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1. Introduction

The relationship between prices and output is crucial to understanding the nature of economic fluctuations and to be able to discriminate between rival macroeconomic models. A

found that they tended to move in the same direction before World War II but in the opposite direction afterwards. Antonakakisa et al. (2017) analysed the time-varying correlation between US output and prices by incorporating short-term interest rates, output and inflation volatilities in the model in order to capture the role of monetary policy, output and inflation uncertainty; they found evidence of time variation and of a predominant role for technology shocks.

Concerning the international evidence, Backus and Kehoe (1992) considered ten countries with data spanning at least a century (Australia, Canada, Denmark, Germany, Italy, Japan, Norway, Sweden, UK and US) and found that prices were more persistent and generally procyclical before World War II (WWII) and countercyclical afterwards; however, in the

period. Fiorito and Kolintzas (1994) also studied the G7 countries and concluded that the correlation between HP-filtered prices and output is negative during the post-WWII period. Finally, Pollin and Zhu (2007) analysed the relationship between inflation and economic growth in 80 countries over the 1961-2000 period; their results vary across countries but suggest a stronger positive correlation during periods characterised by more active demand management policies.

This paper aims to provide new evidence on the relationship between prices and output in both the US and the UK by focusing on the long-run rather than on cyclical fluctuations. For this purpose, it applies fractional integration and long

time. However, most macroeconomic series appear to be nonstationary. A standard approach to remove nonstationarity is to take first differences on the assumption that the differenced series will be stationary I(0). In such a case the original series is said to be integrated of order 1 or I(1). Following the seminal work of Nelson and Plosser (1982) many papers have therefore carried out standard unit root tests.[1] However, it is now well known that such tests have very low power under fractional alternatives,[2] as it is possible for a series to be neither I(0) nor I(1) but instead integrated of order d, where d can be any fractional value in the interval between 0 and 1, or even to be above 1. Gil-Alana and Robinson (1997) examined an updated version of Nelson and Plosser's (1982) dataset consisting of fourteen US macro variables and found that all except one were I(d) with 0 < d < 1. Since then, fractional integration has been widely used for the analysis of macro series (see, e.g., Mayoral, 1996; Chambers, 1998; Michelacci and Zaffaroni, 2000; Caporale and Gil-Alana, 2013; Abritti et al., 2016).

Therefore, the model estimated in the empirical section is of the following form:

$$, \quad (1 \quad L)^d x_t \qquad u_t, \qquad\qquad t \qquad 1, 2, ..., \qquad\qquad (1)$$

where x_t stands for either the observed data or the errors in a regression model that may include deterministic terms such as a constant or a linear time trend or weakly exogenous variables, d is a parameter to be estimated from the data providing a measure of persistence, and u_t is an I(0) process that is assumed to be in turn a white noise or exhibiting (weak) autocorrelation.

Fractional cointegration is the extension of the concept of fractional integration to the multivariate case. A necessary condition to test for it in a bivariate context such as

[1] The most commonly used are the ADF test (Dickey and C Tf1 d3 1 228.me

ours is that the two individual series display the same degree of integration. Since this condition is not satisfied in our case (see below) cointegration tests cannot be performed; instead we analyse the relationship between the two variables by treating each of them in turn as weakly exogenous, i.e. the other variable is regressed against lagged values of the weakly exogenous one; we carry out the analysis by using a simple version of the tests of Robinson (1994) that allows the inclusion of deterministic or weakly exogenous regressors in a model where the errors are potentially I(d) and d may be any real value.

3. Data and Empirical Results

The quarterly series used are the consumer price index (CPI) and real gross domestic product (GDP) (index 2015=100) for both the UK and the US. The sample period goes

As explained before, the fact that the price and output series do not have the same order of integration in either country implies that cointegration analysis cannot be carried out. Therefore, we examine their linkages by estimating a model in which one of the two is treated as weakly exogenous as follows:

$$y_{1t} \quad y_{2t\ k} \quad x_t, \quad (1 \quad L)^d x_t \quad u_t, \quad t \quad 1,2,..., \qquad (3)$$

for k = 1, 2, and 3, where y_{1t} and y_{2t} stand for log CPI and log real GDP respectively in Tables 5 and 6 for both the UK and the US, whilst in Tables 7 and 8 the opposite holds, namely y_{1t} stands for log real GDP and y_{2t} for log CPI respectively for both countries.

[Insert Tables 5 and 6 about here]

Concerning the regression of log CPI on log real GDP, in the UK case (Table 5) the estimated values of d are much higher than 1 regardless of the lag length and range between 1.43 (k = 3 and 4 with white noise errors) and 1.73 (k = 1 with Bloomfield errors); however, the slope coefficient is not significantly different from zero in any single case

4. Conclusions

This paper applies a fractional integration approach to UK and US quarterly data on prices and output from 1975Q1 to 2020Q2 to analyse the stochastic behaviour of these two variables and their long-run relationship in both economies - unlike most of the existing literature that focuses instead on their correlation over the business cycle. The univariate analysis indicates that all series are highly trended and persistent, exhibiting high degrees of integration, especially in the case of CPI.

As for their linkages, since the two variables have different degrees of integration in each of the two countries, fractional cointegration tests cannot be carried out. We assume instead weak exogeneity of each of them in turn and examine causality by testing for the significance of the lagged values of the variable treated as exogenous. We find that the only significant relationship implies the existence of a lagged effect of prices on output in the case of the US, which suggests a dominant role for demand shocks. An alternative approach could be based on the AutoRegressive Distributed Lag (ARDL) model (see Pesaran and Shin, 1999), which does not require the assumption of equal orders of integration. However, this framework has yet to be extended to the case of fractional integration. Work in this direction is currently in progress.

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Table 1: Estimates of d: raw data

i) White noise errors			
Series	No deterministic terms	An intercept	An intercept and a linear trend
CPI UK	1.05 (0.94, 1.17)	1.17 (1.09, 1.26)	**1.12 (1.06, 1.20)**
CPI USA	0.98 (0.87, 1.13)	1.21 (1.07, 1.36)	**1.16 (1.04, 1.30)**

Table 3: Estimates of d: Logged data

i) White noise errors			
Series	No deterministic terms	An intercept	An intercept and a linear trend
CPI UK	1.01 (0.92, 1.13)	1.45 (1.37, 1.55)	**1.35 (1.29, 1.41)**
CPI USA	0.98 (0.89, 1.10)	1.56 (1.48, 1.67)	**1.46 (1.39, 1.56)**
REAL GDP UK	0.98 (0.88, 1.11)	0.83 (0.77, 0.91)	**0.81 (0.74, 0.90)**
REAL GDP USA	0.98 (0.89, 1.11)	0.97 (0.87, 1.11)	**0.98 (0.91, 1.09)**
ii) Autocorrelated errors			
CPI UK	1.00 (0.84, 1.18)	1.55 (1.43, 1.70)	**1.40 (1.32, 1.54)**
CPI USA	0.95 (0.80, 1.15)	1.56 (1.43, 1.73)	**1.44 (1.33, 1.60)**
REAL GDP UK	0.89 (0.73, 1.09)	1.01 (0.87, 1.19)	**1.00 (0.83, 1.19)**
REAL GDP USA	0.93 (0.77, 1.13)	1.10 (0.92, 1.31)	**1.05 (0.90, 1.23)**

In parentheses, the 95% confidence band for the non-rejection values of d at the 95% level. In bold, the selected model for each series on the basis of the statistical significance of the regressors.

Table 4: Estimated coefficients of the models from Table 3

i) White noise errors			
Series	d	Intercept (t-statistic)	Time trend (t-statistic)
CPI UK	1.35 (1.29, 1.41)	2.6008 (276.26)	0.0197 (5.41)
CPI USA	1.46 (1.39, 1.56)	3.0861 (632.66)	0.0120 (4.13)
REAL GDP UK	0.81 (0.74, 0.90)	3.6998 (184.80)	0.0050 (8.15)
REAL GDP USA	0.98 (0.91, 1.09)	3.4544 (298.65)	0.0066 (8.55)
ii) Autocorrelated errors			
CPI UK	1.40 (1.32, 1.54)	2.5973 (262.91)	0.0232 (4.93)
CPI USA	1.44 (1.33, 1.60)	3.0862 (613.01)	0.0118 (4.22)
REAL GDP UK	1.00 (0.83, 1.19)	3.7020 (181.57)	0.0047 (3.13)
REAL GDP USA	1.05 (0.90, 1.23)	3.4534 (299.08)	0.0066 (6.12)

In parentheses in column 2 the 95% confidence band for the non-rejection values of d, and in columns 3 and 4 the t-statistics.

P / log CPI _ (UK

	Time trend (t-statistic)
	0.1301 (1.37)
	0.1345 (1.45)
	0.1355 (1.46)
	0.1164 (1.25)
	0.1382 (1.50)
	0.1345 (1.45)
	0.1355 (1.46)
	0.1165 (1.25)

of d, and in columns 3

P / log CPI_(US

	Time trend (t-statistic)
	0.2275 (2.32)
	0.2561 (2.60)
	0.2859 (2.89)
	0.1890 (1.87)