Changes in the relationship between short-term interest rate, inflation and growth: Evidence from the UK, 1820-2014

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Abstract

This paper examines the dynamic relationship between interest rates, inflation and economic growth using the longest available dataset for the UK and a vector autoregression (VAR). The approach adopted enables structural breaks to be identified in the dynamic system. It then can ascribe breaks in covariance to changes in volatility or to changes in correlation. Our empirical findings indicate several structural breaks in the relationship, which lead to very different inference compared to a constant parameter model. For example, interest rates respond much more strongly to growth or inflation over recent decades. Furthermore, our evidence suggests that all variables become more persistent after the classical gold standard ended with the onset of WW1.

JEL classification: JEL C12, C32, E20, G12, G15.

Keywords: Short-term Interest Rate, Inflation, Growth, VAR, Structural Breaks

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1 Introduction

The short-term interest rate is a key variable in Financial Economics. It is closely related to the economic policy rate set by central banks (or governments). Arguably, the primary tool used by policymakers to manage the macroeconomy is the interest rate. Theoretically there is widespread support that interest rates respond to both inflation and economic activity (interest rates lag these variables) but also that interest rates themselves impact inflation and economic activity (interest rates lead these variables). This paper seeks to model the dynamic behaviour of the short-term interest rate, inflation and growth by implementing recently developed techniques. In particular our paper seeks to identify if there has been any structural breaks in the linkages between these variables in the context of the UK over the period 1820-2014. Second, it sheds light on the timing of these structural breaks; do they, for example, occur when there is a change in monetary policy regime? Third, we examine if each structural break is the result of a shift in volatility or a change in correlation using the recently developed methodology of Bataa, Osborn, Sensier and van Dijk (2013). An important contribution of this paper is the finding that structural breaks in the contemporaneous linkages between the short term interest rate, inflation and growth were almost all volatility breaks up until 1951 but there was a correlation break in 1990. Finally, we analyse if the series exhibit changes in persistence.

The primary analysis is conducted using a vector autoregression (VAR) model and tests for structural breaks in the parameters of the model. Our empirical analysis employs the generalized VAR methodology of Koop, Pesaran and Potter (1996) and Pesaran and Shin (1998), and the forecast error variance decomposition analysis of which is updated in Diebold and Yulmaz (2009, 2012, 2014). More importantly, we implement generalized versions of the impulse responses and the forecast error variance decompositions that are not sensitive to variable ordering. We also allow for structural breaks using the methodology of Bataa, Osborn, Sensier and van Dijk (2013,2014). This methodology allows for changes in both the VAR parameters as well as the contemporaneous covariance matrix, but can allow for breaks to each at different time periods. Further, the covariance matrix breaks can be identified as either a variance break or a correlation break. Our UK data sample is the lengthiest available, covering about two centuries, which is an important consideration given the importance of timespan for this type of time-series analysis (Shiller and Perron, 1985).

To preview some of our results, we find that: 1) most breaks in the last two centuries can be primarily attributed to changes in volatility; 2) after the collapse of the classical gold standard the persistence of all three variables increases[1]; 3) there is very clear evidence that the dynamic relationship between inflation and the short rate changes after the gold standard ended; 4) two changes occurred in recent decades- first, since 1975 the short term interest rate starts to respond very strongly to the past growth rates and second, after 1990 the magnitude of the short term interest rate response to past inflation greatly increases; 5) our results confirm earlier research that a great moderation occurred but suggests that the structural change is apparent by 1952 in UK annual data. We do find a covariance matrix break in 1989 but this can be attributed to a break in correlation rather than a break in volatility. One other important finding of this study is that short term interest rates primarily lag growth and inflation but do not lead.

The outline of the paper is as follows. Section 2 provides a short review of previous literature. A description of the data is presented and unit root analysis is carried out in section 3. In this section we also explain the method that we use to identify changes in the univariate behaviour of the short-term interest rate, inflation and economic growth. Then in section 4 we discuss our multivariate econometric techniques that deals with structural breaks in the relationship between these variables. The empirical results are presented in section 5 and we conclude in section 6.

2 Review of the Empirical Literature

2.1 Inflation leading nominal interest rates

There is a large literature that surveys the Fisher effect-the contemporaneous relationship between inflation and nominal interest rates (see for example Mishkin, 1992). We are concerned with the leading or lagging relationship between inflation and nominal interest rates in this paper and so do not survey this literature in detail here. Nonetheless, there is much literature that finds the response of interest rates to inflation to be less than unity in the short-run; this implies a slow adjustment of interest rates to inflation (Mishkin, 1992; Evans and Lewis, 1995) although Crowder and Hoffman (1996) provide some counter-evidence that the coefficient is approximately unity (after-tax) in the long-run.

[1]The classical gold standard broke down during World War I, as major belligerents resorted to inflationary finance, and was briefly reinstated from 1925 to 1931 as the Gold Exchange Standard. Between 1946 and 1971, countries operated under the Bretton Woods System. Under this further modification of the gold standard, most countries settled their international balances in U.S. dollars, but the U.S. government promised to redeem other central banks holdings of dollars for gold at a fixed rate of thirty-five dollars per ounce.

Rapach (2003) investigates if inflation has a long-run impact (if there is long-run money superneutrality). Alternative theories suggest that inflation could have a permanent effect on the real interest rate and on real output. Rapach (2003) estimates the long-run responses of the real interest rate and real output for 14 industrialised countries using post-war annual data. In all countries a positive inflation shock leads to a permanent fall in the real interest rate and in some countries a positive inflation shock leads to an increase in real output. Hence the evidence is not very supportive of long-run money superneutrality.

Rapach and Wohar (2005) find extensive evidence of structural breaks in the real interest rate and in inflation for an international sample of monthly post-war data. Generally the breaks coincide with increases (decreases) in inflation being associated with a decrease (increase) in the real interest rate.

2.2 The Price Puzzle: Nominal interest rates leading inflation

Historical data shows that there is a positive relationship between inflation and the federal funds rate, which has been dubbed the "price puzzle" (Bernanke and Blinder 1992; Christiano, Eichenbaum, and Evans 1994, and Sims 1992). The reason it is a puzzle is because an unexpected contraction of money supply (i.e. an unexpected increase in the federal funds rate) is expected to be followed by a decrease in the price level, rather than an increase.

Sims (1992), using a recursive identification scheme, finds a positive effect on the price level from a negative monetary shock (e.g. from an increase in interest rates). This is contrary to traditional macroeconomic models where a fall in interest rates, increases money holdings and increases the price level. Sims (1992) argues that this anomaly, termed the price puzzle, is due to an omitted variable problem. There have been a number of different explanations for this "Price Puzzle". These include a misspecification of VARs (Giordani, 2004; Bernanke, Boivin, and Eliasz 2005) and theoretical models which try and justify the observed rise in prices (Barth and Ramey, 2002; Rabanal, 2007). Rusnak, Havranek, and Horvath (2013) conduct a meta analysis through the collection of all published research (103 studies for 31 countries) examining monetary transmission within a VAR framework. Their analysis also suggests that the price puzzle is due to model mis-specification. Nevertheless, the evidence so far seems to rest primarily on recent history and on monthly or quarterly data. In this paper we take a much longer perspective and examine if this so-called anomaly is apparent in annual data.

Balke and Emery (1994) confirm a price puzzle in the US, a positive correlation between federal funds rate increases and subsequent increases in prices. They find that the strength of this correlation is not uniform over the postwar period. In particular, the price puzzle broke down in the 1980s with a correlation between fed funds and inflation that is close to zero; but the correlation is still not negative, as traditional theory would predict.

Balke and Emery (1994) argue that a plausible explanation why the price puzzle was strong during the 1960s and 1970s was that the Federal Reserve responded to supply shocks by raising the federal funds rate but not enough to eliminate the inflationary consequences of the shock. This will result in a positive correlation between the federal funds rate and inflation. They also suggest the price puzzle is much weaker during the 1980s either because the Federal Reserve responded more vigorously to inflationary shocks, or there simply have been fewer large inflationary shocks to the economy. Balke and Emery's (1994) VAR evidence also suggests that supply shocks play an important role and that accounting for these (via modelling the term spread) can reduce the price puzzle even for the pre-1980 period. Thus, the implication is that the Federal Reserve responds to negative supply shocks by increasing the funds rate, but not by enough to fully offset the inflationary implications of the shock.

Lee (1992) conducts a VAR analysis of the relationship between interest rates, inflation, real activity and stock returns from January 1947 through December 1987. He finds interest rates explain a substantial fraction of the variation in inflation.

2.3 Nominal interest rates leading growth

A large body of research (e.g. Bernanke and Blinder, 1992) has demonstrated a strong correlation between measures of short-term nominal interest rates and real output. Such a correlation is puzzling, as most textbook theoretical models link long-term real interest rates to output. These models also argue that there is at best a weak link between short-term interest rates and long-term interest rates.

However, Fuhrer and Moore (1995) developed a rational expectations structural model of the US macro economy. They found that their model matches the dynamic properties of some important macroeconomic variables. The model also allows for a structural interpretation of the well documented, but puzzling, negative correlation between real output and the short-term interest rate. They argue that the systematic way in which monetary policy has leaned against the wind, in combination with a sluggish inflation adjustment and a structural IS curve that relates output to the rationally expected long-term real interest rate, has made the sample path of long-term real interest rates look like the short-term nominal interest rate. Fama (1981) argues there is a negative relationship between stock returns and inflation because inflation proxies for expectations of future real output.

Estrella and Hardouvelis (1991) present evidence that the real short-term interest rate and especially the term spread, provide information about future economic growth. The real short-term interest rate is negatively related to future growth, while the term spread is positively related to future growth. Plosser and Rouwenhorst (1994) however, suggest it is information in the long-end of the term structure that has predictive power for future industrial production.

Haug and Dewald (2011) examine the relationship between money growth and inflation and real output growth for 11 industrialised countries from 1880-2001. They find money growth leads or is contemporaneously related to inflation, however, money growth does not affect real output growth in the long term. They also find that there is no structural change in the longer term relationship between money growth and inflation even though changes in policy regime did occur. Their analysis applies band-pass filters to extract business cycle and longer-term cycles from each of the variables of interest.

2.4 Inflation and growth

There is an extensive amount of empirical studies documenting a negative growth effect of inflation across countries (Gylfason & Herbertsson, 2001; Barro, 2013; among others). Based on a large panel of 170 countries, Gylfason and Herbertsson (2001) finds a stronger negative growth effect of inflation using the Penn World data relative to that found when the World Bank data is used after controlling for other traditional growth determinants. Stronger negative effects on growth at lower levels of inflation are documented by Gillman, Harris, & Matyas (2004) while Guerrero (2006), using countries' past hyperinflation experience as the instrument, confirms previous OLS-based studies', that find growth cost of inflation to be robust. Though robust, recent studies report that it is only at a high inflation rate where results are significant, while at the low ranges of inflation rates results are ambiguous (Sarel, 1996; Ghosh & Phillips, 1998; Khan & Senhadji, 2001; Bick, 2010; Omay & Kan, 2010; Lopez-Villavicencio & Mignon, 2011). These findings depend on the precise inflation threshold which is a function of the countries being examined (i.e. developing, developed or mixed of both countries). For example, Khan and Senhadji (2001) and Lopez-Villavicencio & Mignon (2011) find that the inflation threshold in emerging and developing countries appears to be much higher than those found in developed countries (see Bick, 2010 and Omay and Kan, 2010, among others). Therefore, pooling all of these countries into one panel data set may bias the empirical finding. Lee (1992) finds inflation explains little of the variation in real activity.

Canova, Gambetti, & Pappa (2007) investigate the nature and causes of structural changes in the dynamics of output growth and inflation in the US, UK and the Euro area. They employ a time-varying coefficient VAR model (including inflation, output growth, short-term interest rate and money growth). The structural model is estimated using Bayesian methods including Markov Chain Monte Carlo methods to estimate the posterior distribution of the parameters of interest. They identify three types of structural shocks (technology, real demand, and monetary disturbance) using robust sign restrictions. They find that there are important similarities in the structural dynamics of inflation across countries over the last 35 years. However, the cross country similarities of output growth dynamics are only seen since the 1990s. Movements in volatility and persistent are accounted for by the combination of all three structural shocks. Sources of inflation persistence and volatility are similar but sources of output growth dynamics vary across countries.

2.5 Structural breaks and the Great Moderation

Alogoskoufis and Smith (1991) show that inflation persistence is significantly higher under managed-exchange-rate regimes than under gold-based, fixed-exchange-rate regime[2]. They use a sample split and RPI data for the UK between 1857 and 1987 and CPI data for the US between 1892 and 1987. According to them the autoregressive parameter (persistence) of an AR(1) model for inflation should increase in 1914 and around the end of the Bretton Woods system in 1968 due to the relaxation of the discipline imposed by these systems. They provide a model demonstrating that greater monetary accommodation of inflation and exchange-rate accommodation of inflation differentials increase inflation persistence. Burdekin and Siklos (1999) revisit their results with a formal structural break test of unknown timing and updated data and confirm a shift in inflation persistence around 1914 but contradict Alogoskoufis and Smith's (1991) premise that there was a further shift in 1968. Burdekin and Siklos (1999) provide references showing that the effective constraints on domestic monetary policy under the Bretton Woods System were much looser than those imposed under the gold standard and the

[2] Alogoskoufis and Smith (1991) and Alogoskoufis (1992) also examine relative inflation persistence, as given by inflation differentials between countries. We limit our analysis to the more basic issue of the inflation persistence present in the UK only.

collapse of that system cannot be as important. Bai and Perron (2003) also revisit Alogoskoufis and Smith's (1991) post World War II data and find the persistence increasing from 0.274 to 0.739, but no change in the intercept, in 1967. The estimate of the break is, however, imprecisely estimated with a 95% confidence interval covering the period 1961-1973. When they used the sequential testing strategy after the overall test they find two breaks, one in 1967 and the other in 1975.

Fuhrer and Moore (1995) documented high inflation persistence in post-WWII U.S. Benati (2008a) studies inflation persistence in an international sample and across differing monetary regimes. Benati (2008a) questions whether high inflation persistence is a pervasive feature in differing settings since in some cases inflation is not highly persistent. Benati (2008a) also notes that shifts in trend inflation and in price stickiness could impact the time series dynamics of inflation.

Work highlighting the Great Moderation and its cause includes Stock and Watson (2002; 2007), Kim and Nelson (1999) for the US and Benati (2008b) and Benati and Surico (2008) for the UK. Benati (2008b) results indicate that the great moderation in the UK was more due to good luck (the absence of major shocks) rather than to good policy; his counterfactual analysis suggests that different policy regimes would not have dramatically impacted the path of inflation or growth during his sample period.

However, none of these UK studies have examined the impact of structural breaks in this context, nor do they estimate impulse responses to such shocks. Several subsequent international studies suggest inflation only affects the short-run dynamics of interest rates, although shocks to interest rates can also have an impact on subsequent inflation.

3 Data and Preliminary Analysis

Our empirical analysis is based upon an extremely long time-series of annual UK data covering the period 1820-2014 from Officer (2015). The main variables of interest are the nominal short-term interest rate, retail price index inflation and real economic growth.[3] There are two main reasons for selecting this data. First, an important consideration for low frequency analysis of financial data is the length of the series. That is, the number of years covered in the sample is

[3]Officer (2015) provides two types of short term interest rates since 1790: one is contemporary and the other is consistent; they are exactly the same after 1919. We choose the contemporary series but their correlation from 1820 to 1918 is quite high, 99.6%.

of much greater consequence than the number of observations per se. (Shiller and Perron, 1985, Davidson and MacKinnon, 1993). Second, given our focus on inflation in this study and the tendency for this to be generally calculated as a year-on-year change due to seasonality issues makes an annual sampling of data preferable. Due to differencing of the log real GDP and log RPI the starting date of our main analysis coincides with the establishment of the formal Gold Standard in 1821, when Britain adopted it following the introduction of the gold sovereign by the new Royal Mint at Tower Hill in 1816.

3.1 Unit root analysis

In light of possible changes in extended historic data, any consideration of unit root properties should also take into account possible trend breaks. Perron (1989) draws attention to the importance of trend breaks for the conduct of unit root tests, with recent developments allowing for possible trend breaks under both the unit root null hypothesis and the trend stationary alternative. Therefore, we apply the procedure of Kejriwal and Perron (2010) that employs sequential hypothesis tests to estimate the number of trend breaks and are robust to the unit root properties of the data, with unit root tests then applied to the appropriately detrended data. This procedure is described in Kejriwal and Lopez (2013). In brief, the stability of the trend function is initially tested against one break in slope and level using the test of Perron and Yabu (2009). If this is rejected, the unit root test allows possible multiple trend breaks under both the null and alternative hypotheses using the procedure of Carrion-i-Silvestre, Kim and Perron (2009). Unit root tests are then based on so-called M-tests analysed by Ng and Perron (2001) with quasi-generalized least squares (GLS) detrending and lag specification as proposed by Qu and Perron (2007).

[INSERT: TABLE 1]

Estimation of trend breaks using the Kejriwal and Perron (2010) procedure requires a priori specification of the maximum number of breaks. Although the appropriate maximum number is unclear, we specify this to be three. Since their method is a sequential procedure based on sample splitting, we allow the 'trimming' parameter required in the structural break procedure to increase as the effective sample size decreases. In particular, we set this to 0.15 when only one break is considered over the whole sample; increase it to 0.20 in the subsequent splitting of the original sample. All tests are conducted at an asymptotic 5% level. We use AIC with the maximum lag length selection rule of $p_{max} = integer(4 \times (T/100)^{1/4})$ to select the appropriate

AR process. It has to be mentioned that the break dates are re-estimated every time after the sequential identification of the additional break.

Numbers in the first row of panel A in Table 1 are the exponential Wald statistics discussed by Perron and Yabu (2009) for one break in both slope and level, valid for both I(0) and I(1) error components, while those in the rows below are the suprema of that statistic when applied to the subsamples defined under the corresponding null, the critical values for such sequentially applied statistics being provided in Kejriwal and Perron (2010). Five percent critical values corresponding 0 vs. 1 break, 1 vs. 2 breaks and 2 vs. 3 breaks hypotheses are 3.12, 3.66 and 3.71, respectively. Statistics in panel B are GLS-detrended M unit root test statistics, analysed by Ng and Perron (2001) under no trend break case and extended by Carrion-i-Silvestre *et al.* (2009) to multiple breaks; bold statistics correspond to the number of breaks selected through the Kejriwal and Perron (2010) procedure. Since the critical values are case-dependent they are provided next to each test statistic in square brackets. Stars indicate the rejection of the relevant null hypothesis (namely, the lower number of indicated breaks in panel A and presence of a unit root in panel B) at the 5% level. Rejection in panel A (panel B) applies when the statistic is greater (less) than the critical value. All critical values are asymptotic and from the indicated studies.

Summary results are reported in Table 1 with panel A showing that the sequential procedure indicates the maximum of three breaks for short term rates and log real GDP and two breaks for log RPI.[4] The rejection of the null of 2 breaks against 3 breaks is marginal in both cases as the sequential test statistic $Seq(3/2)$ of 4.3 and 4.44, respectively, are closest to its critical value of 3.71. Panel B provides unit root test statistics for all numbers of breaks from zero to three. Overall, the conclusion that log real GDP and log RPI are I(1) is robust. Results for short term rates depend on the number of breaks allowed. I(0) is preferred if one assumes two breaks. We consider interest rates in level. Economically such an inference is also justified. Differenced log real growth can be seen to be I(0). Differenced log RPI appears to contain unit root if one allows for 2 and 3 trend breaks, but this also can be a manifestation of outliers or volatility breaks; otherwise it appears to be stationary.

[4]Further results relating to trend breaks, together with a range of unit root test statistics, are available from the authors on request.

3.2 Univariate analysis

Since our annual data span almost two hundred years, during which time potentially many structural changes and extreme events happened, it is interesting to consider univariate properties of each series. To that end we use the univariate iterative testing procedure of Bataa, Osborn, Sensier and van Dijk (2014), which decomposes observed variable Y_t into components capturing level (L_t), outliers (O_t) and dynamics (y_t) using

$$Y_t = L_t + O_t + y_t \tag{1}$$

where structural breaks are permitted in all components (except O_t). L_t models the mean which is allowed to change over time. Dynamics are captured through an AR model without an intercept, with breaks permitted, not only in the AR coefficients, but also in the disturbance variance. All breaks (level, dynamics and variance) can be at distinct points of time, and hence regimes are defined as specific to these individual components, while outliers are detected allowing for any mean shifts[5].

[INSERT: TABLE 2 here]

Formal testing is done in an iterative fashion and results from the convergence of this procedure are shown in Table 2[6]. Tests are conducted at 5% significance level, with 10% trimming.

In panel A the null hypothesis of no break against an alternative of up to a maximum 8 breaks is tested using the $WDmax$ test (often termed as Overall test). When $WDmax$ statistic exceed its 95% critical value, which is provided in the table footnote, the null is rejected and the numerical value of the statistic is marked with stars. Otherwise we conclude that there is no break. If $WDmax$ statistic rejects, the identification of the exact number of breaks proceeds in a sequential manner as in Bai and Perron (1998). We first test the null of one break against an alternative of two breaks using $Seq(2/1)$ statistic. If it exceeds its 95% critical value, which is also provided in the table footnote, we reject the null and the numerical value of the statistic is marked with stars. Otherwise we conclude that there is just one break. Such sequential tests are applied until we fail to reject the null or we reach the maximum possible number of breaks.

[5]Our definition of an outlier is based on 5 times Interquartile Range (IQR) away from the median (after removing the mean breaks) as in Bataa *et al.* (2014). For robustness, we also tried 7 times IQR but the results did not change much.

[6]Number of iterations required to converge is shown in the bottom of the table. The main loop iterates between outliers, and level and dynamic breaks, while the sub loop iterates between break dates in AR coefficient breaks and AR model residual variance breaks

Once the number of breaks are identified the locations of these breaks are estimated using the quasi maximum likelihood procedure of Qu and Perron (2007)and reported in bold, along with their 90% confidence intervals. Numbers below these dates are means in the sub-sample, defined by above break dates. We also report the whole sample mean, ignoring the breaks, in brackets.

Although we allow for up to 8 breaks we find no break in the level of mean growth rate and only one break in inflation in 1914 and two breaks in short term rates; break dates are rather precisely estimated with relatively tight confidence intervals. The inflation break coincides well with the collapse of Gold Standard. Deflation is found to be a regular feature of the UK economy during its reign. Indeed the annual average inflation rate is -0.06% a year prior to the break compared to 2.6% a year in the whole sample. Most periods of deflation also appear to have been largely unanticipated, with short term interest rates rarely approaching their zero lower bound and no corresponding break is found around 1914. Short term interest rate and inflation are lower in general in the earlier period, with short rate reaching exceptionally low level of 0.54% a year after the pound left the gold standard in 1931 and until 1951[7].

We find two deflation outliers, dated at 1921 and 1922, in panel B. These are related with the end of World War I and follow the depression of 1919-1921.

Panel C reports on break test results associated with the dynamic component. Explanations of $WDmax$ and $Seq(2/1)$ statistics are the same as in panel A. AIC, HQ and SIC are the lag length chosen by the respective information criteria and we use AIC to ensure no residual serial correlation. If there are breaks in the autoregressive parameters we estimate persistence measures, the sum of autoregressive parameters in this case, across across regimes, i.e. subsamples, and ignoring the breaks (in brackets) in the row below.

Among the persistence measures that of inflation is hotly debated. It measures speed that inflation converges to its mean after a shock to inflation process. Our univariate analysis fails to detect by only a small margin any changes in this tendency which prevents inflation from reverting to its previous path has been stable; $WDmax$ test is 10.50 while the critical value is 10.67. Although our more powerful multivariate analysis in the next section will overturn this conclusion, this is worth noting that a failure to reject the null hypothesis of constant inflation persistence around the collapse of the Classical Gold Standard in 1914 seems to be at odds with

[7]Besides happening close to the 30% devaluation of the pound sterling in 1949, and the start of the post-war socio-economic developments, 1951 is also close to the reopening of the free gold market after the war. Gold is often regarded as a safe investment. Thus as a legitimate investment alternative to cash saving opens up the interest rates has to increase. However 19 years is the shortest possible regime since in order to statistically detect a break we need at least 10% of our total number of 193 observations to lie between any two breaks.

the earlier economic and econometric literature[8].

We find no persistence break in growth rates. Short rate persistence emerges in 1940, but the 90% confidence interval is rather imprecise: between 1926 and 1954. Failure to acknowledge a break is nontrivial however. The persistence is 0.54 prior to the break and it behaves almost like a unit root process afterwards, in contrast to 0.87 estimated over the whole sample. This is after allowing for the mean breaks in panel A.

Panel D provides results on the volatility break tests. In addition to test statistics and break dates, it also reports estimates of the residual standard deviations of the AR processes, across the volatility regimes defined by the break dates and also ignoring these breaks and using the whole sample residual (but taking into account of breaks in the other components: level and AR coefficients).

It is interesting to see that volatility of all three series remain constant for at least 100 years until the WWI unsettles the growth volatility. Although we have another growth volatility break after the WWII their 90% confidence intervals are so wide that easily accommodate one another. Nonetheless post WWII period is characterized by a historic stability.

Short rate volatility almost disappears at exactly the same year as its level did. But there is a resurrection of not only its level but also volatility in 1951. In fact it is at historic high level of 1.56% a year since then but the actual beginning of this active short rate is again rather imprecise: its 90% confidence interval indicates it could have started as late as 1986. Inflation volatility was at least twice more volatile prior to 1980 which seem to coincide with the beginning of the Thatcher Government disinflation policies; but the confidence interval is extremely wide again.

Although Bataa *et al.* (2014) warn that confidence intervals in the context of iterative methodology are only indicative, the sheer width of these intervals for the dynamic and volatility breaks necessitate a more powerful multivariate set up, such as a VAR model, to better pin down the break dates. Indeed given the theoretical connection between these variables, such an analysis allows us to exploit the cross-sectional dimension for obtaining more precise inference on these structural breaks (see Bai, Lumsdaine & Stock, 1998 and Qu and Perron, 2007).

Lutkepohl (1989) show that Granger causality (GC) tests over-reject in finite samples when

[8]However when we use 10% significance level in the iterative procedure we do find a break in 1914 with 90% confidence interval between 1911 and 1917. One seemingly surprising finding is that there is no evidence of the second break in early 1990s that is often expected with the adoption of inflation targeting regime and to have marked a reversal of inflation persistence back to it pre-1914 patterns.

mean breaks are ignored. Moreover because the mean breaks in short rate and inflation series are both economically and statistically very significant they might coincide with tests for breaks in VAR dynamic coefficients and variance-covariance matrix. There is no prior information whether they coincide with potential VAR breaks and ignoring them is not a wise option either. Indeed Ng and Vogelsang (2002) demonstrate that inference based on the estimated VARs, forecasts and impulse response functions are generally invalid when mean shifts are omitted. They provide two choices: (i) remove the breaks before estimating the dynamic parameters, and (ii) estimate the mean shifts simultaneously with the autoregressive parameters. We choose the former as simulations in Ng and Vogelsang (2002) show that this is the best empirical approach when the level breaks in individual series do not coincide and when the breaks are so-called "additive outlier" type. The latter condition is already embedded in Bataa *et al.* (2014) procedure that we use. Blanchard and Quah (1989) and Gambe and Joutz (1993) also make the same choice. Thus, we take out 3 "local" means from the short-rate (1820-1930, 1931-1950, 1951-2014), 2 "local" means from the inflation rate (1820-1913 and 1914-2014). Since there is no mean break in the growth rate we remove its whole sample mean before our multivariate analysis.

4 Multivariate analysis

This section first describes the generalized VAR methodology of Koop, Pesaran and Potter (1996) and Pesaran and Shin (1998), the forecast error variance decomposition (FEVD) analysis of which was recently updated in Diebold and Yulmaz (2009, 2012, 2014). Given our data are annual and cover a long time span it is important to allow for correlated shocks. We rely on both the generalized impulse responses (GIRF) and the generalized forecast error variance decomposition (GFEVD) to derive relationship between growth, inflation and interest rates that is not sensitive to variable ordering. The GVEFD we use is the normalized version as in Diebold and Yuilmaz (2009, 2012, 2014), where the variance decompositions sum up to 100, and is explained is subsection 4.1. It then outlines the iterative structural break testing method of Bataa, Osborn, Sensier and van Dijk (2013) that would allow us to extend the Diebold and Yulmaz (2009, 2012, 2014) methodology by calculating confidence intervals for impulse responses and various spillover indices that are allowed to be different over time.

4.1 Innovation accounting

Koop, Pesaran and Potter (1996) and Pesaran and Shin (1998), referred to KPPS as in Diebold and Yulmaz (2012), provide a framework for studying impulse responses and forecast error variance decompositions that are insensitive to the variable ordering in a VAR model. Let us consider the following conventional VAR system for n demeaned variables:

$$\mathbf{y}_t = \sum_{i=1}^{p} \mathbf{\Phi}_i \mathbf{y}_{t-i} + \mathbf{u}_t \tag{2}$$

where $\mathbf{y}_t$ is a vector of stationary annual series. The error term $\mathbf{u}_t$ in (2) has mean zero and covariance matrix $E(\mathbf{u}_t\mathbf{u}_t') = \mathbf{\Sigma}$, and is temporally uncorrelated. The corresponding moving average form is

$$\mathbf{y}_t = \sum_{i=1}^{\infty} \mathbf{A}_i \mathbf{u}_{t-i} \tag{3}$$

Based on KPPS we can obtain the generalized impulse response function of the one unit shock by

$$\psi_j(H) = \sigma_{jj}^{-1} \mathbf{A}_H \mathbf{\Sigma} \mathbf{e}_j, \quad H = 0, 1, 2 \ldots \tag{4}$$

where σ_{jj} is j^{th} residual variance, and $\mathbf{e}_j$ is a selection vector with unity in the j^{th} element and zero otherwise.

This GIRF shows the effect of a shock to the jth element of $\mathbf{y}_t$,equal to one percentage in magnitude, on the vector $\mathbf{y}_{t+h}$ for $h = 0, 1, ...$

The variance decomposition that allows to assess the fraction of the H step-ahead error variance in forecasting y_i that is due to shocks to y_j, $\forall j \neq i$, for each i is the following:

$$\theta_{ij}(H) = \frac{\sigma_{jj}^{-1} \sum_{h=0}^{H-1} (\mathbf{e}_i' \mathbf{A}_h \mathbf{\Sigma} \mathbf{e}_j)^2}{\sum_{h=0}^{H-1} \mathbf{e}_i' \mathbf{A}_h \mathbf{\Sigma} \mathbf{A}_h' \mathbf{e}_j} \quad i, j = 1, \ldots, n \tag{5}$$

Notice that $\mathbf{\Sigma} = \mathbf{DPD}$, where $\mathbf{D}$ is a diagonal matrix of standard deviations and $\mathbf{P}$ is a correlation matrix; thus both the volatility and the correlations of the VAR residuals play a role in calculation of the IRF and FEVD alongside the VAR parameters.

Based on the normalized version of above Diebold and Yulmaz (2009, 2012, 2014) define total, directional, bivariate and net spillover measures. Given that their spillover measures are already very popular we omit their details here[9].

[9] In addition to the 3 references here they apply their methodology in three other articles, have one Oxford University Press published book and a website http:financialconnectedness.org. The above 3 papers only have

Our main interest is to test the stability of the above model. If the parameter stability is rejected we would like to see the consequences of those changes in the IRF, FEVD and the spillover measures. Parameter instability could originate from those in $\mathbf{\Phi}_i$, and thus in $\mathbf{A}_i$, $\mathbf{D}$ and $\mathbf{P}$. Therefore we use Bataa *et al.*'s (2013) iterative structural break testing method. Space considerations preclude a discussion of their method here but a full exposition may be found in the original study.

Their method, also used in Bataa *et al* (2015), allows one to test for and identify which of its individual elements are behind an identified break in $\mathbf{\Phi}_i$. Conditional mean specification (VAR) can affect the estimations of the variance-covariance matrix, and ultimately the FEVD and various spillover measures that are based on it. Indeed in a related literature, Cheung and Ng (1996) and Hong (2001) point out that causality in mean may affect tests for causality in variance. That is because causality in mean affects the structure of the disturbance terms. Moreover, Lumsdaine and Ng (1999), Mikosch and Starica (2004) and Blake and Kapetanios (2007) show that misspecification of the conditional mean, including structural breaks, may lead to spurious rejection of the null hypothesis of no ARCH. This stresses the importance of not only modeling all important breaks but also not imposing spurious ones.

5 Results

This section considers whether the parameters and the variance-covariance matrix of the VAR model are constant over time and the implications of that time-variation. In particular, subsection 5.1 examines possible structural changes in both the autoregressive parameters and the variance-covariance matrix of a VAR through Bataa *et al.*'s (2013) procedure, while subsections 5.2 and 5.3 identify the sources of these system-wide breaks. Finally, sections 5.4 and 5.5 examine the implications of the detected breaks through impulse responses and forecast error variance decomposition, respectively. To facilitate comparisons, all results are expressed in percentage terms.

5.1 Breaks in the VAR

Owing to the large number of autoregressive coefficients in the VAR, a maximum of three breaks is permitted ($M_c = 3$), with a minimum of 20% of the sample required in each regime while

678 citations as of July 2015 according to Google Scholar.

we allow for a maximum of eight breaks for the variance covariance matrix ($M_v = 8$ with 10% trimming).

The estimation of the VAR with 2 lags, guided by the information criterion and to ensure absence of residual autocorrelation, implies that two years of data are 'lost', so that the sample period for estimation starts in 1823.

The first column of Table 3 contains results for breaks in the autoregressive parameters of equation (2), while the second column provides structural break test results for the variance-covariance matrix. Explanations of the overall test (WDmax), sequential tests and break date estimations are the same as for Table 2, thus omitted here[10].

In contrast to Table 2 we have a new line called bootstrap p-value under the estimated break dates to check the finite sample validity of the asymptotic inference since we have much more parameters compared to our 192 observations. These bootstrap p-values evaluate the significance of the each break and we conclude that a certain break is significant only if it is less than 5%.

$WDmax$ statistic for the VAR coefficients are 232.86 compared to an asymptotic 5% critical value of 42.63. The asymptotic sequential Seq(2/1) test also rejects the null of one break (against two). Seq(3/2) is rejected and we conclude using the asymptotic distribution that there are two coefficient breaks. The bootstrap procedure confirms this conclusion (p-values of 0.04% and 1.07%). It is worth noting that the persistence break detected for the short rate at 1940 in Table 2 is completely explained away by the multivariate set-up.

As for the variance-covariance matrix the $WDmax$ test statistic is again highly significant and the sequential tests indicate five breaks. The bootstrap inference also confirms. It is noteworthy the variance-covariance matrix breaks are essentially those identified in the univariate analysis of subsection 3.2, with two exceptions. We did not find the breaks in XXth century in the univariate analysis and this shows that the multivariate analysis can indeed be more powerful in detecting breaks.

It is reassuring that although we allowed for a maximum of 3 breaks in the VAR coefficient and 8 breaks in the variance covariance matrix we do not reach this limit. This also casts some doubt on the suitability of TVP-VAR techniques which forces, without testing, each parameter

[10]Qu and Perron's (2007) test statistics have the same limit distributions as those in Bai and Perron (1998), who tabulate critical values up to 10 parameters. Although Bai and Perron (2003) provide response surfaces for estimating critical values, Hall and Sakkas (2013) show these surfaces can lead to misleading inferences with a large number of parameters. Therefore we simulate critical values suitable for our cases; details are available from the authors on request.

to change each period, in this case. 20% and 10% trimmings used in this case implies 38 and 19 observations, respectively, have to lie between any two breaks. However none of the breaks are subject to this boundary problem, except the second covariance matrix break in 1869.

Even after removing the level shifts/low frequency movements as discussed in Section 3.2. we find breaks in the dynamic relationships, i.e. in VAR parameters. In Section 3.2. level shifts occurred in the short rate in 1931 and 1951, and in inflation in 1914. These are removed accordingly and the VAR is estimated with demeaned data. But we still find VAR coefficient breaks in 1916 and 1975. Thus the outbreak of WWI, and the subsequent burst in both mean and volatility of inflation appears to have changed the dynamic relationship permanently. The second break could be associated with two large oil shocks in the 1970s.

The first covariance matrix break in 1850 very likely reflects a "revolution" in British free trade policy. There were many important events around this time, well known one is the Repeal of the Corn Laws in 1846 (see Irwin, 1988, among others). In the subsequent years the UK enjoyed a period of historic stability (panel B of Table 5). 1869 break could be associated with the US signal about future world monetary regime. The 1869 Public Credit Act of the US stated that bondholders who purchased bonds to help finance the Civil War (1861-1865) would be paid back in gold. Prior to that US was undecided as to whether it should operate on a greenback, gold, or bimetallic standard. Thus the Law was an early indicator that the country was moving toward reinstating the gold standard. After this break the volatility of both inflation and short rate declined (panel B of Table 5). 1951 covariance break marks the end of uncertainty and volatility associated with world wars and marks the beginning of the post-WWII stability.

Although our model is too simple to explain causality, the 1989 break perhaps reflects changes in monetary policy not only in the UK but also around the world. The confidence interval is admittedly rather loose, reflecting the turmoil of that period. In 1993, the then chairman of the US Federal Reserve, Alan Greenspan told Congress: "The historical relationships between money and income, and between money and the price level have largely broken down, depriving the aggregates of much of their usefulness as guides to policy." The Bank of England tried to target money supply and exchange rates in the 1980s and early 90s. In 1990 UK joined the European exchange rate mechanism (ERM), a system that kept currency fluctuations within limited bands. But on September 16, 1992 UK was forced to withdraw the pound sterling from the ERM and in few weeks came the adoption of inflation targeting.

The above results are based on VAR(2) model. When we use HQ to select the lag length this is estimated to be 1. The identified break datas from this smaller model are 1913 and 1977 for the VAR coefficients and 1885, 1913, 1932, 1951 and 1990 for the variance covariance matrix[11]. In fact, given the number of parameters is twice fever than VAR(2) the maximum number of breaks was set to 5 with 15% trimming, not 3 as in VAR(2). However, one lag was too short to take account of the whole dynamic structure: wild-bootstrapped *p-value* for the absence of 4^{th} order residual autocorrelation is 7% for the model estimated using data 1822-1913 and *p-value* for the absence of 1^{st} order residual autocorrelation is 1.3% between 1914-1977. Thus we had to increase the lag order to 2 which helps remove all the evidence of residual autocorrelation.

It may be noted that Bataa *et al.*'s (2013) iterative procedure converges quickly, after just 3 iterations.

5.2 Breaks in persistence and Granger causality

Conditional on the VAR coefficient breaks in the earlier table, Panel A of Table 4 reports on the origins of these breaks. Columns under the heading "Dependent variable" represent VAR equations. The first value (in bold) of each cell (in each column) reports the difference between the sum of the autoregressive coefficients after and before the break date, with this placed against the dates of the second subsample used in the comparison. Subsamples are those implied by the estimated structural break dates of Table 3. The first value in each cell in panel B reports the sum of estimated coefficients over the indicated subsample. In both cases, the values below are bootstrap p-values (expressed as percentages) for the null hypothesis that the corresponding true value is 0. If an individual coefficient break is not significant at 15% in panel A, the corresponding subsample coefficients are restricted to be equal in panel B. Two and one stars and a spade indicate significance at 5%, 10% and 15% respectively using wild bootstrapped p-value.

We see that persistence increases for all three series in 1916, which is around WWI. The difference between the sum of autoregressive coefficient after and before this date is highly significant for short rate ($p-value$ of insignificance is 3.34%) and is significant at 10% for growth rate and inflation. Persistence, defined as the sum of autoregressive coefficients, becomes significant only after that break, as can be seen from Panel B. Increased persistence in the short rate is in line with our univariate analysis in section 3.2 although the multivariate analysis picks up the break earlier. Notice that the univariate analysis fails to detect a persistence break in

[11]Not reported in the table but available upon request.

both inflation and growth rate. There is no statistically significant evidence that the persistence further changing in 1975. This is in contrast to a large literature that documents a decline in inflation persistence in 1970s using monthly CPI (see e.g. Benati, 2008b, and Bataa, *et al* 2013, 2014, among others).

In 1916 the way inflation is Granger-caused by growth changes, the causality becoming insignificant from being highly inflationary. There is no reverse Granger causality. Short rate Granger causality by growth looses significance between 1917 and 1975. The causality is very strong since then, perhaps reflecting an activist monetary policy aimed at supporting real economic growth. On the other hand higher short rate always affects growth negatively. Short rate starts to Granger cause inflation since 1917. On the other hand there is quite strong evidence against inflation Granger causing short rate since 1976 [12] .

5.3 Breaks in volatility and correlation

Table 5 identifies the origins of the covariance matrix breaks, whether they are from volatility breaks and/or correlation breaks and provides corresponding subsample estimates. The values reported are the final ones computed in the respective general to specific procedures. Conditional on the significant correlation breaks the table also provides instantaneous causality test results. Individual VAR coefficient breaks that are significant at least at 15% are allowed to change while the others are estimated using the whole sample in estimating the VAR residuals and, subsequently, the variance covariance matrix.

The first column of panel A reports the significance of structural break tests for diagonal elements of the covariance matrix of the VAR (volatility) and the second column for off-diagonal elements i.e., the correlation matrix, showing bootstrap p-values for the test of no change over adjacent covariance matrix subsamples identified in Table 3, with the result placed against the dates of the covariance matrix break date.

There is overwhelming evidence that the latest of the variance-covariance matrix breaks

[12] When we enforce all the coefficients break at the system-wide break dates the growth shows no persistence until as late as 1976 and only since then there is Granger causality from the short rate, although it is significant only at 10%. This contradicts with the inference derived allowing for only significant breaks, where the growth persistence is detected earlier and the short rate always Granger causes it. Moreover the Granger causality from the short rate to the inflation is detected only since 1975, much later than in B. If one ignores all the breaks then firstly economically and statistically important persistence changes will also be ignored. Moreover, a strong Granger causality from the growth to the inflation prior to the WWI will be diluted. A period when growth did not Granger cause short rate will also be missed. Furthermore, that fact that short rate is caused by inflation between 1917 and 1975 will disappear. Finally, inflation will never be Granger caused by the short rate, which is hard to justify given the activist monetary policy in recent period.

is not a volatility break. But there is equally strong evidence that this break in 1989 is an important correlation break. 1914, as a correlation break, is significant only at 10% while the remaining variance-covariance matrix breaks are unlikely to be associated with correlation[13].

Panels B and C show sup-sample estimates of residual standard deviations and correlations and also those ignoring the breaks [in square brackets]. Between 1915 and 1951, a period that includes both world wars, growth and inflation are most volatile, but post war stability is remarkable (standard deviations are 1.49% and 2.15% respectively), even compared with their long term values (2.51% and 2.83%). In contrast short rate volatility almost doubles: standard deviation is 1.14% after 1951 compared to 0.77% immediately before it and 0.67% during the classical Gold Standard. Correlation between growth and inflation more than doubles after 1990 and that of between growth rate and short rate triples. Most strikingly, the correlation between the short rate and inflation increases ten-fold from a very low base (from 4.8% to 48.1%) and is extremely high by historical standard (9.4%).

Panel D contains no-instantaneous causality test results before and after 1990 and also those ignoring this significant correlation break. Growth is contemporaneously uncorrelated with both inflation and short rate before this break. If this break is ignored one would have missed this important fact. Both inflation and short rate however are always instantaneously caused by others in the system.

5.4 Impulse responses

A useful qualitative comparison is provided by the impulse responses computed across subsamples. What is allowed to change across subsamples is however important. In terms of the VAR coefficients we are going to be quite liberal and allow for breaks that are significant at 15% in Panel A of Table 4. With regards to the five variance-covariance matrix breaks we are not considering 1989 as a volatility break, but it is the only correlation break, as discussed in the previous sub-section.

Figures 1-3 show generalized impulse responses of three variables over ten years after a one percentage point shock. Rows indicate the variables that respond to the shock. Dotted and

[13] 1914 break looses its 10% significance as a correlation break when Granger causality is imposed. It becomes impossible to make a clear-cut inference about the sources of the variance-covariance matrix if one unnecessarily splits the sample or ignores altogether important structural breaks in the VAR coefficients. It appears imposing constant VAR parameters results in spurious correlation breaks more often than such volatility breaks. Further imposition of Granger causality also appears to play some role with respect to inferences about correlation breaks. Nonetheless it was very clear even from these analyses that correlation from 1990 onward is very different from than before. Details are available from the authors upon request.

dashed lines (in blue) with shaded confidence intervals assume parameter constancy. A response in solid (red) line with dotted confidence interval is specific to the sub-regime specified on the top. For each plot two-standard-deviation wide confidence interval is obtained using a recursive design wild bootstrap procedure. The columns represents the sub-samples defined by the break dates in Tables 4 and 5.

5.4.1 Growth shock

The first thing to note from Figure 1 is that a growth shock does not exhibit strong persistence using the full-sample model with no breaks. There is some increase in the effect of the shock from 1% to about 1.25% for year 1 and 2 following the shock. However, the impact of this growth shock dissipates quickly and the net effect of the shock is little more than the initial shock of 1%. Hence the only long run impact on output is a shift equal to the magnitude of the initial shock but no change in the long-run growth rate of output.

The persistence of a growth shock differs somewhat across regimes when structural breaks are accounted for. The pre-1917 responses are weaker and the post-1975 responses are somewhat stronger. Pre-1917, the initial 1% shock partially reverses and the cumulative response by year 4 is about 0.6%, where the response stabilises; the long-run effect (by year 10) is statistically less than the initial shock and statistically less than the full-sample point estimate. For the post-1975 sample the response follows a similar path to the full-sample in the first couple of years, but from year 5 onwards there is a clear positive trend to the response; the post-1975 responses also have wide confidence intervals and hence the estimated response is not different to that of the full sample.

Our empirical analysis finds that a real growth shock has a positive impact on the short-term interest rate for the full sample period (using the simple / baseline model). The magnitude of the effect is very modest, it is cumulatively about 0.7% (1.1%) after 5 (10) years, which equates to an average rise of about 0.12% (0.11%) in the level of short rate from an initial 1.0% growth shock. However, after allowing for structural breaks our analysis paints rather a different picture for the relationship between real growth and short-term interest rates. Firstly, there is a very weak response of short-term interest rates to growth in the pre-1975 sub-samples; here the cumulative impact on the short rate is less than 0.6%, which implies an average rise of not more than 0.12% in the short rate for years 1-5 and no impact on the short rate for years 6-10. Secondly, the effects of growth shocks on interest rates are much more apparent for the 1976-1989 and 1990-2014 sub-periods. The magnitude of the effect for 1990-2014 peaks at approaching 7% in year 10 and is about 4.8% by year 5; this indicates an average rise in the short rate of 0.96% for the first 5 years which declines to about an average of 0.44% for years 6-10. This suggests that short rates are substantially more responsive to growth shocks in the most recent sub-sample (1990-2014) compared to the full sample. The strong response of short-term rates to growth during the 1990-2014 period is also somewhat surprising given that this period is mainly characterised by the Bank of England following an inflation targeting regime during which it is primarily supposed to be focusing on maintaining price stability rather than responding strongly to fluctuations in growth! However, we should note that it is plausible that this partly reflects the anticipated impact of a growth shock on inflation. Further, the 1990-2014 sub-sample responses are somewhat stronger but not significantly different from the 1976-1989 period suggesting that the responsiveness of interest rates to growth shocks has not increased dramatically since the 1970s.

[INSERT: FIGURE 1]

We now examine the impact of a real growth shock on inflation. Using the simple model estimated over the full sample period the IRFs indicate that a real growth shock has a positive cumulative impact on the inflation for the full sample period (using the simple / baseline model), this can be interpreted as leading to a permanent increase in the price level. The magnitude of the effect is fairly moderate; it is cumulatively about 1.25% (1.5%) after 5 (10) years, which equates to an average rise of about 0.25% (0.15%) in the level of inflation from an initial 1.0% growth shock. Thus it seems that a growth shock does stimulate some demand-pull inflation which is apparent up until around year 5. This is interesting given the more modest impact of the growth shock on growth itself after year 2.

It seems as if the general effect of demand-pull inflation is evident during all the sub-sample periods. The point estimate effect stabilises more quickly for the pre-1917 sub-periods (by year 4) at typically a slightly weaker magnitude for the full sample. For the 1917-1951 and 1952-1975 the response cumulates to about 2% and continues to increase, albeit modestly, through until year 10. However, the sub-sample with the largest difference to the full-sample is 1990-2014; here the response of inflation to growth shocks is much stronger initially and peaks at around 3% in year 5-6 but then slightly reduces over the last few years. Especially in the years immediately following the shock the effect is almost double for 1990-2014 compared to the full sample; the (cumulated) inflation response reaches almost 2.5% by the end of year 3 which equates to inflation on average being more than 0.83% higher during years 1-3. Thus overall there is evidence consistent with demand-pull inflation being at work in the UK economy, which is especially strong during the 1990-2014 sub-period.

5.4.2 Inflation shock

Figure 2 depicts IRFs to a 1% inflation shock. UK inflation shocks for the full-sample model are quite persistent for the first five years but then have a modest additional impact beyond year five. In particular in the baseline model the shock seems to have a cumulated increase on inflation (an impact on the price level) of about 2.5% by year 5, with the effect seeming to get smaller each year. After year 5 there the cumulated response of inflation rises modestly, by about 0.4% in total over years 6-10 (i.e. about a 0.1% increase in inflation per year).

However, there is some time-variation in the effect. Prior to 1917, the effect of an inflation shock on inflation is substantially weaker and peaks in year 1 at a cumulated increase of about 1.5% before slightly dissipating. Hence, during this period there appears to be no long-term effect on the price level. However, in the later sub-samples the an inflation shock has more of a persistent effect and peaks around year 5 at a cumulated response of about 3% and then is approximately flat for years 6-10. The main message from this is that inflation shocks in general primarily have a medium-term effect on subsequent inflation after the initial shock, i.e. the second round and third round consequences of inflation shocks tend to be rather substantial. However, after year 5 there is little evidence that inflation continues to cumulate upwards.

In contrast, for post 1917 sub-samples the response of inflation to an inflation shock is somewhat stronger than the full sample. In the first 5 years the impact exceeds 3% and continues rising (albeit at a slower rate) until at least year 8. Nevertheless the post 1917 responses are not statistically different from the full sample. This evidence corroborates the US results of Fuhrer and Moore (1995) that there is high inflation persistence post-WWII and we provide novel evidence from IRFs. Our results suggest that this high inflation persistence actually dates back to 1917 in the UK; this high persistence of annual inflation over the last century is apparent across different monetary regimes and hence is not entirely consistent with Benati (2008a) who using quarterly change in inflation reports some variation across different settings e.g. less persistence in the UK following the introduction of inflation targeting. The last 14 years of the Gold standard (1917-1931), we find are characterized by higher inflation persistence; however, consistent with Benati we find that inflation persistence was lower in earlier periods prior to

1917, when the Gold standard was also in force. Hence our results broadly suggest that it is not a change in monetary regime that is the source of high inflation persistence.

Our empirical analysis finds that an inflation shock has a tiny impact on the short-term interest rate for the full sample period (using the simple / baseline model). The magnitude of the effect is very modest, reaching a cumulated effect of about 0.4% by year 5, and gradually increasing to about 0.6% by year 10. Thus the response of the nominal short-term interest to an inflation shock is much smaller than one would expect under the Fisher hypothesis where the effect should be 1:1 (if the shock were permanent); for example a 1% inflation shock leads to an increase in the nominal short rate of less than 0.1% p.a. This appears to be consistent with prior literature that the response on interest rates to inflation is less than unity (for example Mishkin, 1992; Evans and Lewis, 1995). This also suggests that this key monetary policy tool is not adjusting sufficiently to counteract the effect of inflation shocks since an interest rate response greater than unity would be necessary to induce a fall in inflation in the future. Finally, there are no significant long-run effects in the UK which appears inconsistent with the post-war results of Rapach (2003) but consistent with money neutrality, this is evidenced by the stabilising of the cumulated response function of the interest rate at an approximately constant level in the later years and this not being statistically different from zero.

Crucially when structural breaks are allowed for then there are key differences. Prior to 1976, IRF's are broadly similar to the full sample period. However, for 1976-1989 and especially for the 1990-2014 sub-period there are striking differences. For 1990-2014, there is a much stronger response of the short rate to the inflation shock than during the full sample period; the cumulated response continues to rise throughout the sample period reaching about 5% by year 10, in contrast to about 0.6% for the full sample model. Thus, there is substantial adjustment of interest rates to inflation during this period. It suggests that during 1990-2014 a 1% inflation shock would lead to about a 0.7% average rise in short rate for the first 5 years and this would remain on average about 0.2% p.a. above the initial level for years 6-10. These results are also statistically different from those reported for the full sample. Nevertheless even during this time money growth targeting (during the 1980s) and inflation targeting (from 1997) were implemented, but yet we find that there is a less than a unity response to the inflation shock in the short-term. This seems rather surprising since one would expect the nominal interest rate would need to adjust by more than unity in order to counter-balance the effect of the inflation shock.

[INSERT: FIGURE 2]

What happens to real growth following an inflation shock? According to the no break model estimated over the full sample period there is a modest impact; the IRFs indicate that an inflation shock has a moderate impact over the first five years leading to an output loss of about 0.2% in total (cumulative growth loss), which are strongest in years 2 and 3. However, from year 5 onwards there is (essentially) no further decline in output but little increase either. These results for the full sample suggest that an inflation shock has a very modest impact on growth in the medium term and leads to a permanent loss of output of less than 0.25%. Furthermore the IRF is not statistically different from 0 at any point during the 10 year period. Thus, there is limited support overall for the view that inflation costs growth from this no break model.

Nevertheless, the results differ somewhat when structural breaks are modelled. For 1952-1975 we find a stronger negative response of cumulated growth to an inflation shock, which hits a trough at about -0.4% (-0.5%) by year 5 (10), although the effect is not statistically distinguishable from 0. In contrast for 1990-2014 there is a very weak response of growth to an inflation shock which never cumulates to a loss of more than 0.2% and indicates an increase in growth from year 3 onwards.

One plausible reason for a positive response of growth to inflation is if the inflation is caused by external i.e. international factors, which indicate strong global demand and hence strong sales of domestic output overseas. This is more likely during the post world war II era given the almost exponential growth in trade and the growing international economic interdependency during this period.

Interestingly, we find evidence generally in favour of long-run money neutrality here since there is no significant long-run effect on the rate of growth and a fairly modest and insignificant long-run effect on the level of output.

5.4.3 Nominal short rate shock

Figure 3 indicates that UK short rate shocks for the full-sample model are rather persistent and the cumulated increase in the short rate continues through until year 10; in other words the interest rate remains above its pre-shock level even after 10 years. However, the short rate does gradually mean-revert, which is evidenced by the increase in the cumulative effect getting smaller as time progresses. The cumulated effect of the shock is about 2.5% after two years, rising to about 4% after five years and about 5% by year ten.

However, the effect does vary over time. Prior to 1917, the effect of a short-term shock on itself is substantially weaker and peaks after just two or three years at a cumulated increase of about 2% and remaining at this level thereafter. This suggests that prior to 1917 the short-term rate itself reverts from the shock within three years; hence during this period there appears to be no long-term effect on the level of the short rate. However, in the other sub-samples the short-term rate shock has a longer lasting impact. For 1917-1951 and 1952-1975 it broadly resembles the full sample, but is slightly weaker, whilst for 1976-1989 it closely resembles the full sample and 1990-2014 the effect is somewhat intensified such that the cumulated response reaches more than 8% by the end of year 10. Thus, it is clear that the short-term rate is highly persistent for the most recent periods but was not very persistent at all prior to WWI.

A nominal short term rate shock has a negative effect on growth for the full sample model; however, the impact is generally weak and statistically insignificant. The effect of a 1% shock to interest rate for the full sample has a cumulative impact on growth which is never greater than 0.5% magnitude in absolute terms. It is also never statistically different from 0, which suggests there are no significant long-run effects. This suggests that the effectiveness of this key monetary policy tool to stimulate economic activity is limited.

However, the sub-sample analysis reveals a somewhat different picture. Our model that accounts for structural breaks indicates that between WWI and the 1970s (for the 1917-1951 and especially 1952-1975 sub-periods) the use of an increase in interest rate to dampen economic activity may have been effective. For the 1952-1975 sub-period in the short-term (for the first two years) there is a sharp negative response of growth to a short rate shock of about -1.5% which then diminishes but the cumulative impact on growth reaches -2.25% by year 5 and -3% by year 10. These results for 1952-1975 are statistically significant as well as economically substantial. Our results suggest there is a statistically significant negative long-run effect apparent after 10 years. Hence, the interest rate could have been an effective tool for managing growth in the UK economy during the aftermath of WWII.

However, in the later part of our sample (1976-1989 and 1990-2014) there is very little evidence that a short rate shock has a substantial negative effect on growth. The point estimates of output losses are never more than 0.6%. Actually for 1990-2014 point estimates suggest output gains not losses from a short rate shock, although these are not statisitcally distinguishable from 0. Hence, unexpected monetary policy shocks are almost totally ineffective over this recent period in stimulating or constraining the economy. Further, it suggests that the perceived wisdom of a strong correlation of short-term interest rates with real output (e.g. Bernanke and Blinder, 1992; Estrella and Hardouvelis, 1991) does not hold over recent years. This raises the question of which policy tools are there nowadays that will actually assist the Bank of England in prudently managing the domestic economy. The finding of no long-run effect of interest rate on growth seems consistent with Haug and Dewald (2011) who report money growth is unrelated to output growth at longer horizons.

Another issue is that for the majority of the 1990-2014 period the UK was following inflation targeting (or monetary growth targeting) and hence growth concerns should not have been forefront in policymakers minds. Nevertheless, we saw in Figure 1 that nominal interest rates respond to growth during all periods especially since the 1970s. Hence the evidence from Figure 3 suggests that if the interest rate was rising out of macroeconomic management concerns then it may not have been meeting its objectives of stabilising growth.

[INSERT: FIGURE 3]

We now examine the impact of an interest rate shock on inflation. The conventional view is that a rise in interest rates leads to lower inflation, ceteris paribus. A shock to the nominal short term rate has a modest positive effect on inflation which peaks within the first few years. The base model estimated over the full sample period shows that the IRFs indicate that the a nominal rate shock has a positive impact on inflation especially during the first three years reaches a cumulative effect of +2% and then peaking at a cumulative effect of about +2.5% by year 10. According to the point estimates this is statistically significant from 0 almost throughout the 10-year horizon examined. This is somewhat surprising given that an increase in the short term rate is conventionally thought to lead to reductions in inflation. However, these results are consistent with the price puzzle (Sims, 1992).

A couple more observations are pertinent which add to price puzzle literature. First, for the pre-1915 sub-samples the price puzzle is weaker. It peaks after the first year at never more than +1.5%. Second for sub-sample periods for 1917-1951, 1952-1975 and 1990-2014 the effect of a short-term interest rate shock are much stronger than for the full sample and reach a peak of at least 5% by year 5 and even approach 10% for the 1917-1951 period. Our results are not entirely consistent with Hanson's (2004) US evidence on the price puzzle which suggests the price puzzle is concentrated in the 1959-1979 sub-period rather than post 1980. Our UK results suggest the price puzzle is strongest for 1917-1951, and of similar magnitude for 1952-1975 and

1990-2014, but weak over 1976-1989. Hence, there appears cross-country variation in the nature of the price puzzle.

Furthermore, the existence of the price puzzle in our data could be attributed as suggested by Balke and Emery (1994) to an insufficient adjustment of the short rate to high inflation. That is the movement in the nominal short rate may not be sufficient to counteract the rise in inflation, in such a scenario the real interest rate will fall (not rise) which may help explain why inflation tends to increase rather than decrease following a nominal short rate rise.

5.5 Forecast error variance decompositions

A comparison of forecast error variance decompositions (FEVD) over sub-periods also sheds light on the the changes in British economy over the last two hundred years. We also provide various connectedness measures based on FEVD.

Table 6 uses generalized VAR for the UK[14] Connectedness tables are provided for two horizons: h =1 and 4 years[15]. The ij^{th} entry of the upper-left 3×3 series submatrix gives the ij^{th} pairwise directional connectedness; i.e., the percent of forecast error variance of series i due to shocks from series j. They are allowed to be different over the sub-samples defined by the statistically significant structural breaks. If bootstrap one (two) standard deviation(s) wide confidence interval does not contain zero they are marked with *(**). The quantities estimated over the whole sample, ignoring the structural breaks are in square brackets. The rightmost (From others) column gives total directional connectedness (from); i.e., row sums (from all others to i). The bottom (To others) row gives total directional connectedness (to); i.e., column sums (to all others from j). Next (Net) row gives the difference in total directional connectedness (to-from). The far-right columns give the total connectedness 'from others' and the rows below the short rate give the total connectedness 'to' others. The bottommost row also gives bi-variate spillovers.

5.5.1 Forecast error variance decomposition results

The first column in each section of Table 6 reports results for the FEVDs when the initial shock occurs (i.e. the period 0 shock). We consider movements of each variable in turn. The point

[14] A table that relies on Cholseky orthogonalization is available from the authors upon request and is omitted here conserve space. The variable ordering is first economic growth, then inflation and finally short term interest rate.

[15] There is not much movement after 4 years of the shock and essentially converges.

estimates is that growth shocks account for 98.9% of the future movement in growth whereas only 1.0% is from inflation and 0.1% from the short rate. Moreover, with two standard deviation confidence intervals provided we cannot reject the hypotheses that i) growth shock accounts for 100% of the movement in growth, ii) inflation accounts for 0% of the movement in growth and iii) the short rate accounts for 0% of movement in growth. This underlines that over the full sample period the evolution of growth is almost entirely determined by shocks to growth itself. The sub-period analysis for growth broadly supports the main findings of the full sample at the 1 year horizon except for the latest sub-period (1990-2014). For 1990-2014, there are a several pertinent differences. First, the amount attributable to the growth shock, falls substantially to 70.5%. Second, the amount attributable to inflation increases to 9.7%, which is more than other sub-periods (2.1%) and the full-sample (1.0%); nevertheless none of these are statistically different from zero when the two standard deviation confidence interval is used. Thirdly, the variance attributable to the short rate is tiny pre-1989 (0.4%) but rises dramatically for the post-1990 sample to 19.8% and becomes statistically different from zero.

Estimates for inflation at the one year horizon indicate that this is also primarily determined by shocks to itself. Point estimates are above 90% for the full sample and above 85% for all sub-periods except 1990-2014. Growth shocks, perhaps surprisingly have very little effect on movements in inflation, at least initially; point estimates are never more than 2% and never statistically different from 0 apart from during 1990-2014, when this rises to 7.3% but is not different from zero using a two standard deviation confidence band. In contrast, short rate shocks do have an impact on inflation where point estimates are always more than 7.5%. Short rate shocks are statistically different from 0 at the 5% level (two SD band) for the full-sample and all sub-periods. Since 1990 short rate shocks can account for 39.6% of the movement in inflation (one year ahead).

Estimates for the short rate indicate that one year ahead deviations from the model forecast is also primarily determined by shocks to itself for the full sample and almost all sub-periods. Point estimates are always above 90%, with the exception of the 1990-2014 sub-sample. Growth shocks, perhaps very surprisingly, have very little effect on movements in the short rate, at least initially; (prior to 1990) point estimates are never more than 0.5% and never statistically different from 0. In contrast, inflation shocks do have an impact on inflation where point estimates are always more than 7.5%. Inflation shocks are statistically different from 0 at the 5% level (two SD band) for the full-sample and all sub-periods. Since 1990 inflation shocks can account for 36.8% of the movement in the short-rate and growth shocks can account for 13.8%; thus, together they can explain 50.6% of the variation in the short-rate since 1990, while the short rate itself explains less than half (49.4%).

The second column of each sub-section of Table 6 reports results for the 4-year forecast horizon (h=4). The empirical results here show some important differences relative to the 1-year forecasts. In general, at the 4-year horizon, less of the variation is determined by the variable itself, although there is more variation in sub-periods, the results for the 1990-2014 sub-period are typically amongst the lowest. For growth, shocks to growth itself explain above 90% of variance for the full-sample (also this declines from 98.9% at h=1 to 90.6% at h=4). In most sub-periods growth shocks explain more than 90% of the variance, although this is below 65% for 1851-1869, 1952-1975 and 1990-2014. These sub-sample reductions for growth are accompanied by an increase in importance of the short rate for each of these sub-periods; in fact the short rate can explain more than 20% of the 4-year FEVD for growth in each of these 3 sub-periods (1851-1869, 1952-1975 and 1990-2014) and these are significantly different from 0 (using two SD band). In contrast, the contribution of inflation is only statistically different from 0 during the 1990-2014 sub-period (using two SD band) when it can explain 17.4% of the variance of growth. Thus, this suggests that for the recent period (1990-2014) both short rate shocks and inflation shocks have important effects on growth during this period and that the central bank should bear this in mind when formulating monetary policy.

The empirical results for inflation FEVD at the 4-year horizon show some key differences relative to the 0-year forecasts. First, less than 85% of the error variance is attributable to inflation shocks itself for the full sample and for all sub-periods. The point estimate is 82.7% for the full sample but is substantially lower at less than 50% for the 1952-1975 and 1990-2014 sub-samples. We can also now reject for the full sample that 0% of the error variance is attributable to the short rate during post 1917 sub-periods and it can explain more than 15% of variance during these sub-periods. Prior to 1916, Growth can explain more than 20% of the 4-year forecast error for inflation in almost all sub-periods (apart from 1851-1869, when it is 7.1%). For 1990-2014, less than 50% of the 4-year inflation FEVD is explained by inflation itself, with 39.8% explained by the short-rate and 12.9% explained in growth. Thus, it appears that short rate shocks are an increasingly important determinant of inflation during this period and that monetary policymakers nowadays may be able to use the short rate to influence inflation. In contrast, prior to 1915 the short-rate had a very modest ability to explain variation in inflation.

[INSERT: TABLE 6]

The FEVD empirical results for the short rate at the 4-year horizon show similarities with those for growth and inflation. Less than 100% of the error variance is attributable to short rate shocks itself for the full sample and for almost all sub-periods. The point estimate is 79.4% for the full sample but shows considerable variation during sub-samples being less than 50% for the 1823-1850, 1915-1916 and 1990-2014 sub-samples. We can also now reject for the full sample that 0% of the error variance is attributable to inflation or to growth. Growth now has a statistically significant impact on the short rate in most sub-samples (especially 1823-1850 and post-1976) and the full sample. Inflation can explains more than 10% of the 4-year forecast error of the short rate during all sub-periods except 1917-1951 and 1952-1975. This seems to suggest in general that the short rate in the medium term is responsive to both growth rate shocks than inflation shocks. Even during the 1990-2014 sub-sample, when the central bank is supposed to have been following an inflation targeting (or money growth targeting) policy, it seems that the short rate has been responding almost equally to both short rate shocks and inflation shocks.

The final columns of Table 6 report the volatility spillovers from the other economic variables. For example, for growth at the 1-year horizon (h=1) only 1.1% of the forecast error variance is attributable to spillovers from other variables only if one ignores the breaks (inflation or the short rate). However, for the 1990-2014 sub-period it is as high as 29.5% demonstrating much larger spillovers in recent years. At the 1-year horizon a similar pattern is also apparent for inflation and short rate; in general about 10% of error variance is attributable to spillovers but this rises to about 50% for 1990-2014.

The "To others" rows of Table 6 give the volatility spillovers from the variable of interest to other variables. There is further evidence that at the 1-year horizon that spillovers greatly increased for the post-1990 sample. Firstly, spillovers to other variables from growth are less than 2.5% in all periods except 1990-2014 where they are more than 20%. Secondly, spillover to other variables from inflation or the short rate are about 10% in all periods except 1990-2014 where they are at least 45%. The "Net" rows give the difference between the "To others" and "From others" segments. In general these values are very close to each other for the 1-year horizon forecast errors, except for the 1990-2014 sub-period. During the latest sub-period it seems as if there are substantially higher spillovers (almost 9%) given to other variables by the short rate than those received from other variables. This is primarily offset by growth giving substantially lower spillovers to other variables than it received.

The "Bi-variate" rows indicate that there is virtually no volatility spillover at 1 year horizon to the growth from either the short rate ore the inflation up until 1990 when the spillover becomes significant. The net bi-variate volatility spillover is from the short rate to the inflation and again this becomes significant only since 1990. These changes are in contrast to the constant parameter model inferences reported in square brackets.

The total spillover index is 5.73% if one ignores all the VAR coefficient and covariance matrix breaks but hugely rises to 42.33% for 1990-2014; this indicates that there were generally very modest spillovers for the full sample but very substantial spillover over recent years. Thus these economic variables are much more inter-connected nowadays and policy-makers should take this into account when implementing macroeconomic management strategies.

The FEVD results at the four year horizon also indicate that there are volatility spillovers which are strong in the post-1990 relationships. However, the main difference between the one year horizon and four year horizon is that in some of the earlier samples there are also substantial spillovers observed at the four year horizon. In the case of the 1976-1989, at the four-year horizon the spillovers are consistently quite high regardless of the variable examined. However, for other samples it tends to be the case that the high spillovers are confined to specific variables; for example in the "from others" section during the 1851-1869 there are relatively large spillovers towards growth but more modest spillovers towards inflation or interest rate during this sub-period. Hence these results suggest that pre-1990 spillover effects depend upon the variables and the sub-periods examined rather than being consistently large as they are in the post-1990 period.

6 Conclusions

The main objective of this paper is to model the dynamic behavior of the short-term interest rate, inflation and economic growth over the period 1820-2014 for the United Kingdom. Initially we examine if there are structural breaks (of persistence, mean or volatility) in individual series using the iterative procedure that was recently developed by Bataa, Osborn, Sensier and van Dijk (2014). We then attempt to identify, in a multivariate context, if there has been any structural breaks in the relationship between these variables using a vector autoregression (VAR) model.

Our primary empirical analysis employs the generalized VAR methodology of Koop, Pesaran and Potter (1996) and Pesaran and Shin (1998), the forecast error variance decomposition analysis of which is updated in Diebold and Yulmaz (2009, 2012, 2014). This is because our data frequency is annual and covers an extended time span; it is important to allow for correlated shocks in this context. We rely on generalized versions of the impulse responses and the forecast error variance decompositions that are not sensitive to variable ordering.

For our multivariate analysis we use the recently developed methodology of Bataa, Osborn, Sensier and van Dijk (2013). The main innovation of their methodology is that it allows for changes in both the VAR parameters as well as the variance covariance matrix, but not necessarily at the same time, a feature that is vital in our long term view. Then we examine if each structural break is the result of a change in persistence, a break in Granger causality, a shift in volatility or a change in correlation. Indeed, we find that one or a combination of these sources are behind the structural breaks in a trivariate VAR of short term interest rates, inflation and growth.

Most covariance matrix breaks can be primarily attributed to changes in volatility, although these mainly occur prior to the end of WWI. We find that the persistence of all three variables increases in 1917. There is also very strong evidence that the dynamic relationship between the inflation and the short rate breaks down during the Great War; there is some evidence Granger causality form the growth rate to the short rate also changed during this period. Our results confirm earlier research that a great moderation occurred but suggests that the structural change is apparent by 1952 in UK annual data. We do find a covariance matrix break in 1989 but this can be attributed to a break in correlation rather than a break in volatility. The finding that there has been no significant structural break in the correlation between the short term interest rate, inflation and growth until 1989 is indeed one of the most interesting.

A key finding from our analysis is that the dynamic behaviour of our variables of interest have been distinctly different in recent decades compared to earlier sub-samples. We find that the short rate responds strongly to both growth post-1975 and to inflation since at least 1990. However, short rate shocks do not have significant negative impacts either upon the growth rate or inflation. The implication is that the effectiveness and role of the short term interest as a tool for policymakers may need to be reconsidered. This is because neither inflation nor growth appear to react significantly to interest rate shocks (at least not in the appropriate direction). In fact we actually provide additional evidence of the price puzzle, a positive relationship between short rate and inflation; this is not confined to post WWII but in our longer sample is apparent since 1917.

Overall there is evidence of substantial structural change over time in the relationship between growth, inflation and the short rate. The primary findings are i) there is an increase in persistence of all variables in 1917, ii) most covariance matrix breaks are attributable to variance shifts and iii) when a correlation break did occur this lead to a substantial intensification of the responsiveness of the short rate to growth and inflation, which is observed for recent decades. Consequently economic policymakers need to carefully monitor the linkages between these variables and be prepared to adjust the way monetary tools are used when faced with structural change.

TABLE 1. UNIT ROOT TEST RESULTS WITH STRUCTURAL BREAKS

	Short term rate	*Log real GDP*	*Log RPI*	Δ *Short term rate*	Δ *Log real GDP*	Δ *Log RPI*
A. Tests for structural breaks in trend						
0 v 1 break	6.11**	28.63**	16.87**	2.43	1.06	4.29**
1 v 2 breaks	12.55**	13.78**	28.97**			9.20**
2 v 3 breaks	4.30**	4.44**	3.49			3.94**
B. GLS-detrended unit root test allowing for trend breaks						
No break	14.77 [5.54]	15.62 [5.54]	62.96 [5.54]	**0.95**** [5.54]	**1.01**** [5.54]	3.40** [5.54]
1 break	16.27 [5.41]	16.86 [6.76]	19.52 [6.67]	1.66** [7.48]	3.19** [6.73]	6.19** [6.73]
2 breaks	6.28** [8.08]	7.87 [7.17]	**9.37** [7.08]	1.96** [6.69]	4.48** [7.78]	8.10 [7.65]
3 breaks	**32.44** [8.40]	**9.80** [7.28]	12.39 [7.39]	2.69** [7.38]	6.17** [8.22]	**9.15** [8.13]

Notes: The test in the first row of panel A is the exponential Wald statistic discussed by Perron and Yabu (2009) for one break in both slope and level, valid for both I(0) and I(1) error components, while those in the rows below are the suprema of that statistic when applied to the subsamples defined under the corresponding null, the critical values for such sequentially applied statistics being provided in Kejriwal and Perron (2010). 5% critical values are 3.12, 3.66 and 3.71, respectively. Statistics in panel B are GLS-detrended M unit root test statistics, analysed by Ng and Perron (2001) for no trend breaks and extended by Carrion-i-Silvestre et al. (2009) to multiple breaks; bold statistics correspond to the number of breaks selected through the Kejriwal and Perron (2010) procedure. Since the critical values are case-dependent they are provided underneath each test in square brackets. *Indicates the rejection of the relevant null hypothesis (namely, the lower number of indicated breaks in panel a and presence of a unit root in panel b) at the 5% level. Rejection in panel A (panel B) applies when the statistic is greater (less) than the critical value. All critical values are asymptotic and from the indicated studies.

TABLE 2. UNIVARIATE ITERATIVE DECOMPOSITION

	Short rate	**Inflation**	**Growth**
A. Outliers			
Dates	none	1921, 1922	none
B. Level component			
Overall test	715.82**	115.37**	4.57
Seq(2/1)	390.16**	2.47	
Seq(3/2)	6.07		
Breaks	**1931** [1929-1933]	**1914** [1910-1918]	none
	1951 [1950-1952]		
Regime means	3.28, 0.54, 6.01	-0.06, 5.10	2.01
(Ignoring breaks)	(3.88)	(2.60)	(2.01)
C. Autoregressive coefficients			
Overall test	21.34** [15.92]	10.50 [10.67]	8.82 [10.67]
Seq(2/1)	4.55 [16.58]		
Breaks	**1940** [1926-1954]	none	none
AIC, HQ, SIC	5,5,1	1,1,1	1,1,1
Regime persistence	0.54, 0.92	0.67	0.25
(Ignoring breaks)	(0.87)	(0.67)	(0.25)
D. Volatility			
Overall test	36.92**	24.95**	21.82**
Seq(2/1)	12.07**	6.89	17.52**
Seq(3/2)	10.58		10.49
Breaks	**1931** [1894-1932]	**1980** [1953-1983]	**1918** [1915-1962]
	1951 [1950-1986]		**1946** [1908-1947]
St.Dev. across regimes	1.05, 0.24,1.56	3.55, 1.67	2.56, 4.60, 1.95
[Ignoring breaks]	[1.19]	[3.31]	[2.78]
E. Number of iterations for convergence			
Main and sub loop	4,3	6,2	3,2

Notes: Decomposition as in (1) using the iterative method of Bataa, Osborn, Seniser and van Dijk (2014), with breaks detected using the Qu and Perron (2007) test using a 5% significance level (trimming 10% and maximum of 8 breaks). 95% critical values are reported in square brackets for the autoregressive component case, as the lag orders are different. For the level and volatility these are uniform, thus not reported. These uniform 95% critical values of WDMax, Seq(2/1), Seq(3/2), Seq(4/3) are 10.67, 10.97, 11.88 and 12.49, respectively. Outliers are defined as being 5 times the interquartile range from the median after removal of level breaks. The AR order of the dynamic component is selected according to the Akaike (AIC) information criterion, and used at entry to the dynamic/volatility sub-loop. Persistence is computed as the sum of the estimated AR coefficients within the relevant dynamic regime, with the standard deviation of disturbances also shown within volatility regimes. When breaks are detected, the estimated breaks dates are shown **in bold** together with their asymptotic 90% confidence intervals in parentheses. Finally, the numbers required to achieve convergence of the loops of the algorithm are shown.

TABLE 3. ITERATIVE STRUCTURAL BREAK TEST RESULTS

VAR Coefficients	**Covariance matrix**
Asymptotic WDmax test statistics [and critical values]	
232.86** [42.63]	77.97** [22.59]
Asymptotic sequential test statistics [and critical values]	
101.15** [41.93]	36.57** [23.23]
19.73 [43.72]	29.26** [24.15]
	32.84** [24.77]
	25.98** [25.26]
	13.54 [25.70]
Break dates, confidence intervals (and bootstrap p-values)	
	1850 [1849-1851]
	(0.00)
	1869 [1866-1870]
	(0.01)
1916 [1909-1921]	**1914** [1913-1929]
(0.04)	(0.05)
	1951 [1949-1952]
	(0.00)
1975 [1971-1978]	**1989** [1972-1990]
(1.07)	(1.88)

Notes: Values are reported are at convergence of the iterative procedure of Bataa, Osborn, Seniser and van Dijk (2013). The overall test (WDMax) examines the null hypothesis of no break against an unknown number of breaks, to a maximum number of breaks indicated in the first row. If the overall statistic is significant at 5%, sequential tests (SeqF) are applied starting with the null hypothesis of one break and continuing until the relevant statistic is not significant. Asymptotic critical values for the 5% significance level are reported [in square brackets] and ** indicates the statistic is significant at this level. The estimated break dates and their 90% confidence intervals [in square brackets] are also reported. Bootstrap p-values corresponding to the null hypothesis that a break that is asymptotically significant is actually not are also reported in brackets.

TABLE 4. INDIVIDUAL COEFFICIENT BREAKS AND GRANGER CAUSALITY

		A. Coefficient Break Tests Dependent variable			**B. Estimated Coefficients** Dependent variable		
Subsample	Explanatory	Growth	Inflation	Short rate	Growth	Inflation	Short rate
1823-1916	Growth				**-0.15**	**0.66****	**0.20****
					31.02	0.03	0.10
1917-1975		**0.61***	**-0.49****	**-0.12**♠	**0.39****	**0.19**	**0.08**
		8.62	2.58	13.97	0.79	84.26	10.26
1976-2014		**0.06**	**-0.06**	**0.55****			**0.55****
		32.69	67.37	1.97			1.01
1823-1916	Inflation				**-0.07**	**0.34**	**0.08**
					89.30	14.12	10.88
1917-1975		**-0.06**	**0.43***	**-0.10****		**0.76****	**-0.02***
		60.11	8.41	0.46		0.00	9.30
1976-2014		**0.14**	**-0.04**	**0.15**♠			**0.12**
		50.72	58.07	14.52			44.71
1823-1916	Short rate				**-0.24****	**0.07**	**0.50**
					3.85	22.12	19.96
1917-1975		**0.15**	**0.21****	**0.31****		**0.13****	**0.81****
		21.37	1.60	3.34		2.58	0.00
1976-2014		**0.31**	**-0.34**	**-0.01**			
		16.70	36.63	52.64			

Notes: Columns represent equations. The first value (in bold) of each cell in panel A reports the difference between the sum of relevant coefficients after and before the break date, with this placed against the dates of the second subsample used in the comparison. The first value in each cell in panel B reports the estimated coefficients over the indicated subsample. In both cases, the values below are bootstrap p-values (expressed as percentages) for the null hypothesis that the coefficients are equal to 0. If an individual coefficient break is not significant at 15% in panel A, the corresponding subsample coefficients are restricted to be equal in panel B. Subsamples are those implied by the estimated structural break dates of Table 3. Significant at **5%, *10%, ♠15%, using wild bootstrap p-value.

TABLE 5. ORIGINS OF COVARIANCE MATRIX BREAKS & CAUSALITY

	A. Significance of Breaks		B. Subsample Residual Standard Deviations			C. Subsample Contemporaneous Correlations			D. No instantaneous Causality tests		
	Volatility	Correlation	Growth	Inflation	Short rate	Gr. & Inf.	Gr. & Sh.r.	Inf. & Sh.r.	Growth	Inflation	Short rate
1823-1850			3.19	2.88	0.66	15.9	28.2	4.8	13.8	0.0**	0.1**
1851-1869	0.05**	36.42	1.35	3.29	1.58	-	-	-	-	-	-
1870-1914	0.00**	50.87	2.22	1.88	0.67	-	-	-	-	-	-
1915-1951	0.05**	9.41*	3.94	4.28	0.74	-	-	-	-	-	-
1952-1989	0.00**	14.55	1.66	2.21	1.43	-	-	-	-	-	-
1990-2014	75.11	0.38**	-	-	-	36.3	88.7	48.1	4.8**	0.0**	0.0**
[Ignoring breaks]			[2.57]	[2.85]	[1.10]	[18.6]	[34.3]	[9.4]	[4.0**]	[0.0**]	[0.0**]

Notes: Panel A reports the significance of structural break tests for the diagonal elements of the covariance matrix of the VAR (Volatility) and for the off-diagonal elements of the correlation matrix (Correlation), showing bootstrap p-values (expressed as percentages) for the test of no change over adjacent covariance matrix subsamples identified in table 3, with the result placed against the dates of the later subsample. The values reported are the final ones computed in the respective general to specific procedures (see Bataa *et al*, 2013). The corresponding subsample residual standard deviations are reported in panel B and subsample contemporaneous residual correlations in panel C. The standard deviations and correlations are computed after merging subsamples based on the respective break test results in panel A (using 5% significance). The final panel reports the bootstrap p-value for a test of the joint hypothesis test that all contemporaneous correlations relating to that series are 0. All results are obtained from a VAR in which the restrictions implied by the results of coefficient breaks and persistence or dynamic interaction tests (at 15% significance) are imposed. **(*) indicates significance at 5% (10%).

TABLE 6. GENERALIZED FEVD AND SPILLOVERS

Series	Regimes	Growth		Inflation		Short rate		**From others**	
		h=1	h=4	h=1	h=4	h=1	h=4	h=1	h=4
Growth	1823-1850	97.5**	94.2**	2.1	3.8*	0.4	2.1*	2.5	5.8*
	1851-1869	97.5**	61.4**	2.1	11.1*	0.4	27.4**	2.5	38.6**
	1870-1914	97.5**	92.6**	2.1	3.9*	0.4	3.5*	2.5	7.4*
	1915-1916	97.5**	93.8**	2.1	4.2*	0.4	2.0*	2.5	6.2*
	1917-1951	97.5**	94.6**	2.1	2.2	0.4	3.2*	2.5	5.4*
	1952-1975	97.5**	55.1**	2.1	5.5*	0.4	39.4**	2.5	44.9**
	1976-1989	97.5**	77.6**	2.1	7.2*	0.4	15.2*	2.5	22.4**
	1990-2014	70.5**	58.2**	9.7*	17.4**	19.8**	24.4**	29.5**	41.8**
	[Ignoring breaks]	[98.9**]	[90.6**]	[1.0]	[4.0*]	[0.1]	[5.4*]	[1.1]	[9.4*]
Inflation	1823-1850	1.9	27.5**	89.8**	67.4**	8.3**	5.1*	10.2**	32.6**
	1851-1869	1.9	7.5*	89.8**	84.6**	8.3**	7.9*	10.2**	15.4**
	1870-1914	1.9	29.6**	89.8**	65.4**	8.3**	4.9*	10.2**	34.6**
	1915-1916	1.9	22.3**	89.8**	72.1**	8.3**	5.6**	10.2**	27.9**
	1917-1951	1.9	9.0*	89.8**	72.6**	8.3**	18.4**	10.2**	27.4**
	1952-1975	1.9	3.8	89.8**	47.4**	8.3**	48.7**	10.2**	52.6**
	1976-1989	1.9	9.0*	89.8**	76.2**	8.3**	14.8**	10.2**	23.8**
	1990-2014	7.3*	12.9*	53.1**	47.3**	39.6**	39.8**	46.9**	52.7**
	[Ignoring breaks]	[0.9]	[7.7*]	[91.6**]	[82.6**]	[7.6**]	[9.7**]	[8.4**]	[17.4**]
Short rate	1823-1850	0.4	37.9**	8.4**	18.8**	91.2**	43.3**	8.8**	56.7**
	1851-1869	0.4	3.9*	8.4**	14.6**	91.2**	81.5**	8.8**	18.5**
	1870-1914	0.4	25.7**	8.4**	15.8**	91.2**	58.6**	8.8**	41.4**
	1915-1916	0.4	38.6**	8.4**	21.9**	91.2**	39.5**	8.8**	60.5**
	1917-1951	0.4	15.0*	8.4**	7.9	91.2**	77.1**	8.8**	22.9*
	1952-1975	0.4	1.4	8.4**	5.1*	91.2**	93.5**	8.8**	6.5*
	1976-1989	0.4	28.2**	8.4**	16.6**	91.2**	55.2**	8.8**	44.8**
	1990-2014	13.8*	28.1**	36.8**	30.9**	49.4**	41.1**	50.6**	58.9**
	[Ignoring breaks]	[0.1]	[11.6**]	[7.6**]	[9.0**]	[92.3**]	[79.4**]	[7.7**]	[20.6**]
To others	1823-1850	2.3	65.4**	10.5**	22.6**	8.7**	7.1*		
	1851-1869	2.3	11.4*	10.5**	25.7**	8.7**	35.3**		
	1870-1914	2.3	55.3**	10.5**	19.6**	8.7**	8.4*		
	1915-1916	2.3	60.9**	10.5**	26.1**	8.7**	7.6**		
	1917-1951	2.3	24.0*	10.5**	10.1*	8.7**	21.6**		
	1952-1975	2.3	5.2	10.5**	10.7*	8.7**	88.2**		
	1976-1989	2.3	37.2**	10.5**	23.9**	8.7**	30.0**		
	1990-2014	21.2*	40.9**	46.5**	48.2**	59.3**	64.3**		
	[Ignoring breaks]	[1.0]	[19.3**]	[8.6**]	[13.0**]	[7.6**]	[15.0**]		
Net	1823-1850	-0.2	59.6**	0.3	-10.0	-0.1	-49.6**		
	1851-1869	-0.2	-27.1*	0.3	10.3	-0.1	16.8		
	1870-1914	-0.2	48.0**	0.3	-15.0*	-0.1	-33.0**		
	1915-1916	-0.2	54.7**	0.3	-1.8	-0.1	-52.9**		
	1917-1951	-0.2	18.6*	0.3	-17.3*	-0.1	-1.3		
	1952-1975	-0.2	-39.7**	0.3	-41.9**	-0.1	81.6**		
	1976-1989	-0.2	14.8	0.3	0.1	-0.1	-14.9		
	1990-2014	-8.3**	-0.9	-0.4	-4.5	8.7**	5.3		
	[Ignoring breaks]	[-0.1]	[10.0]	[0.1]	[-4.4]	[-0.1]	[-5.5]		
Bi-variate		Growth & Inflation		Growth & Short rate		Inflation & Short rate		**Index**	
	1823-1850	0.2	-23.7**	0.0	-35.8**	-0.1	-13.8*	7.2**	31.7**
	1851-1869	0.2	3.6	0.0	23.5*	-0.1	-6.7	7.2**	24.2**
	1870-1914	0.2	-25.8**	0.0	-22.2**	-0.1	-10.8*	7.2**	27.8**
	1915-1916	0.2	-18.1*	0.0	-36.6**	-0.1	-16.3*	7.2**	31.5**
	1917-1951	0.2	-6.8	0.0	-11.8	-0.1	10.5	7.2**	18.6**
	1952-1975	0.2	1.7	0.0	38.0**	-0.1	43.6**	7.2**	34.7**
	1976-1989	0.2	-1.8	0.0	-13.0	-0.1	-1.9	7.2**	30.3**
	1990-2014	2.4**	4.5	5.9**	-3.6	2.8*	9.0*	42.3**	51.1**
	[Ignoring breaks]	[0.1]	[-3.7]	[0.0]	[-6.2]	[-0.1]	[0.7]	[5.7**]	[15.8**]

Notes: Connectedness table for two horizons: h =1,2. The ij^{th} entry of the upper-left 3×3 series submatrix gives the ij^{th} pairwise directional connectedness; i.e., the percent of forecast error variance of series i due to shocks from series j. They are allowed to be different over the sub-samples defined by the statistically significant structural breaks. If bootstrap one (two) standard deviation(s) wide confidence interval does not contain zero they are marked with *(**). The quantities estimated over the whole sample, ignoring the structural breaks are in square brackets. The rightmost (From others) column gives total directional connectedness (from); i.e., row sums (from all others to i). The bottom (To others) row gives total directional connectedness (to); i.e., column sums (to all others from j). Next (Net) row gives the difference in total directional connectedness (to-from). The bottom-right numbers are total connectedness (mean 'from' connectedness, or equivalently, mean 'to' connectedness). The bottommost row also gives bi-variate spillovers.

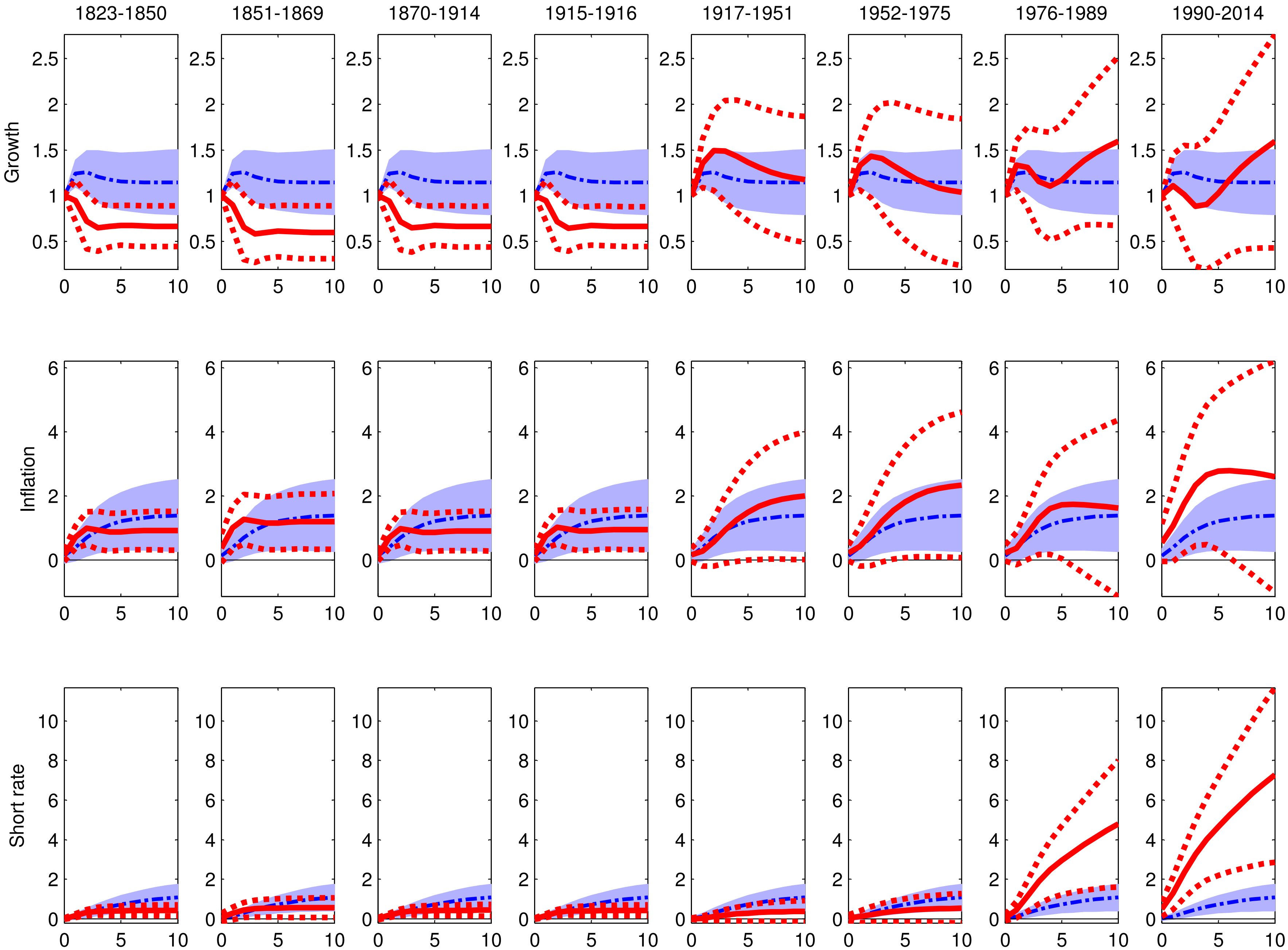

Figure 1: Notes: Generalized impulse responses of three variables over ten years after one percentage point growth shock. Rows indicate the variables that respond to the shock. Blue dotted and dashed lines with shaded confidence intervals assume parameter constancy. A response in solid red line with dotted confidence interval is specific to the sub-regime specified on the top. For each plot two-standard-deviation wide confidence interval is obtained using a bootstrap procedure explained in the text.

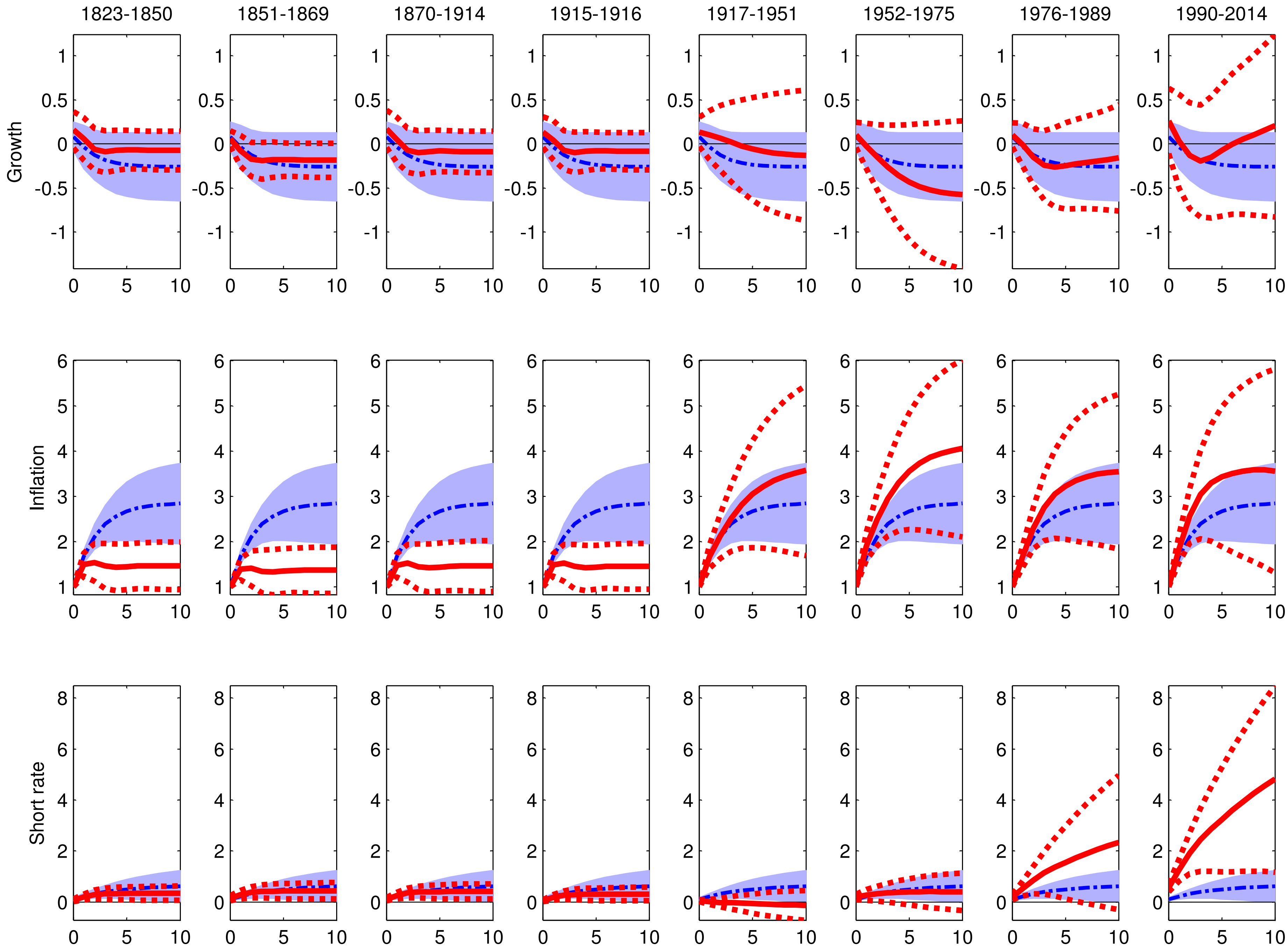

Figure 2: Notes: Generalized impulse responses of three variables over ten years after one percentage point inflation shock. See Notes to Figure 1.

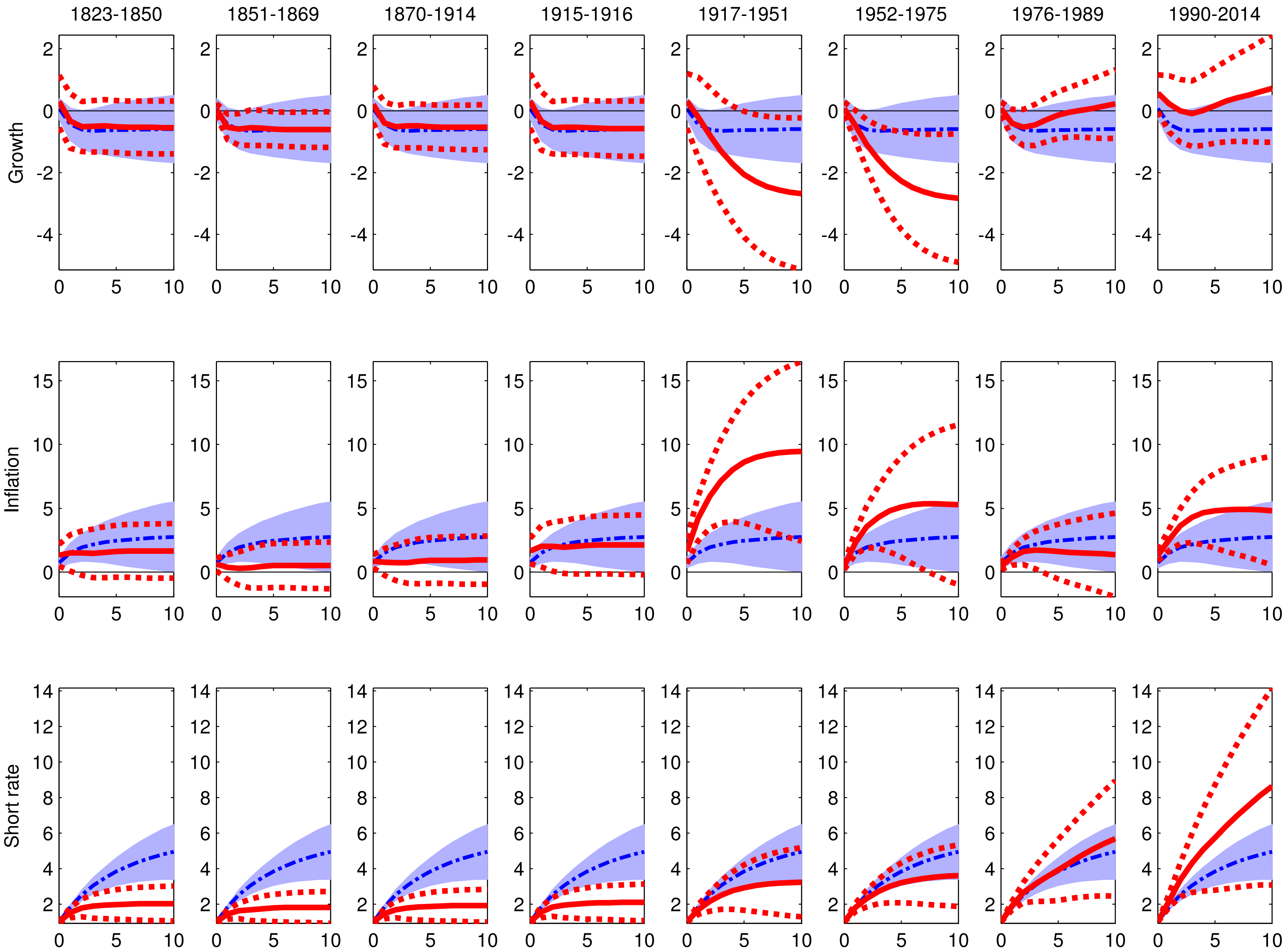

Figure 3: Notes: Generalized impulse responses of three variables over ten years after one percentage point short rate shock. See Notes to Figure 1.

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