

Regime Changes in International Real Interest Rates: Are They a Monetary Phenomenon?

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Abstract

In this paper, we use the Bai and Perron (1998, 2001, 2003) methodology to test for multiple structural breaks in the mean real interest rate for 13 industrialized countries. We find extensive evidence of structural breaks in the mean real interest rate for all 13 countries. In an attempt to explain the breaks in international real interest rates, we also test for multiple structural breaks in the mean inflation rate for the 13 countries. Once again, we find extensive evidence of structural breaks in the mean inflation rate for all of the countries. Interestingly, the breaks in inflation rates and real interest rates often coincide, with increases (decreases) in the mean inflation rate as we move from one regime to the next typically associated with decreases (increases) in the mean real interest rate.

The time-series properties of the U.S. real interest rate have received considerable attention in the extant empirical literature. This is not surprising, as the real interest rate plays a key role in theoretical models of consumption and investment decisions, the valuation of financial assets, and the monetary transmission mechanism. The extant empirical literature has followed an interesting path. In a well-known paper, Fama (1975) presented evidence that the U.S. real interest was constant over the 1953-1971 period. This result was subsequently challenged in a number of studies (Fama and Schwert, 1977; Nelson and Schwert, 1977; Fama, 1981; Fama and Gibbons, 1982), and Mishkin (1981) rejected a constant U.S. real interest rate over the 1953-1979 and 1931-1952 periods. In the wake of these findings, studies such as Huizinga and Mishkin (1986) and Bonser-Neal (1990) did not assume a constant real interest rate in their empirical analyses. However, they did assume a stationary real interest rate, an assumption called into question by Schwert (1986) and Antoncic (1986). In support of a nonstationary real rate, Rose (1988) presented evidence that the null hypothesis of a unit root could not be rejected for the U.S. real interest rate, as well as the real interest rate for a large number of other industrialized countries.¹ Building on the insights in Perron (1989), Perron (1990) and Garcia and Perron (1996) argued that the U.S. real interest rate is better described as a stationary process around an infrequently shifting mean. Such infrequent regime changes in the mean real interest rate make it difficult to reject the unit root null hypothesis for the real interest rate.

Two recent studies continue in the direction of Perron (1990) and Garcia and Perron (1996). Caporale and Grier (2000) and Bai and Perron (2003) use the powerful Bai and Perron (1998, 2001, 2003) methodology to test for multiple structural breaks in the mean of the U.S. real interest rate. Both of these studies find significant evidence of three structural breaks (four regimes) in the mean of the *ex post* U.S. real interest rate over the 1961:1-1986:3 period.² In an interesting exercise, Caporale and Grier (2000) also find that these structural breaks are consistent with changes in political regimes—more specifically, changes in party control of the U.S. presidency or either branch of Congress. Caporale and Grier (2000) thus take an important step in trying to account for the apparent regime changes in U.S. real

interest rates.

In the present paper, we extend the recent empirical literature on the real interest rate in two important ways. First, while Caporale and Grier (2000) and Bai and Perron (2003) focus exclusively on the U.S. real interest rate, we examine whether structural breaks in the real interest rate are limited to the U.S. or whether real interest rates in other industrialized countries are also characterized by breaks in their mean. To this end, we use the Bai and Perron methodology to test for multiple structural breaks in the *ex post* real interest rate for 13 industrialized countries using quarterly postwar data. Interestingly, we find that structural breaks in the mean *ex post* real interest rate are not an exclusively U.S. phenomenon, as almost all of the other countries that we consider exhibit significant evidence of multiple structural breaks. In fact, the timing and direction of the structural breaks are often similar across countries, indicating that the breaks may have a common source or that countries have changed policies at about the same time. Our extensive evidence of structural breaks suggests that infrequent shifts in the mean of the real interest rate are a stylized fact of international macroeconomic data.

Our second extension of the recent empirical literature is in the spirit of Caporale and Grier (2000). While Caporale and Grier (2000) focus on changes in party control of the presidency or either branch of Congress as a potential source of structural breaks in the U.S. real interest rate, we consider an alternative source: regime changes in the process governing the inflation rate. Using postwar quarterly data for the same 13 industrialized countries that we examine in our real interest rate analysis, we again use the Bai and Perron methodology to test for structural breaks in the mean inflation rate. This exercise is similar to Burdekin and Siklos (1999), who test for structural breaks in the inflation rate using over a century of U.K., U.S., Canadian, and Swedish data. We find significant evidence of multiple structural breaks in the mean inflation rate for all 13 of the countries that we consider.³ Interestingly, for many countries, the dates for the inflation and real interest rate regime changes are very near to one another (in a number of instances, the regime change dates match exactly). In these cases, we also find that increases (decreases) in the mean inflation rate from one regime to another are always associated with decreases

(increases) in the mean real interest rate. If we assume that breaks in inflation rates are determined by exogenous shifts in the nature of the monetary regime, these results are in apparent conflict with inflation neutrality (monetary superneutrality), and they suggest that monetary forces are responsible for many of the shifts in international real interest rates that we detect over the postwar era. While the finding that persistent increases in inflation are associated with persistent decreases in the real interest rate may be surprising—as it violates a fundamental neutrality proposition embodied in the Fisher effect—it is consistent with other recent findings that employ a structural vector autoregression framework (King and Watson, 1997; Koustas and Serletis, 1999; Rapach, 2003). However, we emphasize that our empirical approach in the present paper does not explicitly identify the underlying economic structure, and we thus cannot conclude on the basis of our structural break tests alone that breaks in real interest rates are primarily a monetary phenomenon. It could be the case, for example, that real shocks to tastes or technologies are responsible for both the real interest rate and inflation breaks. Nevertheless, we interpret our results as being highly suggestive that changes in the monetary regime bring about persistent changes in real interest rates. Even if one does not agree with our interpretation, we offer our results concerning concurrent breaks in international inflation rates and real interest rates as a stylized fact of international macroeconomic data that requires a coherent explanation.

The rest of the paper is organized as follows. Section 1 outlines the Bai and Perron methodology. Section 2 reports the results of tests for multiple structural breaks in the mean *ex post* real interest rate for 13 industrialized countries, while Section 3 reports the results of tests for multiple structural breaks in the mean inflation rate for the same 13 industrialized countries. In Section 3, we also link the structural breaks detected in inflation rates to the breaks reported in Section 3 for the real interest rates. Section 4 concludes.

1. ECONOMETRIC METHODOLOGY

We use the Bai and Perron (1998, 2001, 2003) methodology (hereafter, BP methodology) to test for infrequent structural breaks in real interest rates, implementing the BP methodology using the procedures described in Bai and Perron (2001, 2003). In Monte Carlo experiments, Bai and Perron (2001) find the BP methodology to be quite powerful in detecting structural breaks. Caporale and Grier (2000) and Bai and Perron (2003) apply the BP methodology to U.S. real interest rates and find evidence of multiple structural breaks in the mean of the real interest rate. We are interested in whether such structural breaks also characterize real interest rates in other industrialized countries. Following Caporale and Grier (2000) and Bai and Perron (2003), we regress the real interest rate on a constant and test for structural breaks in the constant. Consider such a regression model with m breaks ($m + 1$ regimes),

$$r_t = \beta_j + \varepsilon_t, \quad t = T_{j-1} + 1, \dots, T_j, \quad (1)$$

for $j = 1, \dots, m + 1$, where r_t is the real interest rate in period t and β_j ($j = 1, \dots, m + 1$) is the mean real interest rate in the j th regime. The m -partition, (T_1, \dots, T_m) , represents the breakpoints for the different regimes (by convention, $T_0 = 0$ and $T_{m+1} = T$.) Bai and Perron (1998) explicitly treat these breakpoints as unknown, and estimates of the breakpoints are generated using the least squares principle. Consider estimating equation (1) via least squares. For each m -partition, (T_1, \dots, T_m) , the least squares estimates of β_j are generated by minimizing the sum of squared residuals,

$$S_T(T_1, \dots, T_m) = \sum_{i=1}^{m+1} \sum_{t=T_{i-1}+1}^{T_i} (r_t - \beta_i)^2. \quad (2)$$

Let the regression coefficient estimates based on a given m -partition, (T_1, \dots, T_m) , be denoted by $\hat{\beta}(\{T_1, \dots, T_m\})$, where $\beta = (\beta_1, \dots, \beta_{m+1})'$. Substituting these into equation (2), the estimated breakpoints are given by

$$(\hat{T}_1, \dots, \hat{T}_m) = \arg \min_{T_1, \dots, T_m} S_T(T_1, \dots, T_m), \quad (3)$$

where the set of admissible m -partitions is subject to a set of restrictions (given below). It is clear from equation (3) that the breakpoint estimators correspond to the global minimum of the sum of squared residuals objective function. With the breakpoint estimates in hand, it is straightforward to calculate the corresponding least-squares regression parameter estimates as $\hat{\beta} = \hat{\beta}(\{\hat{T}_1, \dots, \hat{T}_m\})$. Bai and Perron (2001) develop an efficient algorithm for the minimization problem in equation (3) based on the principle of dynamic programming.

Bai and Perron (1998) consider testing procedures aimed at identifying the number of structural breaks (m) in equation (1). They begin by specifying a statistic for testing the null hypothesis of no structural breaks against the alternative that there are $m = b$ breaks. Let (T_1, \dots, T_b) be a partition such that $T_i = [T\lambda_i]$ ($i = 1, \dots, b$). Also define R such that $(R\beta)' = (\beta_1 - \beta_2, \dots, \beta_b - \beta_{b+1})$. Bai and Perron (1998) specify the following statistic,

$$F_T(\lambda_1, \dots, \lambda_b) = \frac{1}{T} \left(\frac{T - (b+1)2}{2b} \right) \hat{\beta}' R' [R\hat{V}(\hat{\beta})R']^{-1} R\hat{\beta}, \quad (4)$$

where $\hat{\beta} = (\hat{\beta}_1, \dots, \hat{\beta}_{b+1})'$ is the vector of regression coefficient estimates, and $\hat{V}(\hat{\beta})$ is a heteroskedastic and autocorrelation consistent estimate of the variance-covariance matrix for $\hat{\beta}$. Bai and Perron (1998) next consider a type of maximum F -statistic corresponding to equation (4),

$$SupF_T(b) = F_T(\hat{\lambda}_1, \dots, \hat{\lambda}_b), \quad (5)$$

where $\hat{\lambda}_1, \dots, \hat{\lambda}_b$ minimize the global sum of squared residuals, $S_T(T\lambda_1, \dots, T\lambda_b)$, under the restriction that $(\hat{\lambda}_1, \dots, \hat{\lambda}_b) \in \Lambda_\zeta$, where $\Lambda_\zeta = \{(\lambda_1, \dots, \lambda_b); |\lambda_{i+1} - \lambda_i| \geq \zeta, \lambda_1 \geq \zeta, \lambda_b \leq 1 - \zeta\}$ for some arbitrary positive number, ζ (the trimming parameter). Bai and Perron (1998) develop two statistics, what they call the “double maximum” statistics, for testing the null hypothesis of no structural breaks against the alternative hypothesis of an unknown number of breaks given an upper bound, M . The first double maximum statistic is given by

$$UDmax = \max_{1 \leq m \leq M} SupF_T(m). \quad (6)$$

The second double maximum statistic, $WDmax$, applies different weights to the individual $SupF_T(m)$ statistics so that the marginal p -values are equal across values of m ; see Bai and Perron (1998, p. 59) for details. Finally, Bai and Perron (1998) specify what they label the $SupF_T(l+1|l)$ statistic to test the null hypothesis of l breaks against the alternative hypothesis of $l+1$ breaks. It begins with the global minimized sum of squared residuals for a model with l breaks. Each of the intervals defined by the l breaks is then analyzed for an additional structural break. From all of the intervals, the partition allowing for an additional break that results in the largest reduction in the sum of squared residuals is treated as the model with $l+1$ breaks. The $SupF_T(l+1|l)$ statistic is used to test whether the additional break leads to a significant reduction in the sum of squared residuals. Bai and Perron (1998, 2003) derive asymptotic distributions for the double maximum and $SupF_T(l+1|l)$ statistics and provide critical values for various values of π and M .

A nice feature of the BP methodology is that it allows for quite general specifications when computing test statistics and confidence intervals for the break dates and regression coefficients. These include autocorrelation and heteroskedasticity in the regression model residuals, as well as different moment matrices for the regressors in the different regimes. We use the most general specification that allows for all of these features in our applications in Sections 2 and 3 below.

Bai and Perron (1998) discuss a sequential application of the $SupF_T(l+1|l)$ statistics—a specific-to-general modeling strategy—as a way to determine the number of structural breaks. While Bai and Perron (2001) find that this procedure performs well in some settings, its performance can be improved upon when multiple breaks are present, as the $SupF_T(1|0)$ statistic can have low power in the presence of multiple breaks. On the basis of extensive Monte Carlo simulations, Bai and Perron (2001) recommend the following strategy to identify the number of breaks. First, examine the double maximum statistics to determine if any structural breaks are present. If the double maximum statistics are significant, then examine the sequence of $SupF_T(l+1|l)$ statistics to decide on the number of breaks. Finally, Bai

and Perron (2001) recommend using a trimming parameter of at least 0.15 (corresponding to $M = 5$) when allowing for heteroskedasticity and serial correlation, and we follow this recommendation for our applications.⁴

2. REGIME CHANGES IN INTERNATIONAL REAL INTEREST RATES

2.1 Data

Our quarterly nominal interest rate and inflation rate data are from the June 2000 International Monetary Fund *International Financial Statistics* (IFS) CD-ROM. We measure the nominal interest rate using the long-term government bond yield (IFS series 61..ZF..). We focus on the long-term government bond yield, as data beginning in the 1960s are available for a much larger number of countries than short-term government bond yield data on the IFS CD-ROM. In addition, it is widely believed that long-term interest rates are the most relevant for saving and investment decisions. We also conducted the analysis for the relatively few countries for which short-term interest rate data are available, and the (unreported) results are similar to those for the long-term bond yields.⁵ The 13 industrialized countries for which we have long-term bond yield data are Australia, Belgium, Canada, Denmark, France, Ireland, Italy, the Netherlands, New Zealand, Norway, Switzerland, the U.K., and the U.S. Our annualized inflation rate data for the same set of countries are based on consumer price indexes (IFS series 64..ZF..). Our sample begins in 1960:4 and extends to 1998:3. We also use the marginal tax rate data in Padovano and Galli (2001) in order to take tax effects into account. Marginal tax rates on a decadal basis are computed as the sum of the values in columns (4) and (6) of Appendix Table A1 in Padovano and Galli (2001), and these provide us with marginal tax rates with a degree of temporal variation. We define the tax-adjusted real interest rate as $r_t = (1 - \tau)i_t - \pi_t$, where τ is the marginal tax rate for the relevant decade, i_t is the nominal interest rate (long-term government bond yield), and π_t is the annualized inflation rate.⁶

2.2 Structural Change Test Results

Table 1 reports Bai and Perron (1998) statistics for tests of structural change in the mean of the real interest rate for each of the 13 countries in our sample.⁷ For all countries, both double maximum statistics are significant at conventional significance levels, so we have strong evidence of structural change in the mean of the tax-adjusted real interest rate for each country. The $F(3|2)$ statistic is significant at the 5% level or higher, while the $F(4|3)$ is insignificant, for Australia, Belgium, Norway, Switzerland, and the U.K. This suggests three structural breaks (four regimes) for the real interest rate for these four countries. The $F(4|3)$ statistic is significant at the 5% level for Italy and the U.S., while the $F(5|4)$ is insignificant for these two countries, indicating four structural breaks (five regimes) for Italy and the U.S. For Canada, France, Ireland, and New Zealand, the $F(2|1)$ statistic is significant at the 1% level, while the $F(3|2)$ is insignificant. This indicates two structural breaks (three regimes) for Canada, France, Ireland, and New Zealand. For Denmark and the Netherlands, the $F(1|0)$ statistic is significant at the 1% level, while the $F(2|1)$ is insignificant. This points to one structural break (two regimes) for Denmark and the Netherlands. The results in Table 1 show that the Caporale and Grier (2000) and Bai and Perron (2003) finding that the U.S. real interest rate is characterized by multiple regime changes over the postwar era also holds for a number of other international real interest rates. Multiple structural breaks in real interest rates appear to be a stylized fact of international macroeconomic data.

Table 2 reports the dates for the structural breaks in the mean real interest rate and their 95% confidence intervals for each of the 13 countries, in addition to the mean real interest rate for each regime. Observe that many countries have structural breaks occurring between the late 1960s and early 1970s, as well as the early 1980s. The breaks in the late 1960s to early 1970s all result in a lower real interest rate, while those in the early 1980s lead to a higher real rate. Australia, Italy, and Switzerland also each have breaks in the early 1990s. These breaks lead to a higher real interest rate in these three countries. Figure 1 plots the real interest rate for each of the 13 countries, along with the shifting mean.⁸ The fact that many of the break dates are quite close across a number of countries suggests that the breaks have a common

cause or that countries have changed policies concurrently.⁹

3. REGIME CHANGES IN INTERNATIONAL INFLATION RATES

In this section, we are interested in explaining the deterministic regime shifts in real interest rates identified in Section 2 above. Caporale and Grier (2000) find that changes in either branch of Congress or the presidency are consistent with the timing of regime shifts in U.S. real interest rates, which is compatible to actual or perceived changes in fiscal policy having important long-lived effects on U.S. real interest rates. We take a somewhat different approach and consider regime shifts in the mean inflation rate. In line with Friedman's (1963) famous dictum, shifts in the mean inflation rate can be regarded as a monetary phenomenon—for example, changes in the central bank's desired inflation rate.¹⁰ We proceed by applying the BP methodology to the inflation rate for each of the 13 countries considered above, and we then examine if structural breaks in inflation rates coincide with the real interest rate breaks identified in Section 2 above.

3.1 Structural Change Test Results

Table 3 reports Bai and Perron (1998) statistics for the inflation rate for each of the 13 countries in our sample. Similar to the results for the real interest rate reported in Table 1, we find that both double maximum statistics are significant at conventional significance levels for all countries. We thus have strong evidence of structural breaks in the means of international inflation rates. The $F(3|2)$ statistic is significant at the 5% level, while the $F(4|3)$ is insignificant, for Australia, Belgium, Denmark, Ireland, Italy, New Zealand, Switzerland, and the U.S. This suggests three breaks (four regimes) for the inflation rate for these eight countries. The $F(4|3)$ statistic is significant at the 5% level for Canada, France, Norway and the U.K., indicating four breaks (five regimes) for these four countries. For the Netherlands, the $F(2|1)$ statistic is significant at the 1% level, while the $F(3|2)$ is insignificant, supporting two breaks

(three regimes) for the Netherlands.

3.2 Coincidental Regime Changes in Inflation Rates and Real Interest Rates

Table 4 reports the break dates and the mean inflation rate for the different regimes according to the BP methodology for each of the 13 countries. For Australia, Belgium, Ireland, Italy, Norway, Switzerland, the U.K., and the U.S., the inflation break dates are typically quite similar to those we found for their respective real interest rates. In some instances, the break dates are actually identical. For Norway and Switzerland, the BP methodology identified three breaks (four regimes) for both inflation and the real interest rate. For each country, the three inflation break dates are identical to the three real interest rate break dates: for Norway, the concurrent dates are 1969:3, 1983:1, and 1988:3; for Switzerland, the concurrent dates are 1970:1, 1975:3, and 1993:1. A similar situation holds for Australia and Belgium, where we found three breaks in both inflation and the real interest rate. In each of these two countries, the dates of the real interest rate and inflation breaks are nearly identical, as they differ by less than one year. For the U.K., we found four breaks for inflation and three breaks for the real interest rate. The first three of the four break dates for inflation are identical to the three break dates for the real interest rate. For the U.S., we found three inflation breaks and four real interest rate breaks. The first set of break dates differ by less than one year for the two series, the second set are identical, and the third set again differ by less than a year.¹¹ Turning to Italy, we also found three inflation breaks and four real interest rate breaks. The three inflation breaks are reasonably close to the last three real interest rate breaks.

Overall, the inflation and real interest rate breaks for Canada, Denmark, France, Ireland, the Netherlands, and New Zealand do not match quite as well as those for the other countries, but a number of the breaks are still closely matched. For Ireland, we identified two breaks for both inflation and the real interest rate; however, the break dates for each series differ by over a year and a half (but less than 3 years). We found four breaks for Canadian inflation but only two breaks for the real interest rate. However, note that the first inflation break date matches almost exactly with the second real interest rate

break date, and the second inflation break date does match the third real interest rate break date exactly. A similar situation holds for France: four inflation breaks and two real interest rate breaks, with the two inflation breaks reasonably close to the second and third real interest rate breaks. For Denmark and the Netherlands, we found three inflation breaks but only one real interest rate break. For New Zealand, we found three breaks for inflation and two for the real interest rate, with only the third inflation break matching well with the second real interest rate break.

Using Tables 2 and 4, we next examine the relationship between the directions of mean changes in inflation rates and real interest rates as we move from one regime to the next. It is remarkable to observe that for every case in which the inflation and real interest rate breaks occur relatively close to one another, increases (decreases) in mean inflation are always associated with decreases (increases) in the mean real interest rate as we move from one regime to the next. As an example, consider the case of Norway, where the breaks in inflation and the real interest rate match exactly. As we move from the first inflation regime to the second, the average inflation rate increases from 3.75% to 8.79%, while the real interest rate decreases from -0.23% on average in the first regime to -4.29% in the second regime. As we move from the second inflation regime to the third, the average inflation rate falls to 6.73, and the real interest rate increases to -0.25% on average in the third regime. Finally, the average inflation rate decreases on average from the third to fourth regimes, while the mean real interest rate increases to 1.39 in the fourth regime. It may also be interesting to consider the case of the U.S. The first regime ends in the mid-1960s for each series, with the mean inflation rate increasing from 1.72 to 4.53 and the mean real interest rate decreasing from 1.54 to 0.41 as we move from the first to second regimes. The second regime for each series ends in the beginning of 1973, with the mean inflation rate increasing again to 8.75, and the mean real interest rate decreasing further to -1.87 . As the third regime ends for each series in the early 1980s, mean inflation falls to 3.33, and the mean real interest rate rises to 5.20. (There is another regime shift in the real interest rate in the mid-1980s that is not associated with a regime break in inflation.)

Table 5 summarizes our findings as they relate to closely dated shifts in the mean inflation rate

and mean real interest rate for each country. It reports the date for any break in the mean inflation rate or real interest rate, and a directional arrow indicates whether the mean inflation rate or real interest rate increases or decreases at the time of the break. We place the dates for an inflation break and a real interest rate break on the same line in Table 5 if the 95% confidence intervals for the inflation rate and real interest rate break dates reported in Tables 2 and 4 overlap. From Table 5, we see that of the 33 real interest rate breaks detected, 28 of these are matched by breaks in the inflation rate in the sense that the 95% confidence intervals for the real interest rate and inflation rate break dates overlap.¹² For these 28 matches, we can also see from Table 5 that increases (decreases) in the mean inflation rate are always associated with decreases (increases) in the mean real interest rate.¹³

3.3 Discussion

If one were to have a strong prior belief that regime changes in the mean inflation rate are primarily due to exogenous changes in monetary policy, our results suggest that a long-run Fisher effect—and thus long-run inflation neutrality—often does not obtain in our international data. Under a full Fisher effect, we would not expect regime shifts in the mean inflation rate to correspond to regime shifts in the mean real interest rate, but our international data are characterized by such corresponding regime shifts when we employ the BP methodology.¹⁴ In addition, our finding that an increase (decrease) in the mean inflation rate is typically associated with a decrease (increase) in the mean real interest rate indicates that the nominal interest rate “under-adjusts” to a change in the inflation regime.¹⁵

Of course, it is important to emphasize that the approach in the present paper is strictly statistical, and so we cannot make strong statements about the underlying economic structure. It could be the case that the concurrent breaks in inflation and the real interest rate are not the primarily the product of exogenous shifts in monetary policy, but of real shocks, such as shocks to tastes or technologies.¹⁶ Nevertheless, we think it is unlikely that non-monetary shocks can provide an explanation for all of the concurrent breaks that we detect in inflation rates and real interest rates. Take the case of Australia as an

example.¹⁷ Australia exhibits three concurrent breaks in inflation and the real interest rate. The first pair of concurrent breaks occur in the early 1970s. In this case, the mean inflation rate increases and the mean real interest rate decreases. One could attribute this decrease in the mean real interest rate to a negative supply (technology) shock corresponding to an increase in the world price of oil around this time. However, the increase in inflation appears to precede the rise in world oil prices by a number of periods, so the timing may be off for this explanation. Another pair of concurrent breaks occur in the early 1980s, with the mean inflation rate decreasing and the mean real interest rate increasing. This increase in the mean of the real interest rate could be attributed to an expansionary fiscal shock, as the government's deficit (as a share of GDP) increases around this time.¹⁸ While non-monetary shocks may provide plausible explanations for the first two pairs of concurrent breaks, the last pair of breaks, occurring in the early 1990s, do not have a ready non-monetary explanation. For this pair of breaks, the mean inflation rate decreases, while the mean real interest rate increases. There is no obvious supply shock to point to during this period, and the government's budget was actually in surplus during the early 1990s, which, if anything, should work toward lowering the real interest rate. The government did adopt inflation targets at the start of 1993 (Siklos, 1999), suggesting a change in the monetary regime in the early 1990s.¹⁹ It seems quite reasonable to interpret the last pair of breaks as the product of an exogenous disinflation that, according to our results, lowered the mean inflation rate and increased the mean real interest rate. While it is beyond the scope of the present paper to provide a complete narrative analysis of each regime break in the mean inflation rate and mean real interest rate that we detect using the Bai and Perron methodology, we reiterate our contention that all of the concurrent breaks are unlikely to be the result of non-monetary shocks.²⁰

It is also instructive to note that three recent empirical studies working in a structural vector autoregression framework find that permanent (exogenous) shocks to inflation lower the long-run real interest rate. These studies use theoretically motivated assumptions in an attempt to plausibly identify structural shocks in vector autoregression models with stochastic trends. King and Watson (1997) find

that the U.S. nominal interest rate under-adjusts in the long run to a permanent inflation shock, so that the long-run real interest rate falls. Using data from eight industrialized countries, Koustas and Serletis (1999) generally reject a long-run Fisher effect, with the long-run nominal interest rate under-adjusting to a permanent increase in inflation and thereby lowering the real interest rate. Rapach (2003) also finds that the real interest rate falls in the long run in response to a permanent inflation shock in 14 industrialized countries.²¹ The assumption of stochastic trends—as in King and Watson (1997), Koustas and Serletis (1999), and Rapach (2003)—and deterministic breaks—as in the present paper—can be viewed as complementary ways of modeling persistent changes in international inflation rates and real interest rates. Both of these approaches indicate that persistent increases in inflation are associated with persistent decreases in the real interest rate.

Furthermore, a number of theoretical models predict that a permanent change in the inflation rate can affect the steady-state real interest rate in a manner consistent with the results in Tables 1-5. For example, the theoretical models of Mundell (1963), Tobin (1965), Weiss (1980), and Espinosa-Vega and Russell (1998a,b) all predict that an increase in the steady-state inflation rate decreases the steady-state real interest rate. In Mundell (1963), a permanent increase in inflation lowers the steady-state real interest rate due to the dependence of saving on real money balances. An increase in inflation increases the nominal interest rate and thus reduces the demand for real money balances; the decrease in real money balances increases saving via a wealth effect, and the real interest rate must fall to restore goods market equilibrium. Tobin (1965) develops an augmented Solow growth model where agents hold both capital and real money balances. An increase in steady-state inflation increases the costs of holding real money balances, leading agents to substitute capital holdings for real money balances in their portfolios. The portfolio reallocation increases the steady-state capital stock and thereby lowers the marginal product of capital, which equals the real interest rate in equilibrium. A potential criticism of the Mundell (1963) and Tobin (1965) models is that they are not based on explicitly optimizing behavior by economic agents. However, Weiss (1980) and, more recently, Espinosa-Vega and Russell (1998a,b) show that the type of

real interest rate response to an increase in the steady-state inflation rate predicted by the models in Mundell (1963) and Tobin (1965) can emerge in models with explicitly optimizing agents in an overlapping generations framework.²² Espinosa-Vega and Russell (1998a,b) show that the decrease in the steady-state real interest rate that results from an increase in the steady-state inflation rate can be sizable.

4. CONCLUSION

In this paper, we extend the empirical real interest rate literature in two ways. First, we examine whether the structural breaks detected by Caporale and Grier (2000) and Bai and Perron (2003) for the U.S. real interest rate also characterize real interest rates in other industrialized countries. We obtain extensive evidence of multiple breaks in the mean postwar real interest rate for 13 industrialized countries. Thus, we find that the structural breaks detected by Caporale and Grier (2000) and Bai and Perron (2003) are not limited to the U.S. In addition, we test for structural breaks in the inflation rate for the same 13 industrialized countries, with an eye toward explaining the real interest rate breaks. We detect structural breaks in all of the inflation rate series, and, interestingly, the inflation breaks often occur quite closely to the real interest rate breaks. Moreover, when the breaks are closely dated, regime shifts that increase (decrease) the mean inflation rate always correspond to regime shifts that decrease (increase) the mean real interest rate.²³ Our results strongly suggest that shifts in the mean of international real interest rates are often a monetary phenomenon.

We conclude by noting that even if a reader does not agree that our results are highly suggestive of the notion that shifts in the mean of international real interest rates are often a monetary phenomenon, we offer our results as a stylized fact of international macroeconomic data. While we are sympathetic to a monetary explanation, we clearly acknowledge that there may be plausible alternatives, and we believe that the testing of possible explanations is a fertile area for future research.

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Table 1: Bai and Perron (1998) double maximum and $SupF_T(l+1|l)$ statistics for tests of multiple structural breaks in the mean of the real interest rate

Country	$UDmax^a$	$WDmax(5\%)^b$	$F(1 0)^c$	$F(2 1)^c$	$F(3 2)^c$	$F(4 3)^c$	$F(5 4)^c$
Australia	33.89**	74.34*	12.31*	12.20*	16.93**	5.11	-
Belgium	147.33**	212.09*	12.32**	13.19*	88.18**	4.58	0.20
Canada	41.85**	49.73*	27.81**	14.56**	2.34	0.26	-
Denmark	31.25**	51.11*	31.25**	2.08	1.28	1.28	-
France	245.92**	354.02*	14.19**	31.66**	4.88	1.86	1.25
Ireland	54.80**	65.91*	49.03**	14.28**	1.70	1.40	0.27
Italy	30.51**	55.53*	6.53	19.57**	31.06**	31.06**	-
Netherlands	26.70**	52.89*	26.70**	2.82	9.64 [†]	0.75	-
New Zealand	169.60**	298.45*	8.30 [†]	46.22**	3.40	4.45	-
Norway	84.58**	100.52*	11.99*	187.34**	18.53**	1.02	-
Switzerland	26.69**	35.86*	26.69**	8.89 [†]	11.79*	0.99	-
U.K.	21.88**	36.03*	8.93*	16.87**	11.17*	11.17 [†]	-
U.S.	202.98**	297.99*	13.60**	68.26**	15.38**	13.77*	0.91

Notes: [†],*,** indicate significance at the 10, 5, and 1 percent levels, respectively; - indicates that there was no more place to insert an additional break given the minimal length requirement.

^aOne-sided (upper-tail) test of the null hypothesis of 0 breaks against the alternative hypothesis of an unknown number of breaks given an upper bound of 5; 10, 5, and 1 percent critical values equal 7.46, 8.88, and 12.37, respectively.

^bOne-sided (upper-tail) test of the null hypothesis of 0 breaks against the alternative hypothesis of an unknown number of breaks given an upper bound of 5; critical value equals 9.91.

^cOne-sided (upper-tail) test of the null hypothesis of l breaks against the alternative hypothesis of $l+1$ breaks; $F(1|0)$, $l=0$; $F(2|1)$, $l=1$;...; $F(5|4)$, $l=4$; 10, 5, and 1 percent critical values are: $F(1|0)$, 7.04, 8.58, and 12.29; $F(2|1)$, 8.51, 10.13, and 13.89; $F(3|2)$, 9.41, 11.14, and 14.80; $F(4|3)$, 10.04, 11.83, and 15.28; $F(5|4)$, 10.58, 12.25, and 15.76.

Table 2: Bai and Perron (2001) estimates of the mean real interest rate and end date for each real interest rate regime

Country	Regime 1	Regime 2	Regime 3	Regime 4	Regime 5
Australia	0.86 (0.55) 1972:4 [1967:4,1975:1]	-4.36 (1.20) 1980:2 [1979:3,1983:4]	1.72 (0.43) 1990:1 [1988:1,1994:2]	3.97 (0.33)	
Belgium	0.97 (0.45) 1972:1 [1971:3,1973:1]	-4.88 (0.30) 1977:4 [1977:2,1978:1]	0.35 (0.19) 1984:1 [1982:4,1984:3]	2.74 (0.17)	
Canada	1.99 (0.35) 1972:3 [1967:1,1975:3]	-0.31 (0.48) 1982:4 [1981:3,1984:2]	4.26 (0.32)		
Denmark	-0.89 (0.38) 1981:4 [1979:4,1987:1]	1.77 (0.27)			
France	0.28 (0.37) 1972:1 [1968:4,1973:1]	-3.60 (0.57) 1983:4 [1983:2,1985:1]	2.48 (0.16)		
Ireland	0.05 (0.47) 1969:4 [1961:4,1971:2]	-3.12 (0.75) 1982:4 [1982:2,1984:2]	4.72 (0.32)		
Italy	0.34 (0.36) 1966:1 [1965:3,1968:1]	2.52 (0.12) 1972:3 [1972:2,1972:3]	-5.54 (0.87) 1981:2 [1980:1,1983:3]	0.65 (0.60) 1991:1 [1990:1,1998:3]	3.00 (0.28)
Netherlands	-0.12 (0.36) 1976:2 [1972:2,1980:2]	2.30 (0.30)			
New Zealand	0.33 (0.59) 1970:2 [1967:2,1971:1]	-6.21 (0.75) 1987:3 [1987:1,1989:1]	1.36 (0.35)		
Norway	-0.23 (0.18) 1969:3 [1968:2,1969:4]	-4.29 (0.28) 1983:1 [1982:3,1984:1]	-0.25 (0.25) 1988:3 [1984:4,1989:4]	1.39 (0.28)	
Switzerland	0.40 (0.44) 1970:1 [1965:4,1973:3]	-2.03 (0.60) 1975:3 [1971:4,1977:4]	1.09 (0.41) 1993:1 [1992:3,1997:4]	2.62 (0.17)	
U.K.	0.05 (0.25) 1968:1 [1964:3,1971:1]	-1.27 (0.30) 1973:3 [1970:4,1973:4]	-5.39 (1.07) 1981:1 [1980:2,1983:2]	1.44 (0.34)	
U.S.	1.54 (0.10) 1966:1 [1960:4,1966:2]	0.41 (0.28) 1973:1 [1972:2,1974:2]	-1.87 (0.17) 1981:2 [1980:4,1981:3]	5.20 (0.29) 1986:4 [1986:3,1987:4]	2.55 (0.12)

Notes: The number of regimes for each country is selected according to the results in Table 1 as described in the text. The first number in each cell is the estimated mean for the regime; standard errors are reported in parentheses. The end date for the regime is below the mean; 95% confidence intervals for the end dates are reported in brackets. The first regime begins in 1960:4, and the last regime ends in 1998:3.

Table 3: Bai and Perron (1998) double maximum and $SupF_T(l+1|l)$ statistics for tests of multiple structural breaks in the mean of the inflation rate

Country	$UDmax^a$	$WDmax(5\%)^b$	$F(1 0)^c$	$F(2 1)^c$	$F(3 2)^c$	$F(4 3)^c$	$F(5 4)^c$
Australia	79.15**	173.69**	7.10 [†]	56.11**	15.51**	9.09	-
Belgium	66.17**	124.42**	8.22 [†]	35.28**	20.21**	1.63	1.63
Canada	1067.84**	2343.24**	20.56**	17.64**	138.74**	103.02**	-
Denmark	157.32**	226.48**	22.48**	44.38**	37.60**	1.38	-
France	685.24**	1178.23**	9.63*	85.93**	32.56**	32.58**	-
Ireland	153.23**	182.09	6.73	58.52**	11.04 [†]	11.04	-
Italy	377.16**	448.22**	8.79*	116.70**	21.82**	1.91	0.69
Netherlands	75.08**	164.76**	20.18**	13.21*	3.67	-	-
New Zealand	145.22**	318.67**	9.10*	42.59**	26.08**	2.30	-
Norway	81.52**	133.11**	13.88**	109.44**	14.83**	58.00**	-
Switzerland	29.79**	33.70**	29.79**	2.07	30.20**	2.71	0.50
U.K.	45.52**	99.89**	2.69	38.05**	23.18**	15.90**	-
U.S.	105.86**	204.32**	16.94**	13.20*	65.51**	9.54	-

Notes: [†],*,** indicate significance at the 10, 5, and 1 percent levels, respectively; - indicates that there was no more place to insert an additional break given the minimal length requirement.

^aOne-sided (upper-tail) test of the null hypothesis of 0 breaks against the alternative hypothesis of an unknown number of breaks given an upper bound of 5; 10, 5, and 1 percent critical values equal 7.46, 8.88, and 12.37, respectively.

^bOne-sided (upper-tail) test of the null hypothesis of 0 breaks against the alternative hypothesis of an unknown number of breaks given an upper bound of 5; critical value equals 9.91.

^cOne-sided (upper-tail) test of the null hypothesis of l breaks against the alternative hypothesis of $l+1$ breaks; $F(1|0)$, $l=0$; $F(2|1)$, $l=1$;...; $F(5|4)$, $l=4$; 10, 5, and 1 percent critical values are: $F(1|0)$, 7.04, 8.58, and 12.29; $F(2|1)$, 8.51, 10.13, and 13.89; $F(3|2)$, 9.41, 11.14, and 14.80; $F(4|3)$, 10.04, 11.83, and 15.28; $F(5|4)$, 10.58, 12.25, and 15.76.

Table 4: Bai and Perron (2001) estimates of the mean inflation rate and end date for each inflation rate regime

Country	Regime 1	Regime 2	Regime 3	Regime 4	Regime 5
Australia	2.56 (0.46) 1971:1 [1969:4,1971:3]	10.11 (0.62) 1983:2 [1981:2,1989:1]	6.98 (0.48) 1990:3 [1989:4,1991:3]	2.01 (0.31)	
Belgium	3.17 (0.46) 1972:1 [1971:1,1973:1]	8.93 (0.44) 1977:3 [1975:3,1979:2]	6.27 (0.38) 1985:1 [1984:1,1985:4]	2.17 (0.28)	
Canada	1.77 (0.17) 1966:1 [1965:3,1966:3]	3.86 (0.11) 1972:4 [1972:2,1973:1]	9.22 (0.39) 1982:4 [1982:2,1983:4]	4.49 (0.21) 1991:3 [1991:1,1992:1]	1.40 (0.15)
Denmark	5.78 (0.46) 1972:4 [1971:4,1974:3]	10.20 (0.37) 1983:1 [1982:2,1983:4]	4.64 (0.39) 1989:3 [1989:1,1991:1]	2.05 (0.14)	
France	3.40 (0.36) 1968:1 [1967:3,1969:1]	5.86 (0.11) 1973:3 [1973:1,1973:4]	10.37 (0.19) 1984:3 [1984:1,1984:4]	3.38 (0.29) 1992:1 [1991:3,1993:4]	1.59 (0.11)
Ireland	5.27 (0.96) 1972:3 [1971:4,1974:3]	13.65 (0.50) 1984:3 [1984:1,1984:4]	2.82 (0.38)		
Italy	4.01 (0.14) 1972:4 [1972:2,1973:1]	15.14 (0.38) 1984:1 [1983:2,1984:2]	6.23 (0.50) 1992:1 [1991:2,1994:3]	3.63 (0.23)	
Netherlands	4.13 (0.63) 1970:3 [1969:2,1973:2]	8.28 (0.48) 1976:4 [1975:2,1979:1]	5.43 (0.38) 1982:3 [1981:2,1983:2]	1.96 (0.24)	
New Zealand	2.70 (0.29) 1966:4 [1964:1,1967:1]	6.46 (0.66) 1973:2 1971:2,1974:2]	12.62 (0.60) 1987:4 [1986:4,1988:3]	2.93 (0.70)	
Norway	3.75 (0.19) 1969:3 [1968:1,1969:4]	8.79 (0.44) 1983:1 [1981:4,1990:2]	6.73 (0.30) 1988:3 [1986:3,1988:4]	2.63 (0.48)	
Switzerland	3.22 (0.26) 1970:1 [1967:4,1970:3]	7.14 (0.69) 1975:3 [1974:4,1977:3]	3.22 (0.20) 1993:1 [1991:2,1994:2]	1.00 (0.34)	
U.K.	3.41 (0.21) 1968:1 [1966:1,1968:2]	7.06 (0.56) 1973:3 [1970:3,1973:4]	14.23 (1.19) 1981:4 [1980:4,1983:4]	5.72 (0.67) 1991:2 [1990:4,1993:3]	2.88 (0.18)
U.S.	1.72 (0.24) 1967:3 [1966:4,1968:2]	4.53 (0.24) 1973:1 [1971:3,1973:2]	8.75 (0.43) 1982:1 [1980:4,1982:3]	3.33 (0.35)	

Notes: The number of regimes for each country is selected according to the results in Table 1 as described in the text. The first number in each cell is the estimated mean for the regime; standard errors are reported in parentheses. The end date for the regime is below the mean; 95% confidence intervals for the end dates are reported in brackets. The first regime begins in 1960:4, and the last regime ends in 1998:3.

Table 5: Inflation rate and real interest rate regime break dates and direction of change in the mean inflation rate and mean real interest rate in each regime

Australia		Belgium		Canada		Denmark		France	
Inflation	Real rate	Inflation	Real rate	Inflation	Real rate	Inflation	Real rate	Inflation	Real rate
1971:1 ↑	1972:4 ↓	1972:1 ↑	1972:1 ↓	1966:1 ↑		1972:4 ↑		1968:1 ↑	
1983:2 ↓	1980:2 ↑	1977:3 ↓	1977:4 ↑	1972:4 ↑	1972:3 ↓	1983:1 ↓	1981:4 ↑	1973:3 ↑	1972:1 ↓
1990:3 ↓	1990:1 ↑	1985:1 ↓	1984:1 ↑	1982:4 ↓	1982:4 ↑	1989:3 ↓		1984:3 ↓	1983:4 ↑
				1991:3 ↓				1992:1 ↓	
Ireland		Italy		Netherlands		New Zealand		Norway	
Inflation	Real rate	Inflation	Real rate	Inflation	Real rate	Inflation	Real rate	Inflation	Real rate
	1969:4 ↓		1966:1 ↑	1970:3 ↑		1966:4 ↑		1969:3 ↑	1969:3 ↓
1972:3 ↑		1972:4 ↑	1972:3 ↓	1976:4 ↓	1976:2 ↑		1970:2 ↓	1983:1 ↓	1983:1 ↑
1984:3 ↓	1982:4 ↑	1984:1 ↓	1981:2 ↑	1982:3 ↓		1973:2 ↑		1988:3 ↓	1988:3 ↑
		1992:1 ↓	1991:1 ↑			1987:4 ↓	1987:3 ↑		
Switzerland		U.K.		U.S.					
Inflation	Real rate	Inflation	Real rate	Inflation	Real rate				
1970:1 ↑	1970:1 ↓	1968:1 ↑	1968:1 ↓		1966:1 ↓				
1975:3 ↓	1975:3 ↑	1973:3 ↑	1973:3 ↓	1967:3 ↑					
1993:1 ↓	1993:1 ↑	1981:4 ↓	1981:1 ↑	1973:1 ↑	1973:1 ↓				
		1991:2 ↓		1982:1 ↓	1981:2 ↑				
					1986:4 ↑				

Notes: Break dates on the same line indicate that the real interest rate and inflation rate break date confidence intervals reported in Tables 2 and 4 overlap; ↑(↓) indicates that the mean in the real interest rate or inflation rate increases (decreases) at the break.

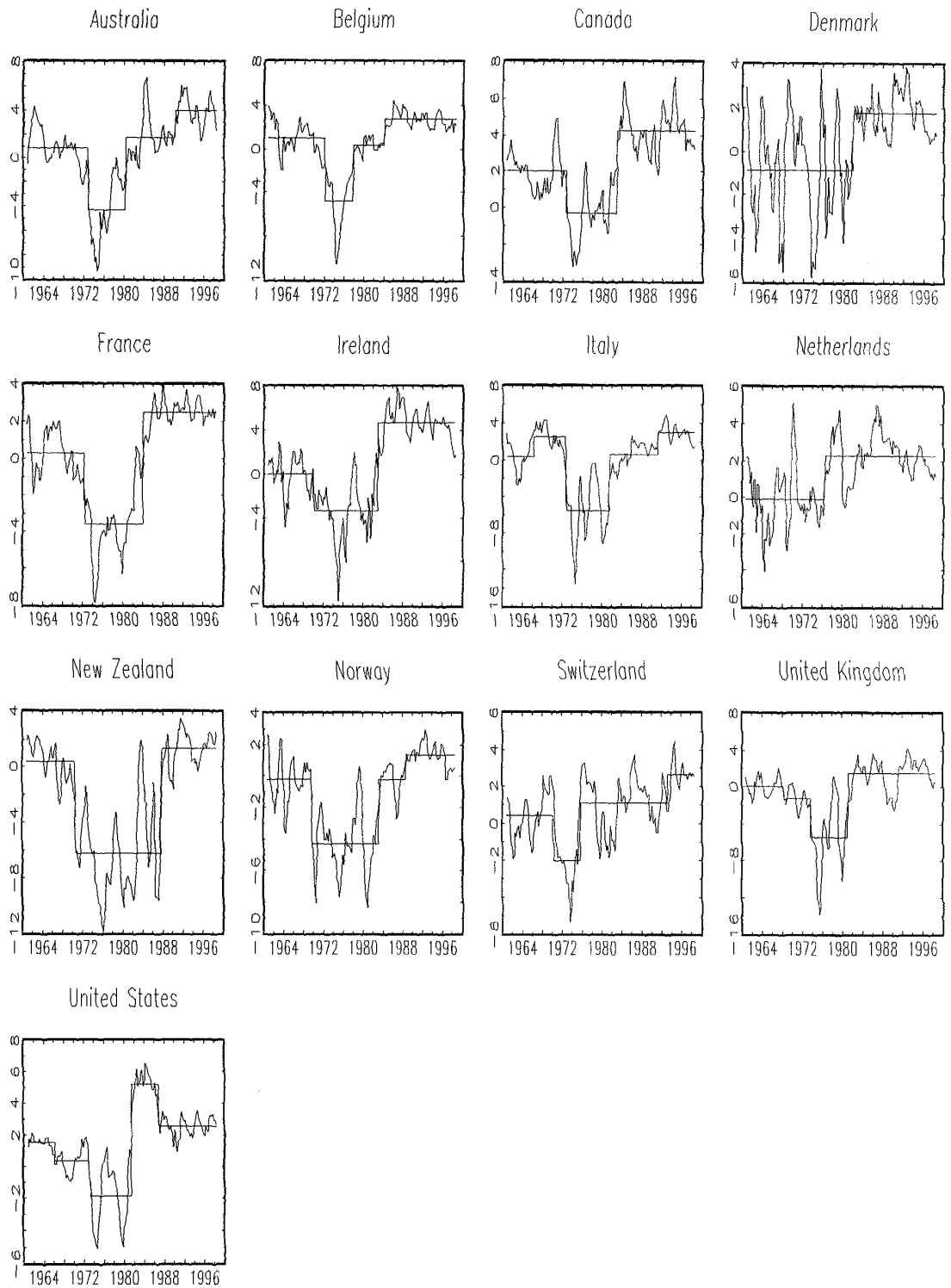


Figure 1: Tax-adjusted *ex post* real interest rates and means for the different regimes indicated by the Bai and Perron methodology.

NOTES

¹ Also see Walsh (1987), who finds evidence of a unit root in the U.S. real interest rate over various sample periods.

² The breakpoints are nearly identical in the two studies: 1967:1, 1972:4, and 1980:2 for Caporale and Grier (2000); 1966:4, 1972:3, and 1980:3 for Bai and Perron (2003). The studies use slightly different real interest rate measures.

³ Burdekin and Siklos (1999) also find evidence of multiple structural breaks in inflation rates for the countries they consider. Levin and Piger (2003) test for a structural break in the inflation rate for 12 industrialized countries over the 1984-2002 period.

⁴ We implement the BP methodology using the GAUSS program available from Pierre Perron's home page at <http://econ.bu.edu/perron/>.

⁵ All unreported results are available at <http://fac-staff.seattleu.edu/rapachd/research.htm>.

⁶ We follow the convention in the literature and match up the nominal interest rate for a given quarter with the inflation rate over the following quarter. We are also following much of the extant literature, including Caporale and Grier (2000) and Bai and Perron (2003), by using the *ex post*, and not the *ex ante*, real interest rate. This allows us to sidestep the thorny issue of specifying explicitly how inflationary expectations are formed and should not be crucial for measuring the average real interest rate over relatively long periods. If expectations are rational, actual and expected inflation will only differ by a white-noise error term. As some of the consumer price indexes display strong seasonal variation, we seasonally adjust the inflation rates by averaging over the current and previous three quarters.

⁷ In our applications of the BP methodology, we allow for the most general specification by setting $cor_u = 1$, $het_u = 1$, and $het_z = 1$ in the notation of BP (2001).

⁸ Note that few of the break dates occur at the beginning of a decade (for example, 1970:1), so that the breaks are not a statistical artifact of using decadal marginal tax rates. We also repeated the analysis using tax-unadjusted real interest rates, and the (unreported) results are similar for most countries.

⁹ It would be interesting in future research to extend the BP methodology to a panel framework in order to potentially identify more precisely break dates that are common across countries.

¹⁰ As argued by McCallum (1990), there is widespread support for Friedman's dictum that "inflation is always and everywhere a monetary phenomenon" when understood in its proper perspective. The dictum should not be interpreted as saying that non-monetary factors have no effect on the inflation rate in the short run, but that the steady-state inflation rate is primarily determined by the growth rate of the money supply. Since we are focusing on relatively infrequent changes in the mean inflation rate, this interpretation of Friedman is appropriate.

¹¹ The break dates found for the U.S. and Canadian inflation rates are close to those found by Atkins (2002). He reports three breaks for U.S. inflation occurring in 1965:4, 1973:1, and 1981:3 and four breaks for Canadian inflation occurring at 1965:4, 1972:4, 1982:3, and 1991:1.

¹² Of the 33 real interest rate breaks, 22 are matched by breaks in the inflation rate occurring within one year.

¹³ Given that we test for breaks in the real interest rate and inflation rate, we theoretically can identify breaks in the nominal rate using the breaks identified for the real interest rate and inflation rate and the definition linking the real rate to the nominal interest rate and inflation rate. Nevertheless, we also used the BP methodology to test for structural breaks in the (tax-adjusted) nominal interest rate. In line with the results reported in Tables 1-5, the (unreported) results show a number of breaks in the nominal interest rate that correspond with the inflation rate and real interest rate breaks. When all of the breaks are closely matched, the mean of the nominal interest

rate typically increases (decreases) when the mean of the inflation rate increases (decreases). However, the change in the nominal interest rate is smaller in magnitude than the change in the inflation rate, and this often shows up as a break in the mean of the real interest rate.

¹⁴ Note that there are cases in Table 5 where regime shifts in the mean inflation rate are not matched by regime shifts in the mean real interest rate. This indicates that some changes in the inflation regime are neutral or that any change in the mean real interest rate is not large enough to show up as a significant break according to the BP methodology.

¹⁵ See footnote 13 above.

¹⁶ A relatively simple story based on persistent changes in real output growth, in the spirit of an endogenous growth model, proceeds as follows. Assume a standard money demand function and that money growth is constant. Also assume that equilibrium is characterized by the equality of the real interest rate and the real output growth rate. In this case, changes in real output growth bring about changes in the real interest rate and the inflation rate, and the directions of the changes are consistent with those in Table 5. This story may be a plausible explanation for some of the concurrent shifts in the mean inflation rate and mean real interest rate and deserves further consideration in future research.

¹⁷ The timing and direction of the real interest rate and inflation breaks in Australia are fairly typical of those in a number of other countries, so we use focus on Australia as an example.

¹⁸ The data on the Australian budget deficit are from the IFS CD-ROM.

¹⁹ We do not necessarily expect the adoption of formal inflation targets by a government to correspond to a regime shift in the mean inflation rate, although this appears to be the case for Australia.

²⁰ A detailed narrative analysis of the breaks is a potentially useful topic for future research.

²¹ See the survey on recent tests of long-run neutrality propositions by Bullard (1999).

²² Tobin (1965) relies on the substitution of real capital for fiat money in agents' portfolios. However, since the fiat money to capital stock ratio is small, the Tobin (1965) effect is unlikely to be important empirically. An attractive feature of the models in Espinosa-Vega and Russell (1998a,b) is that they rely on the ratio of currency-seigniorage-revenue to bond-seigniorage-revenue. This ratio is larger than the fiat money to capital stock ratio by an order of magnitude and is thus much more likely to be empirically relevant.

²³ Note that our results help explain the recent findings of Hafer (2002). In a comment on Caporale and Grier (2000), Hafer (2002) regresses the U.S. real interest rate on a set of dummy variables corresponding to the political regimes in Caporale and Grier (2000), as well as the inflation rate. With the inflation rate included as a regressor in the regression model, some of the political regime changes in Caporale and Grier (2000) are no longer significant, and inflation enters significantly into the model. Given that we find inflation breaks occurring almost concurrently with real interest rate breaks in the U.S., it is not surprising that some of the Caporale and Grier (2000) political regime changes become insignificant and that inflation enters significantly in Hafer's (2002) regression model.