Are UK inflation expectations rational?

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Abstract

This paper tests for unbiasedness in inflation expectations drawn from a survey of UK employees by Gallup. It focuses on the econometric difficulties presented by having a small sample, there being overlapping forecast horizons, and by trying to make inference when the data appear to be non-stationary. Applying a method of inference suggested by Inder (1993) the paper concludes that measured expectations systematically overstate inflation. The paper checks the robustness of this result by looking at alternative survey data and by using alternative techniques for modelling the long run.

1 Introduction

The influence of the concept of rational expectations on macroeconomics has been profound. Since the development of the natural rate hypothesis (Phelps (1967), Friedman (1968)), we have known that rational agents will adjust their inflation expectations in line with policy, so frustrating active stabilisation. Barro and Gordon (1983) pushed the state of knowledge out further: they used the nuts and bolts of the (1977) Kydland and Prescott paper to point out that if expectations are rational, there may be an inflationary bias in the economy. In short, exactly how inflation expectations behave is of paramount importance in determining economic outcomes.

Given the predictions of these seminal papers, we could simply turn to our macroeconomic aggregates for output and prices to test whether expectations are 'rational' in some loose sense: in the long run, it seems that inflation has not increased output (see, for example, a survey by Briault (1995)). This econometric fact is consistent with (although not proof of) rational expectations. But a parallel literature has developed which, rather than inferring the behaviour of expectations from macroeconomic outcomes, seeks to test the rational expectations hypothesis (REH) directly on measures of individuals' expectations.

The literature is now large — so much so that it is cumbersome to provide a comprehensive survey of it. But notable examples are Visco (1984), who concluded that the inflation forecasts of professional economists in the Netherlands were not rational; Baghestani (1992), who makes the same observation using US data collected by the University of Michigan Survey Research Center; Batchelor and Dua (1987), who show that the REH fails for survey data in the United Kingdom; and Pacquet (1992), who could not reject the hypothesis that expectations from the Livingston Survey (also in the United States) were rational.(1) Keane

(1) Some papers have tested the REH using *qualitative* data on inflation expectations. For example, Engsted (1991) uses a technique derived from Carlson and Parkin (1975) and Pesaran (1984) to convert forecasts of CBI survey respondents in the United Kingdom categorised as 'up', 'down' or 'stay the same' into quantitative forecasts for output price inflation. He shows that these expectations appear to be unbiased and efficient. There are also examples of REH tests in the finance literature. For a comprehensive survey of tests of rational expectations in survey data in general, see Zimmermann (1997). For example, Ngama (1994) examines whether the forward exchange rates in the Chicago market are

and Runkle (1990) use panel data on the inflation expectations of individual professional forecasters in the United States: they find that they cannot reject rational expectations. Razzak (1997) also cannot reject rational expectations in New Zealand survey data.

In this paper, we use a monthly survey of UK employees' inflation expectations conducted by Gallup, from 1984–1996. This is the longest quantitative series on inflation expectations in the United Kingdom. We explain how the REH is formulated for econometric testing, and discuss the many econometric problems that need to be addressed to obtain reliable results. In particular, we explain that since inflation expectations and inflation appear to be non-stationary, the REH implies that the two series should cointegrate: that there should be no long-run difference between them, and that innovations in one should be matched by innovations in the other. (In other words, the cointegrating vector should have a zero constant and a unit coefficient.) We then apply Inder's (1993) recommendation for modelling the long run: using a single equation to model the long run and the dynamics together as suggested by Bewley (1979), and applying a Phillips-Hansen (1990) correction to the standard errors to conduct inference and so test for rational, unbiased expectations. We show that these employees' expectations are biased forecasts of actual inflation.

We gauge the robustness of our results by applying Inder's technique to inflation expectations data taken from the Barclays Basix survey. This quarterly survey has recorded quantitative inflation expectations of the general public and various specialist groupings since 1986 Q4. These data also suggest that inflation expectations are biased forecasts of actual inflation.

We also test to see how robust our inferences are about the long run to alternative estimators that do not rely on knowledge of the exogeneity/endogeneity nor the integration properties of our data. We conduct cointegration tests suggested by Pesaran *et al* (1996) and estimate the long-run coefficients using a method proposed by Pesaran and Shin (1995). Our estimate of the forecast bias in measured expectations appears to be robust.

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unbiased predictors of the future spot rate. Dutt and Ghosh (1995) apply the REH test to survey data on expectations of future exchange rates.

The focus of our investigation is on the potential for small-sample bias to contaminate our test of the REH. This is important not only because our data sample is ‘small’ in a statistical sense, but because any apparent bias could come about either wholly or partly because agents are simply taking time to learn about a change from a high to a low-inflation regime. We discuss this possibility in the final section of the paper.

2 An econometric test of the REH

For inflation expectations to be rational requires that they are unbiased and efficient predictors of actual inflation. In other words, actual inflation should be equal to expected inflation on average, and equal to expected inflation plus a random forecast error period by period. This forecast error should not correlate with other data in the agents’ information set.

Formally, actual and expected inflation should be related such that:

$$\pi^e_{t|t-n} \equiv \alpha + \beta\pi_t + u_t \qquad \textbf{(1)}$$

where π_t = the inflation rate at time t

$\pi^e_{t|t-n}$ = the expectation of inflation at time t, formed at time t-n

u_t = a white-noise error

and α, β are parameters.

The REH requires on the one hand that expectations are unbiased (ie that α, β = (0,1)), and on the other that they are also efficient: that the u_t’s are not autocorrelated, or correlated with other information that agents have in their information set. In this paper we have focused on the test for unbiasedness.

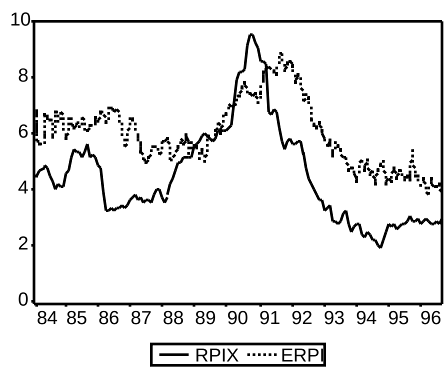
As we shall show, it appears that our expected and actual inflation data are I(1) – difference stationary over our sample period. Engsted (1991) and Paquet (1992) discuss how in this context testing for rational expectations amounts to testing first that actual and expected inflation cointegrate: if expectations are rational, there should be no persistent divergences between actual and expected inflation, ie expectational errors should be stationary. Second, the test for unbiasedness, assuming that the series cointegrate, is that the cointegrating vector has no constant and that this stationary combination involves equal and opposite coefficients on expected and actual inflation.

The remainder of the paper is therefore structured as follows. First, we discuss the data. Second, we test for non-stationarity. Third, we test for cointegration. Fourth, we discuss and then implement the most appropriate method for modelling the long run in this context, testing for unbiasedness in the process.

3 The data

The chart below plots our data for actual and expected inflation.

Chart 1: actual and expected (RPIX) inflation



The expectations are the mean of one year ahead inflation forecasts of respondents to a monthly Gallup survey of some 1,000 employees (drawn from a stratified sample of the population of Great Britain). Respondents are asked to forecast in the following ranges (in percentages): 0, 1−2, 3−4, 5−6, 7−8, 9−10, 11−12, 13−14, 15−20, 20+. Gallup calculate an average by taking the mid-point of each range and weighting by the number of respondents within it. The 20+ range mid-point is assumed to be 24%. There are clearly problems with these data: respondents are not offered the chance to forecast falls in prices; the ranges do not change over the sample period; the assumption that the mean of each range will lie at the mid-point may well not be valid (the within-band distributions could be skewed, for example). But they are problems we cannot do a great deal about, at least not without data on individual agents' expectations period by period, which is not available.

Pesaran (1987) discusses how measurement errors in the inflation expectations data can invalidate inference, but this is in the context of converting qualitative data into quantitative forecasts (in particular, using measures generated by ordinary least-squares regressions). For us, since we have quantitative data, measurement error is a second-order problem, at least relative to Pesaran's work. We can take comfort from two further sources. First, Lee (1994) argues that cointegration-based tests of the REH are still valid, as long as the measurement error is itself stationary; if this is the case, any rejection of the REH because of non-cointegration is indeed due to irrational expectations. Second, Engsted (1991) argues that the test that actual and expected inflation are cointegrated with zero constant term and unit coefficient is effectively a test of the joint hypothesis of unbiasedness and negligible measurement errors. This is because if inflation expectations were actually biased, it is highly unlikely that the measurement error would cause the otherwise not cointegrated variables to cointegrate with exactly the cointegrating vector (1,-1). Of course, if our test of the REH is rejected, one cannot say whether it is because of biased expectations or non-negligible measurement errors, or both.

So we proceed with our measure of inflation expectations, albeit cautiously. But which measure of actual inflation is appropriate? The series we chose was RPIX: that is, the Retail Price Index excluding mortgage interest payments, published by the Office for National Statistics. We could have used an all-inclusive index (the RPI). After all, the question the survey respondents were asked was 'over the next twelve months, what percentage increase in prices do you expect?' where 'prices' is probably best interpreted as 'the average of *all* prices'. In fact, we also perform our unbiasedness tests using RPI data as a check on robustness.

But there are two reasons for excluding mortgage interest payments (MIPs). The first is that that they are related to the monetary policy instrument itself. The errors in forecasting an all-inclusive index will include the errors in accounting for the transitory effect of a change in monetary policy on the mortgage interest rate. We took the view that agents would most probably not include the temporary, perverse effects of future increases in interest rates on the all-items RPI. A second reason is that including MIPs might induce an endogeneity bias into our test for the REH. This is because MIPs themselves embody inflation expectations, in two ways. First, MIPs are related to the nominal interest rates charged by lenders, which are in turn affected by the inflation expectations that influence the equilibrium in the market for housing (and other) finance via the Fisher equation. Second, MIPs are related to house prices. And house prices will include expectations about future real returns from buying a house, plus expectations about future increases in the general price level. So we exclude MIPs on the grounds that if we tested for the REH by regressing an all-items RPI on expected inflation, we would in effect be testing whether current expectations are rational forecasts of future inflation expectations, which is not the hypothesis we want to test.

As we can see from Chart 1, there is some evidence that RPIX and ERPIX (inflation expectations) move together, but it is also apparent that there are notable differences between them. On average, inflation expectations seem to overstate actual inflation.

4 The integration properties of actual and expected inflation

Table A shows the results of augmented Dickey-Fuller (ADF) tests on RPIX and ERPIX. First, note that these tests are carried out both with and without a trend (although our prior is that inflation and inflation expectations are highly unlikely to be trended) and that in either case both RPIX and ERPIX appear to be I(1). Tests were run with varying numbers of lags (from 1-36). In principle, the preferred number of lags is selected by inspection of the Akaike Information Criterion and the results of Ljung-Box and Breusch-Godfrey LM tests for serial correlation in the residuals. In fact these do not make a great deal of difference to our conclusion as to the order of integration of our series.(2)

Table A: ADF tests, 1987 M2-1996 M9

	C					***C,T***			
(Lags)	**ADF(a)**	**Akaike(b)**	**B-G(c)**	**L-B(d)**	**(Lags)**	**ADF**	**Akaike**	**B-G**	**L-B**
(1) RPIX									
(1)	-1.01	-2.34	0.37	0.03	(1)	-1.67	-2.35	0.39	0.01
(6)	-1.75	-2.36	0.69	0.26	(6)	-2.21	-2.36	0.61	0.20
(12)	-1.06	-2.33	0.88	0.84	(12)	-1.54	-2.33	0.80	0.76
(18)	-2.28	-2.34	0.95	0.98	(18)	-2.70	-2.35	0.86	0.98
(24)	-2.04	-2.26	0.97	0.99	(24)	-2.58	-2.27	0.90	0.97
(30)	-2.07	-2.21	0.97	1.00	(30)	-2.51	-2.22	0.90	0.99
(36)	-1.42	-2.14	0.80	1.00	(36)	-1.77	-2.16	0.51	0.99
(2) ERPIX									
(1)	-0.67	-1.99	0.89	0.82	(1)	-1.05	-1.99	0.83	0.69
(6)	-0.77	-1.92	0.72	0.92	(6)	-1.10	-1.92	0.41	0.83
(12)	-1.80	-1.92	0.95	0.98	(12)	-2.15	-1.92	0.85	0.96
(18)	-1.18	-1.84	0.84	0.98	(18)	-1.52	-1.83	0.65	0.95
(24)	-2.11	-1.81	0.88	1.00	(24)	-2.39	-1.81	0.76	1.00
(30)	-0.95	-1.74	0.96	1.00	(30)	-1.17	-1.74	0.81	1.00
(36)	-0.52	-1.66	0.97	1.00	(36)	-0.49	-1.66	0.85	1.00

Notes:

(a) MacKinnon critical values at 5% are -2.8% for models without trend, -3.45 for models without.
(b) Akaike in formation criterion.
(c) Breusch-Godfrey LM test for serial correlation in the ADF residuals; probability reported when the null hypothesis is that all lagged residuals' coefficients are zero (up to 36 lags).
(d) Ljung-Box statistics; probabilities when the null hypothesis is that all autocorrelations (up to 36 lags) are zero.

Table B reports Phillips and Perron (1988) tests, which rather than trying to clean the residuals in the ADF regression of serial correlation by including extra lag terms, involve a non-parametric correction to the t-statistics in a Dickey-Fuller-type regression. These test results — consistent with the ADF tests — also show that RPIX and ERPIX are I(1) over our sample period.

(2) We ignored the Schwartz criterion, since it is better to include too many than too few lags in the ADF and the Schwartz Criterion imposes a larger penalty for the use of extra regressors. Fox (1997) confirms this.

Table B: Phillips-Perron tests: the integration properties of annual inflation,1984 M1–1996 M9

	RPIX		ERPIX	
Model	*c*	*c,t*	*c*	*c,t*
lags				
l=12	-1.39	-1.58	-2.18	-2.52
l=6	-1.67	-1.88	-1.63	-1.97

Critical values

	c	*c,t*
1%	-3.47	-4.02
5%	-2.88	-3.44
10%	-2.58	-3.14

We proceed on the basis that RPIX and ERPIX are I(1) and turn next to testing for cointegration. (Later we will test for cointegration (and estimate the long-run coefficients) using a methodology proposed by Pesaran *et al* (1996) that allows for uncertainty as to whether the variables of interest are I(0) or I(1).)

5 Modelling the long run: testing for cointegration

Residual-based tests for cointegration

We test for cointegration – a necessary but not sufficient condition for the REH to hold – by examining whether the residuals from the estimation of **(1)** are stationary, as in Engle and Granger (1987). The results are shown in Table C; as we shall see, the evidence for cointegration is mixed.

Table C: Residual-based cointegration tests

$\Pi_e^t = \beta_0 + \beta_1 \Pi_t + u_t$

Sample 1984 M1 - 1996 M9;

$\hat{\beta}_0$	3.74 (0.478)		
$\hat{\beta}_1$	0.50 (0.099)		
R^2	0.523		
$\bar{R}^2$	0.520		
$\hat{\sigma}$	0.827		
$\chi^2_{sc}(12)$	115.679*		
$\chi^2_{ff}(1)$	14.797*		
$\chi^2_{n}(2)$	7.017*		
$\chi^2_{h}(1)$	5.208*		
ADF	c	c,t	-
l = 12	-1.76	-1.92	-1.78**
l = 24	-2.64	-2.73	-2.70*
Phillips-Perron			
l = 12	-3.54*	-3.67**	-3.56*
l = 24	-3.97*	-4.07*	-3.98*
l = 4	-3.29**	-3.41*	-3.31*
Hall IV *12 lags*			
Res(-14)	-1.75	-1.84	-1.78
Res(-18)	-0.34	-0.38	-0.38
Res(-24)	-0.40	-0.38	-0.60
24 lags			
Res(-26)	-1.15	-1.19	-1.23

Notes:

1 Newey-West standard errors in brackets, computed with uniform window of size 13, equal to the length of the MA error process.
Rejection level of Ho: *at 5%; **at 10% (ADF and Hall tests use Mackinnon's (1991) critical values; Phillips-Perron test uses Phillips-Ouliaris (1990) critical values)

We report three sets of tests: conventional ADF tests, using Mackinnon's (1991) critical values; Phillips-Perron tests, using Phillips-Ouliaris (1990) critical values; and Hall (instrumental variables) tests, again using Mackinnon values. These approaches are designed to accommodate moving-average processes in the error term, which we expect because of the fact that our forecast horizons overlap, and of which there is some (limited) evidence in the autocorrelellograms (not reported).

The contribution of the Hall (1989) tests are that they are shown to be less prone to size distortions in the presence of (negative) moving-average errors, unlike – as noted by Schwert (1989) – the Phillips and Perron tests. The approach is straightforward. Note that the forecast data are collected in the first week of the month, for the change in the price level between that month and the level in twelve months' time. At the time the forecasts are made, the only published data available are the prices dated two months previous to the forecast. Any unexpected shocks to inflation that are realised between the publication of the most recent price data and the end of the forecast period (14 months) will affect future forecasts – in fact all 13 future forecasts – made before that time. So the forecast error at time *t*+13 will be a function of the forecast error at *t*+12, and so on back to the current forecast at *t*. In other words:

$$\hat{e}_t = \alpha_0 \hat{e}_{t-1} + \alpha_1 \hat{e}_{t-2} + \alpha_2 \hat{e}_{t-3} + \ldots + \alpha_{12} \hat{e}_{t-13} \qquad \textbf{(2)}$$

where $\hat{e}_t$ are the synthetic forecast errors, the residuals from the estimation of **(1)**. So the forecast errors follow an MA(13) process, implying that $\text{cov}(\hat{e}_i, \hat{e}_j) \neq 0$ for some values of *i*,*j*, which makes normal inference on the coefficients in **(1)** problematic. But **(2)** also implies that the residual term u_t in the ADF test for the stationarity of the forecast error (**(3)** below) is also a moving average process. If we have:

$$\Delta\hat{e}_t = \beta\hat{e}_{t-1} + u_t \qquad \textbf{(3)}$$

then rearranging **(3)** for u_t and substituting in **(2)**, we have that:

$$u_t = \alpha_0 \hat{e}_{t-1} + (\alpha_1 - \alpha_0[1-\beta])\hat{e}_{t-2} + (\alpha_2 - \alpha_1[1-\beta])\hat{e}_{t-3} + \ldots + \alpha_{12}[1-\beta]\hat{e}_{t-14} \qquad \textbf{(4)}$$

which is also an MA(13) process.

Following Hall's advice, we can estimate **(3)** by finding an instrument for $\hat{e}_{t-1}$ that is not correlated with u_t. We do this by estimating an auxiliary regression for $\hat{e}_{t-1}$ using lags $\hat{e}_{t-14} \ldots \hat{e}_{t-(14+n)}$ as regressors, and then estimating **(3)** by replacing $\hat{e}_{t-1}$ with $\hat{e}^{IV}{}_{t-1}$.

Comparing the various tests for stationarity in the residuals of the long run, it seems that the evidence for cointegration is mixed. The Phillips-Perron tests seem to suggest that the synthetic residuals from **(1)** are stationary, implying cointegration; the ADF tests without constant or trend also support cointegration, but those with a constant and/or a trend, and the Hall tests, all reject cointegration. We know from the autocorrelellograms that there was only a weak trace of the moving average process in the forecast error, so we are cautious in interpreting the Hall tests: if in fact we have instrumented for $\hat{e}_{t-1}$ unnecessarily, then the instrument should show a higher variance than $\hat{e}_{t-1}$ itself, implying that we are more likely to conclude that (perhaps falsely) $\hat{\beta}$ is insignificantly different from zero, since the standard error of $\hat{\beta}$ will be larger, and so find that the residuals are non-stationary.

Since our tests are inconclusive, and the economic content of the paper is heavily dependent on the results of the empirics, we explore other avenues for cointegration testing.

Kremers et al *ECM test for cointegration*

Kremers, Ericcson and Dolado (1992) have argued that residual-based cointegration tests are less powerful than tests based on testing the significance of the ECM term in the dynamic model. The reason is that the ADF test on the synthetic residuals imposes a common factor restriction that the long run and short-run elasticities are equal.[3] In general this will not hold. And the test is less powerful, because it throws away potentially useful information in the short-run dynamics. We fitted the following dynamic model for expected inflation ERPIX using a conventional 'general to specific' approach:

$$\Delta\Pi_t^e = \underset{(0.02)}{-0.16}\left(\Pi_t^e - 0.50\,\Pi_t - 3.74\right)_{t-1} \underset{(0.09)}{-0.27}\,\Delta\Pi_{t-1}^e \underset{(0.09)}{-0.24}\,\Delta\Pi_{t-4} \qquad \textbf{(5)}$$

(R^2 = 0.22; R(bar)^2 = 0.21; S.e.e. = 0.35; $\chi^2_{sc}(12)$ = 11.05; $\chi^2_{ff}(1)$ = 0.41 $\chi^2_n(2)$ = 1.79; $\chi^2_h(1)$ = 1.29)

This model appears to pass the usual diagnostic tests. Recursive least squares estimation suggests that there are no serious structural breaks over the sample period:

The t-statistic on the ECM term in **(5)** is equal to 7.5.[4] Making inference on this statistic is not straightforward. Kremers *et al* show that the distribution of this t-statistic lies somewhere between a standard normal and a Dickey-Fuller distribution. If we use Dickey-Fuller values (loading the dice a little against finding cointegration), it is clear that a t-statistic of 7.50 will reject the null of no cointegration at conventional significance levels. (This test requires that inflation be *strongly exogenous* with respect to expected inflation, this turns out to be the case, as we report later in the paper.)

We also fitted a dynamic model with an unrestricted ECM term as follows:

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(3) This is clear when we note that the Dickey-Fuller test on the synthetic residuals (forecast errors) $\hat{e}_t$, which takes the form: $\Delta\hat{e}_t = \gamma\hat{e}_{t-1}$ can be re-written, by substituting in the long run, as follows: $\Delta(\Pi_t^e - \hat{\alpha} - \hat{\beta}\Pi_t) = \gamma(\Pi_{t-1}^e - \hat{\alpha} - \hat{\beta}\Pi_{t-1})$.

(4) We calculated this using Newey-West (1987) standard errors, calculated with a uniform window of size 13 (as recommended by Pesaran (1987)).

$$\Delta\Pi_t^e = \alpha + \beta\Delta\Pi_t + \gamma\Pi_{t-1}^e + \delta\Pi_{t-1} + \varphi(L)\Delta\Pi_t^e + \phi(L)\Delta\Pi_t \qquad \mathbf{(6)}$$

And obtained **(7)**:

$$\Delta\Pi_t^e = \underset{(0.09)}{0.40} - \underset{(0.02)}{0.16}\,\Pi_{t-1}^e + \underset{(0.01)}{0.12}\,\Pi_{t-1} - \underset{(0.09)}{0.29}\,\Delta\Pi_{t-1}^e - \underset{(0.08)}{0.28}\,\Delta\Pi_{t-4} \qquad \mathbf{(7)}$$

($R^2 = 0.25$; R(bar)$^2 = 0.23$; S.e.e. = 0.35; $\chi^2_{sc}(12) = 9.82$ $\chi^2_{ff}(1) = 0.72$
$\chi^2_n(2) = 1.93$; $\chi^2_h(1) = 0.29$)

Our test for cointegration is now a test of the hypothesis that γ is non-zero. In fact the t-ratio on this parameter is 8.35, which suggests that the actual and expected inflation series are cointegrated. Once again, recursive estimates showed no signs of any serious structural break.

Exogeneity

Before we go on to find ways of conducting inference on the parameters in the long run, we check that the exogeneity assumptions needed for the Kremers *et al* tests to be valid do actually hold. In particular, we need to establish that actual inflation is *strongly exogenous*.(5)

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(5) The literature on testing for rational expectations seems to be divided on whether to assume that actual or expected inflation is exogenous. Under the null of REH, it should not make any difference: (1) is an identity and the error term u_t will be uncorrelated with the regressand and the regressor whichever way round the equation is written. We could simply assume that inflation expectations are weakly exogenous by using the temporal ordering of the data: we know that inflation expectations are formed before inflation is realised. However, if REH fails, inflation expectations will not incorporate all the information necessary to model (predict) actual inflation, and a regression with actual inflation on the left-hand side will be misspecified: the equation will be missing some other variable, correlated with actual inflation. This would affect our estimates of the bias in expectations formation. We decided to let the data decide.

We tested for exogeneity following the recommendations of Charemza and Deadman (1992), in the following way.

First, we tested for *weak* exogeneity by testing whether the ECM term in equation **(5)** is insignificant in a dynamic model for *actual* inflation (again searching for the particular dynamic structure using a general to specific methodology). This gave the following equation:

$$\begin{aligned}\Delta\Pi_t = &\underset{(0.03)}{-0.05}\left(\Pi_t^e - 0.50\Pi_t - 3.74\right) + \underset{(0.03)}{0.21}\,\Delta\Pi_{t-1} \\ &- \underset{(0.08)}{0.31}\,\Delta\Pi_{t-12}\end{aligned} \qquad \textbf{(8)}$$

(R^2 = 0.1668; $R(bar)^2$ = 0.1547; S.e.e. = 0.2857; $\chi^2_{sc}(12)$ = 16.7111; $\chi^2_{ff}(1)$ = 13.6132*: $\chi^2_n(2)$ = 644.1947*; $\chi^2_h(1)$ = 7.8141*)

Recursive least squares estimates show that there is no significant structural break over the sample period. The t-statistic on the ECM term is 1.99 (calculated, as before, using Newey-West standard errors with uniform window of 13 months) and is therefore (at conventional levels) insignificant. We re-ran this regression using the unrestricted long run embedded in **(7)** and **(8)**, and obtained the following:

$$\begin{aligned}\Delta\Pi_t = &-\underset{(0.04)}{0.02}\ \underset{(0.04)}{-0.07}\left(\Pi_t^e - 0.75\Pi_t - 2.48\right) \\ &+ \underset{(0.05)}{0.35}\,\Delta\Pi_{t-1} - \underset{(0.07)}{0.17}\,\Delta\Pi_{t-7} - \underset{(0.09)}{0.42}\,\Delta\Pi_{t-12} \\ &- \underset{(0.05)}{0.15}\,\Delta\Pi^e_{t-12}\end{aligned} \qquad \textbf{(9)}$$

(R^2 = 0.40; $R(bar)^2$ = 0.37; S.e.e. = 0.33; $\chi^2_{sc}(12)$ = 22.9; $\chi^2_{ff}(1)$ = 8.2: $\chi^2_n(2)$ = 18.8; $\chi^2_h(1)$ = 3.00)

The t-statistic on the ECM term in (**9**) is 1.68, which is insignificant at the 5% level, although significant at 10%. This is supportive, although not conclusive, evidence of the weak exogeneity of the long-run component of actual inflation.

We then, again following Charemza and Deadman (1992), turned to test for the weak exogeneity of the short-run component of actual inflation. We took the residuals from **(9)**, - with the ECM term omitted, and ran **(7)** with the residuals included as an additional variable. These residuals had no additional explanatory power in (**7**), which implies that the short-run movements in actual inflation are weakly exogenous.

For actual inflation to be *strongly exogenous* requires *weak* exogeneity, plus the requirement that (in our case) actual inflation is not Granger-caused by expected inflation. We test for this by estimating a vector error-correction in differences of actual and expected inflation, and testing to see whether the error-correction term can be excluded from the equation for differences in actual inflation.[6] Beginning with a model with twelve lags of the difference terms, we find that the t-statistic on the error-correction term in the equation for actual inflation is 1.50, while the t-statistic for the error-correction term in the equation for expected inflation is 3.51; this is evidence supporting our assumption that while inflation Granger-causes expected inflation, expected inflation does not Granger-cause inflation, which we need in order for our inference in the Kremers *et al* tests for cointegration to be valid.[7] To satisfy readers that our inferences about cointegration are not sensitive to these causality tests, however, we shall present results from a cointegration test suggested by Pesaran *et al* (1996) that are not reliant on knowledge of the endogeneity/exogeneity properties of the data.

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(6) Note that Toda and Phillips (1993) suggest that non-causality should be tested for using VECMs.

(7) We should note that the test results are a little sensitive to the lag-length included in the VECM; the table below summarises.

	t-statistics on ECM in equation for:	
Number of lags	Actual inflation	Expected inflation
6	-2.30	4.21
9	-1.15	4.34
12	1.50	3.51
24	-0.07	1.89

We now move on to test for unbiasedness, proceeding on the basis that the two series are indeed cointegrated, retaining our focus on the small-sample properties of the estimator, and its robustness to the presence of moving-average errors in the cointegrating relationship.

6 Testing for unbiasedness -- making inference on the coefficients in our long run

(i) Phillips-Hansen (1990): correcting for endogeneity and moving averages

Since the variables estimated in **(1)** are difference-stationary − I(1) − we cannot conduct formal inference on the coefficients in the long run estimated in Table C, which showed that the constant bias in expectations was 3.7 percentage points, and the coefficient on inflation was a long way from unity, at around 0.5. The literature has thrown up two solutions to this problem: either to estimate **(1)** in a dynamic form, so that all terms are I(0), or to devise some correction to the variance-covariance matrix in the VAR to correct for violations of Gaussian assumptions about the error term. The methodology we use confronts two further problems: first, the difficulty caused by the fact that there is (or at least we have a strong prior that there is) a moving average process embedded in the forecast error, which breaches the usual Gaussian assumptions. Second, that we have a small sample and that our coefficient estimates may be biased because of this.

One approach to inference is to use the VAR methodology proposed by Johansen (1988). But this requires Gaussian error terms in each equation of (in our case) the two-equation system, and further that the errors from each equation are uncorrelated with each other. In an unrestricted VAR, it is unlikely that the errors in the underlying VAR will not be correlated if there is a moving-average error in the cointegrating relationship.[8]

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(8) We are grateful to our anonymous referee for pointing out that we could estimate a restricted underlying VAR, excluding lag lengths less than 2, and ensure that the moving-average error in the cointegrating relation does not cause cross-equation error correlation. But we proceed with the Phillips-Hansen approach anyway, since (i) restricting the underlying VAR in this way is not attractive, and (ii) as Moore and Copeland (1995) put it ‘there is no obvious advantage to be gained... from the fact that the Johansen approach can accommodate more than one cointegrating vector in the system...’

For this reason, the methodology proposed by Phillips and Hansen (1990) has become popular in the single-equation literature. For example, Moore and Copeland (1995) use this procedure to test for efficiency in the foreign exchange markets; Ngama (1994) applies Phillips-Hansen to test whether forward exchange rates are unbiased predictors of the spot rate. The Phillips-Hansen procedure corrects for the bias due to the autocorrelation in the residuals of the long-run regression in a manner akin to the Phillips-Perron (1988) methodology for testing for unit roots in a univariate context. It also corrects for any possible endogeneity of the RHS variables with respect to the LHS, by estimating the long-run covariance between the regressor and the regressand.

Phillips-Hansen estimation yields the following long-run equation for inflation expectations (ERPIX).(9)

$$\Pi_t^e = \underset{(0.37)}{3.36} + \underset{(0.08)}{0.58}\,\Pi_t \qquad \textbf{(10)}$$

($R^2 = 0.5104$)

The joint restriction of a unit coefficient and zero constant is easily rejected: the associated Wald statistic $\chi_W^2(2) = 155.306$. We can also easily reject these restrictions if we impose them separately. For the test of a constant coefficient, we find that $\chi_W^2(2) = 83.425$; for the unit coefficient restriction, $\chi_W^2(2) = 30.0419$. This amounts to a conclusive rejection of unbiasedness of expectations and hence of the REH.

(ii) Small-sample bias: Inder's (1993) modification of Phillips-Hansen

It is has been known for some time that static estimates of the long run of the sort proposed by Engle and Granger (1987) are subject to small-sample bias: this point was first made by Banerjee *et al* (1986) using a

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(page 132) since we know that with two I(1) variables, there should be a maximum of one cointegrating vector.

(9) Assuming that at least one regressor is I(1) with drift, and using a uniform window of 13 to calculate the adjusted standard errors, as recommended in Pesaran (1987).

Monte Carlo study. The intuition is that in small samples, small enough that the system may not have had time to equilibrate after a shock, such shocks may pollute estimates of the long run. And in our case, we might well expect the bias on the coefficient on inflation to be downwards and the constant term to be biased upwards. This would be the case if inflation expectations, although unbiased in the long run, adjust sluggishly in the short run.[10] Suppose that inflation expectations evolve according to the following process:

$$\Pi_t^e = b_0\Pi_t + b_1\Pi_{t-1} + \ldots + b_n\Pi_{t-n}$$

where $\sum_{i=1}^{n} b_i = 1$ and $b_i > 0, \quad i$. Suppose further that actual inflation is on average equal to the inflation target:

$$\Pi_t = \Pi_t^T + \varepsilon_t$$

where Π_t^T is the inflation target in period *t* and ε_t is white noise. Inflation expectations are unbiased in the long run — on average expected inflation is equal to the target:

$$\Pi_{lr}^e = b_0\Pi_{lr} + b_1\Pi_{lr} + \ldots + b_n\Pi_{lr} = \left(b_0 + b_1 + \ldots + b_n\right)\Pi_{lr} = \Pi_{lr}$$

But if the sample period used in estimating the levels regression is not long enough, the coefficient on the inflation term in equation (1), β may only pick up a combination of the b_i s, which sum to less than unity. And then the estimated constant term, $\alpha = \bar{\Pi}^e - \beta\bar{\Pi}$, although zero in the long run (when $\beta = 1$), might turn out to be positive if $\beta < 1$. This bias might be a particular problem towards the end of the sample

(10) See Caballero (1994) who argues that sluggishness of the capital stock imparts a similar downwards bias to estimates of the long run response of capital to changes in the cost of capital.

period, when an inflation target was operating in the United Kingdom. Individuals may be taking time to adjust their expectations following the introduction of the inflation target that might be expected to reduce the inflation rate in the long run.

One solution is to model the static and dynamic terms simultaneously, along the lines of Bewley (1979). Pacquet (1992) uses cointegration to test for REH on Livingston survey (US) data, and shows that including dynamic terms in the static regression can affect the estimates of the long-run parameters considerably. The Kremers *et al* test that we performed in **(7)** also models the long and the short run together. From our estimate of **(7)**, we can recover a long-run expression for inflation expectations that looks like this:

$$\Pi_t^e = 2.48 + 0.75\,\Pi_t \qquad \textbf{(11)}$$

Note that the coefficient on inflation, at almost three quarters, is quite different from that in Table C (about one half), and that the constant terms differ, too: the bias is 3.7% in the Engle and Granger equation, and 2.5% in the Kremers *et al* test. This is closer to the values demanded by the REH. The presence of a downward bias on the coefficient on inflation and an upward bias on the constant term is consistent with inflation expectations adjusting sluggishly in the short run.

Although modelling the dynamics and the long run jointly accounts for small-sample bias, it does not enable us to test hypotheses about the individual parameters in the long run (because the individual components of the long-run solution are all I(1) variables).

Inder (1993) proposed a way round this, which we apply to our data. It is basically an amalgam of Phillips-Hansen (1990) and Bewley (1979). Inder (1993) describes the methodology formally, but an intuitive explanation is as follows. First, we estimate **(12)** below using the IV procedure suggested by Bewley.

$$\Pi_t^e = \alpha + \beta\Pi_t + \lambda_1(L)\Delta\Pi_t^e + \lambda_2(L)\Delta\Pi_t \qquad \textbf{(12)}$$

We then define:

$$\overset{*}{\Pi}{}_t^e = \Pi_t^e - (\lambda_1'(L)\Delta\Pi_t^e + \lambda_2'(L)\Delta\Pi_t) \qquad \textbf{(13)}$$

where denotes estimated values. In other words, $\overset{*}{\Pi}{}_t^e$ is expected inflation minus the estimated dynamics. We then estimate **(14)** using Phillips-Hansen corrections to make inference on the vector ρ'.

$$\overset{*}{\Pi}{}_t^e = \rho_1 + \rho_2 \Pi_t \qquad \textbf{(14)}$$

In other words, we first estimate the long and the short run together, to minimise problems arising from small-sample bias; we then take the actual levels series, subtract the estimated dynamics, and use this as the dependent variable to regress on actual inflation (in a levels equation like **(14)**). Finally, we make inference on the new modified long run using Phillips-Hansen corrections to the variance-covariance matrix. Inder (1993) conducts a Monte-Carlo study to show the benefits of increased precision of estimates when using this two-stage process.

The estimated long run from **(14)** in our case looks like this:

$$\overset{*}{\Pi}{}_t^e = \underset{(.22)}{2.64} + \underset{(.05)}{0.70}\,\Pi_t \qquad \textbf{(15)}$$

This is reassuringly close to the long run implied by the unrestricted ECM estimated in the KED test for cointegration. When we test for unbiasedness, we find that the data reject the hypotheses of a zero constant, a unit coefficient on actual inflation, and the two imposed jointly:

$$\chi_1^2(\rho_0 = 0) = 138.2$$
$$\chi_1^2(\rho_1 = 1) = 41.6$$
$$\chi_1^2(\rho_0 = 0, \rho_1 = 1) = 263.8$$

We therefore conclude that while there is evidence that the inflation expectations of employees surveyed by Gallup cointegrate with actual inflation, these expectations are not unbiased.

7 Some tests for robustness

In this section, we examine how robust our conclusions about the rationality of inflation expectations are, in particular to (a) different measures of observed inflation; (b) different measures of inflation expectations; and (c) uncertainty about the integration properties of observed and expected inflation.

(a) Robustness to alternative measures of actual inflation

In Section 3 we discuss why we have opted to use RPIX as our measure of inflation. Recall that we put forward two arguments for not using the all-items RPI. The first was that mortgage interest payments (MIPs), since they are correlated with the nominal interest rate, would contaminate inflation forecast errors with errors associated with forecasting changes in monetary policy. The second argument was that MIPs are correlated with house prices, which, since they are asset prices, would contain information about not only expected relative returns to houses, but also expected inflation. For readers who are not convinced by these arguments, we look to see how robust our results are to using the all-items RPI as our measure of observed inflation. When we performed Inder's procedure, we found that the estimated long run in this case is:

$$\Pi_t^{*e} = \underset{(.16)}{355} + \underset{(.03)}{052}\,\Pi_t \qquad \textbf{(16)}$$

Comparing this with equation (15), it appears that the deviation from the REH is even more pronounced if we use RPI as our measure of inflation. Formally, the restrictions of a zero constant, a unit coefficient on actual inflation, and the two imposed jointly are easily rejected by the data.

(b) Robustness to alternative measures of inflation expectations

The conclusion that inflation expectations are biased may be specific to the Gallup data used in this paper. To get a handle on robustness, we apply Inder's procedure to expectations series taken from the Barclays Basix survey. This is a *quarterly* survey going back to 1986 Q4. Specifically, it surveys one year ahead and two year ahead inflation expectations of the general public (GP), academic economists (AE), business economists (BE), investment analysts (IA), finance directors (FD) and trade unions (TU). The sample sizes vary greatly between groupings: the 1997 Q2 survey asked 1,891 members of the general public, 36 academic economists, 72 business economists, 14 investment analysts, 50 finance directors and 15 trade unions.

Charts 2-7 plot one year ahead inflation expectations against RPIX inflation outturns for the different groups.

Chart 2

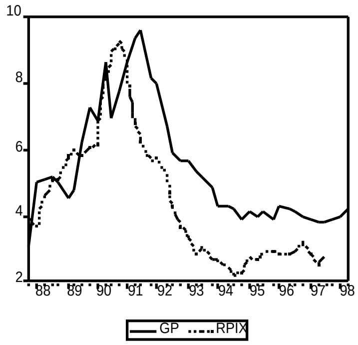


Chart 3

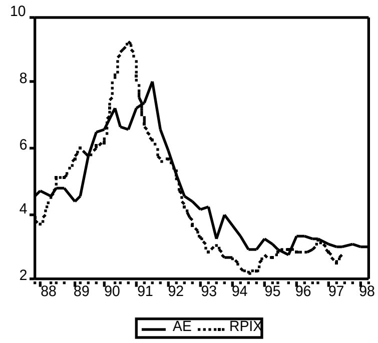


Chart 4

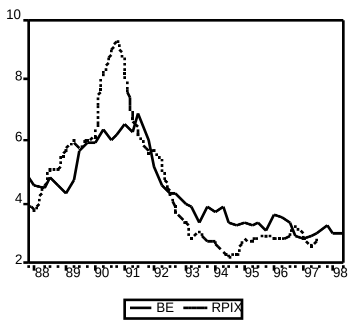


Chart 5

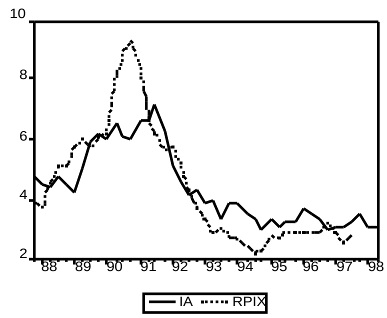


Chart 6

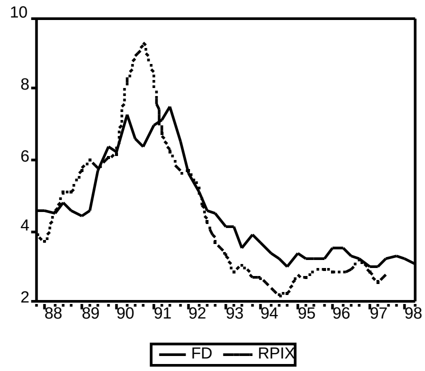
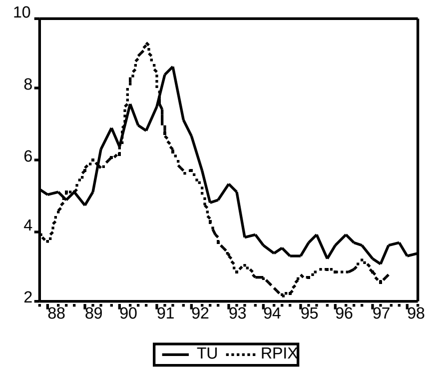


Chart 7



Stationarity tests suggest that all the series are non-stationary over the sample period. In fact, the small number of observations (reflecting the fact that the sample period is shorter than in the Gallup survey and the frequency is lower) means that it is not possible to reject even the hypothesis that actual and expected inflation rates are I(3)! Following the stationarity tests using the Gallup survey, we proceed on the assumption that the series are all I(1), and further assume that the two series are cointegrated.

We conducted Inder's procedure using the Basix data, which gave the following long-run relationships:

$$GP = \underset{(0.14)}{1.57} + \underset{(0.03)}{0.90}\,RPIX \qquad \textbf{(17)}$$

$$AE = \underset{(0.09)}{0.98} + \underset{(0.018)}{0.80}\,RPIX \qquad \textbf{(18)}$$

$$BE = \underset{(0.18)}{1.53} + \underset{(0.038)}{0.64}\,RPIX \qquad \textbf{(19)}$$

$$FD = \underset{(0.09)}{1.29} + \underset{(0.018)}{0.74}\,RPIX \qquad \textbf{(20)}$$

$$IA = \underset{(0.12)}{1.60} + \underset{(0.02)}{0.64}\,RPIX \qquad \textbf{(21)}$$

$$TU = \underset{(0.08)}{1.34} + \underset{(0.02)}{0.82}\,RPIX \qquad \textbf{(22)}$$

In all cases the data reject the restrictions that the constant term equals zero, the RPIX coefficient is unity and that both hold together. This supports the conclusion that inflation expectations are biased estimators of actual inflation. Interestingly, the constant terms are in this case closer to zero and the coefficients are closer to unity. So the deviation from the REH, at least on these criteria, appears to be less marked in the Basix survey than in the Gallup survey. The reported expectations in the survey are the arithmetic mean of individual survey responses.

(c) Robustness to uncertainty about the integration properties and exogeneity/endogeneity of observed and expected inflation

Our cointegration tests and our hypothesis tests on the long-run coefficients are all contingent on our assumptions about the integration properties of actual and expected inflation, namely, that they are both I(1). Our result that the series were I(1) − documented in Tables A and B − were not absolutely conclusive, and it is possible that they are contaminated by small sample biases, or structural changes in the data-generating processes (for example, monetary regime changes). Our cointegration tests also required that actual inflation was strongly exogenous with respect to expected inflation. While we established *weak* exogeneity without difficulty, our results regarding whether actual inflation was not Granger-caused by expected inflation were more equivocal.

Here we test for the existence of a long-run relation between actual and expected inflation using the methodology proposed by Pesaran *et al* (1996), and then estimate the coefficients in the long-run relationship using the ARDL (Autoregressive Distributed Lag) procedure described in Pesaran and Shin (1995); these methods allow us to make inferences that are robust to whether our data are I(1) or I(0). Moreover, they do not rely on assumptions about the exogeneity or otherwise of actual or expected inflation.[11]

(11) The anonymous referee pointed out that the ARDL estimates have better small-sample properties than the Phillips-Hansen estimator; we do not know of any evidence as to whether the ARDL estimates are more robust in small samples than Inder's modification of Phillips-Hansen.

We describe these procedures only briefly here: a more rigorous account can be found in the source papers. Following Pesaran *et al* (1996), we estimate an unrestricted dynamic equation for expected inflation (using the Gallup measure of inflation expectations, and RPIX to measure inflation) as follows:

$$\Delta\Pi_t^e = \alpha + \beta\Pi_{t-1}^e + \gamma\Pi_{t-1} + \delta(L)\Delta\Pi_t^e + \phi(L)\Delta\Pi_t \qquad \textbf{(23)}$$

We test for cointegration by inspecting the F-statistic from a hypothesis test that the levels coefficients are zero, varying the number of dynamic lags in **(23)** between 3-12, since it is known that inferences can be sensitive to the number of lags included. These F-statistics are reported in Table D below. We compare them with two critical values (at the 5% level of significance), one appropriate under the assumption that both variables are I(0), the other under the assumption that both are I(1).

Table D: Cointegration tests à la Pesaran *et al* (1996)

No. of lags in dynamic model	F-stat
3	8.47
6	8.91
9	9.29
12	6.62

(Critical values at 5% level of significance: 4.93 if variables are I(0); 5.76 if variables are I(1). See Pesaran *et al* (1996) for more details.)

We reject the hypothesis that the levels terms are insignificant with reasonable confidence, and can therefore conclude that actual and expected inflation are cointegrated.

We go on to estimate the long-run coefficients using the ARDL procedure suggested by Pesaran and Shin (1995). Table E below sets out the long-run coefficient estimates that result from estimating the ARDL with the number of lags determined by either the Akaike or the Schwartz/Bayesian Criterion. In the right-hand panels, we report the results of hypothesis tests on the coefficients.

Table E: ARDL estimates of the long run, à la Pesaran and Shin (1995)
(up to 8 lags chosen in ARDL to limit the number of ARDL models estimated)

$$(\Pi_t^e = \alpha + \beta \Pi_t)$$

Criterion used to choose lag in ARDL	Akaike		Schwartz/Bayes	
	α	β	α	β
Long-run estimates:	2.54	0.74	2.36	0.77
Asymptotic st. errors:	0.48	0.10	0.89	0.18
Hypothesis tests:	$\chi^2_{1,2}$	p-value	$\chi^2_{1,2}$	p-value
(1) $\alpha = 0$	27.88	0.00	6.97	0.01
(2) $\beta = 1$	7.10	0.01	1.60	0.21
(3) $\alpha = 0, \beta = 1$	82.54	0.00	23.30	0.00

These results are broadly similar to those we got from using Inder's modified Phillips-Hansen estimates. Once again we find that the average bias in inflation expectations is about 2.5 percentage points, and that the responsiveness of expectations to a unit change in inflation is about three quarters. The data appear to reject unbiasedness; the exception is that if we chose the optimal lag length in the ARDL model using the Schwartz/Bayes Criterion, we cannot reject the hypothesis that inflation expectations move one-for-one with actual inflation. We should not read too much into this: the Schwartz/Bayes criterion chooses shorter lag lengths in the ARDL, and automatically generates larger standard errors around the coefficient estimates and correspondingly less precise inference.

8 Conclusions and interpretation

Summary of econometric results

Our econometric evidence casts doubt on the notion that inflation expectations in the United Kingdom are 'rational'. We first tested for cointegration, since both actual and expected inflation were I(1) over our sample period. Residual-based tests for cointegration and a Kremers *et al* (1992) test suggested that the series did indeed cointegrate, although the test results were not unambiguous. Second, assuming cointegration, we moved on to make inference on the parameters in the long run. We applied the Phillips-Hansen (1990) correction to the variance-covariances, because of the violations of Gaussian assumptions about the error term caused by the overlapping observations, and concluded that expectations were indeed biased. A modification of this technique, suggested by Inder (1993), and applied to a long run modelled jointly with the short-run dynamics, also suggested that unbiasedness did not hold. **Broadly, it seems that expected inflation overstates actual inflation by about 2.5 percentage points, and that a 1 percentage point increase in actual inflation causes expected inflation to increase by about 0.75 percentage points.**

These coefficients differ markedly from those obtained from our Engle-Granger estimate, where the bias in expectations was about 3.7 percentage points and the coefficient on actual inflation was about 0.5. Our analysis confirm that these estimates are subject to small-sample bias. We showed that a bias of this type is consistent with the idea that individuals' inflation expectations - although unbiased in the long run — adjust sluggishly in the short run. This might be a particular problem towards the end of the sample period: individuals may be taking time to adjust their expectations following the introduction of the inflation target, that might be expected to reduce the inflation rate in the long run. But our econometric analysis suggests that this does not explain all the bias in inflation expectations.

We also showed that our estimates of the apparent bias in expectations were robust to using alternative measures of expectations, using RPI instead of RPIX as our measure of inflation, and to using alternative methodologies for testing for the existence of a long-run relationship and estimating long-run coefficients.

Intepretation

Taken at face value, the results suggest that respondents are making systematic errors in forecasting inflation: they appear to be over-predicting. But we need to recall that there are other ways of interpreting our results.

First, the results are consistent with there being either systematic under-estimation of inflation (which we deem unlikely: see Cunningham (1996), for example); or systematic over-estimation of inflation expectations. The robustness of our results to changes in the inflation and expectations measures used is a partial, but not complete, defence.

Second, it is also possible that the apparent positive bias in inflation expectations is a small-sample phenomenon, that agents are simply taking time to learn about a change in regime over the sample period from a high inflation regime to a low inflation regime. This remains a possibility, despite our attempts to allow for small-sample bias in the econometric analysis. Nevertheless, we make a further observation. The outturns for inflation over our sample period do not support the idea that we are measuring expectations over a period of uniform disinflation. Inflation started out in 1984 at 5.1% and fell to 2.4% in 1986; it rose thereafter to a peak of about 10.9% and had fallen to 2.1% by mid 1996.

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