Inflation Targeting and Inflation Persistence,

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Abstract

This paper argues that the adoption of an inflation target reduces the persistence of inflation. We develop the theoretical literature on inflation persistence by introducing a Taylor rule for monetary policy into a model of persistence and showing that inflation targets reduce inflation persistence. We investigate changes in the time series properties of inflation in seven countries that introduced inflation targets in the late 1980s or early 1990s. We find that the persistence of inflation is greatly reduced or eliminated following the introduction of inflation targets.

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1. INTRODUCTION

This paper argues that the persistence of inflation is lower when there is an inflation target. This implies that inflation is more responsive to monetary policy when inflation is the main focus of policy.

The idea that inflation persistence may depend on macroeconomic institutions or policy regimes, of which inflation targets are a recent example, is well established in the literature. Alogoskoufis and Smith (1991) and Alogoskoufis (1992) argue that inflation is less persistent with fixed exchange rates. Other authors, for example Siklos (1999) and Burdekin and Siklos (1999), argue that other factors, such as wars, supply shocks or Central Bank reforms, also affect persistence.

We extend the theoretical literature by introducing a Taylor-rule representation of monetary policy (Taylor, 1993) into an otherwise standard model of inflation persistence, similar to Taylor (1979), Alogoskoufis and Smith (1991), Agenor and Taylor (1992) and Alogoskoufis (1992). We show that inflation persistence is affected by the parameters of the monetary policy rule. An increased weight on the price-level target in the monetary policy rule reduces persistence. As a result, persistence is lower when there is an inflation target.

We test our model by investigating changes in the time series properties of inflation persistence in seven countries that adopted inflation in the late 1980s or early 1990s (Australia, Canada, Finland, New Zealand, Spain, Sweden and the UK). We find that the persistence of inflation is greatly reduced or even eliminated following the introduction of inflation targets. Using annual data we cannot reject the hypothesis that inflation persistence has been eliminated in any country. Using quarterly data, we can only reject this hypothesis for the UK.

The paper is structured as follows. Section 2) contains our theoretical model. Section 3) contains our empirical results. Section 4) summarises and concludes.

2. THE MODEL

In this section we present our model. We begin by considering aggregate demand. We then develop the supply side of the model before finally analysing inflation persistence.

2a) Aggregate Demand

We assume that aggregate demand is given by

$$(1) \qquad y_t = \bar{y} - g\,(i - E_{t-1}p_{t+1} + E_{t-1}p_t) + v_t$$

where y is the natural logarithm of output, $\bar{y}$ is an exogenous component of demand, i is the nominal interest rate, p is the natural logarithm of the price level, $E_{t-1}p_{t+1}$ is the expected price level in period (t+1) using information available at time t-1, v_t is a white noise demand shock and t indexes time. We assume that monetary policy is conducted through an interest rate Taylor-type rule,

$$(2) \qquad i_t = i^* + f\,(p_t - p^*) + y\,(y_t - y^*)$$

where i* is a constant, *p** is the log of the policymaker's target for the price level and *y** is the log of the policymakers target level of output. The policy parameters f and y describe the responsiveness of nominal interest rates to deviations of inflation and output from their respective targets. Our model extends the literature by introducing a familiar Taylor rule description of monetary policy (Taylor, 1993). In the existing literature, Taylor (1979) assumes aggregate demand depends on the real money supply and that the nominal money supply is proportional to the price level. Alogoskoufis and Smith (1991), Agenor and Taylor (1992) and Alogoskoufis (1992) have an aggregate demand relationship similar to (1), but model interest rates using a money demand equation and again assume the nominal money supply is proportional to the price level. In essence, these models are equivalent to y=p*=y*=0 in (2).

The introduction of a Taylor rule allows us to analyse various policy regimes. If $f \rightarrow \infty$ and $y \rightarrow 0$, the over-riding priority of monetary policy is to achieve a price level of p_t^*. This is equivalent to an inflation target of π_t^*, where $p_t^* = \pi_t^* + p_{t-1}$. If $y \rightarrow \infty$ and $f \rightarrow 0$, there is an output target. If $y = f$, there is a target for nominal GDP. The parameters of the Taylor rule affect the extent to which monetary policy accommodates inflation. With an inflation target, policy does not accommodate inflation as real interest rates increase whenever inflation rises above the target. With an output target, changes in the price level do not change the real interest rate and so monetary policy fully accommodates inflation.

Substituting (2) into (1), we can summarise aggregate demand as

$$(3) \qquad y_t = y_c - \frac{gf}{1+gy} p_t + \frac{g}{1+gy}(E_{t-1}p_{t+1} - E_{t-1}p_t) + \frac{v_t}{1+gy}$$

where $y_c = \{\bar{y} - g(i^* - f p^* - y y^*)\}/(1+gy)$. The slope of this aggregate demand curve depends on the policy regime. The curve is horizontal if there is an inflation target, is vertical if there is an output target and has a conventional negative slope if there is no target (see also, Taylor, 1999a).

2b) Aggregate Supply

We use a standard model of aggregate supply. We assume there are a large number of identical monopolistically competitive firms. Each firm's technology is described by a simple production function,

$$(4) \qquad y_{jt} = a + \ell_{jt} + x_t \,,$$

where ℓ is employment, x is a supply shock, a is a constant and j indexes the firm. We follow the literature (eg. Alogoskoufis, 1992 and Bleaney, 2001) in assuming that the supply shock follows a random walk, $x_t = d_x + x_{t-1} + e_t$. The demand for each firm depends positively on aggregate demand and negatively on its relative price:

$$(5) \qquad y_{jt} = y_t - h(p_{jt} - p)$$

where y_t is given by (3) From a standard profit maximisation problem and using equations (4) and (5), the price chosen by each firm is:

(6) $$p_{jt} = m + w_{jt} - a - x_t$$

where w_j is the nominal wage and $m = (1-1/h)^{-1}$ is the mark-up of price over marginal cost in firm *j*.

We assume that wage adjustment is staggered and described by a discrete time, Calvo-type utility-maximising wage contract model, (Calvo, 1983). At any given time, the wage at each firm has a fixed probability ***d*** of being adjusted to a new value of $\hat{w}$, and a fixed probability, *(1-d)*, of remaining fixed at the previous period's wage. The aggregate wage is given by the sum of all wage contracts still in force. With $d(1-d)^s$ being the fraction of wage contracts adjusted *s* periods before *t*, the aggregate wage is given by:

(7) $$w_t = d\sum_{s=0}^{\infty}(1-d)^s \hat{w}_{t-s} = d\hat{w}_t + (1-d)w_{t-1}$$

The wage chosen when adjustment occurs is forward-looking:

(8) $$\hat{w}_t = [1-(1-d)b]\sum_{s=0}^{\infty}(1-d)^s b^s E w^*_{t+s}\hat{w}_t = [1-(1-d)b]w^*_t + (1-d)b E_t \hat{w}_{t+1}$$

where w^*_t is the optimal wage common to all union that adjust their wage contracts in period t. We assume that this is given by,

(9) $$w^*_t = w^* + E_{t-1}p_t + s\, E_{t-1}(y_t - y^*)$$

where w^* is desired real wage growth (assumed constant for simplicity), y^* is a reference level of output and s measures the elasticity of real wages with respect to output. Equation (9) can be derived from almost any model of wage formation.

We then use (8) to express wage contracts in terms of w_t and use (6) to express w_t in terms of prices. This gives the following equation for the aggregate price level,

$$
\begin{aligned}
p_t = {} & \frac{(1-d)\hat{d}(w^* + m - a)}{1+(1-d)^2 b} + \frac{(1-d)}{1+(1-d)^2 b}(p_{t-1} + \hat{d} E_{t-1} p_t + b E_{t-1} p_{t+1}) \\
& + \frac{(1-d)\hat{d} s}{1+(1-d)^2 b} E_{t-1}(y_t - y^*) + \frac{(1-d)}{1+(1-d)^2 b}(x_{t-1} + b E_{t-1} x_{t+1}) - x_t
\end{aligned} \tag{10}
$$

where $\hat{d} = d(1/(1-d) - b) > 0$ and increasing in d. Defining inflation as $p_t = p_t - p_{t-1}$ and taking expectations, we can summarise the supply side of our model as

$$
\begin{aligned}
E_{t-1} p_t = {} & b E_{t-1} p_{t+1} + \hat{d}(w^* + m - a) + \hat{d} s \, E_{t-1}(y_t - y^*) \\
& + x_{t-1} - \left(\frac{1+(1-d)^2 b}{1-d} \right) E_{t-1} x_t + b E_{t-1} x_{t+1}
\end{aligned} \tag{11}
$$

This aggregate supply or Phillips curve is similar to others in the literature (e.g Taylor, 1999b, Mankiw, 2000 or Holden and Driscoll, 2001).

2c) Inflation Persistence

We substitute the aggregate demand curve, equation (3), into the aggregate supply curve, equation (11), to obtain

(12) $$p_t = \frac{(1-d)}{1+b(1-d)^2}\left(\overline{p} + \left(1 - \frac{gs(1+f)}{1+gy}\right)\hat{d}E_{t-1}p_t + \left(1 + \frac{gs\hat{d}}{b(1+gy)}\right)bE_{t-1}p_{t+1} + p_{t-1} + s_t \right)$$

where, $\overline{p} = \hat{d}(w^* + m - a + s(y_c - y^*))$ and $s_t = x_{t-1} - \left(\frac{1+b(1-d)^2}{1-d}\right)x_t + bE_{t-1}x_{t+1}$.

Forming expectations and rearranging equation (12) we obtain,

(13) $$E_{t-1}p_{t+1} - \frac{1}{q}\left(\frac{1+b(1-d)^2}{(1-d)} + \hat{d}(1 - \frac{gs(1+f)}{1+gy})\right)E_{t-1}p_t + \frac{p_{t-1}}{q} = -\frac{(\overline{p} + E_{t-1}s_t)}{q}$$

where, $q = b + \frac{gs\hat{d}}{1+gy} > 0$. Equation (13) can be written as

(14) $$\left(F^2 - (l_1 + l_2)F + l_1 l_2\right)LE_{t-1}p_t = -l_1 l_2(\overline{p} + E_{t-1}s_t)$$

where l_1 and l_2 are the smaller and larger roots respectively of (14), L is the lag operator and F is the forward operator. Expressing (14) as

(15) $$E_{t-1}p_t = l_1 p_{t-1} + l_1 \sum_{i=0}^{\infty} l_2^{-i}(\overline{p} + E_{t-1}s_{t+i})$$

substituting (15) into (12) and then taking first differences, we obtain

(16) $$p_t = l_1 p_{t-1} + \frac{l_1 b}{1-ql_1}d_x + \frac{l_1}{1-ql_1}\left(1 - \frac{q(1-d)}{1+b(1-d)^2}\right)e_{t-1} - \frac{(1-d)}{1+b(1-d)^2}e_t$$

where

(17) $$l_1 = \frac{1}{2q}\left(\frac{1+b(1-d)^2}{(1-d)} - \hat{d}\left(1-\frac{gs\,(1+f)}{1+gy}\right)\right)\left(1-\sqrt{1-\frac{4q\,(1-d)^2}{1+b\,(1-d)^2-(1-d)\hat{d}\,(1-\frac{gs\,(1+f)}{1+gy})}}\right)$$

From (17), we can show (see appendix for details)

(18) $$\frac{d\,l_1}{df} < 0; \quad \frac{d\,l_1}{dy} > 0\,.$$

Equations (16)-(18) comprise our model of the persistence of inflation. We find that the parameters of a Taylor rule for monetary policy affect the persistence of inflation (although the targets p* and y* do not). Inflation is less persistent when policymakers place a greater emphasis on the price level or a lesser emphasis on output. We therefore predict that inflation will be less persistent with an inflation target.

3) Empirical Evidence

In this section we present evidence on how the persistence of inflation is affected by inflation targeting. We consider those OECD economies that adopted inflation targets in the late 1980s or early 1990s, namely New Zealand (adopted inflation targeting in 1989Q3), Australia (1993Q2), Canada (1991Q1), Sweden (1993Q1) and the UK (1992Q3). We also consider Finland and Spain, which adopted inflation targets (in 19932Q1 and 1994Q1, respectively) but abandoned these upon entering EMU in 1998Q2 (see Bernanke at al, 1999 for further institutional details). We use the consumer price index to measure prices throughout and use both annual and quarterly data to ensure our findings are robust.

We first examine the time series properties of our data, testing for unit roots. For our quarterly data, we test for seasonal unit roots, using the HEGY test (Hylleberg et al, 1990). The HEGY test identifies the precise nature of seasonal integration and allows us to model any seasonal unit roots accordingly. The following auxiliary regression is undertaken:

$$j\,(L)z_{4t} = p_{1}z_{1,t-1} + p_{2}z_{2,t-1} + p_{3}z_{3,t-2} + p_{4}z_{3,t-1} + e_{t} \qquad (19)$$

where $j\,(L) = 1 - \sum_{j=1}^{n} j_{\,j} L^{j}$ is a stationary autoregressive polynomial of order n in L. Deterministic variables are left out of the equation for simplicity but are included in the empirical estimates. The *z*-variables are given by:

$$z_{1t} = (1 + L + L^{2} + L^{3})\,p_{t}$$
$$z_{2t} = -(1 - L + L^{2} - L^{3})\,p_{t}$$
$$z_{3t} = -(1 - L^{2})\,p_{t}$$
$$z_{4t} = (1 - L^{4})\,p_{t}$$

where p is the log of consumer prices. One-period lags of the dependent variable are included in the tests.

The results of the HEGY tests are presented in Table 1). The null hypotheses of $p_{2} = 0$, and $p_{3} \cap p_{4} = 0$ are rejected at the 5%-level for all countries, implying the absence of semiannual, complex and annual unit roots. However, the null hypothesis of $p_{1} = 0$ cannot be rejected at the 5%-level for any of the countries. This suggests that consumer prices in quarterly data contain a zero-frequency unit root and therefore that first-differences is the appropriate filter for making the series stationary. We examined the time series properties of our annual data using simple ADF tests. We found that prices were clearly I(1) in each country (the results are not presented to save space, but are available from the authors on request). We therefore define the rate of inflation using both annual and quarterly data as $\pi_t = p_t - p_{t-1}$, where p is the log of the consumer price index.

To examine the impact of inflation targets on inflation persistence, we consider simple regression models of the form

(20) $$\pi_t = \alpha + (\beta_1 + \beta_2 FX_t + \beta_3 IT_t)\, \pi_{t-1} + u_t$$

where FX is an indicator variable that equals unity during periods of fixed exchange rates and equals zero in other periods; IT is an indicator variable that equals unity during periods where an inflation target was in operation and equals zero in other periods and u is an error term. We use the White (1980) procedure to correct our estimated standard errors for heteroskedasticity and the Newey-West (1987) estimator to adjust standard errors for serial correlation.

Equation (20) is similar to other models in the literature on inflation persistence. These models typically interact lagged inflation with indicators of institutional presence or economic events. For example, Alogoskoufis and Smith (1991), Alogoskoufis (1992) and Bleaney (2001) use indicators of fixed exchange rates, corresponding to β_3=0 in (20). Other authors, eg Burdekin and Siklos (1999) also include indicators of other events, for example oil shocks and structural changes at Central Banks.

Estimates of (20) are presented in table 2). In every country, the estimate of β_3 is negative and significantly different from zero using both annual and quarterly data. Indeed, we can only reject the hypothesis that inflation targets have eliminated inflation persistence (H_0: $\beta_1+\beta_3=0$) in the case of the UK using quarterly data and cannot reject the hypothesis for any country when using annual data. These finding provide strong evidence in favour of our hypothesis that adopting an inflation target will reduce the persistence of inflation. The only other evidence on this is in Siklos (1999), who finds more ambiguous results using data up to 1997 (this may be because we have more observations from the inflation targeting regime).

The impact of exchange rate regimes on inflation persistence is less clear. We find a significantly lower rate of persistence during fixed exchange rates for Spain and the UK, which is consistent with Alogoskoufis and Smith (1991) and Alogoskoufis (1992), but no consistent significant effect in Australia, New Zealand, Finland and Sweden. This is broadly consistent with the results in Burdekin and Siklos (1999), who argue that wars, oil shocks or changes in Central Bank statutes have at least as great an impact on inflation persistence as exchange rates regimes.

We investigated the robustness of these findings in several ways (the results of these lengthy experiments are not reported but are available from the authors). First, we used alternative measures of inflation. We estimated (20) where inflation was defined as $\pi_t = p_t - p_{t-2}$ and $\pi_t = p_t - p_{t-4}$. We found broadly similar results; in particular, we continued to find large and significant reduction in inflation following the introduction of an inflation target. Second, we estimated an augmented model that allowed the intercept to vary between policy regimes, to allow for changes in the equilibrium inflation rate between regimes (Bleaney, 2001). We continued to find that the persistence of inflation was lower when there was an inflation target, although there was an effect on estimates for the fixed exchange rate regime, similar to Bleaney (2001). Third, we included the measures of oil shocks and changes in Central Bank statutes that were identified as significant by Burdekin and Siklos (1999) and Siklos (1999). We again continued to find that the adoption of inflation targets lead to a reduction in inflation persistence, although most our estimates again became less well determined. Fourth, we assessed the importance of mispecification apparent in the estimates in Table 2). Such mispecification is not surprising as we estimate a very simple model; similar findings are reported in the literature (eg Burdekin and Siklos, 1999). Our use of the White (1980) and Newey-West (1987) corrections should ensure that our estimates are robust to this. We found we could eliminate mispecification by including more lags of the dependent variable and by including dummies for time periods associated with marked volatility. We again continued to find that inflation targets are associated with less inflation persistence. Fifth, we estimated the model $\pi_t = \alpha_t + \beta_t \pi_{t-1} + u_t$ using both Kalman Filter and rolling window techniques. Although our estimates were not as precise as those reported in Table 2, we continued to find that inflation persistence was lower in the 1990s than in the preceding two decades. Overall, therefore, it seems that our conclusions are robust.

Finally, we summarise our findings by presenting estimates of the pooled model

$$\pi_{it} = \alpha + (\beta_1 + \beta_2 FX_{it} + \beta_3 IT_{it}) \pi_{it-1} + u_t \tag{21}$$

where i indexes the country and t indexes time. Our estimates, presented in Table 3, confirm the results of the country-by-country estimates in Table 2. The introduction of inflation targets leads to a large reduction in the persistence of inflation. Using annual data, the persistence of inflation fell from 0.54 to 0.16 following the introduction of inflation targets. We cannot reject the hypothesis that the persistence of inflation was eliminated. Using quarterly data, the persistence of inflation falls from 0.62 to 0.21, although in this case we can reject the hypothesis that inflation persistence is eliminated.

4) Conclusion

This paper has argued the persistence of inflation is lower when there is an finflation target, so inflation is more responsive to monetary policy when inflation is the main focus of policy. We presented a model in which inflation targeting reduces inflation persistence by reducing the extent to which monetary policy accommodates inflation. We then presented evidence from seven countries that adopted inflation targets in the late 1980s-early 1990s. We showed that the persistence of inflation did indeed fall sharply after the introduction of an inflation target.

Inflation targets are a relatively recent innovation in monetary policy. Over time, as data accumulates, it will become possible to analyse the impact of inflation targets in greater detail, investigating different aspects such as the choice of a target value as opposed to a target range, or different definitions of inflation to be targeted. We intend to consider these issues in future work

Table 1)
HEGY unit root tests

	$t(\hat{p}_1)$	$t(\hat{p}_2)$	$F(\hat{p}_3, \hat{p}_4)$
Canada	-3.31	-7.91	38.17
Australia	-2.45	-4.75	42.93
New Zeal.	-2.07	-6.49	47.42
Finland	-1.22	-5.32	55.33
Spain	-1.71	-5.42	50.91
Sweden	-2.30	-5.91	55.37
UK	-2.07	-5.44	36.95

A time-trend, constant term, seasonal dummies and a lagged dependent variable are included in the estimates. Estimation period: 1946.Q2-2001.Q2, which yields 220 observations. Critical values for T = 200 at the 5% level are (see Hylleberg *et al*, 1990: $t(p_1)$ = -3.49, $t(p_2)$ = -2.91, and $F(p_3, p_4)$ = 6.57.

Table 2)

Parameter estimates of inflation persistence

$$\pi_t = \alpha + (\beta_1 + \beta_2 FX_t + \beta_3 IT_t)\, \pi_{t-1} + u_t$$

(a) quarterly data

sample 1945Q1-2001Q2

	β_1	β_2	β_3	DW	R^2	HET	RES	SC	Ho:(β_1=β_3)
Canada	0.69(13.0)	-0.40(5.10)	-0.56(3.60)	2.17	0.46	4.79	1.19	1.41	0.79
Australia	0.65(9.15)	0.03(0.44)	-0.47(2.68)	2.42	0.49	19.5	0.58	2.61	1.71
New Zeal.	0.57(3.83)	0.14(0.95)	-0.40(2.44)	2.25	0.48	21.4	1.51	1.15	2.48
Finland	0.51(6.18)	-0.06(0.46)	-0.76(3.19)	2.29	0.26	18.2	1.41	3.24	2.11
Spain	0.61(9.16)	-0.31(3.49)	-0.79(4.41)	2.25	0.42	19.2	0.20	2.38	0.61
Sweden	0.37(4.83)	-0.05(0.49)	-0.43(3.25)	2.21	0.21	12.7	0.40	2.46	0.11
UK	0.68(6.09)	-0.42(3.30)	-0.30(2.06)	2.20	0.55	63.3	0.18	1.22	7.32

(b) annual data

sample 1946-2001

	β_1	β_2	β_3	DW	R^2	HET	RES	SC	Ho:(β_1=β_3)
Canada	0.88(5.38)	-0.45(3.14)	-0.57(2.34)	1.85	0.52	6.64	1.14		0.20
Australia	0.79(11.1)	-0.19(0.99)	-0.50(2.06)	1.71	0.54	11.4	0.14		1.32
New Zeal.	0.77(10.0)	-0.12(0.91)	-0.54(2.71)	2.10	0.54	6.84	1.13		1.24
Finland	0.63(4.32)	-0.34(1.86)	-1.90(2.93)	2.01	0.41	8.78	1.55		3.35
Spain	0.72(10.3)	-0.24(1.83)	-0.63(3.42)	1.77	0.54	3.42	0.22		0.17
Sweden	0.68(7.48)	-0.28(2.93)	-0.68(2.64)	1.99	0.47	1.14	1.40		0.00
UK	0.77(8.46)	-0.44(3.16)	-0.50(3.34)	1.77	0.68	15.1	0.14		2.34

Notes: The numbers in parentheses are absolute *t*-statistics. Constants and seasonal dummies are included in the estimates but not shown. The *t*-values are based on White's heteroscedasticity consistent covariance matrix. Estimation period: 1946.Q2-2001.Q2. DW = Durbin-Watson test for first order serial correlation, HET = Breusch-Pagan LM test for heteroscedasticity, and is distributed as $c^{2}(6)$ under the null hypothesis of no heteroscedasticity, RES is Ramsey's RESET test with the predicted value squared as additional regressor, and is distributed as *F*(1,213) under the null hypothesis of no functional form problems, and *SC* is a LM test for 1-4 order serial correlation and is distributed as *t*(217) under the null hypothesis of no 1-4 order serial correlation. Ho:(β_1=β_3) is a test of the null hypothesis β_1=β_3. It is distributed as $c^{2}(1)$ under the null

Table 3)
Pooled parameter estimates of inflation persistence

$$\pi_{it} = \alpha + (\beta_1 + \beta_2 FX_{it} + \beta_3 IT_{it}) \pi_{it-1} + u_t$$

(a) quarterly data
sample 1945Q1-2001Q2

β_1	β_2	β_3	DW	R^2	Ho:(β_1=β_3)
0.62(27.0)	-0.22(6.70)	-0.41(4.85)	2.15	0.64	6.08

(b) annual data
sample 1946-2001

β_1	β_2	β_3	DW	R^2	Ho:(β_1=β_3)
0.54(1501)	-0.39(6.17)	-0.38(2.48)	1.96	0.62	1.05

see notes to table 2)
Additional Notes: R^2 is based on Buse's raw-moment R^2.

Appendix

Derivation of equation (18)

From equation (17) and using the definitions of $l_1 l_2$ and $l_1 + l_2$, we can show that

(A1) $$\frac{d\,l_1}{df} = -\frac{1}{2q}\left(\frac{\hat{d}sg}{1+gy}\right)\left(\frac{l_1 + l_2}{l_2 - l_1} - 1\right) < 0\,.$$

Since $l_2 > l_1$ and $l_1 + l_2 > l_2 - l_1$ we also have $\frac{l_1 + l_2}{l_2 - l_1} > 1$. Therefore, for any value of $d < 1$, we obtain $q, \hat{d} > 0$, and hence, $\frac{d\,l_1}{df} < 0$ and so inflation targeting reduces inflation persistence.

Conversely, we can show that the effect of output stabilisation on inflation persistence is positive,

(A2) $$\frac{d\,l_1}{dy} = \left(\frac{\hat{d}sg^2}{(1+by)^2}\right)\left(\frac{l_1(1 - l_1 + f)}{(l_2 - l_1)(l_1 + l_2)^2}\right) > 0\,.$$

Since for convergence the small root of the dynamic equation is required to be less than unit $l_1 < 1$ then $1 - l_1 + f > 0$ and so given $l_2 > l_1$ for any value of $d < 1$, $\frac{d\,l_1}{dy} > 0$.

References

Agenor, P-R and M Taylor, 1992, "Testing for Credibility Effects", *IMF Staff Papers*, 39, pp 545-71

Alogoskoufis, G, 1992, "Monetary Accommodation, Exchange Rate Regimes and Inflation Persistence", *Economic Journal*, 102, pp 461-80

Alogoskoufis, G, and R Smith, 1991, "The Phillips Curve, the Persistence of Inflation and the Lucas Critique: Evidence from Exchange Rate Regimes", *American Economic Review*, 81, pp 1254-71

Bernanke, B., T. Laubach, F. Mishkin and A. Posen, (1999), *Inflation targeting: lessons from the International Experience*, Princeton University Press.

Bleaney, M, 2001, "Exchange Rate Regimes and Inflation Persistence", *IMF Staff Papers*, 47, pp 387-402

Burdekin, R and P Siklos, 1999, "Exchange Rate Regimes and Shifts in Inflation Persistence: Does Nothing Else Matter?", *Journal of Money Credit and Banking*, 31, pp 235-47

Calvo, G, 1983, "Staggered Contracts in a Utility-Maximizing Framework," *Journal of Monetary Economics*, 12, pp 383-98.

Holden, S and J Driscoll, 2001, "A Note on Inflation Persistence", NBER working paper 8690

Hylleberg, S, R Engle, C Granger and B Yoo, 1990, “Seasonal Integration and Cointegration,” *Journal of Econometrics*, 44, 159-176.

Mankiw, G, 2000, ""The Inexorable and Mysterious Tradeoff Between Inflation and Unemployment", *Economic Journal*, 111, pp C45-61

Newey, W and K West, 1987, "A Simple Positive Semi-Definite Heteroskedasticity and Autocorrelation Consistent Covariance Matrix", *Econometrica*, 55, pp 703-8

Siklos, P, 1999, "Inflation-Target Design: Changing Inflation Performance and Persistence in Industrial Countries", *Federal Reserve Bank of St Louis Review*, March-April, pp 47-58

Taylor, J, 1979, "Staggered Wage-Setting in a Macro Model", *American Economic Review, Papers and Proceedings*, 69, pp 108-13

Taylor, J, 1993, "Discretion versus policy rules in practice", *Carnegie-Rochester Conference Series on Public Policy*, 39, pp 195-214.

Taylor, J, 1999a, “The Robustness and Efficiency of Monetary Policy Rules as Guidelines for Interest Rate Setting by the European Central Bank”, *Journal of Monetary Economics*

Taylor, J, 1999b, "Staggered Price and Wage Setting in Macroeconomics", in Taylor and Woodford (eds.), *Handbook of Macroeconomics*, Vol 1B, Elsevier, 1999

White, H, 1980, "A Heteroskedasticity-Consistent Covariance Estimator and a Direct Test for Heteroskedasticity", *Econometrica*, 48, pp 817-38