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Long-run macroeconomic dynamics: applications of fractional integration and cointegration

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DOCTORAL THESIS

Long-Run Macroeconomic Dynamics: Applications of Fractional Integration and Cointegration

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"Live as if you were to die tomorrow. Learn as if you were to live forever."

– Mahatma Gandhi

"Somewhere, something incredible is waiting to be known."

– Carl Sagan

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Preface

I would like to begin by expressing my heartfelt gratitude to my supervisors, Emilio Congregado and Luis A. Gil-Alana, for believing in me from the beginning, teaching me so much, and serving as a source of inspiration. Emilio, your leadership and dedication to this group, always considering each and every one of us, along with your constant support, insightful advice, and your constructive critiques, have been fundamental. Luis, your constant support and availability, no matter the time or place, have been invaluable. From the bottom of my heart, I deeply appreciate the guidance and kindness you have both shown me throughout this journey, as well as the exceptional example you set, both professionally and personally.

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Mª Inmaculada Vílchez Trujillo

Part I: Introduction

Chapter 1: Introduction

1.1. Introduction

The phenomenon of long memory in economic time series has garnered increasing attention over the past few decades. Understanding such processes is critical for accurately forecasting economic dynamics, informing policy decisions, and exploring various subfields, including energy, finance, entrepreneurship, labor economics, and unemployment.

Understanding long memory processes has significant implications for economic policymaking. Economic shocks often exhibit persistent effects that influence the formulation of monetary and fiscal policies. Utilizing long memory models in forecasting enhances predictions during periods of economic turbulence, which is vital for strategic planning and resource allocation. Policymakers can benefit from integrating long memory modeling into their economic forecasting practices, ultimately promoting better intervention strategies during economic crises.

The property of hysteresis in economic systems refers to how certain transitory shocks can lead to lasting effects on key variables, both real and nominal. This concept is crucial in time series analysis, as economists seek to understand not only the immediate responses to a shock but also how these responses can influence the future trajectory of the system. For example, an economic crisis can result in diminished productive capacity and increased unemployment, with recession effects persisting even after economic conditions have improved. This underscores the need to consider the long-term properties of time series when assessing economic resilience and the effectiveness of recovery policies.

Transitory shocks, such as temporary increases in energy prices, may initially impact production and employment. However, if firms do not adequately adjust to these fluctuations, or if workers are laid off and struggle to find new employment, the repercussions may become permanent. This phenomenon resembles the behavior of materials exhibiting hysteresis, where deformation does not fully recover after the load is removed. In economic terms, when disruptions to employment lead to permanent changes in workforce skills, a similar resilience phenomenon is observed. Therefore, understanding how and why some shocks can alter the long-term equilibrium of the labor market is crucial.

The implications of hysteresis on economic dynamics emphasize the necessity for economic policies that address not only immediate crises but also long-term consequences. For instance, investing in entrepreneurship promotion programs among the unemployed can help mitigate the permanent effects of a recession by enabling individuals to secure employment. Accordingly, the study of time series properties and hysteresis becomes a valuable tool for designing strategies that foster a more resilient economy, minimizing shock impacts and encouraging sustainable recovery.

In general, macro-level analyses tend to prevail over the use of microdata in discussions related to resilience or survival properties of phenomena or states. With few exceptions1, the historical observation of individual units (workers, households, or firms) often lacks a sufficiently

[1] For instance, works exploring long-term evidence on the causal effects of programs and shocks by using propensity score matching techniques.

prolonged follow-up over time to assess the long-term effects of specific treatments or shocks. Thus, long-term evaluations are typically approached through analyzing the long-memory properties of time series or panel data, where observation units are generally aggregated units such as countries, regions, or sectors.

Traditional econometric approaches for identifying hysteresis in economic time series have typically relied on unit root tests, examining the persistence of shocks -whether transitory shocks become permanent. These methods implicitly assume that time series represents a single process, interpreting a high degree of persistence in deviations from a mean (a unit root) as evidence of hysteresis, or the long-lasting impact of shocks on the natural rate. 2 The prevailing literature often employs unit root tests for assessing the presence of hysteresis, including adaptations for panel data or those that account for non-linearity (see Mitchell, 1993; Røed, 1996; 1997; León-Ledesma and McAdam, 2004; Lee and Chang, 2008; Romero-Ávila and Usabiaga, 2008; Akdoğan, 2017; Meng, Strazicich, and Lee, 2017; Khraief, 2020; Clavijo-Cortes, 2023; Congregado et al., 2025; and Furlanetto et al., 2025, among others). Another widely applied approach in studying long memory properties in labor economics is the use of fractional integration as an alternative to traditional unit root tests, to infer the degree of persistence (Caporale and Gil-Alana, 2003, 2007, 2008; 2018; Cuestas and Gil-Alana, 2024; Gil-Alana and Brian Henry, 2003; or Gil-Alana, González-Blanch, and Poza, 2024, among others). Specifically, this approach employs the number of differences required to render the series stationary as a determinant of the series' degree of dependency; the greater the number, the longer the duration of the shocks. 3

The combination of fractional integration and cointegration are the techniques that will be employed throughout the five papers comprising this thesis.

1.2. Data and methods

This thesis employs aggregate time series and time series analysis approaches. It includes five papers addressing three different themes (financial, commodity price, and entrepreneurship) with the common denominator of using time series and the aggregated nature of the labor series utilized. In this latter case, many of the series used are original and generated through sample identification and the subsequent application of population weights, along with the aggregation of

[2] Jaeger and Parkinson (1994) argue that this approach suffers from several limitations. To circumvent these weaknesses, they propose isolating the natural and cyclical components using some type of unobserved components model estimated via the Kalman filter. This approach explicitly decomposes the time series into two unobserved components: a stationary cyclical component and a non-stationary natural rate component. The key innovation lies in modeling the dynamic relationship between these components. This two-component model offers a more nuanced perspective than univariate analysis, as it can quantify the impact of cyclical shocks on the natural rate, providing a precise measure of hysteresis. Essentially, this approach not only delivers a binary test for the presence of hysteresis but also provides a quantitative measure of its magnitude, enabling a more detailed and informative analysis of the persistence of any shock affecting the time series of interest. This approach has been used to revisit the issue of hysteresis in unemployment (Jaeger and Parkinson, 1994), extended to a non-linear framework by di Sanso and Pérez Alonso (2010).

[3] More recently, Yaya et al. (2021) and Furuoka et al. (2024) developed a new unit root test, in which they combine fractional integration and autoregressive neural networks to revisit hysteresis in unemployment.

values and the generation of series by applying this procedure to the various available waves. Only for the self-employment series from the UK was this task unnecessary.

Methodologically, and in alignment with the objectives of each study, the thesis utilizes different strategies for analyzing the long-memory properties of the series under investigation, along with cointegration and causality analyses conducted in two of the chapters. We utilize fractional integration and cointegration methods, acknowledging the long memory properties often found in macroeconomic time series. These approaches provide a more accurate representation of the data's persistence and long-run relationships compared to traditional methods. It specifically employs fractional integration, causality analysis and a fractionally cointegrated vector autoregressive (FCVAR) model.

1.3. Contributions

This dissertation makes several key contributions to the study of self-employment and entrepreneurship. It introduces a theoretical framework that clearly distinguishes between necessity and opportunity entrepreneurs, enriching the understanding of their respective dynamics and behaviors throughout the economic cycle. It examines the causal relationships between self-employment, unemployment, venture capital, and other economic factors, providing valuable insights for designing effective policies. It develops a dynamic entrepreneurship model that integrates labor market conditions and capital factors, allowing for a better understanding of how these elements influence self-employment.

It highlights the consequences of self-employment dynamics for economic policy formulation, suggesting that promoting self-employment among the unemployed is not always an effective strategy for reducing unemployment.

Our studies examine the long-term dynamics of self-employment across sectors and regions, providing insights into how these dynamics differ between distinct groups of self-employed individuals.

It employs fractional integration models to analyze the persistence or hysteresis in self-employment, addressing potential asymmetries among different types of self-employed individuals and across various sectors and regions.

The research investigates whether shocks to self-employment result in permanent or transitory impacts, contributing to the understanding of how external factors influence self-employment dynamics.

The outcomes underscore the importance of designing effective public policies that consider the diverse experiences of necessity versus opportunity entrepreneurs, with a focus on enhancing labor resilience and sustainable conditions.

This dissertation also contributes to the analysis of the long-run relationship between loans to non-financial corporations (NFCs), real gross domestic product (RGDP), real gross fixed capital formation (RGFCF), the cost of borrowing differential between long- and short-term rates (DIFF), and a proxy for the cost of debt, securities, and equity issuance (DEBT) in four Eurozone countries (Germany, France, Italy, and Spain) using fractional integration and cointegration methods.

The analysis reveals heterogeneous short-run and long-run dynamics across the four countries. The highest degrees of integration are found in the loan-to-NFC series. The study finds evidence of a single fractional cointegration vector in each country, indicating a long-run equilibrium relationship amongst the variables. The speed of adjustment to this equilibrium, however, differs across countries.

Finally, the last contribution of this dissertation is identifying a significant shift in the equilibrium model linking spot and future oil prices due to a global exogenous shock. This shift suggests the existence of two different, time-dependent equilibria: a backwardation equilibrium before the break and a near-unity contango after the break. This evidence demonstrates how global shocks can affect the spot-futures oil price relationship, inducing equilibrium switches.

1.4. Chapters overview

This thesis consists of five self-contained essays structured as follows.

Chapter 1 revisits the hypothesis of persistence in entrepreneurship at the macro-level, exploring the potential asymmetries across sectors. The persistence is explored by employing traditional unit roots tests and fractional integration techniques to infer the degree of persistence in UK self-employment rates, by industry, using quarterly data from 1997:01 to 2022:04. The results indicate an order of integration substantially above 1 in the case of total self-employment, but mean reversion is found in a number of sectors, including those related with agriculture, wholesale and retail, accommodation and food services, financial and insurance activities, real state, public administration and defence, human health and other services. On the other hand, construction displays a value above 1, and the unit root hypothesis cannot be rejected in another six sectors. The results demonstrate the importance of advancing knowledge on the long-term effects of shocks in entrepreneurship by exploring different sources of heterogeneity. These findings and implications involve making precise adjustments to policies to better align with specific targets or conditions.

Chapter 2 analyses whether shocks in self-employment generate permanent or transitory impacts in Spain, using fractional integration models to differentiate between self-employed individuals with and without employees across various sectors and regions. Most of the analysed series exhibit mean reversion, indicating that shocks are transitory. However, self-employed individuals with employees show greater persistence, particularly in the construction sector, where hysteresis is observed in nearly half of the country. For self-employed individuals without employees, shocks are more persistent in the commerce, transport, and services sectors, especially in northern Spain, while the heavy and energy industries demonstrate a faster recovery nationwide. Additionally, positive temporal trends were identified in the services sector and negative trends in the agroindustry. These findings underscore the need for policies tailored to sectoral and regional characteristics, promoting sustainable conditions and labour resilience.

Chapter 3 uses fractional integration and cointegration methods to analyse the long-run relationship between loans to non-financial corporations, real gross domestic product, real gross fixed capital formation, the cost of borrowing differential between long- and short-term rates, and a proxy for the cost of debt, securities, and equity issuance. The analysis includes four Eurozone countries, namely Germany, France, Italy, and Spain, and spans the most recent decades. More precisely, fractional integration and cointegration models are estimated to investigate the persistence of the series as well as their long-run relationships and short-run dynamics using both unrestricted and restricted specifications. The univariate results are heterogeneous, the highest degrees of integration being found in the case of loans to non-financial corporations, whilst the multivariate ones provide evidence of a single fractional cointegration vector as well as of a lower adjustment speed to the long-run equilibrium compared to previous studies in all four countries.

Moreover, both the short- and long-run response of loans to exogenous shocks to real gross domestic product and the cost of borrowing differentials differs across countries because of country-specific factors.

Chapter 4 examines the long-run relationship between spot and future oil prices by considering nonlinearity, a potentially non-integer degree of integration, and allowing both canonical and fractional cointegration methods to compete for the best data fit. We also apply an ex-ante unit root test with endogenous break detection and a pre-break and post-break approach to compare canonical and fractional vector error correction models. Bridging a gap between nonlinear and fractional applications, this methodology uniquely detects nonlinearities in the long-run equilibrium without resorting to state-space or threshold models. Using weekly observations of spot and future prices from the U.S. Energy Information Administration, our estimates identify a structural break around 2004. The canonical contango situation detectable across the whole series breaks down upon examining the data's memory processes before and after this break. Our results also reveal a significant shift in the equilibrium model linking spot to future oil prices due to a global exogenous shock, suggesting two different, time-dependent equilibria: a backwardation equilibrium before the break and a near-unity contango after. This evidence demonstrates that global shocks can affect the spot-futures oil price relationship, inducing equilibrium switches.

Chapter 5 conjectures that necessity entrepreneurship is associated with the labour market, while opportunity entrepreneurship is related with the capital market. Thus, we analyse the relationship between necessity entrepreneurship and unemployment, as well as opportunity entrepreneurship and venture capital, gross fixed capital formation, and trademarks, designs, and applications in the UK from 2005Q3 to 2023Q2. First, we conduct a causal analysis, finding that unemployment both causes necessity entrepreneurship and is, in turn, caused by it; opportunity entrepreneurs cause venture capital; opportunity entrepreneurs are caused by trademarks, designs, and international applications; and finally, opportunity entrepreneurship and gross fixed capital formation mutually cause each other. We also study the four cointegration relationships, finding a very rapid positive fractional cointegration relationship between necessity entrepreneurs and the unemployed. In contrast, the remaining relationships exhibit slower cointegration, which are also positive but very weak in the case of venture capital.

Table 1. Thesis structure

CHAPTER	OBJETIVES	SCOPE	DATA AND SOURCES	ECONOMETRIC FAMEWORK
2	Revising the hypothesis of persistence in entrepreneurship at the macro-level, exploring the potential asymmetries across sectors.	UK	Entrepreneurship for 13 sectors. 1997Q1-2022Q4 (quarterly data). ONS Statistics UK	Traditional and fractional integration
3	Analysing whether shocks in self-employment generate permanent or transitory impacts, differentiating between self-employed individuals with and without employees across various sectors and regions.	Spain	Self-employed with and without employees, disaggregated by autonomous community and sector of activity. 1999Q1-2024Q2 (quarterly data). Labor Force Survey (EPA)	Fractional integration
4	Examining the long-run relationship between loans to non-financial corporations, real gross domestic product, real gross fixed capital formation, the cost of borrowing differential between long- and short-term rates, and a proxy for the cost of debt, securities, and equity issuance.	France, Germany, Italy and Spain	Loans to non-financial corporations, real GDP, real GFCF, cost of borrowing differential between long- and short-term rates and a proxy for the cost of debt, securities, and equity issuance. 2003Q1-2022Q4 (quarterly data). FRED and ECB database.	Fractional integration and cointegration

5	Evaluating the long-run relationship between spot and future oil prices by considering nonlinearity.		Spot and futures prices of WTI. 01/03/1981-12/15/2023 (weekly data). U.S. Energy Information Administration	One-break test. Traditional and fractional integration and cointegration.
6	Finding the relationship between necessity entrepreneurship and unemployment, as well as opportunity entrepreneurship and venture capital, gross fixed capital formation and trademarks, designs and applications.	UK	Necessity and opportunity entrepreneurs, unemployed, venture capital, GFCF and trademarks, designs and applications. 2005Q2- 2023Q2 (quarterly data). LFS, Dealroom.co, EUIPO and FRED.	Traditional and fractional integration and cointegration. Granger Causality

1.5. Publications

Some chapters of this PhD thesis are based on pieces of research published or submitted to academic journals. The chapters can be read independently of each other.

Chapter 2, coauthored by Emilio Congregado and Luis Alberiko Gil-Alana, was presented at the XVI Labour Economics Meeting in 2024 and has been submitted to Structural Change and Economic Dynamics.

Chapter 3, a work jointly done with Luis A. Gil-Alana, is accepted and forthcoming in issue 183 of Papeles de Economia Española.

The contents of chapter 4 correspond with the article published in 2024 in International Advances in Economic Research, jointly with Guglielmo Maria Caporale, Luis Alberiko Gil-Alana and Nicola Rubino.

Chapter 5 was done jointly with Luis Alberiko Gil-Alana and Nicola Rubino. This work was submitted to Journal of International Money & Finance.

Finally, chapter 6 was initiated during my stay as a visiting researcher at Florida Atlantic University, invited by Professor Joao Faria. Therefore, this work was carried out jointly with him, Emilio Congregado y Luis Alberiko Gil-Alana. It represents the first draft and is pending presentation and submission.

Artículos

Algunos de los trabajos científicos que forman parte de la tesis, debido a restricciones relativas a derechos de autor, han sido retirados. En sustitución de los artículos ofrecemos la siguiente información: referencia bibliográfica, enlace a la revista y resumen, si estuviera publicado.

- Gil-Alana. L.A., Vílchez Trujillo, I. "Persistencia en el autoempleo: asimetrías regionales y sectoriales". Papeles de Economía Española. ISSN 0210-9107, Nº 183, 2025 (Ejemplar dedicado a: Función empresarial y desarrollo regional: hacia una estrategia de cualificación del tejido empresarial), págs.. 90-110.

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Resumen:

Este estudio examina si los shocks en el autoempleo tienen impactos permanentes o transitorios utilizando series temporales españolas e integración fraccio- nal, considerando asimetrías regionales, sectoriales y según el tipo de autoempleo. Se distingue entre los trabajadores por cuenta propia que contratan a otros (empleadores) y aquellos que no lo hacen. El análisis re- vela que la mayoría de las series muestran reversión a la media, lo que indica efectos temporales de los shocks.

Sin embargo, los shocks que afectan a los empleadores tienden a mostrar cierta persistencia. Por sectores, la construcción es más propensa a experimentar efectos permanentes de shocks transitorios, observándose histéresis en casi la mitad de las regiones de España.

Para los autoempleados sin asalariados, los shocks per- sistentes se encuentran principalmente en el comercio, el transporte y los servicios, particularmente en las regiones del norte. En contraste, los shocks en las in- dustrias pesadas y energéticas revierten rápidamente a la media a nivel nacional. Además, se observan tenden- cias temporales positivas en los servicios, mientras que en la agroindustria aparecen tendencias negativas. Los resultados sugieren la necesidad de políticas específi- cas por sector y región para fomentar la sostenibilidad y la resiliencia.

- How Labor and Capital Markets Shape Necessity and Opportunity Entrepreneurs: Unpacking Their Causal Drivers. It represents the first draft and is pending presen-tation and submission.

Part II: Persistence and Long-run Relationships

Chapter 2: Seeking fine-tuned entrepreneurship policies: long-lasting effects of changes in entrepreneurship across sectors

This paper investigates the significance of sector-specific hysteresis effects on entrepreneurship dynamics at the macro level, building upon previous research that highlights the impact of hysteresis on aggregate self-employment. The persistence is explored by employing traditional unit roots tests and fractional integration techniques to infer the degree of persistence in UK self-employment rates, by industry, using quarterly data from 1997:01 to 2022:04. The results indicate an order of integration significantly above 1 in the case of total self-employment; however, mean reversion is observed in several sectors, including agriculture, wholesale and retail, accommodation and food services, financial and insurance activities, real estate, public administration and defense, human health, and other services. In contrast, the construction sector exhibits a value above 1, and the unit root hypothesis cannot be rejected in six other sectors. The results demonstrate the importance of advancing knowledge on the long-term effects of shocks in entrepreneurship by exploring different sources of heterogeneity, specifically sector-specific hysteresis effects on self-employment dynamics. These findings and implications involve making precise adjustments to policies to better align with specific targets or conditions.

2.1 Introduction

The success of entrepreneurship promotion policies is often associated with various outcomes achieved among the treated group. A vast literature analyzes the determinants of success in terms of earnings achieved (Hamilton, 2000; Dillon and Stanton, 2017), innovation (Koellinger, 2008), the emergence of high-growth firms -gazelles and unicorns- (Autio and Rannikko, 2016; Gonzalez-Uribe and Reyes, 2021), job satisfaction (Millan et al., 2013; Block et al., 2015; Johansson Sevä et al., 2016), and the job creation capacity of nascent entrepreneurs (Cowling, Taylor and Mitchell, 2004; Henrekson and Johansson, 2010; Haltiwanger, Jarmin and Miranda, 2013). However, most of this type of literature in entrepreneurship aims to analyze the fundamental factors affecting self-employment survival, implicitly considering continued self-employment as a desired outcome associated with entrepreneurial success (see, e.g., Dvouletý and Lukeš, 2016, for a survey). Consequently, a considerable part of the empirical literature on the factors influencing entrepreneurial success, as well as numerous evaluations and pseudo-evaluations of entrepreneurship promotion policies, inherently links self-employment survival with economic performance (Gimeno et al., 1997; Georgellis, Sessions and Tsitsianis, 2005; Burke et al., 2008; Caliendo and Kritikos, 2010; Millan, Congregado and Roman, 2012; Caliendo et al., 2015).[1] In this type of literature on the causal processes of entrepreneurship, micro-level analyses are predominant in the dynamics, and their findings serve as the foundation for informing the design of public policies.

[1] In general, it is common to this body of literature associating survival with better economic performance. However, we will agree that not all continuations are voluntary, nor are all exits from self-employment involuntary or necessarily imply a worsening of the situation for the agent who decides to leave self-employment (Boeri et al., 2020; Henley, 2021; O'Donnell et al., 2024). One exception is the work of Van Praag (2003), who analyzes the survival of entrepreneurs in the United States, distinguishing between those that are voluntary and those that are not.

However, as a caveat we note that the scope of these results may be limited and some of them biased due to at least some of the following circumstances. Firstly, tracking episodes of self-employment using individual data often does not allow for the monitoring of long-term career paths or capture long-term trends and structural changes, which hinders the assessment of the lasting effects of policies, as durations are frequently censored beyond a certain value[2]; secondly, self-employment survival does not have to be a desired outcome[3]; and thirdly, a substantial share of previous entrepreneurship research seems to have overlooked the differences across sectors.[4]

Instead, we explore the persistence into self-employment at the macro-level by using long aggregate time series, exploring the macro-dynamics of self-employment and the differences across-sectors. Our approach allows evaluating the long run effects of entrepreneurship exogenous shocks that are not accurately captured by conventional evaluations (Congregado, Golpe and Parker, 2012). At the core of this inquiry lies the question of whether self-employment follows a trend-stationary or non-stationary time-series process. If entrepreneurship is trend-stationary, economic and policy shocks can be seen as transitory in aggregate terms, as the self-employment rate eventually reverts to its long-term rate. Therefore, if the self-employment rate is trend-stationary or exhibits a broken trend, policy or economic shocks will only have temporary effects at the aggregate level. Conversely, if the self-employment rate is non-stationary, such shocks will have permanent effects. Following a common strategy in analyzing long-memory properties of time series, we examine persistence and hysteresis in entrepreneurship across different sectors. Our approach tests for unit roots and combines traditional analysis with other fractionally integrated methods.

The research employs quarterly data from the UK (1997-2022), examining fifteen activity sectors. The United Kingdom presents a unique context for the study of self-employment, characterized by rapidly increasing self-employment rates, a diverse array of self-employment forms, a flexible labor market, and a varied landscape of firm sizes (Henley, 2021). These interrelated factors create an exceptional case for analyzing self-employment dynamics. The complex interactions among these elements, along with the resulting heterogeneity within the self-employed population, provide valuable insights into the underlying causes and consequences of self-employment. Such findings are not only pertinent to the UK context but also offer significant implications for a wider international audience interested in the dynamics of entrepreneurship and labor market trends (Boeri et al., 2020).

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[2] The works of Caliendo and Künn (2011) and Caliendo and Tübbicke (2020, 2022) are some of the scarce evaluations conducted over the medium to long-term using microdata. However, these studies often track individuals over a limited time frame, which typically allows for an examination of short-term effects but not medium- or long-term outcomes. Consequently, the findings from this empirical literature at the micro level inform a considerable portion of the design of business promotion policies, yet they may not fully capture the sustained impacts of these interventions.

[3] In the case of policies promoting self-employment among the unemployed, the objective can be purely transitory, meaning that the self-employment spell helps alleviate the unemployment situation while contributing to enhancing the employability of the unemployed individual and facilitating their return to wage employment. In this case, survival as a self-employed individual is not the primary objective of the policy, at least in the medium and long term. Something similar applies in the case where self-employment survival may become in a kind of "poverty trap" whereby those workers, marginally attached to self-employment, survive due to the lack of alternative wage employment options, without it being associated with any success factors (Shane, 2009; Congregado, Golpe and Carmona, 2010; Henley, 2021). On the contrary, in the case of entrepreneurial promotion policies targeting opportunity entrepreneurs, survival is a target outcome.

[4] Survival rates across sectors differ due to several factors. Smaller firms generally face higher risks, especially in sectors with a larger minimum efficient scale and high turnover rates. Conversely, favorable growth dynamics and R&D intensity can enhance survival. Additionally, higher proportions of necessity entrepreneurs in a sector tend to lower survival probabilities, particularly for voluntary survival, as these entrepreneurs often have lower success rates compared to their opportunity counterparts.

This study represents a significant contribution to the literature by being the first to evaluate the long-term effects of shocks on entrepreneurship at a macro level, specifically by sector. Our findings reveal evidence of persistence in several sectors, although most exhibit mean reversion, with the construction sector standing out as the only exception that does not revert to the mean. This highlights the presence of sector-specific factors that influence long-term survival rates.

The implications of these results are profound for the formulation of entrepreneurship promotion policies. The paper emphasizes the critical importance of conducting long-term sector-specific evaluations when designing policies aimed at supporting entrepreneurship. Understanding the diverse impacts of hysteresis across different sectors allows policymakers to tailor their interventions effectively, addressing the unique characteristics that determine firm survival and resilience. Analyzing these sectoral differences not only enhances our understanding of the dynamics of self-employment but also guides the creation of targeted strategies that can mitigate the long-term consequences of economic shocks.

The rest of the paper proceeds as follows. In Section 2, the aim is to establish connections between the content of our work and the existing literature. In Section 3, we describe the context. In Section 4, we detail the methodology and data. We present and discuss the results, in Section 5, and offer concluding remarks and some avenues for further research in Section 6.

2.2 Related literature

This research relates to four main bodies of literature. The first addresses the determinants of survival, particularly in the context of an increasing body of literature on the longitudinal dynamics of self-employment that explores survival in self-employment over the medium term (see, e.g., Caliendo, Kritikos and Stier, 2023). In parallel, recent studies suggest that an entrepreneur's decision regarding persistence must be made repeatedly. New business cycle conditions, new regulations, and the emergence of alternative forms of employment versus the traditional employer-employee relationship can induce significant changes in occupational choice decisions (Caliendo, Goethner and Weißenberger, 2020).

The second issue pertains to the recognition of heterogeneity within self-employment. Self-employment is highly heterogeneous, and failure to acknowledge this has significant consequences, not only in terms of potential bias and poor inference in empirical literature. Disregarding the existence of different groups of self-employed individuals and the need for differentiated policies for each group in public policy design can negatively impact policy effectiveness. Differences based on start-up motivation, the distinction between solo self-employed individuals and job creators, incorporated versus unincorporated status, independent and dependent self-employment, as well as variations across sectors, are among the sources of heterogeneity that have been extensively explored in the literature –Millan, Congregado and Román (2012, 2014); Congregado, Golpe and Parker (2012); Gil-Alana and Payne (2015); Cieślik and Dvouletý (2019); Fairlie and Fossen (2020).

The third body of literature is linked to macro-level research devoted to evaluating the degree of persistence in entrepreneurship using time series analysis. The works of Congregado, Golpe, and Parker (2012), Parker, Congregado and Golpe (2012a), Gil-Alana and Payne (2015), and Lopez-Perez, Rodriguez-Santiago and Congregado (2020) investigate the persistence of entrepreneurship to determine whether "policy" shocks (resulting from government actions) or "economic" shocks (stemming from technological changes) have permanent or transitory effects on entrepreneurship. If entrepreneurship exhibits a unit root, shocks will have permanent effects, leading to hysteresis in entrepreneurial behavior. Conversely, if entrepreneurship is stationary, shocks will result in temporary deviations from the long-term growth trajectory or "natural rate." Congregado, Golpe and Parker (2012) employed unobserved component models to examine the presence of hysteresis in entrepreneurship in Spain and the United States. Using quarterly data from January 1987 to April 2008, they found evidence of hysteresis in Spain but not in the United States. Similarly, Parker, Congregado and Golpe (2012a) utilized panel unit root tests with structural breaks for 23 OECD countries and annual data from 1972 to 2006. Their findings suggest a high degree of persistence in shocks. Specifically, if entrepreneurship exhibits a unit root, policy shocks such as loan guarantee programs, subsidized start-up support services, technology transfer and innovation programs, changes in business start-up regulations, and favorable tax treatment for start-ups will have enduring effects on future levels of entrepreneurial activity. Conversely, if policy or economic shocks are temporary, their effects will diminish over time. Gil-Alana and Payne (2015) employed fractional integration techniques with seasonal adjustments to assess the degree of persistence in business ownership entrepreneurship in the United States, finding evidence of nonstationary yet mean-reverting behavior. More recently, Lopez-Perez, Rodriguez-Santiago and Congregado (2020) provided robust evidence of hysteresis in UK entrepreneurship, testing for unit roots and employing unobserved components models to separate the natural rate from cyclical fluctuations. Regarding the existence of differences among types of self-employed individuals, Congregado, Golpe and Parker (2012) pointed out that the long-term effects of exogenous shocks exhibit substantial differences among solo self-employed individuals and employers. Recently, Congregado, Fossen, Rubino and Troncoso (2024) considered differences between necessity and opportunity entrepreneurs.

The fourth body of literature examines hysteresis through a sector-specific lens. In the field of employment, the literature presents a series of arguments that would justify the existence of substantial heterogeneity in hysteresis effects on employment across sectors. The industrial sector exhibits the most significant hysteresis effects, consistent with findings from other studies that link hysteresis to high levels of capital investment and technological diffusion. In contrast, the services and wholesale/retail trade sectors demonstrate relatively weaker hysteresis effects. This divergence likely results from differences in sector-specific adjustment costs, the presence of non-convex costs related to capital and labor, the impact of economic uncertainty on investment decisions, and the interplay between labor demand and supply factors (Mota & Vasconcelos, 2024). However, these factors cannot be extrapolated to the field of self-employment. In the field of the economics of self-employment, differences in survival rates across sectors may be attributed to several circumstances. On one hand, sectors dominated by smaller firms appear to face a higher survival risk; however, this does not apply to the mature stages of the product life cycle or to technologically intensive products (Agarwal & Audretsch, 2001). In sectors where the minimum efficient scale is larger, smaller firms confront a greater risk of exit (Strotmann, 2007). Additionally, sectors characterized by high turnover, with elevated entry and exit rates, negatively impact survival (Ejermo & Xiao, 2014). Conversely, the growth dynamics of a sector tend to favor survival. For instance, Cefis and Marsili (2006) suggest that the intensity of R&D within a sector affects survival rates. Furthermore, it appears that belonging to highly agglomerated sectors increases the likelihood of exit (Aksaray & Thompson, 2018; Wennberg & Lindqvist, 2010). Lastly, if a sector exhibits a higher proportion of necessity entrepreneurs compared to opportunistic entrepreneurs, the survival probability—particularly in terms of voluntary survival—is likely to be lower. This hypothesis is supported by the findings of Caliendo and Kritikos (2010), Block and Sandner (2009), and Caliendo et al. (2015), which indicate that necessity entrepreneurs and subsidized start-ups out of unemployment demonstrate lower survival rates for their businesses compared to opportunity or regular entrepreneurs.

In sum, despite previous efforts to explore hysteresis in entrepreneurship at the macro level, we still lack a comprehensive understanding of whether differences exist among distinct subgroups of self-employed workers. Limited knowledge persists regarding potential disparities across sectors, and these distinctions may be significant. On one hand, the impacts of new labor market trends—particularly those related to task fragmentation, outsourcing, and on-demand work—are more pronounced in certain sectors than in others. These organizational and structural changes can influence occupational decisions. On the other hand, there are sectoral differences in the costs associated with entering and exiting self-employment, which can either facilitate or hinder transitions between self-employment and wage employment, whether such transitions are voluntary or involuntary, to varying degrees. Finally, sectoral differences will considerably affect the permeability between sectors, meaning the ability to make intersectoral transitions in response to sector-specific shocks or shocks that impact sectors with varying intensities.

This research contributes to three main areas of literature. Firstly, it aligns with the self-employment longitudinal dynamics literature, focusing on the determinants of survival in the medium and long-term. This piece of research connects with macro-level studies assessing the persistence of entrepreneurship over time, exploring the impacts of policy and economic shocks on entrepreneurship and hysteresis. This paper analyzes the role of sector-specific hysteresis effects in shaping self-employment dynamics, building upon prior research that has underscored the influence of hysteresis on overall self-employment trends.

Secondly, the study emphasizes the critical issue of recognizing the heterogeneity within self-employment, pointing out the potential consequences of neglecting such diversity on empirical literature and public policy design. Few previous studies investigating persistence in entrepreneurship at the macro level have taken into consideration that different types of entrepreneurs might exhibit different macro-dynamics. Thus, while heterogeneity in the long-term macrodynamics of self-employment has been studied concerning the characteristics of entrepreneurs, there has been a lack of analysis aimed at detecting the potential presence of distinct sector-specific hysteresis effects. This is precisely the aim of this paper.

2.3. Institutional setting

The UK offers a compelling case study for analyzing self-employment due to several significant recent shifts in its labor market. During the recovery period following the Great Recession, self-employment rates in the United Kingdom reached record levels, peaking at 5.0 million people at the end of 2019, a figure that represented 15.3% of total employment (ONS, 2022). This was contrary to expectations based on previous episodes of crisis, which suggested that self-employment would decrease and exhibit a countercyclical behavior characterized by a return to wage employment for marginally attached entrepreneurs –see, e.g. the empirical studies conducted by Faria et al. (2009, 2010), Carmona, Congregado and Golpe (2012), Carmona et al. (2016), Koellinger and Thurik (2012), Parker, Congregado & Golpe (2012b), Constant and Zimmermann (2014), Galindo and Mendez (2014), Gohmann and Fernandez (2014), Yu, Orazem and Jolly (2014), Fritsch and Kritikos (2016) and Fossen (2021).

However, despite very low unemployment rate, a situation of technically full employment, the levels of self-employment in the United Kingdom exhibited a certain persistence that did not diminish until the onset of the coronavirus pandemic.[5] This surprising trend has attracted the attention of analysts and scholars who have sought to explain this phenomenon and investigate the type of self-employment that experienced an upturn during the crisis and stability during the recovery.

Among the early hypotheses suggested, one explored whether the recovery lacked sufficient intensity to generate full-time employment opportunities that would enable a transition from self-employment to wage employment for marginally attached individuals, which could be associated with the existence of labor market slack in the UK, as stated by Bell and Blanchflower (2014). Another hypothesis examined the consequences of certain structural changes in the labor market, specifically the emergence of new forms of self-employment in parallel with the increasing prevalence of platform-mediated employment. This includes the so-called "GIG sector" in which traditional employer-employee relationships are progressively replaced by various forms

[5] Since then, the number of self-employed individuals has experienced a significant decline, dropping to 4.2 million, which has resulted in a self-employment rate of 13.2% in 2022.

of on-demand employment under employer-self-employed relationships (Boeri et al., 2020; Henley, 2021; Congregado et al., 2022).

Under these latter two hypotheses, and within the context of the entrepreneurship literature, one would expect that necessity entrepreneurs -i.e., those who entered self-employment due to a lack of wage employment opportunities during the recession- would have played a prominent role in these growth phenomena (see, e.g., the studies of Fossen, 2021, for the US, and Neymotin, 2021, for Canada). However, this hypothesis is refuted by Henley (2017) for the British case.

A third argument suggests that a combination of incentives promoting self-employment, the absence of retention mechanisms for employees, particularly in small and medium-sized enterprises, and low job protection for wage employment have led to a structural variation in the parameters governing occupational choices. As a result, an increasing number of individuals are willing to experience spells of wage employment and self-employment as normal situations throughout their working lives (Henley, 2021).

This recent evolution of entrepreneurship in the UK has also drawn the attention of analysts due to its implications in terms of the effectiveness of public policies promoting self-employment. Firstly, the persistence of self-employment rates in the post-Great Recession recovery could be interpreted in light of the long-term positive effects of self-employment promotion policies among the unemployed, which characterize entrepreneurial promotion during recessions. Secondly, if such persistence exists, it may be attributed to a combination of incentive schemes that, in conjunction with a particular framework of labor relations, generate structural changes with implications for occupational choices[6]. Thirdly, the effectiveness of job retention policies, particularly the mechanisms designed to support the survival of self-employed individuals during the COVID-19 crisis, is largely determined by their long-term effects, as they represent shocks with permanent consequences.

Overall, the UK's unique circumstances—including high and growing self-employment rates, diverse self-employment types, a flexible labor market, and a wide range of firm sizes—make it an exceptional context for examining self-employment. Analyzing self-employment through a sectoral lens offers fertile ground for understanding the underlying factors influencing survival rates and occupational choices, ultimately yielding insights that are valuable not only for the UK but also for a broader international audience interested in entrepreneurship dynamics.

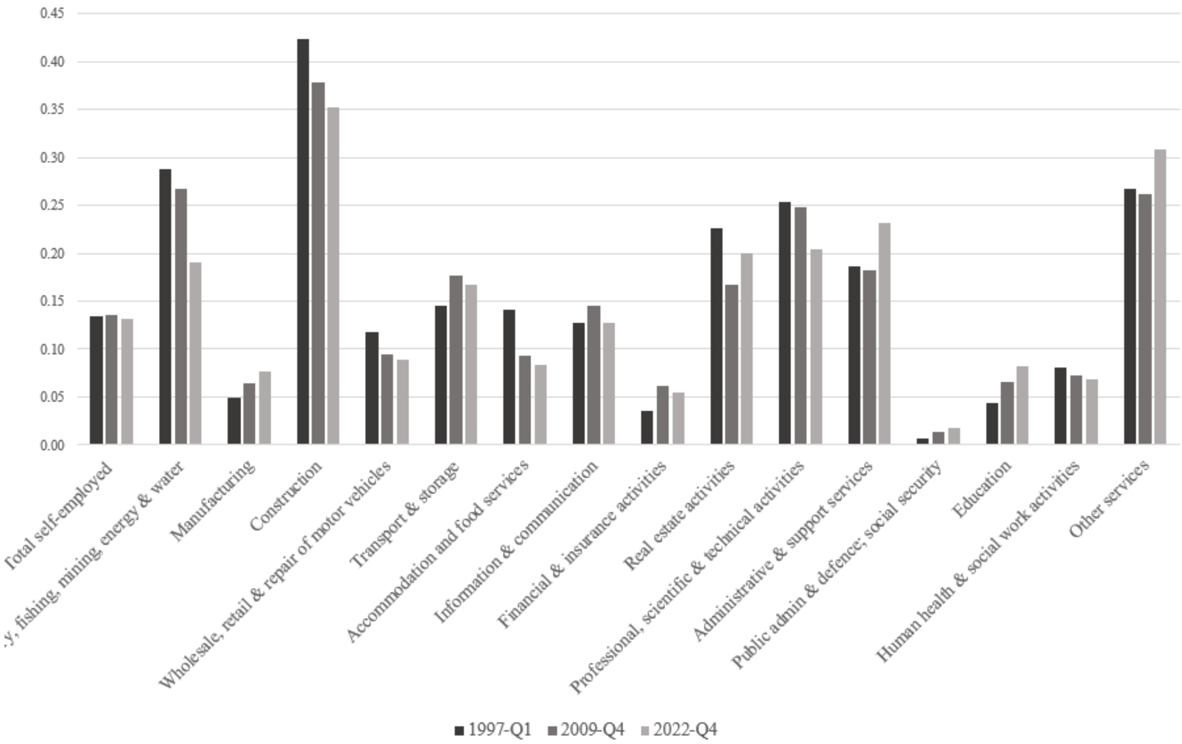
2.4.Data

We adopt the traditional operationalization of entrepreneurship at the macro level, using the ratio of self-employment to the labor force within a sector as a proxy for sectoral entrepreneurship. Data for the UK were obtained from the Office for National Statistics (ONS), covering the quarterly period from 1997:1 to 2022:4 across 13 sectors. As presented in Figure 1, the average self-employment ratio is 0.1352. We categorize sectors into those with ratios above or below this average. The sectors exceeding the average include construction; other services; agriculture, forestry, fishing, mining, energy, and water (AGRIC); professional, scientific, and technical activities (PROF); real estate activities (RSTATE); administrative and support services (ADMIN); and transport and storage (TRANSP). Conversely, the sectors below the average are wholesale, retail, and repair of motor vehicles (WHOLE); accommodation and food services (ACCOM); human health and social work activities (HUMAN); manufacturing (MANUF); financial and insurance activities (FINANC); and education, public administration, defense, and social security (EDUC).

[6] This context encompasses both frameworks that promote the emergence of opportunities and those that result in distortions and forms of dependent self-employment.

Information and communication (INF & COM) could be included in the first group, although in some periods it falls below the average.

Figure 1. Ratio self-employed by sectors (1997:1-2022:4)



2.5. Methodology

Dealing with the methodology we first conducted unit root tests. In particular, we perform Augmented Dickey Fuller (ADF) test (Dickey and Fuller, 1979) and Kwiatkowski et al. (1992) (KPSS test, hereafter) in order to test the possibility of a unit root or the existence of stationarity in the different sectors of self-employment in the UK.

The ADF test evaluates the null hypothesis that a time series, y_t, follows a unit root process (integrated of order 1, denoted as I(1)), as opposed to the alternative hypothesis that it is stationary (integrated of order 0, denoted as I(0)). The ADF test involves estimating the test regression:

$$\Delta y_t = (\emptyset_1 - 1)y_{t-1} + \sum_{i=1}^{p-1} \gamma_i \, \Delta y_{t-1} + \varepsilon_t \tag{1}$$

In our case, we have implemented this test considering three different alternatives: without constant or trend, with only a constant, and incorporating both constant and trend.

Before applying this test, it is necessary to define the optimal number of lags. Schwert (1987) suggests choosing the maximum number of lags $pmax = 12(\frac{T}{100})^{1/4}$. Once this maximum lag is defined, we will use the Akaike information criteria (AIC) and Bayesian information criteria (BIC) criteria to determine the optimal lag to use.

On the other hand, we validated the results of the ADF test by subsequently employing KPSS test in which the presence of a unit root is posited within the alternative hypothesis, while the null hypothesis is to be stationary. This test is based on the model:

$$y_t = \varepsilon_t + p_t + \theta_t, \tag{2}$$

$$and\ p_t = p_{t-1} + \varphi_t$$

where $\varphi_t \sim iid(0, \sigma_\varphi^2)$. The test relies on the premise that when the variance of the random walk p_t equals zero, the walk reduces to a constant. This suggests that the series exhibits stationarity with respect to a specified trend (preferably non-zero, but potentially identical to it with $\varepsilon = 0$. The authors also explore this scenario. Consequently, the relevant Lagrangean Multiplier test statistic will be predicated on the hypothesis $\sigma_\varphi^2 = 0$.

Nevertheless, the above unit root tests are very restrictive firstly because of the lack of flexibility in relation with the order of differentiation, being this exclusively 0 for stationarity and 1 or 2 for nonstationary series, and thus, not allowing for fractional numbers. In fact, it is a well-known stylized fact that these unit root methods have extremely low power if the Data Generating Process is fractionally integrated. Because of this, in the empirical application carried out in the following section, we consider the following model,

$$y(t) = \alpha + \beta t + x(t), \quad (1-B)^d x(t) = u(t), \quad u(t) = \rho u(t-1) + \varepsilon(t), \qquad (3)$$

where *y(t)* refers to the observed time series; α and β are unknown coefficients determined within the model, and $(1-L)^d$ being the fractional integration polynomial in L with fractional degree d, which makes the model integrated of order I(0). Note that based on the quarterly nature of the series examined, a seasonal AR(1) model is employed to describe this component. The estimation is based on a frequency expression of the Whittle function, throughout a testing procedure described in Robinson (1994). This is a Lagrange Multiplier (LM) test that report confidence bands for the non-rejection values of d. In the application we consider three potential scenarios: with no deterministic terms (i.e., $\alpha = \beta = 0$), including the intercept (i.e., $\alpha \neq 0,\ \beta = 0$), and with a linear trend (i.e., $\alpha \neq \beta \neq 0$).

2.6. Results

The results of the ADF and KPSS tests are presented in Table 1. It is important to evaluate these results jointly due to the tendency of the KPSS test to frequently reject the null hypothesis. As shown in Table 1, the aggregate self-employment rate exhibits a unit root (TOTAL), and this result is robust, i.e., it holds regardless of the unit root test employed. Across sectors, our findings are more diverse. Both tests indicate the existence of stationarity in the series for the following sectors: TRANS, ACCOM, RSTATE, and ADMIN. Furthermore, the ADF analysis extends the stationarity to four additional industries: MANUF, WHOLE, EDUC, and OTHER.

Table 1. Augmented Dickey-Fuller and KPSS test for integration order

Series	ADF				KPSS	
	No det. terms	A constant	A constant and a time trend	Decision	Test Statistic	Decision
TOTAL	-0.035	-1.371	-1.556	I(1)	0.226*	I(1)
AGRIC	-1.246	**-2.463**	-2.430	I(1)	0.215*	I(1)
MANUF	0.463	-1.696	**-3.514***	I(0)	0.155*	I(1)
CONSTR	0.249	-1.422	-0.073	I(1)	0.250*	I(1)
WHOLE	-1.171	**-2.417**	**-3.683***	I(0)	0.358*	I(1)
TRANSP	-0.093	**-1.986**	**-3.864***	I(0)	0.144	I(0)
ACCOM	-1.412	**-3.921***	**-4.625***	I(0)	0.120	I(0)
INF & COM	-0.304	-1.899	-1.687	I(1)	0.240*	I(1)
FINANC	-0.200	-1.721	-2.000	I(1)	0.221*	I(1)
RSTATE	-0.419	**-3.457***	-3.620*	I(0)	0.063	I(0)
PROF	-0.744	-1.297	-1.874	I(1)	0.247*	I(1)
ADMIN	0.653	-1.281	**-3.154**	I(1)	0.123	I(0)
PUBLIC	0.129	-1.556	-1.253	I(1)	0.172*	I(1)
EDUC	0.574	-1.016	**-3.713***	I(0)	0.203*	I(1)
HUMAN	-0.600	**-2.777**	-2.883	I(1)	0.189*	I(1)
OTHER	0.345	-1.505	**-3.546***	I(0)	0.192*	I(1)

*Null rejection at 5% level using the critical values of each distribution. AGRIC: Agriculture, forestry, fishing, mining, energy & wate; MANUF: Manufacturing; CONSTR: Construction; WHOLE: Wholesale, retail & repair of motor vehicles; ACCOM: Accommodation and food services; INF & COM: Information & communication; TRANSP: Transport & storage; FINANC: Financial & insurance activities; RSTATE: Real estate activities; PROF: Professional, scientific & technical activities; ADMIN: Administrative & support services; PUBLIC: Public admin & defence; social security; EDUC: Education; HUMAN: Human health & social work activities; OTHER: Other services

We next applied the fractional integration approach, and the results are presented in Tables 2 and 3.[7] Table 2 shows the estimates of *d* and their corresponding 95% confidence intervals for the three cases mentioned above: 1) no terms, 2) with a constant, and 3) with a constant and a linear time trend. Table 3 reports the estimated coefficients of the selected model for each series. We first observe that the time trend is necessary in several cases: MANUF, WHOLE, ACCOM, PUBLIC, EDUC, and OTHER. This coefficient is significantly positive for MANUF, PUBLIC, EDUC, and OTHER, while it is negative for WHOLE and ACCOM.

[7] Our objective is to determine whether the respective time series are I(0) stationary ($d = 0$); stationary with long memory ($0 < d < 0.5$); non-stationary but mean-reverting ($0.5 \leq d < 1$); or non-stationary and non-mean reverting ($d \geq 1$).

Table 2. Estimates of the differencing parameter in the model given by Eq. (3)

Series	No det. terms	A constant	A constant and a tin
TOTAL	0.93 (0.68, 1.12)	**1.24 (1.09, 1.49)**	1.23 (1.09, 1.48)
AGRIC	0.94 (0.77, 1.13)	**0.85 (0.73, 0.98)**	0.85 (0.7,5 0.98)
MANUF	1.02 (0.86, 1.24)	0.74 (0.61, 0.99)	**0.75 (0.58, 1.00)**
CONSTR	0.78 (0.51, 0.99)	**1.20 (1.01, 1.43)**	1.19 (1.01, 1.43)
WHOLE	0.98 (0.81, 1.16)	0.63 (0.53, 0.82)	**0.65 (0.53, 0.83)**
TRANSP	0.91 (0.74, 1.13)	**0.82 (0.68, 1.07)**	0.82 (0.65, 1.07)
ACCOM	0.91 (0.79, 1.07)	0.72 (0.51, 0.94)	**0.77 (0.64, 0.95)**
INF & COM	0.90 (0.72, 1.10)	**0.91 (0.77, 1.09)**	0.91 (0.77, 1.09)
FINANC	0.86 (0.68, 1.09)	**0.81 (0.69, 0.99)**	0.81 (0.69, 0.99)
RSTATE	0.81 (0.66, 0.97)	**0.42 (0.29, 0.60)**	0.41 (0.26, 0.61)
PROF	0.96 (0.76, 1.16)	**0.82 (0.69, 1.01)**	0.81 (0.68, 1.01)
ADMIN	0.91 (0.76, 1.08)	**0.87 (0.76, 1.02)**	0.86 (0.74, 1.02)
PUBLIC	0.75 (0.56, 0.97)	0.77 (0.65, 0.95)	**0.76 (0.63, 0.95)**
EDUC	0.92 (0.77, 1.10)	0.89 (0.78, 1.09)	**0.87 (0.74, 1.09)**
HUMAN	0.88 (0.69, 1.06)	**0.76 (0.62, 0.94)**	0.77 (0.64, 0.94)
OTHER	0.91 (0.73, 1.09)	0.71 (0.61, 0.86)	**0.67 (0.52, 0.85)**

AGRIC: Agriculture, forestry, fishing, mining, energy & wate; MANUF: Manufacturing; CONSTR: Construction; WHOLE: Wholesale, retail & repair of motor vehicles; ACCOM: Accommodation and food services; INF & COM: Information & communication; TRANSP: Transport & storage; FINANC: Financial & insurance activities; RSTATE: Real estate activities; PROF: Professional, scientific & technical activities; ADMIN: Administrative & support services; PUBLIC: Public admin & defence; social security; EDUC: Education; HUMAN: Human health & social work activities; OTHER: Other services

Table 3. Estimated coefficients on the selected models

Series	No det. terms	Constant (tv)	Time trend (tv)	Seasonality
TOTAL	1.24 (1.09, 1.49)	0.1345 (73.46)	------	0.116
AGRIC	0.85 (0.73, 0.98)	0.2853 (24.78)	------	-0.008
MANUF	0.75 (0.58, 1.00)	0.0492 (15.17)	0.00024 (2.05)	-0.027
CONSTR	1.20 (1.01, 1.43)	0.4332 (51.65)	------	0.149
WHOLE	0.65 (0.53, 0.83)	0.1181 (42.34)	-0.00029 (-3.96)	-0.087
TRANSP	0.82 (0.68, 1.07)	0.1453 (20.14)	------	-0.013
ACCOM	0.77 (0.64, 0.95)	0.1384 (22.51)	-0.00051 (-2.11)	-0.078
INF & COM	0.91 (0.77, 1.09)	0.1278 (15.58)	------	0.032
FINANC	0.81 (0.69, 0.99)	0.0366 (6.92)	------	0.044
RSTATE	0.42 (0.29, 0.60)	0.1989 (23.88)	------	0.061
PROF	0.82 (0.69, 1.01)	0.2527 (33.38)	------	0.135
ADMIN	0.87 (0.76, 1.02)	0.1838 (24.78)	------	-0.283
PUBLIC	0.76 (0.63, 0.95)	0.0674 (3.73)	0.00012 (1.73)	0.079
EDUC	0.87 (0.74, 1.09)	0.0434 (11.79)	0.00038 (1.82)	-0.241
HUMAN	0.76 (0.62, 0.94)	0.0790 (31.17)	------	-0.019
OTHER	0.67 (0.52, 0.85)	0.2670 (36.18)	0.00052 (2.50)	-0.010

AGRIC: Agriculture, forestry, fishing, mining, energy & wate; MANUF: Manufacturing; CONSTR: Construction; WHOLE: Wholesale, retail & repair of motor vehicles; ACCOM: Accommodation and food services; INF & COM: Information & communication; TRANSP: Transport & storage; FINANC: Financial & insurance activities; RSTATE: Real estate activities; PROF: Professional, scientific & technical activities; ADMIN: Administrative & support services; PUBLIC: Public admin & defence; social security; EDUC: Education; HUMAN: Human health & social work activities; OTHER: Other services.

Focussing on persistence, starting with TOTAL, we see that the estimated value of *d* is equal to 1.24, with a confidence interval of (1.04, 1.49), rejecting thus the unit root null hypothesis in favour of values of *d* above 1, and finding thus strong evidence against the hypothesis of mean reversion. However, if we look now at the disaggregated data by sectors, we observe that reversion to the mean is satisfied in eight of the series. For AGRIC (0.85), WHOLE (0.65), ACCOM (0.27), FINANC (0.81), RSTATE (0.42), PUBLIC and HUMAN (0.76) and OTHER (0.67). For MANUF (0.75), PROF (0.82) and ADMIN (0.87), the unit root cannot be rejected though it is just at the upper bound of the interval; for INF & COM (0.91) and EDUC (0.87) the unit root cannot be rejected, and it is decisively rejected in favour of *d* above 1 for CONSTR (1.20). On the other hand, seasonality does not seem to be a relevant issue in any of the series, the highest values observed at ADMIN (-0.283) and EDUC (-0.241).

Looking for robustness, we also perform some alternative methods based on semiparametric techniques. We use the log-periodogram regression, initially proposed by Geweke and Porter-Hudak (1983) and the subsequently modified and refined version of Robinson (1995). The results are respectively presented in Tables 4 and 5. To derive these estimates, we employ a bandwidth of *m=T^0.6* for the GPH log-periodogram regression and *m=T^0.85* for the Robinson log-periodogram regression to calculate the corresponding values of d for each series.[8] Finally, to obtain the appropriate value of *d*, we add 1 to the calculated estimates. In these results we have also considered the trend, therefore we work with the residuals after applying OLS. Once more, the series corresponding to the highest level of persistence is CONSTR.

Table 4. Estimates of *d* using the GPH log-periodogram regression

$m = T^{0.6}$			
Sector	$\hat{d}$	***Standard Error***	***P-Value***
TOTAL	0.314	0.305	0.305
AGRIC	-0.004	0.218	0.985
MANUF	-0.377	0.125	0.009
CONSTR	0.254	0.153	0.119
WHOLE	-0.192	0.343	0.590
TRANSP	0.289	0.196	0.162
ACCOM	-0.255	0.160	0.134
INF & COM	-0.134	0.187	0.484
FINANC	-0.047	0.258	0.857
RSTATE	0.175	0.213	0.424
PROF	-0.316	0.176	0.094
ADMIN	-0.253	0.240	0.309
PUBLIC	0.024	0.254	0.925
EDUC	-0.372	0.172	0.049
HUMAN	-0.107	0.305	0.731
OTHER	-0.161	0.148	0.296

AGRIC: Agriculture, forestry, fishing, mining, energy & wate; MANUF: Manufacturing; CONSTR: Construction; WHOLE: Wholesale, retail & repair of motor vehicles; ACCOM: Accommodation and food services; INF & COM: Information & communication; TRANSP: Transport & storage; FINANC: Financial & insurance activities; RSTATE: Real estate activities; PROF: Professional, scientific & technical activities; ADMIN: Administrative & support services; PUBLIC: Public admin & defence; social security; EDUC: Education; HUMAN: Human health & social work activities; OTHER: Other services.

[8] The estimation process is performed using first-differenced data, as this method requires limiting the results to the range of $-0.5 < d < 0.5$.

Table 5. Estimates of *d* using the Robinson log-periodogram regression

	$m = T^{0.85}$		
Sector	$\hat{d}$	***Standard Error***	***P-Value***
TOTAL	0.235	0.110	0.038
AGRIC	-0.165	0.090	0.071
MANUF	-0.089	0.096	0.358
CONSTR	0.223	0.090	0.017
WHOLE	-0.339	0.100	0.001
TRANSP	-0.039	0.108	0.720
ACCOM	-0.128	0.078	0.110
INF & COM	-0.032	0.089	0.721
FINANC	-0.143	0.113	0.214
RSTATE	0.425	0.120	0.001
PROF	-0.086	0.102	0.403
ADMIN	-0.070	0.120	0.561
PUBLIC	-0.114	0.117	0.336
EDUC	0.004	0.103	0.966
HUMAN	-0.189	0.118	0.115
OTHER	-0.145	0.093	0.126

AGRIC: Agriculture, forestry, fishing, mining, energy & wate; MANUF: Manufacturing; CONSTR: Construction; WHOLE: Wholesale, retail & repair of motor vehicles; ACCOM: Accommodation and food services; INF & COM: Information & communication; TRANSP: Transport & storage; FINANC: Financial & insurance activities; RSTATE: Real estate activities; PROF: Professional, scientific & technical activities; ADMIN: Administrative & support services; PUBLIC: Public admin & defence; social security; EDUC: Education; HUMAN: Human health & social work activities; OTHER: Other services.

In sum, the findings indicate that for eight sectors, shocks will have lasting (persistent) effects, although they will eventually revert to their respective long-run means. This suggests that the degree of persistence associated with shocks varies across sectors, leading to differing impacts of entrepreneurship policies.

2.7. Concluding Remarks

Despite previous attempts to examine hysteresis in entrepreneurship at the macro level, little is known about sector-specific differences. In this paper, we have sought to bridge, at least partially, this gap by exploring sectoral differences for gaining a better understanding of the new dynamics exhibited by self-employment in an advanced economy like the UK. To this end, this empirical work examined whether shocks to entrepreneurship have a permanent or a transitory impact in the UK across sectors. For this purpose, and looking for robustness, we applied alternative strategies for exploring unit roots, including fractional integration and seasonal autoregressive components. Our results support the hypothesis of a high level of persistence for the total self-employment series and our findings indicate that in nine sectors, shocks will yield lasting (persistent) effects, although they will ultimately revert to their respective long run means in eight of these sectors. This indicates that the level of persistence associated with shocks varies across sectors, resulting in different impacts of entrepreneurship policies. The results not only highlight the importance of considering the long term but also advise against the implementation of horizontal policies. Consequently, policymakers should recognize that certain temporary incentive schemes introduced to address specific adverse shocks (such as promoting self-employment among the unemployed during a recession) will produce lingering effects due to the persistence and long memory behavior observed in many sectors.

Moreover, it is important to note that, despite many sectors showing lasting but dissipating effects, the transition to self-employment is often irreversible at an aggregate level. Decisions regarding occupational choice can become permanent, aligning with the effects identified in a significant number of sectors, as persistence in self-employment can facilitate intersectoral flows. In this context, we can conjecture that entrepreneurial skills and the increasingly blurred boundaries between the skills required to perform entrepreneurial functions across different sectors may enable these sectoral flows.

The construction sector stands out as the only sector exhibiting clear persistence and non-mean reverting behavior. By identifying sectors most susceptible to hysteresis effects, policymakers can implement preventive measures to mitigate or minimize excessive market shocks, even those initially perceived as temporary. The varying intensity of hysteresis across sectors necessitates a sector-specific analysis to clarify these differences and comprehend the underlying mechanisms driving such variations.

In sum, the results not only highlight the importance of considering the long term but also advise against the implementation of horizontal policies. This contributes to a more holistic understanding of economic dynamics, particularly in analyzing sector-specific hysteresis's role in elucidating the long-term impacts of significant economic downturns, such as those experienced during the COVID-19 pandemic, which had disparate initial effects across different sectors. Such an understanding is vital for designing effective policies aimed at alleviating the long-term consequences of economic shocks.

Ultimately, policies that do not consider sector-specific hysteresis may lack efficacy and could even prove counterproductive. A sector-specific approach to policymaking enhances the likelihood of implementing successful mitigation and recovery strategies. In sum, this paper posits that sector-specific analyses of hysteresis are essential for grasping the long-term consequences of economic shocks and for crafting more effective and targeted economic policies. The heterogeneous nature of impacts across sectors underscores the necessity for a more nuanced approach to analysis and policymaking. In sum, our results might contribute to fine-tuned policies, involving making precise adjustments or optimizations to policies, strategies, or models to better align with specific requirements or conditions.

Despite the interest of our results, we cannot overlook potential limitations of our findings. Future research should emphasize examining the dynamics of hysteresis across different countries and investigate additional sectors that relate to self-employment, as this would enhance the generalizability and robustness of the findings. Furthermore, subsequent analyses could explore the micro-foundations of sector-specific hysteresis and the inter-sectoral dynamics, particularly in the context of the UK's labor market. Given the substantial shifts in self-employment trends and the rise of the gig economy, examining these transitions can enhance our understanding of long-term employment dynamics. Studying the interactions between hysteresis and other macroeconomic variables could provide valuable insights.

Finally, we emphasize the importance of interpreting the results carefully. Without data that differentiate between voluntary and non-voluntary survival in self-employment, persistence may indicate both success and the possibility of being trapped in self-employment. This idea is likely to inspire research and discussion on business survival and its various manifestations.

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Chapter 4: Modelling Loans to Non-Financial Corporations in the Eurozone: A Long-Memory Approach

This paper uses fractional integration and cointegration methods to analyse the long-run relationship between loans to non-financial corporations, real gross domestic product, real gross fixed capital formation, the cost of borrowing differential between long- and short-term rates, and a proxy for the cost of debt, securities, and equity issuance. The analysis includes four Eurozone countries, namely Germany, France, Italy, and Spain, and spans the most recent decades. More precisely, fractional integration and cointegration models are estimated to investigate the persistence of the series as well as their long-run relationships and short-run dynamics using both unrestricted and restricted specifications. The univariate results are heterogeneous, the highest degrees of integration being found in the case of loans to non-financial corporations, whilst the multivariate ones provide evidence of a single fractional cointegration vector as well as of a lower adjustment speed to the long-run equilibrium compared to previous studies in all four countries. Moreover, both the short- and long-run response of loans to exogenous shocks to real gross domestic product and the cost of borrowing differentials differs across countries because of country-specific factors.

4.1. Introduction

Credit plays an important role in the economy. In particular, the amount of loans provided to non-financial corporations (NFCs) is an indicator of the investment and spending decisions of the banking sector and thus also provides useful information to policy makers. Within the Eurozone in particular, bank lending is one of the major sources of financing to NFCs, with European firms heavily relying on bank lending to finance investment, especially in the case of small and medium-sized enterprises that have few alternatives to address their external financing needs (EIB, Revoltella et al., 2014). Credit is normally found to be highly correlated with asset prices and hence can help understand financial cycles. It also has an important role in the transmission of monetary policy to the real side of the economy. Loans are a key component on the asset side of the balance sheet of Eurozone banks, and thus a significant counterpart to monetary aggregates. Consequently, corporate lending and financing to NFCs are important measures to consider for assessing the monetary policy stance.

Detailed knowledge of the factors determining corporate loan decisions is therefore crucial for understanding monetary developments and the setting of monetary policy in the Eurozone. Credit growth for NFCs in this area has been trending downwards, especially because of the accession of new member states that have been more severely affected by the global financial crisis (GFC) of 2007-8. Stagnation of bank lending can be a severe constraint on economic growth in Europe where it plays a much more important role in financing the corporate sector than, for instance, in the United States (US). In the wake of the 2007-8 crisis the capacity of many banks to lend to relatively high-risk sectors and to young, innovative firms has been seriously hindered by capital constraints and a strong deterioration in the quality of the assets on their balance sheets. However, prior to the 2019 coronavirus disease outbreak (Covid 19) lending to NFCs was showing clear signs of recovery in all the four largest Eurozone countries because of the decreasing influence of various demand-side and supply-side factors related to the global financial crisis of

2007-8 which had depressed lending levels. Given their importance in the European context, this paper analyses the determinants of the amount of loans provided to NFCs in four countries belonging to the Eurozone, namely Germany, France, Italy, and Spain. The selection of countries has been made on the basis of the availability of relatively long time series, which are required to identify long-run relationships, and also to make our results comparable to those of other studies that have provided evidence on the same set of countries, which include Focarelli and Rossi (1998), Levieuge (2017) and Dajcman (2023).

The empirical framework is based on fractional integration and cointegration methods since most macroeconomic series appear to exhibit long-memory or long-range dependence (namely, their autocorrelations do not decay exponentially but rather according to a hyperbolic shape), which makes I(d) processes the most appropriate to model them since it allows shocks to have long-lasting effects.

After testing for the degree of fractional integration of loans to NFCs and their main determinants the long-run equilibrium relationships are examined using both unrestricted and restricted Fractionally Cointegrated Vector Autoregressive (FCVAR) models as in Johansen and Nielsen (2010, 2012), since these are shown to outperform standard Cointegrated Vector Autoregressive (CVAR) ones.

Literature review

Since the early 1990s a vast literature has developed on modelling credit to the private sector, especially within the central banking community because of its policy relevance. A common feature of these studies is the econometric framework used. Owing to the typically non-stationary nature of loans and their determinants, a Vector Error Correction Model (VECM) has normally been estimated.

Sørensen et al. (2010) were the first to use Johansen's (1992) methodology to explain the long-term behaviour of loans to NFCs in the Eurozone and identified three cointegrating relationships. Previous studies had generally modelled credit to the private sector. For instance, Hofmann (2001) estimated a 4-variable VECM for eight Eurozone countries from 1980 to 1998 and was unable to detect any cointegration relationships. Hülsewig (2003) analyzed German data using a 5-variable VECM. He found two cointegrating relationships which he interpreted as the credit demand and the credit supply equilibria, with credit demand reverting rather slowly to its long-run equilibrium and supply effects through their impact on lending rates being insignificant. Calza et al. (2006) estimated a 4-variable VECM for the Eurozone and detected one cointegration relationship interpreted as the credit demand equilibrium. Gambacorta and Rossi (2010) investigated possible non-linearities in the response of bank lending to monetary policy shocks in the Eurozone over the period 1985-2005 by means of an Asymmetric Vector Error Correction Model (AVECM) involving four endogenous variables; they found that the effect on credit, GDP and prices of a monetary policy tightening was larger than that of a monetary policy easing. This result supported the existence of an asymmetric credit channel in the Eurozone.

Other studies have focused on business lending in individual Eurozone countries. Focarelli and Rossi (1998) specified a 5-variable VECM model and found three cointegrating relationships, namely loan demand, a relationship between investment and borrowing requirements and the lending rate equaling risk-free government bond yields. Bridgen and Mizen (1999) investigated the interactions between investment, money holding and bank borrowing by private NFCs and identified long-run relationships for investment, money and borrowing, with the dynamics indicating the existence of feedback from money and credit disequilibria onto investment. Bridgen and Mizen (2004) found equilibrium relationships for investment, lending and money with causal linkages running from money and lending to investment and from money to lending in a dynamic model. Kakes (2000) analysed the role of bank lending in the monetary transmission mechanism in Germany following a sectoral approach and distinguishing between corporate lending household lending; they reported that banks respond to a monetary contraction by adjusting their security holdings rather than by reducing their loans portfolio. Finally, Plasil et al. (2012) showed that Czech banks had to restrict credit significantly when the financial crisis hit.

Other studies, such as Busch et al. (2010) and Tamasi and Vilagi (2011), estimate Vector Autoregressive (VAR) models with theory-based restrictions imposed on the impulse response functions to identify difference types of shocks. Ferrari et al. (2013) presented evidence suggesting survey indicators of credit conditions can be useful for macroprudential purposes. De Bondt et al. (2010) examined the information content of the Eurozone Bank Lending Survey for aggregate credit and output growth, which suggests that both price and non-price conditions and terms of credit matter for credit and business cycles.

More recently, Levieuge (2017) estimated both VAR and VECM specifications for bank loans to NFCs in France and reported that the former outperforms the latter and that the growth rate of equity prices is the best predictors of such loans. In another recent paper, Pitoňáková (2018) investigated the factors influencing the demand for private sector loans in the euro area, and concluded that loans are negatively related to the producer price index and real interest rates, while they are positively related to the industrial production index. Finally, Dajcman (2023) examined the impact of uncertainty shocks on the demand for business loans in individual euro area countries by estimating impulse response functions in the context of a Bayesian vector autoregression (VAR) model.

4.2. Methodology

The analysis involves two steps. First, the stochastic properties of loans to NFCs and their determinants are examined by means of both standard unit root tests and fractional integration methods (specifically, Sowell's (1992) exact maximum likelihood (EML) estimator and Robinson's (1994) tests based on the Lagrange multiplier (LM) principle). Second, the economic relationships linking them are investigated in the context of both standard and fractional cointegration multivariate models (in the latter case, the recently introduced FCVAR approach of Johansen and Nielsen (2012) is implemented). These methods are outlined below.

Long Memory Processes

An important characteristic of many economic and financial time series is their non-stationary nature, which can be described by a variety of models. Until the 1980s the standard approach was to use deterministic (linear or quadratic) functions of time, thus assuming that the residuals from the regression model were I(0) stationary. Later on, and especially after the seminal work of Nelson and Plosser (1982), a general consensus was reached that the non-stationary component of most series was stochastic, and unit roots (or first differences, I(1)) were most appropriate for them. However, the I(1) case is merely one particular model that can describe such behavior. In fact, the number of differences required to achieve I(0) stationarity is not necessarily an integer but could be any point on the real line, including fractional values. In the latter case, the process is said to be fractionally integrated or I(d). Long memory is a feature of observations that are far apart in time but highly correlated. This can be captured by fractionally integrated or I(d) models as the one given by Eq. (1) below, where d can be any real value, L is the lag-operator

$(Lx_t = x_{t-1})$ and u_t is I(0), defined for the purposes as a covariance stationary process with a spectral density function that is positive and finite at the zero frequency.

$$(1 - L)^d x_t = u_f, t = 0, \pm 1, \ldots \quad (1)$$

Although fractional integration can also occur at other frequencies away from zero, as in the case of seasonal and cyclical fractional models, the series used for this analysis do not have these features. Hence, standard I(d) models are estimated as shown in Eq. (1). The idea of fractional integration was introduced by Granger and Joyeux (1980), Granger (1980,1981) and Hosking (1981), although Adenstedt (1974) had already shown awareness of its representation. The polynomial $(1 - L)^d$ in Eq. (1) can be expressed in terms of its binomial expansion, such that, for all real d, x_t depends not only on a finite number of past observations but on the whole of its past history. In this context, d plays a crucial role since it indicates the degree of dependence of the series: the higher the value of d is, the higher the level of association between the observations will be.

More precisely, one can distinguish between several cases depending on the value of d. Specifically, if $d = 0$, $x_t = u_t$, x_t is said to be "short memory" or I(0), and if the observations are (weakly) autocorrelated (e.g. AR), then the values in the autocorrelations decay exponentially fast; if $d > 0$, x_t is said to exhibit long memory, so called because of the strong association between observations far apart in time. In this case, if d belongs to the interval $(0, 0.5)$, x_t is still covariance stationary, while $d > 0.5$ implies non-stationarity. Finally, if $d < 1$, the series is mean-reverting and therefore the effects of shocks disappear in the long run, whilst if $d \geq 1$, they persist forever. Hence the value of this parameter provides very useful information to policy makers.

There exist several methods to estimate and test the fractional differencing parameter d. Some of them are parametric while others are semi-parametric and can be specified in the time or in the frequency domain. Sowell (1992) analyzed the exact maximum likelihood (EML) estimator of the parameters of the autoregressive fractionally integrated moving average (ARFIMA) model in the time domain using a recursive procedure that allows a quick evaluation of the likelihood function. Doornik and Ooms (2003) refined this likelihood-based procedure, and Doornik and Ooms (2004) then applied this method to model inflation data in the United Kingdom (UK) and the United States (US).

Other parametric methods to estimate d in the frequency domain were proposed by, among others, Fox and Taqqu (1986) and Dahlhaus (2006). The small sample properties of these and other estimators were examined by Hauser (1999). A semi-parametric frequency domain estimator is the log-periodogram estimator proposed by Geweke and Porter-Hudak (1983), and other semi-parametric methods have been put forward by Velasco (1999a, 1999b) and Phillips and Shimotsu (2004, 2005) among others. Another approach widely employed in the empirical literature and in the present study is the parametric testing procedure of Robinson (1994), which is a Lagrange Multiplier (LM) test based on the Whittle function in the frequency domain. Robinson (1994) showed that, under certain very general regularity conditions, its LM-based statistic converges asymptotically to a standard $N(0,1)$ distribution, and this limit behaviour holds independently of the use of exogenous regressors (or deterministic terms) and the specific modelling assumptions about the I(0) disturbances. The tests of Robinson (1994) were applied to an extended version of the Nelson and Plosser's (1982) dataset in Gil-Alana and Robinson (1997) to test for unit roots and other long-memory processes when the singularity at the spectrum occurred at the zero frequency, as in the case in the series analysed here. Such tests have not been previously used to analyse the provision of loans to NFCs as in the present study. The results of the fractional integration analysis are reported in Table 2.

Fractional Cointegration

The Fractionally Cointegrated Vector Autoregressive (FCVAR) model was introduced by Johansen (2008) and further developed by Johansen and Nielsen (2010, 2012). It is a generalization of Johansen´s (1995) Cointegrated Vector Autoregressive (CVAR) model which allows for fractional processes of order d that cointegrate with order d-b. To introduce the FCVAR model one can start with the well-known, non-fractional, CVAR model. Let $Y_t, t = 1, \dots, T$ be a p-dimensional I(1) time series. The CVAR model can then be expressed as in Eq. (2):

$$\Delta Y_t = \alpha\beta' Y_{t-1} + \sum_{i=1}^{k} \Gamma_i \Delta Y_{t-i} + \varepsilon_t = \alpha\beta' L Y_t + \sum_{i=1}^{k} \Gamma_i \Delta L^i Y_t + E_t \qquad (2)$$

The simplest way to derive the FCVAR model is to replace the difference and lag operators Δ and L in Eq. (2) with their fractional counterparts, Δ^b and $L_b = 1 - \Delta^b$ respectively, to obtain the fractionally differentiated model given by Eq. (3) below. By considering the cointegration order as in $Y_t = \Delta^{d-b} X_t$, one obtains a cointegrated model where ϵ_t is a p-dimensional independent and identically distributed vector with mean zero and covariance matrix Ω (Eq. (4)):

$$\Delta^b Y_t = \alpha\beta' L_b Y_t + \sum_{i=1}^{k} \Gamma_i \Delta^b {L_b}^i Y_t + \varepsilon_t \qquad (3)$$

$$\Delta^d X_t = \alpha\beta' L_b \Delta^{d-b} X_t + \sum_{i=1}^{k} \Gamma_i \Delta^b L_b^i X_t + \varepsilon_t \qquad (4)$$

The parameters have the same interpretation as in the CVAR model. In particular, α and β are pxr matrices, where$0 \leq r \leq p$. The columns of β are the cointegrating relationships in the system corresponding to the long-run equilibria. The parameters Γ_i govern the short-run behaviour of the variables and the coefficients α represent the speed of adjustment towards equilibrium for each of the variables. Thus, the FCVAR model allows simultaneous modelling of the long-run equilibria, the adjustment responses to deviations from those and the short-run dynamics of the system. As an intermediate step towards the final model, a version of the model in Eq. (4) with $d = b$ and a constant mean term inside the cointegrating vector (Eq. (5)) is considered.

$$\Delta^{\mathrm{d}} \mathrm{X_t} = \alpha(\beta' \mathrm{L_d X_t} + \rho') + \sum_{\mathrm{i}=1}^{\mathrm{k}} \Gamma_{\mathrm{i}} \Delta^{\mathrm{d}} {\mathrm{L_d}}^{\mathrm{i}} \mathrm{X_t} + \varepsilon_{\mathrm{t}} \qquad (5)$$

Johansen and Nielsen (2012) and Nielsen and Morin (2014) discuss estimation and inference of this model, the latter providing computational codes for the calculation of estimators and test statistics. It is noteworthy that fractional differencing is defined in terms of an infinite series, but any actual sample will include only a finite number of observations. To calculate the fractional differences, one should assume that X_t was zero before the start of the sample. The bias introduced by this assumption is analysed by Johansen and Nielsen (2014) using higher-order expansions. They showed that it can be completely avoided by including a level parameter μ that shifts each of the series by a constant. The final estimated model, which will be tested against the standard CVAR one (Eq. (2)) by means of a log-likelihood test is given by Eq. (6):

$$\Delta^{\mathrm{d}}(\mathrm{X_t} - \mu) = \mathrm{L_d} \alpha\beta'(\mathrm{X_t} - \mu) + \sum_{\mathrm{i}=1}^{\mathrm{k}} \Gamma_{\mathrm{i}} \Delta^{\mathrm{d}} \mathrm{L_d^i}(\mathrm{X_t} - \mu) + \varepsilon_{\mathrm{t}} \qquad (6)$$

The asymptotic analysis in Johansen and Nielsen (2016) shows that the maximum likelihood estimators of $(d, \alpha, \Gamma, \dots . \Gamma_2)$ are asymptotically normal, while the maximum likelihood estimator of (β, ρ) is asymptotically mixed normal when $d_0 < 1/2$ and asymptotically normal when $d_0 > 1/2$. FCVAR models have recently been estimated for forecasting commodity returns by Dolatabadi et al. (2018); for forecasting political opinion polls by Nielsen and Shibaev (2018) and Jones et al. (2014); for commodity futures markets by Dolatabadi et al. (2014).

4.3. Data

The analysis focuses on four Eurozone economies, namely Germany, France, Italy and Spain. The following quarterly series are examined: real loans to Non-Financial Corporations (NFC), real gross capital fixed formation (RGFCF), real GDP (RGDP), the differential between the cost of borrowing for new long- and short -term loans in the Euro Area (DIFF), and the cost of debt, security and equity issuance (DEBT). The sample starts in 2003Q1 and ends in 2022Q4. The series were obtained from the Federal Reserve Economic Data (fred.stlouisfed.org, 2024), with the exception of the cost of borrowing for new long- and short-term loans used to calculate DIFF, which was taken from the European Central Bank Database (data.ecb.europa.eu, 2024).

Figures 1, 2, 3 and 4 display the time series plots of the series for each of the four countries examined. Real NFC loans generally trend upwards throughout the sample period, except in Spain, where they have decreased since 2010. Real Gross Fixed Capital Formation has been increasing steadily in France and Germany but has declined in Italy and Spain since the middle of the first decade of this century, and it fluctuates most in Germany and Italy. The behavior of real GDP shows the effects of the 2007-8 global financial crisis (GFC): after rising steadily in all four countries it fell significantly during the crisis before gradually recovering. Finally, long and short borrowing rates and private sector debt have been steadily declining in the past two decades, though since 2021 long term rates have spiked. Table 1 reports some descriptive statistics. It can be seen that RGFCF is a very sizeable component of RGDP; also, as a percentage of RGDP loans are at their highest in France and their lowest in Spain, the cost of borrowing and of debt are highest in France and Italy respectively, and lowest in Germany in both cases. Further, RGFCF and RGDP are the most volatile series, and there is evidence of either negative or positive skewness and of leptokurtosis in all cases.

The aim of the analysis is to test for the existence of a long-run relationship linking these variables, namely:

$$LOANS_t = f(RGDP, RGFCF, DIFF, DEBT) \qquad (7)$$

where RGDP and RGFCF capture the overall state of the economy and correspond to structural and cyclical components respectively; DIFF measures the cost of short- vis-à-vis long-term lending; DEBT captures the cost of alternative sources of financing including equity issuance. The priors of the analysis are the following: RGDP and RGFCF should have a positive effect on loans, the latter being a component of the former and acting as a scale variable (see, e.g., Focarelli and Rossi, 1998); DIFF, which measures the cost of borrowing (including bank lending), is expected to have a negative impact (see Calza et al., 2003), whilst DEBT should have a positive one, since an increase in the cost of alternative sources of financing provides an incentive to resort to bank loans instead (see Friedman and Kuttner, 1993).

4.4. Empirical Results

Order of integration

As a first step, standard unit root tests are conducted, specifically the ADF test developed by Dickey and Fuller (1979) and the stationary tests of Kwiatkowski et al. (KPSS). The results (not reported to save space) support the hypothesis of unit roots or nonstationarity in virtually all cases. However, it is well known that unit root tests have very low power against specific alternatives such as structural breaks (Campbell and Perron, 1991); trend-stationary models (DeJong et al., 1992), regime-switching (Nelson et al., 2001), or fractional integration (Diebold and Rudebusch, 1991; Hassler and Wolters, 1994; Lee and Schmidt, 1996). Therefore, in the present study a more general fractional integration framework is considered, which includes the classic unit root models as a particular case of interest.

The fractional integration parameter d is obtained by estimating Eq. (8), where y_t is the observed time series; β_0 and β_1 are the coefficients on the intercept and the linear time trend respectively, and the disturbance term u_t is $I(0)$ and assumed to be a white noise.[1]

$$y_t = \beta_0 + \beta_1 t + x_t, \ (1-L)^d x_t = u_t, \ t = 1,2,... \qquad (8)$$

Table 2 displays the Whittle estimates of d along with the 95% confidence intervals of the non-rejection values of d using Robinson's (1994) parametric approach. The estimates of d are reported for the three standard cases of no regressors in the undifferenced regression (i.e., $\beta_0 = \beta_1 = 0$ in Eq. (8)), an intercept ($\beta_0$unknown and $\beta_1 = 0$), and an intercept with a linear time trend (both β_0 and β_1 unknown). The coefficients in bold are those for the selected model according to the statistical significance of the regressors; note that Robinson's (1994) approach is based on the null differenced model, which is $I(0)$ by construction, and thus the t-values are still valid in the differenced regression model. In other words, under the null hypothesis in Eq. (9), Eq. (8) becomes $\tilde{y}_t = \beta_0 \tilde{1}_t + \beta_1 \tilde{t}_t + u_t$, where $\tilde{t}_t = (1-L)^{d_0} 1_t$, and $\tilde{t}_t = (1-L)^{d_0} t_t$, and , given that u_t is I(0) by construction, standard t-tests apply.

$$\mathrm{H_o}: d = d_o \qquad (9)$$

In the case of LOANS, the estimates of d are all significantly above 1 in the four countries examined, which suggests the presence of a unit root or a higher integrated process.

Concerning RGDP the time trend coefficient is significant in all four countries and the estimates of d are all in the $(0, 1)$ interval, ranging from 0.36 for Germany to 0.56 for Spain. Concerning Germany, France and Italy, the results support the stationary assumption since d is found to be significantly below 0.5; finally, for Spain the confidence intervals include both stationary and nonstationary values.

Regarding RGFCF, the time trend coefficient is statistically significant in the case of Germany and France but insignificant for the other two countries (Italy and Spain). The estimated value of the fractional coefficient d is heterogeneous across countries: 0.22 for Germany and 0.46 for France, in both cases supporting the hypothesis of stationarity with long memory; 0.57 for Italy, which implies non-stationarity, and finally 1.13 for Spain, where the unit root null hypothesis could not be rejected.

For DIFF, evidence of unit roots is found in the case of Germany and France, while mean reversion (i.e., d < 1) occurs in Italy ($d = 0.64$) and Spain (d = 0.30). Finally, in the case of DEBT the time trend is insignificant, and the unit root null cannot be rejected for Germany, France and Spain, while Italy is the only country with evidence of d > 1.

Fractional Cointegration

Next, the estimates the FCVAR models are examined. As mentioned in Section 3.2, to account for the initial bias value of zero, the approach of Johansen and Nielsen (2014) is followed by incorporating a level parameter (μ in Eq. (6)), which shifts each series by a nonzero constant to

[1] Note that the I(0) u_t term also allows for (weakly) ARMA-types of autocorrelations, with very similar results in this case.

be estimated. Furthermore, following Johansen and Nielsen (2012) and Nielsen and Morin (2014), and given the heterogeneous results of the fractional tests, the fractional parameter d and the fractional exponent b are restricted to be equal, so that the cointegrating order of the fractional processes of the series analysed is given by $CI(d-b=0)$, as shown in Eq. (5).

First, the number of lags is chosen using the Akaike Information Criterion (AIC) and Bayesian Information Criterion (BIC), both of which suggest a parsimonious model with no more than two lags in each of the four estimations. Serial correlation is tested using a White Noise Test. Next, the number of rank relations between the variables is determined using a Likelihood Ratio (LR) test for the cointegrating rank, as outlined in Johansen (1995) for the CVAR case and Johansen (2008) for the FCVAR case. Then, LR tests are performed of the adequacy of the CVAR model (where the null hypothesis is $d=b=1$, i.e., the order of cointegration is *CI(0)* and fractional differentiation is not necessary) against the FCVAR one (the alternative hypothesis being $d=b\neq 1$, i.e., fractional differentiation is required to find a stationary long-run equilibrium of order $CI(1)$).

Finally, LR tests of a set of overidentifying restrictions are carried out for both the FCVAR and CVAR models in order to choose the best specification. In particular, the null that α and β are equal to 0 is tested. Failure to reject this hypothesis in the case of α implies that the corresponding variable adjusts over time towards the long-run equilibrium, whilst in the case of β the implication is that the variable enters the long-run equilibrium relationship.[2]

France

The results for the LR fractional cointegration test for France, as well as the estimated fractional parameter and the analysis of the residuals for the unrestricted model, are shown in Table 3. Table 4, column 1, reports the LR test statistic for comparing a standard CVAR against the FCVAR alternative in the case of France. This implies that the null of no statistically significant difference between the likelihood of the two competing models should be rejected, and that the FCVAR should be preferred. The unrestricted FCVAR results (its estimated αs and βs) are reported in Eq. (10). [3] The LR test for the cointegrating rank in Table 4 implies the existence of at most one cointegrating relationship. Equation (10) shows that loans are the only variable converging towards the long-run equilibrium after an exogenous shock has occurred, though at a very slow rate (α=0.021), while the other diverge as indicated by their positive α coefficients.

By normalizing with respect to loans, it is obtained the long-run equilibrium vector given by Eq. (11), with positive coefficients on RGDP and RGFCF as expected, and a positive one on DIFF, which is in contrast to the prior.

$$\Delta^{\hat{d}}\begin{bmatrix}\mathrm{LOANS}_t-\mu_1\\ \mathrm{RGDP}_t-\mu_2\\ \mathrm{RGFCF}_t-\mu_3\\ \mathrm{DIFF}_t-\mu_4\end{bmatrix}=L_{\hat{d}}\begin{bmatrix}-0.021\\ 0.092\\ 0.225\\ 0.356\end{bmatrix}'\begin{bmatrix}1.000\\ -4.256\\ -0.264\\ -0.190\end{bmatrix}(X_t-\hat{\mu})+\sum_{i=1}^{2}\hat{\Gamma}_i\Delta^{\hat{d}}L_{\hat{d}}^{i}(X_t-\hat{\mu})+\hat{\varepsilon}_t \quad (10)$$

$$\mathrm{LOANS}_t=mu+4.256\mathrm{RGDP}_t+0.264\mathrm{RGFCF}_t+0.190\mathrm{DIFF}+\nu_t \quad (11)$$

[2] Figures reporting a graphical rendition of the characteristic polynomials for both the CVAR and the FCVAR models in each country analysed are reported in the online appendix.

[3] Note that p-values are not reported in the case of the unrestricted models but only for the selected final specification.

Restricted model for France

This section reports a set of LR tests for restrictions concerning the long-run significance of the determinants of loans being considered (setting the β coefficient on each of the covariates equal to 0) and the weak exogeneity of each series (setting each α coefficient equal to 0). The final specification is given by Eq. (12), which shows the estimates of the α and β parameters as well as the corresponding p-values (in parentheses) of the LR tests of the null of a zero coefficient. The normalized long-run equilibrium is given by Eq. (13). The significant coefficients are also shown in Table 4, Columns 2 and 3, alongside the CVAR-FCVAR test.

The long-run relationship now only includes RGDP and DIFF as determinants of loans, the latter now having a negative coefficient as expected, though of a small magnitude. The adjustment coefficient on LOANS is almost double in size, having increased from -0.021 to -0.049. However, it is still much lower than the corresponding estimate of -0.286 reported by Levieuge (2017) on the basis of standard CVAR analysis. This is in fact the only negative adjustment coefficient, which implies that loans are the only endogenous variable in the model.

$$\Delta^{\hat{d}}\begin{bmatrix}\mathrm{LOANS}_t-\mu_1\\ \mathrm{RGDP}_t-\mu_2\\ \mathrm{RGFCF}_t-\mu_3\\ \mathrm{DIFF}_t-\mu_4\end{bmatrix}=L_{\hat{d}}\begin{bmatrix}-0.049\\(0.001)\\0.000\\(0.408)\\0.094\\(0.064)\\0.459\\(0.397)\end{bmatrix}'\begin{bmatrix}1.000\\(0.000)\\-4.215\\(0.048)\\0.000\\(0.091)\\0.088\\(0.008)\end{bmatrix}(X_t-\hat{\mu})+\textstyle\sum_{i=1}^{2}\hat{\Gamma}_i\Delta^{\hat{d}}L_{\hat{d}}^{i}(X_t-\hat{\mu})+\hat{\varepsilon}_t \qquad (12)$$

$$\mathrm{LOANS}_t=mu+4.215\mathrm{RGDP}_t-0.088\mathrm{DIFF}+\nu_t \qquad (13)$$

Germany

The results for the LR test for model selection for Germany, as well as the estimated fractional parameter and the analysis of any correlation left in the residuals of the unrestricted model estimates, are reported in Table 5. Only 1 lag is required in the German case, there is no evidence of serial correlation, and the LR test supports an FCVAR specification rather than a CVAR one, as can be seen in Table 6, Column 1. The rank test implies that there are up to three cointegrating vectors, but only the one corresponding to a demand-driven equilibrium relationship is selected. The estimated value for both d and b is 0.517, which implies slow convergence towards the long-run equilibrium and the existence of an I(0) cointegrating relationship.

Equation (14) shows that all the α coefficients are positive, but none of them are statistically significant. However, the normalized cointegrating vector indicates that there exists a long-run equilibrium relationship with a positive coefficient on RGDP and DEBT and a negative one on DIFF, consistently with the priors.

$$\Delta^{\hat{d}}\begin{bmatrix}\mathrm{LOANS}_t-\mu_1\\ \mathrm{RGDP}_t-\mu_2\\ \mathrm{DIFF}_t-\mu_3\\ \mathrm{DEBT}_t-\mu_4\end{bmatrix}=L_{\hat{d}}\begin{bmatrix}0.018\\0.011\\0.187\\1.404\end{bmatrix}'\begin{bmatrix}1.000\\2.127\\0.315\\-0.330\end{bmatrix}(X_t-\hat{\mu})+\textstyle\sum_{i=1}^{2}\hat{\Gamma}_i\Delta^{\hat{d}}L_{\hat{d}}^{i}(X_t-\hat{\mu})+\hat{\varepsilon}_t \qquad (14)$$

$$\mathrm{LOANS}_t=mu+2.127\mathrm{RGDP}_t-0.315\mathrm{DIFF}_t+0.330\mathrm{DEBT}_t+\nu_t \qquad (15)$$

Restricted model for Germany

Again, the starting point is to test for the significance of the long run β coefficients as well of the dynamic adjustment α coefficients to obtain a restricted model which only includes the relevant determinants of loans and considers the weak exogeneity properties of the variables of interest. Table 6 reports, in Columns 2 and 3, the LR tests for the α and β parameters where the 0 null could not be rejected, while Eq. (16) shows the final restricted model together with the p-values from the LR tests in parentheses, and Eq. (17) displays the final equilibrium relationship normalised on LOANS.

Evidence is found of short-run exogeneity for DIFF (since the null that the corresponding α coefficient is equal to 0 cannot be rejected), while RGDP and DEBT exhibit high speeds of convergence (-0.297 and -0.272 respectively), and the corresponding estimate for loans, equal to -0.079, is much closer to the values reports by previous studies on Europe (e.g. Calza et al. (2006) with a -0.075).

Concerning the long-run equilibrium determinants, the estimated coefficient of 0.726 for RGDP is only marginally significant at the 10% level with a p-value equal to 0.51, whjle the coefficient on DEBT is equal to 0.108, and is highly significant s indicated by its p-value. However, both these coefficients are inconsistent with the priors. DEBT does not adjust towards the long-run equilibrium and thus can be considered weakly exogenous.

$$\Delta^{\hat{d}}\begin{bmatrix}\mathrm{LOANS}_t-\mu_1\\ \mathrm{RGDP}_t-\mu_2\\ \mathrm{DIFF}_t-\mu_3\\ \mathrm{DEBT}_t-\mu_4\end{bmatrix}=L_{\hat{d}}\begin{bmatrix}-0.079\\(0.001)\\-0.297\\(0.000)\\0.000\\(0.823)\\-0.272\\(0.828)\end{bmatrix}'\begin{bmatrix}1.000\\(0.010)\\0.726\\(0.051)\\0.000\\(0.673)\\0.108\\(0.000)\end{bmatrix}(X_t-\hat{\mu})+\textstyle\sum_{i=1}^{2}\hat{\Gamma}_i\Delta^{\hat{d}}L_{\hat{d}}^{i}(X_t-\hat{\mu})+\hat{\varepsilon}_t \qquad (16)$$

$$\mathrm{LOANS}_t = mu - 0.726\mathrm{RGDP}_t - 0.108\mathrm{DEBT}_t + \nu_t \qquad (17)$$

Italy

In Table 7, Column 1, the LR test implies again that an FCVAR specification is preferable to a CVAR one, and again there is no evidence of serial correlation. Furthermore, the rank test suggests that there exists a single cointegrating vector. Table 7, as before, also reports the residual analysis of the unrestricted model for Italy. As in the case of Germany, there is evidence of $I(1)$ behaviour for the individual series and $I(0)$ for the long-run relationship, the value of d and b being 0.655.

In the unrestricted model given by Eq. (18), loans adjust to the long-run equilibrium rather slowly, the corresponding α coefficient being equal to -0.010; note that this result implies a much slower convergence rate compared to the adjustment coefficient of -0.072 estimated by Calza et al. (2003) for Italy, and of -0.060 reported by Casolaro et al. (2006) for the Euro Area. DIFF is the only variable with a positive α coefficient.

$$\Delta^{\hat{d}}\begin{bmatrix}\mathrm{LOANS}_t-\mu_1\\ \mathrm{RGDP}_t-\mu_2\\ \mathrm{RGFCF}_t-\mu_3\\ \mathrm{DIFF}_t-\mu_4\\ \mathrm{DEBT}_t-\mu_5\end{bmatrix}=L_{\hat{d}}\begin{bmatrix}-0.010\\-0.116\\-0.300\\0.026\\-0.845\end{bmatrix}'\begin{bmatrix}1.000\\1.687\\-1.662\\-0.276\\0.148\end{bmatrix}(X_t-\hat{\mu})+\textstyle\sum_{i=1}^{2}\hat{\Gamma}_i\Delta^{\hat{d}}L_{\hat{d}}^{i}(X_t-\hat{\mu})+\hat{\varepsilon}_t \qquad (18)$$

$$\text{LOANS}_t = mu - 1.687\text{RGDP}_t + 1.662\text{RGFCF}_t + 0.276\text{DIFF}_t - 0.148\text{DEBT}_t + \nu_t \quad (19)$$

Restricted model for Italy

In Table 8, the final specification is discussed based on tests for the significance of the long-run parameters and of the adjustment ones, the latter being weak exogeneity tests. The estimates for the restricted model together with the p-values of the LR tests are shown in Eq. (20), whilst the normalized long-run equilibrium is given by Eq. (21). In this case RGDP does not enter the cointegrating relationship, whilst its component RGFCF does, and DIFF and DEBT do as well, with a positive and negative coefficient respectively. The estimated α on loans is equal to -0.013, which implies slightly faster convergence compared to the unrestricted model, whilst the adjustment coefficient on DIFF is positive and only slightly bigger in absolute terms compared to the unrestricted model, and the biggest negative coefficient (-0.940) is the one on DEBT, with DIFF being the only exogenous variable. Note that the adjustment coefficient on LOANS is very low and insignificant (the p-value for the LR test is 0.317).

$$\Delta^{\hat{d}}\begin{bmatrix}\text{LOANS}_t - \mu_1\\ \text{RGDP}_t - \mu_2\\ \text{RGFCF}_t - \mu_3\\ \text{DIFF}_t - \mu_4\\ \text{DEBT}_t - \mu_5\end{bmatrix} = L_{\hat{d}}\begin{bmatrix}-0.013\\(0.317)\\-0.111\\(0.000)\\-0.305\\(0.000)\\0.000\\(0.590)\\-0.940\\(0.042)\end{bmatrix}'\begin{bmatrix}1.000\\(0.069)\\0.000\\(0.302)\\-1.289\\(0.007)\\-0.248\\(0.024)\\0.140\\(0.001)\end{bmatrix}(X_t - \hat{\mu}) + \sum_{i=1}^{2}\hat{\Gamma}_i\Delta^{\hat{d}}L_{\hat{d}}^{i}(X_t - \hat{\mu}) + \hat{\varepsilon}_t \quad (20)$$

$$\text{LOANS}_t = mu + 1.289\text{RGFCF}_t + 0.248\text{DIFF}_t - 0.140\text{DEBT}_t + \nu_t \quad (21)$$

Spain

The results for the LR tests for Spain, as well as the estimated fractional parameter of the unrestricted model and the serial correlation tests are reported in Table 9. Again, the LR test supports the FCVAR model (Table 10, column 1) and the rank test suggests a single cointegrating vector. Finally, the order of integration (d and b) is equal to 0.633.

In the unrestricted model given by Eq. (22), the adjustment coefficients are negative in the case of LOANS and RGFCF, the former estimate being close to previous euro area estimates and on those in the literature based on CVAR specifications (see, for instance, Calza et al. (2006), who model the stock of private sector bank loans for the Euro Area as a function of an inflation index and the cost of borrowing, estimating an error correction coefficient of around -0.075). In the normalized cointegrating vector given by Eq. (23), RGDP and DIFF have a positive and sizeable impact, while the coefficient on DEBT (also positive) is rather small, and rather puzzlingly RGFCF appears to have a negative effect.

$$\Delta^{\hat{d}}\begin{bmatrix}\text{LOANS}_t - \mu_1\\ \text{RGDP}_t - \mu_2\\ \text{RGFCF}_t - \mu_3\\ \text{DIFF}_t - \mu_4\\ \text{DEBT}_t - \mu_5\end{bmatrix} = L_{\hat{d}}\begin{bmatrix}-0.086\\0.198\\-0.142\\0.799\\0.258\end{bmatrix}'\begin{bmatrix}1.000\\-1.666\\0.465\\-0.214\\-0.006\end{bmatrix}(X_t - \hat{\mu}) + \sum_{i=1}^{2}\hat{\Gamma}_i\Delta^{\hat{d}}L_{\hat{d}}^{i}(X_t - \hat{\mu}) + \hat{\varepsilon}_t \quad (22)$$

$$\text{LOANS}_t = mu + 1.666\text{RGDP}_t - 0.465\text{RGFCF}_t + 0.214\text{DIFF}_t + 0.006\text{DEBT}_t + \nu_t \quad (23)$$

Restricted model for Spain

Once again, the starting point is to test for the significance of both the long-run β coefficients and the adjustment α ones to obtain a restricted specification with a better fit. Both are found to be insignificant in the case of DEBT, whilst the estimated adjustment coefficient is now slightly higher (-0.082) in the case of loans and insignificant in the case of RGFCF.

The long-run relationship includes RGDP, DIFF and RGFCF, the latter again with a negative coefficient. Note that the former two have sizeable and positive adjustment coefficients (0.189 and 0.789 respectively), which implies divergence from the long-run equilibrium.

$$\Delta^{\hat{d}}\begin{bmatrix}\text{LOANS}_t - \mu_1\\ \text{RGDP}_t - \mu_2\\ \text{RGFCF}_t - \mu_3\\ \text{DIFF}_t - \mu_4\\ \text{DEBT}_t - \mu_5\end{bmatrix} = L_{\hat{d}}\begin{bmatrix}-0.082\\ (0.002)\\ 0.189\\ (0.005)\\ -0.136\\ (0.013)\\ 0.789\\ (0.007)\\ 0.000\\ (0.690)\end{bmatrix}'\begin{bmatrix}1.000\\ (0.000)\\ -1.529\\ (0.029)\\ 0.419\\ (0.243)\\ -0.228\\ (0.011)\\ 0.000\\ (0.724)\end{bmatrix}(X_t - \hat{\mu}) + \sum_{i=1}^{2}\hat{\Gamma}_i\Delta^{\hat{d}}L_{\hat{d}}^{i}(X_t - \hat{\mu}) + \hat{\varepsilon}_t \quad (24)$$

$$\text{LOANS}_t = mu + 1.529\text{RGDP}_t - 0.419\text{RGFCF}_t + 0.228\text{DIFF}_t + \nu_t \quad (25)$$

4.5. Conclusions

This paper investigates the determinants of loans to NFCs in four countries belonging to the Euro Zone (Germany, France, Italy and Spain). The findings are of interest not only to academics, but also to practitioners and European policy makers since this type of financing is much more important for the corporate sector in Europe than elsewhere. The modelling approach is based on fractional integration and cointegration methods, which have the advantage of considering the possible long-memory properties of the series of interest. Specifically, univariate models are estimated for the individual series and both unrestricted and restricted FCVAR models to examine linkages between them.

The univariate results are very heterogeneous across the variables and the countries, with the highest degrees of integration being found in the case of loans to NFCs. As for the multivariate results, evidence of fractional cointegration is found in the case of Germany, France, Spain, and Italy. The estimated speed of adjustment of loans in the restricted models (based on appropriate overidentifying restrictions) is either very close to (Germany and Spain) or slightly slower (France and Italy) than the estimates reported in previous studies based on standard CVAR models at both the national and European level (see, e.g., Calza et al. 2003, Calza et al. 2006, Levieuge 2017). Regarding the factors driving loans to NFCs in the long run, the evidence varies across countries, with DEBT being the only determinant in the case of Germany, RGDP playing the main role in the case of France, and a range of factors being relevant in the case of both Italy and Spain. In particular, the cost of bank lending relative to other sources of financing clearly matters, and thus policy measures affecting either can have an impact on loans.

Future research should estimate systems with a wider set of variables capturing both the demand and supply side as well as the impact of policy measures (Gambacorta and Rossi 2010). Moreover, structural breaks should be investigated using a variety of methods allowing for a gradual evolution or sudden shifts of the parameters, and nonlinearities reflecting the existence of thresholds should also be examined.

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Appendix

Tables

Table 1: Descriptive statistics of the variable used in the integration and cointegration analysis

FRANCE	Mean	Median	Min	Max	St.dev	Skewness	Kurtosis
LOANS	3886.590	3699.860	2625.900	5402.100	863.996	0.287	1.898
RGFCF	116321.200	114094.100	98549.580	143627.700	10514.050	0.618	2.765
RGDP	515267.200	513036.000	448128.600	581943.000	32302.950	0.078	2.339
DIFF	0.596	0.585	-0.380	1.930	0.674	0.189	1.720
DEBT	2.272	2.389	-0.305	4.690	1.605	-0.145	1.487
GERMANY	**Mean**	**Median**	**Min**	**Max**	**St.dev**	**Skewness**	**Kurtosis**
LOANS	3714.445	3613.115	3440.240	4264.580	238.765	1.068	2.832
RGFCF	135932.900	137145.500	97610.150	164697.100	16375.620	-0.258	2.399
RGDP	675497.900	673721.300	576658.900	757540.100	53016.200	-0.049	1.806
DIFF	0.137	0.145	-0.650	1.180	0.425	0.359	2.626
DEBT	1.900	1.690	-0.607	4505	1.707	0.056	1.468
ITALY	**Mean**	**Median**	**Min**	**Max**	**St.dev**	**Skewness**	**Kurtosis**
LOANS	1885.446	1958.985	0.000	2164.050	349.938	-3.916	21.268
RGFCF	78893.290	77708.200	59923.640	96631.590	9627.918	0.073	1.885
RGDP	400654.900	401645.300	336892.800	431420.300	15555.780	-0.765	5.229
DIFF	0.209	0.170	-1.020	1.380	0.513	0.072	2.790
DEBT	3.342	3.833	0.616	6.537	1.472	-0.225	2.019
SPAIN	**Mean**	**Median**	**Min**	**Max**	**St.dev**	**Skewness**	**Kurtosis**
LOANS	2019.355	1917.875	1324.090	2612.160	3430.205	0.245	2.116
RGFCF	5.61E + 10	5.51E + 10	4.39E + 10	7.24E + 10	7.91E + 09	0.509	2.430

RGDP	269158.500	268978.500	230586.300	304440.000	17008.606	-0.018	2.572
DIFF	0.153	0.070	-0.820	1.380	0.377	0.943	4.855
DEBT	3.038	3.681	0.078	6.795	1.711	-0.159	1.874

Note: The table shows the descriptive statistics for all the covariates employed in the analysis (Loans to Non-Financial Corporations (LOANS), real gross capital fixed formation (RGFCF), real GDP (RGDP), the differential between the cost of borrowing for new long- and short -term loans in the Euro Area (DIFF), and the cost of debt, security and equity issuance (DEBT)). The sample ranges from 2003Q1 to 2022Q4.

Table 2: Long Memory-Fractional Integration Analysis

RGDP	**No regressors**	**With Intercept**	**With Intercept and Trend**
Germany	0.96 (0.84,1.10)	0.65 (0.59,0.69)	$\mathbf{0.36^{*}}$ $(\mathbf{0.31}, \mathbf{0.46})$
France	0.83 (0.77,0.94)	0.62 (0.60,0.74)	$\mathbf{0.41^{*}}$ $(\mathbf{0.38}, \mathbf{0.49})$
Italy	0.96 (0.83,1.11)	0.44 (0.34,0.55)	$\mathbf{0.44^{*}}$ $(\mathbf{0.37}, \mathbf{0.58})$
Spain	0.84 (0.76,0.97)	0.57 (0.53,0.67)	$\mathbf{0.56^{*}}$ $(\mathbf{0.47}, \mathbf{0.72})$
RGFCF	**No regressors**	**With Intercept**	**With Intercept and Trend**
Germany	0.77 (0.64,0.92)	0.44 (0.38,0.53)	$\mathbf{0.22^{*}}$ $(\mathbf{0.17}, \mathbf{0.33})$
France	0.80 (0.72,0.88)	0.66 (0.55,0.68)	$\mathbf{0.46^{*}}$ $(\mathbf{0.40}, \mathbf{0.55})$
Italy	0.91 (0.80,0.99)	$\mathbf{0.57^{*}}$ $(\mathbf{0.49}, \mathbf{0.67})$	0.59 (0.49,0.69)
Spain	1.07 (0.94,1.21)	$\mathbf{1.13^{*}}$ $(\mathbf{1.04}, \mathbf{1.22})$	1.16 (1.04,1.23)
LOANS	**No regressors**	**With Intercept**	**With Intercept and Trend**
Germany	1.34 (1.28,1.37)	1.34 (1.28,1.44)	$\mathbf{1.41}$ $(\mathbf{1.28}, \mathbf{1.48})$
France	1.33 (1.26,1.37)	1.39 (1.27,1.47)	$\mathbf{1.36}$ $(\mathbf{1.29}, \mathbf{1.45})$
Italy	1.39 (1.31,1.39)	1.36 (1.33,1.41)	$\mathbf{1.39}$ $(\mathbf{1.30}, \mathbf{1.38})$
Spain	$\mathbf{1.60^{*}}$ $(\mathbf{1.55}, \mathbf{1.67})$	1.60 (1.52,1.64)	1.62 (1.54,1.67)
DIFF	**No regressors**	**With Intercept**	**With Intercept and Trend**
Germany	1.01 (0.84 ,1.22)	$\mathbf{1.00}$ $(\mathbf{0.82}, \mathbf{1.23})$	1.0 (0.83,1.23)
France	1.01 (0.84 ,1.25)	$\mathbf{1.11}$ $(\mathbf{0.91}, \mathbf{1.39})$	1.11 (0.92, 1.39)
Italy	0.65 (0.50 ,0.87)	$\mathbf{0.64}$ $(\mathbf{0.49}, \mathbf{0.85})$	0.67 (0.55,0.85)
Spain	0.30 (0.19 ,0.45)	$\mathbf{0.30}$ $(\mathbf{0.20}, \mathbf{0.45})$	0.31 (0.20,0.46)

DEBT	No regressors	With Intercept	With Intercept and Trend
Germany	1.20 (1.04 ,1.40)	**1.17 (0.98, 1.45)**	1.17 (0.97, 1.45)
France	1.20 (1.04 ,1.41)	**1.19 (0.99, 1.45)**	1.19 (0.99, 1.44)
Italy	1.23 (1.07 ,1.42)	**1.25 (1.05, 1.50)**	1.25 (1.05, 1.50)
Spain	1.18 (0.98 ,1.36)	**1.14 (0.99, 1.33)**	1.14 (0.99, 1.34)

Note: This table reports the estimated values of d with their corresponding confidence intervals in brackets for each of the specifications considered (Loans to Non-Financial Corporations (LOANS), real gross capital fixed formation (RGFCF), real GDP (RGDP), the differential between the cost of borrowing for new long- and short -term loans in the Euro Area (DIFF), and the cost of debt, security and equity issuance (DEBT)). The sample ranges from 2003Q1 to 2022Q4. The values in bold are those for the models selected on the basis of the statistical significance of the regressors.

Table 3: Fractional cointegration analysis for France

Likelihood Ratio Tests				
Rank	d	b	LR statistic	P-value
0	0.766	0.766	53.378	0.005
1	0.780	0.780	20.036	0.358
2	1.149	1.149	8.372	0.882
3	1.272	1.272	1.756	0.880
4	1.193	1.193	...	...
White Noise Tests				
Variable	Q	P-value	LM	P-value
All variables	22.290	0.899	-	-
LOANS	3.121	0.210	4.260	0.119
RGDP	0.047	0.977	0.025	0.987
RGFCF	0.784	0.676	0.877	0.645
DIFF	0.241	0.887	1.684	0.431

Value for $d = b$		
	estimate	s.e.
	0.736	0.082

Note: The upper part of the table shows the results for the Trace LR statistic estimated for $d = b$, together with the relevant probability values. The middle part shows the white noise test results on the residuals of the model, with a null of absence of residual autocorrelation. The lower part of the table shows the final estimate for the fractional parameters and their statistical precision. The variables employed are Loans to Non-Financial Corporations (LOANS), real GDP (RGDP), the differential between the cost of borrowing for new long- and short -term loans in the Euro Area (DIFF), and the cost of debt, security and equity issuance (DEBT). The sample ranges from 2003Q1 to 2022Q4.

Table 4: Hypothesis testing for France

	H_1^{FCvsC}	$H_2^{\beta^{DIFF}}$	$H_3^{\alpha^{GDP}}$
df	1	1	1
LR	5.504	2.855	0.686
P-value	0.019	0.091	0.408

Note: The table shows the results for the Likelihood Ratio tests. The first column compares the likelihood of an FCVAR against a canonical CVAR. The subsequent columns show all the LR tests for the α and β parameters where the 0 null could not be rejected. The variables employed are Loans to Non-Financial Corporations (LOANS), real GDP (RGDP), the differential between the cost of borrowing for new long- and short -term loans in the Euro Area (DIFF), and the cost of debt, security and equity issuance (DEBT). The sample ranges from 2003Q1 to 2022Q4.

Table 5: Fractional cointegration analysis for Germany

Likelihood Ratio Tests				
Rank	d	b	LR statistic	P-value
0	0.964	0.964	87.712	0.000
1	0.518	0.518	45.218	0.000

2	0.387	0.387	27.440	0.000
3	0.151	0.151	1.427	0.232
4	0.183	0.183	-	-
White Noise Tests				
Variable	Q	P-value	LM	P-value
All variables	11.826	0.756	-	-
LOANS	0.297	0.597	0.194	0.660
RGDP	0.000	0.984	0.000	0.986
DIFF	0.677	0.411	0.331	0.565
DEBT	1.197	0.274	1.670	0.196
Values for $d = b$				
	estimate	s.e.		
	0.517	0.053		

Note: The upper part of the table shows the results for the Trace LR statistic estimated for $d = b$, together with the relevant probability values. The middle part shows the white noise test results on the residuals of the model, with a null of absence of residual autocorrelation. The lower part of the table shows the final estimate for the fractional parameters and their statistical precision. The variables employed are Loans to Non-Financial Corporations (LOANS), real GDP (RGDP), the differential between the cost of borrowing for new long- and short -term loans in the Euro Area (DIFF), and the cost of debt, security and equity issuance (DEBT). The sample ranges from 2003Q1 to 2022Q4.

Table 6: Hypothesis testing for Germany

	H_1^{FCvsC}	$H_2^{\beta^{DIFF}}$	$H_3^{\alpha^{DIFF}}$
df	1	1	1
LR	24.111	0.179	0.050
P-value	0.000	0.091	0.823

Note: The table shows the results for the Likelihood Ratio tests. The first column compares the likelihood of an FCVAR against a canonical CVAR. The subsequent columns show all the LR tests for the α and β parameters where the 0 null could not be rejected. The variables employed are Loans to Non-Financial Corporations (LOANS), real GDP (RGDP), the differential between the cost of borrowing for new long- and short -term loans in the Euro Area (DIFF), and the cost of debt, security and equity issuance (DEBT). The sample ranges from 2003Q1 to 2022Q4.

Table 7: Fractional cointegration analysis for Italy

Likelihood Ratio Tests				
Rank	d	b	LR statistic	P-value
0	0.803	0.803	114.333	0.000
1	0.655	0.655	67.491	0.000
2	0.304	0.304	20.127	0.017
3	0.010	0.010	4.447	0.349
4	0.295	0.295	0.081	0.776
5	0.309	0.309	-	-
White Noise Tests				
Variable	Q	P-value	LM	P-value
All variables	24.695	0.480	-	-
LOANS	0.311	0.577	0.726	0.394
RGDP	2.055	0.152	0.486	0.485

RGFCF	1.224	0.269	0.624	0.430
DIFF	0.000	0.996	0.009	0.925
DEBT	0.165	0.685	0.113	0.737
Values for $d = b$				
	estimate		s.e.	
	0.655		0.005	

Note: The upper part of the table shows the results for the Trace LR statistic estimated for $d = b$, together with the relevant probability values. The middle part shows the white noise test results on the residuals of the model, with a null of absence of residual autocorrelation. The lower part of the table shows the final estimate for the fractional parameters and their statistical precision. The variables employed are Loans to Non-Financial Corporations (LOANS), real gross capital fixed formation (RGFCF), real GDP (RGDP), the differential between the cost of borrowing for new long- and short -term loans in the Euro Area (DIFF), and the cost of debt, security and equity issuance (DEBT). The sample ranges from 2003Q1 to 2022Q4.

Table 8: Hypothesis testing for Italy

	H_1^{FCvsC}	$H_2^{\beta^{GDP}}$	$H_3^{\alpha^{\mathrm{DIFF}}}$
df	1	1	1
LR	19.373	1.067	0.291
P-value	0.000	0.302	0.590

Note: The table shows the results for the Likelihood Ratio tests. The first column compares the likelihood of an FCVAR against a canonical CVAR. The subsequent columns show all the LR tests for the α and β parameters where the 0 null could not be rejected. The variables employed are Loans to Non-Financial Corporations (LOANS), real gross capital fixed formation (RGFCF), real GDP (RGDP), the differential between the cost of borrowing for new long- and short -term loans in the Euro Area (DIFF), and the cost of debt, security and equity issuance (DEBT). The sample ranges from 2003Q1 to 2022Q4.

Table 9: Fractional cointegration analysis for Spain

Likelihood Ratio Tests				
Rank	d	b	LR statistic	P-value
0	0.804	0.804	105.667	0.000

1	0.632	0.632	55.122	0.000
2	0.541	0.541	37.969	0.000
3	0.392	0.392	23.278	0.000
4	0.010	0.010	0.055	0.814
5	0.010	0.010	-	-
White Noise Tests				
Variable	Q	P-value	LM	P-value
All variables	14.621	0.958	-	-
LOANS	0.890	0.345	0.721	0.396
RGDP	0.048	0.826	0.011	0.916
RGFCF	0.691	0.406	0.097	0.755
DIFF	0.014	0.907	0.010	0.921
DEBT	0.123	0.726	0.107	0.744
Values for $d = b$				
	estimate		s.e.	
	0.633		0.048	

Note: The upper part of the table shows the results for the Trace LR statistic estimated for $d = b$, together with the relevant probability values. The middle part shows the white noise test results on the residuals of the model, with a null of absence of residual autocorrelation. The lower part of the table shows the final estimate for the fractional parameters and their statistical precision. The variables employed are Loans to Non-Financial Corporations (LOANS), real gross capital fixed formation (RGFCF), real GDP (RGDP), the differential between the cost of borrowing for new long- and short -term loans in the Euro Area (DIFF), and the cost of debt, security and equity issuance (DEBT). The sample ranges from 2003Q1 to 2022Q4.

Table 10: Hypothesis testing for Spain

	H_1^{FCvsC}	$H_2^{\beta^{DEBT}}$	$H_3^{\alpha^{DEBT}}$
df	1	1	1
LR	18.571	0.125	0.159
P-value	0.000	0.724	0.690

Note: The table shows the results for the Likelihood Ratio tests. The first column compares the likelihood of an FCVAR against a canonical CVAR. The subsequent columns show all the LR tests for the α and β parameters where the 0 null could not be rejected. The variables employed are Loans to Non-Financial Corporations (LOANS), real gross capital fixed formation (RGFCF), real GDP (RGDP), the differential between the cost of borrowing for new long- and short -term loans in the Euro Area (DIFF), and the cost of debt, security and equity issuance (DEBT). The sample ranges from 2003Q1 to 2022Q4.

Figures

Figure 1: Time series for France.

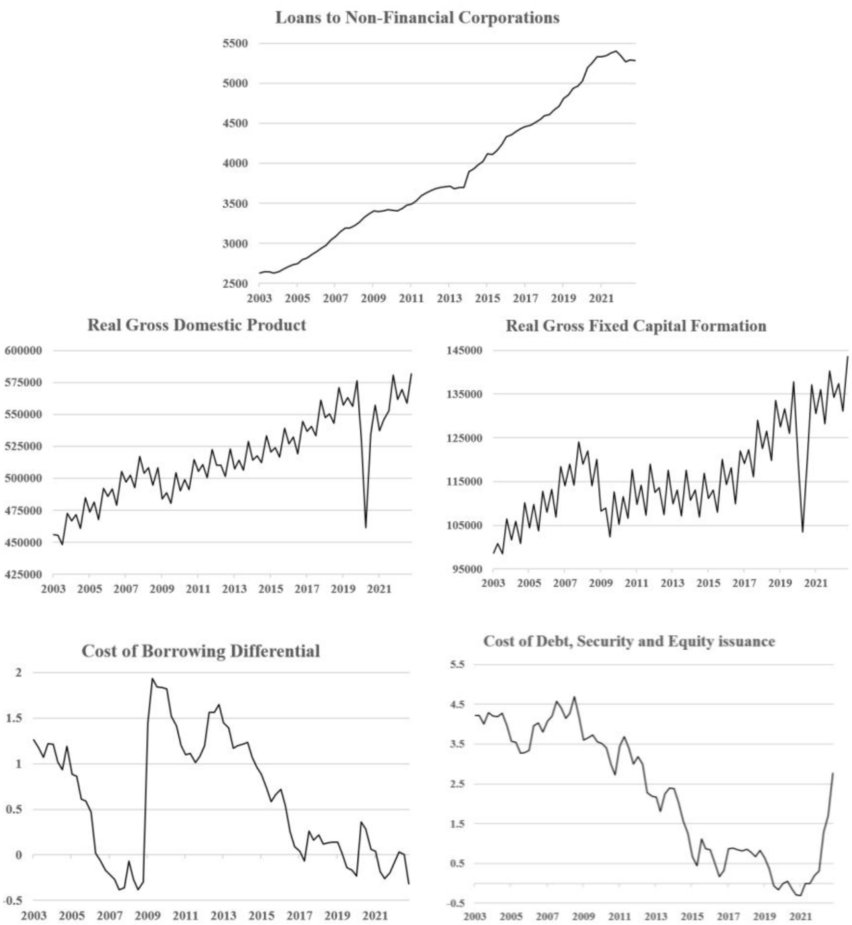


Figure 2: Time series for Germany.

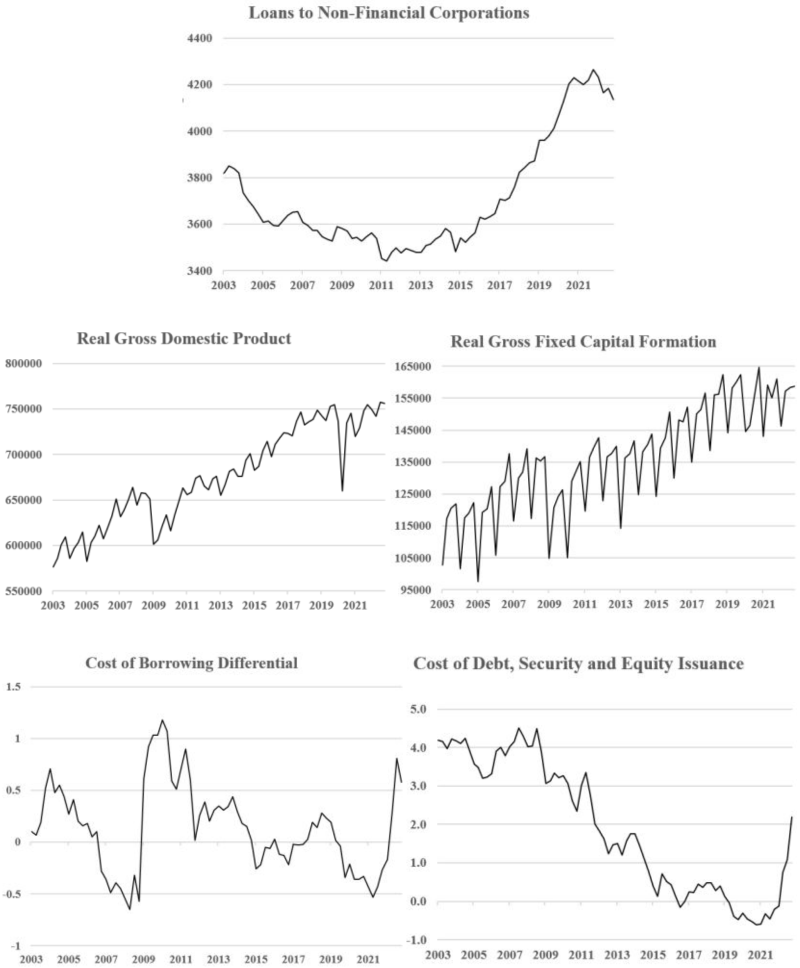
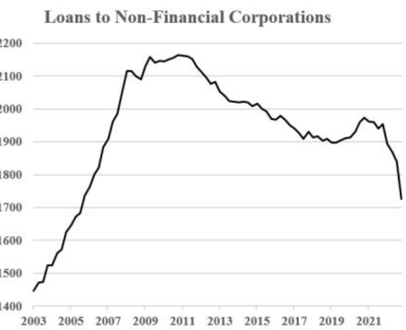


Figure 3: Time series for Italy.



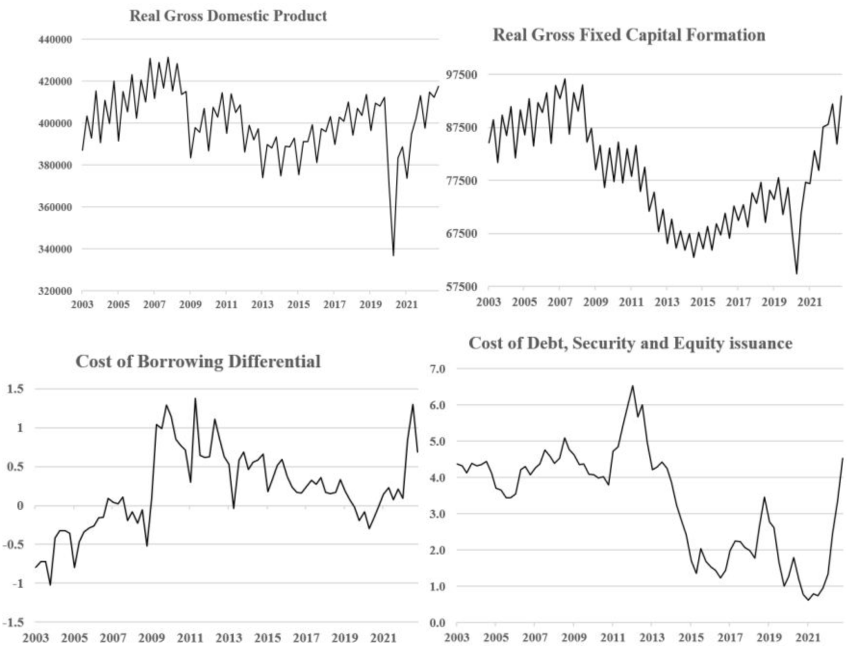
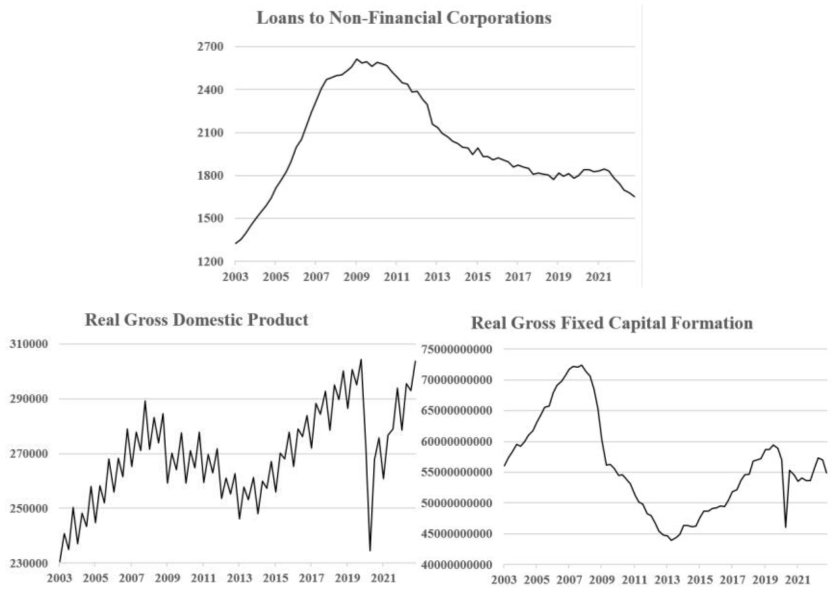
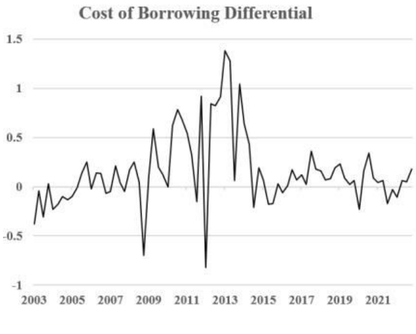
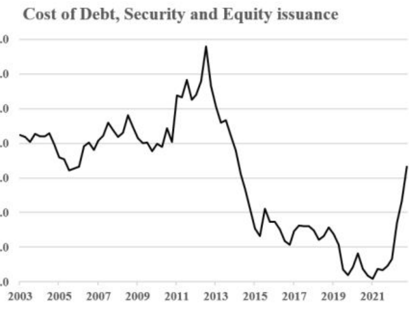


Figure 4: Time series for Spain.




Cost of Borrowing Differential
1.5
1
0.5
0
-0.5
-1
2003
2005
2007
2009
2011
2013
2015
2017
2019
2021


Cost of Debt, Security and Equity issuance
7.0
6.0
5.0
4.0
3.0
2.0
1.0
0.0
2003
2005
2007
2009
2011
2013
2015
2017
2019
2021

Chapter 5: Nonlinearities and Potential Fractional Equilibria in the Spot/Future Oil Prices Relationship

In this paper, we examine the long-run relationship between spot and future oil prices by considering nonlinearity, a potentially non-integer degree of integration, and allowing both canonical and fractional cointegration methods to compete for the best data fit. We also apply an ex-ante unit root test with endogenous break detection and a pre-break and post-break approach to compare canonical and fractional vector error correction models. Bridging a gap between nonlinear and fractional applications, this methodology uniquely detects nonlinearities in the long-run equilibrium without resorting to state-space or threshold models. Using weekly observations of spot and future prices from the U.S. Energy Information Administration, our estimates identify a structural break around 2004. The canonical contango situation detectable across the whole series breaks down upon examining the data's memory processes before and after this break. Our results also reveal a significant shift in the equilibrium model linking spot to future oil prices due to a global exogenous shock, suggesting two different, time-dependent equilibria: a backwardation equilibrium before the break and a near-unity contango after. This evidence demonstrates that global shocks can affect the spot-futures oil price relationship, inducing equilibrium switches.

5.1. Introduction

Oil is one of the most important sources of energy in the world, so oil price and strategic decisions regarding oil affect the global economy. Moreover, trying to explain the relationship between spot and future prices of oil has been and still remains a topic of great interest in financial economics (Zavadska et al., 2018). In this sense, this long-run relation between spot and futures prices of oil provides valuable information about the current and future conditions of the oil market, as well as opportunities for investors and market actors to make well-informed decisions (Ellwanger and Snudden, 2023; Yu et al., 2023).

Market actors are especially interested in knowing if the relationship between spot and futures could be classified as a situation of contango, where spot prices are higher than future prices, or backwardation, the opposite situation. The canonical situation, regardless of the market of interest and of the short or long span targets of any empirical analysis, would be contango since that would necessarily imply that the analyzed market is well stocked (Considine and Aldayel, 2020). However, during the last two decades the oil market has been affected by three important shocks: the Covid-19 pandemic, the oil price war in the Organization of Petroleum Exporting Countries (OPEC) and the Russia-Ukraine war. Greater volatility movements have been observed during these stressful periods, with both spot and futures markets of West Texas Intermediate (WTI, hereafter)[1] playing a significant role in the transmission of volatility (Apostolakis et al., 2024).

[1] In this context, it is pertinent to define the two primary world reference point oil prices: West Texas Intermediate (WTI) and Brent Blend (see Ederington et al., 2019), using the first in our study.

The relationship between oil spot and futures prices has already been studied using various methodologies, including causality and cointegration, employing different tests and yielding varied results (see, e.g., Bekiros and Diks, 2008; Huang et al., 2009; Chang, and Lee, 2015; Holmes and Otero, 2019). This paper contributes to the studies on the long run relationship between spot and future crude oil prices by evaluating the possibility of either a suitable cointegrated VAR or fractionally cointegrated VAR, and taking into account possible nonlinearities by assessing endogenously the potential existence of deterministic, historically identifiable breaks in the spot prices through unit root testing with endogenous break retrieval. Historical literature on the topic has long since focused on either a nonlinear or a tentative fractional representation of the relationship. As a first ever contribution to this methodological gap, we bridge the two approaches by focusing ex-ante on nonlinearities through unit root analysis, embodying a possible break in the alternative model, and then developing a fractional model based on the results of the outcome of the former analysis.

By using fractional integration, we allow a much richer dynamic modelling specification in the series, and the differencing parameter can be taken as an indicator of the degree of persistence of the data. Through the FCVAR (Fractionally Cointegrated Vector Autoregressive) approach, we investigate the existence of a long run equilibrium relationship between the two variables. In traditional cointegration models, the individual time series are expected to be nonstationary $I(1)$, while a linear combination of the two is expected to be stationary $I(0)$. The FCVAR model is more flexible and permits that the degree of differentiation of the series may be $I(d)$, so that the variables are cointegrated of order $d - b$, with $b > 0$, where d and b may be real values.

By testing whether a traditional or a fractional cointegrated model better suits the data and the underlying theoretical model, we thus not only test the model hypotheses in the long run but were also able to determine to what extent the average speed of adjustment to the equilibrium of the chosen variable after a shock would converge at a generally faster (CVAR) or lower (FCVAR) speed. Furthermore, as a preliminary inspection of the series gave us a clear indication of the possible existence of a jump in the mean of the series, we choose to run a series of unit root tests, which would allow us to contemporaneously test for a unit root null hypothesis and detect endogenously the existence of a possible break in the mean (Perron and Vogelsang, 1992). When a break in the mean is detected, we split the sample in two different regimes, to check whether and by what degree the long run equilibria and adjustment speeds change.

First, our analysis addresses, to our knowledge for the first time, an ex-ante analysis of linearity in the spot-futures series by considering potential endogenous breaks as an alternative to the null hypothesis of a unit root. Previous literature has highlighted the high persistence of energy price series (Ozdemir et al., 2013), but until now, it has relied on state-space applications (Allen et al., 2018) or threshold modelling (Wang and Wu, 2013) within a canonical error correction mechanism, disregarding the benefits of combining preliminary break analysis with a fractionally cointegrated specification.

Second, in our application, we choose not to predetermine the degree of cointegration for the relationship by selecting an a-priori adequate model for spot and forward prices. Instead, we allow the canonical CVAR and FCVAR models to compete, letting the data "speak for themselves." Although this decision might seem straightforward, past literature has generally neglected choices based on the goodness of fit of competing long-run models, typically opting for either CVAR modelling (Figuerola-Ferretti and Gonzalo, 2010) or FCVAR modelling (Rossi and Santucci-De Magistris, 2013) when estimating the long-run elasticities and adjustment speeds of commodity and currency markets.

Finally, estimating an a-priori break and dividing the whole series into two sub-samples allow us to estimate directly linear, fractionally, or non-fractionally cointegrated branches of the overall series without the concern of introducing potential nuisance parameters (such as breaks in intercepts, trends, or other deterministic terms in the specification of the cointegrating vector). This approach also avoids threshold modelling, which would theoretically require at least a weakly exogenous threshold variable to create regime branches, thus detaching ourselves from the theoretical model presented in Section 2.

Our results highlight that the long-term relationship between oil spot and futures prices is characterized by a contango situation, but point out that such an overall steady state condition, once a time-consistent (deterministic) exogenous break has been taken into account, represents the result of the linear combination of two different states: a pre-break steady state, characterized by a slight degree of backwardation, and a post-break state, characterized by a weak long run contango equilibrium which, after testing, would statistically not suggest either the presence of contango or of backwardation after the identified exogenous shock has taken place.

Our study is perhaps a first in time application, given the considerable length of the series. Comparing the first branch of studies, Huang et al. (2009) obtained similar results as they worked with three periods between 1986 and 2007, finding backwardation in the first two periods (up to 2001) and contango in the third period. Dolatabadi, Narayan, Nielsen & Xu, (2018) analyzed 17 countries between 1983 and 2012 for different alternative commodities, finding again backwardation. Earlier, the same authors (Dolatabadi, Nielsen & Xu, 2016) looked for the same occurrence in metal commodities in an FCVAR setup, this time finding no conclusive evidence on either contango or backwardation. The rest of the notable literature (see, e.g., Vides, Golpe & Iglesias, 2020, on price differentials or Wu, Xu, Zheng, & Chen, 2021, on contango/backwardation in the bitcoin markets) would either target a different objective variable, financial definition or simply be incomparable given the time mismatch in the series analyzed.

The remainder of the paper is organized as follows: Section 2 presents the literature review; Section 3 describes the data and the methodology; Section 4 is devoted to the empirical results, while Section 5 concludes the paper.

5.2. Literature Review

Crude oil is one of the most important natural resources, the main source of non-renewable energy and the commodity with the biggest impact on the world. Crude oil and its derivatives are used in different economic activities such as transportation or industrial production. The fluctuations in the price of crude oil have notable impacts on economies and financial markets (Jia et al., 2017). These fluctuations depend, largely, on demand and supply changes, economic cycles (Yao et al., 2017), the global financial crises and their aftermath, and military conflicts and political events (Monge et al., 2017). Thus, it is essential to central banks, policy institutions and oil-related industries, to know how oil prices evolve across time and how they are related (Bai et al., 2020).

In the crude oil market, we differentiate two kind of prices that define two types of markets, spot (the negotiation of the sale price and quantity of crude oil and the exchange of the latter, take place at the same time, immediately) and futures (the contract for purchase or sale at a future date, agreeing in the present on the price quantity and maturity date). Investigating the relationship between spot and futures prices is crucial for understanding the interaction dynamics between these two markets. (Shao et al., 2019). This relationship may give rise in the long term to two different situations. Under typical market conditions, when the crude oil market is balanced, prices usually exhibit contango. In this state, futures prices are higher than spot prices, reflecting the costs associated with storing the commodity, including warehousing expenses, foregone interest, and a convenience yield on inventories. Conversely, when futures prices are lower than spot prices, the market is in backwardation. This indicates a situation where immediate demand outweighs the benefits of holding inventories, leading to higher current prices compared to future prices (Considine et al., 2021).

Giving value to the use of time series techniques on this topic, there exists an extensive literature that relates spot and futures prices of different commodities.[2] One of the first pioneering works on this topic was that of Silvapulle and Moosa (1999). They consider the relationship between spot and futures oil prices using linear and non-linear Granger causality tests, concluding that spot and futures markets react to new information. Another way to relate spot and future prices commonly employed in the first generations of studies tackling this topic is based on the lead-lag methodology. This approach is followed by Brooks et al. (2001), who use spot index and the contract for the FTSE 100. They employ the cointegration and error correction model (ECM, hereafter) of Engle and Granger (1987), determining that spot returns are led by futures returns. Likewise, Floros and Vougas (2007) studied the lead-lag relationship between spot and futures traded in the Athens Derivatives Exchange (ADEX). They found that spot and futures are cointegrated and a lead-lag relationship exists between the two. Kaufmann and Ullman (2009) also assessed spot and futures oil prices, using the two-step dynamics ordinary least squares (DOLS) of Stock and Watson (1993) and the vector error correction model (VECM, hereafter) of Johansen (1988) and Johansen and Juselius (1990), finding out that sometimes innovations appearing on Dubai-Fateh spot prices and other times on WTI spread to other spot and futures prices. Ozdemir et al. (2013) analysed the persistence in crude oil spot and futures prices using the Hansen (1999) grid-bootstrap procedure. They concluded that the degree of persistence in Brent spot and future prices series of crude oil is high if structural breaks are disregarded, and not very persistent in the opposite case. Moreover, Fernandez (2016) also used cointegration and causality test to relate spot and future prices, in this case with daily observations prices of aluminium, copper, lead, nickel, tin and zinc. Their analysis shows that correlations are not always higher during contango than during backwardation. Chang and Lee (2015) used wavelet transformations and employed spot and future prices of crude oil to study their movements over time. They concluded that if the maturity of the contracts increases, the variability of futures prices and the speed of adjustment of futures/spot prices differentials increase too. However, Holmes and Otero (2019), replicating and eventually expanding upon these latter results, added that futures prices do not respond to long-run disequilibrium and that short-run causality is also present at the longer maturities. Pradhan et al. (2021) analysed the relation between spot and futures prices of aluminium, copper, crude oil, gold, nickel, and silver employing the autoregressive distributed lag (ARDL) bounds-testing technique. They found a cointegration relationship between the spot and futures prices of these commodities, while the causality analysis discloses uni-directional causality in the long run from the spot to the futures price for aluminium and copper, and both bi-directional and uni-directional causality in the short run between the two prices for aluminium, copper, gold, and silver.

Within these prices' relationships, some authors focus on those which deal with the existence of contango or backwardation. Examples are Huang, Yang and Hwang (2009), Botterud et al. (2010), Gulley and Tilton (2014), Fernandez (2016), among others. Nonetheless, the long-run relationship between prices of commodities may contain additional information apart from the existence of contango or backwardation. In this regard, Bekiros and Diks (2008) studied cointegration, linear and non-linear causality and non-causality to analyse the relationship between crude oil spot and futures prices. They concluded that spot and futures returns can reveal asymmetric GARCH effects and/or statistically significant higher order conditional moments. Lee and Zeng (2011) employed quantile cointegration techniques and found out that when any shock is applied on 1-month futures oil prices, high spot prices are more affected than low spot prices. Wang and Wu (2013) used cointegration and the non-linear threshold vector error correction model (TVECM) by Balke and Fomby (1997), observing that the relationships between spot and future prices are not equal in the short-term and in the long-term. Allen, Chang, McAleer and Singh (2018) worked with spot and futures prices of agricultural commodities (corn, sugar, and

[2] For an exhaustive chronological list of the most recent literature focusing on the backwardation/contango hypotheses across an extensive array of different commodities through cointegration techniques, see Table 4 in the Appendix of this article.

wheat), ethanol and crude oil. They made use of cointegration and partial cointegration techniques and found evidence that a long run relationship should indeed exist in some of the markets reviewed. Another way to look at cointegration is presented by the Park and Hahn (1999) test. This is implemented by Hu, Hou and Oxley (2019), who worked with spot and futures markets of bitcoin, and-conclude that cointegration exists between those two variables. In a similar fashion, Lee, Meslmani and Switzer (2020) also investigated the relation between spot and futures prices of bitcoin. They found a long-run relationship between spot and futures and conclude that spot prices are predicted by futures. Linking our reasoning to the literature we have already mentioned in the preceding pages, we do not aim to confute the results from past literature, which would linger from evidence of backwardation to complete equilibrium (i.e., rejection of both the backwardation and the contango states) but rather offer a solution that explains systematically, merging nonlinear studies with conventional cointegrating and fractional ones, the different streams of results obtained up until now.

As we mentioned in the introduction, this paper employs a CVAR/FCVAR mixed modelling strategy which follows an analysis of possible structural breaks which allows us to divide our series into two independent regimes.[3] We follow the Johansen's (2008a, 2008b) methodology, also developed by Johansen and Nielsen (2012). In this regard, Gil-Alana, Yaya and Awe (2017) studied the fractional cointegration between gold and oil prices, finding a long memory and mean reverting, with an order of cointegration of around 0.46, a result implying that the convergence of these variables is slower compared to the results from standard cointegration methods. Furthermore, Dolatabadi, Nielsen, and Xu (2015, 2016) and Dolatabadi, Nayaran, Nielsen, and Xu (2018) arguably found out that the FCVAR model has more advantages over the CVAR model and showed that when they analyse the price discovery process for commodity markets, spot prices are slightly more critical than futures. All this follows Figuerola-Ferretti and Gonzalo (2010), who performed the same study, but employing a canonical CVAR model. The above-mentioned authors adapted this theoretical model to the FCVAR to examine price discovery and the existence of contango and backwardation in a fractional environment. Rossi and Santucci de Magistris (2013) used spot and futures of S&P500 and the FVECM of Johansen (2008a) to analyse the dynamic relationship between those prices. They concluded that it is important to incorporate the long run equilibrium to obtain the best results. FCVAR has also been used by Vides et al. (2020) and Wu et al. (2021) among many others.

A last, important branch of the literature we need to address explores structural breaks as a prime source of non-linearity in commodity markets. It is well known how exogenous events can shift oil price co-movements. Just to mention a few examples, Maslyuk and Smyth (2008) fitted a deterministic break into the cointegrating equilibrium between crude oil spot and futures with different maturities in a Engle-Granger based framework, checking for differences in the speed of convergence of spot prices with respect to a residual error correction term. Arouri et al. (2012) found that GARCH models nested with endogenously retrieved structural breaks have greater forecasting accuracy of oil prices compared to linear models. Lastly, Noguera (2013) identified up to 12 shifts in real oil prices across the 1961–2011 period.

Looking at the short term, another source of potential (short run) non-linearity would be the continuous regime shifts between price contango and price backwardation. As we will clarify illustrating the theoretical model in Section 3, when a discount rate and a storage cost is assumed, contango stands as an almost physiological condition, and thus futures prices would normally be higher than spot prices. Explanations such as Pindyck (2001) theory of real options can explain, at least in the short term, periods of backwardation. In general, the underlying determinants of energy prices and their respective correlations can change greatly when the relationship between

[3] A CVAR or FCVAR are chosen alternatively based on the goodness of fit of the resulting model.

spot and futures switches from contango to backwardation (see, for instance, Nomikos and Pouliasis, 2011), affecting forecasting and elasticities in the short run.

Based on previous empirical literature, our study bridges a significant gap in understanding the impact of non-linearity on the long-run relationship between spot and futures prices. Our approach includes a preliminary univariate analysis to identify the most statistically significant break in the series, followed by a multivariate cointegration strategy to confirm the shock's relevance and magnitude in altering the equilibrium. If no regime shift between contango and backwardation is detected after the shock, we conclude that exogenous regime shifts are incapable of significantly affecting storage costs or discount rates to alter the cointegrating relationship between spot and future prices. Thus, non-linearities in the spot-futures relationship are predominantly influenced by short-term changes in underlying fundamentals.

5.3. Data and Methodology

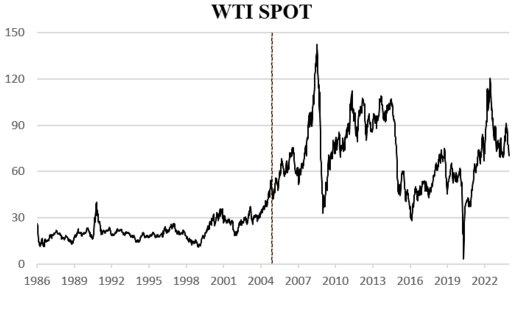
5.3.1. Data

Our data comprises weekly observations from the U.S. Energy Information Administration, including spot and futures prices of WTI crude oil from January 3, 1986 to December 15, 2023. (https://www.eia.gov). The sample period has 1981 observations. For the futures series, we must point out that four types of contracts were selected according to the delivery date; Contract 1 expires on the third business day prior to the 25th calendar day of the month preceding the delivery month. If the 25th calendar day of the month is a non-business day, trading ceases on the third business day prior to the business day preceding the 25th calendar day. After a contract expires, Contract 1 for the remainder of that calendar month is the second following month. Contracts 2 - 4 represent the successive delivery months following Contract 1.
The plots of the series are displayed across Figure 1. We observe a similar evolution of spot and future prices, showing a change in trend of prices that was more stable until the Iraq War broke out (March 2003), after which a significant increase in oil prices is clearly visible, starting from around 2004. Since then, crude oil prices have been affected by different recessions and booms which have caused high increases and descents in those prices.

TABLES 1 – 3 ABOUT HERE

In Tables 1, 2, and 3, the differences in the values of the descriptive statistics for the total series and the two segments before and after 2004 can be observed, respectively. We notice a sharp increment of the mean value of the series, which pivots from around 21 to around 71 across the two samples, with a net increment in precision given by a sharp decrease of both standard deviations of the two branches with respect to the standard deviation of the whole series relative to their mean. The preliminary visual analysis of Figure 1, together with the stark descriptive changes in Tables 1 to 3, suggest the necessity to explore the order of integration of the series allowing for a possible break (albeit deterministic) to enter the series.

Figure 1. Time series plots



5.3.2. *Theoretical underpinning*

The theoretical framework adopted in this article is based on the equilibrium model of spot and futures proposed by Garbade and Silver (1983) and Figuerola-Ferretti and Gonzalo (2010). More specifically, we rely on the Dolatabadi et al. (2015) specification.

First, we define an equilibrium model of spot and futures prices with infinitely elastic arbitrage. In this model we assume that there are no taxes or transaction costs, no limitations on borrowing, only costs to finance a futures position along with storage costs and without limitations on short sale in the spot market. The equilibrium model is:

$$f_t = s_t + r_t + c_t , \quad (1)$$

where f_t denotes the log-futures price, s_t represents the log-spot prices, r_t is the continuously compounded interest rate, and c_t is the storage cost.

We make three assumptions: (1) $r_t = \overline{r} + u_{rt}$, where $\overline{r}$ is the mean of r_t and u_{rt} is a $I(0)$ process with mean zero and finite positive variance; (2) $c_t = \overline{c} + u_{ct}$, where $\overline{c}$ is the mean of c_t and u_{ct} is an $I(0)$ process with mean zero and finite positive variance; and (3) Δs_t is the first difference of log-spot price and denotes an $I(0)$ process with mean zero and finite positive variance. The model is converted to:

$$f_t - s_t = \overline{r} + \overline{c} + u_{rt} + u_{ct}. \quad (2)$$

In our case we assume standard no-arbitrage-cost conditions, where spot and futures prices have a unit root, so that s_t and f_t are $I(1)$ processes and cointegrate to $C(1,0)$ with a cointegrating vector to be estimated close to $(1, -1)$.

We also define the equilibrium model of spot and futures prices with finite elasticity and other factors that make arbitrage transactions risky. Figuerola-Ferratti and Gonzalo (2010) use the convenience yield as the main cost of arbitrage (see Kaldor, 1939, and Brennan and Schwartz, 1985). They define backwardation as "*the present value of the marginal convenience yield of the commodity inventory*". The opposite case, when this value is negative, is the situation of contango.

The equilibrium model changes with respect to the no-arbitrage condition by adding the convenience yield, y_t:

$$f_t + y_t = s_t + r_t + c_t. \tag{3}$$

The convenience yield is approximated by a linear combination of s_t and f_t as $y_t = \gamma_1 s_t - \gamma_2 f_t$, where $\gamma_i \in (0,1)$ and $i = 1, 2$. Appling this condition, the equilibrium model becomes:

$$s_t = \beta_2 f_t + \beta_3 + u_{rt} + u_{ct} \tag{4}$$

where $\beta_2 = \frac{\gamma_2 - 1}{\gamma_1 - 1}$ and $\beta_3 = \frac{\bar{r} + \bar{c}}{\gamma_1 - 1}$.
β_2 can take different values:

- β_2 >1: spot price is higher than futures price, so there is a long-run backwardation.
- β_2 <1: spot price is lower than futures price, so there is a long-run contango.
- β_2 =1: there is no contango or backwardation in the long run.

We test the conditions proposed by Dolatabadi et al. (2015): (1) $r_t = \bar{r} + v_{rt}$, where $\bar{r}$ is the mean of r_t and v_{rt} is an $I(1-b)$ process with $b > 1/2$, mean zero, and finite positive variance; (2) $c_t = \bar{c} + v_{ct}$, where $\bar{c}$ is the mean of c_t and v_{ct} is an $I(1-b)$ process with $b > 1/2$, mean zero, and finite positive variance, and (3) Δs_t denotes an $I(0)$ process with mean μ and finite positive variance. The equilibrium relationship thus becomes:

$$s_t = \beta_2 f_t + \beta_3 + v_{rt} + v_{ct} \, . \tag{5}$$

Using this framework, we allow for a fractional degree of differentiation. In the empirical application carried out in the following section, we do not impose any a priori assumption about the degree of integration of the series, permitting both the individual series and the errors to be fractionally integrated.

5.3.3. Empirical strategy

This subsection provides a concise description of the specification and various steps that make up the econometric strategy used in this work. As is common, we start employing the Augmented Dickey Fuller (ADF) test from Dickey and Fuller (1979) and the Kwiatkowski, Phillips, Schmidt and Shin (KPSS) test (1992) to detect the presence of unit roots in the series. The one-break Perron and Vogelsang (1992) test is then employed to detect the presence of a unit root and contemporaneously retrieve endogenously a possible break in the series, in accordance with the descriptive analysis from the previous paragraph. Subsequently, we test for the presence of some degree of fractional integration in the whole series and in the two autonomous branches with the log-periodogram regression method of Geweke and Porter-Hudak (1983) and Robinson (1995) and with a testing procedure developed in Robinson (1994). This would serve as both a confirmation and a refinement of the canonical unit root test results obtained in the previous test. As a last step, we test and compare the goodness of fit of two alternative models, a canonical cointegrated VAR (CVAR) and a fractional cointegrated VAR (FCVAR) as proposed by Johansen and Nielsen (2012) and Morin et al., (2021).[4] This would allow for the order of integration of the individual series to be fractional and cointegrate with an order $d–b$, with $b > 0$. Further details on the methodologies are available in the appendix.

5.4. Results

5.4.1. Unit root tests

We begin showing the results of the ADF and the KPSS tests. In Table 4, we present the results of these tests applied to the overall series. First, we check if the series are integrated of order 2 using the ADF test in first differences, rejecting the null in all cases. We then perform the ADF test for the series in levels and we are unable to reject the presence of a unit root. In the case of the KPSS test, we reject the stationarity of the series, reaching the same conclusion as with the ADF.

Table 4: Unit root tests. Whole Series

Series	**ΔADF**		**ADF**		**KPSS**	
	Test Statisti	Decision	Test Statisti	Decision	Test Statisti	Decision
SPOT	-10.174*	Not I(2)	-2.416	I(1)	6.920*	I(1)
F1	-10.070*	Not I(2)	-2.422	I(1)	6.930*	I(1)
F2	-10.131*	Not I(2)	-2.313	I(1)	7.000*	I(1)
F3	-10.246*	Not I(2)	-2.219	I(1)	7.050*	I(1)
F4	-11.068*	Not I(2)	-2.145	I(1)	7.080*	I(1)

Note: ADF denotes the Augmented Dickey-Fuller and ΔADF in the first differences. KPSS is Kwiatkowski-Phillips Schmidt-Shin. The sample size is 1981. * Null rejection at 5% level

From the literature, we can deduce that there is a change in prices in late 2004. Therefore, we apply the test proposed by Perron and Vogelsang (1992) to verify if indeed there is a break in the data. We find a break on November 5, 2004, for the spot price, while for futures, the break is marked on October 29, 2004, just one week earlier. The results are shown in Table 5. Evidently, even after a break is nested into the alternative hypothesis, the null hypothesis of a unit root cannot be rejected. To explore independently the persistence of the series before and after the shock, we repeat the ADF and the KPSS unit root tests for the two branches before and after the break. Given the closeness of the breaks detected, we decided to use the break date provided by the spot price, November 5, 2004.

[4] Both models can be thought of as a generalization of the CVAR model. For a complete presentation of the baseline CVAR model, see Johansen (1995).

Table 5: Perron and Vogelsang Structural Break Unit Root Test Results

Series	Test Statistic	Date
SPOT	-4.585	5 nov 2004
F1	-4.581	29 oct 2004
F2	-4.873	29 oct 2004
F3	-4.848	29 oct 2004
F4	-4.819	29 oct 2004

Note: The sample size is 1981.

In Tables 6 and 7, we show the ADF and KPSS results for the first and second branch respectively. We observe that for the first segment, the conclusions are the same as those obtained from the total series; all series are integrated of order 1 for both tests. However, for the second segment, ADF indicates that the series are stationary, while KPSS suggests the presence of a unit root. Nevertheless, the KPSS test tends to over-reject the null hypothesis. Whatever the case, the results of this section stand as a confirmation that the stochastic component affecting the two branches of the series might indeed be two separate processes and thereby warrant a deeper exploration of the order of integration of the series that we carry out in the next section, devoted to the fractional results.

Table 6: Unit root tests. First Branch

Series	**ΔADF**		**ADF**		**KPSS**	
	Test Statisti	Decision	Test Statisti	Decision	Test Statisti	Decision
SPOT	-10.346*	Not I(2)	-0.605	I(1)	3.340*	I(1)
F1	-8.611*	Not I(2)	-0.710	I(1)	2.580*	I(1)
F2	-8.572*	Not I(2)	-0.450	I(1)	2.660*	I(1)
F3	-10.246*	Not I(2)	-0.188	I(1)	2.720*	I(1)
F4	-11.068*	Not I(2)	0.052	I(1)	2.750*	I(1)

Note: ADF denotes the Augmented Dickey-Fuller and ΔADF in the first differences. KPSS is Kwiatkowski-Phillips Schmidt-Shin. The sample size is 984. * Null rejection at 5% level

Table 7: Unit root tests. Second Branch

Series	**ΔADF**		**ADF**		**KPSS**	
	Test Statisti	Decision	Test Statisti	Decision	Test Statisti	Decision
SPOT	-7.022*	Not I(2)	-3.299*	I(0)	0.550*	I(1)
F1	-8.795*	Not I(2)	-3.534*	I(0)	1.060*	I(1)

F2	-8.595*	Not I(2)	-3.530*	I(0)	1.180*	I(1)
F3	-8.523*	Not I(2)	-3.514*	I(0)	1.300*	I(1)
F4	-8.488*	Not I(2)	-3.497*	I(0)	1.420*	I(1)

Note: ADF denotes the Augmented Dickey-Fuller and ΔADF in the first differences. KPSS is Kwiatkowski-Phillips Schmidt-Shin. The sample size is 997. * Null rejection at 5% level

5.4.2. Fractional Integration results

Tables 8 and 9 report the estimates of the differencing parameter, d, for the whole series and the two branches we identified endogenously by recursive estimation of the most significant break dates through the Perron and Vogelsang's (1992) test. In Table 8, the estimates of the white noise model show that the whole series would show a higher than $I(1)$ behaviour. By examining the two sub-samples, the first one would show similar results to the whole series, with an exception made for SPOT whose d estimate points at a lower confidence interval exactly equal to 1. The second sub-sample would instead clearly show a set of d estimates of higher magnitude, where the value 1 would not be contained in the confidence intervals. The same conclusions as with the white noise case hold in the autocorrelation case. The estimates of d increase as we move from SPOT to F2, F3 or F4. Using the whole sample, the estimates of d are significantly above 1 in all series; however, for the data using only the first subsample, the values are within the $I(1)$ interval; for the second subsample, they are above. Thus, the results obtained for the whole sample are clearly influenced by those of the second subsample. It is thus paramount, given this evidence, to consider not only the possibility of a contango/backwardation phenomenon in the series as a whole, but to ask whether or not the equilibrium, in terms of elasticities and speed of adjustment, might have changed after the break, and to what degree.

Table 8: Estimates of d. White noise. Original data

i) Total (03 January 1986 – 15 December 2023)			
Series	No terms	An intercept	A linear trend
SPOT	1.10 (1.06, 1.14)	**1.11 (1.08, 1.15)**	1.11 (1.08, 1.15)
F1	1.11 (1.07, 1.15)	**1.12 (1.09, 1.16)**	1.12 (1.09, 1.16)
F2	1.12 (1.08, 1.16)	**1.14 (1.10, 1.17)**	1.14 (1.10, 1.17)
F3	1.12 (1.08, 1.16)	**1.14 (1.10, 1.18)**	1.14 (1.10, 1.18)
F4	1.12 (1.09, 1.16)	**1.14 (1.10, 1.18)**	1.14 (1.10, 1.18)
ii) First branch (03 January 1986 – 05 November 2004)			
Series	No terms	An intercept	A linear trend
SPOT	0.98 (0.93, 1.03)	**1.05 (1.00, 1.11)**	1.05 (1.00, 1.11)
F1	0.98 (0.94, 1.03)	**1.06 (1.01, 1.12)**	1.06 (1.01, 1.12)
F2	1.00 (0.95, 1.05)	**1.10 (1.04, 1.16)**	1.10 (1.04, 1.16)
F3	1.00 (0.95, 1.05)	**1.10 (1.04, 1.16)**	1.10 (1.04, 1.16)
F4	1.00 (0.95, 1.05)	**1.10 (1.05, 1.16)**	1.10 (1.05, 1.16)
iii) Second branch (12 November 2004 – 15 December 2023)			

Series	No terms	An intercept	A linear trend
SPOT	1.08 (1.04, 1.13)	**1.12 (1.07, 1.18)**	1.12 (1.07, 1.18)
F1	1.09 (1.04, 1.14)	**1.13 (1.08, 1.18)**	1.13 (1.08, 1.18)
F2	1.10 (1.05, 1.15)	**1.14 (1.09, 1.19)**	1.14 (1.09, 1.19)
F3	1.09 (1.05, 1.15)	**1.14 (1.09, 1.19)**	1.14 (1.09, 1.19)
F4	1.10 (1.05, 1.15)	**1.14 (1.09, 1.20)**	1.14 (1.09, 1.20)

Note: This table reports the estimated values of d with their corresponding confidence intervals in brackets for ea of the specifications. The values in bold are those for the models selected on the basis of the statistical significan of the regressors.

Table 9: Estimates of d. Autocorrelation. Original data

i) Total (03 January 1986 – 15 December 2023)			
Series	No terms	An intercept	A linear trend
SPOT	1.04 (0.98, 1.10)	**1.06 (1.01, 1.12)**	1.06 (1.01, 1.12)
F1	1.04 (0.99, 1.10)	**1.09 (1.01, 1.14)**	1.09 (1.01, 1.14)
F2	1.06 (0.99, 1.11)	**1.09 (1.02, 1.14)**	1.09 (1.02, 1.14)
F3	1.05 (0.99, 1.11)	**1.07 (1.02, 1.14)**	1.07 (1.02, 1.14)
F4	1.06 (1.00, 1.12)	**1.08 (1.03, 1.15)**	1.08 (1.03, 1.15)
ii) First branch (03 January 1986 – 05 November 2004)			
Series	No terms	An intercept	A linear trend
SPOT	0.92 (0.86, 1.00)	**0.95 (0.88, 1.04)**	0.95 (0.87, 1.05)
F1	0.93 (0.87, 1.01)	**0.98 (0.90, 1.08)**	0.98 (0.90, 1.08)
F2	0.94 (0.88, 1.02)	**0.97 (0.90, 1.07)**	0.97 (0.89, 1.07)
F3	0.96 (0.90, 1.04)	**1.00 (0.91, 1.10)**	1.00 (0.90, 1.10)
F4	0.98 (0.92, 1.06)	**1.00 (0.93, 1.10)**	1.00 (0.93, 1.10)
iii) Second branch (12 November 2004 – 15 December 2023)			
Series	No terms	An intercept	A linear trend
SPOT	1.03 (0.97, 1.11)	**1.07 (1.00, 1.16)**	1.07 (1.00, 1.16)
F1	1.05 (0.98, 1.14)	**1.08 (1.01, 1.17)**	1.08 (1.01, 1.17)
F2	1.05 (0.88, 1.14)	**1.08 (1.01, 1.17)**	1.08 (1.01, 1.17)
F3	1.05 (0.98, 1.14)	**1.08 (1.01, 1.17)**	1.08 (1.01, 1.17)
F4	1.04 (0.97, 1.12)	**1.09 (1.00, 1.16)**	1.09 (1.00, 1.16)

Note: This table reports the estimated values of d with their corresponding confidence intervals in brackets for ea of the specifications. The values in bold are those for the models selected on the basis of the statistical significan of the regressors.

There are many procedures for the estimation of the differencing parameter in a semiparametric context. As we have already stated in the methodology section, we report the results from the log-periodogram regression proposed by Geweke and Porter-Hudak (1983) as modified by Robinson (1995). The fractional parameter d for each univariate series is estimated and presented in Table 10. As suggested by the authors and by past literature, we use a bandwidth equal to $m = T^{0.5}$ to calculate the values of d for each series. We work with the first-differenced data, since this test requires that the results are limited to the interval $-0.5 < d < 0.5$. Subsequently, to obtain the proper value of d we add 1.

Table 10: Estimates of *d* using the log-periodogram regression. Whole Series

	GPH estimates $m = T^{0.5}$				
	SPOT	**F1**	**F2**	**F3**	**F4**
$\hat{d}$	0.778	0.779	0.785	0.789	0.790
	(0.086)	(0.086)	(0.084)	(0.083)	(0.083)

Note: GPH denotes the Geweke and Porter-Hudak semiparametric log-periodogram regression estimator. Standa errors are given in parenthesis beneath estimates of *d*. The sample size is 1942

Table 11: Estimates of *d* using the log-periodogram regression. First Brach

	GPH estimates $m = T^{0.5}$				
	SPOT	**F1**	**F2**	**F3**	**F4**
$\hat{d}$	0.483	0.454	0.564	0.647	0.713
	(0.178)	(0.185)	(0.166)	(0.160)	(0.161)

Note: GPH denotes the Geweke and Porter-Hudak semiparametric log-periodogram regression estimator. Standa errors are given in parenthesis beneath estimates of *d*. The sample size is 984

Table 12: Estimates of *d* using the log-periodogram regression. Second Brach

	GPH estimates $m = T^{0.5}$				
	SPOT	**F1**	**F2**	**F3**	**F4**
$\hat{d}$	0.814	0.814	0.824	0.825	0.824
	(0.084)	(0.082)	(0.082)	(0.081)	(0.080)

Note: GPH denotes the Geweke and Porter-Hudak semiparametric log-periodogram regression estimator. Standa errors are given in parenthesis beneath estimates of *d*. The sample size is 997

Although offering a set of slightly differenced, less integrated results compared to the previous methods, indicating some degree of quasi-unit root behaviour instead of a full differentiation (d less than 1 but with values generally close to it) the GPH offers us additional confirmation and useful insight on the stark difference between the two branches of the data.

In particular, and consistently with the previous estimates, the estimates of d for the first branch appear to be much less persistent than the estimates of d in the second branch, the values of which appear to be much closer to 1 and, more importantly, even closer to the estimates of d for the whole series. This is, once again, evidence that the results for the whole sample are mainly influenced by those of the second subsample.

5.4.3. FCVAR results

Following Johansen and Nielsen (2012), we test for a possible FCVAR with an unrestricted constant, ξ , which would allow a better interpretability of the coefficients of the cointegrating vector. We have also imposed the restriction $d \geq b \geq 1$. In Table 13 we define the steps to determine the value of β_2, that is, to know if in the long-run relationship between spot and future prices of crude oil there exists contango, backwardation or neither.

Table 13: Steps of the Empirical Research

Procedure
1.Lag Selection[5]
2.Cointegration Rank Testing
3.Test restriction that *d=b=1* (CVAR vs FCVAR)
4.Estimation of $d = b, \beta_{2,} \beta_2^*, \ \alpha_s$ and α_f

For the four cointegration relationships under study, we use three lags, based on the criterion of parsimony, which proves sufficient in all cases and allows us to confirm the absence of any residual correlation in the post-estimation white noise test on the residuals of the models.

The next step entails testing the cointegration rank to know whether there exists a cointegrating relationship between the variables. In all four specifications, the Johansen trace test confirms the existence of at most one cointegrating relationship. Before estimating β_2 we must test whether the FCVAR model is more suitable than CVAR model by testing the null $d = 1$. We find that in all cases the most appropriate model is the FCVAR, except for the SPOT-F2, SPOT-F3, and SPOT-F4 relationships in the first branch, where CVAR is preferred.

Finally, we test the null hypothesis of $\beta_2 = 1$. We obtain the value of beta to ascertain whether it equals 1 or there is contango or backwardation, the values of d and b indicating the degree of integration and cointegration, and the values of alpha to determine the speed of adjustment of both prices. We show these results in Tables 14, 15 and 16. The p-value of β_2 in the tables indicates the outcome of the test determining whether β_2 would be statistically equal to 1 or not. For the sake of clarity, we also define β_2* which would be equal to 1 in the case of being unable

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5 Before deciding the optimal number of lags for the estimations, we define a maximum order of lags using the measure proposed by Schwert (1987): $K_{max} = 12 * \frac{T}{100}^{1/4}$, where T is the sample size. So, our K_{max} is 25.

to reject H_0 or equivalent to β_2 otherwise.

Table 14: Cointegration Results Whole Series

	SPOT-F1	SPOT-F2	SPOT-F3	SPOT-F4
$\boldsymbol{d = b}$	0.944	0.716	0.730	0.743
S. e.	(0.032)	(0.038)	(0.036)	(0.037)
$\boldsymbol{\beta_2}$	-0.707	-0.663	-0.655	-0.647
P-value	[0.824]	[0.003]	[0.003]	[0.003]
$\boldsymbol{\beta_2 *}$	-1	-0.663	-0.655	-0.647
$\boldsymbol{\alpha_s}$	-0.981	-0.297	-0.157	-0.106
P-value	[0.000]	[0.071]	[0.000]	[0.000]
$\boldsymbol{\alpha_f}$	-0.692	-0.217	-0.097	-0.057
P-value	[0.002]	[0.020]	[0.001]	[0.004]

Note: $\beta_2 *$ denotes the value of β_2 once the null hypothesis of β_2=1 is rejected or not. Standard errors are given in brackets and p-values in squared brackets. The sample size is 1942

Analszing the results of the entire series, we observe that the elasticity of the relationship between the spot price and F1 price appears to be close to 1. After testing for the unitary null hypothesis, we are finally unable to reject the null hypothesis of absence of contango or backwardation. As for the relationship between the spot price and the other futures, there is a situation of increasing contango as the future price rises, and the estimates remain similar across all three models with elasticities ranging from 0.663 to 0.647. The same pattern is reflected in the values of d and b, as well as in the alphas.

As the maturity for the future contracts gets delayed in time, the speed of adjustment steadily decreases. We observe that both prices, the spot and futures, converge to the unique equilibrium relationship estimated, with the spot prices reverting back to the equilibrium after a shock at a faster pace compared to the future prices. For future 1, the long run coefficient estimate is 0.944, very close to 1, while for the other futures, the long-term equilibrium elasticities are slightly lower, at 0.716, 0.730, and 0.743, respectively. Regarding the alphas, the short-term adjustment of both variables is much faster for future 1 than for the other three futures, gradually slowing down as the future increases.

The results for the overall model would thus confirm a generalized, long run contango situation. However, given the existence of the structural break we previously detected in the data and the different time series properties of the two isolated sub-samples, an additional step in the cointegration analysis, entailing a separate modelling before and after break, is clearly required.

Table 15: Cointegration Results First Segment

	SPOT-F1	SPOT-F2	SPOT-F3	SPOT-F4
$\boldsymbol{d = b}$	0.710	1	1	1

S. e.	(0.047)	-	-	-
$\boldsymbol{\beta_2}$	-0.681	-1.019	-1.029	-1.036
P-value	[0.178]	[0.000]	[0.000]	[0.000]
$\boldsymbol{\beta_2 *}$	-1	-1.019	-1.029	-1.036
$\boldsymbol{\alpha_s}$	-1.263	-0.234	-0.149	-0.114
P-value	[0.023]	[0.000]	[0.000]	[0.000]
$\boldsymbol{\alpha_f}$	-0.725	-0.074	-0.050	-0.035
P-value	[0.264]	[0.201]	[0.131]	[0.130]

Note: $\beta_2 *$ denotes the value of β_2 once the null hypothesis of β_2=1 is rejected or not. Standard errors are given in brackets and p-values in squared brackets. The sample size is 984

We present the results for the first segment in Table 15. For the cointegration relationship between the spot price and futures 2, 3, and 4, a CVAR is preferred over FCVAR, resulting in $d = b = 1$.

The branch 1 model, everything else equal, clearly hints at backwardation, as visible by the slightly above one estimates for all the β_2 besides the one characterizing the relationship between SPOT and F1, where the long-run elasticity is found to be nonsignificant and the speed of adjustment overshoots the equilibrium (-1.263), thus implying no feasible error correction representation. Indeed, after having tested for the unitary null hypothesis across all relationships, we find it is rejected across the board, so that $\beta_2 = \beta_2^*$ in the SPOT-F2, SPOT-F3 and SPOT-F4 estimates of Table 15, and the slight degree of backwardation detected in the unrestricted estimates can thus finally be confirmed.

As for the remaining $\alpha_s s$ and $\alpha_f s$, the spot price is the only one that adjusts to the long run equilibrium, at a constantly slower rate as the distance between spot and future prices increases (from -0.234 in SPOT-F2 to -0.114 in SPOT-F4). The first branch of the analysis, frames what would be considered the canonical state of backwardation, with forward prices generally lower than spot prices at any moment of time in a steady state condition.[6]

Table 16: Cointegration Results Second Segment

	SPOT-F1	SPOT-F2	SPOT-F3	SPOT-F4
$\boldsymbol{d = b}$	1.006	0.582	0.605	0.623
S. e.	(0.039)	(0.047)	(0.041)	(0.041)

[6] We recall that the elasticities of the cointegrating vector refer to long run equilibria: arguably, although commodity trading, in any form, would entail continuous regime switches from a situation of contango to a situation of backwardation with a high frequency, our estimates refer to a statistical long run relationship were spot prices, being relatively more subject to variation than future prices, tend to converge faster to a steady state equilibrium (no α_1 we estimated was lower than its relative α_2 in any of the models employed).

$\boldsymbol{\beta_2}$	-0.686	-0.737	-0.739	-0.747
P-value	[0.056]	[0.313]	[0.277]	[0.178]
$\boldsymbol{\beta_2}$ *	-1	-1	-1	-1
$\boldsymbol{\alpha_s}$	-1.021	-0.334	-0.169	-0.106
P-value	[0.005]	[0.003]	[0.006]	[0.009]
$\boldsymbol{\alpha_f}$	-0.731	-0.244	-0.101	-0.055
P-value	[0.069]	[0.014]	[0.044]	[0.071]

Note: $\beta_2 *$ denotes the value of β_2 once the null hypothesis of β_2=1 is rejected or not. Standard errors are given in br kets and p-values in squared brackets. The sample size is 997

For the second segment, in all cases, FCVAR estimation is generally preferred. Similarly to the first segment, the relationship between spot and future 1 appears to be statistically significant considering the value of Beta, but in terms of error correction it overshoots an equilibrium which thus does not appear to exist. Regarding the fractional cointegration between the spot price and futures 2, 3, and 4, our results clearly show that the long run unitary elasticity hypothesis for the second branch of the series cannot be rejected, pointing after the break, to an equilibrium characterized by the absence of contango or backwardation. As for the adjustment speeds, the estimated values are perhaps comparable to those obtained in the first branch of the series. As before, the further away in time spot prices are from forward prices, the slower the adjustment is. As with the first branch, the $\alpha_s s$ appear to be mostly significant at 1%, while the comparatively slower adjustment speeds $\alpha_f s$ of the future price equations are less precisely estimated, being only significant above the 1% p-value threshold.

5.5. Conclusions

We have contributed in this article to the literature on the long run relationship between spot and futures oil prices by implementing a straightforward pre-break and post-break approach to put canonical and fractional vector error correction models to compete in order to verify time persistence of the contango/backwardation states. As opposed to the rest of the literature, this methodology allows for detection of non-linearities in the long-run equilibrium and subsequent identification of its effect on the equilibrium; we are estimating without resorting to either space-state models or threshold models.

We have found that a contango relationship has substantially held its significance across time in the overall sample. However, global exogenous shocks can affect long run relationships, changing their path or statistical significance across time. As we detected a break in the neighbourhood of 2004, we found out that the presence of a global exogenous shock in the series would eventually alter the long run relationship between spot and future prices, causing an equilibrium regime shift.

Our estimates prove that the contango between spot and future prices, which is clearly present as we take the whole sample into consideration, would break down as we examine the memory processes of the data before and after the break: the pre-break estimate would present some lesser degree of backwardation, and the post-break estimate an overall contango situation, which however, once tested, resulted in a unitary elasticity between the two prices, implying, on average, absence of arbitrage across the oil market after the 2004 break.

This conclusive evidence allows us to clarify that, in the long run, non-linearities due to global exogenous shocks are, similarly to sudden changes in fundamentals in the short run, capable of affecting single components of the spot-futures oil price relationship such as storage costs and can determine equilibrium switches.

Our results also qualify an important additional piece of evidence: the necessity to bridge a gap between nonlinear analyses and more refined methods that take more into account the degree of memory in the analyzed prices. We thus offer a solution to both the strand of past literature that would exclusively focus on nonlinearities in a canonical cointegrated threshold VAR or in any other type of state-space model and to a parallel strand that would select discretionally linear fractional cointegrating methods to model the price equilibria, merging these two visions.

Future research would logically need to endeavour to use our strategies with alternative commodity markets than those we have seen in the literature we considered (metals, agricultural or other energy commodities) and, perhaps more importantly, try to extend it to other long run equilibria definitions where potential linearity and strong memory processes are present (other kinds of price differentials, exchange rate parities, interest rates differentials, just to name a few cases).

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Appendix A: Methodology

A.1 Linear Unit root testing

To test the time series properties of our data, we conduct a preliminary analysis with the Augmented Dickey Fuller (ADF) and the Kwiatkowski, Phillips, Schmidt and Shin (KPSS) tests. As we are dealing with nominal prices, in the ADF test we assume a maximum order of integration which would entail the presence of at most two unit roots. The canonical test equation for the ADF would as such be:

$$\Delta y_t = \alpha + \beta t + \gamma y_{t-1} + \delta_1 \Delta y_{t-1} + \cdots + \delta_{p-1} \Delta y_{t-p+1} + \varepsilon_t \tag{6}$$

with $y = (SPOT, F1, F2, F3, F4)$, where the statistic based on the statistical significance of γ would present a null of unit root and an alternative hypothesis of absence of a unit root:

$$\mathrm{DF}_\tau = \frac{\hat{\gamma}}{\mathrm{SE}(\hat{\gamma})}\,. \tag{7}$$

To test up to $I(2)$, in the first step, the canonical equation above is modified to accommodate the first order difference of the series. That would entail the following test equation:

$$\Delta^2 y_t = \alpha + \beta t + \gamma \Delta y_{t-1} + \delta_1 \Delta^2 y_{t-1} + \cdots + \delta_{p-1} \Delta^2 y_{t-p+1} + \varepsilon_t\,, \tag{8}$$

where, once again, $y = (SPOT, F1, F2, F3, F4)^{\mathrm{T}}$, with an equivalent test statistic to the one showed above. Upon rejection of the null of $I(2)$, the testing continues with the un-differenced series, and the test equation reduces back to first one above.

To corroborate the linear results from the ADF, we also employ the KPSS test, where the null of the test would imply stationarity, while the alternative non-stationarity. Based on an autoregressive specification, the test is a Lagrange Multiplier (LM) test that evaluates statistically significant differences between the variances of the error of a stationary specification against a nonstationary one. Given the following process random walk process:

$$\mathrm{y}_t = x_t + \gamma_t + u_t \tag{9}$$

with $x_t = x_{t-1} + e_{i,t}$,

if $Var(e_t) = 0$, then $x_t = x_{t-1}$, then the potentially nonstationary component of the series collapses to a constant, which will be estimated together with intercept γ_t. The test statistic would thus be:

$$\vartheta = \frac{\mathrm{var}(e_{i,t})}{\mathrm{var}(u_{i,t})}\,. \tag{10}$$

When its value is above 0, the null of stationarity can thus be rejected, and the variable can be considered integrated.

A.2. Non-linear unit root testing

As stated in Section 3.1, both visual inspection of the series and an a-priori analysis of their descriptive statistics suggest us that the spot and forward contracts' prices might have been subject to a strong exogenous shock, which probably originated in the neighborhood of 2004. This might have modified the long run equilibrium relationship of the theoretical model, and at the same time it represents a source of disturbance for canonical unit root tests which might be uncapable of rejecting the null of unit root when they should, because of the presence of a constant, deterministic upward shift in the series. To consider such possibility, and to retrieve endogenously a specific date for the sudden change in the series, we employ the Perron and Vogelsang (PV) one-break unit root test.

Under the null hypothesis of presence of a unit root, the test would imply:

$$y_t = \delta D(TB)_t + y_{t-1} + w_t, t = 2, \ldots, T, \tag{11}$$

where $D(TB)_t$ is an impulse dummy taking value 1 in the period immediately after the break and 0 otherwise. The alternative would instead be:

$$y_t = c + \delta DU_t + v_t, t = 2, \ldots, T\ , \tag{12}$$

where DU_t is a shift dummy taking value one for all the periods immediately after the break, and 0 in any other period. The autoregressive test equation would thus be:

$$\tilde{y}_t = \sum_{i=0}^{k} \omega_i D(TB)_{t-i} + \alpha \tilde{y}_{t-1} + \sum_{i=1}^{k} c_i \Delta \tilde{y}_{t-i} + e_t, \quad t = k+2, \ldots, T. \tag{13}$$

where $\tilde{y}_t$ represents the residuals of a first step regression of the analysed series on an intercept and the shift dummy DU_t. This version of the PV test, the additive outlier model, is perhaps mostly suited for our data as it adapts well to sudden changes in the DGP of a series. The test statistic of α will finally be chosen after a recursive estimation of a trimmed subsample of the series, where the statistic evaluated will be the one where the intercept coefficient is the most negative of all the ones calculated in the recursive window.

3.2.1 Fractional integration

To deepen and disambiguate the results from the traditional unit root tests, we employ a testing approach developed in Robinson (1994) and widely employed in empirical application related with long memory and fractionally integrated processes.
The model under examination is

$$y(t) \ = \ \alpha \ + \ \beta t + \ x(t), \quad (1-B)^d x(t) \ = \ u(t), \qquad t = 1, 2, 3, \ldots., \tag{14}$$

where y(t) refers to the observed time series, α and β are unknown coefficients determined within the model, and $(1-L)^d$ being the fractional integration polynomial in L with fractional degree d, which makes the model integrated of order $I(0)$. It tests the null hypothesis:

$$H_o: \ d = d_o, \tag{15}$$

in (14) for any real value d_o, including thus values outside the stationary region, i.e., $d_o \geq 0.5$. In the empirical application carried out in the following section we perform the test for a range of d_o-values from -1 to 2 with 0.01 increments, choosing the band of values where H_o cannot be rejected at the 95% level. The estimate of d is then chosen as the value of d_o that produces the lowest statistic in absolute value, this value being almost identical to the one obtained by maximizing the likelihood function as described in the Appendix. This method, widely employed in empirical applications (see, e.g., Gil-Alana and Robinson, 1997) has a standard null and local limit distribution, and that limit behaviour holds independently of the use of deterministic terms and the modelling of the $I(0)$ disturbance term u_t in (14).

The results reported across the tables are the estimates of the differencing parameter d along with the 95% confidence bands for the estimates of d for three different model specifications, namely i) with no deterministic terms, i.e. imposing $\alpha\ \beta\ =\ 0$ in (14) (results reported in column 2); ii) with a constant, i.e., $\beta\ =\ 0$ in (14) (in column 3); and iii) and with a constant and a linear time trend (in the last column of the tables). The coefficients marked in bold in the table are those from the model selected in each case on the basis of the statistical significance of the coefficients.

We also analyse the order of integration of the price series with the semi-parametric log-periodogram method proposed by Geweke and Porter-Hudak (1983) (GPH) and later improved by Robinson (1995). The GPH test is based on the maximum likelihood estimation of the model below:

$$\ln\ I(\omega_k) = \alpha_0 + \alpha_1 \ln\left[2\sin\left(\frac{2\omega_k}{2}\right)\right]^2, \tag{16}$$

where I(ω) is a log-linear approximation of the periodogram of the series with ω_k frequencies, while α_0 and α_1 represent the components of the slope of the periodogram.
By reparametrizing the slopes we would obtain the Hurst Coefficient, $H = 1 - \alpha_1 \backslash 2$, and the order of fractional integration d, as $-d = \alpha_1$.

A.3 The CVAR and FCVAR Model

We begin detailing the non-fractional CVAR model. Let $Y_{t,} t = 1, \ldots, T$, be a p-dimensional $I(1)$ time series. Then the CVAR model is:

$$\Delta Y_t = \alpha\beta' L Y_t + \textstyle\sum_{i=1}^{k} \Gamma_i\, \Delta L^i Y_t + \varepsilon_t\ , \tag{17}$$

where ε_t, the error term, is a p-dimensional disturbance, identically and independently distributed, with mean zero and covariance matrix Ω. To obtain the FCVAR model, we must replace the difference and lag operators Δb and $Lb = (1 - \Delta b)$, respectively. We would thus get:

$$\Delta^b Y_t = \alpha\beta' L_b Y_t + \textstyle\sum_{i=1}^{k} \Gamma_i\, \Delta^b L_b^i Y_t + \varepsilon_t\ , \tag{18}$$

and to derive the FCVAR model, we apply to $Y_t = \Delta^{d-b} X_t$,

$$\Delta^d X_t = \alpha\beta' L_b \Delta^{d-b} X_t + \textstyle\sum_{i=1}^{k} \Gamma_i\, \Delta^d L_b^i X_t + \varepsilon_t. \tag{19}$$

FCVAR model is interpreted with the same parameters from the CVAR model. In particular, the vectors α and β are p x r matrices where $0 \leq r \leq p$. The columns of β are the cointegration vectors, and $\beta' X_t$ represent the long-run equilibrium relations. The coefficients α are the adjustment coefficients that reflect the speed of adjustment to the equilibrium. The parameters Γ_i denote the short-run dynamics of the variables.

The FCVAR model has two additional parameters compared with the CVAR model. Those parameters are d, that determines the fractional integration order of the time series and b, that represents the degree of fractional cointegration, so that the series in question are cointegrated of order d – b. For $0 \leq (d-b) < 0.5$, the error correction term follows a stationary process and when $0.5 \leq (d-b) < 1$, there is a non-stationary unit root process. Note that, for $d = b = 1$, the FCVAR model collapses to the CVAR model, that is, it is a special case where fractionality is not considered in the FCVAR model.

Johansen and Nielsen (2012) introduce a restricted constant *ρ,* and an unrestricted constant, *ξ*, is proposed by Dolatabadi et al. (2015). So, it implies the general error correction form of the model:

$$X_t = \alpha L_b \Delta^{d-b}(\beta' X_t + \rho') + \textstyle\sum_{i=1}^{k} \Gamma_i\, \Delta^d L_b^i X_t + \xi + \varepsilon_t \tag{20}$$

where, *ρ* denotes the mean level of equilibrium relation in the long run as $\mathrm{E}[\beta' X_t + \rho'] = 0$ and, *ξ* includes the linear trend of the variables.

When cointegration exists, a fractionally vector error correction model (FVECM) with one lag in its short-term structure can be depicted as follows:

$$\binom{\Delta s}{\Delta f} = \binom{\alpha_s}{\alpha_f}\ (s_{t-1} - \beta_2 f_{t-1} - c) + \textstyle\sum_{i=1}^{n} \Gamma_i \binom{\Delta s_{t-1}}{\Delta f_{t-1}} + \binom{w_{1t}}{w2t}, \tag{21}$$

where α_s and α_f are the adjustment coefficients and β_2 is the long run elasticity of forward prices with respect to spot prices. c is the restricted constant, n the lag length, and w are the errors. The short-run dynamics are gathered in Γ_i, a set of $(2 * 2)$ matrices, while α_s and α_f determine the speed of adjustment. Our objective is, once again, to determine which between the spot and the future contracts represents the driving factor of the price generating process, acting as an attractor for the other series.

Appendix B: Figures

This appendix reports the inverse roots of the characteristic polynomial for the SPOT-F1. SPOT-F2, SPOT-F3 and SPOT-F4 whole series models (Figure 1A to Figure 4A). To guarantee the stability and invertibility of the model, the roots should lie outside the unit circle. The black circle corresponds to the canonical CVAR unitary bounds. The red circle represents the modified stability boundary for the FCVAR method. Additional stability checks for the remaining branching models are available on request.

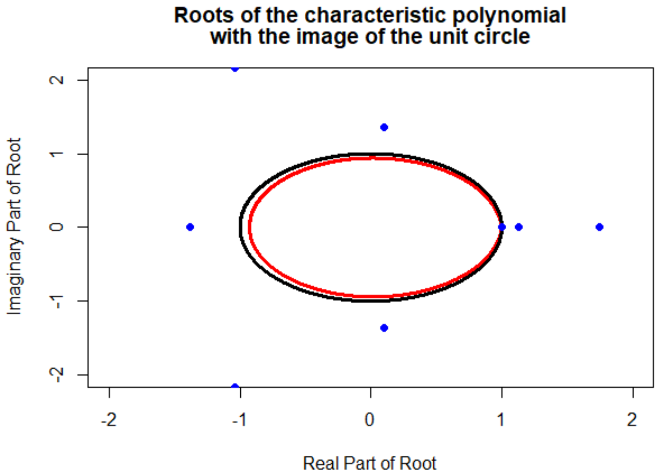

Figure 1: Roots of the characteristic polynomial for SPOT-F1 whole series

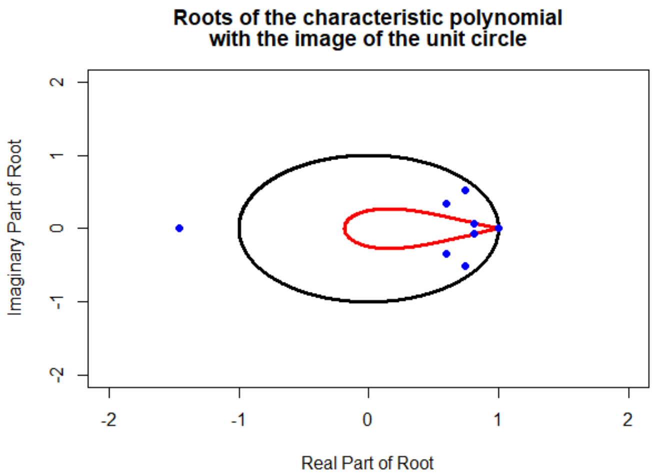

Figure 2: Roots of the characteristic polynomial for SPOT-F2 whole series

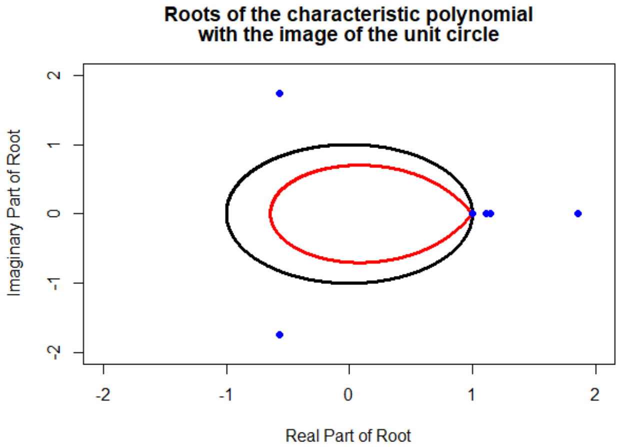

Figure 3: Roots of the characteristic polynomial for SPOT-F3 whole series

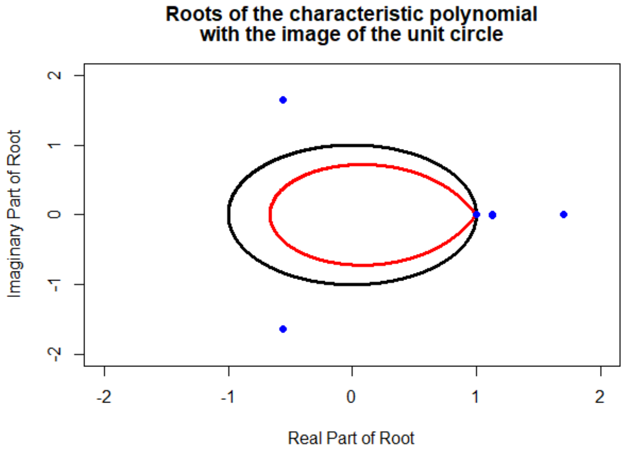

Figure 4: Roots of the characteristic polynomial for SPOT-F4 whole series

Appendix B: Tables

Table 1: Descriptive Statistics

Whole series (03 January 1986 – 15 December 2023)							
	Mean	Median	Min	Max	St.dev	Asymmetry	Kurtosis
Spot	46.867	37.700	3.320	142.520	29.615	0.686	2.313
F1	46.869	37.860	3.920	142.460	29.603	0.683	2.307
F2	46.941	38.520	10.920	143.040	29.639	0.664	2.269
F3	46.946	38.830	11.050	143.390	29.646	0.649	2.238
F4	46.905	39.100	11.230	143.650	29.633	0.637	2.215
First Branch (03 January 1986 – 05 November 2004)							
	Mean	Median	Min	Max	St.dev	Asymmetry	Kurtosis
Spot	22.026	20.060	11.000	54.430	6.799	1.373	5.513
F1	21.996	20.045	11.090	54.300	6.774	1.376	5.537
F2	21.790	19.930	10.920	53.730	6.569	1.444	5.896
F3	21.596	19.765	11.050	53.170	6.353	1.503	6.253
F4	21.417	19.695	11.230	52.520	6.143	1.561	6.601
Second Branch (12 November 2004 – 15 December 2023)							
	Mean	Median	Min	Max	St.dev	Asymmetry	Kurtosis
Spot	71.384	69.480	3.320	142.520	22.059	0.282	2.568
F1	71.417	69.440	3.920	142.460	21.964	0.288	2.572
F2	71.765	69.880	19.070	143.040	21.493	0.323	2.557
F3	71.965	69.780	22.670	143.390	21.118	0.341	2.582
F4	72.062	70.030	24.830	143.650	20.812	0.356	2.618

This section reports the estimates of the white noise and autocorrelation specifications for the Robinson (1994) model with logged data (Table 1A: white noise model. Table 2A: Autocorrelation model). Although not essential for the analysis, as we are not strictly interested in evaluating the absolute magnitude of the log run elasticities of the cointegrating vectors, such results are reported for the sake of completeness and stand in line with the results obtained with the unlogged raw data in the main body of the article.

Table 2: Estimates of d. White noise. Logged data

Series	No terms	An intercept	A linear tren
i) Whole series (03 January 1986 – 15 December 2023)			
SPOT	0.92 (0.88, 0.95)	**0.89 (0.85, 0.92)**	0.89 (0.85, 0.92)
F1	0.93 (0.90, 0.97)	**0.92 (0.88, 0.95)**	0.92 (0.88, 0.95)
F2	1.00 (0.96, 1.04)	**1.13 (1.09, 1.17)**	1.13 (1.09, 1.17)
F3	1.00 (0.96, 1.04)	**1.13 (1.09, 1.17)**	1.13 (1.09, 1.17)
F4	1.00 (0.96, 1.04)	**1.13 (1.09, 1.17)**	1.13 (1.09, 1.17)
ii) First branch (03 January 1986 – 05 November 2004)			
Series	No terms	An intercept	A linear tren
SPOT	0.97 (0.92, 1.02)	**1.05 (1.00, 1.10)**	1.05 (1.00, 1.10)
F1	0.97 (0.92, 1.02)	**1.06 (1.01, 1.12)**	1.06 (1.01, 1.12)
F2	0.98 (0.93, 1.03)	**1.07 (1.04, 1.15)**	1.09 (1.04, 1.15)
F3	0.98 (0.93, 1.03)	**1.10 (1.05, 1.15)**	1.10 (1.05, 1.15)
F4	0.98 (0.93, 1.03)	**1.10 (1.05, 1.15)**	1.10 (1.05, 1.15)
iii) Second subsample (12 November 2004 – 15 December 2023)			
Series	No terms	An intercept	A linear tren
SPOT	0.94 (0.90, 0.99)	**0.83 (0.78, 0.88)**	0.83 (0.78, 0.88)
F1	0.95 (0.91, 1.00)	**0.87 (0.82, 0.91)**	0.87 (0.82, 0.91)
F2	1.00 (0.96, 1.05)	**1.15 (1.10, 1.21)**	1.15 (1.10, 1.21)
F3	1.00 (0.96, 1.05)	**1.14 (1.09, 1.20)**	1.14 (1.09, 1.20)
F4	1.00 (0.95, 1.05)	**1.14 (1.09, 1.20)**	1.14 (1.09, 1.20)

Table 3: Estimates of d. Autocorrelation. Logged data

Series	No terms	An intercept	A linear tren
i) Total (03 January 1986 – 15 December 2023)			
SPOT	0.90 (0.74, 0.95)	**0.91 (0.86, 0.99)**	0.91 (0.86, 0.99)
F1	0.90 (0.85, 0.96)	**0.94 (0.89, 1.01)**	0.94 (0.89, 1.01)
F2	0.92 (0.88, 0.98)	**1.01 (0.95, 1.07)**	1.01 (0.95, 1.07)
F3	0.94 (0.88, 0.98)	**1.02 (0.97, 1.08)**	1.02 (0.97, 1.08)
F4	0.94 (0.89, 0.99)	**1.04 (0.97, 1.09)**	1.04 (0.97, 1.09)
ii) First subsample (03 January 1986 – 05 November 2004)			

Series	No terms	An intercept	A linear tren
SPOT	0.94 (0.87, 1.01)	**0.98 (0.98, 1.08)**	0.98 (0.98, 1.08)
F1	0.94 (0.87, 1.00)	**0.99 (0.92, 1.09)**	0.99 (0.92, 1.09)
F2	0.93 (0.87, 1.02)	**0.99 (0.92, 1.09)**	0.99 (0.92, 1.09)
F3	0.95 (0.88, 1.02)	**1.02 (0.93, 1.12)**	1.02 (0.93, 1.12)
F4	0.96 (0.89, 1.02)	**1.02 (0.95, 1.12)**	1.02 (0.95, 1.12)
iii) Second subsample (12 November 2004 – 15 December 2023)			
Series	No terms	An intercept	A linear tren
SPOT	0.95 (0.88, 1.03)	**0.89 (0.81, 0.97)**	0.89 (0.81, 0.97)
F1	0.95 (0.90, 1.03)	**0.91 (0.83, 0.99)**	0.91 (0.83, 0.99)
F2	0.98 (0.92, 1.05)	**1.00 (0.93, 1.09)**	1.00 (0.93, 1.09)
F3	0.99 (0.91, 1.05)	**1.02 (0.93, 1.08)**	1.02 (0.93, 1.08)
F4	0.99 (0.93, 1.06)	**1.01 (0.95, 1.11)**	1.01 (0.95, 1.10)

Table 4: Chronological literature on cointegrating methodologies in contango/ backwardation studies

PAPER	MAIN IDEA	METHODOLODY	DATA	RESULTS
Brooks et al. (2001)	Relate spot and future prices of spot index and the contract for the FTSE 100	Cointegration and ECM	June 1996 - June 1997	Spot returns are leaded by futures returns. β=0.983
Bekiros & Diks (2008)	Relationship between crude oil (WTI) spot and futures prices	Cointegration, linear and nonlinear causality	Period 1: 10/21/91 - 10/29/99. Period 2: 11/01/99 - 10/30/07	Different results, but the most important: bi-directional causality
Huang et al. (2009)	Relate spot and futures of WTI crude oil	VECM and MVTAR	Period 1: 01/02/86 - 02/28/91; Period 2: 03/31/91 - 08/31/01; Period 3: 09/01/01 - 04/30/07	For periods 1 and 2 exist backwardation and in period 3, contango
Kaufmann & Ullman (2009)	Assess spot and futures oil prices, trying to know where innovations are originated and how they are spread	Two-step DOLS ECM and VECM	Different dates depend on WTI, Brent, Bonny Light or Dubai-Fateh.	Sometimes innovations appear on Dubai-Fateh spot prices and other times on WTI, in both cases, then spread to other spot and futures prices
Maslyuk & Smyth (2009)	Cointegration between spot and	Residual-based Gregory and	January 1991 -	Spot and futures prices of the same and different grades are

	futures prices of oil crude of the same grade or different grade	Hansen (1996) cointegration methodology	November 2008	cointegrated and there exists a break
Lee & Zeng (2011)	Relationship between spot and futures oil prices	Quantile cointegration regression and causality	01/02/86 - 07/06/09	The length of futures contracts has an influence on cointegrating relationships between spot and futures oil prices. In most of the results, spot oil prices Granger cause futures oil prices
Arouri et al. (2012)	Investigate the volatility spillovers between oil and stock markets in Europe (Dow Jones (DJ) Stoxx Europe 600 index and seven DJ Stoxx sector indices)	VAR-GARCH	01/01/98 - 12/31/09	Significant volatility spillovers between oil price and sector stock returns
Rossi & Santucci de Magistris (2013)	The no-arbitrage relation between futures and spot prices of S&P500 index	FVECM	November 27, 1998 to August 25, 2008	β is close to -1. Spot log-range converges faster to the long-run equilibrium than the futures log-range
Wang & Wu (2013)	Relationships between spot and future prices of WTI	TVECM	01/03/86 - 02/18/11	In the short-term, futures price plays the major role in the formation of long-run equilibrium. In the long-term, both spot and futures prices contribute to the dynamics of long-run equilibrium.
Chang & Lee (2015)	Relation between spot and future prices of crude oil	Cointegration, VEC and Wavelet transform	January 1986 - February 2014	Long-run cointegration relationship. If maturity of the contracts increases, the variability of futures prices and the speed of adjustment of futures/spot prices differentials increase too
Dolatabadi et al. (2015)	Analyse contango or backwardation for aluminum, copper, lead, nickel, and zinc	FCVAR	January 1989 - October 2006	There is not contango or backwardation
Dolatabadi et al. (2016)	Analyse contango or backwardation for aluminum, copper, lead, nickel, and zinc	FCVAR	January 1989 - October 2006	There is not contango or backwardation
Gil-Alana et al. (2017)	Cointegration between gold and oil prices	FCVAR	April 1968 - June 2015	Order of integration in the long run relationship of about 0.46
Allen et al., (2018)	Relationships spot - futures prices of corn, sugar, wheat, ethanol and crude oil	MSVEC	07/19/06 - 07/02/15	Exists a long relationship between some of the markets reviewed. Crude oil: β=1.0027
Dolatabadi et al. (2018)	Analyse contango or backwardation for 17 commodity spot and futures markets	FCVAR	03/30/83 - 11/12/12	In 14 of these 17 commodities exist backwardation. Crude oil: β=1.007

Holmes & Otero (2019)	Replicate and expand Chang, C. P., & Lee	Cointegration, VEC and Wavelet transform	January 1986 - February 2014	Futures prices do not respond to long-run disequilibrium and that short-run causality is also present at the longer maturities.
Hu et al. (2019)	Relationships between spot and futures markets of bitcoin	Cointegration with Park and Hahn (1999) test	Spot prices: 12/25/17 - 07/29/19. Future prices: 12/18/17 - 06/16/19	Exists cointegration between those variables
Vides et al. (2020)	Long-run relationship between short- and long-term interest rates and spread persistence together	FCVAR	October 1993 - December 2018	The persistence of spread is the stronger the larger the difference in maturity is between considered interest rates
Pradhan et al. (2021)	The relation between spot and futures prices of aluminum, copper, crude oil, gold, nickel, and silver	ARDL and VECM	01/01/09 - 12/31/20	All variables: I(1), cointegration between spot and futures prices. There is no causality between the spot and futures prices for crude oil.
Wu et al. (2021)	Analyse contango or backwardation for bitcoin	FCVAR	12/18/17 - 07/31/20	Bitcoin futures market follows a long-run contango

Part III: Concluding remarks

Chapter 7: Conclusions

7.1. Conclusions

The second chapter analyzes whether shocks to self-employment in the UK have permanent or transitory effects, with a focus on sectoral differences. Robust methods, such as fractional integration and seasonal autoregressive components, were used to assess the persistence of these effects. The results show a high level of persistence in the total self-employment series, with nine sectors where shocks have lasting effects, although in eight of them, they eventually revert to the long-term mean. This suggests that the impact of shocks varies across sectors, implying that entrepreneurship policies should be adapted to these differences and avoid horizontal approaches. The study highlights that, although many sectors exhibit persistent but transitory effects, self-employment at the aggregate level tends to be irreversible. Entrepreneurial skills and the convergence of competencies across sectors can facilitate intersectoral mobility. The construction sector stands out as the only one showing clear persistence without mean reversion, making it particularly vulnerable to hysteresis effects. These findings underscore the need for sector-specific policies to mitigate the impacts of economic shocks, such as those experienced during the COVID-19 pandemic, which affected different sectors unevenly. Failing to consider sectoral hysteresis can reduce the effectiveness of policies and, in some cases, prove counterproductive. The study acknowledges limitations, such as the need to investigate hysteresis in other countries and sectors related to self-employment, as well as to explore intersectoral dynamics and the impact of the gig economy on the UK labor market. It is also suggested to examine the relationship between hysteresis and other macroeconomic variables. Finally, it is cautioned that persistence in self-employment may reflect both success and a lack of alternatives, requiring a deeper analysis of whether the permanence in self-employment is voluntary or involuntary.

The third chapter studies explored whether shocks to self-employment in Spain and its autonomous communities are permanent or transitory, using fractional integration to assess potential asymmetries between self-employed individuals with and without employees across various sectors. The findings indicate that most series show mean reversion, suggesting that shocks are generally transitory. However, differences emerge between employers and solo self-employed individuals, as well as across sectors and regions. Shocks are more persistent for employers in the commerce, transportation, and services sectors, particularly in northern regions, while the heavy and energy industry sector shows the quickest mean reversion nationwide. The construction sector, on the other hand, exhibits signs of hysteresis in nearly half of the country. These results emphasize the need for economic policies tailored to sectoral, regional, and self-employment-specific characteristics, with a focus on medium- and long-term impacts. A key limitation of the study is the need to distinguish whether persistence in self-employment reflects voluntary engagement due to favorable conditions or involuntary factors like structural barriers, lack of alternatives, or job insecurity. Future analyses should consider factors like job satisfaction, income, market rigidities, regional inequalities, and opportunities for sectoral mobility. Public policies promoting economic diversification, continuous training, and sustainable working conditions could help ensure that persistence in self-employment results from resilience and success rather than exclusion or limitation.

The fourth chapter examines the factors influencing loans to non-financial corporations (NFCs) in Germany, France, Italy, and Spain—countries where such financing plays a critical role in the corporate sector. Using fractional integration and cointegration methods, which account for potential long-memory properties in the data, both univariate and multivariate models are estimated to analyze individual series and their interconnections. The univariate analysis reveals considerable heterogeneity across countries and variables, with the highest levels of integration observed in loans to NFCs. Multivariate results show evidence of fractional cointegration in all four countries. The speed of adjustment of loans in restricted models is comparable to previous studies for Germany and Spain, while slightly slower in France and Italy. Long-term determinants of loans differ across countries: DEBT is the sole factor in Germany, RGDP plays the primary role in France, and a combination of factors influences Italy and Spain. Notably, the cost of bank lending relative to alternative financing sources is significant, suggesting that policy measures targeting these factors could affect loan dynamics. Future research should expand the analysis to include more variables addressing both demand and supply factors and the influence of policy interventions. Additionally, exploring structural breaks and nonlinearities, such as threshold effects, could provide deeper insights into the dynamics of loans to NFCs.

Chapter five contributes to the literature on the long-term relationship between spot and futures oil prices by using a simple pre-break and post-break approach, comparing canonical and fractional vector error correction models to verify the temporal persistence of contango and backwardation states. Unlike previous studies, this methodology allows for the detection of nonlinearities in the long-term equilibrium without resorting to state-space or threshold models. The results show that the contango relationship has remained significant over time in the general sample. However, global exogenous shocks can alter these long-term relationships, modifying their trajectory or statistical significance. A break was identified around 2004, where a global exogenous shock changed the relationship between spot and futures prices, causing a regime shift in the equilibrium. Before the break, a slight backwardation is observed, while after the break, a contango situation predominates. However, tests reveal unit elasticity between the two prices after the 2004 break, implying, on average, the absence of arbitrage opportunities in the oil market. These findings confirm that nonlinearities caused by global exogenous shocks can affect specific components of the spot-futures relationship, such as storage costs, and cause changes in equilibrium. Additionally, the study highlights the need to integrate nonlinear analysis with more sophisticated methods that consider the degree of memory in the analyzed prices, merging traditional approaches with fractional models. Future research should apply these strategies to other commodity markets, such as metals, agricultural products, or energy, and extend them to other definitions of long-term equilibrium where strong memory processes and linearity may be present, such as price differentials, exchange rate parities, or interest rate differentials.

The last chapter explores the dynamics of self-employment, focusing on necessity and opportunity entrepreneurs, and contributes to the understanding of entrepreneurship macrodynamics. It highlights the significant heterogeneity among self-employed individuals, which has been overlooked in previous literature. The study finds that necessity entrepreneurs, driven by adverse labor market conditions, exhibit countercyclical behaviors, while opportunity entrepreneurs, influenced by access to capital and support, show procyclical patterns. A key contribution is the development of a theoretical model that considers both necessity and opportunity entrepreneurship, providing insights into their impact on the labor market and economic conditions. The study also emphasizes the importance of understanding the motivations behind entrepreneurship for assessing business economic performance. The research offers valuable insights for policymakers, suggesting that necessity entrepreneurship does not always lead to sustainable business performance and highlighting the need for strategies that effectively stimulate self-employment and entrepreneurship. It also points out the need for further exploration to understand the relationship between start-up motivations and business performance. In conclusion, the chapter underscores the complexities of entrepreneurship and the importance of ongoing investigation to refine strategies that promote productive self-employment and economic growth.

7.2. Conclusions in Spanish

El segundo capítulo de esta tesis analiza si los shocks en el autoempleo en el Reino Unido tienen efectos permanentes o transitorios, con un enfoque en las diferencias sectoriales. Se utilizaron métodos robustos, como la integración fraccional y componentes autorregresivos estacionales, para evaluar la persistencia de estos efectos. Los resultados muestran un alto nivel de persistencia en la serie total de autoempleo, con nueve sectores donde los shocks tienen efectos duraderos, aunque en ocho de ellos eventualmente se revierte a la media a largo plazo. Esto sugiere que el impacto de los shocks varía entre sectores, lo que implica que las políticas de emprendimiento deben adaptarse a estas diferencias y evitar enfoques horizontales. El estudio destaca que, aunque muchos sectores muestran efectos persistentes pero transitorios, el autoempleo a nivel agregado tiende a ser irreversible. Las habilidades emprendedoras y la convergencia de competencias entre sectores pueden facilitar la movilidad intersectorial. El sector de la construcción es el único que presenta una persistencia clara sin reversión a la media, lo que lo hace especialmente vulnerable a los efectos de histéresis. Estos hallazgos subrayan la necesidad de políticas específicas por sector para mitigar los impactos de los shocks económicos, como los experimentados durante la pandemia de COVID-19, que afectaron de manera desigual a distintos sectores. No considerar la histéresis sectorial puede reducir la eficacia de las políticas y, en algunos casos, resultar contraproducente. El estudio reconoce limitaciones, como la necesidad de investigar la histéresis en otros países y sectores relacionados con el autoempleo, así como explorar las dinámicas intersectoriales y el impacto de la economía gig en el mercado laboral del Reino Unido. También se sugiere examinar la relación entre la histéresis y otras variables macroeconómicas. Finalmente, se advierte que la persistencia en el autoempleo puede reflejar tanto el éxito como la falta de alternativas, lo que requiere un análisis más profundo sobre la naturaleza voluntaria o involuntaria de la permanencia en el autoempleo.

El tercer capítulo explora si los shocks en el autoempleo en España y sus comunidades autónomas son de carácter permanente o transitorio, utilizando la integración fraccional para evaluar posibles asimetrías entre los autoempleados con y sin asalariados en distintos sectores. Los resultados indican que la mayoría de las series presentan reversión a la media, lo que sugiere que los efectos de los shocks son, en general, transitorios. Sin embargo, se observan diferencias entre los empleadores y los autoempleados sin asalariados, así como entre sectores y regiones. Los shocks son más persistentes para los empleadores en los sectores de comercio, transporte y servicios, especialmente en las regiones del norte del país, mientras que el sector de la industria pesada y energética muestra la reversión a la media más rápida a nivel nacional. Por otro lado, el sector de la construcción presenta signos de histéresis en casi la mitad del país. Estos resultados destacan la necesidad de diseñar políticas económicas adaptadas a las características específicas de los sectores, las regiones y los tipos de autoempleo, con un enfoque en los impactos a medio y largo plazo. Una limitación clave del estudio es la necesidad de distinguir si la persistencia en el autoempleo refleja una participación voluntaria debido a condiciones favorables o si responde a factores involuntarios como barreras estructurales, falta de alternativas o inseguridad laboral. Los futuros análisis deberían considerar factores como la satisfacción laboral, los ingresos, las rigideces del mercado, las desigualdades regionales y las oportunidades de movilidad sectorial. Las políticas públicas que promuevan la diversificación económica, la formación continua y condiciones laborales sostenibles podrían ayudar a garantizar que la persistencia en el autoempleo sea el resultado de la resiliencia y el éxito, y no de la exclusión o las limitaciones.

El cuarto capítulo examina los factores que influyen en los préstamos a las sociedades no financieras (SNF) en Alemania, Francia, Italia y España, países donde este tipo de financiación desempeña un papel fundamental en el sector corporativo. Utilizando métodos de integración y cointegración fraccionaria, que tienen en cuenta posibles propiedades de memoria larga en los datos, se estiman modelos univariantes y multivariantes para analizar las series individuales y sus interconexiones. El análisis univariante revela una considerable heterogeneidad entre los países y las variables, observándose los niveles más altos de integración en los préstamos a las SNF. Los resultados multivariantes muestran evidencia de cointegración fraccionaria en los cuatro países. La velocidad de ajuste de los préstamos en los modelos restringidos es comparable a estudios previos para Alemania y España, mientras que en Francia e Italia es ligeramente más lenta. Los determinantes a largo plazo de los préstamos varían entre países: la DEUDA es el único factor relevante en Alemania, el PIB real (RGDP) desempeña el papel principal en Francia, y una combinación de factores influye en Italia y España. Es especialmente relevante el costo de los préstamos bancarios en comparación con otras fuentes de financiación, lo que sugiere que las medidas políticas que afecten estos factores podrían influir en la dinámica de los préstamos. Futuras investigaciones deberían ampliar el análisis para incluir más variables que aborden tanto los factores de demanda como de oferta, así como el impacto de las políticas públicas. Además, sería útil explorar rupturas estructurales y no linealidades, como efectos umbral, para obtener una comprensión más profunda de la dinámica de los préstamos a las SNF.

El capítulo quinto aporta a la literatura sobre la relación a largo plazo entre los precios spot y futuros del petróleo mediante el uso de un enfoque simple de pre-ruptura y post-ruptura, comparando modelos canónicos y de corrección de errores vectoriales fraccionarios para verificar la persistencia temporal de los estados de contango y backwardation. A diferencia de estudios previos, esta metodología permite detectar no linealidades en el equilibrio a largo plazo sin recurrir a modelos de espacio-estado o modelos de umbral. Los resultados muestran que la relación de contango se ha mantenido significativa a lo largo del tiempo en la muestra general. Sin embargo, los shocks exógenos globales pueden alterar estas relaciones a largo plazo, modificando su trayectoria o significancia estadística. Se identificó una ruptura alrededor de 2004, donde un shock exógeno global cambió la relación entre los precios spot y futuros, provocando un cambio de régimen en el equilibrio. Antes de la ruptura, se observa un leve backwardation, mientras que después de la ruptura predomina una situación de contango. Sin embargo, las pruebas revelan una elasticidad unitaria entre los dos precios tras la ruptura de 2004, lo que implica, en promedio, la ausencia de oportunidades de arbitraje en el mercado petrolero. Estos hallazgos confirman que las no linealidades causadas por shocks exógenos globales pueden afectar componentes específicos de la relación spot-futuro, como los costos de almacenamiento, y provocar cambios en el equilibrio. Además, el estudio destaca la necesidad de integrar análisis no lineales con métodos más sofisticados que consideren el grado de memoria en los precios analizados, fusionando enfoques tradicionales con modelos fraccionarios. Futuras investigaciones deberían aplicar estas estrategias a otros mercados de materias primas, como metales, productos agrícolas o energéticos, y extenderlas a otras definiciones de equilibrio a largo plazo donde puedan estar presentes procesos de memoria fuerte y linealidad, como diferenciales de precios, paridades cambiarias o diferenciales de tasas de interés.

El último capítulo explora la dinámica del autoempleo, centrándose en los emprendedores por necesidad y oportunidad, y contribuye a la comprensión de la macrodinámica del emprendimiento. Destaca la significativa heterogeneidad entre los individuos autoempleados, que ha sido pasada por alto en la literatura previa. El estudio encuentra que los emprendedores por necesidad, impulsados por condiciones adversas del mercado laboral, exhiben comportamientos contracíclicos, mientras que los emprendedores por oportunidad, influenciados por el acceso al capital y el apoyo, muestran patrones procíclicos. Una contribución clave es el desarrollo de un modelo teórico que considera tanto el emprendimiento por necesidad como por oportunidad, proporcionando ideas sobre su impacto en el mercado laboral y las condiciones económicas. El estudio también enfatiza la importancia de comprender las motivaciones detrás del emprendimiento para evaluar el desempeño económico empresarial. La investigación ofrece valiosas ideas para los responsables de políticas, sugiriendo que el emprendimiento por necesidad no siempre conduce a un desempeño empresarial sostenible y destacando la necesidad de estrategias que estimulen efectivamente el autoempleo y el emprendimiento. También señala la necesidad de una mayor exploración para comprender la relación entre las motivaciones para iniciar un negocio y el desempeño empresarial. En conclusión, el capítulo subraya las complejidades del emprendimiento y la importancia de la investigación continua para refinar estrategias que promuevan el autoempleo productivo y el crecimiento económico.

Abstract

This thesis consists of five self-contained essays structured as follows.

Chapter 2 revisits whether shocks in self-employment generate permanent or transitory impacts by using Spanish time series and fractional integration, taking into account the possible existence of asymmetric results across regions, sectors, and types of self-employed individuals. Specifically, it distinguishes between self-employed workers who hire others, becoming employers, and those who do not. From the analysis of the long-memory properties of these series, it can be anticipated that most of them exhibit mean reversion, indicating that they generally experience only a transitory or temporary effect when affected by a shock of any kind. However, alongside this general finding, shocks affecting employer series tend to show some degree of persistence. In terms of sectors, the construction sector appears to have a higher probability of shocks having permanent effects, as there is evidence of hysteresis in this sector in almost half of the country's regions. For self-employed individuals without employees, persistent shocks are observed, to some extent, only in the commerce, transportation, and services sectors, particularly in the northern regions of the country. In contrast, when shocks affect the heavy and energy industries, mean reversion is relatively quick across all regions. In addition to these findings, positive temporary trends are identified in the services sector, while negative trends are observed in the agro-industrial sector. These results highlight the need for policies tailored to sectoral and regional characteristics, promoting sustainable conditions and labor resilience.

Chapter 3 analyses whether shocks in self-employment generate permanent or transitory impacts in Spain, using fractional integration models to differentiate between self-employed individuals with and without employees across various sectors and regions. Most of the analysed series exhibit mean reversion, indicating that shocks are transitory. However, self-employed individuals with employees show greater persistence, particularly in the construction sector, where hysteresis is observed in nearly half of the country. For self-employed individuals without employees, shocks are more persistent in the commerce, transport, and services sectors, especially in northern Spain, while the heavy and energy industries demonstrate a faster recovery nationwide. Additionally, positive temporal trends were identified in the services sector and negative trends in the agroindustry. These findings underscore the need for policies tailored to sectoral and regional characteristics, promoting sustainable conditions and labor resilience.

Chapter 4 uses fractional integration and cointegration methods to analyse the long-run relationship between loans to non-financial corporations, real gross domestic product, real gross fixed capital formation, the cost of borrowing differential between long- and short-term rates, and a proxy for the cost of debt, securities, and equity issuance. The analysis includes four Eurozone countries, namely Germany, France, Italy, and Spain, and spans the most recent decades. More precisely, fractional integration and cointegration models are estimated to investigate the persistence of the series as well as their long-run relationships and short-run dynamics using both unrestricted and restricted specifications. The univariate results are heterogeneous, the highest degrees of integration being found in the case of loans to non-financial corporations, whilst the multivariate ones provide evidence of a single fractional cointegration vector as well as of a lower adjustment speed to the long-run equilibrium compared to previous studies in all four countries. Moreover, both the short- and long-run response of loans to exogenous shocks to real gross

domestic product and the cost of borrowing differentials differs across countries because of country-specific factors.

Chapter 5 examine the long-run relationship between spot and future oil prices by considering nonlinearity, a potentially non-integer degree of integration, and allowing both canonical and fractional cointegration methods to compete for the best data fit. We also apply an ex-ante unit root test with endogenous break detection and a pre-break and post-break approach to compare canonical and fractional vector error correction models. Bridging a gap between nonlinear and fractional applications, this methodology uniquely detects nonlinearities in the long-run equilibrium without resorting to state-space or threshold models. Using weekly observations of spot and future prices from the U.S. Energy Information Administration, our estimates identify a structural break around 2004. The canonical contango situation detectable across the whole series breaks down upon examining the data's memory processes before and after this break. Our results also reveal a significant shift in the equilibrium model linking spot to future oil prices due to a global exogenous shock, suggesting two different, time-dependent equilibria: a backwardation equilibrium before the break and a near-unity contango after. This evidence demonstrates that global shocks can affect the spot-futures oil price relationship, inducing equilibrium switches.

Chapter 6 explores the dynamics of self-employment by examining the distinct roles of necessity and opportunity entrepreneurs in the context of the UK labor market. Utilizing a dynamic model of entrepreneurship, the study reveals significant heterogeneity among self-employed individuals, highlighting that necessity entrepreneurs exhibit countercyclical behaviors driven by adverse labor market conditions, while opportunity entrepreneurs are influenced by capital availability and intensity of trademark registration. The research contributes to existing literature by offering a theoretical framework that simultaneously analyzes both types of entrepreneurship for the first time. It underscores the importance of understanding causal relationships among necessity entrepreneurship, unemployment, and opportunity entrepreneurship. The findings suggest that promoting self-employment among the unemployed may not be a universally effective strategy for reducing unemployment.

Abstract in Spanish

Esta tesis consta de cinco ensayos autocontenidos estructurados de la siguiente manera.

El capítulo 2 revisita si los shocks en el autoempleo generan impactos permanentes o transitorios haciendo uso de series españolas e integración fraccional, teniendo en cuenta la posible existencia de resultados asimétricos entre regiones, sectores y entre tipos de autoempleados, concretamente al distinguir entre los trabajadores por cuenta propia que contratan a otros convirtiéndose en empleadores de aquellos otros que no lo hacen. Del análisis de las propiedades de larga memoria de estas series, cabe anticipar que la mayor parte de ellas muestran reversión a la media, lo que indica que la mayor parte de ellas solo sufren un efecto transitorio o temporal cuando se ven afectadas por un shock de cualquier naturaleza. Sin embargo, y junto a este resultado general, los shocks que afectan a las series de empleadores exhiben por regla general cierta persistencia, mientras que, por sectores, es en el sector de la construcción en el que parece haber una mayor probabilidad de que shocks adquieran efectos permanentes, ya que se obtiene evidencia de histéresis en este sector para casi la mitad de las regiones del país. Para los autoempleados sin asalariados, en cambio, los shocks persistentes sólo se observan, en cierta medida, en los sectores de comercio y del transporte y servicios, especialmente en las regiones del norte del país, mientras que cuando los shocks afectan a en la industria pesada y energética la reversión a la media es bastante rápida en cualquiera de las regiones del país. Junto a estos resultados, se identifican tendencias temporales positivas en el sector servicios y negativas en la agroindustria. Estos resultados apuntan hacia la necesidad de políticas adaptadas a las características sectoriales y regionales, promoviendo condiciones sostenibles y resiliencia laboral.

El capítulo 3 analiza si los choques en el autoempleo generan impactos permanentes o transitorios en España, utilizando modelos de integración fraccional para diferenciar entre trabajadores autónomos con y sin empleados en diversos sectores y regiones. La mayoría de las series analizadas exhiben reversión a la media, lo que indica que los choques son transitorios. Sin embargo, los trabajadores autónomos con empleados muestran una mayor persistencia, particularmente en el sector de la construcción, donde se observa histéresis en casi la mitad del país. Para los autónomos sin empleados, los choques son más persistentes en los sectores de comercio, transporte y servicios, especialmente en el norte de España, mientras que las industrias pesadas y energéticas demuestran una recuperación más rápida a nivel nacional. Además, se identificaron tendencias temporales positivas en el sector servicios y tendencias negativas en la agroindustria. Estos hallazgos destacan la necesidad de políticas adaptadas a las características sectoriales y regionales, promoviendo condiciones sostenibles y resiliencia laboral.

El capítulo 4 utiliza métodos de integración y cointegración fraccional para analizar la relación a largo plazo entre los préstamos a las corporaciones no financieras, el producto interno bruto (PIB) real, la formación bruta de capital fijo real, el diferencial de coste de endeudamiento entre tasas de interés a largo y corto plazo, y un indicador representativo del coste de emisión de deuda, valores y acciones. El análisis incluye cuatro países de la Eurozona: Alemania, Francia, Italia y España, y abarca las décadas más recientes. Más específicamente, se estiman modelos de integración y cointegración fraccional para investigar la persistencia de las series, así como sus relaciones a largo plazo y dinámicas a corto plazo, utilizando especificaciones tanto restringidas como no restringidas. Los resultados univariados son heterogéneos, encontrándose los mayores

grados de integración en el caso de los préstamos a las corporaciones no financieras. Por su parte, los resultados multivariados proporcionan evidencia de un único vector de cointegración fraccional, así como una velocidad de ajuste al equilibrio de largo plazo menor en comparación con estudios previos en los cuatro países. Además, tanto la respuesta a corto como a largo plazo de los préstamos frente a choques exógenos en el PIB real y los diferenciales de coste de endeudamiento varía entre países debido a factores específicos de cada uno.

El capítulo 5 examina la relación a largo plazo entre los precios spot y futuros del petróleo considerando la no linealidad, un posible grado de integración no entero y permitiendo que los métodos de cointegración canónica y fraccional compitan por el mejor ajuste de los datos. También aplicamos una prueba de raíz unitaria ex ante con detección de quiebres endógenos, así como un enfoque de pre-quiebre y post-quiebre para comparar modelos de corrección de errores vectoriales canónicos y fraccionales. Al cerrar una brecha entre aplicaciones no lineales y fraccionales, esta metodología detecta de manera única las no linealidades en el equilibrio a largo plazo sin recurrir a modelos de espacio de estados o de umbral. Utilizando observaciones semanales de precios spot y futuros de la Administración de Información Energética de los EE. UU., nuestras estimaciones identifican un quiebre estructural alrededor de 2004. La situación de contango canónico detectable en toda la serie se desmorona al examinar los procesos de memoria de los datos antes y después de este quiebre. Nuestros resultados también revelan un cambio significativo en el modelo de equilibrio que vincula los precios spot con los futuros debido a un choque exógeno global, lo que sugiere dos equilibrios diferentes y dependientes del tiempo: un equilibrio de backwardation antes del quiebre y un contango cercano a la unidad después. Esta evidencia demuestra que los choques globales pueden afectar la relación entre los precios spot y futuros del petróleo, induciendo cambios en los equilibrios.

El Capítulo 6 explora la dinámica del autoempleo al examinar los roles distintos de los emprendedores por necesidad y por oportunidad en el contexto del mercado laboral del Reino Unido. Utilizando un modelo dinámico de emprendimiento, el estudio revela una heterogeneidad significativa entre los trabajadores por cuenta propia, destacando que los emprendedores por necesidad exhiben comportamientos contracíclicos impulsados por condiciones adversas del mercado laboral, mientras que los emprendedores por oportunidad están influenciados por la disponibilidad de capital y la intensidad del registro de marcas. La investigación contribuye a la literatura existente al ofrecer un marco teórico que analiza simultáneamente ambos tipos de emprendimiento por primera vez. Además, subraya la importancia de comprender las relaciones causales entre el emprendimiento por necesidad, el desempleo y el emprendimiento por oportunidad. Los resultados sugieren que fomentar el autoempleo entre las personas desempleadas puede no ser una estrategia universalmente eficaz para reducir el desempleo.