U.S. and U.K. In‡ation: Evidence on Structural Change in the Order of Integration

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Abstract

We employ smooth transition models to test the null hypothesis of a unit root in time series on U.S. and U.K. monthly in‡ation beginning in 1957. Under the alternative hypothesis the test allows for structural change from level-stationarity to di¤erence-stationarity. For both countries the hypothesis of a unit root is rejected and it is estimated that rapid structural change began in 1970:6 in U.K. in‡ation and 1973:6 in U.S. in‡ation.

Acknowledgements

The views expressed in this paper belong to the author and do not necessarily re‡ect the view of the Department of Economics, TCD.

1 Introduction

Although the hypothesis of a unit root in in‡ation rates is still widely supported in the empirical literature, for details see Culver and Papell (1997), Barsky (1987) found distinctive changes in the degree of persistence of U.S. and U.K. in‡ation between 1870 and 1979, corresponding to di¤erent monetary regimes. He argues that the correlation of nominal interest rates and in‡ation predicted by the Fisher e¤ect is not always found in empirical work because of the changing statistical properties of in‡ation over time. Barsky's evidence suggests that post-Second World War in‡ation in the U.S. and U.K. may have contained a unit root after 1960, but before 1960 it was level-stationary.

Further evidence on structural change in the statistical properties of U.S. and U.K. in‡ation has been presented since Barsky (1987). For example using data from 1946 to 1992 and employing a Markov switching model, Evans and Wachtel (1993) calculate the probability of U.S. in‡ation being a random walk process, …nding it to be much higher between the early 1970s and mid-1980s than at any other time over their sample period. Note that Evans and Wachtel (1993) do not discuss speci…c economic reasons for this type of structural change however Alogoskou…s and Smith (1991), who propose that structural change in the persistence of U.S. and U.K. in‡ation took place around the early 1970s, believe the collapse of the Bretton Woods system of …xed exchange rates to be the cause. They argue that under ‡oating exchange rates monetary policy in the U.S. and U.K. has been more accommodating of shocks to in‡ation than under the Bretton Woods system, hence the increase in persistence of those shocks.

As well as being relevant to decisions regarding in‡ation policy, the presence of structural change in the in‡ation rate has important implications for the understanding of in‡ation forecasts. Evans and Wachtel (1993) note that the systematic di¤erences between survey forecasts of in‡ation and actual in‡ation have in the past been interpreted by some authors as evidence against rationality. However Evans and Wachtel (1993) point out that if forecasts are made across di¤erent regimes, or if they are made whilst learning about past regime changes or anticipating future regime changes, then whilst forecasters may be acting rationally their forecast errors might still be serially correlated.

Barsky (1987) draws his conclusions regarding changes in the persistence of in‡ation mainly from the analysis of cumulative periodgrams, spectral, and autocorrelation properties of the data. Evans and Wachtel (1993) employ a Markov switching model allowing for a stationary AR(1) process in one regime and a random walk in the other. In this paper we focus on statistically testing the null hypothesis that the monthly inflation rate in the U.S. and U.K. was difference-stationary throughout 1957:1 - 1998:12, against the alternative hypothesis that it began the period as a level-stationary process and at some point structural change to difference-stationarity occurred.[1] The test that we employ was proposed by Newbold et al. (2000) and the techniques involved allow for smooth or discrete structural change under the alternative hypothesis. Furthermore a graphical representation of the structural change is obtained in terms of a transition in the parameters of the testing model.

The Markov switching model estimated by Evans and Wachtel (1993) suggests the possibility of structural change in U.S. inflation from difference-stationarity back to level-stationarity in the mid-1980s. Furthermore the policy of inflation targeting adopted by the U.K. in 1992 suggests that for this country the possibility of a return from difference-stationarity back to level-stationarity in the 1990s is plausible. With monetary policy in the U.K. after 1992 being increasingly focused on maintaining inflation within a target range, the persistence of inflation in addition to its level may have been affected. To investigate whether inflation in the U.S. and U.K. did return to level-stationarity in the 1980s or 1990s, we use data from 1974:1 - 1998:12 and test the null hypothesis of difference-stationarity throughout this sub-sample, against the alternative hypothesis of structural change from difference-stationarity to level-stationarity. The period 1974:1 - 1998:12 was chosen specifically because it is believed that by 1974:1, inflation rates in the U.S. and U.K. had become difference-stationary processes.

From the analysis of our full sample of U.S. and U.K. inflation (1957:1 - 1998:12) strong rejections of the unit root null hypothesis are obtained. It is estimated that structural change in the U.K. inflation rate from level-stationarity to difference-stationarity began in 1970:6 and was completed by 1971:1. For the U.S. it is estimated that structural change from level-stationarity to difference-stationarity began in 1973:6 and was completed by 1974:1. Our results appear to be robust to the removal of noticeable outliers and to whether raw data or seasonally adjusted data are used. The investigation of U.S. and U.K. inflation over the period 1974:1 - 1998:12 for possible structural change back to level-stationarity, reveals no significant evidence against the unit root hypothesis.

[1] Assuming an autoregressive generating process the hypothesis of difference-stationarity implies that shocks to the inflation rate have a permanent effect. In finite samples the correlogram dies out very slowly. The hypothesis of level-stationarity implies that shocks to the inflation rate decay over time, the pattern of the decay depending on the size and sign of the coefficients of the generating process.

In the next section of the paper we outline the test proposed by Newbold et al. (2000) of the null hypothesis that a time series is difference-stationary, I(1), with the alternative hypothesis of structural change from level-stationarity to difference-stationarity, I(0) to I(1). It is also explained how this test can be used to test the same null hypothesis against the alternative hypothesis of structural change from I(1) to I(0). In section 3 the empirical results on possible structural change from I(0) to I(1) in U.S. and U.K. inflation 1957:1 - 1998:12 are discussed. In section 4 the empirical results on possible structural change from I(1) to I(0) in U.S. and U.K. inflation 1974:1 - 1998:12 are discussed. Section 5 concludes.

2 Testing for Structural Change in the Order of Integration of the Inflation Rate

Our full data set for the empirical work in this paper consists of 504 monthly observations on the consumer price inflation rate in the U.S. and U.K. from 1957:1 to 1998:12. The inflation rate for each country is computed by taking the first difference of the natural logarithm of the consumer price index reported in the International Monetary Fund's database on International Financial Statistics. In addition to the raw data we also investigate seasonally adjusted data derived from the application of monthly seasonal dummies. The U.S. and U.K. raw data is plotted as Figure 1(a) and Figure 1(b) respectively and from these graphs it is immediately clear that visually, the inflation rates in these countries strongly resemble I(0) processes over certain periods. However a feature of unit root processes with zero drift is that periods of the process might appear to be I(0).

Consider the following model for time series on the inflation rate y_t

$$y_t = \alpha_1 + \alpha_2 S_t(\gamma;\tau) + \beta_1 y_{t-1} + \beta_2 S_t(\gamma;\tau) y_{t-1} + \varepsilon_t \qquad (1)$$

where $\varepsilon_t \sim iid(0;\sigma_\varepsilon^2)$ and $S_t(\gamma;\tau)$ is the logistic function based on a sample of size T,

$$S_t(\gamma;\tau) = [1 + \exp\{-\gamma(t - \tau T)\}]^{-1}. \qquad (2)$$

Assuming γ and τ are parameters and that $\gamma > 0$ the logistic function moves from 0 to 1 monotonically as $t \to \infty$. Thus (1) represents a transition from one AR(1) process

$$y_t = \alpha_1 + \beta_1 y_{t-1} + \varepsilon_t \qquad (3)$$

to another

$$y_t = (\alpha_1 + \alpha_2) + (\beta_1 + \beta_2)y_{t-1} + \varepsilon_t \tag{4}$$

as $t \to \infty$. The parameter τ determines the mid-point of the transition since when $t = \tau T$, $S_t(\gamma; \tau) = 0.5$. The parameter γ determines the speed of the transition, with larger values of γ corresponding to a faster transition. Note that as $\gamma \to \infty$ the transition from 0 to 1 becomes instantaneous at time $t = \tau T$ whilst if $\gamma = 0$, $S_t(\gamma; \tau) = 0.5 \ \forall \ t$.

The restriction $\beta_1 + \beta_2 = 1$ imposes difference-stationarity as $t \to \infty$: With this restriction imposed (1) can be re-arranged as

$$y_t - S_t(\gamma; \tau)y_{t-1} = \alpha_1 + \alpha_2 S_t(\gamma; \tau) + \beta_1(y_{t-1} - S_t(\gamma; \tau)y_{t-1}) + \varepsilon_t. \tag{5}$$

Depending on the value of β_1 the model given by (5) is consistent with y_t being $I(1)$ from $t = 1, 2, \ldots, T$ or with y_t initially being $I(0)$ with a transition to $I(1)$. Difference-stationarity throughout the sample implies $\beta_1 = 1$, whereas structural change from $I(0)$ to $I(1)$ implies $\beta_1 < 1$.[2]

Note that the model given by equation (5) can be augmented in the same way as the standard Dickey-Fuller test to account for additional serial correlation in the data by adding lags of Δy_t

$$\begin{aligned} y_t - S_t(\gamma;\tau)y_{t-1} &= \alpha_1 + \alpha_2 S_t(\gamma;\tau) + \beta_1(y_{t-1} - S_t(\gamma;\tau)y_{t-1}) \\ &\quad + \sum_{i=1} \phi_i \Delta y_{t-i} + \varepsilon_t. \end{aligned} \tag{6}$$

We might also want to impose restrictions on the intercepts α_1 and α_2. For example to restrict the transition to be from an $I(0)$ process to an $I(1)$ process with zero drift, we can impose $\alpha_1 + \alpha_2 = 0$ in (6), which can then be re-arranged as

$$\begin{aligned} y_t - S_t(\gamma;\tau)y_{t-1} &= \alpha_1(1 - S_t(\gamma;\tau)) + \beta_1(y_{t-1} - S_t(\gamma;\tau)y_{t-1}) \\ &\quad + \sum_{i=1} \phi_i \Delta y_{t-i} + \varepsilon_t. \end{aligned} \tag{7}$$

The formal test of the null hypothesis that the inflation rate y_t is $I(1)$ throughout the sample period against the alternative hypothesis of a transition from $I(0)$ to $I(1)$ is given

[2] If $\beta_1 = 1$ then note that the $S_t(\gamma;\tau)y_{t-1}$ terms on the left and right-hand-side of (5) cancel out, leaving a unit root process with a possible transition in the drift term from α_1 to $\alpha_1 + \alpha_2$. If $\beta_1 < 1$ then go back to equation (3) for confirmation that y_t will initially be level-stationary.

by the t-statistic for testing $\beta_1 = 1$ in (6) or (7); that is

$$t = \frac{\hat{\beta}_1 - 1}{Se(\hat{\beta}_1)} \tag{8}$$

where $\hat{\beta}_1$ is the nonlinear least squares (NLS) estimator of β_1. From hereafter the test calculated from the estimation of (6) is referred to as t_a and the test calculated from the estimation of (7) as t_b. Newbold et al. (2000) provide simulated critical values for t_a and t_b for the empirical sample size $T = 1300$ which are presented in our Table 1. We also simulated critical values for $T = 300$ as this is the size of the smallest empirical sample that is considered in this paper, 1974:1 - 1998:12, simulating under the null hypothesis of a random walk with iid standard normal error terms. Clearly from Table (1) there is little difference between the critical values simulated for $T = 1300$ by Newbold et al. (2000) and ours simulated for $T = 300$. Thus for the tests based on the full sample of 504 observations the critical values simulated by Newbold et al. (2000) are used.

As mentioned in the introduction, in addition to the analysis of structural change from I(0) to I(1) in our full sample of U.S. and U.K. inflation, the possibility of structural change in U.S. and U.K. inflation from I(1) to I(0) in the 1980s or 1990s is also investigated, using data from 1974:1 to 1998:12. Recall that the alternative hypothesis of the original test is structural change from I(0) to I(1). To test the null hypothesis of a unit root against the alternative hypothesis of structural change from I(1) to I(0), we apply the original test proposed by Newbold et al. (2000) to our data ordered in reverse. To clarify, define

$$x_t = y_{T+1-t} \tag{9}$$

and assume that $y_{1974:1}$ is an observation from an I(1) process, then the models for testing the null hypothesis that the inflation rate between 1974:1 and 1998:12 is I(1), against the alternative hypothesis of structural change from I(1) with drift to I(0), or I(1) without drift to I(0), are

$$\begin{aligned} x_t - S_t(\gamma;\tau)x_{t-1} &= \alpha_1 + \alpha_2 S_t(\gamma;\tau) + \beta_1(x_{t-1} - S_t(\gamma;\tau)x_{t-1}) \\ &\quad + \sum_{i=1} \theta_i \Delta x_{t-i} + \varepsilon_t \end{aligned} \tag{10}$$

and

$$
\begin{aligned}
x_t - S_t(\gamma,\tau)x_{t-1} &= \alpha_1(1 - S_t(\gamma,\tau)) + \beta_1(x_{t-1} - S_t(\gamma,\tau)x_{t-1}) \\
&\quad + \sum_{i=1}^{k} \theta_i \Delta x_{t-i} + \varepsilon_t
\end{aligned} \tag{11}
$$

respectively. As before the test statistic used is the t-statistic for testing $\beta_1 = 1$ compared to the critical values given in Table 1.

3 Empirical Results: Structural Change from I(0) to I(1) in U.S. and U.K. Inflation 1957:1 - 1998:12

Since it is not thought that U.S. and U.K. inflation has ever been nonstationary with drift over this period, only the results from estimating the model given by (7) are reported.[3] In all cases estimation was by NLS employing the OPTMUM sub-routine in GAUSS, and the general-to-specific strategy was used for choosing the value of k in (7), starting with $k = 12$ and including all lags up to the last lag significant at the 10% level.[4] Our results for the U.S. inflation rate are given in Table 2 and for the U.K. inflation rate in Table 3. The key estimated parameters in addition to the calculated test statistics t_b from the estimation of (7) for both the raw and seasonally adjusted data are included in these tables.

The second column of Table 2 gives the results from the estimation of (7) for the raw data on the U.S. inflation rate 1957:1 - 1998:12. Using the appropriate critical values given in Table 1 for large T, the calculated test statistic $t_b = -3.703$ means that the null hypothesis of the data being I(1) throughout the sample can be rejected at the 5% level of significance. The third column of Table 2 gives the results from the estimation of (7) for the seasonally adjusted U.S. inflation rate. In this case the calculated test statistic $t_b = -4.217$ means that the null hypothesis of the data being I(1) throughout the sample can now be rejected at the 1% level of significance. From the model estimated for the

[3] Note however that there are only trivial changes in our results if we compute tests from the estimation of (6).

[4] To avoid local minimisation of the sum of squares function we employ a fine grid-search over starting values for both γ and τ in the numerical optimisation procedure. To reduce the computational burden involved in estimating our structural change models we can concentrate the sum of squares function entirely with respect to the nonlinear parameters γ and τ. More details on the estimation of these types of structural change models are given in Sollis (1999).

raw data $\hat{\beta}_1 + S_t(\hat{\gamma}; \hat{\tau})\hat{\beta}_2$ is plotted as Figure 2(a), this is the estimated transition from $\hat{\beta}_1$ to $(\hat{\beta}_1 + \hat{\beta}_2)$ and reveals the exact timing and speed of the estimated structural change from I(0) to I(1). It is estimated that rapid structural change began in 1973:6 and ended in 1974:1. Note that the same transition is found in the seasonally adjusted data and thus no diagram is presented for this case. The identical nature of the structural change estimated in the raw and seasonally adjusted data can be confirmed by the similarity of the estimated parameters given in Table 2. The timing of the structural change that we find in U.S. inflation is consistent with the timing of the structural change suggested by the previous authors mentioned.

The structural change in our raw and seasonally adjusted data on U.S. inflation is estimated to have begun in 1973:6. As can be seen from Figure 1(a) there is a noticeable spike in the inflation rate close to this time (specifically in 1973:8) corresponding to the oil price shock. To check that our rejection of the unit root null hypothesis is not dependent on this single observation, t_b is also computed for the raw and seasonally adjusted U.S. inflation rate with this observation set to zero. Our rejection of the null hypothesis is robust to the removal of this spike since in this case $t_b = -4.511$ and $t_b = -4.391$ for the raw and seasonally adjusted data respectively, which are both rejections at the 1% level of significance.

Our results for the U.K. inflation rate 1957:1 - 1998:12 are given in Table 3. The test statistic and key estimated parameters from (7) estimated for the raw data are given in the second column and for seasonally adjusted data in the third column. As in our investigation of U.S. inflation, experimenting with dummy variables to remove any noticeable outliers in our U.K. data revealed no substantial changes in the calculated test statistic or the pattern of structural change estimated. For both the raw data and the seasonally adjusted data the null hypothesis that the inflation rate contains a unit root can be rejected in favour of a rapid transition beginning in 1970:6 and ending in 1971:1. For the raw data $t_b = -3.939$ and for the seasonally adjusted data $t_b = -4.534$, thus the unit root null hypothesis can be rejected at the 5% and 1% levels of significance respectively. For comparison with Figure 2(a) we plot $\hat{\beta}_1 + S_t(\hat{\gamma}; \hat{\tau})\hat{\beta}_2$ from (7) estimated for the raw U.K. data as Figure 2(b).[5] Interestingly the estimated structural change in U.K. inflation is seen to take the same amount of time to be fully completed as that estimated to have occurred in the U.S. series, although it is estimated to have taken

[5] Again the estimated transition from $\hat{\beta}_1$ to $(\hat{\beta}_1 + \hat{\beta}_2)$ for the seasonally adjusted data is very similar to that estimated using the raw data hence only the latter is presented.

place three years earlier.

4 Empirical Results: Structural Change from I(1) to I(0) in U.S. and U.K. In‡ation 1974:1 - 1998:12

The Markov switching model of Evans and Wachtel (1993) suggests the possibility of structural change from I(1) to I(0) in U.S. in‡ation around the mid-1980s, and there is reason to believe that U.K. in‡ation may have become level-stationary in the 1990s. If this were true for our U.S. and U.K. data series then the models given by (6) and (7) are mis-speci…ed as they allow for structural change in only one direction, from I(0) to I(1). To investigate possible structural change from I(1) to I(0) in U.S. and U.K. in‡ation in the mid-1980s, sub-samples from our full data set are analysed, speci…cally the in‡ation rates over the period 1974:1 - 1998:12.

We believe that in‡ation in the U.S. and U.K. became I(1) at some point in the early 1970s, our estimation employing the full-sample of data suggesting that this structural change was completed by 1971:1 in the U.K. and 1974:1 in the U.S.. Thus 1974:1 appears to be a sensible starting point for the analysis of possible structural change in U.S. and U.K. in‡ation back from I(1), to I(0). Of course if there was structural change back to level-stationarity in the data then our full sample results are from a mis-speci…ed model, and therefore the …ndings of di¤erence-stationarity by 1971 and 1974 for the U.K. and U.S. respectively cannot be trusted. However experimenting with sub-samples starting at various di¤erent dates in the early 1970s reveals that the results in this section are robust to these changes.

To test the null hypothesis that in‡ation in the U.S. and U.K. was I(1) throughout the period 1974:1 - 1998:12, against the alternative hypothesis that it began the period as an I(1) process and at some point there was structural change to I(0), for each country the appropriate test t_b is computed for the respective sub-sample of data ordered in reverse. Thus the null of I(1) is tested against the alternative hypothesis of structural change from I(0) to I(1) in the reversed data, an equivalent test to that of the null of I(1) against the alternative hypothesis of structural change from I(1) to I(0) in the data ordered naturally.

For both countries the values of t_b obtained are given in Table 4. Comparing Table 4 with the appropriate critical values given in Table 1 reveals only weak support for the hypothesis of structural change back to level-stationarity in either the raw or seasonally adjusted time series on U.S. in‡ation. Similarly the null hypothesis that the U.K. in‡ation rate was I(1) throughout 1974:1 - 1998:12 cannot be rejected at the 10% level of signi…cance, irrespective of whether the raw or seasonally adjusted data is tested. Thus despite the visual appearance of level-stationarity in the …nal section of our samples on U.S. and U.K. in‡ation, the test employed reveals only weak statistical evidence against the hypothesis that the in‡ation rate in these countries remained I(1) after a break from I(0) in the early 1970s.

5 Conclusion

The techniques employed in this paper con…rm that over the period 1957:1 - 1998:12 the underlying statistical properties of in‡ation rates in the U.S. and U.K. were not stable. However rather than simply describing the possibility of structural change in in‡ation over this period, we statistically test the null hypothesis of a unit root (a hypothesis widely supported in the existing empirical literature) against the alternative hypothesis of a speci…c form of structural change, level-stationarity to di¤erence-stationarity. Our techniques allow for the endogenous estimation of structural change and for the structural change to be gradual or discrete. It is found that structural change in the U.S. in‡ation rate from I(0) to I(1) occurred in approximately six months, beginning in 1973:6. Statistically this model is preferred to the hypothesis of a unit root over the period 1957:1 - 1998:12. The possibility of structural change in the reverse direction, from I(1) to I(0), between 1974:1 and 1998:12 is also investigated. Our results show that over this period the null hypothesis of a unit root in U.S. in‡ation cannot be rejected at the 10% level of signi…cance.

For the U.K. in‡ation rate between 1957:1 and 1998:12 the null hypothesis of I(1) is also rejected in favour of structural change from I(0) to I(1). It is estimated that this structural change began in 1970:6 and was completed by 1971:1. Our investigation of in‡ation in the U.K. between 1974:1 and 1998:12 reveals no statistically signi…cant evidence of structural change back from I(1) to I(0).

It is often the case that practitioners will use a …xed parameter autoregressive model to test hypotheses concerning the time series properties of post-Second World War consumer price in‡ation, or indices of consumer prices. See for example the many empirical studies of money demand, the Fisher e¤ect, or the purchasing power parity theory. In these studies the …nding that in‡ation is either I(0), or I(1), is then typically followed by the estimation of long-run models and error correction models of short-run adjustment, containing inflation or prices as dependent or explanatory variables. Our evidence suggests that for the U.S. and U.K., the assumption of a constant level of integration when testing for the presence of a unit root in inflation and when subsequently estimating long-run and short-run models, is not tenable.

Table 1

Simulated Critical Values for Structural Change Tests

Significance	T = 1300		T = 300	
	t_a	t_b	t_a	t_b
1%	-4.403	-4.145	-4.481	-4.219
5%	-3.777	-3.611	-3.818	-3.645
10%	-3.748	-3.361	-3.750	-3.391

Table 2

Estimated Parameters and Tests for Structural Change from I(0) to I(1) in U.S. Inflation 1957:1 - 1998:12[6]

	U.S.	U.S. (seasonally adjusted)
t_b	-3.703^{*}	-4.217^{**}
$\hat{\alpha}_1$	.00077	-.00023
	(3.313)	(-1.387)
$\hat{\beta}_1$	.730	.699
	(10.035)	(9.805)
$\hat{\delta}$	1.209	1.310
	(4.634)	(6.792)
$\hat{\rho}$	.388	.388
	(4.892)	(5.246)
k	12	12
RSS	.00223	.00202

[6] * and ** denote significance at the 5% and 1% levels respectively, t-statistics for the estimated parameters are in parentheses and RSS is the residual sum of squares.

Table 3

Estimated Parameters and Tests for Structural Change from I(0) to I(1) in U.K. Inflation 1957:1 - 1998:12[7]

	U.K.	U.K. (seasonally adjusted)
t_b	-3.939*	-4.463**
b_1	.00112	-.00096
	(2.357)	(-2.257)
b_1	.639	.574
	(6.966)	(6.020)
b	1.109	.810
	(4.357)	(5.864)
b	.313	.315
	(4.367)	(5.224)
k	12	11
RSS	.0119	.0096

[7] * and ** denote significance at the 5% and 1% levels respectively, t-statistics for the estimated parameters are in parentheses and RSS is the residual sum of squares.

Table 4

Estimated Tests for Structural Change from I(1) to I(0)

in U.S. and U.K. In‡ation 1974:1 - 1998:12

	t_b	k
U.S.	-2.995	10
U.S. (seasonally adjusted)	-2.329	8
U.K.	-2.079	12
U.K. (seasonally adjusted)	-2.435	11

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Figure 1(a). U.S. Inflation 1957.1 – 1998.12

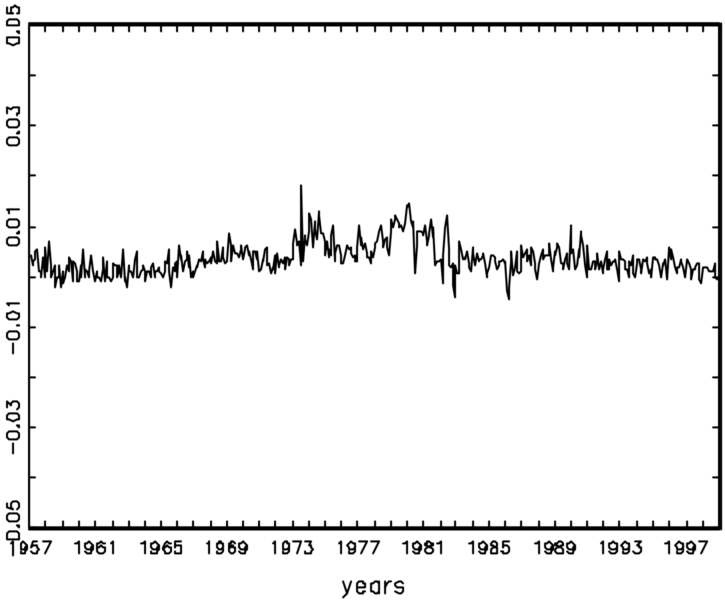

Figure 1(b). U.K. Inflation 1957:1 – 1998:12

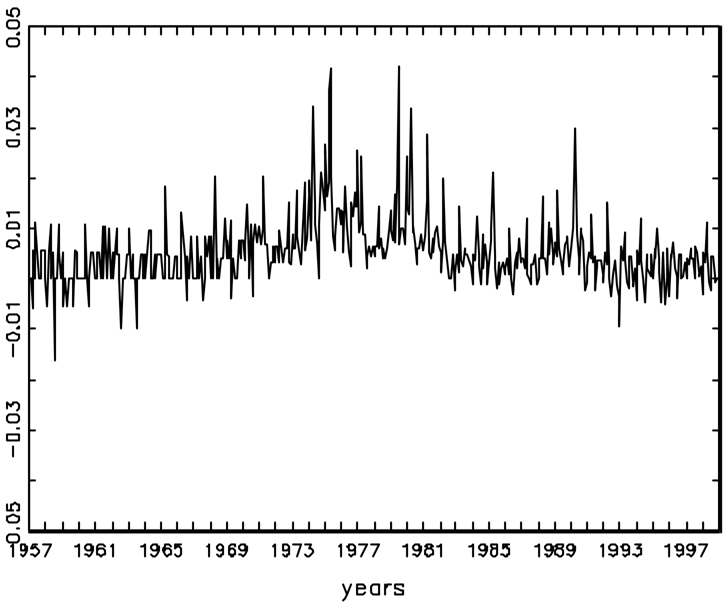

Figure 2(a). Estimated Transition from I(0) to I(1) in U.S. Inflation 1957:1 – 1998:12

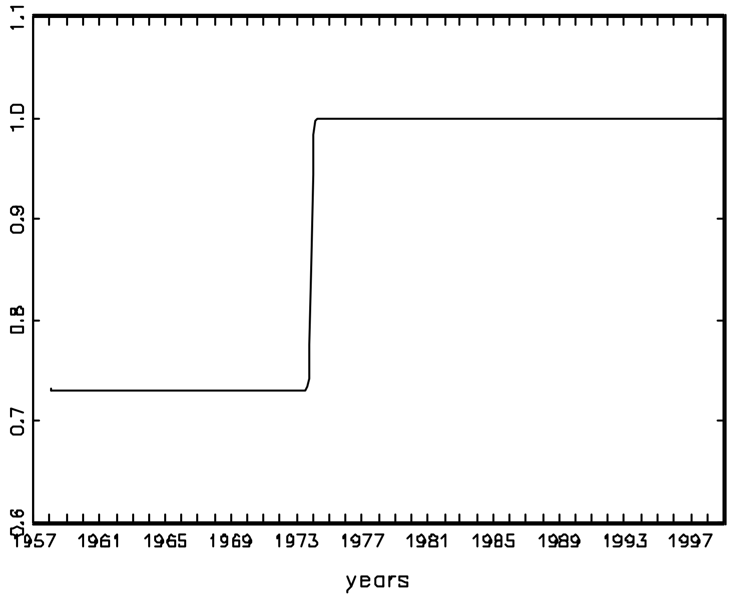

Figure 2(b). Estimated Transition from I(0) to I(1) in U.K Inflation 1957:1 – 1998:12