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On structural shifts and stationarity of the ex ante real interest rate

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Abstract

New evidence supporting the Fisher effect is obtained from tests for long-run reversion in the ex ante real interest rate, constructed using the Michigan Survey data on inflation expectations. This study finds that the real interest rate may appear nonstationary when in fact it is stationary and a process shift is responsible. Interestingly, the allowance for simply a mean shift is adequate to reject the unit-root hypothesis for both real pre- and after-tax interest rates. The mean shift coincides with the dramatic reversal of inflation psychology around late 1980 or early 1981. The timing and the cause of the structural shift identified for ex ante real rates differ from that for ex post real rates discussed in previous studies.

JEL classification: F43; F32
Keywords: Fisher effect; Inflation expectations; Real interest rate; Tax effect; Mean shift

1. Introduction

The Fisher theory (Fisher, 1930) suggests that changes in inflation expectations are fully reflected in nominal interest rate adjustment. The implied one-for-one long-run relationship between the nominal interest rate and expected inflation has been keenly investigated in research studies. For the Fisher

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relationship to hold, the ex ante real interest rate—the difference between the nominal interest rate and the expected rate of inflation—should display long-run reversion. However, the postulated mean-reverting property of the real interest rate has been called into question since the unit-root finding reported by Rose (1988). Rose points out that the presence of a unit root in the real interest rate contradicts not only the Fisher relationship but also Lucas-type consumption-based asset pricing models (Breedon, 1979; Hansen & Singleton, 1983; Lucas, 1978). These asset-pricing models imply that the consumption growth rate and real asset returns should share similar time series properties, but the consumption growth rate has been found to not contain a unit root in contrast to real Treasury bill rates.

Since Rose's (1988) unit-root finding, many studies have reexamined the Fisher relationship and found it rather hard to establish the long-run Fisher effect empirically. They include, among others, Atkins and Coe (2002), Choi (1994), Coppock and Poitras (2000), Crowder and Hoffman (1996), Daniels, Nourzad, and Toutkoushian (1996), Evans and Lewis (1995), Koustas and Serletis (1999), Lai (1997), Lanne (2001), Lee, Clark, and Ahn (1998), and Mishkin (1992, 1995). These studies mostly report less than supportive evidence for the Fisher effect. Although the nominal interest rate and the inflation rate do tend to move in the same direction over time, the former does not fully adjust to the latter; consequently, they fail to maintain a one-for-one adjustment relationship over the long run. In other words, the real interest rate, which measures the deviation from the equilibrium Fisher relationship, exhibits little long-run reversion.

This study presents new evidence in support of long-run reversion in the real interest rate. Ex ante interest rate data are constructed using the Michigan Survey data on inflation expectations. A common problem with Fisher-effect tests is that expected inflation is not directly observable. Many studies use realized inflation as a proxy for expected inflation and analyze the ex post real interest rate. As noted by Evans and Lewis (1995), the use of realized inflation could generate serious biases in estimating the Fisher relationship. The market might rationally anticipate possible shifts in the inflation process, which did not materialize eventually. This would result in persistent deviations between expected inflation and realized inflation, thereby creating the appearance of permanent shocks to the real interest rate even when none was truly present. Instead of using realized inflation as the proxy, this study uses the inflation expectations data gathered by the University of Michigan's *Survey of Consumers*—a nationally representative survey conducted every month based on a random sample of U.S. households. The Michigan Survey data are widely utilized by professional investors, financial institutions, and government agencies.

The analysis here allows for possible process shifts in the ex ante real interest rate. Such process shifts can be induced by infrequent but significant economic changes such as oil price shocks and regime changes in monetary or fiscal policy. Recognizing the possible structural shifts is important because they can confound unit-root tests and bias them toward finding nonstationarity. Evans and Lewis (1995) estimate the Fisher relationship based on inflation forecasts generated by a Markov switching model of mean shifts. The generated inflation forecasts are shown to yield coefficient estimates that are closer to unity than realized inflation rates, albeit the improved estimates are still noticeably less than unity. Recent findings reported by Bekdache (1999), Johnson and Garcia (2000), and Malliaropulos (2000) further indicate that the real interest rate process may have experienced some structural breaks. To account for the structural-break possibility, the present study will analyze the ex ante real interest rate data using a variety of unit-root tests that can allow for either a mean shift or a trend shift or both at an unknown date.

2. The Fisher relationship and tax adjustment

According to Fisher (1930), the one-period nominal interest rate at time t, denoted by i_t, can be decomposed into two main components as follows:[1]

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

where r_t is the ex ante real interest rate and π_t^e is the expected inflation rate. If changes in π_t^e have no permanent impact on r_t in accordance with long-run neutrality of money, those changes in expected inflation should be reflected fully in subsequent movements of the nominal interest rate over time. This implies a one-to-one relationship between the nominal interest rate and expected inflation in the long run. For the long-run Fisher relationship to hold, the ex ante real interest rate—the difference between the nominal interest rate and expected inflation—should be a mean-reverting stationary process.

To allow for possible tax effects, this study will examine both real pre- and after-tax interest rates. The tax-adjusted version of the Fisher relationship (see, e.g., Crowder & Hoffman, 1996; Engsted, 1996) can be described by

$$(1-\tau)i_t = r_t^\tau + \pi_t^e \tag{2}$$

where τ is the marginal tax rate on interest income. Accordingly, r_t^τ, measured by $(1-\tau)i_t - \pi_t^e$, gives the real after-tax rate of return. To the extent that taxes are not fully indexed to real returns, changes in taxes may induce instability in the real interest rate. Consequently, if tax changes were nonstationary, they could lead to rejections of the Fisher effect by unit-root tests. In the empirical analysis here, real after-tax interest rates are constructed based on marginal federal income tax rates faced by four-person families earning either median or twice median income. This study finds evidence in support of the Fisher effect, even in its tax-adjusted form.

To test for the Fisher effect requires some measure of expected inflation. This study uses the inflation expectations data gathered by the University of Michigan's *Survey of Consumers*. Respondents are asked to state their expected percent change in consumer prices, on the average, during the next 12 months. The Michigan Survey has reported the median of the survey participants' inflation rate forecasts in each monthly survey since 1978. We will examine ex ante 1-year real interest rates (i.e., 1-year Treasury bill rates minus 1-year expected inflation rates) covering the sample period from January 1978 through December 2002. The Treasury bill data are from the *Federal Reserve Economic Data* (FRED).

Each ex ante real interest rate series is first tested for stationarity using regular unit-root tests that allow for no structural breaks: the augmented Dickey–Fuller test and a modified Dickey–Fuller test. Their corresponding test statistics are denoted by ADF and DF-WS. The DF-WS test, which employs weighted symmetric least squares estimation, is devised by Park and Fuller (1995) as an efficient unit-root test.[2]

[1] Strictly speaking, the exact Fisher equation is represented by $i_t = r_t + \pi_t^e(1 + r_t)$. Since $\pi_t^e r_t$ is relatively small in magnitude, the Fisher relationship can be closely approximated by Eq. (1).

[2] As illustrated by Pantula, Gonzalez-Farias, and Fuller (1994), the DF-WS test shows good power performance.

Table 1
Results of unit-root tests that allow for no structural shifts

Test type	ADF	DF-WS
Non-tax-adjusted real rate		
Without trend	− 2.403	− 2.544
With trend	− 3.150	− 2.941
Tax-adjusted real rate (Series A)		
Without trend	− 2.667	− 2.409
With trend	− 2.917	− 2.479
Tax-adjusted real rate (Series B)		
Without trend	− 2.647	− 2.523
With trend	− 2.613	− 2.520

Each sample series consists of 300 observations. Tax-adjusted interest rate series are constructed based on marginal federal income tax rates faced by four-person families: Series A is for families earning median income, and Series B is for families earning twice median income. The lag parameters of both ADF and DF-WS tests are selected based on the SIC. The null hypothesis of a unit root is tested against the one-sided alternative of no-unit root. For the ADF test, the critical values are − 2.849 (in the case without trend) and − 3.397 (in the case with trend) when using the 5% significance level (Cheung & Lai, 1995). For the DF-WS test without trend, the critical values are − 2.59 (in the no-trend case) and − 3.20 (in the trended case) when using the 5% significance level (Park & Fuller, 1995). None of the test statistics reported above is statistically significant.

Table 1 contains the unit-root test results from tests with and without a linear trend included. For both ADF and DF-WS procedures, lag selection is based on the usual Schwarz information criterion (SIC). In no case can we find significant evidence to reject the unit-root hypothesis. Tests with several different lags were also tried, and the results were found to be not sensitive to the lag specification. Next, we turn to exploring the trend-break possibility.[3]

3. Unit-root tests with structural-break alternatives

To study the relevance of the structural-break hypothesis, a variety of sequential unit-root tests will be performed. These sequential tests extend the ADF test by accounting for a possible mean or trend shift in the data process, with no prior knowledge of the break date. The treatment of an unknown break date is desirable since any arbitrarily fixed date can be subject to criticism of data mining. In fact, no theory seems able to offer information on exactly when a trend break will occur. This problem is addressed by estimating the likely breakpoint directly from the data using a sequential testing procedure.

[3] Conventional unit-root tests may have low power against slow-reverting alternatives, including fractionally integrated processes (Tsay, 2000). To investigate this possibility, the spectral GPH test (Geweke & Porter-Hudak, 1983) was applied. According to the GPH test results, the allowance for fractionally integrated dynamics failed to help detect long-run reversion in the ex ante real interest rate.

3.1. The different approaches

Two approaches for modeling structural breaks in time series have been considered in the literature (Banerjee, Lumsdaine, & Stock, 1992; Perron & Vogelsang, 1992; Vogelsang & Perron, 1998; Zivot & Andrews, 1992). One is the additive outlier (AO) approach that views the break as happening instantly and the other is the innovational outlier (IO) approach that allows the break to take place gradually over time.

There are three AO models capturing different types of instant shifts:

$$y_t = \mu + \lambda t + \eta DU_t^k + z_t \tag{3a}$$

$$y_t = \mu + \lambda t + \gamma DT_t^k + z_t \tag{3b}$$

$$y_t = \mu + \lambda t + \eta DU_t^k + \gamma DT_t^k + z_t \tag{3c}$$

where $DU_t^k = I(t > k)$ and $DT_t^k = (t - k)I(t > k)$, with $I(\cdot)$ being the indicator function. The innovation process z_t z_t is defined by $(1 - \rho L)A(L)z_t = B(L)v_t$, where $A(L)$ and $B(L)$ are lag polynomials with stable roots and v_t is white noise. When $\rho = 1$, y_t has a unit root. Model (3a) permits an abrupt mean shift to occur at time $t = k$. By contrast, a trend shift (i.e., a change in the trend slope) is allowed for at time $t = k$ under model (3b). Specification (3c) admits both mean and trend shifts. Without imposing a priori the specific form of the structural break, these different models are all applied and tested.

To test for a unit root in the AO models, detrended series are first obtained as follows:

$$y_t = \mu + \lambda t + \eta DU_t^k + \tilde{y}_t^a \tag{4a}$$

$$y_t = \mu + \lambda t + \gamma DT_t^k + \tilde{y}_t^b \tag{4b}$$

$$y_t = \mu + \lambda t + \eta DU_t^k + \gamma DT_t^k + \tilde{y}_t^c \tag{4c}$$

where $\tilde{y}_t^a$, $\tilde{y}_t^b$, and $\tilde{y}_t^c$ are residual series giving the detrended values of y_t under different models. Next, tests for $\alpha_0 = 0$ under the null hypothesis of a unit root are performed using the following regressions:

$$(1 - L)\tilde{y}_t^a = \sum_{j=0}^{p} \omega_j D_{t-j}(k) + \alpha_0 \tilde{y}_{t-1}^a + \sum_{j=1}^{p} \alpha_j (1 - L)\tilde{y}_{t-j}^a + u_t \tag{5a}$$

$$(1 - L)\tilde{y}_t^b = \alpha_0 \tilde{y}_{t-1}^b + \sum_{j=1}^{p} \alpha_j (1 - L)\tilde{y}_{t-j}^b + u_t \tag{5b}$$

$$(1 - L)\tilde{y}_t^c = \sum_{j=0}^{p} \omega_j D_{t-j}(k) + \alpha_0 \tilde{y}_{t-1}^c + \sum_{j=1}^{p} \alpha_j (1 - L)\tilde{y}_{t-j}^c + u_t \tag{5c}$$

where u_t is the error term. The one-time dummy variables, $D_{t-j}(k) = I(t = k + j + 1)$ for $j = 0, \ldots, p$, are included in cases involving a mean shift to ensure the test robustness with respect to the error correlation structure (for further discussion, see Vogelsang & Perron, 1998). The possible breakpoint, k, needs to be estimated from the data. By varying k in each regression over the sample period, k will be chosen to maximize over a sequence of F statistics testing for the significance of the break parameters: $\eta=0$ in regression (4a), $\gamma=0$ in (4b), and $\eta=0=\gamma$ in (4c). This method has been shown to perform well in picking the true breakpoint. With T being the sample size, the t statistics for testing $\alpha_0=0$ computed at where $F(\tilde{k})=\max_{h \leq k \leq T-h} F(k/T)$ are denoted by $\tau_{DF}(\text{AO}, \tilde{k}, i)$, where $i=a$, b, and c for regressions (5a)–(5c). The trimming parameter, h, is set equal to the integer part of $0.1 \times T$, similar to Banerjee et al. (1992).

In contrast to the AO approach, the IO approach entertains situations in which the break occurs gradually and slowly over time. There are three possible IO gradual-break model specifications:

$$y_t = \mu + \lambda t + \varphi(L)(\eta DU_t^k + v_t) \tag{6a}$$

$$y_t = \mu + \lambda t + \varphi(L)(\gamma DT_t^k + v_t) \tag{6b}$$

$$y_t = \mu + \lambda t + \varphi(L)(\eta DU_t^k + \gamma DT_t^k + v_t) \tag{6c}$$

where $\varphi(L)$ is a lag function, through which the lingering effects of a gradual structural break are captured. The long-run impact of a mean shift is measured by $\varphi(1)\eta$, whereas the long-run impact of a trend shift is measured by $\varphi(1)\gamma$. Parallel to the AO models, Eq. (6a) represents the mean-shift model, Eq. (6b) describes the trend-shift one, and Eq. (6c) gives the hybrid model.

Following Vogelsang and Perron (1998), the regressions for unit-root tests are given by

$$(1-L)y_t = \mu + \lambda t + \omega D_t(k) + \eta DU_t^k + \alpha_0 y_{t-1} + \sum_{j=1}^{p} \alpha_j (1-L) y_{t-j} + \varepsilon_t \tag{7a}$$

$$(1-L)y_t = \mu + \lambda t + \gamma DT_t^k + \alpha_0 y_{t-1} + \sum_{j=1}^{p} \alpha_j (1-L) y_{t-j} + \varepsilon_t \tag{7b}$$

$$(1-L)y_t = \mu + \lambda t + \omega D_t(k) + \eta DU_t^k + \gamma DT_t^k + \alpha_0 y_{t-1} + \sum_{j=1}^{p} \alpha_j (1-L) y_{t-j} + \varepsilon_t \tag{7c}$$

where ε_t is the error term. Regression (7b) has been used by Banerjee et al. (1992) and Zivot and Andrews (1992). These two studies also consider regressions similar to (7a) or (7c) but without the one-time dummy variable, $D_t(k)$. Like that in the AO models, the breakpoint in the IO models will be selected based on the maximum of the sequential F statistics which test for $\eta=0$ in regression (7a), $\gamma=0$ in (7b), and $\eta=0=\gamma$ in (7c). In each regression, the t statistic for testing $\alpha_0=0$ is computed at the identified breakpoint and the resulted statistic is denoted by $\tau_{DF}(\text{IO}, \tilde{k}, i)$ with $i=a$, b, or c.

3.2. Sequential unit-root test results

Table 2 summarizes the results of sequential unit-root tests under different instant-break alternatives. In performing each of the tests, the lag specification is determined based on the SIC, as used for both ADF and DF-WS tests. As evidenced by the test results, the allowance for a structural break helps uncover long-run reversion in the ex ante real interest rate. When either a mean or a trend shift is permitted, the unit-root null hypothesis can be rejected in favor of the no-unit-root alternative at the 5% significance level. Allowing for the joint alternative of both mean and trend shifts also yields statistically significant evidence against a unit root. These results hold true for both pre- and after-tax rates.

Table 3 presents the results of sequential unit-root tests under different gradual-break alternatives. The various models are able to capture and reveal the no-unit root behavior of the ex ante real interest rate. When either a mean shift or a trend shift is admitted under the alternative hypothesis, the unit-root null hypothesis can be rejected at the 5% significance level. The same results apply, again, to pre- and after-tax rates alike. Indeed, accounting for tax adjustment seems to make the test statistics even more significant in rejecting the unit-root hypothesis.

Combining all the test results together, there is considerable evidence supporting that the ex ante real interest rate can be well characterized as stationary around a structural break. While a trend shift may exist, the allowance for a mean shift is sufficient to reject the unit-root hypothesis. In addition, tax effects alone cannot explain the mean-shift results here, given that a mean shift is identified in pre- and after-tax rates alike. The structural-break evidence bears on statistical testing of the Fisher effect. The structural break, if it is not properly accounted for, can induce tests to

Table 2
Results of sequential unit-root tests with instant-break alternatives

	Mean-shift model $\tau_{DF}(AO, \tilde{k}, a)$	Trend-shift model $\tau_{DF}(AO, \tilde{k}, b)$	Combined model $\tau_{DF}(AO, \tilde{k}, c)$
Non-tax-adjusted real rate			
Test statistic value	−4.242**	−4.364**	−4.648**
Estimated break date	November 1980	August 1981	July 1982
Tax-adjusted real rate (Series A)			
Test statistic value	−5.073**	−4.938**	−5.192**
Estimated break date	November 1980	March 1982	February 1981
Tax-adjusted real rate (Series B)			
Test statistic value	−5.763**	−5.386**	−5.805**
Estimated break date	March 1981	September 1982	March 1981

Tax-adjusted interest rate series are constructed based on marginal federal income tax rates faced by four-person families: Series A is for families earning median income, and Series B is for families earning twice median income. The null hypothesis of a unit root is tested against the one-sided alternative hypothesis of no-unit root but an instant structural break. The lag parameter of each model is determined using the SIC. Critical values for individual tests are tabulated by Vogelsang and Perron (1998). For the mean-shift model, the critical value is −4.17 for the 5% significance level. For the trend-shift model, the critical value is −4.34 for the 5% significance level. For the hybrid model, the critical value is −4.61 for the 5% significance level.

** Statistically significant at the 5% level.

Table 3
Results of sequential unit-root tests with gradual-break alternatives

	Mean-shift model $\tau_{DF}(IO, \tilde{k}, a)$	Trend-shift model $\tau_{DF}(IO, \tilde{k}, b)$	Combined model $\tau_{DF}(IO, \tilde{k}, c)$
Non-tax-adjusted real rate			
Test statistic value	− 5.582**	− 4.583**	− 5.539**
Estimated break date	July 1980	July 1980	March 1981
Tax-adjusted real rate (Series A)			
Test statistic value	− 6.411**	− 5.183**	− 6.401**
Estimated break date	July 1980	June 1981	July 1980
Tax-adjusted real rate (Series B)			
Test statistic value	− 6.795**	− 5.615**	− 6.794**
Estimated break date	July 1980	February 1982	July 1980

Tax-adjusted interest rate series are constructed based on marginal federal income tax rates faced by four-person families: Series A is for families earning median income, and Series B is for families earning twice median income. The lag parameter of each model is chosen based on the SIC. The null hypothesis of a unit root is tested against the one-sided alternative of no-unit root but a gradual structural break. For the mean-shift model, the critical value is − 4.79 for the 5% significance level (Banerjee et al., 1992). For the trend-shift model, the critical value is − 4.39 for the 5% significance level (Banerjee et al., 1992). For the hybrid model, the critical value is − 5.16 for the 5% significance level (Vogelsang & Perron, 1998).

** Statistically significant at the 5% level.

spuriously find unit-root nonstationarity, even though the real interest rate actually displays long-run reversion.

4. On the timing of the mean shift

In analyzing the dynamics of ex post real interest rates, several studies have reported evidence of infrequent but significant process shifts. Instructively, the timing of such shifts may be linked to specific economic events. For example, the structural instability has been ascribed to possible regime shifts in

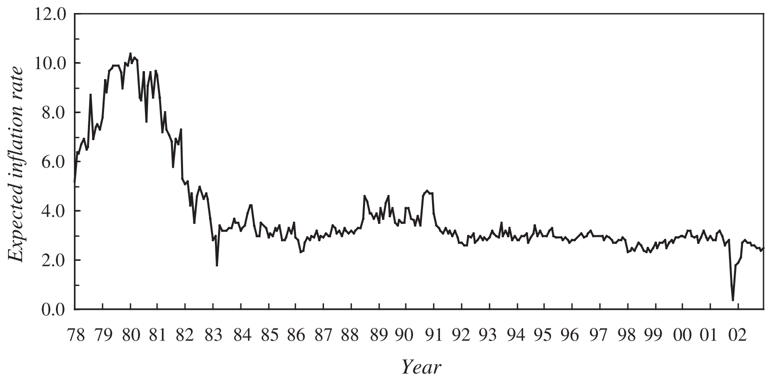

Fig. 1. Expected rate of inflation (in annual percentage rate).

Table 4
Differences in expected inflation rates (in percentage points) between subperiods

	Pre-1981	1981–2002	Difference in mean
Mean expected inflation	8.556	3.601	4.954 (.236)**
Mean expectation error	3.162	0.006	3.156 (.310)**

S.E. for the mean difference is given in parenthesis. The forecast error is measured as the expected inflation rate minus the actual inflation rate. Accordingly, a positive error indicates an overprediction of actual inflation.

** Statistically significant at the 5% level.

monetary or fiscal policy. Huizinga and Mishkin (1986) identify the occurrence of shifts in the ex post real interest rate process in October 1979 and October 1982. These dates coincide with the well-documented changes in the Federal Reserve's operating procedures to money supply targeting. Malliaropulos (2000) also attributes the located shifts to the change in monetary policy. According to Walsh (1988), the shift in October 1982 might actually correspond to a shift in the inflation process and that the rise in the federal budget deficit in early 1982 could be a contributing factor as well. Garcia and Perron (1996) report that their Markov switching model estimates of break dates are more in line with the sharp rise in the federal budget deficit in the later part of 1981 and the beginning of 1982 than with the change in monetary policy. Evans and Lewis (1995) explore an alternative explanation for structural breaks and investigate the potential impact of expected inflation shifts that might not show up in actual inflation data.

In contrast to other studies, this study analyzes the break dates for ex ante real interest rates and obtains interestingly different estimates. According to the mean-shift breakpoint estimates, the structural break is estimated to occur between mid-1980 and early 1981. The timing of the mean shift does not match directly with the specific economic events mentioned earlier, suggesting that there may be some other factor at work here.

The key difference between ex ante and ex post real interest rates is that the former is tied to inflation expectations while the latter is not. It is thus useful to look more closely at the changes in inflation expectations. Fig. 1 plots the expected rate of inflation over the sample period. A visual inspection reveals that expected inflation rates during the pre-1981 period were significantly higher than those during the post-1981 period. Further computation also shows that expected inflation consistently and substantially exceeded actual inflation during the pre-1981 period (see Table 4). The rising inflation expectations in the pre-1981 period were buoyed in large part by sharp increases in oil prices in the mid- and late 1970s and in part by lack of credibility in monetary policy (Broaddus, 1995; Dotsey & DeVaro, 1995).[4] The expected inflation rate peaked at more than 10% in the first quarter of 1980. By the end of 1980, a significant reversal of inflation psychology began after the U.S. economy experienced a steep recession in mid-1980, with real GDP falling at an annual rate exceeding 8% in the second quarter of 1980. Indeed, the downward shift in inflation expectations lasted throughout 1981 and the expected inflation rate dropped to about 5% by the end of the year. Since then, the expected inflation rate has always stayed below 5%.[5] In sum, the mean shift detected in the ex ante interest rate reflects closely the dramatic downward shift in inflation expectations around late 1980 or early 1981.

[4] As part of the credibility problem, Dotsey and DeVaro (1995) point out that while a change in operating procedures was announced in October 1979, the deep recession in 1980 caused the Federal Reserve to back off from its disinflationary policy in mid-1980. According to these researchers, the change in monetary policy was not fully credible until mid-1982.

[5] In fact, the severe recession in 1980 was quickly followed by another recession that began in mid-1981 and lasted till the fourth quarter of 1982. The significant contraction in economic activity continued to dampen inflationary expectations.

5. Conclusion

A common problem associated with testing for the Fisher effect concerns the lack of a direct measure of inflation expectations. Restricted by data availability, previous studies often use actual inflation as a proxy for anticipated inflation; consequently, they analyze ex post real interest rates under some rationality assumption about expectations. In this paper, the empirical relevance of the Fisher effect has been reexamined based on ex ante real interest rates. Expected inflation is measured directly using the Michigan Survey data on inflation expectations, which have not been explored much in previous studies. Without allowing for any structural breaks, regular unit-root tests consistently fail to uncover stationarity in the real interest rate. When unit-root tests permitting either a mean shift or a trend shift are applied, however, significant evidence in favor of no-unit root can be unveiled, strongly rejecting the absence of no long-run reversion. It follows that the ex ante real interest rate may appear nonstationary when in fact it is stationary and a process shift is responsible. Interestingly, the allowance for simply a mean shift is adequate to reject the unit-root hypothesis. Moreover, the mean shift located in the ex ante interest rate coincides with the dramatic downward reversal in inflation psychology around late 1980 or early 1981 after a deep recession in mid-1980. The timing of the identified mean shift is different from other break date estimates reported previously for ex post real interest rates.

In finding no-unit root in the real Treasury bill rate, the results here resolve the puzzling inconsistency—as observed by Rose (1988)—concerning the behavior of ex ante real asset returns implied by consumption-based asset pricing models. The stationarity finding also supports that the deviation between the nominal interest rate and expected inflation displays long-run reversion, as predicted by the long-run Fisher relationship.

A final remark about the empirical approach is in order. Following a common approach in the recent literature, this study focuses its analysis on the long-run Fisher relationship. The no-unit root finding confirms the long-run Fisher effect, supporting that higher nominal interest rates reflect higher inflationary expectations over the long run. A stronger version of the Fisher relationship, on the other hand, suggests short-run comovement. It posits that a change in the nominal interest rate is associated one-to-one with an immediate change in the expected inflation rate. Under this interpretation, nonstationarity in the real interest rate can still be consistent with the Fisher hypothesis; consequently, the no-unit root finding here says little about the validity of this alternative version of the Fisher relationship.[6]

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[6] The author owes this point to an anonymous referee.

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