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LONG-RUN, PERSISTENCE AND COMMON TRENDS

AUTOR

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ESSAYS IN FINANCIAL ECONOMETRICS: LONG-RUN, PERSISTENCE AND COMMON TRENDS.

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A thesis submitted in fulfillment of the requirements for the degree of Doctor of Philosophy

in the

Doctoral Programme: Economics, Business, Finance and Computing Science

October 22, 2019

Declaration of Authorship

I, José Carlos VIDES, declare that this thesis titled, "ESSAYS IN FINANCIAL ECONOMETRICS: LONG-RUN, PERSISTENCE AND COMMON TRENDS." and the work presented in it are my own. I confirm that:

- This work was done wholly or mainly while in candidature for a research degree at this University.
- Where any part of this thesis has previously been submitted for a degree or any other qualification at this University or any other institution, this has been clearly stated.
- Where I have consulted the published work of others, this is always clearly attributed.
- Where I have quoted from the work of others, the source is always given. With the exception of such quotations, this thesis is entirely my own work.
- I have acknowledged all main sources of help.
- Where the thesis is based on work done by myself jointly with others, I have made clear exactly what was done by others and what I have contributed myself.

Signed:

Date:

"*If I have seen further than others, it is by standing upon the shoulders of giants.*"

Isaac Newton

INTERNATIONAL UNIVERSITY OF ANDALUSIA - UNIVERSITY OF HUELVA

Abstract

Doctoral Programme: Economics, Business, Finance and Computing Science

Doctor of Philosophy

ESSAYS IN FINANCIAL ECONOMETRICS: LONG-RUN, PERSISTENCE AND COMMON TRENDS.

by José Carlos VIDES

This thesis recognizes that the premises of standard cointegration ($I(1)/I(0)$ dichotomy) are too restrictive. In this sense, the empirical literature has shown that many economic and financial time series hold long-range dependence in the autocorrelation function but do not precisely exhibit a unit root process, i.e., the long memory process. For this reason, traditional cointegration assumptions that the time series may follow the dichotomy $I(0)/I(1)$ are discard, in favour that they follow a fractional process $I(d)$. We also shed the notion that the error term follows a stationary process ($I(0)$) in cases of cointegration of both variables. In turn, the rigidity of the traditional approach is overcome in favour of allowing for the series to be cointegrated, and the error term does not necessarily need to be $I(0)$; for example, the error term may be cointegrated in order $I(d-b)$, unlike other techniques that assume the error term is $I(0)$. In this sense, the Fractionally Cointegrated Vector Autorregressive (FCVAR) model is an expansion of the traditional cointegrated VAR (CVAR) model, and it allows to determine the number of equilibrium relations via cointegrating rank testing to estimate memory parameters, long-run cointegrating relations with adjustment parameters, and short-run lagged dynamics. To this end, in the current dissertation, we develop empirical analysis to demonstrate the properties of the time series under the fractional cointegration assumptions. In chapters 2 and 3 we consider the cointegrating relation and adjustment dynamics amongst four major stock markets for the Eurozone, and the five major stock markets for Latin America, respectively. The results evidence that there is a full financial integration in both economic regions, despite of the financial crisis occurred in recent years (this case is studied in chapter 2). The following 4 chapters are devoted to study the term structure of interest rates. Indeed, we summarize an empirical review of the Expectations Hypothesis of the Term Structure (EHTS) aiming to establish the adequate procedures for its measurement by using time series and evidencing the linearity restrictions associated with the traditional approaches used in time series applications on term structure (chapter 4). Furthermore, it is also analyzed the relationship between the European Over Night Index Average (Eonia) rate and 3-month Euribor rate (chapter 5). In chapter 6, we apply a pairwise estimation to a wide sample consisting on 9 different maturities of Treasury Constant interest rates. Otherwise, in chapter 7, we use two historical databases for the USA in order to check the behavior of short- and long-term interest rates. In this four chapters, we demonstrate the fractional properties of the cointegrating relations subject to the EHTS conditions. Additionally, we study how the spread is resulting both interest rates, jointly, analyzing the long memory in the spread that has implications for the monetary transmission mechanism and its effectiveness. In chapter 8, the US debt sustainability is analyzed taking into account the Intertemporal Budget Constraint conditions. We propose different scenarios in which the deficit, i.e., the difference between revenues and expenditures, possess different features, providing significant implications for policy makers. Then, the dominance between revenues and expenditures in the common trend is shown. Finally, in chapter 9, concerning the crude oil market, we test if the relationship between West Texas Intermediate and Brent crude oil is globalized or regionalized. Besides, the difference between both crude oils may be an indicator of forecasting, depending the value of its degree of integration and to finish, the driver of the relationship is defined, which may be an indicator for business operators, arbitrageurs, economic agents and policy makers.

Abstract in Spanish

Esta tesis reconoce que las premisas de la cointegración estándar (dicotomía $I(1)/I(0)$) son demasiado restrictivas. En este sentido, la literatura empírica ha demostrado que muchas series económicas y financieras tienen una dependencia de largo alcance en la función de autocorrelación, pero no exhiben con precisión un proceso de raíz unitaria, es decir, el proceso de memoria larga. Por esta razón, los supuestos tradicionales de cointegración de que las series puedan seguir la dicotomía $I(1)/I(0)$ se descartan, a favor de que sigan un proceso fraccional $I(d)$. También eliminamos la noción de que el término de error sigue un proceso estacionario ($I(0)$) en casos de cointegración. A su vez, la rigidez del enfoque tradicional se supera a favor de permitir que la serie se cointegra, y el término de error no necesariamente tiene que ser $I(0)$; por ejemplo, el término de error puede cointegrarse en orden $I(d-b)$, a diferencia de otras técnicas que suponen que el término de error es $I(0)$. Así, el Vector Autorregresivo Fraccionalmente Cointegrado (FCVAR) es una expansión del Vector Autoregresivo Cointegrado (CVAR), y permite determinar el número de relaciones de equilibrio mediante tests de ranking de cointegración, relaciones de cointegración a largo plazo con parámetros de ajuste y dinámica a corto plazo. Para ello, en esta tesis, desarrollamos un análisis empírico para demostrar las propiedades de las series temporales bajo los supuestos de cointegración fraccional. En los capítulos 2 y 3 consideramos la relación de cointegración y la dinámica de ajuste entre las cuatro principales bolsas de la Eurozona y los cinco principales de América Latina, respectivamente. Los resultados muestran que existe una integración financiera total en ambas regiones económicas, a pesar de la crisis financiera ocurrida en los últimos años (este caso se estudia en el capítulo 2). Los siguientes 4 capítulos están dedicados a estudiar la estructura temporal de los tipos de interés. Realizamos una revisión empírica de la Hipótesis de las Expectativas de la Estructura de Temporal (EHTS, por sus siglas en inglés) con el objetivo de establecer los procedimientos adecuados para su medición mediante el uso de series temporales y evidenciando las restricciones de linealidad asociadas con los enfoques tradicionales utilizados en sus aplicaciones en la estructura temporal (capítulo 4). También se analiza la relación entre el tipo intradiario (Eonia) y el Euribor a 3 meses (capítulo 5). En el capítulo 6, aplicamos una estimación por pares a una muestra amplia que consta de 9 vencimientos diferentes de tipos de interés. Por otro lado, en el capítulo 7, utilizamos dos bases de datos históricas de los EE.UU. para verificar el comportamiento de los tipos de interés a corto y largo plazo. En estos capítulos, testamos las propiedades fraccionales de las relaciones de cointegración sujetas a las condiciones de la EHTS. Además, estudiamos cómo es el diferencial entre ambos tipos de interés, analizando su memoria larga, con implicaciones para el mecanismo de transmisión monetaria y su efectividad. En el capítulo 8, se analiza la sostenibilidad de la deuda de los EE.UU. Teniendo en cuenta las condiciones de Restricción Presupuestaria Intertemporal. Proponemos diferentes escenarios en los que el déficit, es decir, la diferencia entre ingresos y gastos, posee características diferentes, lo que proporciona implicaciones significativas para los responsables políticos. Luego, se muestra cuál, entre ingresos y gastos, predomina en la tendencia común. Finalmente, en el capítulo 9, relativo al mercado del petróleo crudo, probamos si el West Texas Intermediate y Brent está globalizados o regionalizados. Además, la diferencia entre ambos crudos puede ser un predictor, dependiendo de su grado de integración y, para finalizar, se define el impulsor de la relación, que puede ser un indicador para operadores comerciales, árbitros, agentes económicos y formuladores de políticas.

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I would like to express my deepest gratitude to my supervisor, Antonio Golpe, for his unwavering support, friendship, patience, and careful supervision during these years. I deeply thank Jesús Iglesias for his helpful support and guidance throughout the thesis and for his friendship. I also thank the remaining members of Spanish Entrepreneurship Research Group who taught and inspired me during my time in University of Huelva. Knowing that the list of aknowledgements is incomplete, I would like to thank José Manuel Bravo, Emilio Congregado, Manuel Emilio Gegúndez, Concepción Román or Luis A. Gil-Alana.

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To Jesús Iglesias

Preface

This PhD thesis investigates the properties of macroeconomic and financial variables by using a novel mathodology, i.e., the Fractionally Cointegrated Vector Autorregressive (FCVAR) model. The origins of this contribution are in 2016, when I finished my Master thesis, regarding the stock market integration in the Eurozone and how the sovereing debt crisis affected this integration at Master in Economics, Finance and Computing Science. Professor Antonio Golpe, who is my thesis director, acted as supervisor.

The use of this methodology came from a part of the master's degree abovementioned when Professor Antonio Golpe imparted the basis of time series and predictive models and later, Professor Luis A. Gil-Alana gave a session on fractional integration and cointegration, within the Predictive Models module, which makes me interested on this topic. Subsequently, due to the transmitted motivation of Professor Emilio Congregado, I decided to start the current research under the supervision of Antonio Golpe, who introduced me in the topic of time series and its applications on different topics in macroeconomics and finance.

This doctoral dissertation, submitted to the International University of Andalusia, was written during a first part in which I was unemployed and later, during my employment as Interim Substitute Professor at the Department of Economics in Huelva, joint to my position in Punta Umbría town hall.

I am very grateful to my supervisor, Antonio Golpe for his consideration, patience, effort, knowledge provided and friendship. I am also grateful to Professor Jesús Iglesias who, with his knowledge and efforts, gives a lot of value to this doctoral dissertation and improved my researching skills.

Finally, I would like to express my gratitude to the Department of Economics of University of Huelva and the Master's faculty staff and colleagues, in particular, Emilio Congregado, José Manuel Bravo, Hector Carcel, Serhiy Lyalkov, Eduardo Thorices, Pedro Toscano, Adam Harrison, Julia Feria, Concepción Román, Raquel Justo or Ana María Rodríguez, among others, for their comments and suggestions. I could not quantify how much I have learned from all of you.

Chapter 1

Introduction

Fractional cointegration is a generalised class of cointegrated systems which provides the possibility to estimate the fractional orders of integration of the time series, rather than fixing the memory parameters to be integer values ($I(1)/I(0)$ dichotomy). Empirical studies found that many macroeconomic and financial variables possess long memory in the long-run; nonetheless, nnot much attention has been paid to the dynamics of short-run adjustment of the fractional cointegration relationship (Cheang, 2018). The feature of nonlinear adjustments in long-run equilibrium relation of cointegrating variables is separately documented in the strand of traditional standard cointegration literature. In particular, the rigidity of the traditional approaches is overcome in favour of allowing the series to be cointegrated or order $I(d)$, and the error term does not necessarily need to be $I(0)$; for example, the error term to be cointegrated in order $I(d-b)$, unlike other techniques that assume the error term is $I(0)$. Under this assumption, we determine a controversy derived from the standard cointegration does not allow the spread to be nonstationary.

The use of the Fractionally Cointegrated Vector Autoregressive (FCVAR, hereafter) model (Johansen, 2008a, 2008b or Johansen and Nielsen, 2012) provides many advantages when estimating a system of fractional time series variables that are potentially cointegrated. This model, which is extended to allow for deterministic trends, has advantages when estimating a system of fractional time series variables that are potentially cointegrated. Additionally, the flexibility of the model allows one to determine the number of equilibrium relations via statistical tests and jointly estimate the adjustment coefficients and cointegrating relations while accounting for short-run dynamics. Each of these features will typically be relevant to the research question in empirical work (Dolatabadi, Nielsen, and Xu, 2016).

1.1 Econometric framework: an overview

As we mentioned before, this study takes a macroeconomics approach based on a new methodology in the literature, focusing on the relationship amongst different variables. Our starting point is the study of possible cointegrating relationships and then, depending on the nature of the variables selected, the degree of cointegration is also studied.

In this respect, the search of cointegrating relationships and the checking of the degree of cointegration is guided by the use of the Fractionally Cointegrated Vector Autorregressive (FCVAR) model (Johansen, 2008a, 2008b; Johansen and Nielsen, 2012) to test for the presence of fractional cointegration, by using a MATLAB program developed by Nielsen and Popiel (2016). This model have been combined with Bai and Perron (2003) test for the study of the presence of possible structural breaks. In some chapters, this methodology allows us to explore the persistence in the error correction term and the adjustment coefficients, attending to the Vector Correction Error Model in a fractional cointegration environment.

Finally, as Dolatabadi et al. (2016), Dolatabadi, Narayan, Nielsen, and Xu (2018) show, the FCVAR model permits to combine the study of the common trends by adapting these estimations to the Gonzalo and Granger (1995)'s Permanent-Transitory decomposition.

1.2 Contributions of this thesis

The contribution of this thesis with respect to previous work regarding the use of time series techniques on macro and financial variables. First, we apply a novel and sophisticated time series methodology, i.e., the FCVAR model, to be able to investigate empirically different topics under the fractional cointegration assumptions. In particular, we have applied this method to weekly, monthly, quarterly and yearly macro and financial time series, showing flexibility and an adequacy for different environments, which enable us to draw inference on the long-run relationship and/or short-run dynamics. Second, when investigating jointly the possible cointegrating relations and the behavior of the error term, we have developed and extended the existing literature, elaborating new scenarios in which policy makers and the concerned parties may have decisions. Third, this methodology allows us to combine with other techniques in order to achieve better results and contributions. These contributions are detailed in next sections and chapters.

1.3 Chapter overview

The dissertation has 4 parts, divided in 9 chapters. Part I includes chapters 2 and 3, we check the possible cointegrating relationship amongst European and Latin American stock markets, respectively. In part II, chapters 4 to 7 addresses the monetary transmission mechanism by using different interest rates maturities and countries, giving monetary policy implications. Part III includes chapter 8, where the budget debt sustainability for the USA is analyzed. Finally, part IV and, therefore, chapter 9 covers the behavior of the relationship between West Texas Intermediate and Brent crude oils and the spread resulting both crude oils.

The convergence of international markets is a phenomenon resulting in multiple confluences of economic, technological and political factors being those which have allowed national and international regulation to become more in line with economic forces and globalization processes. Indeed, in chapter 2, the financial integration among stock markets[1] in the Eurozone is checked by using Fractionally Cointegrated Vector Autoregressive (FCVAR) model, showing a perfect and complete Euro financial integration. Considering the possibility of the existence of structural breaks, we apply the test for structural breaks proposed by Bai and Perron (2003) that the which would provide a better empirical description of the European market integration. Among the breaks identified, the first regime (1998:01 until 2001:04) is in the way to the introduction of the single currency thus the markets were regulating to the new financial context. The second regime (2001:05–2007:06) would correspond to the economic growth and expansion period of the countries of the stock markets selected. In the third regime (2007:07 until 2012:04), according to the European Area Business Cycle Dating Committee, there was the financial crisis and the sovereign debt crises. In this respect, we also examine the fractional cointegration in each regime, showing that the Euro financial integration is very robust but in the financial and sovereign debt crises regime, IBEX 35 appears to be the weak link in the integration, unless the results show that when this period is finished, the Euro financial integration returns to be full. Chapter 3 proposes a wide review of the most relevant papers in the field of financial market integration by regions. In addition, it is intended to take a further step in the investigation of the long-term relationship using an expansion of cointegration, that is, fractional cointegration (FCVAR model). Nevertheless, the aim of this chapter is to test the possibility of a financial integration among five Latin American stock markets[2] as possible evidence for their economic development, and the possible expansion of the Latin American Integrated Market from a novel econometric perspective, by using a monthly sample which spans from September 2004 to June 2019. Hereby, the analysis suggests that there are four cointegrating vectors among the five equity markets, suggesting that Latin American stock markets are fully and perfectly integrated. Furthermore, the estimate of common order of integration of five stock markets shows that the stochastic trend is

[1]For our empirical analysis, we use a monthly sample of closing stock market prices for the period of January 1998 to September 2016 of the four major stock markets of the Eurozone, namely, Germany (DAX), France (CAC), Spain (IBEX) andItaly (FTSE MIB).

[2]The markets used are MERVAL, from Argentina; BVSP, from Brazil; IPSA, from Chile; IGBC, from Colombia; and IPC, from Mexico.

fractionally nature and possess stationarity with long memory.

In part II, chapter 4 summarizes an empirical review of the Expectations Hypothesis of the Term Structure (EHTS hereafter) aiming to establish the adequate procedures for its measurement by using time series. On one hand, the chapter discusses the main findings in the literature in the USA and the EMU and, on the other hand, analyses the linearity restrictions associated with the traditional approaches used in time series applications on term structure. The use of FCVAR model represents a novel procedure to solve the linearity restrictions. Finally, this application allows the economic policies that derive from its results to be more appropriate for the objectives of the design of monetary policies. Chapter 5, aiming to analyse the expectations hypothesis of term structure (EHTS), persistence in the European OverNight Index Average (Eonia) spread and permanent-transitory decomposition using a novel approach. We use a monthly frequency sample for the 3-month Euribor rate and Eonia rate, covering the period from January 1999 to February 2019. The results obtained confirm the EHTS and show evidence of a high persistence of the spread, which means that shocks may impede effectiveness in monetary policy and that the European Central Bank (ECB) loses control over interest rates. Additionally, according to Gonzalo and Granger (1995)'s Permanent-Transitory decomposition, we determine that the Eonia rate has a permanent component and thus dominates the common trend in the cointegration system. In chapter 6, we consider the possibility that the FCVAR model could serve as a novel empirical tool for examining the US term structure of interest rates. This econometric approach allows one to test the existence of a long-run relationship between short- and long-term interest rates[3] and spread persistence together in a pairwise estimation. As one of the main contributions of this paper, we elaborate on new scenarios of the degree of noncontemplative EHTS fulfillment. The results obtained contribute new scenarios not previously presented in the literature. We also find that the persistence of spread is the stronger the larger the difference in maturity is between considered interest rates, revealing a long memory process, which implies consequences of controlling power over interest rates by FED. Additionally, we try to explain how the Quantitative Easing program and its impact on the long-run relationship between each pair of maturities so, we apply the FCVAR model in both regimes obtaining different results. On the one hand, for the first regime, the resultsshow steady behavior, where most interest rate maturity pairs analyzed are cointegrated in a(1, -1) vector, and the spread follows a stationary process and, on the other hand, according to the Regime II estimations, this regime covers the aftermath of the global financial crisis and government efforts to allay the impact of this quarrelsome period. Consequently, the results obtained are very similar to those of the original sample, i.e., two pairs of maturities follow a nonstationary but mean-reverting process. Finally, in chapter 7, we check the fulfillment of the EHTS throughout the last century and half in the USA. For this reason, we use two types of database, i.e., the Jordà-Schularick-Taylor Macrohistory Database (from 1870 to 2013 on an annual basis) and Shiller's database (it begins in 1871 and finishes in 2011), finding similar results. In both estimations, we cannot reject the EHTS in this time period, and more importantly, according to the FVECM, the coefficients associated with short-term rates are significant, which implies that the spread has prediction power in the bearing of futures short-term rates. We also find that the long-term rate drives the long-run relationship, contributing to the total proportion to the common trend, and the persistence of the spread shows control power over interest rates by Fed.

In part III, in chapter 8, we analyze the US debt sustainability by applying the FCVAR model following the Intertemporal Budget Constraint (IBC), which is generally based on the analysis of the past behavior of the fiscal variables. This chapter shows a new approach in the literature to provide additional evidence on the long-run sustainability of the US government fiscal policy. In this sense, we confirm the existence of a cointegration relationship between expenditures and revenues and provides evidence that the US budget deficit shows strong sustainability over the period that covers, in a quarterly sample, 1947Q1 to 2019Q2. Furthermore, focusing on the degree of persistence of the budget balance, it is a key question for fiscal

[3]For our empirical analysis, we employ a monthly sample of Treasury Constant interest rates of 9 different maturities for the period of October 1993 to December 2018. The data correspond to 3-month, 6-month, 1-year,2-year, 3-year, 5-year, 7-year, 10-year and 20-year constant maturity rates

policy management. The results also support that the budget deficit follows a non-stationary process but reverts to its mean, which could suggest a slow speed adjustment towards the equilibrium of the public accounts. Consequently, strong measures would be necessary to neutralize exogenous shocks and to support the fiscal balance adjustment when those shocks affect it negatively, particularly troublesome. Furthermore, attending to the FVECM and the Permanent-Transitory decomposition and subsequently the component share, we have found that expenditures and revenues are permanent components in the common trend and that expenditures are sensitive to revenues in a similarly manner to how revenues are sensitive to expenditures.

Finally, chapter 9, in part IV, analyses the possible relationship between two of the main indicators of the oil market, the North Sea Brent (Brent) and West Texas Intermediate (WTI) crude oil prices, employing a weekly sample of the Brent and WTI crude oil prices over the period from 15^{th} May 1987 to 19^{th} April 2019 and by using the FCVAR model to determine whether these markets are regionalized or globalized. We propose to measure the cointegration and the stationary simultaneously, which would allow us to study new scenarios in which both prices could be cointegrated but the spread could be nonstationary. Additionally, this model allows us to identify other points of interest, such as the price structure and the persistence of the spread between each one. Although the FCVAR shows that these markets are strongly globalized, attending to the stationary of the spread, this shows a long memory process, and consequently, the shocks are long-lived. This result is novel in the literature, since until now, the globalization or regionalization of markets has been defined from these perspectives individually. In addition, the results confirm that Brent drives the price structure.

1.4 Conclusions of the thesis

Combining the empirical evidence provided in the various chapters in this book, we can formulate some conclusions as regards the application of the FCVAR model in variables of macroeconomics and financial nature, discarding the traditional assumption of the cointegration techniques, i.e., the dichtomy $I(0)/I(1)$. The empirical evidence presented in this book is diverse, although they deepen the topics analyzed and fill in a gap in the different selected literature.

Chapter 2 shows the financial integration in the Eurozone which, following Kasa (1992), perfect and fully. However, considering the existence of structural breaks, the Bai–Perron test is applied, detecting 3 structural breaks and then testing the FCVAR model in each of four regimes. The results for the different regimes show that integration of the European markets has been complete however, during the sovereign debt crisis, this full integration disappeared because IBEX 35 index went out of long-run equilibrium, which could mean that this index was more sensitive during this period, being the weak link in the integration. Once this turbulent period ended, full Euro financial integration resumed. Financial integration is attributable to technological advances during recent decades, which has reduced transaction costs and allowed for greater access to information. It has thus contributed to more sustainable economic growth. The findings of the paper have important implications for investors and policy makers.For investors, the high degree of integration implies a more attractive place for investment. However, as stock market prices are interrelated, the possibility of strong impacts from external shocks is not reduced. In this line, cointegration may imply perfect spillover. For policy makers, market integration in the Eurozone has led to various debates. Market integration has increased competition and market efficiency and led to greater interdependence between the Eurozone markets; this may require increased supervision and securities market oversight.

Chapter 3 shows a similar work than in the previous chapter. In this chapter the stock market integration in Latin America is checked, proposing a new focus in the field. The reason for selecting this region is because of its rapid economic growth and its opening up as a market for foreign investors. It is evidenced that these stock markets are fully and perfectly integrated and this perfect integration show a stationary, long memory behaviour. Attending to the long-run relationship and the absence of weak exogeneity of each stock market, as their prices are interrelated with themselves so, the possibility of exposure from external shocks is not diminished. This results also entails implications for policy makers and investors. On the one hand, the financial integration would contribute greater stability and would allow each country to be more competitive and efficient in the region. Whereas, on the other hand, investors would prefer to invest in markets characterized by increasing growth, which will give them more investment options and risk diversification opportunities. Finally, regional financial integration could foster the development of Latin American indices and preserve financial experience and innovation in the region.

Chapter 4 corresponds a deep review of the evidences in accordance to the EHTS in different regions, mainly the USA anf Europe. In this chapter, the term structure of interest rates is analyzed under non-linearity assumptions based on the premises of fractional cointegration as a way to avoid the restrictions of the standard cointegration.

Chapter 5 shows a long-run relationship between Eonia rate and the 3-month Euribor and its spread follows anon-stationary but mean-reverting process. As a consecuence, the greater persistence in money market rates may indicate the difficulty for monetary policy signals to be transmitted along the money market yield curve. Furthermore, the lasting impact of shocks may impede the transparency of policy signals and therefore, the ECB would suffer a gradual loss of control power over interest rates. Thus, our political recommendation is that, although the ECB has monetary policy tools linked to interest rates, the transmission mechanism of these policies is not guaranteed to be immediate. Indeed, if the ECB wants to keep the interest rate under control, it must contemplate the evolution of the Eonia rate.

Chapter 6 evinces persistence in the spread of each pair of interest rates maturities, which possess some important implications for monetary policy. This persistence may affect the Fed's control of long-term interest rates and of the yield curve. To address this issue, the Fed should increase the frequency of money market interventions. Additionally, as the Fed only has power over shorter-end interest rates, its manipulation may influence other short-term interest rates and thus may be necessary for the application of measures affecting longer-term rates when the monetary policy transmission mechanism predicted by the EHTS is not met. Policies oriented over time, such as the Quantitative Easing program, would thus be necessary to maintain this transmission mechanism or the substitutability of interest rates.

Chapter 7 displays that across the last century and a half, despite of there were wars, economic crises and/or changes in economic policy in the USA, the EHTS is supported, and thus, the Federal Reserve has a control power over monetary policy. In addition, the spread persistence gets a value below 0.5, which could be an indicator that the Fed already has control over term structure. If the spread is stationary, the long-and short-term rates are driven by a common stochastic trend and do not allow arbitrage opportunities because market forces adjust to correct any temporary disequilibrium. Overall, the results endorse the creation of a figure such as the Federal Reserve, which has maintained the effectiveness of monetary policy.

Chapter 8 presents a novel empirical strategy that detects the different types of sustainability based on the values of the cointegrating vector and the degree of integration of the error term, i.e., the deficit. In our understanding, the prism under which the cointegration approach has been applied to the debt sustainability analysis has been very limited. However, the FCVAR breaks this assumption so that although a unitary relationship between income and expenses exists, their cointegration relationship could be long-lived and even non-stationary. In other words, the strong sustainability concept proposed by the IBC theory should be taken with caution when it is tested empirically in the sense that, despite contemplating cointegration between expenses and income, any shock could have long-lived temporary effects. Consequently, strong measures would be necessary to neutralize exogenous shocks and to support the fiscal balance adjustment. Thus, if the US government aspired to achieve a strong sustainability, the burden of correcting budgetary disequilibria is entirely carried out via policy mixes. One plausible measure could be the treatment of expenditures because expenditure programs can be handled more easily than complex tax legislation.

Finally, chapter 9 treats the controversy in the existing literature concerning the treatment of the crude oil market, the fractional cointegration model voids most of the problems raised in this literature, proposing to measure the cointegration and the stationary simultaneously, which would allow us to study new scenarios in which both prices could be cointegrated but the spread could be nonstationary. Our results support several implications for business operators, arbitrageurs, economic agents and policy makers. First, a globalized market determines the price configuration of the Brent and WTI oil markets, assuming that oil markets have linked prices moving closely together. However, the spread follows a long memory process, which impede the immediate adjustments and increases the arbitrage opportunities. Nonetheless, business operators could use this spread persistence for investment provisions. Finally, government policies will have a long-lived effect that is, the effect will not be immediate. Indeed, it is possible that erroneous signals to the monetary policy authority are sent, which could feel the need to affect interest rates to mitigate the impact of oil prices on the economy.

1.5 Publications

The following publications emerged as a result of this dissertation:

- **Chapter 2**: Vides, J. C., Golpe, A. A., and Iglesias, J. (2018). How did the Sovereign debt crisis affect the Euro financial integration? A fractional cointegration approach. *Empirica*, 45(4), 685-706.
- **Chapter 4**: Vides, J. C., Iglesias, J., and Golpe, A. A. (2018). The Term Structure Under Non-linearity Assumptions: New Methods in Time Series. In *New Methods in Fixed Income Modeling* (pp. 117-136). Springer, Cham.
- **Chapter 5**: Golpe, A. A., Iglesias, J., and Vides, J. C. (Forthcoming). The role of EONIA in the dynamics of short-term Interbank rates. *Panoeconomicus*.
- **Chapter 6**: Vides, J. C., Golpe, A. A., and Iglesias, J. The EHTS and the persistence in the spread reconsidered. A fractional cointegration approach. (Under review in *International Review of Economics and Finance*, 2^{nd} phase)
- **Chapter 9** Bravo-Caro, J. M., Golpe, A. A., Iglesias, J., and Vides, J. C. (Forthcoming). A new way of measuring the WTI - Brent spread. Globalization, shock persistence and common trends. *Energy Economics*.

Part I

Financial market integration

Chapter 2

How did the Sovereign debt crisis affect the Euro financial integration? A fractional cointegration approach

2.1 Introduction

The convergence of international markets has resulted from multiple confluences of economic, technological and political factors that have allowed national and international regulations to increasingly align with economic forces and globalization processes. The formation of the Euro was an effort to enhance synergies of member countries, creating highly favourable conditions in which capital markets could develop important similarities between them (Salgado, Saldivar, and Ríos, 2015).

Relationships between stock markets have been widely studied from different perspectives. Using techniques such as EMH (Kim, Stern, and Stern, 2009), CAPM (Heimonen, 2002) and/or GARCH (Illueca and Lafuente, 2002), conclusions about relationships, convergence or co-movements among markets have been reached. Furthermore, several techniques have been used to apply time series data (see Brooks, 2014) to integration and cointegration among different global economic regions, mainly the USA-EU (see Caporale, Gil-Alana, and Orlando, 2015, among others), and intraregional markets, such as members of the EMU (Da Fonseca, 2013).

The aim of this paper is to study financial integration among the four major stock markets in the Eurozone (Germany, France, Spain and Italy) for the period of January 1998 to September 2016 from an econometric perspective.[1] This paper presents a novel approach to the integration of stock markets, filling a gap in the literature with regard to time series analysis of market cointegration. In this sense, our paper contributes to previous literature on the analysis of the integration of stock markets from a fractional cointegration vector autoregressive perspective. Although fractional cointegration had been used in previous studies, the approach proposed by Johansen (2008a) and Johansen and Nielsen (2012) is novel to the literature. This model, which is extended to allow for deterministic trends, has advantages when estimating a system of fractional time series variables that are potentially cointegrated. Additionally, the flexibility of the model allows one to determine the number of equilibrium relations via statistical tests and jointly estimate the adjustment coefficients and cointegrating relations while accounting for short-run dynamics. We use data with a monthly frequency

[1] The stock markets studied include the German stock market, the behavior of which is reflected in the DAX index; the French stock market, reflected in the CAC 40 index; the Italian stock market, as indicated by the FTSE MIB index; and the Spanish stock market, as shown by the IBEX 35 index. The choice of a stock market is based on the size of the respective national economy and the capitalization of the stock markets, which are the major ones in the Eurozone.

to estimate the model, then perform statistical tests of cointegration, exclusion and weak exogeneity. We then apply the Bai and Perron (2003) test for structural breaks and use the FCVAR model to examine each break detected.

The remainder of the chapter is organized as follows. Section 2.2 provides a review of the literature, focusing initially on the techniques used to study stock markets and subsequently on the application of the integration and cointegration test in different economic regions. Section 2.3 presents the methodology applied. Section 2.4 discusses the empirical results, and conclusions are presented in Sect. 2.5.

2.2 Literature review

Some measure of market development is essential in making intertemporal comparisons. For this reason, the treatment of such variables can explain the relationship between markets in the same economic region or, conversely, whether markets in different regions exhibit similar behaviour. As a result of computerized trading systems, markets can operate simultaneously. This allows for the study of the integration of stock markets, whose interrelations had previously been studied in various ways, e.g., using financial techniques such as the Efficient Market Hypothesis (EMH) or the Capital Asset Pricing Model (CAPM), until econometric models such the unit root test, GARCH and cointegration tests became available. The EMH is based on return predictability, as seen in the past price history of a market (Fama 1970, 1991), combined with other techniques such as the unit root test (Kim et al., 2009) or the variance ratio test[2] (Huang, 1995; Smith, 2007).

In contrast to previous research that has sought to explain intra-market behaviour, new research exploring this link has emerged, using other techniques, such as the international Capital Asset Pricing Model (CAPM) (Sharpe, 1964), which proposes that stock market returns are affected by interest rates movements. Thus, for an investor in international markets, excess returns are related to changes in exchange rates (Heimonen, 2002). Moreover, Yang (2012) combined the CAPM and cointegration to explain how benchmark markets are integrated with the global market. Over the decades, researchers have found the study of integration to be a useful approach to the study of the behaviour of inter-markets.[3] To illustrate the concept of integration, we note that markets are integrated when investors can pass from one market to another at no extra cost and when possibilities for arbitrage ensure the equivalence of share prices in both markets (Jawadi and Arouri, 2008). Early papers, seeking to demonstrate integrated markets, proposed techniques such as correlation tests to explain short-run portfolio diversification (Solnik, 1974; Longin and Solnik, 1995).

Nevertheless, in reviewing the existing literature, we found that most studies examined the integration of world stock markets only in a linear framework, using correlation tests as a tool of data analysis. Examples include Hamao, Masulis, and Ng (1990) and Markellos and Siriopoulos (1997). Hence, some researchers have confirmed the existence of relationships using the GARCH model to explore co-movements[4] among stock markets (Illueca and Lafuente (2002); Chouliaras, Christopoulos, Kenourgios, and Kalantonis (2012); Da Fonseca (2013) and Lee and Mercurelli (2014)), assuming that positive and negative error terms have symmetric effects on volatility. In more recent times, some researchers have utilized a variance of cointegration technique, specifically, fractional cointegration. For example, Caporale et al. (2015) use this technique to analyse linkages among US and European markets. They indicate that shocks that affect long-run relationships vanish at a very slow rate. Gagnon, Power, and Toupin (2016) also use this method to study the cointegration of risk-neutral moments of five major stock markets in Europe, showing that there is strong financial integration and concluding that such integration is partial when anticipations are considered.

[2] Lo and MacKinlay (1988) examined the predictability of time series by comparing the variances of differences in the data calculated over different intervals.

[3] Henceforth, we consider the relationships denoted by inter-markets to be the relationships among markets.

[4] Forbes and Rigobon (2002) explained co-movement as contagion, i.e., as a significant increase in cross-market linkages after a shock to one country or group of countries.

2.2.1 Empirical cointegration approach for the stock market analysis

This section explores the targets of the cointegration analysis that has been applied to stock markets. Research into integration and cointegration has employed several techniques, such as unit root tests of Dickey and Fuller (1979, 1981), used to establish the order of integration. Although in these papers, the authors provide one of the most influential works in the field of unit root tests, the test has low power because long memory processes cannot be explained by this test (Caporale et al., 2015). Subsequently, the cointegration of the variables was analysed, using the multivariate cointegration test of Johansen (1988, 1991), which enables testing of the cross-country market efficiency hypothesis. The Johansen cointegration test is used to show common stochastic trends across stock markets, and for this purpose, this test affords more robust results than other cointegration tests when there are more than two variables (Gonzalo, 1994). According to this idea, since the seminal paper of Kasa (1992), who studied the financial integration of five developed markets, applying common stochastic trends in these series. As a consequence, this methodology has led to numerous studies that find long-run co-movements between international stock markets, using univariate or multivariate cointegration models—for instance, Kenourgios, Samitas, and Paltalidis (2009), Yang, Kolari, and Min (2003) and Tian (2007).

Stock market analysis has been applied to different regions of the world, but most relevant studies have focused on the USA and Europe and their relations. Many strands of research, using cointegration tests, have obtained mixed results regarding market relationships. One strand focuses on US stock markets; Gil-Alana, Cunado, and Perez de Gracia (2013) observed very similar patterns in US stock markets for daily prices during the 1971–2007 period. Granger and Hyung (2004) and Mikosch (2000), using different techniques, explained the cointegration through structural breaks, showing long memory dependence. Conversely, Alvarez-Ramirez, Alvarez, Rodriguez, and Fernandez-Anaya (2008) demonstrated a shift in long-term behaviour—that is, a random walk. Additionally, empirical studies of relationships among international stock markets have focused on the United States. For example, Francis and Leachman (1998) and Richards (1995) both examined the existence of cointegration relationships between the developed European and U.S. markets. The first demonstrated long-run equilibrium among markets, whereas the second showed that national return indices are not cointegrated. Caporale et al. (2015) used fractional cointegration to find linkages between US and European stock markets, contrasting different recovery paths due to monetary policy pursued in the two economies. Studies have also shown relations between US or European markets and Asian markets. For example, Wong, Agarwal, and Du (2004) utilized fractional cointegration, reporting linkages between India, the USA, the UK and Japan. While this approach is extensively used in the literature, another strand in the literature focuses on stock markets within Europe. Taylor and Tonks (1989) and Corhay, Rad, and Urbain (1993) found strong evidence for cointegration among several major European stock markets in the late 1970s and 1980s. In an international context, Bessler and Yang, 2003 sought to demonstrate interdependence among nine major stock markets, finding that they are not fully integrated, and Darrat and Zhong (2005) studied cointegration between NAFTA countries, showing stable long-run linkage between the three stock markets. In addition, Kasa (1992) noted a common stochastic trend in the equity index prices of five developed countries, while Dickinson (2000) found that a cointegrating relationship between the major European stock markets exists and may be partly driven by the long-run relationships of macroeconomic fundamentals among these countries, possibly through indirect channels of international interaction.

Overall, a growing literature is emerging, one that seeks to explain the process of market integration due the convergence, using cointegration and taking into account endogeneity issues (Chouliaras et al., 2012; Syriopoulos, 2007; Bley, 2009; Mylonidis and Kollias, 2010; Lee and Mercurelli, 2014) and/or structural breaks (Kim, Moshirian, and Wu, 2006; Demian, 2011; Karmann and Ludwig, 2014). However, Da Fonseca (2013), using a VAR model, demonstrated that the major stock markets in the Euro area were not perfectly integrated during the first decade of the EMU. In sum, this technique provides a mode of demonstrating different ways of explaining market integration in different contexts. Caporale et al. (2015) recently showed that cointegration has also been used to determine whether there are diversification benefits from investing in different stock markets.

If cointegration does not hold, markets are not linked in the long run, and therefore, it is possible to gain from diversification. For this reason, testing for cointegration and any changes over time in its degree is important. For example, Richards (1995) demonstrated a lack of cointegration among various stock markets and hence the existence of diversification benefits for investors. From a theoretical perspective, applying the fractional cointegration technique (FCVAR model), which is an expansion of the CVAR approach (see Johansen, 1995), is adequate to provide more information about the cointegrating rank, the adjustments of the coefficients and long-run relationships among different variables—which in the present case are financial markets (see, Gagnon et al. (2016)).

2.3 Methodology

Our econometric strategy involves analysis of stock price data at monthly frequency. Once we have our model estimation, we perform statistical tests of cointegration, exclusion and weak exogeneity. We then apply the Bai and Perron (2003) test for structural breaks and use the FCVAR model to examine each break detected.[5]

2.3.1 Fractional cointegration model: FCVAR methodology

Our objective is to study the interdependence of the major Euro stock markets. In this paper, the FCVAR model allows us to study the common long-run equilibrium relationship between market indices. The model is a generalization of Johansen's (1995) cointegrated vector autoregressive (CVAR) model to allow for fractional processes of order d that co-integrate to order $d-b$. This model has the advantage of being used for stationary and non-stationary time series. This model is presented in Johansen (2008a, 2008b) and further developed in Johansen and Nielsen (2012) and Nielsen and Popiel (2016), and is gaining traction in finance (Bollerslev, Osterrieder, Sizova, and Tauchen (2013) and Gagnon et al. (2016)).

To introduce the FCVAR model, we begin with the well-known, non-fractional, CVAR model. Being $Y_t = 1, \ldots, T$ a p-dimensional $I(1)$ time series. So, the CVAR model is:

$$\Delta Y_t = \alpha\beta' Y_{t-1} + \sum_{i=1}^{k} \Gamma_i \Delta Y_{t-i} + \varepsilon_t = \alpha\beta' L Y_t + \sum_{i=1}^{k} \Gamma_i \Delta L^i Y_t + \varepsilon_t \tag{2.1}$$

The fractional difference operator introducing persistence in the model is Δ and the fractional lag operator is $\Delta = (1-L)$. Replacing lags operators in by their fractional counterparts Δ^b and $\Delta^b = (1-L^b)$, we obtain:

$$\Delta^b Y_t = \alpha\beta' L_b Y_t + \sum_{i=1}^{k} \Gamma_i \Delta^b L_b^i Y_t + \varepsilon_t, \tag{2.2}$$

we apply to $Y_t = \Delta^{d-b} X_t$, such that:

$$\Delta^d X_t = \alpha\beta' L_b \Delta^{d-b} X_t + \sum_{i=1}^{k} \Gamma_i \Delta^d L_b^i X_t + \varepsilon_t. \tag{2.3}$$

As always, ε_t is p-dimensional independent and identically distributed with mean zero and covariance matrix Ω. The parameters α and β are $p \times r$ matrices, where $0 \leq r \leq p$. In matrix β the columns are the cointegrating relationships and $\beta' X_t$ are the stationary combinations, i.e., the long-run equilibrium. We follow the assumption derived from the seminal paper of

[5] An alternative to our application is to take into account structural breaks, aiming to control the dynamics. As suggested by Johansen (2014), in practice, it is important to check the breaks in the dynamics. From this perspective, Hansen and Johansen (1999) proposed the theory of recursive estimation in the standard cointegration model.

Kasa (1992) about linearity in the relationship. However, on this linearity in our approach, once we are subject to this condition, seeks the study of changes in the behavior of the series through the analysis of structural breaks proposed by Bai and Perron (2003) as above mentioned, which allows us measure possible non-linearity in the time horizon of the relationship. The coefficients in α correspond the speed of adjustment unto equilibrium. Therefore, $\alpha\beta'$ is the adjustment long-run and Γ_i represents the short-run behavior of the variables.

Considering $d = b$ as an assumption of no persistence in the cointegration vectors and a constant mean term for the cointegrating relations, we reach an intermediate step before the final model. That is:

$$\Delta^d X_t = \alpha \left(\beta' L_d X_t + \rho'\right) + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta^d L_d^i X_t + \varepsilon_t. \tag{2.4}$$

We consider the simple model as:

$$\Delta^d (X_t - \mu) = L_d \alpha\beta' \left(X_t - \mu\right) + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta^d L_d^i (X_t - \mu) + \varepsilon_t, \tag{2.5}$$

where the variable μ is a level parameter that shifts each of the series by a constant in the way to avoid the bias related to the starting values in the sample (Johansen and Nielsen, 2016). $\beta'\mu = \rho'$ defines the mean stationary cointegrating relations. Johansen and Nielsen (2012) show that the maximum likelihood estimators $(d, \alpha, \Gamma_i, \ldots, \Gamma_k)$ are asymptotically normal and the maximum likelihood estimator of $(\beta,\, \rho)$ is asymptotically mixed normal.
For testing the hypotheses on the model parameters we use FCVAR model which is almost equal to CVAR (Johansen, 1995). We test if a market is a part of a cointegrating relationship and is included in a long-run equilibrium. Hypotheses on β can be formulated:

$$\beta = H\varphi, \tag{2.6}$$

where H is a matrix of dimension $p \times s$ and contains the restrictions and φ is a matrix of free parameters with dimension $s \times r$. The degrees of freedom are given by $df = (p - s)r$. If $r > 1$, the degrees of freedom of the test is $df = \sum_{i=1}^{r} (p - r - s_i + 1)$ (Jones, Nielsen, and Popiel, 2014).

With the test of hypotheses α, we test the weak exogeneity as:

$$\alpha = A\psi, \tag{2.7}$$

where A is a matrix of dimension $p \times m$ and ψ is a $m \times r$ matrix of free parameters with $m \geq r$ (Jones et al., 2014). The degree of freedom of the test is given by $df = (p - m)\, r$. If a row of α is zero, the associated variable is weakly exogenous. Note that matrix α and β are normalized separately in the same way for the CVAR model because the degrees of freedom are non-standard.

To sum up, by estimating the FCVAR model, we extract richer information from what was mentioned in previous sections. Importantly, by separately parameterizing the long-run and the short-run dynamics of the series, the model is able to accommodate empirically realistic I(d) long-memory and their fractional cointegration, while maintaining that the returns are $I(0)$ (Bollerslev et al., 2013).

2.4 Empirical analysis

2.4.1 Data description

For our empirical analysis, we use a sample of closing stock market prices of the four major stock markets of the Eurozone, namely, Germany (DAX), France (CAC), Spain (IBEX) and Italy (FTSE MIB). The data are collected from Yahoo! Finance. Our series are monthly and run from January 1998 to November 2016 (amounting 227 observations). Our analysis begins after converting all series to natural logarithms.

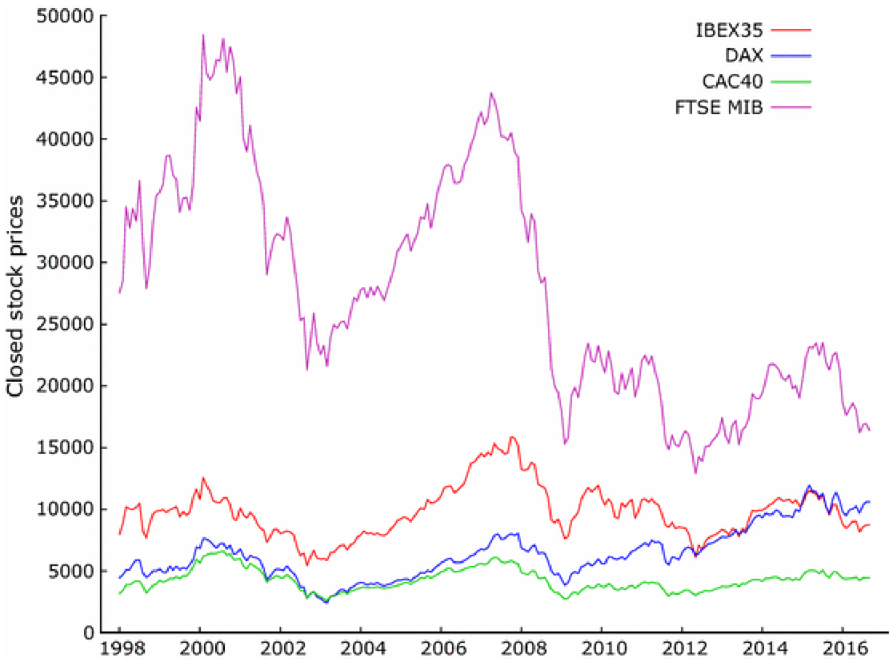
In Table 2.1 and Fig. 2.1, we present descriptive statistics and the dynamics of our series. The descriptive statistics associated with the closing prices of each index, shown in Table 2.1, reveal that the FTSE MIB index has the highest volatility, while the CAC40 has the lowest, and IBEX and DAX have similar volatility coefficients. For its part, Fig. 2.1 presents the time series dynamics for all indices in terms of how the series move; a common trend emerges among the monthly closing prices of these indices.

TABLE 2.1: Descriptive statistics for the options data

	DAX	**CAC 40**	**IBEX 35**	**FTSE MIB**
Mean	6440.8	4283.2	9797.4	27678.0
Median	6123.3	4229.4	9741.5	25919.0
Min	2423.9	2618.5	5431.7	12874.0
Max	11966.0	6625.4	15890.0	48479.0
SD	2131.0	891.35	2071.5	9137.1

Note: From 01/1998 to 11/2016

FIGURE 2.1: Time series plot for closing stock market prices of the four stock markets.



2.4.2 Testing for fractional cointegration

This section analyses the fractional cointegration of two paths: Univariate analysis is presented as an introduction to the second, multivariate analysis.

Univariate analysis

To determine whether the FCVAR model is appropriate to our data, we examine each of our series individually before conducting the multivariate analysis. In general, if both stationarity tests and unit root tests of a time series are rejected, that implies that the time series is likely a fractional time series. Therefore, before obtaining estimates of d, we perform augmented Dickey–Fuller (ADF) and Ng–Perron (2001) tests for unit roots on each of our individual series. The results are shown in Table 2.2. All tests reject stationarity, and tests of stock markets do no reject the presence of a unit root.

Table 2.2: Ng–Perron and Augmented Dickey–Fuller unit root tests for the stock markets

	Parameter	**DAX**	**CAC**	**IBEX**	**FTSE**
Ng - Perron	$\overline{M}Z_{\alpha}^{GLS}$	7.079	5.552	7.458	8.617
	$\overline{M}Z_{t}^{GLS}$	1.854	1.166	1.195	2.046
	$\overline{M}SB^{GLS}$	0.262	0.300	0.257	0.237
	$\overline{M}PT^{GLS}$	12.919	16.411	12.256	10.687
ADF	Statistic	1.891	2.313	2.457	2.236
Critical values (%)	Ng–Perron				ADF
	$\overline{M}Z_{\alpha}^{GLS}$	$\overline{M}Z_{t}^{GLS}$	$\overline{M}SB^{GLS}$	$\overline{M}PT^{GLS}$	α
1	23.800	3.420	0.143	4.030	-3.999
5	17.300	2.910	0.168	5.480	-3.413
10	14.200	2.620	0.185	6.670	-3.139

The critical values for the Ng–Perron test are tabulated in Ng and Perron (2001). The *MAIC* information criteria is used to select the autoregressive truncation lag, k, as proposed in Perron and Ng (1996)
$***$ Rejects null hypothesis at 1% significance level
$**$ Rejects null hypothesis at 5% significance level
$*$ Rejects null hypothesis at 10% significance level

There are several procedures for estimating the fractional differencing parameter in semiparametric contexts. Although the semiparametric log-periodogram regression proposed by Geweke and Porter-Hudak (1983) is the most used, this method was modified and further developed by Robinson (1995) and has been analysed by Velasco (1999) and Shimotsu and Phillips (2002), among others. Next, we proceed to the estimation of the fractional parameter d for each univariate series, with results presented in Table 2.3. The first three columns are semiparametric log-periodogram regression estimates from Geweke and Porter-Hudak (1983), here labelled GPH, computed with bandwidths $m = T^{0.4}$, $m = T^{0.5}$, and $m = T^{0.6}$, respectively.[6]

Table 2.3: Univariate analysis

	GPH estimates		
	$m = T^{0.4}$	$m = T^{0.5}$	$m = T^{0.6}$
	$\hat{d}$	$\hat{d}$	$\hat{d}$
DAX	1.051	1.056	1.165
	(0.223)	(0.108)	(0.132)
CAC 40	0.912	0.909	0.968
	(0.555)	(0.260)	(0.153)
IBEX 35	1.021	1.000	0.941
	(0.383)	(0.212)	(0.128)
FTSE MIB	1.189	1.050	1.170
	(0.169)	(0.168)	(0.195)

Note: GPH denotes the Geweke and Porter-Hudak semiparametric log-periodogram regression estimator. Standard errors are given in parenthesis beneath estimates of d. The sample size is 227

[6] In order to test the presence of unit roots, the estimates were obtained using first-differenced data, because the original series might be above 0.5 and this test requires that the results are limited to the interval $-0.5 < d < 0.5$, then adding 1 to obtain the proper estimates of d.

Statistical and hypothesis test

First, we determine the number of stationary cointegrating relations, following the hypotheses of the rank test based on a series of LR tests: $H_0 : rank = r$, against the alternative: $H_1 : rank = p$ for $r = 0, 1, \ldots$ (See Johansen, 1995).

The LR test statistics are provided in Johansen and Nielsen (2012), and the P values are available from MacKinnon and Nielsen (2014), based on their numerical distribution functions. The estimated rank is the first non-rejected value of the test, and when this rank is different from zero, we can also conclude that there exists a long-run equilibrium in the stock markets.

Once the rank cointegration test is established, we estimate the model parameters, using several hypothesis of interest[7] (Table 2.4). The first hypothesis is H_1^d , which examines whether fractional integration is more appropriate than traditional cointegration. The null hypothesis is $d = 1$, and its rejection implies that the FCVAR model is more suitable than a CVAR model. The remaining hypotheses can be divided into tests of a cointegrated relationship (β parameters) and tests for weak exogeneity of the variables (α parameters). The parameters in α and β are not identified without additional normalization restrictions; see Johansen (1995).

Table 2.4: Key for hypothesis test

H_1^d	The fractional parameter, d, is equal to one
H_1^β	FTSEMIB index does not enter the cointegrating relation(s)
H_2^β	IBEX 35 index does not enter the cointegrating relation(s)
H_3^β	CAC 40 index does not enter the cointegrating relation(s)
H_4^β	DAX index does not enter the cointegrating relation(s)
H_1^α	FTSEMIB index is weakly exogenous
H_2^α	IBEX 35 index is weakly exogenous
H_3^α	CAC 40 index is weakly exogenous
H_4^α	DAX index is weakly exogenous

Our primary interest in the cointegrating vectors concerns whether our variables form a stationary long-run equilibrium. The hypotheses H_1^β, H_2^β, H_3^β, H_4^β are used to test whether a given stock market is part of a cointegrating relationship and existing long-run equilibrium. If we reject these hypotheses, we can conclude that a long-run equilibrium relationship does not exist. The hypotheses H_1^α, H_2^α, H_3^α, H_4^α are used to test whether each variable is individually weakly exogenous. If a row of is zero, the variable does not respond to disequilibrium in the relationship. A rejection of the null hypothesis implies that a market index adjusts towards the long-run equilibrium after a shock.

Multivariate analysis

To complete our econometric strategy, we apply a multivariate analysis that allows us to estimate the possible relations among the variables used and test the different hypotheses. At the same time, the univariate analysis provides the value of the fractional integer. In this sense, Table 2.5 presents the estimation results for the FCVAR model applied to stock market prices. The null hypothesis of standard cointegration H_1^d is rejected with a P value of 0.000, suggesting that a fractional cointegration model is more appropriate. First, to establish the lag selection, we apply BIC criteria (see the "Appendix", Table A.1), selecting a lag length of one. To determine whether there is a long-run relationship among the stock markets selected, we test the cointegration rank (see Table A.2 in the "Appendix") before testing the hypotheses and find that the number of cointegrating vectors is three. We test hypotheses H_1^β, H_2^β, H_3^β, H_4^β to verify that our variables are in the cointegrating relations, using the 10% level of significance to reject a given null hypothesis (Jones et al., 2014). The results

[7]Hypothesis testing is explained in paragraph 3, Methodology

presented for β confirm that we strongly reject the null hypothesis of the non-existence of a long-run equilibrium, with a P value of 0.000, except in the cases of the FTSE MIB and IBEX 35, which do not share a long-run relationship. Indeed, stock markets that are cointegrated have a long-run relationship, so long-run correlations are higher than short-run correlations. If n variables have p cointegrating relationships, they have $n-p$ common trends. When $n-p=1$, as in the case studied, the individual stock markets are completely and perfectly integrated. Moreover, the test of weak exogeneity suggests that the selected stock markets are not weakly exogenous.[8]

TABLE 2.5: Estimated result for FCVAR

Lags	**1**		
Coint. relation (β)	1	2	3
FTSE MIB	1.000	0	0
IBEX 35	0	1.000	0
CAC 40	0	0	1.000
DAX	0.380	1.143	-0.612
Adjustment matrix (α			
FTSE MIB	-0.169	0.008	0.014
IBEX 35	-0.129	-0.002	0.091
CAC 40	0.082	-0.025	-0.046
DAX	0.323	0.073	0.308
Hypothesis test	df	LR statistics	P value
H_1^d	1	25.422	0.000
H_1^β	3	4.798	0.187
H_2^β	3	3.719	0.237
H_3^β	3	27.186	0.000
H_4^β	3	97.504	0.000
H_1^α	3	17.904	0.000
H_2^α	3	31.168	0.000
H_3^α	3	9.730	0.021
H_4^α	3	30.797	0.000

The top part of the table indicates the optimal number of lags representing the short run dynamics and the estimations of β and α as well as their associated standard error in parenthesis. The bottom part of the table reports the P values for the test of exclusion and weak-exogeneity tested in the Hypothesis test. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion and weak-exogeneity tests. The sample size is 227

2.4.3 Testing the fractional cointegration by structural breaks

We consider the possibility that the existence of structural breaks would provide a better empirical description of the European market integration. We now apply the test for structural breaks proposed by Bai and Perron, 2003 with a 15% trimming, which limits the maximum number of breaks allowed under the alternative hypothesis to 3. Among the breaks identified, the first regime (1998:01 until 2001:04) is in the way to the introduction of the single currency thus the markets were regulating to the new financial context. The second regime (2001:05–2007:06) would correspond to the economic growth and expansion period of the countries of the stock markets selected. In the third regime (2007:07 until 2012:04), according to the European Area Business Cycle Dating Committee, there was the financial crisis and the sovereign debt crises. Finally, the fourth regime (2012:05–2016:11) would be the end of the sovereign crisis until today. Tables 2.6, 2.7, 2.8, 2.9 and 2.10 shows the results for each regime.

[8]If a stock market is weakly exogenous, anticipations in this stock market do not adjust to shifts in anticipations for other markets.

Table 2.6: Bai-Perron tests of multiple structural changes in the relationship between the European stock markets

Statistics						
UDmax	WDmax	$\text{SupF}_t\,(1)$	$\text{SupF}_t\,(2)$	$\text{SupF}_t\,(3)$	$\text{SupF}_t\,(4)$	$\text{SupF}_t\,(5)$
256.711***	493.187***	125.278***	246.153***	221.508***	214.213***	256.711***
$\text{SupF}_t\,(2/1)$	$\text{SupF}_t\,(3/2)$	$\text{SupF}_t\,(4/3)$	$\text{SupF}_t\,(5/4)$			
231.393***	42.498***	44.156*	13.411			
Break dates estimates						
T_1		2001:4			[2000:03–2001:11]	
T_2		2007:6			[2007:05–2007:10]	
T_3		2012:4			[2012:01–2012:05]	

*, **, and *** denote significance at the 10, 5 and 1% levels, respectively. The critical values are taken from Bai and Perron (1998), Tables 1 and 2; and from Bai and Perron (2003), Tables 1 and 2. The number of breaks has been determined according to the sequential procedure of Bai and Perron (1998), at the 1% size for the sequential test. 90% confidence intervals for T_1 in square brackets

Once the structural breaks are defined, we proceed to use the FCVAR model to test each regime for cointegration (see tables A.3 to A.6 in appendix) and weak exogeneity. As can be seen in Table 2.7, the P value indicates that the null hypothesis of standard cointegration is rejected, suggesting that a fractional cointegration model is more appropriate. Applying the rank test (which is at most three), the number of cointegrating vectors is three; in other words, DAX, CAC 40, IBEX 35 and FTSE MIB are fully integrated. In view of the Hypothesis test, the results confirm a long-run equilibrium relationship among these variables. Based on the weak-exogeneity test, we accept the null hypothesis, with the IBEX 35 index and the CAC 40 index having P values of 0.109 and 0.205, respectively. Indeed, anticipations in these stock markets do not adjust to shifts that occur in the long-run relationship. The empirical results suggest that some linkage has existed over time, i.e., there is strong integration among the selected stock indices.

Turning to the second regime, Table 2.8 shows the results of the FCVAR model. It is observed that the null hypothesis of standard cointegration is strongly rejected. The behaviours of the cointegrating vectors match the results of the model applied to the original time series; we choose one lag to test the rank of the cointegrating vectors, finding three. Testing the β hypotheses, we determine that the null hypothesis of the non-existence of a long-run equilibrium is rejected in all cases, and we also reject the hypothesis of weak exogeneity. In sum, in this regime, the cointegrating vectors exhibit the same behaviour as in the original sample, implying that the stock indices are fully and perfectly integrated.

For the third regime (Table 2.9), which corresponds to the financial and European sovereign debt crisis period, we also strongly reject the null hypothesis of standard cointegration, with a P value of 0.000. Additionally, using the rank test, we find that there are two cointegrating vectors. Therefore, following Kasa (1992), the market integration is neither complete nor perfect. An explanation of this result is that this was a convulsive and uncertain period, and as we can see, the IBEX 35 index does not belong to the long-run relationship, perhaps owing to the observed integration weakness. Thus, the weak-exogeneity test shows that all markets adjust to shifts in anticipation of other markets. With respect to the IBEX 35 index, we appreciate that unless this market is not in the long-run relation, it is affected by such a relationship.

To complete our review of the regimes, the application of the FCVAR model to the fourth regime is shown in Table 2.10. First, as we have done previously, we test the hypothesis of standard cointegration, which is strongly rejected, with a P value of 0.000. Then, we test the rank of the cointegrating vector, finding three, which means that once the sovereign debt crisis ended, Euro market integration again became complete. In the case of the weak-exogeneity test, we observe that in none of the cases of the selected markets is the null hypothesis rejected, which means that anticipations in these stock markets do not adjust to shifts in the long-run relationship. The results obtained are similar to those for regime 2.

Table 2.7: Estimated result for FCVAR (Regime 1)

Lags	**1**		
Coint. relation (β)	1	2	3
FTSE MIB	1.000	0	0
IBEX 35	0	1.000	0
CAC 40	0	0	1.000
DAX	-1.497	-0.234	-1.609
Hypothesis test	df	LR statistics	P value
H_1^d	1	42.259	0.000
H_1^β	3	17.304	0.001
H_2^β	3	9.679	0.022
H_3^β	3	13.822	0.003
H_4^β	3	10.378	0.016
H_1^α	3	9.837	0.020
H_2^α	3	6.058	0.109
H_3^α	3	4.582	0.205
H_4^α	3	7.626	0.054

The top part of the table indicates the optimal number of lags representing the short run dynamics and the estimations of β and α as well as their associated standard error in parenthesis. The bottom part of the table reports the P values for the test of exclusion and weak-exogeneity tested in the Hypothesis test. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion and weak-exogeneity tests. The sample size is 41

Table 2.9: Estimated result for FCVAR (Regime 3)

Lags		**1**	
Coint. relation (β)		1	2
FTSE MIB		1.000	0
IBEX 35		0	1.000
CAC 40		1.842	1.963
DAX		0.719	0.146
Hypothesis test	df	LR statistics	P value
H_1^d	1	21.353	0.000
H_1^β	3	8.673	0.013
H_2^β	3	1.255	0.534
H_3^β	3	7.738	0.021
H_4^β	3	6.762	0.024
H_1^α	3	31.754	0.000
H_2^α	3	15.369	0.000
H_3^α	3	32.219	0.000
H_4^α	3	15.353	0.000

The top part of the table indicates the optimal number of lags representing the short run dynamics and the estimations of β and α as well as their associated standard error in parenthesis. The bottom part of the table reports the P values for the test of exclusion and weak-exogeneity tested in the Hypothesis test. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion and weak-exogeneity tests. The sample size is 58

Table 2.8: Estimated result for FCVAR (Regime 2)

Lags	**1**		
Coint. relation (β)	1	2	3
FTSE MIB	1.000	0	0
IBEX 35	0	1.000	0
CAC 40	0	0	1.000
DAX	2.263	-1.758	1.234
Hypothesis test	df	LR statistics	P value
H_1^d	1	29.503	0.000
H_1^β	3	15.874	0.001
H_2^β	3	20.799	0.000
H_3^β	3	22.958	0.000
H_4^β	3	52.133	0.000
H_1^α	3	39.118	0.000
H_2^α	3	17.714	0.001
H_3^α	3	36.883	0.000
H_4^α	3	38.889	0.000

The top part of the table indicates the optimal number of lags representing the short run dynamics and the estimations of β and α as well as their associated standard error in parenthesis. The bottom part of the table reports the P values for the test of exclusion and weak-exogeneity tested in the Hypothesis test. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion and weak-exogeneity tests. The sample size is 74

Table 2.10: Estimated result for FCVAR (Regime 4)

Lags	**1**		
Coint. relation (β)	1	2	3
FTSE MIB	1.000	0	0
IBEX 35	0	1.000	0
CAC 40	0	0	1.000
DAX	0.014	0.129	-0.694
Hypothesis test	df	LR statistics	P value
H_1^d	1	34.563	0.000
H_1^β	3	14.667	0.002
H_2^β	3	22.645	0.000
H_3^β	3	7.039	0.071
H_4^β	3	8.971	0.030
H_1^α	3	27.253	0.000
H_2^α	3	26.205	0.001
H_3^α	3	24.059	0.000
H_4^α	3	16.717	0.001

The top part of the table indicates the optimal number of lags representing the short run dynamics and the estimations of β and α as well as their associated standard error in parenthesis. The bottom part of the table reports the P values for the test of exclusion and weak-exogeneity tested in the Hypothesis test. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion and weak-exogeneity tests. The sample size is 54

2.5 Conclusion

In this paper, we have studied European stock market cointegration, using a fractionally cointegrated vector autoregressive (FCVAR) model applied to the closing prices of the major four stock market indices in the Eurozone. Despite controversy in the existing literature regarding treatment of this issue, the fractional cointegration model avoids most of the problems raised in the literature. Additionally, this model allows us to identify financial integration and weak exogeneity in our monthly time series.

Our equilibrium is characterized by three cointegrating vectors, which, following Kasa (1992), suggests that the individual stock markets are fully and perfectly integrated. However, to improve the analysis, we consider the existence of structural breaks, applying the Bai–Perron test and then testing the FCVAR model in each of four regimes—regimes that correspond to the introduction of the Euro currency, the financial crisis, the end of the sovereign debt crisis and a final period that runs through November 2016. The FCVAR model indicates some significant differences in patterns of convergence throughout the original sample as a function of the regime studied. The results for the different regimes show that, for the most part, integration of the European markets has been complete but also that, during the sovereign debt crisis, full integration of these indices disappeared. The reason for this development is that the IBEX 35 index went out of long-run equilibrium, which could mean that this index was more sensitive during this quarrelsome period, while the other markets were more robust—i.e., that the IBEX 35 index is the weak link in the integration. We therefore wish to emphasize the case of the Italian market (FTSE MIB), which, like the others, suffered from a sovereign debt crisis but, in contrast to the others, remained in the long-run relationship. Once this turbulent period ended, full Euro financial integration resumed, as we see in the fourth regime, although interest rates spreads, notably those of Italy, started to increase again in the second half of 2016. Financial integration is attributable to technological advances during recent decades, which has reduced transaction costs and allowed for greater access to information, notably reducing differences between national and international financial transactions. It has thus contributed to more sustainable economic growth.

The findings of the paper have important implications for investors and policy- makers. For investors, the high degree of integration implies a more attractive place for investment. However, this equilibrium also implies that portfolio diversification will be less effective. As stock market prices are interrelated, the possibility of strong impacts from external shocks is not reduced. In this line, cointegration is not the same as contagion. This is because cointegration may imply perfect spillover or, alternatively, no spillover at all if the variables are driven by a common third factor, which may be a global factor (Belke, Gros, and Osowski, 2017). For policy makers, market integration in the Eurozone has led to various debates. Market integration has increased competition and market efficiency and led to greater interdependence between the Eurozone markets; this may require increased supervision and securities market oversight, as Mylonidis and Kollias (2010) and Fratzscher (2002) find in their studies. Therefore, investors will prefer to invest in markets characterized by increasing growth, which will give them more investment options and risk diversification opportunities (e.g., buying stocks in two submarkets). There is thus potential gain through a focus on local rather than global factors. Future research into long-run relationships among the selected stock markets may focus on cycles to find possible synchronicity among markets. In addition, testing for breaks in the dynamics may be a new analytical approach to understanding the integration of markets. That is, future research could be oriented to the study of breaks in the dynamics of a Fractional Cointegration Approach, for instance, applying recursive estimation or rolling cointegration.

2.6 Appendix

Table A.1: Lag length selection

k	LR statistics	AIC	BIC
1	56.88	-3537.50	**-3410.78**
2	27.76	-3533.27	-3531.75
3	87.67	-3588.94	-3352.62
4	36.42	-3593.36	-3302.24
5	-2.01	-3559.35	-3213.43
6	108.55	-3635.90	-3235.18

The table shows lag length selection and bold indicates lag order selected. The sample size is 227

Table A.2: Cointegration rank test

Rank	Log-likelihood	LR statistics	*P*value
0	1779.254	52.996	0.060
1	1788.489	34.525	0.033
2	1792.018	27.468	0.003
3	1805.231	1.042	0.307
4	1805.752	—	—

The table shows the rank test. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is 227

Regime 1

Table A.3: Cointegration rank test (Regime 1)

Rank	Log-likelihood	LR statistics	*P*value
0	294.552	56.201	0.000
1	307.954	29.397	0.001
2	317.988	9.330	0.053
3	321.685	1.934	0.164
4	322.652	—	—

The table shows the rank test. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is 41

Regime 2

Table A.4: Cointegration rank test (Regime 2)

Rank	Log-likelihood	LR statistics	*P*value
0	583.557	84.570	0.000
1	598.455	54.774	0.000
2	617.983	15.718	0.003
3	625.788	0.109	0.741
4	322.652	—	—

The table shows the rank test. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is 74

Regime 3

Table A.5: Cointegration rank test (Regime 3)

Rank	Log-likelihood	LR statistics	*P*value
0	436.957	37.846	0.001
1	442.765	26.231	0.001
2	454.411	2.939	0.568
3	455.608	0.545	0.460
4	455.880	—	—

The table shows the rank test. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is 58

Regime 4

Table A.6: Cointegration rank test (Regime 4)

Rank	Log-likelihood	LR statistics	*P*value
0	442.787	51.266	0.000
1	444.708	47.425	0.000
2	462.616	11.608	0.020
3	467.517	1.806	0.178
4	468.420	—	—

The table shows the rank test. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is 54

Chapter 3

Long memory linkages among Latin American stock markets. A fractional cointegration approach

3.1 Introduction

Financial integration would benefit the region through more efficient allocation of capital, a higher degree of risk diversification and a more robust market framework Umutlu, Akdeniz, and Altay-Salih (2010). Furthermore, in developing countries, financial markets play an important role in economic development, for instance by acting on the saving rate or on the portion of savings channelled to investment with the latter leading to the formation of new markets (see Pagano (1993) and Greenwood and Smith (1997)).

Plenty of empirical studies have tested long-run relationships amongst the emerging financial markets and major developed markets, focusing on the extent to which stock markets are internationally integrated and, consequently, have important implications for the potential diversification of stock markets. Moreover, relationships between stock markets and market integration have been widely studied among different global economic regions. For instance, from an interregional point of view, focusing on The United States - European Union link (see Caporale et al. (2015), among others), and on intraregional markets, such as members of the European Monetary Union (Vides, Golpe, and Iglesias, 2018), market integration was evidenced in both cases. It also has been studied in other emerging regions such as The North American Free Trade Agreement (Lahrech and Sylwester, 2013), the Association of Southeast Asian Nations (Yu, Fung, and Tam, 2010) or Africans regions, highlighting Onyuma (2006) (for a survey) or Ncube and Mingiri (2015). There is also evidence for South America (see Diamandis, 2009; and Chuliá, Guillén, and Uribe (2017). Therefore, a novel table is elaborated in order to synthetize the studies regarding the market integration around the world. However, despite there not being many papers which propose a deepened research into Latin American financial integration, these papers could be underlined as examples of this topic, finding different results, showing different co-dependencies or integration degree using time series data (see Chen, Firth, and Rui (2002), Diamandis (2009), Romero-Alvarez, Atehortúa, and Guzmán-Aguilar (2013) or Chuliá et al. (2017), for instance).

Although the existing literature regarding the Latin American stock markets is limited; this paper proposes a wide review of the most relevant papers in the field of financial market integration by regions. In addition, it is intended to take a further step in the investigation of the long-term relationship using an expansion of cointegration, that is, fractional cointegration. Nevertheless, the aim of this paper is to study the possibility of a financial integration among the Latin American countries, and the possible expansion of the Latin American Integrated Market (MILA, here after) from a novel econometric perspective. This novelty is based on filling a gap in the literature of the Latin American integration regarding time series analysis of market cointegration. Indeed, this paper contributes to previous literature on the analysis of the integration of stock markets from a fractionally cointegrated vector autoregressive perspective. Although fractional cointegration had been used in previous studies, the approach developed by Johansen and Nielsen (2012) and further developed by Nielsen and Popiel (2016) is novel to the literature. This model, which is extended to allow for deterministic trends, has advantages when estimating a system of fractional time series variables that are potentially cointegrated. Additionally, the flexibility of the model allows for determining the number of equilibrium relations via statistical tests and jointly estimating the adjustment coefficients and cointegrating relations while accounting for short-run dynamics. Then, the choice of a stock market is based on the size of the respective national economy and the capitalization of the stock markets, which are the major ones in this region, so the markets used are MERVAL, from Argentina; BVSP, from Brazil; IPSA, from Chile; IGBC, from Colombia; and IPC, from Mexico and the sample covers the period of September 2004 to June 2019, amounting to 178 observations with a monthly frequency and converted in a common currency in order to alleviate exchange rate noise. Hence, equilibrium is found, which implies that portfolio diversification will be less effective. Although, financial integration among economies helps to improve their capacity to absorb shocks, the possibility of external shocks is not reduced because of stock market prices would be interrelated.

The remainder of the chapter is organized as follows. Section 3.2 provides a review of the literature, differentiating the studies of market integration in different economic regions in the world and subsequently on the application of the integration and cointegration test as a procedure. Section 3.3 presents the methodology applied. Section 3.4 discusses the empirical results, and conclusions are presented in section 3.5.

3.2 Literature review

Attending to the theoretical framework, the financial integration is the process whereby two or more countries or markets to become more connected to each other, acquiring a regional or global dimension whether these markets are closer to their neighbours or to global institutions, respectively. Hence, to be able to illustrate the concept of integration, Jawadi and Arouri (2008) explain that two or more markets are integrated if investors could pass from one market to another without paying any extra costs and when possibilities for arbitrage ensures the equivalence of share prices in both markets.

Despite there being a controversy about the treatment of the concept of financial integration, from a theoretical stand point, it may be signed by the convergence of assets' prices with the same characteristics. Thus, a perfect integration exists if similar assets have the same price even if they are traded on different markets however, it is necessary to attend the two main criteria to establish an adequate meaning of financial integration: First, any barrier or market access that does not allow the free movement of capital would constrain the integration so; the cross-border financial activity criterion is the first issue to consider. Second, it is possible that two or more markets can be open to each other and be imperfectly integrated due to them having different market structures, and then free access is not a sufficient condition for integration (IMF, 2016). In this case, the degree of convergence across markets is also essential in this issue. Nevertheless, from a practical point of view, this financial integration is always imperfect due to the particularities of each market or country such as the economic, technological and political factors (Salgado et al., 2015). So, it could say that the markets are segmented and local investors are restricted to investing in local markets and foreign investors are not allowed to invest in the local market (Bekaert and Harvey, 2003).

The rest of this section is devoted to exploring two main questions. Firstly, the importance of the cointegration as the methodology usually applied in this issue. Secondly, the empirical evidence on integration markets are deeply explained to detail the empirical puzzle generated around this topic. Finally, covering these two main objectives, a table is elaborated summarizing different empirical studies, methodologies and integration evidence by region.

3.2.1 Cointegration as an efficient way to test the market integration

The empirical evidence of market integrations by regions has been developed from multiple approaches (Lim and Brooks, 2011); however, we pay the attention to the cointegration to dissect the analysis that has been applied to stock markets. Research into integration and cointegration have employed several techniques, such as unit root tests of Dickey and Fuller (1979, 1981), used to establish the order of integration, although this test has low power because long memory processes cannot be explained by this test (Caporale et al., 2015). Consequently, the cointegration of the variables was analysed, using the multivariate cointegration test of Johansen (1988, 1991), which shows common stochastic trends across stock markets, giving more robust results than other cointegration tests where there are more than two variables (Gonzalo, 1994). According to this idea, since the seminal paper of Kasa (1992), who studied the financial integration of five developed markets, applying common stochastic trends in these series.

As it is shown before, the integration across the different regions in the world is an important topic and cointegration methodologies may be crucial for its testing and if cointegration does not hold, markets are not linked in the long-run, and so, it is possible to gain from diversification. For this reason, testing for cointegration and any changes over time is crucial.

Thus, the fractional cointegration is the next step of the cointegration techniques and it is possible to find determinant examples such as, Wong et al. (2004) who used fractional cointegration, reporting linkages among India, the USA, the UK and Japan. For its part, Caporale et al. (2015) used fractional cointegration to find linkages between US and European stock markets and determined that there are diversification benefits from investing in different stock markets.

From a theoretical perspective, applying the Fractionally Cointegrated Vector Autoregressive (FCVAR, here after) model developed by Johansen and Nielsen (2012) and Nielsen and Popiel (2016), is adequate to provide more information about the cointegrating rank, the adjustments of the coefficients and long-run relationships among different variables with different order of integration, i.e. the dichotomy $I(0)/I(1)$. This novel methodology is presented in a few studies about financial markets (see Bollerslev et al., 2013; Gagnon et al. (2016) or Vides, Golpe, and Iglesias (2018), for instance).

3.2.2 Empirical evidences concerning integrated markets

This body of market integration literature presents empirical evidence of market integration in different regions throughout the world. The most relevant examples are based on many papers about the European Monetary Union (EMU) and the European Union (EU) to test the long-term relationship (see Kim et al. (2006); Syriopoulos (2007); Bley (2009); Mylonidis and Kollias (2010); Demian (2011); Chouliaras et al. (2012); Da Fonseca (2013); Lee and Mercurelli (2014); Karmann and Ludwig (2014); and Vides, Golpe, and Iglesias (2018), for instance). In these papers, the market integration is supported and demonstrated with different techniques. Moreover, this region has been analysed with other regions of the world, trying to obtain evidence about long-run relationship among worldwide regions. For instance, highlighting studies based on the possible relationship between Europe and The United States of America (The USA, here after) to obtain mixed results regarding market relationship. For example, Gilmore and McManus (2002) and Caporale et al. (2015) tested the existence of long-run relationships between the developed European and U.S. markets.

Starting with one of the most studied regions in the world because it has the most studied and analysed country, i.e. The USA, this is focused on the North American Free Trade Agreement (NAFTA), Darrat and Zhong (2005), Aggarwal and Kyaw (2005) or Lahrech and Sylwester (2013) who studied the possibility of market integration between the North American Free Trade Agreement (NAFTA) countries, showing an increasing integration among the three stock markets, in the most of cases. Paying the attention to the Association of Southeast Asian Nations (ASEAN) and formed by Malaysia, Indonesia, Brunei, Vietnam, Cambodia, Laos, Burma, Singapore, Thailand and the Philippines, there are several studies which assess the integration among those countries that are part of this region. Thus, there are examples such as Yu et al. (2010) for instance, where they find evidence of long-run linkages in this region.

Considering the African continent, although there are different regions such as, mainly the East African Community (EAC), the Economic Community of West African States (ECOWAS), the Economic Community of Central African States (ECCAS), the Common Market for Eastern and Southern Africa (COMESA) and Southern African Development Community (SADC) among others, there are studies, which are focused on the possible linkages within different regions, showing different results depending on the region analysed. In spite of the recent interest in researching emerging markets, there is not much applied on African stock markets, finding disparate results. Wang, Yang, and Bessler (2003) find African stock markets have integration which appears to have declined after the 1997/98 crisis. Though, Onyuma (2006) (see for a survey) proposes reasons to consider a continental integration once regional alliances haven been established. Nevertheless, Ncube and Mingiri (2015) evidenced a fragmentation of African markets. For its part, Esso (2010) demonstrates evidence of long-run relationship among ECOWAS countries. However, Gebrehiwot and Sayim (2015) find that the level of financial market integration in the COMESA region is not significant, and most of the markets are fragmented.

Focusing on the Latin American markets, some authors have also made studies focused on exploring the existence of a possible financial integration in Latin America, finding controversy in the results. In these analyses, three important regions could be appreciated, i.e. the Bolivarian Alliance for the Peoples of Our Americas (ALBA), most of the countries of the Southern Common Market (Mercosur) and the Pacific Alliance (PA, here after), the latter being an alternative attempt of regional market integration in the Americas. The PA has triggered out reactions from ALBA countries, as well as from Brazil and some of its Mercosur partners. For Brazil, its concerns lie with losing control in its environment as Mexico tries to get importance in the region. Moreover, the PA increases the attraction forces in Mercosur. With this scenario, the development of the Latin American Integrated Market (MILA[1]) could be crucial in the growth and the possible integration in the region due to various PA plans for integration, the MILA initiative seeks to establish a unified capital market. The pioneer research applied in Latin American regions is Chen et al. (2002) where presented common features and linkages among six Latin American countries. Nevertheless, Hunter (2006) assesses the level of integration among Argentina, Chile and Mexico markets, concluding that these markets have not become integrated and they are segmented. Thus, Diamandis (2009) gives evidence of a partial integration among Argentina, Brazil, Chile and Mexico. For its part, Romero-Alvarez et al. (2013) explain that a possible financial integration would affect diversification benefits for investors of the member countries of the MILA. Bolanos, Burneo, Galindo, and Berggrun (2015) show negative results regarding the integrating process before the integration of Mexico to the MILA. More recently, Espinosa-Méndez, Gorigoitıa, and Vieito (2017) find long-term relationship among MILA members and Chuliá et al. (2017) show differences between Latin American markets. On the one hand, Chile and Colombia represent a good path in diversification meanwhile; on the other hand, Peru and Argentina present high co-dependences. In the table 3.1 is presented a deep summary of the literature about this topic, showing that around the 70% of the selected studies evidence a fully integration among stock markets using time series techniques, such as GARCH (and its variations), VAR (and its developments or variations), VECM, Fractional Cointegration and/or the study of the causality.

[1]For its Spanish initials.

TABLE 3.1: Summary of market integration studies

Region	Author/s	Data Sample	Countries	Technique	Evidence
European Monetary Union (EMU)	Kim et al. (2006)	Daily data from March 2,1994 to September 19, 2003	France, Germany, Italy and Spain and UK, Japan and The USA	GARCH model	Real economic integration and the reduction in currency risk have generally had the desired effect on financial integration
	Syriopoulos (2007)	Daily data from January 1, 1997 to September 20, 2003	Poland, Czech Republic, Hungary, Slovakia, Germany and The USA	Johansen's VAR model, VEC model and Granger Causality	The Central European markets follow a common path of growth, and become gradually more integrated with the international developed markets
	Bley (2009)	Daily data from January 1998 to September 20, 2006	Austria, Belgium, Germany, Spain, Finland, France, Greece, Ireland, Italy, Netherlands, Portugal, UK and The USA	Johansen's VAR model, VEC model and Granger Causality	Euro markets became more integrated between 1998 and 2003 but an ever increasing level of financial market integration in the Euro zone should not be taken for granted
	Mylonidis and Kollias (2010)	Daily data from January 4, 1999 to July 23, 2009	Germany, France, Spain and Italy	Rolling cointegration and Structural breaks	Although some convergence has taken place over time, it is still in the process of being achieved. Also, the German and French markets appear to be the ones with a higher degree of convergence while Germany presents the dominant position
	Demian (2011)	Daily data from January 1, 2001 to May 31, 2009	Hungary, Poland, Czech Republic, Slovakia, Estonia, Romania, Germany, France, UK and Italy	Johansen's VAR model, VECM and Granger Causality	EU accession plays a minor direct role in the development of these links, cointegration being driven more by financial and economic factors as opposed to explicit political actions
	Chouliaras et al. (2012)	Daily data from February 1, 2005 to June 30, 2011	Portugal, Italy, Ireland, Greece and Spain	Johansen cointegration, Granger causality, Gregory and Hansen residuals cointegration with regime shifts, fully modified ordinary least squares, and a multivariate GARCH model	An existence of cointegrating relationships among these stock markets while there are volatility spillovers between Greece and the rest of the countries

Table 3.1 continued from previous page

Region	Author/s	Data Sample	Countries	Technique	Evidence
	Da Fonseca (2013)	Daily data from January 1, 2001 to December 31, 2011	France, Germany, Holland, Italy and Spain	Johansen's VAR model and GARCH model	The five biggest stock markets of the Euro area have not been perfectly integrated during the first decade of the European Monetary Union.
	Lee and Mercurelli (2014)	Monthly data from January 1992 to June 2012	France, Germany and Italy	Structural VAR, sliding window technique and DCC models	The adoption of the euro has increased the symmetry of underlying shocks and accelerated the convergence process within this group.
	Karmann and Ludwig (2014)	Monthly data from January 1960 to October 2010	France, Germany and UK	Rolling Cointegration and Structural Breaks	There seems to be an increased interconnectedness of all three stock markets, i.e. there are almost identical reactions of these markets to shocks
	Vides, Golpe, and Iglesias (2018)	Monthly data from January 1998 to November 2016	France, Germany, Italy and Spain	Fractionally Cointegrated Vector Autoregressive (FCVAR) model	There is a perfect financial integration among these countries despite in the financial and sovereign debt crises this perfect cointegration disappears
EMU – The USA	Gilmore and McManus (2002)	Weekly data from July1, 1995 to August 1, 2001	Czech Republic, Hungary, Poland and The USA	Johansen (1988) Cointegration test and Granger Causality test	The results suggest that US investors can obtain benefits from international diversification into these markets.
	Caporale et al. (2015)	Monthly data from December 1986 to December 2013	The USA and EuroStoxx Index for EMU	Robinson's Whittle semiparametric approach	There is evidence of fractional cointegration over the subsample from December 1996 to March 2009, indicating that the effects of shocks affecting the long-run relationship vanish at a very slow rate
North American Free Trade Agreement (NAFTA)	Darrat and Zhong (2005)	Daily data from June 1, 1989 to April 10, 2002	Canada, Mexico and The USA	Johansen and Juselius Cointegration test and Variance Decomposition	The evidence proves robust and consistently indicates intensified equity market linkage since the NAFTA accord
	Aggarwal and Kyaw (2005)	Daily, weekly, and monthly data from January 1988 to December 2001	Canada, Mexico and The USA	Johansen's VAR model	US stock prices are more integrated with both Canadian and Mexican stock prices after the passage of NAFTA

Table 3.1 continued from previous page

Region	Author/s	Data Sample	Countries	Technique	Evidence
	Lahrech and Sylwester (2013)	Weekly data from December 30, 1988 to July 21, 2006	Canada, Mexico and The USA	DCC Models	NAFTA increased linkages between U.S. and Mexican equity markets and between Canadian and Mexican markets
Association of Southeast Asian Nations (ASEAN)	Yu et al. (2010)	Daily data from March 16, 1994 to December 19, 2008	Japan, Mainland China, Hong Kong, Taiwan, South Korea, Singapore, Malaysia, Thailand, Indonesia, and the Philippines	Dynamic Cointegration and Dynamic Conditional Correlation (DCC) model	The process is not complete and the degrees of integration between mature and emerging equity markets are different
Africa	Wang et al. (2003)	Weekly data from January 1, 1996 to May 31, 2002	South Africa, Egypt, Morocco, Nigeria and Zimbabwe and The USA	Johansen's VAR model	African stock markets have integration which appears to have declined after the 1997/98 crisis
	Onyuma (2006) (Survey)	Explains in a theoretical way how is the behaviour in the African regions			Proposes reasons to consider continental integration once regional alliances have been established
	Ncube and Mingiri (2015)	Monthly data from February 2000 to September 2008	South Africa, Botswana, Namibia, Mauritius and Nigeria and Germany, Japan and the USA	Johansen's VAR model and Granger Causality	There is a fragmentation of African markets
Economic Community of West African States (ECOWAS)	Esso (2010)	Yearly data from 1960 and 2005	Benin, Burkina Faso, Cape Verde, Cote d'Ivoire, Gambia, Ghana, Guinea, Guinea-Bissau, Liberia, Mali, Niger, Nigeria, Senegal, Sierra Leone, and Togo	Gregory and Hansen Threshold Cointegration and Toda-Yamamoto causality test	There is a long-run relationship between finance and growth and test for cointegration in presence of breakpoint
Common Market for Eastern and Southern Africa (COMESA)	Gebrehiwot and Sayim (2015)	Monthly data from January 2005 to December 2013	Kenya, Egypt, Madagascar, Mauritius, Malawi, Rwanda, Swaziland, Seychelles, Uganda, Zambia and The USA and China	Johansen's VAR model	The level of financial market integration in the COMESA region is not significant, and most of the markets are fragmented

Table 3.1 continued from previous page

Region	Author/s	Data Sample	Countries	Technique	Evidence
Latin America	Chen et al. (2002)	Daily data from February 1, 1995 to June 30, 2000	Argentina, Brazil, Chile, Colombia, Mexico and Venezuela	Johansen's VAR model	There is one cointegrating vector which appears to explain the dependencies in prices
	Hunter (2006)	January 1992 for Mexico, August 1993 for Argentina and January 1994 for Chile to December 1999	Argentina, Chile, and Mexico	GARCH	These markets have not become integrated and they are segmented
	Diamandis (2009)	Weekly data from January 1988 to July 2006	Argentina, Brazil, Chile and Mexico (MEX) and a mature market, New York Stock Exchange (US)	Johansen's VAR model	Although cointegration exists there are small long-run benefits from international portfolio diversification since the stock prices adjust very slowly to these common trends
Latin American Integrated Market (MILA)	Romero-Alvarez et al. (2013)	Daily data from December 2009 to June 2012	Chile, Colombia and Peru	CAPM and Principal Components Analysis (PCA)	A possible financial integration would affect diversification benefits for investors of the member countries of the MILA
	Bolanos et al. (2015)	Monthly data from November 2008 to August 2013	Chile, Colombia and Peru	Different measures of financial variables	Negative results regarding the integrating process before the integration of Mexico to the MILA
	Espinosa-Méndez et al. (2017)	Daily data from July 16, 2002 to July 29, 2016	Chile, Colombia, Mexico and The USA	DCC – GARCH model, Johansen's VAR model	Findings suggest that in an integration process such as MILA, as stock market members differ, in terms of stock market development, the markets will benefit from the integration. However, in the long term these benefits dissipate over time
	Chuliá et al. (2017)	Daily data from June 30, 1995 and June 30, 2015	Argentina, Brazil, Chile, Colombia, and Peru, United Kingdom, Canada, Germany, France, Italy and Japan and The USA	Multivariate quantile models (MVMQ) and Pseudo impulse - response functions (PIRFs)	On the one hand, Chile and Colombia represent a good path in diversification meanwhile, on the other hand, Peru and Argentina present high co-dependences

3.3 Methodology

Our objective is to study the interdependence of the major Euro stock markets. In this paper, the FCVAR model allows us to study the common long-run equilibrium relationship between market indices. The model is a generalization of Johansen (1995) cointegrated vector autoregressive (CVAR) model to allow for fractional processes of order d that co-integrate to order $d-b$. This model has the advantage of being used for stationary and non-stationary time series. This model is presented in Johansen (2008a, 2008b) and further developed in Johansen and Nielsen (2012) and Nielsen and Popiel (2016), and is gaining traction in finance (Bollerslev et al., 2013; Gagnon et al., 2016; or Vides, Golpe, and Iglesias, 2018).

To introduce the FCVAR model, we begin with the well-known, non-fractional, CVAR model. Being $Y_t = 1, \ldots, T$ a p-dimensional $I(1)$ time series. So, the CVAR model is:

$$\Delta Y_t = \alpha\beta' Y_{t-1} + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta Y_{t-i} + \varepsilon_t = \alpha\beta' L Y_t + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta L^i Y_t + \varepsilon_t \tag{3.1}$$

The fractional difference operator introducing persistence in the model is Δ and the fractional lag operator is $\Delta = (1-L)$. Replacing lags operators in by their fractional counterparts Δ^b and $\Delta^b = (1-L^b)$, we obtain:

$$\Delta^b Y_t = \alpha\beta' L_b Y_t + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta^b L_b^i Y_t + \varepsilon_t, \tag{3.2}$$

we apply to $Y_t = \Delta^{d-b} X_t$, such that:

$$\Delta^d X_t = \alpha\beta' L_b \Delta^{d-b} X_t + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta^d L_b^i X_t + \varepsilon_t. \tag{3.3}$$

As always, ε_t is p-dimensional independent and identically distributed with mean zero and covariance matrix Ω. The parameters α and β are $p \times r$ matrices, where $0 \leq r \leq p$. In matrix β the columns are the cointegrating relationships and $\beta' X_t$ are the stationary combinations, i.e., the long-run equilibrium. We follow the assumption derived from the seminal paper of Kasa (1992) about linearity in the relationship. The coefficients in α correspond the speed of adjustment unto equilibrium. Therefore, $\alpha\beta'$ is the adjustment long-run and Γ_i represents the short-run behavior of the variables.

Considering $d = b$ as an assumption of no persistence in the cointegration vectors and a constant mean term for the cointegrating relations, we reach an intermediate step before the final model. That is:

$$\Delta^d X_t = \alpha\left(\beta' L_d X_t + \rho'\right) + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta^d L_d^i X_t + \varepsilon_t. \tag{3.4}$$

We consider the simple model as:

$$\Delta^d (X_t - \mu) = L_d \alpha\beta' \left(X_t - \mu\right) + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta^d L_d^i (X_t - \mu) + \varepsilon_t, \tag{3.5}$$

where the variable μ is a level parameter that shifts each of the series by a constant in the way to avoid the bias related to the starting values in the sample (Johansen and Nielsen, 2016). $\beta'\mu = \rho'$ defines the mean stationary cointegrating relations.

In order to determine the number of stationary cointegrating relations following the hypotheses in the rank test based on a series of LR tests. In the FCVAR model, we test the hypothesis $H_0 : rank(\Pi) = r$, against the alternative: $H_1 : rank(\Pi) = p$. Being $L(d, b, r)$ the profile likelihood function is given a rank r, where (α, β, Γ) has been reduced by rank regression (see Johansen and Nielsen, 2012).Maximizing the profile likelihood distribution under both hypothesis, the LR test statistics are now $LR_t(q)$. The asymptotic distribution of $LR_t(q)$ depends on the parameter b and on $q = n - r$. MacKinnon and Nielsen (2014), based on their numerical distribution functions, provide asymptotic critical values of the LR rank test. In the case of weak cointegration, i.e., $0 < b < 1/2$, $LR_t(q)$ has a standard asymptotic distribution,$LR_t(q)$ $LR_t(q) \xrightarrow{D} \chi^2(q^2)$.

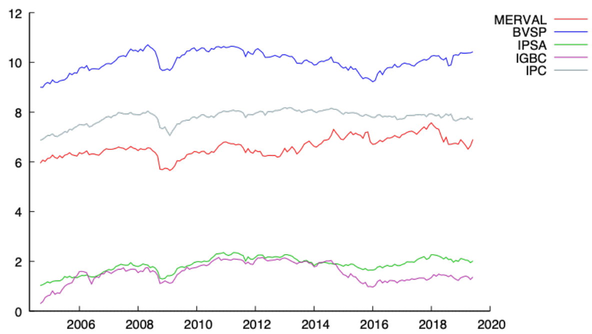
To sum up, by estimating the FCVAR model, we extract richer information from what was mentioned in previous sections. Importantly, by separately parameterizing the long-run and the short-run dynamics of the series, the model is able to accommodate empirically realistic $I(d)$ long-memory and their fractional cointegration, while maintaining that the returns are $I(0)$ (Bollerslev et al. 2013).

3.4 Data and results

3.4.1 Data description

For our empirical analysis, we use a sample of closing stock market prices of the five major stock markets of Latin America, namely, Argentina (MERVAL), Brazil (BVSP), Chile (IPSA), Colombia (IGBC) and Mexico (IPC). The data is collected from Yahoo! Finance in a monthly frequency and runs from September 2004 to June 2019 (amounting 178 observations). Following the example of Pukthuanthong and Roll (2009), a common currency was used to alleviate exchange rate noise as such conversions represent a ubiquitous practice in studies of international financial markets so, all the indices are in US dollars and dividend-adjusted. The analysis begins after converting all series to natural logarithms. In Figure 3.1 and Table 3.2, the dynamics of the series selected and the descriptive statistics are showed. Figure 3.1 shows the time series dynamics for all indices in terms of how the series move. In order to provide a better representation, the series are plotted in their logarithm values and, as it is shown, a common behaviour emerges among the logarithms of monthly closing prices of these indices.

FIGURE 3.1: Time series plot for the five major stock markets of Latin America.



For its part, the descriptive statistics associated with the closing prices of each index, shown in table 3.2, reveal that the BVSP index has the highest volatility, while the IPSA has the lowest. It also could be observed how the markets of Chile and Colombia present similar descriptive statistics.

As explained above, the main interest of this paper consists in testing the stock market integration in Latin America. With this aim, it is necessary to test some hypotheses under the requirements of the fractional cointegration (FCVAR model).

The next table 3.3 shows the procedure that is followed to check whether market integration is possible. The econometric exercise starts by studying the possibility that the fractional cointegration would be more appropriate than the traditional one. Once this is done, testing if

TABLE 3.2: Descriptive statistics for the market data (2004:09 - 2019:06)

	Mean	**Median**	**Min.**	**Max.**	**SD**
MERVAL	793.2	696.4	286.1	1941	322.5
BVSP	24811	24035	8062	44631	9178
IPSA	6.69	6.60	2.79	10.53	1.91
IGBC	5.03	4.67	1.35	8.64	1.88
IPC	2509	2630	962.9	3579	604.3

any market enters (or not) in the long-run relationship[2] is a necessary condition and then, the examination of the weak exogeneity[3] would give some evidence if any market is dominating the step 3.

TABLE 3.3: Strategy of empirical research

Procedure		**Hypotheses**
Step 1	Fractional cointegration?	H_1^d: Is the fractional cointegration more appropriate that traditional cointegration?
Step 2	Estimation of β	H_1^β: Does any market enter in the long-run relationship?
Step 3	Estimation of adjustment coefficients	H_1^α: Does any market show weak-exogeneity?

The next subsection is devoted to analysing whether fractional cointegration is allowed to be divided two ways. On the one hand, a univariate analysis is presented as an introduction to the second, the multivariate analysis, which would allow us to know if the series are cointegrated and allows the hypotheses testing.

3.4.2 Univariate analysis

Aiming to determine whether the FCVAR model is appropriate to the main purpose, each of the series is examined individually before conducting the multivariate analysis. In general, if both stationarity tests and unit root tests of a time series are rejected, that implies that the time series is likely a fractional time series. Despite there being several procedures for estimating the fractional differencing parameter in a semiparametric context. Though the semiparametric log-periodogram regression proposed by Geweke and Porter-Hudak (1983) is the most used, this method was modified and further developed by Robinson (1995) and has been analysed by Velasco (1999) or Shimotsu and Phillips (2002), among others. Then, we proceed to the estimation of the fractional parameter d for each univariate series, with results presented in Table 3.4. The first three columns are semiparametric log-periodogram regression estimates from Geweke and Porter-Hudak (1983), here labelled GPH[4], computed with bandwidths $m = T^{0.4}$, $m = T^{0.5}$, and $m = T^{0.6}$, respectively. The remaining columns in Table 3.4 present FAR (k) estimates with $r = 0$ and k lags, see Johansen and Nielsen (2010). Results are shown for $k = 0$, $k = 1$ and $k = 2$, and the associated Ljung-Box Q-test statistics, labelled $Q_{\hat{\varepsilon}}$, for serial correlation up to lag 12 in the residuals are also given.

As we can see, the GPH estimates have large standard errors, making it difficult to draw any conclusions but supporting the idea that the fractional cointegration could be appropriate

[2]The hypotheses H_1^β are used to test whether a given stock market is part of a cointegrating relationship and existing long-run equilibrium. If these hypotheses are rejected, we can conclude that a long-run equilibrium relationship does not exist.

[3]The hypotheses H_1^α is used to test whether each variable is individually weakly exogenous. If a row of α is zero, the variable does not respond to disequilibrium in the relationship. A rejection of the null hypothesis implies that a market index adjusts towards the long-run equilibrium after a shock.

[4]In order to test the presence of unit roots, the estimates were obtained using first-differenced data, because the original series might be above 0.5 and this test requires that the results are limited to the interval $-0.5 < d < 0.5$, then adding 1 to obtain the proper estimates of d. According to the literature, the bandwidth size spans from 0.25 to 0.8. In our study, the three bandwidths selected are 0.4, 0.5 and 0.6.

TABLE 3.4: Univariate analysis

	GPH estimates			FAR(k) estimates					
	$m = T^{0.4}$	$m = T^{0.5}$	$m = T^{0.6}$	$k = 0$		$k = 1$		$k = 2$	
	$\hat{d}$	$\hat{d}$	$\hat{d}$	$\hat{d}$	$Q_{\hat{\varepsilon}}$	$\hat{d}$	$Q_{\hat{\varepsilon}}$	$\hat{d}$	$Q_{\hat{\varepsilon}}$
MERVAL	0.256	0.691	0.739	1.008	9.111	0.675	8.093	0.382	8.498
	(0.196)	(0.237)	(0.145)	(0.073)	(0.693)	(0.123)	(0.778)	(0.073)	(0.745)
BVSP	0.771	0.761	1.040	1.061	6.542	0.951	5.850	0.291	6.144
	(0.372)	(0.184)	(0.186)	(0.065)	(0.892)	(0.243)	(0.923)	(0.039)	(0.909)
IPSA	1.434	1.108	1.093	1.054	11.407	1.062	11.346	0.312	11.827
	(0.560)	(0.267)	(0.197)	(0.064)	(0.494)	(0.140)	(0.499)	(0.037)	(0.460)
IGBC	1.144	1.155	1.056	1.103	7.785	1.079	7.750	0.278	9.494
	(0.215)	(0.173)	(0.138)	(0.063)	(0.802)	(0.122)	(0.804)	(0.032)	(0.660)
IPC	0.958	0.884	1.118	1.034	9.092	0.992	8.974	0.292	9.408
	(0.253)	(0.165)	(0.175)	(0.064)	(0.695)	(0.145)	(0.705)	(0.033)	(0.668)

GPH denotes the Geweke-Porter-Hudak semiparametric log-periodogram regression estimator and FAR(k) denotes the fractional AR model with $r = 0$ and k lags. $Q_{\hat{\varepsilon}}$ denotes the Ljung-Box Q-test statistic for the residuals, computed with 12 lags because monthly data is used. Standard errors are given in parentheses beneath estimates of d and P values are in parentheses beneath $Q_{\hat{\varepsilon}}$ tests. The sample size is T = 178.

for this issue. For the FAR (k) models show that the residuals are well behaved and the estimates of d are in line with or similar to those for the GPH estimates but their standard errors are lower.

3.4.3 Multivariate analysis

To complete our econometric strategy, a multivariate analysis is applied, which allows an estimation of possible relations among the series used and test the different hypotheses, which are abovementioned. At the same time, the univariate analysis provides the value of the fractional integer. In this sense, Table 3.5 determines the estimation results for the Fractionally Cointegrated Vector Autoregressive (FCVAR) model applied to Latin America stock market prices.

First, to specify the model Dolatabadi et al. (2016), Dolatabadi et al. (2018) are followed and, initially the lag selection is established by using the Bayesian Information Criteria (BIC) (see the appendix, Table A.1), selecting a lag length of one. To determine whether there is a long-run relationship among the stock markets selected, we test the cointegration rank before testing the hypotheses and find that the number of cointegrating vectors is four. Following Kasa (1992), if n variables have p cointegrating relationships, they have $n - p$ common trends. When $n - p = 1$, as occurred in this case, the individual stock markets are completely and perfectly integrated. Finally, the residuals appear well-behaved with no evidence of serial correlation; getting a P value of 0.412 (see table A.2 and A.3 in appendix).

TABLE 3.5: Cointegrating rank test

Rank	**Log-likelihood**	**LR statistics**	***P* value**
1	1189.512	48.366	0.000
2	1196.002	35.385	0.000
3	1204.056	19.278	0.001
4	1209.643	8.104	0.004
5	1213.695	—	—

The table shows FCVAR cointegration rank. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is T = 178.

The exercise continues in tables 3.6 and 3.7, testing the hypothesis of fractional cointegration versus standard cointegration. Attending to H_1^d, where is the convenience of FCVAR is checked, the LR test and its P value of 0.000 reveal that the null hypothesis of d is equal to 1 is rejected so, the FCVAR is more appropriate for this study. We test hypotheses H_1^β, H_2^β,

H_3^β, H_4^β and H_5^β (MERVAL, BSVP, IPSA, IGBC and IPC, respectively) in order to examine if the variables selected enter in the long-run relationship, using the 10% level of significance to reject a given null hypothesis as Jones et al. (2014) establish in their paper.

TABLE 3.6: CVAR vs. FCVAR (H_1^d)

Unrestricted log-likelihood	1209.643
Restricted log-likelihood	1201.345
LR statistic	16.595
P value	0.000

Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is T = 178.

TABLE 3.7: Hypothesis test

	H_1^β	H_2^β	H_3^β	H_4^β	H_5^β	H_1^α	H_2^α	H_3^α	H_4^α	H_5^α
LR	9.953	17.542	7.794	13.241	8.347	16.207	18.047	11.988	31.516	22.116
P value	0.041	0.002	0.099	0.010	0.080	0.003	0.001	0.017	0.000	0.000

Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is T = 178.

The results presented for hypotheses β confirm that we strongly reject the null hypothesis of the non-existence of a long-run equilibrium, with a P value around 0.000 in almost cases, which means that the stock markets selected share a long-run relationship. Indeed, stock markets that are cointegrated have a long-run relationship, so long-run correlations are higher than short-run correlations. Nevertheless, it should be noted the case of Chilean IPSA market (see H_3^β), which gets a P value very close to the 10% significance level. This may be possible due to Chile is a leader of foreign capital receipt, compared to other countries of the region, supporting the development of its financial markets. Furthermore, the estimate of common order of integration of five stock markets is 0.411 (see table A.2 in appendix), showing that the stochastic trend is fractionally nature and possess stationarity with long memory. Moreover, the test of weak exogeneity, which corresponds to hypotheses H_i^α, suggests that the given stock markets do not show weak exogeneity.

TABLE 3.8: Long-run fractional cointegration (β)

	1	**2**	**3**	**4**
MERVAL	1.000	0	0	0
BVSP	0	1.000	0	0
IPSA	0	0	1.000	0
IGBC	0	0	0	1.000
IPC	-1.917	-1.212	-2.313	-1.319

TABLE 3.9: Estimated speed of adjustment vector (α)

MERVAL	-0.234	-0.486	0.308	0.241
BVSP	-0.060	-0-448	0.181	0.137
IPSA	0.017	-0.204	0.092	0.199
IGBC	-0.127	-0.193	0.280	0.018
IPC	0.004	-0.299	0.096	0.362

Since there are four cointegrating vectors in this estimation, the adjustment dynamics are more complex. As an example, we consider again a one percentage point rise in the stock price that pushes the system out of equilibrium and again ignore the short-run dynamics in

Γ_i. Holding everything else constant, the effect on the equilibrium errors is (see tables 3.8 and 3.9):

$$u_{1t} = 1.917,\ u_{2t} = 1.212,\ u_{3t} = 2.313,\ u_{4t} = 1.319$$

where the magnitudes correspond to the coefficient on each variable in each equilibrium relation. The adjustment associated with (fractional differences of) MERVAL is:

$\alpha_{11}u_{1t} + \alpha_{12}u_{2t} + \alpha_{13}u_{3t} + \alpha_{14}u_{4t} = -0.234(1.917) - 0.486(1.212) + 0.308(2.313) + 0.241(1.319) = -0.008$

which is negative and implies that the series moves the system back towards equilibrium, i.e. it pushes u_{1t} back down towards zero. For BVSP, the adjustment is -0.059, which is also negative and pushes u_{2t} back down towards zero. For IPSA, the adjustment is 0.259, which is positive and pushes u_{3t} back up towards zero. Then, for IGBC, the series moves the system back towards equilibrium, being the adjustment 0.194 and pushes u_{4t} back up towards zero. Finally, for IPC, the adjustment is 0.344 and pushes u_{5t} back up towards zero.

3.5 Conclusion

In this paper, the stock market integration in Latin America has been checked, using a fractionally cointegrated vector autoregressive (FCVAR) model applied to the monthly closing prices of stock market indices in Latin America. Latin America is the focus of this study because of its rapid economic growth and its opening up as a market for foreign investors. The existing gap in the literature concerning this issue in different emerging regions in the world in general, and in Latin American in particular, this paper fills the gap proposing a new focus in the field, i.e. the fractional cointegration model which avoids most of the problems raised in the standard cointegration literature. Hence, this framework is more general than the standard approach based on the $I(0)/I(1)$ dichotomy, which is more restrictive.

First, it is shown that the five Latin America emerging stock markets are cointegrated. Although the relationship is characterized by four cointegrating vectors then, following Kasa (1992), the stock markets are fully and perfectly integrated and this perfect integration show a stationary, long memory behaviour. Attending to the long-run relationship and the absence of weak exogeneity of each stock market, as their prices are interrelated with themselves so, the possibility of exposure from external shocks is not diminished and spillovers may be implicit although financial integration among economies helps to improve their capacity to absorb those shocks.

The findings of the paper have important implications for investors and policy makers. For investors, there are better alternatives of financial instruments, enlarging the possibilities of investment and developing new portfolios, implying a more attractive and competitive place for investment since there is a financial integration. In fact, Hunter (2006) shows that the integration has some implications for foreign investors seeking international portfolio diversification by investing in the emerging markets, whose security prices are influenced by international factors. Thus, investors will prefer to invest in markets characterized by increasing growth, which will give them more investment options and risk diversification opportunities. For policy makers, the financial integration would contribute greater stability in the rules of the game and would allow each country to be more competitive and efficient in the region.

In sum, regional financial integration could foster the development of Latin American indices, which in turn may support the engineering of specific financial products Latin American; preserve financial experience and innovation in the region and regulatory expertise and surveillance of regional agents. Besides, larger regional markets are likely to attract extra-regional flows, thus promoting regional and global integration (IMF, 2016).

3.6 Appendix

TABLE A.1: Lag length selection

Lags	AIC	BIC
1	-2305.39	**-2111.30**
2	-2334.29	-2060.65
3	-2401.67	-2048.49
4	-2425.62	-1992.90
5	-2445.94	-1933.67
6	-2383.94	-1792.13

The table shows lag length selection and bold indicates lag order selected

TABLE A.2: FCVAR Order of integration

$\hat{d}$	0.411 (0.038)

This table shows the estimation of the order of integration of the FCVAR model for the countries analyzed in this study. Standard error appears in parenthesis

TABLE A.3: Serial correlation LM-test

LM	12.429 (0.412)

Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. P values are in parenthesis below LM test values

Part II

The monetary transmission mechanism

Chapter 4

The Term Structure under non-linearity assumptions: New methods in time series

4.1 Introduction

The relationship between interest rates or bond yields and different maturities has been widely studied in the finance literature from a theoretical and empirical point of view. The so-called term structure of interest rates has always been of fundamental importance to financial economists, investors and practitioners (see Campbell and Shiller (1991) for an influential approach or Gürkaynak and Wright (2012) for a survey).[1] In particular, understanding the term structure of interest rates is essential for the assessment of the effects of monetary and macroeconomic policies (Mankiw, Summers, and Weiss, 1984) in the context of monetary policy as an indicator of market expectations (Rudebusch, 1995); the term structure of interest rates contains useful information regarding future real activity and inflation and has prediction power (Estrella and Mishkin, 1997).

According to this framework, the term structure should move in line with the predictions of the expectations hypothesis of the term structure (EHTS) so that returns respond to international market forces. Given the connection of the implications of monetary policy decisions to the future of financial markets, the literature has emphasized the understanding of how this relationship works in the long term. Bernanke and Blinder (1992) supported that this relationship among short and long-term inter- est rates implies that their spread contains significant information on future changes in the short-term rates and has an important role in the potential effectiveness of monetary policy. Similarly, Holmes, Otero, and Panagiotidis (2015) sustained that if a monetary policy is effective, changes in short-term policy interest rates should have an impact on long-term ones.

Empirical works have examined this hypothesis in different regions, focusing on the analysis of the relationship in the long term and, consequently, on the study of the linearity of this relationship by using cointegration analysis tools. However, a controversial framework is derived for this empirical review due to recent arguments that question the usefulness of linear cointegration because it provides less power and fails to detect a long-run relationship among short and long-term interest rates (Araç and Yalta, 2015). Perhaps the most determinant contribution about the treatment of these series in the long-run is the one made by Hassler and Nautz (2008), which evidenced the presence of the fractional $I(d)$ process in the long-run relationship between interest rates, creating a novel path in the treatment of the fractional time series.

Concerning the United States of America (USA), the EHTS is frequently accepted as a forecasting tool (Poole, Rasche, and Thornton, 2002), and its implications in the monetary policy are also incorporated in the design of the fiscal policy (see Weber and Wolters, 2012, 2013). Nevertheless, Mili, Sahut, and Teulon (2012) show nonlinearities in the relationship

[1]The study of term structure has been done for a long time, going back to Macaulay (1938)

between interest rates in the USA. On the other hand, concerning the Euro Zone or the European Monetary Union (EMU), in the seminal paper of Hassler and Nautz (2008), the long-run relationship among European rates is explained by a fractional perspective, giving interpretations as a measure of the central bank's ability to control the overnight rate. Cossetti and Guidi (2009) denote that the actions of the European Central Bank (ECB) in monetary policy do not have substantial impacts on the yield curve; Nautz and Scheithauer (2011) also indicate that the monetary policy design determines the strength of the relationship between the overnight rate and the central bank's policy rate. Finally, Tamakoshi and Hamori (2014) reject the presence of linearity in the Eonia—3-month Euribor rate relationship.

In this regard, a new approach in the study of the relationship between short- and long-term interest rates has arisen in the existing literature, considering that the standard unit root and cointegration test might be too restrictive ($I(1)/I(0)$ dichotomy). Indeed, the rejection of the assumption that both short- and long-term interest rates follow the dichotomy $I(1)/I(0)$ displaying the long-memory process ($I(d)$-type) in the case of the cointegration of both interest rates. The spread could also be measured as $I(d - b)$. To the best of our knowledge, the rigidity of the traditional approach, the linear cointegration, is broken to allow the series to be cointegrated, and the spread does not necessarily need to be stationary-$I(0)$. Overall, this new approach consists of the fractionally cointegrated vector autoregressive (FCVAR) model (Johansen, 2008a, 2008b) and Johansen and Nielsen (2012), which was further developed by Nielsen and Popiel (2016).

The rest of the chapter aims to establish an empirical framework that is useful to analyse the EHTS under improved tools at the time that this approach is implemented in a monetary policy portfolio. Therefore, the next section 4.2 presents a review and a definition of the EHTS, and section 4.3 shows the empirical evidence by region. Later, in section 4.4, we develop the FCVAR model. Additionally, the monetary policy and controllability of interest rates are discussed in section 4.5, and the conclusions are discussed in section 4.6.

4.2 The Term Structure and the Expectations Hypothesis of the Term Structure

The term structure of interest rates analyses the relationship between the time remaining until the expiration of the various obligations or bonds and their returns during that period, provided that they all have the same degree of risk, liquidity and tax (Schaefer, 1981). It is also called yield curve. The most well-known term structure is formed by financial assets issued by the state because of (a) care solvency risk and (b) problems for caregiver country titles when the market for such assets for the liquidity is problematic.

The term structure of interest rates has multiple applications, which can be divided into four large groups. First the financial economy allows the evaluation of multiple financial assets and the design of the investment or hedging strategies (Bansal and Shaliastovich, 2013). Second economic theory allows for the study of issues such as the formation of expectations and the relationship between short- and long-term interest rates and the transmission of the monetary policy to the relevant macroeconomic variables (Mankiw et al., 1984). Third, the Treasury contributes to analysing the constraints of funding. Finally, the term structure is an indicator for monetary policy that is useful for analysing, along with other tools, the conditions in which this theory acts, the prospects of achieving the target set, the perception of the tone of politics in the monetary policy and the degree of confidence in maintaining it in the future (Cassola and Morana, 2008).

As it has been mentioned previously, one of the applications of the term structure is in the formation of expectations. In this sense, one of the most influential theories of term structure emerged as a way to explain the possible relationship between short- and long-term interest rates. This hypothesis that we are introducing is the EHTS, which establishes that an average of the current and expected short-term rates determines long-term rates with an inter-temporal term premium (Bekaert, Hodrick, and Marshall, 1997) and has economic implications in macroeconomics or finance and in the shape of the yield curve (see Shiller and McCulloch, 1990 for a survey).

This hypothesis was initially defined by Lutz (1940), although it was also con- firmed by different authors recently. He starts with the hypothesis that investors have homogeneous but not identical expectations and that the interest rates can be predicted with certainty. Thus, the basic hypotheses formulated by Lutz are as follows:

(a) The markets are efficient; the new information is rapidly reflected in the share prices.
(b) The investors maximize their expected profit by using short- and long-term securities.
(c) There are no transaction costs, and there is freedom of capital movements.
(d) Both the payment of the coupons and the return of the principal are known with certainty.

This hypothesis also explains the behaviour of the yield curve, since an upward sloping yield curve implies that future short-term rates are expected to rise. Conversely, with a downward sloping yield curve, the future short-term rates are expected to fall, i.e., the slope of the yield curve is an important source of information on the real economy evolution. In consequence, Estrella and Hardouvelis (1991) found that a positive curve slope is associated with future increases in real economic activity using macroeconomic variables and by providing a significant predictive power. One implication of the EHTS stated by Fama (1984) and Fama and Bliss (1987) is that the forward rate is an unbiased predictor of future short-term rates. Another implication of this hypothesis is that the spread between the long-term interest rate and the short-term interest rate—the term spread—is an unbiased predictor of future short-run changes in long-term rates (Mankiw, 1986; Campbell and Shiller, 1991; Campbell, 1995). The potential effectiveness of the monetary policy is revealed by this relationship, which consists of the control of short-term policy rates by central banks (Bernanke and Blinder, 1992) that will be explained in the next sections.

The fundamental equation of the EHTS of an $n > 1$ period bond R_t (i.e., long-term interest rate) is equal to an average of the current and expected r_t (i.e., short-term interest rate) set of a $n \leq 1$ period plus a constant term. The relationship can be expressed in the following form:

$$R_t = \frac{1}{n} \sum_{k=0}^{n-1} E_t[r_{t+k}] + \phi_t^*, \tag{4.1}$$

where ϕ_t^* is a possible stationary term and E_t is the expectation operator at time t for the evolution of short-term interest rates driving long-term interest rates.

4.3 Evidence by Region

The term structure and the EHTS have been analysed in different contexts and economic regions, although the USA and the EMU are the main regions studied. In this section, we summarize the body of empirical papers that have arisen in the literature supporting (or not) the EHTS, distinguishing between both regions.

4.3.1 The USA

Concerning the studies on the EHTS in the USA, these works have been reviewed, and it is possible to find studies showing certain controversy in relation to the confirmation of the EHTS. Several studies find evidence in support of the EHTS (e.g., Campbell and Shiller, 1987; Hamilton, 1988; Hall, Anderson, and Granger, 1992). The evolution of this analysis has aimed to explore changes in the analysed periods. Engsted and Tanggaard (1994) and Enders and Granger (1998) show asymmetries in the movements towards the long- run equilibrium relationship. Additionally, Hansen (2003), Hansen and Seo (2002), Seo (2003), Junker, Szimayer, and Wagner (2006), Clarida, Sarno, Taylor, and Valente (2006) and Mili et al. (2012) support the EHTS using cointegrating techniques, showing evidence supporting the nonlinearity in the term structure of interest rates. In this context, Sarno and Thornton (2003) used non-linear error-correction equations, finding that the adjustment of the overnight rate to the Treasury bill rate is asymmetric. There is also evidence in support of the EHTS in the relation between short- and long-term rates among the European and the USA rates (Lanne, 2003; Brüggemann and Lütkepohl, 2005) and in the combining of yield factors and macroeconomic variables to relate with the EHTS, which is evidence in favour of certain regimes (Diebold, Rudebusch, and Aruoba, 2006). Weber and Wolters (2012, 2013) applied the vector error correction model (VECM) to US term structure in order to contribute an economic explanation of the deviations from the EHTS. Recently, Kishor and Marfatia (2013) showed that the future rate is cointegrated with the 3- month rate. Holmes et al. (2015) also examined the term structure of interest rates using a pairwise stationary approach supporting that the EHTS holds in the long-run, i.e., the short-run policy changes affect the long-term rates.

However, some reasons to reject the EHTS validation for the USA have also emerged in the recent literature. According to Bekaert and Hodrick (2001) (see for a survey), there are three potential reasons for the rejection of the EHTS:

First, the EHTS is based on the assumption of rational expectations and unlimited arbitrage. It may be that irrational investors make systematic forecast errors, and the ability of rational investors to profit from this situation is limited by their risk aversion. Second, the presence of time-varying risk premiums means that standard tests of the EHTS omit the variables capturing the risk premium. Whether these variables are related with interest rates, the estimated coefficients would be pulled away from those implied by the EHTS. Third, the tests themselves may lead to false rejections because of their poor properties in finite samples. (p. 1358)

In this respect, against the fulfillment of the EHTS in the USA, Engle, Lilien, and Robins (1987) demonstrated the failure of the EHTS. In a wide sample examining the term spread of G7 countries, Hardouvelis (1994), who used a VAR model that attempted to forecast changes in long-term interest rates, showed that EHTS is supported in all countries except the USA, while Bekaert and Hodrick (2001) also showed evidence against the EHTS using different methodologies. Nevertheless, Thornton (2005) tested the EHTS in federal fund rates in order to determine whether the market's expectation is less able to forecast the federal fund rates. Sarno, Thornton, and Valente (2007) also find mixed results in a bi-variate analysis; therefore, using maturities from one month to 10 years and a powerful test (Lagrange Multiplier test), the EHTS is rejected. Conversely, Guidolin and Thornton (2010) concluded that future short-term rates have deep implications for policy makers, suggesting that whether or not EHTS is true, the inability to predict the future short-term rate would imply that both long-term and short-term rates are equal, suggesting that this relation would be inconsistent, hence, the conventional theory of the term structure of interest rates is threatened. Finally, Bulkley, Harris, and Nawosah (2011, 2015) determined the failure of the EHTS using bond yields on US Treasury securities. Overall, this subsection has summarized the empirical puzzle that is derived from the review of this literature in the USA.

4.3.2 European Monetary Union

In the European Monetary Union, the empirical framework is similar to that of the USA, accentuated by the numerous applications dedicated to each of the different countries that make up this region. This topic has been studied in different ways, highlighting the papers that relate long-term interest rates, i.e., sovereign bonds or interest rates, with longer maturities and short-term interest rates; some authors believe that the Euro OverNight Index Average (Eonia) rate is crucial for the signalling and transmission of the ECB monetary policy, using the Eonia as an indicator of the behaviour of the interest rates (Benito, León, and Nave, 2007). For almost all central banks, the inter-bank money market for overnight lending is the key channel through which the monetary policy is executed. In this sense, overnight rates are the operational target of monetary policy that anchors the term structure of interest rates (Nautz and Offermanns, 2007).

Initially, the study of this topic in the Eurozone was limited because studies were focused on specific countries rather than the whole region. Therefore, papers based in some European countries are noteworthy. Hurn, Moody, and Muscatelli (1995) obtain results in favour of the EHTS from interest rates in the UK interbank market. Dahlquist and Jonsson (1995) were unable to reject the EHTS based on interest rates from Sweden. In the case of German studies, Hardy (1998), Hammersland and Vikøren (1997) and Hafer, Kutan, and Zhou (1997), demonstrated that German term structure occupies a dominant position in the future EMU. Gerlach and Smets (1997) find empirical support for the EHTS for Belgium, France, Germany, Italy, and Spain. For the Eurozone, we have to go back to Gerlach and Smets (1997), who did not reject the EHTS in a sample of 17 Euro-countries. However, this issue acquired relevance once the European Monetary Union was born. Thus, Ayuso and Repullo (2003) show that non-symmetric adjustment of the Eonia would also be induced by an asymmetric loss of function of the central bank. This latter issue will be important for the development of future studies concerning the term structure associated with the monetary policy in the Eurozone. Meanwhile, Nautz and Offermanns (2007) confirmed the expectations with an asymmetric response for the Euro area. Importantly, we find a seminal paper (Hassler and Nautz, 2008) that explained the long-run relationship between European rates from a fractional perspective and gives some advice about the controllability of interest rates by Central Banks, changing the traditional assumption of the term structure treatment and using the cointegration techniques. For its part, Cossetti and Guidi (2009) denote that the actions of the ECB in monetary policy do not have substantial impacts on the yield curve because the presence of cointegration was rejected for maturities longer than six years, which means that for shorter rates, the presence of expectations would not be rejected. Regarding the pressure on the Eonia, Linzert and Schmidt (2011) show that the rate expectations are not relevant in a scenario with a reduction of liquidity. Otherwise, Nautz and Scheithauer (2011) indicate that the strength of the relation between the overnight rate and the central bank's policy rate is determined by monetary policy design. Similarly, Belke, Beckmann, and Verheyen (2013) used a linkage between short-term interbank interest rates, i.e., the Eonia and the 3-month Euribor rate, to study the persistence of the spread due to the importance of the market expectations of the European monetary policy attitudes in the near future. Likewise, Tamakoshi and Hamori (2014) rejected the presence of linearity in the Eonia—3-month Euribor rate relationship using a threshold cointegration. They also determined that the Eonia plays a crucial role in signalling the target of the monetary policy. Araç and Yalta (2015) consider whether the recent financial and debt crises may have affected the relation between short-term and long-term interest rates, indicating that the EH holds in Greece, Ireland and Portugal. Meanwhile, for the other countries in the sample, there is no cointegration between short- and long-term interest rates.

Finally, in order to simplify and clarify this section, a summary is provided in Table 4.1 in order to show the authors, year of publication, the concerning country or region, the technique used and whether the EHTS is supported (or not).

Table 4.1: Summary of EHTS evidence by regions

Authors, year	**Region/Country**	**Technique**	**EHTS**
Araç and Yalta (2015)	Eurozone	Nonlinear cointegration	Only in Greece, Ireland and Portugal
Ayuso and Repullo (2003)	European Monetary Union	Generalized Method of Moments	Yes

Table 4.1 continued from previous page

Authors, year	Region/Country	Technique	EHTS
Baillie and Bollerslev (1994a)	Canada, West Germany, Japan, United Kingdom, France, Italy and Switzerland	Fractional Cointegration	Yes
Bekaert and Hodrick (2001)	The USA, Germany and United Kingdom	VAR and Lagrange Multiplier	No
Bekaert et al. (1997)	(Bootstrap approach)	VAR – GARCH	No
Belke et al. (2013)	European Monetary Union	VAR and Lagrange Multiplier	Yes
Benito et al. (2007)	European Monetary Union	ARCH-Poisson-Gaussian process	
Brüggemann and Lütkepohl (2005)	The USA and the Eurozone	VECM	Yes
Bulkley et al. (2011)	The USA	VAR	No
Bulkley et al. (2015)	The USA	The Law of Small Numbers	No
Busch and Nautz (2010)	European Monetary Union	Fractional Integration	Yes
Camarero, Ordóñez, and Tamarit (2008)	European Monetary Union	Pooled and Panel Cointegration	Yes
Campbell (1995)	The USA	Regressions of long-run changes	Yes
Campbell and Shiller (1987)	The USA	CVAR	Yes
Campbell and Shiller (1991)	The USA	VAR	Yes
Clarida et al. (2006)	The USA, Germany and Japan	Nonlinear VECM	Yes
Cömert (2012)	The USA	Simple OLS and Generalized Method of Moments	Yes
Cossetti and Guidi (2009)	European Monetary Union	EGARCH and Cointegration	Yes
Dahlquist and Jonsson (1995)	Sweden	Cointegration and ECM	Yes
Diebold et al. (2006)	The USA	Non-structural VAR	Yes
Enders and Granger (1998)	The USA	Momentum Threshold Autoregressive (M-TAR) and ECM	Yes
Engle et al. (1987)	The USA	ARCH	No
Engsted and Tanggaard (1994)	The USA	VECM	Yes
Estrella and Hardouvelis (1991)	The USA	OLS	Yes
Estrella and Mishkin (1997)	The USA and Germany	VAR	Yes
Evans and Marshall (1998)	The USA	VAR and Impulse – Response Functions	Yes
Fama (1984)	The USA	OLS	Yes
Fama and Bliss (1987)	The USA	OLS	Yes

Table 4.1 continued from previous page

Authors, year	Region/Country	Technique	EHTS
Gerlach and Smets (1997)	Europe and The USA	Cross – sectional regressions	Yes
Guidolin and Thornton (2010)	The USA	Diebold and Li Model and OLS	No
Hafer et al. (1997)	Belgium, France, Germany and Netherlands	VAR and Permanent-Transitory decomposition	Yes
Hall et al. (1992)	The USA	Cointegration	Yes
Hamilton (1988)	The USA	Markov processes and OLS	Yes
Hansen and Seo (2002)	The USA	Threshold VECM	Yes
Hansen (2003)	The USA	VAR	Yes
Hardouvelis (1994)	G7 countries	VAR	Yes (except The USA)
Hardy (1998)	Germany	OLS	Yes
Hassler and Nautz (2008)	European Monetary Union	Fractional Integration	Yes
Holmes et al. (2015)	The USA	Pair-wise Cointegration	Yes
Hurn et al. (1995)	United Kingdom	VAR	Yes
Junker et al. (2006)	The USA	Copula functions	Yes
Kishor and Marfatia (2013)	The USA	Dynamic OLS and VECM	Yes
Lanne (2003)	The USA	Markov Switching Model	Yes
Mankiw (1986)	The USA, Canada, United Kingdom and Germany	GLS	Yes
Mankiw et al. (1984)	The USA	WLS	No
Mili et al. (2012)	The USA	Parametric Nonlinear Inference Approach	Yes
Nautz and Offermanns (2007)	European Monetary Union	Nonlinear Cointegration	Yes
Nautz and Scheithauer (2011)	European Monetary Union, The USA, United Kingdom and Switzerland	Fractional Integration	Yes
Poole et al. (2002)	The USA	OLS and Poole/Rasche and Kuttner methodology	No
Sarno and Thornton (2003)	The USA	Non-Linear Asymmetric Vector Equilibrium Correction Model	Yes
Sarno et al. (2007)	The USA	Lagrange Multiplier	No
Seo (2003)	The USA	Threshold cointegration	Yes
Strohsal and Weber (2014)	The USA	GARCH and Cointegration	Yes

Table 4.1 continued from previous page

Authors, year	Region/Country	Technique	EHTS
Tamakoshi and Hamori (2014)	European Monetary Union	Threshold Cointegration	Yes
Thornton (2005)	The USA		No
Weber and Wolters (2012)	The USA	VECM	Yes
Weber and Wolters (2013)	The USA	VECM	Yes

4.4 Time Series applications of the Term Structure: The FCVAR

Regarding the methodology used in the previous section, the majority of the literature has shown that it is possible to establish a relationship between short- and long-term rates using, mainly, cointegration techniques. Engle and Granger (1987) developed this concept. Initially, there were studies focused on whether interest rates can be characterized as an $I(0)$ or $I(1)$ series. For instance, Cox, Ingersoll Jr, and Ross (1985) concluded that the short-term nominal interest rate is a stationary and mean-reverting $I(0)$ process, whereas Campbell and Shiller (1987) assumed a unit root. To solve this restriction, many authors used threshold cointegration, such as Hansen and Seo (2002) and Seo (2003), who showed evidence supporting the nonlinear mean-reversion in the term structure of interest rates. Nevertheless, as it has been measured traditionally, we believe that the standard unit root and cointegration test might be too restrictive ($I(1)/I(0)$ dichotomy); the choice of such models is hotly debated, since it is unclear whether $I(0)$ or $I(1)$ processes are more appropriate (Caporale and Gil-Alana, 2016). In this sense, different authors have indicated that term structure could display long- memory processes. In this regard, Hassler and Nautz (2008), Cassola and Morana (2008), Busch and Nautz (2010), Caporale and Gil-Alana (2016) and Nautz and Scheithauer (2011) determine that an $I(d)$ process could provide additional flexibility to the relationship behaviour, with d values different from 0 or 1.

According to this idea, a novel methodology emerges in order to avoid the problems with the axioms of traditional cointegration associated with rigidity, rejecting the assumption that both short- and long-term interest rates follow the dichotomy $I(1)/I(0)$, and the spread follows a stationary process ($I(0)$), in line with Perez-Quiros and Mendizábal (2006) or Nautz and Offermanns (2007). The model is a generalization of Johansen (1995)'s cointegrated vector autoregressive (CVAR) model to allow for fractional processes of order d that co-integrate to order d - b. This model has the advantage of being used for stationary and non-stationary time series and is presented by Johansen (2008a, 2008b) and further developed by Johansen and Nielsen (2012) and Nielsen and Popiel (2016).

To introduce the FCVAR model, we begin with the well-known, non-fractional, CVAR model. Being $Y_t = 1, \ldots, T$ a p-dimensional $I(1)$ time series. Therefore, the CVAR model is:

$$\Delta Y_t = \alpha\beta' Y_{t-1} + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta Y_{t-i} + \varepsilon_t = \alpha\beta' L Y_t + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta L^i Y_t + \varepsilon_t \tag{4.2}$$

The fractional difference operator introducing persistence in the model is Δ and the fractional lag operator is $\Delta = (1 - L)$. Replacing lags operators in by their fractional counterparts Δ^b and $\Delta^b = (1 - L^b)$, we obtain:

$$\Delta^b Y_t = \alpha\beta' L_b Y_t + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta^b L_b^i Y_t + \varepsilon_t, \tag{4.3}$$

and we apply this equation to $Y_t = \Delta^{d-b} X_t$, such that:

$$\Delta^d X_t = \alpha\beta' L_b \Delta^{d-b} X_t + \sum_{i=1}^{k} \mathit{\Gamma}_i \Delta^d L_b^i X_t + \varepsilon_t. \tag{4.4}$$

As always, ε_t is p-dimensional independent and identically distributed with mean zero and covariance matrix Ω. The parameters α and β are $p \times r$ matrices, where $0 \leq r \leq p$. In matrix β the columns are the cointegrating relationships and $\beta' X_t$ assumes the existence of a common stochastic trend, which is integrated with order d. The short-term parts from the long-run equilibrium are integrated in order $d-b$, but if $d-b < 1/2$, then it is asymptotically a zero-mean stationary process. The coefficients in α correspond to the speed of adjustment of the equilibrium. Therefore, $\alpha\beta'$ is the adjustment long-run, ρ' is the restricted constant term, and Γ_i represents the short-run behaviour of the variables. We reach the final model

$$\Delta^d X_t = L_d \alpha(\beta' X_t + \rho') + \sum_{i=1}^{k} \Gamma_i \Delta^d L_d^i X_t + \varepsilon_t. \tag{4.5}$$

When the VAR model is in the case of $d = b = 1$ (CVAR), X_t is integrated with order d, and b is the strength of the cointegrating relationships (as the value of b is higher, the persistence is lower in the cointegrating relationships). The error correction term is integrated with order $(d-b)$, which is $I(0)$ in this case. However, in the fractional cointegration, these axioms are relaxed because $(d-b) = 0$, i.e., the error correction term shows a short-run stationary behaviour or $(d-b) > 0$, i.e., there is a long memory process, and the error correction term will revert in the long run.

Johansen and Nielsen (2012) show that the maximum likelihood estimators $(d, \alpha, \Gamma_i, \ldots, \Gamma_k)$ are asymptotically normal and the maximum likelihood estimator of (β, ρ) is asymptotically mixed normal.

To determine the number of stationary cointegrating relations following the hypotheses in the rank test based on a series of LR tests, in the FCVAR model, we test the hypothesis $H_0 : rank(\Pi) = r$, against the alternative: $H_1 : rank(\Pi) = p$. As $L(d, b, r)$, the profile likelihood function is given a rank r, where (α, β, Γ) has been reduced by rank regression (see Johansen and Nielsen, 2012). In the case of a model with a constant, we test $H_0 : rank(\Pi, \mu) = r$, against the alternative: $H_1 : rank(\Pi, \mu) = p$, and the profile likelihood function given rank r is $L(d, r)$, where the parameters $(\alpha, \beta, \rho, \Gamma)$ have been focused. Note that matrix α and β are normalized separately in the same way for the CVAR model because the degrees of freedom are non-standard.

Maximizing the profile likelihood distribution under both hypothesis, the LR test statistics are now $LR_t(q)$. The asymptotic distribution of $LR_t(q)$ depends on the parameter b and on $q = n - r$. MacKinnon and Nielsen (2014), based on their numerical distribution functions, provide asymptotic critical values of the LR rank test. In the case of "weak cointegration", i.e., $0 < b < 1/2$, $LR_t(q)$ has a standard asymptotic distribution,$LR_t(q)$ $LR_t(q) \xrightarrow{D} \chi^2(q^2)$. To summarize,by estimating the FCVAR model,we extract richer information from what was mentioned in the previous sections. Importantly, by separately parameterizing the long-run and the short-run dynamics of the series, the model is able to accommodate empirically realistic $I(d)$ long-memory and fractional cointegration while maintaining that the returns are $I(0)$ (Bollerslev et al. 2013).

As we said, this methodology allows testing of the long-run relationship between interest rates with different maturities, the measurement of the spread[2] and the implications for monetary policy in a joint estimation. For this reason, Table 4.2 proposes a possible strategy of empirical research, allowing for the study of long-run relationships of interest rates and testing the spread persistence in order to achieve monetary policy conclusions.

[2]When the relationship between interest rates with different maturities is supported, it has to be restricted by a cointegrating vector of (1, 1), then, the difference between those interest rates are interpreted as the spread.

4.5 Monetary policy and controllability of interest rates

Studies concerning the term structure of interest rates have tried to evaluate their impact and how they are affected by the monetary policy of Central Banks. The term structure has long been established as reflecting economic agents' anticipations of future events and as an indicator of monetary policy, as seen in the volume of academic articles written over the past century dealing with term structure, which is testimony to both the practical importance of the topic as well as its intrinsic academic appeal (see Vetzal, 1994 for a survey). In consequence, changes in the economy could affect the EHTS; therefore, if a variation in short-term policy impacts the long term, monetary policy is effective (Holmes et al., 2015).

TABLE 4.2: Strategy of empirical research

	Procedure	Hypotheses
Step 1	Fractional cointegration?	H_1^d: Is the fractional cointegration more appropriate that traditional cointegration?
Step 2	Estimation of β	H_1^β: Cointegrating vector is (1, -1)
Step 3	Estimation of adjustment coefficients (α_R, α_r)	$H_1^\beta \cap H_1^{\alpha_i}$: The interest rates are weakly exogenous under the restriction of the cointegrating vector (1, 1)
Step 4	Degree of spread persistence, i.e., order of integration $(d-b)$	H_1^{d-b}: Is the spread a long memory process?

As mentioned in the previous sections, the potential effectiveness of monetary policy is revealed by this relationship, which consists of the control of short-term policy rates by central banks. In principle, since authorities can control the path of short-term interest rates, they should also be able to influence the wide movements in long-term interest rates sufficiently for policy objectives, provided traditional term structure relations hold up reasonably well (Christiansen and Pigott, 1997). This does not require that the classical term structure theory holds exactly but only that expectations about future short-term interest rates have a major influence on long-term rates, as is suggested by traditional studies of the term structure (Shiller and McCulloch, 1990). Although the connection between monetary policy and long-term interest rates appears to be weaker and less reliable, monetary policy can readily influence short-term interest rates (Roley and Sellon, 1995; Camarero et al., 2008). In this sense, monetary policy shocks primarily affect short-term interest rates with a diminishing effect on longer-term rates, which can be explained by the EHTS (Evans and Marshall, 1998). Hence, under the expectations hypothesis, changes in the term structure can be used to infer changes in investors' expectations concerning the path of monetary policy. If, in addition, the central bank's rule relating monetary policy to macroeconomic conditions is known by those investors, then we could also read off changes in their expectations of the state of the economy (see Gürkaynak and Wright, 2012 for a survey).

More importantly, several studies have treated the spread between long- and short- term interests. These studies have focused on different regions in the world; for the EMU, the impact of monetary policy shocks on bond yields declines with the maturity of the bonds, and this impact is significantly lower when the shock stems from a monetary policy meeting of the ECB (Perez-Quiros and Sicilia, 2002). Regarding policy implications, Hassler and Nautz (2008), Cassola and Morana (2008) and Nautz and Scheithauer (2011) reveal that the Eonia spread is $I(0)$ before but fractionally integrated with long memory when the order of fractional integration d has increased to approximately 0.25. Since $d < 0.5$, the Eonia is still under the ECB's control. Additionally, the increased persistence of the Eonia spread suggests that the degree of controllability of the Eonia spread may have declined. Meanwhile, Busch and Nautz (2010) estimated a long memory process and found that the persistence of deviations in long-term money market rates from the European central bank's policy rate has decreased, implying that monetary policy has become more effective in controlling interest rates. Caporale and Gil-Alana (2016) suggest that the ECB, and member state central banks, have controlled money market rates in a strict way, particularly at the short-end of their term structure. In the case of the USA, the work of Strohsal and Weber (2014) and Holmes et al. (2015) supports the EHTS; however, the degree of integration of the spread would be different from $I(0)$. Previously, Cömert (2012) related overnight interest rates and long-term rates in the US and offered evidence that the Fed has been gradually losing its control over long-term interest rates.

Another tool to control the interest rates is the study of the spread. In this regard, Bernanke and Blinder (1992) showed that this relationship among short- and long-term interest rates implies that their spread contains significant information on future changes in short-term rates and has an important role in the potential effectiveness of monetary policy, which consists of the control of short-term policy rates by central banks. The economy is affected by the monetary impulses through long-term interest rate movements. Therefore, if spreads are highly persistent, the lasting impact of shocks may impede the transparency of policy signals and, thus, the central bank's impact on longer-term rates. In this respect, Nautz and Offermanns (2007) found evidence that the reaction of the Eonia rate to the spread is non-symmetric but, interestingly, the ECB did not lose control over the Eonia rate. From a fractional integration perspective, the spread could exhibit long memory ($d > 0$), but non-stationarity ($d \geq 0.5$) can be rejected in most cases. An explicit test for a change in the order of fractional integration is provided by Sibbertsen and Kruse (2009). Additionally, Baillie and Bollerslev (1994a), Tkacz (2001)[3] and Cassola and Morana (2008), among others, suggest that the spread could follow a fractional order of integration, which could be an indicator of the power that the authorities have over the interest rates.

4.6 Conclusion

Corresponding to the EHTS, long-term rates could explain changes in future short-term rates. Understanding the term structure of interest rates has always been viewed as crucial to assess the impact of monetary policy and its transmission mechanism. Indeed, if a monetary policy is effective, changes in short-term policy interest rates should impact long-term ones. Despite this hypothesis being widely known, major contributions arose in the end of the last century and the beginning of the 21^{st} century. Notably, the studies carried out by Campbell and Shiller (1987) and Fama and Bliss (1987) contributed to establishing the main implications of the EHTS. More recently, several papers have examined the existence of the EHTS by using time series methodologies, using different perspectives, i.e., selecting different maturities for interest rates and/or different countries, and providing conclusions for investors and policy makers.

Initially, the research concerning the EHTS was focused on the study of the interest rates under the lens of the existence (or not) of unit roots, i.e., the series would be $I(0)/I(1)$. Nonetheless, authors such as Mili et al. (2012) and Hassler and Nautz (2008), for instance, showed the presence of non-linarites and a fractional $I(d)$ process in the long-run relationship between interest rates, respectively. According to these results, a novel technique in the treatment of the fractional time series emerges, i.e., the FCVAR applied to the long-run relationship between short- and long-term interestrates. Under the FCVAR assumptions,it could be considered that the standard unit root and cointegration test might be too restrictive ($I(1)/I(0)$ dichotomy). Indeed, the rejection of the assumption that both short- and long-term interest rates follow the dichotomy $I(1)/I(0)$ displaying the long-memory process ($I(d)$-type) is similar to the case of the cointegration of both interest rates. Additionally, the spread could be measured as $I(d-b)$. Therefore, the rigidity of the traditional approach is broken in favour of allowing the series to be cointegrated, and the spread does not necessarily need to be $I(0)$.

[3]Following Tkacz (2001), when $(d-b) = 0$, the spread follows a stationary process and the shock duration is short-lived, i.e. this means that a shock would show a slow return towards the long-run equilibrium. If $0 < (d-b) < 0.5$, there is a stationary process, and the shock duration is long-lived, and finally, if $0.5 < (d-b) < 1$, the spread follows a non-stationary process, although it is mean-reverting and the shock duration is long-lived.

Finally, by testing the term structure of interest rates, it is possible to study the behaviour of the long-run relationship between interest rates and how the term structure would change in the time after a shock. According to this idea, involving the concept of long memory and fractional integration and cointegration, the joint estimation of the long-run relationship and the study of the persistence of the spread are possible. In this respect, the long memory of the spread holds adequate forecasting power over longer horizons (Baillie and Bollerslev, 1994a). Otherwise, another factor plays an important issue in the design of the monetary policy, i.e., the persistence of the spread, in which a greater persistence may indicate that it is more difficult for monetary policy signals to be transmitted along the money market yield curve. Additionally, if spreads are highly persistent, the lasting impact of shocks may impede the transparency of policy signals and, thus, the central bank's impact on longer-term rates, implying a gradual loss of control over interest rates by Central Banks (see Cassola and Morana, 2008; Hassler and Nautz, 2008; Cömert, 2012 for Europe and the USA).

Chapter 5

The role of EONIA in the dynamics of short-term Interbank rates

5.1 Introduction

Interest rates play an important role in the monetary policy defined by central banks, joining the short- and longer-term interest rates to predict the behaviour of the financial markets and the economy. In particular, the term structure has long been established as reflecting economic agents' anticipation of future events and an indicator for policy makers, as evidenced by the volume of academic literature over the past century dealing with the term structure (see Vetzal (1994) for a survey).

According to this framework, the financial environment is competitive, and the term structure should move in assembly with the predictions of the expectations hypothesis of term structure (EHTS hereafter); thus, returns respond to international market forces, and considering the term structure of interest rates has always been viewed as crucial to assess the impact of monetary policy and its transmission mechanism. Indeed, Bernanke and Blinder (1992) supported that this relationship between short- and longer-term interest rates implies that their spread contains significant information on future changes in short-term rates and plays an important role in the potential effectiveness of monetary policy. Cossetti and Guidi (2009) denote that the actions of the European Central Bank (ECB hereafter) in monetary policy do not have substantial effects on the yield curve, and Nautz and Scheithauer (2011) indicate that the monetary policy design determines the strength of the relation between the overnight rate and the central bank's policy rate.

In this context, we apply the fractionally cointegrated vector autoregressive (FCVAR hereafter) model combined with Permanent-Transitory decomposition (P-T decomposition hereafter) (Gonzalo and Granger, 1995). We test for the existence of a long-run relationship between short- and long-term interest rates, as combined spread persistence, and also provide evidence that interest rate has the dominant position in the common trend. The chapter is structured as follows. Section 5.2 presents the literature selected. Section 5.3 introduces the data selected and the econometric strategy as well as the methodology used to determine the results, which are shown in section 5.4. Finally, in section 5.5, we summarize and establish the conclusions.

5.2 Literature review

This body of literature has been supported by the expectations hypothesis of term structure (EHTS), which consists of the study of this linkage among overnight rates and short-term rates to explain the monetary policy in the Eurozone, establishing that longer-term interest rates are determined by the expected short-term rates plus a constant term and thus that both interest rates show a long-run relationship (see Campbell and Shiller, 1987). In other words, if the EHTS is confirmed, the spread is a predictor of the changes in the relationship

(Mankiw, 1986; Campbell and Shiller, 1991;Campbell, 1995). In this sense, the vast literature has focused on the study of the EHTS, assuming that the spread follows a stationary process as a condition to contrast this issue. Nevertheless, there are authors who have countered this assumption about the possible non-stationarity or persistence of the spread but obviating the existence of the EHTS.

In this regard, Bernanke and Blinder (1992) supported that this relationship among short- and longer-term interest rates implies that their spread contains significant information on future changes in short-term rates and plays an important role in the potential effectiveness of monetary policy, which consists of the control of the short-term policy rate by central banks; the economy is affected by monetary impulses through long-term interest rate movements. More recently, Hassler and Nautz (2008) have revealed an important result: they expose that if the persistence of the Eonia spread is too high, it means that the central bank would lose control over interest rates due to the perdurable impact of shocks, avoiding the signalling role of the Eonia rate. For its part, Cossetti and Guidi (2009) denote that the actions of the ECB in monetary policy do not have substantial effects on the yield curve. Linzert and Schmidt (2011) analyze how a reduction in liquidity could alleviate pressure on the Eonia spread according to the monetary policy designed, and Nautz and Scheithauer (2011) indicate that the monetary policy design determines the strength of the relation between the overnight rate and the central bank's policy rate. In this line of research, the linkage among short-term interbank interest rates in European banks, i.e., the Eonia and the 3-month Euribor rates, to study the persistence of the spread due to the importance of market expectations of the European monetary policy attitude in the near future, has been recently established by Belke et al. (2013). Furthermore, according to Tamakoshi and Hamori (2014), the Eonia rate plays a crucial role in signalling the target of monetary policy, while the Euribor rate provides outstanding interest rates for various financial products, i.e., the 3-month Euribor rate is used because it has been a focus in recent studies of interbank money markets. Finally, Hauck and Neyer (2014) explain how the Eurosystem's liquidity measures to reactivate the interbank market could conflict with aims from the monetary policy perspective and financial stability perspective. Our empirical setup for the analysis of the dynamics in the relationship between the overnight rate and the short-term interest rates is given. Soares and Rodrigues (2013) warn that changes in official interest rates impact banks' funding costs and bank loans' interest rates. In this sense, they also support that given that central bank reference rates are transmitted along the yield curve and other asset prices, the central bank can influence investment and consumption decisions and, ultimately, consumer prices. Furthermore, bearing in mind the dynamics between these two interest rates is of crucial importance for the implementation of monetary policy by the ECB since one of its main objectives is to influence the interest rates in the short term in the interbank money market (Hassler and Nautz, 2008).

According to the previous scenario, one of the main results regarding these policy implications of spread persistence has been shown by Hassler and Nautz (2008), Cassola and Morana (2008) and Nautz and Scheithauer (2011) in Europe. They reveal that the Eonia spread is $I(0)$ before but fractionally integrated with long memory when the order of fractional integration d has increased to approximately 0.25. Since $d < 0.5$, the Eonia is still under the ECB's control. Additionally, the increased persistence of the Eonia spread suggests that the degree of controllability of the Eonia spread may have declined, while Busch and Nautz (2010) estimated a long memory process and found that the persistence of deviations in longer-term money market rates from the European Central Bank's policy rate has decreased, implying that monetary policy has become more effective in controlling interest rates. Overall, in relation to having control of monetary policy, another strand of the literature has focused on Permanent-Transitory decomposition (Gonzalo and Granger, 1995) to explain the information contained in the common trend, which is useful in the long run and for expectations about the course of government policies, i.e., to identify and estimate the common trend that drives the cointegrating relation between the interest rates. One first application of this methodology is by Hafer et al. (1997), who demonstrated that German term structure occupies a dominant position in the future EMU.

The FCVAR model has been employed in reference to financial markets and political economics. Caporin, Ranaldo, and De Magistris (2013) applied the FCVAR model on high and low prices to predict stock prices. For its part, Rossi and De Magistris (2013) applied this methodology to study the relationship between spot and futures markets, and Jones et al. (2014) checked the fractional long-run relationship between Canadian political support and macrovariables. Additionally, Dolatabadi et al. (2016) and Dolatabadi et al. (2018) applied the FCVAR model for the analysis of price discovery in commodity markets, and more recently, Maciel (2017) modelled and forecasted daily high and low asset prices. Few studies in the literature have addressed the application of this methodology in the interest rates. Such studies are Abbritti, Carcel, Gil-Alana, and Moreno (2018), who studied the US term premium under fractional cointegration conditions, and Gil-Alana and Carcel (2018), who performed the same for exchange rates. This methodology is useful in that it allows us to test for cointegration between interest rates of different maturities and spread stationarity simultaneously, unlike what is possible with the traditional cointegration method; with the traditional method, different studies have executed this exercise separately.

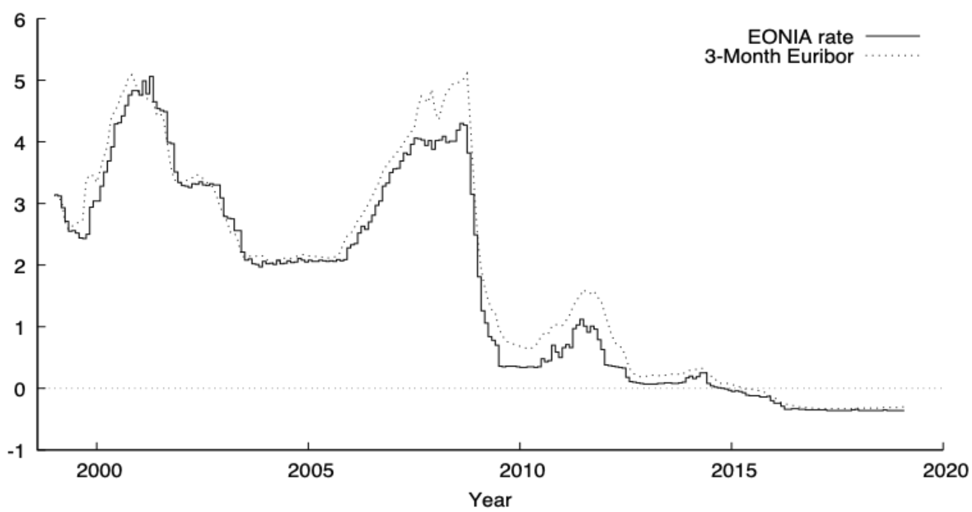
This paper is novel in this body of literature in that it recognizes that the premises of standard cointegration testing ($I(1)/I(0)$ dichotomy) time-series variables, integrated at order one and comoved at order zero, are too restrictive, i.e., linear combinations of $I(1)$ nonstationary processes are $I(0)$ stationary. In this sense, the empirical literature has shown that many economic and financial time series hold long-range dependence in the autocorrelation function but do not precisely exhibit a unit root process, i.e., the long memory process. For this reason and according to our research, we reject traditional cointegration assumptions that all interest rates cannot move away from one another for long periods of time and that they are unit roots or $I(1)$; they follow dichotomy $I(0)/I(1)$, such that they follow a fractional process $I(d)$. We also discard the notion that the error term follows a stationary process ($I(0)$) (in line with Perez-Quiros and Mendizábal (2006) or Nautz and Offermanns (2007), who assume that the Eonia spread is stationary) in cases of the cointegration of both variables. In turn, the rigidity of the traditional approach is overcome in favour of allowing for the series to be cointegrated, and the error term does not necessarily need to be $I(0)$; for example, we allow for the error term to be cointegrated in order $I(d$ - $b)$, unlike other techniques, which assume that the error term is $I(0)$. To the best of our knowledge, the relationship between shorter- and longer-term interest rates follows a $I(0)/I(1)$ process; however, fractional cointegration may refute this assumption such that, in the presence of a unitary long-run relationship between interest rates with different maturities, shocks that affect this cointegration relationship can be long-lived and even non-stationary. Indeed, the study of the long-run relationship and the behaviour of the error term may be analysed jointly, which is one of the main advantages of this methodology. Therefore, our new approach uses the FCVAR model developed by Johansen and Nielsen (2012) and Nielsen and Popiel (2016), which is an expansion of the traditional cointegrated VAR (CVAR) model proposed by Johansen (1995), enabling us to establish the number of equilibrium relations via cointegrating rank testing to estimate memory parameters, long-run cointegrating relations with adjustment parameters, and short-run lagged dynamics. In this respect, our purpose is to analyze the dynamics of the short-term side of the yield curve, i.e., the relationship between Eonia rate and short-term interbank rates (Euribor rate) as well as the repercussion that the behaviour of the spread between both interest rates may affect the monetary policy and its implications simultaneously. Overall, the FCVAR model allows for several scenarios not considered until now to be determined. Finally, using P-T decomposition, we provide evidence that the interest rate has the dominant position in the common trend.

5.3 Data and Econometric approach

For our empirical analysis, we employ a monthly sample of short-term interest rates of the Eurozone over the period from January 1999 (this is the date that the euro currency was introduced) to February 2019 (totalling 242 observations for each interest rate series). The data correspond to the 3-month Euribor (R_t) interest rate and Eonia (r_t) rate measured in percentages. Euribor is the rate at which euro interbank term deposits are offered by one prime bank to another prime bank within the EMU zone and is based on market criteria that include those banks that adequately reflect the diversity of the euro money market. Meanwhile, the Eonia rate is the 1-day interbank interest rate for the Eurozone, and it is computed as a weighted average of all overnight unsecured lending transactions in the interbank market. In other words, it is the rate at which banks provide loans to each other with a duration of 1 day. Therefore, the Eonia rate could be considered the 1-day Euribor rate. The data are collected from the EUROSTAT website. First, it should be noted that these interest rates were chosen because, following the study by Tamakoshi and Hamori (2014), on the one hand, the ECB Governing Council is responsible for regulating the official interest rates in the Eurozone, which operates as a benchmark for interbank market interest rates. This agrees with the first step of the monetary policy transmission mechanism (Cossetti and Guidi, 2009). According to the EHTS, the long-term interest rate should reflect the contemporary level of the very short- or short-term interest rate and its expectations over the maturity of the long-term investment. Consequently, it is the shortest maturity interest rate, i.e., the overnight interest rate, and the expectations on this rate that establish the remaining interest rates. It is important to appreciate how the Eurosystem stimuluses the market interest rate, i.e., the Euro Overnight Index Average (EONIA) rate plays a benchmark role in the Eurozone (Soares and Rodrigues, 2013). In this sense, the Eonia rate not only contains information on market expectations about the position of monetary policy in the near future but also anchors interest rates of greater maturity, and as it has been contended, the ECB influences short-term rates such as the 3-month Euribor rate by monitoring the Eonia rate, which should shift nearby MRO. Furthermore, Cossetti and Guidi (2009) show that the Eonia rate is highly correlated with the monetary policy rate, i.e., the Eonia rate could be a proxy for the monetary policy rate. On the other hand, Euribor rates are also important because they provide leading interest rates for various financial products, including interest and futures rate swaps. Euribor rates, such as the 3- and 6-month Euribor rate, which are widely used as an index for interest rates on bank loans in several Eurozone countries, are influenced by expectations of shorter-term interest rates and by liquidity and credit risk premium. Therefore, a change in official interest rates may affect the funding costs of banks and interest rates of bank loans.

As a preview of the variables selected, figure 5.1 presents a graphical analysis of the time-series dynamics plotted for the Eonia rate and 3-month Euribor rate. This plot shows a similar behaviour in both variables that could confirm our subsequent results. In fact, the fluctuations in the Eonia rate suggest liquidity conditions that are temporarily relaxed or restrictive on the money market. Table 5.1 shows the descriptive analysis associated with each interest rate. Both rates show similar values for the different measures. For instance, in terms of volatility, as we can see, both interest rates exhibit a very similar behaviour.

FIGURE 5.1: Eonia rate and 3-month Euribor rate dynamics.



Our empirical procedure consists of several steps. On the one hand, we apply the FCVAR

Table 5.1: Descriptive statistics for the options data

	Mean	Median	Min.	Max.	SD
3-Month Euribor rate	1.835	1.985	-0.330	5.110	1.726
Eonia rate	1.626	1.535	-0.360	5.060	1.676

The data sample covers from January 1999 to February 2019

model proposed by Johansen (2008a, 2008b) and Johansen and Nielsen (2012), aiming to contrast the EHTS and the possible existence of spread persistence. On the other hand, we study Permanent-Transitory decomposition (Gonzalo and Granger, 1995) to determine which interest rate drives the common trend.

To test the EHTS in the context of cointegration theory, the commonly linear model is as follows:

$$R_t = c + \beta r_t + \varepsilon_t \tag{5.1}$$

According to Campbell and Shiller (1987), R_t and r_t should be non-stationary and related through a cointegration relationship with parameters (1,-1). These results imply that β_R and β_r are the cointegrated constants and that their combination is a stationary process, and the spread of the interest rate follows a mean-reverting process. If the spread is stationary, the short- and long-term rates are driven by a common stochastic trend and do not allow for arbitrage opportunities because market forces adjust to correct any temporary disequilibrium.

5.3.1 Fractionally cointegrated vector autoregressive (FCVAR) model

The model is a generalization of Johansen (1995)'s cointegrated vector autoregressive (CVAR) model to allow for fractional processes of order d that co-integrate to order $d-b$. This model has the advantage of being used for stationary and non-stationary time series and is presented by Johansen (2008a, 2008b) and further developed by Johansen and Nielsen (2012) and Nielsen and Popiel (2016).

As always, ε_t is p-dimensional independent and identically distributed with a mean of zero and covariance matrix Ω. The parameters α and β are $p \times r$ matrices, where $0 \leq r \leq p$. In matrix β, the columns are the cointegrating relationships, and $\beta' X_t$ assumes the existence of a common stochastic trend, which is integrated of order d, and the short-term parts from the long-run equilibrium are integrated of order $d-b$; however, if $d-b < 1/2$, then it is asymptotically a zero-mean stationary process. The coefficients in α correspond to the speed of adjustment until equilibrium. Therefore, $\alpha\beta'$ is the long-run adjustment, ρ' is the restricted constant term, Γ_i represents the short-run behaviour of the variables, and the fractional difference operator introducing persistence in the model is Δ. Meanwhile, the fractional lag operator is $\Delta = (1-L)$. Replacing lags operators with their fractional counterparts Δ^b and $\Delta^b = (1 - L_b$, we obtain the final model:

$$\Delta^d X_t = \alpha\beta' L_b \Delta^{d-b} X_t + \sum_{i=1}^{k} \Gamma_i \Delta^b L_b^i Y_t + \varepsilon_t \tag{5.2}$$

When the VAR model is in the case of $d = b = 1$ (CVAR), X_t is integrated of order d, and b means the strength of the cointegrating relationships (as the value of b is higher, the persistence is lower in the cointegrating relationships). The error correction term is integrated of order $(d-b)$, that is, $I(0)$ in this case. However, in fractional cointegration, these axioms are relaxed because $(d-b) = 0$, i.e., the error correction term shows a short-run stationary behaviour, or $(d-b) > 0$, i.e., there is a long memory process, and the error correction term will revert in the long run.

Johansen and Nielsen (2012) show that the maximum likelihood estimators $(d, \alpha, \Gamma_i, \ldots, \Gamma_k)$ are asymptotically normal and that the maximum likelihood estimator of (β, ρ) is asymptotically mixed normal.

To determine the number of stationary cointegrating relations following the hypotheses in the rank test based on a series of LR tests. In the FCVAR model, we test the hypothesis $H_0 : rank(\Pi) = r$, against the alternative: $H_1 : rank(\Pi) = p$. $L(d, b, r)$ is the profile likelihood function given a rank r, where (α, β, Γ) has been reduced by rank regression (see Johansen and Nielsen, 2012), and the profile likelihood function given rank r is $L(d, r)$, where the parameters $(\alpha, \beta, \rho, \Gamma)$ have been excluded.

Maximizing the profile distribution under both hypothesis, the LR test statistics are now $LR_t(q)$. The asymptotic distribution of $LR_t(q)$ depends on the parameter b and on $q = n - r$. MacKinnon and Nielsen (2014), based on their numerical distribution functions, provide asymptotic critical values of the LR rank test. In the case of "weak cointegration", i.e., $0 < b < 1/2$, $LR_t(q)$ has a standard asymptotic distribution,$LR_t(q)$ $LR_t(q) \xrightarrow{D} \chi^2(q^2)$.

According to the existing literature, cointegration implies a FVECM such as the following:

$$\begin{pmatrix} \Delta R_t \\ \Delta r_t \end{pmatrix} = \begin{pmatrix} \alpha_R \\ \alpha_r \end{pmatrix} (R_{t-1} - \beta r_{t-1} - c) + \sum_{i=1}^{n} \Gamma_i \begin{pmatrix} \Delta R_{t-i} \\ \Delta r_{t-i} \end{pmatrix} + \begin{pmatrix} w_{1t} \\ w_{2t} \end{pmatrix} \tag{5.3}$$

with adjustment parameters α, cointegration coefficient β, restricted constant (c), lag length (n) and errors w. Γ_i are 2×2 parameter matrices in the short-run dynamics. The adjustment coefficients α_R and α_r capture the speed of adjustment of R_t and r_t towards equilibrium.

5.3.2 Permanent-transitory (P-T) decomposition in the FCVAR model

According to Gonzalo and Granger (1995)'s P-T decomposition, we let $X_t = (R_t, r_t)'$, where R_t and r_t denote the 3-month Euribor rate and Eonia rate, respectively. In P-T decomposition, X_t can be decomposed into a transitory (stationary) part, βX_t, and a permanent part, $W_t = \alpha'_\perp X_t$, where $\alpha'_\perp \alpha = \alpha' \alpha_\perp = 0$. W_t is the common permanent component of X_t, and it is interpreted as the dominant rate, where the information that does not affect W_t will not have a permanent effect on X_t. To determine which parameter contributes to each market (Eonia and Euribor), we attend to the key parameter $\alpha_\perp$. Following the mirror hypothesis, the linear hypothesis on $\alpha_\perp$ can also be tested directly on $\alpha_\perp$ or alternatively on α itself using the values of the LR tests in each hypothesis, and critical values can be taken from the χ^2 distribution for testing. For example, to test the hypothesis that the dominant rate is the 3-month Euribor rate, i.e., $\alpha_\perp = (0, a)'$, we can equivalently test the mirror hypothesis, $H_0 : \alpha = (\gamma, 0)'$. Similarly, to test the hypothesis that the dominant rate is the Eonia rate, i.e., $\alpha_\perp = (a, 0)'$, we test the mirror hypothesis, $H_1 : \alpha = (0, \gamma)'$ (see Dolatabadi et al. (2018), which first combined the FCVAR model with P-T decomposition).

An interpretation of coefficient α is that an adjustment coefficient measures how disequilibrium errors could be affected by current changes in X_t. Under this interpretation, we wonder whether any coefficients in α are zeros, i.e., the variable in question is weakly exogenous. For example, under hypothesis H_1, parameter $\alpha = 0$, such that the Eonia rate does not react to the disequilibrium error, i.e., the transitory component, implying that the Eonia rate is the main contributor to the common trend.

To determine the proportion, i.e., the component share that each parameter has in the long-run relationship, we follow Baillie, Booth, Tse, and Zabotina (2002), who notice that since $\alpha' \alpha_\perp = 0$, it may also be expressed in terms of the elements of the error correction vector α. To interpret this, we let $\alpha = (\alpha_1, \alpha_2)'$ and $\alpha_\perp \alpha = (\alpha_{\perp,1}, \alpha_{\perp,2})'$. Afterwards,

$\alpha'_{\perp}\alpha = \alpha_{\perp,1}\alpha_1 + \alpha_{\perp,2}\alpha_2 = 0$ implies that $\alpha_{\perp,1} = -\alpha_{\perp,2}\alpha_2/\alpha_1$, and thus, component share (CS hereafter) may be expressed as

$$CS_1 = \frac{\alpha_2}{\alpha_2 - \alpha_1}, CS_2 = \frac{-\alpha_1}{\alpha_2 - \alpha_1} \tag{5.4}$$

In this respect, the CS for variable 1 reflects how sensitive variable 2 is relative to variable 1, and vice versa.

Finally, in table 5.2, we present the strategy followed in our empirical research. Using this strategy, we also show the questions that we try to answer from an econometric approach based on the previously developed methodology. In this sense, the first step is testing the existence of a long-run relationship between the Eonia rate and the 3-month Euribor rate, and we study whether fractional cointegration is more appropriate than standard cointegration. In the second step, we study the possible relation one to one, i.e., the cointegrating vector (1, -1); this is the existence of the EHTS. The next step consists of analysing the adjustment coefficients; in the fourth step, we examine the fractional cointegration degree (persistence of the Eonia spread). The last step consists of applying P-T decomposition to determine which interest rates 'drive' the common trend.

TABLE 5.2: Strategy of empirical research

	Procedure	Hypotheses
Step 1	Fractional cointegration?	H_1^d: Is the fractional cointegration more appropriate that traditional cointegration?
Step 2	Estimation of β	H_1^β: Cointegrating vector is (1, -1)
Step 3	Estimation of adjustment coefficients (α_R, α_r)	$H_1^\beta \cap H_1^{\alpha_i}$: The interest rates are weakly exogenous under the restriction of the cointegrating vector (1, 1)
Step 4	Degree of spread persistence, i.e., order of integration $(d-b)$	H_1^{d-b}: Is the spread a long memory process?
Step 5	Permanent - Transitory decomposition	$H_1^\beta \cap H_{1\perp}^{\alpha EUR/Eon} \equiv H_1^\beta \cap H_1^{\alpha Eon/EUR}$ (mirror): Euribor and/or Eonia has a permanent component in the common trend

5.4 Results

As a preliminary step, we estimate the order of fractional integration of the interest rates. To motivate an FCVAR model, we first discuss the univariate results, observing long memory, and then, we proceed with the estimation of the fractional parameter d for each univariate series; the results are presented in table 5.3. The three columns are semiparametric log-periodogram regression estimates from Geweke and Porter-Hudak (1983), computed with bandwidths $m = T^{0.4}$, $m = T^{0.5}$, and $m = T^{0.6}$. Although the semiparametric log-periodogram regression proposed by Geweke and Porter-Hudak (1983) is the most used, this method was modified and further developed by Robinson (1995) and has been analysed by Velasco (1999) and Shimotsu and Phillips (2002), among others. The estimates are consistent with the joint estimates presented later. As we can see in table 5.3, the values for d increase as the bandwidth increases, becoming a mean-reverting value of approximately 1. To test the presence of unit roots, the estimates were obtained using first-differenced data because the original series might be above 0.5, and this test requires that the results are limited to the interval $-0.5 < d < 0.5$, then adding 1 to obtain the proper estimates of d. We can also see that both results—those for the Eonia rate and 3-month Euribor rate—are very similar and in line with the results shown later.

Table 5.3: Univariate analysis. GPH estimates

GPH estimates			
	$m = T^{0.4}$ $\hat{d}$	$m = T^{0.5}$ $\hat{d}$	$m = T^{0.6}$ $\hat{d}$
Eonia rate	0.745 (0.343)	1.028 (0.204)	1.337 (0.188)
3-month Euribor rate	0.652 (0.201)	1.130 (0.191)	1.199 (0.129)

GPH denotes the Geweke and Porter-Hudak semiparametric log-periodogram regression estimator. Standard errors are given in parenthesis beneath estimates of d.The sample size is 242

This section presents the results obtained corresponding to the study of a fractional cointegration analysis from a multivariate perspective. First, in table 5.4, the lag length selected under the Bayesian information criterion (BIC criterion) is one; as we can see, there is evidence that the number of cointegrating vectors is also one.

Table 5.4: Lag length selection and Rank test

Lags	**AIC**		**BIC**	
1	**-842.07**		**-800.20**	
2	-835.93		-780.11	
3	-835.11		-765.33	
4	-832.68		-748.95	
5	-837.73		-740.04	
6	-834.53		-722.89	
Rank	**LR Statistics**	**CV 1%**	**CV 5%**	**CV 10%**
0	20.462	24.151	19.342	17.065
1	9.586	11.461	7.895	6.303

Bold indicates lag length order selected. The bottom of the table shows the LR statistics and Critical Values (CV). The sample size is 242.

In the first step, we also reveal that fractional cointegration is more appropriate than traditional cointegration ($H_1^d = 0.089$), as shown in table 5.5. To verify the EHTS, we follow the next approach. First, aiming to estimate the long-run relationship between long- and short-term rates, it can be observed that parameter β is close to 1. As we cannot reject the hypothesis that the cointegrating vector is (1, -1) (H_1^{β}), the EHTS is supported by this result, and thus, we can interpret the difference between the 3-month Euribor rate and the Eonia rate as the spread, i.e., $(R_t - r_t)$.

Analogously and even more importantly, in table 5.5, we show a very interesting result about the spread persistence (step 4); thus, we can explain the difference $(d - b)$ as the order of integration of the spread. Hypothesis H_1^{d-b} determines the degree of spread persistence, which reaches a value of 0.705. According to table 1 in Tkacz (2001), when $(d - b) = 0$, the spread follows a stationary process, and the shock duration is short-lived, i.e., this means that a shock would show a slow return towards the long-run equilibrium. If $0 < (d-b) < 0.5$, there is a stationary process, and the shock duration is long-lived; finally, if $0.5 < (d - b) < 1$, the spread follows a non-stationary process, although it is mean-reverting, and the shock duration is long-lived. This implies a long memory process, and the series demonstrates stationary but mean-reverting behaviour with long-lived shock duration. As we have previously warned, the results could allow us to study the persistence of the spread due to the importance of the market expectations related to European monetary policy attitude in the near future. In this sense, according to Cassola and Morana (2008) and Hassler and Nautz (2008), if the ECB wants to direct Eonia, the order of integration of the Eonia spread should be less than 0.5.

TABLE 5.5: Fractional cointegration test and results

Hypothesis	**LR statistics**	***P* value**
H_1^d	2.894	0.089
	Cointegrating vector	(1.000, -1.019)
	$\hat{d}$	1.413 (0.134)
	$\hat{b}$	0.708 (0.145)
Hypothesis	**LR statistics**	***P* value**
H_1^β	0.134	0.714
H_1^{d-b}	0.705	

The top of the table shows the LR statistics and P values. Standard errors are in parenthesis below values of $\hat{d}$ and $\hat{b}$. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is 242.

Therefore, in the latter study, the authors used another short-term rate, i.e., the key policy rate. However, in our case, we obtain a higher value than that proposed by them, which implies that there is a loss in the control of these monetary policies. In addition, our results have been obtained, including a larger time horizon, which covers a period that spans until today. In sum, table 5.6 summarizes all of the abovementioned scenarios, describing the new scenarios not considered until now.

TABLE 5.6: Policy implementations scenarios

	Order of integration of the error correction term (ECT)		
Value of β	$I(d-b) = I(0)$	$I(0) < I(d-b) < I(0.5)$	$I(0.5) < I(d-b) < I(1)$
$\beta = 1$	The policies duration is **short-lived.**	The policies duration is **long-lived.**	The ECT follows a **non-stationary** process, although mean-reverting and policies durations are **long-lived.**

The shaded area corresponds to the best scenario for policy implementations. As $\beta = 1$, the ECT is assumed as the spread between both interest rates.

The next step, according to the existing cointegration literature, consists of testing the significance of the adjustment coefficients in the joint hypothesis, $H_1^\beta \cap H_1^{\alpha_i}$, using an LR test, and we find that only the coefficients associated with short-term rates (α_{EON}) are significant (table 5.7), which implies that the spread has a prediction power in the behaviour of the future short-term rates, which is consistent with the EHTS.

Finally, in step 5, referring to table 5.7, where the mirror hypothesis is shown, we decompose the common trend to determine if the 3-month Euribor rate or Eonia rate drives the common trend. In our case, the 3-month Euribor rate does not contribute to the long-run rate because the parameter α_{EUR} is not zero. On the other hand, the parameter $\alpha_{EUR} = 0$, such that the Eonia rate is weakly exogenous, is a permanent component, which implies that this rate drives the common trend. Thus, movements in the Eonia rate can precipitate a change in the 3-month Euribor rate until a new common trend is established. This finding conforms to previous empirical findings proposed by Cossetti and Guidi (2009) and Tamakoshi and Hamori (2014), who support the existence of a long-run relationship between both interest rates.

Regarding P-T decomposition, a shock in the Eonia rate will have a permanent (long-run) effect on Eonia and Euribor, but a shock in the 3-month Euribor rate, with no movement in

TABLE 5.7: FVECM results under constrained parameter (1, -1)

Hypothesis	LR statistics	P value
$H_1^{\beta} \cap H_1^{\alpha_{EUR}} \equiv H_1^{\beta} \cap H_1^{\alpha_{\perp Eon}}$	10.732	0.001
$H_1^{\beta} \cap H_1^{\alpha_{Eon}} \equiv H_1^{\beta} \cap H_1^{\alpha_{\perp EUR}}$	0.922	0.337
α_{EUR}	-0.723	
α_{Eon}	0.160	
Component Share		
$CS_{EUR}(\alpha_{\perp EUR})$	0.181	
$CS_{Eon}(\alpha_{\perp Eon})$	0.819	

In the field of hypothesis we reference the mirror hypothesis. The top of the table shows the LR statistics and P values. Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is 242. $\alpha_{\perp EUR}$ and $\alpha_{\perp Eon}$ are normalized such that the two elements add to one.

the Eonia rate, is completely transitory. In addition, we found that the Eonia rate remains fixed at any change in the 3-month Euribor rate, so this change will affect the spread $(R_t - r_t)$ only through z_t (transitory component) and, therefore, will only have transitory effects. In sum, we show that the 3-month Euribor rate does not contribute to the long-run rate, so the Eonia rate is the dominant rate. This can also be interpreted as both interest rates contributing to the common trend. As we can see in the bottom of table 5.7, where the component share is presented, common trend proportions are estimated at 18.1% and 81.9% for the 3-month Euribor rate and Eonia rate, respectively. Thus, we find that the Eonia rate dominates in the common trend.

5.5 Conclusion

It is well known that the Eonia rate plays a crucial role in signalling the target of monetary policy, while the Euribor rate provides outstanding interest rates for various financial products. In this sense, we first aimed to contrast the usefulness of monetary policies through the relationship between 3-month Euribor interest rate and the Eonia rate. Our approach using the Eonia rate could be used in the ECB's policy. In this context, Cossetti and Guidi (2009), among others, warn that the Eonia rate marks the first step of the monetary policy transmission process, and they show that the Eonia rate is highly correlated with the monetary policy rate, i.e., the Eonia rate could be a proxy of it. To complete our empirical strategy, we have analysed the Eonia spread using a novel approach, i.e., we use a FCVAR model to determine the long-run relation between these two interest rates and to find monetary policy evidence according to the persistence of the Eonia spread. We also analysed the effect of monetary policy using P-T decomposition.

Accordingly, the Eonia rate acts as a useful tool for the implementation of monetary policy, which would allow for checking the correct functioning of the monetary policy transmission mechanism. In this respect, the proposal of Soares and Rodrigues (2013), who warn about the usefulness of the 3-month Euribor to contrast the real effects of changes in rates on the real economy, is recalled. Essentially, the measure of a short-term rate versus a very short-term rate would allow for testing the validity of the EHTS; the EHTS argues that different types of maturities are related to each other. Regarding this, the FCVAR model is the only technique that permits the testing of the relationship between interest rates with different maturities, which could be long memory and even nonstationary, providing a novel set of results on this topic. In addition to measuring the long-run relationship and its characteristics, by using Permanent-Transitory decomposition, the interest rate that drives the relationship is known.

Overall, our primary results support that the EHTS is confirmed, denoting a long-run relationship between the Eonia rate and the 3-month Euribor. Subsequently, and even more importantly, the spread between the Eonia rate and 3-month Euribor rate follows a non-stationary but mean-reverting process, which shows that any shock over this would be long-lived. In other words, any shock affecting this relationship will involve more extensive adjustment processes over time. The greater persistence in money market rates may further indicate that it is more difficult for monetary policy signals to be transmitted along the money market yield curve. If policy spreads are highly persistent, the lasting impact of shocks may impede the transparency of policy signals and, thus, the central bank's impact on longer-term rates, implying a gradual loss of control power over interest rates by the ECB. Thus, our political recommendation derived from these results warns that, although the ECB has monetary policy tools linked to interest rates, the transmission mechanism of these policies is not guaranteed to be immediate. Indeed, the analysis of Permanent-Transitory decomposition reveals that the Eonia rate is the dominant rate in the relationship, i.e., it drives the common trend. This result has an important implication for policy makers because if the ECB wants to keep the interest rate under control, it must contemplate the evolution of the Eonia rate.

Going forward, future research concerning this topic might be concerned with the implications of the implementation of Quantitative Easing by the ECB in 2015. However, central banks can use rates to promote lending and prevent inflation by reducing rates. Nevertheless, the results of this measure would be very different from those of conventional credit expansion policy. In this sense, Herbst, Wu, and Ho (2014) asseverate that, after the global financial crisis, this issue occurred, known as a "reserve trap".

Chapter 6

The EHTS and the persistence in the spread reconsidered. A fractional cointegration approach.

6.1 Introduction

Since the financial crisis suffered in the USA at the end of the 2000s, the country's economic growth has been increasing steadily. Moreover, while the Fed has established an interest rate of 2.25% (December 2018) to inject more liquidity into the economy and to stimulate prices, the dollar has suffered downward pressure. Nonetheless, a gradual incremental increase in interest rates would involve the announcement of monetary normalization and would therefore denote a positive signal emerging from an extremely unique situation. In this case, rates would increase through an upward effect on short-term rates, which would gradually spread to long-term interest rates. If this rise in interest rates corresponds to inflationary pressure, negative impacts should be mitigated.

According to this framework, in a competitive financial environment, the term structure should move in assembly with predictions of the Expectation Hypothesis of term structure[1] (EHTS hereafter) such that returns respond to international market forces, as the term structure of interest rates has always been viewed as crucial to assessing the impact of monetary policy and its transmission mechanisms. Indeed, when a monetary policy is effective, changes in short-term policy interest rates should impact long-term ones (Holmes et al., 2015). However, the EHTS can be varied due to changes in the economy, i.e., changes in monetary policy or the financial market. Thus, potential unsteadiness in the relationship between short- and long-term interest rates could produce confusing results (Esteve, Navarro-Ibáñez, and Prats (2013)).

The literature of this topic has attempted to demonstrate the EHTS based on different contexts and methodologies, and the data reveal contradictory results for the USA. Thus, the EHTS is accepted as a forecasting tool (Poole et al., 2002) for its economic implications for monetary policy (Weber and Wolters, 2012, 2013). However, there is evidence of cases for which this hypothesis does not hold; for example, when applied to G7 countries, the EHTS is supported for all countries except for the USA (Hardouvelis, 1994). Finally, it is well known that linear cointegration provides less power and fails to detect a long-run relationship between short- and long-term interest rates (Araç and Yalta, 2015).

The EHTS is often tested using cointegration techniques or by imposing stationary conditions on the spread (Li and Davis, 2017). However, the fractionally cointegrated VAR (FCVAR hereafter) allows one to sidestep some limitations of standard cointegration and of the stationarity of the spread, i.e., it is integrated at order zero. The FCVAR model has been

[1] In this chapter, we use the term EHTS as Expectation Hypothesis of term structure, following the common use in the literature.

applied in reference to financial markets and political economics. Rossi and De Magistris (2013) used this methodology to study the relationship between spot and futures markets, and Caporin et al. (2013) applied the FCVAR model on high and low prices to predict stock prices. Jones et al. (2014) examined the fractional relationship between Canadian political support and macroeconomic variables. Dolatabadi et al. (2016) and Dolatabadi et al. (2018) applied the FCVAR model for the analysis of price discovery in commodity spot and futures markets and, more recently, Gil-Alana and Carcel (2018) did the same for exchange rates. This methodology is useful in that it allows us to test cointegration between interest rates of different maturities and spread stationarity simultaneously, unlike what is possible with the traditional cointegration method; with the traditional method, different studies have executed this exercise separately; they study cointegration relationships between interest rates or the stationarity of the spread (see Vides, Iglesias, and Golpe (2018) for a survey).

Thus, regarding the Expectations Hypothesis of Term Structure, this is the first study in which the FCVAR model is applied, providing significant results for practitioners and policy makers. This paper is novel in that it recognizes that premises of standard cointegration testing ($I(1)/I(0)$ dichotomy) time series variables integrated at order one and comoved at order zero are too restrictive, i.e., linear combinations of $I(1)$ nonstationary processes are $I(0)$ stationary. In this sense, the empirical literature has shown that many economic and financial time series hold long-range dependence in the autocorrelation function but do not precisely exhibit a unit root process, i.e., the long memory process.[2] For this reason and according to our research, we reject traditional cointegration assumptions that all interest rates cannot move away from one another for long periods of time and that they are unit roots or $I(1)$; they follow dichotomy $I(0)/I(1)$, such that they follow a fractional process $I(d)$. We also discard the notion that the error term follows a stationary process ($I(0)$) in cases of the cointegration of both variables. In turn, the rigidity of the traditional approach is overcome in favor of allowing the series to be cointegrated, and the error term does not necessarily need to be $I(0)$; for example, we allow the error term to be cointegrated in order $I(d - b)$, unlike other techniques which assume that the error term is $I(0)$ such as Hansen and Seo's Threshold cointegration or Gregory-Hansen test, the Kejriwal-Perron test, the Hatemi-J test and the Arai-Kurozumi test for structural breaks. To the best of our knowledge, the EHTS based on investors exploiting arbitrage opportunities under empirical approaches usually assumes that the relation between short- and long-term interest rates follows a $I(0)/I(1)$ process; however, fractional cointegration may contradict this assumption such that, in the presence of a unitary long-run relationship between interest rates with different maturities, shocks that affect this cointegration relationship can be long-lived and even non-stationary. In turn, this methodology allows one to establish different scenarios of EHTS fulfillment. Therefore, our new approach uses the FCVAR model developed by Johansen and Nielsen (2012) and Nielsen and Popiel (2016), which is an expansion of the traditional cointegrated VAR (CVAR hereafter) proposed by Johansen (1995), enabling us to establish the number of equilibrium relations via cointegrating rank testing to estimate memory parameters, long-run cointegrating relations with adjustment parameters, and short-run lagged dynamics. Finally, our primary results suggest that the EHTS is accepted, but it is accepted to different degrees depending on the maturity of interest rates based on different scenarios when maturity values are greater or less than 5 years. These results highlight the need to relax the rigid assumptions of traditional cointegration (CVAR), as they have led us to noncontemplative scenarios.

The rest of the chapter is organized as follows. The following section 6.2 presents a brief review of the literature, and section 6.3 describes the methodology used for the work. Section 6.4 discusses our empirical results and conclusions, from which we identify the economic policy implications given in section 6.5.

[2] Baillie and Bollerslev (1994a), Baillie (1996), Dueker and Startz (1998) and Mohanty, Peterson, and Smith (1998) among others noted that fractional processes may better describe the long memory of macroeconomic variables.

6.2 A brief literature review

The term structure has long been established as reflecting economic agents' anticipations of future events and as an indicator of monetary policy in academic articles on term structures written over the past century, which serves as testament to the practical importance of this topic and to its intrinsic academic appeal (see Vetzal (1994) for a survey).

The EHTS shows the relationship between short- and long-term interest rates and is the most influential theory explaining term structure relations. The hypothesis states that long-term interest rates are determined by an average of the current and expected short-term interest rate. Thus, this relationship between short- and long-term interest rates implies that their spread contains significant information on future changes in short-term rates and plays an important role in the potential effectiveness of monetary policy, which involves the control of short-term policy rates by central banks and economy affected by monetary impulses through long-term interest rate movements (Bernanke and Blinder, 1992).

As it has been mentioned, the EHTS has economic implications for macroeconomics and finance and for the shape of the yield curve (see Shiller and McCulloch (1990) for a survey). According to the EHTS, an upward sloping yield curve implies that future short-term rates are expected to rise. Conversely, under a downward sloping yield curve, future short-term rates are expected to fall; for example, the slope of the yield curve serves as an important source of information on real economy evolution. As a consequence, Estrella and Hardouvelis (1991) found that a positive curve slope is associated with future increases in real economic activity using macroeconomic variables and providing significant predictive power. One implication of the EHTS as noted by Fama (1984) and Fama and Bliss (1987) relates to the fact that the forward rate is an unbiased predictor of future spot rates. As another implication of this hypothesis, the spread between the long-term interest rate and short-term interest rate –the term spread– is an unbiased predictor of future short-run changes in long-term rates (Mankiw, 1986; Campbell and Shiller, 1991; Campbell, 1995).

In focusing in the most important economy in the world, i.e., the USA, empirical research on the EHTS has been long been conducted, and two strands of inquiry provide support or arguments on this issue. Several studies find evidence in support of the EHTS, e.g., Campbell and Shiller (1987) found partial support for the present value of the term structure of interest rates by examining the statistical significance of the EHTS. Additionally, changes in regimes of short- and long-term rates of US treasury bills support the EHTS (Hamilton, 1988), cointegrating US treasury bills among interest rates (Hall et al., 1992). Further, Poole and Rasche (2000) and Poole et al. (2002) demonstrated that the market is able to anticipate changes in the FED's target federal fund rate. There is also evidence in support of the EHTS on the relationship between short- and long-term rates for Europe and the USA (Lanne, 2003; Brüggemann and Lütkepohl, 2005), thus combining yield factors and macroeconomic variables to relate to the EHTS, serving as evidence in favor of certain regimes (Diebold et al. (2006)). Weber and Wolters (2012, 2013) applied the vector error correction (VEC) for the US term structure to offer an economic explanation for deviations from the EHTS. More recently, Holmes et al. (2015) examined the term structure of interest rates using a pairwise stationary approach and showed that the EHTS holds over the long-term, i.e., short-run policy changes affect long-term rates.

However, others present arguments against the EHTS for the USA such as Hardouvelis (1994), who used a VAR model designed to forecast changes in long-term interest rates using the term spread of G7 countries and who showed that the EHTS is supported in all countries except for the USA. Nevertheless, Thornton (2005) tested the EHTS for federal funds rates to find that the market's expectations are less able to forecast federal funds rates. Conversely, Guidolin and Thornton (2010) concluded that future short-term rates have deep implications for policy makers, suggesting that whether or not the EHTS is true, an inability to predict the future short-term rate would imply that both long- and short-term rates are equal and that this relation is inconsistent. In turn, the conventional theory of the term structure of interest rates is threatened. Finally, Bulkley et al. (2011, 2015) identified the failure of the EHTS based on bond yields of US Treasury securities. Despite initial controversy, this literature has shown that it is possible to establish a relationship between short- and long-term rates.

From an empirical point of view, the issue has best been explained through the concept of cointegration provided by Engle and Granger (1987). Several studies have adopted a linear adjustment process, i.e., interest rates maintain a long-term equilibrium relationship such that the interest spread does not tend to increase or decrease over time, reverting to its mean (see Engle and Granger (1987); Campbell and Shiller (1987); and Shea (1992), for instance). Engsted and Tanggaard (1994) studied the EHTS for the US term structure and concluded that the EHTS cannot be rejected using a VECM. Hansen and Seo (2002) also used a threshold VECM to demonstrate that their results are roughly consistent with the term structure selected, and Seo (2003) utilized a trivariate threshold VECM and found evidence in support of nonlinear mean-reversion in the term structure of interest rates. Meanwhile, Esteve et al. (2013) studied cointegration with multiple structural breaks for Spain, and evidence of the EHTS was not found. It is well known that linear cointegration provides less power and fails to detect a long-run relationship between short- and long-term interest rates (Araç and Yalta, 2015); thus, several authors, such as Clarida et al. (2006) and Mili et al. (2012), show nonlinearities in the relationship between interest rates. Finally, Lange (2018) shows that the Canadian term spread is mean-stationary and related to macroeconomics. From this premise, our objective is to test the EHTS with a nonlinear cointegrating approach and to demonstrate robustness with different tests.

In focusing on the spread there are some arguments in its treatment in the sense that there are two main ways to check it. On one half, Baillie and Bollerslev (1994a, 1994b), Tkacz (2001) and Cassola and Morana (2008) among others suggest that the spread could follow a fractional order of integration. On the other half, Strohsal and Weber (2014) and Holmes et al. (2015), for instance, show that the spread degree of integration could be different from $I(0)$ but while supporting the EHTS. Although we have studied cointegration and spread from a fractional perspective separately, our approach allows us to analyze them together.

In other words, the analysis demonstrate a new approach that involves using the fractionally cointegrated vector autoregressive (FCVAR) model developed by Johansen and Nielsen (2012) and Nielsen and Popiel (2016), which is an expansion of the CVAR proposed by Johansen (1995) and which enables us to establish the number of equilibrium relations with a cointegrating rank test to estimate memory parameters, long-run cointegrating relations with adjustment parameters, and short-run lagged dynamics. This econometric approach allows for jointly testing the existence of a long-run relationship between short- and long-term interest rates and spread persistence.

6.2.1 Did Quantitative Easing affect the term structure of interest rates?

This subsection focuses on events that have recently occurred in the USA and on measures applied by the Fed. At the start of the 2000s, the longest growth period in the history of the United States ended due to a fall in investment caused by the collapse of the dot-com bubble and the September 11 attacks. This situation was reversed with the implementation of painful fiscal adjustments and from the costs of wars in Afghanistan and Iraq (Kraay and Ventura, 2005). Afraid that the economy would slip back into recession, the Fed kept the federal fund rate extremely low, reaching a low of 1% by the middle of 2003. As the momentum and prices of joint expansion began to rise, the target of federal funds increased slowly in a series of movements to 5.25% in mid-2006 (Labonte and Makinen, 2008). Finally, by the end of 2007 the subprime mortgage market collapsed and quickly spread to the rest of the world. The US government responded with a fiscal stimulus package and unprecedented bank bailout; in spite of the NBER declaring a recession for more than a year after the end date (June 2009), the measures applied by the Fed according to the Federal Reserve of Sant Louis correspond to Quantitative Easing (QE, hereafter) programs, which focused primarily on the type and quantity of asset acquisition that would affect financial market conditions, inflation and ultimately economic activity (Williamson, 2017). This QE program was announced on November 25, 2008, and it involved three main phases (or QE) occurring between 2009 and 2014, as follows:

- The QE1 lasted from December 2008 to March 2010, and $175 billion in agency securities and $1.25 trillion in agency mortgage-backed securities (MBS) were purchased.
- The QE2 spanned November 2010 to June 2011, during which time $600 billion in long-maturity Treasury securities were acquired. From September 2011 to December 2012, this acquisition was applied to the so-called *Operation Twist*. The measure involved a swap of more than $600 billion consisting of the acquisition of Treasury securities with maturities of six to thirty years and the sales of Treasury securities with maturities of three years or less.
- Finally, the QE3 covered the period of September 2012 to October 2014 and was based on the purchase of MBS and long-term Treasury securities initially set at $40 billion per month for MBS and $45 billion per month for values of long-term Treasury securities.

Until September 2017 the Fed began to very gradually reduce the balance sheet to a more typical value. However, Wright (2012) has shown that these announcement effects were short lived, lasting only a few months.

Simultaneously the Fed faced the subprime mortgage crisis by reducing the federal funds target from 5.25% to a range of 0% to 0.25% in December 2008, which economists call the zero lower bound. From this monetary policy measure, economic expansion and the unemployment rate were valued at close to the Fed's estimate of full employment when it began raising rates in December 2015. The Fed has since continued to raise rates more slowly than it initially intended over a series of steps to incrementally tighten monetary policy. The Fed raised rates once in 2016 and three times in 2017 by 0.25 percentage points each time. Fed has raised the federal funds rate three times this year to a range between 2% and 2.25%.

Table 6.1 summarizes the abovementioned events to provide clearer review of the discussed measures.

Table 6.1: Summary of Fed measures

Year	**Event**	**Measure**
Early 2000s	Dot-com bubble and September 11^{th} attacks	A painful fiscal adjustment due to the cost of the Afghanistan and Iraq wars. Fed causes federal fund rates to reach a low of 1% by mid-2003
Mid-2006	Economic expansion	Fed increases federal funds to 5.25%
Mid-2007	Subprime mortgage crisis	Fiscal stimulus and bank bailout
Nov. 2008		Announcement of Quantitative Easing program
Dec. 2008		Fed establishes federal fund target to a range of 0 – 0.25% (zero lower bound)
Dec. 2008	Quantitative Easing 1	Purchase of $175 billion in agency securities and of $1.25 trillion in agency MBS
Nov. 2010	Quantitative Easing 2	Acquisition of $600 billion in long-maturity Treasury securities
Sept. 2012	Quantitative Easing 3	Purchase of $40 billion per month for MBS
Dec. 2015	Economic expansion and unemployment rate close to the Fed's estimate	Fed raises rates once in 2016 and three times in 2017 by 0.25 percentage points each time. At present, federal fund rates range from 2 – 2.25%.

6.3 Methodology

Our econometric strategy involves obtaining and analyzing the model estimation at a monthly frequency, and we then perform statistical tests of cointegration, exclusion and weak exogeneity based on the fundamental equation for the EHTS in an econometric context.

6.3.1 The EHTS model

The fundamental equation of the EHTS of an $n > 1$ period bond R_t (i.e., long-term interest rate) is equal to an average of the current and expected r_t (i.e., short-term interest rate) set of $n \leq 1$ period plus a constant term. The relationship can be expressed in the following form:

$$R_t = \frac{1}{n}\sum_{k=0}^{n-1} E_t[r_{t+k}] + \phi_t^*, \tag{6.1}$$

where ϕ_t^* is a possible stationary term and E_t is the expectation operator at time t for the evolution of short-term interest rates driving long-term interest rates. To test the EHTS in the context of cointegration theory, the common linear mode used is:

$$R_t = c + \beta r_t + \varepsilon_t \tag{6.2}$$

In agreement with Campbell and Shiller (1987), R_t and r_t should be non-stationary and related through a cointegration relationship with parameters (1,-1). This implies that β_R and β_r are cointegrated constants and that their combination involves a stationary process while the spread of the interest rate reverts to the mean. When the spread is stationary, the long- and short-term rates are driven by a common stochastic trend and do not allow for arbitrage opportunities because market forces adjust to correct any temporary disequilibrium. As the EHTS suggests, the interest rate spread is an optimal forecast[3] of future changes in long-term interest rates. Thus, the market's expectations of the short-rate developments in the bond yield are reflected in the slope of the term structure with a one-to-one relation, $\beta = 1$.[4]

According to the existing literature, cointegration implies a VECM such as:

$$\begin{pmatrix} \Delta R_t \\ \Delta r_t \end{pmatrix} = \begin{pmatrix} \alpha_R \\ \alpha_r \end{pmatrix} (R_{t-1} - \beta r_{t-1} - c) + \sum_{i=1}^{n} \Gamma_i \begin{pmatrix} \Delta R_{t-i} \\ \Delta r_{t-i} \end{pmatrix} + \begin{pmatrix} w_{1t} \\ w_{2t} \end{pmatrix} \tag{6.3}$$

with adjustment parameters α, cointegration coefficient β, restricted constant (c), lag length (n) and errors w. Γ_i are 2×2 parameter matrices in the short-run dynamics. The adjustment coefficients α_R and α_r capture the speed of adjustment of R_t and r_t towards equilibrium.

For this work the FCVAR model allows us to study the common long-run equilibrium relationship between long- and short-term interest rates. The model is a generalization of Johansen (1995)'s cointegrated vector autoregressive (CVAR) model to allow for fractional processes of order d that cointegrate to order $d\breve{}b$. As we conduct our analysis using a bivariate fractional cointegration approach, we recognize that the standard unit root and cointegration test may be too restrictive ($I(1)/I(0)$ dichotomy). The CVAR and FCVAR differ in part because for the CVAR, the error correction term (the spread when the EHTS is supported) is $I(0)$, while for the FCVAR, this assessment is not restricted (the integration order may be different from zero, reflecting a long-memory process); thus, the assumption that the spread term of $I(0)$ could reflect an $I(d)$ process is rejected. A more generalized $I(d)$-type specification has been adopted based on the possibility of fractional orders of integration.

[3]Baillie and Bollerslev (1994a) discovered that cointegrating relationships may not be precisely valued at $I(0)$, implying that a fractionally cointegrated relationship may yield noticeable gains in forecast accuracy only within the context of a longer-term forecast.

[4]When $\beta = 1$ we assume that the difference between short- and long-term interest rates is the term spread.

Cointegration without these values is unrestricted.

Accordingly, the FCVAR model allows us to identify several degrees or scenarios of EHTS fulfillment. Once the test shows that there is cointegration, the degree of integration in the spread allows us to identify up to three different scenarios. This idea is illustrated more fully in the following subsection and is illustrated in table 6.3, in which new possibilities allowed through the application of the FCVAR as a generalization of traditional approaches are broken down.

TABLE 6.2: New possibilities arising when applying the FCVAR

		Cointegration/Long-run relationship	
		Yes	*No*
Spread	Stationary	**Several degrees of the EHTS (see table 6.3)**	Controversy
	Nonstationary		No EHTS

6.3.2 Fractional cointegration model – FCVAR methodology

This model is presented in Johansen (2008a, 2008b) and is further developed in Johansen and Nielsen (2012) and Nielsen and Popiel (2016). It offers the advantage of being applicable to stationary and non-stationary time series. Our objective is to study the EHTS under fractional cointegration conditions.

To introduce the FCVAR model, we begin with the well-known, non-fractional, CVAR model. Being $Y_t = 1, \ldots, T$ a p-dimensional $I(1)$ time series. Therefore, the CVAR model is:

$$\Delta Y_t = \alpha\beta' Y_{t-1} + \sum_{i=1}^{k} \varGamma_i \Delta Y_{t-i} + \varepsilon_t = \alpha\beta' L Y_t + \sum_{i=1}^{k} \varGamma_i \Delta L^i Y_t + \varepsilon_t \tag{6.4}$$

The fractional difference operator is Δ, and the fractional lag operator is $\Delta = (1 - L)$. We replace lags operators with fractional counterparts Δ^b and $\Delta^b = (1 - L^b)$, and apply $Y_t = \Delta^{d-b} X_t$ such that:

$$\Delta^b Y_t = \alpha\beta' L_b Y_t + \sum_{i=1}^{k} \varGamma_i \Delta^b L_b^i Y_t + \varepsilon_t, \tag{6.5}$$

As always, ε_t is p-dimensional independent and identically distributed with mean of zero and covariance matrix Ω. Parameters α and β are $p \times r$ matrices, where $0 \leq r \leq p$. In matrix β, the columns denote cointegrating relationships and $\beta' X_t$ assumes the existence of a common stochastic trend integrated at order d. Short-term parts from the long-run equilibrium are integrated at order $d - b$, but when $d - b < 1/2$ asymptotically, a zero-mean stationary process occurs. The coefficients of α correspond the rate of adjustment to equilibrium. Therefore, $\alpha\beta'$ is the adjustment long-run, ρ' is the restricted constant term, and Γ_i represents the short-run behaviour of the variables. We in turn the final model:

$$\Delta^d X_t = L_d \alpha(\beta' X_t + \rho') + \sum_{i=1}^{k} \Gamma_i \Delta^d L_d^i X_t + \varepsilon_t. \tag{6.6}$$

When the VAR model is applied with $d = b = 1$ (CVAR), X_t is integrated at order d, and b denotes the strength of the cointegrating relationships (as the value of b increases persistence declines in cointegrating relationships). The error correction term is integrated from order $(d-b)$, which is $I(0)$ in the case of standard cointegration $(d = b = 1)$. However, in fractional cointegration these axioms are relaxed because $(d-b) = 0$, i.e., the error correction term shows a short-run stationary behaviour, or because $(d - b) > 0$, i.e., there is a long memory process, and the error correction term will revert to its mean over the long run. As the cointegrating vector is $(1, -\beta)$, we can interpret the difference $(d - b)$ as the order of integration for the cointegrating error or as the degree of persistence (H_1^{d-b}). According to Table 1 given in Tkacz (2001), when $(d-b)=0$, the cointegrating error follows a stationary process, and the shock duration is short-lived. When $0<(d-b)<0.5$, a stationary process occurs, and the shock duration is long-lived. Finally, when $0.5<(d-b)<1$, the cointegrating error involves a non-stationary process while mean-reverting, and the shock durations are long. Thus, one of the main contributions of this paper lies in its elaboration of new conditions of the degree of EHTS fulfillment while establishing β as a "different condition of relationship strength" and while focusing on the persistence of the error term. These conditions are synthesized in table 6.3.

TABLE 6.3: EHTS scenarios

Order of integration of the error correction term	**Value of β (assuming cointegration)**	
	$\beta=1$	$0<\beta<1$
$I(d-b)=I(0)$	Theoretical EHTS and the shock duration is **short-lived.**	Weak EHTS and the shock duration is **short-lived.**
$I(0)<I(d-b)<I(0.5)$	EHTS and the shock duration is **long-lived.**	Weak EHTS and the shock duration is **long-lived.**
$I(0.5)<I(d-b)<I(1)$	EHTS and the ECT follows a **non-stationary** process, although mean-reverting and shock durations are **long-lived.**	Weak EHTS and the ECT follows a **non-stationary** process, although mean-reverting and shock durations are **long-lived.**

The shaded area corresponds to the traditional EHTS.

Johansen and Nielsen (2012) show that the maximum likelihood estimators $(d,\alpha,\Gamma_i,\ldots,\Gamma_k)$ are asymptotically normal and the maximum likelihood estimator of $(\beta,\,\rho)$ is asymptotically mixed normal.

To determine the number of stationary cointegrating relations following the hypotheses of the rank test based on a series of LR tests, in the FCVAR model, we test the hypothesis $H_0 : rank(\Pi) = r$, against the alternative: $H_1 : rank(\Pi) = p$. $L(d,b,r)$ is the profile likelihood function of rank r, where (α,β,Γ) is reduced by rank regression (see Johansen and Nielsen, 2012). For the model with a constant, we test $H_0 : rank(\Pi,\mu) = r$ against the alternative: $H_1 : rank(\Pi,\mu) = p$, and the profile likelihood function given rank r is $L(d,r)$, where parameters $(\alpha,\beta,\rho,\Gamma)$ are focused out.

To maximize the profile likelihood distribution under both hypothesis, the LR test statistics are now $LR_t(q)$. The asymptotic distribution of $LR_t(q)$ depends on parameter b and on $q=n-r$. MacKinnon and Nielsen (2014) based on their numerical distribution functions on the asymptotic critical values of an LR rank test. In cases of "weak cointegration", i.e., $0<b<1/2$, $LR_t(q)$ has a standard asymptotic distribution,$LR_t(q)$ $LR_t(q) \xrightarrow{D} \chi^2(q^2)$.

6.4 Empirical analysis

The purpose of the present study is to test the existence of EHTS values with 3 to 240 months of maturity. The first step here involves testing the existence of a common trend, i.e., a long-run relationship between short- and long-term rates. For this cointegration analysis, we determine if fractional cointegration is more appropriate than standard cointegration. As a second step, we study the potential relation one to one, i.e., the cointegrating vector (1,-1). The next step involves analyzing the adjustment coefficients, and finally we examine the fractional cointegration degree (persistence).

TABLE 6.4: Empirical research method

	Procedure	Hypotheses
Step 1	Fractional cointegration?	H_1^d: Is the fractional cointegration more appropriate that traditional cointegration?
Step 2	Estimation of β	H_1^{β}: Cointegrating vector is (1, -1)
Step 3	Estimation of adjustment coefficients (α_R, α_r) (FVECM)	$H_1^{\beta} \cap H_1^{\alpha_i}$: Variables are weakly exogenous under the restrictions of the cointegrating vector (1, 1)
Step 4	Degree of spread persistence, i.e., order of integration $(d-b)$	H_1^{d-b}: Is the spread a long memory process?

6.4.1 Data description

For our empirical analysis, we employ a monthly sample of Treasury Constant interest rates of 9 different maturities for the period of October 1993 to December 2018 (amounting 303 observations for each interest rate series). The data correspond to 3-month, 6-month, 1-year, 2-year, 3-year, 5-year, 7-year, 10-year and 20-year constant maturity rates. The data are gathered from Federal Reserve Economic Data (FRED) assembled by the Economic Research Division of the Federal Reserve Bank of St. Louis. As 1-month Treasury Constant maturity rate data are only available from January 2001, we use these maturities to determine the availability of consistent interest rate data for the period studied. Interest rates are measured as percentages and are shown in figure 6.1. We consider 3-month, 6-month and 1-year periods as short-run periods. A period of 1 year is defined as short-term to render our estimation more robust.[5] On the other hand, we define the rest of the maturity rates as long-run. Table 6.5 shows descriptive statistics associated with each interest rate for different maturities. In terms of volatility, the variables exhibit similar behavior and figure 6.1 presents a graphical analysis of time series dynamics traced for all maturities.

TABLE 6.5: Descriptive statistics for the data

	3-month	6-month	1-year	2-year	3-year	5-year	7-year	10-year	20-year
Mean	2.459	2.597	2.729	3.005	3.217	3.607	3.922	4.168	4.696
Median	1.760	1.970	2.220	2.610	2.870	3.360	3.790	4.150	4.810
Min	0.010	0.040	0.100	0.210	0.330	0.620	0.980	1.500	1.820
Max	6.360	6.510	7.140	7.590	7.710	7.780	7.830	7.960	8.200
S.D.	2.180	2.221	2.217	2.193	2.109	1.918	1.783	1.566	1.566

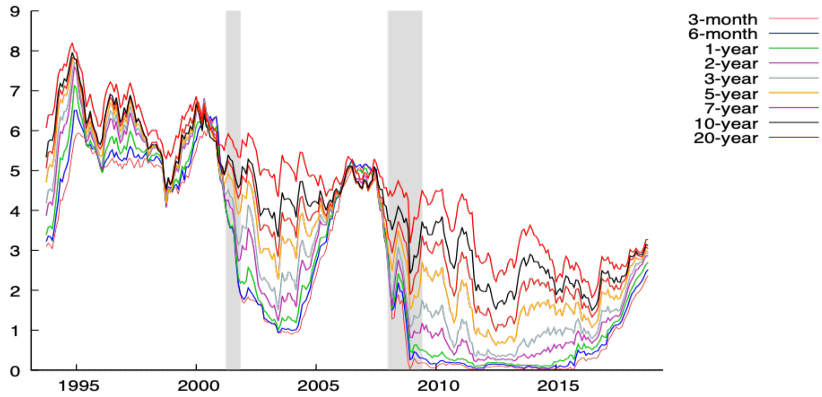
From 10/1993 to 12/2018

6.4.2 Univariate analysis

As a preliminary step we estimate the order of the fractional integration of the interest rates. There are several means of estimating the fractional differencing parameter in semiparametric contexts. Although the semiparametric log-periodogram regression method proposed by Geweke and Porter-Hudak (1983) is the most widely used, the method was modified and further developed by Robinson (1995) and has been analyzed by Velasco (1999) and Shimotsu and Phillips (2002) among others. To develop a fractionally cointegrated model, we first discuss long memory univariate results; then, we proceed to the estimation of fractional parameter d for each univariate series with results presented in table 6.6. The three columns show semiparametric log-periodogram regression estimates drawn from Geweke and Porter-Hudak

[5] We also estimate pairs of short-run maturities, i.e., 3-month – 6-month, 3-month – 1-year and 6-month – 1-year.

FIGURE 6.1: Time series traced for Treasury Constant interest rates of different maturities.



(1983) computed with bandwidths of[6] $m = T^n$, which is equivalent to Fourier frequencies, where m is the integer, T is the number of observations and n is the bandwidth size. The estimates are consistent with joint estimates presented below. As is shown in table 6.6, the values for d decrease as maturities increase, becoming mean-reverting values.

TABLE 6.6: Univariate analysis. GPH estimates

	3-month	6-month	1-year	2-year	3-year	5-year	7-year	10-year	20-year
$m = [T^{0.4}] = 9$	1.216	1.266	1.313	1.335	1.243	1.018	0.871	0.678	0.658
	(0.298)	(0.316)	(0.306)	(0.319)	(0.247)	(0.187)	(0.233)	(0.389)	(0.213)
$m = [T^{0.5}] = 17$	1.491	1.480	1.377	1.247	1.191	1.069	0.936	0.775	0.776
	(0.183)	(0.189)	(0.163)	(0.168)	(0.162)	(0.163)	(0.157)	(0.208)	(0.157)
$m = [T^{0.6}] = 30$	1.467	1.388	1.309	1.151	1.069	0.968	0.889	0.794	0.838
	(0.107)	(0.107)	(0.099)	(0.101)	(0.098)	(0.098)	(0.098)	(0.118)	(0.106)

GPH denotes the Geweke and Porter-Hudak semiparametric log-periodogram regression estimator. Standard errors are given in parenthesis beneath estimates of d.The sample size is 303

6.4.3 Cointegration analysis

In this subsection, we present the procedure used to derive the results shown below. First, to select the FCVAR lag length, we use AIC criteria so that the lag length selected is different for each pair of variables studied. As can be observed when the short-term interest rate referenced is for a 3-month period, the optimal lag length is six in all cases (table A.1a of the appendix). When 6 months is the reference short-term period, the result is diverse and for 1 year the result ranges between four and five (tables A.1b and A.1c of the appendix).

We then determine that there is a long-run relationship between each pair of maturities selected, and we test the cointegration rank before testing the hypothesis for the fractional parameter. We in turn find that the number of cointegrating vectors is one in almost most of cases (table A.2 of the appendix). Once a rank cointegration test is conducted, we test hypothesis H_1^d, which tests whether fractional cointegration is more appropriate than traditional cointegration (the CVAR model), i.e., the null hypothesis is $d = 1$, and its rejection implies that the FCVAR model is more suitable than traditional cointegration. Accordingly, by assuming $I(1)$ cointegration or an $I(0)$ VAR model, we may be misspecifying the model estimates, parameters, test restrictions and implied dynamics, such as the term spread. In contrast, whether there is a stochastic trend of an order lower than unity, economic shocks

[6]To test the presence of unit roots, estimates were obtained from first-differenced data, as the original series may exceed 0.5, and the test requires that results are limited to an interval of $-0.5 < d < 0.5$ while adding 1 to obtain the proper estimates of d. According to the literature, the bandwidth size ranges from 0.25 to 0.8. For our study, the three bandwidths selected are valued at 0.4, 0.5 and 0.6.

have temporary mean-reverting effects on the relevant variables. This allows for more flexibility when theoretical and macrodynamic principles are applied, and the stochastic long-run term structure trend can in turn be established by shocks with transitory effects on interest rates (Abbritti et al., 2018). In this sense, Table 6.7 shows the results of an LR test and shows that the CVAR is rejected in favor of the FCVAR, i.e., fractional cointegration is appropriate for this study.

TABLE 6.7: H_1^d: LR test, CVAR vs. FCVAR

				Maturities				
3-month vs.	6-month	1-year	2-year	3-year	5-year	7-year	10-year	20-year
H_1^d	10.220 (0.001)	12.808 (0.000)	17.197 (0.000)	28.038 (0.000)	23.276 (0.000)	22.781 (0.000)	20.125 (0.000)	20.387 (0.000)
6-month vs.		1-year	2-year	3-year	5-year	7-year	10-year	20-year
H_1^d		25.650 (0.000)	29.549 (0.000)	25.468 (0.000)	20.811 (0.000)	15.407 (0.000)	19.705 (0.000)	16.308 (0.000)
1-year vs.			2-year	3-year	5-year	7-year	10-year	20-year
H_1^d			16.667 (0.000)	15.051 (0.000)	16.418 (0.000)	16.202 (0.000)	16.536 (0.000)	18.463 (0.000)

Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is 303. LR statistics and P values are in parenthesis below LR test values.

We next estimate the long-run relationship between long- and short-term rates (see equation 6.2). The estimated values are shown in table 6.8. It can be observed that parameter is close to 1. [7]

TABLE 6.8: Cointegrating vector $(1, -\beta)$

				Maturities				
3-month vs.	6-month	1-year	2-year	3-year	5-year	7-year	10-year	20-year
	[1,-1.029]	[1,-1.048]	[1,-1.085]	[1,-1.089]	[1,-1.130]	[1,-1.122]	[1,-1.142]	[1,-1.309]
6-month vs.		1-year	2-year	3-year	5-year	7-year	10-year	20-year
		[1,-1.025]	[1,-0.981]	[1,-0.998]	[1,-1.143]	[1,-1.144]	[1,-1.156]	[1,-1.263]
1-year vs.			2-year	3-year	5-year	7-year	10-year	20-year
			[1,-1.094]	[1,-1.121]	[1,-1.162]	[1,-1.144]	[1,-1.151]	[1,-1.266]

Recalling that the EHTS implies that series are cointegrated while the cointegrating vector between each variable is constrained in (1,-1), H_1^β, we must test the existence of this vector. From the LR test illustrated in table 6.9, we do not reject this parameter restriction when a given maturity exceeds 5 years when using any short-term rate. However, when we select a maturity of 2 or 3 years, the parameter restriction is rejected.[8] The traditional prism of the EH establishes that there must be a linear combination between short- and long-term interest rates constrained by a (1,-1) vector. Nonetheless, from tables 6.2 and 6.3 presenting data on different degrees of maturity, it may be observed that for a given combination of interest rates, the EHTS illustrates a different vision from a traditional perspective. Then, the EH is referred to as "weak", as Esteve et al. (2013) shows. In our study, a "weak-EH" refers to combinations of interest rate pairs with values of $0 < \beta < 1$ based on new scenarios in the conception of this hypothesis. Our results may be interpreted in two ways. On one hand, maturities of over 5

[7]For every estimation, we check residuals for serial correlations using a multivariate Ljung-Box Q-test $Q_{\hat{\varepsilon}}$ with $h = 12$ lags because our data are monthly. The results show no evidence of serial correlations in the residuals for every estimation, and the Ljung-Box Q-test shows no signs of misspecification, indicating that the model is well specified (see table A.3 of the appendix).

[8]We also reject combinations of short-term rates with an LR test.

years meet the abovementioned requirements such that the EHTS is strongly supported due to this level of interest rate maturity. On the other hand, for maturities of less than 5 years, i.e., 1, 2 and 3 years, following Esteve et al. (2013), such results support a weak version of the EHTS for interest rates. For this reason, with the FCVAR as shown in table 6.3, we establish different scenarios.

TABLE 6.9: H_1^{β}: LR test statistics of the hypothesis $\beta = 1$ (cointegrating vector is (1,-1)

	Maturities							
3-month vs.	6-month	1-year	2-year	3-year	5-year	7-year	10-year	20-year
LR test	8.615	8.506	7.300	3.881	1.319	0.746	0.502	1.123
P value	(0.003)	(0.004)	(0.007)	(0.049)	(0.251)	(0.388)	(0.478)	(0.289)
6-month vs.		1-year	2-year	3-year	5-year	7-year	10-year	20-year
LR test		5.245	20.036	9.572	1.193	0.728	0.518	0.710
P value		(0.022)	(0.000)	(0.002)	(0.275)	(0.393)	(0.472)	(0.399)
1-year vs.			2-year	3-year	5-year	7-year	10-year	20-year
LR test			5.185	2.721	1.357	0.769	0.498	0.762
P value			(0.023)	(0.099)	(0.244)	(0.380)	(0.480)	(0.383)

Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is 303.

As our next step, according to the existing cointegration literature, we estimate an FVECM (see equation 6.3) testing the significance of adjustment coefficients of the joint hypothesis, $H_1^{\beta} \cap H_1^{\alpha_i}$, as shown in table 6.10,[9] using an LR test, and we find that only those coefficients associated with short-term rates (α_r) are significant, implying that the spread offers predictive power on the behavior of future short-term rates consistent with the EHTS. Finally, as expected, adjustment coefficients of the short-term rate are positive, which serves as extra evidence in support of the EHTS; conversely, adjustment coefficients of long-term rates are much lower in magnitude than those short-term rates, but the adjustment coefficients are insignificantly different from zero (this finding is based the results of Hansen and Seo (2002)).

[9]We focus on maturities where the EHTS is supported according to results given in table 6.9.

TABLE 6.10: $H_1^{\beta} \cap H_1^{\alpha_i}$: Adjustment coefficients under cointegration vector (1,-1)

					Maturities			
3-month vs.	6-month	1-year	2-year	3-year	5-year	7-year	10-year	20-year
$H_1^{\beta} \cap H_1^{\alpha_R}$	—	—	—	—	0.006	0.032	0.267	0.709
					(0.937)	(0.858)	(0.605)	(0.400)
$H_1^{\beta} \cap H_1^{\alpha_r}$	—	—	—	—	15.440	11.904	12.735	11.740
					(0.001)	(0.001)	(0.000)	(0.001)
α_R	—	—	—	—	0.003	-0.005	-0.014	-0.250
α_r	—	—	—	—	0.124	0.076	0.072	0.083
6-month vs.		1-year	2-year	3-year	5-year	7-year	10-year	20-year
$H_1^{\beta} \cap H_1^{\alpha_R}$		—	—	—	0.097	0.020	0.132	0.738
					(0.755)	(0.887)	(0.716)	(0.390)
$H_1^{\beta} \cap H_1^{\alpha_r}$		—	—	—	10.798	13.407	9.427	12.172
					(0.001)	(0.000)	(0.002)	(0.000)
α_R		—	—	—	0.014	-0.015	-0.009	-0.118
α_r		—	—	—	0.101	0.277	0.056	0.401
1-year vs.			2-year	3-year	5-year	7-year	10-year	20-year
$H_1^{\beta} \cap H_1^{\alpha_R}$			—	—	0.018	0.065	0.306	0.777
					(0.893)	(0.799)	(0.580)	(0.378)
$H_1^{\beta} \cap H_1^{\alpha_r}$			—	—	7.246	7.748	7.698	8.115
					(0.006)	(0.005)	(0.006)	(0.004)
α_R			—	—	0.018	-0.024	-0.051	-0.109
α_r			—	—	0.279	0.202	0.201	0.313

Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is 303.P values are in parenthesis below LR test values.

Finally, as the cointegrating vector is (1, -1), we can interpret the difference $(d-b)$ as the order of integration of the spread or as the degree of persistence (H_1^{d-b}). As stated in the methodology section, when $(d-b) = 0$, the spread follows a stationary process, and the shock duration is short-lived. When $0 < (d-b) < 0.5$, there is a stationary process, and the shock duration is long-lived. Finally, when $0.5 < (d-b) < 1$, the spread follows a non-stationary but mean-reverting process, and the shock duration is long-lived. As we show in table 6.11, there are two sources of evidence of this difference. On one hand, the order of integration of the spread is noticeably above zero, reflecting a long-memory process. On the other hand, most maturities follow a stationary process; thus the duration of the shock is long-lived. Meanwhile, when the maturity of interest rates is 20 years, the process switches to a non-stationary but mean-reverting process, and the effect of the shock declines at a slower rate than the previous maturities. This result is in line with the results of Weber and Wolters (2012) and Holmes et al. (2015).

TABLE 6.11: H_1^{d-b}: Spread degree of persistence

	Maturities							
3-month vs.	6-month	1-year	2-year	3-year	5-year	7-year	10-year	20-year
$\hat{d}$					1.057	0.989	1.037	1.116
$\hat{b}$					0.808	0.837	0.821	0.795
H_1^{d-b}					*0.239*	*0.151*	*0.216*	*0.321*
6-month vs.		1-year	2-year	3-year	5-year	7-year	10-year	20-year
$\hat{d}$					1.058	1.132	1.022	1.231
$\hat{b}$					0.813	0.640	0.824	0.577
H_1^{d-b}					*0.250*	*0.492*	*0.198*	*0.654*
1-year vs.			2-year	3-year	5-year	7-year	10-year	20-year
$\hat{d}$					1.128	1.091	1.108	1.197
$\hat{b}$					0.646	0.655	0.633	0.584
H_1^{d-b}					*0.482*	*0.436*	*0.475*	*0.613*

Fractional order of integration of the explanatory variables and the errors cointegrating are denoted by d and b

According to section 6.2.1, we attempt to answer the following question: Did the QE affect the term structure of interest rates? For this reason, we check if the QE program had any impact on the long-run relationship between each pair of maturities by applying our methodology, i.e., the FCVAR model. We apply November 2008, the date marking as the start of the QE program (as Holmes et al. (2015) did), as the breakpoint of our sample. We then apply two regimes: the first regime covers October 1993 to November 2008 and the second covers December 2018 to December 2018. We next apply the FCVAR and obtain the following results[10] (see tables 6.12a and 6.12b). On one hand, for the first regime, the results show steady behavior, where most interest rate maturity pairs analyzed are cointegrated in a (1, -1) vector, and the spread follows a stationary process. On the other hand, according to the Regime II estimations, this regime covers the aftermath of the global financial crisis and government efforts to allay the impact of this quarrelsome period. Consequently, the results obtained are very similar to those of the original sample, in which the majority of interest rate pairs apply to another scenario, thus establishing places where the cointegrating vector is not (1, -1), although the spread follows a stationary process. We also find that two pairs of maturities follow a nonstationary but mean-reverting process.

TABLE 6.12A: Summary of results Regime I

Order of integration of the error correction term (ECT)	Value of β (assuming cointegration)	
	$\beta = 1$	$0 < \beta < 1$
$I(d-b) = I(0)$	[3M – 1Y] [3M – 2Y] [3M – 3Y] [3M – 5Y] [3M – 7Y] [3M – 10Y][3M – 20Y] [6M – 1Y] [6M – 3Y] [6M – 5Y] [6M – 7Y] [6M – 10Y] [6M – 20Y] [1Y – 2Y] [1Y – 3Y] [1Y – 20Y]	[3M – 6M] [6M – 2Y]
$I(0) < I(d-b) < I(0.5)$	[1Y – 5Y] [1Y – 7Y] [1Y – 10Y]	
$I(0.5) < I(d-b) < I(1)$		

The shaded area corresponds to the traditional EHTS. The sample covers October 1993 to November 2008.

[10]For reasons of space the results are available upon request.

TABLE 6.12B: Summary of results Regime I

Order of integration of the error correction term (ECT)	Value of β (assuming cointegration) $\beta = 1$	$0 < \beta < 1$
$I(d-b) = I(0)$	[1Y – 3Y]	[3M – 6M] [3M – 1Y] [3M – 2Y] [3M – 3Y] [6M – 1Y] [6M – 2Y] [6M – 3Y] [1Y – 2Y]
$I(0) < I(d-b) < I(0.5)$	[3M – 5Y] [3M – 7Y] [3M – 10Y] [3M – 20Y] [6M – 5Y] [6M – 7Y] [6M – 10Y] [1Y – 5Y] [1Y – 7Y] [1Y – 10Y]	
$I(0.5) < I(d-b) < I(1)$	[6M – 20Y] [1Y – 20Y]	

The shaded area corresponds to the traditional EHTS. The sample covers December 2008 to December 2018.

Tables 6.13 and 6.14 provide a summary of results, in which whether hypotheses are accepted or rejected is indicated with a tick or cross symbol, respectively, in the results column. In this table, we also present data that are valuable to highlight. From table 6.14, we observe that each pair of interest rates (short- vs. long-term interest rates) is disposed under a different scenario. As stated above, the results obtained for the second regime are similar to these. At first glance, the results reveal that, while none of the interest rate pairs checked occupy the theoretical EHTS zone, they spread to a weak EHTS or to situations in which the spread follows a long-memory process. If we were to execute this exercise with a traditional cointegration[11] approach, the results would be different because they would occupy the theoretical EHTS zone. Thus, the FCVAR model allows us to avoid rigidity in the stationarity of the spread, showing that we adhere to scenarios not appreciated in traditional cointegration.

TABLE 6.13: Summary of results I

	Hypotheses	Results	Observations
Step 1	H_1^d: Is the fractional cointegration more appropriate that traditional cointegration?	✓	
Step 2	H_1^β: Cointegrating vector is (1, -1)	✓	Only for 5-year or more interest rates
Step 3	$H_1^\beta \cap H_1^{\alpha_i}$: Variables are weakly exogenous under the restrictions of the cointegrating vector (1, 1)	✓	Only for short-term interest rates (α_r)
Step 4	H_1^{d-b}: Does the spread involve a long memory process?	✓	- Stationary (5, 7, 10-year) - Non-stationary but mean-reverting (20-year)

[11] A summary of estimations of the CVAR is included in the Appendix as Table A.4

TABLE 6.14: Summary of results II

Order of integration of the error correction term (ECT)	Value of β (assuming cointegration)	
	$\beta = 1$	$0 < \beta < 1$
$I(d-b) = I(0)$		[3M – 6M] [3M – 1Y] [3M – 2Y] [3M – 3Y] [6M – 1Y] [6M – 2Y] [6M – 3Y] [1Y – 2Y] [1Y – 3Y]
$I(0) < I(d-b) < I(0.5)$	[3M – 5Y] [3M – 7Y] [3M – 10Y] [3M – 20Y] [6M – 5Y] [6M – 7Y] [6M – 10Y] [1Y – 5Y] [1Y – 7Y]; [1Y – 10Y]	
$I(0.5) < I(d-b) < I(1)$	[6M – 20Y] [1Y – 20Y]	

The shaded area corresponds to the traditional EHTS. The sample size is 303 and the sample covers October 1993 to December 2018.

6.5 Conclusion

With the EHTS, long-term rates can explain changes in future short-term rates. Understanding the term structure of interest rates has always been viewed as crucial to determining the impact of monetary policies and their transmission mechanisms; it also plays an important role in macroeconomic predictions and in portfolio analysis (Li and Davis, 2017). Indeed, when monetary policy is effective, changes in short-term policy interest rates should impact long-term ones. Based on a fractionally cointegrated VAR model, our analysis considers both cointegration between short- and long-term interest rates and the long memory of their linear combination, i.e., the spread. We describe the spread as the difference between long- and short-term rates. The proposed methodology affords us the opportunity to reject the apriorism of incompatibility, whereby interest rates are cointegrated and the term spread is rendered non-stationary. We also manage to alter the EHTS by extending opportunities raised in the literature.

We use US monthly interest rates for nine different maturities running from October 1993 to December 2018 Our results provide evidence in accordance with the EHTS for 5- to 20-year interest rates; the spread between the short rate and long end of the term structure was found to be an optimal predictor of future short rates from a maturity of a 5-year horizon. Importantly, we find evidence that the spread presents a long memory process, in contrast to the usual assumption of $I(0)$.

Due to the global financial crisis of 2008, the Fed initiated a series of incentives to face such a crisis, i.e., the QE program, establishing this point as a breakpoint and analyzing the resulting regimes. In this regard, we find that the behaviors of short- and long-term rates differ depending on the observed regime, and we highlight the change in pattern occurring from the first regime, where relations appear stable; from the crisis, we observe how this stability disappears a priori, giving rise to a situation very similar to that observed for the entire sample. In sum, these results show that our results may be motivated by the application of the QE program.

According to these empirical results, we reveal persistence in the spread, and we consequently outline some important implications for monetary policy. On one hand, Baillie and Bollerslev (1994a) noted that a long memory spread offers adequate forecasting power at longer horizons. On the other hand, and even more importantly, in line with Cassola and Morana (2008), Hassler and Nautz (2008) and Cömert (2012), we show that the persistence of the spread implies a gradual loss of control power over interest rates of the Fed, particularly when the maturity is 20 years. As persistence in term spread increases, the gap between maturities also increases. This persistence might limit the amount of information contained in short-term rates for future monetary policy, which may affect the Fed's control of long-term interest rates and of the yield curve. To address this issue, the Fed should increase the frequency of money market interventions.

We suggest that the growing difference between short- and long-term interest rates creates a vulnerable link as the term spread increases. In line with Blinder, Ehrmann, Fratzscher, De Haan, and Jansen (2008), Ben Bernanke's 2012 Jackson Hole speech and Li and Davis (2017), this connection is essential for further guidance. Furthermore, since the Fed only has power over shorter-end interest rates, its manipulation may influence other short-term interest rates and thus may be necessary for the application of measures affecting longer-term rates when the monetary policy transmission mechanism predicted by the EHTS is not met. Policies oriented over time, such as the QE program, would thus be necessary to maintain this transmission mechanism or the substitutability of interest rates.

6.6 Appendix

Table A.1a: Lag length selection when 3-month is reference of short-term

Lags	6-month	1-year	2-year	3-year	5-year	7-year	10-year	20-year
1	-1063.84	-799.23	-575.68	-481.68	-395.14	-372.49	-373.13	-403.58
2	-1068.39	-812.54	-585.67	-489.85	-396.41	-372.20	-371.58	-401.48
3	-1077.42	-823.24	-588.59	-490.21	-397.68	-375.51	-379.19	-409.37
4	-1070.59	-818.39	-585.15	-488.32	-396.18	-373.35	-378.05	-409.51
5	-1070.95	-820.87	-587.62	-491.14	-401.54	-378.50	-382.91	-412.81
6	**-1090.25**	**-830.45**	**-596.63**	**-500.08**	**-410.40**	**-386.98**	**-393.02**	**-420.13**

Bold indicates lag order selected

Table A.1b: Lag length selection when 6-month is reference of short-term

Lags	1-year	2-year	3-year	5-year	7-year	10-year	20-year
1	-1139.94	-743.54	-606.96	-487.99	-450.04	-438.71	-455.01
2	-1156.86	-754.32	**-615.36**	-488.79	-447.51	-435.74	-452.09
3	-1158.26	-751.87	-611.76	-481.50	-445.15	-436.79	-452.14
4	-1158.30	**-756.91**	-608.64	-480.72	-444.23	-438.85	-454.56
5	**-1167.29**	-748.48	-611.52	-492.80	**-455.94**	-447.08	**-462.09**
6	-1165.55	-749.57	-612.59	**-493.00**	-455.95	**-447.33**	-460.16

Bold indicates lag order selected

Table A.1c: Lag length selection when 1-year is reference of short-term

Lags	2-year	3-year	5-year	7-year	10-year	20-year
1	-948.32	-734.64	-549.23	-487.93	-460.95	-459.51
2	-951.53	-738.20	-548.39	-486.72	-459.99	-457.47
3	-949.54	-734.80	-547.46	-489.07	-467.25	-465.37
4	-953.47	-739.57	-551.39	-492.83	-470.00	-467.57
5	**-953.82**	**-742.17**	**-561.41**	**-503.20**	**-482.01**	**-479.59**
6	-952.56	-740.18	-559.52	-500.47	-480.72	-477.21

Bold indicates lag order selected

Table A.2: Rank test

Rank test when 3-month is reference of short-term								
rank	6-month	1-year	2-year	3-year	5-year	7-year	10-year	20-year
0	27.562	10.987	17.532	24.088	19.407	17.384	17.551	16.119
	(0.003)	(0.027)	(0.011)	(0.010)	(0.001)	(0.013)	(0.089)	(0.140)
1	4.544	1.981	1.052	0.632	2.017	1.493	2.665	2.348
	(0.288)	(0.691)	(0.860)	(0.940)	(0.627)	(0.747)	(0.515)	(0.566)
Rank test when 6-month is reference of short-term								
rank		1-year	2-year	3-year	5-year	7-year	10-year	20-year
0		34.212	16.412	17.255	9.882	19.056	8.663	16.484
		(0.000)	(0.003)	(0.002)	(0.042)	(0.046)	(0.267)	(0.105)
1		2.658	1.784	0.372	1.575	1.107	2.169	1.502
		(0.479)	(0.182)	(0.542)	(0.723)	(0.652)	(0.613)	(0.522)
Rank test when 1-year is reference of short-term								
rank			2-year	3-year	5-year	7-year	10-year	20-year
0			19.975	15.612	15.615	15.609	14.825	13.930
			(0.054)	(0.124)	(0.122)	(0.126)	(0.154)	(0.205)
0			0.627	0.488	0.929	0.812	1.131	0.891
			(0.775)	(0.842)	(0.703)	(0.742)	(0.644)	(0.684)

The top of the table shows the LR statistics and P values are in parenthesis. We follow the rank test procedure for small samples (the sample size is 303) suggested by Jones, Nielsen, and Popiel (2014)

Table A.3: Ljung-Box Q-test

Maturities								
3-month vs.	6-month	1-year	2-year	3-year	5-year	7-year	10-year	20-year
$Q_{\hat{\varepsilon}}$	59.483	43.407	26.770	22.143	19.726	20.548	22.471	20.366
	(0.124)	(0.661)	(0.994)	(0.999)	(1.000)	(1.000)	(0.999)	(1.000)
6-month vs.		1-year	2-year	3-year	5-year	7-year	10-year	20-year
$Q_{\hat{\varepsilon}}$		60.490	73.098	67.618	30.501	38.576	33.013	40.404
		(0.107)	(0.011)	(0.032)	(0.977)	(0.832)	(0.951)	(0.774)
1-year vs.			2-year	3-year	5-year	7-year	10-year	20-year
$Q_{\hat{\varepsilon}}$			49.023	44.599	42.825	42.219	41.013	39.010
			(0.432)	(0.613)	(0.684)	(0.708)	(0.752)	(0.819)

Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. The sample size is 303. LR statistics and P values are in parenthesis below LR test values.

Table A.4: CVAR results

Order of integration of the error correction term (ECT)	Value of β (assuming cointegration)	
	$\beta = 1$	$0 < \beta < 1$
$I(d-b) = I(0)$	[3M – 2Y] [3M – 3Y] [3M – 5Y] [3M – 7Y] [3M – 10Y] [6M – 1Y] [6M – 2Y] [6M – 3Y] [6M – 5Y] [6M – 7Y] [6M – 10Y] [1Y – 2Y] [1Y – 3Y] [1Y – 5Y] [1Y – 7Y] [1Y – 10Y]	[3M – 6M] [3M – 1Y]
$I(0) < I(d-b) < I(0.5)$		
$I(0.5) < I(d-b) < I(1)$		

The shaded area corresponds to the traditional EHTS.

Chapter 7

The impact of the term spread in US monetary policy from 1870 – 2013.

7.1 Introduction

Over the years, central banks have sought to maintain an efficient monetary policy in order to keep price stability and economic development under control despite the different cycles. This power to control is due to the spread between long- and short-term rates and offers significant information regarding the effects of shocks in the long run (Estrella and Hardouvelis, 1991). For this reason and based on the Expectations Hypothesis of Term Structure (EHTS here after), the term structure should respond to international market forces due to the fact that interest rates have been viewed as crucial in monetary policy and its transmission mechanism; thus, changes in short-term interest rates usually impact on long-term interest rates, and if this occurs, we can conclude that the monetary policy is effective (Holmes et al., 2015). According to this idea, we revise recent US economic history (since 1870), aiming to explain how different impacts in the economy have affected both interest rates and monetary policy and the measures that have been implemented by the authorities.

The body of selected literature for this topic has focused on the EHTS directed in the United States of America (the USA hereafter), and results differ. On the one side, the EHTS is accepted as a predictive instrument (Mankiw and Miron (1986); Poole et al. (2002) or Adrian and Estrella (2008)) or because it has deep implications for monetary policy (e.g., Weber and Wolters (2012, 2013). On the other side, we find evidence against the EHTS that fails when short-term maturities are used (see Sarno et al. (2007) or Bulkley et al. (2011, 2015) or when its ability to forecast short-term rates has been reduced (Guidolin and Thornton, 2010). Finally, attending to the difference between long- and short-term interest rates, i.e., the spread, some authors support that the degree of integration would be different from $I(0)$ (Strohsal and Weber (2014) or Holmes et al. (2015), for instance), and this has implications for the control of interest rates (Cömert, 2012).

In this sense, this paper analyses the possible relationship between short- and long-term interest rates based on a new approach in the literature. The novelty of our approach consists in the Fractionally Cointegrated Vector Autoregressive (FCVAR) model (Johansen and Nielsen (2012) and Nielsen and Popiel (2016)) because it is considered that the axioms of traditional cointegration may be too restrictive; i.e., with respect to the dichotomy ($I(0)/I(1)$), the series would follow an $I(d)$ process. So, this new methodology allows us to break with the rigidity of traditional approaches in favour of letting the series be cointegrated, and the spread does not necessarily need to be $I(0)$; and the assumption that interest rates could follow the dichotomy $I(0)/I(1)$ is rejected. Using two types of database, i.e., the *Jordà-Schularick-Taylor Macrohistory Database* and *Shiller's database*, we find similar results. In both estimations, we cannot reject the EHTS in this time period, and more importantly, according to the FVECM, the coefficients associated with short-term rates are significant, which implies that the spread has prediction power in the bearing of futures short-term rates. We also find that the long-term rate drives the long-run relationship, contributing to the total proportion to the common trend, and the persistence of the spread shows control power over interest rates by Fed. In sum, we have applied a novel econometric approach joint with an exhaustive revision of the main events in the history of US monetary policy focused on interest rates as a tool; two databases are used that allow us to detect interest rate behaviour throughout the recent history of the USA and the implications for monetary policy.

The rest of the chapter is as follows. The next section 7.2 presents a tour of recent US economic history and focuses briefly on the empirical literature; section 7.3 shows the methodology used in the chapter. Then, section 7.4 discusses the empirical results and conclusions, leading to some of the economic policy implications shown in section 7.5.

7.2 A brief review of monetary policy history

7.2.1 Monetary policy through the last century and a half

The effects of monetary policy over the last century and a half have been different due to the efforts of the Fed to maintain equilibrium between economic growth and market forces based on interest rates. For this reason, the distinction of diverse eras is necessary to support and explain the behaviour of the term structure and the Fed's actions regarding monetary policy using interest rates as a mechanism of control. In this sense, several authors have tried to show historical evidence of changes in monetary policy. Nonetheless, we focus on two relevant papers: Taylor (1999), which divides recent history into three main periods, and Darné and Charles (2011), which explains events. We follow an economic classification organization such as the National Bureau of Economic Research as well. Therefore, based on an in-depth depth review of the literature, we build a table that summarizes US monetary policy from recent history.

Aiming to provide a deep overview, Taylor (1999), in his study about the history of monetary policy, suggests different periods, which span from the end of the nineteenth century to the end of the twentieth century. The earlier period covers from 1879 to 1914 and follows the classical international gold standard era; the latter period extends from 1955 to 1997 and covers the era of Bretton Woods, when the exchange rate was fixed, and the modern flexible exchange era. In this paper, Taylor (1999) also argues about the type of Fed actions in the last years of that period because this policy rule is different from that applied by the gold standard, Bretton Woods or the early part of the flexible exchange rate era. But, as is well known, various events in recent history played a significant role in monetary policy and the treatment of interest rates. Thus, Darné and Charles (2011) identified several episodes that help to understand and explain changes in US monetary policy; we use these to refine the different periods proposed by Taylor (1999).

The early period – The gold standard

The backgrounds of these episodes are defined by the economic cycles that have marked the measures in different ways. Following the National Bureau of Economic Research (NBER hereafter) dating cycles, we start with the end of the Civil War that devastated the USA, which links with the beginning of the dataset used in this paper, i.e., 1870. Following Kindleberger (2000), at this time, there were financial difficulties due to the fact that debts were very high, and the objective of the administration was to sell Treasury gold to pay off the national debt, stabilize the dollar, and improve the economy. In 1873, the 'Panic of 1873' and Long Depression occurred, prompted by a drop in silver demand and subsequent downward pressure on the value of silver. For this reason, the US government moved to the gold standard; silver prices fell, and the domestic money supply was also reduced. The perception of instability in US monetary policy caused investors to withdraw from long-term obligations, particularly long-term bonds (Bordo and Kydland, 1995). After a period of economic expansion came the 'Panic of 1893', where silver was undervalued due to overproduction and the U.S. Treasury was forced to borrow $65 million in gold to support the gold standard. In response, foreign investors sold American stocks to obtain American funds supported by gold. As a result of the retraction of market liquidity in the 'Panic of 1907', a commission was established to investigate the crisis and propose future solutions, leading to the creation of the Federal Reserve System. Due to the entrance of the USA into World War I, financial inflation was high due to huge gold imports from the European belligerents who bought war material (Bordo and Haubrich, 2004).

Pre-World War II and the road to the end of XX century

In 1928, the stability and progress of the economy were threatened, so the Fed initiated a tight monetary policy in order to avoid a stock market bubble. This tight policy led to the stock market crash of October 1929 and was the beginning of the Great Depression (Orphanides, 2003). This period is characterized by the repeated failures of the Federal Reserve System to balance the monetary collapse (Friedman and Schwartz, 1963). The contraction came to an end because linkage with the gold standard was broken and there was a program of reflation for Treasury gold and silver (Bordo and Haubrich, 2004). Before World War II, in the New Deal period, there was another depression due to the application of different economic measures, but the tight monetary policy carried out by the Federal Reserve was a crucial measure.

By the end of World War II, the USA and other countries joined a new international monetary system, the Bretton Woods system. It involved a much less direct link to gold as a nominal anchor than had existed during either the inter-war or pre-war periods of the gold standard. As is well known, after the war, investment in the arms industry brought subsequent periods of recession, highlighted by the post-Korean War period, in which the Federal Reserve changed to a more restrictive monetary policy because of fears of inflation and the formation of an economic bubble. Between 1968 and 1970, the Bretton Woods system began collapsing due to the inappropriate policies of its members, evasion of capital controls, and the abandonment of the responsibility to maintain price stability (Bordo, 1993). The US government tried to finance social programs and the Vietnam War, using an expansionary monetary policy. This led to the recession of 1970, where the Federal Reserve raised interest rates, i.e., monetary tightening. After the 1973 oil crisis, the Fed increased interest rates to solve the problem of stagflation (Bernanke, Gertler, Watson, Sims, and Friedman, 1997); the early 1980s were characterized by of the raising of interest rates to fight inflation by the Federal Reserve under the direction of Paul Volcker. Campbell and Clarida (1987) explain the shocks that occurred in the 1980s as a federal budget deficit.

The early 2000s

In the early 2000s, the longest growth period in the history of the US ended owing to a fall in investment caused by the collapse of the dot-com bubble and the September 11^{th} attacks. This situation was reversed by the implementation of painful fiscal adjustment and also through the cost of the Afghanistan and Iraq wars (Kraay and Ventura, 2005). Finally, by the end of 2007, the subprime mortgage market collapsed and quickly spread to the rest of the world. The US government responded with an unprecedented bank bailout and fiscal stimulus package. The NBER declared the recession over more than a year after the end date (June 2009).

Figure 7.1 shows the *Jordà-Schularick-Taylor Macrohistory Database* as a synthesis of all previously explained cycles, with the grey bands determining the cycle moments provided by NBER and the short- and long-term interest rates. Figure 7.2 corresponds to *Shiller's database* and shows another similar synthesis for the events explained before. Also, the grey bands explain the cycle moments and are provided by NBER. Both databases will be explained in a later section.

FIGURE 7.1: Time series plot for *Jordà-Schularick-Taylor Macrohistory Database*

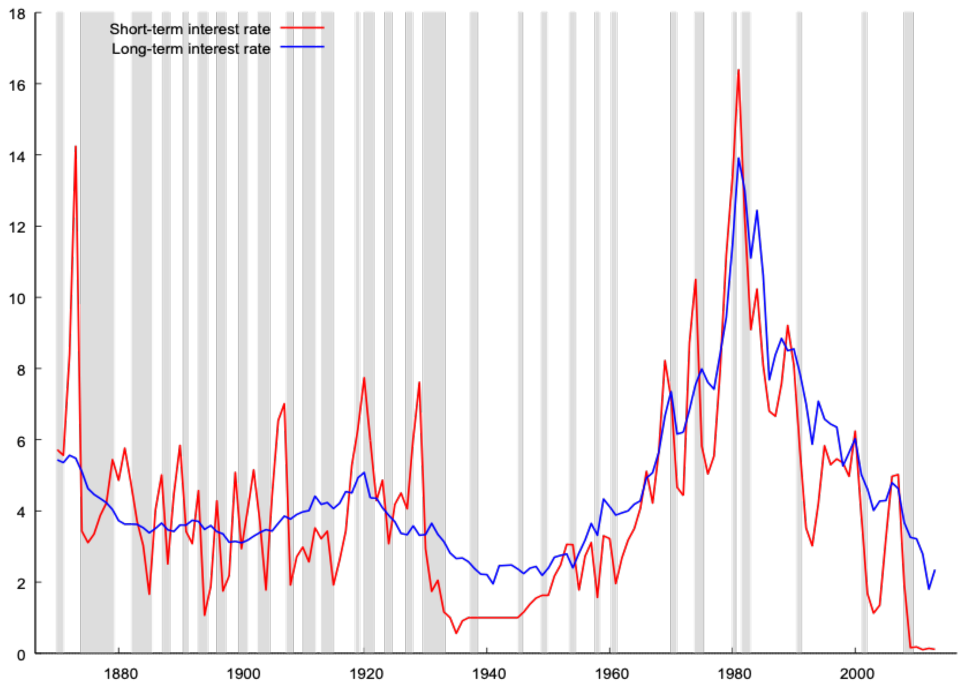
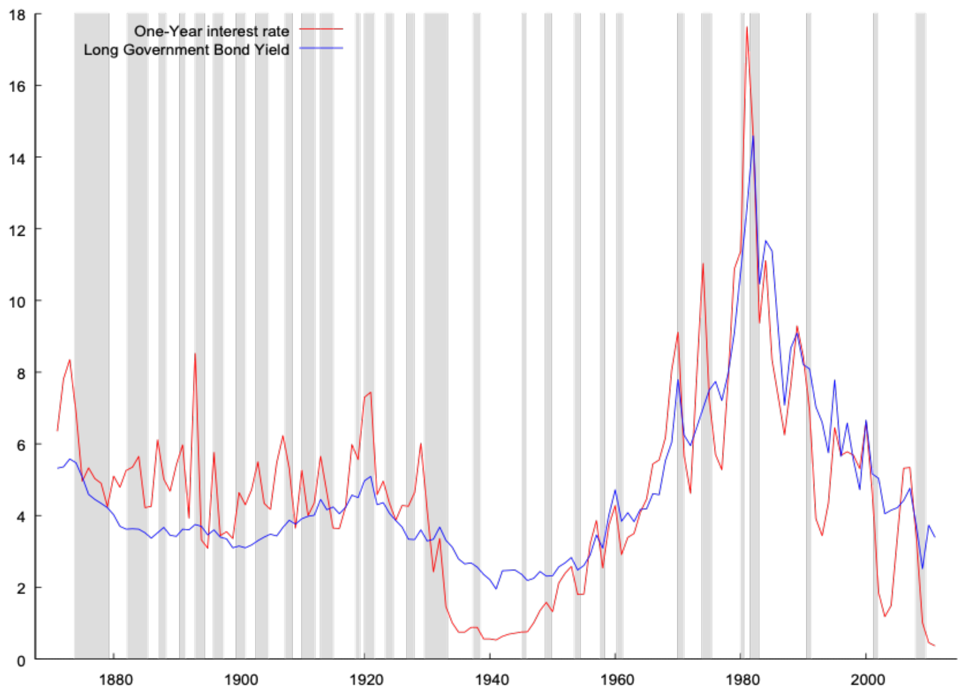


FIGURE 7.2: Time series plot for *Shiller's database*



Interest rates allow economists and policy makers to predict cycles, which is crucial in the study of yield curves to forecast the behaviour of the term structure of interest rates. For this reason, the EHTS plays an important role in the linkage among short- and long-term rates. Adrian and Estrella (2008) reveal the predictive power of the yield curve so that an inverted yield curve signals a recession; meanwhile, a positive sloped yield curve is often a harbinger of inflationary growth.

Aiming to clarify the proposed events by several authors and sources, in table 7.1, we show a summary of all events cited previously in a novel contribution in the literature review. For the elaboration of this table, we have used the information provided in Darné and Charles (2011) and different studies, and the dates are checked with the dates provided by the NBER.

TABLE 7.1: Summary of events

Reference	Date	Event	Reason	Economic Measure
NBER; Kindleberger (2000)	1870	Post-USA Civil War	Very high debts	Treasury gold was sold to pay and stabilize the dollar
Bordo and Kydland (1995)	1873	'Panic of 1873'	Downward pressure on the silver value	The USA moved to gold standard
Friedman and Schwartz (1963); Bordo and Haubrich (2004)	1893	'Panic of 1893'	Overproduction of silver	The government borrow \$65M in gold to support gold standard
Darné and Charles (2011)	1907	'Panic of 1907'	Retraction of market liquidity	Creation of the Federal Reserve system
Bordo and Haubrich (2004); Darné and Charles (2011)	1914-1918	World War I	High financial inflation	Tight central bank policies around the world
Darné and Charles (2011)	1928	Developing market bubble		Tight monetary policy
Friedman and Schwartz (1963); Orphanides (2003); Bordo and Haubrich (2004)	1929-1933	Great Depression	Failures of Federal Reserve system to balance the monetary collapse	The linkage with the gold standard was broken and program to reflation of Treasury gold and silver
Bordo and Haubrich (2004)	Pre-World War II	New Deal	Depression of the economy	Tight monetary policy
Bordo (1993)	Post-World War II		A much less direct link to gold as a nominal anchor than in previous periods	Joining to Bretton Woods system
Darné and Charles (2011)	1950-1957	Post-Korean War	Inflation scare and bubble forming	Fed changed to a restrictive monetary policy
Bordo (1993)	1968	Bretton Woods system collapse	Inappropriate policies	Abandonment of this monetary system
Bordo and Haubrich (2004); Darrat and Zhong (2005)	1968 – 1971	Vietnam War	Financing social programs and Vietnam War	Expansionary monetary policy
Bernanke, Gertler, Watson, Sims, and Friedman (1997)	1973	Oil crisis	Stagflation caused by shocks in oil prices	Rising of interest rates
Campbell and Clarida (1987); Bordo and Haubrich (2004); Darné and Charles (2011)	End of 1970's and beginning 1980's	Volcker's direction	Fear to inflation; Federal budget deficit	Rising of interest rates by Fed
Kraay and Ventura (2005)	Beginning of 2000's	Dot-com bubble and September 11^{th} attacks		A painful fiscal adjustment due to the cost of Afghanistan and Iraq wars.
NBER	2007	Subprime mortgage crisis		Fiscal stimulus and bank bailout

Own elaboration.

7.2.2 The Expectations Hypothesis of the Term Structure. How does EHTS affect monetary policy?

Studies concerning the term structure of interest rates have tried to evaluate the impact and how they are affected by the monetary policy of Central Banks. However, changes in the economy could affect the EHTS, so if a variation in short-term policy impacts on the long term, monetary policy is effective (Holmes et al., 2015).

The EHTS is the most influential theory of term structure relations, and it has been studied as a way to show the relationship between short- and long-term rates. This hypothesis establishes that an average of the current and expected short-term rates determines long-term rates and the spread; differences in long- and short-term rates implies crucial information on future changes in short-term rates. So, the potential effectiveness of monetary policy is revealed by this relationship, which consists of the control of short-term policy rates by central banks (Bernanke and Blinder, 1992). Two papers in the literature, Fama (1984) and Fama and Bliss (1987), explain that the long-term rate is an unbiased predictor of future short-term rates[1]. Another implication of the EHTS is that the spread between the long-term rate and the short-term rate is an unbiased predictor of future short-run changes in the long-term rates, as Mankiw (1986), Campbell and Clarida (1987), Campbell and Shiller (1991) and Campbell (1995) evidenced in their works.

From the contrast of this hypothesis shown by the studies above, studies have further focused their efforts on testing the EHTS and its economic implications in the USA or providing evidence of the forecasting power of the term structure. In this literature, mixed empirical evidence has been shown. This controversy is motivated by several reasons, such as the data selected, the techniques applied and the time period studied. Regarding the technique for testing the EHTS, we find, as an initial example, the study proposed by Mankiw and Miron (1986), which related 3-month and 6-month rates and longer-term rates in order to find evidence of predictive ability of interest rates before the establishment of the Federal Reserve in the marketplace. For this reason, we show below different studies that treat the EHTS in different contexts.

On one hand, from a technical perspective, we find in support of EHTS several studies that used the cointegration concept of Engle and Granger (1987), such as Campbell and Clarida (1987), which examined the statistical significance of the EHTS in the January 1959 to October 1983 period. Hamilton (1988) used a Markov Switching model in order to explain the changes in regime for short- and long-term rates of US Treasury Bills for the January 1961 − March 1987 period. Hall et al. (1992) used the same variables with a different time series (they selected the period from January 1970 to December 1988) to support the EHTS using a cointegration. For its part, Shea (1992) selected data for the period December 1951 to February 1987 in order to determine that the interest maintains a long-term relationship; therefore, the spread would not have a tendency to increase or decrease over time. Engsted and Tanggaard (1994) tested the implications of the EHTS on US term structure, finding that for the period January 1952 to February 1987, the zero-sum restriction on the cointegration vectors implied that the EHTS cannot be rejected following a variation of the Vector Error Correction Model (VECM hereafter), i.e., the threshold VECM. Moreover, Longstaff (2000) also used cointegration techniques, concretely the Vector Autoregressive (VAR) model, in a daily sample of extreme short-term rate and longer-term rates supporting the EHTS. In addition, Poole and Rasche (2000) and Poole et al. (2002) demonstrated that the market could anticipate changes in the FED's target during the period October 1988 through February 2000 using the federal funds rate. Hansen and Seo (2002) and Seo (2003) found roughly consistent results with the term structure prediction in the period that they chose (January 1952 to December 1992 and January 1960 to December 1999, respectively). Diebold et al. (2006) verified the EHTS in certain periods but not in the entire sample. Using another methodology such as the Autoregressive Conditional Heteroskedasticity (GARCH), Mili et al. (2012) explain specific interest rates dynamics in order to show the nonlinearities in the relationship between interest rates in the daily period from July 2001 to April 2011. For its part, Weber

[1] These works use the terms forward rate for the long-term rate and spot rate for the short-term rate.

and Wolters (2012, 2013) applied the VECM in the US term structure in order to contribute an economic explanation of the deviations from EHTS in a similar data sample. Also, Kishor and Marfatia (2013) showed that the FF futures rate is cointegrated with the 3-month T-bill rate, and they move together in the long run in a daily sample that spans from May 1989 to June 2008. More recently, Holmes et al. (2015) supports that the EHTS holds in the long run, i.e., short-run policy changes effects on long-term rates for the period October 1993 until April 2015.

On the other hand, there is evidence against the EHTS, such as Sarno et al. (2007), which rejects the EHTS when short maturities from 1952 to 2003 are used. Thornton (2005) and Guidolin and Thornton (2010) tested the EHTS, determining less ability to forecast short-term rates, which has deep implications for policy makers; thus, conventional theory of the term structure of interest rates is threatened, and Bulkley et al. (2011) and Bulkley et al. (2015) evidenced the failure of EHTS using Treasury securities in a similar monthly sample that starts in 1952.

More importantly, several studies have treated the spread between long- and short-term interest rates in the sense that this spread contains information about the term structure. These studies have focused on different regions in the world, but in the case of the USA, the work of Strohsal and Weber (2014) and Holmes et al. (2015) supports the EHTS; however, the degree of integration of the spread would be different from $I(0)$. Previously, Cömert (2012) related overnight interest rates and long-term rates in the US from 1983 to 2007 and presented evidence that the Fed has been gradually losing its control over long-term interest rates.

In sum, despite the initial controversy, this literature has shown that it is possible to establish a relationship between short- and long-term rates across the last century and a half in The USA. It has also addressed the importance of the spread between interest rates as a tool to determine the efficiency of monetary policy and forecasting power. Finally, unlike the studies previously considered, which have analyzed the cointegration and spread, to the best of our knowledge, our approach is the only one that makes it possible to analyze both jointly. Hence, our econometric method permits us to test the existence of a long-run relationship between the interest rates selected and the persistence of the spread based on a new approach in the existing literature, a Fractionally Cointegrated Vector Autoregressive (FCVAR) model provided by Johansen and Nielsen (2012); this model makes it possible to avoid the problems with the axioms of traditional cointegration associated with rigidity.

7.3 Methodology

Our econometric strategy is based on obtaining and analyzing at a yearly frequency the model estimation; we then perform statistical tests of fractional cointegration and weak exogeneity based on the fundamental equation of the EHTS in an econometric context in order to explain possible monetary policy efficiency.

7.3.1 The EHTS model

The fundamental equation of the EHTS of an $n > 1$ period bond R_t (i.e., long-term interest rate) is equal to an average of the current and expected r_t (i.e., short-term interest rate) set of $n \leq 1$ period plus a constant term. The relationship can be expressed in the following form:

$$R_t = \frac{1}{k}\sum_{i=0}^{k-1} E_t[r_{t+n}] + \pi_t^*, \tag{7.1}$$

where π_t^* is a possible stationary term and E_t is the expectation operator at time t for the evolution of short-term interest rates driving the long-term interest rates. In order to test the EHTS in the context of cointegration theory, the linear mode used is:

$$R_t = c + \beta r_t + \varepsilon_t \tag{7.2}$$

Agreeing with Campbell and Shiller (1987), R_t and r_t should be non-stationary and related through a cointegration relationship with parameters (1,-1)), i.e., a one-to-one relation, $\beta = 1$. These implies that β_R and β_r are cointegrated constants, and its combination is a stationary process, and the spread of the interest rate could revert to the mean. If the spread is stationary, the long- and short-term rates are driven by a common stochastic trend and do not allow arbitrage opportunities because market forces adjust to correct any temporary disequilibrium. As the EHTS suggests, the interest rate spread is an optimal forecast[2] of future changes in long-term interest rates.

In this paper, the fractionally cointegrated vector autoregressive (FCVAR) model allows us to study the long-run equilibrium relationship between long- and short-term interest rates. This model allows fractional processes of order d that cointegrate to order $d-b$; conducting our analysis using a bivariate fractional cointegration approach, we consider that the standard unit root and cointegration test might be too restrictive ($I(1)/I(0)$ dichotomy). In the FCVAR model, the error correction term (the spread when EHTS is supported) is different from $I(0)$; i.e., this assessment is not restricted (the integration order could be different from zero and thus show a long-memory process), rejecting the general assumption that the spread is $I(0)$ and could be an $I(d)$ process. More general $I(d)$-type specification has been adopted, considering the possibility of fractional orders of integration – cointegration without these values is unrestricted.

7.3.2 Fractional cointegration model – FCVAR methodology

This model is provided by Johansen (2008a, 2008b) and developed in Johansen and Nielsen (2012) and Nielsen and Popiel (2016); it has the advantage of being used for stationary and non-stationary time series. Our objective is to study the EHTS under fractional cointegration conditions.

To introduce the FCVAR model, we introduce the fractional difference operator to the CVAR model, which Δ and the fractional lag operator is $\Delta = (1-L)$. Replacing lags operators with fractional counterparts Δ^b and $\Delta^b = (1-L^b)$, and applying $Y_t = \Delta^{d-b}X_t$, such that:

$$\Delta^d X_t = \alpha\beta' L_b \Delta^{d-b} X_t + \sum_{i=1}^{k} \Gamma_i \Delta^d L_b^i X_t + \varepsilon_t, \tag{7.3}$$

As always, ε_t is p-dimensional independent and identically distributed with mean of zero and covariance matrix Ω. The parameters α and β are $p \times r$ matrices, where $0 \leq r \leq p$. In matrix β, the columns are the cointegrating relationships and $\beta' X_t$ assumes the existence of a common stochastic trend integrated at order d and b, indicating the strength of the cointegrating relationships, and the short-term parts from the long-run equilibrium being integrated of order $d-b$. The coefficients of α correspond the rate of adjustment unto equilibrium. Hence, $\alpha\beta'$ is the long-run adjustment, ρ' is the restricted constant term, and Γ_i represents the short-run behaviour of the variables. We reach the final model:

$$\Delta^d X_t = L_d \alpha(\beta' X_t + \rho') + \sum_{i=1}^{k} \Gamma_i \Delta^d L_d^i X_t + \varepsilon_t. \tag{7.4}$$

The error correction term is integrated from order $(d\check{\ }b)$, which is $I(0)$ in this case. However, in fractional cointegration, these axioms are relaxed because $(d-b) = 0$; i.e., the error correction term shows a short-run stationary behavior; or $(d-b) > 0$; i.e., there is a long memory process, and the error correction term will revert to its mean in the long run.

In order to determine the number of stationary cointegrating relations, the hypotheses in the rank test are followed based on a series of LR tests. In the FCVAR model, we test the

[2] Baillie and Bollerslev (1994a) discovered that a cointegrating relationships may not be precisely $I(0)$, implying that a fractionally cointegrated relationship may generate noticeable gains in forecast accuracy only within the context of a longer-term forecast.

hypothesis $H_0 : rank(\Pi) = r$, against the alternative: $H_1 : rank(\Pi) = p$, $L(d, b, r)$ being the profile likelihood function of rank r, where (α, β, Γ) have been reduced by rank regression (see Johansen and Nielsen, 2012).

Maximizing the profile likelihood distribution under both hypothesis, the LR test statistics are now $LR_t(q)$. The asymptotic distribution of $LR_t(q)$ depends on parameter b and on $q = n - r$. MacKinnon and Nielsen (2014) based on their numerical distribution functions on the asymptotic critical values of an LR rank test. In cases of "weak cointegration", i.e., $0 < b < 1/2$, $LR_t(q)$ has a standard asymptotic distribution,$LR_t(q)$ $LR_t(q) \xrightarrow{D} \chi^2(q^2)$.

According to the latter study, fractional cointegration implies a FVECM such as:

$$\begin{pmatrix} \Delta R_t \\ \Delta r_t \end{pmatrix} = \begin{pmatrix} \alpha_R \\ \alpha_r \end{pmatrix} (R_{t-1} - \beta r_{t-1} - c) + \sum_{i=1}^{n} \Gamma_i \begin{pmatrix} \Delta R_{t-i} \\ \Delta r_{t-i} \end{pmatrix} + \begin{pmatrix} v_{1t} \\ v_{2t} \end{pmatrix} \tag{7.5}$$

with adjustment parameters α, cointegration coefficient β, restricted constant (c), lag length (n) and errors v. Γ_i are 2×2 parameter matrices in the short-run dynamics. The adjustment coefficients α_R and α_r capture the speed of adjustment of R_t and r_t towards equilibrium. Additionally, according to EHTS, the absolute values of the estimates of α_R are much smaller than α_r; we suggest that the correction in the equation for the short term of interest rates exceeds the long-run equilibrium, i.e., the spread defined by the difference between long-term and short-term interest rates.

7.3.3 Permanent-Transitory (P-T) decomposition in the FCVAR model

According to Gonzalo and Granger (1995)'s Permanent-Transitory decomposition, we let $X_t = (R_t, r_t)'$, where R_t and r_t represent the long-term rate and short-term rate, respectively. In Permanent-Transitory decomposition, X_t can be decomposed into a transitory (stationary) part, βX_t, and a permanent part, $W_t = \alpha'_\perp X_t$, where $\alpha'_\perp \alpha = \alpha' \alpha_\perp = 0$. W_t is the common permanent component of X_t, and it is interpreted as the dominant rate, where the information that does not affect W_t will not have a permanent effect on X_t. We focus on the key parameter $\alpha_\perp$ can also be tested directly on $\alpha_\perp$ or alternatively on α itself using the values of LR tests in each hypothesis, and critical values can be taken from the Ξ^2 distribution for testing. For example, to test the hypothesis that the dominant rate is the long-term rate, i.e., $\alpha_\perp = (0, a)'$, we can equivalently test the mirror hypothesis, $H_0 : \alpha = (\gamma, 0)'$. Similarly, to test the hypothesis that the dominant rate is the short-term rate, i.e., $\alpha_\perp = (a, 0)'$, we test the mirror hypothesis, $H_1 : \alpha = (0, \gamma)'$ (see Dolatabadi et al. (2018), which first combined the FCVAR model with Permanent-Transitory decomposition).

An interpretation of coefficient α is that an adjustment coefficient measures how disequilibrium errors could be affected by current changes in X_t. Under this interpretation, we wonder whether any coefficients in α are zeros, i.e., the variable in question is weakly exogenous. For example, under hypothesis H_1, parameter $\alpha = 0$, such that the short-term rate does not react to the disequilibrium error, i.e., the transitory component, implying that the short-term rate is the main contributor to the common trend.

In order to determinate the magnitude of each variable in the long-run, we use the Component Share (CS), as Baillie et al. (2002) note that since $\alpha' \alpha_\perp = 0$, it may also be expressed in terms of the elements of the error correction vector α. To interpret this, we let $\alpha = (\alpha_1, \alpha_2)'$ and $\alpha_\perp \alpha = (\alpha_{\perp,1}, \alpha_{\perp,2})'$. Afterwards, $\alpha'_\perp \alpha = \alpha_{\perp,1}\alpha_1 + \alpha_{\perp,2}\alpha_2 = 0$ implies that $\alpha_{\perp,1} = -\alpha_{\perp,2}\alpha_2/\alpha_1$, and thus, component share (CS hereafter) may be expressed as

$$CS_1 = \frac{\alpha_2}{\alpha_2 - \alpha_1}, CS_2 = \frac{-\alpha_1}{\alpha_2 - \alpha_1} \tag{7.6}$$

7.4 Data and results

7.4.1 Data description

To study the long-run dynamics of term structure, we need a dataset as long as it allows us to check if the Fed has kept interest rates under control in recent history. Furthermore, we choose two different databases, and we use short- and long-term interest rates. One of them is provided by the *Jordà-Schularick-Taylor Macrohistory Database*, which covers 17 advanced economies from 1870 to 2013 on an annual basis. It includes 25 real nominal variables, but among these, we have selected financial variables such as short- and long-term interest rates for the USA[3]. The other data base selected is available in Chapter 26 from *Shiller's book Market Volatility* (1989) and is revised and updated from Robert Shiller's website. This set is formed with One-Year interest rate and Long Government Bond Yield (10-years of maturity) and is useful in order to examine long-run historical trends, as it begins in 1871 and finishes in 2011; our purpose, as we said previously, is the study of the term structure behaviour and monetary policy across time according to EHTS.

TABLE 7.2: Main statistics for both database

	Mean	**Median**	**Min.**	**Max.**	**SD**
Short-term rate	4.181	3.615	0.100	16.390	2.895
Long-term rate	4.635	3.921	1.802	13.911	2.272
One-Year interest rate	4.715	4.620	0.365	17.630	2.793
Long Government Bond Yield	4.654	3.980	1.950	14.590	2.246

From 1870 to 2013 for two first rows and from 1871 to 2011 for the last two rows.

We study the essential statistics as a preliminary procedure to know the data. In table 7.2, we show both databases selected for our empirical issue as explained previously. In this table, we combine the short- and long-term interest rates from *Jordà-Schularick-Taylor Macrohistory Database* and the One-Year interest rate and Long Government Bond Yield from Shiller's website. This table also shows that these statistics are very similar for both databases when short- and long-term rates are compared, so it could be a prelude to our results.

7.4.2 Econometric Strategy

The purpose of the present study is to test the existence of EHTS based on the previous methodology, and it is based on the treatment of two historical databases for US interest rates explained in the next subsection applying different tests. The first step will reside in checking if the fractional cointegration, i.e., the FCVAR model, is more appropriate than the standard cointegration. If we accept this step, we move to a second step, which involves the estimation of β under the hypothesis that the cointegrating vector is (1, -1). Continuing with that restriction, i.e., the cointegrating vector is (1, -1), we estimate the adjustment coefficients under a Fractional Vector Error Correction Model (FVECM). These adjustment coefficients would provide us information about the Permanent-Transitory decomposition. Finally, with the aim of knowing if the spread is a long memory process, we will establish the order of integration or degree of spread persistence as the difference between order of integration (d) and strength of cointegration (b). Table 7.3 performs a summary of the econometrical strategy and the order of the proposed results.

[3]The short-term interest rates references to a maturity of 3 months; meanwhile, the long-term is 10-year. For more details, visit http://www.macrohistory.net/data/

Table 7.3: Strategy of empirical research

	Procedure	Hypotheses
Step 1	Fractional cointegration?	H_1^d: Is the fractional cointegration more appropriate that traditional cointegration?
Step 2	Estimation of β	H_1^β: Cointegrating vector is (1, -1)
Step 3	Estimation of adjustment coefficients (α_R, α_r)	$H_1^\beta \cap H_1^{\alpha_i}$: Variables are weakly exogenous under the restriction of the cointegrating vector (1, 1)
Step 4	Permanent - Transitory decomposition	$H_1^\beta \cap H_{1\perp}^{\alpha_{long/short}} \equiv H_1^\beta \cap H_1^{\alpha_{short/long}}$ (mirror): Long-term rate and/or short-term rate has a permanent component in the common trend
Step 5	Degree of spread persistence, i.e., order of integration $(d-b)$	H_1^{d-b}: Is the spread a long memory process?

7.4.3 Univariate analysis

Before the application of the FCVAR model, in a preliminary step, we estimate the order of fractional integration of the historical interest rates. In order to motivate a fractionally cointegrated model, we consider univariate results observing long memory, and then we proceed to the estimation of the fractional parameter d for each univariate series, with results presented in table 7.4. These three columns are semiparametric log-periodogram regression estimates from Geweke and Porter-Hudak (1983) computed with bandwidths $m = T^{0.5}$, $m = T^{0.6}$, and $m = T^{0.7}$. The estimates are consistent with the joint estimates presented later. As we can see in table 7.4, the values for d are similar when we check it in the same time slot, i.e., if this test is applied on short- or long-term rates. We can observe as the values of d decrease when we pass from a bandwidth $m = T^{0.5}$ to $m = T^{0.6}$, and then, these values begin to increase again when bandwidth $m = T^{0.7}$. For shorter-term interest rates, the values of d are always between 0.5 and 1, which means that these processes are stationary and becoming mean-reverting values. The same occurs for longer-term interest rates, which gets values of around 1. In other words, these values suggest that the fractional cointegration is more appropriate to our approach.

Table 7.4: Univariate analysis. Geweke and Porter-Hudak (GPH) estimates

	$m = T^{0.5}$ $\hat{d}$	$m = T^{0.6}$ $\hat{d}$	$m = T^{0.7}$ $\hat{d}$
Short-term rate	0.838	0.564	0.608
	(0.246)	(0.193)	(0.125)
Long-term rate	1.138	0.970	1.080
	(0.264)	(0.167)	(0.131)
One-Year interest rate	0.884	0.551	0.616
	(0.252)	(0.167)	(0.125)
Long Government Bond Yield	1.129	1.006	1.055
	(0.137)	(0.104)	(0.108)

GPH denotes the Geweke and Porter-Hudak semiparametric log-periodogram regression estimator. Standard errors are given in parenthesis below estimates of d. The sample size is 144 for *Jordà-Schularick-Taylor Macrohistory Database*, i.e. Short and long term rates and 141 for Shiller's Database, i.e. One-Year interest rate and Long Government Bond Yield respectively.

7.4.4 Cointegration analysis

This subsection is devoted to steps 1 to 5 contained in table 7.3. First, under the Bayesian Information Criteria (BIC), we establish the optimal lag length for better accuracy in our estimation under the assumptions of the FCVAR model; there is a chance that the lag length selected would be different in each database studied. As can be observed, when we attend to *Jordà-Schularick-Taylor Macrohistory Database*, the optimal lag length is 1; meanwhile, for the Shiller's database, it is 1 as well[4]. Once the lag length is determined, we proceed to the first step, which reveals our premise about the capability of fractional cointegration in our estimation. For this, we test the cointegrating rank, evidencing that the number of cointegrating vectors is one in both databases. Subsequently, when the rank test is finalized, we test the hypothesis H_1^d, which displays if the fractional cointegration is more appropriate than traditional cointegration. The rejection of this hypothesis implies that fractional cointegration is appropriate for this study.

In our case, as we can see at the bottom of table 7.5, this hypothesis could be rejected; so we continue our estimation under the fractional cointegration premises. We note that traditional cointegration has limitations, so we consider that the shocks on our series could be persistent, following a long memory process in the residuals of the cointegrating relationship that exists; thus, a slow reversion towards the long-run equilibrium can take place. A fractional cointegration approach allows us to capture the relationships between the short- and long-term interest rates by considering that the spread could follow a fractional process $I(d-b)$; this is a long memory process and contrary to traditional cointegration, which forces this process to be $I(0)$.

TABLE 7.5: Rank test and Fractional Cointegration test

	Jordà-Schularick-Taylor	*Shiller's*
Optimal lag length	1	1
Rank test		
0	26.610 (0.000)	13.836 (0.008)
1	6.133 (0.013)	2.991 (0.252)
$\hat{d}$	0.758 (0.103)	1.041 (0.094)
$\hat{b}$	0.384 (0.198)	0.556 (0.187)
H_1^d	8.130 (0.004)	7.974 (0.005)

In this table is shown the estimations for each database in different columns. It also shows the values of LR Statistics and the P values are in brackets. For the parameters d and b we show their values and the standard deviations values are in parenthesis. The significance level is set to 10% for exclusion following Jones, Nielsen, and Popiel (2014)). The sample size is 144 for *Jordà-Schularick-Taylor Macrohistory Database* and 141 for Shiller's Database.

The next issue consists in the study of the long-run equilibrium between the short- and long-term interest rates. The estimated values are shown in table 7.6. It can be observed that parameter β is close to 1.[5] Observing that the EHTS implies that the series are cointegrated but the cointegrating vector between each variable is restricted in (1,-1), H_1^β, we must test the existence of this vector. Using an LR test as we can see in table 7.6, we do not reject this parameter restriction, concluding that the EHTS could not be rejected. This result reveals that in the last century, long-term interest rates are determined by short-term interest rates. Nonetheless, despite this result being well-known in the previous literature, this is the first time that a fractional cointegration is applied to confirm this relationship and that the database spans a long range of time.

[4]The lag length values are shown in table A.1 in the appendix.

[5]In every estimation, we check the residuals for serial correlation using a multivariate Ljung-Box Q-test, $\hat{\varepsilon}$ with h = 12 lags. The results show no evidence of serial correlation of the residuals in every estimation, and the Ljung-Box Q-test shows no signs of misspecification, which indicates that the model is well specified (see table A.2 in appendix).

TABLE 7.6: Estimates of β and H_1^{β}: Cointegrating vector restriction (1, -1)

	Jordà-Schularick-Taylor	*Shiller's*
Vector β	1.000 -1.087	1.000 -0.992
H_1^{β}	0.290 (0.590)	0.003 (0.957)

In this table is shown the estimations for each database in different columns. It also shows the values of LR Statistics and the P values are in brackets. The significance level is set to 10% for exclusion following Jones, Nielsen, and Popiel (2014). The sample size is 144 for *Jordà-Schularick-Taylor Macrohistory Database* and 141 for Shiller's Database.

The next step consists in the estimation of the FVECM (see equation 7.5); the significance of the adjustment coefficients in the joint hypothesis $H_1^{\beta} \cap H_1^{\alpha_i}$ is tested as shown in table 7.7. Using an LR test, we find that only the coefficients associated to short-term rates (α_r) are significant, which implies that the spread has a prediction power in the behavior of the futures short-term rates, which is consistent with EHTS. Finally, as expected, the adjustment coefficients of the short-term rate are positive, which is extra evidence in support of the EHTS; conversely, the adjustment coefficients of the long-term rates are much smaller in magnitude than short-term rates although insignificantly different from zero (this result is according to the results obtained by Hansen and Seo (2002)).

Finally, we decompose the FVECM in order to see which interest rate has a permanent component in the common trend. This is potentially useful information for the design and adjustment of monetary policy. To do this, we follow the methodology provided by Gonzalo and Granger (1995), i.e., the Permanent-Transitory decomposition, to establish the common trend in order to determine whether the long-term rate or short-term rate drives the common trend. In our case, both short-term rates do not contribute to the long run because the parameter α_r is different than zero. On the other hand, the parameter $\alpha_R = 0$, such that both long-term rates are weakly exogenous, being the permanent component, and this implies that this rate drives the common trend.

This can also be interpreted as proportions of short- and long-term contributions to the common trend. As we can appreciate at the bottom of table 7.7, in both databases, the longer-term rates contribute around 100% in the common trend.

TABLE 7.7: $H_1^{\beta} \cap H_1^{\alpha_i}$: Adjustment coefficients under constrained parameter (1,-1)

	Jordà-Schularick-Taylor	*Shiller's*
$H_1^{\beta} \cap H_{1,\perp}^{\alpha_{long/short}} \equiv H_1^{\beta} \cap H_1^{\alpha_{short/long}}$	0.005 (0.945)	0.415 (0.520)
$H_1^{\beta} \cap H_{1,\perp}^{\alpha_{short/long}} \equiv H_1^{\beta} \cap H_1^{\alpha_{long/short}}$	18.808 (0.000)	10.584 (0.001)
α_R	0.014	-0.127
α_r	2.436	1.408
$\alpha_{R\perp}$	1.006	0.917
$\alpha_{r\perp}$	-0.006	0.083

In this table is shown the estimations for each database in different columns. It also shows the values of LR Statistics and the P values are in brackets. The significance level is set to 10% for exclusion following Jones, Nielsen, and Popiel (2014). The sample size is 144 for *Jordà-Schularick-Taylor Macrohistory Database* and 141 for Shiller's Database. $\alpha_{R\perp}$ and $\alpha_{r\perp}$ are normalized such that the two elements add to one.

In the last step, as the cointegrating vector is (1, -1), we can interpret the difference $(d-b)$[6] as the order of integration of the spread, that is, the degree of persistence (H_1^{d-b}). According to Table 1 in Tkacz (2001), when $(d-b) = 0$, the spread follows a stationary process and the shock duration is short-lived[7]. If $0 < (d-b) < 0.5$, there is a stationary process, and the shock duration is long-lived, and finally, if $0.5 < (d-b) < 1$, the spread follows a non-stationary process, although it is mean-reverting and the shock duration is long-lived. As we can see in table 7.8, there is two-way evidence in the knowledge of this difference. On the one hand, clearly the order of integration of the spread is distinctly higher than zero and then exhibits a long-memory process. On the other hand, both databases follow a stationary process, and thus, the duration of the shock is long-lived. This result could be in line with the results that Weber and Wolters (2012) and Holmes et al. (2015) proposed in their studies.

TABLE 7.8: H_1^{d-b}: Degree of persistence of the spread

	Jordà-Schularick-Taylor	*Shiller's*
H_1^{d-b}	0.374	0.485

In this table is shown the estimations for each database in different columns. It also shows the values of LR Statistics and the P values are in brackets. The significance level is set to 10% for exclusion following Jones, Nielsen, and Popiel (2014). The sample size is 144 for *Jordà-Schularick-Taylor Macrohistory Database* and 141 for Shiller's Database.

7.5 Conclusion

The role of central banks according to monetary policy through recent history implies the application of different economic measures to fight market inefficiency in the form of economic cycles. For this reason, the treatment of the term structure, i.e., the grouping and checking of short- and long-term interest rate behavior due to the term structure of interest rates, has always been viewed as crucial to assess the impact of monetary policy and its transmission mechanism. In this situation, if a monetary policy is effective, changes in short-term policy interest rates should impact on long-term rates.

Combining two historical databases with a recent methodology, i.e., the fractionally cointegrated vector autoregressive (FCVAR) model, our analysis is capable of taking both the cointegration among short- and long-term interest rates and the long memory of the spread. We explain the spread as the difference between long- and short-term rates. This methodology provides us with the opportunity to reject the apriorism of the incompatibility of interest rates being cointegrated and the term spread non-stationary. Overall, we additionally propose a complete overview concerning the last and half century monetary policy events, aiming to clarify the proposed events by several authors and sources, in addition to a novel econometric approach that would fill the gap in the literature.

The results provided by *both databases* evidence robustness in the results in accordance with EHTS; i.e., the spread between the short-term rates and long-term rates of the term structure was found to be an optimal predictor of future short rates. Importantly, we find evidence that the spread shows a long memory process, in particular a stationary process, in contrast to the usual assumption of $I(0)$. According to the P-T decomposition, a shock in the long-term rate will have a permanent (long-run) effect on the short-term rate and long-term rate, but a shock in the short-term rate, with no movement in the long-term rate, is completely transitory. In addition, we found that the long-term rate maintains fixed with any change in the short-term rate, so this change will affect the spread $(R_t - r_t)$ only through z_t (transitory component) and therefore will only have transitory effects. In sum, we evidence

[6]According to our methodology, d and b denote the fractional order of integration of the explanatory variables and the strength of cointegration, respectively.

[7]This means that a shock would show a slow return towards the long-run equilibrium.

that the short-term rate does not contribute to the long- run, so the long-term rate is the dominant rate.

In line with the results obtained, we show long memory in the spread, which has deep implications for monetary policy. We support the study elaborated by Baillie and Bollerslev (1994a), where it was established that a long memory spread holds adequate forecasting power at longer horizons and, even more importantly, the persistence of the spread implies a gradual loss of control over interest rates by the Fed as Cömert (2012) suggests. In our case, as the estimations in each database show, the persistence is below 0.5, which could be an indicator that the Fed already has control over term structure. If the spread is stationary, the long- and short-term rates are driven by a common stochastic trend and do not allow arbitrage opportunities because market forces adjust to correct any temporary disequilibrium. Although, the P-T decomposition shows that the long-term rate is the dominant in the common trend and its proportion is around 100%. We could conclude that the long-term rate would drive the long-run behavior of this relationship.

Overall, the results evidence that across the last century and a half, there were wars, economic crises and/or changes in economic policy in the USA. In general, the EHTS is supported, and thus, there is control power over monetary policy. In addition, the results endorse the creation of a figure such as the Federal Reserve, which has maintained the effectiveness of monetary policy.

7.6 Appendix

TABLE A.1: Bayesian Information Criterion. Lag length selection

Lags	*Jordà-Schularick-Taylor*	*Shiller's*
1	**879.60**	**840.95**
2	891.09	854.76
3	902.60	853.38
4	906.64	882.61
5	918.50	857.11
6	920.48	851.96

Bold indicates lag order selected

TABLE A.2: Ljung-Box Q-test

	Jordà-Schularick-Taylor	*Shiller's*
$Q_{\hat{\varepsilon}}$	45.798 (0.564)	39.667 (0.798)

Following Jones, Nielsen, and Popiel (2014), the significance level is set to 10% for exclusion. P values are in parenthesis below LR test values.

Part III

Budget debt sustainability

Chapter 8

U.S. budget deficit sustainability revisited: Long run, persistence and common trend.

8.1 Introduction

Governments' interest in controlling public accounts and maintaining a balanced budget has been one of the key elements in their policies, and this interest has been fundamental since the financial crisis of a decade ago. The austerity scenario that has engulfed the developed countries has given rise to a broad debate on the sustainability of public accounts in the long term given the importance in developed countries of budgets for economic development and maintenance of the welfare state. In the US case, the budget debate has been intense in recent years, even more after causing the shutdown in the 2013 budget, which caused a sharp decline in income that was recovered within two weeks (Gelman, Kariv, Shapiro, Silverman, and Tadelis, 2015).

The literature devotes no less intense attention to explaining the effects that the management of public budgets has on economic development and more specifically to examining whether the level of primary surpluses responds positively to marginal changes in the debt/GDP ratio (Bohn, 1998). In particular, interest in knowing the sustainability of the public debt has clearly emerged. In this sense, although the US has suffered from extended periods of primary deficits, the seminal paper of Bohn (1998) shows that US fiscal policy has been historically sustainable. Subsequent empirical evidence strongly supports this sustainability of fiscal policy (see Martin (2000); Cuñado, Gil-Alana, and de Gracia (2004); Bohn (2007); Berenguer-Rico and Carrion-i-Silvestre (2011) or Cascio (2015), among others). However, there are studies that question the sustainability of US fiscal policy (e.g., see Hamilton and Flavin, 1986). Recently, Nguyen, Suardi, and Chua (2017) highlighted that the US public debt was sustainable until 2005, when the primary surplus to GDP reacted negatively to the debt/income ratio.

Nevertheless, in this body of literature, the degree of control of this sustainability has been only briefly studied, although there are studies that have tested the relationship of cointegration in the long term (Quintos (1995); Bohn (2007); or Irandoust (2018); among others). In addressing this issue, we will study the long-term relationship, and at the same time, we will study whether the degree of persistence of the budget balance is important for fiscal policy management. Persistence in the budget balance may stem from the immediate budgetary effects of shocks not being reversed in subsequent periods, but it may also stem from a 'snowball effect' via changes in future interest payments (Friedman, 2006). Furthermore, in this gap in the literature, the behavior of public revenues and expenditures has not been investigated from the common trend approach proposed by Gonzalo and Granger (1995). This approach can not only test the relationship between revenue and expenditure but can also determine which of the two is dominant in the common trend. In other words, it can reveal which indicators, i.e., public revenues and/or expenditures, should be adjusted to maintain the sustainability of the public budget.

In the described empirical scenario, in our chapter we propose a novel approach considering that the standard unit root and traditional cointegration test might be too restrictive ($I(1)/I(0)$ dichotomy). Indeed, we reject the assumption that both expenditures and revenues follows the dichotomy $I(1)/I(0)$, so they follow a fractional process $I(d)$, and the error term follows a stationary process ($I(0)$), in case of cointegration of both variables. Thereby, the rigidity of the traditional approach is broken in favour to allow the series to be cointegrated and the error term is does not necessarily need to be $I(0)$, i.e. we let that it could be cointegrated in order $I(d-b)$. To the best of our knowledge, our new approach consists in the Fractionally Cointegrated Vector Autoregressive (FCVAR) model developed by Johansen and Nielsen (2012) and Nielsen and Popiel (2016), which is an expansion of the CVAR, proposed by Johansen (1995). Combining this econometric approach with the conditions of Intertemporal Budget Constraint literature, we elaborate a new assumption in the fiscal policy sustainability. Additionally, we also measure the Permanent-Transitory decomposition following the Gonzalo and granger proposal in order to explore what indicators, public revenues or expenditures, or both, should be adjusted in order to maintain the sustainability of the public budget.

In the next section 8.2, this chapter proposes a review of the literature concerning the long-run relationship in the public debt. Section 8.3 details the IBC framework that allows us to construct our empirical strategy shown in section 8.4. Section 8.5 presents the results obtained by our econometric exercises that will be the basis of the main conclusions and policy implications proposed in section 8.6.

8.2 Literature review

Governments apply different fiscal policy measures to manage the earnings and the public expenditures to achieve different goals such as accelerating economic growth, fully utilizing all productive resources of society and maintaining price stability (Auerbach, Kotlikoff, et al., 1987). This approach implies the importance of paying attention to the sustainability term that represents the government's need to guarantee the capacity to complete its functions. For this reason, a sustainable fiscal policy is one that would cause the discounted value of debt to go to zero at the limit so that the present value borrowing constraint would hold, and it also requires intertemporal balancing of the government budget, setting the current value of debt as the discounted sum of expected future surpluses (Berenguer-Rico and Carrion-i-Silvestre, 2011).

Early studies concerning fiscal policy and its implications present different hypotheses such as the tax-and-spend hypothesis, which states that changes in expenditures are led by changes in tax revenues (Friedman, 1978). There is also a reverse relation, which implies the spend-and-tax hypothesis (Roberts, 1978), wherein expenditures are made prior to tax collection. Hence, the fiscal synchronization hypothesis establishes the simultaneity of the spending and revenue decisions (Meltzer and Richard, 1981). However, an important issue in the study of fiscal policy is raised by the paper published by Sargent and Wallace (1984), which examined the implications of the government's budget constraint in the measures carried out by economic authorities. This chapter is an initial step in the IBC conception and provides a basis for numerous studies that assume the existence of such constraints.

In this sense and agreeing with the IBC literature, this approach requires that the current debt be financed by surpluses in future periods, as Trehan and Walsh (1988, 1991), among others, found in their survey. This condition holds if a long-run cointegrating relationship exists, i.e., stationarity of the error term, between public expenditures and revenues. In particular, we can determine two types or degrees of sustainability depending on the cointegrating vector. Therefore, from a wide perspective, Quintos (1995) establishes a "strong" and a "weak" sustainability to categorize whether treatment of the coefficient that relates the two variables is necessary. As a result, it could be mentioned that there is a "strong" sustainability if the difference between expenditures and revenues is 1 (or not significantly different from 1), and "weak" sustainability would exist when the coefficient is strictly smaller than 1. More recently, Bohn (2007) evidenced that the IBC conditions are compatible with any order of integration of the variables involved.

The existing literature follows the hypotheses explained previously and is based on whether fiscal policy is consistent with the IBC, recognizing two research strands depending on the techniques used. The first strand is based on a long-run cointegrating relationship, while the second strand considers a non-linear relationship. Thus, according to our goal, we follow the strand in the literature that tests the presence of cointegration between government revenues and expenditures (see Trehan and Walsh (1988, 1991), Hakkio and Rush (1991), Haug (1991), and Quintos (1995), for instance). In the particular case of US fiscal policy, there are influential papers with the same goal, i.e., contrasting the existence of a long-run relationship (such as Bohn (1998), Martin (2000), Cuñado et al. (2004) or Cascio (2015)), determining the existence of a threshold in the relationship (see Arestis, Cipollini, and Fattouh, 2004), analyzing structural breaks (see, e.g., Bajo-Rubio, Dıaz-Roldán, and Esteve (2008) or Berenguer-Rico and Carrion-i-Silvestre (2011), Afonso and Jalles (2014)) or using a quantile cointegration (Chen, 2016) or a Markov switching cointegration approach (Gabriel and Sangduan, 2011).

This issue has also been applied in different countries and confirmed that the IBC could be consistent in different environments to test fiscal policy sustainability (see, for instance, Bajo-Rubio, Dıaz-Roldán, and Esteve (2010), Bajo-Rubio, Diaz-Roldan, and Esteve (2014), Afonso, Agnello, Furceri, and Sousa (2011), Paleologou (2013), Camarero, Carrion-i-Silvestre, and Tamarit (2015) or Irandoust (2018)). Hence, if the linearity assumption is given, some implications should be considered. We refer to the symmetric responses of expenditures and revenues to the fiscal policy actions irrespectively of their initial conditions (see Giavazzi, Jappelli, and Pagano (2000) or Cipollini (2001), where they evidenced that this is empirically unlikely).

In sum, although these studies depend on the different econometric techniques used and the countries analyzed, one remarkable conclusion that arises from these studies is that we can identify deficit sustainability under the IBC and under different approaches, such as structural breaks or asymmetry tests (Chen, 2016). Overall, the literature has been based on cointegration techniques; we have not found many studies focused on fractional cointegration. We follow the methodology developed by Johansen and Nielsen (2012) and Nielsen and Popiel (2016), which is an expansion of the CVAR proposed by Johansen (1995), and combine this econometric approach with the conditions of the IBC literature, which allows a new assumption in fiscal policy sustainability. This could be a next step in this field.

8.3 Deficit and debt sustainability framework

This section is devoted to the explanation of the IBC literature, which requires that the current debt be financed by surpluses in future periods. Our starting point follows the IBC version and assumes that budget deficits are financed using bonds; therefore, the one-period government budget constraint is:

$$\Delta B_t = G_t - R_t \tag{8.1}$$

where B_t denotes the real market value of government debt, G_t is real government expenditures inclusive of interest payments andR_t represents real tax revenues and $\Delta = (1 - L)$ is the first difference operator. Indicating i_t as the real interest rate and accepting i_t to be stationary around a mean i as Hakkio and Rush (1991) did, we can describe EXP_t as the real expenditure exclusive of interest payments and $i_t B_{t-1}$ as the interest payments on the level of debt accumulated at the end of the previous period. We obtaint he following:

$$G_t = EXP_t + i_t B_{t-1} \tag{8.2}$$

Equation (8.3) is the expression of the debt and is shown as

$$B_t = (1+i)B_{t-1} + EXP_t - R_t \tag{8.3}$$

where

$$B_t = \left(\frac{1}{1+i}\right)(R_{t+1} - EXP_{t+1}) + \left(\frac{1}{1+i}\right)B_{t+1} \tag{8.4}$$

As this restriction remains in different periods, we solve for B_t via forward substitution and we obtain

$$B_t = \sum_{j=0}^{\infty}\left(\frac{1}{1+i}\right)^{j+i}(R_{t+j+1} - EXP_{t+j+1}) + \lim_{j\to\infty}\left(\frac{1}{1+i}\right)^{j+1}B_{t+j+1} \tag{8.5}$$

The budget balance across time holds if and only if the current value of outstanding debt is equal to the present value of future budget surpluses. Denoting E_t as the expectation operator, conditional on information at time t, fiscal sustainability implies:

$$\lim_{j\to\infty} E_t\left(\frac{1}{1+i}\right)^{j+1}B_{t+j+1} = 0 \tag{8.6}$$

If we satisfy condition (8.6), we could assume that the budget deficit is sustainable; thus, the stock of debt held by the public is expected to grow no faster than the growth rate of the economy. Taking first differences on equation (8.4), we obtain the following:

$$\Delta B_t = \sum_{j=0}^{\infty}\left(\frac{1}{1+i}\right)^{j+i}(\Delta R_{t+j+1} - \Delta EXP_{t+j+1}) + \lim_{j\to\infty}\left(\frac{1}{1+i}\right)^{j+1}\Delta B_{t+j+1} = 0 \tag{8.7}$$

and indeed, sustainability requires the condition

$$\lim_{j\to\infty} E_t\left(\frac{1}{1+i}\right)^{j+1}\Delta B_{t+j+1} = 0 \tag{8.8}$$

As mentioned in the literature review above, the empirical approximation that we follow is based on the cointegration-based approach, which measure whether R_t and G_t are cointegrated. This approach, based on Hakkio and Rush (1991) and Quintos (1995), among others, allows us to obtain a better interpretation of the analysis in terms of the degree of sustainability. We can also distinguish between two degrees of sustainability, as we explained previously: the strong and weak sustainability of the budget deficit. Under the restriction $\beta = 1$ we could contrast the existence of strong sustainability when R_t and G_t are cointegrated; if $0 < \beta < 1$ the fiscal deficit would be weakly sustainable and however, if $\beta \leq 0$ the fiscal deficit would be unsustainable. Finally, the regression model would be:

$$R_t = \alpha + \beta G_t + \varepsilon_t \tag{8.9}$$

8.4 Econometric approach

This section gives a specification for testing sustainability under fractional cointegration conditions. This model is provided by Johansen and Nielsen (2012) and further developed by Nielsen and Popiel (2016) and has the advantage of being used for stationary and non-stationary time series. To introduce the FCVAR model, we introduce the fractional difference operator into the CVAR model, which is Δ, and the fractional lag operator is $\Delta^b = (1 - L_b)$. We replace lag operators by their fractional counterparts Δ^b and $L_b = (1 - \Delta^b)$, we obtain:

$$\Delta^b Y_t = \alpha\beta' L_b Y_t + \sum_{i=1}^{k}\Gamma_i \Delta^b L_b^i Y_t + \varepsilon_t \tag{8.10}$$

and applying $Y_t = \Delta^{d-b}X_t$, we obtain the FCVAR model:

$$\Delta^d X_t = \alpha\beta' L_b \Delta^{d-b} X_t + \sum_{i=1}^{k} \Gamma_i \Delta^d L_b^i X_t + \varepsilon_t \tag{8.11}$$

where ε_t is p-dimensional independent and identically distributed with mean zero and covariance matrix Ω. The parameters α and β are $p \times r$ matrices, where $0 \leq r \leq p$. In matrix β the columns are the cointegrating relationships and $\beta' X_t$ assume the existence of a common stochastic trend that is integrated of order d and b, indicating the strength of the cointegrating relationships, and the short-term parts from the long-run equilibrium being integrated of order $(d - b)$. The coefficients in α correspond to the speed of adjustment to equilibrium. Hence, $\alpha\beta'$ is the long-run adjustment, ρ' is the restricted constant term and Γ_i represents the short-run behavior of the variables. Johansen and Nielsen (2012) considered the insertion of the restricted constant term ρ in the long-run cointegrating relation. Dolatabadi et al. (2018) suggested an unrestricted constant ξ as the linear trend of the fractionally integrated processes. The following specification shows a more general form:

$$\Delta^d X_t = L_d \alpha(\beta' X_t + \rho') + \sum_{i=1}^{k} \Omega_i \Delta^d L_d^i X_t + \xi + \varepsilon_t. \tag{8.12}$$

where ρ is denoted as the restricted constant term, i.e., the mean level of equilibrium relation, and ξ is the unrestricted constant term that generates a deterministic trend in the levels of the variables (Dolatabadi et al., 2018).

There are two additional parameters in the FCVAR model compared with the CVAR model. The parameter d represents the order of fractional integration of the observable time series. The parameter b determines the degree of fractional cointegration, that is, the reduction in fractional integration order of $\beta' X_t$ compared to X_t itself. The relevant ranges for b are (0, 1/2), in which case the equilibrium errors are fractional of order greater than 1/2 and are therefore non-stationary although mean reverting, and (1/2, 1], in which case the equilibrium errors are fractional of order less than 1/2 and are stationary (Dolatabadi et al., 2016). The error correction term is integrated from order $(d - b)$, that is *I(0)* in this case. However, in fractional cointegration, these axioms are relaxed because $(d - b) = 0$, i.e. the error correction term shows a short-run stationary behaviour; or $(d - b) > 0$, i.e. there is a long memory process, and the error correction term will revert to its mean in the long run. As the cointegrating vector is $(1, -\beta)$, we can interpret the difference $(d - b)$ as the order of integration of the deficit, that is the degree of persistence (H_1^{d-b}).[1] According to Table 1 in Tkacz (2001), when $(d - b) = 0$, cointegrating error follows a stationary process, and the shock duration is short-lived. If $0 < (d - b) < 0.5$, there is a stationary process and the shock duration is long-lived, and finally, if $0.5 < (d - b) < 1$, cointegrating error follows a non-stationary process, although mean-reverting and the shock duration is long-lived. Thus, the main contribution of this chapter is the elaboration of new conditions of the degree of fiscal sustainability based on IBC, establishing β as a previous condition of sustainability and keeping the attention in the persistence of the error term. These conditions are synthesized and represented in table 8.1.

To determine the number of stationary cointegrating relations, the hypotheses in the rank test are followed based on a series of LR tests. In the FCVAR model, we test the hypothesis $H_0 : rank(\Pi) = r$ against the alternative $H_1 : rank(\pi) = p$, $L(d, b, r)$ being the profile likelihood function given a rank r, where (α, β, Γ) have been reduced by rank regression (see Johansen and Nielsen, 2012). Maximizing the profile likelihood distribution under both hypotheses, the LR test statistic is now $LR_t(q)$. The asymptotic distribution of $LR_t(q)$ depends on the parameter b and on $q = n - r$. MacKinnon and Nielsen (2014), based on their numerical distribution functions, provide asymptotic critical values of the LR rank test. In the case of $0 < b < 1/2$, $LR_t(q)$ has a standard asymptotic distribution, $LR_t(q) \xrightarrow{D} \chi^2(q^2)$.

According to the latter literature, fractional cointegration implies a FVECM such as:

[1] Agreeing to our methodology, d and b means the fractional order of integration of the explanatory variables and the strength of the cointegration, respectively.

$$\begin{pmatrix} \Delta R_t \\ \Delta G_t \end{pmatrix} = \begin{pmatrix} \alpha_R \\ \alpha_G \end{pmatrix} (R_{t-1} - \beta G_{t-1} - c) + \sum_{i=1}^{n} \Gamma_i \begin{pmatrix} \Delta R_{t-i} \\ \Delta G_{t-i} \end{pmatrix} \begin{pmatrix} u_{1t} \\ u_{2t} \end{pmatrix} \tag{8.13}$$

With adjustment parameters α, cointegration coefficients β, restricted constant (c), lag length (n) and errors u. The Γ_i are 2 x 2 parameter matrices in the short-run dynamics. The adjustment coefficients α_R and α_G capture the speed of adjustment of R_t and G_t towards equilibrium.

Table 8.1: Sustainability of public debts

Order of integration of the error correction term (ECT)	**Value of β**	
	$\beta = 1$	$0 < \beta < 1$
$I(d-b) = I(0)$	Strong sustainability and the shock duration is **short-lived.**	Weak sustainability and the shock duration is **short-lived.**
$I(0) < I(d-b) < I(0.5)$	Strong sustainability and the shock duration is **long-lived.**	Weak sustainability and the shock duration is **long-lived.**
$I(0.5) < I(d-b) < I(1)$	Strong sustainability and follows a **non-stationary** process, although mean-reverting and the shock duration is **long-lived.**	Weak sustainability and follows a **non-stationary** process, although mean-reverting and the shock duration is **long-lived.**

The first row corresponds to the two traditional cases of the IBC. If $\beta = 1$, the error correction term could be interpreted as the budget deficit.

To determine the magnitude of each variable in the long run, we use the Permanent-Transitory decomposition (Gonzalo and Granger, 1995)[2] component share (CS). This method has been developed and applied formerly by Dolatabadi et al. (2016), Dolatabadi et al. (2018) for the fractionally cointegrated VAR model. As Baillie et al. (2002) note that since which notice that since $\alpha'\alpha_\perp = 0$, may also be expressed in terms of the elements of the error correction vector α. To interpret this, we let $\alpha = (\alpha_R, \alpha_G)'$ and $\alpha \perp \alpha = (\alpha_{\perp,R}, \alpha_{\perp,G})'$. Afterwards, $\alpha'_\perp \alpha = \alpha_{\perp,R}\alpha_R + \alpha_{\perp,G}\alpha_G = 0$ implies that $\alpha_{\perp,R} = -\alpha_{\perp,G}\alpha_G/\alpha_R$ and so, Component Share (CS here after) may be expressed as

$$CS_R = \frac{\alpha_G}{\alpha_G - \alpha_R}, CS_G = \frac{\alpha_1}{\alpha_G - \alpha_R} \tag{8.14}$$

In this respect, the CS for variable 1 reflects how sensitive variable 2 is, relative to variable 1 and vice-versa.

8.4.1 Data and empirical strategy

Data

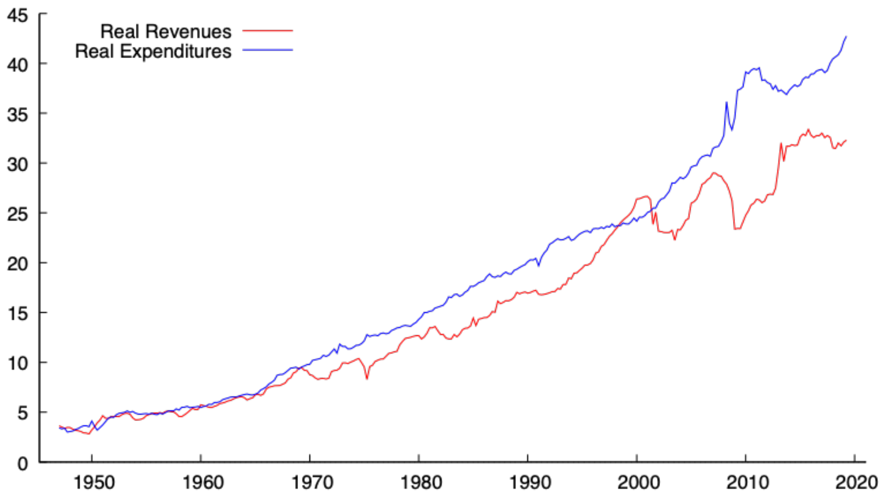
In this section, we provide a test of the sustainability of US fiscal policy over the period that covers, in a quarterly sample, 1947Q1 to 2019Q2, amounting to 290 observations.[3] The dataset selected corresponds to real expenditures inclusive of interest paid on debt, and real revenues. The data are collected from the US Bureau of Economic Analysis in Tables I.1.9 and III.1. In Figure 8.1, the evolution of the series, i.e., real expenditures and revenues of the USA, are plotted. As can be observed, since the Vietnam War in the middle of the 1960s,

[2] See Appendix A for a further develop of this technique.

[3] See Table A.4 in the Appendix B for a summary of the main policy measures and events that occurred during the analyzed period.

expenditures are always higher than revenues in almost all quarters, as expenditures came to increase notably relative to revenues to finance social programs and the war, although this situation reverted at the end of the 1990s and the beginning of the 2000s because this decade was a period of very high economic growth in the USA.

FIGURE 8.1: Time series plot for US government real expenditures and revenues



Empirical strategy

As we explain in earlier sections, our goal consists in testing the fiscal policy sustainability for The USA. Of course, we follow the IBC conditions, i.e. we check if R_t and G_t are cointegrated and testing if $\beta = 1$ in order to define how is the sustainability degree of fiscal policy under the assumptions of the fractional cointegration (FCVAR model). This approach gives a new view in the procedure as we represent in table 8.1. The next table 8.2 exhibits the procedure that we follow to check the type of sustainability of the US fiscal policy which is based in the assumption established in table 8.1.

TABLE 8.2: Strategy of empirical research

	Procedure	Hypotheses
Step 1	Fractional Cointegration?	H_1^d: Is the fractional cointegration more appropriate than traditional cointegration?
Step 2	Estimation of β	H_1^β: How sustainable is the US fiscal policy?
Step 3	Degree of deficit persistence, i.e. order of integration *(d-b)*	H_1^{d-b}: Is the deficit a long memory process?
Step 4	Estimation of adjustment coefficients (α_R, α_G)(FVECM)	$H_1^\beta \cap H_1^{\alpha_i}$: Variables are weakly exogenous under the restriction of the cointegrating vector (1, -1)
Step 5	Permanent-Transitory decomposition	$H_1^\beta \cap H_{1\perp}^{\alpha_{G/R}} \equiv H_1^\beta \perp H_1^{\alpha_{R/G}}$(mirror): Expenditures and/or revenues represent a permanent component in the common trend.

We start our econometric exercise studying the possibility that the fractional cointegration would be more appropriate than the traditional one. In this regard, we consider that the standard unit root and traditional cointegration test might be too restrictive ($I(1)/I(0)$ dichotomy). Indeed, the estimation methods of traditional cointegration were developed assuming that $d = b = 1$, i.e. the linear combination is weakly dependent stationary. Then, we reject the assumption that both expenditures and revenues follows the dichotomy $I(1)/I(0)$, so they follow a fractional process $I(d)$, and the error term follows a stationary process

$I(d-b)$, in case of cointegration of both variables, allowing that d and b to be real numbers. So, the rigidity of the traditional approach is broken and enables substantially more flexible modelling of long-run relationships between time series. Once this is done, we test the degree of sustainability of the US fiscal policy and then, under this estimation, we examine if the deficit shows a long memory process in the step 3. Finally, in steps 4 and 5, under Fractionally Vector Error Correction Model (FVECM) premises, we study which of our variables has a permanent behaviour in the common trend.

8.5 Empirical results

Before the application of steps 1 to 5, a univariate analysis is checked to study the suitability of the fractional procedure. There are several procedures for estimating the fractional differencing parameter in semiparametric contexts. Although the semiparametric log-periodogram regression proposed by Geweke and Porter-Hudak (1983) is the most used, this method was modified and further developed by Robinson (1995) and has been analyzed by Velasco (1999) and Shimotsu and Phillips (2002), among others. Next, we proceed to the estimation of the fractional parameter d for each univariate series, and the results are presented in Table 8.3. The first three columns are semiparametric log-periodogram regression estimates from Geweke and Porter-Hudak (1983), here labelled GPH, computed with bandwidths $m = T^{0.4}$, $m = T^{0.5}$, and $m = T^{0.6}$.[4]

TABLE 8.3: Univariate analysis

	GPH estimates		
	$m = T^{0.4}$ $\hat{d}$	$m = T^{0.5}$ $\hat{d}$	$m = T^{0.6}$ $\hat{d}$
Real Revenues	0.268 (0.400)	0.617 (0.224)	0.772 (0.137)
Real Expenditures	1.007 (0.141)	1.077 (0.085)	1.325 (0.098)

GPH denotes the Geweke and Porter-Hudak semiparametric log-periodogram regression estimator. Standard errors are given in parenthesis beneath estimates of d.The sample size is 290

This subsection is dedicated to show the steps 1 to 5 detailed in table 8.2. First, we establish the optimal lag length for a better accuracy in our estimation under the assumptions of FCVAR model and under the Bayesian Information Criteria (BIC); the lag length order selected is 1. Once the lag length is determined, we test the cointegrating rank, evidencing that the number cointegrating vectors is one[5]. Next, we proceed to the first step which shows whether our premise about the property of fractional cointegration in our estimation. For this, when the rank test is finished, we test the hypothesis H_1^d which exposes if the fractional cointegration is more appropriate that traditional cointegration. The rejection of this hypothesis implies that the fractional cointegration is appropriate for this study.

In our case, as we can see at the top of table 8.4, this hypothesis could be rejected; therefore, we continue our estimation under the fractional cointegration premises. We remember that traditional cointegration has restrictions, mainly the dichotomy $I(0)/I(1)$ that we explained in the previous sections. Considering that shocks on our series could be persistent, i.e., a long memory process in the residuals of the cointegrating relationship exists, a reversion towards the long-run equilibrium can occur. The FCVAR approach allows us to describe the relationships between the real expenditures and revenues by considering that the error correction term could follow a fractional process $I(d-b)$, that is, a long memory process,

[4] To test the presence of unit roots, the estimates were obtained by using first-differenced data because the original series might be greater than 0.5 and because this test requires that the results be limited to the interval $-0.5 < d < 0.5$ and then by adding 1 to obtain the proper estimates of d.

[5] The lag length values and the rank test are shown in tables A.1 and A.2 in the Appendix B, respectively.

and contrary to the traditional cointegration, which requires the process to be $I(0)$.

The next issue consists in the study of the long-run equilibrium between the real expenditures and revenues. It can be observed that the parameter beta is close to 1.[6] We constraint the relationship to vector (1,-1) to determine how sustainable US fiscal policy is, and contrasting with H_1^{β}, we test the existence of $\beta = 1$. Using a LR test, we cannot reject the null H_1^{β} that $\beta = 1$, concluding that the fiscal policy is strongly sustainable, according to the IBC suggestions, as is shown in table 8.1. Given that the cointegrating vector is (1, -1), we can interpret the difference $(d - b)$ as the order of integration/degree of persistence (H_1^{d-b}) of the error correction term, in our case, the budget deficit. As previously mentioned, when $(d - b) = 0$, the error follows a stationary process, and the shock duration is short-lived. If $0 < (d - b) < 0.5$, there is a stationary process, and the shock duration is long-lived. Finally, if $0.5 < (d - b) < 1$, the error follows a non-stationary process, although it is mean-reverting and the shock duration is long-lived. At the bottom of this table 8.4, we can see that the parameters of interest ($\hat{d}$ and $\hat{b}$) are significantly different from zero and different from each other (thus $d \neq b$). Therefore, the hypothesis H_1^{d-b} reveals evidence of the possible persistence of the error (as we impose the (1, 1) restriction on the cointegrating vector, the error correction term is the difference between revenues and expenditure, i.e., the budget deficit in our case). The order of integration of the error (budget deficit) reaches a value of 0.183, below 0.5, which implies that the US budget deficit follows a long-memory process. Indeed, the adjustment process towards equilibrium will take a long time. This finding implies a long run equilibrium relationship between public expenditures and public revenues (Cuñado et al., 2004).

The empirical results may also possess some lessons for the long-term viability of the public finance stance. Accordingly, our results reveal that while the US budget is strongly sustainable according to the IBC theory, it follows a non-stationary process, although mean-reverting, and the shock will be long-lived, as is shown in the empirical strategy supported in this article (see Table 8.1). The most important fact derived from this result is that the US fiscal budget presents persistence, and consequently, it would be difficult for the economic authority to control it. This is potentially useful information for the design and adjustment of fiscal policy. If the budget balance is not stationary, shocks affect the fiscal balance in the long term; in this case, shocks affecting the fiscal balance negatively might be particularly troublesome (Cuestas and Staehr, 2013).

TABLE 8.4: Fractional cointegration test and results

Hypothesis	**Cointegrating vector**	**LR statistics**	***P* value**
H_1^d		68.395	0.000
H_1^{β}	(1, -0.926)	1.663	0.197
Restricted model. Cointegrating vector (1, -1)			
	Value	$\hat{d}$	$\hat{b}$
H_1^{d-b}	0.183	1.003 (0.075)	0.820 (0.157)

The top of the table shows the LR statistics and P values. Standard errors are in parenthesis below values of $\hat{d}$ and $\hat{b}$.

At this point, the question we propose as fundamental to this design of fiscal policies is which of the two components, expenditures or revenues, could be adjusted with the aim of reversing this scenario. Thus, we follow the Permanent-Transitory decomposition by Gonzalo and Granger (1995) to establish the common trend in order to determine whether expenditures

[6]In every estimation, we check the residuals for serial correlation using a multivariate Ljung-Box Q-test,$Q_{\hat{\varepsilon}}$ with $h = 12$ lags. The results show no evidences of serial correlation of the residuals in every estimation; the Ljung-Box Q-test shows no signs of misspecification, which indicates that the model is well specified (see table A.3 in Appendix B).

or revenues drive the common trend. In addition, following Putniņš (2013), an interpretation of the permanent and transitory component applied to public revenues and expenditures allows us to know how their management would facilitate long-term equilibrium, since the permanent component in the public budget by analogy could be interpreted as the *common efficient budget*. Consequently, in the long term, budget management must involve the permanent component and try to avoid shocks that generate transitory components. Thus, policymakers must design public budgets without forgetting the basic budgetary principles; that is, they must contemplate at all times the resources that have a permanent component in the US economy.

Thus, in table 8.5, in the last step, we decompose the FVECM to see which element has a permanent component in the common trend. In our case, according to subsection 8.4.1, real expenditures and revenues contribute in the long run because neither parameter is different from zero. We also find that, as the two selected variables are permanent and have significant importance, the US government must take them into consideration when elaborating the budget because both indicators lead the common trend. Attending to the CS, this can also be interpreted considering the proportions of expenditures and revenues that contribute to the common trend; as we can see, both indicators are similar in percentage but slightly higher in the case of expenditures, evidencing as we said previously that the common trend is share by both expenditures and revenues.

TABLE 8.5: FVECM results (restricted model)

Hypothesis		**LR-Statistics**	**P-value**
$H_1^{\beta} \cap H_1^{\alpha_G} \equiv H_1^{\beta} H_1^{\alpha_R}$		8.407	0.004
$H_1^{\beta} \cap H_1^{\alpha_R} \equiv H_1^{\beta} H_1^{\alpha_G}$		2.701	0.099
Adjustment coefficients and Component Share			
α_R	-0.079	CS_R	44.76%
α_G	0.064	CS_G	55.24%

We reference the mirror hypothesis. The rest of the table shows the LR statistics and P values. Component shares are normalized such that the two elements sum to one.

8.6 Conclusion

The empirical testing of the sustainability of the government following the conditions of the IBC is generally based on the analysis of the past behavior of the fiscal variables. Otherwise, we may argue that the fact that the public finances have been consistently conducted in a sustainable way, it could provide a good indicator for its future behavior. This chapter shows a new approach in the literature to provide additional evidence on the long-run sustainability of the US government fiscal policy, using quarterly data covering a period from 1947Q1 to 2019Q2. Using the new approach of the FCVAR model provides us significant advantages concerning the degree of sustainability. Overall, this chapter presents a novel empirical strategy that detects the different types of sustainability based on the values of beta and the difference of $(d-b)$. In our understanding, the prism under which the cointegration approach has been applied to the debt sustainability analysis has been very limited. The concept of strong sustainability usually assumes that the relation of income and expenses are stationary and any shock is short-lived. However, the FCVAR breaks this assumption so that although a unitary relationship between income and expenses exists, their cointegration relationship can be long-lived and even non-stationary. In other words, the strong sustainability concept proposed by the IBC theory should be taken with caution when it is tested empirically in the sense that, despite contemplating cointegration between expenses and income, any shock could have long-lived temporary effects.

This chapter confirms the existence of a cointegration relationship between expenditures and revenues and provides evidence that the US budget deficit shows strong sustainability.

Furthermore, focusing on the degree of persistence of the budget balance, it is a key question for fiscal policy management. The results also support that the budget deficit follows a non-stationary process but reverts to its mean, which could suggest that the impact of a shock on the budget deficit would be long-lived; i.e., showing a slow speed adjustment towards the equilibrium of the public accounts. Consequently, strong measures would be necessary to neutralize exogenous shocks and to support the fiscal balance adjustment when those shocks affect it negatively, particularly troublesome.

Furthermore, attending to the FVECM, the Permanent-Transitory decomposition and subsequently the component share, we have found that expenditures and revenues are permanent components in the common trend and that expenditures are sensitive to revenues in a similarly manner to how revenues are sensitive to expenditures. Therefore, we can conclude that the two variables influence each other in the same way, that is that the US government has been equilibrating the flows of revenues and expenditures in the long-run, implying a link in the behavior of the deficit with the fiscal policy, achieving a fiscal improvement. Some policy implications derived from the results could help to design a better portfolio for the US government. Primary, the control power over the budget is decreased by economic authorities, whose economic measures should consider that both indicators, i.e., expenditures and revenues, have almost the same importance in the long run. Hence, if the US government aspired to achieve a strong sustainability and avoid long-lived shocks, the burden of correcting budgetary disequilibria is entirely carried out via policy mixes, i.e., by treatment of the expenditures and revenues, and never on one or the other independently. Nevertheless, if it had to discriminate between the two, budget design should begin by reviewing expenditure items but not forget the revenue structure. As expenditure programs can be handled more easily than complex tax legislation, we would expect the slightly expenditure dominance to increase with the shortening of the planning horizon. Finally, even more importantly, this chapter warns that the budget design must principally consider the permanent component to obtain a *common efficient budget*.

8.7 Appendix A

According to the Gonzalo and Granger (1995) Permanent-Transitory decomposition is defined as $X_t = (R_t, G_t)'$, where R_t represents real taxes revenues and G_t is real government expenditures inclusive of interest payments, respectively. In the Permanent-Transitory decomposition X_t can be decomposed into a transitory (stationary) part, $\beta' X_t$, and a permanent part, $W_t = \alpha'_\perp X_t$ where $\alpha'_\perp \alpha = \alpha' \alpha_\perp = 0$. W_t is the common permanent component of X_t and it is interpreted as the dominant indicator, where the information that does not affect W_t will not have a permanent effect on X_t. We focus to the key parameter $\alpha_\perp$ in order to know which indicator contributes to the common trend. Following the mirror hypothesis, the linear hypothesis on $\alpha_\perp$ can also be tested directly on $\alpha_\perp$ or alternatively on α itself using the values of LR tests in each hypothesis and critical values can be taken from the χ^2 distribution for testing. For example, to test the hypothesis that the dominant parameter is the real taxes revenues, i.e. $\alpha_\perp = (0, a)'$ we can equivalently test the mirror hypothesis $H_0 : \alpha = (\gamma, 0)'$. Similarly, to test the hypothesis that the dominant parameter is the real government expenditures, i.e. $\alpha\perp = (a, 0)'$, we test the mirror hypothesis $H_1 : \alpha = (0, \gamma)'$ (see Dolatabadi et al. (2016), Dolatabadi et al. (2018), who first combined the FCVAR with permanent-transitory decomposition).[7]

An interpretation of the coefficient α is that an adjustment coefficient measures how disequilibrium errors could be affected in current changes in X_t. Under this interpretation, we wonder if any coefficients in α are zeros, i.e. the variable in question is weakly exogenous. For example, under the hypothesis H_1, the parameter $\alpha = 0$ means that the real government expenditures do not react to the disequilibrium error, i.e., the transitory component, implying that real government expenditures is the main contributor to the common trend.

[7]The interpretation of component is linked to the concept of weak exogeneity for the cointegrating parameters (Zivot, 2000).

8.8 Appendix B

Table A.1: Lag length selection

Lags	**AIC**	**BIC**
1	-2507.32	**-2455.94**
2	**-2514.31**	-2348.25
3	-2510.68	-2429.94
4	-2508.32	-2412.90
5	-2504.42	-2394.32
6	-2500.15	-2375.38

Bold indicates lag order selected

Table A.2: Rank test

Rank	$\hat{d}$	$\hat{b}$	**Log-likelihood**	**LR statistics**	***P* value**
0	0.867	0.711	1257.420	20.483	0.000
1	1.071	1.071	1265.662	3.999	0.432
2	1.069	1.069	1267.661	—	—

Following Jones et al. (2014), the significance level is set to 10% for exclusion.

Table A.3: Ljung-Box Q-test

$Q_{\hat{\varepsilon_R}}$	12.699 (0.391)
$Q_{\hat{\varepsilon_G}}$	10.397 (0.581)

Following Jones et al. (2014), the significance level is set to 10% for exclusion. *P* values are in parenthesis below LR test values

TABLE A.4: Summary of events

Reference	Date	Event	Reason	Implication
Weidenbaum (1986)	Post-World War II	Council of Economic Advisers (CEA) creation	Provide advice and facilitate the application of a wide range of national and international economic policy issues	Replacement from a "cyclical model" of the economy to a "growth model"
Figlio and Fletcher (2012)	1944-1952	Servicemen's Readjustment Act	Provide the WWII soldiers an access to the financing of technical or university studies, joint to with a pension that would help their subsistence by one year and also granted the soldiers facilities to obtain loans to acquire homes or start a business on their own	
Naya (1971)	1955-1975	Vietnam War	Increase of spending to finance social programs and the war	
Buckley and Cleary (2010)	1972	Veterans Readjustment Benefits Act	Similar to Servicemen's Readjustment Act (1944) but focused to Vietnam War soldiers and adjusted to the cost of living and the demands of the labor market	
Hogan (1985)	1974	Congressional Budget and Impoundment Control Act	Allow Congress to challenge the president's budget more easily	Deficits became increasingly difficult to control
Ramey (2011)	1981	The Soviet invasion of Afghanistan	A significant reversal in U.S. defense policy	A significant and prolonged increase in defense spending
Bordo (2011)	1990	Budget Enforcement Act	Fiscal deficits reduction from close to 5% to a surplus	Combination of cuts in government spending and rise in tax rates
Bordo (2011)	1993	Omnibus Budget Reconciliation Act		
Kraay and Ventura (2005)	Beginning of 2000's	Dot-com bubble and September 11^{th} attacks		A painful fiscal adjustment due to the cost of Afghanistan and Iraq wars
NBER	2007	Subprime mortgage crisis		Fiscal stimulus and bank bailout
Wilson (2012)	2009	American Recovery and Reinvestment Act (ARRA)	Developed in response to the Great Recession (save existing jobs and create new ones, investing in infrastructure, education, health, and renewable energy)	Economic stimulus package
Gelman, Kariv, Shapiro, Silverman, and Tadelis (2015)	2013	Government shutdown	A series of legislative battles surrounding the Affordable Care Act (ACA) (Obamacare)	Congress did not pass legislation to appropriate funds for fiscal year 2014

Own elaboration

Part IV

Energy commodities pricing

Chapter 9

A new way of measuring the WTI – Brent spread. Globalization, shock persistence and common trends.

9.1 Introduction

This chapter analyses the possible relationship between two of the main indicators of the oil market, the North Sea Brent (Brent) and West Texas Intermediate (WTI) crude oil prices, by using the Fractional Cointegration Vector Autoregressive (FCVAR hereafter) model to determine whether these markets are regionalized or globalized. We also apply the P-T decomposition in order to identify which crude oil drives the common trend and the Brent-WTI price structure. The methodological approaches recently taken to analyse this long-run relationship usually apply the standard cointegration techniques or analyse the persistence of the spread (or relationships) in order to determine whether the markets are globalized or regionalized. However, a controversy emerges between these empirical approaches because it is assumed, on one hand, that the markets are globalized when there is cointegration in the relationship and, on the other hand, that the markets are regionalized when the spread is stationary. In that sense, this article allows us to analyse a previously uncontemplated scenario, the possibility that the series are cointegrated in the globalized market but that the spread is a nonstationary regionalized market. The use of the FCVAR provides important advantages compared to previous approaches. In particular, the rigidity of the traditional approaches is overcome in favour of allowing the series to be cointegrated, and the error term does not necessarily need to be $I(0)$; for example, we allow the error term to be cointegrated in order $I(d-b)$, unlike other techniques that assume the error term is $I(0)$. Under this assumption, we determine a controversy derived from the standard cointegration does not allow the spread to be nonstationary. The FCVAR model breaks these restrictions, providing a more reliable framework for determining if markets are regionalized or globalized.

This empirical interest stems from the well-known strategic role of the crude oil market in the world economy. Crude oil is a widely traded commodity that often affects other commodity prices and other financial markets, and oil price shocks can precipitate macroeconomic adjustments in some countries (Ji and Fan, 2012; or Coronado, Fullerton Jr, and Rojas, 2017). Indeed, in developed countries, crude oil has become the main source of energy, accounting for approximately 40% of energy sources, and is present in all productive sectors. Several factors in the past decades have generated important imbalances in this market. On one hand, there is a growing demand for oil from emerging countries such as China and India. On the other hand, the technological innovation of combining horizontal drilling with hydrofracturing has created an oil boom within the United States (Feyrer, Mansur, and Sacerdote, 2017). In addition, events in unstable producing countries such as Nigeria or Iran, wars in countries such as Syria or Iraq, and attacks on the oil infrastructure in countries such as Saudi Arabia have set a new framework in which the adjustments between supply and demand are becoming increasingly more noticeable in the evolution of oil prices.

Consequently, the configuration of crude oil prices has been a focus of energy economics literature research, producing a generous number of investigations. Weiner (1991) stated that the prices of crude oils with same quality move closely together all the time when sustained by the concept of "globalization" against the hypothesis that states that the oil market is regionalized (see Reboredo, 2011). Therefore, according to Weiner's point of view, the Brent-WTI spread could supposedly be nearly constant over time (Fattouh, 2010). Furthermore, both components of the Brent-WTI price spread are light sweet crude oils, and they are almost identical in physical composition. As such, any substantial deviation in the price between these crude oils can only be a consequence of the spatial price equilibrium and not differences in their intrinsic values (Bennett and Yuan, 2017). However, increasing interest is emerging in the literature to understand the behaviour of the spread between the two indices, and many explanations for this spread are emerging in this literature. This interest is based on risk management, especially after the world's crude oil market began to diverge at the end of 2010 (Ji and Fan, 2015), and derives from the increasing average distance between the different pricing behaviours of this spread that were observed in the period from 2011 to 2013, when the Brent-WTI spread widened to as much as $25 per barrel.

The remainder of this chapter is organized as follows. Section 9.2 reviews the empirical findings concerning the WTI-Brent spread from the long-run perspective. Section 9.3 develops the empirical strategy followed in this paper. Section 9.4 presents the results derived from the econometric application, and section 9.5 presents the main conclusions and policy implications.

9.2 Theoretical and empirical background

This section summarizes the main drivers of the Brent-WTI spread, which are described in subsection 9.2.1. Subsection 9.2.2 shows the empirical approaches that have been used to measure the relationships between crude oil markets, and describes how the literature has defined two possible scenarios to understand the Brent-WTI spread, namely, the "globalized" and "regionalized" market scenarios. Then, subsection 9.2.3 focuses on understanding several methods involving fractional models to explore time series and some previous applications of the FCVAR model. Finally, the new possibilities for FCVAR model application are shown in subsection 9.2.4.

9.2.1 The drivers of the Brent-WTI spread

The body of this literature that has studied the Brent-WTI spread has been devoted to breaking down the causes of the fluctuations in price and analyzing the relationship over the long run. The pioneer's efforts to explain the Brent-WTI spread's drivers appear linked to the transportation cost literature.[1] Although recently, Bennett and Yuan (2017) maintained that markets that are geographically adjacent to each other tend to be more highly integrated than are markets separated by distance, they recognized that institutional barriers such as exchange rates still cause the no-arbitrage condition to fail in close commodity markets. In Bennet and Yuan's work, they propose an empirical analysis that confirms many of the previous findings of Büyüksahin, Lee, Moser, and Robe (2013), where they constructed a tractable theoretical model that allows one to identify the causes of the changes in the Brent-WTI price spread over time, and it is flexible to other commodities markets' price spread patterns. This set of arguments found in the literature may explain the asymmetric adjustment back to the equilibrium position, and the paper of Milonas and Henker (2001) summarized these as temporary demand/supply divergences, seasonal factors, transportation costs, convenience yields and the volatility of the underlying cash commodity.

[1] For a wide explanation concerning transportation costs, see for instance Dumas, Desrosiers, Gelinas, and Solomon (1995) or Sercu, Uppal, and Van Hulle (1995)

Since the Brent-WTI price spread is the underlying crude oil futures market of the New York Mercantile Exchange (NYMEX) and the Intercontinental Exchange (ICE), the good performance of the spread underlying both guarantees to the economic agents of the world the minimization of the risks associated with price fluctuations in their oil investments around the world. Nevertheless, Silverio and Szklo (2012) explain how over time the markets for benchmark oils have become more sophisticated and complex due to two main factors. On the one hand, the political and market conditions at that time that are the result of instantaneous and decentralized assessments of the market conditions by the participants produce uncertainty, volatility and, consequently, greater risks for the participants. On the other hand, the physical markets in which the benchmark crudes are traded depend on a relatively broad base to exist, and the spread of new information about the market in terms of prices has been impaired.

The close links with other crude oil markets result in (financial) risk-aversion where market participants with heterogeneous expectations or some compulsive or noisy trading activities may also cause an asymmetric adjustment process where a price is pushed up (or down), causing the spread to widen (or narrow) until informed traders react to the temporary deviation and push prices back to the equilibrium position (Cootner, 1962). Financial market frictions, futures contracts availability, and institutional and regulatory constraints are also major factors in oil market price mechanisms and may affect the convergence to the equilibrium. In particular, the financial markets determine oil prices in recent years, facilitate the price discovery and offer a means of transferring risk Silverio and Szklo (2012).[2] Consequentially, price discovery is the process of uncovering an asset's full information or permanent value, and the unobservable permanent price reflects the fundamental value of the stock or commodity.

All these factors play roles in the observable price, which can be decomposed into its fundamental value and transitory effects. The latter consists of price movements due to factors such as the bid-ask bounce, temporary order imbalances or inventory adjustments (see Figuerola-Ferretti and Gonzalo, 2010). Furthermore, the benchmarks are economically important because they are traded in the commodity centers, and the spreads between the two benchmarks are also traded and are vital in the price discovery process of crude oil and its derivatives in order to be able to maintain a balanced pricing relationship among the different grades of crude in their categories (Hammoudeh, Ewing, and Thompson, 2008). Since the extent to which the benchmarks were used by economic agents was unreliable, this gap can be deepened. For this reason, it has become a matter of priority for researchers, economic agents and policy makers to know the evolution of the Brent-WTI price spread in order to guarantee the stability of the oil market, as instrumented by the crude oil futures market.

Overall, in the last two decades, a large number of both theoretical and empirical studies have emerged that have provided solid arguments concerning these factors that drive the Brent-WTI spread (see among others Bacon and Tordo, 2004; Lanza, Manera, and Giovannini, 2005; Hammoudeh et al., 2008; Schmidbauer and Rösch, 2012; Büyüksahin et al., 2013; Liao, Lin, and Huang, 2014; Balcilar, Demirer, and Hammoudeh, 2014; Mensi, Hammoudeh, Nguyen, and Yoon, 2014; Giulietti, Iregui, and Otero, 2014; Borenstein and Kellogg, 2014; Deeney, Cummins, Dowling, and Bermingham, 2015; Dowling, Cummins, and Lucey, 2016; Loutia, Mellios, and Andriosopoulos, 2016; Zhang and Yao, 2016; or Bennett and Yuan, 2017). Among the other factors at play in the relationship that drive to stationary crude oil differentials (Giulietti et al., 2014) are the role that the system of OPEC has on the volatility of prices, especially for WTI during low prices (see Schmidbauer and Rösch, 2012; or Mensi et al., 2014; and Loutia et al., 2016). In this line, a strand of literature also tries to explain these factors from a behavioral perspective (see Deeney et al., 2015; or Dowling et al., 2016) or speculation (Balcilar et al., 2014; or Zhang and Yao, 2016). Overall, this body of literature has studied the Brent-WTI spread in a wide context, but unfortunately, the 2010 dramatic change in the spatial price spreads of tradable commodities (e.g., oil) cannot be readily explained with standard models from the economic literature (Bennett and Yuan, 2017).

[2]Price discovery refers to the use of futures prices for pricing cash market transactions (Figuerola-Ferretti and Gonzalo, 2010).

9.2.2 Brent-WTI in the long run; *'regionalized'* or *'globalized'* market

Leaving aside the factors that contribute to determining the spread fluctuations, the long-term relationship has been considered as the main objective in the articles of this literature. This vast empirical evidence has resulted in a large number of investigations that have supported the idea that the adjustment process moves under the 'one great pool', 'integrated markets' or 'globalized market' hypotheses as opposed to the hypothesis of 'regionalized markets'. Therefore, the key factor in many recent articles is devoted to exploring the long-term spread behaviors and detecting the stability of the long-term relationship. Some papers show evidence of greater light crude oil market integration (see for instance Ji and Fan, 2012), even in the 1990s, when there were significant transaction costs between oil markets (Kleit, 2001). Nevertheless, no consensus exists regarding explaining the long-run relationship. In recent years, strong evidence supports that the Brent-WTI crude oil price spreads changed from a stationary time series to a nonstationary time series in 2010 due to two breakpoints in 2008 and 2010, thus motivating a large number of studies that have recently focused on the behaviors of this long-term relationship (see Büyüksahin et al., 2013; Chen, Huang, and Yi, 2015; or Liu, Wang, Wu, and Wu, 2016).

This long-run analysis has been applied using several approaches, such as aiming to develop the Granger causality method for testing the relationship among the spread indices (see Kolodziej and Kaufmann, 2013; Berk, 2016; or, more recently, Coronado et al., 2017). Under a very general vision, the studies measuring this long-term relationship have studied the dynamic of the relationship (see, for instance, Ghoshray and Trifonova, 2014). Additionally, recently, Narayan and Narayan (2007) and Jia, An, Sun, Huang, and Wang (2017) focused on the volatility and time-varying market integration and diversification during different typical stages of the global oil price. Nevertheless, the distribution of articles devoted to testing this long-term relationship can be grouped into two large groups, including the studies analyzing the cointegration of both markets and the studies dedicated to studying the persistence of this relationship, as a way of demonstrating whether it is a globalized or regionalized market (Gülen, 1997, 1999).

From this long-run analytical perspective, one of the first works dedicated to the analysis of cointegration is that of Ardeni (1989), who uses tests of nonstationarity and cointegration for a group of commodities and shows that the law of one price fails in the long-run relationship. He argues that the failure of the law of one price can be rationalized with two factors, namely, the high costs of arbitrage (Richardson, 1978) and institutional barriers, a conclusion that was supported more recently by Goldberg and Verboven (2005), Fattouh (2010) and Olsen, Mjelde, and Bessler (2015). Other authors highlighted the asymmetric adjustment process in the long-run equilibrium (Hammoudeh et al., 2008) or that these markets were not totally integrated (Milonas and Henker, 2001). Furthermore, there was strong evidence of threshold effects in the adjustments to long-run equilibrium, which implies that markets are not necessarily integrated in every time period, which was demonstrated by Milonas and Henker (2001). In this context, the threshold cointegration analysis has also been implemented by Chen, Finney, and Lai (2005), Ewing, Hammoudeh, and Thompson (2006) and Mann and Sephton (2016), where they found strong evidence of asymmetric adjustments. This long-run relationship has also been emphasized by the work Liu et al. (2016) that analyzed the dynamics of the Brent-WTI price spread using a procedure suggested by Bai and Perron (1998, 2003) to test the structural breaks in the spread. They found that the Brent-WTI price spread changed from a stationary time series to a nonstationary time series in December 2010. Additionally, they show how the Brent-WTI price spread responds to different shocks in the physical market, including shocks to the WTI supply, Brent supply, US demand and international demand (see also Scheitrum, Carter, and Revoredo-Giha, 2018). For its part, Ye, Karali, et al. (2016) analyzed the structural breaks in the long-run relationship during the 1993-2016 time period, revealed that the spread is found to experience multiple structural changes during the sample period, and found that the price impact of the breaks that occurred in the later time period were larger. In this structural change analysis, Zavaleta, Walls, and Rusco (2015) also evidenced that a structural break during the financial crisis of

2008 changed the long-run equilibrium price relationships and the short-run price dynamics. However, in this set of empirical studies against the intuition and theory, Azar and Salha (2017) support that the samples chosen in their paper do not contain calendar structural breaks and that the regionalization of the oil market is strongly denied. These results reject the underlying theory and set the stage for a possible financial anomaly.

Considering the analysis of the persistence of the spread, Liao et al. (2014) investigate the spread of the relationship by distinguishing the quantiles with structural breaks of the Brent-WTI price spread crude oil prices as a benchmark, and they find that the spreads contain a unit root in the lower quantiles but display mean reversion behaviors in the upper quantiles. Other papers analyze the comovements of different crude oil prices and markets as well as the deviation of these comovements. In particular, Klein (2017) identifies high but volatile correlations, thus indicating that the long-term movements of the Brent-WTI price spread are driven by the same dynamics. Indeed, he also confirms the 'globalized market' hypothesis and the leading effects of the WTI over Brent using short-term trends of several days, especially in the negative direction.

Finally, the leading role in the crude oil market has also been recently analyzed. The paper of Ji and Fan (2015) also determined that WTI behaved as the price setter before 2010, while Brent has played the leading role in the crude oil market since 2011. For its part, from a different focus using a wavelet-based complex network, Jia et al. (2017) recently determined the multiperiod evolution characteristics of leading oil prices and the key spreading paths that play a special role in driving the global oil price comovement tendency towards globalization and regionalization. In particular, the authors support that there are various typical evolutionary features during the changes in leading oil prices and key spreading paths from the weekly cycle to the long yearly cycle in different volatile stages.

9.2.3 Several methods that deal with fractional models to explore time series

The empirical approaches used in this body of empirical literature are very extensive. This subsection provides a brief review of the applications of fractional models in oil markets and contrast them with empirical applications of the FCVAR model for these types of time series. It is well known that fractional models are suitable for studying long-term memory in commodity and energy prices. Studies have used fractional models, including the ARFIMA model, to evidence the level of integration of the oil market and their long-term equilibrium relationships (see Bachmeier and Griffin, 2006; or Coakley, Dollery, and Kellard, 2011, for instance), the local Whittle estimator to test the relationship between financial and physical oil markets (Ghorbel, Souissi, et al., 2016), and decompositions of the price discovery to determine which crude oil holds the dominant position as a benchmark in the crude oil market (Elder, Miao, and Ramchander, 2014; and Liu, Schultz, and Swieringa, 2015). Predictability and market efficiency have also been analysed through price volatility, revealing different events affecting long-term memory and persistence in the series, by using the local Whittle estimator, the ARFIMA model and the FIAPARCH model (Wang and Wu, 2012, 2013), Chkili, Hammoudeh, and Nguyen (2014), or David, Quintino, Inacio Jr, and Machado (2018), respectively). Finally, in this focus on long-term equilibrium perspectives, the cyclicity, persistence and/or structural breaks of selected commodity price series have been studied in the empirical literature by using different fractional integration approaches, such as the ARFIMA model or Whittle functions (see Gil-Alana and Gupta, 2014); Gil-Alana, Chang, Balcilar, Aye, and Gupta, 2015; or Monge, Gil-Alana, and de Gracia, 2017, for instance).

Regarding the empirical application of the fractional CVAR model, the major empirical studies have been performed in the areas of financial markets and macroeconomics. In particular, this approach has been applied to analyse the relationship between spot and futures markets (Rossi and De Magistris (2013) and to predict the stock prices by connecting high and low prices (Caporin et al., 2013). Similarly, Baruník and Dvořáková (2015) applied the FCVAR model to contrast the volatility in the selected stock markets. Gagnon et al. (2016)

used the FCVAR model with regard to the Brent and WTI relationship, focusing on imperfect integration during the Cushing bottleneck period and the Brent and WTI linked in risk anticipation. Additionally, other recent studies have applied the FCVAR to examine the fractional relationship between political support and macroeconomic variables (Jones et al. (2014); Nielsen and Shibaev (2018)) and to investigate the inflation hedging ability of gold from 1257 to 2016 (Aye, Carcel, Gil-Alana, and Gupta, 2017), exchange rates (Yaya and Gil-Alana, 2018) and the behaviour of high and low prices of four commodities (Gil-Alana and Carcel, 2018). Finally, Dolatabadi, Nielsen, and Xu (2015), Dolatabadi et al. (2016), Dolatabadi et al. (2018) used the FCVAR model to analyse and forecast the commodity market. In this context, the FCVAR model is applied in the current study to provide a previously uncontemplated view in the crude oil market literature. The insights for this topic are developed in the following subsection, where that previously uncontemplated view is detailed, highlighting the added value of this study.

9.2.4 New possibilities by applying the FCVAR

Our new approach uses the FCVAR model developed by Johansen and Nielsen (2012, 2016) and further developed by Nielsen and Popiel (2016). The FCVAR model is an expansion of the traditional cointegrated VAR (CVAR) model proposed by Johansen (1995), and it allows us to determine the number of equilibrium relations via cointegrating rank testing to estimate memory parameters, long-run cointegrating relations with adjustment parameters, and short-run lagged dynamics. In this respect, our purpose is to analyse the dynamics of the crude oil market, i.e., the relationship between Brent and West Texas Intermediate, aiming to determine if these markets are regionalized or globalized by testing the long-run relationship between both crudes and the spread simultaneously. This paper recognizes that the premises of standard cointegration testing ($I(1)/I(0)$ dichotomy) time-series variables, integrated at order one and comoved at order zero, are too restrictive, i.e., linear combinations of $I(1)$ nonstationary processes are $I(0)$ stationary. In this sense, the empirical literature has shown that many economic and financial time series hold long-range dependence in the autocorrelation function but do not precisely exhibit a unit root process, i.e., the long memory process. For this reason and according to our research, we discard traditional cointegration assumptions that crude oil prices cannot move away from one another for long periods of time and that they are unit roots or $I(1)$; they follow dichotomy $I(0)/I(1)$ such that they follow a fractional process $I(d)$. We also shed the notion that the error term follows a stationary process ($I(0)$) in cases of cointegration of both variables. In turn, the rigidity of the traditional approach is overcome in favour of allowing for the series to be cointegrated, and the error term does not necessarily need to be $I(0)$; for example, we allow for the error term to be cointegrated in order $I(d-b)$, unlike other techniques that assume the error term is $I(0)$. Indeed, the study of the long-run relationship and the behaviour of the error term may be analysed jointly, which is one of the main advantages of this methodology. Overall, the FCVAR model allows several previously unconsidered scenarios to be determined (see table 9.3). As we have seen in the previous subsections, the empirical review of the literature on this long-term relationship reveals two fundamental blocks of approaches. On the one hand, there are studies that measure the cointegration of the Brent-WTI Market, and on the other hand, there are studies that measure the persistence of the Brent-WTI price spread. In both cases, the approximations are understood as a globalized or regionalized market, but no study applies both approaches jointly. To try to understand both approaches made in the literature, table 9.1a is presented, which aims to illustrate the empirical approaches. As seen in table 9.1a, it presents separate evidence of the cointegration and spread, and it shows that the traditional approaches can generate two types of controversies in the interpretation of results depending on the applied empirical approach.

In response to that in table 9.1a, our approach seeks to respond to controversy 2 detected in the traditional approaches. This controversy is given because the analysis of the cointegration does not allow the spread to be nonstationary.[3] To this end, the empirical framework could be changed by the application of an FCVAR model, which would allow us to break from these

[3] Additionally, the stationary of the spread must be cointegrated, so controversy 1 in our application is not studied.

TABLE 9.1A: Standard empirical approaches

		Cointegration /Long-run relationship	
		Yes	No
Spread	Stationary	Globalization	*Controversy 1* Globalization (Not Cointegrated) / Regionalization (Spread)
	Nonstationary	***Controversy 2*** Globalization (Cointegrated) / Regionalization (Spread)	Regionalization

restrictions, thus providing a more reliable framework when making decisions on the adopted policies to take control of the spread of oil crude prices.

To the best of our knowledge, derived from the assumptions of the traditional cointegration approach, when the series are cointegrated, the relationship is persistent, and the fractional cointegration could solve this rigidity by allowing for intermediate stages. That is, even when having cointegrated series, any shock could be long-lived, and this possibility has not been studied before in this literature. Therefore, the FCVAR model allows us to identify several degrees of globalization. Once the testing shows that there is cointegration, the degree of integration of the spread permits us to detect up to three different degrees of globalization. This idea is illustrated more fully in the next section and is illustrated in table 9.1b, where the new possibilities allowed by the application of the FCVAR as a generalization of traditional approaches are broken down.

TABLE 9.1B: New possibilities by applying the FCVAR

		Cointegration /Long-run relationship	
		Yes	No
Spread	Stationary	**Several degrees of globalization (see table 9.3)**	Controversy 1 Globalization (Not Cointegrated) / Regionalization (Spread)
	Nonstationary		Regionalization

Whole, this approach allows both the cointegration and the stationarity of the spread to be simultaneously analyzed by studying the order of integration of the error correction term. Consequently, new scenarios could be researched. In this regard, in the following section, by means of the development of the empirical approach of the FCVAR, the new possibilities and the set of new scenarios aforementioned by controversy 2's ideas are more fully detailed and summarized according to several degrees of globalization. Likewise, this approximation has also the advantage of understanding what relative weights each of the indices has with respect to the behaviors of the other by following a similar approach as that proposed by Ji and Fan (2016), which is a factor that is understood as fundamental in the study of the relation. In this context, the empirical evidence presents a fault in the long-run Brent-WTI spread.

9.3 Data and Methodology

9.3.1 Data

For our empirical analysis, we employ a weekly sample of the Brent and WTI crude oil prices over the period from 15^{th} May 1987 to 19^{th} April 2019 (amounting to 1667 observations for each crude oil series). The data correspond to the Brent (B_t) crude oil and West Texas Intermediate (W_t) crude oil measured in (US \$) prices. The data are collected from the US Energy Information Administration (EIA).

As a preview on the selected variables, Figure 9.1 presents a graphical analysis of the time series dynamics plotted for Brent and West Texas Intermediate. This plot shows a similar behavior in both variables, which could confirm our subsequent results.

Figure 9.1: Dynamics of Brent and West Texas Intermediate

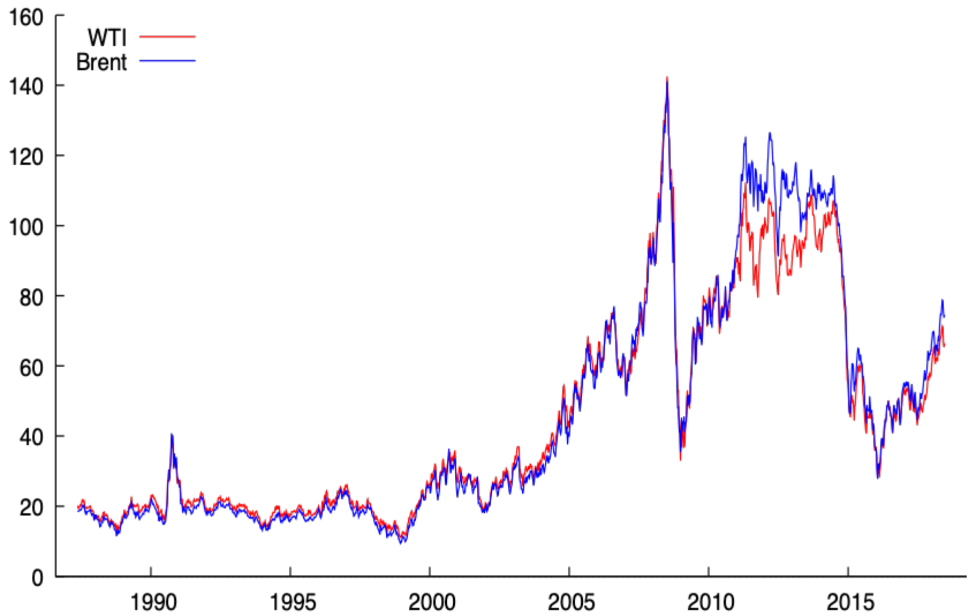


Table 9.2 shows several descriptive statistics for the two crude oil prices. These statistics corroborate the tendencies observed in Figure 9.1. A similar variation in sample mean, standard deviation, asymmetry, and kurtosis is found, which suggests that these commodities can offer investors quite different risk-return trade-offs when considered from an investment portfolio point of view.

Table 9.2: Descriptive statistics for the data

	Mean	Median	Min.	Max.	SD	Asym	Kurtosis
WTI	45.105	31.820	142.52	11	29.455	0.846	-0.441
Brent	46.167	30.530	141.07	9.44	32.973	0.900	-0.429

The data spans from 15^{th} May 1987 to 19^{th} April 2019.

9.3.2 Methodology

Our empirical procedure consists of several steps. First, we apply the fractionally cointegrated vector autoregressive (FCVAR) model proposed by Johansen and Nielsen (2012) in order to contrast the possible existence of the spread's persistence. Then, we study the permanent-transitory decomposition (Gonzalo and Granger, 1995; and Figuerola-Ferretti and Gonzalo, 2010) in order to determine which crude oil drives the common trend. Thus, in the context of cointegration theory, the commonly linear model is as follows:

$$B_t = c + \beta W_t + \varepsilon_t \tag{9.1}$$

According to this expression, B_t are the weekly spot prices of Brent at time t, and W_t represent the weekly spot prices of WTI. Both spot prices should be nonstationary and related through a cointegration relationship with the parameters $(1, -\beta)$. Following the work of Growitsch, Stronzik, and Nepal (2015), the coefficient β_t represents the strength of price globalization. If $\beta_t = 0$, it implies that there is no relation between the markets and that they are completely decoupled. If prices have globalized and markets are perfectly integrated and competitive, β_t should be equal to 1. If this difference is stationary, Brent and West Texas Intermediate are driven by a common stochastic trend and do not allow for arbitrage opportunities because the market forces adjust to correct any temporary disequilibrium.

Moving on to the empirical procedure, the next application of the model is a generalization of Johansen (1995)'s cointegrated vector autoregressive (CVAR) model to allow the fractional processes of order d that cointegrate to order $d-b$. The fractional cointegrated vector autoregressive (FCVAR) model has the power to be used for stationary and nonstationary time series and is settled in Johansen and Nielsen (2012) and Nielsen and Popiel (2016). To introduce the FCVAR model, first, we must refer to the CVAR model. Letting Y_t, $t=1,\ldots,T$ be an $I(1)$ time series, the CVAR model is:

$$\Delta Y_t = \alpha\beta' L Y_t + \sum_{i=1}^{k} \Gamma_i \Delta L^i Y_t + \varepsilon_t \tag{9.2}$$

To derive the FCVAR model, we begin by introducing the fractional difference operator to the CVAR model, Δ, which inserts persistence in the model, and the fractional lag operator is $L = (1-\Delta)$. Replacing the lag operators with their fractional counterparts Δ^b and $L_b = (1-\Delta^b)$, respectively, we obtain

$$\Delta^b Y_t = \alpha\beta' L_b Y_t + \sum_{i=1}^{k} \Gamma_i \Delta^b L_b^i Y_t + \varepsilon_t \tag{9.3}$$

Applying $Y_t = \Delta^{d-b} X_t$, we obtain the following FCVAR model:

$$\Delta^d X_t = \alpha\beta' L_b \Delta^{d-b} X_t + \sum_{i=1}^{k} \Gamma_i \Delta^d L_b^i X_t + \varepsilon_t \tag{9.4}$$

As usual, ε_t is a p-dimensional i.i.d. variable with mean zero and covariance matrix Ω. The parameters α and β are $p\times r$ matrices, where $0 \leq r \leq p$. The columns in matrix β are the cointegrating vectors, and $\beta' X_t$ assumes the existence of a common stochastic trend, which is integrated to order d, and the short-term parts from the long-run equilibrium are integrated to order $d-b$. The speed of the adjustment to the equilibrium coefficients is reflected in α. Thus, $\alpha\beta'$ is the long-run adjustment, and Γ_i represents the short-run dynamics of the variables.

There are two additional parameters in the FCVAR model compared with the CVAR model. The parameter d represents the order of fractional integration of the observable time series. The parameter b determines the degree of fractional cointegration, that is, the reduction in fractional integration order of $\beta' X_t$ compared to X_t itself. The relevant ranges for b are (0, 1/2), in which case the equilibrium errors are fractional of order greater than 1/2 and are therefore non-stationary although mean reverting, and (1/2, 1], in which case the equilibrium errors are fractional of order less than 1/2 and are stationary (Dolatabadi et al., 2016). Note that for $d=b=1$, the FCVAR model is reduced to the CVAR model, which is thus nested in the FCVAR model as a special case. Johansen and Nielsen (2012) show that the maximum likelihood estimators $(b, \alpha, \Gamma_i, \ldots, \Gamma_k)$ are asymptotically normal and that the maximum likelihood estimator of (β, ρ) is asymptotically mixed normal when $b > 1/2$ and asymptotically normal when $b < 1/2$. The important implication is that the standard asymptotic inference can be applied to all these parameters.

We now determine the number of stationary cointegrating relations following the hypotheses of the rank test based on a series of LR tests. In the FCVAR model, we test the hypothesis $H_0: rank(\Pi) = r$ against the alternative $H_1: rank(\Pi) = p$ for $r = 0, 1, \ldots$ "estimated" rank is then the first non-rejected value in the sequence of tests. Being $L(d, b, r)$is the profile likelihood function given a rank r, where (α, β, Γ) have been reduced by rank regression (see Johansen and Nielsen, 2012). The asymptotic distributions of these LR test statistics are non-standard and are derived in Johansen and Nielsen (2012). We use the P values obtained from computer programs made available by MacKinnon and Nielsen (2014) based on their numerical distribution. Maximizing the profile likelihood distribution under both hypotheses, the LR test statistics are now $LR_t(q)$. The asymptotic distribution of $LR_t(q)$ depends on the parameter b and on $q = n-r$. MacKinnon and Nielsen (2014) based on their numerical distribution functions, provide asymptotic critical values of the LR rank test. In the case of "weak cointegration", i.e., $0 < b < 1/2$, $LR_t(q)$ has a standard asymptotic distribution $LR_t(q) \xrightarrow{D} \chi^2(q^2)$.

The specification in (9.4) is the so-called restricted constant version of the model by Johansen and Nielsen (2012), which is also used by Dolatabadi et al. (2016). Deterministic trends may be assumed in the FCVAR model in several ways. Johansen and Nielsen (2012) considered the insertion of the restricted constant term ρ in the long-run cointegrating relation. Dolatabadi et al. (2016) suggested an unrestricted constant ξ as the linear trend of the fractionally integrated processes. The following specification shows a more general form:

$$\Delta^d X_t = \alpha L_b \Delta^{d-b}(\beta' X_t + \rho') + \sum_{i=1}^{k} \Gamma_i \Delta^d L_b^i X_t + \xi + \varepsilon_t \tag{9.5}$$

where ρ is denoted as the restricted constant term, i.e., the mean level of equilibrium relation, and ξ is the unrestricted constant term that generates a deterministic trend in the levels of the variables (Dolatabadi et al., 2016).

Therefore, the FCVAR model allows simultaneous modelling of the long-run equilibria, the adjustment reactions to deviations from the equilibria and the short-run dynamics of the system. Johansen and Nielsen (2012) and Nielsen and Popiel (2016) provide estimation and inference explanations for the model, and the latter study specifies MATLAB computer programs for the calculation of estimators and test statistics.

When the VAR model encounters the case of $d = b = 1$ (CVAR), the error correction term is integrated of order $(d - b)$, which is $I(0)$ in this case. However, in the fractional cointegration, these axioms are relaxed because $(d - b) = 0$, which means that the error correction term shows short-run stationary behaviour, or $(d - b) > 0$, which in turn means that there is a long memory process and that the error correction term will revert in the long run.

Connecting with the previous idea explained in subsection 9.2.3 and according to table 1 in Tkacz (2001), when $(d - b) = 0$, the error correction term follows a stationary process, and the shock duration is short-lived. If $0 < (d - b) < 0.5$, there is a stationary process, and the shock duration is long-lived. Then, if $0.5 < (d - b) < 1$, the error correction term follows a nonstationary process, although it is mean-reverting, and the shock duration is long-lived. Finally, when $(d - b) = 1$, the error correction term follows a unit root process. These new scenarios derived from the degree of integration of the spread are represented in table 9.3.

TABLE 9.3: Degree of globalization of the Brent-WTI price differential by applying the FCVAR

Order of integration of the error correction term (ECT)	**Long-run relationship (Value of β)** $\beta = 1$	$0 < \beta < 1$
$I(d - b) = I(0)$	**Strong globalization** and the shock duration is **short-lived.**	**Weak globalization** and the shock duration is **short-lived.**
$I(0) < I(d - b) < I(0.5)$	**Strong globalization** and the shock duration is **long-lived.**	**Weak globalization** and the shock duration is **long-lived.**
$I(0.5) < I(d - b) < I(1)$	**Strong globalization** and follows a **non-stationary** process, although mean-reverting and the shock duration is **long-lived.**	**Weak globalization** and follows a **non-stationary** process, although mean-reverting and the shock duration is **long-lived.**

The first row corresponds to the two traditional cases of the standard approaches. If $\beta = 1$, the error correction term could be interpreted as the *Brent-WTI* price spread/differential.

Maximizing the profile likelihood distribution under both hypotheses, the LR test statistic is now $LR_t(q)$. The asymptotic distribution of $LR_t(q)$ depends on the parameter b and on $q = n - r$. MacKinnon and Nielsen (2014) was based on these numerical distribution functions and provided asymptotic critical values of the LR rank test. According to the existence literature, cointegration implies a fractionally vector error correction model (FVECM) such as the following:

$$\begin{pmatrix} \Delta B_t \\ \Delta W_t \end{pmatrix} = \begin{pmatrix} \alpha_B \\ \alpha_W \end{pmatrix} (B_{t-1} - \beta W_{t-1} - c) + \sum_{i=1}^{n} \Gamma_i \begin{pmatrix} \Delta B_{t-i} \\ \Delta W_{t-i} \end{pmatrix} \begin{pmatrix} u_{1t} \\ u_{2t} \end{pmatrix} \tag{9.6}$$

This model includes the adjustment parameters α, the cointegration coefficient β, the restricted constant (c), the lag length (n) and the errors u. Γ_i are 2×2 parameter matrices in the short-run dynamics. The adjustment coefficients α_B and α_W capture the speed of the adjustment of the Brent (B_t) and West Texas Intermediate (W_t) towards the equilibrium.

Permanent-Transitory (PT) decomposition in the FCVAR model

According to the Gonzalo and Granger (1995) and Figuerola-Ferretti and Gonzalo (2010) P-T decomposition, we let $X_t = (B_t, W_t)'$, where B_t and W_t denote Brent crude oil and West Texas Intermediate crude oil, respectively. In the P-T decomposition X_t can be decomposed into a transitory (stationary) part, $\beta' X_t$, and a permanent part, $Z_t = \alpha'_{\perp} X_t$ where $\alpha'_{\perp}\alpha = \alpha'\alpha_{\perp} = 0$. Z_t is the common permanent component of X_t and it is interpreted as the dominant crude oil, where the information that does not affect Z_t will not have a permanent effect on X_t. To know which parameter contributes to each market (Brent or WTI), we attend to the key parameter $\alpha_{\perp}$. Following the mirror hypothesis, the linear hypothesis on $\alpha_{\perp}$ can also be tested directly on $\alpha_{\perp}$ or alternatively on α using the values of LR tests in each hypothesis, and critical values can be taken from the χ^2 distribution for testing. For example, to test the hypothesis that the dominant crude oil is the Brent, which means that $\alpha_{\perp} = (0, a)'$, we can equivalently test the mirror hypothesis $H_0 : \alpha = (\gamma, 0)'$. Similarly, to test the hypothesis that the dominant crude oil is WTI, i.e. $\alpha_{\perp} = (a, 0)'$, we test the mirror hypothesis $H_1 : \alpha = (0, \gamma)'$ (see Dolatabadi et al. (2018) that first combined the FCVAR with P-T decomposition in the commodity market)

An interpretation of the coefficient α is that it is an adjustment coefficient that measures how disequilibrium errors could be affected by the current changes in X_t. Under this interpretation, we wonder whether any coefficients in α are zero, which means that the variable in question is weakly exogenous. For example, under hypothesis H_1, the parameter $\alpha = 0$ such that the WTI does not react to the disequilibrium error and it is the transitory component, thus implying that the WTI is the main contributor to the common trend.

To determine the magnitude of each variable in the long run, we use the component share (CS). As Baillie et al. (2002) notes, since $\alpha'\alpha_{\perp} = 0$, it may also be expressed in terms of the elements of the error correction vector α. To interpret this, we let $\alpha = (\alpha_b, \alpha_w)'$ and $\alpha \perp \alpha = (\alpha_{\perp,b}, \alpha_{\perp,w})'$. Afterwards, $\alpha'_{\perp}\alpha = \alpha_{\perp,b}\alpha_b + \alpha_{\perp,w}\alpha_w = 0$ implies that $\alpha_{\perp,b} = -\alpha_{\perp,w}\alpha_w/\alpha_b$. Therefore, the component share (CS) may be expressed as

$$CS_B = \frac{\alpha_w}{\alpha_w - \alpha_b}, CS_W = \frac{\alpha_b}{\alpha_w - \alpha_b} \tag{9.7}$$

where B and b and W and w corresponds to Brent and WTI crude oil, respectively.

9.3.3 Model specification

According to Dolatabadi et al. (2016), we have followed the model specification proposed by them. In this respect, before estimating the FCVAR model and the hypotheses of interest, there are three additional elements in the specification of the FCVAR model: the lag length (k), the deterministic components, and the cointegration rank (r).

First, regarding the selection of the lag length, we meticulously apply some sources of information, including the Bayesian information criterion (BIC), the LR test statistics for significance of Γ_k, and the tests for serial correlation in the residuals. In each case, that are based on the model that includes all the deterministic components considered and has full rank $r = p$. Indeed, for our series, we first use the BIC as a starting point for the lag length, and from there we find the nearest lag length that satisfies the criteria. Second, we check whether Γ_k is significant based on the LR test. Third, we check that the tests for serial correlation in the residuals do not show signs of misspecification.

After selecting the lag length, we need to select the deterministic components and the cointegrating rank (r). For the former, this work considers that the restricted constant, $\rho\pi_t$, is present following the methodology framework. Otherwise, the selection of deterministic components is concentrated into the absence or presence of the unrestricted constant[4], that is, the trend component. As Dolatabadi et al. (2016) state, because the limit distribution of the cointegration rank test depends on the actual cointegration rank and the presence or absence of the trend, we must simultaneously decide the cointegration rank and whether the trend is included. The testing of both hypotheses jointly is discussed in depth in Johansen (1995).

9.4 Results

This section shows the results of applying a FCVAR model to simultaneously assess the long run and the persistence of the relationship between the Brent-WTI price spread. That is, the model allows us to discriminate if the markets are globalized or regionalized, which would allow us to study a scenario not contemplated until now. The application of the FCVAR model, which is a new procedure to accomplish this goal, is summarized in table 9.3. We start our econometric exercise by studying the possibility that the fractional cointegration would be more appropriate than the standard one. Once this step is done, we test the degree of the Brent-WTI spread persistence. Then, under this estimation, we examine if the error term shows a long memory process in step 2 and 3. Finally, in step 4, using the FVECM and, subsequently, the P-T decomposition, we assess which of our variables has a permanent behavior in the common trend, which allows us to know which variable is the price setter.

TABLE 9.4: Strategy of empirical research

	Procedure	**Hypotheses**
Step 1	Standard Cointegration vs. Fractional Cointegration	H_1^d: Is the fractional cointegration more appropriate than traditional cointegration?
Step 2	*Cointegrating vector (1, -1)*	H_1^β: Strong or weak globalization?
Step 3	Degree of Brent-WTI spread persistence, i.e., order of integration *(d – b)*	H_1^{d-b}: How long is the duration of the shock (short-lived or long-lived)?
Step 4	Permanent-Transitory decomposition	$H_1^\beta \cap H_{1\perp}^{\alpha_{B/W}} \equiv H_1^\beta \perp H_1^{\alpha_{W/B}}$ (mirror): What drives the Brent-WTI price structure?

Before testing the possible long-run relationship in the Brent-WTI crude oil price spread and aiming to decide if the FCVAR model is suitable to the main purpose, each of the series is examined singly before driving the multivariate analysis. Broadly, if both stationary tests and unit root tests of a time series are rejected, which suggests that the time series is likely a fractional time series, despite the fact that there are considerable procedures for estimating the fractional differencing parameter in a semiparametric context. Though the semiparametric log-periodogram regression recommended by Geweke and Porter-Hudak (1983) is the most used, this method was varied and deeper developed by Robinson (1995) and has been analyzed by Velasco (1999) and Shimotsu and Phillips (2002), among others. Then, the estimation of the fractional parameter d is determined for each univariate series, with the results presented

[4]We have specified the model with and without the presence of the unrestricted constant. The achieved results are practically identical regardless of the presence or absence of the unrestricted constant. The results are available upon request.

in table 9.5. The first three columns correspond to the semiparametric log-periodogram regression estimates from Geweke and Porter-Hudak (1983), which are labeled here as GPH[5] and are computed with the bandwidths $m = T^{0.4}$, $m = T^{0.5}$, and $m = T^{0.6}$, respectively. The remaining columns in table 9.5 present the FAR (k) estimates with $r = 0$ and the k lags, such as in Johansen and Nielsen (2010). The results are shown for $k = 0$, $k = 1$ and $k = 2$, and the associated Ljung-Box Q-test statistics, which are labeled as $Q_{\hat{\varepsilon}}$, for the serial correlation up to a lag of 12 in the residuals are also given. In this sense, by conducting the univariate analysis, the GPH estimates support the idea that the fractional cointegration could be appropriate for this issue. The FAR (k) models show that the residuals are well behaved and that the estimates of d are in line with or similar to those for the GPH estimates, although their standard errors are lower.

Table 9.5: Univariate analysis

	GPH estimates			FAR(k) estimates					
	$m = T^{0.4}$	$m = T^{0.5}$	$m = T^{0.6}$	$k = 0$		$k = 1$		$k = 2$	
	$\hat{d}$	$\hat{d}$	$\hat{d}$	$\hat{d}$	$Q_{\hat{\varepsilon}}$	$\hat{d}$	$Q_{\hat{\varepsilon}}$	$\hat{d}$	$Q_{\hat{\varepsilon}}$
WTI	0.699	0.763	1.062	1.086	25.464	1.009	22.839	1.119	14.580
	(0.204)	(0.125)	(0.079)	(0.021)	(0.013)	(0.037)	(0.029)	(0.035)	(0.265)
Brent	0.847	0.814	1.043	1.139	17.092	0.652	14.631	0.538	13.892
	(0.160)	(0.094)	(0.069)	(0.022)	(0.146)	(0.035)	(0.262)	(0.101)	(0.308)

GPH denotes the Geweke-Porter-Hudak semiparametric log-periodogram regression estimator, and FAR(k) denotes the fractional AR model with $r = 0$ and k lags. $Q_{\hat{\varepsilon}}$ denotes the Ljung-Box Q-test statistic for the residuals, computed with 12 lags because monthly data is used. Standard errors are given in parentheses beneath estimates of d and P values are in parentheses beneath $Q_{\hat{\varepsilon}}$ tests. The sample size is T = 1667.

In this section, we have shown the procedure that we will perform for a battery of results shown below. Following the model specification proposed by Dolatabadi et al. (2016), we follow a path to determine the optimal specification of our model and we chose one lag (see Appendix). In this respect, although DeJong, Nankervis, Savin, and Whiteman (1992) have shown that both the BIC and AIC criteria and their estimates notably differ, we must consider that too long of a lag length would distort the data and lead to a decrease in the estimation power. Additionally, MacKinnon and Nielsen (2014) reveal that one lag is sufficient to whiten the residuals in the FCVAR model. Then, once the lag length is selected, we determine if there is a long-run relationship between the variables that are chosen. For this reason, we test the cointegration rank before testing the hypothesis of the fractional parameter and evidence that the number of cointegrating vectors is one in our case (table A.2 in the 9.6). Once the rank cointegration test is established, we test the hypothesis H_1^d, which tests whether the fractional cointegration is more appropriate than standard cointegration is. Table 9.6 shows that, in our case, we reject the null hypothesis of $d = 1$, and its rejection implies that the FCVAR model is more suitable than a traditional cointegration therefore, fractional cointegration is appropriate for this study. The next issue consists of estimating the long-run relationship between Brent and WTI (see equation 9.1). As it can be observed, the parameter β is close to 1, which will be crucial for our purpose. For this reason, we test the hypothesis H_1^β, and with a P value of 0.207, it supports the existence of a long-run relationship, thus implying that both crude oils are strongly globalized. In summation, step 1 and step 2 under our empirical proposal reveal that fractional cointegration is more appropriate than standard cointegration, while also showing that Brent-WTI are strongly globalized.

[5]For testing the presence of unit roots, the estimates were obtained using first-differenced data because the original series may be above 0.5. This test expects that the results are limited to the interval $-0.5 < d < 0.5$, and then, 1 is added to obtain the appropriate estimates of d.

TABLE 9.6: Fractional cointegration test and results

Hypothesis tests:			$H_1^d : d = b = 1$		$H_1^\beta : \beta = (1, -1)$
		LR	2.565		2.280
		P value	0.109		0.131
Cointegration vector:			$\beta = (1, -1.098)$		
	$\hat{d} =$ 1.034 (0.049)	$\hat{b} =$ 0.796 (0.097)	$Q_{\hat{\varepsilon}}(10) = 0.466$	$Q_{\hat{\varepsilon}}(10) = 0.277$	Log(Ł) $= -5827.33$)
Restricted cointegration vector:			$\beta = (1, -1)$		
	$\hat{d} =$ 1.060 (0.061)	$\hat{b} =$ 0.649 (0.092)	$Q_{\hat{\varepsilon}}(10) = 0.481$	$Q_{\hat{\varepsilon}}(10) = 0.316$	Log(Ł) $= -5828.97$)

Standard errors are in parenthesis below the values of $\hat{d}$ and $\hat{b}$. The sample size is 1667.

Following our application, in order to complete the third step, by assuming that the cointegrating vector is (1, -1), we can interpret the difference $(d - b)$ as the order of integration of the Brent–WTI price spread, which is the degree of persistence (H_1^{d-b}).[6] In this case, this hypothesis receives a value of 0.411 $(\hat{d} - \hat{b})$, thus implying that the Brent–WTI price spread follows a long memory process, which suggests potential forecasting power at longer horizons (Baillie and Bollerslev, 1994a). This implication may be important to the design of investment or hedging strategies in the futures market. Therefore, this value also implies that the duration of the shock is long-lived. Taking into account that shocks are long-lived, one difference between globalized markets and regionalized markets is that there are more players involved in a globalized market, so it is natural for a globalized market to take longer to respond to shocks than regionalized markets. Moreover, regulations could indirectly affect the transmission mechanism in both markets and also it is a possible clue about how a shock is spread and how it affects the concerned parties, including producers, suppliers and investors.

Brent crude oil is the benchmark oil for the European market and for 65% of global crude oil types that are referenced, whose prices are expressed as a premium or a discount against Brent. Several factors differentiate Brent from other crude oils, including its representative quality standard, which facilitates the appraisal process of other grades; the proximity of the North Sea to an important region of oil consumption and the main refining centres of Europe and the USA; and stable and favourable fiscal regulation (from the perspective of the producers), a solid legal regime and relatively low political risk in the United Kingdom, whose government supervises the benchmark index. The status of Brent crude has also been driven by the diverse ownership of production. Diverse ownership greatly reduces the likelihood of market interference and price manipulation compared to a monopolistic structure. This feature has greatly facilitated the willingness of market participants to adopt Brent as a point of reference. From a geopolitical point of view, due to the 'Arab spring' and Libyan crisis, which have decreased the supply of light, sweet crude in the European region, the prices of both crude oils began to mirror each other (see Difiglio, 2014; or Baumeister and Kilian, 2016), although Brent kept its hefty premium. Finally, another reason emerging to explain why Brent is driving the oil price formation is the supply glut at the main storage facility of WTI in Oklahoma; the premium/discount situation has flipped and now Brent is more expensive than WTI.

At last, table 9.7 shows the FVECM and, subsequently, the Permanent–Transitory decomposition. Regarding the price adjustments to the short-run disequilibrium, attending to the joint hypothesis, we find that the Brent index is weakly exogenous (P value 0.728), which further corroborates the evidence that this market is the driver in global oil markets. To check this premise, we apply P-T decomposition, which also suggests that the Brent crude oil is dominant in the common trend and drives the Brent-WTI price structure. This finding is partially confirmed by the estimates of the component shares of the Brent and WTI prices, which show that the Brent price constitutes almost 90% of the price structure. For example,

[6] According to our methodology, d and b represent the fractional order of integration of the explanatory variables and the cointegrating error, respectively.

as Brent is the dominant crude oil in the common trend and drives the price structure between Brent and WTI, the stakeholders must consider that when a shock occurs, it primarily and directly affects Brent crude oil instead of WTI; that is, any change would affect Brent itself and the Brent-WTI relationship. Otherwise, if the shock occurs over WTI, the shock would only affect WTI and Brent would remain inherent; that is, Brent could be assumed a leading indicator ahead of WTI, i.e., what happens to Brent will also happen to WTI. In this regard, the stakeholders must design strategies in order to take advantage of this situation. Additionally, diverse events (such as institutional forces, market demand and supply, price volatilities or exogenous shocks) are absorbed differently into crude oil markets.

TABLE 9.7: FVECM results under constrained parameters (1, -1)

Hypothesis tests:		$H_1^{\beta} \cap H_1^{\alpha Brent} \equiv H_1^{\beta} \cap H_1^{\alpha \perp WTI}$	$H_1^{\beta} \cap H_1^{\alpha WTI} \equiv H_1^{\beta} \cap H_1^{\alpha \perp Brent}$
	LR	2.565	2.280
	P value	0.109	0.131
Speed of adjustment:			
		$\alpha_{Brent} = -0.017$	$\alpha_{WTI} = 0.144$
Component share:			
		$CS_{Brent} = 0.894$	$CS_{WTI} = 0.106$

With respect to the hypothesis, we reference the mirror hypothesis. The sample size is 1667. CS_{Brent} and CS_{WTI} denote the component shares of Brent and WTI, respectively, and both are normalized such that the two elements add to one.

Finally, next table 9.8 summarizes the set of results showed by our empirical application. This table shows the main information derived of the application of the FCVAR model and the P-T decomposition to the Brent-WTI price spread.

TABLE 9.8: Summary of results

Steps	**Hypotheses**
Step 1	H_1^d: The fractional cointegration is more appropriate than traditional cointegration
Step 2	H_1^{β}: The Brent-WTI market is strongly globalized
Step 3	H_1^{d-b}: The Brent–WTI price spread follows a long memory process (long-lived shocks)
Step 4	$H_1^{\beta} \cap H_{1\perp}^{\alpha B/W} \equiv H_1^{\beta} \perp H_1^{\alpha W/B}$ (mirror): Brent drives the Brent-WTI price structure

9.5 Conclusion

In this paper, we have studied the possible relationship between two of the main indicators of the oil market, the Brent and WTI crude oil prices, using the FCVAR model. Despite the controversy in the existing literature concerning the treatment of this topic, the fractional cointegration model voids most of the problems raised in this literature. In particular, in this article, we propose to measure the cointegration and the stationary simultaneously, which would allow us to study new scenarios in which both prices could be cointegrated but the spread could be nonstationary. Additionally, this model allows us to identify other points of interest, such as the price structure and the persistence of the spread between each one.

The application of the FCVAR allows us analyze the order of integration of the error correction term, thereby revealing several important results. Firstly, important considerations can be taken into account in relation to the evidence held so far since although the series do have a long-term relationship, the spread shows a long memory process, and consequently, the shocks are long-lived. This result is novel in the literature, since until now, the globalization or regionalization of markets has been defined from these perspectives individually. In addition, the results confirm that Brent drives the price structure. Although the FCVAR shows that these markets are strongly globalized, attending to the stationary of the spread, we reveal that this spread shows a long memory process.

These results support several implications for business operators, arbitrageurs, economic agents and policy makers. First, a globalized market determines the price configuration of the Brent and WTI oil markets. This concept assumes that oil markets have linked prices moving closely together. However, as we reveal that the spread is a long memory process, this scenario is not acceptable, given that price adjustments will not be immediate. Our finding indicates that arbitrage opportunities are increased, which implies that the oil market and the energy futures markets, due to their high liquidity, have increased the ability of market agents to arbitrage immediately. The results also have implications for other stakeholders. On one hand, business operators in their hedging strategies need to take into account the persistence of the spread when considering the adjustment period for investment provisions. On the other hand, government policies will have a long-lived effect, and the effect will not be immediate. A high degree of persistence is likely to send erroneous signals to the monetary policy authority, which could feel the need to affect interest rates to mitigate the impact of oil prices on the economy, thinking that the effect of oil prices will last longer than in fact they do (Gil-Alana and Gupta, 2014).

Finally, focusing on the driver in global oil markets also yields interesting results. The central banks closely monitor the oil price for CPI calculations and global growth projections (Mann and Sephton, 2016). In this context, the decision between using Brent or WTI prices is not insignificant. In this paper, we demonstrate that the Brent is the benchmark price, so banks should incorporate it into their forecasts. Indeed, the price of oil is believed to be a leading indicator of growth and inflation in the economy (Stock and Watson, 2003). Additionally, if policy makers seek to guarantee the symmetry of information in these markets in order to evade the risky markets, it would be advisable to avoid the intervention via taxes, imports or exports, the environment, or changing the quantity of production.

9.6 Appendix

Table A.1: Lag length selection

Lags	AIC	BIC
1	11670.02	**11745.77**
2	11676.83	11774.23
3	11668.05	11787.09
4	11652.83	11793.51
5	11638.82	11801.14
6	11650.54	11834.52

Bold indicates lag order selected

Table A.2: Cointegration Rank test

Rank test	Log-likelihood	LR statistics
0	-5835.768	29.519
1	-5827.833	**13.650**
2	-5821.008	—

In bold the number of cointegration relations

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