Contributions

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Dating Structural Changes in UK Monetary Policy

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Abstract: The UK historical monetary policy experience is rich of institutional changes, but it remains unclear which of these many events dominated the policy actions and what timing characterised the inception of different policy regimes. We develop a new empirical approach to answer these questions and we identify in particular the historical institutional events that effectively translated into a shift of the systematic actions of the UK monetary authorities. We find that not all institutional events triggered a contemporaneous change in the actual policy conduct, although a coherent evolution in phases is evident since 1978, when a significant monetary policy rule emerges. These occasional but not sporadic regime changes explain a considerable share of the movements in the official interest rate, as well as an overstatement of the importance of policy inertia.

Keywords: endogenous regressors, monetary policy rule, regime change, structural change, structural stability, United Kingdom

1 Introduction

The importance of having an accurate description of the historical evolution of monetary policy in a country is widely recognised in the macroeconomic literature. In particular, the timing with which this evolution took place is crucial information to be able to compare the performance of alternative monetary institutions, but also to understand the influence monetary policy can have on the economy. For the first objective, indeed, we need accurate dates to assign different periods to specific policy regimes in order to perform counterfactual simulations that highlights the role of a change in the stance of monetary policy. For the second objective, we

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need the same information to correctly isolate from the systematic policy function the exogenous policy shocks, which allow us to measure the response of the macroeconomic variables of interest to government intervention.

The UK history offers a unique setting to study how different institutional developments have shaped the monetary policy actions. In Table 1 we list a set of historical events that may have had an impact on the systematic policy decisions of the British monetary authorities. Also to reflect their different scope, we distinguish: important institutional events that are very likely to have influenced the overall monetary policy strategy (*broad institutional changes*); modifications that involved more operational aspects of the monetary policy decisions (*technical modifications*); and particular circumstances in the world economy (*international events*). This long list of events highlights one aspect that the existing empirical work has not fully addressed, which is the difficulty in dating, based only on *a priori* information, the periods in which the systematic actions of the monetary authorities were effectively different.

A great deal of empirical work has used simple policy instrument equations either to assess the importance of specific regime changes over certain historical periods, or to capture a generic and continuous random shift in the policy reaction function (see the literature review in the next section).[1] Little attention has been devoted to three important aspects that instead we address in this paper. First, it is important to have an accurate estimate of the dates in which an effective change in the stance of monetary policy took place, along with a rigorous statistical assessment of such evidence. This evidence would help us to identify the relative importance of different institutional events in terms of their ability to influence the actual policymaking. Second, it is undoubtedly of interest to identify in a robust fashion which of the macroeconomic priorities of the monetary authorities did effectively shift at each policy regime change. Third, because a simple policy instrument equation is an empirical approximation to a complex decision-making process that makes use of a small number of parameters, we do not know *a priori* when such an empirical approximation is adequate and when it is not.

To investigate these three aspects of the historical evolution of the monetary policy in the UK, we focus on the systematic actions of the monetary authorities which are typically captured by a policy function describing the endogenous policy response to the dynamics of the economy, and we develop a flexible empirical approach that builds on a set of tools provided by the econometrics of structural stability suitably adapted to our research question. This approach delivers three

1 We notice that for the UK there is a documented historical continuity in the use of the interest rate as the principal tool of monetary policy as commonly perceived by policy makers, academics and the wider public since the 1950s (Batini and Nelson 2005).

Table 1: List of historical events.

Broad institutional changes	
1976 Jul:	Introduction of monetary target on M3
1979 May:	Election of Margaret Thatcher
1986 Oct:	Big Bang
1987 Mar:	Shadowing the German Mark, as a prelude to enter ERM
1990 Oct:	Joining ERM
1992 Sep:	UK exit from ERM
1992 Oct:	Inflation targeting with RPIX band of 1–4%
1997 May:	BoE operational independence
2009 Mar:	Start of quantitave easing
Technical modifications	
1973 Dec:	Corset introduced
1979 Oct:	Abolition of exchange controls
1980 Mar:	Announcement of the medium term financial strategy
1980 Jun:	End of corset
1981 Aug:	Suspension of the official discount window and its minimum lending rate
1982 Mar:	Introduction of additional money supply targets
1985 Oct:	M3 money target abandoned
1993 Feb:	Introduction of the inflation report
1994 Apr:	Decision to publish the minutes of the monthly interest rate meetings
1995 Jun:	Chancellor announces an inflation target of 2.5% or less for the RPIX
2003 Jun:	Chancellor announces changeover from RPIX 2.5% to CPI target 2%
2003 Dec:	Change of inflation target: 2% for CPI (with a band of $\pm$1%)
International events	
1973 Oct to 1974 Mar:	Arab oil embargo
1976 Nov:	IMF loan to government after Sterling devaluation
1978 Nov to 1980:	Oil price shock, Iranian revolution and Iraq-Iran war
1979 Aug:	Volcker appointed Fed chairman
1987 Oct:	World stock market crash
2000 Mar:	Dot-com bubble burst
2008 Sep:	Financial crisis

ERM is the European Exchange Rate Mechanism; CPI is the Consumer Price Index; RPIX is the Retail Price Index excluding mortgage interest rate payments; Big Bang refers to the series of reforms that largely deregulated the London's financial sector.

main empirical findings: 1) many important historical events were not accompanied by a simultaneous shift in policy behaviour; 2) a clear evolution of the priorities of the monetary authorities emerges with inflation goals in particular becoming gradually more important from the 1970s up to early 2000s, and the output gap dominating after the 2008 financial crisis; 3) regime changes intended as shifts in the systematic actions of the monetary authorities explain a considerable share of the observed movements in the official interest rate, while disregarding such regime changes leads to an overestimation of the importance of policy inertia. The first two points in particular suggest that dating policy regimes based on individual institutional events can be misleading, and that instead a better characterization of the UK monetary policy is in terms of evolution in phases with the change in regime determined by the accumulation of many institutional events.

The empirical approach we propose consists of three steps. In the first step, we ask ourselves if there is any substantial evidence of a change in the long-run targets of the policy makers, and in their ability to influence private inflation expectations and short-term market interest rates. To this aim we perform stability tests on the mean of the variables representing the macroeconomic targets and on the term spread that distinguishes a short-term market rate from the policy rate under the direct control of the policy makers. In the second step, we explore the existence of changes in the systematic actions of the monetary authorities. With this objective we perform a comprehensive investigation of the parameter shifts in a standard monetary policy instrument rule, addressing in particular the problems produced by the presence of forward-looking variables and the possibility of structural changes in marginal distributions, two issues that can confound the whole estimation and inference on parameter change (Hansen 2000). To tackle these two issues, we employ the method developed by Hall, Han, and Boldea (2012) that hinges on a careful testing procedure to model multiple breaks at unknown location in the presence of correlation between regressors and the error term. Finally, in the third step, we characterize each estimated regime change by identifying which of the macroeconomic priorities did effectively undergo a revision. Here, we employ a modified version of the testing approach suggested by Inoue and Rossi (2011) to estimate the unstable subset of parameters.

Our approach presents several advantages over existing techniques that have been used to capture regime changes in monetary policy. Firstly, we highlight how treating the break dates as unknown parameters to estimate represents a clear advantage over an analysis that imposes the time location of the regime change based on *a priori* information. This latter practice, indeed, is exposed to the risk of being arbitrary and misleading because we often do not know for sure when exactly the announced institutional changes were effectively enforced, as a result of delays and anticipations, and also which specific events were statistically important for the actual policy actions. Secondly, because it can reasonably be argued that policy regime changes are occasional events, it seems appropriate to allow policy coefficients to shift a finite number of time within a sample of typical size to capture the sequence of phases that characterized the evolution of policy over time. While it is important to avoid imposing constancy of coefficients over the whole sample, on the other hand there is no need to assume they vary at every observations, as in a time-varying-parameter (TVP) framework. The transition from one regime to the other might occur in a gradual fashion in some circumstances, but if regime changes can in general be approximately described as discrete occurrences there is no need to assume a monetary policy rule that is continuously drifting over time. Moreover, compared to TVP models our approach is better designed to identify the timing with which subsequent regimes was established, especially by providing a measure of the statistical evidence for the estimated dates of change. This feature makes our approach more informative about the likely historical events behind these changes, but also facilitates the partition of history in different time intervals characterized by a specific regime, something which is useful for counterfactual policy simulations. In particular, by treating the dates of the shifts as unknown, we are also able to let the data select the specific time periods when a policy rule effectively existed and when instead it does not provide a satisfactory description of the policy actions. Finally, compared to regime-switching frameworks we do not restrict the regimes to a finite set of recurrent states since this would not be an appropriate assumption when the aim is to capture the possibly involved history of the policy instrument setting over a period almost as long as half a century. Changes in the general monetary policy strategy are better described as possibly unique and occasional events.

The time span we consider starts from the termination of the Bretton Woods system, a drastic institutional event which we take as an evident regime change, to recent years when the monetary authorities have had to rely on unconventional policy tools in a Zero Lower Bound (ZLB) environment. Because the actual policy rate has been held almost constant at 0.5% since 2009Q1, we employ for this period the shadow rate series constructed by Wu and Xia (2016), which can be regarded as the natural extension of the actual official bank rate in representing the monetary policy stance during ZLB periods. The resulting policy rate series allows us to explore the timing with which a new regime were established during the onset of the financial crisis, as well as the possibility that more than one regime have characterized the ZLB policy environment. As this rather long time span covers many potential episodes of change in the policy rule, we inevitably have to deal with small intervals containing few observations. Hence, even though Monte Carlo studies have demonstrated the good small-sample performance of the econometric tools we employ, a limitation remains in our inference about structural changes due to our relying on asymptotic critical values.

The structure of the paper is the following. In the next Section we provide a brief overview of how the existing empirical literature has addressed the issue of a changing monetary policy. In Section 3 we illustrate our methodology. In Section 4 we present the results from applying this methodology to the UK. In Section 5 we discuss the estimates in light of the UK history, and we compare them with results from previous studies. Section 6 offers some concluding remarks. More details about the results, the testing procedure and our implementation strategy are presented in the Supplementary Material.

2 Existing Work on Monetary Policy Rules

The fact that the monetary authorities' reaction function is likely to have been different across time has been typically addressed by presenting different estimates obtained after splitting the sample in specific segments that are justified by historical events. This is the approach pioneered by Clarida, Galì, and Gertler (2000), when they compare the responsiveness of the Federal Reserve under subsequent chairmen, with a particular focus on the period before and after the appointment of Volcker in 1979. To explore the influence of institutional changes of the Bank of England on its monetary policy, Adam, Cobham, and Girardin (2005) compare the coefficients of a forward-looking Taylor rule over subsequent subsamples delimitated by a series of events that are expected to be relevant for the monetary policy conduct. They observe a changing role played by domestic versus international factors in the interest rate setting, and a far stronger evidence of a substantial change in the policy reaction function after the introduction of operational independence in 1997, rather than with the adoption of the inflation targeting in 1992.

Nevertheless, the common approach based on sample splitting is subject to two major criticisms: there is a risk that the exact dates used to define the subsamples are arbitrary; we cannot be certain that the policy function is actually stable within the selected segments. As a consequence, this crude way of recognizing a changing value for the coefficients of the policy reaction function has been replaced by more statistically rigorous methods to capture either time-varying coefficients or non-linear responses of the interest rate to macroeconomic indicators.[2] A first available option is offered by Markov-switching models, a notable example of which is Sims and Zha (2006), who use alternative specifications of a multivariate model of the US economy to find that instability in the monetary policy rule had a far smaller influence on the macroeconomic fluctuations than the shifts in the error variance. Assenmacher-Wesche (2006) estimates a policy function for a set of three countries that includes the UK, employing a Markov-switching model where both the response coefficients and the error variance are subject to independent changes. An improved ability to describe the US history of monetary policy is shown by Castelnuovo, Greco, and Raggi (2014) to be the main feature of a policy equation where the inflation target changes, while parameters and the error variance are subject to a two-regime switching. Taylor and Davradakis (2006) fit a threshold model with two

2 Because an observed variation over time of the policy coefficients may reflect non-linear effects in the policy equation, the presence of this non-linearity has been also justified on the ground of theoretical models, where the objectives of the central bank are asymmetric with respect to positive and negative deviations (Surico 2007), or as a result of a non-linear relationship in the economy (Dolado, María-Dolores, and Naveira 2005).

regimes triggered by the level of inflation to underline the existence of a non-linearity in the monetary policy conduct of the UK. The interest rate setting displays an asymmetry, in the sense that an aggressive response to inflation deviations emerges only after it has significantly overcome its target. By contrast, using instrumental variables to estimate a threshold model for a forward-looking Taylor rule, Koustas and Lamarche (2012) find that they are not able to reject the hypothesis of a linear policy rule for the UK.

Even though regime-switching models improve the fit of estimated policy equations, there are reasons to believe that regime changes tend to occur in a gradual fashion, rather than in the form of abrupt shifts and some evidence seems to confirm this picture (e.g. Canova and Gambetti 2009; Koop, Leon-Gonzalez, and Strachan 2009).[3] Smooth Transition Regression Models have been used by Castro (2011) and Kharel, Martin, and Milas (2010) to study respectively the role played by a financial condition index in the UK and the European Monetary Union, and the response of the UK nominal interest rate to the real exchange rate. Boivin (2006) employs a Time-Varying Parameters (TVP) model to estimate a forward-looking version of the Taylor rule on US real-time data from 1969 to 1998, and discovers that the coefficients have been subject to changes that are gradual, in particular around the appointment of Volcker in 1979, and often non synchronised. He infers then that previous conflicting results on the same period must be due to the failure of simple models that impose a single regime change at a fixed date to capture the actual complex pattern of the monetary policy actions. Same econometric technique is applied by Trecroci and Vassalli (2010) to a group of countries that includes the UK, confirming the improved statistical performance of a time-varying Taylor rule over the classical fixed-parameter specification. After developing a procedure to address the problem of endogeneity in time-varying parameter models, Kim and Nelson (2006) employs it to estimate a monetary policy rule for the US, while Kishor (2012) applies the same approach to a set of countries that includes the UK.

3 Methodology

As anticipated in the Introduction, the empirical approach we propose consists of three steps. In the first step we look for evidence within our sample of possible shifts in the implicit long-run target of the monetary authorities, as well as changes

3 For instance, the need for a transparent and predictable monetary policy conduct favors a practice in which sudden shifts in the responsiveness to macroeconomic variables are avoided. This is very likely the case of the adoption of an inflation targeting regime, which often has been anticipated by large reductions in the inflation rate before its official implementation.

in their procedure to control market short-term interest rates. In the second step, we perform a comprehensive investigation of possible changes in the systematic actions of the policy makers by examining a standard policy instrument rule. Finally, in the third step, we identify more clearly which policy objectives have shifted in correspondence with each estimated regime change.

In all stability tests we rely on the asymptotic distribution of the test statistics, which is inevitably an imperfect approximation of the actual finite-sample distribution, especially in small subsamples like the ones we end up with in the second step as a result of the relatively numerous episodes of structural change we obtain. This implies an inevitable risk of overfitting breaks due to size distortions, as well as underfitting breaks due to lack of power. It is important to mention this limitation of the inference we produce in the second step, although the econometrics papers developing these tests document satisfactory finite-sample performance. Moreover, to increase test power and decrease size distortion, we devote particular attention on some specific technical choices during the execution of the tests.[4]

3.1 Changes in Targets and Procedures

While it can be argued that other macroeconomic aggregates have been targeted by the monetary authorities in specific historical intervals, it is safe to claim that stabilization of the inflation rate has represented an important goal for most of the period under analysis. An estimate of the implicit inflation target can be obtained from the value of the intercept in a monetary policy rule if we have an estimate of the equilibrium real interest rate. However, the possible changes over time in the real interest rate and the potential presence of omitted variables, which represent policy factors that were important in the decision-making of the monetary authorities in particular historical circumstances, makes the interpretation of the level of such intercept rather dubious. An alternative method is to consider the actual level of the inflation rate over the middle-long horizon, which reflects the policy makers' inflation target perceived by the private sector if the economic system admits a unique stationary solution. We adopt this latter method and derive the implicit long-

4 To ensure that we do not miss existing structural breaks, we rely on tests for multiple breaks rather than single ones, we set a sufficiently high trimming, and we stop the sequential testing procedure to estimate the number of breaks only if the segment under analysis does not have enough observations to ensure parameter identification. To avoid over-estimation of breaks, we analyse ex-post whether there has been a change in the error variance that might have produced a spurious break in parameters, and we give precedence to F-type tests over Wald-type tests as this latter is based on a general covariance matrix that we know is poorly estimated in small samples. See the Supplementary Material for more details.

run inflation target of the monetary authorities from the sample average of the observed inflation rate. Even if the value of such average may not represent exactly the target level, it is still likely to signal any sizeable shift that occurred over time in the target set by the policy makers and its credibility to the public.

Testing for a structural change in the level of a univariate series is strictly related with the determination of the nature of its trend, and various tests have been proposed in the literature. One way to go is to first run a standard ADF test to ascertain the statistical nature of the trend of inflation; then, if such test fails to reject the null, explore the possibility that this outcome is the consequence of a break in the trend function of inflation. With this aim we employ the method of Perron and Vogelsang (1992) to test the unit root hypothesis allowing for one break in the level of inflation at an unknown location. If we reject the null, we conclude for the existence of a shift in the implicit long-run target for inflation at the time location estimated by the test.[5]

The second aspect that we explore is the link between the official bank rate, which is the most direct measure of the policy decisions of the monetary authorities, and the short-term market interest rates, as represented for instance by the three-month rate on Treasury Bills. The spread between these two interest rates reflects an expectations and a term premium components, which are possibly driven by a multitude of factors, but a substantial shift in this spread may reflect in particular a change in the ability of the monetary authorities to control the market interest rates, or the consequences of a modification in its operational procedures. We apply the test for multiple breaks of Bai and Perron (1998) on an autoregressive model of the spread to verify the possible existence of such a change and to estimate the date of its occurrence.

3.2 Changes in Systematic Actions

We define as regime change in monetary policy a significant shift in the parameters of a policy instrument rule (or Taylor-type of rule). In the following we describe the exact specification of the policy rule equation and the estimation approach. Further technical details on our implementation strategy can be found in the Supplementary Material.

5 An alternative option is to run the test of Harvey, Leybourne, and Newbold (2004) for a break in the level, which is robust to the order of integration.

3.2.1 Model Specification

The monetary policy rule is a standard equation that defines the short-term nominal interest rate as a linear function of the deviations of the expected future inflation and output gap from their respective targets

$$r_t = (rr^\star + \pi^\star) + \beta_2[E(\pi_{t+n}|\mathscr{F}_{t-1}) - \pi^\star] + \beta_3\left[E\left(y_{t+n}\middle|\mathscr{F}_{t-1}\right) - y^\star\right] + \varepsilon_t, \tag{3.1}$$

where π_{t+n} and y_{t+n} indicate the inflation rate and the output gap at time $t + n$, ε_t is an error term, $E(\cdot|\mathcal{F}_{t-1})$ denotes the expectations conditional on the available information set that contains all past values up to $t - 1$, rr^* is the equilibrium real interest rate, and the remaining starred variables represent the target values for the monetary authorities. Following standard practice we assume $y^* = 0$, and we set $n = 0$ as a compromise between a backward and a forward looking specification, which are more suitable to describe the policy actions respectively in the first part of the sample and after the introduction of inflation targeting.

Under a rational expectations approximation for the forecasts of the policy makers, we adopt the usual errors-in-variables approach that replaces the expected values with the actual ones minus a measurement error: $E(\pi_t|\mathcal{F}_{t-1}) = \pi_t - \eta_t$, $E(y_t|\mathcal{F}_{t-1}) = y_t - v_t$.[6] As a result, the equation to be estimated is

$$r_t = \beta_1 + \beta_2\pi_t + \beta_3 y_t + u_t, \tag{3.2}$$

where $\beta_1 = rr^\star + (1 - \beta_2)\pi^\star$ and $u_t = \varepsilon_t - \beta_2\eta_t - \beta_3 v_t$.

Since such a specification is likely to produce serially correlated residuals, two alternative approaches have been proposed to model such persistence, one explicitly considering the monetary authorities preference for a gradual adjustment of the policy rate, and one highlighting the effects of persistent omitted factors. While we do not deny the importance of both sources of interest rate persistence, we opt for the second approach for two reasons. First, we prefer to sacrifice a detailed description of the dynamics of the interest rate setting for a more parsimonious specification that does not undermine the power of our stability tests, especially considering the potential small size of the intervals between

6 As for the output gap, an alternative approach is to use the forecasts that the policy makers were actually making at the time they set the interest rate (real-time data), but this type of data is not available for the period we analyse. Moreover, it is very likely that the full sample information we use to estimate the policy makers expectations largely offsets the lack of other sources of information that the policy makers had available but that clearly cannot be included in a simple policy equation. As argued in Adam, Cobham, and Girardin (2005), calculating the output gap using full-sample data may provide a more accurate measure of what the policy makers thought at the time than using real-time data, which carry the risk of understating their ability to recognize which phase of the cycle the economy was in.

break dates. Second, we are interested in exploring the possibility of an additional third source of persistence, the one that emerges as a result of neglected shifts in the policy coefficients, especially in the presence of persistent regressors, as different monetary regimes are established over time.[7] To this aim we exploit the size of the residual correlation to assess the consequences of ignoring the existing breaks in policy for the observed policy inertia. To evaluate, on the other hand, the reverse risk of spurious breaks resulting from autocorrelated errors we will also examine the outcome of tests that are robust to autocorrelation.

3.2.2 Estimation Procedure

Estimation and testing of multiple structural breaks in equations where regressors are potentially correlated with the error is not a straightforward task. Indeed, a standard GMM approach is not a valid choice when we are interested in uncovering the location of the break points.[8] Hall, Han, and Boldea (2012) and Boldea, Hall, and Han (2012)[9] developed an econometric procedure to consistently estimate number and location of breaks in parameters in the presence of endogenous regressors. The validity of their procedure is established both in terms of asymptotics and in finite samples by a Monte Carlo simulation.

Their approach is based on 2SLS and consists in performing a distinct analysis of the stability of the first and second stage of the 2SLS estimation. If the reduced-form system in the first stage turns out to be stable, inference about instability in the structural equation can proceed using the limiting distributions calculated in Bai and Perron (1998).[10] If the reduced-form is unstable, on the contrary, such distributions of the relevant test statistics cannot be used, and calculating the fitted values of the regressors ignoring the breaks in the reduced-form can lead to seriously misleading results, where these breaks emerge from the stability tests as if they were breaks associated with the parameters of the structural equation. To

7 Such a possibility was suggested by Rudebusch (2002), and is similar to the argument claiming that the apparent policy inertia is due to policy responding to other persistent omitted factors (English, Nelson, and Sack 2003; Rudebusch 2002, 2006). Evidence in favour of an explicit interest rate smoothing component in the US monetary policy is provided by Castelnuovo (2003, 2006), and more recently by Coibion and Gorodnichenko (2012). A small Monte Carlo simulation we ran on a bivariate model subject to one break and a persistent regressor showed that the residual autocorrelation from ignoring the break is a function of the magnitude of the parameter change as well as the regressor's degree of persistence.

8 Hall, Han, and Boldea (2012) prove that under general conditions a GMM estimate of the break point is inconsistent, and they show that this outcome arises from the particular structure of the very criterion function that is at the core of the GMM estimation method.

9 HHB hereafter.

10 BP hereafter.

prevent the breaks in the reduced-form from contaminating the break inference in the structural equation HHB propose to split the sample in segments that have a stable reduced-form, and then perform the break estimation within each of these segments after the fitted values have incorporated the estimated reduced-form instability.[11]

In our case, the structural equation with m asymptotically distinct breaks can be written as

$$r_t = \beta_{1,j} + \beta_{2,j}\pi_t + \beta_{3,j}y_t + u_t, \qquad j = 1, \dots, m+1, \qquad t = T^*_{j-1}+1, \dots, T^*_j \tag{3.3}$$

where, by convention, $T^*_0 = 0$ and $T^*_{m+1} = T$, which is the sample size. We assume that π_t and y_t are potentially correlated with the error term u_t. The reduced form system estimated in the first stage regression is characterized by h asymptotically distinct breaks, and can be written as

$$\pi_t = z_t^{'}\Delta_{\pi,i} + v_t^{\pi}, \;\; i = 1, \dots, h+1, \;\; t = T_{i-1}+1, \dots, T_i \tag{3.4}$$

$$y_t = z_t'\Delta_{y,i} + v_t^{y}, \tag{3.5}$$

where $T_0 = 0$, $T_{h+1} = T$, z_t is a $q \times 1$ vector of instruments that includes an intercept, and we define $x_t = (\pi_t, y_t)$.

We extend the HHB procedure by adding two additional steps, one consisting in a refinement of the estimates of the policy coefficients by performing a final partial-sample GMM estimation, and one in which we verify *ex-post* that the detected episodes of parameter instability are not the result of shifts in the error variance. As a result, our extended procedure consists of the following seven steps.

1. Perform a break estimation on the two reduced-form Eqs. (3.4) and (3.5) using the Least Squares method of BP.
2. Divide the whole sample into $\widehat{h}+1$ stable segments, and calculate the fitted values $\widehat{x}_t$ incorporating the estimated reduced-form breaks.
3. In each stable segment, perform a break estimation on the structural Eq. (3.3) using the HHB method. The resulting estimated break dates are those idiosyncratic to the structural equation.
4. Test each estimated reduced-form break to ascertain whether it is significant also in the structural equation (common break).

11 An alternative option to the HHB procedure is to follow the suggestion of Perron and Yamamoto (2013), which amounts to use the OLS approach of BP to estimate the break dates ignoring the correlation between regressors and the error term, and then, conditionally on these dates, estimate the regression parameters using IVs. While this approach ensures gains in efficiency, it is vulnerable to the risk that a break in the marginal distribution of the regressors or in their correlation with the error term produces a spurious break in the policy equation, or, conversely, offsets the evidence of a shift in the policy coefficients hiding an existing genuine break.

5. Estimate the parameters of Eq. (3.3) by 2SLS, conditional on the set of all significant breaks.
6. Test for heteroskedasticity and autocorrelation, and in case of rejection perform a partial-sample GMM estimation of Eq. (3.3) conditional on the estimated set of breaks.
7. Check for the presence of shifts in the error variance of Eq. (3.3) by running a Goldfeld-Quandt test.

The tests for stability and the sequential test to estimate the number of breaks in step 1 are those proposed in BP and are formally defined in the Supplementary Material. Once we have defined the set of stable segments from step 2, we analyse each of them separately. The fitted values are defined as

$$\hat{x}_t = z_t'\hat{\Delta}_i = z_t'\left(\sum_{t=\hat{T}_{i-1}+1}^{\hat{T}_i} z_t z_t'\right)^{-1} \sum_{t=\hat{T}_{i-1}+1}^{\hat{T}_i} z_t x_t \tag{3.6}$$

$$i = 1, \ldots, \hat{h}+1, \quad t = \hat{T}_{i-1}+1, \ldots, \hat{T}_i,$$

where $\hat{\Delta}_i = [\hat{\Delta}_{\pi,i}, \hat{\Delta}_{y,i}]$, while the second stage equation for segment i is

$$r_t = \beta^i_{1,j} + \beta^i_{2,j}\hat{\pi}_t + \beta^i_{3,j}\hat{y}_t + \tilde{u}_t, \;\; j = 1, \ldots, m_i+1, \;\; t = T^\star_{j-1}+1, \ldots, T^\star_j, \tag{3.7}$$

with $T^\star_0 = \hat{T}_{i-1}$, $T^\star_{m_i+1} = \hat{T}_i$, m_i indicating the number of breaks in segment i, and $\tilde{u}_t$ representing the second stage residuals evaluated at the true parameters value. Eq. (3.7) is first tested for stability over segment i using the second stage equivalent of the BP stability tests. If we reject the null of stability, we proceed to estimate the number of breaks using either an information criterion or a sequential testing procedure. To estimate the number of breaks using a sequential testing procedure, we make use of the following test statistic[12]

$$F_{\overline{T}_i}(l+1|l) = \max_{1\le j\le l+1}\left\{\frac{SSR_l\left(\hat{T}^\star_1, \ldots, \hat{T}^\star_l\right) - \inf_{\tau\in\Lambda_{j,\eta}} SSR_{l+1}\left(\hat{T}^\star_1, \ldots, \hat{T}^\star_{j-1}, \tau, \hat{T}^\star_j \ldots, \hat{T}^\star_l\right)}{\hat{\sigma}^2_j}\right\}, \tag{3.8}$$

where l indicates the step in the sequence, $\overline{T}_i = T_i - T_{i-1}$ is the size of segment i,

$$SSR_l\left(\hat{T}^\star_1, \ldots, \hat{T}^\star_l\right) = \sum_{j=1}^{l+1} \sum_{t=\widehat{T^\star_{j-1}}+1}^{\widehat{T^\star_j}} \left(r_t - \hat{\beta}^i_{1,j} - \hat{\beta}^i_{2,j}\hat{\pi}_t - \hat{\beta}^i_{3,j}\hat{y}_t\right)^2, \tag{3.9}$$

12 For the sake of brevity, we show here the definition only of the F-type tests, while the general formulae for the corresponding *Wald*-type tests can be found in HHB.

$$\widehat{\sigma}_j^2 = \sum_{t=\widehat{T}_{j-1}^{\star}+1}^{\widehat{T}_j^{\star}} \left(r_t - \widehat{\beta}_{1,j}^i - \widehat{\beta}_{2,j}^i \widehat{\pi}_t - \widehat{\beta}_{3,j}^i \widehat{y}_t \right)^2 \Big/ \left(\widehat{T}_j^{\star} - \widehat{T}_{j-1}^{\star} - 3 \right), \tag{3.10}$$

and $\widehat{T}_{l+1}^{\star} = \widehat{T}_i$, while the set of admissible partitions considered at each step is

$$\Lambda_{j,\eta} = \left\{ \tau : \widehat{T}_{j-1}^{\star} + \left(\widehat{T}_j^{\star} - \widehat{T}_{j-1}^{\star} \right) \eta \le \tau \le \widehat{T}_j^{\star} - \left(\widehat{T}_j^{\star} - \widehat{T}_{j-1}^{\star} \right) \eta \right\}. \tag{3.11}$$

The sequential procedure is the following. We impose the break date that maximises the F test for one break, and we test for an additional break using the conditional sup$F(2|1)$. If we reject, we impose the first and the second estimated breaks, and we test for an additional break using the conditional sup$F(3|2)$. This procedure continues, every time adding the dates that sequentially maximise the sup$F$ statistic, until we fail to reject or we reach the maximum number of estimable breaks. At that point, the estimated number of breaks is given by the number of rejections. Once we have determined the number of breaks $\widehat{m}_i$, we estimate their location using the dates that minimised the sequential F statistic at each step. To verify whether each of the estimated reduced-form breaks are significant also in the policy rule equation we apply the Wald test for common breaks constructed by HHB (see their Theorem 9). The union of all the idiosyncratic breaks found in every segments, along with the reduced-form breaks that turned out to be significant also in the second stage, constitutes the total set of $\widehat{m}$ breaks of the structural equation.

Finally, we explicitly deal with the possibility of non-spherical errors. We introduce a refinement of the HHB procedure in the case there is substantial evidence of heteroskedasticity or autocorrelation by implementing an additional stage where we apply the partial-sample GMM estimator of the regression parameters (see Andrews 1993), conditional on the estimated set of break dates obtained from the HHB procedure. Moreover, we address the limitation that the HHB procedure does not explicitly consider breaks in the error variance. Pitarakis (2004) has underlined that the break inference can be substantially distorted if there are changes in the error variance that are being ignored during the testing, a problem that cannot be solved by resorting to robust statistics since these perform poorly in small samples. The risk is to detect spurious breaks in parameters due to size distortions, or disregard existing ones because of the low power produced by changes in the error variance.[13] For this reason we test

13 Lubik and Surico (2010) show the advantages in terms of power of using heteroskedasticity-robust tests when testing for breaks a reduced-form model fitted on data generated by a DSGE model subject to a break in monetary policy.

whether there are significant signs of changes in the error variance between each pair of adjacent segments using the Goldfeld-Quandt statistic.[14]

3.3 Changes in Priorities

The outcome from the previous estimation procedure will produce a set of $\hat{m}$ break dates, which defines $\hat{m}+1$ monetary policy regimes. Because the whole approach is devised to estimate pure structural breaks, we still do not know which specific parameters of the policy rule equation did significantly shift in the transition from one regime to the other.[15] It is important to identify the unstable subset of parameters in each estimated break date for two reasons. Firstly, it is interesting to know whether the regime change involved a different attitude of the policy makers with respect to one policy goal rather than the other or both of them. We recognize, though, that such a change in the policy stance can reflect either a shift in the policy makers preferences or the optimal policy adjustment to a structural change in the parameters of the economy. Secondly, it is also important to verify whether the source of the detected instability is merely a shift in the intercept because this would imply a more difficult interpretation of the estimated regime change and a risk that the observed parameter variation is in fact just a symptom of model misspecification due to omitted variables.

We apply a modified version of the method developed by Inoue and Rossi (2011) to our set of estimated break dates $\{\hat{T}_j^\star\,; j=1,..,\hat{m}\}$ for the policy rule equation.[16] More precisely, we implement the following algorithm in five steps to identify the stable subset of parameters for all estimated break dates:

1. select $j=1,..,\hat{m}$;
2. for each individual parameter β_i, with $i=1,2,3$, calculate the p-value of the test statistic for one partial break at $\hat{T}_j^\star$, conditional on one break in the remaining parameters, using the segment $(\hat{T}_{j-1}^\star, \hat{T}_{j+1}^\star)$;

14 Perron and Zhou (2008) underline the size distortions present in this procedure when estimating the break dates in the first step, and propose instead a joint approach based on a quasi-likelihood ratio test.

15 The HHB approach can be modified to consider structural breaks that involve only one specific subset of parameters (partial structural break model), but this subset would be constant over time and selected based on *a priori* information.

16 Inoue and Rossi (2011) show that their algorithm estimates the true subset of stable parameters with asymptotic probability equal to $1-\alpha$ if the break is partial, and equal to 1 when the break is pure, with α indicating the significance level. A satisfactory finite sample performance is confirmed by them via a Monte Carlo Simulation.

3. drop the parameter with the lowest p-value and test for stability the remaining set of parameters, conditional on one break in the dropped parameters, using the segment $(\widehat{T}^{\star}_{j-1}, \widehat{T}^{\star}_{j+1})$;
4. if we reject at step 3, repeat the same step until we fail to reject or the set of parameters is empty;
5. repeat steps 1–4 for all j.

As for the test statistic we use the F-type of test, coherently with our choice in the HHB procedure.[17] To verify the robustness of the results from the Inoue-Rossi method we compare them with the subset of unstable parameters estimated by two information criteria, BIC and HQ, calculated over all possible combinations of stable and unstable parameters.

4 Results

We now present the results we obtained from implementing the methodology described in the previous section on UK data, with a focus on the main relevant technical aspects. An economic interpretation of these results, along with a discussion of the historical context, is offered in the subsequent Section 5. Several of the more technical details of the estimation are shown in the Supplementary Material.

4.1 Univariate Analysis

The data we use for our estimation have quarterly frequency, from 1972Q3 to 2016Q2, and consist of: the annual inflation rate calculated using the last month of the quarter of the Consumer Price Index (CPI); the official interest rate that the Bank of England charges on secured overnight lending; the shadow bank rate calculated by Wu and Xia (2016); the interest rate on the three-month Treasury Bills; the output gap, calculated as the residuals from the regression of real Gross Domestic Product (GDP) in logarithms on a deterministic linear-quadratic trend. [18,19]

17 See discussion in the Supplementary Material. Since the break date is fixed, being the result of the HHB procedure on pure breaks, we refer to the standard F distribution for calculating the p-values.

18 The CPI is the typical measure of aggregate price dynamics and it is strongly correlated with the RPIX, which was the official target of the central bank in the period 1995–2003.

19 Although the output gap can be calculated using many different criteria, we follow an approach that is fairly common in applied work allowing a closer comparison with existing results (e.g. Adam, Cobham, and Girardin 2005; Clarida, Galì, and Gertler 1998; Cobham 2006).

We choose the official bank rate as dependent variable in the policy rule equation because this is the most direct measure of the policy decisions of the British monetary authorities. Any short-term market interest rates, which should strictly speaking be considered as "operational targets", may introduce the influence of other factors like additional noise from market arbitrage activities or changes in the technical operating procedures of the central bank, with the risk of confounding our stability testing and break estimation with a measurement error. To assess whether this distinction is effectively important we examine the spread between the official bank rate and the three-month interest rate on Treasury bills, and we find that its correlation with the inflation rate is highly significant, and more importantly we obtain strong evidence of a break in 1980Q3 at 1% significance level using a *Wald* version of the BP test on an autoregressive model for the spread. This break point highlights a substantial downward shift in the level of the spread, from 0.51 to 0.11, which might be explained by the institutional changes that were taking place in those years and that were likely to affect both the strategy and the practice of the monetary authorities.[20]

When we examine the possibility of a shift in the long-run mean of the inflation rate, we obtain that while the ADF tests are clearly unable to reject the null of unit root, the outcome is opposite when we allow for a break in the trend function. The PV test supports the stationarity of inflation once a shift in the level of the series is modelled, which the same procedure estimates to occur at 1980Q2.[21] Hence, we conclude that in 1980Q2 a significant shift occurred in the monetary authorities' implicit inflation target perceived by the private sector, followed one quarter later by a sizeable reduction in the level of the interest rate spread. In addition to the changes in the operating procedures of the central bank taking place in those years, this break coincidence also suggests an improved ability of the monetary authorities to control market interest rates that is likely related to a stronger influence on private expectations, reflected by the fall in the long-run equilibrium level of inflation.

Following the financial crisis since 2009 the monetary policy in the UK has been characterized by the Zero Lower Bound (ZLB) and the resort to unconventional policy tools by the central bank. Though the official bank rate has been kept constant at 0.5% for most of the time since then, it remains possible that more than one regime have characterized the stance of the monetary authorities with respect to its macroeconomic goals. To explore this possibility we employ a "shadow rate" series, which has been advocated as a valid measure to summarize the stance of the monetary

20 In particular, a greater recourse to open market operations rather than discount window lending is decided in November 1980, which is likely the main reason the policy rate became closer to the interbank market rates.

21 The PV statistic is –4.91, which is significant at 2.5% (critical value is –4.74 from PV, Table 1, for T = 150 and k = 0).

policy when its actions are constrained by the ZLB (Krippner 2012; Wu and Xia 2016). Wu and Xia (2016), in particular, show that their measure extracted from a multifactor shadow rate term structure model displays the same dynamic correlations with the relevant macroeconomic variables as the actual policy rate under normal times, and so can be treated as its natural extension.[22] Following their suggestion, we build a new series for our dependent variable in the monetary policy equation by splicing the actual official bank rate from 1972Q3 to 2009Q1 and the corresponding shadow rate from 2009Q2 to 2016Q2, calculated for the UK by Wu and Xia (2016).

4.2 Estimation of Policy Regimes

In this section we present first the results for the reduced-form equations of inflation and the output gap (first stage), then those for the policy rule equation (second stage). The strategy we follow for selecting the instruments is discussed in the Supplementary Material.[23]

4.2.1 Reduced-Form

The outcome of the break inference on the inflation and the output gap equations is displayed in Tables 2 and 3, while detailed results on the estimation of the breaks can be found in the Supplementary Material. As for the inflation equation stability is unequivocally rejected at 1% by both the F and the *Wald* versions of the tests against a fixed number of breaks and by the *UD* max test. Our favoured method, where the sequential testing uses its own break dates, suggests the presence of one break at 1980Q2, which is estimated with fairly low uncertainty considering that the 90% confidence interval covers only 6 and 10 observations using respectively the OLS and White standard errors.[24] The relatively high accuracy of the break date estimate is confirmed by plotting the F statistic for one break evaluated at each possible date (see the Supplementary Material). This graph exhibits, indeed, a

22 It is known that the shadow rate is different depending on the model used (e.g. Christensen and Rudebusch 2014), but Wu and Xia (2016) show that this concerns the level and not the dynamics of the shadow rate, which is what matters for our purposes.

23 As additional estimations we also tried a policy rule specification including a smoothing component, in the form of one and two lags of the official bank rate, and also the OLS approach suggested by Perron and Yamamoto (2013). In both cases policy coefficients turned out to be highly implausible in light of commonly accepted accounts of UK monetary policy history and previous empirical estimates.

24 Confidence intervals for the break dates are calculated using the asymptotic theory developed by Bai (1997).

Table 2: Break interval, inflation eq.

Break date: 30 (1980Q2)				
	Confidence level			
Std err	**90%**	**95%**	**97.5%**	**99%**
White	28–32	27–33	26–34	24–36
LS	28–32	28–32	28–32	27–33

Table 3: Break interval, output gap eq.

Break date: 31 (1980Q3)				
	Confidence level			
Std err	**90%**	**95%**	**97.5%**	**99%**
White	25–37	22–40	19–43	14–48
LS	28–34	26–36	25–37	23–39
Break date: 143 (2008Q2)				
	Confidence level			
Std err	**90%**	**95%**	**97.5%**	**99%**
White	141–143	141–143	141–143	140–144
LS	141–143	140–144	140–144	140–144

sharp peak at 1980Q2, which also matches the timing of the shift in the mean of the marginal distribution of inflation, estimated at 1980Q2 (see Section 4.1).

As for the output gap equation stability is unanimously rejected by both the F and *Wald* tests against a fixed number of breaks, and also taking its unweighted maximum. Our preferred criterion, the sequential testing starting with the $UD \max F$ at 1% significance level, suggests two breaks at 1980Q3 and 2008Q2. While the confidence intervals of the first break date are very wide, those associated with the second break date are extremely narrow, denoting a high accuracy in the identification of the timing of this episode of instability. The large uncertainty about the location of the break at 1980Q3 is confirmed by the graph of the F test statistic for one break over the whole sample (see the Supplementary Material), which is, indeed, beyond the 1% critical value for most of the observations until 1994Q2. Nevertheless, both estimated break dates correspond to local maxima of the F statistic.

We now have to select the specific set of reduced-form breaks that will be incorporated in the fitted values employed in the second stage of the HHB procedure. We decide to model a single common break at 1980Q2 since this break date was found for the inflation equation but is only one observation distant from the first break date in the output gap equation, it has a confidence interval that is entirely encompassed by the one calculated for the output gap, and the F test for one break in the output gap equation remains very large in a wide neighbourhood of its peak. Because we need to define a common partition for the break estimation in the second stage we impose the second break in the output gap equation, at 2008Q2, also on the inflation equation, without noticeable changes in the parameter estimates.

4.2.2 Structural Equation

As a consequence of the two structural breaks detected in the reduced-form system we have a partition of three stable segments that we have to examine individually in the second stage. The confidence intervals of the estimated break dates that are idiosyncratic to the structural equation are displayed in Table 4, while detailed results on the break estimation is presented in the Supplementary Material. Given that segments are small, we show the confidence intervals calculated using both the White and the Least Squares (LS) estimate of the covariance matrix.

All three segments turn out to be undoubtedly unstable. Using our favourite criterion that applies the UD max F test statistic in the first step with 1% significance level, we find that the first segment (1973Q2–1980Q2) features one break at 1978Q3. The evidence for this instability is quite strong, considering that the same conclusion is reached across all four criteria, including the Wald version of the testing procedure, and that the confidence interval for the break location is fairly narrow. As for the second segment (1980Q3–2008Q2), we find three breaks at 1988Q2, 1991Q2 and 2001Q2, while the Wald version of the tests suggests an additional fourth break at 1989Q3. The evidence for the instability in 1991Q2 is very strong if we take into account that this is the first break found in the sequential procedure, with a huge value of the test statistic, and with a particularly tight confidence interval (with 95% probability the break date is within the year 1991). In the third segment (2008Q3–2016Q2), we obtain one break at 2011Q3, which is with 99% confidence within only three quarters. To assess the relevance of the two reduced-form breaks in the policy rule equation we apply the Wald test for common breaks constructed by HHB. We obtain that the break at 1980Q3 is not significant since the statistic is 8.16 with a p-value of 0.04, while the break at 2008Q2 is highly significant with a statistic that is 118.71 and a virtually zero p-value.

Overall, we observe that, though the sample period is characterized by numerous structural changes, the break inference establishes their location in time with a high level of accuracy (see Table 4 and Figure 1). The relevance of the estimated break dates as symptoms of distinct episodes of parameter instability is especially evident by the

Table 4: Break interval, policy rule eq.

Break date: 23 (1978Q3)				
	Confidence level			
Std err	**90%**	**95%**	**97.5%**	**99%**
White	21–25	21–25	20–26	20–26
LS	21–25	21–25	20–26	19–27
Break date: 62 (1988Q2)				
	Confidence level			
Std err	**90%**	**95%**	**97.5%**	**99%**
White	60–64	60–64	60–64	59–65
LS	60–64	60–64	60–64	59–65
Break date: 74 (1991Q2)				
	Confidence level			
Std err	**90%**	**95%**	**97.5%**	**99%**
White	73–75	73–75	72–76	72–76
LS	73–75	72–76	72–76	72–76
Break date: 144 (2001Q2)				
	Confidence level			
Std err	**90%**	**95%**	**97.5%**	**99%**
White	112–116	112–116	111–117	110–118
LS	112–116	112–116	111–117	111–117
Break date: 155 (2011Q3)				
	Confidence level			
Std err	**90%**	**95%**	**97.5%**	**99%**
White	154–156	154–156	154–156	154–156
LS	154–156	154–156	153–157	153–157

fact that no confidence interval features overlapping parts, neither using the widest option of 99%. Since both the Breusch-Pagan and the Breusch-Godfrey test statistics reject with virtually zero p-values, we refine our estimates of the policy rule equation by performing a GMM estimation, conditional on our estimated set of structural breaks.[25] The GMM covariance matrix is obtained via the method of Newey-West to select the bandwidth combined with a Bartlett kernel. The results of this estimation are displayed in Table 5.

We notice that until 1978Q4 there is no statistically significant policy rule, but thereafter all three coefficients of Eq. (3.2), the intercept β_1, the inflation coefficient β_2 and the output gap coefficient β_3, are significant across all segments, with the only exception of β_3 in segment 2 and 5 and β_2 in segment 6. If we estimate the policy rule equation under the assumption of constant parameters, we obtain a significant value of 0.48 and −0.26 for respectively the inflation and the output gap coefficient. We highlight the size of the biasedness in such estimates, which is the consequence of ignoring the existent parameter instability, especially in the case of the output gap coefficient, which turns negative and significant, while in the model that allows for structural breaks a negative output gap coefficient occurs only in one segment and without being significant (see also Figure 1).[26]

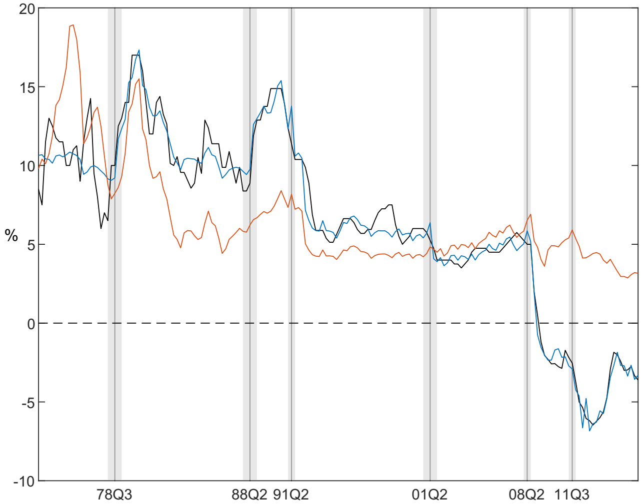

Figure 1: Actual interest rate (black line) and predicted levels from the break model (blue line) and from the constant-coefficient model (red line), with estimated break dates indicated by vertical lines, along with their 90% confidence intervals.

25 The Breusch-Pagan statistic is 41.81, with 12 degrees of freedom, while the Breusch-Godfrey F statistic is 16.89, with 4 and 145 degrees of freedom.

26 In the Supplementary Material we illustrate the importance of modelling the structural breaks that exist in the reduced-form representation of the endogenous regressors both for the break inference and the parameter estimates of the policy rule.

Table 5: GMM policy rule estimates.

Segment	*Period*	β_1	β_2	β_3
1	1973Q1–1978Q3	**8.14**	0.11	0.16
		(2.15)	(0.09)	(0.22)
2	1978Q4–1988Q2	**8.02**	**0.44**	−0.02
		(0.34)	(0.02)	(0.05)
3	1988Q3–1991Q2	**4.09**	**1.27**	**0.78**
		(1.27)	(0.17)	(0.10)
4	1991Q3–2001Q2	**4.19**	**0.96**	**0.18**
		(0.19)	(0.07)	(0.05)
5	2001Q3–2008Q2	**2.96**	**0.63**	0.08
		(0.30)	(0.23)	(0.11)
6	2008Q3–2011Q3	**1.71**	−0.11	**1.25**
		(0.48)	(0.14)	(0.06)
7	2011Q4–2016Q2	**4.99**	**0.26**	**2.96**
		(0.52)	(0.09)	(0.21)

In bold coefficients significant at 1%; standard errors in brackets.

As a quantitative assessment of the importance of the shifts in the policy coefficients for the overall variability of the interest rate, we use the second stage residuals from the 2SLS estimation to decompose the total sum of squares in the explained sum of squares from the no-break model, ESS_c, which can be thought of as the variability in the interest rate that is explained by its average constant relationship with inflation and the output gap, the additional explanatory power produced by the break model, ESS_b–ESS_c, which represents the contribution to the interest rate variability of the shifts in the parameters of this relationship, and the residual TSS–ESS_b. What is remarkable in this accounting exercise is that not only does the explicit modelling of the shifts in parameters provide additional explanatory power, but it even outweighs substantially the overall fit provided by a stable policy rule equation. While the share of *TSS* explained by a constant policy rule is only 36%, modelling discrete changes in the parameters allows to explain an additional 61% of all the movements in the official interest rate over the whole period.[27]

The overall validity of the model is confirmed by Hansen's test for overidentifying restrictions, which, as expected, does not signal a substantial difference between the models with and without breaks (see the Supplementary Material). On the contrary, a considerable difference between the two models emerges in terms of residual autocorrelation. While the break model retains some of the significant

27 We have to acknowledge that this gain in fit is achieved at the cost of introducing 18 new regressors. In fact, however, these additional regressors are not truly new variables but simply the product of existing regressors and dummies that reflect specific segments.

autocorrelation, which prevents us from discarding the other two potential sources of interest rate persistence, there is a massive reduction in its magnitude compared to the no-break model if we look at the values of the Breusch-Godfrey F statistics. This is strongly confirmed by the fact that if we re-estimate the two models including as an additional regressor one lag of the dependent variable, the estimate of its coefficient falls by almost half once we control for the existent structural breaks, that is from 0.99 to 0.55, both with a p-value of 0.009. This result highlights how neglecting existent parameter instability in policy rule equations may translate into spurious evidence of policy inertia, or at least lead to substantially overstate its importance in the actual setting of the policy instrument. On the other hand, the fact that the testing procedure based on the autocorrelation-robust Wald statistic produces the same list of break dates, with the exception of one extra break in segment 2, suggests no substantial risk of spurious breaks induced by the autocorrelated errors. When we look for significant signs of piecewise heteroskedasticity, which might have confounded our break inference on the policy rule equation, we find that the Goldfeld-Quandt statistic features very large p-values in correspondence of four of the six estimated breaks, with a change in variance occurring only in the second and fifth segments, which suggests a very low risk that our estimated breaks are in fact spurious. Finally, we perform two robustness checks by running the break estimation after replacing the chosen output gap measure with an alternative one constructed by the OECD and using the RPIX as inflation index. The timing of the instability episodes remains largely unchanged, with only minor differences in the values of the policy coefficients, which confirms our findings about the evolution of the UK monetary policy (see Supplementary Material).

4.3 Identification of Changing Priorities

The outcome of our empirical approach highlights the existence of seven distinct policy regimes in the history of the UK monetary policy. However, we still do not know which specific parameters did significantly change in the transition from one regime to the other. For this reason, we apply our modified version of the Inoue-Rossi method, as described in Section 3.3. Results are displayed in Table 6, organized in six panels, one for each break date.

It is noticeable how the inflation coefficient, β_2, has significantly changed in all but two of the estimated episodes of instability, 2001Q2 and 2011Q3, which are also the only two instances when the intercept breaks. In three episodes there is a significant shift in the policy makers stance with respect to both the inflation rate and the output gap. Hence, we can conclude that in the almost 45 years that followed the collapse of the Bretton Woods system the UK experienced seven distinct monetary

Table 6: Identification of the unstable subset of parameters.

1978Q3						1988Q2					
Individual stat			Subset stat			Individual stat			Subset stat		
Param	**F**	**p-value**	**Subset**	**F**	**p-value**	**Param**	**F**	**p-value**	**Subset**	**F**	**p-value**
β_1	0.089	0.768	β_1, β_3	0.045	0.956	β_1	1.395	0.245	β_1, β_2	6.694	0.003
β_2	0.792	0.383				β_2	3.563	0.067	β_1	1.395	0.245
β_3	0.069	0.795				β_3	10.652	0.002			
BIC	β_1, β_3		***Unstable subset***			BIC	β_1		**Unstable subset**		
HQ	β_1, β_3		β_2			HQ	β_1		β_2, β_3		

1991Q2						2001Q2					
Individual stat			Subset stat			Individual stat			Subset stat		
Param	**F**	**p-value**	**Subset**	**F**	**p-value**	**Param**	**F**	**p-value**	**Subset**	**F**	**p-value**
β_1	0.008	0.932	β_1, β_2	3.348	0.044	β_1	4.682	0.034	β_2, β_3	0.561	0.574
β_2	0.592	0.446	β_1	0.007	0.931	β_2	0.506	0.480			
β_3	7.274	0.010				β_3	0.035	0.853			
BIC	β_1		**Unstable subset**			BIC	β_2, β_3		**Unstable subset**		
HQ	β_1		β_2, β_3			HQ	β_2, β_3		β_1		

2008Q2						2011Q3					
Individual stat			Subset stat			Individual stat			Subset stat		
Param	**F**	**p-value**	**Subset**	**F**	**p-value**	**Param**	**F**	**p-value**	**Subset**	**F**	**p-value**
β_1	1.569	0.218	β_1, β_2	5.514	0.008	β_1	4.980	0.034	β_1, β_2	3.965	0.031
β_2	5.033	0.031	β_1	1.569	0.218	β_2	1.449	0.239	β_2	1.449	0.239
β_3	47.907	0.000				β_3	16.885	0.000			
BIC	β_2, β_3		**Unstable subset**			BIC	β_1, β_3		**Unstable subset**		
HQ	$\beta_1, \beta_2, \beta_3$		β_2, β_3			HQ	β_1, β_3		β_1, β_3		

Under "Individual stat" we list the F statistics that test the stability of an individual parameter conditional on all the other being unstable; under "Subset stat" we show the F statistics that test the stability of a subset of parameters, and which are used in the Inoue-Rossi sequential procedure to identify the stable subset; the stable subset estimated by information criteria is indicated by BIC and HQ; the final subset of unstable parameters that is estimated by the Inoue-Rossi procedure is presented under "Unstable subset."

policy regimes, which were effectively different in terms of the attitude of the policy makers towards the two main policy goals.[28]

5 Discussion

A look at our estimation results in Table 5 reveals two key findings. First, many of the historical dates that we expected to be important, and that we included in Table 1, surprisingly do not come up as significant break points, but the overall evolution confirms existing accounts of the UK monetary policy. What emerges is the picture of a monetary policy that undergoes a sequence of phases over time, not strictly matching the timing of the official institutional events, and thus suggesting an important role played by anticipations, early transitions and delays in the actual enforcement of the institutional regime changes. Second, a policy instrument rule did not exist until the end of the 1970s, and since then a significant positive response to inflation has characterized monetary policy until 2016, with the only exception of the interval between 2008 and 2011. As can be seen from Figure 1, the fit of our estimated policy rule is indeed disastrous up to 1978Q3, but then it improves substantially during the 1980s, and it is able to capture very well the movements of the interest rate from 1988Q2 up to 2016Q2.

This rather sharp evidence confirms previous understandings of the UK history, and in particular the claim that exactly in that period a paradigm change about the transmission mechanism took place within academic and policy circles, which led to the abandonment of incomes policy and price regulation in favour of a monetary approach to the control of the inflation process (e.g. Batini and Nelson 2005).

This shift in the general monetary policy strategy is testified by a series of events such as the abolishment of the exchange rate controls in October 1979, the start of the Medium Term Financial Strategy (an explicit multi-year monetary target) in March 1980 and the end of the “corset” scheme to limit the banks’ lending in June 1980. The scope of the transformations taking place during this period is corroborated by the large downward shift in the mean of the inflation process we found in 1980Q2, along with a break in the reduced-form equation of both inflation and the output gap, which is accompanied in the subsequent quarter by a

28 Obviously, we are not able to understand whether this changed attitude reflects shifts in policy makers preferences or the optimal response to structural breaks in the parameters of the economy. Answering this question would require integrating the current analysis with an estimation of the structural model of the economy.

significant fall in the level of the interest rate spread. These pieces of evidence all confirm an increased credibility of the government policy in its effort to curb inflation, as well as a stronger ability of the monetary authorities to influence the market short-term interest rates.

As for the magnitude of the policy response to inflation (β_2) and output gap (β_3), it can be argued that the positive and significant inflation coefficient in the second segment reflects the restrictive macroeconomic strategy adopted to curb inflation, but otherwise the relatively low fit of the policy rule is the consequence of a series of events that affected the interest rate decisions beyond what could be expected from the mere consideration of the dynamics of prices and economic activity.[29]

An estimated inflation coefficient that is greater than one between 1988Q3 and 1991Q2 captures the implementation of a very tight monetary policy in the face of an increasing inflation, whereas a response smaller than one in the subsequent period, 1991Q3–2001Q2, was arguably the consequence of an inflation rate that was falling throughout the whole segment, combined with an asymmetric response of the interest rate that was deliberately stronger with respect to increases than decreases in inflation. Since this fall in inflation started in 1991Q3, being the product of the past restrictive monetary policy but also of a more credible nominal anchor in place, the inflation targeting period from 1992Q3 appears *de facto* as a continuation of the same regime. Hence, both the adoption of the inflation targeting in 1992 and the established central bank independence in 1997 appear as part of a regime change that started earlier in time. The relatively more stable pattern for both inflation and the output gap is plausibly the reason why for the period 2001Q3–2008Q2 we obtain a smaller but still significant coefficient on inflation and a not significant output gap coefficient. The reverse in the statistical significance of the two coefficients starting from 2008Q3, with a strong and significant response only with respect to the output gap up to 2011Q2, reflects the priorities of the quantitative easing program, which starts officially in March 2009. Finally, the break we found in 2011Q3, which is accompanied by a massive increase in the response to output gap, from 1.25 to 2.96, as well as by the inflation coefficient becoming significant, matches the timing with which the second round of quantitative easing is announced as a means to ward off the risk of a double-dip recession. Indeed, this event translates into a sharp fall in the shadow rate from –2.53% to –3.70%. The fact that this announcement starts off a new regime rather than representing a single episode is confirmed by the numerous subsequent

29 Just to mention a few of these events: the repeated failures to hit the monetary targets and their abandonment in October 1985; the financial sector deregulation in 1986; the world stock market crash in 1987; the exchange rate policy before and after entering the Exchange Rate Mechanism (ERM) in October 1990.

interventions in the same direction.[30] It is also worth noting that since we obtained a reduced-form break at 2008Q2, the use of a shadow rate since 2009Q1 does not affect the break inference in the first part of the sample, where we use the actual official rate.

The Taylor principle ensuring a unique stationary equilibrium for the economy, as stated in Bullard and Mitra (2002), turns out to be satisfied only over the period 1988Q3–2001Q2. More precisely, it is clearly met in segment 3, given that the inflation coefficient is above 1, but also in segment 4 using a plausible range of values for the discount factor and the slope of the New Keynesian Phillips curve. This latter needs to be smaller than 0.045 if we use a discount factor of 0.99, which is not implausible for the UK (see e.g. estimates in Batini, Jackson, and Nickell 2005).

When we compare our results with those from previous studies, we notice that Clarida, Galì, and Gertler (1998) found that during the 1979–1990 period the response to inflation was 0.98, which is consistent with an average of the two values we obtain in segment 2 and 3. Comparing the results from a policy rule estimated in three *a priori* selected periods, Adam, Cobham, and Girardin (2005) obtained that from 1985 to 1990 the interest rate was exclusively determined by the US and German interest rates, and it is only with the introduction of the inflation targeting in 1992 that inflation and the output gap became significant factors, but they dominated the interest rate setting in the subsequent 1997–2002 period. While they interpret this outcome as a confirmation of the importance of the institutional change that attributed operational independence to the Bank of England in 1997, we find, on the contrary, that neither 1992 or 1997 are picked by the tests as significant break dates, suggesting an evolution of the effective systematic policy decisions that does not match exactly the historical timing of the institutional changes. In this respect our results are closer to Kishor (2012), who obtained from a TVP policy rule that the inflation coefficient reaches the highest level around 1990 and then slowly declines towards 1998, although his estimated response is greater than one already from 1983.

6 Conclusion

We studied the evolution of the UK monetary policy by developing a coherent empirical approach to estimate and accurately date its historical regime changes,

30 Increases in asset purchases are announced by the Bank of England in October 2011 (from 200bn to 275bn), February 2012 (additional 50bn), and July 2012 (additional 50bn). Moreover, several statements about the intention to keep the official bank rate low for a prolonged period are issued in 2013 and 2014; restrictions on bank capital are relaxed in July 2016; the official bank rate is cut to 0.25% in August 2016.

building on a set of tools that have recently been advanced in the econometric literature on structural stability. The advantages of our approach compared to existing methods to capture regime changes are mainly two. It is specifically designed to provide a rigorous estimation of the timing with which policy regimes have been established, and it is able to identify which priorities of the policy makers have changed in the transition from one regime to another.

Our approach delivers a plausible description of the evolution of the UK monetary policy over time. We found that, in the almost 45 years following the collapse of the Bretton Woods system, regime changes have been frequent, they are responsible for a substantial share of the observed variation in the policy instrument, but their timing relative to the causing institutional change is far from obvious. As a result of an accumulation of institutional changes, the systematic policy actions of the monetary authorities underwent a sequence of phases, characterized by a different priority assigned to the main policy goals, but following a coherent evolution over time, with inflation becoming gradually more important from the 1970s up to early 2000s, and the output gap dominating after the 2008 financial crisis.

We showed how ignoring these frequent regime changes when estimating a policy rule produces highly biased parameter estimates and an overstatement of the importance of policy inertia. Overall, our results provide a clear warning to empirical studies that, for instance, build policy counterfactuals in an attempt to disentangle the effects of monetary policy on the economy, either taking for granted the timing with which institutional policy reforms exert their effects or imposing a generic continuous adjustment over time in the policy coefficients.

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References

Adam, C., D. Cobham, and E. Girardin. 2005. “Monetary Frameworks and Institutional Constraints: UK Monetary Policy Reaction Functions, 1985–2003.” *Oxford Bulletin of Economics and Statistics* 67 (4): 497–516.

Andrews, D. W. K. 1993. “Tests for Parameter Instability and Structural Change with Unknown Change Point.” *Econometrica* 61: 821–56.

Assenmacher-Wesche, K. 2006. “Estimating Central Banks’ Preferences from a Time-Varying Empirical Reaction Function.” *European Economic Review* 50 (8): 1951–74.

Bai, J. 1997. “Estimating Multiple Breaks One at a Time.” *Econometric Theory* 13 (03): 315–52.

Bai, J., and P. Perron. 1998. “Estimating and Testing Linear Models with Multiple Structural Changes.” *Econometrica* 66: 47–78.

Batini, N., and E. Nelson. 2005. “The U.K.’s Rocky Road to Stability.” *Working Papers*, 2005-020. Federal Reserve Bank of St. Louis.

Batini, N., B. Jackson, and S. Nickell. 2005. “An Open-Economy New Keynesian Phillips Curve for the UK.” *Journal of Monetary Economics* 52 (6): 1061–71.

Boivin, J. 2006. “Has U.S. Monetary Policy Changed? Evidence from Drifting Coefficients and Real-Time Data.” *Journal of Money, Credit and Banking* 38 (5): 1149–73.

Boldea, O., A. R. Hall, and S. Han. 2012. “Asymptotic Distribution Theory for Break Point Estimators in Models Estimated via 2SLS.” *Econometric Reviews* 31 (1): 1–33.

Bullard, J., and K. Mitra. 2002. “Learning about Monetary Policy Rules.” *Journal of Monetary Economics* 49 (6): 1105–29.

Canova, F., and L. Gambetti. 2009. “Structural Changes in the US Economy: Is There a Role for Monetary Policy?” *Journal of Economic Dynamics and Control* 33 (2): 477–90.

Castelnuovo, E. 2003. “Taylor Rules, Omitted Variables, and Interest Rate Smoothing in the US.” *Economics Letters* 81 (1): 55–9.

Castelnuovo, E. 2006. “The Fed’s Preference for Policy Rate Smoothing: Overestimation Due to Misspecification?” *The B.E. Journal of Macroeconomics (Topics)* 6 (2). https://doi.org/10.2202/1534-5998.1416.

Castelnuovo, E., L. Greco, and D. Raggi. 2014. “Policy Rules, Regime Switches, And Trend Inflation: An Empirical Investigation For The United States.” *Macroeconomic Dynamics* 18 (4): 920–942.

Castro, V. 2011. “Can Central Banks’ Monetary Policy be Described by a Linear (Augmented) Taylor Rule or by a Nonlinear Rule?” *Journal of Financial Stability* 7 (4): 228–46.

Christensen, J. H., and G. D. Rudebusch. 2014. “Estimating Shadow-Rate Term Structure Models with Near-Zero Yields.” *Journal of Financial Econometrics* 13 (2): 226–59.

Clarida, R., J. Galì, and M. Gertler. 1998. “Monetary Rules in Practice: Some International Evidence.” *European Economic Review* 42: 1033–67.

Clarida, R., J. Galí, and M. Gertler. 2000. “Monetary Policy Rules and Macroeconomic Stability: Evidence and Some Theory.” *The Quarterly Journal of Economics* 115 (1): 147–80.

Cobham, D. 2006. “Using Taylor Rules to Assess the Relative Activism of the European Central Bank, the Bank of England and the Federal Reserve Board.” In CDMA Conference Paper Series, 602.

Coibion, O., and Y. Gorodnichenko. 2012. “Why are Target Interest Rate Changes So Persistent?” *American Economic Journal: Macroeconomics* 4 (4): 126–62.

Dolado, J. J., R. María-Dolores, and M. Naveira. 2005. “Are Monetary-Policy Reaction Functions Asymmetric?: The Role of Nonlinearity in the Phillips Curve.” *European Economic Review* 49 (2): 485–503.

English, W. B., W. R. Nelson, and B. P. Sack. 2003. “Interpreting the Significance of the Lagged Interest Rate in Estimated Monetary Policy Rules.” *Contributions in Macroeconomics* 3 (1). https://doi.org/10.2202/1534-6005.1073.

Hall, A. R., S. Han, and O. Boldea. 2012. “Inference Regarding Multiple Structural Changes in Linear Models with Endogenous Regressors.” *Journal of Econometrics* 170 (2): 281–302.

Hansen, B. E. 2000. “Testing for Structural Change in Conditional Models.” *Journal of Econometrics* 97 (1): 93–115.

Harvey, D. I., S. J. Leybourne, and P. Newbold. 2004. “Tests for a Break in Level When the Order of Integration is Unknown.” *Oxford Bulletin of Economics and Statistics* 66 (1): 133–46.
Inoue, A., and B. Rossi. 2011. “Identifying the Sources of Instabilities in Macroeconomic Fluctuations.” *Review of Economics and Statistics* 93 (4): 1186–204.
Kharel, R. S., C. Martin, and C. Milas. 2010. “The Complex Response of Monetary Policy to the Exchange Rate.” *Scottish Journal of Political Economy* 57 (1): 103–17.
Kim, C. J., and C. R. Nelson. 2006. “Estimation of a Forward-Looking Monetary Policy Rule: A Time-Varying Parameter Model Using Ex Post Data.” *Journal of Monetary Economics* 53 (8): 1949–66.
Kishor, N. K. 2012. “A Note on Time Variation in a Forward-Looking Monetary Policy Rule: Evidence from European Countries.” *Macroeconomic Dynamics* 16 (3): 422–37.
Koop, G., R. Leon-Gonzalez, and R. W. Strachan. 2009. “On the Evolution of the Monetary Policy Transmission Mechanism.” *Journal of Economic Dynamics and Control* 33 (4): 997–1017.
Koustas, Z., and J. F. Lamarche. 2012. “Instrumental Variable Estimation of a Nonlinear Taylor Rule.” *Empirical Economics* 42 (1): 1–20.
Krippner, L. 2012. “Modifying Gaussian Term Structure Models when Interest Rates are Near the Zero Lower Bound.” *Reserve Bank of New Zealand Discussion Paper*, 2012/02.
Lubik, T., and P. Surico. 2010. “The Lucas Critique and the Stability of Empirical Models.” *Journal of Applied Econometrics* 25 (1): 177–94.
Perron, P., and T. J. Vogelsang. 1992. “Nonstationarity and Level Shifts with an Application to Purchasing Power Parity.” *Journal of Business and Economic Statistics* 10 (3): 301–20.
Perron, P., and Y. Yamamoto. 2013. “Using OLS to Estimate and Test for Structural Changes in Models with Endogenous Regressors.” *Journal of Applied Econometrics* 30 (1): 119–44.
Perron, P., and J. Zhou. 2008. “Testing Jointly for Structural Changes in the Error Variance and Coefficients of a Linear Regression Model.” *Unpublished Manuscript*. Department of Economics, Boston University.
Pitarakis, J. Y. 2004. “Least Squares Estimation and Tests of Breaks in Mean and Variance Under Misspecification.” *Econometrics Journal* 7 (1): 32–54.
Rudebusch, G. D. 2002. “Term Structure Evidence on Interest Rate Smoothing and Monetary Policy Inertia.” *Journal of Monetary Economics* 49 (6): 1161–87.
Rudebusch, G. D. 2006. “Monetary Policy Inertia: Fact or Fiction?” *International Journal of Central Banking* 2 (4): 85–135.
Sims, C. A., and T. Zha. 2006. “Were There Regime Switches in US Monetary Policy?” *The American Economic Review* 96 (1): 54–81.
Surico, P. 2007. “The Fed’s Monetary Policy Rule and US Inflation: The Case of Asymmetric Preferences.” *Journal of Economic Dynamics and Control* 31 (1): 305–24.
Taylor, M., and E. Davradakis. 2006. “Interest Rate Setting and Inflation Targeting: Evidence of a Nonlinear Taylor Rule for the United Kingdom.” *Studies in Nonlinear Dynamics and Econometrics* 10 (4): 1–20.
Trecroci, C., and M. Vassalli. 2010. “Monetary Policy Regime Shifts: New Evidence from Time-Varying Interest Rate Rules.” *Economic Inquiry* 48 (4): 933–50.
Wu, J. C., and F. D. Xia. 2016. “Measuring the Macroeconomic Impact of Monetary Policy at the Zero Lower Bound.” *Journal of Money, Credit and Banking* 48 (2–3): 253–91.

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