Multiple Regimes in U.S. Monetary Policy? A Nonparametric Approach*

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Abstract

We use two different nonparametric methods to determine whether there were multiple regimes in U.S. monetary policy over the period 1955–2003. We model monetary policy using two different versions of Taylor's rule for the nominal interest rate target. By contrast with parametric tests for regime changes, the nonparametric methods we use allow the *data* to determine the dimensions on which to split the sample for purposes of estimating the coefficients of the Taylor rule. We find evidence for a few structural breaks and consistent agreement between our two nonparametric methods on the dating of those breaks.

keywords: monetary policy, regime change, structural break, Taylor Rule, nonparametric methods.

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1 Introduction

Eight times a year, approximately every six weeks and sometimes more often, the U.S. Federal Reserve (the Fed) announces its target for the key interbank borrowing rate known as the federal funds rate. The Fed's target for the "fed funds" rate is one of the most anticipated and influential decisions regularly affecting world financial markets. Not surprisingly, there have been numerous efforts to assess the process by which this decision is made and whether this process has changed over time.[1] In this paper, we employ a standard, benchmark, non–forward looking model of the federal funds rate target – namely Taylor's rule (Taylor, 1993, 1999)– and using versions of this rule, we focus attention on the second issue of whether there have been regime changes in U.S. monetary policy over the 48 year period 1955:Q3–2003:Q3. We make use of new, nonparametric methods that allow the *data* to determine whether there have been any regime changes; in particular, we do not rely in any way on the history of monetary policy decision–making or on our knowledge of other relevant historical events. This data–based determination of regime changes in U.S. monetary policy is what sets our paper apart from previous studies.

Indeed the standard approach to studying regime changes is to examine various different regimes separately; for example, Fair (2001) treats the period of October 1979–July 1982, when the Fed experimented with targeting money growth rates, as a single monetary regime. Alternatively, one might consider the terms of Fed chairmen as defining different regimes (e.g. Judd and Rudebusch (1998), Clarida et al. (1998, 2000), Taylor (1999)). A difficulty with this approach is that the various policy regimes have to be identified *a priori*. The typical procedure is to pick regime dates based on some known features of the available data and then use tests of parameter constancy, e.g. Chow tests, to justify the dates chosen. However, as Hansen (2001) observes, if the breakpoints are not known a priori, then the Chi-squared critical values for the Chow test are inappropriate. Furthermore, using known features of the data (e.g. the Volker policy experiment of 1979–1982) to determine break points can make these candidate break dates endogenous – correlated with the actual data – leading to incorrect inferences about the significance of those candidate break dates.

For these reasons, we propose the use of nonparametric methods that allow the *data* to determine whether any regime changes have occurred using a simple, common framework for the modeling of monetary policy.[2] Specifically we use versions of Taylor's rule (Taylor (1993)) for the federal funds rate target as our model of monetary policy. While Taylor's rule was originally intended to describe monetary policy in the Greenspan era, 1987–present, other researchers, e.g. Clarida et al. (1998, 2000) Judd and Rudebusch (1998), Orphanides (2003), even Taylor (1999) himself, have used versions of this rule to characterize monetary policy prior to 1987; Taylor goes the furthest, back to 1879. [3]

We use two different nonparametric methods in an effort to let the data detect changes in Taylor's rule over the sample period 1955:3–2003:3.[4] In particular we use piecewise–linear, classification and regression tree (CART) methods (Breiman et al. (1984)) and sequential tests for structural breaks where the break dates are unknown variables to be estimated (Bai and Perron (1998)). Our use of two different nonparametric methods allows us to assess the robustness of our findings. We apply these two methods to two popular specifications of Taylor's rule, the original Taylor (1993) specification and the Taylor (1993) specification with a lagged dependent variable (an "interest rate smoothing" version). Both methods suggest that either version of the Taylor rule, augmented to take account of just a few breaks or *regime changes*, does a remarkably good job of characterizing the evolution of the federal funds rate target over the sample period we examine, 1955:3–2003:3.

Our main findings are as follows. For the original Taylor rule specification, both nonparametric methods detect a strong, tightly identified break at 1979 quarter 3 (1979:3), the famous date associated with the start of Paul Volker's Fed chairmanship and strong disinflationary policy. The two methods also agree on two other breakdates in 1968:1 and 1986:4.

We interpret the 1968:1 date as a shift in Fed policy toward inflation in response to deficit financing of the Vietnam war and Great Society programs. The 1986:4 date is more difficult to interpret. For a policy-smoothing version of the Taylor rule, both nonparametric methods also agree on three break dates; these three dates differ by only a few quarters from the three break dates identified using the original Taylor rule specification. Finally, what we *do not* find is also interesting, for example, we find no evidence of changes in policy across different Fed chairmanships, with the important exception of the Volker regime change in 1979:3.

2 Taylor's Rule

Taylor's rule (Taylor, 1993) specifies how the Federal Reserve should set its target at time t for the nominal federal funds rate, i. Taylor's original (1993) rule is:

$$i_t = r + \pi_t + \lambda_\pi(\pi_t - \overline{\pi}) + \lambda_y y_t,$$

which may be rewritten as:

$$i_t = r - \lambda_\pi \overline{\pi} + (\lambda_\pi + 1)\,\pi_t + \lambda_y y_t.$$

Here, r is the target real federal funds rate (assumed to be unchanging and known to the Fed), $\overline{\pi}$ is the target inflation rate (also assumed to be unchanging and known to the Fed), π_t is the four-quarter moving average of the inflation rate as of time t, and y_t is the time t output gap, defined as [(real GDP - real potential GDP)/ real potential GDP]. In this paper, we suppose that potential GDP is determined exogenously, and we make use of the Congressional Budget Office estimates of this series, which is the standard series used to estimate Taylor rules. All other data are standard quarterly series, available from CITIBASE and other data sources. In particular, the federal funds rate is taken to be the quarterly average target rate,[5] inflation is measured by the four–quarter moving average of the implicit GDP Deflator, and real GDP is the standard NIPA series in chained 1996 dollars.

It is a simple matter to use OLS to estimate the parameters of the Taylor rule specification using data on i_t, π_t and y_t. The linear regression model is:

$$i_t = \alpha + \beta_1 \pi_t + \beta_2 y_t + \epsilon_t, \tag{1}$$

where ϵ is assumed to an i.i.d. process with mean zero and constant variance. Note that in specification (1), α is an estimate of the difference $r - \lambda_\pi \overline{\pi}$ and β_1 is an estimate of $1 + \lambda_\pi$ (and *not* λ_π). Taylor (1999) argues that across most monetary regimes in place over our sample (e.g. fixed money growth regime, international gold standard) one would expect estimates of $\beta_1 > 1$ and estimates of $\beta_2 > 0$.[6]

Researchers who have estimated Taylor–type rules for monetary policy have considered two main variations on the benchmark specification (1).[7] In the first, "forward–looking" variation, the time t inflation and output gap variables are replaced with *expected values* of those variables, where the expectations are conditional on information available through time $t-1$. In the second variation, a lagged dependent variable is included on the right hand side. The latter specification is usually written as

$$i_t = (1 - \rho)\left[\alpha + \beta_1 \pi_t + \beta_2 y_t\right] + \rho i_{t-1} + \epsilon_t, \tag{2}$$

where the parameter $\rho < 1$ captures policy inertia or a more deliberate policy of interest rate smoothing. Sometimes both forward–looking and policy-smoothing versions are combined, as in Clarida et al. (2000).

Here, in addition to specification (1), we consider only the lagged dependent variable specification (2) alternative form of the Taylor rule.[8] To estimate the forward-looking specification would require generalized method of moments (GMM) estimation techniques, and the data–driven (nonparametric) procedures we use are not equipped to handle such estimation methods.

3 CART

The first nonparametric procedure we employ is the Classification and Regression Tree (CART) approach of Breiman et al. (1994).[9] In CART, one provides the model specification and the dimensions along which changes in the model's coefficient estimates may occur. In our case, the natural dimension on which to partition the data is according to *time* though one could also partition the data according to realizations of the explanatory variables, e.g., threshold values for the inflation and output gaps.

A brief description of the CART procedure can be given here; for details see Breiman et al. (1994). The algorithm begins by running a regression of the model using all the data in the given sample. Then, the data are split into two subsamples and a regression is run on each subsample, the first subsample consisting of all data up to and including the candidate breakdate and the second subsample consisting of the data after the candidate breakdate. This procedure is repeated for each possible division of the data. The division that results in the maximal improvement in the total sum of squared errors relative to the full sample regression becomes the first 'root' split of the sample. Then for each of the these two subsamples, the same procedure is repeated. This partitioning process continues until there are too few observations left to run a least squares regression (0 degrees of freedom). The visual representation of this repeated partitioning process is a tree where the terminal nodes contain the same number of observations as there are coefficients to estimate using ordinary least squares (in the case of our Taylor rule specification each terminal nodes of this initial tree will involve 3 observations). Once this initial fully-grown tree is constructed, the second, "pruning" stage involves selection of the best subtree to describe the data. Since each split of the data reduces the overall sum of squared errors of the model, a cost for tree complexity is imposed which is proportional to the number of splits in the tree, forcing a trade–off between overall tree fitness and tree size. The optimal cost for complexity is the one that minimizes a jacknife estimate of the out–of–sample predictive variance of the model.

Figure 1 shows the tree resulting from application of the CART procedure to the Taylor rule specification (1)

[Insert Figure 1 here].

In Figure 1, each non–terminal node of the tree, or "split" is represented by the dating scheme, year:quarter. For example, the top–most split divides the data into two subsamples: all data points from the beginning of the sample *through* the split date, 1979:3, fall to the lower left descendent node, while the remaining data points fall to the lower right descendent node. At each of these subsequent nodes, the data may be further divided in the same manner, with the date of the split indicated at each node. Terminal nodes of the tree are depicted by numbers in square brackets; these numbers correspond to those used in Table 1, which reports regression results for the 8 final, CART-determined subsamples.

In the tree shown in Figure 1, the root split occurs at 1979:3. This split coincides with Paul Volker's appointment as chairman (actually on August 6, 1979) and the decision at the October 6, 1979 FOMC meeting to switch the focus of monetary policy to tighter control of the monetary base in an effort to bring down the high U.S. inflation rate. This is perhaps the most frequently cited date in the literature examining regime changes in U.S. monetary policy. Here, the date appears at the root of the tree, as determined by the data *alone*.

Another notable feature of the CART splits is that, with the exception of the first splits (node 1), the coefficient estimates on inflation and the output gap, β_1 and β_2, are both positive and significantly different from zero at the 5% level of significance (see Table 1).[10] Recall that β_1 is actually an estimate of $1 + \lambda_\pi$, so estimates for β_1 that are less than

Table 1: Coefficient Estimates and Standard Errors For the Regression Tree with Splits on Time, Specification (1)

Terminal Node	Subsample: Dates	Estimates of α	β_1	β_2	Adj. R^2	Nobs
1	1955:3–1968:1	3.309	-0.193	0.319	0.593	51
		(0.264)	(0.128)	(0.053)		
2	1968:2–1970:4	-10.254	3.348	0.459	0.797	11
		(2.116)	(0.437)	(0.056)		
3	1971:1–1974:3	-3.58	1.711	0.695	0.824	15
		(0.674)	(0.099)	(0.107)		
4	1974:4–1979:3	1.397	**0.870**	0.788	0.825	20
		(1.058)	(0.149)	(0.125)		
5	1979:4–1981:1	-18.478	**3.688**	1.044	0.829	6
		(8.384)	(0.936)	(0.276)		
6	1981:2–1986:4	4.115	1.466	**0.190**	0.908	23
		(0.496)	(0.097)	(0.097)		
7	1987:1–1998:3	2.809	1.380	0.805	0.873	47
		(0.344)	(0.107)	(0.077)		
8	1998:4–2003:3	1.381	**1.386**	0.871	0.945	20
		(0.490)	(0.274)	(0.049)		

Note: the regressions were performed with Newey–West standard errors to reflect first order autocorrelated disturbances and heteroskedasticy. Boldface indicates that a 95% confidence interval includes 1 and 0 for the estimates of β_1 and β_2, respectively.

unity imply that the coefficient on the inflation gap, λ_π is negative, and that policy for that subsample can be regarded as destabilizing. Indeed, estimates of the coefficient on the inflation gap, β_1 are significantly greater than unity in just 4 of the 8 subsamples. Estimates of β_1 for which a 95 percent confidence interval includes unity are indicated in bold. The estimates of the coefficient on the output gap, β_2, generally vary around Taylor's choice of $\lambda_y = 0.5$; with one exception, again indicated in bold, a 95 percent confidence interval around each β_2 estimates excludes 0.

The negative estimate of the coefficient on the inflation gap for the first, twelve-year subsample (node 1) suggests that up through 1968:1, the Fed did not attach much significance to *changes* in the inflation rate from its target value as the Taylor rule emphasizes. A negative estimate for λ_π also obtains in the 1974:4–1979:3 subsample. The negative estimates for λ_π over the subsamples 1955:3-1968:1 and 1974:4–1979:3 are consistent with the findings of Taylor (1999), whose estimate of $1 + \lambda_\pi$ over the much longer period of 1960:1–1979:4, is 0.813.

A final observation concerning the results presented in Table 1 is that there appear to have been *two* major efforts at disinflationary policy, defined by increases in the estimated coefficient attached to the inflation gap variable. The first begins after 1968:1 and the second begins after 1979:3. While the latter date is well known as discussed earlier, the 1968:1 date is less well known as the beginning of a new monetary regime. We associate the change in policy in 1968:1 with a change in inflationary pressures the Fed faced in the late 1960s as the result of the federal government's deficit financing of the Vietnam war and Great Society programs. Indeed, Meulendyke (1998) notes that by the mid-1960s, the Fed was generally content with the basic course of monetary policy as price stability had been largely achieved, and recessions had been relatively mild. However by the late 1960s, the picture had abruptly changed. More precisely, as DeLong (1997) observes,

> "In prices, as measured by the GDP deflator, the major jump in inflation occurred after 1968: from 5% in 1968 to the peak of just over 10% in 1981. In wages the major jump has already occurred by 1968..." (DeLong (1997), p. 257)

Mayer (1999) further observes that the Fed seemed overly concerned with the *nominal* funds rate target prior to 1968 and tended to ignore the possibility that nominal rates carried an expected inflation premium. Mayer (1999 p. 55) also observes that "academic economists had rediscovered the Fisher effect only around 1968." These observations provide some corroboration of the 1968:1 breakdate and especially of the change in the estimated value of λ_π brought about by that break. As we shall see, both the 1968:1 and the more famous 1979:3 breakdates will continue to appear as prominent breakpoints using the sequential

Table 2: Coefficient Estimates and Standard Errors For the Regression Tree with Splits on Time, Lag Specification (2)

Terminal Node	Subsample: Dates	Estimates of α	β_1	β_2	ρ	Adj. R^2	Nobs
1	1955:3–1969:1	0.759	-0.118	0.593	0.772	0.904	54
		(0.220)	(0.269)	(0.138)	(0.063)		
2	1969:2–1979:3	0.027	**1.121**	1.351	0.653	0.870	42
		(0.542)	(0.236)	(0.295)	(0.094)		
3	1979:4–1980:3	5.486	-0.314	5.276	0.758	1.000	4
		(-)	(-)	(-)	(-)		
4	1980:4–1984:3	5.560	**1.162**	0.401	0.190	0.887	16
		(1.256)	(0.145)	(0.172)	(0.136)		
5	1984:4–2003:3	0.386	1.667	1.124	0.767	0.960	76
		(0.161)	(0.254)	(0.164)	(0.043)		

Note: the standard errors reflect a correction for first order autocorrelated disturbances as well as for heteroskedasticy using White's correction method. Boldface indicates that a 95% confidence interval includes 1 and 0 for the estimates of β_1 and β_2, respectively.

search method.

The other major break-date identified by the CART procedure is at 1986:4. The identification of this date as a break-date appears to be tied up with a big change in the coefficient attached to the output gap; the estimated coefficient value is significantly lower in the 1981:2–1986:4 subsample than in prior or subsequent subsamples (see Table 1). However, the historical significance of this break-date is difficult to ascertain.[11]

The results from applying the CART procedure to the Taylor rule specification (2) with a lagged dependent variable are shown in Figure 2 and the regression estimates are given in Table 2.

[Insert Figure 2 here.]

Rather remarkably, the change in the specification of the monetary policy rule to specification (2) results in three main splits that are not far away in time from the three main splits obtained from applying the CART procedure to specification (1). The root split now occurs at 1980:3, and the next two splits are at 1969:1 and at 1984:3. Furthermore, the 1979:3 date resurfaces again as a break–date, though it is not as prominent a break-date as it was found to be in using specification (1).

The coefficient estimates for ρ in Table 2 are, with one exception, close to .6 to .7 and significantly different from zero which is in line with other empirical studies that focus primarily on the Greenspan era.[12] This finding implies that the Fed is adjusting toward its target level at a rate of only 30-40 percent per quarter, a seemingly slow rate of adjustment. For the sixteen quarters between 1980:4–1984:3, the coefficient estimate of ρ is low and not significantly different from zero, a finding which suggests that the Fed was not beholden to a smoothing policy over this period; there appears to be little in the historical record that would substantiate such a claim.

The large coefficients on the lagged funds rate imply that variations in inflation and the output gap have little effect on the federal funds rate target. Nevertheless, the estimates of β_1 and β_2 are, with few exceptions, statistically different from zero. We note also that the estimates reported in Table 2 for the subsample 1979:4–1980:3 are highly implausible. While the standard errors are 0 (the adjusted $R^2 = 1$) the four coefficient estimates are based on only 4 observations. Still, the period identified by the CART procedure, 1979:4–1980:3 is well known to have been a period during which the Fed was experimenting with both its targets and operating procedures.[13]

In summary, the application of CART to specifications (1) and (2) suggests that a substantial change in Fed monetary policy occurred in the late 1970s, 1979:3 (or the early 1980s 1980:3) and that other major changes occurred in the late 1960s and mid 1980s. These findings are consistent with historical accounts. Indeed, the "Great Inflation" is often dated as beginning in the late 1960s and continuing through the 1970s.[14]

4 Sequential Test For Structural Change

We again treat the breakdates as unknown, but we now follow the methodology outlined in Hansen (2001). The first step is to look for visual evidence of breaks. We consider the estimate of the residual variance of the model (i.e., the sum of squared errors divided by the sample size) at all possible breakdates. This is done by splitting the sample at each possible breakdate and estimating the model specification over the two subsamples. In the absence of any structural breaks, the residual variance should vary randomly across the candidate breakdates; local minima provide evidence of breaks.

Figures 3a and 3b illustrate the residual variance (i.e., the total sum of squared errors divided by sample size) obtained from allowing a single breakdate in the estimated Taylor rule at each quarter over the period 1956:3–2002:3, for each of the two Taylor Rule specifications. Figure 3a clearly reveals a global minimum in the residual variance for specification (1) at 79:3, with a V-shape that provides good identification of the break; in the absence of a break we would expect to see the variance randomly vary across the sample. Consistent with the CART findings, there is also a local minimum at 68:1 but there are several other local minima evident in Figure 3a as well. Figure 3b reveals a notch shape for lagged specification (2), with a global minimum at 1980:3. This breakdate is also consistent with the root identified by the CART procedure for the lagged specification.[15]

[Insert Figures 3a-3b here.]

The next step is the application of the sequential test for structural change of Bai and Perron (1998), henceforth denoted BP, to find breakdates and their confidence intervals.[16] For a model with ℓ breaks, BP applies a test of the null hypothesis of no structural break vs. the alternative of a single break to each of the $\ell + 1$ mutually exclusive segments of data in the model. Beginning with a model of no breaks, the test is applied to each possible segment of data associated with a model containing an additional break. Model estimation includes a correction for serially correlated errors if warranted. Rejection of the null hypothesis of ℓ breaks in favor of a model with $\ell + 1$ breaks is warranted if the overall minimal value of the sum of squared residuals is sufficiently smaller with $\ell + 1$ breaks than the value obtained from the model with ℓ breaks. The break selected is the one associated with the overall minimum.

By contrast with the CART procedure, which grows the largest possible tree and then prunes it back based on out–of–sample predictive performance, the BP procedure *sequentially* tests for ℓ vs. $\ell + 1$ structural breaks (starting with $\ell = 0$) stops when the null hypothesis (of no break) cannot be rejected. Thus, while the CART procedure can be thought of as a *general–to–specific* methodology the sequential test is by way of contrast, a *specific–to–general* approach.

The result of applying the sequential procedure to Taylor rule specification (1) is that at either the 10%, 5% or 1% levels of significance, only three break dates are found. These three break dates, 1968:1, 1979:3 and 1986:4 correspond precisely to the root and second level splits identified using the CART procedure (see Figure 1). Table 3 presents the estimated coefficients of the four regimes. Note that the estimated coefficients for the first regime 1955:3–1968:1, are the same as for terminal node 1 of the CART procedure applied to Taylor rule specification (1), (see Table 1) since the subsamples over which the least squares regressions are conducted are exactly the same. Consistent with the CART findings, the results in Table 3 reveal that increasing weight is given to the inflation gap over the four regimes. Notice also that the weight on the output gap increases after 1968:1 but decreases after 1979:3. This evidence suggests that, while the Fed responded to the rising inflation of the 1960s, it continued to give weight to output gap deviations, perhaps to the detriment of successfully containing inflation. It is not until after the 1979:3 break that the estimated weight on the inflation gap ($\lambda_\pi = \beta_1 - 1$) exceeds that on the output gap.

An advantage of the Bai and Perron test over the CART procedure is that the former

Table 3: Coefficient Estimates and Standard Errors For the Sequential Split Method, Specification (1)

Terminal Node	Subsample Dates	Estimates of α	β_1	β_2	Adj. R^2	Nobs
1	1955:3–1968:1	3.309	-0.193	0.319	0.593	51
		(0.264)	(0.128)	(0.053)		
2	1968:2–1979:3	-0.307	**1.126**	0.692	0.659	46
		(1.000)	(0.162)	(0.097)		
3	1979:4–1986:4	5.029	**1.104**	**0.051**	0.798	29
		(0.681)	(0.440)	(0.110)		
4	1987:1–2003:3	1.414	1.770	0.746	0.827	67
		(0.365)	(0.112)	(0.065)		

Note: the regressions were performed with Newey–West standard errors to reflect first order autocorrelated disturbances and heteroskedasticy. Boldface entries indicate that a 95% confidence interval includes 1 for estimates of β_1 and 0 for estimates of β_2.

approach allows us to assign 95 percent confidence intervals to each break point. For the first break at 1968:1, this confidence interval is (1967:2,1968:4). For the second break at 1979:3, the confidence interval is (1978:4, 1980:1). Finally, for the third break at 1986:4, the confidence interval is (1984:2, 1987:2). Of the three breakdates, the 1979:3 date is the most tightly measured according to these confidence intervals, and this finding is consistent with the appearance of the 1979:3 date as the root split of the tree using the CART methodology. Other break-dates for specification (1) identified using the CART procedure, 1970:4, 1974:4, 1981:1 and 1998:3, do not emerge as break–dates from application of Bai and Perron's test; the sequential search procedure suggests that these further subsamples are statistically unwarranted.

Applying the BP sequential procedure to Taylor rule specification (2) we find evidence for four breaks at 1969:1, 1980:3, 1987:3 and 1996:3. Notice that two of these break dates, 1969:1 and 1980:3, correspond precisely to break dates identified from applying the CART procedure to Taylor rule specification (2). The 95 percent confidence intervals for the break

Table 4: Coefficient Estimates and Standard Errors For the Sequential Split Method, Lag Specification (2)

Terminal Node	Subsample: Dates	Estimates of α	β_1	β_2	ρ	Adj. R^2	Nobs
1	1955:3–1969:1	0.759	-0.118	0.593	0.772	0.904	54
		(0.220)	(0.269)	(0.138)	(0.063)		
2	1969:2–1980:3	-0.493	**1.488**	1.820	0.736	0.882	46
		(0.608)	(0.339)	(0.493)	(0.079)		
3	1980:4–1987:3	2.658	1.453	**0.167**	0.293	0.917	28
		(0.670)	(0.129)	(0.131)	(0.113)		
4	1987:4–1996:3	1.442	**0.694**	1.347	0.727	0.975	36
		(0.272)	(0.363)	(0.226)	(0.061)		
5	1996:4–2003:3	0.673	**-0.464**	1.604	0.835	0.958	28
		(0.327)	(1.51)	(0.523)	(0.089)		

Note: Boldface entries indicate that a 95% confidence interval includes 1 for estimates of β_1 and 0 for estimates of β_2.

dates identified using Bai and Perron procedure are as follows. For the first break at 1969:1 the confidence interval is (1967:1,1969:2); for the second at 1980:3 it is (1979:4, 1980:4); for the third break at 1987:3 it is (1987:1, 1988:3); and for the fourth break at 1996:4 it is (1995:3, 1997:1). The 1980:3 breakdate is the most tightly identified, and again, this finding is consistent with this break–date appearing as the root split of the tree using the CART methodology.

Comparing the results in Table 4 with those in Table 2, we see that the BP procedure identifies the period 1969:2–1980:3 as a single regime; the 1979:3 break–date, and the troublesome 1979:4–1980:3 subsample identified by the CART procedure, disappears. Further the BP procedure identifies two further breakdates following the 1980:3 break–date (1987:3 and 1996:3) while the CART procedure identifies only one (1984:3). Consequently, the coefficient estimates for these latter subsamples are quite different. In particular, for the BP–determined subsamples, the coefficient estimates for ρ are greater and the coefficient estimates for β_1 fall below unity following the 1987:3 break.

5 Conclusion

We have used two different, but complementary nonparametric methods in an effort to assess the structural stability of two versions of a Taylor rule specification for the U.S. federal funds rate over the period 1955:3–2000:2. This exercise is important in that many researchers working with Taylor rule specifications have simply asserted the existence of various different monetary regimes without asking whether the dates that define those regimes survive when subjected to tests for structural breaks. The nonparametric techniques we employ are relatively new and allow the *data* to determine whether there have been different regimes in monetary policy, and if so, how confident we can be in the dates that define those regime periods.

For both specifications of the Taylor rule, (1) or (2), application of the CART procedure or Bai and Perron's sequential test for structural breaks result in remarkably similar break dates. For specification (1), the three main break-dates using the CART procedure are the same dates identified using the sequential test. However, these three dates — 1968:1, 1979:3, 1986:4 – occur earlier in time than breakdates identified under (2), e.g., by the sequential test – 1969:1, 1980:3, 1987:3. The root or main break–dates identified by the CART procedure, 1979:3 for specification (1) and 1980:3 for specification (2), also happen to be the most tightly identified break-dates using the sequential search procedure.

The well–known start date of the Volker era, 1979:3, is frequently chosen as a breakdate in empirical analyses of U.S. monetary policy. Here, it emerges from application of two different nonparametric methods to Taylor rule specification (1). If monetary policy is well–captured by Taylor-type rules, and if it is believed that a major change in U.S. monetary policy began following the start of Paul Volker's term in the third quarter of 1979, then our results using nonparametric methods imply that the specification of the Taylor rule that would most reliably give rise to such a breakdate is the original specification proposed by Taylor (1993) and not the interest–rate smoothing/policy inertia formulation that is commonly used in practice. Indeed, a 95 percent confidence interval surrounding the 1980:3 breakdate identified for specification (2) does not include 1979:3. However, it does seem that either policy rule specification, augmented by one break in the late 1960s and by another one in late 1979 or 1980, can provide a very good approximation of monetary policy over the post-WWII period.

Importantly, our analysis rules out certain dates of historical significance as marking changes in monetary policy. In particular, we note that with the important exception of the start of the Volker chairmanship, the start and end dates of Fed chairmanships are not among the major break-dates for monetary policy identified by both procedures, nor are other dates often thought to be important, e.g., the abandonment of M1 targeting by the Fed in October, 1982.

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Notes

[1]See, e.g., Brunner (1994), Clarida et al. (2000), Feinman (1993), Taylor (1999) and Hamilton and Jorda (2002).

[2]Even if the break dates were known a priori, nonparametric methods can be used to validate the a priori breakdates.

[3]Without a simple and common model of monetary policy over the entire sample period, we would be unable to employ nonparametric methods as we would quickly confront the curse of dimensionality. A more complex approach to understanding U.S. monetary policy over the post-WWII period is pursued by Cogley and Sargent (2002, 2005). They estimate a reduced-form vector autoregression (VAR) model for macroeconomic variables with drifting coefficients, and following criticism by Sims (2002), they also allow for stochastic volatility in the VAR shocks. Using this model, they examine the extent to which changes in monetary policy, as characterized by a Taylor-type rule, should be attributed to changes in the preferences of policymakers or to changes in the parameters or variances of the macro data generating processes.

[4] Hendry (2000) reminds us of the difficulty in detecting breaks in time series using conventional methods. For this reason we use two different tests that are based on different assumptions to establish the robustness of our results.

[5]Our main empirical findings are robust to using the end-of-quarter federal funds rate target as well. As English et al. (2003) observe, the quarterly average rate is more consistent with the measures of inflation and the output gap which are based on activity over the quarter, and for this reason, we use the quarterly average target data.

[6]Clarida et al. refer to policy rules that satisfy these restrictions as "stabilizing" while those that don't are called "destabilizing" or "accommodative." For example if $\lambda_\pi < 0$, as reflected by a estimate of $\beta_1 < 1$, the Fed would be reacting to inflation in excess of its

target by reducing the nominal rate target (and therefore the real rate of interest), which could lead to further inflation.

[7]See, for example, Judd and Rudebusch (1998), Kozicki (1999), Clarida et al. (2000) Rudebusch (2002) and English et al. (2003). See also chapter 1 of Taylor (1999) for a description of the various calibrated specifications of the Taylor rule used in the papers collected in that conference volume.

[8] A variant of this lagged specification, which allows for autocorrelated errors, has been proposed by English et al. (2003).

[9]The CART procedure has been used in other economic applications by Durlauf and Johnson (1995), Cooper (1998) Kelly and ÓGráda (2000) and Johnson and Garcia (2000), but has not been applied to our task of searching for monetary policy regimes.

[10] We focus on the coefficient estimates on inflation and the output gap (β_1, β_2) rather than on the intercept, α, as it is difficult to disentangle whether changes in α are due to changes in the targets, r or $\overline{\pi}$. Kozicki and Tinsley (2003) suggest that much of the variation in the intercept coefficient might arise from large sustained shifts in $\overline{\pi}$ as variations in r appear to be small and transitory.

[11] For instance, this breakdate occurs after the September 1985 Plaza Accord initiated a multi-year devaluation of the dollar against major currencies and prior to the start of Alan Greenspan's term as Fed chairman in August 1987.

[12] See, e.g. Kozicki (1999) and Amato and Laubach (1999).

[13]This period of experimentation – in which the Fed switched from using the Federal Funds rate to using nonborrowed reserves as its operational target and set desired growth rates for the money aggregates M1 and M2 – is usually dated as lasting from 1979:3–1982:3 (Meulendyke (1998)), as the Fed did not change its operating procedures, in particular, it did not deemphasize growth rates for money aggregates, especially M1, until October 1982. The 1980:3 date also corresponds to the single break in the level of the ex ante real interest

rate identified by Antoncic (1986) using a linear time series model and monthly data; she identified the break in the level of the ex ante real interest rate as occurring at the end of October 1980.

[14]See, e.g. Taylor (1999) who refers to this period as the second worst for monetary policy in twentieth century U.S. history, with the Great Depression regime being the worst.

[15]Both these candidate breakdates and the CART estimated breakdates are the same by definition because they both select the break resulting in the minimum total sum of squared errors.

[16]We used the software of Bai and Perron, and thank Jushan Bai for making it available to us.

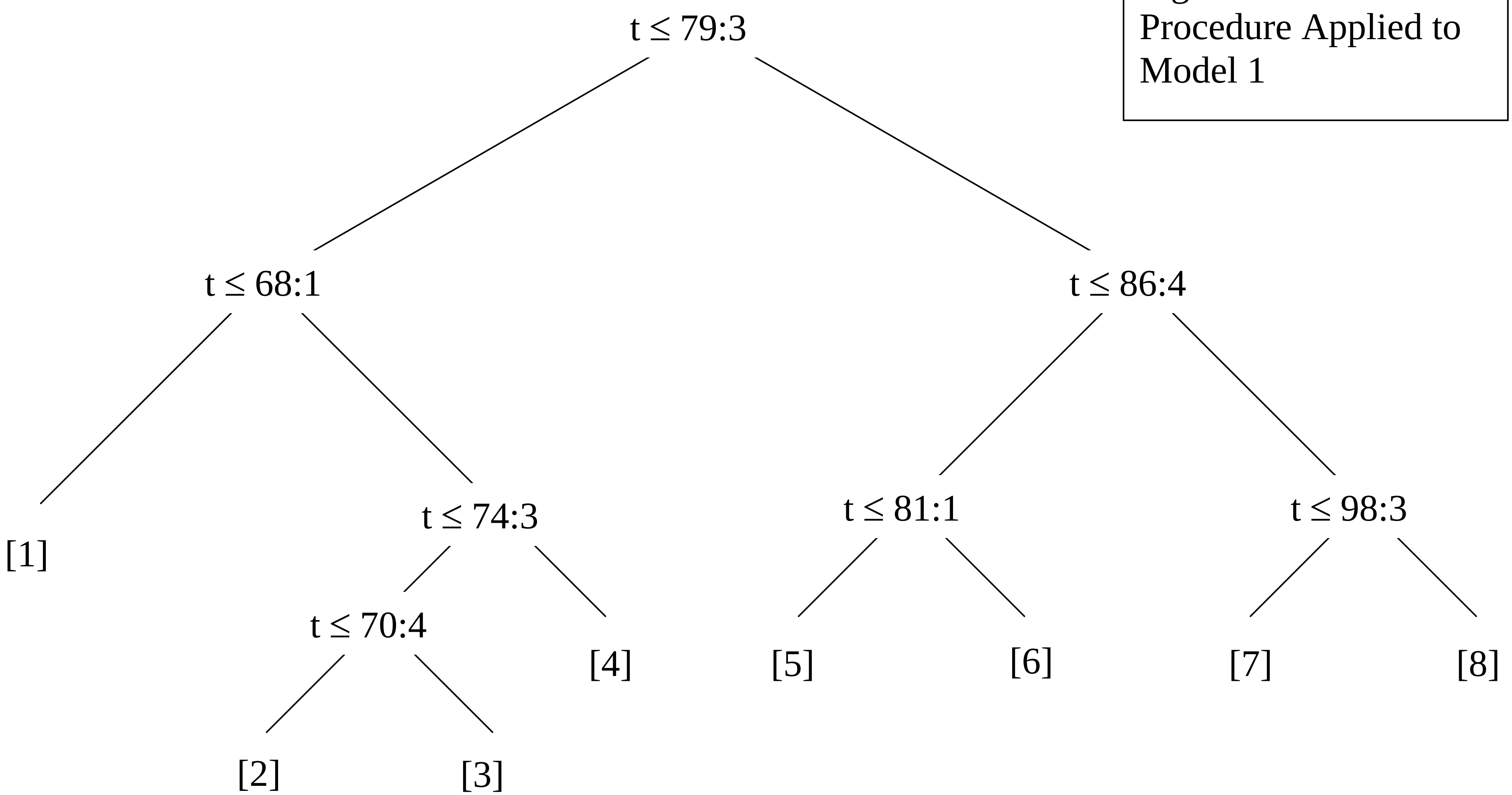

Figure 1: CART Procedure Applied to Model 1

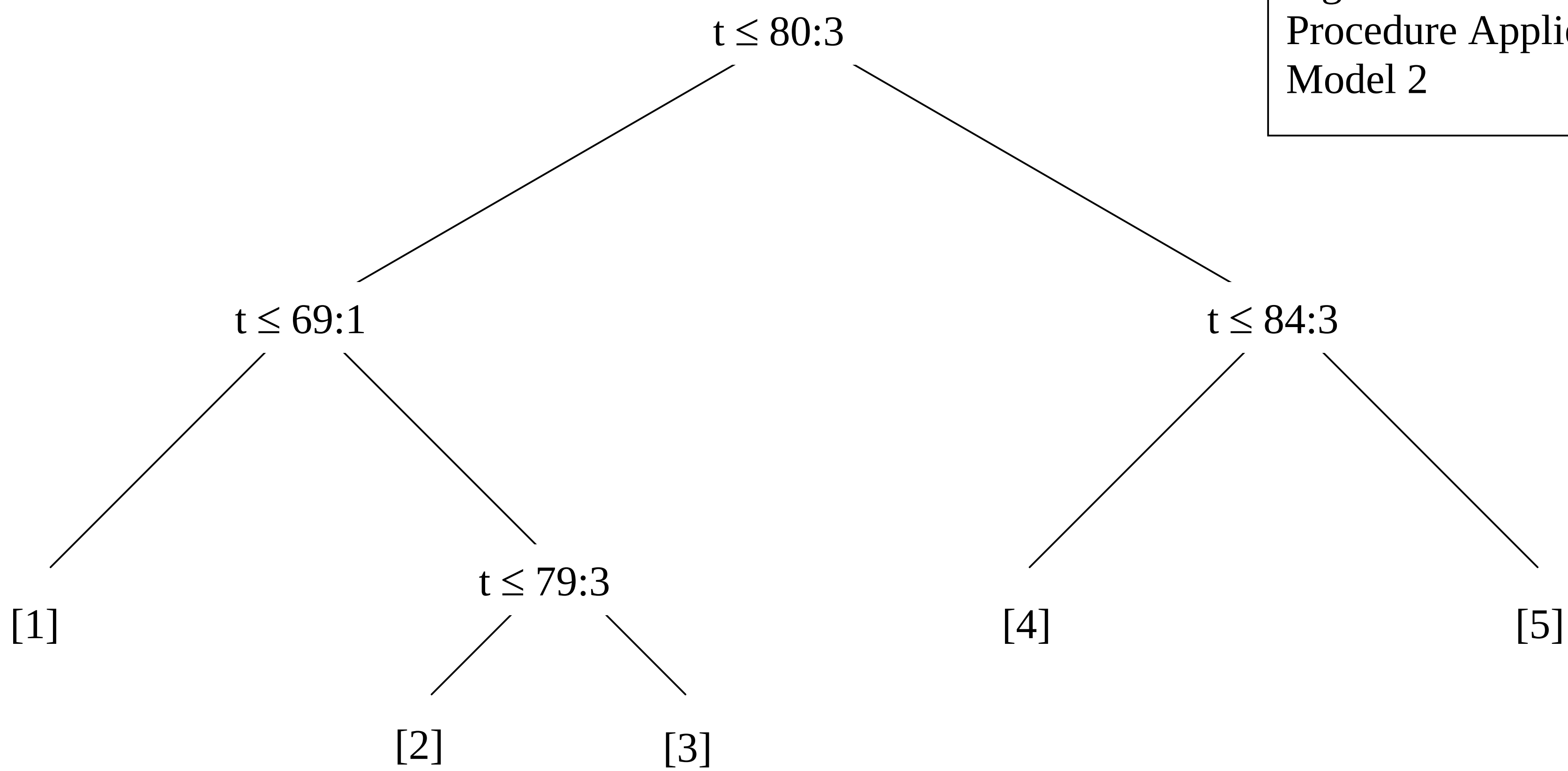

Figure 2: CART Procedure Applied to Model 2

Figure 3a: Estimated Variance for a Single Break Model (Specification 1)

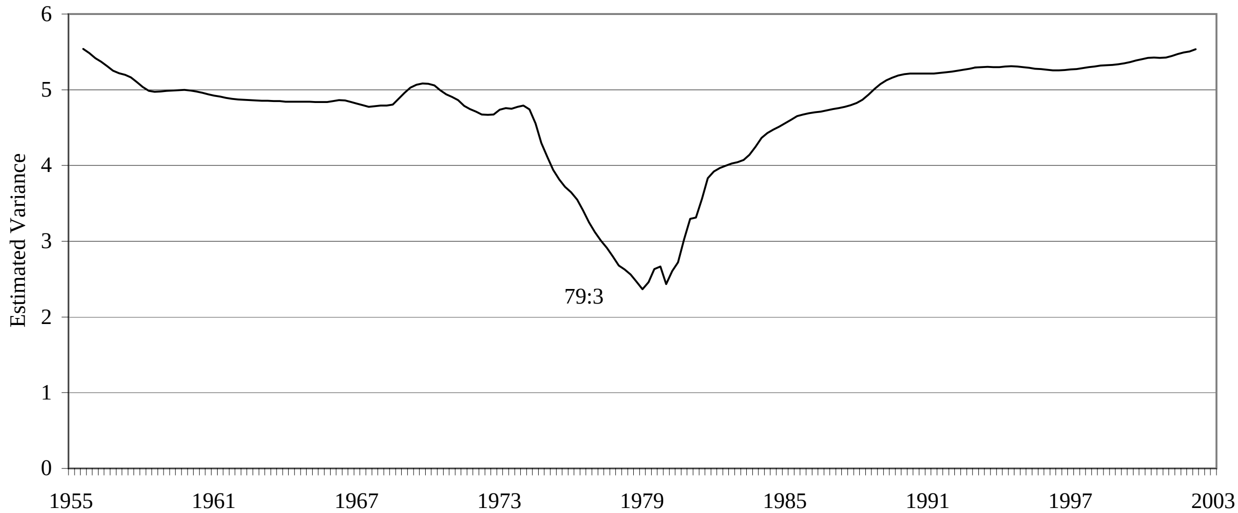

Figure 3b: Estimated Variance for a Single Break Model (Specification 2)