Structural Breaks in Inflation Dynamics

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(Very preliminary: comments welcome)

Abstract

Are there structural breaks in the dynamics of the inflation process? Does inflation possess a unit root? Is inflation a highly persistent process? We use tests for multiple structural breaks at unknown points in the sample, and a newly developed test for unit roots allowing for up to m structural breaks, to investigate breaks in inflation dynamics for 23 inflation series from 18 countries (plus the eurozone), and their implications for the serial correlation properties of inflation. All inflation series display structural breaks, often highly suggestive as they appear to broadly coincide with readily identifiable macroeconomic events, like the breakdown of Bretton Woods, the Volcker disinflation in the U.S., and the introduction of inflation targeting in several countries. Allowing for structural breaks, the null of a unit root can be strongly rejected for the vast majority of the series. Finally, evidence seems to suggest that, in general, inflation is not a highly persistent process.

We discuss the implications of our rejection of a unit root for Mishkin's explanation of time-variation in the extent of the Fisher effect. We argue that Mishkin's theory, based on the notion that inflation and interest rates are cointegrated, is difficult to defend in the light of our evidence against a unit root for almost all inflation series. The alternative Ibrahim and Williams (1978)-Barthold and Dougan (1986)-Barsky (1987) explanation, based on the notion of changes in the extent of inflation forecastability along the sample, is on the other hand compatible with our findings.

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1 Introduction

Are there structural breaks in the dynamics of the inflation process? Does inflation possess a unit root? Is inflation a highly persistent process? We use tests for multiple structural breaks at unknown points in the sample, and a newly developed test for unit roots allowing for up to m structural breaks, to investigate breaks in inflation dynamics for 23 inflation series from 18 countries (plus the eurozone), and their implications for the serial correlation properties of inflation[1]. We document structural breaks in all the series we analyse. For many countries/series, structural breaks appear to be clustered around the beginning of the 1970s (16 series for 14 countries), of the 1980s (14 series for 13 countries), and of the 1990s (14 series for 11 countries). Further, in several cases estimated break dates are highly suggestive, as they appear to broadly coincide with readily identifiable macroeconomic events, like the breakdown of Bretton Woods, the Volcker disinflation in the U.S.[2], and the introduction of inflation targeting in several countries. For the U.K., for example, estimated break dates are 1991:1 based on the CPI, and 1992:3 based on the GDP deflator[3]. Canada has an estimated break date in 1991:2 (based on the CPI) [4]. New Zealand, which adopted inflation targeting in February 1990, has a break date in 1989:4. Allowing for structural breaks, our new unit root test allows us to increase the number of rejections, compared to standard Dickey-Fuller tests. Finally, conditional on the estimated breaks, inflation series exhibit, in general, little persistence, with the exception of a few countries–for example, the U.S. and the U.K.–around the time of the Great Inflation. [Such a conclusion, however, has to be considered for the time being as tentative. We need more statistical tests on this.]

We discuss an implication of our findings for the Fisher effect. We argue that Mishkin's explanation for the well-known, puzzling time variation in the extent of the Fisher effect seen in the data, based on the notion that inflation and nominal interest rates are cointegrated, is difficult to defent in the light of our rejection of a unit root for the vast majority of inflation series. The alternative Ibrahim and Williams (1978)-Barthold and Dougan (1986)-Barsky (1987) explanation, based on

[1] A related recent paper is Levin and Piger (2002), who investigate inflation dynamics in 12 industrial countries over the period 1983-2001 by means of both classical and Bayesian methods. For all series, they find strong evidence of a single structural break in both the intercept and the innovation variance, but no evidence of a break in the autoregressive coefficients. Conditional on the identified breaks in the intercept and innovation variance, all inflation series exhibits very little persistence. As we discuss more extensively below, their results are broadly similar to ours.

[2] Estimated break dates are 1981:4 based on the CPI, and 1981:2 based on the GNP deflator (see tables 3 and 18)

[3] In a broader investigation of changes in the stochastic properties of U.K. macroeconomic time series over the last decade, compared to the previous post-WWII era, Benati and Talbot (2002) detect similar breaks in the rate of growth of several national accounts deflators around the period of the adoption of an inflation targeting regime, in November 1992.

[4] On changes in the dynamics of Canadian inflation around the time of the adoption of inflation targeting, in 1991, see also Ravenna (2000).

the notion of changes in the extent of inflation forecastability along the sample, is on the other hand compatible with our findings, and is explored in related work in progress[5].

The paper is organised as follows. The next section describes our dataset. Section 3 reports results from tests for multiple structural breaks at unknown points in the sample, based on the Andrews (1993), Andrews and Ploberger (1994), Bai (1994), and Bai (1997) methodology. Section 4 describes our new test for a unit root allowing for up to m structural breaks, and illustrates the results. Section 5 discusses implications of our findings for the Fisher effect. Section 6 concludes.

2 The data

Our dataset contains 23 inflation series from 18 countries, with markedly different sample periods[6]. For the U.K., both the GDP deflator and the CPI[7] are from the *Office of National Statistics*. The sample periods are 1955:1-2002:2 and, respectively, 1947:1-2002:3. For the U.S., the CPI is from *U.S. Department of Labor, Bureau of Labor Statistics*, and is available from 1947:1 to 2002:3. The GNP deflator series is from Balke and Gordon (1986) for the period 1875:1-1950:1, and from *U.S. Department of Commerce, Bureau of Economic Analysis*, for the period 1950:2-2001:1. For New Zealand, the CPI is from *Statistics New Zealand*, and is available from 1925:3 to 2002:2. For Sweden, Australia, Finland, Norway, Portugal, Spain, and Ireland the CPI is from the OECD database, and is available from 1962:1 to 2001:2. For France and Australia the GDP deflator is from the OECD database, and is available from 1962:1 to 2001:2. For Canada, the CPI is from *Statistics Canada*'s website on the internet. The sample period we consider is 1947:1-2002:2. For Germany, the CPI, available from 1950:1 to 2002:2 is from the *Bundesbank* monthly bullettin. [Here investigate the story of the reunification: it is not clear how it is taken care of in this series. If we can't find what they exactly did, we have to drop these results]. The French CPI, available from 1951:1 to 2001:2, is from *INSEE*. For Italy, the CPI excluding tobacco items is from *ISTAT*, and is available from 1947:1 to 2002:3, but we only consider the period 1948:1-2002:3 to prevent our results from being distorted by the high inflation and subsequent stabilisation of 1947. The Swiss CPI, available from 1947:1 to 2002:2, is from *Ufficio Federale di Statistica*. For Belgium, the CPI, available from 1947:1 to 2002:2, is from the Belgian central bank. For the Netherlands, the CPI is from *CBS*, the Dutch statistical office, and is available from 1945:4 to 2002:3. For Austria, the CPI is from *Statistik Austria*, and the sample period is

[5] Benati and Kapetanios (2002)

[6] We thank Graham Howard of the Reserve Bank of New Zealand, Peter Stadler of the Swiss central bank, Raf Wouters of the Belgian central bank, Cees Ullersma of the Dutch central bank, and Fabio Rumler of the Austrian central bank for kindly providing data for their respective countries.

[7] More precisely, RPIX, the consumer price index excluding mortage interest payments, which is the price index targeted by the Bank of England under the current monetary framework.

1950:1-2002:2. For the eurozone, both the GDP deflator and the HICP are from the ECB website. The sample period is 1970:1-1998:4 for the HICP, and 1970:1-1998:3 for the GDP deflator. Only six series–Belgium's CPI, the U.S. GNP deflator, the GDP deflators for France, the U.K., and the eurozone, and the U.S. CPI–are seasonally adjusted. For all the other series we use seasonal dummies. On the other hand, in the present version of the paper we do not use dummies in order to control for specific events like the Nixon price controls in the U.S., or the introduction of the poll tax in the U.K. in April 1990. Although we plan to do this at a later stage, we regard as extremely unlikely that our results may have been significantly distorted by our lack of controlling for such one-off events. [Or we may even drop the whole thing: Fuhrer and Moore, Nelson, and Mankiw and Reis don't do it, so for compatibility of the results we may just as well drop it]

Finally, the commercial paper rate series used in section 5.2, available for the period January 1857-present, has been constructed by linking the commercial paper rate series from the *NBER Historical Database* (commercial paper rates, New York City[8]; NBER series: 13002), available for the period January 1857-December 1971, to the commercial paper rate series from the Federal Reserve database (3-month prime commercial paper rate, averages of daily figures; acronym: CP3M), available for the period April 1971-present.

3 Results from Tests for Multiple Structural Breaks at Unknown Points in the Sample

In this section we use the methods of Andrews (1993), Andrews and Ploberger (1994), Bai (1994), and Bai (1997) to test for multiple structural breaks at unknown points in the sample in univariate representations for inflation for the 23 inflation series in our dataset. For each series we estimate the following AR(k) model:

$$y_t = \mu + \phi_1 y_{t-1} + \phi_2 y_{t-2} + ... + \phi_k y_{t-k} + \epsilon_t \tag{1}$$

via OLS, and we test for a structural break at an unknown point in the sample in the intercept, the AR coefficients, and the innovation variance. Following Andrews (1993), we assume that the break did not occur in either the first, or that last 15%

[8] Sources: for the period January 1857-January 1937: Macaulay (1938), pp. A142-161. For the period February 1937-1942: computed by the NBER based on weekly data in *Bank and Quotation Record, Commercial and Financial Chronicle*. For the period 1943-1971: Federal Reserve Board. Data represent 60-90 day prime endorsed bills for 1858-1859; prime 60-90 day double name for 1860-1923; prime four-six months, double and single names thereafter. Data for 1857 are from rates given in a Treasury report in bankers' magazine; rates for 1858 are from the New York Chamber of Commerce Report, 1858, p. 9; rates for 1859-June 1862 are arithmetic averages between the monthly averages of *Hunt's Merchants* magazine and of Bankers'; rates for July 1862-1865 are estimated from a table of daily rates from different New York newspapers. Data for 1942-1971 are averages of daily offerings rates of dealers 60-90 day prime bills.

of the sample. For each possible breakdate we compute the relevant Wald[9] statistic, and we compare the maximum Wald statistic with the 10% asymptotic[10] critical values tabulated in Andrews (1993). If the null of no structural break is rejected, we proceed to estimate the breakdate by minimizing the residual sum of squares. The sample is then split in correspondence to the estimated breakdate, and the same procedure is repeated for each subsample. If the null of no structural break is not rejected for either subsample, the procedure is terminated. Otherwise, we estimate the new breakdate(s), we split the relevant subsample(s) in correspondence to the estimated breakdate(s), and we proceed to test for structural breaks for hierarchically obtained subsamples[11]. The procedure goes on until, for each hierarchically obtained subsample, the null of no structural break is not rejected at the 10% level. Following van Dijk, Osborne, and Sensier (2002), throughout the whole procedure we impose that at least 15% of the sample lies between two consecutively identified breakdates. After estimating the number of breaks, and getting preliminary estimates of the breakdates, each breakdate is re-estimated according to the modification of the Bai (1997) 'refinement' procedure proposed by van Dijk, Osborne, and Sensier (2002)[12]. Finally, we estimate the model conditional on the identified breakdates. Throughout the whole process, the lag order for each model is chosen based on the Schwartz information criterion[13]. Although, at the stage of making the choice between rejecting or accepting the null of no structural break for a specific (sub)sample, we uniquely focus on the *sup*-Wald statistic, we also report, for each estimated breakdate, Andrews and Ploberger (1994)'s average and exponential Wald statistics, defined as:

[9] A key reason for focusing on the *sup*-, *ave*-, and *exp*-Wald versions of the Andrews (1993) and Andrews and Ploberger (1994) tests is that, as conjectured by Boivin (1999), the -Wald versions of these tests, compared to the likelihood ratio and to the Lagrange multiplier versions, may exhibit more power. Cogley and Sargent (2002b) provide preliminary corroborating evidence that this may indeed be the case for one specific alternative, that of random-coefficients with stochastic volatility.

[10] [Here discuss the issue of small-sample critical values, and why we don't even try to compute them (just a computational nightmare). Stress that this is what people do when they have lots of series. McConnell and Peres-Quiros have only one series, so they can do it.]

[11] Here we follow Bai (1997) in estimating multiple structural breaks one at a time.

[12] Specifically, each of the n estimated break dates is re-estimated conditional on the remaining n-1 break dates. In implementing the van Dijk, Osborne, and Sensier (2002) modification of the Bai (1997) procedure, we adopt the following iterative approach. We start by taking the first-stage estimated break dates as our initial conditions. Then, we re-estimate each break date conditional on the remaining n-1 break dates. These re-estimated break dates then become the initial conditions for the next iteration, and so on. The procedure is terminated when, from one iteration to the next, there is no difference in estimated break dates, so that we have reached a sort of 'econometric Nash equilibrium'.

[13] In particular, SIC is applied to the model estimated over the whole sample, conditional on the identified breakdates. Ideally, we would like to apply SIC to each identified sub-sample, therefore allowing each subsample to have a different lag order. Such a strategy, however, presents the drawback of dramatically complicating the econometrics, given that within such an approach the problem of selecting the lag order for each identified subsample becomes inextricably intertwined with the issue of testing for structural breaks.

$$Ave - Wald = \frac{1}{(N_2 - N_1 + 1)} \sum_{t=N_1}^{N_2} Wald\,(t) \tag{2}$$

$$Exp - Wald = \ln \left\{ \frac{1}{(N_2 - N_1 + 1)} \sum_{t=N_1}^{N_2} \exp\left[Wald\,(t)\right] \right\} \tag{3}$$

For each estimated breakdate, and for both the *sup*-, the *ave*-, and the *exp*-Wald statistics, we report approximated asymptotic critical values computed according to Hansen (1997). Finally, for each identified subsample, we report the estimated unconditional mean of the process, the sum of the AR coefficients, and the estimated innovation variance, together with their estimated standard errors. Estimated standard errors for the unconditional mean—a non-linear function of the estimated parameters—have been computed according to the delta method described, for example, in Campbell, Lo, and MacKinlay (1997).

As we discuss more extensively in what follows, for all the series we detect evidence of multiple structural breaks. A rejection of the joint hypothesis of constancy in the intercept, the AR coefficients, and the innovation variance, however, could in principle be due to a break uniquely in the intercept, uniquely in the innovation variance, and so on. Unfortunately, understanding what exactly is driving the strong rejctions we obtain is, in general, not easy[14]. A first possibility is to run separate tests for structural breaks for the three sets of coefficients, under the assumption that the remaining coefficients do not experience any break. Although in what follows we report results from such tests for the AR coefficients, the intercept, and the innovation variance taken separately, these results should be considered with extreme caution. As shown for example by Hansen (1992) in the context of the Nyblom-Hansen test, structural break tests for individual (sets of) coefficients may have a very low power when the remaining coefficients, whose stability is not being tested and is instead assumed, may in fact be subject to breaks as well. A second possibility could to use a sequential procedure, starting (say) by testing for structural breaks for the innovation variance (the parameter for which, as we discuss in what follows, based on individual break tests we find overwhelming evidence of instability) under the assumption of no breaks in either the intercept or the AR coefficients. After estimating the break dates for the innovation variance, a second-stage test for structural breaks in either the mean or the AR coefficients will not follow the Andrews (1993) anymore, due to the distortion induced by the first-stage testing, but critical values could be computed via a bootstrap procedure.

Tables 1-23 illustrate the results for the 23 inflation series in our dataset, while figures 1-9 show, for some selected series, estimates for the unconditional mean of the process, the innovation variance, and the sum of the AR coefficients for each identified

[14]We thank Ken West for extremely helpful discussions on this issue. The problem could be easily solved only under the unrealistic assumption that the log-likelihood is block-diagonal in the three sets of parameters. Ruling out such an assumption, there seems to be no easy solution.

sub-sample, together with 90% confidence bands. It is important to stress that, as a consequence of the first-stage test, standard errors (reported in the tables in parentheses) *cannot* be use to perform tests of equality/inequality of parameters *across* sub-samples, and are only valid *within* sub-samples. This implies, for example, that the confidence bands are only valid within a specific sub-sample, while the indication they give of parameter equality/inequality across sub-samples may be misleading. For all the series we detect evidence of multiple structural breaks at the 10% level, but the Hansen (1997) p-values are, in the vast majority of cases, extremely small, thus indicating very strong identification of the break dates. Specifically, for five series we identify two breaks; for twelve series we identify three breaks; and for the remaining nine series we identify four breaks. The vast majority of break dates appears to be clustered around the beginning of the 1970s (16 series, for 14 countries[15], have a break between 1969:3 and 1973:3[16]), of the 1980s (14 series, for 10 countries[17] plus the eurozone, have a break between 1980:1 and 1983:4), and of the 1990s (14 series, for 10 countries[18] plus the eurozone, have a break between 1990:1 and 1993:4[19]). Although the interpretation of such purely statistical evidence is clearly contentious, the concentration of so many break dates, for so many countries, around the time of the collapse of the Bretton Woods regime is highly suggestive. For all the countries and the series in this group, the unconditional mean of the process is estimated to have increased. In some cases the increase is particularly marked. For New Zealand, for example, the unconditititional mean jumps from 0.026 to 0.12. For the U.K., the increase is from 0.044 to 0.144, based on the CPI, and from 0.063 to 0.165 based on the GDP deflator. For the U.S., based on the CPI, it is from 0.045 to 0.116. For Italy, based on the CPI, it is from 0.088 to 0.174. For Germany and Switzerland, on the other hand, the increase is much milder. Based in both cases on the CPI, the estimated unconditional mean increases from 0.025 to 0.052 and, respectively, from 0.035 to 0.048. Second, for several countries/series in this group (but not for all) we estimate a significant increase in the innovation variance. For the U.K., the estimated standard deviation of the innovation increases from 0.0279 to 0.0729, based on the GDP deflator, and from 0.0198 to 0.0664, based on the CPI. For Australia, based on the CPI, it increases from 0.0207 to 0.0603. For Italy, based on the CPI, it increases from 0.0137 to 0.0582. For Portugal, based on the CPI, it increases from 0.0518 to 0.1852. Finally, as for the sum of the AR coefficients, we estimate relatively

[15] New Zealand, Sweden, Australia, Germany, France, Italy, Switzerland, the U.K., Belgium, Norway, Portugal, Spain, Ireland, and the U.S..

[16] We have chosen a four-year interval centered in 1971:3, the quarter of the collapse of Bretton Woods.

[17] Canada, Germany, France, Italy, the U.S., the U.K., Switzerland, Norway, Spain, and Ireland.

[18] Sweden, Australia, Canada, Germany, France, the Netherlands, the U.K., the U.S., Switzerland, and Spain.

[19] Our results are therefore broadly in line with those of Levin and Piger (2002), who, for all the series in their sample, detect evidence of a single structural break in the mean and innovation variance at the beginning of the 1990s.

small changes for all countries and series, with the only exception of Germany, Italy, Switzerland, and Portugal (all based on the CPI).

Tentatively interpreting the second clustering of break dates is less straightforward, but the beginning of the 1980s is the period in which central banks around the industrialised world decisively shifted their policies towards inflation-fighting. All the series displaying a break around this period, indeed, show a decrease in their estimated unconditional mean, sometimes–like in the case of the U.S. and Italian CPI, and of the U.K. GDP deflator–particularly marked. Finally, interpreting the third clustering appears as even more difficult. Four countries which exhibit breaks at the very beginning of the 1990s–the U.K., Sweden, Australia, and Canada–adopted around those years inflation targeting regimes. For the other countries–Germany, France, the Netherlands, Switzerland, Spain, the U.S., and the eurozone considered as a whole–the interpretation is not clear at all, with the possible exception of Germany, which in those years experienced the reunification shock.

[here discuss results for individual countries, and the issue of persistence]

Let's now consider results from structural breaks tests for individual sets of parameters[20]. For each series, we estimated an AR(k) model by OLS, selecting the lag order based on SIC, and we started by performing three Andrews (1993) and Andrews and Ploberger (1994) tests for structural breaks in the intercept, in the innovation variance, and in the AR coefficients considered as a whole, under the assumption that the sets of parameters which were not being tested for breaks remained constant along the sample. Tests for the stability of the innovation variance are remarkably uniform in detecting strong evidence of structural breaks. Based on the *sup*-Wald statistic, for only one series (Swedish CPI inflation) we cannot reject the null of stability, with p-values being, in most cases, extremely small. Based on the *ave*- and the *sup*-Wald statistics[21],on the other hand, we can reject the null of stability for all series except the Swedish and Australian CPI, and the eurozone GDP deflator and, respectively, the Swedish, Australian, Canadian CPI, and the eurozone's HICP inflation. Tests for the stability of the intercept[22], on the other hand, are even more uniform in *not* rejecting the null of stability: based on the *ave*- and the *exp*-Wald statistics, not in a single instance we reject stability at the 10% level, while based on the *sup*-Wald test, we reject stability in only four cases (the GDP deflator and HICP for the eurozone, and the CPI for Austria and Finland). [here explain why this is not surprising, given the low power of these individual tests] Finally, as for the autoregressive coefficients considered as a whole, based on the *ave*- and the *exp*-Wald statistics only in one case we can reject the null of no break (in both cases for Belgium's CPI), while based on

[20] The *sup*-, *ave*-, and *exp*-Wald statistics for testing the stability of the innovation variance are not reported here, but are available upon rquest.

[21] The *sup*-, *ave*-, and *exp*-Wald statistics for testing the stability of the innovation variance are not reported here, but are available upon rquest.

[22] Both for the intercept, and for the AR coefficients considered as a whole, test statistics have been computed with a Newey-West correction for the covariance matrix.

the *sup*-Wald statistic we can reject in 17 cases. The series for which, based on the *sup*-Wald statistic, we have no rejection are the CPI for the Netherlands, Norway, the U.S., Australia, Germany, France, and Italy, and the HICP for the eurozone.

Second, we performed Nyblom-Hansen tests[23] for stability in the intercept, in the innovation variance, and in the AR coefficients considered as a whole, once again based on the previously estimated AR(k) model for each series (with the lag order selected based on SIC)[24]. Results are reported in Table 24. Stability in the variance was not rejected, even at the 90%, only for 6 series. For five series the rejection was at the 5% level, while for all the remaining series stability in the innovation variance was rejected at least at the 1% level. A significant difference compared to the single-parameter Andrews (1993) and Andrews and Ploberger (1994) tests is given by the intercept. For 6 series, the rejection is at the 5% level; and for 5 series, the rejection is at the 10% level. [here try to make sense of this difference with the Andrews test] As for the sum of the AR coefficients [here write the program and discuss the results]

4 Results from Unit Root Tests Allowing for Up to m Structural Breaks

Does inflation possess a unit root? Nelson and Plosser (1982) was extremely influential in establishing the conventional wisdom notion of inflation as a highly persistent process, possibly possessing a unit root. In the light of our previous evidence in favor of multiple structural breaks in all the inflation series we analyse, however, the notion that inflation may possess a unit root should be seen with suspicion. As first discussed by Perron (1990), failure on the part of a researcher to control for possible structural breaks in the unconditional mean of a process will spuriously increase its estimated extent of persistence. In the limit, even a white noise process with an unconditional mean shifting according to (say) a Markov-switching process will look very much like a unit root process.

In this section we therefore proceed to re-examine the evidence in favor of a unit root in inflation based on a newly developed test for unit roots allowing fo up to m structural breaks. The test we propose follows from the sequential DF t-statistics proposed by Banerjee, Lumsdaine, and Stock (1992) and Zivot and Andrews (1992) for the case of a single break. The following model forms the basis of our investigation.

$$y_t = \mu_0 + \mu_1 t + \alpha y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \sum_{i=1}^{m} \phi_i DU_{i,t} + \sum_{i=1}^{m} \psi_i DT_{i,t} + \epsilon_t \tag{4}$$

[23] Specifically, we implemented the Lagrange multiplier version of the Nyblom-Hansen test as described in Hansen (1992).

[24] The Nyblom-Hansen procedure tests the null hypothesis that the parameter(s) of interest are constant, against the alternative that they follow a martingale. As discussed for example by Nyblom (1989), such an alternative comprises a number of cases of interest, among them random-coefficients, and a one-time shift at an unknown point in the sample.

$1-\gamma(L)$ has all its roots outside the unit circle, where $\gamma(L)=\gamma_1 L+\ldots+\gamma_k L^k$. We denote the probability limit of the estimated covariance matrix of the vector $(\Delta y_{t-1},\ldots,\Delta y_{t-k})$ by Σ. $DU_{i,t}$ and $DT_{i,t}$ are intercept and trend break dummy variables respectively defined by :

$$DU_{i,t}=1(t>T_{b,i}),\quad DT_{i,t}=1(t>T_{b,i})(t-T_{b,i})$$

where $T_{b,i}+1$ denotes the date of the i-th structural break and 1(.) is the indicator function taking the value of 1 if the argument of the function is true and 0 otherwise.

To facilitate the analysis we follow Banerjee, Lumsdaine, and Stock (1992) and Lumsdaine and Papell (1997) and define the following vector of regressors:

$$\boldsymbol{z}_t=(1,t+1,y_t-\bar{\mu}t,DU_{1,t+1},\ldots,DU_{m,t+1},DT_{1,t+1},\ldots,DT_{m,t+1},\Delta y_t-\bar{\mu},\ldots,\Delta y_{t-k+1}-\bar{\mu})'$$

where $\bar{\mu}=E(\Delta y_t)$. Then, $y_t=\boldsymbol{z}_{t-1}'\boldsymbol{\theta}$ where

$$\boldsymbol{\theta}=(\mu_0+(\gamma(1)-\alpha)\bar{\mu},\mu_1+\alpha\bar{\mu},\alpha,\phi_1,\ldots,\phi_m,\psi_1,\ldots,\psi_m,\gamma_1,\ldots,\gamma_k)'$$

. The sequence of errors is assumed to be a martingale difference sequence with finite conditional $4+\xi$, $\xi>0$, moments. The second conditional moment is denoted by σ^2. Denoting the number of observations for model (4) by T, we rewrite the break dates as $T\delta_1,\ldots,T\delta_m$ where $0<\delta_i<1$, $i=1,\ldots,m$ are the break fractions. We also define the scaling matrix

$$\Xi_T=diag(T^{1/2},T^{3/2},T,\underbrace{T^{1/2},\ldots,T^{1/2}}_{m}\underbrace{T^{3/2},\ldots,T^{3/2}}_{m},\underbrace{T^{1/2},\ldots,T^{1/2}}_{k})$$

partitioned conformably to $\boldsymbol{z}_t$. We define the OLS estimator for model (4) and given break dates as

$$\hat{\theta}(\delta_1,\ldots,\delta_m)=\Psi_T(\delta_1,\ldots,\delta_m)^{-1}\boldsymbol{\zeta}_T(\delta_1,\ldots,\delta_m)$$

where $\boldsymbol{\zeta}_T(\delta_1,\ldots,\delta_m)=\Xi_T^{-1}\sum_{i=1}^{T}\boldsymbol{z}_{t-1}(\delta_1,\ldots,\delta_m)y_t$ and

$$\boldsymbol{\Psi}_T(\delta_1,\ldots,\delta_m)=\Xi_T^{-1}\sum_{i=1}^{T}\boldsymbol{z}_{t-1}(\delta_1,\ldots,\delta_m)\boldsymbol{z}_{t-1}(\delta_1,\ldots,\delta_m)'\Xi_T^{-1}$$

We also define $\boldsymbol{\varphi}_T(\delta_1,\ldots,\delta_m)=\Xi_T^{-1}\sum_{i=1}^{T}\boldsymbol{z}_{t-1}(\delta_1,\ldots,\delta_m)\epsilon_t$

In order to construct our test we define the following alternative hypotheses:

$$H_i:\alpha<1,\phi_{i+1}=\ldots=\phi_m=\psi_{i+1}=\ldots=\psi_m=0,\quad i=1,\ldots,m-1$$

$$H_m:\alpha<1$$

As usual, we denote the null hypothesis $\alpha = 1, \mu_1 = \phi_1 = \ldots = \phi_m = \psi_1 = \ldots = \psi_m = 0$ by H_0. Clearly, previous testing procedures concentrated on testing H_0 against H_1 or H_2. Our aim is to construct a test of H_0 against $\cup_{i=1}^{m} H_i$. The most straightforward method involves constructing the relevant t-statistics on the estimate of α for all possible break partitions for a given break number and all break numbers from 1 to m and taking the infimum of the set of these t-test statistics. Let us denote the set of all possible break partitions for a given number of breaks by $\mathcal{T}_i$, $i = 1, \ldots, m$ and their union over i by $\mathcal{T}$. The distribution under the null hypothesis for a t-test statistic given the number of breaks and the break fractions follows from Proposition 1 of Kapetanios (2002) and Remark 1 of Lumsdaine and Papell (1997). The distribution of the infimum of the t-test statistics, over $\mathcal{T}$, under the null hypothesis follows directly from Lemma A.4 of Zivot and Andrews (1992). The consistency of the test is guaranteed by the consistent estimation of the break fractions and the other coefficients under the alternative of structural breaks proven by, among others, Bai and Perron (1998). Note that the results of Bai and Perron (1998) concerning consistency of the estimated coefficients allows for deterministic trends. Nevertheless, such an approach is unnecessarily computationally intensive[25]. By Bai and Perron (1998, pp. 64) we have that a sequential procedure would allow consistent estimation of break fractions, and therefore consistent estimation of the whole model under the alternative hypothesis, with only $O(T)$ least squares operations for any given number of breaks. We can therefore construct a consistent and less computationally intensive test using the t-statistics from these least squares operations.

We therefore propose constructing a test using the following grid search scheme following Bai and Perron (1998).

1. For a given maximum number of breaks, m, start by searching for a single break and store the t-statistics of the hypothesis $\alpha = 1$ for all possible partitions over the sample. Denote the set of all possible partitions as $\mathcal{T}_1^a$. Also, denote the set of t-test statistics by $\boldsymbol{\tau}^1$.

2. Choose the break date associated with the minimum sum of squared residuals (SSR) given by

$$SSR = \sum_{t=k+2}^{T} (y_t - \hat{\mu}_0 - \hat{\mu}_1 t + \hat{\alpha} y_{t-1} + \sum_{i=1}^{k} \hat{\gamma}_i \Delta y_{t-i} + \hat{\phi}_1 DU_{1,t} + \hat{\psi}_1 DT_{1,t})^2$$

   where k is assumed known.

3. Imposing the estimated break date on the sample, start looking for the next break over all possible partitions in the resulting subsamples. Denote the set

[25] An alternative procedure to estimate multiple breaks with reduced computational burden has recently been suggested by Bai and Perron (2000). This procedure could be used instead of the sequential procedure we suggest in this context.

of all possible partitions by $\mathcal{T}_2^a$. Obtain the set of t-statistics of the hypothesis $\alpha = 1$ over all possible partitions and denote this by $\boldsymbol{\tau}^2$. Append $\boldsymbol{\tau}^2$ to $\boldsymbol{\tau}^1$ to obtain $\boldsymbol{\tau}_1^2 = \boldsymbol{\tau}^1 \cup \boldsymbol{\tau}^2$.

4. Choose the break with the minimum SSR as the next estimated break.

5. Repeat steps 3 and 4 until m break dates have been estimated. Denote the resulting sets of all possible partitions as $\mathcal{T}_i^a$, $i = 3, \ldots, m$.

6. Adopt as the test statistic, $\tau_{\min}^m$, the minimum t-statistic over the set $\boldsymbol{\tau}_1^m = \boldsymbol{\tau}^1 \cup \boldsymbol{\tau}^2 \cup \ldots \cup \boldsymbol{\tau}^m$.

Before we discuss the asymptotic distribution of this test statistic we note that we do not look for consecutive breaks or for breaks at the end or beginning of the sample. Each estimated break is assumed to lie between two subsamples whose size goes to infinity with rate T as the sample size increases. In other words we impose a nonzero trimming parameter, ε on each break search. Under the null hypothesis of a unit root, the test statistic will have a well defined distribution which will be the same as that of the minimum of $\int_0^1 W_i^*(\hat{\boldsymbol{\delta}}_i, r)dW(r) / \left(\int_0^1 W_i^*(\hat{\boldsymbol{\delta}}_i, r)dr\right)^{1/2}$ over $\hat{\boldsymbol{\delta}}_i$ where $\hat{\boldsymbol{\delta}}_1 = \hat{\delta}_1$, $\hat{\boldsymbol{\delta}}_i = (\hat{\delta}_1, \ldots, \hat{\delta}_{i-1}, \delta_i)$, $i = 2, \ldots, m$ and $W_i^*(\boldsymbol{\delta}_i, r)$, $\boldsymbol{\delta}_1 = \delta_1$, $\boldsymbol{\delta}_i = (\delta_1, \ldots, \delta_i)$, $i = 2, \ldots, m$, is the continuous time residual from the projection of a Brownian motion onto the functions $[1, r, 1(r > \delta_1), (r - \delta_1)1(r > \delta_1), \ldots, 1(r > \delta_i), (r - \delta_i)1(r > \delta_i)]$. Note that in $\hat{\boldsymbol{\delta}}_i$ the only parameter that varies with the minimization is δ_i. The rest of the break fractions are given and have been estimated from previous SSR minimisations. This distribution merits further discussion. We firstly note that obviously the set over which we take the infimum, $\mathcal{T}^a \equiv \cup_{i=1}^m \mathcal{T}_i^a$, is a subset of the set $\mathcal{T}$, over which the infimum would have been taken had we simply extended the method used by Lumsdaine and Papell (1997) to more than two breaks. Therefore, the uniform convergence in distribution of the test statistics over $\mathcal{T}^a$ follow straightforwardly from extending the results of Zivot and Andrews (1992) and Lumsdaine and Papell (1997). The asymptotic behaviour of the estimates $\hat{\delta}_i$ depend crucially on whether $\varepsilon = 0$ or not. If $\varepsilon = 0$, $\hat{\delta}_1 = 0$ or 1 with equal probability. Otherwise, $\hat{\delta}_1$ converges to some random variable. For more details see Nunes, Kuan, and Newbold (1995) and Bai (1998). It is clear that the conditional distribution of $\hat{\delta}_i$ given $\hat{\delta}_1, \ldots, \hat{\delta}_{i-1}$ is the same as that of $\hat{\delta}_1$. The marginal distribution is however clearly not the same. In any case the distribution of break fractions and the test statistic is likely to depend on the trimming parameter, ε. In conclusion, the asymptotic distribution is quite complex and will be approximated by simulation similarly to previous work in the literature. Under the alternative hypothesis of up to m structural breaks, the break fractions and therefore the coefficients of the model are estimated consistently according to Bai and Perron (1998) and consequently the statistic goes off to minus infinity providing a consistent test.

We note the following. Firstly, we distinguish between three cases. The first assumes that $\psi_1 = \ldots = \psi_m = 0$ under both the null and the alternative. This case will be denoted as case A. The second assumes the same for $\phi_1, \ldots, \phi_m$. This will be denoted as B. The third considers the general model (4) under the alternative and will be denoted as C. Secondly, we assume that k is known. This assumption is not crucial to the analysis and may easily be dropped if the results of Ng and Perron (1995) are taken into account. Their work assumes that the error term in the unit root model follows an ARMA process but that ADF tests are used. Then, it is shown that if a data dependent procedure is used to determine k and this data dependent procedure allows k to rise within specified rates then the distribution of the ADF tests do not change. Both standard information criteria (AIC, BIC) and sequential testing procedures are shown to satisfy the required conditions.

The critical values of the test for cases A,B and C are presented in Table ?? for up to $m = 5$ and $\varepsilon = 0.05$. For higher m, results are available upon request. The critical values have been computed by simulation where standard random walks are generated and used to estimate the relevant model for each case. The errors are standard normal and generated using the GAUSS pseudo-random number generator. For all simulations the number of observations for the random walks is set to 250 and the number of replications to 1000. The test statistics for the unit root tests are presented in Table ??. The above results make interesting reading. The Dickey-Fuller test statistics reject the null hypothesis of a unit root in favour of stationarity in about half of the series considered. The tests that incorporate the possibility of a break under the alternative hypothesis clearly rejects for a much larger number of series indicating the possible presence of a break distorting the analysis according to the Dickey-Fuller test. Further, increasing the number of potential breaks considered we see that in a majority of cases especially for models A and C the number of series for which the null hypothesis of a unit root is rejected increases.

5 Inflation Persistence, Inflation Forecastability, and the Time-Varying Fisher Effect

Despite being one of the cornerstones of monetary economics, as documented for example by Ibrahim and Williams (1978), Barthold and Dougan (1986), and as discussed at length by Barsky (1987), evidence in favor of the Fisher effect is entirely absent from the pre-Bretton-Woods period, and it only appears after about 1960[26]. As stressed for example by Mishkin (1992), evidence *pro*-Fisher has essentially disap-

[26] Lack of evidence in favor of the Fisher effect was stressed by Irving Fisher himself, who, in the *Theory of Interest*, proposed an explanation based on the notion that agents form inflation expectations based on a long distributed lag of past inflation. In the end, however, Fisher himself was dissatisfied with his own theory–see Fisher (1930), pp. [here put exact references].

peared after the beginning of the 1980s. Figure 10 plots U.S. GNP deflator quarterly inflation (quoted at an annual rate), and the 3-month U.S. commercial paper rate, for the period 1875:2-2001:1. Based largely on Meltzer (1986), we divide the monetary history of the United States since 1875 into the following regimes/historical periods: the 'greenback period', prevailing until 1878:4; the Classical Gold Standard regime (1879:1-1914:4); the regime Meltzer labels as a 'gold exchange standard with a central bank', between 1915:1 and 1932:4[27]; the period between 1933:1 and 1941:4, with 'no clear standard'[28]; the period of pegged interest rates, between 1942:1 and 1951:1; the Bretton Woods regime (1951:2-1971:3); the period from the collapse of Bretton Woods to the end of the Volcker disinflation (1971:4-1982:4); and the most recent period, after the Volcker disinflation (1983:1-2001:1). The visual impression from Figure 10 is of a substantial lack of a correlation between movements in inflation and movements in the commercial paper rate up until the 1950s; of a strong correlation between the two series between the beginning of the 1950s and the end of the Volcker disinflation; and of a less clear pattern over the most recent period. Figure 11 shows results from rolling Fama (1976)-tipe regressions of the *ex-post* quarterly inflation rate on a constant and the 3-month commercial paper rate prevailing over the same quarter, for a rolling window of 20 years[29]. Specifically, the figure shows rolling estimates of the coefficient on the 3-month commercial paper rate, together with 90% confidence bands. (Confidence bands have been computed by means of a Newey and West (1987) correction.) The rationale behind Fama (1976) regressions–the methodology traditionally employed to investigate the Fisher effect– is that, under rational expectations, and assuming the Fisher hypothesis to be true, the nominal interest rate prevailing over a specific time period should contain information on the inflation rate which will prevail over the same period. In particular, assuming the *ex-ante* real interest rate to be constant[30], the estimate of the coefficient on the nominal interest rate should not be significantly different from one, thus implying that movements in expected inflation translate one-to-one into movements in nominal interest rates. A number of things are readily apparent from the graph. First, a significant difference between the years up until mid-1960s and the subsequent period, as far as the width of the confidence bands is concerned, with the later period being characterised by a much smaller extent of econometric uncertainty. Second, although for the period up until mid-1960s it is often not possible to reject, at the 90% level, the null that the coefficient on the 3-month commercial paper rate is equal to one, rolling estimates are almost invariably way off the mark, being around zero during the Classical Gold

[27]Meltzer (1986) takes the departure of Great Britain from the Gold Standard (in the third quarter of 1931) as the event marking the end of the interwar gold standard. Given our exclusive focus on the United States, we take instead the first quarter of 1933, when the United States allowed the dollar to float.

[28]See Meltzer (1986), table 4.1.

[29]Very similar results, based on rolling windows of 15 and 25 years, are available upon request.

[30]An assumption which, needless to say, is very much at odds with the recent macroeconomics literature.

Standard period, and being systematically *negative* over the period between 1914 and mid-1960s[31]. After mid-1960s, rolling estimates of the coefficient on the commercial paper rate gradually increase, taking, over most of the 1970s, values not significantly different from one. After about 1980 estimates decrease, stabilising, after the end of the Volcker disinflation, around 0.5, and being significantly different from one.

[here add a section with a discussion of the results from Fama (1976) regressions in the spirit of Fama (1984), with a decomposition similar to the one he used to discuss the Fama puzzle in the FX market]

Currently, there are two leading explanations for such a puzzling time-variation in the extent of the Fisher effect[32]. First, Barthold and Dougan (1986) and Barsky (1987) attribute changes in the extent of the Fisher effect to changes in the extent of inflation forecastability along the sample. To take an extreme case, if inflation is completely unforecastable in the R^2 sense, Fama (1976)-type regressions will fail to uncover evidence *pro*-Fisher even in a world in which the Fisher effect holds *ex-ante* by assumption/construction. The evidence produced by Benati (2003) of dramatic changes in the stochastic properties of inflation both in the U.S. and in the U.K. over the last several decades, and in particular of wide fluctuations in inflation persistence in both countries–which, as first stressed by Barsky (1987), implies equally marked fluctuations in the extent of inflation forecastability–is clearly compatible with such an explanation. A second explanation, put forward by Mishkin in a series of papers[33], is based on the notion that inflation and interest rates are cointegrated. During certain historical periods they share strong stochastic trends, thus making the Fisher effect apparent. Over different historical periods, on the other hand, the stochastic trends they have in common are much more subdued, thus causing the Fisher effect to all but disappear. In the light of the evidence we have produced in the previous pages, we regard the Mishkin explanation as unpersuasive, for the simple reason that, for two series to be cointegrated, they first have to be individually I(1). Although an investigation of the issue of whether nominal interest rate do contain a unit root once one allows for possible structural breaks in their unconditional mean is beyond the scope of this paper, the evidence we have produced against a unit root in almost

[31] This puzzling pattern, and possible explanations for it, are discussed in Benati and Kapetanios (2002). In particular, building on the work of Ibrahim and Williams (1978), Barthold and Dougan (1986), and Barsky (1987), we argue that for the Fisher effect to be detectable via Fama (1976)-type regressions, two things have to hold. First, inflation has to be forecastable in the R^2 sense. Second, there must be sufficient amount of variation in both inflation and nominal interest rates. For a number of various, sometimes highly specific, historical reasons, the period between mid-1960s up until the end of the Volcker disinflation appears to be the only one during which both conditions held.

[32] Here we rule out the Friedman and Schwartz (1976) explanation–based on the notion that economic agents only gradually 'learned their Fisher'–on purely logical grounds. The partial disappearance of a Fisher effect in recent years documented in the previous paragraph would indeed imply that, over the last two decades, economic agents have somehow 'unlearned their Fisher', which appears as implausible to us.

[33] See for example Mishkin (1992).

all inflation series appears to us to rule out, on purely logical grounds, the Mishkin explanation. This leaves open the possibility of the alternative Ibrahim and Williams (1978)-Barthold and Dougan (1986)-Barsky (1987) explanation, which we explore in related work in progress[34].

6 Conclusions

In this paper, we have applied tests for multiple structural breaks at unknown points in the sample, and a newly developed test for unit roots allowing for up to m structural breaks, to investigate break in inflation dynamics for 23 inflation series from from 18 countries (plus the eurozone), in order to produce empirical evidence relevant to the following three questions: *(1) Are there structural breaks in the dynamics of the inflation process? Does inflation possess a unit root? Is inflation a highly persistent process?* We have documented structural breaks in all the series we have analysed. For many countries/series, structural breaks appear to be clustered around the beginning of the 1970s (16 series for 14 countries), of the 1980s (14 series for 13 countries), and of the 1990s (14 series for 11 countries). Further, in several cases estimated break dates are highly suggestive, as they appear to broadly coincide with readily identifiable macroeconomic events, like the breakdown of Bretton Woods, the Volcker disinflation in the U.S., and the introduction of inflation targeting in several countries. Allowing for structural breaks, our new unit root test allows us to increase the number of rejections, compared to standard Dickey-Fuller tests. Finally, conditional on the estimated breaks, inflation series exhibit, in general, little persistence, with the exception of a few countries–for example, the U.S. and the U.K.–around the time of the Great Inflation. As we have stressed, however, such a conclusion has to be considered as tentative, given the intrinsic difficulty of understanding what exactly is driving our rejections of the joint hypothesis of constancy in the innovation variance, the intercept, and the AR coefficients in the autoregressive representations for inflation series we use.

We have discussed an implication of our findings for the Fisher effect. We have argued that Mishkin's explanation for the well-known, puzzling time variation in the extent of the Fisher effect seen in the data, based on the notion that inflation and nominal interest rates are cointegrated, is difficult to defent in the light of our rejection of a unit root for the vast majority of inflation series. The alternative Ibrahim and Williams (1978)-Barthold and Dougan (1986)-Barsky (1987) explanation, based on the notion of changes in the extent of inflation forecastability along the sample, is on the other hand compatible with our findings.

[34]Benati and Kapetanios (2002)

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Table 1 Estimated structural breaks in UK CPI inflation, 1947:1-2002:3

	Estimated structural breaks				
	1959:2	1971:2	1981:3	1991:1	
Sup-Wald	83.881	97.368	32.185	153.153	
(p-value)	(1.47E-11)	(3.00E-14)	(0.0171)	(0)	Lag order: 6
Ave-Wald	43.075	52.291	21.553	68.244	SIC: -6.248
(p-value)	(3.62E-09)	(4.76E-12)	(4.14E-03)	(0)	AIC: -6.416
Exp-Wald	38.577	44.453	13.330	71.549	
(p-value)	(1.51E-12)	(0)	(8.92E-03)	(0)	
	Sub-periods				
	1947:1-1959:1	1959:2-1971:1	1971:2-1981:2	1981:3-1990:4	1991:1-2002:3
Unconditional	0.044	0.044	0.144	0.054	0.023
mean	(0.033)	(0.048)	(0.077)	(0.031)	(4.28E-03)
Sum of the	0.445	0.669	0.597	0.599	0.329
AR coefficients	(0.209)	(0.174)	(0.224)	(0.151)	(0.106)
Innovation	1.00E-03	3.88E-04	5.09E-03	5.46E-04	7.91E-05
variance	(2.47E-04)	(8.90E-05)	(1.29E-03)	(1.46E-04)	(1.84E-05)

Table 2 Estimated structural breaks in UK GDP deflator inflation, 1955:2-2002:2

	Estimated structural breaks				
	1964:2	1972:4	1981:1	1992:3	
Sup-Wald	33.955	31.835	87.574	34.989	
(p-value)	(8.25E-05)	(2.08E-04)	(5.6E-16)	(5.23E-05)	Lag order: 3
Ave-Wald	15.337	18.480	31.743	(14.558)	SIC: -5.675
(p-value)	(6.78E-04)	(6.85E-05)	(2.51E-09)	(1.18E-03)	AIC: -5.761
Exp-Wald	13.131	13.475	39.051	13.698	
(p-value)	(7.13E-05)	(5.09E-05)	(2.22E-16)	(4.09E-05)	
	Sub-periods				
	1955:2-1964:1	1964:2-1972:3	1972:4-1980:4	1981:1-1992:2	1992:3-2002:2
Unconditional	0.031	0.063	0.165	0.058	0.025
mean	(5.57E-03)	(0.013)	(0.030)	(5.44E-03)	(1.66E-03)
Sum of the	-0.464	0.612	0.574	0.195	-0.759
AR coefficients	(0.344)	(0.217)	(0.195)	(0.168)	(0.310)
Innovation	2.18E-03	7.81E-04	5.31E-03	8.16E-04	3.41E-04
variance	(5.72E-04)	(2.02E-04)	(1.39E-03)	(1.78E-04)	(8.04E-05)

Table 3 Estimated structural breaks in US CPI inflation, 1947:2-2002:3

	Estimated structural breaks				
	1958:4	1973:1	1981:4	1990:4	
Sup-Wald	66.661	26.185	27.959	50.021	
(p-value)	(1.84E-11)	(2.25E-03)	(1.08E-03)	(5.35E-08)	Lag order: 3
Ave-Wald	25.487	15.393	9.924	19.676	SIC: -6.816
(p-value)	(3.36E-07)	(6.51E-04)	(0.027)	(2.81E-05)	AIC: -6.893
Exp-Wald	28.697	10.990	10.850	20.455	
(p-value)	(9.09E-12)	(5.66E-04)	(6.47E-04)	(4.58E-08)	
	Sub-periods				
	1947:2-1958:3	1958:4-1972:4	1973:1-1981:3	1981:4-1990:3	1990:4-2002:3
Unconditional	0.018	0.035	0.101	0.040	0.025
mean	(0.012)	(0.014)	(0.015)	(4.42E-03)	(2.40E-03)
Sum of the	0.502	0.860	0.678	0.179	0.300
AR coefficients	(0.179)	(0.101)	(0.152)	(0.204)	(0.152)
Innovation	1.44E-03	1.66E-04	6.78E-04	4.60E-04	1.28E-04
variance	(3.27E-04)	(3.23E-05)	(1.72E-04)	(1.15E-04)	(2.74E-05)

Table 4 Estimated structural breaks in French CPI inflation, 1951:2-2002:2

	Estimated structural breaks				
	1959:3	1973:2	1981:4	1991:1	
Sup-Wald	123.751	30.579	87.974	59.907	
(p-value)	(0)	(0.028)	(2.33E-12)	(4.84E-07)	Lag order: 6
Ave-Wald	65.147	11.988	40.236	32.010	SIC: -6.701
(p-value)	(0)	(0.336)	(2.65E-08)	(6.91E-06)	AIC: -6.879
Exp-Wald	57.815	12.113	39.426	26.218	
(p-value)	(0)	(0.021)	(6.70E-13)	(1.69E-07)	
	Sub-periods				
	1951:2-1959:2	1959:3-1973:1	1973:2-1981:3	1981:4-1990:4	1991:1-2002:2
Unconditional	0.039	0.045	0.116	0.024	0.015
mean	(0.107)	(0.023)	(0.019)	(0.049)	(0.015)
Sum of the	0.602	0.623	0.511	0.809	0.610
AR coefficients	(0.263)	(0.213)	(0.213)	(0.058)	(0.175)
Innovation	3.84E-03	4.60E-04	5.75E-04	1.67E-04	8.19E-05
variance	(1.32E-03)	(9.71E-05)	(1.66E-04)	(4.54E-05)	(1.93E-05)

Table 5 Estimated structural breaks in German CPI inflation, 1950:1-2002:2

	Estimated structural breaks				
	1960:2	1969:3	1981:3	1993:2	
Sup-Wald	132.678	165.230	145.042	154.202	
(p-value)	(0)	(0)	(0)	(0)	Lag order: 1
Ave-Wald	48.559	74.157	63.791	76.372	SIC: -6.936
(p-value)	(0)	(0)	(0)	(0)	AIC: -7.032
Exp-Wald	61.344	78.788	67.988	72.315	
(p-value)	(0)	(0)	(0)	(0)	)
	Sub-periods				
	1950:1-1960:1	1960:2-1969:2	1969:3-1981:2	1981:3-1993:1	1993:2-2002:2
Unconditional mean	0.021	0.025	0.052	0.029	0.017
	(0.023)	(6.43E-03)	(0.021)	(0.022)	(8.79E-03)
Sum of the AR coefficients	0.451	-0.035	0.611	0.669	0.202
	(0.123)	(0.154)	(0.114)	(0.124)	(0.125)
Innovation variance	1.47E-03	2.79E-04	2.02E-04	3.40E-04	1.53E-04
	(3.51E-04)	(6.97E-05)	(4.35E-05)	(7.42E-05)	(3.82E-05)

Table 6 Estimated structural breaks in Swedish CPI inflation, 1962:2-2001:2

	Estimated structural breaks				
	1970:1	1977:3	1984:1	1993:3	
Sup-Wald	41.590	42.774	42.014	113.742	
(p-value)	(1.31E-04)	(8.06E-05)	(1.10E-04)	(0)	Lag order: 4
Ave-Wald	23.135	20.792	19.715	26.728	SIC: -5.954
(p-value)	(2.96E-04)	(1.24E-04)	(2.36E-03)	(3.00E-05)	AIC: -6.129
Exp-Wald	17.209	17.709	17.825	52.167	
(p-value)	(0.860)	(0.551)	(0.497)	(0)	
	Sub-periods				
	1962:2-1969:4	1970:1-1977:2	1977:3-1983:4	1984:1-1993:2	1993:3-2001:2
Unconditional mean	0.041	0.096	0.100	0.062	0.016
	(0.031)	(0.028)	(0.031)	(0.028)	(0.105)
Sum of the AR coefficients	0.491	0.361	0.242	0.464	0.724
	(0.302)	(0.377)	(0.233)	(0.238)	(0.217)
Innovation variance	6.56E-04	1.47E-03	9.28E-04	1.43E-03	2.21E-04
	(2.13E-04)	(4.44E-04)	(3.09E-04)	(3.70E-04)	(6.37E-05)

Table 7 Estimated structural breaks in Swiss CPI inflation, 1947:1-2002:2

	Estimated structural breaks				
	1961:2	1970:3	1983:4	1993:3	
Sup-Wald	98.074	29.002	68.010	54.339	
(p-value)	(0)	(0.026)	(4.27E-09)	(1.64E-06)	Lag order: 5
Ave-Wald	44.069	19.082	31.897	30.899	SIC: -6.623
(p-value)	9.85E-10	8.20E-03	3.59E-06	6.84E-06	AIC: -6.776
Exp-Wald	44.562	12.516	29.416	23.861	
(p-value)	(0)	(9.68E-03)	(4.76E-09)	(7.51E-07)	
	Sub-periods				
	1947:1-1961:1	1961:2-1970:2	1970:3-1983:3	1983:4-1993:2	1993:3-2002:2
Unconditional mean	0.012	0.035	0.048	0.033	8.98E-03
	(0.016)	(7.20E-03)	(0.054)	(0.041)	(6.35E-03)
Sum of the AR coefficients	0.507	-0.293	0.712	0.711	-0.106
	(0.165)	(0.375)	(0.190)	(0.154)	(0.301)
Innovation variance	2.98E-04	4.00E-04	1.43E-03	2.30E-04	1.78E-04
	(6.43E-05)	(1.07E-04)	(3.04E-04)	(5.95E-05)	(4.85E-05)

Table 8 Estimated structural breaks in Portugal's CPI inflation, 1962:2-2001:2

	Estimated structural breaks				
	1973:1	1979:3	1986:2	1995:2	
Sup-Wald	83.845	19.605	119.392	87.075	
(p-value)	(1.00E-14)	(0.058)	(0)	(0)	Lag order: 1
Ave-Wald	48.475	10.613	61.892	48.680	SIC: -4.923
(p-value)	(2.00E-14)	(0.040)	(0)	(2.00E-14)	AIC: -5.040
Exp-Wald	38.792	7.301	55.048	39.856	
(p-value)	(0)	(0.037)	(0)	(0)	
	Sub-periods				
	1962:2-1972:4	1973:1-1979:2	1979:3-1986:1	1986:2-1995:1	1995:2-2001:2
Unconditional mean	0.065	0.238	0.210	0.089	0
	(0.028)	(0.061)	(0.082)	(0.046)	(5.97E-03)
Sum of the AR coefficients	0.296	-0.249	0.401	0.565	-0.021
	(0.156)	(0.212)	(0.190)	(0.142)	(0.162)
Innovation variance	2.68E-03	0.034	6.37E-03	9.32E-04	2.23E-04
	(6.24E-04)	(0.011)	(1.92E-03)	(2.37E-04)	(7.05E-05)

Table 9 Estimated structural breaks in Spain's CPI inflation, 1962:2-2001:2

	Estimated structural breaks				
	1973:3	1980:1	1986:3	1992:4	
Sup-Wald	20.081	34.570	73.176	47.332	
(p-value)	(0.090)	(4.19E-04)	(8.43E-12)	(1.65E-06)	Lag order: 2
Ave-Wald	10.263	14.587	33.407	25.005	SIC: -5.987
(p-value)	(0.104)	(9.54E-03)	(2.92E-08)	(1.10E-05)	AIC: -6.123
Exp-Wald	7.109	13.456	32.350	20.207	
(p-value)	(0.080)	(4.01E-04)	(4.01E-12)	(7.09E-07)	
	Sub-periods				
	1962:2-1973:2	1973:3-1979:4	1980:1-1986:2	1986:3-1992:3	1992:4-2001:2
Unconditional	0.073	0.184	0.117	0.059	0.035
mean	(0.027)	(0.039)	(0.052)	(9.60E-03)	(0.012)
Sum of the	0.459	0.295	0.603	0.251	0.560
AR coefficients	(0.183)	(0.266)	(0.189)	(0.162)	(0.183)
Innovation	2.18E-03	4.38E-03	7.56E-04	2.17E-04	2.16E-04
variance	(5.13E-04)	(1.39E-03)	(2.39E-04)	(7.03E-05)	(5.57E-05)

Table 10 Estimated structural breaks in French GDP deflator inflation, 1962:2-2001:2

	Estimated structural breaks			
	1972:1	1983:3	1993:2	
Sup-Wald	21.762	61.904	18.906	
(p-value)	(0.0131)	(1.84E-10)	(0.038)	Lag order: 3
Ave-Wald	11.461	32.706	11.209	SIC: -6.763
(p-value)	(0.010)	(1.17E-09)	(0.012)	AIC: -6.859
Exp-Wald	7.864	28.126	7.742	
(p-value)	(0.010)	(1.65E-11)	(0.012)	
	Sub-periods			
	1962:2-1971:4	1972:1-1983:2	1983:3-1993:1	1993:2-2001:2
Unconditional	0.046	0.109	0.028	0.012
mean	(6.88E-03)	(6.13E-03)	(8.46E-03)	(2.93E-03)
Sum of the	0.193	0.259	0.662	0.594
AR coefficients	(0.276)	(0.202)	(0.111)	(0.167)
Innovation	1.10E-03	9.27E-04	2.21E-04	4.46E-05
variance	(2.77E-04)	(2.02E-04)	(5.29E-05)	(1.17E-05)

Table 11 Estimated structural breaks in Italy's CPI inflation 1948:1-2002:3

	Estimated structural breaks			
	1964:4	1973:3	1982:4	
Sup-Wald	88.040	43.152	103.879	
(p-value)	(8.30E-14)	(6.89E-05)	(0)	Lag order: 4
Ave-Wald	40.403	26.957	54.770	SIC: -6.050
(p-value)	(2.71E-09)	(2.59E-05)	(8.63E-14)	AIC: -6.189
Exp-Wald	40.058	18.184	48.010	
(p-value)	(2.26E-14)	(3.60E-05)	(0)	
	Sub-periods			
	1948:1-1964:3	1964:4-1973:2	1973:3-1982:3	1982:4-2002:3
Unconditional	0.036	0.088	0.174	0.036
mean	(0.018)	(0.114)	(0.023)	(0.025)
Sum of the	0.108	0.954	0.056	0.789
AR coefficients	(0.216)	(0.124)	(0.240)	(0.045)
Innovation	2.96E-03	1.89E-04	3.39E-03	2.06E-04
variance	(5.64E-04)	(5.15E-05)	(8.89E-04)	(3.44E-05)

Table 12 Estimated structural breaks in Canada's CPI inflation 1947:1-2002:2

	Estimated structural breaks			
	1961:3	1982:3	1991:2	
Sup-Wald	41.356	72.893	87.481	
(p-value)	(2.34E-05)	(9.68E-12)	(7.33E-15)	Lag order: 2
Ave-Wald	20.611	35.447	47.163	SIC: -6.217
(p-value)	(2.13E-04)	(6.68E-09)	(1.17E-12)	AIC: -6.325
Exp-Wald	17.61541163	32.2737608	39.45630524	
(p-value)	(8.49E-06)	(4.33E-12)	(2.78E-15)	
	Sub-periods			
	1947:1-1961:2	1961:3-1982:2	1982:3-1991:1	1991:2-2002:2
Unconditional	0.023	0.075	0.047	0.018
mean	(0.038)	(0.073)	(0.010)	(5.85E-03)
Sum of the	0.705	0.844	0.354	-0.138
AR coefficients	(0.106)	(0.075)	(0.166)	(0.158)
Innovation	1.60E-03	7.11E-04	2.41E-04	3.89E-04
variance	(3.21E-04)	(1.14E-04)	(6.33E-05)	(8.81E-05)

Table 13 Estimated structural breaks in Belgium's CPI inflation 1947:1-2002:3

	Estimated structural breaks			
	1955:4	1971:1	1985:2	
Sup-Wald	90.640	20.950	47.860	
(p-value)	(0)	(0.018)	(1.48E-07)	Lag order: 3
Ave-Wald	49.173	10.095	18.809	SIC: -5.885
(p-value)	(0)	(0.025)	(5.36E-05)	AIC: -5.961
Exp-Wald	41.290	7.026	19.602	
(p-value)	(0)	(0.022)	(1.09E-07)	
	Sub-periods			
	1947:1-1955:3	1955:4-1970:4	1971:1-1985:1	1985:2-2002:3
Unconditional	0.024	0.026	0.078	0.021
mean	(0.029)	(4.72E-03)	(0.011)	(2.23E-03)
Sum of the	0.377	0.288	0.597	0.048
AR coefficients	(0.237)	(0.227)	(0.149)	(0.183)
Innovation	0.010	6.89E-04	1.12E-03	3.08E-04
variance	(2.79E-03)	(1.29E-04)	(2.17E-04)	(5.37E-05)

Table 14 Estimated structural breaks in Dutch CPI inflation 1945:4-2002:3

	Estimated structural breaks			
	1961:2	1974:2	1991:3	
Sup-Wald	89.415	140.495	70.575	
(p-value)	(2.78E-15)	(0)	(2.98E-11)	Lag order: 2
Ave-Wald	26.430	83.525	29.924	SIC: -5.754
(p-value)	(4.10E-06)	(0)	(3.52E-07)	AIC: -5.860
Exp-Wald	40.313	66.152	31.435	
(p-value)	(1.11E-15)	(0)	(1.01E-11)	
	Sub-periods			
	1945:4-1961:1	1961:2-1974:1	1974:2-1991:2	1991:3-2002:3
Unconditional	0.044	0.057	0.030	0.028
mean	(0.025)	(0.021)	(0.034)	(0.014)
Sum of the	0.048	0.164	0.832	0.092
AR coefficients	(0.190)	(0.167)	(0.073)	(0.203)
Innovation	7.63E-03	1.25E-03	4.08E-04	2.50E-04
variance	(1.47E-03)	(2.60E-04)	(7.27E-05)	(5.66E-05)

Table 15 Estimated structural breaks in Norway's CPI inflation 1962:2-2001:2

	Estimated structural breaks			
	1970:1	1982:2	1988:3	
Sup-Wald	41.775	44.292	120.523	
(p-value)	(1.951E-05)	(6.44E-06)	(0)	Lag order: 2
Ave-Wald	18.611	25.857	54.444	SIC: -6.228
(p-value)	(7.84E-04)	(6.10E-06)	(4.88E-15)	AIC: -6.365
Exp-Wald	16.978	19.212	56.122	
(p-value)	(1.550E-05)	(1.85E-06)	(0)	
	Sub-periods			
	1962:2-1969:4	1970:1-1982:1	1982:2-1988:2	1988:3-2001:2
Unconditional	0.038	0.094	0.070	0.027
mean	(0.014)	(0.019)	(0.030)	(7.01E-03)
Sum of the	0.066	0.238	0.559	0.387
AR coefficients	(0.270)	(0.198)	(0.171)	(0.144)
Innovation	1.00E-03	2.00E-03	3.50E-04	2.09E-04
variance	(2.96E-04)	(4.31E-04)	(1.14E-04)	(4.36E-05)

Table 16 Estimated structural breaks in Finland's CPI inflation 1962:2-2001:2

	Estimated structural breaks			
	1976:4	1984:3	1994:1	
Sup-Wald	91.235	74.739	106.017	
(p-value)	(3.33E-16)	(1.30E-12)	(0)	Lag order: 1
Ave-Wald	40.843	31.023	47.278	SIC: -6.331
(p-value)	(9.76E-12)	(2.04E-08)	(5.94E-14)	AIC: -6.448
Exp-Wald	41.999	33.819	48.752	
(p-value)	(1.11E-16)	(3.81E-13)	(0)	
	Sub-periods			
	1962:2-1976:3	1976:4-1984:2	1984:3-1993:4	1994:1-2001:2
Unconditional	0.087	0.090	0.040	0.017
mean	(0.039)	(0.044)	(0.024)	(0.010)
Sum of the	0.685	0.607	0.687	0.420
AR coefficients	(0.103)	(0.145)	(0.123)	(0.195)
Innovation	2.14E-03	4.92E-04	2.14E-04	2.18E-04
variance	(4.21E-04)	(1.36E-04)	(5.28E-05)	(6.16E-05)

Table 17 Estimated structural breaks in Austrian CPI inflation 1950:1-2002:2

	Estimated structural breaks			
	1957:3	1967:1	1984:2	
Sup-Wald	109.429	34.815	104.076	
(p-value)	(0)	(1.55E-04)	(0)	Lag order: 1
Ave-Wald	62.438	19.278	54.910	SIC: -5.417
(p-value)	(0)	(1.27E-04)	(1.10E-16)	AIC: -5.513
Exp-Wald	50.358	14.536	47.881	
(p-value)	(0)	(6.27E-05)	(0)	
	Sub-periods			
	1950:1-1957:2	1957:3-1966:4	1967:1-1984:1	1984:2-2002:2
Unconditional	0.091	0.038	0.056	0.024
mean	(0.110)	(0.024)	(0.010)	(5.55E-03)
Sum of the	0.389	-0.290	0.252	-0.135
AR coefficients	(0.181)	(0.158)	(0.125)	(0.106)
Innovation	0.026	4.90E-03	9.23E-04	4.14E-04
variance	(7.49E-03)	(1.21E-03)	(1.63E-04)	(7.11E-05)

Table 18 Estimated structural breaks in US GNP deflator inflation 1875:2-2002:2

	Estimated structural breaks			
	1921:1	1952:4	1981:2	
Sup-Wald	62.75063849	234.1384532	47.08569746	
(p-value)	(1.22505E-10)	(0)	(2.11811E-07)	Lag order: 3
Ave-Wald	42.51248691	136.6287018	21.50355586	SIC: -5.559409798
(p-value)	(4.54636E-13)	(0)	(7.10303E-06)	AIC: -5.517833688
Exp-Wald	28.0683342	113.0183165	19.0299126	
(p-value)	(1.75168E-11)	(0)	(1.95792E-07)	
	Sub-periods			
	1875:2-1920:4	1921:1-1952:3	1952:4-1981:1	1981:2-2002:2
Unconditional	0.020522024	0.021550585	0.066355167	0.022907437
mean	(0.013307948)	(0.016773979)	(0.044698179)	(0.003513926)
Sum of the	0.40573414	0.69865533	0.958596364	0.732876942
AR coefficients	(0.105088887)	(0.068697235)	(0.04861721)	(0.0504)
Innovation	0.011248818	0.003199415	0.000205444	6.12803E-05
variance	(0.001199128)	(0.000407974)	(2.7702E-05)	(9.62928E-06)

Table 19 Estimated structural breaks in Australia's CPI inflation 1962:2-2001:2

	Estimated structural breaks			
	1970:4	1977:1	1991:1	
Sup-Wald	49.713	34.823	55.208	
(p-value)	(1.99E-07)	(1.54E-04)	(1.54E-08)	Lag order: 1
Ave-Wald	22.973	19.518	29.107	SIC: -5.941
(p-value)	(8.72E-06)	(1.07E-04)	(8.81E-08)	AIC: -6.058
Exp-Wald	20.865	14.049	23.770	
(p-value)	(1.48E-07)	(9.86E-05)	(8.63E-09)	
	Sub-periods			
	1962:2-1970:3	1970:4-1976:4	1977:1-1990:4	1991:1-2001:2
Unconditional	0.027	0.120	0.083	0.023
mean	(7.97E-03)	(0.043)	(0.011)	(0.011)
Sum of the	0.080	0.416	0.299	0.149
AR coefficients	(0.183)	(0.216)	(0.106)	(0.146)
Innovation	4.29E-04	3.64E-03	7.38E-04	8.09E-04
variance	(1.15E-04)	(1.15E-03)	(1.46E-04)	(1.88E-04)

Table 20 Estimated structural breaks in New Zealand's inflation 1925:4-2002:2

	Estimated structural breaks		
	1970:1	1989:4	
Sup-Wald	62.937	139.113	
(p-value)	(3.93E-10)	(0)	Lag order: 1
Ave-Wald	38.459	39.885	SIC: -5.267
(p-value)	(6.36E-11)	(2.08E-11)	AIC: -5.341
Exp-Wald	27.282	64.557	
(p-value)	(2.67E-10)	(0)	
	Sub-periods		
	1925:4-1969:4	1970:1-1989:3	1989:4-2002:2
Unconditional	0.026	0.121	0.020
mean	(0.012)	(0.025)	(7.91E-03)
Sum of the	0.395	0.500	0.379
AR coefficients	(0.070)	(0.098)	(0.096)
Innovation	2.43E-03	2.70E-03	2.89E-04
variance	(2.63E-04)	(4.43E-04)	(6.03E-05)

Table 21 Estimated structural breaks in Ireland's CPI inflation 1962:2-2001:2

	Estimated structural breaks		
	1973:1	1983:4	
Sup-Wald	41.486	93.788	
(p-value)	(2.21E-05)	(3.3E-16)	Lag order: 2
Ave-Wald	20.300	50.732	SIC: -5.417
(p-value)	(2.61E-04)	(8.04E-14)	AIC: -5.552
Exp-Wald	17.548	43.382	
(p-value)	(9.05E-06)	(0)	
		Sub-periods	
	1962:2-1972:4	1973:1-1983:3	1983:4-2001:2
Unconditional	0.0614	0.157	0.031
mean	(0.023)	(0.035)	(0.012)
Sum of the	0.295	0.221	0.561
AR coefficients	(0.219)	(0.203)	(0.110)
Innovation	2.01E-03	6.18E-03	4.36E-04
variance	(4.79E-04)	(1.40E-03)	(7.65E-05)

Table 22 Estimated structural breaks in Eurozone's GDP deflator inflation 1970:2-1998:3

	Estimated structural breaks		
	1984:2	1992:2	
Sup-Wald	32.462	62.962	
(p-value)	(1.09E-05)	(1.75E-12)	Lag order: 1
Ave-Wald	13.657	24.095	SIC: -6.318
(p-value)	(1.60E-04)	(3.60E-08)	AIC: -6.390
Exp-Wald	12.223	27.115	
(p-value)	(2.00E-04)	(9.82E-04)	
		Sub-periods	
	1970:2-1984:1	1984:2-1992:1	1992:2-1998:3
Unconditional	0.126	0.077	0.042
mean	(7.38E-03)	(3.48E-03)	(3.59E-03)
Sum of the	0.468	-0.252	0.126
AR coefficients	(0.119)	(0.165)	(0.178)
Innovation	8.48E-04	6.05E-04	2.52E-04
variance	(1.65E-04)	(1.56E-04)	(7.29E-05)

Table 23 Estimated structural breaks in Eurozone's HICP inflation, 1970:2-1998:4

	Estimated structural breaks		
	1981:2	1993:2	
Sup-Wald	109.033	69.747	
(p-value)	(0)	(4.45E-11)	Lag order: 2
Ave-Wald	38.970	29.309	SIC: -7.542
(p-value)	(5.10E-10)	(5.44E-07)	AIC: -7.71
Exp-Wald	50.123	30.986	
(p-value)	(0)	(1.60E-11)	
		Sub-periods	
	1970:2-1981:1	1981:2-1993:1	1993:2-1998:4
Unconditional mean	0.098	0.039	0.015
	(0.035)	(0.034)	(0.022)
Sum of the AR coefficients	0.708	0.858	0.664
	(0.134)	(0.053)	(0.168)
Innovation variance	3.99E-04	1.03E-04	5.41E-05
	(9.40E-05)	(2.24E-05)	(1.85E-05)

Table 24 Results from Nyblom-Hansen tests for individual sets of parameters

	Intercept	*Variance*	*AR coeffs*
UKCPI	NR	NR	
SWICPI	NR	NR	
BELCPI	5%	1%	
NETHCPI	NR	1%	
AUSTRIACPI	5%	5%	
FINCPI	5%	1%	
NORCPI	10%	10%	
PORCPI	10%	5%	
SPACPI	10%	1%	
IRECPI	NR	5%	
FRAGDPDEF	10%	1%	
UKGDPDEF	NR	5%	
USCPI	NR	1%	
EUROGDPDEF	5%	NR	
EUROHICP	10%	5%	
NZCPI	5%	NR	
SWECPI	NR	NR	
AUCPI	NR	NR	
CANCPI	NR	1%	
GERCPI	NR	1%	
FRACPI	NR	1%	
ITACPI	NR	1%	
USGNPDEF	5%	1%	

For details, see text. NR=no rejection even at the 10% level; 1%=rejection at the 1% level; 5%=rejection at the 5% level; 10%=rejection at the 10% level.

Table 25 Critical values for __ for models A, B, and C					
		Significance level			
Model	m	0.1	0.05	0.025	0.01
A	1	-4.661	-4.938	-5.173	-5.338
	2	-5.467	-5.685	-5.965	-6.162
	3	-6.265	-6.529	-6.757	-6.991
	4	-6.832	-7.104	-7.361	-7.560
	5	-7.398	-7.636	-7.963	-8.248
B	1	-4.144	-4.495	-4.696	-5.014
	2	-4.784	-5.096	-5.333	-5.616
	3	-5.429	-5.726	-6.010	-6.286
	4	-5.999	-6.305	-6.497	-6.856
	5	-6.417	-6.717	-6.998	-7.395
C	1	-4.820	-5.081	-5.297	-5.704
	2	-5.847	-6.113	-6.344	-6.587
	3	-6.686	-7.006	-7.216	-7.401
	4	-7.426	-7.736	-7.998	-8.243
	5	-8.016	-8.343	-8.593	-9.039

Table 26 Tests Statistics for Unit Root test with breaks for model A					
Country	*Maximum break number*				
	1	2	3	4	5
New Zealand	-6.098**	-6.885**	-6.885*	-9.935**	-10.361**
Sweden	-4.180	-4.935	-5.290	-5.668	-5.727
Australia	-4.943*	-5.579	-6.609*	-7.802**	-8.105*
Canada	-7.854**	-9.137**	-9.137**	-9.137**	-9.137**
Germany	-7.522**	-8.150**	-8.356**	-8.735**	-9.031**
France	-4.995*	-5.974*	-7.329**	-9.230**	-9.230**
Italy	-4.896	-6.705**	-11.064**	-11.521**	-11.725**
U.S.	-11.057**	-12.914**	-12.914**	-13.520**	-13.949**
U.K.	-7.331**	-8.937**	-8.937**	-8.937**	-9.062**
Switzerland	-4.637	-5.257	-6.348	-6.348	-6.671
Belgium	-8.592**	-9.550**	-10.014**	-11.107**	-11.640**
Netherland	-5.880**	-7.484**	-8.130**	-8.130**	-9.448**
Austria	-6.177**	-6.195**	-6.195	-6.195	-6.195
Finland	-4.905	-5.552	-5.991	-6.298	-6.563
Norway	-4.883	-6.090*	-6.594*	-8.041**	-8.675**
Portugal	-7.267**	-8.102**	-14.064**	-15.230**	-17.477**
Spain	-6.308**	-7.199**	-7.821**	-8.829**	-13.416**
Ireland	-4.065	-5.304	-5.777	-6.757	-7.164
France	-3.904	-7.689**	-13.886**	-14.653**	-14.653**
U.K.	-4.448	-7.621**	-7.621**	-7.621**	-14.540**
U.S.	-5.917**	-7.692**	-8.483**	-9.720**	-9.951**
Euro1	-3.586	-4.195	-4.728	-5.415	-5.415
Euro2	-6.733**	-7.180**	-7.622**	-8.827**	-9.380**
No. of rejections (5%)	2	2	3	0	1
No. of rejections (1%)	12	15	14	17	16
Single stars indicate rejection at the 5% significance level.					
Double stars indicate rejection at the 1% significance level.					

Table 27 Tests Statistics for Unit Root test with breaks for model A

Country	*Maximum break number*				
	1	2	3	4	5
New Zealand	-6.098**	-6.885**	-6.885*	-9.935**	-10.361**
Sweden	-4.180	-4.935	-5.290	-5.668	-5.727
Australia	-4.943*	-5.579	-6.609*	-7.802**	-8.105*
Canada	-7.854**	-9.137**	-9.137**	-9.137**	-9.137**
Germany	-7.522**	-8.150**	-8.356**	-8.735**	-9.031**
France	-4.995*	-5.974*	-7.329**	-9.230**	-9.230**
Italy	-4.896	-6.705**	-11.064**	-11.521**	-11.725**
U.S.	-11.057**	-12.914**	-12.914**	-13.520**	-13.949**
U.K.	-7.331**	-8.937**	-8.937**	-8.937**	-9.062**
Switzerland	-4.637	-5.257	-6.348	-6.348	-6.671
Belgium	-8.592**	-9.550**	-10.014**	-11.107**	-11.640**
Netherland	-5.880**	-7.484**	-8.130**	-8.701**	-9.448**
Austria	-6.177**	-6.195**	-6.195	-6.195	-6.195
Finland	-4.905	-5.552	-5.991	-6.298	-6.563
Norway	-4.883	-6.090*	-6.594*	-8.041**	-8.675**
Portugal	-7.267**	-8.102**	-14.064**	-15.230**	-17.477**
Spain	-6.308**	-7.199**	-7.821**	-8.829**	-13.416**
Ireland	-4.065	-5.304	-5.777	-6.757	-7.164
France	-3.904	-7.689**	-13.886**	-14.653**	-14.653**
U.K.	-4.448	-7.621**	-7.621**	-7.621**	-14.540**
U.S.	-5.917**	-7.692**	-8.483**	-9.720**	-9.951**
Euro1	-3.586	-4.195	-4.728	-5.415	-5.415
Euro2	-6.733**	-7.180**	-7.622**	-8.827**	-9.380**
No. of rejections (5%)	2	2	3	0	1
No. of rejections (1%)	12	15	14	17	16

Single stars indicate rejection at the 5% significance level.
Double stars indicate rejection at the 1% significance level.

Table 28 Tests Statistics for Unit Root Test with Breaks for Model B

Country	*Maximum break number*				
	1	2	3	4	5
New Zealand	-5.343**	-5.725**	-8.653**	-9.233**	-13.336**
Sweden	-4.464	-4.715	-4.833	-5.033	-5.452
Australia	-4.289	-5.021	-5.694	-5.694	-6.032
Canada	-7.016**	-7.698**	-9.182**	-9.182**	-9.182**
Germany	-7.466**	-7.686**	-8.018**	-8.080**	-8.127**
France	-4.954*	-5.381*	-5.901*	-6.695*	-7.393**
Italy	-4.909*	-5.468*	-10.550**	-10.618**	-11.202**
U.S.	-10.715**	-10.753**	-10.849**	-11.183**	-16.528**
U.K.	-6.691**	-7.896**	-8.505**	-8.898**	-8.898**
Switzerland	-4.833*	-4.927	-5.642	-5.922	-6.471
Belgium	-8.556**	-8.934**	-9.647**	-9.853**	-9.853**
Netherland	-5.210**	-5.879**	-7.104**	-7.250**	-7.751**
Austria	-7.491**	-9.507**	-10.038**	-10.236**	-10.763**
Finland	-3.482	-5.183*	-6.004*	-6.474*	-9.094**
Norway	-4.619*	-4.908	-5.381	-5.980	-6.806*
Portugal	-13.046**	-13.334**	-13.556**	-13.662**	-14.079**
Spain	-5.584**	-6.698**	-11.810**	-12.421**	-12.692**
Ireland	-3.122	-3.741	-4.362	-4.429	-4.528
France	-6.681**	-6.965**	-11.843**	-12.172**	-12.628**
U.K.	-4.704*	-6.290**	-6.696**	-7.203**	-8.596**
U.S.	-6.135**	-6.453**	-7.540**	-8.082**	-8.447**
Euro1	-3.570	-3.689	-4.529	-5.429	-6.280
Euro2	-6.508**	-6.671**	-6.859**	-7.184**	-7.596**
No. of rejections (5%)	6	3	2	2	1
No. of rejections (1%)	13	14	15	15	17

Single stars indicate rejection at the 5% significance level.
Double stars indicate rejection at the 1% significance level.

Table 29 Tests Statistics for Unit Root Test with Breaks for Model C					
Country	*Maximum break number*				
	1	2	3	4	5
New Zealand	-6.032**	-6.817**	-6.817	-7.905*	-8.572*
Sweden	-4.639	-5.047	-6.114	-6.807	-7.898
Australia	-5.483*	-6.389*	-6.757	-7.444	-7.927
Canada	-7.970**	-9.510**	-9.510**	-9.510**	-9.510**
Germany	-7.877**	-8.495**	-8.912**	-9.227**	-9.484**
France	-5.547*	-5.547	-8.784**	-9.132**	-9.132**
Italy	-10.374**	-10.897**	-11.684**	-11.684**	-11.684**
U.S.	-11.575**	-12.881**	-12.881**	-13.755**	-14.437**
U.K.	-7.167**	-7.784**	-7.829**	-8.344**	-8.356*
Switzerland	-5.337*	-6.235*	-6.639	-7.580	-8.670*
Belgium	-9.250**	-9.250**	-9.250**	-9.250**	-9.250**
Netherland	-5.867**	-8.103**	-8.910**	-9.328**	-9.410**
Austria	-8.400**	-10.208**	-10.592**	-10.881**	-11.269**
Finland	-5.053	-5.450	-8.307**	-8.766**	-8.772*
Norway	-5.020	-5.742	-6.515	-6.515	-6.515
Portugal	-13.826**	-15.168**	-16.469**	-16.544**	-17.491**
Spain	-6.523**	-11.926**	-11.926**	-11.926**	-11.926**
Ireland	-4.672	-5.795	-6.623	-7.335	-7.335
France	-7.291**	-12.577**	-13.519**	-13.519**	-13.519**
U.K.	-6.736**	-8.089**	-13.631**	-14.897**	-16.856**
U.S.	-7.018**	-7.405**	-8.010**	-8.010*	-8.010
Euro1	-3.686	-4.102	-4.102	-4.102	-4.777
Euro2	-6.701**	-7.626**	-8.305**	-8.578**	-9.894**
No. of rejections (5%)	3	2	0	2	4
No. of rejections (1%)	15	15	16	15	13
Single stars indicate rejection at the 5% significance level.					
Double stars indicate rejection at the 1% significance level.					

Table 30 Tests Statistics for Dickey-Fuller Unit Root test			
Country	*Maximum break number*		
	1	2	3
New Zealand	-2.643**	-3.687**	-3.863*
Sweden	-1.209	-2.367	-2.690
Australia	-1.231	-2.523	-2.721
Canada	-4.800**	-5.769**	-5.812**
Germany	-3.759**	-6.549**	-6.471**
France	-2.111*	-3.361*	-3.440*
Italy	-2.186*	-3.247*	-3.226
U.S.	-6.398**	-10.369**	-10.555**
U.K.	-4.647**	-5.553**	-5.797**
Switzerland	-2.363*	-3.617**	-3.585*
Belgium	-4.258**	-6.168**	-6.181**
Netherland	-2.885**	-4.743**	-4.831**
Austria	-4.588**	-5.875**	-5.909**
Finland	-1.281	-2.159	-2.645
Norway	-1.260	-2.785	-3.168
Portugal	-1.548	-2.753	-2.893
Spain	-1.244	-2.028	-2.343
Ireland	-1.164	-2.128	-2.513
France	-1.222	-1.406	-1.739
U.K.	-2.013*	-3.031*	-3.056
U.S.	-2.629**	-3.926**	-4.043**
Euro1	-1.034	-1.024	-2.861
Euro2	-1.309	-2.593	-5.588**
No. of rejections (5%)	4	3	5
No. of rejections (1%)	9	10	9

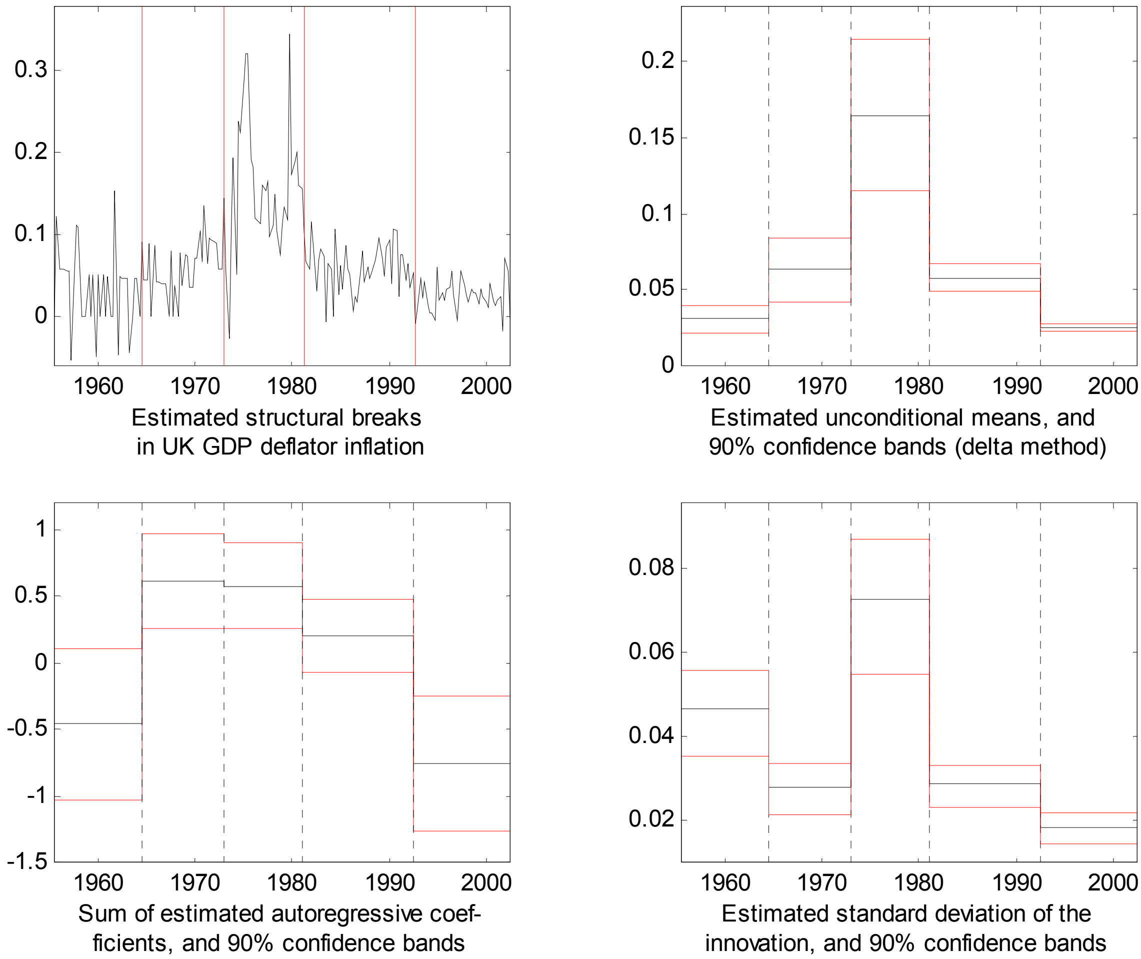

Figure 1: UK GDP deflator inflation (1955:2-2002:2), estimated structural breaks in the mean, innovation standard deviation, and AR coefficients

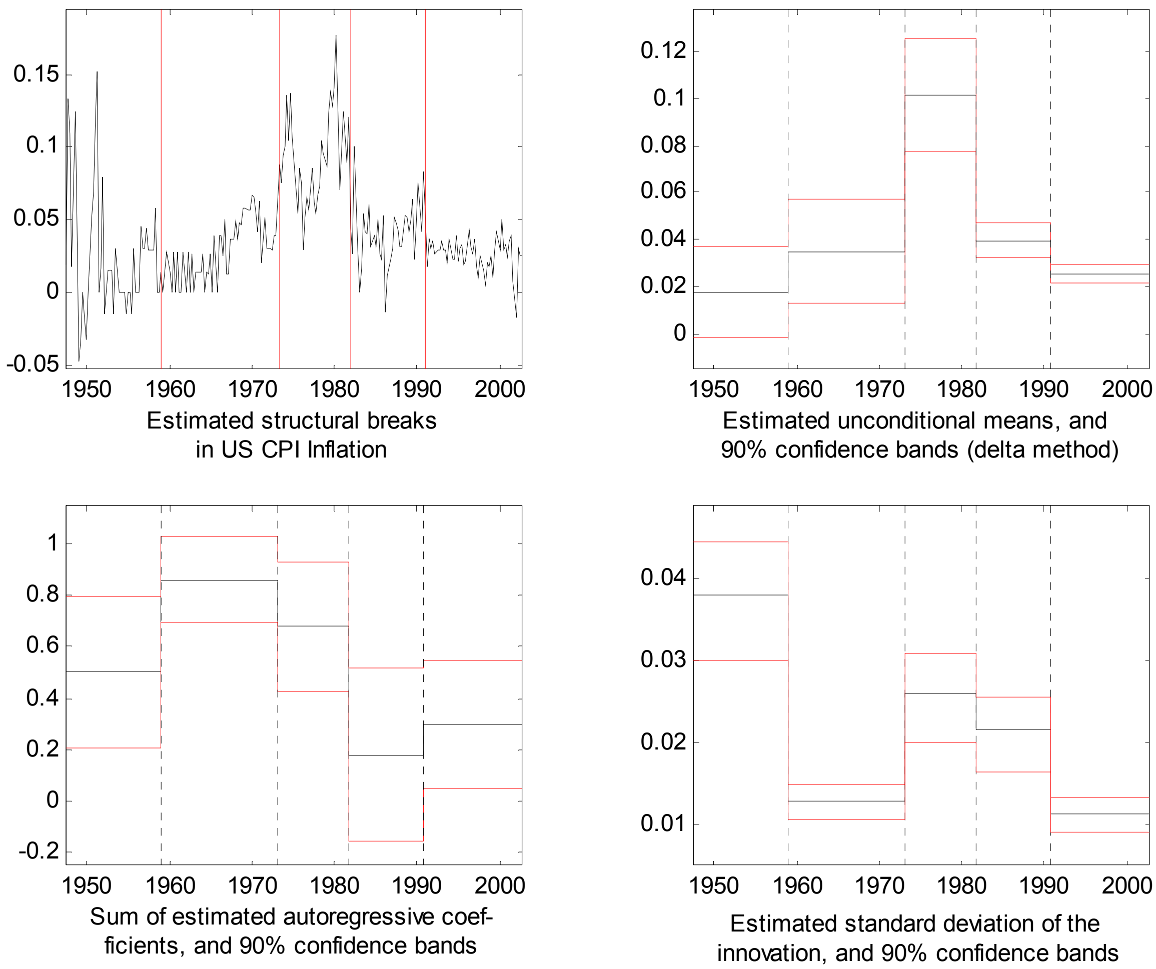

Figure 2: US CPI inflation (1947:2-2002:3), estimated structural breaks in the mean, innovation standard deviation, and AR coefficients

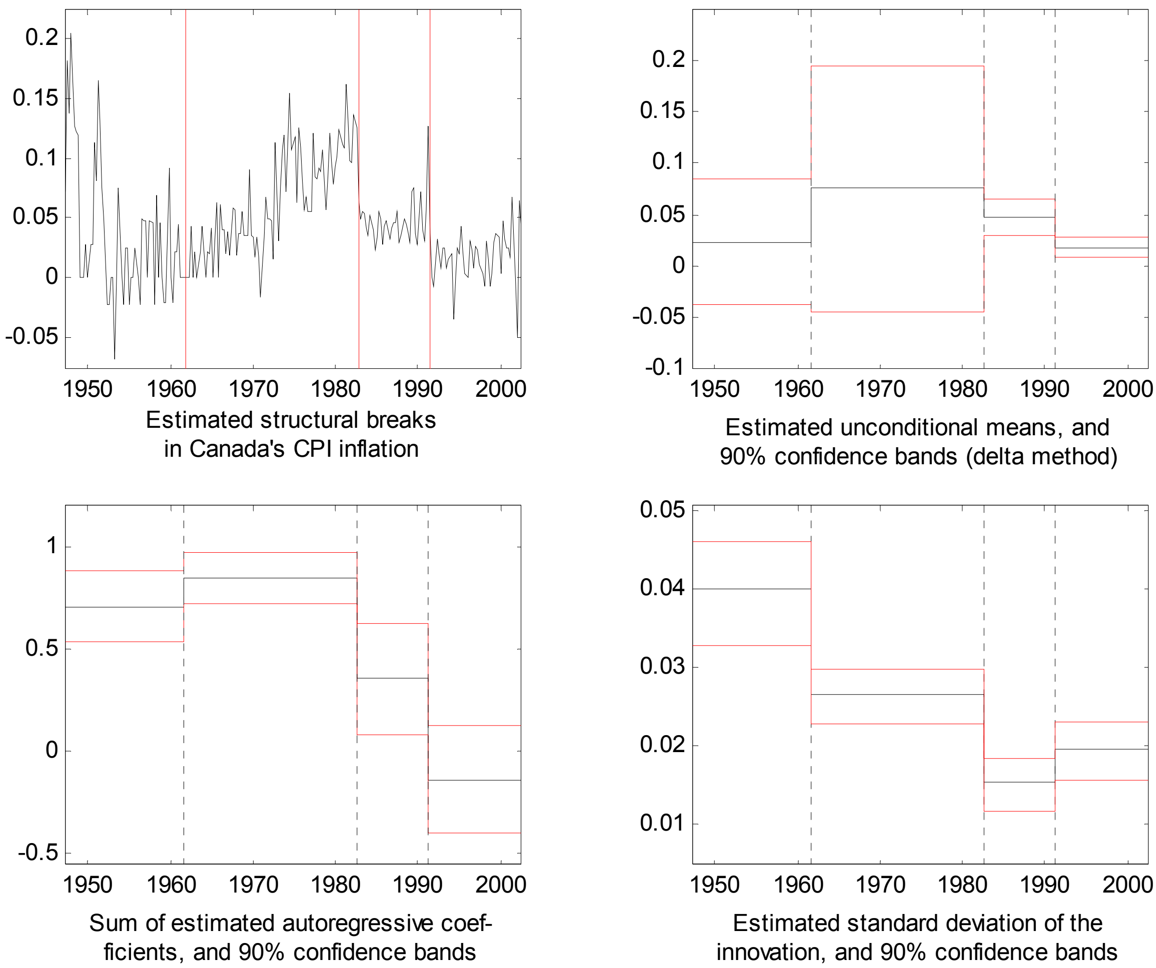

Figure 3: Canada's CPI inflation (1947:1-2002:2), estimated structural breaks in the mean, innovation standard deviation, and AR coefficients

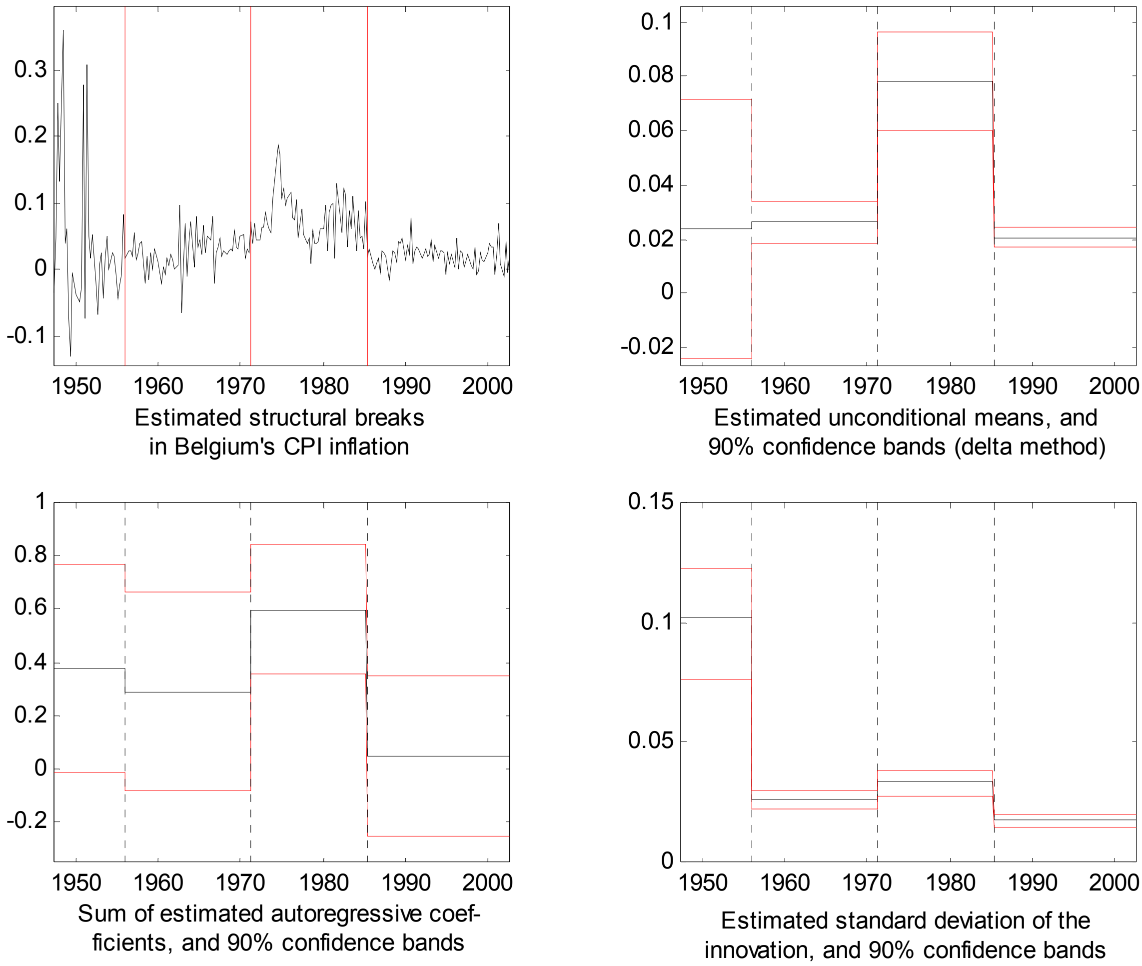

Figure 4: Belgium's CPI inflation (1947:1-2002:3), estimated structural breaks in the mean, innovation standard deviation, and AR coefficients

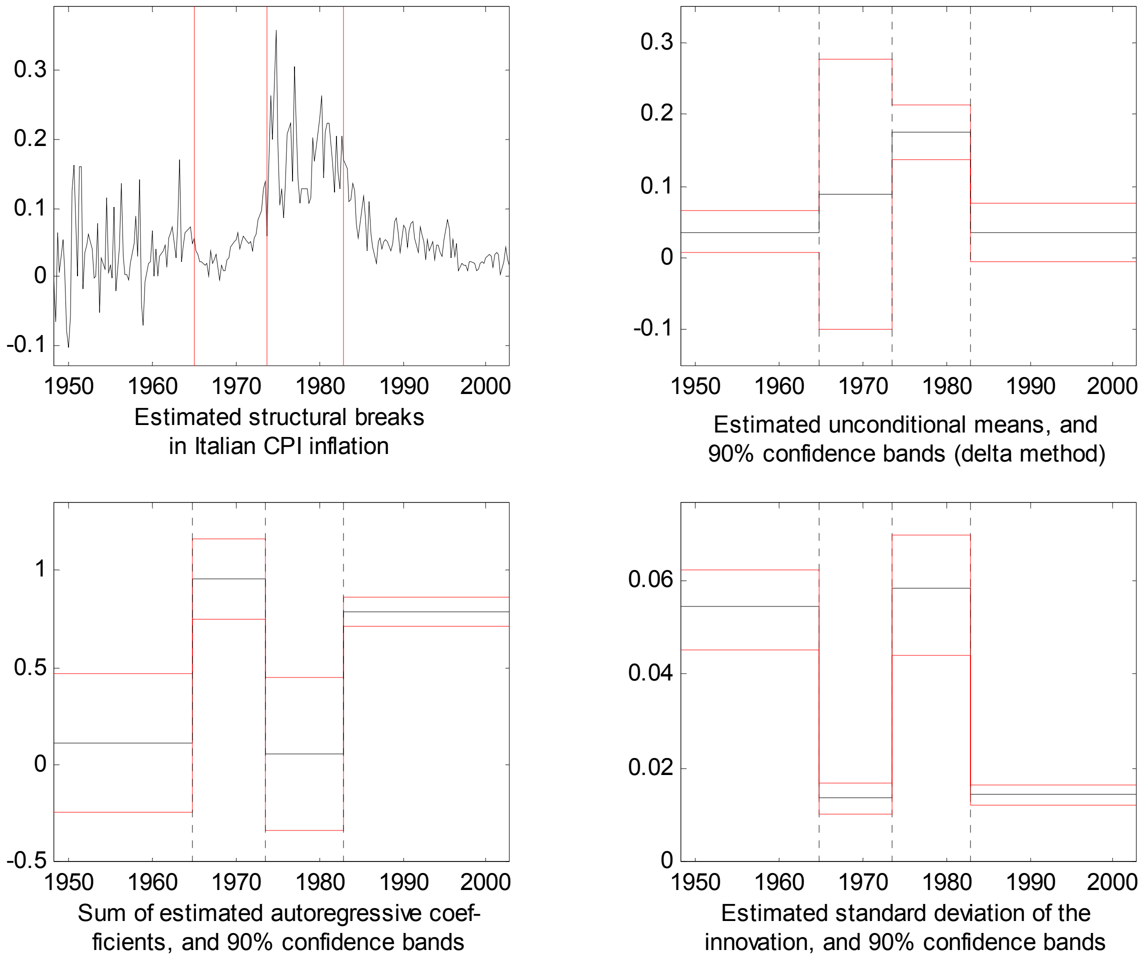

Figure 5: Italian CPI inflation (1948:1-2002:3), estimated structural breaks in the mean, innovation standard deviation, and AR coefficients

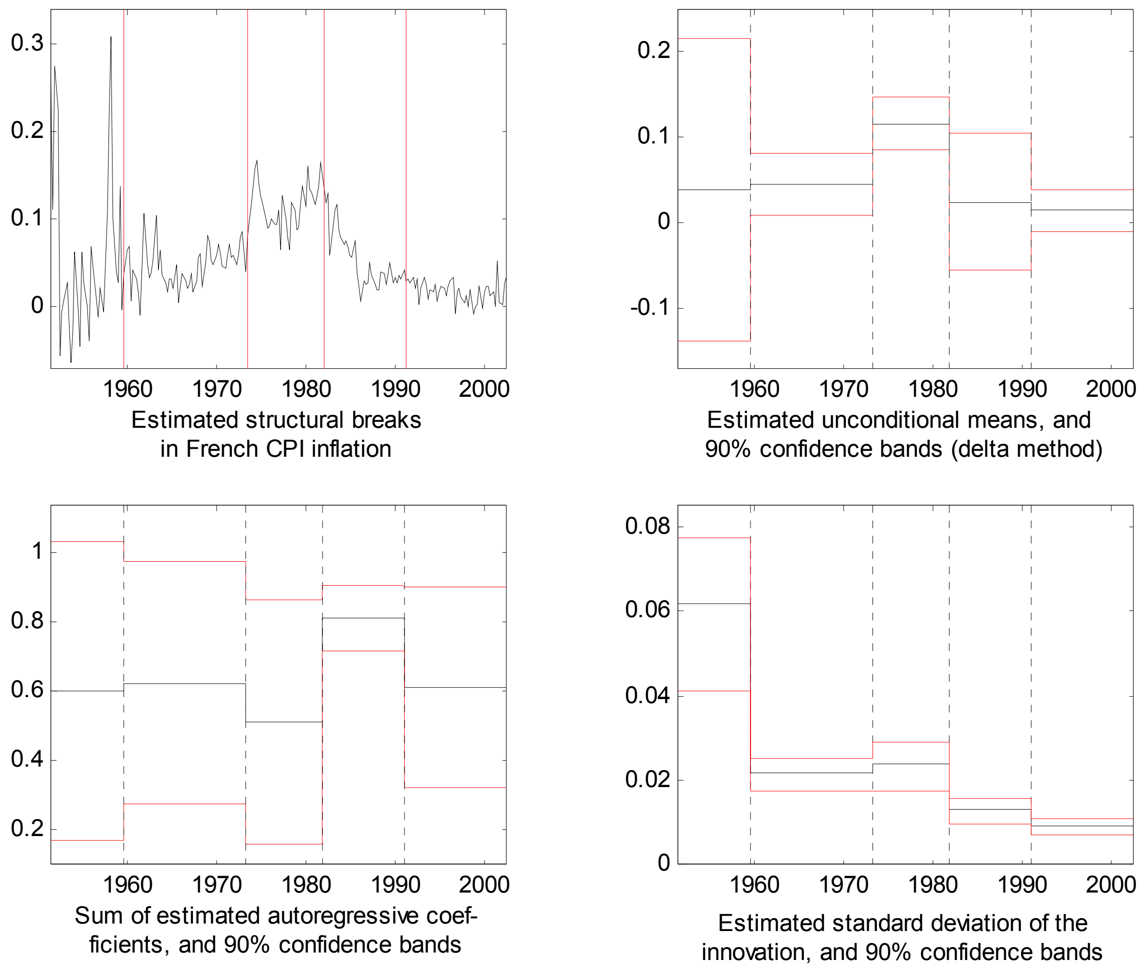

Figure 6: French CPI inflation (1951:2-2002:2), estimated structural breaks in the mean, innovation standard deviation, and AR coefficients

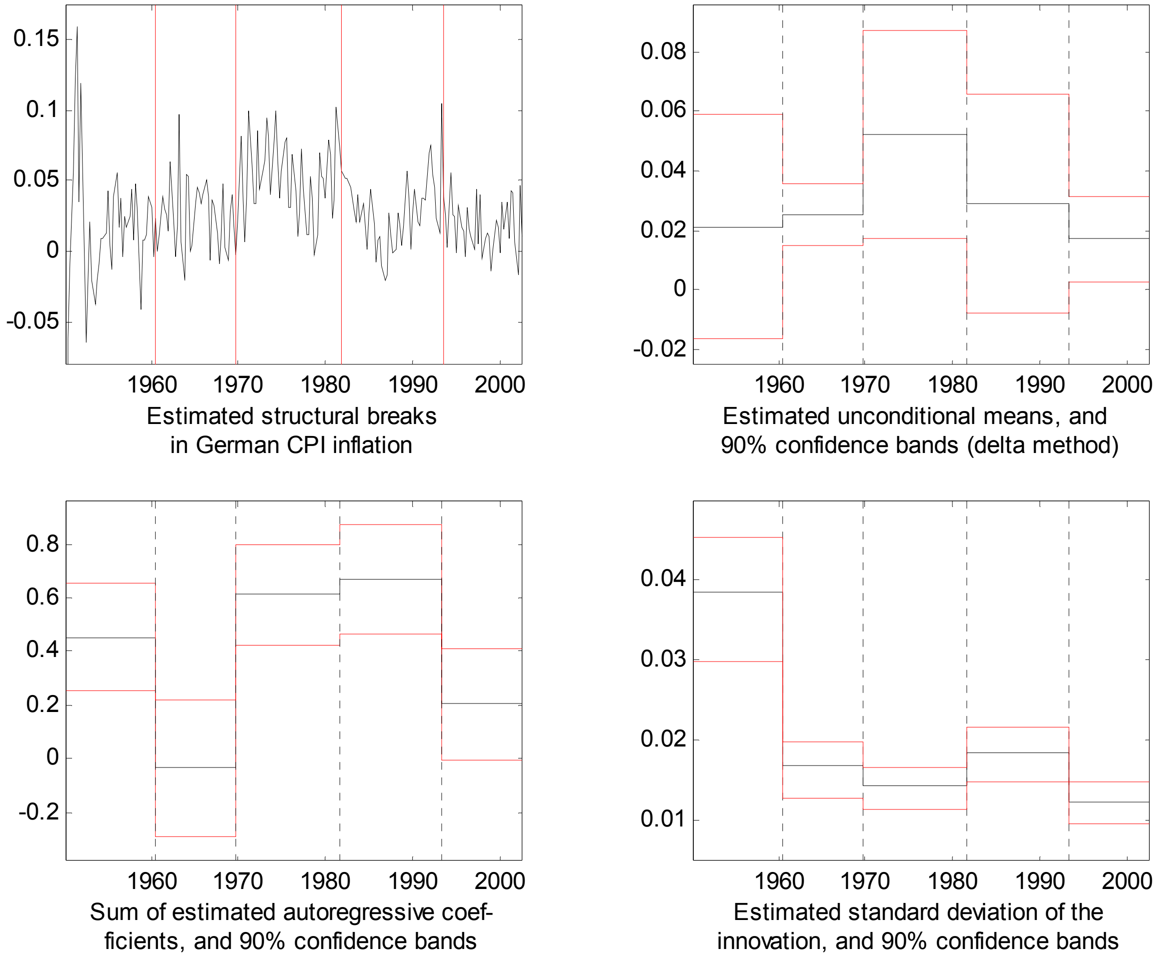

Figure 7: German CPI inflation (1950:1-2002:2), estimated structural breaks in the mean, innovation standard deviation, and AR coefficients

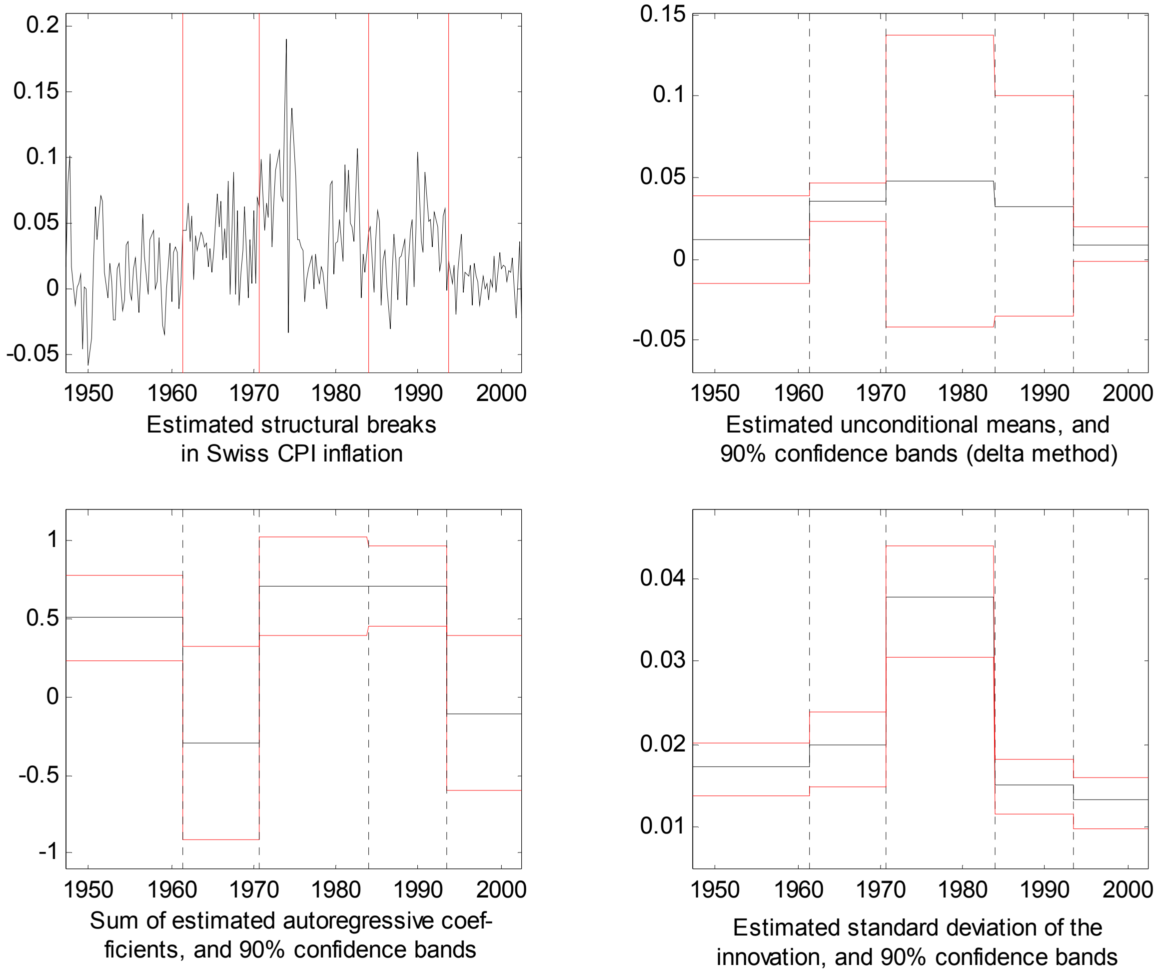

Figure 8: Swiss CPI inflation (1947:1-2002:2), estimated structural breaks in the mean, innovation standard

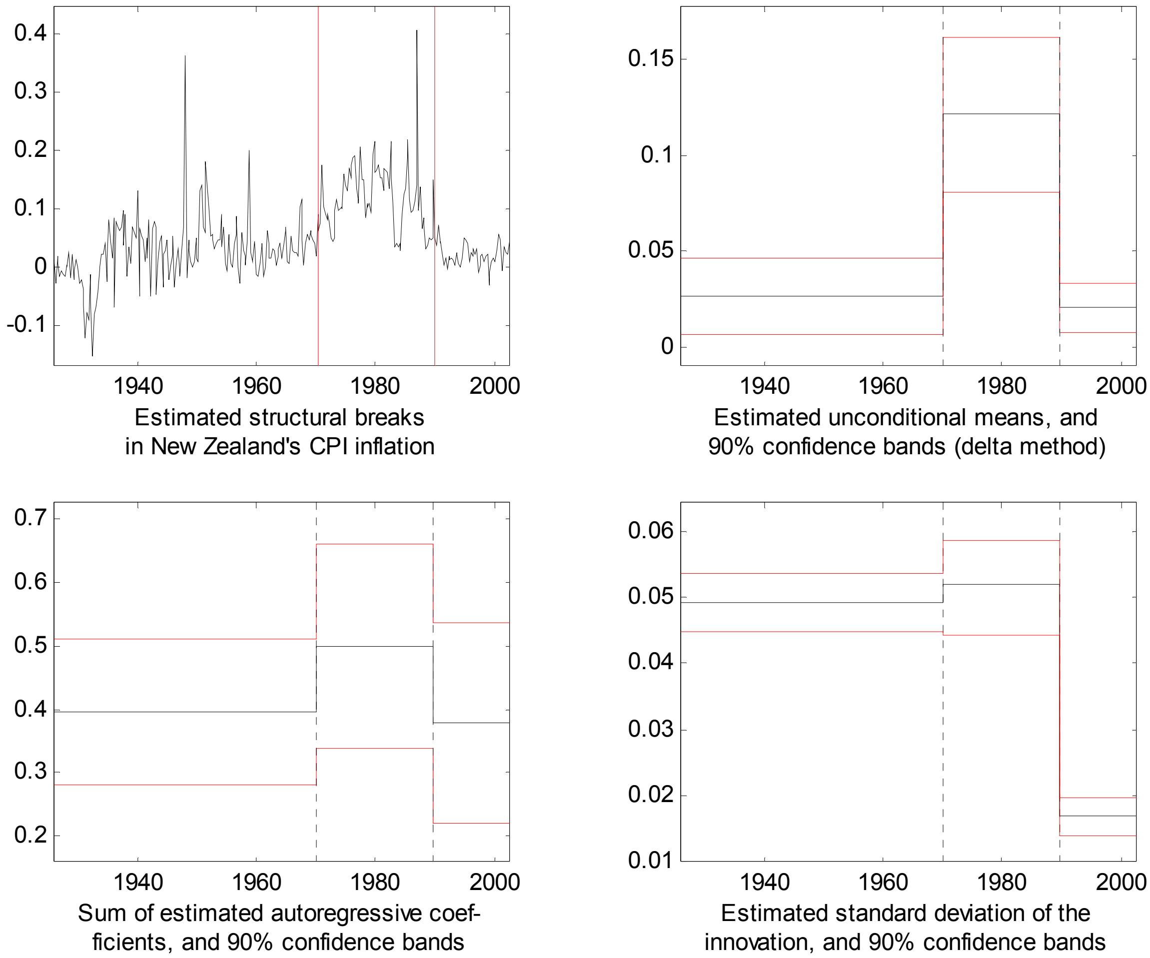

Figure 9: New Zealand's CPI inflation (1925:4-2002:2), estimated structural breaks in the mean, innovation standard deviation, and AR coefficients

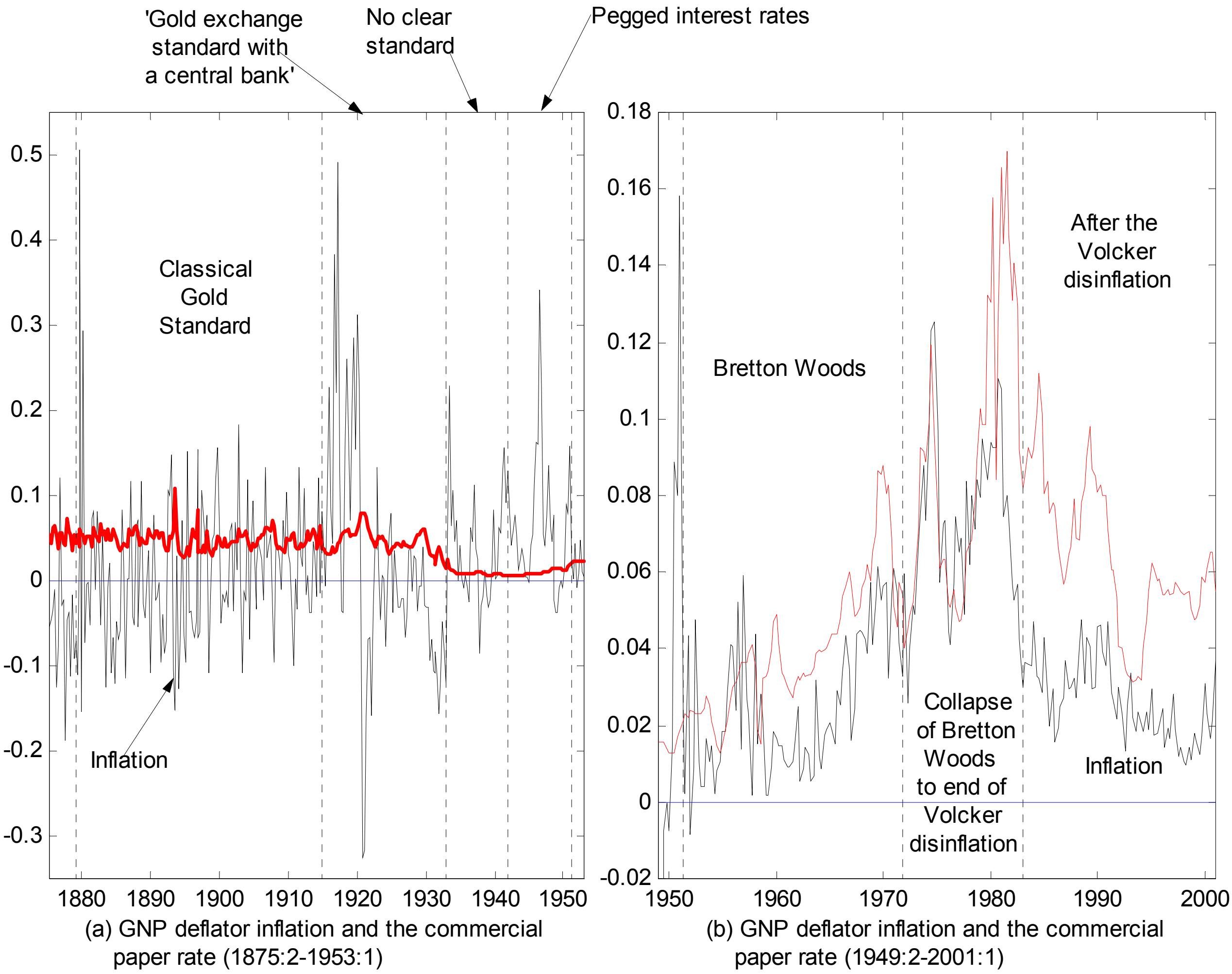

Figure 10: U.S. GNP deflator inflation and the commercial paper rate, 1875:2-2001:1

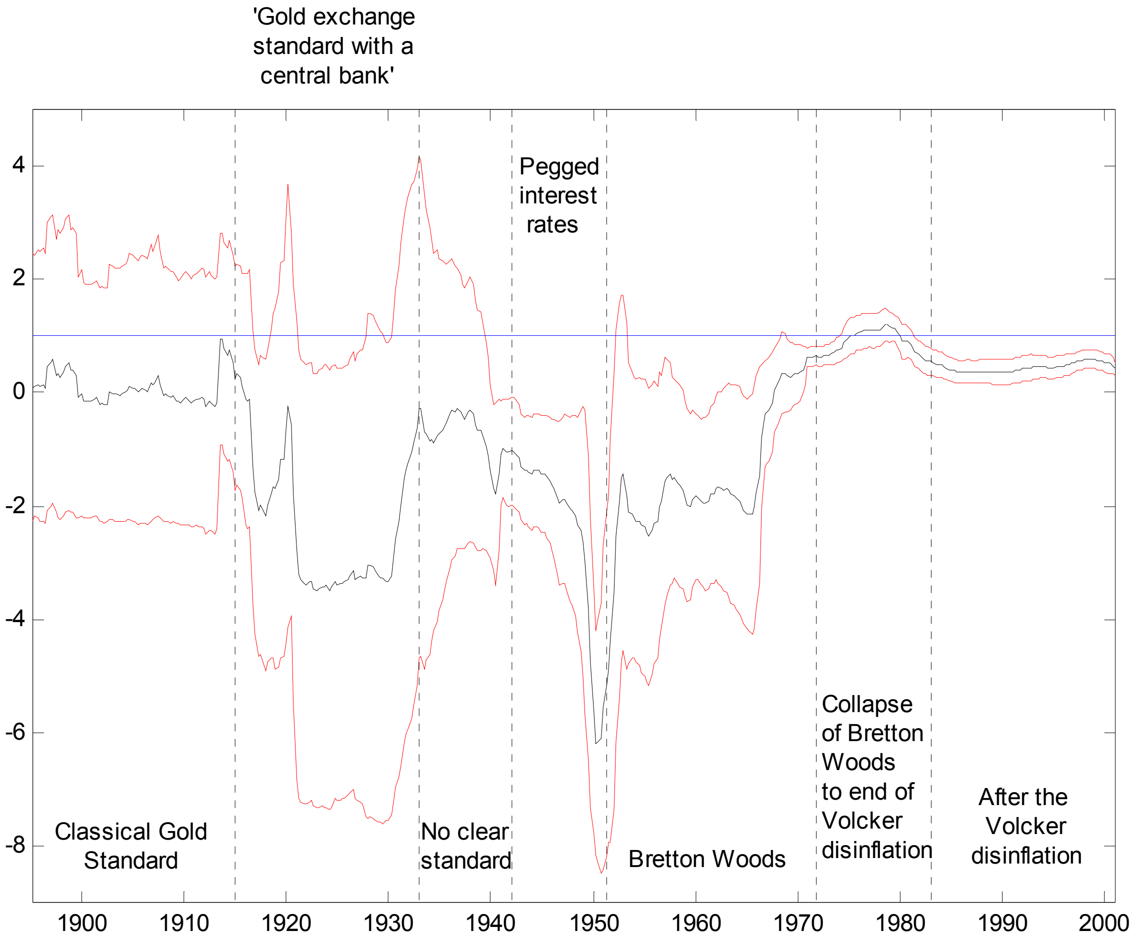

Figure 11: Evidence of time-variation in the extent of the Fisher effect in the U.S.: rolling estimates of beta from Fama (1976) regressions (rolling window: 20 years)