0.15	1.53	0.17
1987	0.73	0.49	1.49	0.19	-0.13	0.90
1988	0.93	0.38	0.71	0.50	-0.16	0.88
1989	0.45	0.67	1.97	0.10	0.19	0.86
1990	-0.29	0.78	1.24	0.26	1.67	0.14
1991	-1.07	0.32	0.53	0.61	-0.02	0.98
1992	0.31	0.77	1.20	0.28	-0.08	0.94
1993	0.50	0.64	1.25	0.26	-0.12	0.90
1994	1.88	0.10	0.33	0.75	-1.09	0.31
1995	0.10	0.92	-0.13	0.90	-1.52	0.17
1996	1.28	0.24	-0.98	0.36	0.42	0.68
Percentage rejection at 5% level		12		15		6
Percentage rejection at 1% level		6		9		6

Notes: The critical value at the 5 percent level are ±2.365 (or ±3.499 at the 1 percent level) for Australia and the US and ±2.447 (or ±3.707 at the 1 percent level) for the UK. * rejection at the 5 percent level and ** rejection at the 1 percent level.

Figure 1. Alternative approach estimates of inflation rates versus their standard errors

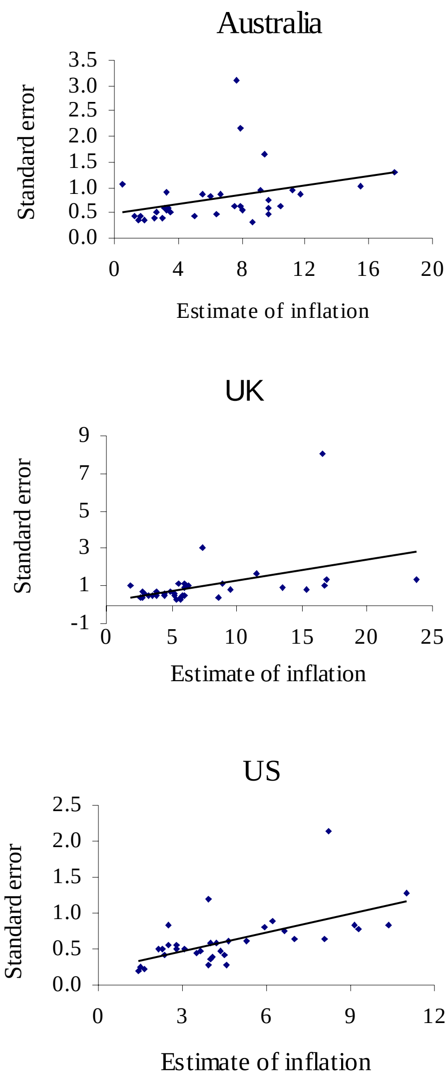
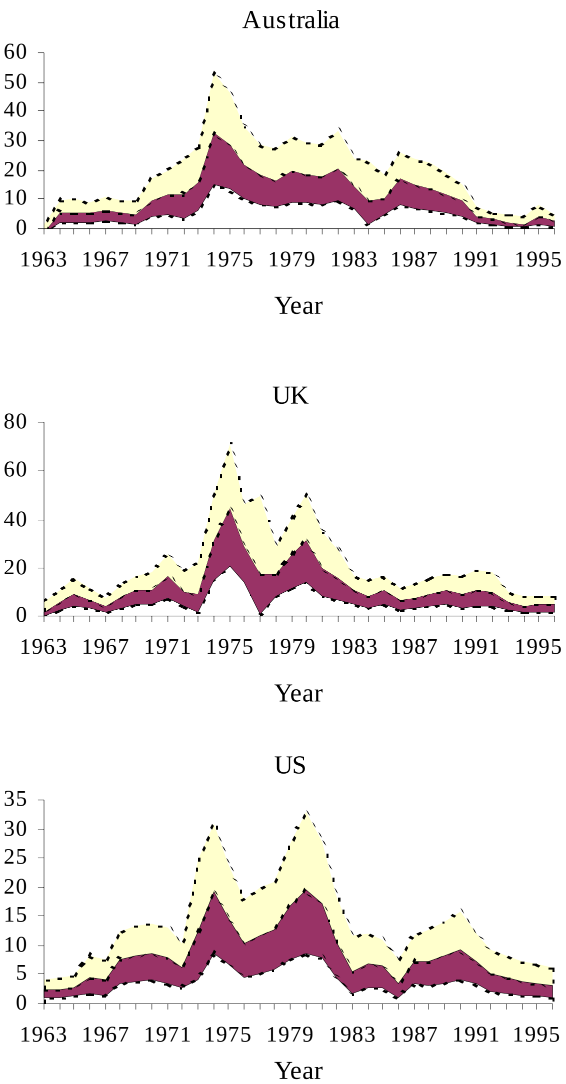


Figure 2. Alternative approach estimates of inflation rates and their 95% confidence bands



V. Concluding comments

In this paper we used a simple regression approach to derive an estimator for the rate of inflation and its standard error. We showed that this estimator is directly related to the Laspeyres and Paasche indices and approximately equal to the Divisia estimator derived by Clements and Izan (1981, 1987). Our estimator and its standard error are simple to estimate and have clear economic interpretation. One of the attractions of our estimator and its standard error is that for computation they require only the base-period budget shares whereas their Divisia counterparts require both the current- and base-period budget shares, which will cause some practical difficulties when evaluating the current rate of inflation. The standard error of the estimator increases with the degree of relative price variability. This agrees with the intuitive notion that when the individual prices move very disproportionately, the overall rate of inflation is less well defined. We also presented an empirical application of the results using expenditure data from three countries, Australia, the UK and the US.

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