On the long-run real implications of short-term interest rates *

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Abstract

Models of business cycles with endogenous growth can generate permanent effects of short-run shocks. This leads to the theoretical possibility that monetary policy indicators, such as interest rates, and output share long-run relationships. We show that this can be the case in a simple endogenous growth model with cash in advance constraints for both households and firms. We then test this proposition using historical data for the UK and US based on an information value approach. Our findings show that both statistics and economic theory-consistent evidence largely indicate the absence of long run relationships between the real output and the most relevant monetary indicator, short term interest rates. These findings are not only a full sample result, but also valid in most of the sub-samples throughout the second half of the 20th century and are robust to the inclusion of possibly omitted real variables and nonlinear functional forms.

Keywords: business cycles, endogenous growth, information value, long term relationship, cointegration, bounds tests.

JEL Codes: E3, E4, E5.

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1 Introduction

The relationship between interest rates and real output and inflation has always been at the core of Monetary Economics research. This is because interest rates may be affected by the monetary authority and hence, changes in monetary policy can affect economic activity if the economy is characterized by some market imperfections. Also, interest rates may serve as forecast indicators as they can contain important information about the future evolution of the economy in the short run. This second aspect is important regardless of whether we consider interest rates monetary policy indicators or simply reflecting equilibrium outcomes of market conditions in the economy[1].

To the best of our knowledge, however, there is no evidence available as to whether interest rates have any information content about output in the long-run. The question is significant for two reasons. First, if we consider the short run interest rate a monetary policy indicator variable, then the existence of a long-run relationship between interest rates and output would imply that the monetary policy authority would have to consider this long-run impact on their policy reaction functions. Secondly, we may consider that interest rates simply reflect market equilibrium conditions and are not controllable by the monetary policy authority. In this case, if the short run interest rate has information content in the long-run, this information should be taken into account in short run forecasting exercises. The existence of model uncertainty about the true nature of the role played by interest rates in macroeconomic models prevents us from knowing which one of these two arguments is practically more important. Either way, however, it is relevant to know if such a long-run statistical relationship exists, especially given the apparent loss of information of other monetary variables such as monetary aggregates since the early 1980's[2]. In this paper we analyze long-term relationships between the short-run interest rate and output. We first present a model of business cycles with endogenous growth that is able to generate, under some cir-

[1] See for instance Walsh (2003) and references therein.

[2] For the loss of information content in monetary aggregates, see for instance Estrella and Mishkin (1997), among others. For a recent counter argument see Aksoy and Piskorski (2006).

cumstances, long-run equilibrium relationships between interest rates and output. We then analyze the statistical evidence regarding this relationship for the U.S. and the U.K for the second half of the 20th century from a pure information content approach as advocated by Sims (1972, 1980)[3].

Standard New Keynesian models argue that interest rates can have business cycle stabilizing properties. This class of models assumes that monetary policy will not have permanent effects on output. However, a large body of theoretical papers has been devoted to analyzing whether cycles and stabilization policies can have a long-run impact on output. These models are built along two different lines. Firstly, stabilization policy can have an impact on output in the long run via learning by doing or via R&D in endogenous growth models. In the case of Stadler (1990), Pelloni (1997), and Blackburn and Pelloni (2005), temporary aggregate demand shocks can have permanent effects on output through a learning-by-doing mechanism. In the models of Fatás (2000, 2001, 2002) and Stiglitz (1993) temporary real shocks can also have a permanent impact on output because they will affect the amount of resources devoted to R&D expenditures. These effects were already noticed in King, Plosser and Rebelo (1988). The models of Caballero and Hammour (1994) and Hall (1991), for instance, postulate that recessions lead to permanent positive effects on productivity because they lead to the destruction of the less productive firms, because during recessions the returns to R&D are higher, or because productive investment is cheaper during recessions. Secondly, the models of Gomme (1993) and Dotsey and Sarte (2000) analyze the impact of inflation and inflation volatility on growth in a simple endogenous growth framework[4]. In Dotsey and Sarte (2000), inflation acts as a tax via the cash in advance constraints on consumption and investment. This model suggests that while real growth and money growth are positively related in the short run, on average money growth adversely affects economic growth. Clearly, these strands of the literature stand at odds with

[3] Much of the monetary literature addressing long run relationships focuses on monetary aggregates neutrality [see, among many others, Bernanke and Mihov (1998), Boschen and Mills (1995) and King and Watson (1997)].

[4] See Gillman and Kejak (2005) for an excellent review of theoretical models of inflation and growth.

the standard business cycle models. Within this framework, business cycles can be correlated with long-run growth. As a consequence, if short-term interest rates serve as either policy instruments or policy indicators that contain important information about output growth in the short-run, these could also have information content about output in the long-run. In that case, any rational policymakers would like to exploit this information as a guide for policymaking.

In this paper we proceed in two steps. We first develop a simple cash-in-advance model in the spirit of Fuerst (1992) and Lucas (1990) that generates a link between money supply shocks and short term interest rates and allows a potential break down of the uncomfortable Fisherian relation. We then provide the stylized growth path characteristics of this simple model in a rather flexible way. Secondly, we provide an empirical analysis about the long term relationship between short term interest rates and real output. To the best of our knowledge, we are not aware of any study that systematically analyses long-term statistical relationships between real output and nominal short-term interest rates[5]. Research by Bernanke and Mihov (1998) probably is the only exception to the literature in that it recognizes a causal role for nominal interest rates in the provision of liquidity into the economy and its implications in the long run. However, their study focuses on a structural model rather than long-run information content approach adopted here. It is important to stress that the information value approach, as a first test of statistical connection between certain variables, is immune to questions of causality, exogeneity or controllability of potential instruments. In other words, as long as long-term swings in the short run interest rate contain information about long term movements in income beyond what is already contained in movements in income itself, the forecaster or policy authority can potentially exploit this regardless of whether the information it contains reflects true causation, reverse causation based on anticipations, or mutual causation by some independent but unobserved influence. Therefore, issues raised by earlier work related

[5] Throughout the paper, we also conduct tests based on U.S. ex-ante real interest rates. Our results confirm close co-movement of these two rates, being therefore statistically indistinguishable.

to structural models and the Lucas critique are not of direct relevance[6].

However, since an assessment of the long term relationships depends crucially on the stochastic properties of the variables, we need to address carefully the issue of the order of integration of real output and interest rates. Although standard univariate analysis cannot reject the nonstationarity of most short-term real or nominal interest rate series, one cannot take this result at face value. Economic intuition suggests that short-term nominal interest rates should rather be stationary if the economy is not going through an episode of hyperinflation.[7] In order to address this uncomfortable statistical feature of short term interest rates we proceed in two steps. First, we take initial simple statistical evidence seriously and provide a series of cointegration tests based on univariate statistical properties of short term interest rates. Cointegration tests based on Johansen's maximum likelihood procedure impose minimal auxiliary assumptions to account for long term relationships. In the second step, we take economic intuition seriously and implement the Pesaran et al. (2001) bounds tests. These bounds tests for long run level relationships do not require non-stationarity of short-term interest rates and, therefore, are economic theory consistent.

Once we establish whether or not there is long-run information content about output in the interest rate series, we address the issue of whether these findings are the result of the omission of important real variables that explain the long-run evolution of output. Finally, we address if our findings are the result of using a simple linear specification by estimating some plausible nonlinear functional forms between nominal interest rates and output.

Our results show that both purely statistical-based and economic theory-consistent evidence largely rejects the existence of long term relationships between short-term interest rates and real output for the US and the UK. This absence of long run relationships is not only a full sample result, but also valid in most of the sub-samples in

[6]For an excellent discussion of the information content approach see for example Friedman and Kuttner (1992).

[7]For recent evidence on the debate of interest rate stationarity see, for instance, Rapach and Wohar (2005).

the post Second World War period and are robust to the inclusion of possible omitted real variables. We only find weak evidence of a possible nonlinear long-run relationship for the US case, but we consider it not to be robust.

The paper is organized as follows. In Section 2 we present a stylized model in which we show how a long-run relationship between interest rates and output can potentially arise. In Section 3 we present the data. Section 4 discusses the relevance of the short run interest rate in forecasting output in the short-run. Section 5 presents univariate time series properties of the variables before conducting long-term tests. In Section 6 we conduct long-term tests based on statistical evidence. We present cointegration results with a particular emphasis on sub-sample stability. In Section 7 we implement economic theory consistent bounds tests with particular emphasis on sub-sample stability. Section 8 addresses the omitted variables and nonlinearity problems. Section 9 concludes.

2 A Simple Cash in Advance Model

There are several potential channels that can theoretically generate a long-run relationship between interest rates and output. All the models in this literature assume that the production technology leads to permanent growth in the long-run as is standard in endogenous growth models. The challenge however is that standard DSGE models with or without cash-in-advance constraints predict a proportional relationship between money supply shocks and nominal interest rates due to Fisherian fundamentals. As the empirical evidence is mainly in favour of a negative relationship between the two, Lucas (1990) suggested an elegant solution by subtly altering the informational structure and by incorporating a financial intermediary in an otherwise standard real business cycle model. Fuerst (1992) further explores, in a general equilibrium setting, conditions where the value of cash in financial markets and goods markets may differ. This setting is able to break the Fisher relation. Therefore, we will present a version of the model in the spirit of Lucas (1990) and Fuerst (1992)[8]. As the production process uses a simple Ak technology, any shock to the equilibrium generates permanent effects on consumption, capital accumulation and output along the balanced growth path.

We start with a slightly modified version of the model presented in Fuerst (1992). The representative agent is a family which consists of a shopper, a financial intermediary and a firm and constitute the entire economy. Shoppers consume consumption goods that are produced by the firms. Firms consume investment goods produced by other firms in the economy. Using the capital stock, firms produce a contemporaneous output. Financial intermediaries manage the loanable funds that consist of previously deposited balances by the savers (consumers) and the newly injected money supply. Thereby, they manage loans between savers and borrowers and they channel cash injections of the monetary authority in to the economy. Note that the financial intermediary is a passive agent in this economy. It takes deposits from the shoppers and gives away loans to firms to maintain the flow of funds. We assume that it is costly for savers to rebalance their saving/consumption decision after each monetary injection. Therefore, the loanable funds supply will be determined by the current injection of the money supply, therefore only borrowers (firms) have direct access to the supply of loanable funds. In a world with uncertainty about the monetary policy actions and shoppers inability to revert their saving/consumption decisions, these unanticipated monetary policy actions will have real consequences via imbalances in the liquidity positions of the savers and borrowers. We will detail the mechanics in the following sections. The important assumption in the model is that both shoppers and firms are subject to cash-in-advance (CIA) constraints.

The timing of the Lucas-Fuerst model is as follows. First shoppers (consumers/savers) have to decide the amount of cash they need for their consumption purchases $(m - d)$ where m represents money inherited from previous period and d represents deposits be-

[8] It has to be noted, however, that there potentially alternative ways of generating this relationship. A simple endogenous growth model with learning-by-doing in the spirit of Stadler (1990) augmented with an interest rate rule is just an example.

ing made by the shoppers at the financial intermediary. Once the shopper irreversibly allocates her income between deposits and cash holdings for consumption purposes, the household separates. The state of the economy is revealed, that means the monetary authority decides whether to inject new, possibly unanticipated, money in to the economy. The shopper travels to the goods market, while the financial intermediary and the firm travel to the financial market. Consumers complete their shopping with their irreversible cash holdings. Next, firms decide on their new investment that should be made with money holdings. Firms finance new capital goods expenses (i) by borrowing (b) from the financial intermediary and therefore are liable to pay an interest (R) on their debt. An unanticipated monetary injection will leave the borrowers more liquid than anticipated, therefore purchasing power will be redistributed in their favor. Given that cash holding decisions of shoppers are irreversible, an excess monetary injection affects both asset prices and real economic activity as borrowers will be more liquid than expected. Borrowers (firms) will demand a decrease in nominal rates; while they increase their real investment demand. Likewise a monetary contraction has opposite implications. Once investment decisions are complete, the household reunites and all remaining cash is pooled, loan repayments are made, goods purchased are consumed. The crucial aspect of the model is the informational asymmetry on the timing of the monetary policy shocks between firms and shoppers.

2.1 Consumers

We assume that the infinitely living representative consumer has preferences over uncertain consumption streams given as:

$$\max E_t \sum_{t=0}^{\infty} \beta^t u(c_t) \tag{1}$$

We assume that the period utility function $u(c)$ is concave, twice continuously differentiable and increasing in c. Our earlier description of the timing of the events motivate two CIA constraints. First, the shopper is subject to CIA constraint in nominal terms such that:

$$p_t c_t \leqslant m_t - d_t \tag{2}$$

where p_t is the price level at time t. Second, the firm has to pay for investment expenditures in advance by borrowing from the financial intermediary. This leads to to the following CIA constraint in nominal terms on the firm's investment:

$$p_t\left[k_{t+1} - (1-\delta)k_t\right] \leqslant b_t \tag{3}$$

where investment is given as $i_t = [k_{t+1} - (1-\delta)k_t]$ and k_t represents capital stock in time t and δ the rate of depreciation of capital. Finally, the cash balances that the household ends the period with are given as:

$$m_{t+1} = m_t + \nu_t + d_t R_t + p_t y_{t+1} - p_t c_t - p_t\left[k_{t+1} - (1-\delta)k_t\right] - b_t R_t \tag{4}$$

where the m_{t+1} is the end of period money balances, ν_t represents the monetary transfers at the end of the period, $d_t R_t$ represents nominal interest income earned for household deposits at the financial intermediary, $p_t y_{t+1}$ represents the output valued at market prices, and $b_t R_t$ represents the firm's repayments on the loan taken form the financial intermediary.

2.2 Firms

Along with consumers, producers consist of numerous identical firms. Firms are risk neutral and they produce output using a simple Ak technology:

$$y_t = Ak_t \tag{5}$$

where the capital accumulation identity is given as:

$$k_{t+1} = (1-\delta)k_t + i_t.$$

2.3 Monetary Authority

We assume that the monetary authority transfers money using the following stochastic process:

$$m_{t+1} = (z + \eta_t)\, m_t \tag{6}$$

where η_t is an identically and independently distributed random variable with $\eta_t \sim (0, \sigma_\eta^2)$. The characterization of the monetary policy allows to make a distinction between the systematic (z) and discretionary (η) parts of the monetary policy conduct.

2.4 The Liquidity Effect

Representative consumer maximizes its lifetime utility (1) subject to (2), (3), (4) and the transversality condition. We denote λ_1 as the Lagrange multiplier associated with the cash in advance constraint for the shopper, λ_2 as the Lagrange multiplier associated with the cash in advance constraint for the firm and λ_3 the Lagrange multiplier associated with the budget constraint. Solving for the first order conditions yields an expression for the nominal interest rates as a function of the Fisher relation. After substituting f.o.c.'s and rearranging we obtain the following expression that characterizes the liquidity effect in equilibrium as discussed by Fuerst (1992):

$$1 + R_t = E_t \left(\frac{\Lambda_t + u'(c_t)/p_t}{\beta u'(c_{t+1})/p_{t+1}} \right) \tag{7}$$

Here, $\Lambda = \lambda_2 - \lambda_1$ represents the liquidity effect that breaks the Fisher relation. The rest of the expression gives the term related to expected inflation.(see Fuerst, 1992, for details). If $\Lambda = 0$, i.e. the value of cash in the goods market (λ_1) is equal the value of cash in financial markets, we obtain $1 + R_t = E_t \left(\frac{u'(c_t)}{\beta u'(c_{t+1})/\pi_{t+1}} \right)$ which is the standard consumption Euler equation that takes into account the expected inflation $(\pi_{t+1} = \frac{p_{t+1}}{p_t})$. In that case, nominal interest rates are simply a reflection of the real interest rates and the expected inflation. Therefore, the Fisher relation is valid. A monetary injection would lead to a proportional increase in the nominal interest rates. If $\Lambda \neq 0$, however, we obtain what Fuerst (1992) calls the liquidity effect. The premium arises when cash in one market is relatively more liquid. If $\Lambda < 0$ (i.e. $\lambda_{2t} < \lambda_{1t}$) financial markets are more liquid, therefore the asset price is higher than the value dictated by the Fisherian fundamentals and vice versa. In this case firms are more liquid than the shoppers. In an ideal world without liquidity constraints, households would like to transfer excess liquidity from the firms to the shoppers after an unexpected monetary expansion is realized. Given the constraint on the shopper and that immediate portfolio allocations towards the shopper are ruled out, financial intermediaries can sell this excess liquidity only to the firm which in turn uses it for investment purposes. Given that the firm is already at a high liquid position relative to the shopper ($\Lambda < 0$), the only way to make these funds attractive for firms is to lower the costs, i.e. R_t falls. On the other hand, if $\Lambda > 0$, the cash value in the financial market is higher than the cash value in the goods market, i.e. the firm is less liquid. Financial intermediaries can price excess liquidity generated by the monetary authority to the firm which are in search of liquidity to fund their investment needs at a higher level, i.e. R_t raises.

We can also show that the liquidity effect is present in the financial market and thereby linked to the consumption process. After some manipulations of the first order conditions we obtain:

$$E_t\left(\frac{u'(c_{t+1})}{\beta}\right) = \beta E_t\left(\frac{\left[A + (1 + R_{t+1})(1-\delta)\right] u'(c_{t+2})\pi_{t+1}}{\pi_{t+2}(1 + R_t)}\right) \tag{8}$$

i.e. the marginal rate substitution between current and future consumption should be equal to relative real interest changes. If nominal interest rates are determined by Fisherian fundamentals, (i.e. expected future inflation and real interest rate), the monetary injection will be reflected proportionally in the expected inflation and nominal interest rates, therefore the liquidity effect will be absent. Otherwise, the liquidity effect distortion will create a nonfundamental link between nominal interest rates and output, when the change in Λ is larger than the change in inflation expectations after monetary injections.

2.5 Growth Path

We are interested in the growth path implications of the liquidity effect induced short run equilibrium condition. Given the AK technology, and given the assumption that the change in Λ is larger than the change in inflation expectations after monetary injections, any shock will have a lasting impact on the balanced growth path when the liquidity constraints bind in equilibrium. For illustration purposes we assume that the utility of the representative household takes the following logarithmic form:

$$u(c_t) = \ln c_t$$

Therefore (7) can be reformulated in terms of the expected consumption growth rate of the economy $\gamma_c^e = E_t\left(\frac{c_{t+1}}{c_t}\right)$:

$$\gamma_c^e = \beta E_t\left(\frac{1+R_t}{\pi_{t+1}\left(p_t c_t \Lambda_t + 1\right)}\right) \tag{9}$$

Let us specify the inverse of inflation as a function of underlying monetary shocks as $\frac{1}{\pi_{t+1}} = \frac{1}{\left(z+\eta_{t+1}\right)}$. Then, condition (9) can be expressed as a function of the monetary injection:

$$\gamma_c^e = \beta E_t\left(\frac{1+R_t}{\left(z+\eta_{t+1}\right)\left(p_t c_t \Lambda_t + 1\right)}\right).$$

Given that $E_t(\eta_{t+1}) = 0$, we can rewrite $\gamma_c^e = \beta\frac{1+R_t}{z(p_t c_t \Lambda_t+1)}$, where expected average growth rate of consumption is related to the level of nominal interest rate, which in turn is determined by the liquidity effect and the Fisherian fundamentals. Note that if $\Lambda = 0$, this expression reduces to $\gamma_c^e = \beta\frac{1+R_t}{z}$ and the expected inflation will be fully reflected in the nominal interest rate. In this case, expected consumption growth rate will be only linked to parameters that determine the real interest rates. Nominal interest rates will be just a reflection of expected inflation in equilibrium.

Under the assumption that the shadow values of cash in the goods and financial markets are different in equilibrium, we obtain a straightforward link between nominal interest rates and the consumption growth. As argued earlier, when $\Lambda_t < 0$, implying more liquid financial markets after discretionary monetary injections, asset prices are higher than the value dictated by the Fisherian fundamentals in equilibrium. The only way to make these funds attractive for firms is to lower their costs, i.e. R_t falls thereby reducing the returns to savings. Similarly, when $\Lambda > 0$, less liquid financial markets after discretionary monetary contractions, financial intermediaries can increase the price of loans, i.e. R_t raises. Note also that with $z > 0$, the expected change in the consumption growth is negatively associated with the change in the liquidity parameter.

$$\frac{\partial \gamma_c^e}{\partial \Lambda_t} = -E_t \left(\frac{p_t c_t \beta \left(1 + R_t\right)}{z \left(c_t p_t \Lambda_t + 1\right)^2} \right) < 0 \tag{10}$$

As the gap between asset price values of the financial liquidity and consumer liquidity increases, firms reduce their investment demand due to higher costs of production finance. This in turn reduces the expected growth rate of the economy and expected consumption growth.

Finally, given the law of motion for capital:

$$\begin{aligned} k_{t+1} &= (1-\delta)k_t + i_t \\ &= (1-\delta)k_t + Ak_t - c_{t.} \end{aligned}$$

and substituting (9) and using the Ak technology we have:

$$k_t = \left[(1-\delta) + A - \beta \frac{1 + R_t}{z \left(p_t c_t \Lambda_t + 1\right)} \right]^{-1} c_t \tag{11}$$

$$y_t = A \left[(1-\delta) + A - \beta \frac{1 + R_t}{z \left(p_t c_t \Lambda_t + 1\right)} \right]^{-1} c_t \tag{12}$$

Equations (11), and (12) show that if liquidity constraints bind in equilibrium a simple endogenous growth model, by its very nature, is able to generate permanent effects on output via monetary policy changes. Liquidity effects impose different growth rates on investment and consumption. Any change in the growth rate of investment due to unexpected monetary shocks alters the production process, thereby implies permanent changes in output and in turn future consumption. The sign of long term output implications depends on whether the shopper or the firm is relatively more liquid in equilibrium. If there are no further unexpected monetary shocks economy grows along the balanced growth path. The empirical question we then need to address is whether or not monetary policy indicators (interest rates) contain information about output in the long run. In the following sections, we will conduct a battery of tests for long term statistical relationship between short term interest rates and output for sufficiently long time series data.

3 Data

We use data for the UK and the US. The annual data for the U.K. covers the period 1873-2001[9]. We will study real output represented by real GNP. This data was obtained from the study of Hendry (2001) [http://www.nuff.ox.ac.uk/users/hendry/]. This study stops in 1991 and hence, from this year onwards we update the data using OECD's Main Economic Indicators and IMF's International Financial Statistics database (IFS). We use the Treasury Bill rate as the short term interest rate measure and 10-years Government Bond yield as long term interest rate as reported by Hendry (2001).

In the case of the U.S. data on output and the Treasury Bill Rate is obtained from the U.S. Federal Reserve. Treasury Bill Rates have missing observations during the end of the 1930s and beginning of WWII, so we could only start in 1941. We also use two long-term interest rates such as the 10-year Government Bond Rate and Moody's AAA Yield Index starting from 1929[10].

As a cross check of our annual data results we also carried out our tests using

[9] Detailed data descriptions and source references are tabulated in the Appendix.

[10] The behaviour of the AAA Yield Index was very close to the one of the 10-year Bond and hence we do not report these results here.

quarterly data from 1960:1 to 2001:2. In this case we used as short term rates the Treasury Bill rate for both UK and US and also the Federal Funds Rate for the US. This quarterly data comes from IFS, OECD and the statistics provided by the U.S. Federal Reserve Board (FRB). We report the quarterly data results only when they yielded substantially different results from the annual data.

Insert Figure 1 about here

Figure 1 plots the annual data on the Treasury Bill rate and the log of GDP for the US and UK. The main feature that arises from both plots is the large and sustained increase in interest rates that reach a peak in the 1980-1982 period of disinflationary policies. In the case of the UK this pattern appears as more accentuated as we can observe practically constant interest rates in the 1873-1929 period. In contrast, output shows the typical upwards trend with few and isolated changes over time.

4 Interest Rates as Monetary Policy Indicators

A long run analysis of policy indicators that are not informative about short term business cycle fluctuations is not useful for our purposes. Before proceeding to the long run analysis we need the sample period that delivers significant and stable information content of short term interest rates to explain business cycle fluctuations in the U.K. and the U.S.

In order to determine the relevant sample size we proceed as follows. We first specify an autoregressive specification for real output changes in the spirit of Sims (1972) that is given by:

$$\Delta y_t = \alpha + \sum_{k=1}^{m} \beta_k \Delta y_{t-k} + \sum_{k=1}^{n} \delta_k \Delta i_{t-k} + \nu_t \tag{13}$$

where Δy and Δi are the growth rates of real output (annual log differences of real GNP) and the change in the short term interest rate (annual log differences of the T-Bill). We then run full sample as well as recursive Granger Causality tests for the policy indicator, short-term interest rates. Results are reported in Table 1 and Figure 2[11].

Insert Table 1 and Figure 2 about here

Our preferred annual data sample for the U.K is 1948-2001 and for the U.S. 1947-2001. For the U.K. there are several earlier episodes in which short-term interest rates contain useful information to explain business cycle fluctuations. However, in periods with major events such as First World War and Great Depression the information content of short-term interest rates vanishes making periods before 1948 redundant for the long-term analysis. In the case of U.S., short-term interest rates do not exhibit stable and significant information content before 1947 therefore we drop these data points from our sample relevant for the long-term analysis.

In Table 1 we present full sample $\chi - Square$ (and $p - values$) for the corresponding interest rate measures. Irrespective of the maturity all interest rate measures are significant for the full sample we choose. In Figure 2 we also present p-values of rolling regressions (with a 30 years window). Here we note that in most of the sub-samples U.K. and U.S. T-Bill rates contain significant information about short run output fluctuations.

Note that given high level of inflation persistence in the U.S. and the U.K., nominal interest rates track very well real interest rates and therefore stand as a reasonable proxy for even ex-ante real interest rates[12].

[11] We select lags based on AIC and SIC. Our preferred specification for the U.S. contains four lags for short-term interest rates and our preferred specification for the U.K. contains one lag for the short-term interest rates. To capture autoregressive dynamics for real output both U.K. and U.S. real output equations contain four lags. In recursive estimates minimum sample size is 30 years. The White test for heteroskedasticity rejected the non-constancy of the residual variance for almost all-financial variables in specification (13). Therefore, the White heteroskedasticity-consistent standard errors are used to derive the corresponding χ-square statistics of the Granger causality tests. Moreover, the relative performance of short term interest rates in terms of the heteroskedasticity consistent Granger causality statistics is very similar to those based on the statistics computed with unadjusted OLS residuals. Finally, we note that the Ljung-Box Q-statistics do not reject the null hypothesis that there is no autocorrelation in the residuals of the equation (13).

[12] We also repeat the same exercise with the use of U.S. ex- ante real interest rates instead of nominal interest rates. In constructing the ex-ante real interest rates based on inflation expectations,

5 Univariate Time Series Properties

We carry out four standard unit-root tests on the data. These are an ADF test of the null of non-stationarity; the KPSS variance ratio test of the null of stationarity; the Modified Phillips-Perron test with GLS de-trending ($M^{\alpha}GLS$) of Ng and Perron (2001) for the null of a unit root; and Elliott et al.'s (1997) most powerful DF-GLS test for the null of a unit root. The lag augmentation was chosen using the Ng and Perron (2001) Modified Information Criteria (MIC)[13]. This method reduces very substantially size distortions. The tests are carried out using a constant term and a constant and a deterministic trend. The results are reported in Tables 2 and 3. They reveal that most of the variables are non-stationary. We cannot reject the non-stationarity hypothesis for all the variables involved except for the US Treasury Bill rate when using quarterly data[14].

Insert Tables 2 and 3 about here

The behavior of the series may have also been characterized by the existence of structural breaks that will affect the power of the previous unit root tests. We hence tested for structural change in the series using the Bai and Perron (1998) technique and found that most interest rates show one structural change around 1981-82 for both countries. When applying unit root tests considering these breaks we found stationarity when we model the break as a trend break with both segments joined at the break time point, but not when using other specifications[15].

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we relied on Federal Reserve Bank of Philadelphia's survey of professional forecasters (for the period of 1970-2001). Our results indicate that the information role of real interest rates is very much in line with the short term nominal interest rates in the U.K. and the U.S.

[13]The results using other information methods such as AIC or a general to specific method (GTS) did not change the conclusions about unit-roots.

[14]For longer term maturities the evidence strongly supports non-stationarity.

[15]Results are available on request.

6 Long Term Relationship Tests: Taking Statistics Seriously

As mentioned earlier, possibly non-stationary interest rates are an uncomfortable result from a theoretical viewpoint, as interest rates have to be stationary for a dynamic general equilibrium to exist. Our results may also reveal the well-known power problems of unit-root tests and/or problems arising from structural breaks. This is a non-trivial problem as cointegration tests such as the Johansen's VAR method rely on the strong assumption that all endogenous variables to the system are strictly I(1). In order to deal with this problem we will proceed to analyze long-run relations between interest rates and output by using two approaches. In the first, we will assume that both variables are I(1) and apply traditional cointegration tests. That is, we rely on the statistical evidence on stationarity. In the second, we will use a bounds tests procedure that is independent of the stationarity results and allows us to be both theoretically and statistically consistent. This allows us to test the implications of the theory model in a way that is robust to the assumptions made about the integration order of the variables.

6.1 Cointegration

Johansen's method of estimating cointegrating vectors is a good starting point for tests of long run relationships. It needs minimal auxiliary assumptions to make tests workable. We do not make any assumptions about the nature of the shocks but rather focus on the long term relationship between the short term interest rates and real output[16]. We are interested in the long term marginal predictive content of short term interest rates in explaining the long term equilibrium output.

[16] Gonzalo (1994) compares ordinary least squares, nonlinear least squares, maximum likelihood in an error correction model, principal components and canonical correlations performance in estimating cointegrating vectors. Based on Monte Carlo simulations, he finds that the estimation of a fully specified error correction model by maximum likelihood as suggested by Johansen procedure performs better even when the errors are non-normal distributed or when the dynamics are unknown.

Insert Table 4 about here

We consider four cases about the deterministic trends present in the relation between output and the interest rate. Case I corresponds to no deterministic trend in the data, and an intercept but no trend in the cointegrating equation. Case II corresponds to a linear trend in the data and an intercept but not a trend in the cointegrating equation. Case III corresponds to a linear trend in the data and both an intercept and a trend in the cointegrating equation and finally Case IV corresponds to a quadratic trend in the data, and both an intercept and a trend in the cointegrating equation. Case I would imply that the first differenced variables share the same mean which, on inspection of Figure 1, is very unlikely as output appears to be heavily trended. Case IV would imply that the first difference of the variables have a deterministic trend. This is again unlikely and very evident in the case of the interest rate that shows no accelerating or decelerating growth over time (see Figure 1 for both countries). A priori, hence, the most likely cases to describe accurately any deterministic trend in the data are cases II and III.

With exception of Case I, full sample cointegration tests reported in Table 4 cannot reject, in general, the hypothesis of no cointegration for the short term interest rate measures for alternative specifications on the cointegrating equation[17].

6.2 Stability of Cointegration Relationships

To analyze the long run stability of the output and interest rate relationships we conduct several exercises based on recursive LR-values.

[17] It is well known that since it is very difficult to distinguish an I(d,d>.5) from an I(1) variable, Johansen LR tests often tend to find spurious cointegration relation even if there is none. Therefore, in our case a Johansen LR test of finding no cointegration should be interpreted as a rather conservative result. Note that we have also tested for long run interest rates. Only in the case of U.S. there is some evidence of cointegration between the long term interest rates and real output if the cointegrating equation can be characterized by an intercept but no trend. All other specifications favor no cointegration between long term interest rates and real output.

We graphically explore the stability of LR-values for at least thirty years long time intervals within which we expect that any monetary impact would disappear. For this purpose we present a series of LR-values of Johansen tests obtained from recursive estimations for real output and interest rates. Three types of recursive estimations are considered. In the first exercise, we implement a rolling sub-samples analysis where we allow for a 30-years window in the recursive estimations. In the second exercise, the beginning of the entire sample period is held fixed. In the third and final exercise endpoint of the entire sample period, 2001, is held fixed[18].

Insert Figures 3 and 4 around here

We first present LR-values for the 30-years window rolling sample (first row in Figures 3 and 4). For the U.K. (U.S.) the first LR-value corresponds to the 1948-1977 (1949-1978) estimation period and the last one to 1972-2001 estimation period. In the case of the U.K. there are several episodes for which the hypothesis of no cointegration can be rejected under alternative cointegrating equations. Particularly, periods corresponding to the loss of independent monetary policy during the participation in the ERM seem to be connected to a violation of no-cointegration relationship. For the U.S., results show we cannot reject the hypothesis of no cointegration in all rolling sub-samples considered[19].

Second, we report recursive cointegration results by fixing the starting point in the second rows of Figures 3 (U.K.) and 4 (U.S.). The first LR-value plotted in the figures displays the test statistics for the sample period 1948-1977, with the last value corresponding to the entire sample period 1948-2001. The two dashed lines correspond

[18] For the sake of comparison we also run cointegration tests for whole available sample period irrespective of whether the short term interest rates are useful policy indictors or not. In that case, in the first exercise (2001) is held fixed, while in the second one the beginning of the entire sample period (1873 for the UK variables, 1941 for the US variables) remains unchanged. Results are available upon request.

[19] We have repeated the same exercise for U.K. and U.S. medium to long term interest rates (Moody's AAA Corporate Bonds, and 10 years Bond yield for the U.S. and 10 years Bond yield for the U.K.) Results do not change substantially. Results for medium to long term interest rates are available upon request from authors.

to the 5% and 1% significance level. The results for the UK show that the hypothesis of no-cointegration in general cannot be rejected for short term interest rates (T-Bill) and output However, Case I indicates high instability in the corresponding LR-values and the null hypothesis is rejected. Similarly, for the U.S., with exception of Test I the hypothesis of no cointegration can not be rejected in general in nearly all subsamples. Some exceptions arise for the early 1980's when the U.S. monetary policymaking has changed drastically.

Finally the third rows of Figure 3 (U.K.) and Figure 4 (U.S.) display recursive LR-values when we fix the endpoint of the sample. The results show that in this case the hypothesis of no-cointegration cannot, in general, be rejected for both the U.K. and U.S. short term interest rate (T-Bill) and real output under the different specifications of the cointegration equation.

Overall, the various tests cannot reject the hypothesis of no cointegration in most sub-samples considered.

7 Bounds Tests: Taking Economic Theory Seriously

Power problems of unit-root tests and theory-based arguments cast doubts about the assumption made earlier that both output and the interest rate are $I(1)$ variables. Pesaran et al. (2001) develop a technique to test for the existence of a long-run relationship between two variables irrespective of whether they are $I(1)$ or $I(0)$. This methodology becomes most useful in our empirical tests where variables with different orders of integration may be involved. Their approach is based on the estimation of an unconstrained dynamic error correction representation for the variables involved and testing whether or not the lagged levels of the variables are significant. In other words, Pesaran et al.'s (2001) test consists of the estimation of the following conditional error correction model (ECM);

$$\Delta y_t = \alpha_0 + \beta_1 y_{t-1} + \beta_2 i_{t-1} + \sum_{k=1}^{m} \varphi_k \Delta y_{t-k} + \sum_{k=1}^{n} \theta_k \Delta i_{t-k} + \omega \Delta i_t + u_t \qquad (14)$$

In order to test for the existence of a long run relationship Pesaran et al. (2001) consider two alternatives. First, an F-statistic test of joint significance of the lagged levels of the variables involved[20]. Second, a $t-$ratio test for the significance of the lagged level of the dependent variable (y_{t-1}). Pesaran et al. provide two sets of critical values assuming that both regressors are $I(1)$ and that both are $I(0)$. These two sets provide a band covering all possible combinations of the regressors into $I(0), I(1)$ or mutually cointegrated[21]. Also, if the F-statistic for the joint null of zero coefficients on y_{t-1} and i_{t-1} shows to be insignificant, then we cannot reject the null hypothesis that the variable it is not a long run forcing variable. By interchanging y_t and i_t as dependent and independent variables in regression (14) we can assess whether y_t is or not a forcing variable. Here we consider the same four cases about deterministic components that were used for cointegration.[22]

Insert Table 5 about here

Table 5 reports the results of the tests together with the 5% critical bounds. If the statistic is below the 5% upper bound we cannot reject the null of no long-run relationship between the variables[23]. We report the tests both assuming that the interest rate is the forcing variable and that output is the forcing variable. The lag order was chosen using the SBC on the ECM model (14). We report both the $F-$tests and the $t-$tests for each of the deterministic component cases. The results reveal a very clear picture. In all the tests we can reject the existence of a long-run relationship

[20]In case that the ECM contains a deterministic trend, the F-test also includes the null of the coefficient on the trend being equal to zero.

[21]We refer to Pesaran et al (2001) for a detailed description of the testing procedure. Note that the critical values provided contain an upper and lower bound outside which inference is conclusive. However, if the F- or t-statistics fall within these bounds, we cannot reach any conclusion unless the cointegration rank of the forcing variable it is known a priori.

[22]Note that only in Cases III and V can we report t-tests as well as F-tests.

[23]Note that, in order to be on the conservative side, we will reject long-run relations even if the statistic lies within the critical bounds.

between output and the interest rate. This was also the case when using the quarterly data[24].

In order to test whether these results are stable and robust to the choice of the sample we carried out three stability testing procedures equivalent to those used for the cointegration analysis. The three different methods will obviously yield different patterns and give a complete overview of the stability of the results[25].

Insert Figures 5, 6 about here

Plots of the $F-$tests are provided in Figures 5, 6 together with the upper 5% bound. We do not present the results fixing the end-point for the sake of space as they always yielded values below the upper 5% bound. If the plot is above the bound there would be evidence of a long-run relation for that recursion. Focusing on the US Treasury Bill and tests FIII and FIV we can see that, despite some variation, the tests are always below the 5% bound with a tendency to decrease in the final years of the sample, and especially after the "Volcker disinflation" period. This is a very similar pattern to that found in the cointegration analysis. For the US, hence, the evidence unequivocally supports the absence of long term relationship between real output and the monetary policy indicator. For the UK our results also show a higher instability and some isolated periods of long term relationship. This set of results, however, support absence of long term relationship much more strongly than the cointegration tests, as most recursions yield statistics below the critical band. When looking at the annual data recursive tests for the UK using the Treasury Bill rate we can observe that the test substantially surpasses the upper bound in some periods which are common to those found when using cointegration tests. This period coincides with the inclusion of the years between 1988 and 1992 and is also reflected, to a lesser extent, in the

[24]The results using all the available sample period also show no long-run relationships.

[25]We also carried out formal tests for parameter stability on the unrestricted error correction model. We applied Hansen's (1992) stability test and found no evidence of individual parameter or joint instability. Instability was higher for the UK, although always below the critical values. For the US there was some evidence of variance instability.

quarterly data estimates. This is the period when the pound sterling first shadowed the DM and then entered the ERM and the subsequent speculative attack that took the pound out of the ERM in September 1992. The loss of monetary policy generated by these events may have had some long-run impact on output. However, this appears as an isolated event not supported by all three methods and should be taken with some degree of caution. For the rest of the observations for the UK the bounds test is below the critical band[26].

8 Omitted variables and nonlinearities

Our evidence so far shows that there is no long-run information content of interest rates for output, favouring the interpretation that the liquidity effects described in the theory model do not have a significant effect and are not sufficiently strong so as to be reflected onto statistical long term relations between output and interest rates. In other words, interest and output do not appear to share a common trend that allows establishing a long-run role of the monetary policy indicator. However, it could be argued that these results may be subject to omitted variables bias or are the result of the assumption that this potential relationship is linear. We address these problems in this section.

8.1 Omitted variables

Assume that long-run real output is determined by an $N-1$ set of real shocks q_t stemming from technology, labor supply, energy prices, etc., plus the interest rate. Then there should be a linear combination $\beta\left(y_t, i_t, q_t\right)'$, with β a $1 \times N+1$ vector of coefficients, which yields stationary errors. However, if the elements q_t are omitted from the relation, we would find that there is no common trend just because these real variables are needed to complete the long-run stationary combination. Hence, we will

[26] When using the full sample available for both the UK and US the results are also similar but also show outliers during the II World War (especially for the UK). Results are available upon request.

now analyze whether the results from the previous sections are robust to the inclusion of variables that share common trends with real output.

The way we proceed is as follows. We first test whether a set of potentially important real variables influencing the long-run evolution of output share common trends. If some of these variables share common trends we can reduce the vector to a sub-set of variables q_t^s that share common trends with the rest but not between them (see Boschen and Mills, 1995). Hence, the long-run impact matrix of this vector q_t^s will have rank 0 ($rank = 0$). We then test if these variables are cointegrated with real output, i.e. if the rank of $\{y_t \; q_t^s\}$ is 1. If so, we can conclude that these variables have long-run information content about output. We then introduce the interest rate in the vector and test for the rank of $\{y_t \; q_t^s \; i_t\}$. If we find that the rank of the long run impact matrix of this vector is still 1, we can then conclude that the interest rate does not enter the common trend of output with the rest of real variables. If the rank was found to be 2, we would then conclude that interest rates do have a long-run relation with output and our previous results are just a consequence of omitting relevant variables. Note that, given that we have previously found that the rank of the long run matrix of $\{y_t \; i_t\}$ is 0, this rules out that finding a cointegration rank of 2 in $\{y_t \; q_t \; i_t\}$is the consequence of cointegration between i_t and q_t. This is because if we find $rank = 2$ and we have that the relations $\{y_t \; q_t\}$ and $\{q_t \; i_t\}$ have $rank = 1$, then this automatically entails a cointegration relation between y_t and i_t. In case we found no long-run relation between y_t and q_t and then by adding i_t found $rank = 1$, then we can again conclude against no long-run information content and our previous results would be a consequence of omitted variables[27].

Insert Table 6 about here

Due to data availability issues we could only use annual data starting in 1949 for

[27] During this exercise we will work assuming that interest rates are I(1). This is because, to our knowledge, there is no system equivalent to the bounds test procedure used in section 5.

the US and the UK and ending in 1999 for the UK. The set of variables q_t included real oil prices (OIL), world income minus US/UK income (WI), a measure of multifactor productivity (MFP), total labor force (LF), real government expenditure (GE) and real total tax revenue (TAX)[28]. These variables would be able to capture shocks stemming from energy prices, external shocks, productivity, demographic change and fiscal policy as in Boschen and Mills (1995). The results from the cointegration tests between these variables in vector q_t showed that for the US the vector of real variables can be reduced to OIL and MFP and for the UK to OIL and LF[29]. For the US, OIL and MFP do not appear to have a common trend when we estimate the model with a trend in the data but not in the cointegration equation, but seem to have one when introducing a trend in the cointegration vector. For the UK, OIL and LF do not have common trends in any specification. We then test for cointegration relations between output and these two q_t^s variables. Table 6 reports the results of the LR test of no-cointegration. Given the behavior of the data involved and for reasons of space we report only Case II and III for deterministic trends. The results show that, for both countries, the real variables share a common trend with output both together and separately (with the exception of MFP in Case II for the US). When we introduce the Treasury Bill rate in the VAR, results reject the existence of more than one cointegrating vector. This lends support to the hypothesis that interest rates and output do not share common trends. That is, the results presented in the previous sections are not driven by the possible bias arising from the exclusion of relevant variables explaining long-run output.

8.2 Nonlinearities

It could be argued that, if any relationship exists in the long-run between these two variables, it may be nonlinear. Output may not respond to interest rate shocks unless these are very large so as to make the liquidity effects evident in the data. The response

[28] The source of the variables is described the Appendix The productivity measure for the UK was not directly available and was estimated as labor productivity not explained by capital deepening as in Hendry (2001).

[29] We do not report these tests here to save space, but results are available on request.

may also be asymmetric if the impact of positive and negative shocks is different. In these cases, linear tests for no long-run relationships would not be powerful against the alternative of a nonlinear long-run relationship. These nonlinearities can arise in two different ways. First, the speed of mean reversion towards the linear long-run relationship may be asymmetric and depend on the on how far away the variables are from their equilibrium relationship. In this case standard non-cointegration tests would have low power against the cointegration alternative. Secondly, the long-run relationship may be nonlinear in itself. In this case, the potential impact of interest rates on output may depend on the level of the interest rate. In this case the linear relationship would not constitute a long-run equilibrium vector and cointegration test would again fail to detect possible long-run effects.

We have thus addressed these two possible forms of nonlinearity. Recent developments on nonlinear error correction models in Kapetanios et al. (2006) allow us to consider the possibility that the speed of convergence towards a (linear) cointegration relation is a smooth function of the error correction term. The battery of tests presented in Kapetanios et al. (2006) constitutes tests for the joint hypothesis of nonlinearity and cointegration. We carried out these tests using the original data, the demeaned data and the demeaned and de-trended data as recommended by Kapetanios et al. Our findings show no support for the existence of an asymmetric mean reverting long-run relationship.[30].

The second form of nonlinearity, a nonlinear cointegration vector, can be addressed by making use of the recent contributions of Choi and Saikkonen (2004) and Saikkonen and Choi (2004).We consider two different functional forms for the nonlinear relationship. The first is a Logistic Smooth Transition model (LSTR) and the second an Exponential Smooth Transition model (ESTR). In the case of the ESTR model, the cointegration vector coefficient changes depending on whether interest rates are close or far away from an endogenously determined threshold, regardless of whether this

[30]We do not present the results of these tests here for reasons of space. All the results for this section, however, are available from the authors.

difference is positive or negative. In the LSTR case we would have that the reaction of output in the long run changes depending on whether interest rates are above or below a threshold. We followed a three-step procedure consisting of testing first for nonlinearity, estimating the nonlinear cointegration vector, and finally testing if this vector yields cointegration and eliminates nonlinearities.[31] The results from this procedure show that the null of linearity could not be rejected in any of the cases for the UK. For the US we find some evidence of nonlinearity. The estimation results show that only the LSTR model using demeaned and detrended data yields evidence of cointegration and no remaining nonlinearity. Several of the estimated coefficients, however, were not statistically significant. This is indicative that any possible nonlinear long-run relationship is indeed a very weak one and the results should be taken with high degree of caution.

9 Conclusions

In this paper we investigate long-term relationships between short term interest rates and real output. We use short term nominal interest rates as the relevant monetary indicator that contains significant and stable information about the U.K. and U.S. business cycle fluctuations in the post-II World War period. We first study a parsimonious endogenous growth model that accounts for the liquidity effects generated by cash in advance constraints. We show that the model can generate permanent effects of monetary shocks that would imply a long-run relationship between output and nominal interest rates. Next we provide evidence on whether or not the channels linking business cycles and monetary policy in models of fluctuations with endogenous growth are empirically relevant. If short-term interest rates have information content about short-run movements of output, an implication of these class of models is that

[31] Tests for nonlinearity and remaining nonlinearity are based on Hamilton (2001). The test for nonlinear cointegration is based on Choi and Saikkonen (2004) and the estimation of the nonlinear vector on Saikkonen and Choi (2004). Full details of the estimation and results are available. We thank In Choi for providing some of the GAUSS codes used in this section.

they could potentially contain information about output movements in the long-run. If this is the case, monetary policy that uses short term interest rates as either instruments or policy indicators would have to use this long-term information content in their information set.

Our various tests favor the absence of long term relationships between real output and nominal interest rates. There is neither significant nor stable long term relationship between short term interest rates and real output in the U.K. and the U.S. in most of the sub-samples considered. These results are robust to the inclusion of a set of real side variables that have common trends with output. For the UK we find some evidence of long term relationship during the ERM period. For the US only a particular nonlinear specification for the US detects a nonlinear long-run relationship, albeit only a weak one. In all, we conclude that there is very little evidence of long term relationships between output and interest rates.

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Appendix: Data sources

Variable	Period	Periodicity	Source
US data			
Real Output	1960-2001	Quarterly	OECD MEI
Treasury Bill (3-month)	1960-2001	Quarterly	IMF-IFS
Federal Funds Rate	1960-2001	Quarterly	FRB
Treasury Bill 3 month	1941-2001	Annual	FRB
Moody's AAA	1929-2001	Annual	http://www.globalfindata.com/
10-years Gov Bond Rate	1929-2001	Annual	http://www.globalfindata.com/
Real Output	1929-2001	Annual	FRB
Real Oil Prices	1949-2001	Annual	IMF-IFS
World Income	1949-2001	Annual	IMF-IFS
Multi-factor productivity	1949-2001	Annual	Bureau of Economic Analysis
Labor force	1949-2001	Annual	Bureau of Economic Analysis
Real Government expenditure	1949-2001	Annual	IMF-IFS
Real tax revenue	1949-2001	Annual	IMF-IFS
UK data			
Real Output	1960-2001	Quarterly	OECD
Treasury Bill 3month	1960-2001	Quarterly	IMF-IFS
10-GovBond	1960-2001	Quarterly	OECD
Treasury Bill	1873-2001	Annual	Hendry (2001) updated with IFS
10-year Gov Bond Rate	1873-2001	Annual	Hendry (2001) updated with OECD
Real Output	1873-2001	Annual	Hendry (2001) updated with OECD
Real Oil Prices	1949-2001	Annual	IMF-IFS
World Income	1949-2001	Annual	IMF-IFS
Productivity measure	1949-1999	Annual	Hendry (2001) updated with OECD and own estimates
Labor force	1947-1999	Annual	Hendry (2001) updated with OECD
Real Government expenditure	1947-2001	Annual	IMF-IFS
Real tax revenue	1947-2001	Annual	IMF-IFS

Table 1: Granger Causality χ-Square Statistics
(OLS Estimates, White Heteroskedasticity Consistent Standard Errors)

	χ-Square (p-values)
U.K. Real Output Equation (1948-2001)	
T-Bill	2.844 (0.091)
Gov Bond	6.669 (0.0098)
U.S. Real Output Equation (1947-2001)	
T-Bill	14.713 (0.0053)
Gov Bond	13.288 (0.0099)

Table 2: Unit root tests on output

		ADF		KPSS		M_α^{GLS}		ERS DFGLS	
	Lag	Const	Trend	Const	Trend	Const	Trend	Const	Trend
US									
1960:1-2001:2	4	-1.002	-3.335	2.825	0.282	1.500	-10.02	-1.752	-2.267
1947A-2001A	0	-0.925	-2.359	1.289	0.200	1.779	-6.734	3.224	-1.983
UK									
1960:1-2001:2	0	-0.482	-2.221	4.154	0.250	1.650	-7.506	2.421	-1.987
1947A-2000A	0	-0.698	-2270	3.645	0.230	1.924	-6.009	3.332	-1.928

NOTES: Bold indicates rejection of the null of a unit root for ADF, DFGLS and M_α^{GLS} and acceptance of the null of stationarity for the KPSS test at the 5% level.

Table 3: Unit root tests on interest rates

		ADF		KPSS		M_α^{GLS}		ERS DFGLS	
	Lag	Const	Trend	Const	Trend	Const	Trend	Const	Trend
US									
Quarterly Data (1960:1-2001:2)									
T-Bill	5	-3.205*	-3.084	0.521	0.442	-16.77*	-26.16*	-2.563*	-2.921*
FedFunds	2	-2.263	-2.159	5.525	0.913	-6.545	-8.708	-1.819	-2.158
Annual Data (1947-2001)									
T-Bill	0	-2.339	-2.233	4.678	0.689	-5.106	-8.821	-1.742	-2.205
Gov Bond	0	-1.839	-1.553	5.744	0.750	-2.891	-5.486	-1.308	-1.605
UK									
Quarterly Data (1960:1-2001:2)									
T-Bill	1	-2.710	-2.624	0.630	0.495	-7.504	-12.55	-1.939	-2.407
Annual Data (1948-2001)									
T-Bill	0	-2.174	-1.945	2.538	0.356	-3.150	-6.194	-1.384	-1.887
Gov Bond	0	-1.578	-0.630	6.225	0.992	-1.380	-1.528	-0.968	-0.679

NOTES: Ibid Table 2

Table 4: Johansen Cointegration Tests Likelihood Ratio Statistics (Full sample)

	Case I	Case II	Case III	Case IV
		U.K. (1948-2001)		
T Bill	40.57614*	12.28785	21.35119	10.04771
10 years Bond	32.83165*	7.828417	24.40714	19.09715*
		U.S. (1947-2001)		
T Bill	23.32090*	15.15664	19.45023	7.996870
Gov Bond	26.72015*	15.07568	21.91855	6.941328
		Critical Values		
5%	19.96	15.41	25.32	18.17
1%	24.60	20.04	30.45	23.46

Case I: no deterministic trend in the data, and an intercept but no trend in the cointegrating equation.
Case II: linear trend in the data and an intercept but not no trend in the cointegrating equation
Case III: linear trend in the data and both an intercept and a trend in the cointegrating equation
Case IV: quadratic trend in the data, and both an intercept and a trend in the cointegrating equation.

Table 5: Bounds Test analysis of long-run relationships

	Lag	F-II	F-III	F-IV	F-V	t-III	t-V
			US (1947-2001)				
T Bill → Y	2	0.788	2.241	2.006	3.554	-1.032	-2.462
Gov Bond → Y	2	1.076	2.375	2.075	3.378	-1.302	-2.331
Y → T Bill	2	2.012	0.639	0.595	0.019	-1.116	0.002
Y → Gov Bond	2	1.556	0.796	0.924	0.203	-0.760	-0.001
			UK (1948-2001)				
T Bill → Y	3	0.565	1.859	1.871	1.313	0.228	-0.503
Gov Bond → Y	3	0.274	1.725	1.722	1.167	0.017	-0.353
Y → T Bill	3	1.768	0.995	1.381	0.000	0.818	0.000
Y → Gov Bond	3	1.503	1.069	1.538	0.008	0.244	-0.008
				5% Critical Bounds			
		3.62	4.94	4.68	6.56	-2.83	-3.41
		4.16	5.73	5.15	7.30	-3.22	-3.69

The table produces tests for the existence of long-run relationships between real output and short term interest rates. It has F-tests and t-tests. There are 4 cases of deterministic components considered (corresponding to Pesaran et al., 2002, cases):

- Case II: restricted intercepts and no trends.
- Case III: unrestricted intercepts and no trends (t-test also reported).
- Case IV: unrestricted intercepts and restricted trends.
- Case V: unrestricted intercepts and unrestricted trends (t-test also reported).

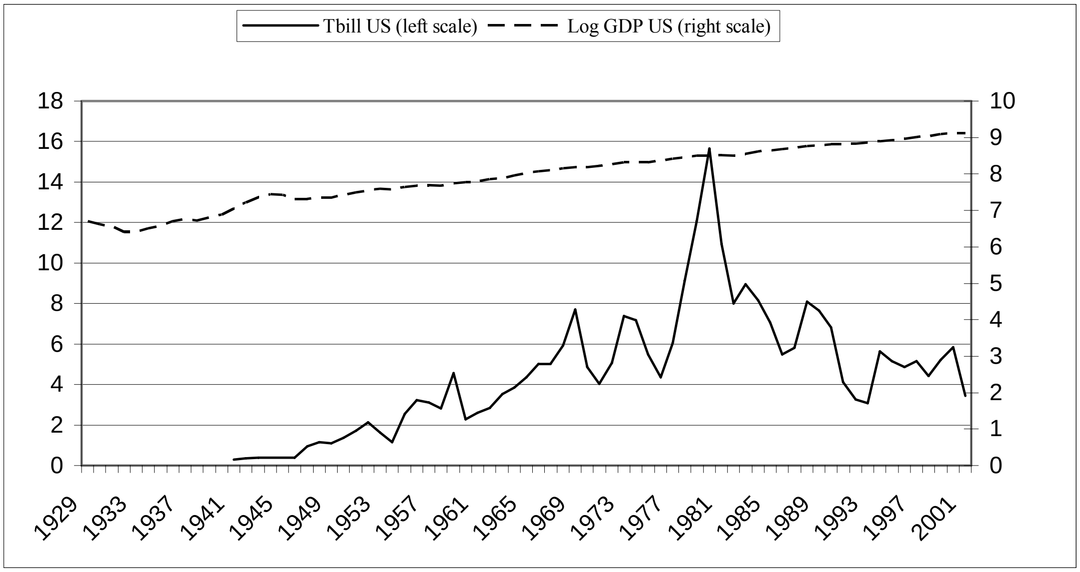
Table 6. LR cointegration tests. Omitted variables.

	Case II			Case III		
	$r \leq 2$	$r \leq 1$	$r = 0$	$r \leq 2$	$r \leq 1$	$r = 0$
	US (1949-2001)					
	Real variables and output					
Y, OIL, MFP	-	8.58	27.80*	-	9.63	38.25**
Y, OIL	-	-	23.20**	-	-	33.40**
Y, MFP	-	-	17.20	-	-	34.75**
	Real variables, output and interest rates					
Y, OIL, MFP, TBill	9.64	15.54	36.61**	9.75	15.71	45.22**
Y, OIL, TBill	-	6.65	26.70*	-	6.67	35.56**
Y, MFP, TBill	-	14.52	31.43**	-	14.80	38.78**
	UK (1949-1999)					
	Real variables and output					
Y, OIL, LF	-	12.82	24.48*	-	12.97	31.50**
Y, OIL	-	-	17.75*	-	-	26.48**
Y, LF	-	-	19.27*	-	-	27.48**
	Real variables, output and interest rates					
Y, OIL, LF, TBill	11.92	16.01	30.82*	11.38	16.92	43.47**
Y, OIL, TBill	-	13.65	24.51*	-	15.60	38.04**
Y, LF, TBill	-	11.87	21.05	-	11.43	33.68**

NOTES:
Case II: linear trend in the data and an intercept but not no trend in the cointegrating equation
Case III: linear trend in the data and both an intercept and a trend in the cointegrating equation

For reasons of space, if the VAR contains N variables, we do not report the $r \leq N-2$ test if the $r \leq N-3$ test already rejects the null.

Figure 1. Treasury Bill Rates and log GDP



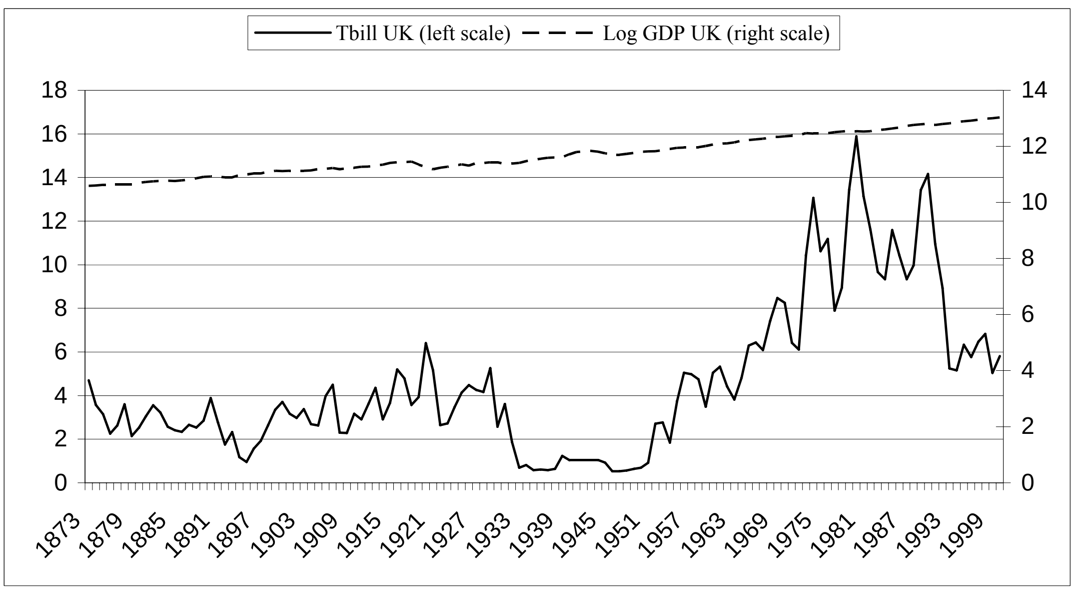


Figure 2: Granger-causality p-values: Rolling Regressions (30 years window)

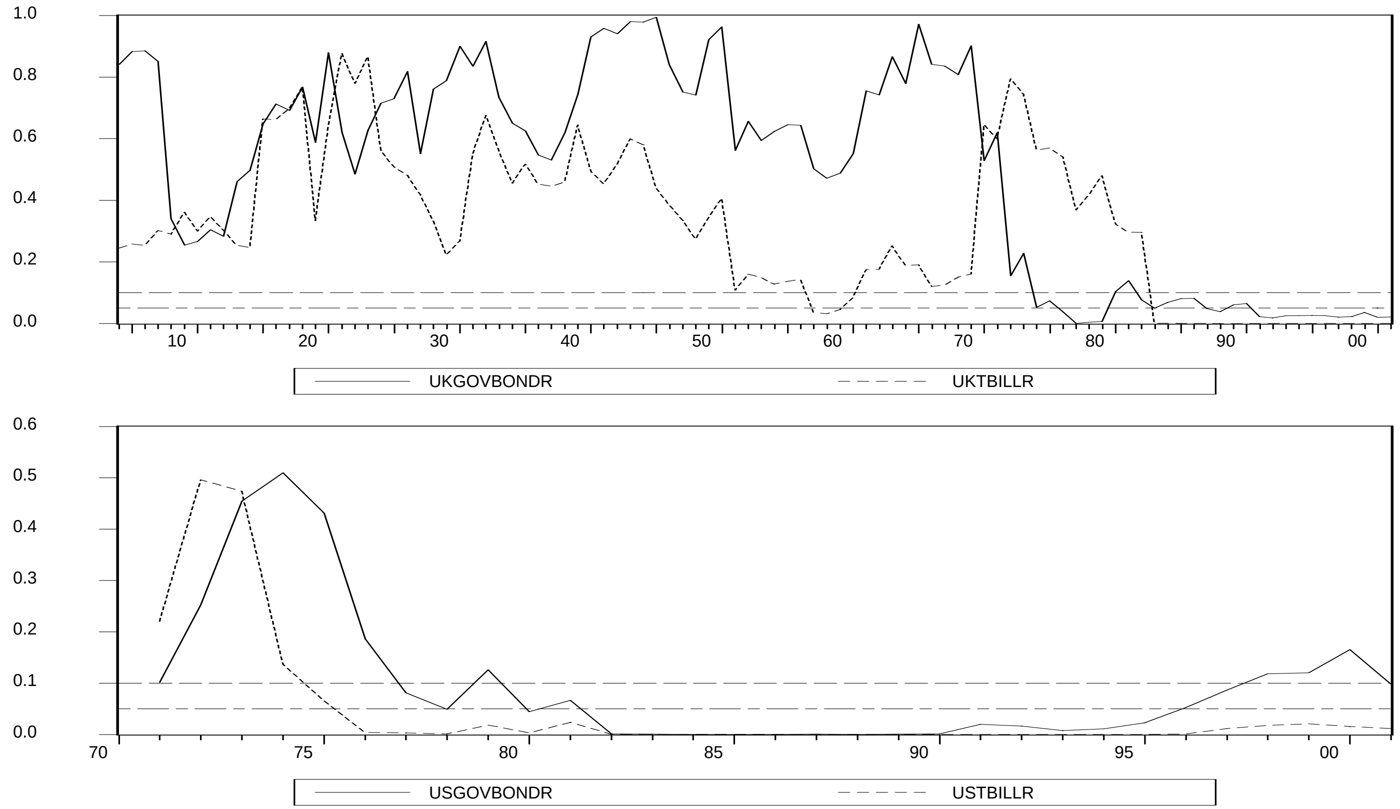


Figure 3: U.K. Cointegration Results: Sub-sample Stability 1948-2001 (likelihood ratio)

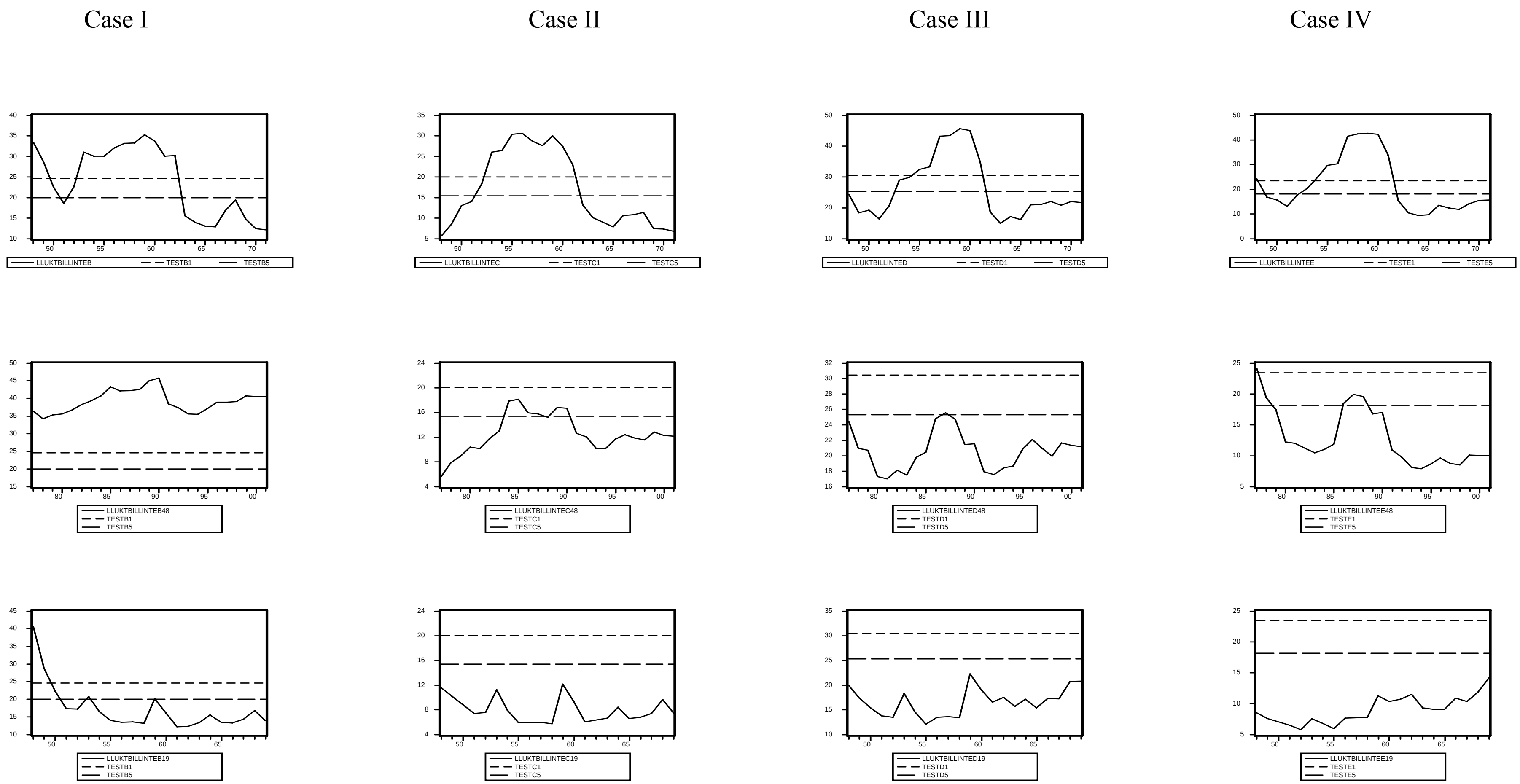


Figure 4: U.S. Cointegration Results: Sub-sample Stability 1947-2001 (likelihood ratio)

Case I Case II Case III Case IV

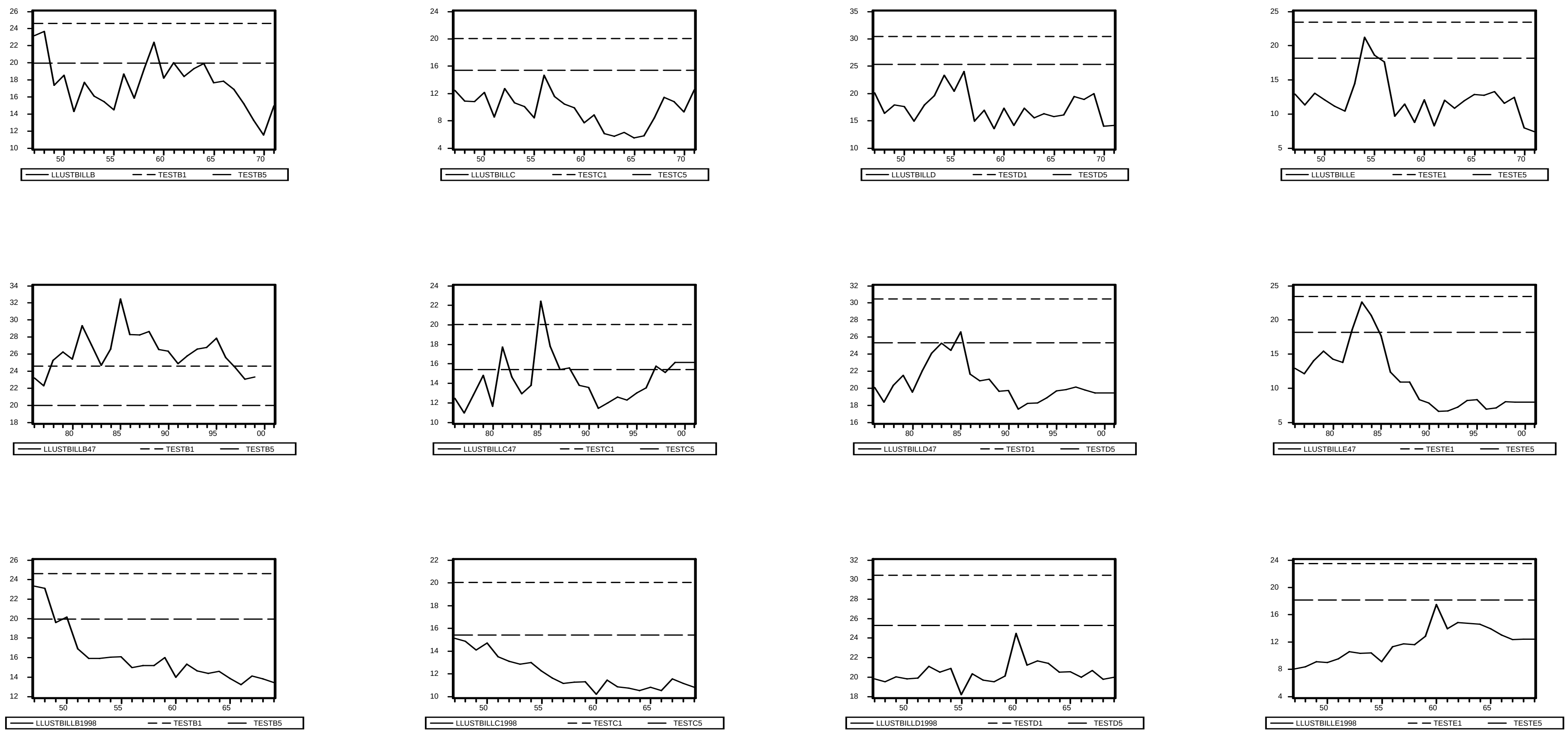
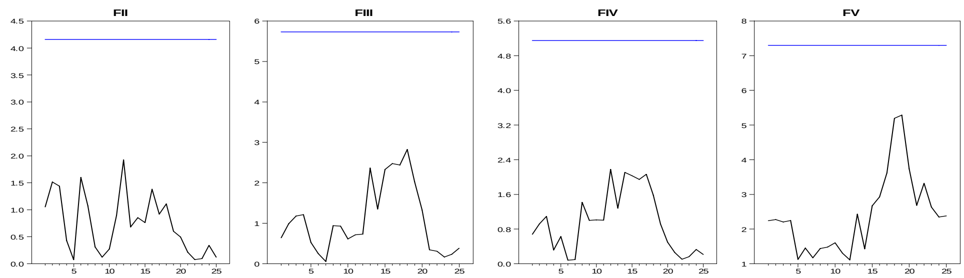
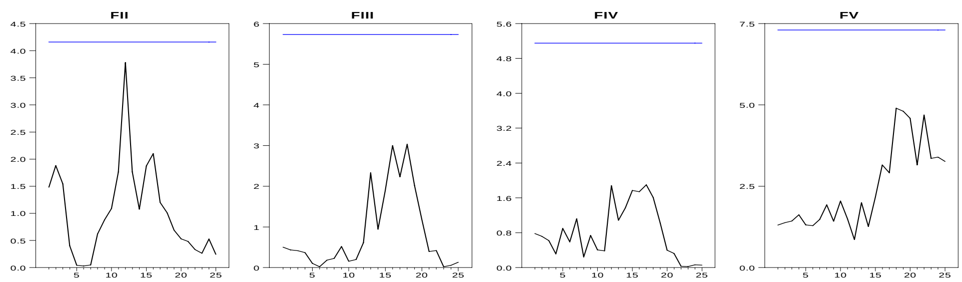


Figure 5: Bounds test results: rolling window estimates

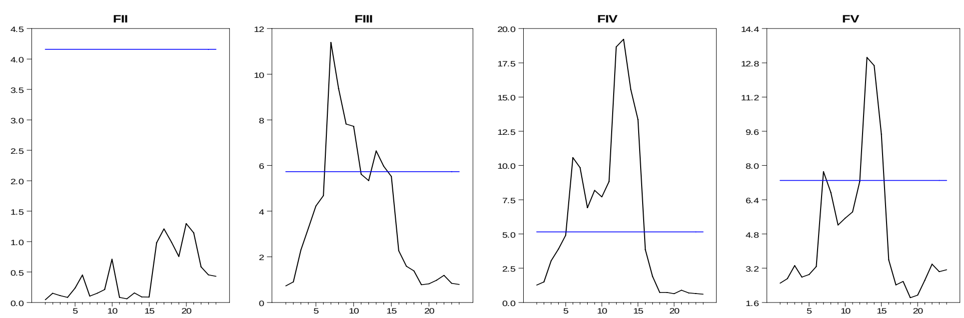
US Treasury Bill Rate



US Government Bond Rate



UK Treasury Bill Rate



UK Government Bond Rate

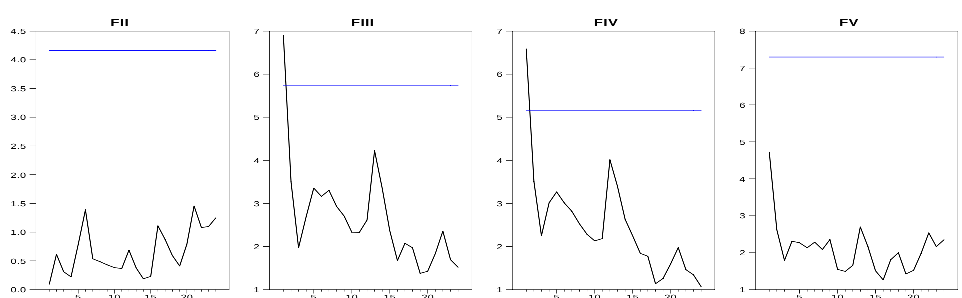
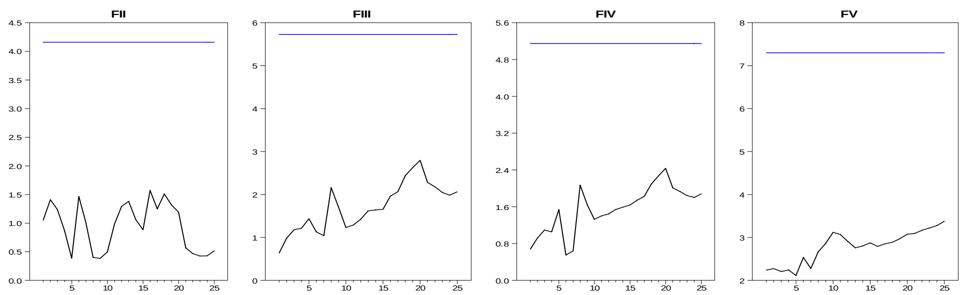


Figure 6: Bounds test results: recursive estimates and fixed initial point

US Treasury Bill Rate



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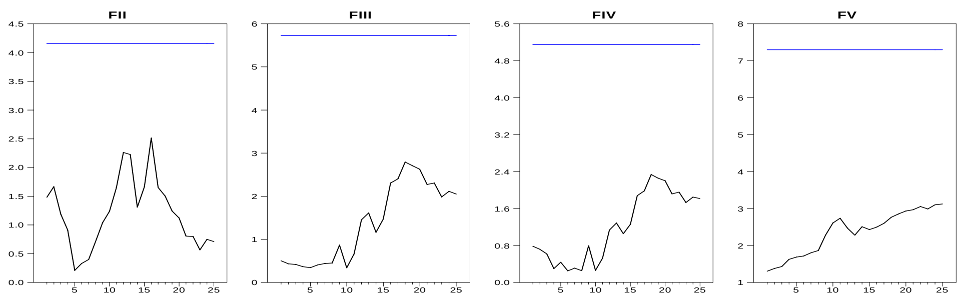

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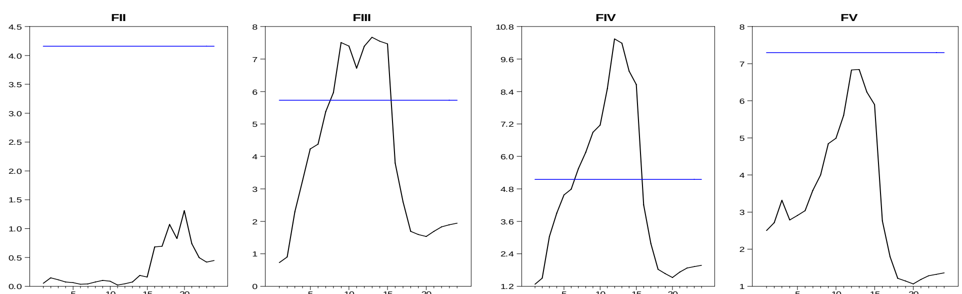

UK Government Bond Rate