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Testing the Fisher effect in the presence of structural change: A case study of the UK, 1966-2007

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ABSTRACT

In this paper, we provide an empirical test of the Fisher effect using cointegration techniques, where the existence of instabilities in the cointegrating or long-run relationship is explicitly tested. The analysis is applied to the UK, a country that has been subject to potentially strong regime shifts, for the period 1966-2007. To this end, we apply some recent econometric techniques aimed to detect eventual structural changes, allowing the instability to occur at an unknown date.

1. INTRODUCTION

EMPIRICAL TESTING OF THE SO-CALLED 'FISHER EFFECT' is a traditional topic in monetary and financial economics. In fact, the greater or lower degree to which nominal interest rates incorporate the expected evolution of the inflation rate, without affecting the real interest rate, is an important issue for a number of relevant questions in both theory and policy. As an example, if the Fisher effect holds, the superneutrality of money would apply, and the nominal interest rate would be a good predictor of future inflation as well as a bad indicator of the kind of monetary policy followed. Furthermore, the Fisher effect would be a necessary condition for the validity of the consumption-based capital asset pricing model or CCAPM (Haliassos and Tobin, 1990).

The hypothesis dates back to Fisher (1896, 1930), who also provided its first empirical test. It is important to note that Fisher's own results showed that the hypothesis associated with his name would be satisfied only partially since, although the interest rate responded to changes in the inflation rate in the sense suggested by the theory, it did so by a smaller amount and with a substantial delay. In addition, Fisher pointed to the existence of money illusion as the ultimate explanation of his results, so that agents would be unable to distinguish changes in nominal values from changes in real values of the economic variables.[2]

The emergence of the literature on unit roots and cointegration has provided an important impulse to the empirical testing of the Fisher effect. Following the early work of Rose (1988), a number of further contributions seeking to test for the Fisher effect using cointegration techniques have appeared, with sometimes conflicting results; a non-exhaustive list would include, among others, Moazzami (1991), Mishkin (1992), Peláez (1995), Crowder (1997), Bajo-Rubio and Esteve (1998), and Koustas and Serletis (1999). More recently, the empirical analysis of the Fisher effect has turned to a nonlinear perspective; see, e.g., Bajo-Rubio, Díaz-Roldán and Esteve (2005), Lanne (2006), Christopoulos and León-Ledesma (2007), and Yoon (2010). A survey is provided in Neely and Rapach (2008).

On the other hand, the presence of structural breaks in the series may be reflected in the parameters of the estimated models that, when used for inference or forecasting, can induce misleading results. In general, structural breaks are a problem for the analysis of financial time series, since they are usually affected by either exogenous shocks or changes in policy regimes. This problem might be even more important for the empirical testing of the Fisher effect since, in the original contribution of Fisher, this is clearly a long-run phenomenon (see, e.g., the discussion in Mishkin, 1992, p.213). Thus, if the required time series were long enough, the possibility that these series are subject to structural changes would be accordingly higher.

There are some studies available analysing the role of structural changes on the evolution of real interest rates, e.g. Garcia and Perron (1996) or Siklos and Skoczylas (2002). However, and as far as we know, this is not the case for the empirical analyses of the Fisher effect using cointegration techniques, with the only exception being Bajo-Rubio and Esteve (1998). There, an empirical analysis of the Fisher effect was performed for the Spanish case over the period 1962-1996, where the possible presence of structural changes in both the trend of the series and the estimated long-run relationship between the nominal interest rate and the inflation rate was tested explicitly. In this paper, we go a step further by applying the more recent procedures developed by Bai and Perron (1998, 2003a, 2003b) for estimating and testing multiple structural break dates in the trend function of the series. A key feature of the Bai and Perron procedure is that it allows testing for multiple breaks at unknown dates, by successively estimating each break point using a specific-to-general strategy in order to determine consistently the number of breaks.

The empirical application is performed on data for the United Kingdom (UK). This country can provide an interesting case study, since it has been subject to potentially stronger regime shifts than elsewhere over recent years. Some candidate shocks would include, e.g., the large changes in the value of the sterling; the two oil crises; the impact of North Sea oil; the 'shadowing' of the German mark before sterling's entry into the exchange rate mechanism (ERM) of the European Monetary System; the withdrawal from the ERM and introduction of inflation targeting in October 1992; and the granting of operational independence to the Bank of England in May 1997. In addition, a significant turning point in macroeconomic policymaking in the UK, which could be taken as a change in policy regime, was the taking office of Margaret Thatcher as Prime Minister after the May 1979 election. Despite the similarity of ideological principles of the Thatcher government with others elected at that time (in particular, US President Ronald Reagan's), the impact on the UK's economy seemed to have been stronger, when compared with the previous situation. Actually, the actions of the 1979 Conservative government were characterised by a design and execution of macroeconomic policy that represented a radical departure from, and even repudiation of, the practice followed after World War II, based on Keynesian demand management and cooperation among the social partners; see, e.g., Buiter and Miller (1983) or Kaldor (1983).

Therefore, the objective of this paper is to provide an empirical test of the Fisher effect using cointegration techniques, for the UK during the period 1966-2007, where the existence of instabilities in the cointegrating or long-run relationship is tested explicitly. To this end, we apply recent econometric techniques aimed at detecting eventual structural changes, allowing the instability to occur at an unknown date. The rest of the paper is organised as follows: the theoretical framework is presented in Section 2, the data and macroeconomic policy background are briefly discussed in Section 3, and the empirical results are shown in Section 4; finally, Section 5 summarises the main results and policy implications.

2. Theoretical framework

In this section, we describe the procedure to test for the Fisher effect outlined in Bajo-Rubio and Esteve (1998). Our starting point will be the well-known Fisher equation, where the nominal interest rate consists of two components, the *ex-ante* real interest rate and the expected inflation rate:

$$i_t = r_t^e + \pi_t^e \tag{1}$$

where i_t is the nominal interest rate in period t, r_t^e is the *ex-ante* real interest rate, and π_t^e is the inflation rate expected in t-1 for the next period. Lenders would require a nominal interest rate to compensate them for any eventual loss in their purchasing power during the life of the loan; such a loss is represented by the expected inflation rate. With no money illusion (in the broad sense defined above; see note 2), a change in the expected inflation rate should be fully transmitted to the nominal interest rate, to keep the *ex-ante* real interest rate approximately constant in the long run.

Hence, the Fisher hypothesis could be tested starting from the following equation:

$$i_t = \alpha + \beta \pi_t^e \tag{2}$$

where the constant term would proxy the *ex-ante* real interest rate, and the lack of rejection of the null hypothesis $\beta=1$ would indicate the presence of a full Fisher effect. In addition, if we make the assumption of rational expectations, the expected inflation rate would match that true inflation rate, π_t, except for a random prediction error ε_t:

$$\pi_t^e = \pi_t + \varepsilon_t \tag{3}$$

so that, replacing (3) into (2), we would get:

$$i_t = \alpha + \beta \pi_t + \eta_t \tag{4}$$

where $\eta_t = \beta\varepsilon_t$.

Since the innovation ε_t (and hence η_t) is stationary,[3] and if i_t and π_t have a unit root, for the Fisher effect to be satisfied both variables should be cointegrated. In particular, an estimate of β not significantly different from one in the cointegrating regression (4) would indicate the presence of a full Fisher effect, so that i_t - π_t would be stationary. Notice also that, by definition:

$$r_t = i_t - \pi_t \tag{5}$$

where r_t is the *ex-post* real interest rate; and replacing (1) and (3) in (5):

$$r_t = r_t^e + \varepsilon_t \tag{6}$$

In other words, given the assumption of rational expectations, the *ex-ante* and *ex-post* real interest rates would only diverge by a random and stationary term, so that stationarity of the former implies stationarity of the latter.

However, it might occur that, in equation (4), the nominal interest rate and the inflation rate were cointegrated, but the estimate of β was significantly different from one. In this case, i_t - $\beta\pi_t$ would be stationary and r_t = i_t - π_t would be I(1); in particular, an estimate of β significantly lower than one would indicate the presence of a partial Fisher effect. On the other hand, such a case would be consistent with the presence of partial money illusion, so we could estimate an equation such as:

$$r_t = i_t - \pi_t = (\alpha + \beta\pi_t + \eta_t) - \pi_t$$

or:

$$r_t = \alpha + \beta'\pi_t + \eta_t \tag{7}$$

where $\beta' = \beta - 1$. In that equation, and assuming that π_t is I(1), r_t and π_t would be cointegrated and $r_t - \beta'\pi_t$ would be stationary. In addition, given that a linear combination of I(0) series is also I(0), $(i_t - \beta\pi_t) - (r_t - \beta'\pi_t) = i_t - r_t - \pi_t$ would be stationary.

Still, if the estimate of β in (4) were not significantly different from zero, then the estimate of β' in (7) would not be significantly different from -1. In this case, the Fisher effect would not apply and money illusion would be complete, since any change in the expected inflation rate would not be transmitted at all by lenders through the nominal interest rate, so this change would be fully reflected in the *ex-ante* real interest rate. Hence, r_t and π_t would be I(1) and cointegrated, and $r_t + \pi_t = i_t$ would be stationary.

Notice, on the other hand, that the above considerations would only apply in the presence of cointegration between it and π_t (or between r_t and π_t). Therefore, when both variables are not cointegrated, even if they are I(1), this would suggest a bad specification of the model and, therefore, the need to include some additional variables when estimating equation (4) (or equation (7)); see Owen (1993).

Summarising the previous discussion, our procedure to test for the Fisher effect will be as follows. First, we will test for the order of integration of the variables nominal interest rate, inflation rate, and *ex-post* real interest rate (where the latter two would proxy, respectively, the expected inflation rate and the *ex-ante* real interest rate, which are both non-observable). Then, if the nominal interest rate and the inflation rate were both I(1), we will estimate equation (4), so that:

- If the two variables were cointegrated and the estimate of β not significantly different from one, there would be a full Fisher effect so that changes in the expected inflation rate would be transmitted one-for-one to the nominal interest rate.

- If the two variables were cointegrated and the estimate of β significantly lower than one, there would be a partial Fisher effect so that changes in the expected inflation rate would be transmitted in a proportion $\beta<1$ to the nominal interest rate, due to the presence of some money illusion.

Notice that these two cases imply that the *ex-post* real interest rate would be either I(0) or I(1), respectively, which would mean an indirect test of the order of integration of that variable. Finally, if i_t and π_t were not cointegrated, introducing some additional variables, presumably influencing the nominal interest rate, would be justified when estimating equation (4).[4]

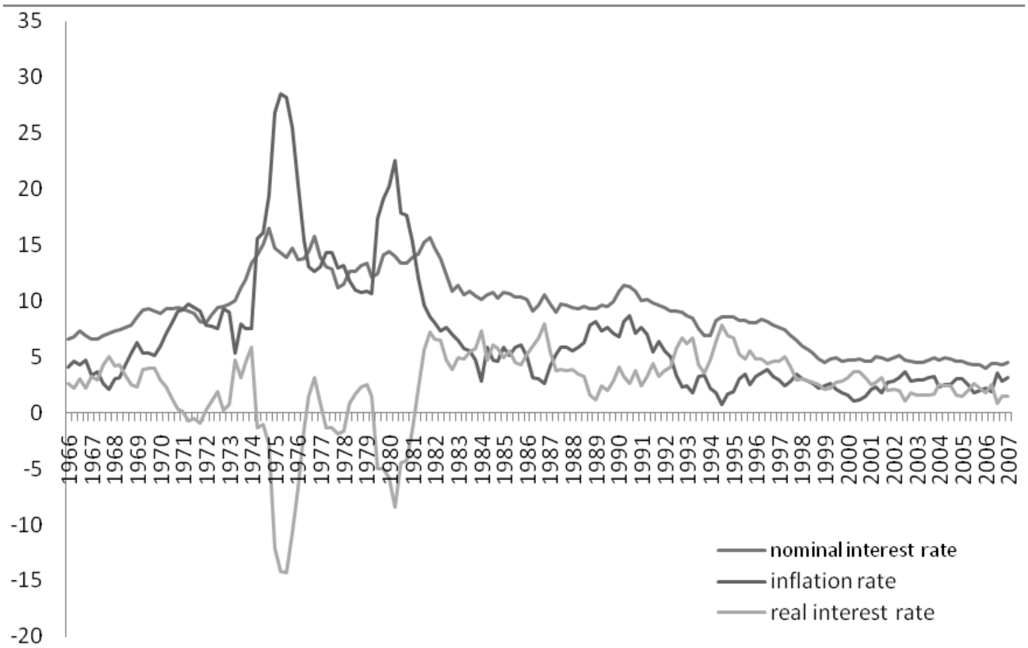
3. Data and macroeconomic policy background

In the empirical application, we will make use of quarterly data for the UK over the period 1966:1-2007:1, taken from OECD (2008). Specifically, we use the series on long-term government bond yields/over 10-year/total, and the annual percentage change of the GDP implicit price deflator (2003=100, at market prices), as proxies for the nominal long-term interest rate and the inflation rate, respectively. From here, the *ex-post* real interest rate is computed as the difference between the interest rate and inflation series. We have chosen the interest rate on long-term government bonds because this is the most standard proxy for the long-term interest rate in empirical analyses of the Fisher effect. Further, using the GDP deflator is usually preferred to other alternatives, such as the Consumer Price Index, since it is not based on a fixed basket of goods and services, so allowing changes in consumption patterns or the introduction of new goods and services to be reflected automatically in the inflation rate. Finally, the choice of the sample period is dictated by the availability of data on our proxy for the long-term real interest rate, which was not available before 1966. The time evolution of the three series is shown in Figure 1.

As mentioned in the Introduction, the sample period was characterised by drastic changes in the conduct of the UK's macroeconomic policies. The year 1972 witnessed a change in the macroeconomic policy stance followed by the then Conservative government, when expansionary monetary and fiscal policies were addressed to stimulate output and employment. Accordingly, incomes policy became the main instrument to fight inflation. This policy stance was mostly followed by the Labour government elected in March 1974; see Nelson (2003). As a result, inflation was high during the 1970s, it was not fully transmitted to the long-run nominal interest rate, and the real interest rate was negative for most of the period.

However, an even greater change in the policy regime occurred after the election of a new Conservative government in May 1979. Cooperation with the social partners was replaced by the government's commitment to an anti-inflationist strategy, whose main instrument was monetary policy through the control of monetary aggregates. This was later reinforced with the announcement in March 1980 of the so called 'Medium Term Financial Strategy', in which a pre-set declining target growth rate for sterling M3 (the target monetary aggregate) was made the centrepiece of the government strategy, coupled with a decrease in public sector borrowing as a percentage of GDP (Goodhart, 1989). Inflation decreased in the following years, at the expense of a deep recession and a very high increase in unemployment. Monetary policy, however, was only

Figure 1: Nominal interest rate, inflation rate and ex-post real interest rate in the UK: 1966(1) - 2007 (1)



partially responsible for that situation, since the strategy of monetary targeting suffered serious problems of instrumentation, in the context of a huge appreciation of sterling in real terms. This led the government to move away gradually from a policy of money supply targeting, towards one of exchange rate targeting; in particular, tying the sterling to the German mark at the beginning of 1987.

The response of monetary policy to inflation remained low until the introduction of inflation targeting in October 1992, following the UK's departure from the ERM the month before (Nelson, 2003). Even though the long-run nominal interest rate appeared to respond more strongly to the inflation rate than before, transmission would still have been incomplete. In fact, real interest rates rose following the adoption of inflation targets (Siklos and Skoczylas, 2002). Finally, the granting of operational independence to the Bank of England in May 1997 seems to indicate the beginning of a new period, characterised by the lowering of inflation expectations at the time of the announcement of the new regime, suggesting a substantial improvement of the credibility of monetary policy (Allsopp, 2002). In terms of the Fisher effect, this would have resulted in a transmission of expected inflation to the long-run nominal interest rate that was not significantly different from zero.

4. Empirical results

In this section, we will provide an empirical test of the Fisher hypothesis following the approach outlined in the Section 2. First, we have tested for the order of integration of the variables, using the tests of Ng and Perron (2001). These authors propose several unit root tests that are modifications of others previously available, in order to improve their performance, i.e. their size and power, in particular in small sample sizes. In general, most of the conventional unit root tests suffer from three problems. First, they have low power when the root of the autoregressive polynomial is close to, but less than, one (De Jong *et al.*, 1992). Second, most of the tests suffer from severe size distortions when the moving-average polynomial of the first-differenced series has a large negative autoregressive root (Schwert, 1989). Third, implementing the unit root tests often implies the selection of an autoregressive truncation lag that is strongly associated with size distortions and/or the extent of power loss (Ng and Perron, 1995).

Trying to address these critiques, Ng and Perron (2001) have proposed a methodology that is robust to the three problems quoted above. This consists of a class of modified tests, namely, $\bar{M}Z_{\alpha}^{GLS}$ and $\bar{M}Z_{t}^{GLS}$, originally developed in Stock (1999) as M tests, with GLS detrending of the data as proposed in Elliott *et al.* (1996). In addition they proposed, following a similar procedure, the test ADF^{GLS} that corrects for the problems associated with the standard Augmented Dickey-Fuller test.

The results from the Ng and Perron tests are shown in Table 1. As can be seen, the existence of two unit roots is clearly rejected for all variables, whereas the null hypothesis of non-stationarity for the series in levels cannot be rejected (with the exception only of the *ex-post* real interest rate in the case of the ADF^{GLS} test). Accordingly, we can conclude that the three series would be integrated of order I(1).

Next, we estimate the long-run or cointegration relationship between the nominal interest rate and the inflation rate shown in equation (4). Given the relatively small sample size and the presence of only one cointegrating relationship, we estimate and test the coefficients of equation (4) by means of the Dynamic Ordinary Least Squares (DOLS) method of Stock and Watson (1993), following the methodology proposed by Shin (1994). This method provides a robust correction to the possible presence of endogeneity in the explanatory variables, as well as of serial correlation in the error terms of the OLS estimation. The first step would consist of estimating a long-run dynamic equation including leads and lags of the (first difference of the) explanatory variable in equation (4):

$$i_t = \alpha + \beta\pi_t + \sum_{j=-q}^{q} \gamma_j \Delta\pi_{t-j} + v_t \qquad (8)$$

and, in a second step, performing Shin's (1994) test from the calculation of Cμ, a LM statistic from the DOLS residuals that tests for deterministic cointegration (i.e., when no trend is present in the regression).

Table 1: Ng-Perron tests for unit roots

I(2) vs. I(1)	$\bar{M}Z_{\alpha}^{GLS}$	$\bar{M}Z_{t}^{GLS}$	ADF^{GLS}	I(1) vs. I(0)	$\bar{M}Z_{\alpha}^{GLS}$	$\bar{M}Z_{t}^{GLS}$	ADF^{GLS}
Δi_t	-76.8*	-6.19*	-9.99*	i_t	-3.33	-1.23	-1.25
$\Delta \pi_t$	-52.8*	-5.13*	-6.31*	π_t	-11.51	-2.39	-2.41
Δr_t	-67.2*	-5.79*	-7.35*	r_t	-16.35	-2.85	-3.01**

Notes:
(i) * and ** denote significance at the 1% and 5% levels, respectively. The critical values are taken from Ng and Perron (2001), Table 1.
(ii) The autoregressive truncation lag has been selected using the modified Akaike information criterion, as proposed by Perron and Ng (1996).

The coefficients from the DOLS regression and the results of the Shin test are reported in Table 2. Since the null of deterministic cointegration is not rejected at the 5 per cent significance level, there would be some evidence of deterministic cointegration between the nominal interest rate and the inflation rate, with a cointegration vector (1, -0.33) that is different from the theoretical values (1, -1). These results would indicate the presence of a partial Fisher effect in the long run, with a transmission to the nominal interest rate of 0.33 points of each point increase in the inflation rate. This suggests lenders would have suffered some money illusion in the sense that the nominal interest rate would have not been adjusted to compensate them fully for higher inflation.

Next, we address whether the previously estimated partial Fisher effect is stable over time, or if it exhibits some structural break, allowing the instability to occur at an unknown date. We will rely on two different approaches to test for structural stability. First, we apply the tests of Hansen (1992) in order to detect any parameter instability in the cointegration relationship. In addition, we compare the results from these tests, which only consider the possibility of *one* change, with those from the tests of *multiple* structural changes of Bai and Perron (1998, 2003a, 2003b).

Table 2: Estimation of the long-run relationship: Stock-Watson-Shin cointegration tests

α	β	Cμ
1.68	0.33	0.273
(45.7)	(17.9)	

Notes: (i) t-statistics in parentheses; (ii) The Cμ statistic is not significant at the 5% level. The critical values are taken from Shin (1994), Table 1, for *m*=1; (iii) The number of leads and lags selected was q=5≈INT($T^{1/3}$), as proposed in Stock and Watson (1993). The long-run variance of the cointegrating regression residuals has been estimated using the Bartlett window with *l*=12≈INT($T^{1/2}$), as proposed in Newey and West (1987).

Hansen (1992) proposes three tests of parameter instability based on the 'fully modified' estimator of Phillips and Hansen (1990), namely, the sup$F$, mean$F$, and L_c test statistics. All of them have the same null hypothesis (i.e., stability of the regression parameters), but differ in the alternative, since the sup$F$ test captures changes in regimes, and the mean$F$ and L_c tests capture instead gradual shifts over time.

Table 3 presents the results from Hansen's instability tests. According to the results, the relationship between the nominal interest rate and the inflation rate seems to be unstable, since the three test statistics are highly significant. As can be seen in Figure 2, the sequence of the F-statistic reaches the 5 per cent critical value of the sup$F$ test at the beginning of 1980; and its maximum values by 1992, when the Bank of England announced the new policy of inflation targeting, and by 1995, immediately before the Bank of England received operational independence.

Table 3: Hansen tests for parameter instability

Lc	mean$F$	sup$F$
1.10	53.23	139.81
(0.01)	(0.01)	(0.01)

Notes: (i) Probability of parameter instability in parentheses. (ii) According to Hansen (1992), a relation is said to be stable if the estimated probability is greater or equal than 20%.

Finally, we analyse the instability of the long-run relationship between the nominal interest rate and the inflation rate given by equation (4), by means of the tests developed by Bai and Perron (1998, 2003a, 2003b). This procedure allows testing for multiple breaks at unknown dates, by successively estimating each break point using a specific-to-general strategy in order to determine consistently the number of breaks. In particular, Bai and Perron (1998) proposed three test statistics to test for multiple breaks: (i) the sup$F_T(k)$ test, a sup$F$-type test of the null hypothesis of no structural break (m = 0) versus the alternative of a fixed (arbitrary) number of breaks (m = k); (ii) the UDmax test, a double maximum test of the null hypothesis of no structural break (m = 0) versus the alternative of an unknown number of breaks given some upper bound M ($1 \leq m \leq M$); and (iii) the supF_T (l +1 | l) test, a sequential test of the null hypothesis of l breaks versus the alternative of l+1 breaks.

The results of applying Bai and Perron's approach are shown in Table 4, where M = 3 (i.e., up to 3 breaks have been allowed). A word of caution in interpreting the results from the Bai and Perron tests is in order here since, despite the quarterly nature of the data, the sample size is not sufficiently long to draw any firm conclusions, especially for the estimates of the β's across the different sub-samples (see below). In any case, according to the results shown

Figure 2: Hansen instability tests

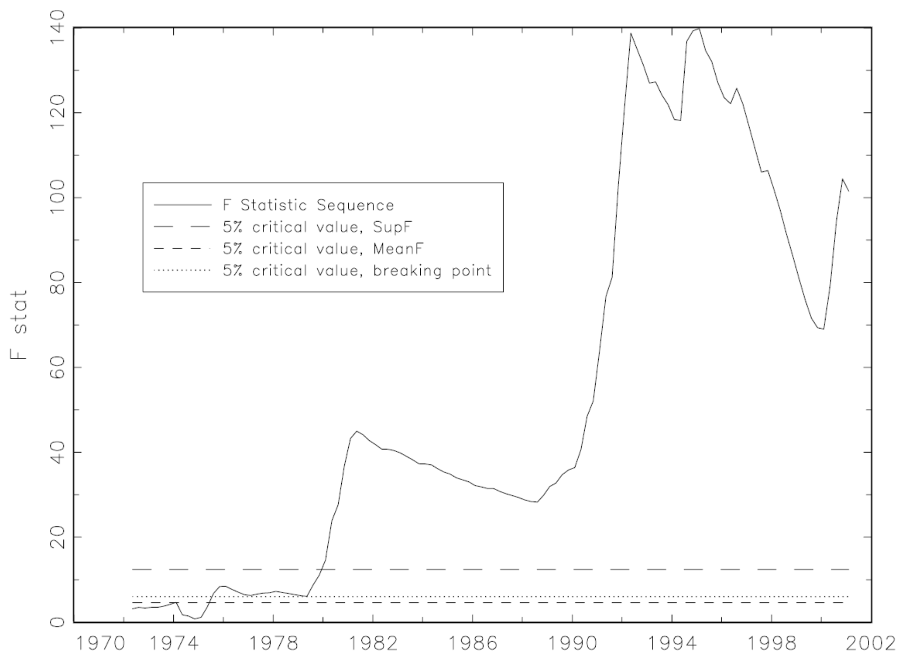


Table 4: Bai-Perron tests for multiple structural changes in the long-run relationship

Test statistics	
$supF_T(1)$	354.0
$supF_T(2)$	202.4
$supF_T(3)$	288.9
UDmax	354.0
$supF_T(2\mid 1)$	44.11
$supF_T(3\mid 2)$	77.70

Break points

$\hat{T}_1$	$\hat{T}_2$	$\hat{T}_3$
1974:4	1983:1	1997:3
[1974:1-1975:2]	[1982:3-1984:2]	[1997:1-1997:4]

Parameter estimates

$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$	$\hat{\beta}_4$
0.54	0.06	0.38	0.19
(12.6)	(2.37)	(5.97)	(0.80)

Notes:
(i) All the tests statistics are significant at the 1% level. The critical values are taken from Bai and Perron (1998), Tables I and II; and from Bai and Perron (2003b), Tables 1 and 2.
(ii) The number of breaks has been determined according to the sequential procedure of Bai and Perron (1998), at the 5% size for the sequential test $supF_T(l+1\mid l)$.
(iii) 90% confidence intervals for (j=1,2,3) in brackets.
(iv) t-statistics, robust to serial correlation, for (i=1,2,3,4) in parentheses.

in Table 4, all the sup$F_T(k)$ tests are significant, so that at least one break would be present in the relationship. In addition, according to the *UD*max test statistic there would appear at least one break in the model; and the sup$F_T(2\,|\,1)$, and sup$F_T(3\,|\,2)$ tests would be significant at the conventional levels. The sequential procedure would detect three breaks, denoted by $\hat{T}_j$ (j = 1, 2, 3) in Table 4, namely, 1974:4 (following the first oil shock), 1983:1 (following the decrease in inflation, after the peak figures recorded the previous years), and 1997:3 (following the operational independence of the Bank of England). Lastly, Table 4 also reports the estimate of the coefficient gauging the size of the Fisher effect, which would decrease after 1974:4 from 0.54 to 0.06; would rise substantially after 1983:1 to 0.38; and would decrease again after 1997:3 to a value of 0.19, not significantly different from zero.

5. Conclusions

In this paper, we have tried to provide some additional insight into the empirical testing of the Fisher effect, by analysing the role of potential structural breaks in the time series of the variables concerned. The analysis was applied to the UK case, a country that has been subject to potentially stronger regime shifts than other countries over the last half-century, for the period 1966:1-2007:1. To this end, we have made use of some recent econometric techniques aimed to detect possible structural changes or instabilities in cointegration regressions, allowing the instability to occur at an unknown date.

First, we found that the three variables under analysis, i.e., the nominal long-term interest rate, the rate of inflation, and the *ex-post* real interest rate were non stationary. Next, we found evidence of deterministic cointegration between the nominal long-term interest rate and the inflation rate, with an estimated coefficient on the latter variable equal to 0.33. There was also strong evidence of the presence of several structural changes in the cointegration relationship. Finally, up to four regimes or periods were detected in the whole sample, with estimated coefficients for the inflation rate of 0.54 for 1966:1-1974:4, 0.06 for 1975:1-1983:1, 0.38 for 1983:2-1997:3, and a non significant coefficient for 1997:4-2007:1.

Overall, our results would support the existence of a partial Fisher effect for the UK economy in the long run; in particular, for each point increase in the inflation rate, one third would have been passed through to a higher nominal interest rate, with the rest being reflected in a lower real interest rate. The ultimate reason for these results would be the presence of some degree of money illusion in the financial markets, a fact already noticed by Fisher (1930), which is not readily susceptible to elimination by market forces (Summers, 1983). As a result, the nominal interest rate would not be a good indicator of the evolution of the inflation rate, and monetary policy would be able to influence the real interest rate in the long run. Finally, the independence of the Bank of England after May 1997 would seem to open a new period, characterised by a strong reduction in inflation expectations and a substantial improvement in the credibility of monetary policy, so that the Fisher effect would seem no longer to be holding.

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ENDNOTES

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2. Notice, however, that the presence of money illusion in the context of the Fisher effect would be equivalent to the case in which the lenders (due to their lack of market power, because of strategic considerations, and the like) choose not fully to transmit to the nominal interest rate any expected change in the inflation rate, even if they anticipate changes in inflation correctly. Accordingly, when we refer in the paper to 'money illusion', it will be in this broad sense.

3. Notice that the error ε_t would be stationary even relaxing the assumption of rational expectations, although in such a case it would not be necessarily white noise.

4. Note that if the nominal interest rate were I(0), and the other two variables I(1) and cointegrated, then the estimate of β' in (7) would not be significantly different from -1, there would be complete money illusion, and changes in the expected inflation rate would be transmitted one-for-one, with the opposite sign, to the *ex-ante* real interest rate. In turn, if the inflation rate were I(0), and the other two variables I(1) and cointegrated, there would be a one-for-one relationship between the nominal interest rate and the *ex-ante* real interest rate.

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