U.S. and U.K. Interest Rates 1890 - 1934: New Evidence on Structural Breaks

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Abstract

This paper presents econometric evidence on whether the founding of the Federal Reserve in 1914 caused a structural change from level-stationarity to difference-stationarity in U.S. and U.K. short-term nominal interest rates. We develop new econometric tests that allow for parameter transitions to test for a break of this kind and undertake a grid search analysis of dates and speeds for the change. We find that U.S. nominal interest rates most likely evolved rapidly to difference-stationarity in June 1917. For the U.K. we fail to reject the null that U.K. interest rate series follow a difference stationary process over the entire period 1890-1934. Our analysis differs from previous research on this topic in that we take care to explore statistical uncertainty around parameter estimates, and incorporate higher order dynamics into our econometric analysis.

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U.S. AND U.K. INTEREST RATES 1890 - 1934: NEW EVIDENCE ON STRUCTURAL BREAKS

The period 1890-1933 was a tumultuous time in financial markets in both the U.S. and the U.K.. The U.S., the U.K., and other European countries suffered through World War I, underwent changes in monetary institutions (with the founding of the Federal Reserve System in the U.S.), and changes in monetary policy regimes and objectives (e.g. the suspension of the gold standard). Monetary economists have empirically investigated the effect of such changes on the data generating process describing nominal interest rate movements. A body of empirical literature has investigated the changing stochastic behavior of short-term nominal interest rates during the period 1890-1933 in the U.S. as well as in the U.K.. This literature has concluded that during the period 1890-1910 short-term nominal interest rates in both the U.S. and the U.K. followed stationary, I(0), processes, whereas for the period 1920-1933, these interest rates had become nonstationary, I(1), processes. The exact time (or date) at which the structural change from stationary to nonstationary behavior took place, and the causes of the structural change in the data generating process describing nominal interest rates, has been an area of controversy. The most notable statement has been the argument, made by Mankiw, Miron, and Weil (1987) that the founding of the Federal Reserve System in the U.S. in 1914 was a catalyst for the structural change in U.S. interest rate behavior. Expanding on this, Barsky, Miron, and Weil (1988) have argued that the Federal Reserve also caused a structural break in the time series behavior of interest rates in the UK. Subsequent empirical research (focusing primarily on the U.S.), questioning previous works on various grounds, have argued against the 1914 break point for U.S. interest rates, finding other break point dates. While little research has been done re-examining the stochastic behavior of U.K. interest rates, there is reason to suspect a structural break in U.K. short-term nominal interest rates. Consider, for example, the evidence that Barsky, Mankiw, Miron and Weil (1988) present on the autocorrelations of each series (their Table I). For the U.S. monthly series, the autocorrelations for the first sub-period (1890-1910) damp quickly relative to those for the second sub-period (1920-1933). For the U.K. series there are significantly stronger and more persistent autocorrelations in the first sub-period. Indeed, Barsky, Mankiw, Miron, and Weil

(1988) point out that “In the first sample the autocorrelations indicate that the short rate is more persistent in Britain than it is in the United States”[1] but they fail to investigate the issue any further. In this paper we provide a more detailed investigation of the above issues.

The purpose of this paper is to present new econometric evidence on the hypothesis of a structural break in the stochastic processes generating both U.S. and U.K. short-term nominal interest rates between 1890:1 and 1934:1. Throughout our analysis we utilize the logistic function to model the structural break as a transition from an I(0) to an I(1) process. This permits scope in assessing the speed as well as the timing of any transition. Section 1 presents a brief review of the literature in this area. Our statistical analysis follows and has three main sections. In Section 2 we develop new procedures for testing for a structural break from I(0) to I(1) and apply them to U.S. and U.K. short-term nominal interest rates. In Section 3 we concentrate on dating the structural break in the U.S. series. In Section 4 we use grid search techniques to illustrate a fundamental difference between the U.S. and U.K. series regarding what types of transitions (in terms of specific dates and speeds) cannot be rejected by likelihood ratio tests. Section 5 concludes the paper.

Our results indicate that, with a fully specified dynamic model, a rapid structural break from I(0) to I(1) most likely occurred in U.S. nominal interest rates in June 1917. In contrast, for the U.K. we fail to find strong evidence supporting any particular type of structural change, or of any change at all. In fact, we find no evidence against the proposition that the U.K. interest rate series were difference stationary over the entire period 1890-1934. One possible explanation for this is the interest rate smoothing behavior of the Bank of England during this period documented by Goodfriend (1988). Our results indicate that we cannot support the Mankiw, Miron, and Weil (1987) argument that the founding of the Federal Reserve System alone represented a new regime, or the proposition of Barsky, Mankiw, Miron, and Weil (1988) who argue that the Fed in some way caused a structural break in U.K. interest rates.

1. Literature Review

It has been proposed that monthly U.S. short-term nominal interest rates underwent a structural break from being level-stationary, I(0), with a dominant first-order autoregressive root of about .75, to approximately difference-stationary, I(1), sometime between the end of

1914 and the middle of 1915; see for example Mankiw, Miron, and Weil (1987), and Barsky, Mankiw, Miron, and Weil (1988). Mankiw, Miron and Weil (1987) argue that the change in the stochastic behavior of interest rates was a result of the founding of the Federal Reserve System in 1914 and its implementation of interest rate smoothing policies.[2] While agreeing that a change in the stochastic behavior of U.S. interest rates took place at some time between 1910 and 1920, several authors have questioned the date; see for example Fishe and Wohar (1990), Fishe (1991), Angelini (1994), and Kool (1995). Fishe and Wohar (1990) find a structural break in either early 1915 (supporting the Mankiw, Miron and Weil (1987) results) or in 1912, depending on whether they examine three-month or six-month interest rates. Fishe (1991) employed weekly data for the U.S., allowing for multiple structural breaks over the sample 1890 to 1933. Break points were reported in January 1908 (a period just following the October 1907 financial crisis), in June 1917 (a date associated with Federal Reserve Amendments that significantly increased the Fed's operating capability), and in January 1930. Kool (1995) employs a recursive method based on Bayesian learning and argues that the results of previous switching regression techniques attempting to date the structural change in interest rates are not robust. His estimation method yields a switch to nonstationarity in late 1917. Angelini re-examines the work of Mankiw, Miron and Weil (1987) and Fishe and Wohar (1990) and concludes that their results are not robust to sample periods and that there is no evidence of structural change in 1914.

While the majority of authors believe that when the structural change in U.S. interest rates did occur, it occurred quite rapidly (for example Mankiw, Miron and Weil (1987) found it highly probable that the structural break in U.S. interest rates took less than a year to be fully complete), other arguments have been proposed regarding the speed of the adjustment to a new regime following the founding of the Fed. Willis (1923) and Wicker (1966) note that the Fed was not very active in its early years, with most of its efforts during the years 1914-1916 being focused on internal organization. After this initial period however the Fed was able to concentrate more resources on their interest rate policy that consequently affected the stochastic behaviour of interest rates. Riefler (1930) and Kool (1995) point out that the Fed did not provide large amounts of liquidity to the economy until after 1917 when war financing became an important concern.

In addition to questions over the timing and speed of the structural change, various alternative arguments have also been presented to explain the reason for the structural change. For example while Fishe and Wohar (1990) support Mankiw, Miron and Weil (1987) and Barsky, Mankiw, Miron and Weil (1988) with respect to the behavior of the U.S. three-month interest rate[3], they suggest that it was not the founding of the Fed that changed the behavior of interest rates, but instead the reopening of the U.S. bond and stock markets in November and December of 1914. Moreover, Angelini (1994) notes that during World War I the New York money market was strongly subjected to the regulation and control of the Money Market Committee. Further, throughout this period major reforms were passed, that according to some writers of the time greatly affected the functioning of the money market and may have resulted in a permanent change in how this market operated and functioned. Clark (1986) puts forward the proposition that the changing behavior of interest rates and inflation was a worldwide phenomenon, resulting from the suspension of the gold standard, beginning in 1914 and ending by September 1917. The World War I suspension of the gold standard was not abrupt, but piecemeal. While the war forced most of the European nations off the gold standard, in the U.S. it was nominally maintained and it was not until September 1917 that the government began to constrain gold exports in an effort to restrict gold outflows. From May 1919 through March 1920 inflationary pressures led to the resumption of gold outflows (see Wicker (1966)). Then as gold reserves approached minimum requirements, the Fed increased the discount rate in January 1920 to 6 percent. Gold inflows followed and in June 1920 the full gold standard was resumed, until 1933 when the U.S. went off the gold standard.

Although it appears quite plausible that the establishment of a new institution such as the Federal Reserve at the end of 1914 could have had a strong and immediate effect on market conditions in the U.S., it seems very unlikely that the introduction of the Fed would have been the catalyst for structural change around the world. Clark (1986) proposes that the structural change in interest rate behavior also took place in European countries where central banks had already been in operation for many years. For example, the Bank of England was established in 1694. Barsky, Mankiw, Miron and Weil (1988) also find a change in the behavior of U.K. interest rates around the same time as their postulated change in the behavior of U.S. interest rates, and while they acknowledge that the U.S. economy was not sufficiently dominant to have altered worldwide interest rates, they argue against Clark (1986) that it was the dissolution of the gold standard that caused this change. They present a theoretical model to support their argument that the founding of the Fed was the ultimate cause of the worldwide change in interest rate behaviour because it "marked the beginning of a new era in which all major countries had a central bank"[4].

Barro (1989) and Kool (1995) suggest a more attractive explanation for the structural change in U.S. and U.K. interest rates, as reflecting the beginning of interest rate targeting. Barro (1989) attributes the regime change in the U.S. to the fact that the central bank had an objective of smoothing interest rates around a random walk target to stabilize the economy at some point following the founding of the Fed. It has been argued by Kool (1995) that interest rate targeting in the U.S. and the U.K. began at different times, albeit for the same reason, namely the financing of military spending through government borrowing at low interest rates. Kool (1995) finds structural breaks in 1915 for the U.K. and 1917 for the U.S..

2. Testing for a transition from I(0) to I(1) in U.S. and U.K. interest rates

Many of the above cited studies suggest reasonable arguments for a structural break occurring in U.S. interest rates sometime during 1917, rather than 1914. Our subsequent empirical analysis will examine this issue. The data for our empirical analysis are taken from the set used by Barsky, Mankiw, Miron and Weil (1988) and Mankiw, Miron and Weil (1987). The U.S. data series consists of 529 monthly observations on the three month time loan rate available at New York City banks, taken from the National Monetary Commission *Financial Review*, updated using the *Commercial and Financial Chronicle*. The U.K. series consists of 529 monthly observations on the three month rate on bankers' bills available in London and taken from the *Economist*. This is the open market rate of bankers' bills, not to be confused with the Bank of England's discount rate, known as the bank rate. Fishe and Wohar (1990), among others, question the reliability of this data series and note a number of problems with the U.S. monthly data set. We therefore follow their strategy of analysing primarily weekly data for the U.S.; a series of 1305 observations beginning in 1909. For completeness, we also report findings for monthly data as well.[5]

Consider for a time series y_t of interest rates modeled as

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

where ε_t are independent, identically distributed (IID) deviates, and $S_t(\gamma,\tau)$ is the logistic function based on a sample of size T,

$$S_t(\gamma,\tau) = [1+\exp\{-\gamma(t-\tau T)\}]^{-1}$$

that monotonically traverses the interval (0,1). Equation (1) represents a model that allows a smooth transition in y_t from one first-order autoregression, as $t \to -\infty$,

$$y_t = \alpha_1 + \beta_1 y_{t-1} + \varepsilon_t$$

to another, as $t \to \infty$,

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

The interpretation of the parameters of $S_t(\gamma,\tau)$ is as follows. The parameter τ determines the timing of the transition midpoint fraction because, for $\gamma > 0$, we have $S_{-\infty}(\gamma,\tau) = 0$, $S_{+\infty}(\gamma,\tau) = 1$ and $S_{\tau T}(\gamma,\tau) = 0.5$. The speed of the transition is then determined by the parameter γ. If γ is small then $S_t(\gamma,\tau)$ takes a long period of time to traverse the interval (0,1), and in the limiting case with $\gamma = 0$, $S_t(\gamma,\tau) = 0.5$ for all t. On the other hand, for large values of γ, $S_t(\gamma,\tau)$ traverses the interval (0,1) very rapidly, and as γ approaches $+\infty$ this function changes value from 0 to 1 instantaneously at time $t = \tau T$. Thus the model allows for no transition, instantaneous transition, and all smooth intermediate cases. Models of this type have been discussed by Granger and Terasvirta (1993) and Lin and Terasivirta (1994), though the test for transition from I(0) to I(1), developed in this section, is new.

In the first-order autoregressive framework of (1), if the final state is difference-stationarity, about which there seems little dispute in the literature, then $\beta_1 + \beta_2 = 1$. This constraint will be imposed in our subsequent analysis. Level-stationarity in the initial state implies $|\beta_1| < 1$. Assuming $\beta_1 + \beta_2 = 1$, then (1) can be re-written as

$$(y_t - S_t(\gamma,\tau) y_{t-1}) = \alpha_1 + \alpha_2 S_t(\gamma,\tau) + \beta_1 (y_{t-1} - S_t(\gamma,\tau) y_{t-1}) + \varepsilon_t \, . \qquad (2)$$

To test the null hypothesis that y_t is I(1) throughout against a transition in y_t from I(0) to I(1) with drift, the relevant hypotheses are,

$$H_0 : \beta_1 = 1, \quad H_1 : \beta_1 < 1.$$

As there is little reason to suspect long term drift in the later, difference-stationary, period, we also consider the same test, but constraining the drift to be zero; that is for a transition from a stationary first-order autoregression with non-zero mean to a random walk with no drift. In this case $\boldsymbol{a}_1 + \boldsymbol{a}_2 = 0$ so that model (2) becomes

$$(y_t - S_t(\boldsymbol{g}, \boldsymbol{t})y_{t-1}) = \boldsymbol{a}_1(1 - S_t(\boldsymbol{g}, \boldsymbol{t})) + \boldsymbol{b}_1(y_{t-1} - S_t(\boldsymbol{g}, \boldsymbol{t})y_{t-1}) + \boldsymbol{e}_t \,. \tag{3}$$

The null and alternative hypotheses on $\boldsymbol{b}_1$ are as before.

The appropriate test of H_0 versus H_1 in both models is based on the t - statistic for testing $\boldsymbol{b}_1 = 1$; that is,

$$t = \frac{\hat{\boldsymbol{b}}_1 - 1}{\hat{s}e(\hat{\boldsymbol{b}}_1)}$$

where $\hat{\boldsymbol{b}}_1$ is the (nonlinear) least squares estimator of $\boldsymbol{b}_1$ in (2) or (3). To obtain critical values for these test statistics, we simulated under the null hypothesis from the random walk model

$$y_t = y_{t-1} + \boldsymbol{e}_t$$

with $\boldsymbol{e}_t$ generated as IID N(0,1). Models (2) and (3) were fitted to this process by nonlinear least squares (NLS), care being taken to ensure that global rather than local minima were found through the use of a grid of starting values for $\hat{\boldsymbol{t}}$. Because models (2) and (3) are linear in the $\boldsymbol{a}$ and $\boldsymbol{b}$ parameters, when estimating, following Leybourne, Newbold, Vougas (1998), we are able to speed up the convergence of the optimization algorithm by concentrating the sum of squares function with respect to these parameters. Thus the NLS estimation problem reduces to minimizing the sum of squares function with respect to the two parameters $\hat{g}$ and $\hat{\boldsymbol{t}}$. Given the greater volatility in the early part of these interest rate series compared to the post-break years, particularly for the U.S., we also simulated critical values assuming heteroskedasticity under the null hypothesis. Specifically these critical values were simulated under the null hypothesis from random walk series of 1305 observations, with variance of $\boldsymbol{e}_t$ in the first 34% of the sample three times that in the final 66% of the sample. (Our analysis suggests that for the U.S. weekly series the most likely abrupt break is 34% of the way through.) Of course these critical values are only directly relevant for the model estimated for this data, hence in Table 1 we present both sets of critical values at the 0.10, 0.05 and 0.01 significance levels for both the tests (unconstrained drift and drift constrained to zero), estimated from 5000 replications.[6]

When calculating the test statistics for each interest rate series, rather than working with the raw data, we corrected for any potential dynamic misspecification in the model (2) or (3) by application of an autoregressive filter to the data prior to the modeling. The filter we used involved first estimating by ordinary least squares the autoregression

$$\Delta y_t = \hat{\alpha} + \sum_{i=1}^{k} \hat{\phi}_i \Delta y_{t-i} + \hat{\varepsilon}_t$$

We then defined $y_t^* = \sum_{h=1}^{t} \hat{\varepsilon}_h$ as our series to be modeled and estimated the models (2) and (3) using the filtered data y_t^* in place of y_t.[7] Treating the autocorrelation dynamics in this way ensures that the asymptotic null distribution of the test statistic is unchanged. Moreover, since nuisance parameters are easily eliminated through ordinary least squares estimation of the autoregression in first differences, only modest additional computation is involved in simulations, through which we were able to verify that the critical values of Table 1 remain appropriate for sample sizes of interest. Table 2 provides the calculated test statistics and estimated parameters for each transition model.

Comparing Table 2 with Table 1, for monthly U.S. short term nominal interest rates over the period 1890:1 to 1934:1, using the critical values simulated under the assumption of heteroskedastic error terms, we can reject the null of no transition from I(0) to I(1) for all of our models at the 5% level of significance. This is so irrespective of whether or not we incorporate dummy variables in the model for points of data irregularity.[8] Even stronger rejections are obtained, at the 1% significance level, when U.S. weekly data beginning in 1909 are used. Note however that irrespective of which critical values are used, for U.K. monthly short-term nominal interest rates over the period 1890:1 to 1934:1, unequivocally we cannot reject the I(1) null hypothesis. Thus, we find, along with previous authors, very strong evidence of transition from level-stationarity to difference-stationarity for the U.S., however evidence for such a transition in the U.K. interest rate series is not significant, even at the 10% level. That is, our analysis fails to reveal strong evidence against the proposition that U.K. interest rates were difference stationary over the entire period.[9]

3. Estimating the structural change in U.S. short term nominal interest rates

In section 2 our tests fail to find strong support for the hypothesis that a structural change from I(0) to I(1) occurred at all in U.K. interest rates over the period 1890:1 - 1934:1. In this section of our empirical analysis we concentrate on the U.S. series. Here and in the following section, we report results on U.S. weekly data, beginning in 1909. In fact, very similar results were found for U.S. monthly data over the period 1890:1 - 1934:1. A detailed analysis of this monthly series is given in Sollis (1999).

To determine the timing and speed at which the structural change in U.S. short term nominal interest rates occurred, we used the approach outlined in section 2, estimating models with and without any constraints on the drift $\boldsymbol{a}_1 + \boldsymbol{a}_2$ in (1) and (2). The model specifications are,

$$(y_t - S_t(\boldsymbol{g},\boldsymbol{t})y_{t-1}) = \boldsymbol{a}_1 + \boldsymbol{a}_2 S_t(\boldsymbol{g},\boldsymbol{t}) + \boldsymbol{b}_1(y_{t-1} - S_t(\boldsymbol{g},\boldsymbol{t})y_{t-1}) + \sum_{i=1}^{k} \boldsymbol{q}_i \Delta y_{t-i} + \boldsymbol{e}_t \tag{4}$$

for unconstrained drift, and for drift constrained to be 0,

$$(y_t - S_t(\boldsymbol{g},\boldsymbol{t})y_{t-1}) = \boldsymbol{a}_1(1 - S_t(\boldsymbol{g},\boldsymbol{t})) + \boldsymbol{b}_1(y_{t-1} - S_t(\boldsymbol{g},\boldsymbol{t})y_{t-1}) + \sum_{i=1}^{k} \boldsymbol{q}_i \Delta y_{t-i} + \boldsymbol{e}_t \tag{5}$$

with, as in the previous section, $k = 5$ in both models. In each case, then, the constraint $\boldsymbol{b}_1 + \boldsymbol{b}_2 = 1$ is imposed on (1), in line with the evidence of the previous section of a transition from I(0) to I(1). Equations (4) and (5) provide a slightly different approach here, where the emphasis is on estimation rather than testing (as was the case in section 2), to the incorporation of dynamics into (2) and (3). For the present purposes, it is convenient to express the model to be estimated as a single equation, using the raw, rather than, the transformed data. The results obtained in this section are, as we shall see, entirely compatible with those of the previous section.

Figure 1 shows the estimated transition, from a nonlinear least squares fit for the constrained drift model. (Virtually identical results were obtained for the unconstrained model (4)). The estimated transition is virtually instantaneous, occurring in June 1917. As reported in Sollis (1999) a similar analysis of monthly data from 1890:1 - 1934:1 also generates a best estimate of an almost instantaneous transition in June 1917. The only apparent difference of substance between results for the two data sets is that while Sollis reports an estimate of 0.74 for the dominant autoregressive parameter in the pre-transition period, as can be seen from Figure 1 the corresponding value for the higher frequency data is 0.93. Because $0.93^4 \simeq 0.75$, these estimates are certainly not incompatible. While our results for weekly data and for monthly data covering a longer time period are in agreement, they differ from those of Mankiw, Miron and Weil (1987), who also used a logistic switching model, estimating a transition, using monthly data, that is essentially completed between December 1914 and June 1915. We can achieve the same results for those data by dropping the dynamics - that is the terms in lagged first differences in (4). It is the inclusion of these terms, then, that accounts for the differences in our point estimates and those of Mankiw, Miron and Weil. However, it must be stressed that the analysis of this section does not statistically exclude an earlier transition as a possibility.

The estimated transitions reported in this section are nonlinear least squares point estimates. In Figure 1 we report as our best estimate a very rapid transition in mid-1917. However, as yet we have not attached uncertainty measures to this estimate, which does not for example exclude on statistical grounds the possibility of an earlier or less rapid transition. This issue is explored in the following section.

4. A grid search analysis of potential parameter transitions.

In our final empirical section we make a more detailed comparison of the U.S. and U.K. interest rate series, from the perspective of estimating parameter transitions to capture potential structural changes from I(0) to I(1). Specifically, we fit transition models to both series and determine the set of parameter values (g, t) - that is, speed of transition and transition midpoint - that cannot be excluded by likelihood ratio tests. Here we report results for the U.S. weekly series beginning in 1909.[10] In section 2 we noted the lack of strong evidence of any such transition for the U.K. series. Nevertheless, failure to reject a null hypothesis (of no transition) does not necessarily constitute strong evidence in favour of that hypothesis, so a transition model is estimated for the U.K. data. In section 3 we noted that our best estimate for the U.S. series is a very rapid transition in June 1917, but it is important to assess what other possibilities cannot be excluded on statistical grounds.

Working with our most general model, given by equation (2), we augment to explicitly take account of higher-order autocorrelations giving model (4), which is then estimated by nonlinear least squares. As before, in that model we set $k = 5$ for the U.S. weekly series and $k = 24$ for the U.K. monthly series. Then for each series we calculate the sum of squared errors (SSE) over a grid allowing g to range from .005 to 5 in steps of .005 and t from .1 to .9 in steps of .01, minimizing SSE, or equivalently maximizing the Gaussian likelihood, for the remaining parameters in the model at each point on the grid. Note that increasing the range of g beyond 5 leads to no information gain because at this stage the transition is already virtually instantaneous.

Defining the natural log of the likelihood function at the global maximum over (g, t) as $L_1(g_1, t_1)$ and the log of the maximized likelihood function at any other point (g_0, t_0) as $L_0(g_0, t_0)$, then at each point in our grid search we obtain a different L_0 for each different g, t combination. Assuming that $l = 2(L_1 - L_0)$ has a chi-squared distribution with 2 degrees of freedom under $(g, t) = (g_0, t_0)$, the set of all g, t combinations that cannot be rejected at the 5% significance level can be obtained by locating the g, t values such that $l < 5.99$. These sets are graphed as Figure 2, with the dashed and solid lines indicating the computed sets for the U.K. and U.S. series respectively. Rather than graph the transition speed g (that has little intuitive interpretation) against t, we convert g into the approximate time that it takes for 90% of the transition process to be completed (45% either side of the midpoint t), using for monthly (weekly) data the approximate conversion formula 90% transition in months (weeks) $\simeq (6/g)$. This approximation appears to work well over the range of values considered here.[11]

From Figure 2, for the U.S. short term nominal interest rate, the acceptable transition midpoints t range from October 1914 to January 1918, while in terms of the transition speed, we can accept speeds from instantaneous to 7.7 years for 90% of the transition to be completed. Yet it can be seen on the same graph that for the U.K. we can accept midpoints earlier than 1909, and either very fast transitions or extremely slow transitions with 90% of the transition taking 100 years. This huge range of acceptable parameter combinations for the

U.K. is compatible with the analysis of section 2, that suggested only weak evidence for any transition in this series.

As an aid to interpreting Figure 2, in Figures 3 and 4 we plot the estimated beta transition ($\hat{b}_1 < 1$ to 1) that maximizes the likelihood function, and for comparison the transition of $\hat{b}_1 < 1$ to 1 that is consistent with a g at the highest point of the acceptance set (a slower transition), but with the same t value as for the transition that maximizes the likelihood function. From Figure 3, it can be seen that the transition that generates the maximum value of the likelihood function for the U.S. series is virtually instantaneous, starting from June 1917. Alternatively the bulk of the transition consistent with a g from the edge of the acceptable g, t set but with the same midpoint (the slowest transition we can accept), takes place between 1914 and 1921.

As shown in Figure 4, for the U.K., the transition that maximizes the value of the likelihood function again occurs quickly, in this case in June 1915. The slowest transition we cannot reject (with the same midpoint) is simply an upward sloping line that in fact starts from .88 in 1890, reaches .91 by 1909 and .96 by the final observation. It can be seen that for the U.K. series we cannot reject an extremely slow transition that is not fully completed within the sample.

For the U.S. series, in terms of the g, t combinations that cannot be rejected at the 5% level of significance, we have a set with midpoints that are tightly clustered between October 1914 and January 1918. For the U.K. however the set of the g, t combinations that cannot be rejected is much larger, with a wide range of acceptable midpoints and acceptable speeds ranging from transitions that are completed in 1 month, to transitions that are still not completed after 100 years!

Clearly in terms of estimating structural change in the form of parameter transitions, the U.S. and U.K. short term interest rate series are fundamentally different. While at the 5% level of significance for the U.S. we cannot reject parameter transitions with midpoints between 1914:10 and 1918:1, this period being a good candidate for a ‘transitional period’ in terms of the historical evidence discussed in the introduction, transitions that are extremely slow can be rejected. For the U.K. at the 5% level of significance there is a much wider range of midpoints that cannot be rejected and in terms of transitional speed we cannot reject the slowest feasible transition, or the fastest. Thus for the U.K. series if a transition is estimated in spite of the test outcomes of section 2, then our statistically acceptable results tell us almost any type of break might have occurred - we cannot find strong support for any specific type of break, and the hypothesis of no break is in this sense further supported.

5. Conclusion

In this paper we have presented evidence on three important empirical issues regarding structural changes from I(0) to I(1) in U.S. and U.K. short term nominal interest rates over the period 1890:1 to 1934:1. Those issues are; whether structural breaks in these series do occur, if structural breaks do occur then when do they most likely take place and how quickly are they completed, and finally what types of breaks cannot be statistically rejected for each series. Given questions about data reliability in the early years, we report an analysis of weekly U.S. data beginning in 1909. The fact that similar conclusions emerge from analyzes of the U.S. monthly series and a higher frequency series covering a shorter time span strengthens the credibility of our results. Developing a new testing procedure for a transition from an I(0) to an I(1) process, we find that we can unequivocally reject the null hypothesis that the U.S. interest rate series is I(1) throughout the sample period in favour of the hypothesis that a structural change from I(0) to I(1) occurred. For the U.K. series we cannot reject the I(1) null hypothesis, even at the 10% level of significance. Our results indicate that our best estimate is that a rapid structural change from I(0) to I(1) occurred in U.S. nominal interest rates in mid-1917.

While we find a different date for the most likely structural break in U.S. interest rates to Mankiw, Miron, and Weil (1987) and therefore find little support for their conclusion that it was the founding of the Federal Reserve System in 1914 that caused the transition from I(0) to I(1), historical evidence suggests that the Fed did play a part in the structural change. In June 1917 financial amendments were passed by President Wilson which not only helped to finance the participation of the U.S. in World War I, but also greatly increased the flexibility of the Federal Reserve. The Amendments allowed the Fed, through the district banks, to operate a more flexible monetary policy. This monetary policy was much more likely to have had the effect of causing a structural change in the stochastic behavior of short-term nominal interest rates of the type observed than anything that happened prior to the Amendments of 1917.[12]

Our results indicate that for the U.K. we fail to find strong evidence supporting any particular type of structural change, or indeed of any change at all. In fact our results indicate that we cannot reject the proposition that UK interest rates are difference stationary over the entire period 1890-1934. There appears to be good economic justification for this finding. While the U.K. was still on the gold standard in the early 1900s, Goodfriend (1988) has shown that a central bank can smooth interest rates even under a gold standard. Hence, the Bank of England could well have been smoothing interest rates prior to the founding of the Federal Reserve in 1914, resulting in the difference stationary behavior of the U.K. interest rate series.

In conclusion, we cannot support Mankiw, Miron and Weil (1987) in their argument that the founding of the Fed alone represented a new economic regime, or the proposition of Barsky, Mankiw, Miron and Weil (1988) that in some way the Fed caused a structural break in U.K. interest rates. There is strong historical evidence in the form of the 1917 Financial Amendments to support our econometric evidence that a structural break in U.S. short term nominal interest rates most likely took place in June 1917. At this date the power of the Federal Reserve System was increased sufficiently to allow them to operate a successful interest rate smoothing policy. The new regime might reasonably be dated from this point in the U.S., though, as might reasonably be expected, no similar evidence for such a rapid impact on U.K. interest rates is apparent.

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Table 1. Simulated critical values for structural change tests

Significance	*T*	$t_{(a_1+a_2)=0}$	*t(h)*	$t(h)_{(a_1+a_2=0)}$
1%	-4.403	-4.145	-5.194	-4.907
5%	-3.777	-3.611	-4.622	-4.376
10%	-3.748	-3.361	-4.568	-4.128

Notes: *t* denotes simulated large sample critical values of the test statistics assuming under the null hypothesis that error terms are IID standard normal. *t(h)* denotes simulated large sample critical values of the test statistics, assuming under the null hypothesis that error terms are normally distributed with variance in the first 34% of the sample three times that in the final 66% of the sample.

Table 2. Estimated parameters and calculated *t* test for structural change

Model	$\hat{a}_1$	$\hat{a}_2$	$\hat{b}_1$	$\hat{g}$	$\hat{t}$	t
U.S., 1890:1 – 1934:1, (m)	-.371 (.153)	.342 (.144)	.895 (.022)	1.60 (.353)	.597 (.054)	**-4.859****
U.S., 1890:1–1934:1, (m) $a_1 + a_2 = 0$	-.372 (.155)	.372	.893 (.022)	14.68 (.384)	.570 (.048)	**-4.945****
U.S., 1890:1 – 1934:1, (m) Dummies	-.370 (.167)	.344 (.166)	.893 (.023)	1.60 (.388)	.597 (.050)	**-4.698****
U.S., 1890:1 – 1934:1, (m) Dummies, $a_1 + a_2 = 0$	-.370 (.144)	.370	.892 (.023)	14.68 (.358)	.570 (.048)	**-4.696****
U.S., 1909:1 – 1934:1, (w)	.073 (.068)	-.076 (.049)	.931 (.013)	51.97 (.393)	.339 (.038)	**-5.410***
U.S. 1909:1 – 1934:1, (w) $a_1 + a_2 = 0$	.073 (.067)	-.073	.931 (.013)	27.61 (.359)	.339 (.038)	**-5.409***
U.K. 1890:1 – 1934:1, (m)	-.154 (.078)	.133 (.076)	.949 (.016)	19.26 (12.63)	.578 (.477)	-3.272
U.K. 1890:1 – 1934:1, (m) $a_1 + a_2 = 0$	-.154 (.078)	.154	.949 (.016)	30.64 (19.94).	.578 (.477)	-3.273

Notes: (m) and (w) denotes monthly and weekly observations respectively. 'Dummies' refers to the inclusion of dummy variables for the months 1907:11, 1907:12 and 1908:1.

**, * denote significance at the 5% and 1% levels respectively. Standard errors are in parentheses.

ENDNOTES

[1] Barsky, Mankiw, Miron and Weil (1988), page 1130.

[2] The Federal Reserve Act was passed on December 23, 1913. The presidents of the banks met for the first time in July 1914 and the banks opened for business on November 16, 1914.

[3] Fishe and Wohar (1990) also find a break in 1912 when employing a six-month interest rate, casting doubt on the MMW hypothesis that the Fed was the primary contributing factor causing the stochastic behavior of interest rates to change.

[4] Barsky, Mankiw, Miron and Weil (1988), page 1125.

[5] For more details on these interest rate series see, Mankiw, Miron and Weil (1987), Barsky, Mankiw, Miron and Weil (1988), and Fishe and Wohar (1990).

[6] For the NLS estimation here and throughout our analysis we employed the OPTMUM subroutine library of GAUSS 3.1. The critical values are virtually identical for series of 529 monthly and 1305 weekly observations. We also simulated innovations from a heavy-tailed distribution, Student's t with 5 degrees of freedom. This had a negligible impact on the critical values. Subsequently we allowed for dynamics by introducing lags. We found that this elaboration of the test has only a modest impact on the critical values.

[7] For the U.S. weekly, U.S. monthly and the U.K. monthly series, using the general to specific testing methodology at the 10% level of significance we identified $k = 5$, $k = 20$ and $k = 24$ respectively. For the U.S. monthly series we also considered the possibility of elaborating our model by incorporating 0-1 dummy variables for the months 1907:11, 1907:12, 1908:1, which were severely affected by the financial crash of 1907. Like Fishe and Wohar (1990) and Fishe (1991) we find that seasonal dummy variables taking the same value across the sample have no impact on the location or speed of the structural breaks found using our models. For this reason we present results excluding seasonal dummies. We also experimented with seasonal dummy variables that took different values in the early and late periods, in the manner of Mankiw, Miron and Weil (1987). However this methodology involves an arbitrary decision as to what sub-samples should be used to calculate the dummy variables, and none of these types of models offer an increase in explanatory power over models excluding seasonal dummy variables but including additional statistically significant dynamics.

[8] For calculation of the test statistic with U.S. monthly data, the relevant transition models were estimated with dummy variables included for observations 1907:11, 1907:12, 1908:1, these observations being severely affected by the financial crash of 1907. It is widely acknowledged that some of the U.S. monthly data is subject to being contaminated with measurement error. For certain months in the years 1903, 1907, 1910 and 1918, usury laws imposed a ceiling on the reported interest rates meaning that the rates reported in these years were not the true market clearing interest rates. While we report results including only dummy variables for observations 1907:11, 1907:12, and 1908:1, incorporating dummy variables into our models

for observations with measurement error did not alter the timing of the estimated transition, or the general conclusions of statistical support for a break of this type in the U.S. series or the finding of no such break in the U.K. series. This is consistent with Mankiw, Miron, and Weil (1990), who, in their response to Fishe and Wohar (1990), include dummy variables to account for suspect dates, and find little change in their break point. Mankiw, Miron, and Weil (1990), write, "There is no evidence that any of the results we reported are attributable to measurement error." (p. 978) See Fishe and Wohar (1990) for more details on the issue of measurement error in these series.

[9]An additional check on the validity of our findings for the U.K. series was carried out by estimating our models with the exponent in the logistic function reversed to $(tT - t)$ and with g replaced by g^2. In this case our restriction $b_1 + b_2 = 1$ forces the model to be I(1) at the start of the series and our standard test becomes one of the null hypothesis that the model remains at I(1) against an alternative hypothesis that at some point there is a transition to I(0). This variant revealed no additional statistical evidence against the I(1) null hypothesis for the U.K. series.

[10] However, Sollis (1999) finds comparable conclusions for the U.S. monthly series over the whole period 1890-1934. In particular, the sharp distinction in inference about the U.S. and U.K. interest rate time series that will be reported in this section holds whichever of the two U.S. series is analyzed.

[11] For example with our monthly data, when $g = .25$, 90% completion occurs in 2 years (24 months), when $g = 5$, 90% completion occurs in 1 month. Our formula gives that for these g values, 90% of completion occurred in 24 months and 1.2 months respectively. Of course , we allow in these calculations for the higher frequency of the U.S. data, and report for comparability number of years for 90% completion of transition. To verify that our model was appropriate over the whole sample span, we calculated residual autocorrelations for the periods before and after our best estimates of the transition midpoints. These were small, giving no evidence of serious misspecification of the model for the complete data sets.

[12] For a more detailed discussion of these Amendments, see Fishe (1991).