ESSAYS IN INTERNATIONAL FINANCE: PRICE PRESSURE IN THE CHILEAN FOREX MARKET, EXCHANGE RATE FORECASTING AND COMMODITY MARKET SPILLOVERS

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Abstract

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Essays in International Finance: Price pressure in the Chilean FOREX market, Exchange rate forecasting and Commodity market spillovers

This thesis analyses how financial advisory firms may impact price movements in the Chilean peso FOREX market, examines the role of tail dependence in explaining the ability of commodity prices to forecast exchange rates in commodity-exporting economies, and revisits evidence of financial contagion between commodity and stock markets in such economies.

The first essay, "Financial advisory firms, asset reallocation and price pressure in the FOREX market", analyses the effect of advisory firm's investment recommendations on the Chilean pension fund market. Such recommendations induce large, coordinated portfolio readjustments and a subsequent reallocation of pension fund holdings across asset classes. We study the potential for these portfolio asset reallocations in the Chilean pension fund industry to act as a mechanism for exerting price pressures in the Chilean FOREX market. We document investment recommendations generate significant price pressures and increase exchange rate volatility. Using a Central Bank of Chile proprietary database, we find other FOREX market participants anticipate portfolio adjustments after recommendations and front-run the pension fund trades. Our evidence suggests financial advisory firms create substantial effects on the Chilean FOREX market, pushing prices and volatility beyond fundamentals. Policy considerations aiming to regulate advisory firms may mitigate any undesirable impact on financial stability mandates.

The second essay, "Asymptotic dependence and exchange rate forecasting", studies the tail dependence between changes in commodity prices and exchange rates. In a sample of commodity-exporting economies, we find that the source of the documented forecasting ability of commodity prices lies in a revealed asymptotic dependence relationship between the two variables at a daily frequency. Our results show that timing is crucial, with exchange rates adjusting quickly to commodity price shocks. Only daily, contemporaneous observations capture both the forecasting ability of commodity prices and the asymptotic dependence between these variables, with the evidence disappearing when using lagged commodity returns or longer frequency (monthly or quarterly) observations. Our evidence suggests that the commodity market news conveying information relevant for the value of the exchange rates is transitory and short-lived.

The third essay "Commodity market spillovers? Revisiting the impact on financial markets" examines the relationship between commodity and stock markets using a sample of nine commodity-exporting economies between 2000-2019. Prior research attributes the increase of comovement between markets to the effect of contagion initiated by commodity price shocks. However, we find that the documented increase in comovement during crisis episodes does not originate from shocks impacting on commodity markets. Indeed, after controlling for the effect of time varying investor risk aversion, we do not find evidence of financial contagion between markets. Our findings suggest that controlling for the effect of time-varying investor risk aversion is a key element in accurately capturing the relationship between asset returns in these markets.

Declaration of originality

I, Francisco Pinto-Ávalos, confirm that no portion of the work referred to in the thesis has been submitted in support of an application for another degree or qualification of this or any other university or other institute of learning.

Copyright statement

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Chapter 1

Introduction

1.1 Motivation

Following the enhanced interconnectedness of the global economy, the degree of financial market integration evident in international asset markets has increased substantially (Bekaert & Mehl, 2019). This phenomenon has arisen concomitantly with enhanced international capital mobility and capital flows, with a direct impact on trading activity in foreign exchange (FOREX) markets. As a result, the global FOREX market has become increasingly active, reaching an average daily trading volume equivalent to $6.6 trillion in 2019 (Bank for International Settlements, 2019). A large literature analyses the effect of investment capital flows on FOREX markets, some adopting a macroeconomic perspective, such as a focus on the resulting current account imbalances (Frankel, 1991; Gourinchas & Rey, 2007), while other studies utilise a FOREX market microstructure approach to investigate the role of private information contained in agents' trading orders (Evans & Lyons, 2002; Hau & Rey, 2006). Financial integration has also prompted international investors to undertake investment diversification in a variety of other asset classes, such us international commodity markets. Indeed, commodity markets have increasingly captured the attention of investors (Cheng & Xiong, 2014), with a global market value totalling $300 billion during the first half of 2021 (Bank for International Settlements, 2021).

In this context, this thesis aims to address a novel set of research questions to further improve the understanding of the effect of international investment capital flows in three specific areas: the effect of the asset allocation decisions of financial investors on currency values in the FOREX market, specifically focusing on the Chilean Peso; the ability of commodity prices to forecast exchange rates in a set of leading commodity-exporting economies; and the nature of the relationship between price movements in commodity and equity markets. Recently, at the Central Bank of Chile (CBCL), economic policy related matters emanating from these areas have attracted the increasing attention of policy-makers. A pattern also reflected in other emerging markets. As an economist at the CBCL, the motivation for the research in this thesis lies in further analysing the effect of international investment capital flows on the domestic economy of emerging markets, and the consequent implications this may have for financial stability and monetary policy objectives, which are the two main mandates of the CBCL.

In our first essay, in research undertaken during my recent visit to the CBCL in 2021, we examine the role of investment recommendations made by advisory firms in the Chilean pension fund industry relating to the portfolio reallocation of investors' pension asset holdings. We demonstrate these recommendations serve as a trigger device leading to large, coordinated portfolio readjustments between domestic and overseas assets. We also show such asset reallocation has a direct impact on the Chilean Peso FOREX market.

The enhanced influence of unregulated financial advisory firms on the Chilean financial markets is currently a trending topic at the CBCL, gaining significant prominence due to its potential impact on the central bank's financial stability objectives. In this thesis, we aim to further understand the implications of such investment recommendations in exerting price pressure on the Chilean exchange rate, influencing its volatility and having an impact on the trading behaviour of other Peso FOREX market participants. Our primary contribution focuses on unveiling its impact on the FOREX market, given there has been no prior attempt to quantify and convincingly document such an effect. Importantly, our findings show there is indeed a significant impact of financial advisory firm investment recommendations on price movements in the Chilean FOREX market. Our evidence suggests unregulated financial advisory firms trigger exchange rate movements which go beyond the expected impact of fundamentals, and which may have a direct impact on financial stability. Further, this research sheds light on policy questions that the CBCL has already mentioned in its financial stability reports, although a robust quantitative analysis to underpin its concerns is absent in such reports. It is worth noting the relevance of our findings apply not only to Chile, but also have implications for other economies adopting comparable pension fund systems.

The second essay in the thesis provides further context and analysis of existing evidence which documents the ability of commodity returns to forecast those for exchange rates. Using a sample of commodity-exporting economies, our focus is upon identifying non-linear relationships between commodity and exchange rate returns and how these relationships help to explain the forecasting ability of the former for the latter. This research idea emanates from prior unpublished research undertaken during my role as an economist at the CBCL. The motivation for further analysing this topic in my Ph.D. thesis is to further understand the role of external shocks, namely news concerning international commodity markets, on the FOREX markets of emerging economies. From the CBCL's perspective, understanding the role of commodity market shocks on exchange rates provides a critical element to monetary policy matters, as exchange rate movements relate closely to imported inflation pressures from external markets. In fact, the CBCL extensively covers this topic in its monetary policy reports and considers it one of critical relevance in conducting its monetary policy.

Our main contribution to the existing literature lies in providing a novel approach, based on extreme value theory, to explain the forecasting ability of commodity returns. We find that tail dependence, a statistical indicator of the relationship between simultaneous extreme movements in exchange rate and commodity returns, is a key element in explaining the forecasting ability of commodity returns. Our results reveal that both tail dependence and the forecasting ability are transitory and short-lived phenomena, as only daily, contemporaneous commodity returns are able to provide superior forecasting ability for exchange rate returns in comparison to the usual benchmark models. Our findings are relevant for policy-makers interested in further understanding the transmission mechanism from commodity price shocks to exchange rates. For instance, for those economies moving from fixed to flexible exchange rate regimes, the conclusions of this research provide useful elements for understanding the tail behaviour of the exchange rate movements in relation to its fundamental determinants. In addition, for those economies adopting inflation-target regimes, our analysis provides a relevant perspective to analyse the elements underpinning extreme exchange rate movements which may further generate external pass-through price pressure into domestic inflation.

Our final essay revisits existing evidence to reassess the findings of recent literature investigating financial contagion between international commodity markets and the equity markets of a set of commodity-exporting economies. We believe that recent studies focusing on emerging economies evidence a lack of robust empirical approaches to this question as they omit critical factors which may explain the comovement between commodity and equity markets. In particular, many prior studies argue that increased market comovement during episodes of financial stress reflects financial contagion between commodity and stock returns. Motivated by this fact and building upon extensive prior methodological discussions, we provide what we claim to be a more robust approach to examine market contagion by including additional factors which may affect the propagation of shock across markets. In particular, unlike existing studies which conclude financial contagion exists between stock and commodity markets, we control for the effect of global investors' risk appetite and how this risk sentiment varies over time.

The motivation for this research is to further extend our understanding and interpretation of the existing financial contagion literature, which can have important consequences for economic policy-making. During my tenure at the CBCL, I frequently confronted this issue, as the Chilean economy is usually categorised as a small open economy exposed to external shocks. Therefore, contributions such as the one in this thesis, which aim to disentangle the evidence of contagion between equity and commodity markets, help policymakers to understand the source of shocks to which emerging market economies are exposed. Our main findings show that our proposed global risk appetite or risk sentiment factor accounts for the majority of the increase in measured market comovement, particularly during episodes of financial distress. In fact, after controlling for investors' risk appetite, we find no conclusive evidence of financial contagion between equity and commodity markets, suggesting the shock transmission between commodity and stock markets plays a less relevant role in comparison to those factors capturing changes in investors' risk aversion. Our findings are also relevant for portfolio managers in their quest for asset diversification opportunities. Our results suggest that equity markets exhibit a milder response to commodity market shocks than claimed in other research, therefore, important portfolio diversification opportunities are still possible across these two asset classes. Moreover, from a Central Bank perspective, understanding the scope and nature of international commodity market shocks, as they transmit to financial markets, also contributes to the monitoring of critical elements relating to any financial stability mandates.

1.2 Thesis overview and summary of main findings

Chapter 2 analyses the influence exerted by prominent financial advisory firms on the Chilean Peso FOREX market. In Chile, a group of seven pension fund companies (PFCs) manage pension fund savings equivalent to $200 thousand million dollars in 2020 (about 80% of the Chilean gross domestic product). Importantly, 45% of the PFCs' balance sheet is held in overseas markets. Since 2010 certain of these advisory firms have attracted a remarkable following and as a consequence of their popularity exerted a significant influence over decisions made by pension fund investors in Chile. One particular company, *Felices y Forrados* (F&F) has become the most influential of these advisory firms, regularly issuing recommendations to investors to actively reallocate their pension savings based on what they claim to be superior short-term investment strategies. We show that this firm's popularity among pension fund investors is such that F&F recommendations often trigger large, coordinated portfolio adjustments which translate into a reallocation of pension fund holdings between foreign and domestic assets. Motivated by these facts, in this chapter we examine the potential of F&F recommendations and the subsequent investment portfolio reallocations to generate price pressure in the Chilean peso FOREX market. Among our main findings, we conclude that F&F recommendations indeed lead to significant price pressure and enhanced volatility in the Chilean peso exchange rate. In addition, utilising a proprietary CBCL dataset, we find that other agents participating in the Peso FOREX market appear to anticipate the direction of PFC adjustments and engage in front-running the pension fund trades in this market. Our findings suggest F&F recommendations generate a significant impact on the Chilean Peso FOREX market which may not be consistent with financial stability goals. Importantly, our evidence contributes to the policy debate on the regulatory implications of financial advisory firms operating in the pension fund industry. Our results not only apply to the case of Chile, but may be relevant for other countries adopting similar pension fund advisory schemes.

Chapter 2 has benefited from valuable feedback received during the 2021 research policy seminars held at the CBCL, particularly from members of the Board. This study is forthcoming as a Central Bank of Chile working paper, and contributes to the ongoing policy discussion concerning the regulation of financial advisory firms operating in the Chilean pension fund industry. This research also benefited from feedback from participants at the annual review presentations held in the Accounting & Finance Department at Alliance Manchester Business School, University of Manchester.

Chapter 3 investigates whether the documented predictive ability of commodity prices for exchange rates applies to a broader sample of commodity-exporting economies, and further explores the frequencies at which it exists. Recent empirical studies maintain that changes in commodity prices contain a degree of predictive power for exchange rate fluctuations (Ferraro et al., 2015; Kohlscheen et al., 2017). Moreover, both the distribution of exchange rates and commodity prices, measured as log-returns, exhibit fat tails, indicating the presence of extreme values in their sample distributions. Measuring tail dependence enables the determination of the extent to which large exchange rate movements relate to their underlying fundamentals, in particular commodity prices. Using multivariate extreme value techniques, we explore the occurrences of large movements in exchange rates and commodity prices and test for tail dependence using a non-parametric measure of asymptotic dependence. We also seek to investigate whether any documented tail dependence underpins the predictive ability of commodity prices for exchange rates.

This study contributes to the literature on extreme value analysis and its forecasting implications by examining the relationship between exchange rates and commodity prices in a tail dependence framework. We believe it is the first attempt to synthesise these two strands of literature. The main results demonstrate that the predictive ability of commodity prices is indeed highly significant when we use contemporaneous daily observations. On the contrary, using lagged or lower frequency observations reduces or eliminates the ability of commodity prices to forecast exchange rates. Moreover, measured asymptotic dependence between commodity prices and exchange rates is generally only significant when prices are measured contemporaneously and at a daily frequency, while it reduces considerably when using lagged commodity prices or lower frequency observations. Collectively, our findings provide evidence showing that any measured asymptotic dependence and evidence for the claim that commodity prices can predict exchange rates are both transitory and short-lived phenomena. Importantly, our evidence suggests that the nature of the documented asymptotic dependence is a key element in evaluating the relationship between the two variables, and in particular it forms the central component when evaluating empirical claims relating to the ability of commodity prices to predict exchange rates.

Chapter 3 has benefited from the valuable suggestions received during the 2021 EFMA conference, particularly the feedback given by its organisers, John Doukas and Gianluca Mattarocci, the participants of the 2021 Research Seminars at the Central Bank of Chile, a referee of the 2019 IFABS conference and the participants in the 2019 Doctoral Seminar at The University of Liverpool Management School. Earlier versions of this chapter also benefited from the feedback received during annual review presentations organised by the Accounting & Finance Department at the Alliance Manchester Business School, University of Manchester. This study was also accepted for presentation at a number of other conferences which were subsequently cancelled due to the COVID-19 pandemic (namely, the 2020 World Finance Conference, the 10th International Conference of the Financial Engineering and Banking Society, the 2020 EcoMod conference, and the 37th International Conference of the French Finance Association).

Chapter 4 revisits the relationship between price movements in commodity and equity markets. Based on a sample of commodity-exporting economies between 2000-2019, we document the existence of time-varying correlation/spillovers between international commodity markets and domestic stock markets. We observe the link between those markets intensifies during episodes of financial distress, such as the 2007-2009 Global Financial Crisis and the 2009-2012 European sovereign debt crisis. Many prior studies argue the observed increase in comovement between stock and commodity markets is evidence of financial contagion (Creti et al., 2013; Mensi et al., 2013; Roy & Roy, 2017; Xu et al., 2019). However, our claim is that these studies implement an empirical approach exhibiting a questionable degree of robustness when measuring the existence of contagion between markets. Importantly, the omission of relevant global factors may bias the results providing misleading conclusions. Motivated by this fact and building upon the vast literature studying financial contagion, we aim to revisit the evidence of financial contagion by incorporating critical global factors which we believe play a crucial role in describing comovement between markets. Our main findings indicate that time variation in global risk aversion (risk sentiment) accounts for the majority of the comovement between commodity and equity markets, particularly during episodes of financial crisis. In fact, after including the effect of global factors, we find no evidence of financial contagion between these two markets. The contribution of this study is to help to overcome the perceived limitations of related research by highlighting the relevance of global factors in driving the comovement between markets. Our evidence suggests that omitting relevant global factors may induce misleading and potentially erroneous conclusions in classifying the nature of shock propagation between markets.

Chapter 4 has benefited from suggestions received during the internal research seminars at the Central Bank of Chile and the feedback from annual review presentations organised by the Accounting & Finance Department and the internal Ph.D. research seminars of the Economics Department, both held at the University of Manchester.

1.3 Thesis structure

This thesis consists of three related essays and conforms with the accepted journal style format of thesis presentation. Each chapter is self-contained, meaning every essay incorporates its own literature review, identifies a set of novel research questions, justifies use of an appropriate methodology and provides a discussion and analysis of results. Page numbering and sections follows a sequential order throughout the whole thesis. The three essays are co-authored with my thesis supervisors, therefore the first-person plural (we, our) is adopted throughout the thesis. As the main contributor to this research, I undertook the major role in the elaboration of the ideas developed in each of the three essays. My contribution includes the introduction and main development of the research ideas, gathering the datasets and undertaking all data management, as well as conducting all the empirical analysis and major drafting of the resulting chapters.

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Chapter 2

Financial advisory firms, asset reallocation and price pressure in the FOREX market

Abstract

Recommendations of financial advisory firms have become increasingly influential in the allocation of pension fund assets in many countries. Such recommendations often elicit large, coordinated portfolio readjustments and a subsequent reallocation of pension fund holdings across asset classes. Using a proprietary database, we analyse the potential for portfolio asset re-allocations in the Chilean pension fund industry to act as a mechanism for exerting price pressures in the Chilean peso FOREX market. We document significant price pressure and enhanced volatility in the nominal exchange rate surrounding pension fund transactions initiated by financial advisory firm recommendations. We provide evidence that other FOREX market participants seek to exploit the anticipated portfolio adjustments following such recommendations by front-running the pension fund trades. The potential for financial asset market volatility and instability this activity creates has regulatory and policy implications for the countries affected.

2.1 Introduction

The pension fund industry in Chile has experienced steady growth since the turn of the millennium, and in 2020 pension fund companies (PFCs) manage around \$200 thousand million dollars of workers' savings, which corresponds to about 80% of the Chilean GDP. During this time, this industry has become increasingly important in fostering economic growth and contributing to the development of domestic financial market activity in Chile (Corbo & Schmidt-Hebbel, 2003). PFCs are now widely recognised as among the most influential institutional investors in the Chilean economy, with recent studies analysing the effect of PFC investment decisions on trading activity and price dynamics in the country's financial markets. Both Da et al. (2018) and Ceballos and Romero (2020) find that the coordinated asset sales/purchases initiated by PFC investment decisions, generate significant price pressure in the Chilean equity and the Chilean government bond market, respectively.

Our focus is on the effect of PFCs portfolio reallocations on trading activity and price movements in the Chilean peso foreign exchange (FOREX) market. The potential importance of PFC trading activity on trading volume and exchange rates arises as PFCs hold around half of their balance sheets in foreign assets, positioning them among the most relevant agents in Chilean peso FOREX trading. Indeed, as a result of the counter-cyclical nature of their investment decisions, PFC trading activity has been ascribed with an active role in dampening excessive volatility in the Chilean peso exchange rate. Historically, during episodes of global financial distress (prosperity), PFCs have exhibited a tendency to invest in safe (risky) assets, corresponding mainly to domestic fixed income securities (foreign equities). By so doing, PFCs trigger a sale (purchase) of FOREX denominated assets and a purchase (sale) of domestic currency securities during such periods of economic downturns (expansions), generating U.S. dollar inflows (outflows) that partially offset the domestic currency depreciation (appreciation) that usually occurs this stage of the economic cycle.

More recently, evidence suggests that the inherent counter-cyclical nature of PFC investment decisions has weakened following the enhanced activity of PFCs in the FOREX market. This coincides with more active and frequent PFC asset reallocations accompanying revisions to short-term investment strategies (Opazo et al., 2014; Zahler, 2005). Indeed, pension fund investment policies seeking enhanced short-term profitability may even have contributed to more pro-cyclically aligned portfolio readjustments, in the process exacerbating asset price volatility (Levy & Zuniga, 2016; OECD, 2020). In particular, since 2011, PFC FOREX trading volume displays a higher level of volatility in comparison to previous years. The episode of greater activity of PFCs in the FOREX market coincides with the arrival of several unregulated, high profile financial advisory firms in the pension fund industry, a group that have subsequently grown in importance. One particularly influential advisory firm, *Felices y Forrados* (F&F), makes recommendations to pension fund investors to actively trade and reallocate their savings based on short-term investment strategies.[1] F&F justifies this advice on the (unfounded) claim that such behaviour generates higher investor returns in comparison to more passive investment strategies, such as buy-and-hold. Since F&F began making recommendations, investors have become more active in switching between their investment portfolios in the pension fund system. Concomitantly, PFC trading activity in the FOREX market has also increased in both magnitude and amplitude, potentially exacerbating rather than helping to mitigate exchange rate volatility.

This study's contribution is to analyse whether F&F recommendations influence either the nature or magnitude of PFC FOREX trading activity, and to determine any potential effects on the pricing dynamics evident in the Chilean peso exchange rate. To the best of our knowledge, no previous evidence documents the effect of financial advisory firms' recommendations on pricing dynamics in the Chilean FOREX market. Such an analysis is revealing not only because of the specific characteristics of the Chilean pension fund industry, but also from a wider international perspective. First, in the Chilean context, since F&F started to make recommendations, pension fund companies have ranked among the biggest institutional investors in Chile making large, coordinated trades in the Chilean peso sector of the FOREX market. As we later document, these PFC FOREX transactions generate significant price pressures on the Chilean exchange rate, and increase its volatility. Second, current regulations in the pension fund industry incorporate legally binding procedures which serve to delay the effective date when pension fund companies can execute asset sales/purchases in the market following receipt of investor mandates to readjust their portfolios. Using a proprietary database of daily FOREX market trading volume, disaggregated by type of agent, we find this delay in trade execution generates strategic trading complementarities where other FOREX market participants benefit from front-running the anticipated PFC portfolio realignment trades.

Examining the potential effect of financial advisors on the FOREX market is of significant interest from an economic policy and financial stability perspective. Previous studies highlight the role of PFCs and the close link between the pension fund industry and the FOREX market, noting that large pension fund flows may pose a threat to the stability of the Chilean

[1] *Felices y Forrados* translates from Spanish to English as Happy and Loaded. Section 2.3 documents the evidence identifying the influence of F&F recommendations upon investor decision-making.

peso FOREX market (Marcel, 2020). Zahler (2005) argues the herd behaviour which characterises PFCs asset reallocation decisions generates significant portfolio flows that may affect the exchange rate, although the study does not quantify the impact of such flows. Central Bank of Chile (2020) discusses how pension fund re-allocations impact asset trading volumes in the Chilean fixed income market. In an international context, OECD (2020) highlights how large, coordinated pension fund readjustments in domestic financial markets may impact asset prices and exacerbate FOREX market volatility. These studies mirror Raffnsøe et al. (2016) who find that the investment decisions of the pension fund companies in Denmark exhibit a significant impact on the Danish krone. The authors discuss the implications of the role of PFCs on the Danish FOREX market and their impact on the exchange rate policy implemented by the Central Bank of Denmark. In this context, our study aims to provide policy-relevant insights which can be used by Central Banks and other regulatory authorities to examine the role played by unregulated financial advisors in triggering asset price movements beyond those mandated by macro fundamentals and further exacerbating asset price volatility. This may possibly lead to greater formalisation and regulation of their activity in financial markets. In comparative terms, financial advisor regulation and associated policy considerations often take place earlier in more developed economies (Hung et al., 2008; Inderst & Ottaviani, 2012). In emerging economies, however, there is often no formal quantification of the potential impact of financial advisors in FOREX markets, and our study aims to provide insights into this important area of activity. Our results contain useful discussions, not only for the Chilean economy, but also for other many other countries which adopt similar pension fund systems.[2]

Our study relates to research analysing the role of institutional investors, highlighting the mechanism through which large, coordinated investment decisions generate sustained price pressure on financial markets. For instance, recent studies document the herding behaviour of institutional investors and the impact it generates on the U.S. corporate bond market (Cai et al., 2019; Ellul et al., 2011; Goldstein et al., 2017). Others focus on the effect of large institutional investors on the U.S. stock market (Gompers & Metrick, 2001; Khan et al., 2012) and the Israel stock markets (Ben-Rephael et al., 2011). From a theoretical perspective, Basak and Pavlova (2013) show that institutional investors' trading decisions exert pressure on the prices of their benchmark equity indices, generate excessive correlation among stocks and increase equity market volatility. More specifically in the context of our research, Greenwood and Vayanos (2010) analyse the case of pension funds and government bonds in the U.S. and the U.K. Similarly, Froot and Ramadorai (2005), based on a sample of eighteen currencies, find that institutional investors play a key role in explaining exchange rates movements in the short-term. While most of the research in this field covers mature financial markets, a few recent studies focus on the case of Chile, paying special attention to whether F&F recommendations influence price movements in domestic financial markets. Da et al. (2018) and Ceballos and Romero (2020) examine their effect on the stock and government bonds markets, respectively. Both studies find that the large, coordinated sales or purchases

[2]Other countries adopting a defined-contribution pension fund scheme in which pension investors can freely choose the level of risk associated with different portfolio allocations are Colombia, Costa Rica, Slovakia, Slovenia, Estonia, Latvia, Lithuania, Mexico, Peru, Poland and Romania.

which occur following F&F recommendations generate significant price pressure within the respective market under study. We extend the scope of these studies to the Chilean FOREX market.

2.2 The Pension fund industry in Chile: Institutional Context

The Chilean pension system is a defined-contribution scheme which compels employees, hereafter (pension fund) investors, to allocate 10% of their wages to their individual pension saving accounts. Pension fund companies (PFCs), private institutions created by law in 1980 and regulated by the *Superintendencia de Pensiones de Chile* (Chilean regulatory body of the pension fund industry) are in charge of managing pension fund savings accounts on behalf of investors. As figure 2.1 shows, the aggregate savings managed by the PCFs exhibit a steady growth following their creation, totalling around $200,000 million U.S. dollars (USD) by 2020, which represents around 80% of the Chilean GDP.

[Figure 2.1 in here]

Currently, seven PFCs operate in the pension fund market in Chile, charging a management fee equivalent to a percentage of an investor's monthly income. Investors can switch from one company to another with no exit fees. A 2002 regulation requires each PFC to offer five types of pension fund portfolios, from which investors can choose up to two portfolios within the same PFC to allocate their pension savings. This regulation aims to provide investment flexibility to investors by enabling them to select portfolios according to their risk preferences. Table 2.1 provides details of the five portfolios, labelled A (highest risk) to E (lowest risk), that PFCs make available for investors to allocate their mandatory savings. Panel A reveals the total USD value of each portfolio in 2020, with portfolio C the medium risk portfolio being the largest. Panel B presents the asset composition of each portfolio, considering both the type of investment assets (equity and fixed income) and their location (domestic or overseas). Portfolio A is characterised as the riskiest portfolio since its investments are focused towards equities and the majority of its asset allocation (84%) is in overseas markets. In contrast, portfolio E provides the safest investments, allocating most of its pension assets (88%) into domestic, fixed income markets.

[Table 2.1 in here]

Current regulations enforce a number of legally binding requirements restricting the asset composition of each of the PFC portfolios. First, the regulations enforce specific limits to equity allocation within each portfolio. Panel C in table 2.1 shows that portfolio A, the riskiest portfolio, may invest up to 80% and no less than 40% of the total portfolio value in equity, while, portfolio E, the safest portfolio, must invest no more than 5% in equities (with no specified minimum). Portfolios B, C, and D correspond to intermediate risk exposure alternatives lying between funds A and E. These legal limits attempt to ensure that funds are differentiated from each other based on their risk exposure. Second, current regulations also penalise PFC portfolio underperformance in comparison to the average returns of the rest of the PFCs. On the basis of these legal requirements, it is perhaps not surprising to discover that specific portfolios across PFCs hold similar compositions of assets in an attempt to avoid their returns departing significantly from the average of other PFCs. It is claimed these sets of regulations generate a pattern of herd-type behaviour in PFC investment decisions (Raddatz & Schmukler, 2008). Third, the regulations also make it compulsory for PFCs to hedge their currency exposures. In particular, PFCs must hedge their currency risk by selling FX forward contracts after buying any FOREX (predominantly USD) in the spot market. Fourth, in terms of enacting investor-mandated portfolio reallocation, PFCs can only execute the portfolio shifts from the fourth working day following receipt of the investors' instructions.[3] Fifth, the regulations also state that PFCs cannot process switching portfolio reallocation requests accounting for more than 5% of the portfolio value on the same day. Any reallocation exceeding 5% of the total portfolio value takes place on the following working day, processed on a first-come, first-served basis.[4] Given the trading delays imposed by these restrictions and the similarities in portfolio composition, we conjecture this may generate incentives for other FOREX market participants to front-run any anticipated coordinated PFC asset sales/purchases. Further, these restrictions also influence investors to act quickly to request changes to their portfolio in an attempt to obtain more favourable prices before other participants trade.

A major consideration in the context of this study relates to the fact that a high proportion of the PFCs balance sheet corresponds to overseas assets. Figure 2.2 shows that PFCs invest around 40% of their assets in overseas markets and that this proportion remains roughly constant over the last ten years. Considering this evidence, and taking in account the relatively large size of pension fund savings, the PFC investment decisions position PFCs as one of the most relevant institutional agents in the Chilean FOREX market. However, the precise impact of the asset sales/purchases by PFCs on the exchange rate depends on the type of the transaction. This is because of the binding requirements that regulations enforce on PFCs in terms of hedging currency risk. For instance, when a pension investor chooses to take a greater exposure to risky assets she instructs her PFC via a switching request to reallocate savings into a portfolio which contains a greater proportion of foreign assets, say portfolio A. Subsequent to this request, the PFC sells domestic assets and uses the sale proceeds to purchase foreign currency in the spot FOREX markets, which is invested in foreign currency denominated risky assets, generating capital outflows from the domestic economy. If this hypothetical portfolio reallocation scenario to riskier asset portfolios is repeated on a large, coordinated scale across several PFCs it will generate a notable increase in overall PFCs demand for foreign currency, typically USD, in the FOREX spot market. If the ensuing order flow is sufficient, this may translate into depreciation pressures on the peso in the spot market.

[3] This delay is justified on the basis that such requests may contain clerical errors, so PFCs use this time window to evaluate the instruction's accuracy and feasibility.

[4] The 5% rule applies to any of the portfolios, either the current portfolio or the one being invested in. This measure was introduced in response to the notion that PFCs may unwittingly initiate undesirable, risky economic and financial developments due to the large value of the pension funds under management (Zahler, 2005).

[Figure 2.2 in here]

However, one also needs to account for the mandatory hedging of currency risk that regulations specify PFCs must undertake, which may serve to partially mitigate the exchange rate depreciation. These regulations require PFCs to offset their purchases of USD in the spot market by selling the dollars forward, which generates appreciation pressures on the value of domestic exchange rate acting to reduce the original tendency towards depreciation. These FOREX hedging regulations also create a potential asymmetry, in that following large coordinated PFC investor requests to switch to less risky portfolios, say portfolio E, PFCs will sell external assets and use the FOREX proceeds to purchase domestic currency in the spot market which is subsequently invested into domestic fixed income assets. This process generates FOREX (mainly USD) inflows to the Chilean economy, which if sufficiently large will induce an appreciation of the domestic currency. In this case, no binding requirement in relation to currency hedging strategies exists, so any appreciation pressure induced by spot sales of FOREX can be fully transmitted to the domestic currency exchange rate. We later document the empirical importance of these effects, but we now proceed to provide context by describing the role of financial advisory firm recommendations as a trigger of pension portfolio readjustments in the pension fund industry.

2.3 Unregulated Pension Advisory Companies: Felices & Forrados

The period since 2010 has witnessed sustained growth in a number of unregulated pension advisory companies operating in the Chilean pension market. One such firm, Felices & Forrados (F&F), has been actively operating in this market since July 2011. For an annual fee (in 2020 this is $30 USD) F&F provides recommendations to investors via email subscription, advising clients into which of the five specific PFC portfolios available they should channel their investments. These recommendations contain little detailed market analysis and relatively few additional explanations justifying the particular strategy F&F recommends. Such pension fund advice is the main service F&F provides to its subscribers, as the firm itself does not manage investors savings. Based on its aggressive marketing strategies, F&F has gained remarkable prominence and popularity in the past ten years, consistently claiming that investors will be better off by following their recommendations in comparison to undertaking alternative investment strategies, such as passively buying and holding a specific PFC portfolio.[5]

[Figure 2.3 in here]

Two striking facts are evident since F&F began making its recommendations in July 2011. First, the number of portfolio readjustments in PFC portfolios dramatically increases in com-

[5] In addition to F&F, three other unregulated financial advisory firms offer similar advisory services in the pension fund market. These are (the year recommendations commence in parentheses), *Fondo Alerta* (2008), *Tiempo para ganar* (2012) and *Previsionarte* (2013). These three companies, however, receive significantly less media attention and they have considerably fewer followers than F&F. Figure A.2.1 in the appendix presents Google Trends data reporting investor interest over time. Based on these Google searches, F&F is by some distance the most popular advisory firm.

parison to previous years. Figure 2.3 displays the net portfolio flows by portfolio through time at the aggregate pension fund industry level. Panel (a) shows a notable increase in portfolio reallocations after F&F begins providing investors with recommendations in mid-2011 (represented by the vertical line). This increased activity is noteworthy even in comparison to previous episodes of severe financial distress, such as the 2008 global financial crisis, and intensifies at the end of 2019 during the episode of civil unrest in Chile. Second, the greatest amount of portfolio switching coincides with F&F recommendation dates. In panel (b) we present the net portfolio switches during 2011-2020, which clearly appear to increase on days F&F makes recommendations (represented by the vertical dotted lines) and tend to remain relatively high for a few days immediately thereafter. Since 2019, when F&F starts making recommendations more often, the average 5-day cumulative value of portfolio switches after recommendations sums to between 15 and 20% of the average value of portfolio E, the least risky portfolio.[6] The observed persistence in portfolio switches after recommendations is a consequence of the previously noted regulatory delay relating to the requirements imposed on PFCs when processing portfolio switch requests. In addition, the spread of recommendation information from F&F subscribers to non-subscribers potentially reinforces this effect. F&F popularity has increased during recent years, despite the Chilean pension fund regulatory body explicitly providing evidence which demonstrates that investors would have been financially better off if they had not followed F&F recommendations.[7] However, claims of outstanding initial performance along with successful media marketing campaigns has kept investors engaged with implementing F&F recommendations. F&F advertisements regularly appear on the internet (social media) and as table A.2.1 in the appendix reveals, F&F followers tend to be somewhat younger and wealthier than the average non-F&F follower. This evidence is consistent with studies documenting that investors hire financial advisors based on elements such as persuasive advertising, familiarity and so-called 'schmoozing', rather than superior financial performance (Gennaioli et al., 2015), and continue to do so even after advisors recommendations lead investors to significantly underperform the market (Foerster et al., 2017). In the next section, we briefly discuss relevant aspects of the Chilean Peso FOREX market and describe the role of F&F recommendations in eliciting PFCs' FOREX trading activity.

2.4 The Chilean peso FOREX market, F&F recommendations and PFC trading

In 2019 the trading volume on the Chilean peso FOREX market totalled around $1,400 billion USD, approximately seven times the Chilean GDP. The depth of the Chilean peso

[6]While portfolio E is not the largest portfolio, it invests a higher proportion of its assets in the Chilean economy. This provides some perspective on the size of the portfolio switches triggered following F&F recommendations.

[7]Since 2013, the *Superintendencia de Pensiones de Chile* (the regulator of the pension fund market in Chile) shows that pension fund investors following F&F recommendations exhibit lower returns in comparison to those taking passive investment strategies, such as buying and holding a specific PFC portfolio (Superintendencia de Pensiones de Chile, 2013, 2020, 2021). The initial popularity of F&F may arise from the apparent success claimed for its early recommendations, which outperform alternative investment strategies during the first year of its operation. However, this claim seems to be spurious as the alleged outperformance is not statistically significant (Da et al., 2018).

FOREX market, measured in terms of the market's size relative to GDP, is lower in comparison to advanced economies (which average around 40 times GDP), albeit it is higher than other economies in Latin American (averaging around 2.8 times GDP). Liquidity in the Chilean FOREX market has remained stable since the early 2000s and is similar in magnitude to other emerging economies in the region, albeit again lower in comparison to advanced economies.[8] Chilean FOREX market activity is mainly allocated across the spot, forward and interbank swap/repo markets, with a very small residual in other derivatives contracts (futures and options). Spot (immediate delivery) trading volumes are around $460 billion USD, with currency forward and FX swaps comprising 95% of the remaining $940 billion USD of trading activity in 2019. As is customary, the most traded counterparty currency to the Chilean peso is the USD, accounting for between 90% to 95% of trades, depending on the market segment.[9] Panel (a) figure 2.4 presents the trading volume in the Chilean FOREX spot market disaggregated by type of agent. The figure reveals the main agents participating in the market during the late 2000s correspond to retail and exporting companies, together with wealth management firms and mutual funds. This pattern holds throughout the sample although the participation of wealth management firms and mutual funds shrinks toward the end. The trading volume of retail and exporting companies broadly involves transactions relating to the international trade of goods and services, while wealth management and mutual fund activity reflect private investment flows in globally diversified portfolios.

[Figure 2.4 in here]

More recently, however, PFCs have increased their active participation in the FOREX market, largely reflecting USD transaction in response to investor portfolio reallocation requests. Indeed, figure 2.5 highlights the increase in PFC trading volume after F&F start making recommendations, and by 2019 PFC trading volume represents 25% of the total trading activity in the FOREX spot market, just behind retail and export companies which together account for 28% of the total trading volume. As depicted in Panel (b) figure 2.4, the main agents active in the forward and FX swap sector of the FOREX market correspond to non-residents and pension fund companies, with around 50% and 25% participation on average since 2019, respectively. The non-resident trading volume relates mainly to foreign banks engaging in interbank swaps. It also includes foreign investors undertaking speculative carry trade strategies using FX derivatives. The pension funds trading volume captures the mandatory currency hedging obligations imposed upon PFCs in accordance with the regulations we previously discuss.

[Figure 2.5 in here]

One issue which we address subsequently is whether the increased portfolio reallocation in the pension fund system following F&F recommendations may exacerbate exchange rate

[8] See Villena and Hynes (2020) who follow the BIS quarterly reporting standards to compute FOREX market depth as trade volume as a proportion of GDP and market liquidity as the bid-ask spread in the FOREX market.

[9] Based on 2019 values, currency trades denominated in euros represent only around 7% and 1% in the spot and other markets, respectively. The remaining currency trades (less than 3%) correspond to other global currencies.

volatility. Table 2.2 presents some preliminary facts relating to volatility, comparing the standard deviation of changes in both the exchange rate and the PFC net trading volume in the FOREX market during different time periods. The first column in table 2.2 reveals that exchange rate volatility increases during the 2008 global financial crisis (GFC) in comparison to the pre-crisis period and remains high after F&F starts issuing recommendations, although falling somewhat from its crisis level. Columns two and three indicate that the volatility of PFC net trading volume in the FOREX spot and forward markets increases during the period of F&F recommendations, being even higher than the amplitude witnessed during the financial crisis.

[Table 2.2 in here]

In section 2.5.2, we explore this relationship more systematically controlling for additional risk factors which may influence the documented relationship. Overall, these preliminary observations suggest F&F recommendations do influence investors' portfolio reallocations and may be relevant as potential catalyst initiating large, coordinated PFC transactions in the Chilean FOREX market. Importantly, the higher volatility in PFC trading volumes since 2019 raises questions in relation to the counter-cyclical role of PFC trades in this market. We later show PFC USD trades not only generate significant price pressures in the Chilean exchange rate, but also they motivate other FOREX market participants to front-run these coordinated PFC transactions. This FOREX evidence we uncover is consistent with two related papers analysing the effect of F&F in other asset markets. Da et al. (2018) find F&F recommendations generate price pressure in the Chilean equity market and, additionally, provide evidence of other market participant front-running PFCs trades in this market. Similarly, Ceballos and Romero (2020) document that F&F recommendations generate price pressure in the Chilean bond market.

2.5 The effect of F&F on the Chilean FOREX market

The prima facie evidence we report in the previous sections suggests F&F recommendations are associated with an increasing number of PFC portfolio switches and enhanced trading volume in the Chilean peso FOREX market. However, to uncover the true nature of the impact of F&F recommendations on the FOREX market, we attempt to identify the news contained within the recommendation announcements to subsequently investigate if F&F recommendations induce price pressures or enhance volatility levels in the market. In this section we undertake four strands of analysis. First, we use an ordered logit model to establish which, if any, economic factors trigger F&F recommendations. Understanding the drivers of their recommendations is a relevant part of our identification strategy, as it allows to capture the specific shock component of F&F announcements on the Chilean peso FOREX market. Second, having identified the news component in F&F recommendations, we use the local projection method to explore the nature of price pressures on the Chilean nominal exchange rate. Third, utilising the same methodological framework, we analyse the impact of F&F news on exchange rate volatility. Finally, we investigate whether F&F recommendations initiate actions by other FOREX market participants, possibly in an attempt to front-run trades arising from the anticipated pension portfolio readjustments.

The data consist of both proprietary and publicly available information obtained from the Central Bank of Chile at a daily frequency. Data on the Chilean pension fund industry, such as, the value of the pension fund industry and the pension investment portfolios are available on the website of the Chilean pension fund regulatory body.[10] Data relating to macroeconomic and financial variables, such as nominal exchange rates, interest rates, VIX, S&P 500 returns, Chilean government bond returns, domestic activity and inflation expectations, and terms of trade are obtained from Bloomberg. Data detailing the daily trading volume by agent in the Chilean Peso FOREX market is a proprietary dataset from the Central Bank of Chile. The sample period under analysis covers the period from 01 March 2012 to 22 October 2020.[11]

2.5.1 Identifying the F&F recommendation news

To understand whether F&F recommendations impact the FOREX market, it is necessary to identify the news (unanticipated shock component) contained in the recommendation announcement. It is also important to ensure that the news pertains to the F&F recommendation and not to other alternative factors which may both impact the exchange rate and influence the recommendation. In this section, we investigate the factors triggering F&F recommendations to later estimate an empirical exchange rate model which includes exchange rate fundamentals and also variables affecting the probability of F&F making a recommendation which may induce exchange rate movements. To establish the factors which may trigger F&F recommendations requires an understanding of the nature of such recommendation announcements. Essentially, F&F recommendations suggest reallocating pension savings after considering the appropriateness of the differentiated investment risk exposure of the PFC portfolios in the current economic environment. The principle F&F follows to deliver recommendations lies in its short-run market timing claim, based on its ability to assess economic/financial risks in the global and domestic economy. However, we note that the F&F definition of its market timing claim is somewhat inconsistent over time, since it iterates between "maximising pension fund profitability" and "reducing the loss of value of pension funds", two goals that are not necessarily compatible and may even call for differentiated investment strategies. As a result, the brief explanations F&F provides to underpin their recommendations accommodate varying circumstances, making it challenging to identify the supporting rationale behind their recommendation announcements.

Specifically, F&F does not disclose the risk assessment model it employs to gauge the overall state of the global/domestic economy. Instead, F&F releases recommendations to subscribers and provides some limited reasoning to contextualise its advice. The most common

[10] Information available on the *Superintendencia de Pensiones de Chile* website (www.spensiones.cl)

[11] Although F&F starts making recommendations in July 2011, domestic economic uncertainty index data is available only from February 2012. Further, Da et al. (2018) note that F&F does not gain marked popularity until early 2012 and Cuevas et al. (2016) document the number of F&F subscribers in 2011 is significantly lower in comparison to the period starting in 2012. Hence, omitting the first four F&F outlier recommendations will have negligible impact on our findings.

factors F&F refers to correspond to: (i) recent economic/financial risks, and (ii) recent developments in global equity and Chilean fixed income markets. In particular, F&F often refers to both the $S\&P500$ and the Chile government bond market performance as the main elements underpinning its market analysis. Figure A.2.2 in the appendix shows that $S\&P500$ returns (panel a) and domestic government bond returns (panel b) are highly correlated with returns of the riskiest and the safest PFC portfolio, respectively. Therefore, a priori, the performance of these markets appears to represent an important component of F&F's risk assessment and constitutes a critical element in understanding their decision to publicise a recommendation.

2.5.1.1 Predicting the content of F&F recommendation announcements

These elements underpin our decision to consider the outcome of an F&F risk assessment exercise as equivalent to an unobservable latent variable, which emanates from F&F's true model as follows:

$$\Delta Y^{*} = X\beta + \varepsilon \tag{2.1}$$

where the vector ΔY^{*}, the unobservable latent variable, represents the change in the F&F risk assessment. X is a vector of variables corresponding to the factors F&F includes in its risk assessment, β is a vector of coefficients to be estimated, and the vector ε is a zero mean, random disturbance term which follows a standard logistic distribution. While ΔY^{*} is an unobservable variable, we observe F&F recommendations. We assume F&F recommendations are a function of the latent variable (i.e.: variations in F&F risk assessment) as follows:

$$Y = \begin{cases} \text{strong} & \text{if } \kappa_2 \leq \Delta Y^{*} \\ \text{moderate} & \text{if } \kappa_1 \leq \Delta Y^{*} < \kappa_2 \\ \text{no recommendation} & \text{if } \Delta Y^{*} < \kappa_1 \end{cases} \tag{2.2}$$

where Y is a vector containing the observed F&F recommendations and the κ_j's, $(j = 1, 2)$, are scalars representing the threshold points of the latent variable. According to equation 2.2, the change in F&F risk assessment determines the intensity and the direction of its recommendations. For instance, a substantial (slight) increase in F&F risk assessment triggers a recommendation suggesting a strong (moderate) change in investment risk exposure towards less risky portfolios. Conversely, a substantial (slight) decrease in F&F risk assessment triggers a recommendation suggesting a strong (moderate) change in investment risk exposure towards riskier portfolios. Otherwise, marginal variations in F&F risk assessment outcomes lead to suggestions of no changes in risk exposure and effectively no recommendations for portfolio readjustment.

We categorise F&F recommendations in terms of the suggested change in risk exposure as follows: A strong change in risk exposure ($y_t = 2$) occurs when a recommendation suggests changing to one extreme portfolio conditional on the existing recommendation allocating investments within the opposite extreme portfolio. For instance, a strong change in risk exposure occurs when F&F recommends allocating either 100% or a fraction of pension savings into portfolio A (E), the riskiest (least risky) portfolio, given the current recommendation is to allocate either 100% or some fraction of the savings into portfolio E (A), the least risky (riskiest) portfolio. We define moderate changes in risk exposure ($y_t = 1$) as those recommendations which suggest an increased allocation to intermediate portfolios (i.e.: portfolios B, C or D), when existing recommendations involve an extreme portfolio allocation. For instance, a moderate change in risk exposure occurs when F&F recommends a saving allocation of 50% into portfolio C and 50% into portfolio E, when the current recommendation is 100% into portfolio E. No change in risk exposure ($y_t = 0$) occurs on a day with no recommendations.

Using daily observations from 01 March 2012 to 22 October 2020, we estimate an ordered logit model to test whether the variables F&F usually cites as underpinning its recommendations actually serve as drivers of the probability of F&F delivering a specific recommendation. The ordered logit model is as follows:

$$P(Y_i > j) = \frac{\exp\left(X\beta_i - \kappa_{i,j}\right)}{1 + \exp\left(X\beta_i - \kappa_{i,j}\right)} \tag{2.3}$$

with $i = [more\ risk,\ less\ risk]$ and $j = [1, 2]$. Y_i corresponds to a time-series, ordered categorical variable capturing both the direction and the intensity of F&F recommendations. In terms of the direction of F&F recommendations, following Da et al. (2018) we estimate the ordered logit model in equation 2.3 separately for those sets of recommendations advocating taking more and less risk exposure ($i = more\ risk, less\ risk$). The intensity of F&F recommendations determines whether the ordered dependent categorical variable takes a value of 1 or 2, corresponding to situations when F&F recommends moderate or strong changes in risk exposure, respectively, and zero otherwise.

We note that our classification departs from Da et al. (2018), given the majority of their analysis focuses on the first fifteen recommendations in their sample, each of which is a strong switch either from portfolio A to E or E to A. However, from their sixteenth recommendation in March 2014, F&F starts to advocate investor allocations to portfolios with intermediate risk exposure (i.e.: portfolios B, C, and D) and splitting savings across more than one PFC portfolio. In our estimation we include all 93 F&F recommendation announcements in the period 01 March 2012 to 22 October 2020, and the classification we propose is able to account for variations in their inherent risk exposure.[12] The vector X represents the set of explanatory variables consisting of the factors F&F often refers to when proposing its recommendations. Following Da et al. (2018), to capture any short-term trend in these variables during the previous month, we include four lags of the cumulative weekly returns of the Chilean nominal exchange rate ($\Delta usdclp$), S&P 500 ($\Delta S\&P500$) and Chilean government bonds ($\Delta Bond$).

[12]Due to data availability (see footnote 11), we exclude the first four F&F recommendations. Table A.2.2 in appendix displays F&F recommendation dates along with the suggested portfolio alongside the classification we use in this section (see column 'Ologit').

In addition, for the purpose of capturing short-term recent economic and financial risks during the previous week, we include five lags of daily changes of domestic inflation expectations ($\Delta\pi$), domestic economic uncertainty (ΔDEU) and VIX index (ΔVIX). Chilean domestic economic uncertainty corresponds with the economic uncertainty measure of Becerra and Sagner (2020). The index tracks economic-related uncertainty based on daily media news coverage. An increase in the index indicates a higher degree of economic uncertainty. Domestic inflation expectations corresponds to the break-even inflation computed as the yield difference between the 2-years nominal government bond and the 2-years inflation-linked government bonds. This inflation measure is available at a daily frequency and it is widely used by Central Banks to track inflation expectations at a high frequency. β_i is a vector of coefficients and $\kappa_{ij},\ (j = 1, 2)$ are scalars representing the thresholds of the latent variable. We estimate the model using the maximum likelihood method.

2.5.1.2 Results

Table 2.3 displays the results of the ordered logit model estimation. The dependent variable of the model in column 'more risk' ('less risk') corresponds to the ordered categorical variable capturing the intensity of F&F recommendation to reallocate investment funds to more (less) risky portfolios.

[Table 2.3 in here]

We find that lagged exchange rate returns exhibit no statistically significant impact on the probability of F&F making a recommendation to re-allocate risk. While positive performance of S&P 500 returns during the previous week significantly decreases the probability of F&F recommending an adjustment to less risky pension portfolios, Chilean government bond returns have no explanatory power for the probability of F&F issuing any recommendation. This result provides some support to the belief that F&F follows short-term trends in equity markets when issuing pronouncements. We also find statistical support for the position that factors capturing short-term economic and financial risks contribute to explain the probability of F&F recommending portfolio risk-adjustments. An increase in expected inflation significantly reduces (increases) the probability of F&F recommending riskier (safer) portfolios. This is consistent with the idea that higher expected inflation makes inflation-linked bonds more attractive. Therefore, a higher expected inflation may lead a rebalancing strategy to safer portfolios allocating most of their assets in fixed income assets at the cost of a lower exposure to risky portfolios.[13] Both enhanced domestic economic uncertainty and increases in global risk aversion significantly reduce the probability of F&F recommending riskier portfolios. The fact that the estimated latent variable thresholds $(\kappa_j,\ j = 1, 2)$ exhibit high significance confirms our choice of the ordered categorical variable (Y) definition given in equation 2.2. Moreover, the statistically insignificant χ^2 statistic testing the parallel regression assumption in both the 'more risk' and 'less risk' models indicates that this assumption

[13] Safe PFC portfolios, particularly portfolio E, mainly allocate assets to both nominal and inflation-linked bonds.

is not violated.[14] This enhances our confidence that the risk exposure classification we implement to categorise F&F recommendations not only captures their economic underpinnings, but it is also validated from a statistical perspective.

This set of results leads to the following conclusions. First, lagged exchange rate returns do not statistically influence the probability of F&F making recommendations. This finding is important for our purposes, as such evidence helps to mitigate endogeneity concerns relating to possible reverse causality issues in the estimation we introduce in section 2.5.2 to model price pressures in the Chilean Peso FOREX market. Second, our findings suggest that to some degree, short-term changes in fundamental economic and financial drivers influence the decision making process of F&F. In particular, factors capturing daily economic and financial risks (VIX, inflation expectations and economic uncertainty) play a primary role in comparison to financial market performance, although short-run equity returns (S&P 500) are also important. Third, despite the statistical evidence, the relatively low explanatory power of the predictive logit regression, as evidenced in the pseudo R2 in table 2.3, indicates there is still a large unexplained component to the F&F recommendation announcements. This suggests that the decision-making process of F&F is also governed by non-fundamental elements captured in the stochastic disturbance term (ε) in equation 2.1. Consequently, F&F recommendations seem to be somewhat arbitrary and less informative of important economic fundamentals, in this sense conveying noisy information to investors. This conclusion is not surprising as it is corroborated by Da et al. (2018) who also document that fundamental variables tend to display weak explanatory power for F&F recommendations. Indeed, evidence indicates the lack of informativeness in F&F recommendations negatively impacts the value of the pension savings of F&F followers, which reveal inferior financial performance when compared to the pension savings of investors who do not follow F&F (Superintendencia de Pensiones de Chile, 2013, 2020, 2021).[15]

Two other important implications emerge from the previous analysis. First, the noisy process generating F&F recommendations based on short-term investment strategies appears to exacerbate the frequency and volatility of portfolio switches, as panel (b) of figure 2.3 depicts. As we later demonstrate, this observed higher volatility in switching between portfolios, which is triggered by F&F recommendations, relates to enhanced exchange rate pressures in the peso FOREX market and increases in exchange rate volatility. Second, the evidence suggesting that fundamental drivers only tangentially influence the F&F decision-making process leads us to interpret the news contained in F&F recommendations as an exogenous shock. From a statistical point of view, this helps to mitigate endogeneity concerns, as the news contained in F&F recommendations is unlikely to be correlated with the error term of the

[14] The null hypothesis in the parallel regression assumption test states there is no statistical difference in the coefficients between models using an alternative binary definition of the dependent variable, such as a model where the dependent variable takes the value of one in the highest category and zero otherwise in comparison to a model where the dependent variable takes the value of one in the second highest category and zero otherwise. In a simplified additional exercise, instead of considering the intensity dimension to categorise the recommended change in risk exposure as in equation 2.2, we treat all recommendations as if they display the same extent of intensity. Under this simplified version of the model, we define a dependent categorical variable which takes the value of 1 (-1) when a recommendation suggests a higher (lower) risk exposure, and zero otherwise. We include the same explanatory variables displayed in equation 2.3. The outcome of this simplified exercise (available upon request) provides similar results in comparison to the model we estimate in this section.

[15] This evidence sparks a debate about the benefits of following noisy F&F recommendations which suggest utilising short-term investment strategies, as opposed to other strategies focused upon generating longer-term profitability.

exchange rate model we introduce in section 2.5.2.

2.5.2 Exchange rate pressure

This section uses a time-series framework to analyse how the magnitude and the persistence of F&F recommendation announcements impact upon the Chilean peso exchange rate. We employ the local projection method (LPM) proposed by Jordà (2005), a methodology which allows the effect of F&F recommendations to be tracked over time through its employment of impulse-response functions. This is especially useful since we expect the impact of F&F recommendations to persist for some days after their actual issuance, given that existing regulations introduce time delays before allowing the PFCs to processes portfolio switching requests. The dissemination of recommendation announcements from F&F subscribers to non-subscribers may also reinforce the persistence of such recommendations. Consequently, we expect some delay following F&F recommendations in the impact of PFC portfolio readjustments on the Chilean peso FOREX market.

2.5.2.1 Local projection model

Our benchmark empirical model in this section follows that of Contreras et al. (2013) who analyse the 2011 Central Bank of Chile FOREX market intervention. In addition to the explanatory variables in Contreras et al. (2013), we include F&F recommendation announcements to project the impact of its news component on the Chilean nominal exchange rate through time as follows:

$$\Delta s_{t+h} = \alpha^h + \beta^h F\&F_t + \sum_{i=1}^{2} \gamma_i^h \Delta s_{t-i} + \sum_{i=0}^{2} \delta_{k,i}^h x_{k,t-i} + \sum_{i=0}^{2} \theta_{m,i}^h z_{m,t-i} + \varepsilon_t^h \tag{2.4}$$

where Δs_{t+h} corresponds to the nominal exchange rate return between $t-1$ and $t+h$, with $h = 0, \ldots, 30$, and t being the day on which F&F issues a recommendation announcement. $F\&F_t$ is the variable capturing the recommendation's impact. Table A.2.2 in appendix displays details of F&F recommendations along with the value taken by $F\&F_t$. We define $F\&F_t$ as the first difference of $finv_t$, where $finv_t = \sum_{i=1}^{5} w_{it} p_{it}$, with $i = 1, 2, 3, 4, 5$ (the five PFC portfolios), w_{it} represents the recommended F&F allocation in portfolio i at time t, and p_{it} represents the percentage of foreign assets invested in portfolio i at time t. $F\&F_t = 0$ during days with no recommendations. By definition, the $F\&F_t$ variable captures the direction and the magnitude of the F&F recommendation announcements on the nominal exchange rate. This variable's construction aims to quantify the pressure PFCs generate in the FOREX spot market, as it captures the proportion of foreign assets in the portfolios to which F&F suggests allocating savings. $x_{k,t}$ corresponds to the k exchange rate fundamental variables, based on Contreras et al. (2013), which consist of returns on the trade-weighted U.S. dollar index (ΔUSD) and Chilean terms of trades (ΔToT), and the change in the interest rate differential between the short-run domestic and the U.S. interest rates ($\Delta(i - i^{*})$).

In order to identify the effect of F&F announcements on exchange rate returns, we control for m additional factors, not only those influencing F&F recommendations but also factors shown to influence exchange rate returns. Following section 2.5.1's discussion of F&F recommendation determinants, the vector $z_{m,t}$ contains the variables that F&F customarily cites as the main drivers of its recommendations, namely: the change in the VIX index (ΔVIX), the change in domestic economic uncertainty (ΔDEU), the change in domestic expected inflation ($\Delta\pi$), and returns on Chilean government bonds ($\Delta Bond$) and the S&P500 index ($\Delta SP500$). On the basis of findings in section 2.5.1 we interpret F&F recommendation announcements as an exogenous shock uncorrelated to the error term of the exchange rate model we introduce in this section. We include the set of variables in x_t and z_t both contemporaneously and with two lags. In addition, the model also includes two lags of the dependant variable to control for the persistence of exchange rate returns. In selecting the number of lags we follow the Bayesian information criterion (BIC) and panel A of table A.2.3 in appendix exhibits the lowest BIC when the specification includes two lags. α^h, β^h, γ_i^h, δ_i^h and θ_i^h are coefficients we estimate using ordinary least squares.

Table 2.4 displays the estimation of equation 2.4 for the period covering 01 March 2012 to 22 October 2020 using a daily observation frequency and setting $h = 1$. The results indicate that the day following a F&F recommendation announcement, the Chilean peso exhibits a significant depreciation of around 0.86% on average. The remaining coefficients, corresponding to exchange rate fundamentals, show the expected sign, a plausible magnitude and with the exception of the change in the interest rate differential are statistically significant. The rest of the control variables again display the expected sign, albeit only some are significant.

[Table 2.4 in here]

In order to analyse the persistence of F&F recommendations on nominal exchange rate returns we project the effect h days ahead in figure 2.6. The solid (blue) line depicts the cumulative response of the nominal exchange rate to F&F announcement news (i.e.: the β^h coefficient of the model in equation 2.4) and the grey area corresponds to 95% confidence interval bands. As documented previously, the exchange rate exhibits an average 0.86% cumulative depreciation the first day after F&F recommendations which increases to a 1.6% and 1.8% cumulative depreciation by the fifth and tenth day, respectively. The statistical effect fully dissipates around eighteen days. The persistence of the shock over time is consistent with the fact that PFCs are mandated to delay portfolio switches to meet regulatory requirements. This evidence shows that F&F announcement news, although noisy in its nature, generates significant pressure on the Chilean peso nominal exchange rate.[16]

[16] As a robustness check, we run the model of equation 2.4 again using the F&F shock definition based on Da et al. (2018). As figure A.2.3 in appendix depicts, the impact on the nominal exchange rate is 0.5%, 1% and 1.2%, after 1, 5 and 10 days. Although similar in terms of significance, the impact on the exchange rate is lower in magnitude compared to our F&F announcement news definition. We conjecture that this difference may arise from the fact that Da et al. (2018) omit the impact of portfolio switches to intermediate risk exposures (i.e.: portfolios B, C and D), therefore the impact on exchange rate is lower.

[Figure 2.6 in here]

In order to control for the effect of overlapping recommendations, we drop any announcements that occur within a twenty-day window of the previous recommendation. Figure 2.7 displays the results. As we can see, after excluding overlapping recommendations, the impact of F&F news on the exchange rate is similar in magnitude, still evidencing a 1% depreciation on the first day after the recommendation. However, in line with expectations, since any overlapping recommendations augment the prior shock and induce a prolonged impact on the exchange rate, the effect now dissipates earlier, becoming insignificant by the tenth day following the initial recommendation day.

[Figure 2.7 in here]

2.5.2.2 Comparison of the effect of F&F news to other FOREX market shocks

To provide further context for our results, we compare the effect on the nominal exchange rate of F&F announcements relative to the impact of historical FOREX market interventions by the Central Bank of Chile (CBCL). In particular, Contreras et al. (2013) find the 2011 FOREX market intervention depreciates the exchange rate by 4.6% and 12% one and five days after the announcement, respectively. The statistical effect of the announcement lasts between fifteen and eighteen days. As no study has quantified the exchange rate effect of the most recent CBCL interventions, we use the observed percentage exchange rate change corresponding to -3% (-5.5%) and 1.4% (1.2%) one (five) day after the 2019 and 2021 CBCL intervention announcement, respectively. The results in table 2.5 indicate that the effect of F&F announcement news on the exchange rate is generally lower in absolute magnitude and less persistent than the impact of all these CBCL FOREX market intervention announcements, albeit comparable in magnitude to those at the lower end of the spectrum.

[Table 2.5 in here]

2.5.2.3 Exchange rate model based on FOREX trading volume

This section estimates whether FOREX market trading volume induced by PFC portfolio reallocations influences the Chilean exchange rate. The results of this section provide an additional benchmark against which to compare our previous findings that suggest significant pressures arise on exchange rate returns subsequent to F&F announcement news. Our approach closely follows Evans and Lyons (2002) who include order flow (signed, net trading volume) as a fundamental driver explaining exchange rate returns, allowing us to directly quantify the effect of PFCs on the peso FOREX market without the need to consider the effect of F&F recommendation news.

Our model specification based on the PFC trading volume in the Chilean peso FOREX market is as follows:

$$Y = NetTrdVol\,\theta + X\beta + \varepsilon \tag{2.5}$$

where Y is a vector containing nominal exchange rate returns. $NetTrdVol$ is a vector containing the change in the PFC net trading volume in both the FOREX spot and forward markets. X is a vector containing the first difference of other exchange rate fundamental variables based on Contreras et al. (2013). ε is a vector representing the error term. β and $\theta = [\theta_{spot}, \theta_{forward}]$ are coefficients we estimate using ordinary least squares.

[Table 2.6 in here]

We estimate the regression model using daily observations over the period 01 March 2012 to 22 October 2020 and present the results in table 2.6. The findings reveal that the change in PFC net trading volume in both the spot and the forward markets are statistically significant and exhibit the expected sign. For each additional thousand million U.S. dollars PFC purchases (sales) in the spot (forward) market the Chilean peso depreciates (appreciates) 0.32% (0.20%).[17] Since 2011, the average two-day cumulative change in PFC net trading volume in the FOREX spot market following an F&F recommendation is around $2,000 millions U.S. dollars. On the basis of the results in table 2.6, we infer the exchange rate depreciates 0.65% (=0.323x2) due to the cumulative direct effect of PFC activity in the FOREX spot market two days after F&F issues a recommendation. Although a little lower in magnitude, the findings in this section are consistent with the F&F exchange rate pressures we estimate using the LPM of equation 2.4, supporting the idea that F&F recommendations trigger much of the PFC net trading volume in the FOREX market which ultimately impacts the nominal exchange rate. The significant estimated coefficients associated with PFC trading volume in the FOREX market (θ) reveal that $\theta_{forward} < 0 < \theta_{spot}$ and $|\theta_{forward}| < \theta_{spot}$. These results provide empirical support for the view that the regulatory mandated hedging strategy PFCs must undertake in the FOREX forward market acts to partially offset the exchange rate depreciation pressures which ensue when these companies purchase U.S. dollars in the FOREX spot market.

2.5.2.4 Asymmetric impact of F&F recommendations

Now we turn to analyse any potential asymmetries induced by F&F recommendations, by separately considering the effect of those announcements suggesting to investors reallocations to portfolios with riskier and less risky exposures. We estimate the LPM in equation 2.4 twice. Initially, we estimate the model using only recommendations which suggest investors enhance their risk exposure. In this case, the F&F variable takes the appropriate value given by the

[17] The change in PFC net trading volume in FOREX spot and forward markets is measured in thousand million U.S. dollars. A positive value in Δ PFC net trading volume in the Peso FOREX spot (forward) market represents net purchases (sales) of U.S.dollars.

$F\&F$ when the recommendation suggests switching from safer to riskier portfolios, and zero otherwise. Subsequently, we re-estimate equation 2.4 using only those announcements which suggest investors to switch from risky to safer portfolios. In this case, the F&F variable takes the value given by $F\&F \times (-1)$ when FF recommends switching from higher to lower risk exposure, and zero otherwise.

[Figure 2.8 in here]

Panels (a) and (b) in figure 2.8 present the cumulative impact on the exchange rate of F&F recommendations which advocate taking on more and less risk exposure, respectively. Risk enhancing (mitigating) recommendations both generate a cumulative depreciation (appreciation) of around 0.8% the day following the recommendations, the former (latter) dissipating six (twenty) days after the recommendation announcement. This asymmetric cumulative impact is illustrated in Figure 2.8 panel (c), in which we multiply the cumulative effect of taking less risk by minus one to facilitate comparisons. This evidence is consistent with the fact that regulations mandate that PFCs must hedge against currency risk by selling currency forwards in the FOREX derivative market after purchasing U.S. dollars in the FOREX spot market, a strategy they initiate only when enhancing their portfolio risk exposure. These results provide evidence that the compulsory forward sales of U.S. dollars partially compensate for the depreciation pressures of PFC purchases of U.S. dollar in the FOREX spot market. The results are fully consistent with the evidence we introduce in section 2.5.2.3.

2.5.3 Exchange rate volatility

In this section, we investigate if the documented pressure that F&F announcement news exerts on exchange rate returns translates into enhanced exchange rate volatility. Related studies argue that pension fund investors focusing on short-term horizon strategies, as characterised by F&F recommendations, tend to exacerbate asset price volatility (Levy & Zuniga, 2016; OECD, 2020). Following Diebold and Yilmaz (2009) we estimate a measure of nominal exchange rate realised volatility using intra-day observations.[18] In particular, we compute our range-based exchange rate volatility measure as follows:

$$\begin{aligned}\hat{\sigma}_t^2 &= 0.511\,(H_t - L_t)^2 - 0.019\,[(C_t - O_t)\,(H_t + L_t - 2O_t)\\ &-2\,(H_t - O_t)\,(L_t - O_t)] - 0.383\,(C_t - O_t)^2\end{aligned} \tag{2.6}$$

where H_t, L_t, O_t, C_t represent the intra-day high, low, open and close price, respectively.

[Figure 2.9 in here]

As shown in figure 2.9 preliminary evidence reveals that our measure displays volatility clusters after mid-2011, and realised volatility in the exchange rate tends to spike during

[18] This range-based implied volatility measure is highly similar in comparison to a simpler volatility measure estimated as the square exchange rate returns. The comparative time-series plot of both volatility measures is available upon request.

the days F&F makes recommendations (vertical dotted lines). In similar fashion to section 2.5.2, we estimate the effect of F&F recommendations on exchange rate volatility using a modification of the LPM in equation 2.4 in which our dependent variable now corresponds to the cumulative change in the natural logarithm of the square root of exchange rate volatility, but the explanatory variables and the F&F news definition remain the same.

[Figure 2.10 in here]

Figure 2.10 panel (a) illustrates the effect of F&F announcement news on the cumulative change in exchange rate volatility, revealing that volatility increases sharply by around 100% during the first day following F&F recommendations. This effect on exchange rate volatility is short-lived. Indeed, the statistical significance of the cumulative volatility impact quickly dissipates after the first day following the recommendation. The robust but short-lived effect on exchange rate volatility is consistent with the findings of section 2.5.2, which indicates the majority of the effects on the exchange rate occur during the first day after F&F recommendations and dissipate thereafter (figure 2.6). Comparing our results to related studies quantifying the impact of other comparable shocks to exchange rate volatility, Fuentes et al. (2014) document that the 2011 CBCL intervention in the Chilean FOREX spot market increases exchange rate volatility by 36%. The effect, however, becomes less significant after controlling for additional factors. The comparative lower impact of CBCL interventions in the FOREX market on exchange rate volatility may relate to the Central Bank's financial stability objective, suggesting FOREX market interventions are intended to reduce rather than to enhance exchange rate volatility.[19]

To validate the robustness of our findings, we analyse the effect of F&F recommendations on the conditional volatility of exchange rate returns using a GARCH model which includes F&F recommendation announcement news as an additional explanatory variable in the variance equation of the model.[20] The mean equation in the GARCH formulation corresponds to equation 2.4, excluding the F&F news from the set of explanatory variables. We model the conditional variance of cumulative exchange rate returns as follows:

$$\sigma^2_{t,h} = \omega^h + \alpha^h \varepsilon^2_{t-1,h} + \beta^h \sigma^2_{t-1,h} + \gamma^h FF_t \tag{2.7}$$

where $\sigma^2_{t,h}$ corresponds to the conditional variance of cumulative exchange rate returns. $\varepsilon_{t,h}$ corresponds to the residuals of the mean equation. FF_t corresponds to the F&F recommendation news defined as in section 2.5.2. ω, α, β and γ are the coefficients we estimate via maximum likelihood. In similar fashion to the previous analysis, we formulate h days ahead projections of the effect of F&F recommendation news on the conditional variance of

[19] Neely (2008) provides a detailed analysis of the literature discussing Central Bank interventions, revealing the lack of consensus in this area.

[20] Other studies also implement this GARCH methodology to estimate the effect of shocks to the FOREX market on exchange rate volatility. For instance, Doroodian and Caporale (2001) find that Central Bank intervention in the FOREX market generates a significant increase in volatility (measured as conditional variance of exchange rate returns) in the yen/dollar and mark/dollar sectors. Using the same methodology, Domac and Mendoza (2004) find that Central Bank interventions reduces exchange rate volatility in the case of both Mexico and Turkey.

cumulative exchange rate returns. Figure 2.10 panel (b) exhibits the results. The solid line represents the cumulative response of nominal exchange rate volatility we obtain from the GARCH model in equation 2.7 to F&F recommendation news (i.e.: the γ coefficient projected h days ahead). The results are consistent with the original exercise, with exchange rate volatility displaying a short-lived increase with most of the effect arising during the first days following the F&F recommendations.

Viewed collectively, the evidence we present in this section suggests that the effect of F&F recommendations is to enhance exchange rate volatility, although its impact is short-lived. Our results are consistent with related studies arguing that investors focusing on short-term horizons make decisions which tend to exacerbate asset price volatility (Levy & Zuniga, 2016; OECD, 2020).

2.5.4 Trading volume in FOREX market

Our findings to this point suggest F&F recommendations generate enhanced pressure on the Chilean peso nominal exchange rate and exacerbate its volatility. Here we analyse the impact of these recommendations on the trading patterns of different classes of investors who engage in trading activity on the Chilean peso spot market. We employ a proprietary dataset from the CBCL containing information of the daily FOREX trading volume of six important classes of market participants: Pension fund companies (PFCs), non-residents, retail companies, insurance companies, stock brokers, and mutual funds. Our hypothesis is informed by how regulations relating to the timing of portfolio readjustments in the pension fund industry induce the possibility that other market participant may benefit by anticipating massive, coordinated PFC trading volume in the FOREX market following F&F recommendations, in particular by front-running the anticipated trades they induce.

To calculate the impact of F&F recommendations on the trading volume of each class of market participant, we estimate the following regression model over the period 01 March 2012 to 22 October 2020:

$$TrdVol_{p,t} = FF_t\theta_p + X_{t-1}\beta_{p,1} + X_{t-2}\beta_{p,2} + \varepsilon_{p,t} \tag{2.8}$$

where $TrdVol_{p,t}$ is a vector containing the time-series of the natural logarithm of trading volume in the Chilean peso FOREX spot market of investor class p. The subscript $p = 1, \dots, 6$ indicates a particular class of the FOREX spot market participants described above. FF_t is a vector which contains h categorical variables taking a value of one h days (with $h = 1, \dots, 10$), following an F&F recommendation and zero otherwise. X_{t-1} and X_{t-2} represent vectors of one and two period lagged control variables, respectively. ε_p is a vector representing the residual term of investor class p. The control variables in X_{t-1} and X_{t-2} correspond to the ones we include in the model of equation 2.4. We use two lags of the control variables following the BIC criteria (see table A.2.3 panel B in the appendix). We use ordinary least squares to estimate θ_p, $\beta_{p,1}$ and $\beta_{p,2}$, the vectors of coefficients.

[Table 2.7 in here]

Table 2.7 displays the results from estimating equation 2.8, with the main findings as follows. First, we observe that PFC trading volume significantly increases the day following an F&F recommendation announcements, reaching its peak four days later and gradually becomes less elevated over the next few days. This pronounced hump-shaped pattern is consistent with the manner in which PFCs are mandated to implement portfolio switches on the basis of the regulatory requirements discussed in section 2.2, which states that PFCs can only reallocate pension assets on the fourth day following receipt of the switching request. Consequently, viewed in isolation, these regulations initiate an expected increase in PFC trading volume in the peso FOREX spot market on and around the fourth day following F&F recommendation announcements.

Interestingly, PFC trading volume starts increasing from time $t + 1$, the day following an F&F recommendation, showing that PFCs anticipate a large number of portfolio switching requests following F&F recommendations and immediately begin to accommodate their FOREX needs. Second, on average we observe a significant increase of 25% and a 35% in the FOREX spot market trading volume of non-resident and mutual fund companies, respectively, after F&F recommendations. Third, we observe no persistent change in the FOREX trading volume of either retail, insurance or stock broker companies following F&F recommendation announcements. Overall, this set of results raises the possibility that F&F recommendation news not only significantly increases PFC trading volume, but that it may also induce certain other classes of FOREX market participants, such as non-resident investors and mutual fund companies, to anticipate such massive and coordinated PFC transactions in the peso FOREX spot market. Our evidence closely relates to the findings in Corsetti et al. (2002) showing that transactions by large, sophisticated investors in the FOREX market exert other market participants to trade in this market more aggressively. Our findings are also consistent with the findings in Da et al. (2018) who document that F&F recommendations generate significant changes in investors' trading patterns in the Chilean stock market. While we do not test the proposition directly, it is conceivable that the regulatory trading restrictions which mandate a delay to PFC portfolio switches, provide both an incentive and an opportunity enabling other market participants to attempt to profit by front-running these anticipated PFC trades. Finally, considering that PFCs and non-residents together represent more than 50% of the total trading volume in the FOREX market, this significant increase in their trading volume after F&F recommendations translates to pressures on the nominal exchange rate. This evidence is consistent with the results in figure 2.6 where we document that the impact on the nominal exchange rate occurs from the day immediately following F&F recommendations, despite the fact that PFCs can process switching portfolio requests only four days after their receipt.

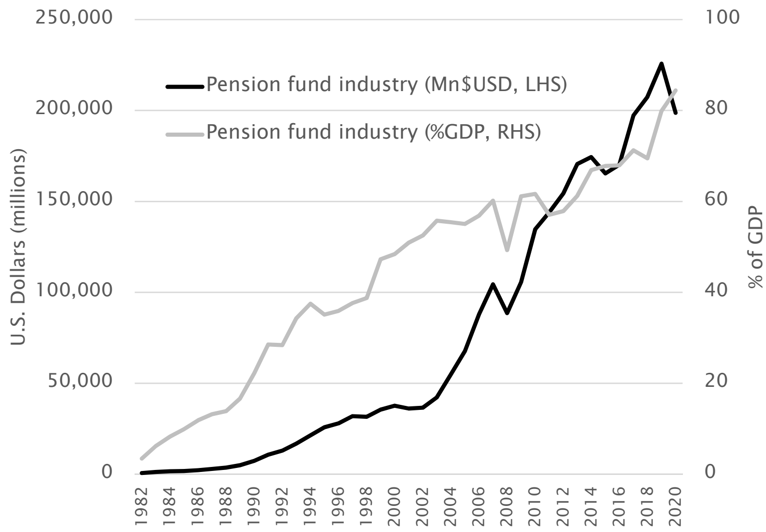
Following the results we present in this section, an additional hypothesis to analyse relates to the effect of exchange rate fluctuations, caused by F&F recommendations, on the valuation of domestic companies' assets. We argue that subsequent changes in the value of domestic firms' assets may increase the investors' demand for shares of certain firms experiencing increases in their market value, potentially leading to effects on the trading pattern of domestic equities. In particular, F&F recommendations generating exchange rate depreciations (appreciations) may induce investors to purchase shares of exporting (importing) companies as the asset value of these firms would benefit from a more favourable exchange rate. Although we do not provide a quantification of this valuation channel affecting domestic equities, this analysis helps to extend Da et al. (2018) findings by further examining the implications of F&F recommendations on the fluctuations of the Chilean peso and the hypothetical, subsequent effect on the trading pattern of domestic stocks.

2.6 Conclusion

The increasing prominence of Felices & Forrados (F&F), a financial advisory firm in the Chilean pension fund industry, has positioned the company as the most relevant firm in the growing market providing pension investment recommendations to investors. Focusing on short-term horizon investment strategies, F&F recommendations trigger large asset reallocations within the pension fund system. This asset reallocation translates into massive, coordinated transactions by pension fund companies in the Chilean peso sector of the FOREX market. In this study, we show that F&F recommendations, although noisy in their nature in the sense they cannot be predicted accurately by macroeconomic or financial market developments, generate a considerable pricing impact on the Chilean peso FOREX market. Among the main results, we show that the Chilean peso depreciates by 0.86% on average after F&F recommendations and the impact persists for ten days. We also find that F&F recommendations exert a substantial increase in exchange rate volatility, although the effect is short-lived and dissipates quickly over time. Collectively, our evidence suggests F&F recommendation announcements generate significant price pressures in the Chilean peso FOREX market. Our findings are consistent with related studies arguing that substantially large, coordinated asset reallocations based on short-term investment strategies tend to impact asset prices, pushing them beyond fundamentals and exacerbating price volatility. Further highlighting our findings which provide evidence that F&F recommendations exert considerable exchange rate pressures, we document that certain other classes of markets participants may anticipate the ensuing large, coordinated transactions of PFCs in the FOREX market, and attempt to profit by front-running these trades, although we leave detailed analysis of this issue to future research. Our results suggest F&F recommendations generate a meaningful impact on the Chilean peso FOREX market that may not be consistent with the CBCL's financial stability objectives. The findings contained in this study contribute to the ongoing policy discussions concerning the appropriate regulation of financial advisory companies operating in the pension fund industry in Chile. Moreover, our analysis also has implications for multiple other jurisdictions which harbour similar pension fund systems, particularly for countries whose regulation allows investors to actively choose investment portfolios based on recommendations of unregulated financial advisory firms.

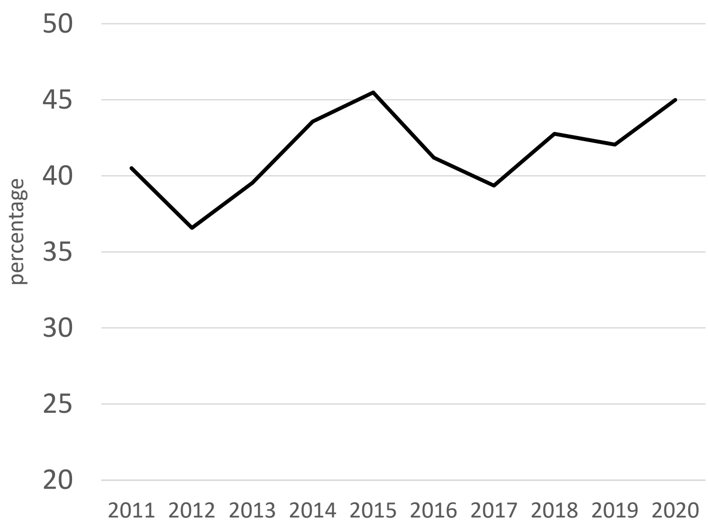
Figures

Figure 2.1. Pension fund industry value



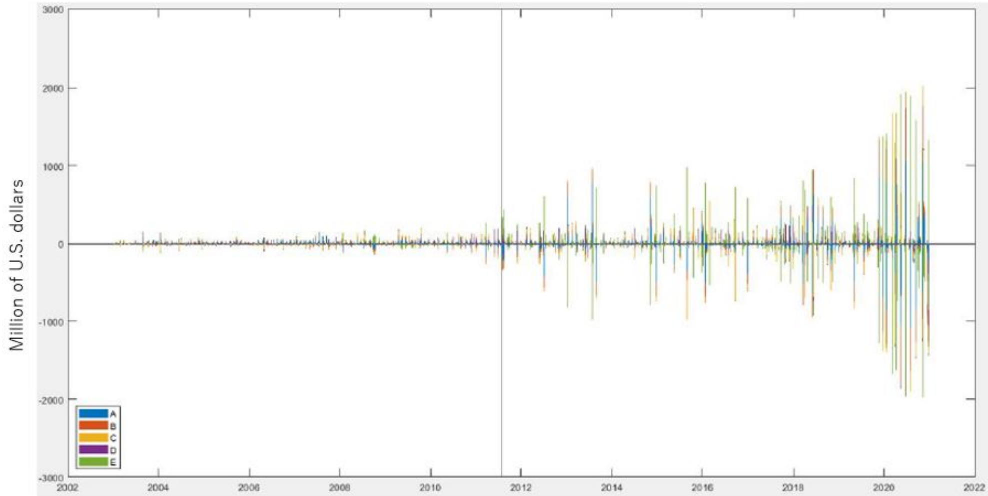
Value of the pension industry in Chile, annual observations, 1982 to 2020. Black line corresponds to the value in million U.S. dollars (left-hand side axis). Grey line corresponds to values as a percentage of the Chilean GDP (right-hand side axis). Source: Chilean regulatory body of the pension fund industry.

Figure 2.2. Percentage of assets invested in foreign markets, aggregate industry

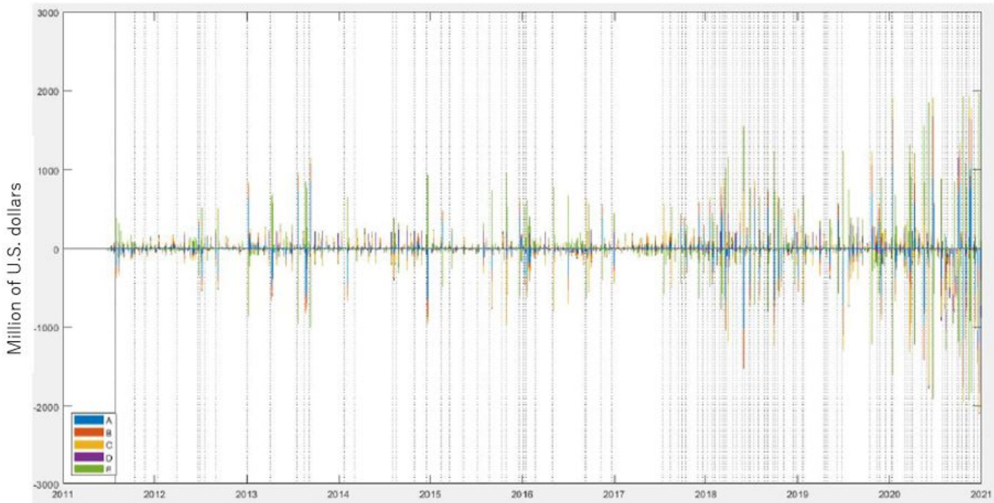


Percentage of assets invested in foreign markets at aggregate industry level (aggregate assets of the PFCs). Annual observations, 2011-2020. Source: Chilean regulatory body of the pension fund industry.

Figure 2.3. Net pension saving flows



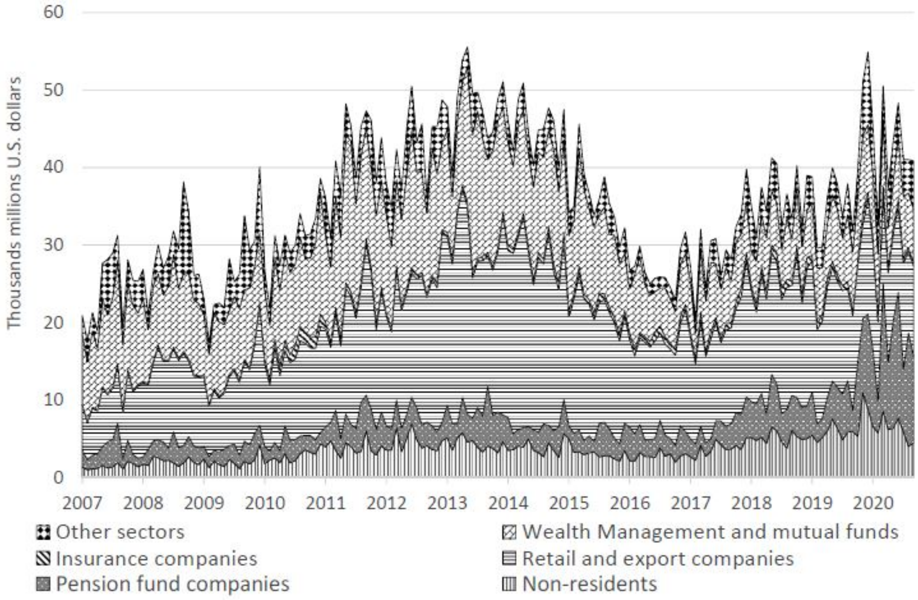
(a) 2003-2020



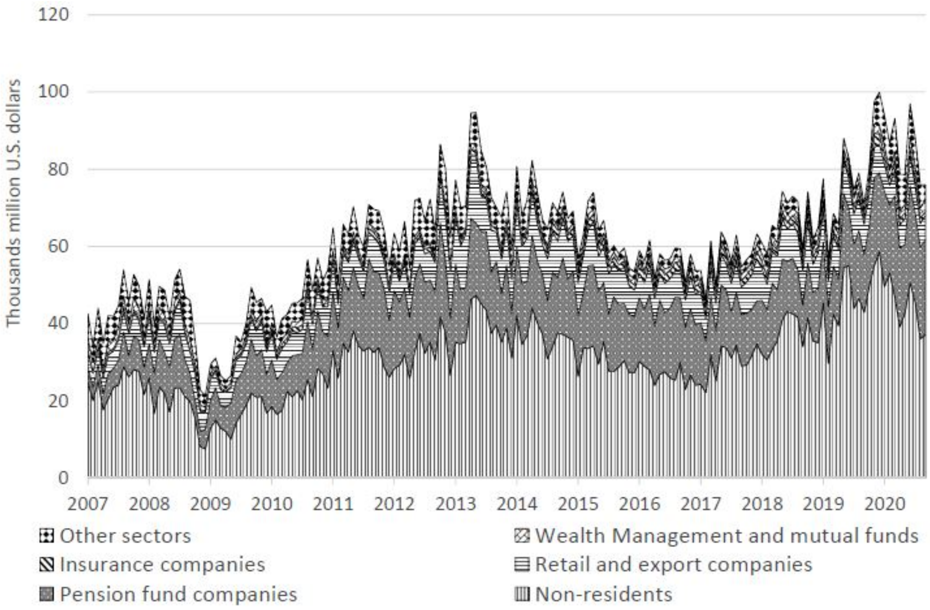
(b) 2011-2020

Net portfolio flows (millions of U.S. dollars) by portfolio type. Portfolio flows correspond to the aggregate daily observations at the industry level. Vertical line in panel (a) indicates the day F&F recommendations commence (July 2011). Vertical lines in panel (b) indicate the dates F&F issues recommendations. Source: Superintendencia de Pensiones de Chile.

Figure 2.4. Trading volume in the Chilean FOREX market by agent



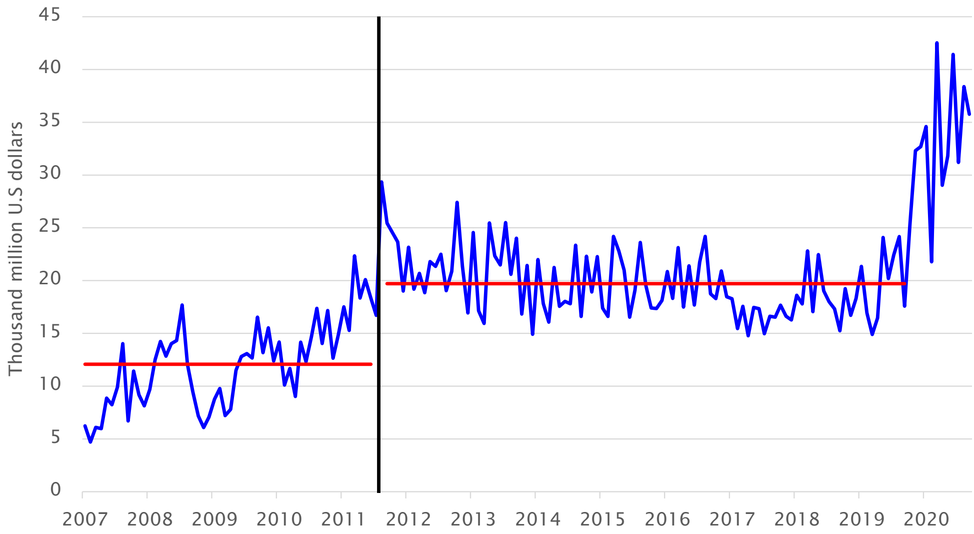
(a) FX spot market



(b) FX derivative market

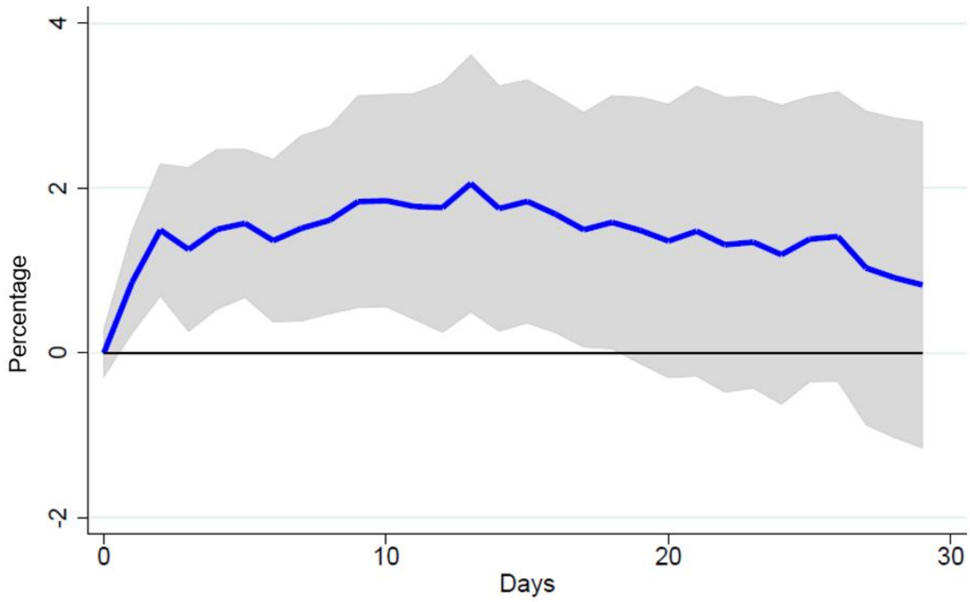
Trading volume in the Chilean FOREX markets by agent. Values in thousands millions U.S. dollars. Monthly observations, 2007 to 2020. Source: Central Bank of Chile.

Figure 2.5. PFC trading volume in the FOREX market



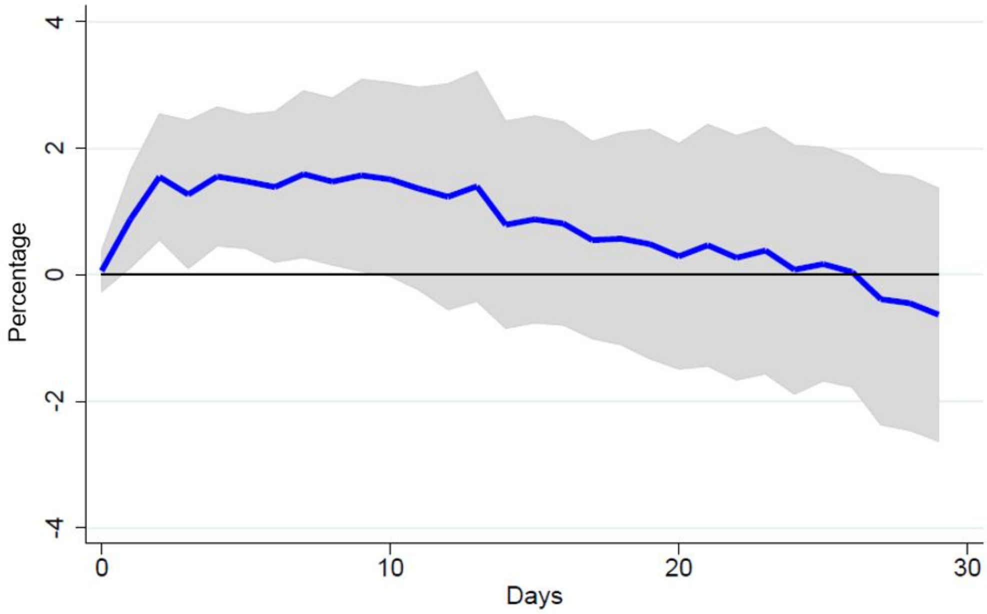
Trading volume of pension fund companies (PFCs) at aggregate industry level in the Chilean FOREX market (spot and derivative), thousands millions of U.S. dollars. Monthly observations, 2007-2020. The vertical black line represents the date F&F commences recommendations. Horizontal red lines denote the average PFC trading volume before and after F&F recommendations. The average PFC trading volume before F&F recommendations corresponds to $12 thousands millions dollars. The average PFC trading volume after F&F recommendations (excluding the period from 2020 onwards) corresponds to $20 thousands millions dollars. Source: Central Bank of Chile.

Figure 2.6. Cumulative response of nominal exchange rate to F&F news



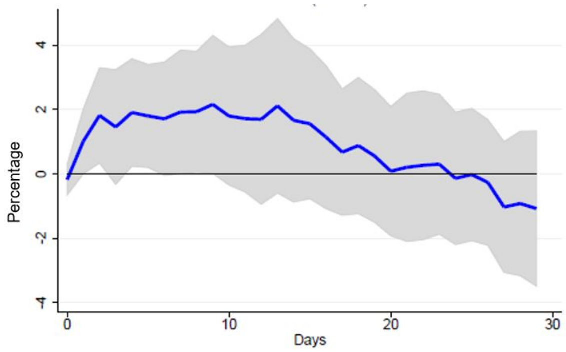
The blue line represents the cumulative response of the $\Delta usdclp$, in percentage change, to F&F news at horizon h (days). An increase indicates a Chilean peso depreciation. The cumulative response of the $\Delta usdclp$ corresponds to the estimated β^h coefficient of equation 2.4, with $h = 0, \ldots, 30$. The grey area represents 95% confidence bands. Daily observations, 01 March 2012 - 22 October 2020. The impact of F&F news is statistically significant when the confidence bands exclude zero. Source: Authors' calculations.

Figure 2.7. Cumulative response of nominal exchange rate to F&F news, excluding overlapped recommendations

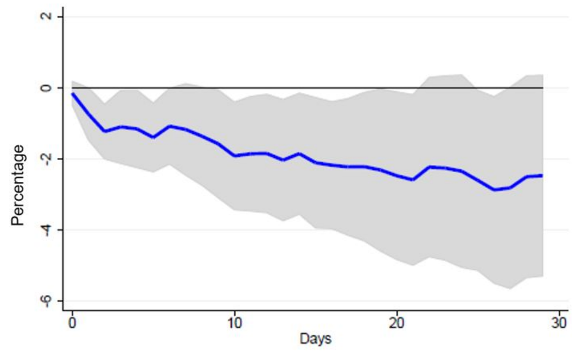


The blue line represents the cumulative response of the $\Delta usdclp$, in percentage change, to FF news at horizon h (days). An increase indicates a Chilean peso depreciation. The cumulative response of the $\Delta usdclp$ corresponds to the estimated β^h coefficient of equation 2.4, with $h = 0, \ldots, 30$. The grey area represents 95% confidence bands. This estimation excludes overlapped recommendations. See table A.2.2 in appendix for details of those overlapped dates, marked with an star (*). Daily observations, 01 March 2012 - 22 October 2020. The impact of F&F news is statistically significant when the confidence bands exclude zero. Source: Authors' calculations.

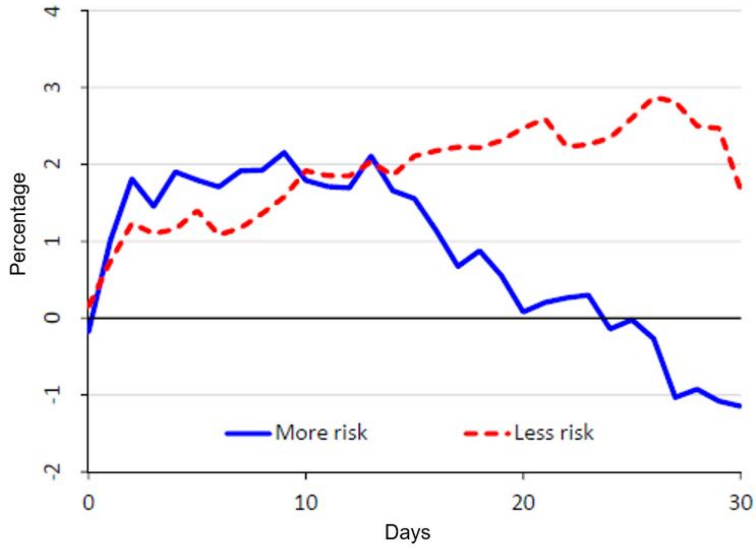
Figure 2.8. Asymmetric impact of F&F recommendations



(a) More risk



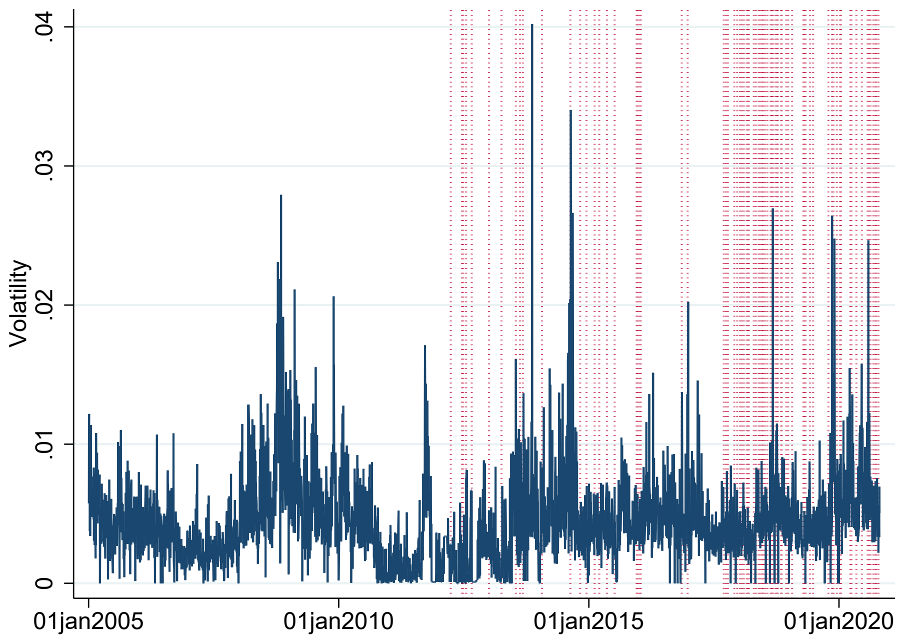
(b) Less risk



(c) Asymmetric impact point estimate

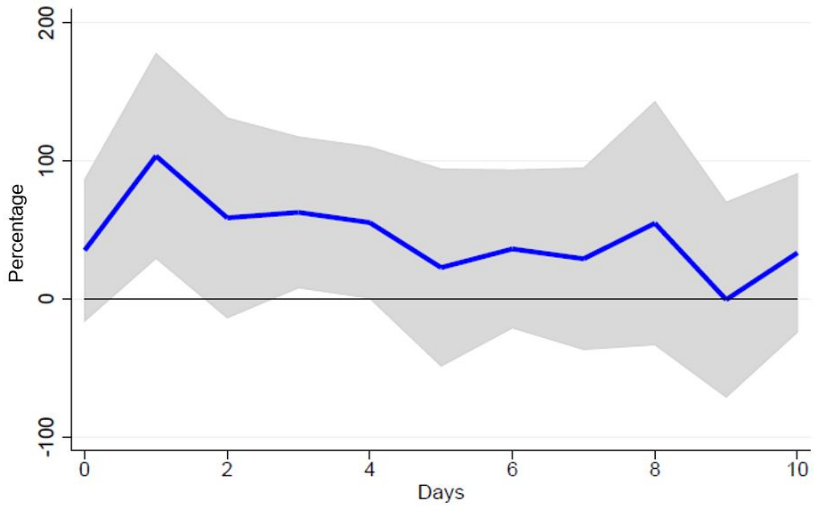
The blue line in panels (a) and (b) represents the cumulative response of the $\Delta usdclp$, in percentage change, to FF news at horizon h (days) suggesting to take more and less risk, respectively. An increase (decrease) indicates a domestic currency depreciation (appreciation). The cumulative response of the $\Delta usdclp$ corresponds to the estimated β^h coefficient of equation 2.4, with $h = 0, \ldots, 30$. The grey area represents 95% confidence bands. The blue line in panel (c) represents the point estimate of panel (a), while the red dashed line represents the point estimate of panel (b), which is multiplied by minus one for illustration purposes only. Daily observations, 01 March 2012 - 22 October 2020. The impact of F&F news is statistically significant when the confidence bands exclude zero. Source: Authors' calculations.

Figure 2.9. Nominal exchange rate volatility

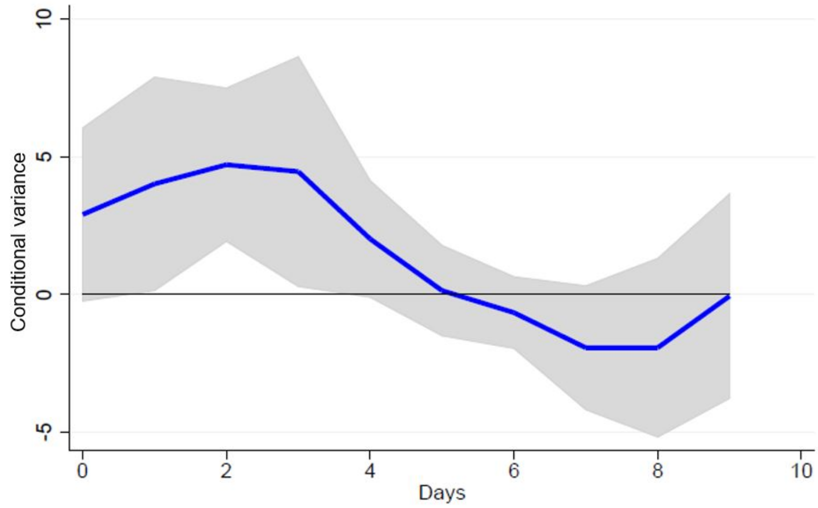


The blue solid line corresponds to the square root of the exchange rate volatility introduced in equation 2.6. Vertical dotted lines denote the days when F&F issues recommendations. Source: Authors' calculations.

Figure 2.10. Cumulative response of exchange rate volatility to F&F news



(a) Exchange rate return volatility



(b) Conditional variance exchange rate returns

The blue line in panel (a) represents the cumulative response of the exchange rate volatility, in percentage change, to F&F news at horizon h. The cumulative response of the exchange rate volatility corresponds to the estimated β^h coefficient of equation 2.4, with $h = 0, \ldots, 10$, where the dependent variable corresponds to the cumulative change in the natural logarithm of the square root of exchange rate volatility. The blue line in panel (b) represents the cumulative response of the conditional variance of nominal exchange rate returns to FF news at horizon h. The cumulative response of the conditional variance of nominal exchange rate returns corresponds to the estimated γ^h coefficient of equation 2.7, with $h = 0, \ldots, 10$. The grey areas represent 95% confidence bands. Daily observations, 01 March 2012 - 22 October 2020. The impact of F&F news is statistically significant when the confidence bands exclude zero. Source: Authors' calculations.

Tables

Table 2.1. Pension fund company portfolio composition

	Portfolio					
	A	**B**	**C**	**D**	**E**	**Total**
Panel A: Value (US$ Mn)						
Value	22,274	26,665	66,024	35,977	48,475	199,416
Panel B: Composition (% of the portfolio)						
Domestic	16	32	51	71	88	55
Equity	13	11	7	3	2	6
Fixed income	4	22	44	68	87	49
Foreign	84	68	49	29	12	45
Equity	66	48	31	15	3	29
Fixed income	18	20	18	14	9	16
Panel C: Equity investment limits (% of the portfolio)						
min.	40	25	15	5	0	–
max.	80	60	40	20	5	–

Panel A displays the value in U.S. million dollars of each portfolio at an aggregate pension fund industry level in 2020. Panel B exhibits the composition of each portfolio considering the location of the invested assets (domestic or foreign markets) and its type (equity or fixed income). Panel C shows the equity investment limits as a percentage of the total portfolio. Source: Superintendencia de Pensiones de Chile (Chilean regulatory body of the pension fund industry).

Table 2.2. Preliminary facts

Episodes	σ usdclp	PFC net trading vol FOREX	
		σ spot	σ forward
Pre GFC	0.55	29.3	59.3
GFC	0.90	91.9	164.7
Since F&F	0.63	174.1	174.7

σ usdclp represents the standard deviation of daily Chilean exchange rate returns. σ spot and σ forward represent the standard deviation of the daily change in the pension fund company trading volume in the Chilean FOREX Spot and Derivative markets, respectively. Dates pre global financial crisis (Pre GFC): January 2003 - July 2007. Global financial crisis (GFC): July 2007 - August 2009. Since F&F: July 2011 - September 2020. Source: Central Bank of Chile.

Table 2.3. Ordered logit model of F&F recommendation determinants

		More risk	Less risk
$\Delta usdclp$	$t-1$	0.180	-0.189
		(0.140)	(0.140)
	$t-2$	-0.194	0.170
		(0.140)	(0.160)
	$t-3$	-0.097	-0.176
		(0.210)	(0.150)
	$t-4$	-0.194	-0.068
		(0.140)	(0.150)
$\Delta SP500$	$t-1$	0.123	-0.255***
		(0.120)	(0.100)
	$t-2$	-0.043	0.178*
		(0.090)	(0.090)
	$t-3$	-0.105	0.013
		(0.120)	(0.090)
	$t-4$	-0.022	0.005
		(0.110)	(0.050)
$\Delta Bond$	$t-1$	-0.237	0.171
		(0.170)	(0.220)
	$t-2$	-0.153	-0.364*
		(0.230)	(0.200)
	$t-3$	0.008	0.264
		(0.270)	(0.250)
	$t-4$	-0.272	-0.001
		(0.240)	(0.340)
$\Delta\pi$	$t-1$	-7.032**	1.168
		(3.450)	(4.160)
	$t-2$	-5.638	-0.636
		(3.760)	(4.080)
	$t-3$	-6.130	11.946**
		(4.290)	(4.780)
	$t-4$	3.865	0.462
		(7.820)	(4.400)
	$t-5$	0.156	-4.120
		(4.290)	(5.420)
ΔDEU	$t-1$	-0.072	-0.024
		(0.060)	(0.040)
	$t-2$	0.058	-0.002
		(0.050)	(0.030)
	$t-3$	-0.110**	-0.020

		(0.050)	(0.040)
	$t-4$	0.022	-0.016
		(0.020)	(0.050)
	$t-5$	-0.015	0.066
		(0.040)	(0.050)
ΔVIX	$t-1$	-0.314**	0.086
		(0.150)	(0.070)
	$t-2$	-0.352***	-0.037
		(0.110)	(0.090)
	$t-3$	0.053	0.019
		(0.080)	(0.060)
	$t-4$	-0.039	-0.081
		(0.090)	(0.080)
	$t-5$	-0.114	0.037
		(0.100)	(0.080)
Latent variable thresholds			
κ_1		4.670***	4.184***
		(0.320)	(0.240)
κ_2		5.089***	4.549***
		(0.380)	(0.280)
Parallel assumption test			
χ^2		22.590	19.120
p-value		[0.707]	[0.866]
N Obs		1511	1511
Pseudo R2		0.15	0.09

Estimation of the ordered logit model of equation 2.3 using daily observations from 01 March 2012 to 22 October 2020. The dependent variable of the model in column 'more risk' ('less risk') corresponds to an ordered categorical variable capturing the intensity of F&F recommendation to take more (less) risk. The ordered dependent variable takes values of 1 and 2 when F&F recommends a moderate and strong change in risk exposure, respectively, and zero otherwise. A strong change in risk exposure occurs when a recommendation suggests changing to one extreme portfolio conditional on the existing recommendation allocating investments within the opposite extreme portfolio. We define moderate changes in risk exposure as those recommendations which suggest an increased allocation to intermediate portfolios (i.e.: portfolios B, C or D), when existing recommendations involve an extreme portfolio allocation. The last column of table A.2.2 in the appendix displays this classification. The set of explanatory variables consists on four lags of the cumulative weekly returns of the Chilean nominal exchange rate ($\Delta usdclp$), S&P 500 ($\Delta S\&P500$) and Chilean government bonds ($\Delta Bond$). In addition, we also include five lags of daily changes of domestic inflation expectations ($\Delta\pi$), domestic economic uncertainty (ΔDEU) and VIX index (ΔVIX). We estimate the models using maximum likelihood. Standard errors in parentheses. P-values of the parallel regression assumption test in square brackets. (*), (**), (***) indicates statistical significance at 10, 5, and 1% level, respectively. Source: Authors' calculations.

Table 2.4. Effect of F&F news on Chilean exchange rate

		$\Delta usdclp_{t+h}$
F&F	t	0.857***
		(0.314)
ΔUSD	t	1.085***
		(0.074)
	$t-1$	0.125
		(0.093)
	$t-2$	-0.038
		(0.088)
ΔToT	t	-0.045***
		(0.007)
	$t-1$	-0.018**
		(0.007)
	$t-2$	-0.008
		(0.007)
$\Delta(i-i^*)$	t	-0.813
		(0.759)
	$t-1$	0.050
		(0.557)
	$t-2$	-0.847
		(0.631)
ΔVIX	t	0.037
		(0.023)
	$t-1$	0.058**
		(0.023)
	$t-2$	0.034
		(0.021)
ΔDEU	t	0.001
		(0.005)
	$t-1$	0.001
		(0.004)
	$t-2$	-0.006
		(0.004)
$\Delta\pi$	t	0.396
		(0.620)
	$t-1$	-0.873
		(0.557)
	$t-2$	0.403
		(0.608)
$\Delta Bond$	t	0.106

		(0.117)
	$t-1$	0.060
		(0.120)
	$t-2$	0.162
		(0.114)
$\Delta SP500$	t	-0.057
		(0.034)
	$t-1$	0.123***
		(0.042)
	$t-2$	0.057
		(0.036)
$\Delta usdclp$	$t-1$	0.036
		(0.049)
	$t-2$	-0.028
		(0.047)
Constant		0.016
		(0.020)
N Obs		1725
Adj. R2		0.24

Estimation of the local projection model of equation 2.4 via OLS, using daily observations, 01 March 2012 - 22 October 2020. The dependent variable corresponds to a time-series of cumulative exchange rate returns at $h = 1$, measured in percentage change. An increase indicates a Chilean peso depreciation. The set of explanatory variables contains F&F news, contemporaneous and lagged observations of the of returns on the trade-weighted U.S. dollar index (ΔUSD) and Chilean terms of trades (ΔToT), the change in the interest rate differential between the short-run domestic and the U.S. interest rates ($\Delta(i - i^*)$), the change in the VIX index (ΔVIX), the change in domestic economic uncertainty (ΔDEU), the change in domestic expected inflation ($\Delta\pi$), and returns on Chilean government bonds ($\Delta Bond$) and the S&P500 index ($\Delta SP500$). Robust standard error in parentheses. (*), (**), (***) indicates statistical significance at 10, 5, and 1% level, respectively. Source: Authors' calculations.

Table 2.5. Comparison of shocks on the Chilean FOREX market

	$t+1$	$t+5$	Length$^{(1)}$
F&F recommendation news	0.86%	1.57%	10-18
Central Bank of Chile FX market interventions			
2011$^{(2)}$ (US dollars purchase)	4.6%	12%	15-20
2019$^{(3)}$ (US dollars sales)	-3.0%	-5.5%	–
2021$^{(3)}$ (US dollars purchase)	1.4%	1.2%	–
Net FX_{PFC} trading volume model	0.65%$^{(4)}$	–	–

Impact of shocks on the Chilean peso exchange rate at one and five days after the events defined in rows. (1) Length corresponds to the number of days the shock displays statistical significant impact on exchange rate returns. (2) Based on Contreras et al. (2013). (3) Nominal exchange rate variation after the Central Bank intervention announcement. (4) Cumulative effect on exchange rate two days after F&F recommendations. Source: Authors' calculations.

Table 2.6. Exchange rate model based on PFCs FOREX trading volume

	$\Delta usdclp_t$
ΔPFC net trading volume	
$Spot_t$	0.323***
	(0.098)
$Forward_t$	-0.200**
	(0.082)
$\Delta(i_t - i_t^*)$	0.384
	(0.562)
ΔToT_t	-0.039***
	(0.004)
ΔUSD_t	1.045***
	(0.042)
Constant	0.009
	(0.011)
N obs	1940
Adj R2	0.35

Results of the estimation of the model in equation 2.5 via OLS using a sample of daily observations, 01 March 2012 - 22 October 2020. The dependent variable corresponds to a time-series of exchange rate returns. An increase indicates a Chilean peso depreciation. The set of explanatory variables contains the change in the FOREX net trading volume of PFCs (ΔPFC net trading volume) in both the spot and forward markets, the change in the interest rate differential between the short-run domestic and the U.S. interest rates ($\Delta(i - i^*)$), the of returns on both Chilean terms of trades (ΔToT) and trade-weighted U.S. dollar index (ΔUSD). Δ PFC net trading volume in Chilean FOREX spot and forward markets are measured in thousand million U.S. dollars. A positive value in Δ PFC net trading volume in Chilean FOREX spot (forward) market represents net purchases (sales) of U.S. dollars. Robust standard error in parentheses. In order to avoid endogeneity issues, we use a measure of PFC net trading volume in the peso FOREX spot and forward markets which is orthogonal to other risk factors that may affect the Chilean exchange rate. Section "Auxiliary PFC FOREX trading volume regressions" on page 68 provides more details about the orthogonalisation of PFC trading volume in the FOREX market. Source: Authors' calculations.

Table 2.7. F&F and FOREX trading volume by agent

	PFC	Non-residents	Retail	Insurance	Brokers	M. Funds
FF						
$t+1$	0.760***	0.233***	0.070	0.144	-0.017	0.332***
	(0.165)	(0.078)	(0.046)	(0.151)	(0.079)	(0.097)
$t+2$	0.939***	0.240**	0.031	0.021	-0.124*	0.198
	(0.155)	(0.104)	(0.051)	(0.165)	(0.075)	(0.125)
$t+3$	1.264***	0.259***	-0.081	0.220	-0.117	0.363***
	(0.153)	(0.081)	(0.065)	(0.165)	(0.073)	(0.120)
$t+4$	1.457***	0.218***	0.000	0.173	-0.086	0.458***
	(0.166)	(0.077)	(0.054)	(0.119)	(0.088)	(0.105)
$t+5$	1.422***	0.140*	0.023	0.128	0.017	0.372***
	(0.139)	(0.079)	(0.066)	(0.154)	(0.070)	(0.119)
$t+6$	0.961***	0.160	-0.031	-0.089	-0.166*	0.345***
	(0.318)	(0.149)	(0.065)	(0.204)	(0.091)	(0.134)
$t+7$	1.154***	0.161*	0.045	0.154	-0.113	0.244*
	(0.106)	(0.085)	(0.067)	(0.142)	(0.078)	(0.128)
$t+8$	0.892***	0.201*	-0.037	-0.120	-0.205*	0.272**
	(0.179)	(0.116)	(0.066)	(0.211)	(0.109)	(0.115)
$t+9$	1.019***	0.251***	0.064	0.055	-0.048	0.362***
	(0.140)	(0.093)	(0.098)	(0.154)	(0.099)	(0.123)
$t+10$	0.561***	0.066	0.012	-0.005	-0.016	0.060
	(0.215)	(0.113)	(0.071)	(0.148)	(0.090)	(0.129)
ΔDEU						
$t-1$	0.009	0.008**	0.002	0.003	0.002	0.002
	(0.006)	(0.003)	(0.002)	(0.005)	(0.003)	(0.004)
$t-2$	0.015**	0.014***	0.002	0.002	-0.002	0.001
	(0.007)	(0.003)	(0.002)	(0.005)	(0.003)	(0.004)
$\Delta\pi$						
$t-1$	-0.192	-0.458	-0.220	0.275	-0.280	0.587
	(0.940)	(0.385)	(0.253)	(0.621)	(0.354)	(0.526)
$t-2$	-0.061	-0.262	-0.201	-0.424	-0.203	-0.340
	(1.021)	(0.372)	(0.258)	(0.612)	(0.336)	(0.532)
$\Delta(i-i^{*})$						
$t-1$	-0.568	0.389	-0.064	0.164	0.225	-0.517
	(0.664)	(0.439)	(0.239)	(0.588)	(0.282)	(0.439)
$t-2$	0.584	0.203	-0.129	0.820	-0.058	0.531
	(0.781)	(0.361)	(0.225)	(0.516)	(0.257)	(0.480)
ΔVIX						
$t-1$	0.009	0.008	-0.003	-0.015	-0.001	0.008
	(0.016)	(0.009)	(0.006)	(0.013)	(0.007)	(0.010)

$t-2$	-0.011	-0.006	-0.007	0.004	0.000	0.004
	(0.016)	(0.008)	(0.006)	(0.011)	(0.006)	(0.011)
$\Delta SP500$						
$t-1$	-2.548*	-0.859	-0.399	-3.104***	-0.894	-2.490***
	(1.523)	(0.758)	(0.503)	(1.129)	(0.680)	(0.914)
$t-2$	0.107	0.569	0.144	-2.177**	-0.777	-0.605
	(1.386)	(0.651)	(0.446)	(1.002)	(0.586)	(0.824)
$\Delta Bond$						
$t-1$	-0.794	2.579	2.632**	1.926	-0.687	1.234
	(4.158)	(2.450)	(1.210)	(3.317)	(1.354)	(2.145)
$t-2$	3.839	-1.654	1.585	0.895	0.183	-3.438
	(4.850)	(2.258)	(1.402)	(3.522)	(1.648)	(2.236)
$\Delta usdclp$						
$t-1$	0.699	-1.909*	-0.574	2.508	0.545	-0.965
	(1.953)	(1.145)	(0.653)	(1.724)	(0.771)	(1.248)
$t-2$	0.976	1.449	0.894	2.404	2.108***	2.240*
	(2.645)	(1.046)	(0.669)	(1.671)	(0.816)	(1.240)
Constant	4.600***	5.135***	6.549***	3.150***	5.970***	3.307***
	(0.039)	(0.016)	(0.011)	(0.026)	(0.014)	(0.020)
N Obs	1762	1770	1771	1771	1771	1771
Adj R2	0.10	0.04	0.00	0.01	0.01	0.03

Results of the estimation of model in equation 2.8 via OLS using daily observations, 01 March 2012 - 22 October 2020. The dependent variable in each column represents the natural logarithm of the trading volume in the peso FOREX spot market by agent. FF_{t+h}, $h = 1, \dots, 10$, is a categorical variable taking the value of one h days after F&F issues a recommendation, and zero otherwise. Additional control variables correspond to lagged observations of the change in domestic economic uncertainty (ΔDEU), the change in domestic expected inflation ($\Delta\pi$), the change in the interest rate differential between the short-run domestic and the U.S. interest rates ($\Delta(i-i^*)$), the change in the VIX index (ΔVIX), returns on the S&P500 index ($\Delta SP500$), returns on Chilean government bonds ($\Delta Bond$), and returns on the Chilean exchange rate. Robust standard errors in parentheses. (*), (**), (***) indicates statistical significance at 10, 5, and 1% level, respectively. Source: Authors' calculations.

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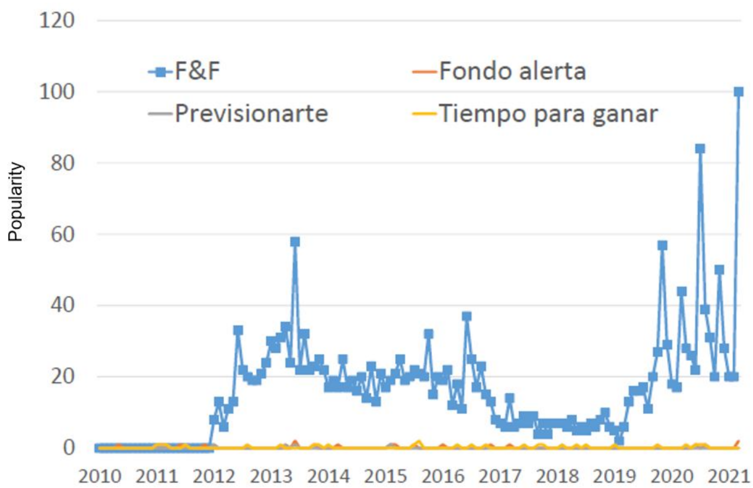
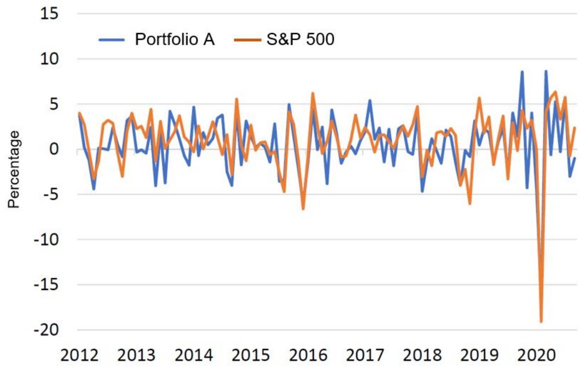
Appendix

Figure A.2.1. Popularity of financial advisory firms in the pension fund market in Chile

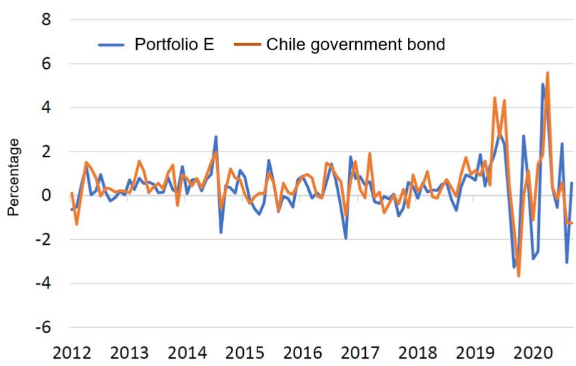


Popularity based on Google trend index. Values in the y-axis capture the relative search interest. A value of 100, 50, and 0, represents the most popular search, half of the most popular, and no popular at all. Source: Google trends.

Figure A.2.2. PFC portfolio returns



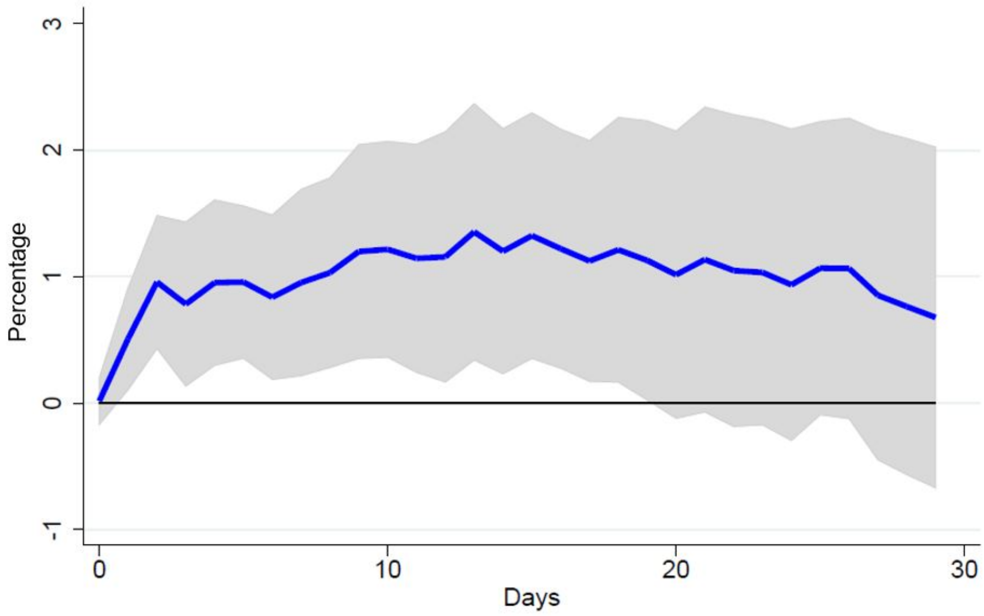
(a) Portfolio A and S&P 500 returns



(b) Portfolio E and Chilean gov. bond returns

Time-series of monthly returns from March 2012 to October 2020. Source: Authors' calculations based on Bloomberg and Superintendencia de Pensiones de Chile.

Figure A.2.3. Impact of F&F recommendations on exchange rate based on Da et al. (2018)



The blue line represents the cumulative response of the $\Delta usdclp$, in percentage change, to FF news at horizon h (days). In this case, F&F news follows the Da et al. (2018) definition. An increase indicates a Chilean peso depreciation. The cumulative response of the $\Delta usdclp$ corresponds to the estimated β^h coefficient of equation 2.4, with $h = 0, \ldots, 30$. The grey area represents 95% confidence bands. Daily observations, 01 March 2012 - 22 October 2020. The impact of F&F news is statistically significant when the confidence bands exclude zero. Source: Authors' calculations.

Table A.2.1. F&F followers statistics

	F&F followers	Non-F&F followers	Difference
Age	38	41	-3***
Savings	50,989	14,896	36,093***
Male	58	52	6**

Table displays average statistics by groups: F&F followers and Non-F&F followers. The last column reports the difference of averages between groups and its statistical significance. Age measured in years. Saving measured in U.S. dollars. Male corresponds to the percentage of males in each group. (*), (**), (***) indicates statistical significance at 10, 5, and 1% level, respectively. Source: Authors' calculations based on *Superintendencia de Pensiones de Chile* (regulator authority of the pension fund market in Chile).

Table A.2.2. F&F recommendations

#	Date	FF recom.	F&F	Ologit
1	27/07/2011	100% E	–	–
2	12/10/2011	100% A	0.627	2
3	22/11/2011	100% E	-0.644	-2
4	11/01/2012	100% A	0.652	2
5	29/03/2012	100% E	-0.645	-2
6	19/06/2012	100% A	0.641	2
7	28/06/2012*	100% E	-0.648	-2
8	19/07/2012	100% A	0.642	2
9	29/08/2012	100% E	-0.640	-2
10	02/01/2013	100% A	0.677	2
11	03/04/2013	100% E	-0.667	-2
12	17/07/2013	100% A	0.688	2
13	16/08/2013	100% E	-0.673	-2
14	06/09/2013	100% A	0.665	2
15	24/01/2014	100% E	-0.654	-2
16	06/03/2014	50% C / 50% E	0.177	1
17	01/08/2014	100% E	-0.134	-1
18	19/08/2014*	50% A / 50% E	0.343	2
19	30/10/2014	100% A	0.353	2
20	15/12/2014	100% E	-0.716	-2
21	12/02/2015	50% A / 50% E	0.363	2
22	18/03/2015	100% A	0.350	2
23	13/05/2015*	50% A / 50% E	-0.352	-2
24	08/07/2015	40% C / 60% E	-0.211	-1
25	24/08/2015	100% E	-0.150	-1
26	13/10/2015	50% C / 50% E	0.184	1
27	26/10/2015	100% E	-0.181	-2
28	16/12/2015	50% A / 50% E	0.349	2
29	22/12/2015*	100% A	0.354	2
30	06/01/2016*	50% A / 50% E	-0.340	-2
31	15/01/2016*	100% E	-0.363	-2
32	22/02/2016	50% C / 50% E	0.179	1
33	29/04/2016	100% E	-0.155	-2
34	06/09/2016*	50% C / 50% E	0.158	1
35	13/09/2016	100% E	-0.155	-2
36	09/11/2016	50% A / 50% E	0.335	2
37	22/12/2016	100% E	-0.346	-2
38	13/07/2017	50% C / 50% E	0.173	1
39	10/08/2017	100% E	-0.168	-2
40	12/09/2017*	50% A / 50% E	0.325	2
41	28/09/2017	100% A	0.326	2
42	12/10/2017*	50% A / 50% E	-0.324	-2
43	28/11/2017	100% A	0.352	2
44	19/12/2017	50% A / 50% E	-0.351	-2
45	09/01/2018	100% A	0.343	2
46	22/01/2018*	50% A / 50% E	-0.339	-2
47	05/02/2018	100% E	-0.348	-2
48	26/02/2018*	50% A / 50% E	0.339	2
49	07/03/2018	100% A	0.350	2
50	14/03/2018*	50% C / 50% E	-0.515	-2

51	23/03/2018*	15% D / 85% E	-0.145	-1
52	19/04/2018*	50% A / 50% E	0.311	2
53	04/05/2018	100% A	0.358	2
54	24/05/2018*	50% C / 50% E	-0.521	-2
55	06/06/2018	60% A / 40% E	0.242	2
56	19/06/2018*	20% A / 80% E	-0.285	-1
57	25/06/2018	100% E	-0.142	-1
58	09/07/2018*	50% A / 50% E	0.355	2
59	27/07/2018	100% E	-0.356	-2
60	20/08/2018*	50% A / 50% E	0.357	2
61	29/08/2018	100% A	0.355	2
62	05/09/2018*	50% A / 50% E	-0.359	-2
63	24/09/2018	100% E	-0.353	-2
64	05/10/2018*	50% A / 50% E	0.354	2
65	11/10/2018*	100% E	-0.360	-2
66	05/11/2018	50% A / 50% E	0.349	2
67	09/11/2018*	100% E	-0.360	-2
68	12/12/2018	50% A / 50% E	0.355	2
69	26/12/2018*	40% C / 60% E	-0.207	-1
70	18/01/2019	100% E	-0.149	-1
71	24/01/2019*	50% A / 50% E	0.351	2
72	16/04/2019	100% E	-0.343	-2
73	23/04/2019*	50% A / 50% E	0.361	2
74	02/05/2019*	100% E	-0.369	-2
75	04/06/2019	50% A / 50% E	0.366	2
76	26/06/2019	100% E	-0.367	-2
77	16/10/2019	50% A / 50% E	0.373	2
78	11/11/2019	100% A	0.402	2
79	22/11/2019*	50% A / 50% E	-0.384	-2
80	16/12/2019	100% E	-0.377	-2
81	09/01/2020	50% A / 50% E	0.399	2
82	16/01/2020*	100% E	-0.374	-2
83	03/03/2020	50% C / 50% E	0.203	1
84	12/03/2020*	100% E	-0.190	-2
85	24/03/2020	40% A / 60% E	0.297	2
86	01/04/2020*	80% A / 20% C	0.356	1
87	07/05/2020	50% C / 50% E	-0.474	-2
88	27/05/2020	100% E	-0.193	-1
89	16/06/2020	50% A / 50% E	0.389	2
90	28/07/2020	25% A / 75% E	-0.187	-1
91	06/08/2020*	50% A / 50% E	0.194	1
92	18/08/2020	25% A / 75% E	-0.181	-1
93	08/09/2020	100% E	-0.172	-1
94	23/09/2020*	30% A / 70% E	0.209	1
95	29/09/2020	60% A / 40% E	0.209	1
96	15/10/2020*	20% A / 80% E	-0.294	-1
97	26/10/2020	10% A / 90% E	-0.074	-1

Column 'Date' corresponds to the date F&F issues a recommendation. A star (*) indicates a overlapping recommendation as there is less than twenty days after the previous recommendation. 'FF recom.' column corresponds to the portfolio allocation F&F suggests in its recommendation. 'F&F' column corresponds to the definition of the recommendation news we introduce in section 2.5.2 and it is computed as the first difference of $finv_t$, where $finv_t = \sum_{i=1}^{5} w_{it} p_{it}$, with $i = 1, 2, 3, 4, 5$ (the five PFC portfolios), w_{it} represents the portfolio allocation recommended by F&F in portfolio i at time t, and p_{it} represents the percentage of foreign investment in portfolio i at time t. 'Ologit' column corresponds to an ordered categorical variable taking the value of 1 and 2 when F&F recommends a moderate and strong change in risk exposure, respectively, and zero otherwise. A strong change in risk exposure occurs when a recommendation suggests changing to one extreme portfolio conditional on the existing recommendation allocating investments within the opposite extreme portfolio. We define moderate changes in risk exposure as those recommendations which suggest an increased allocation to intermediate portfolios (i.e.: portfolios B, C or D), when existing recommendations involve an extreme portfolio allocation. Source: Authors' calculations based on Superintendencia de Pensiones de Chile.

Table A.2.3. Information criteria

Panel A

	1 Lag	2 Lags
LPM	4522	4273

Panel B

	1 Lag	2 Lags
PFC	6471	6140
Non-residents	3326	3108
Retail	1981	1941
Insurance	5235	4922
Brokers	2865	2703
M. Funds	4237	4007

Table contains the Bayesian Information Criterion (BIC) for two different specifications using two alternative lag order for the independent variables. Panel A and B displays the BIC for the model in equation 2.4 and 2.8, respectively. Source: Authors' calculations.

Auxiliary PFC FOREX trading volume regressions

Exchange rate models based on trading volume in FOREX market, as the one we propose in equation 2.5, may potentially suffer endogeneity issues due to simultaneity bias. In order to mitigate endogeneity concerns we use a measure of PFC net trading volume in the FOREX market which is orthogonal to risk factors that may also relate to Chilean exchange rate movements. In particular, the measure of PFC net trading volume orthogonal to risk factors corresponds to the error term ε_i of the following equation:

$$Trd\,Vol_i^{PFC} = X\beta_i + \varepsilon_i \tag{A.2.1}$$

Where $Trd\ Vol_i^{PFC}$ corresponds to the PFC net trading volume in FOREX market i, with $i =$ [Spot, Forward]. X is a vector containing three categories of explanatory variables: global and domestic risks, economic surprises, and terms of trades. ε_i, the residual term, corresponds to the variable we use as a measure of orthogonal PFC net trading volume in the model of equation 2.5.

Table A.2.4 displays the results of the auxiliary models of PFC net trading volume in the FOREX spot market of equation A.2.1. As explanatory variable we include a set of variables tracking domestic economic uncertainty (ΔDEU), domestic inflation ($\Delta\pi$) and external risk factors (ΔVIX). We also include indices tracking domestic and world economic activity surprises along the Chilean terms of trades (ΔToT). The model includes all variables in first difference, excluding ΔToT which corresponds to percentage change. The results show that only external risk factors, captured by the VIX, statistically influence PFC net trading volume in the Peso FOREX market, while the rest of the variables controlling for domestic risk elements, surprises and terms of trades display no statistical significance. Moreover, the adjusted R2 of the models depicts low for all models suggesting the omission of this endogeneity analysis may not generate severe issues in the exchange rate model based on PFC net trading volume in the Peso FOREX market. The orthogonal measure of PFC net trading volume in the FOREX market we include in the model of equation 2.5 corresponds to the residual of model (1) in table A.2.4. Same procedure and conclusions apply for the case of orthogonal PFC net trading volume in the peso FOREX derivative market (results available upon request).

Table A.2.4. Auxiliary PFC FOREX trading volume regressions

		Δ *PFC net FOREX trading volume*			
		(1)	(2)	(3)	(4)
Risk	ΔVIX_t	0.005**	0.005**	0.005**	0.005**
		(0.002)	(0.002)	(0.002)	(0.002)
	ΔDEU_t	0.0008	0.0008	0.0008	0.0008
		(0.001)	(0.001)	(0.001)	(0.001)
	$\Delta \pi_t$	-0.044	-0.044	-0.044	-0.046
		(0.097)	(0.097)	(0.097)	(0.097)
Surprises	Δ Global Economy		-0.00001		
			(0.000)		
	Δ G10 Economy			-0.00001	
				(0.000)	
	Δ Emerging market Ec.			0.00001	
				(0.000)	
ToT	Δ *Terms of Trades*				-0.0015
					(0.001)
	Constant	-0.0005	-0.0005	-0.0004	-0.0004
		(0.004)	(0.004)	(0.004)	(0.004)
	N obs	1994	1994	1994	1994
	Adj. R2	0.002	0.001	0.001	0.002

Results of the estimation of model in equation A.2.1 via OLS using daily observations, 01 March 2012 - 22 October 2020. The dependent variable corresponds to the change of PFC net trading volume in the Chilean FOREX spot market. The set of explanatory variables corresponds to external risk factors, proxied by the VIX index (ΔVIX), domestic economic uncertainty (ΔDEU), and domestic inflation ($\Delta \pi$). We also include indices tracking domestic and world economic activity surprises along the Chilean terms of trades (ΔToT). The model includes all variables in first difference, excluding ΔToT which corresponds to percentage change. Robust standard errors in parentheses. (*), (**), (***) indicates statistical significance at 10, 5, and 1% level, respectively. Source: Authors' calculations.

Chapter 3

Asymptotic dependence and exchange rate forecasting

Abstract

Recent literature documents a non-linear relationship in commodity and exchange rate returns. This paper analyses the relevance of this finding for the documented ability of commodity returns to predict exchange rate returns. In a sample of commodity-exporting economies, we find that the source of the forecasting ability attributed to commodity returns lies in a revealed asymptotic dependence relationship between the two variables returns at a daily frequency. Timing plays a key role, since both the forecasting ability of commodity returns and the asymptotic dependence between the returns of these variables tend to be short-lived. Exchange rates adjust quickly to commodity return shocks, implying only contemporaneous observations capture a robust relationship. Both the asymptotic dependence property and the predictive ability of commodity returns disappears when we include lagged commodity returns in the analysis. Further, the relationship between commodity and exchange rate returns only appears significant at a daily frequency, disappearing when we use monthly or quarterly observations. We interpret these findings as manifesting the effect of commodity market news that conveys information relevant for exchange rate valuation in commodity-exporting economies.

3.1 Introduction

Recent empirical studies (Ferraro et al., 2015; Foroni et al., 2015) provide evidence that changes in commodity prices contain a degree of predictive power for exchange rate fluctuations. Moreover, both the log-return distributions of exchange rates and commodity prices exhibit heavy or fat tails, indicating the presence of extreme values in their sample distributions (for exchange rate returns see Patton (2006), Dias and Embrechts (2010) and Yang and Hamori (2014). For commodity returns see Algieri and Leccadito (2020), Hussain and Li (2020) and Kittiakarasakun (2014)). Motivated by these two sets of observations, this paper explores whether an asymptotic dependence between the two variables underpins the documented predictive ability of commodity returns for exchange rate returns. We seek to synthesis these two findings and, to the best of our knowledge, this paper is the first to study the empirical link between exchange rate and commodity returns utilising a tail dependence framework. The novelty of our contribution lies in measuring the extent of asymptotic dependence and how any such relationship contributes to explaining the ability of commodity returns to forecast exchange rates. We investigate whether the documented predictive ability of commodity returns applies across a sample of nine commodity exporting economies, and the forecast frequencies, if any, at which it exists. This analysis of the tail behaviour of exchange rate and commodity returns may be relevant in terms of policy considerations for the subset of commodity exporting economies which document a close relationship between exchange rate and commodity returns. For instance, understanding the transmission mechanism from commodity price shocks to exchange rates may help with policy management for developing countries moving from fixed-exchange to a more flexible rate regimes, or for those commodity exporting countries who have adopted inflation targeting monetary policies.

A number of recent studies analyse the tail behaviour of financial variables.[1] Research addressing the nature of the information contained in the tails of the distribution focuses on exploring the relationship between variables when linear correlation fails to detect the full extent of the association between them. For example, Cumperayot and De Vries (2017) demonstrate that measuring the asymptotic dependence between exchange rates and classic monetary variables enables one to explain how large swings in exchange rates maybe related to sharp movements in monetary fundamentals.

This paper contributes to the literature on asymptotic dependence and its forecasting implications by examining the relationship between exchange rate and commodity returns using a tail dependence framework. To our knowledge, it is the first paper which attempts to empirically synthesise these two literatures. We highlight the additional information contained in the tails of the exchange rate returns distribution and the potential role of asymptotic dependence as the central component of the documented ability of commodity returns to predict exchange rate returns. Specifically, our approach employs multivariate extreme value techniques which enable us to detect the occurrences of large movements in exchange and commodity returns and test if they are asymptotically dependent. Alternatively stated, we investigate if large shocks to commodity returns are transmitted to exchange rate returns and if such dependence is the source of any documented forecasting ability of the former for the latter. Measuring tail dependence enables us to determine the extent to which large movements in exchange rate returns relate to their underlying fundamentals, in this case, commodity returns. Our empirical approach is particularly relevant to variables which are known to exhibit fat-tailed distributions.

Our main results demonstrate that the asymptotic dependence between commodity and exchange rate returns is significant when both returns are measured contemporaneously at a daily frequency, while it reduces considerably when we use lagged returns or sample returns at a lower frequency. Moreover, we also show that the predictive ability of commodity returns for exchange changes is highly significant when we use contemporaneous daily observations. Using lagged or lower frequency observations reduces this ability of commodity returns to forecast exchange returns. We maintain that the nature of the documented asymptotic dependence is a key element in evaluating the relationship between the two variables. In particular, our results suggest it may be a necessary element to incorporate when analysing the role of commodity returns in predicting exchange rate changes.

[1] See Patton (2006) for a survey of studies implementing a copula approach. Patton (2012) surveys the methodology and approaches to modelling the tail behaviour of financial variables.

3.2 Related literature

3.2.1 Exchange rates and commodity prices

We now contextualise the contribution of this study to the existing literature. Rossi (2013) comprehensively surveys the large research literature on nominal exchange-rate forecasting, including models which use commodity prices. She concludes that the random walk (RW) model without drift remains a particularly difficult benchmark to outperform (Meese & Rogoff, 1983), reporting that while linear models appear to be the most successful, results vary depending upon the set of predictors, the sample period, the forecast evaluation method, and the forecast horizon.

In this vein, several studies focus upon the analysis of the statistical relationship between commodity prices and both real and nominal exchange rates at a variety of frequencies. Conducting in-sample exercises using quarterly frequency observations, Chen and Rogoff (2003) claim strong correlation and cointegration between commodity prices and real exchange rates of several developed country commodity exporters. Importantly for the purpose of our study, they argue that commodity price shocks are considered exogenous shocks useful to explain the exchange rate of commodity-exporting economies. Further, Cashin et al. (2004) provide evidence of the in-sample predictive power of commodity export prices to explain real exchange rates. They find evidence of correlation and cointegration in around one-third of their sample of 58 economies using observations at the monthly frequency. Amano and van Norden (1995, 1998a, 1998b) provide evidence in favour of the ability of commodity prices to explain exchange rates at a monthly frequency using cointegration analysis applied to a subset of advanced economies. Amano and van Norden (1995) present empirical evidence linking the Canada-US real exchange rate with the terms of trade, reporting that the real exchange rate is cointegrated with terms-of-trade variables (price of commodity exports relative to the price of manufactured imports), and that causality runs from the terms of trade to the exchange rate. Moreover, a simple exchange rate equation performs better than a random walk in post-sample forecasting exercises. Amano and van Norden (1998a) documents a robust relationship between the real domestic price of oil and real effective exchange rates for Germany, Japan and the United States. They attribute this effect to the real oil price capturing exogenous terms-of-trade shocks and explain why these shocks may determine long-term real exchange rates. Chen (2002) finds that including commodity prices improves the out-of-sample forecasting ability of fundamental-based models for nominal exchange rates at the monthly frequency for Australia, Canada and New Zealand, although the evidence is not completely robust for the entire sample period under analysis. In contrast, Chen et al. (2010) using both in-sample Granger-causality tests with time-varying parameters and out-of-sample forecasting with rolling windows, document that nominal exchange rates (for commodity currencies) help forecast commodity prices. They also document some evidence showing the impact of commodity prices to exchange rates, although the results are less robust. Their study analyses nominal exchange rates of Canada, Australia, New Zealand, South Africa, and Chile (each relative to the US dollar) along with export-earnings-weighted commodity prices for each country, at a quarterly frequency. They rationalise their findings in the context of a present value model following Engel and West (2005).

While much of the earlier literature is based on low frequency sampling, some recent studies investigate the forecasting ability of commodity prices using daily and higher frequency data. Zhang et al. (2016), using daily data from a sample of four economies (Australia, Canada, Chile, Norway), document the in-sample ability of three commodity prices (crude oil, gold, copper) to explain exchange rates. They employ both conditional and unconditional causality measures and consider non-USD exchange rates, noting that the documented relationship is stronger at short horizons and runs in the direction of commodity prices to exchange rates. Similarly, Foroni et al. (2015), using a mixed frequency estimation based on both daily and monthly observations for a sample of advanced economies, show that incorporating commodity prices improves the forecasting ability of fundamentals-based exchange rate models.

Other recent papers extend the analysis to a broader sample of countries at a daily frequency. Ferraro et al. (2015) document the ability of oil prices to predict exchange rates in a one-step ahead, out-of-sample forecasting exercise for five commodity exporter economies (Australia, Canada, Chile, Norway, and South Africa). However, Akram (2004) shows that the value of the Norwegian krone value against a European basket of currencies is uncorrelated with the oil price. Kohlscheen et al. (2017) find a strong correlation between changes in the nominal exchange rate and a daily index of export commodity prices for 11 countries using a panel dataset.[2] Both studies provide evidence of the forecasting ability of contemporaneous commodity prices to beat a RW model in out-of-sample exercises using observations at a daily frequency. They also show that this forecasting ability tends to be weaker using monthly data, and completely disappears at a quarterly frequency. In addition, both studies find that only contemporaneous commodity prices outperform the RW model and confirm that there is little evidence of out-of-sample predictability using lagged commodity prices. In particular, Ferraro et al. (2015) argue that any forecasting ability comes from the short-lived relationship between commodity prices and exchange rates, therefore, including contemporaneous observations at a daily frequency is a crucial component in capturing the relationship between these variables. Our paper is distinct from these studies in that importantly, none of the aforementioned papers investigate the source of their documented ability of commodity prices to contemporaneously forecast exchange rates, which is the motivation and key contribution of the present paper.

3.2.2 Higher moments in exchange rate distributions

An alternative strand of literature to which we also empirically contribute, analyses the role of higher moments in exchange rate distributions by modelling tail relationships using extreme value theory, often adopting a copula methodology. Patton (2006) studies the tail behaviour

[2] Australia, Canada, Norway, Brazil, Chile, Colombia, Mexico, Peru, South Africa, Russia and Malaysia.

of the Japanese yen and the German mark. His results indicate exchange rate movements located in the tails of the distribution are asymmetric, and that the degree of association between these currencies tends to be higher during periods of currency depreciation in comparison to episodes of appreciation. Similarly, Yang and Hamori (2014) also find asymmetric effects during periods of appreciation and depreciation when analysing the tail behaviour of the euro, Japanese yen and British pound in relation to the gold price. Using a time-varying copula approach to analyse the tail relationship between the Japanese yen and the euro, Dias and Embrechts (2010) argue that time-varying estimation provides additional information about the tail dependence between the variables. While this literature exploits information in the tails for examining the relationship between exchange rates and other financial assets, Cumperayot and De Vries (2017) explore the ability of monetary fundamentals to explain exchange rates. Their results indicate that information located in the tails of the distributions of both exchange rates and monetary fundamentals contributes to explaining the relationship between them, via transmission of the large shocks that impact fundamentals to exchange rates. Our paper demonstrates that asymptotic dependence between commodity prices (our fundamentals) and exchange rates maybe a crucial element to explain the documented ability of the former to predict movements in the latter at the daily time horizon. Such a dependency disappears at lower sample frequencies, such as monthly or quarterly, and so does the predictive ability of commodity prices.

3.3 Commodity exporting economies and commodity prices

Shocks to commodity prices play a key role in the transmission of global demand and supply disturbances to domestic economies. As discussed by Agénor and da Silva (2019), the effect of commodity prices is particularly relevant for the economic outlook of commodity exporting countries. In the short run, a negative commodity price shock reduces a country's export revenues which translates to a lower foreign currency inflow, typically U.S. dollars (USD), flowing into the economy. Consequently, the domestic currency's exchange rate tends to depreciate. Over a longer time horizon, a negative commodity price shock may be interpreted as contributing to a worsening in the economic outlook of the commodity exporting country. As International Monetary Fund (2017) among others points out, the deterioration in the future economic growth expectations in such economies discourages both domestic and inward investment from overseas, and can result in capital outflows and a further tendency for their exchange rate to depreciate, portending the enhanced possibility of a future financial crisis.

The relationship between exchange rates and commodity prices is particularly noteworthy during the commodity price super-cycle which occurs in the first decade of the 2000s. De Gregorio (2012) shows that, during prolonged episodes of high commodity prices, commodity exporting economies exhibit large and persistent current account deficits arising from massive capital inflows targeted at investments in the commodity sector. In this case, the large capital inflows lead to exchange rate appreciations. This logic suggests that shocks generated in commodity markets produce changes in the expectations of the domestic economic performance of commodity exporting economies. Subsequently, a change in the future outlook of the commodity sector drives further capital flow movements which ultimately impacts exchange rates. This documented relationship between commodity prices and capital flows is consistent with the finding of related papers, such as Reinhart and Reinhart (2009) and Byrne and Fiess (2016).

Given our focus on the relationship between commodity prices and exchange rates, the sample of countries we choose to include in this paper must satisfy two conditions: first, commodities exports must represent a significant proportion of the total country's exports, and second, countries must have a free floating exchange rate regime. It follows from these criteria that the exports of these countries in our sample are somewhat poorly diversified in the sense they are mainly dependant upon one or two main commodities. In addition, all countries in the sample adopt a free floating exchange rate regime during the sample period.

[Table 3.1 in here]

Table 3.1 reports the relevant descriptive statistics for the nine economies we eventually include in our final sample. For each country the table shows a set of commodity related indicators during three different time periods: 2000-2005, 2006-2011, and 2012-2017. The indicators highlight the relevance of commodity exports for the economies in the sample. As we observe, commodity exports represent a high percentage of the total exports for all of the countries. Moreover, the commodity sector is a highly relevant one for the whole economy. On average, around 15% of the respective countries GDP is represented by the commodity sector. Additionally, the table also shows the main product exported by each country, which constitutes a high proportion of the commodity exports. Overall, the information exhibited in table 3.1 reveals the relevance of commodity exports for the countries under analysis. It is worth noting that, although, oil seems to be less important for Brazil, that country is one of the top ten oil producers around the globe and is the biggest regional oil producer.[3] The oil industry in Brazil is also an important driver of domestic investment and attracts foreign capital into the country.

3.3.1 Relevant empirical facts about returns

It is a well established empirical fact that the returns of many asset classes exhibit a tendency to follow a leptokurtic distribution, one broadly characterised as displaying higher kurtosis and a greater likelihood of observing extreme values in comparison to the normal distribution. Likewise, distributions describing exchange rate log-returns and commodity log-returns exhibit heavy or fat tails, indicating a higher probability of observing extreme values. Cumperayot and De Vries (2017) and Patton (2006) explore these properties in the case of exchange rate log-returns. Other authors (Buyuksahin & Robe, 2014; UNCTAD, 2011), analysing the

[3]According to the U.S. Energy Information Administration (EIA). See table A.3.1 in appendix for more details about the World's Top Oil Producers.

enhanced degree of financialisation of commodity markets, point out that commodity log-returns also exhibit fat-tailed distributions, claiming that investors interpret commodities as another asset class seeking to actively include them as such in their investment portfolios.

[Figure 3.1 in here]

By way of illustration, figure 3.1 plots the histograms of daily observations of the US dollar - Chilean peso (USDCLP) exchange rate in log-returns (left-hand side panel) and the copper price in log-returns (right-hand side panel) from Jan-2000 to July-2018. We include a theoretical normal distribution (solid line) in each plot for comparison purposes. The figure shows the empirical distributions of both Chilean peso and copper returns depart from a normal distribution, as both reveal positive excess kurtosis and skewness different to zero. Indeed, the Shapiro-Wilk test's null hypothesis of normality is rejected for both returns. This preliminary evidence suggests that the returns of this commodity and exchange rate pair both follow fat-tailed distributions, where we can expect to observe a large number of extreme observations distant from the mean. It is worth noting, however, that return distributions departing from normality is a necessary but not sufficient condition to observe return distributions displaying extreme values. Consequently, we undertake a formal analysis, based on the Hill and Dekkers tail indices we introduce in section 3.4.3, to determine the heavy tail property of the distributions. The subsequent analysis of the results in section 3.5.4., based on the extreme value analysis, reveals the conclusion in this section is also valid for the remaining sample of exchange rate and commodity returns we analyse in this chapter.

In this context, in standard structural asset pricing models of the exchange rate, shocks that drive fundamentals also drive the exchange rate. Cumperayot and De Vries (2017) show that except for the influence of exogenous noise, the properties of the distribution of the fundamental variables, in our case commodity prices, determine the nature of the distribution of exchange rate returns. Of particular importance for our argument is their demonstration that in a linear specification of the exchange rate relationship, if the distribution of commodity prices (fundamentals) has fat tails, then we can expect both fat tails in the exchange rate distribution, and also the exchange rate returns and the commodity prices to exhibit asymptotic dependence. That is, as shocks become more extreme, in the limit the probability of large currency movements conditioned on large fundamental shocks is positive. They also show that there can still be asymptotic dependence even if the correlation coefficient is equal to 0. However, as they note, one cannot simply conclude that if the distributions have fat tails, the random variables are necessarily asymptotically dependent. They provide an example of Student-t distributed random variables combined with a Gaussian copula, which are correlated but asymptotically independent. They demonstrate that it is the linearity of the model when combined with the marginal heavy tail property of the distributions which induces the asymptotic dependency between exchange rates and commodity prices and their contemporaneous shocks.

If one finds no support for asymptotic dependence between the exchange rate returns and economic fundamentals, this may be attributable to one of the following two explanations.

One possibility is that the fundamentals-based exchange rate model does not apply, suggesting the noise is exogenous and is unrelated to commodity prices. Alternatively, even if two random variables are (imperfectly) correlated, but follow a certain class(es) of distribution, for example a multivariate normal distribution, then all dependency vanishes asymptotically. If tests reveal that we cannot reject asymptotic dependence, this at least suggests a strong inter-linkage between the commodity price and exchange rate returns transmitted through the economic relevance of the shocks that drive both variables.

3.4 Data and methodology

3.4.1 Data

The data corresponds to daily observations of nominal exchange rates and commodity prices, both measured in log-returns. We collect all series from Bloomberg and the sample is fully available from 03 January 2000 to 18 July 2018. Bloomberg's primary source of information for copper and gold prices corresponds to the price quoted at the London Metal Exchange (LME), while oil price corresponds to the West Texas Intermediate (WTI) Cushing, Oklahoma crude oil. The LME is widely known as the most liquid and influential global market for metal-type commodities and its quoted prices serve as a reference for global commodity markets, in particular for gold and copper (Ntim et al., 2015; Park & Lim, 2018). Following Ferraro et al. (2015), we choose the WTI oil price due to its relevance as a benchmark price for global oil markets. The countries under analysis and the country-specific commodity prices are shown in table 3.1. Initially, the nominal exchange rate is defined using the U.S. dollar as the base currency, but we also report results using the euro (EUR) and pound sterling (GBP) as an alternative specification in robustness check exercises. The prices for the set of commodities we include in the analysis are all denominated in U.S. dollars.

3.4.2 Forecasting model

In order to test for any inherent forecasting ability of commodity returns, we carry out a routine forecasting evaluation exercise adopted in the literature which we now briefly describe. Our candidate baseline model is a linear regression for each country in which there is only one explanatory variable: a country-specific commodity return. The choice of this model specification follows the comprehensive review of exchange rate forecasting methods in Rossi (2013) documenting that the linear, single-equation model is the most successful type of model to forecast exchange rates. In addition, related studies (Ferraro et al., 2015; Kohlscheen et al., 2017) document the enhanced forecasting ability of single-variable commodity-based models to predict exchange rates in comparison to models including other exchange rate fundamental variables. Table 3.1 reports the country-specific commodity for each country. Equation 3.1

gives the benchmark forecasting model.

$$\Delta s^f_{t+h} = \hat{\alpha}_t + \hat{\beta}_t \Delta p_{t+h}, \quad t = R, R+1, ..., T-h. \tag{3.1}$$

Where Δs^f_{t+h} represents the h-periods-ahead out-of-sample forecast of exchange rate log-returns.[4] Δp_{t+h} denotes the country-specific commodity log-returns. We estimate coefficients $\hat{\alpha}_t$ and $\hat{\beta}_t$ of equation 3.1 using the OLS method and a sample window of length R ($R = 2135$ observations), from 03/01/2000 to 09/04/2009, corresponding to half the total sample T ($T = 4270$ observations) and we produce h-steps ahead forecasts. We then roll forward the window one observation, re-estimate equation 3.1 over the window 04/01/2000 to 10/04/2009 and generate new h-step ahead forecasts. We repeat this process P times ($P \equiv T - R + 1$) up to $T - h$ to produce a series of forecasts for the full out-of-sample period. Following the methodology proposed by Meese and Rogoff (1983), we use perfect foresight data, meaning that we include contemporaneous realised values of commodity returns in the forecasting exercises. For this reason, the approach is also known as an *out-of-sample fit*, since in real life situations it is not possible to know tomorrow's value of commodity returns. We set short-term forecast horizons at $h = 1, 2, 3, 4, 5, 10$ periods ahead. Following the exchange rate forecasting literature, related studies using daily observation mostly focus on short-term horizons (e.g.: Ferraro et al. (2015) and Kohlscheen et al. (2017) use 1-step ahead forecasts, employing daily commodity prices), while others studies using lower frequency monthly, quarterly or annual observations, set longer forecast horizons (see Rossi, 2013).

We select the RW model with and without drift as the benchmarks against which to compare our commodity-based forecasts. According to Rossi (2013), the RW model without drift is the toughest benchmark to beat in out-of-sample forecast exercises. Under the RW logic, the best forecast of tomorrow's exchange rate value is today's exchange rate value. In other words, the exchange rate forecast of the RW model without drift is given by $\Delta s^{RW}_{t+h} = 0$. Where Δs^{RW}_{t+h} represents the h-period-ahead forecast of exchange rate log-returns based on the RW model. To assess the out-of-sample forecasting ability of the models, we statistically compare their root mean squared forecast error (RMSFE) using the Diebold-Mariano (DM) test statistic. Giacomini and White (2006) demonstrate the validity of the DM test when comparing the out-of-sample forecasts of two nested models when the length of the estimation windows is constant. We compute the RMSFE of the models as follows:

$$RMSFE = \sqrt{\frac{1}{P} \sum_{t=R}^{T} \left(\varepsilon^f_{t+h} \right)^2}$$

where ε^f_{t+h} denotes the h-period-ahead out-of-sample forecast error of the model. For the case of the commodity-based model in equation 3.1, the forecast error is obtained as $\varepsilon^f_{t+h} = \Delta s_{t+h} - \Delta s^f_{t+h}$, where Δs_{t+h} corresponds to the observed exchange rate return. While for the

[4] In this study the nominal exchange rate is defined as the value of one U.S. dollar (USD) in terms of the domestic currency. To eliminate concerns regarding the sensitivity of our findings to the choice of the USD as the numeraire (base) currency, we conduct robustness checks using the euro and pound sterling as base currencies in section 3.5.1

case of the RW model, the forecast error is obtained as $\varepsilon_{t+h}^{RW} = \Delta s_{t+h} - \Delta s_{t+h}^{RW}$. Subsequently, to statistically test the out-of-sample forecasting ability of the models, we evaluate the RMSFE ratio between the commodity-based model (numerator) and a RW model (denominator), i.e.: $RMSFE\,ratio \equiv RMSFE_{commodity-based}/RMSFE_{RW}$. If the RMSFE ratio displays a value statistically lower than one, then, we interpret the evidence as the superior forecasting ability of the commodity-based model. The forecasting analysis in section 3.5.1 displays the results of the RMSFE ratios.

It is worth noting that the out-of-sample forecasting analysis we undertake in this chapter follows the customary tools the forecasting literature utilises. The standard framework in this case focuses on analysing the predictive ability of models by comparing their out-of-sample forecast errors (Rossi, 2013). Conversely, the in-sample analysis covering the discussion of the result estimation of the α_t and β_t coefficients of equation 3.1 becomes less relevant since it departs of the purpose of this research, therefore the in-sample results are not reported.

3.4.3 Fat-tailed distribution in returns: Tail indexes

Before investigating the extent of any potential asymptotic dependence between two random variables, the variables under analysis must be shown to exhibit fat-tailed distributions. In order to test that both exchange rate and commodity returns follow fat-tailed distributions, we implement two non-parametric approaches: the Hill tail index (Hill, 1975) and the Dekkers et al. (1989)'s tail index indicator. We now describe the approach to estimate both non-parametric indices.

Consider order statistics of a random variable X of sample length n, such that X is sorted in descending order, as follows

$$X_{(1)} \geq X_{(2)} \geq ... \geq X_{(n)}$$

Equations 3.2 and 3.3 denote the Hill tail index (H) and its variance, respectively.

$$H = \frac{1}{k}\sum_{i}^{k} log\left(\frac{X_{(i)}}{X_{(k)}}\right) \tag{3.2}$$

$$M = \frac{1}{k}\sum_{i}^{k}\left[log\left(\frac{X_{(i)}}{X_{(k)}}\right)\right]^2 \tag{3.3}$$

where $X_{(i)}$ corresponds to the $i-th$ ordered statistic of the random variable X and $X_{(k)}$ indicates the threshold given by a positive integer number k representing the number of observation above the threshold. Equation 3.4 displays the Dekkers et al. (1989)'s tail indicator which we denote by γ.

$$\gamma = 1 + H + \frac{1}{2}\left(\frac{\frac{M}{H}}{H - \frac{M}{H}}\right) \tag{3.4}$$

where H and M corresponds to the Hill tail index and Hill tail index variance we obtain in equations 3.2 and 3.3, respectively. The variance of γ is given by $1 + \gamma^2$.

The γ tail indicator corresponds to the inverse of the tail index α (i.e.: $\gamma = 1/\alpha$), where α corresponds to the shape parameter determining the fatness of the tail of a random variable distribution. A higher value of γ (equivalent to a lower value of α), indicates fatter tails in the distribution of a random variable. In this study, we interpret statistically positive values of H and γ as evidence suggesting log-return of commodity and exchange rate follow fat-tailed distributions.

The Hill tail index is an unbiased estimator and is also more efficient in comparison to alternative tail index indicators as pointed out by both Tsay (2010) and Cumperayot and De Vries (2017). However, the indicator assumes that the data comes from a fat-tailed distribution. In contrast, the γ index is more flexible since it does not assume *a priori* any specific distribution in the data, enabling us to estimate the existence of fat tails without imposing any priors.

3.4.4 Asymptotic dependence

In order to analyse the information contained in the tails of the log-return distributions (i.e. extreme values) and how those observations are related in a multidimensional framework, much related literature utilises the concept of tail dependence. Tail dependence measures the probability that extreme values of one random variable occur given that extreme values of another random variable also arise simultaneously. In other words, it is a measure of the joint probability of the contemporaneous occurrence of extreme changes in two random variables. Formally, let X and Y be two random variables, the upper tail dependence, denoted by λ_{upper}, is given as follows:

$$\lambda_{upper} = \lim_{q \to 1} \Pr\left\{F_X(X) > X_q \mid F_Y(Y) > Y_q\right\} \tag{3.5}$$

where F_X and F_Y represent the cumulative distribution functions, and X_q and Y_q denote threshold values of X and Y, respectively. Thresholds are defined as a function of the q-th percentile containing the most extreme observations. In particular, λ_{upper} in equation 3.5 represents the probability of occurrence of high extreme values in X (i.e.: those above the threshold X_q) conditional to the occurrence of high extreme values in Y (i.e.: those above the threshold Y_q). Analogously, we can define lower tail dependence as follows:

$$\lambda_{lower} = \lim_{q \to 0} \Pr \left\{ F_X(X) < X_q \mid F_Y(Y) < Y_q \right\} \tag{3.6}$$

where λ_{lower} denotes the probability of occurrence of low extreme values in X (i.e.: those below the given threshold X_q) conditional to the occurrence of low extreme values in Y (i.e.: those below the given threshold Y_q). In the same fashion, tail dependence between high (low) extreme values of X and low (high) extreme values of Y can be defined as the probability of simultaneously observing values of X above (below) X_q and values of Y below (above) Y_q.

In order to test for existence of a relationship between exchange rate and commodity extreme returns this paper computes an asymptotic dependence indicator (ADI) which is a non-parametric estimator of the tail dependence depicted in equation 3.5. Let X and Y be constituents of the time-series of commodity and exchange rate returns, respectively, both of sample size n, then the non-parametric tail dependence estimator we implement in this study, following de Haan and Ferreira (2007), is given by:

$$\text{ADI} = \frac{1}{k} \sum_{i=1}^{n} \mathbf{1}_{\{X_i \geq X_k, Y_i \geq Y_k\}} \tag{3.7}$$

where X_i and Y_i denotes the ith observation of commodity and exchange rate returns, and X_k and Y_k correspond to the selected threshold of commodity and exchange rate returns, respectively. k represents the number of observations above the threshold and is determined by the choice of a specific percentile capturing the extreme returns of each variable (i.e. those returns located in the upper tail of the distribution of each variable which represent the most extreme positive returns). $\mathbf{1}_{\{\cdot\}}$ is an indicator function of size n which takes the value of one when X_i and Y_i are simultaneously greater than their thresholds X_k and Y_k, respectively, and takes the value of zero otherwise. A statistically positive value of the ADI in equation 3.7, suggests that two random variables are asymptotically dependent. In our context, the ADI captures the extent of tail dependence between extreme exchange rate and commodity returns. By definition, the specific value of the ADI is interpretable as the probability of observing that a pair of observations of commodity and exchange rate returns both simultaneously lie above their given thresholds.

The relevance of asymptotic dependence analysis arises from its contribution to explaining the relationship between two variables by focusing upon the link between them on the basis of the information contained in the tails of their distributions. This is particularly relevant when two (or more) variables appear to exhibit a low degree of correlation, but nevertheless there is an underlying relationship which is driven by the information contained in the tails. This situation appears to characterise the relationship between the log-returns of both exchange rates and commodity prices.

[Figure 3.2 in here]

Consider figure 3.2 showing a scatter plot of Chilean peso returns (y-axis) and copper returns (x-axis) using daily data from Jan-2000 to July-2018. As the figure indicates, the majority of observations in the scatter plot tend to concentrate around the origin with no clear emergent relationship evident. However, there are some extreme observations in both variables located in the tails of the distribution, whose contemporaneous occurrence may help to explain the statistically documented negative relationship between these variables. Importantly, this negative relationship is precisely what one would expect to observe in the presence of an asymptotic dependence between the two variables' returns in the analysis we undertake in the following sections. In such a case, it is the information contained in the tails of the distributions which is crucial in describing the link between two variables and in transmitting shocks impacting one variable to the other. In contrast, two random variables that both exhibit heavily-tailed distributions are not necessarily asymptotically dependent.

Several other studies analysing tail dependence also adopt the asymptotic dependence indicator we implement in this research. In particular, Cumperayot and De Vries (2017) examine the relationship between large swings of exchange rate and classic monetary fundamental, Cizek et al. (2005) study tail dependence in the context of financial risk management, specifically an empirical analysis of FOREX and stocks markets, Abberger (2005) employs an ADI measure to analyse tail dependence among German stock returns, and Fisher and Switzer (2001) use it to study the properties of the relationship between extreme observations based on simulated data.

The asymptotic dependence estimator we implement in this study (also known as the empirical copula), is a specific way to analyse the probability of the simultaneous occurrence of extreme movements in two variables, and is also an alternative standard procedure to model the tail behaviour of two random variables. The advantage of the present indicator results from its simplicity, particularly the fact that unlike other methods, we do not need to assume a specific functional form for the distributions of exchange rate and commodity returns. We note that it is also implementable when the scale of the variables under analysis is different.

Related studies provide alternative procedures to estimate the tail dependence between two random variables. A common approach in empirical finance focuses on modelling the joint distribution of two or more variables using the copula methodology, in an attempt to capture the entire dependence structure between two random variables (Cizek et al., 2005; Ruppert, 2011). However, the copula approach makes strong assumptions about the joint cumulative distribution function (CDF) of random variables. As a result, estimates of the degree of tail dependence are highly sensitive, and may vary significantly according to different assumptions made regarding the appropriate functional form (type of copula) selected to model the CDF of the relevant returns (Frahm et al., 2005). Poon et al. (2004) offer an alternative approach to capture tail dependence enabling one to measure the extreme linkage between two random variables and quantify the degree of association. Similar to the copula estimation, this method also requires strong assumptions to be satisfied by the distributions of the variables under analysis. Moreover, using Monte Carlo simulations, Fernandez (2008) highlights certain questionable performance issues surround the Poon et al. (2004) measure in detecting asymptotic dependence when the relationship between two variables is weak, and concludes that the ADI approach (empirical copula analysis) is a more suitable method for measuring the degree of asymptotic dependence between two random variables.

3.5 Results

This section reports the results of both the out-of-sample forecast and the asymptotic dependence measure.

3.5.1 Out-of-sample forecasts

This section presents the results of the out-of-sample exercises using the baseline forecasting model we introduce in equation 3.1.

[Table 3.2 in here]

Table 3.2 provides a comparison of the RMSFE ratio of the commodity-based model (numerator) and the RW model (denominator) for the countries under analysis (rows) and for different forecast horizons (columns). We use two version of the RW model: without and with drift in panels A and B, respectively. The results indicate that in all cases and for every forecast horizon, the commodity-based model generates superior forecasts to both RW models. In addition, the Diebold-Mariano test indicates that the RMSFE of the commodity-based model is significantly lower than the RMSFE of both comparator RW models at the 1% level. These findings are consistent with previous studies. For instance, Ferraro et al. (2015) and Kohlscheen et al. (2017) show that commodity-based models prove to be superior to the RW models with and without drift, using daily observations.

In addition, our results appear robust to the inclusion of what the literature terms, unobservable global factors, which influence financial returns. Related studies maintain that the VIX index, by accurately capturing global risk appetite, is an appropriate gauge of those unobservable global factors.[5] They argue that changes in global risk appetite affect investors' decisions which feed through to impact exchange rate and commodity returns. Adrian et al. (2009) and Kohlscheen et al. (2017) emphasise the role of global risk appetite, captured by the VIX index, in forecasting exchange rates. Other authors discuss how exchange rate movements may be linked to global risk factors such as the VIX index (Gourio et al., 2013) and equity market volatility (Lustig et al., 2011) in the context of carry trade strategies. Importantly for empirical implementation, the VIX index is a variable that is available at a daily frequency, so we are able to capture the short-lived impact of global risk aversion on the relationship between exchange rate and commodity returns. In order to include the effect of global common factors, we obtain exchange rate and commodity returns orthogonal to

[5]The VIX is an index computed by the Chicago Board Options Exchange (CBOE) that estimates the implied market volatility obtained from option prices for equities traded on the S&P 500.

the VIX index by running the following regression for each country and commodity in the sample:

$$r_t = \alpha_0 + \alpha_1 dVIX_t + \nu_t \tag{3.8}$$

where r_t corresponds to the appropriate time-series of exchange rate or commodity log-returns, $dVIX$ is the change in the VIX index, and α_0 and α_1 are coefficients we estimate using ordinary least squares (OLS). We interpret the resulting error terms arising from equation 3.8 (ν_t) as the relevant country's exchange rate and commodity log-returns which are orthogonal to changes in the VIX index. In order to test for the potential influence of this global factor in driving our benchmark findings, we replicate the previous out-of-sample exercise using those exchange rate and commodity returns which are orthogonal to the change in the VIX index. The results are given in table 3.2 panel C and reveal that after controlling for the potential effect of the global common factor, the conclusion remains unaffected. Commodity returns provide superior forecasts to a random walk model without drift for all the countries in the sample. These results help alleviate concerns regarding simultaneity bias that may arise due to the potential presence of omitted factors in our benchmark forecasting regressions.

Furthermore, we implement the out-of-sample forecasting exercises of table 3.2 panel A using alternative definitions of the base currency to the USD, namely the euro (EUR) and pound sterling (GBP). This will help dispel any suspicions that the results arise owing to a "dollar effect". Panels D and E in table 3.2 report the results using the EUR and GBP as the numeraire currency, respectively. The results generally confirm our findings with the forecasting ability of commodity-based models remaining significant even using alternative base currencies. Some minor exceptions appear. In the case of Peru, whether the base currency is GBP or EUR, the evidence is now insignificant, although the relevant RMSFE ratio still remains below 1 when using EUR as the base currency. For South Africa, when using EUR as the base currency, statistical significance is now only evident at the 10% level. Overall, the results of this simple robustness exercise confirm that the forecasting ability of commodity returns is not restricted to using the USD as a base currency, suggesting the predictive power goes beyond a mere dollar effect.

3.5.2 Lagged commodity returns

This section undertakes *truly* out-of-sample forecasting exercises using lagged commodity returns rather than contemporaneous returns. In this process we replace equation 3.1 with equation 3.9.

$$\Delta s^f_{t+h} = \hat{\alpha}_t + \hat{\beta}_t \Delta p_t, \quad t = R, R+1, ..., T-h. \tag{3.9}$$

where Δs^f_{t+h} represents the *h*-periods-ahead forecast of exchange rate log-returns and

Δp_t denotes the country-specific commodity log-returns. We generate OLS estimates of the coefficients $\hat{\alpha}_t$ and $\hat{\beta}_t$ in equation 3.9 using a sample window of length R ($R = 2135$ observations), spanning 04/01/2000 to 10/04/2009, which corresponds to half the total sample of T observations ($T = 4270$), and then produce h-step ahead forecasts. We then roll forward the window one observation, re-estimate equation 3.9 over the window from 05/01/2000 to 13/04/2009 and generate a new series of h-step ahead forecasts. We repeat this process up to $T - h$ to produce a series of forecasts for the full out-of-sample period.

[Table 3.3 in here]

Table 3.3 depicts the results comparing the commodity-based models with a RW without drift and with drift in panels A and B, respectively. As the results in panel A show, the forecasting ability of commodity returns disappears when the explanatory variable is replaced by its lagged values. This evidence reveals that the commodity-based model using lagged commodity returns is unable to outperform the driftless RW, although when testing the commodity-based model using lagged commodity returns against a RW model with drift the evidence still supports a superior forecasting ability attributable to lagged commodity returns. Table 3.3 panel B indicates that the commodity based model using lagged commodity returns forecasts better than a RW with drift, and the results are statistically significant.

Additionally, in order to control for the effect of global common factors, we replicate the estimation in table 3.3 panel A, but now using exchange rate and commodity returns which are orthogonal to changes in the VIX index. Table 3.3 panel C displays the results utilising the driftless RW model as the benchmark. The evidence shows that after correcting for the effect of global factors (changes in the VIX index), the forecasting ability of our commodity-based model once again vanishes, similar to the results in table 3.3 panel A. The evidence of the forecasting ability of our commodity-based model corroborates the findings of Ferraro et al. (2015) and Kohlscheen et al. (2017). First, they similarly highlight that lagged commodity returns exhibit a lower forecasting ability in comparison to contemporaneous values when the benchmark is the driftless RW. Second, as in our study, they demonstrate that there is still some evidence suggesting commodity-based models continue to forecast better than the RW with drift. However, they do not appear robust to including controls for the potential impact of other global factors as captured in the VIX index.

3.5.3 Using low frequency data

This section analyses the forecasting ability of commodity based models using lower frequency observations. Following the standard procedures outlined in Ferraro et al. (2015), we compute monthly and quarterly observations using the end-of-sample daily frequency. According to Rossi (2013), in comparison to computing a monthly or quarterly average from daily observations, using end-of-sample observations makes the task of finding evidence of forecasting ability more onerous.

[Table 3.4 in here]

Table 3.4 presents a comparison of the results of the model in equation 3.1 using contemporaneous commodity returns at a monthly frequency against the RW model as a benchmark. As we observe in panel A, where the benchmark corresponds to the driftless RW model, the forecasting ability of the commodity-based model decreases in comparison to the daily data case for the majority of the countries in our analysis. In general terms, there is either no statistical evidence or it is only marginally significant at a 10% level, in favour of commodity returns. There is still some predictive ability in evidence for commodity returns at 5% level in the cases of Canada, Chile and Norway. We observe similar results when comparing the predictive ability of our commodity-based model against a RW model with drift (Panel B). Even though the reduction in the forecasting ability decreases in comparison to the daily frequency, there are still some highly significant cases in favour of commodity returns generating superior forecasts to the RW with drift, such as Canada and Norway, while other countries, namely Chile, Colombia and Russia are significant at 5% level. However, as we note previously, the statistical significance in these cases may derive from the fact the benchmark model, the RW with drift, it is not the most difficult benchmark to beat (Rossi, 2013).

As with the previous daily observations, we subsequently undertake a *truly* out-of-sample forecast exercise, by including lagged commodity returns as the main explanatory variable, estimating equation 3.9. Table 3.4 panels C and D present the results using as benchmarks the RW model with and without drift, respectively. No matter which benchmark model we employ, we find that the forecasting ability of the commodity-based model completely disappears for every country.

Finally, in order to account for the effect of global factors, we replicate the analysis with monthly frequency observations using exchange rate and commodity returns orthogonal to the VIX index. Table 3.4 panels E and F displays the results against a benchmark RW without drift model. Table 3.4 panel E displays the results using contemporaneous commodity returns, and shows that the forecasting ability of such returns either disappears or it is only marginally significant at the 10% level, apart from Canada, where the commodity-based models display some forecasting ability at the 5% level. The results in table 3.4 panel F indicate that, after controlling for the effect of the VIX, the forecasting ability of commodity-based models using lagged commodity returns again completely disappears at the monthly frequency.

In addition, we replicate the previous forecast exercises using quarterly observations and reach the following conclusions.[6] First, by using contemporaneous commodity return observations, the forecasting ability of commodity-based models declines even further relative to those generated by daily and monthly frequency estimations. Second, in comparison to our daily and monthly frequency results, the forecasting ability of lagged commodity returns disappears completely for all countries in the sample. Both these sets of results obtain whether our benchmark model is defined as a RW with or without drift and also after controlling for the effect of global factors as represented by the VIX index.

[6] Results are available upon request.

The results of this section highlight the relevance of the data frequency in forecasting exchange rates using commodity-based models. We demonstrate that after reducing the frequency of the data, from daily to either monthly or quarterly observations, the forecasting ability of commodity returns decreases in both *pseudo* out-of-sample and *truly* out-of-sample exercises. The results hold for both benchmark models we utilise, namely the RW with and without drift. Importantly, controlling for the effect of global factors appears relevant when using observations at a lower frequency. Our evidence indicates that, after controlling for changes in the VIX index, the forecasting ability of commodity returns declines even further in comparison to cases which use observations at a daily frequency.

Our results are consistent with recent studies (Ferraro et al., 2015; Kohlscheen et al., 2017) and reinforce the idea that using observations at a daily frequency is crucial in capturing any dependent relationship between the variables which underpins the model's forecasting success. Contemporaneous commodity returns exhibit a superior forecasting ability in comparison to lagged commodity returns. Thus, there appears to be a short-lived relationship between the variables which is mostly captured when analysing the contemporaneous relationship between commodity and exchange rate returns. Moreover, by lowering the data frequency, any relationship between the variables tends to vanish and as a result, the forecasting power of commodity-based models also decreases. This evidence highlights the relevance of daily data observation frequency in forecasting exchange rates. In this sense, commodity price shocks impacting exchange rates are transitory and tend to dilute quickly over time as economic agents internalise new information. Low frequency observations appear unable to capture such transitory information flows, consequently commodity returns sampled at a lower frequency are not useful in predicting exchange rates.

3.5.4 Fat-tailed distributions of log-returns

We now proceed to investigate the central focus of this paper, the role of any measured asymptotic dependence in explaining the performance of commodity returns in forecasting exchange rate movements. Initially, we seek to determine if the distributions of log-returns of commodity prices and exchange rates both exhibit fat-tails. Establishing this fact is a necessary but not sufficient condition to analyse if asymptotic dependence is underpinning the documented ability of commodity returns to predict exchange rate variations.

[Table 3.5 in here]

Table 3.5 panel A reports the results of estimating the Hill tail index we define in equation 3.2 for both the lower and upper log-returns tails, representing the most negative and positive log returns, respectively. Confidence intervals at the 99% level are also included in parentheses. As shown, in all cases and also for both upper and lower tails, the indicator is positive and statistically different from zero indicating that the distribution of log-returns of each variable follows a fat-tailed distribution.

A more conservative evaluation of fat-tailed distributions is also undertaken using the Dekkers et al. (1989) index. Table 3.5 panel B presents the results of the γ tail index estimator in equation 3.4. As the table shows, the majority of currencies and commodity returns exhibit heavily-tailed distributions at least in one of the tails. Overall, despite the occurrence of some exemptions and acknowledging its relative inefficiency compared to the Hill tail index, we interpret the γ tail index estimator as supporting the presence of heavy tails in both the log-returns commodity and exchange rate distributions, allowing us to conclude that there may be relevant information contained in the tails of the distribution, a possibility we examine further by undertaking asymptotic dependence analyses.

3.5.5 Asymptotic dependence

In this section we present the results of our selected asymptotic dependence measure introduced in section 3.4.4. Before analysing the results, it is appropriate to consider a taxonomy of the important cases we wish to analyse, which is informed by economic intuition. As discussed in section 3.4.4, the ADI captures the asymptotic dependence between our chosen pair of random variables, the log-returns of nominal exchange rates and commodity prices. When combined with the nature of our study, this leaves four scenarios we need to investigate, corresponding to different combinations of the two distributional tails of each of the random variables under analysis. We discuss the economic interpretation and relevance of these cases when we present the results of our investigations. In the meantime, the four cases of interest are the following:

- **Case 1**: An increase in the country-specific commodity price (upper tail of commodity log-return distribution) and a nominal exchange rate appreciation (lower tail of exchange rate log-return distribution).
- **Case 2**: A reduction in the country-specific commodity price (lower tail of commodity log-return distribution) and a nominal exchange rate depreciation (upper tail of exchange rate log-return distribution).
- **Case 3**: A reduction in the country-specific commodity price (lower tail of commodity log-return distribution) and a nominal exchange rate appreciation (lower tail of exchange rate log-return distribution).
- **Case 4**: An increase in the country-specific commodity price (upper tail of commodity log-return distribution) and a nominal exchange rate depreciation (upper tail of exchange rate log-return distribution).

Table 3.6 panel A reports the results of the ADI measures for each country using daily contemporaneous commodity returns and exchange rate returns for the period from 03 Jan 2000 to 18 Jul 2018. We compute confidence bands using bootstrap procedures over 5000 re-sampling iterations. As the table shows, the asymptotic dependence index is positive and statistically significant for all countries in cases 1 and 2, with the single exception in case 1 for Peru. In contrast, for cases 3 and 4, the index both reduces in magnitude and becomes statistically insignificant for all countries, with the exception of South Africa.

[Table 3.6 in here]

Our asymptotic dependence measure also appears robust to a set of alternative specifications. First, both our commodity prices and exchange rates in these initial estimates are denominated in USD. As such, it is important to establish that the asymptotic relationship between the variables does not simply reflect a dollar effect. Table 3.6 panels B and C present the ADI results using euros and pound sterling as base currencies for the bilateral exchange rates. Our results show that the asymptotic dependence between exchange rates and commodity returns remains generally significant in cases 1 and 2 for the majority of counties, while those for cases 3 and 4 are again mostly statistically insignificant. Second, our estimation of the ADI is robust after controlling for the effect of both global factors and heteroskedasticity in log-returns. Following the discussion in section 3.5.1, we highlight the relevance of controlling for the effect of global risk appetite which may influence the dynamic of commodity and exchange rate returns. In order to control for this global factor, as before, we use commodity and exchange rate log-returns that are orthogonal to changes in the VIX index, following the same procedure we implement in section 3.5.1, and estimating equation 3.8 in order to obtain time-series of commodity and exchange rate returns free of the effect of global risk appetite. In addition, due to the intrinsic tendency of financial time-series of log-returns to display volatility clustering, i.e.: large (small) returns to be followed by large (small) returns, there is a possibility that heteroskedasticity is responsible for generating some degree of dependence in the time-series of log-returns which may impart bias to the result of the asymptotic dependence analysis (Kearns & Pagan, 1997; McNeil & Frey, 2000). To control for heteroskedasticity issues, we estimate our ADI measure using standardised residuals for both commodity and exchange rate returns orthogonal to the VIX index, thereby controlling for possible heteroskedastic effects. This involves estimating the following ARCH(1) model for each series of commodity and exchange rate log-returns:

$$r_t = \beta + u_t \tag{3.10}$$

$$\sigma_t^2 = \omega + \alpha u_{t-1}^2 \tag{3.11}$$

Equation 3.10 corresponds to the univariate mean equation modelling log-returns orthogonal to changes in the VIX index and u_t denotes the residual of the mean equation. Equation 3.11 represents the univariate conditional variance of log-returns orthogonal to changes in the VIX index. $\alpha, \beta,$ and ω are coefficients to be estimated using maximum likelihood. Finally, we compute the standardised residuals as $\varepsilon_t = u_t/\sigma_t$. Table 3.6 panel D depicts the results of the ADI estimation using contemporaneous commodity returns, controlling for the effect of both global factors and heteroskedasticity. Our results reveal that for all countries, our estimates of ADI in cases 1 and 2 are statistically significant, while in cases 3 and 4 they are statistically insignificant. This evidence supports two important considerations. First, our findings suggest that the asymptotic dependence measure goes beyond a mere global risk factor that may affect both exchange rate and commodity returns. These results help to alleviate simultaneity bias concerns that may arise due to the potential for omitted factors. Second, we note that our results controlling for the effect of global factors and heteroskedasticity are somewhat lower in magnitude in comparison to the ADI estimates which do not control for these phenomena (table 3.6 panel A). These findings are consistent with related studies documenting that the extent of asymptotic dependence is overestimated when failing to take into account the effect of heteroskedasticity (Kearns & Pagan, 1997; McNeil & Frey, 2000). They also confirm that our ADI measure appears to be more than a simply spurious relationship generated by the effect of heteroskedasticity in log-returns.[7]

Collectively, these results provide empirical support for the intuitive conjectured economic relationship between commodity and exchange rate returns that we discuss in section 3.3. Extreme commodity price shocks generate expectations of changes in real economic activity in commodity-exporting economies which transmits through to impact exchange rates. In particular, a sudden large decrease (increase) in the price of the major country-specific export commodity is associated with a deterioration (improvement) in the economic outlook of that commodity exporting economy. As a result of this sudden deterioration (improvement) in confidence relating to future economic prospects, a sharp depreciation (appreciation) of the nominal exchange rate occurs.

If this economic intuition underpins the presence of asymptotic dependence between returns as a necessary condition for forecasting success, it will only make sense when it is detected in cases 1 (exchange rate appreciation concomitant with an increase in commodity prices) and case 2 (exchange rate depreciation together with a reduction in the commodity's price), and not in cases 3 and 4. This expected negative relationship between the variables is precisely that which receives empirical support in the data, as we illustrate in figure 3.2. Importantly, cases 3 and 4 report low ADI values and reveal no evidence of statistical significance for any of the countries after correcting for the effect of global factors and heteroskedasticity in log-returns.

We interpret our findings as reflecting the impact of news and its ability to convey price relevant information from one market to another. As widely reported, news often exerts a significant impact on asset return volatility, with episodes characterised by frequent news arrival being associated with periods of high return volatility. Prior evidence reveals the same logic also applies to commodity markets, with several studies (Caporale et al. 2017; Frankel and Hardouvelis 1985; Roache and Rossi 2010), documenting that the arrival of news significantly impacts commodity price volatility. According to these studies, it is news affecting commodity markets which represents the key driver of volatility in commodity prices which translate into leptokurtic commodity log-return distributions in which the extreme observa-

[7]Following the ARCH(1) model estimation, Engle's Lagrange multiplier test for standardised residuals reveals there is no remaining heteroskedasticity after modelling the conditional variance with this model specification. Therefore, this approach reduces concerns relating to the bias that heteroskedasticity may generate to our results. Results of this Engle test are available from the authors on request.

tions generate fat tails.

In particular, our interpretation of the asymptotic dependence measure relates to the arrival of unexpected commodity relevant information generating sudden changes in commodity prices. These external shocks convey information relevant to the economic prospects of commodity exporting economies, changing investor perceptions of both future exports and economic activity, and are subsequently transmitted into their exchange rates. Our measure of tail dependence allows us to more precisely quantify how extreme values of exchange rate log-returns relate to extreme values of commodity log-returns. Consequently, the measured relationship between the tails of the relevant return distributions provides us with an estimate of the extent to which the arrival of news relating to commodity markets is transmitted into exchange rates. The fact this occurs and is measured in the variables contemporaneously at a daily interval suggests this incorporation occurs fairly quickly (within the daily interval), testimony to the documented efficiency of foreign exchange markets. Our stance is consistent with Ferraro et al. (2015) who also interpret commodity price shocks as the mechanism conveying information about macroeconomic news that may affect exchange rates. The ADI results we document in this study provide empirical support for this previously investigated conjecture.

3.5.5.1 A case study of Chile

As a further revealing of anecdotal illustration of the above mechanism, we now provide a country-specific example using the relationship between copper and Chilean peso exchange rate returns. Chinese demand for copper is considered as the most important single factor driving the international copper price, as China accounts for approximately half of global copper imports.[8] As Bailliu et al. (2019) demonstrate, the sustained economic expansion of both the housing market and infrastructure investment in China, two copper intensive sectors, underpins the country's copper demand. Indeed, Kruger et al. (2016) estimate that the Chinese housing market development accounts for 85% of the copper price increase during the first decade of the 2000s. Moreover, Chinese economic activity correlates closely with developments in the global copper market. Kolerus et al. (2016) estimate that a 1% increase in China's industrial production generates an increase in global metal prices (iron, copper, nickel, lead, and tin) of between 5% and 7% over a one-year horizon. Importantly for our purpose, they empirically estimate that the impact on global metal prices of news concerning China's industrial activity is comparable in magnitude to the effect of U.S. industrial production announcements. This evidence strongly suggests news about the current economic situation in China tends to impact the global price of copper.

[Figure 3.3 in here]

Figure 3.3 plots the value of the asymptotic dependence indicator for our case 2 (solid blue line), namely a large fall in copper prices and a large depreciation in the Chilean exchange

[8] According to the international trade statistics of the United Nations (Comtrade), China's copper demand corresponds to 49.4% of the world copper imports in 2018, followed by Japan (14.3%) and South Korea (6.6%).

rate, while the shaded grey areas correspond to those periods when negative economic news (negative economic surprises) occurs in relation to China. We define the negative economic news episodes as those corresponding to the Citigroup China Economic Surprise Index we obtain from Bloomberg. As figure 3.3 indicates, during episodes of negative economic news in China, we customarily observe an increase in our asymptotic dependence measure. This signals that the negative news from China, which is associated with a decrease in copper prices, in turn is correlated with large depreciations of the Chilean nominal exchange rate. This pattern is consistent with the view that certain aspects of this negative news from China, reflecting decreases in the demand for copper, are also transmitted to the Chilean exchange rate via the effect of resulting changes in the copper price.[9]

3.5.5.2 Asymptotic dependence using lagged commodity returns

We now examine the effect of the timing of information availability on our asymptotic dependence measure. Recall that previously we show that it is only when utilising contemporaneous, as opposed to lagged, commodity returns that we uncover any improved forecasting performance of such returns for exchange rate changes. If asymptotic dependence underpins this superior performance, we should find that evidence for its existence is largely absent when we estimate our ADI using one period lagged commodity returns in conjunction with current period exchange rate changes.

[Table 3.7 in here]

Table 3.7 panel A shows the ADI estimation using one period lagged commodity returns at a daily frequency. The results show that in cases 1 and 2 the magnitude of asymptotic dependence reduces considerably relative to when using contemporaneous commodity returns, and indeed the indicator is no longer statistically different in most situations. We obtain similar results for cases 3 and 4, which mainly reveal statistically insignificant indicators. Results using one period lagged commodity returns at a daily frequency and correcting for the effect of both global risk appetite and heteroskedasticity display an even clearer picture. As table 3.7 panel B depicts, the ADI decreases even further in magnitude in comparison to the ADI when using contemporaneous commodity returns, and it is insignificant for the majority of cases and countries with very few exceptions. Overall, on the basis of this evidence, we conclude that any measured asymptotic dependence between nominal exchange rate returns and lagged commodity returns at a daily frequency is much weaker if it exists at all after controlling for both the effect of global risk appetite and heteroskedasticity. This is especially evident when comparing the ADI results using lagged commodity returns with the documented presence of ADI in cases 1 and 2 using contemporaneous commodity returns.

[Table 3.8 in here]

[9]Naturally, although China is an important global agent influencing demand in the copper market, there is an imperfect correlation between events in figure 3.3 and our asymptotic dependence measure, as news about economic activity in China only represents a fraction of the potential shocks affecting the copper market, with other major elements also playing a role in explaining international price changes.

3.5.5.3 Asymptotic dependence using low frequency data

We also analyse the effect of data frequency on the relationship between exchange rate and commodity returns, controlling for the effect of both global risk appetite and heteroskedasticity. Table 3.8 panel A presents a comparison of the asymptotic dependence measure between exchange rate returns and contemporaneous commodity returns computed at daily, monthly and quarterly frequency. As we previously document, the asymptotic dependence measure is statistically significant at a daily frequency only for cases 1 and 2, while cases 3 and 4 are statistically insignificant. However, when computing our measure at a monthly frequency, almost all evidence of asymptotic dependence vanishes; Chile in case 2 provides the sole exception. Likewise, when we replicate the exercise using a quarterly frequency, our measure reveals no statistical significance whatsoever within the entire sample. The results we obtain using lower frequencies contrast with the evidence provided in Cumperayot and De Vries (2017), who show that the asymptotic dependence between classical monetary fundamentals and exchange rates is still present when they use data at a quarterly frequency. In the present analysis, large commodity returns measured at lower frequency, either at a monthly or quarterly frequency, tend to be unrelated to large exchange rate movements. Similarly, when using lagged commodity returns compiled utilising low frequency data to estimate our asymptotic dependence measure we obtain analogous results. Table 3.8 panel B reveals that after controlling for the effects of both global risk appetite and heteroskedasticity, evidence of asymptotic dependence completely disappears for all countries in the sample at monthly and quarterly frequency.

3.5.5.4 Exchange rate forecasting and asymptotic dependence

Our results allow us to draw two main conclusions. First, timing plays a key role in describing the relationship between exchange rate and commodity returns. Any forecasting ability of commodity returns for exchange rates and also the asymptotic dependence between the variables tends to be short-lived, specifically only contemporaneous observations can capture that relationship. The out-of-sample forecasting ability of commodity returns and the asymptotic dependence between exchange rate and commodity returns are highly significant in contemporaneous terms, while both tend to disappear when lagged commodity returns are included in the analysis. This suggests foreign exchange markets process and incorporate new information into exchange rates within an intraday time period. Second, any asymptotic dependency between commodity and exchange rate returns is transitory in the data sample. The forecasting ability of commodity returns and the asymptotic dependence between the variables is highly significant when observations are sampled at a daily frequency. In contrast, both the forecasting ability and the asymptotic dependence tend to disappear when we utilise lower frequencies, either monthly or quarterly observations. Overall, our evidence suggests that the documented ability of commodity returns to predict exchange rate returns is closely associated to the asymptotic dependence between those variables. We show that a highly significant asymptotic dependence between commodity and exchange rate returns, i.e.: a measure of the link between the information contained in the tail of commodity and exchange rate return distributions, is a key and necessary component underpinning the forecasting ability of commodity returns. On the contrary, in the absence of a clear asymptotic dependence relationship, we observe a substantially reduced or insignificant ability of commodity returns to predict exchange rate movements.

On the basis that large swings in commodity prices are driven by the arrival of unexpected news concerning market supply and demand conditions, our ADI measure captures how such information, which underpins large swings in commodity prices, also relates to large movements in exchange rates. The impact of unexpected news is transitory as economic agents trade to quickly internalise such surprises in prices.[10] As we document, both the asymptotic dependence and the forecasting ability of commodity returns are highly significant when we use contemporaneous daily observations, while there is a reduction in the statistical significance of both elements when employing lagged observations or those at a low frequency. Following our evidence, the information contained in the tails of the distributions appears to be the key component of the ability of commodity returns to forecast exchange rate movements. In the absence of any measured asymptotic dependence the forecasting ability of commodity returns is no longer evident.

It is worth noting that the proposed transmission mechanism provides a general framework to explain empirically how commodity price shocks are transmitted to exchange rates. This mechanism has recently benefited from the robust theoretical underpinnings outlined in Cumperayot and De Vries (2017). Our findings incorporating the effect of global risk appetite (VIX index) help alleviate concerns relating to omitted variable bias issues, where an unobserved third factor may drive the dynamics of both variables. Moreover, if commodity and FOREX markets are segmented then it is less likely that such omitted factors drive our results.[11] In the same vein, demonstrating causality between variables is beyond the scope of this research, in which we focus on the importance of asymptotic dependence as a key constituent of the ability of commodity returns to forecast movements in exchange rates in out-of-sample tests.[12]

[10]Related literature supports the fact that the effect of unexpected news on financial variables tend to happen in the short run and vanishes over time. For example, Chaboud et al. (2008) show that U.S. macroeconomic news has a significant impact on the euro and the Japanese yen only at an intraday frequency, while Kilian and Vega (2011) show that U.S. macroeconomic news measured at a monthly frequency show no impact on oil prices.

[11]Related studies support the idea that commodity are actually segmented markets. For instance, in the context of asset pricing models, Daskalaki et al. (2014) document that foreign exchange risk factors are not able to price the cross-section of a set of commodity returns. The authors also show that commodity markets are segmented from one another, meaning that each commodity follows its own price dynamics and its price movements are not generally related to another commodity type. The authors conclude that the models usually employed to price assets, such equity, are not able to explain commodity returns. Similarly, Skiadopoulos (2013), shows that there is no common factor between commodity futures prices and other financial assets such as bonds and equities.

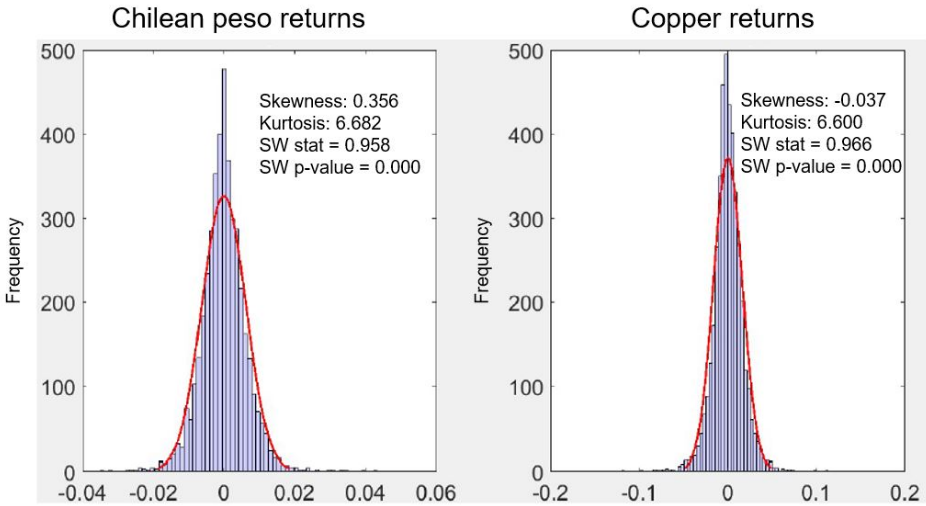
[12]Most studies argue that the causal effect goes from commodity prices to exchange rates highlighting the exogeneity of commodity prices (Cashin et al., 2004; Chen & Rogoff, 2003; Ferraro et al., 2015; Kohlscheen et al., 2017). Related articles analyse the direction of the causal relationship in more detail. For instance, Zhang (2017) documents that causality goes from commodity to exchange rate returns, particularly at short horizons. Similarly, Ahmed (2019) shows that in the short-run, using high frequency data, the causation goes from oil prices to exchange rates. Chen et al. (2010) is an exception suggesting that exchange rate changes may actually predict commodity returns. Their analysis bases on observations at a quarterly frequency. However, as we show in our study, most of the relation between exchange rate and commodity returns take place at daily frequency. In particular, the asymptotic dependence between commodity and exchange rate returns (which we argue is the central element underpinning the forecasting ability of commodity returns) takes place at daily frequency and disappears at lower frequency.

3.6 Conclusions

In this study we analyse the non-linear relationship between commodity and exchange returns using a sample of nine commodity-exporting economies. Our contribution to the literature lies in both quantifying the degree of non-linear relationships between exchange rate and commodity returns, and explaining how this association contributes to our understanding of the documented ability of commodity returns to forecast exchange rate movements. To the best of our knowledge, this is the first study analysing the link between exchange rate and commodity markets based on a tail dependence framework. Our evidence suggests that the non-linear relationship between the variables is a key element that underpins the ability of commodity returns to contemporaneously forecast exchange rate returns. We capture this non-linear relationship using an asymptotic dependence measure, which appears robust to a number of alternative specifications, including biases arising from omitted factors and heteroskedasticity, and the use of different numeraire currencies. Among our main results, we find that any measured asymptotic dependence between exchange rate and commodity returns, as well as the revealed forecasting ability of commodity returns for exchange rates, is transitory and short-lived. Indeed, only contemporaneous information at a daily frequency can empirically capture any non-linear relationship. We observe no asymptotic dependence between the variables, nor do we detect any robust forecasting ability using lagged commodity returns or when sampling data at a lower frequency. Asymptotic dependence also appears to be a necessary condition in order to generate significant contemporaneous forecasts. We conjecture that the central transmission mechanism behind the forecasting ability of commodity returns lies in the macroeconomic importance of the information revealed as a result of commodity price shocks which transmits through to exchange rates movements. We believe this is the phenomenon we capture using our asymptotic dependence measure. The arrival of such unexpected commodity relevant information changes investor expectations of the future real economic prospects of commodity exporting countries, and subsequently manifests in movements in their exchange rates.

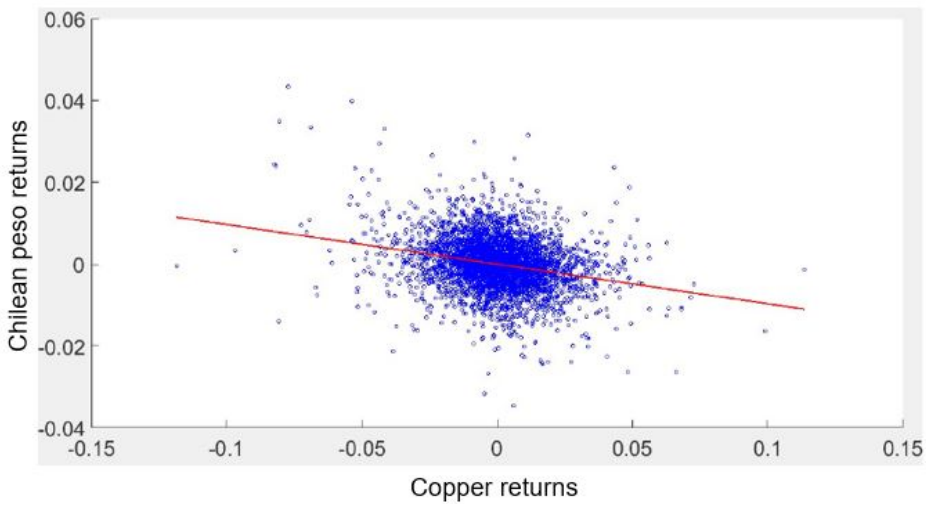
Figures

Figure 3.1. Histogram of Chilean peso and copper daily returns



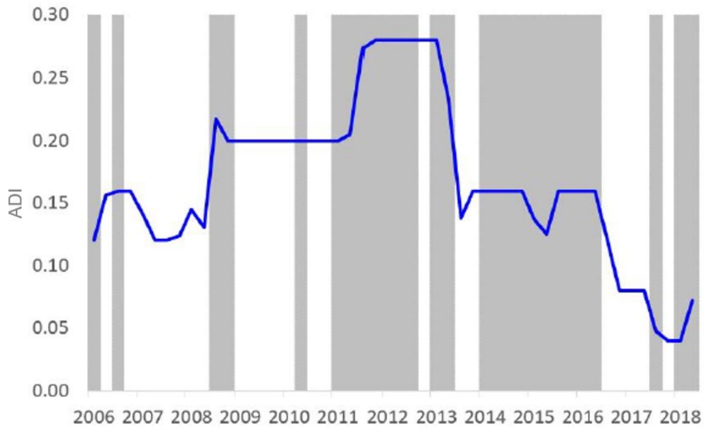
Histogram of Chilean peso returns (left-hand side panel) and copper returns (right-hand side panel) using daily observations from 03/Jan/2000 to 18/Jul/2018. Red solid line represents a theoretical normal distribution. Each figure includes the following statistics describing the distrubution of returns: skewness, kurtosis, the statistic of the Shapiro-Wilk normality test (SW stat) and its p-value (SW p-value). Shapiro-Wilk null hypothesis states that returns are normally distributed.

Figure 3.2. Scatter plot of Chilean peso and copper daily returns



Scatter plot of copper returns (x-axis) and Chilean peso returns (y-axis) using daily observations from 03/Jan/2000 to 18/Jul/2018. Red solid line represents a linear fit estimated using OLS.

Figure 3.3. Asymptotic dependence and the China economic surprise index



The blue solid line represents the asymptotic dependence indicator (ADI) case 2, correcting for the effect of both global risk appetite and heteroskedasticity of log-returns. The shaded areas correspond to periods when the China economic surprise index exhibit negative economic surprises. Asymptotic dependence indicator computed using a rolling window with a width of 1000 daily observations and 2.5% as the tail percentile. Estimation sample corresponds to daily data from 03/Jan/2000 to 18/Jul/2018. The China surprise index corresponds to the Citigroup China Economic Surprise Index obtained from Bloomberg.

Tables

Table 3.1. Commodity exporting economies

	Main export	Commodity exports (% of total exports)			Commodity exports (% of GDP)			Main commodity export (% of commodity exports)		
		2000-05	2006-11	2012-17	2000-05	2006-11	2012-17	2000-05	2006-11	2012-17
Brazil	Oil	46	58	64	5	6	6	9	16	13
Canada	Oil	36	49	50	11	12	12	22	33	37
Chile	Copper	84	87	86	24	31	23	55	70	63
Colombia	Oil	65	71	81	9	11	11	43	46	55
Mexico	Oil	18	26	21	4	7	7	59	59	42
Norway	Oil	79	82	80	25	26	22	64	54	44
Peru	Copper	83	88	89	13	23	18	42	49	46
Russia	Oil	76	83	81	22	20	18	57	64	63
S. Africa	Gold	49	55	57	9	12	14	51	62	59

Column "main export" denotes the main commodity export by country. The rest of the columns report the average percentage value of exports for each of three periods: 2000-2005, 2006-2011, and 2012-2017 for each country. Source: United Nations Conference on Trade and Development (UNCTAD), website (https://unctadstat.unctad.org/EN/Index.html).

Table 3.2. Panel A: Contemporaneous commodity returns vs. RW model without drift

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	0.919***	0.920***	0.920***	0.919***	0.919***	0.919***
stat	-4.867	-4.821	-4.818	-4.844	-4.841	-4.824
Canada	0.812***	0.812***	0.812***	0.813***	0.813***	0.813***
stat	-6.945	-6.928	-6.921	-6.907	-6.900	-6.862
Chile	0.905***	0.905***	0.905***	0.906***	0.907***	0.909***
stat	-3.792	-3.800	-3.808	-3.835	-3.880	-4.036
Colombia	0.856***	0.856***	0.856***	0.855***	0.854***	0.854***
stat	-5.112	-5.116	-5.123	-5.135	-5.173	-5.189
Mexico	0.905***	0.905***	0.905***	0.905***	0.904***	0.904***
stat	-5.595	-5.595	-5.592	-5.588	-5.567	-5.545
Norway	0.880***	0.880***	0.880***	0.880***	0.880***	0.881***
stat	-5.375	-5.353	-5.347	-5.337	-5.329	-5.271
Peru	0.976***	0.976***	0.976***	0.976***	0.976***	0.976***
stat	-4.113	-4.090	-4.090	-4.101	-4.105	-4.059
Russia	0.857***	0.857***	0.857***	0.857***	0.858***	0.858***
stat	-4.879	-4.878	-4.882	-4.885	-4.888	-4.905
S. Africa	0.944***	0.944***	0.944***	0.944***	0.944***	0.943***
stat	-4.015	-4.042	-4.045	-4.057	-4.061	-4.089

The table reports the RMSFE ratio between the commodity-based model in equation 3.1 (numerator) and a RW model (denominator). Base currency: USD. Benchmark model: RW without drift. "stat" corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Perfect foresight information (realised observations of explanatory variable). Sample: Daily data from 03/Jan/2000 to 18/Jul/2018 (4270 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.2. Panel B: Contemporaneous commodity returns vs. RW model with drift

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	0.892***	0.892***	0.892***	0.892***	0.892***	0.892***
stat	-5.644	-5.616	-5.604	-5.624	-5.621	-5.607
Canada	0.783***	0.783***	0.783***	0.783***	0.783***	0.783***
stat	-7.380	-7.375	-7.375	-7.371	-7.372	-7.343
Chile	0.881***	0.881***	0.881***	0.881***	0.882***	0.880***
stat	-4.303	-4.315	-4.304	-4.331	-4.336	-4.377
Colombia	0.824***	0.824***	0.824***	0.823***	0.822***	0.822***
stat	-5.509	-5.514	-5.526	-5.556	-5.609	-5.596
Mexico	0.875***	0.874***	0.874***	0.875***	0.873***	0.874***
stat	-6.614	-6.621	-6.610	-6.608	-6.600	-6.560
Norway	0.849***	0.848***	0.849***	0.849***	0.849***	0.85***
stat	-6.430	-6.413	-6.406	-6.397	-6.391	-6.326
Peru	0.952***	0.952***	0.952***	0.952***	0.952***	0.952***
stat	-4.510	-4.484	-4.495	-4.500	-4.510	-4.490
Russia	0.841***	0.841***	0.841***	0.841***	0.841***	0.842***
stat	-4.872	-4.870	-4.876	-4.877	-4.880	-4.895
S. Africa	0.917***	0.917***	0.916***	0.916***	0.916***	0.916***
stat	-5.585	-5.573	-5.579	-5.603	-5.615	-5.636

The table reports the RMSFE ratio between the commodity-based model in equation 3.1 (numerator) and a RW model (denominator). Base currency: USD. Benchmark model: RW with drift. “stat” corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Perfect foresight information (realised observations of explanatory variable). Sample: Daily data from 03/Jan/2000 to 18/Jul/2018 (4270 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.2. Panel C: Contemporaneous commodity returns vs. RW model without drift, using returns orthogonal to VIX index

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	0.964***	0.965***	0.965***	0.964***	0.964***	0.964***
stat	-3.063	-2.964	-2.961	-3.020	-3.012	-3.017
Canada	0.850***	0.850***	0.851***	0.851***	0.851***	0.852***
stat	-6.386	-6.367	-6.359	-6.342	-6.333	-6.286
Chile	0.926***	0.926***	0.926***	0.926***	0.927***	0.928***
stat	-3.970	-3.967	-3.969	-3.977	-4.007	-4.138
Colombia	0.892***	0.892***	0.892***	0.892***	0.891***	0.892***
stat	-4.840	-4.844	-4.852	-4.859	-4.892	-4.899
Mexico	0.953***	0.953***	0.953***	0.953***	0.953***	0.953***
stat	-4.680	-4.683	-4.681	-4.678	-4.656	-4.631
Norway	0.900***	0.900***	0.900***	0.900***	0.901***	0.901***
stat	-4.848	-4.825	-4.818	-4.808	-4.798	-4.734
Peru	0.981***	0.981***	0.981***	0.981***	0.981***	0.982***
stat	-3.855	-3.833	-3.827	-3.842	-3.843	-3.785
Russia	0.883***	0.883***	0.883***	0.883***	0.883***	0.884***
stat	-4.686	-4.685	-4.691	-4.693	-4.697	-4.719
S. Africa	0.936***	0.935***	0.935***	0.935***	0.935***	0.935***
stat	-4.998	-5.040	-5.052	-5.052	-5.059	-5.079

The table reports the RMSFE ratio between the commodity-based model in equation 3.1 (numerator) and a RW model (denominator). We obtain exchange rate and commodity returns orthogonal to the VIX index by running the following regression per country and commodity: $r_t = \alpha_0 + \alpha_1 dVIX_t + \nu_t$, where r_t corresponds to each time-series of exchange rate and commodity log-return, $dVIX_t$ is the change in the VIX index, and α_0 and α_1 are coefficients to be estimated using the OLS method. We interpret the error term of above regression (ν_t) as the exchange rate and commodity log-returns that are orthogonal to changes in the VIX index. Base currency: USD. Benchmark model: RW without drift. "stat" corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Perfect foresight information (realised observations of explanatory variable). Sample: Daily data from 03/Jan/2000 to 18/Jul/2018 (4270 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.2. Panel D: Contemporaneous commodity returns vs. RW model without drift, using Euro (EUR) as a base currency

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	0.969***	0.968***	0.968***	0.969***	0.969***	0.968***
stat	-3.043	-3.051	-3.055	-3.014	-3.007	-3.053
Canada	0.953***	0.954***	0.954***	0.954***	0.954***	0.954***
stat	-3.408	-3.407	-3.414	-3.400	-3.409	-3.385
Chile	0.995***	0.995***	0.995**	0.995**	0.995**	0.995**
stat	-2.660	-2.578	-2.546	-2.507	-2.506	-2.465
Colombia	0.973***	0.973***	0.973***	0.974***	0.974***	0.975***
stat	-3.245	-3.255	-3.268	-3.259	-3.253	-3.269
Mexico	0.980***	0.980***	0.980***	0.980***	0.980***	0.981***
stat	-2.774	-2.768	-2.773	-2.777	-2.792	-2.800
Norway	0.913***	0.913***	0.913***	0.913***	0.913***	0.913***
stat	-4.784	-4.781	-4.783	-4.793	-4.795	-4.751
Peru	0.995	0.995	0.995	0.995	0.995	0.994
stat	-0.736	-0.731	-0.696	-0.684	-0.658	-0.785
Russia	0.952***	0.953***	0.953***	0.953***	0.954***	0.955***
stat	-3.346	-3.341	-3.342	-3.348	-3.347	-3.357
S. Africa	0.995*	0.995*	0.995*	0.994*	0.994**	0.994**
stat	-1.868	-1.868	-1.912	-1.940	-1.963	-2.040

The table reports the RMSFE ratio between the commodity-based model in equation 3.1 (numerator) and a RW model (denominator). Base currency: Euro. Benchmark model: RW without drift. "stat" corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Perfect foresight information (realised observations of explanatory variable). Sample: Daily data from 03/Jan/2000 to 18/Jul/2018 (4270 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.2. Panel E: Contemporaneous commodity returns vs. RW model without drift, using Pound Sterling (GBP) as a base currency

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	0.983**	0.983**	0.983**	0.983**	0.983**	0.983**
stat	-2.212	-2.201	-2.197	-2.205	-2.196	-2.190
Canada	0.968***	0.968***	0.968***	0.968***	0.969***	0.969***
stat	-2.875	-2.851	-2.850	-2.844	-2.824	-2.772
Chile	0.993**	0.993**	0.993**	0.993**	0.993**	0.993**
stat	-2.513	-2.454	-2.426	-2.421	-2.428	-2.297
Colombia	0.975***	0.975***	0.975***	0.975***	0.975***	0.975***
stat	-3.094	-3.095	-3.100	-3.108	-3.105	-3.108
Mexico	0.988**	0.988**	0.988**	0.988**	0.988**	0.988**
stat	-2.264	-2.257	-2.260	-2.263	-2.252	-2.250
Norway	0.963***	0.963***	0.963***	0.963***	0.963***	0.964***
stat	-3.005	-2.988	-2.991	-3.007	-2.997	-2.934
Peru	1.002	1.002	1.002	1.002	1.003	1.002
stat	0.383	0.311	0.296	0.345	0.457	0.369
Russia	0.960***	0.961***	0.961***	0.961***	0.962***	0.962***
stat	-3.051	-3.038	-3.035	-3.033	-3.021	-3.025
S. Africa	0.985***	0.985***	0.985***	0.985***	0.985***	0.985***
stat	-3.580	-3.617	-3.642	-3.663	-3.681	-3.765

The table reports the RMSFE ratio between the commodity-based model in equation 3.1 (numerator) and a RW model (denominator). Base currency: Pound sterling. Benchmark model: RW without drift. "stat" corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Perfect foresight information (realised observations of explanatory variable). Sample: Daily data from 03/Jan/2000 to 18/Jul/2018 (4270 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.3. Panel A: Lagged commodity returns vs. RW model without drift

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	1.002	1.003*	1.002	1.003**	1.002*	1.002
stat	1.140	1.732	1.414	2.176	1.671	1.464
Canada	1.001	1.001	1.003***	1.001	1.001*	1.002*
stat	0.329	0.578	2.703	1.322	1.683	1.804
Chile	1.002	1.002*	1.001	0.999	1.002	1.001
stat	1.389	1.769	0.276	-0.143	0.680	0.440
Colombia	0.999	1.003***	1.001	1.002***	1.002***	1.002***
stat	-0.907	3.019	1.454	2.667	2.595	2.790
Mexico	1.003	1.003**	1.006***	1.001	1.001	1.001
stat	1.122	2.170	3.216	1.087	1.020	1.089
Norway	0.998	1.001	1.000	1.001	1.000	1.002*
stat	-0.384	1.278	0.537	0.782	0.678	1.895
Peru	1.005***	1.001	1.001	1.002	1.000	1.001
stat	3.035	1.111	0.867	1.225	0.149	1.170
Russia	0.992*	1.000	0.998**	1.002	1.001	1.001
stat	-1.682	0.311	-2.468	1.227	0.771	0.668
S. Africa	1.001	1.001	1.001	1.001	1.001	1.001
stat	0.664	0.734	0.614	0.606	0.573	0.478

The table reports the RMSFE ratio between the commodity-based model in equation 3.9 (numerator) and a RW model (denominator). Base currency: USD. Benchmark model: RW without drift. “stat” corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Lagged observations of explanatory variable. Sample: Daily data from 03/Jan/2000 to 18/Jul/2018 (4270 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.3. Panel B: Lagged commodity returns vs. RW model with drift

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	0.972***	0.973***	0.972***	0.973***	0.972***	0.972***
stat	-3.923	-3.628	-3.764	-3.864	-3.880	-3.793
Canada	0.964***	0.964***	0.966***	0.964***	0.964***	0.964***
stat	-4.177	-4.206	-4.326	-4.376	-4.426	-4.242
Chile	0.975***	0.975***	0.974***	0.972***	0.974***	0.968***
stat	-3.228	-2.834	-3.215	-3.207	-4.228	-3.374
Colombia	0.962***	0.966***	0.964***	0.964***	0.964***	0.964***
stat	-3.858	-3.463	-3.787	-3.633	-3.634	-3.609
Mexico	0.969***	0.97***	0.973***	0.967***	0.967***	0.968***
stat	-3.866	-3.776	-3.655	-4.199	-4.278	-4.109
Norway	0.963***	0.965***	0.965***	0.965***	0.964***	0.967***
stat	-4.588	-5.268	-5.278	-5.249	-5.240	-5.070
Peru	0.98**	0.977***	0.976***	0.978***	0.975***	0.976***
stat	-2.519	-2.751	-2.818	-2.723	-2.954	-2.826
Russia	0.973***	0.981***	0.979***	0.982***	0.982***	0.982***
stat	-2.864	-3.414	-3.267	-3.004	-3.004	-2.773
S. Africa	0.972***	0.972***	0.972***	0.971***	0.971***	0.973***
stat	-4.927	-4.611	-4.999	-4.995	-5.046	-4.543

The table reports the RMSFE ratio between the commodity-based model in equation 3.9 (numerator) and a RW model (denominator). Base currency: USD. Benchmark model: RW with drift. "stat" corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Lagged observations of explanatory variable. Sample: Daily data from 03/Jan/2000 to 18/Jul/2018 (4270 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.3. Panel C: Lagged commodity returns vs. RW model without drift, using returns orthogonal to VIX index

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	1.004**	1.003**	1.002	1.003**	1.002	1.002*
stat	2.363	2.003	1.305	2.067	1.436	1.785
Canada	1.001	1.001	1.002**	1.001	1.001	1.002*
stat	0.485	0.592	2.267	1.054	1.136	1.825
Chile	1.001	1.002*	1.001	1.000	1.002	1.000
stat	0.418	1.802	1.408	-0.103	0.592	-0.149
Colombia	0.997	1.003***	1.001	1.002**	1.002**	1.002**
stat	-1.329	2.710	1.033	2.254	2.268	2.480
Mexico	1.001	1.002**	1.005***	1.001	1.001	1.001
stat	1.644	2.041	2.752	1.400	1.217	0.830
Norway	0.999	1.001	1.001	1.001	1.001	1.003*
stat	-0.287	1.049	0.663	0.728	1.065	1.896
Peru	1.005***	1.001	1.001	1.002	1.000	1.001
stat	2.684	1.082	0.797	1.264	0.089	1.278
Russia	0.990*	1.001	0.998	1.002*	1.001	1.002
stat	-1.871	0.764	-1.589	1.936	1.534	1.198
S. Africa	1.001	1.001	1.002	1.001	1.001	1.001
stat	1.263	0.661	1.079	1.004	1.032	0.701

The table reports the RMSFE ratio between the commodity-based model in equation 3.9 (numerator) and a RW model (denominator). We obtain exchange rate and commodity returns orthogonal to the VIX index by running the following regression per country and commodity: $r_t = \alpha_0 + \alpha_1 dVIX_t + \nu_t$, where r_t corresponds to each time-series of exchange rate and commodity log-return, $dVIX_t$ is the change in the VIX index, and α_0 and α_1 are coefficients to be estimated using the OLS method. We interpret the error term of the regression above (ν_t) as the exchange rate and commodity log-returns that are orthogonal to changes in the VIX index. Base currency: USD. Benchmark model: RW without drift. "stat" corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Lagged observations of explanatory variable. Sample: Daily data from 03/Jan/2000 to 18/Jul/2018 (4270 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.4. Panel A: Monthly frequency, contemporaneous commodity returns vs. RW model without drift

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	0.873	0.881	0.878	0.872	0.873	0.880
stat	-1.494	-1.326	-1.361	-1.405	-1.402	-1.302
Canada	0.734**	0.755**	0.732**	0.706**	0.706**	0.699**
stat	-2.563	-2.089	-2.200	-2.257	-2.269	-2.299
Chile	0.690**	0.689**	0.688**	0.679**	0.667**	0.659**
stat	-2.228	-2.194	-2.161	-2.256	-2.382	-2.304
Colombia	0.802*	0.802*	0.799*	0.799*	0.807*	0.824*
stat	-1.749	-1.701	-1.737	-1.806	-1.826	-1.897
Mexico	0.912	0.916	0.910	0.910	0.907	0.916
stat	-1.000	-0.938	-0.984	-0.988	-1.021	-0.955
Norway	0.726**	0.734**	0.735**	0.729**	0.727**	0.731**
stat	-2.082	-1.976	-1.983	-2.011	-2.042	-1.990
Peru	0.931	0.932	0.926	0.929	0.937	0.927
stat	-1.356	-1.332	-1.458	-1.423	-1.215	-1.526
Russia	0.653	0.667	0.671	0.676	0.678	0.696*
stat	-1.621	-1.592	-1.605	-1.614	-1.634	-1.647
S. Africa	0.890*	0.899	0.897	0.895	0.899	0.894
stat	-1.772	-1.581	-1.542	-1.562	-1.474	-1.370

The table reports the RMSFE ratio between the commodity-based model in equation 3.1 (numerator) and a RW model (denominator). Base currency: USD. Benchmark model: RW without drift. “stat” corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Perfect foresight information (realised observations of explanatory variable). Sample: Monthly data from Jan/2000 to Jun/2018 (222 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.4. Panel B: Monthly frequency, contemporaneous commodity returns vs. RW model with drift

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	0.844**	0.851*	0.848*	0.840*	0.842*	0.847*
stat	-2.052	-1.856	-1.876	-1.952	-1.930	-1.856
Canada	0.700***	0.720**	0.699***	0.671***	0.672***	0.663***
stat	-3.065	-2.565	-2.658	-2.728	-2.736	-2.793
Chile	0.650**	0.650**	0.650**	0.641**	0.630**	0.621**
stat	-2.361	-2.311	-2.265	-2.348	-2.460	-2.367
Colombia	0.787**	0.786**	0.784**	0.782**	0.791**	0.806**
stat	-2.088	-2.025	-2.050	-2.124	-2.136	-2.191
Mexico	0.855	0.863	0.858	0.859	0.855	0.866
stat	-1.617	-1.493	-1.532	-1.532	-1.557	-1.497
Norway	0.693***	0.701**	0.701**	0.695***	0.693***	0.697***
stat	-2.673	-2.546	-2.569	-2.604	-2.650	-2.619
Peru	0.926	0.928	0.924	0.926	0.930	0.915
stat	-1.056	-1.027	-1.056	-1.033	-0.938	-1.123
Russia	0.621**	0.636**	0.640**	0.644**	0.646**	0.665**
stat	-2.054	-2.009	-2.031	-2.052	-2.084	-2.116
S. Africa	0.846*	0.857*	0.856	0.854	0.858	0.853
stat	-1.825	-1.648	-1.617	-1.637	-1.574	-1.465

The table reports the RMSFE ratio between the commodity-based model in equation 3.1 (numerator) and a RW model (denominator). Base currency: USD. Benchmark model: RW with drift. "stat" corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Perfect foresight information (realised observations of explanatory variable). Sample: Monthly data from Jan/2000 to Jun/2018 (222 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.4. Panel C: Monthly frequency, lagged commodity returns vs. RW model without drift

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	1.016	0.962	1.014	1.037	1.041	1.024
stat	0.926	-0.523	0.322	1.452	1.032	0.931
Canada	1.044*	1.029	1.047	1.024**	1.059	1.025*
stat	1.773	1.242	0.903	2.455	1.306	1.820
Chile	1.031	1.023	1.018	1.021	1.017*	1.026
stat	1.185	1.523	1.036	0.856	1.647	1.029
Colombia	1.019*	1.018	1.046	1.045*	1.053	1.019
stat	1.887	0.438	0.559	1.840	1.169	1.058
Mexico	1.035	1.025	1.016	1.002	1.037	1.004
stat	0.824	1.316	0.355	0.144	0.977	0.252
Norway	1.025**	1.025	1.002	1.019	1.051*	1.019*
stat	2.082	0.933	0.113	1.535	1.897	1.723
Peru	1.033	1.031	1.005	1.034	1.030	1.030
stat	0.701	1.241	0.171	0.941	1.096	0.866
Russia	0.936	1.050**	1.042	1.015	1.008	1.007
stat	-0.744	2.484	0.794	0.824	0.340	0.278
S. Africa	1.019	1.013	1.015	1.011	1.012	0.968
stat	0.521	0.562	0.524	0.432	0.396	-0.675

The table reports the RMSFE ratio between the commodity-based model in equation 3.9 (numerator) and a RW model (denominator). Base currency: USD. Benchmark model: RW without drift. "stat" corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Lagged observations of explanatory variable. Sample: Monthly data from Jan/2000 to Jun/2018 (222 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.4. Panel D: Monthly frequency, lagged commodity returns vs. RW model with drift

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	0.982	0.929	0.977	1.000	1.001	0.986
stat	-0.456	-1.619	-0.617	-0.009	0.008	-0.326
Canada	0.996	0.981	0.994	0.973	1.006	0.972
stat	-0.211	-0.662	-0.106	-0.726	0.103	-0.716
Chile	0.973	0.966	0.961	0.963	0.96	0.967
stat	-1.283	-1.065	-1.166	-0.745	-1.091	-1.092
Colombia	1.000	0.999	1.024	1.024	1.029	0.998
stat	-0.004	-0.031	0.254	0.383	0.345	-0.038
Mexico	0.975*	0.966	0.958	0.945	0.976	0.951
stat	-1.759	-0.831	-0.523	-1.375	-0.335	-1.271
Norway	0.978	0.977	0.956	0.972	1.001	0.971
stat	-0.883	-1.159	-1.011	-0.843	0.020	-0.837
Peru	1.028	1.028	1.001	1.026	1.016	1.016
stat	0.336	0.312	0.009	0.272	0.196	0.174
Russia	0.892	1.001	0.993	0.967	0.963	0.962
stat	-1.584	0.016	-0.098	-0.755	-1.068	-0.998
S. Africa	0.972	0.966	0.968	0.964	0.966	0.924
stat	-0.472	-1.168	-0.983	-0.975	-0.935	-1.376

The table reports the RMSFE ratio between the commodity-based model in equation 3.9 (numerator) and a RW model (denominator). Base currency: USD. Benchmark model: RW with drift. “stat” corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Lagged observations of explanatory variable. Sample: Monthly data from Jan/2000 to Jun/2018 (222 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.4. Panel E: Monthly frequency, contemporaneous commodity returns vs. RW model without drift, using returns orthogonal to VIX index

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	0.943	0.947	0.947	0.945	0.947	0.948
stat	-0.941	-0.829	-0.832	-0.861	-0.864	-0.847
Canada	0.784**	0.804*	0.774*	0.747*	0.747**	0.739**
stat	-2.180	-1.757	-1.907	-1.957	-1.968	-2.031
Chile	0.840*	0.838*	0.841*	0.835*	0.825*	0.824*
stat	-1.725	-1.715	-1.667	-1.760	-1.901	-1.813
Colombia	0.898	0.899	0.899	0.904	0.917	0.945
stat	-1.550	-1.507	-1.517	-1.541	-1.493	-1.352
Mexico	0.949	0.951	0.947	0.946	0.942	0.945
stat	-0.792	-0.759	-0.790	-0.826	-0.878	-0.841
Norway	0.756*	0.761*	0.760*	0.753*	0.751*	0.753*
stat	-1.890	-1.837	-1.851	-1.874	-1.911	-1.884
Peru	0.969	0.968	0.962	0.966	0.974	0.964
stat	-0.758	-0.766	-0.904	-0.845	-0.629	-0.932
Russia	0.692	0.710	0.715	0.720	0.723	0.743
stat	-1.575	-1.541	-1.557	-1.572	-1.595	-1.617
S. Africa	0.885*	0.891*	0.894	0.893	0.897	0.898
stat	-1.777	-1.680	-1.537	-1.584	-1.456	-1.354

The table reports the RMSFE ratio between the commodity-based model in equation 3.1 (numerator) and a RW model (denominator). We obtain exchange rate and commodity returns orthogonal to the VIX index by running the following regression per country and commodity: $r_t = \alpha_0 + \alpha_1 dVIX_t + \nu_t$, where r_t corresponds to each time-series of exchange rate and commodity log-return, $dVIX_t$ is the change in the VIX index, and α_0 and α_1 are coefficients to be estimated using the OLS method. We interpret the error term of the regression above (ν_t) as the exchange rate and commodity log-returns that are orthogonal to changes in the VIX index. Base currency: USD. Benchmark model: RW without drift. "stat" corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Perfect foresight information (realised observations of explanatory variable). Sample: Monthly data from Jan/2000 to Jun/2018 (222 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.4. Panel F: Monthly frequency, lagged commodity returns vs. RW model without drift, using returns orthogonal to VIX index

	h=1	h=2	h=3	h=4	h=5	h=10
Brazil	1.055*	0.975	1.041	1.045	1.057	1.061
stat	1.700	-0.356	1.065	1.444	1.599	1.462
Canada	1.068	1.020	1.062	1.034	1.043**	1.040*
stat	1.614	0.611	0.943	1.304	1.977	1.660
Chile	1.020	1.027**	1.031	1.020	1.022	1.025
stat	1.245	2.362	1.387	1.387	1.354	0.992
Colombia	1.038**	1.022	1.064	1.036	1.048	1.032
stat	2.329	0.562	0.701	1.179	1.413	1.312
Mexico	1.078**	1.033	1.036	0.998	1.032	1.030
stat	2.143	1.304	0.566	0.000	1.439	1.381
Norway	1.042**	1.022	1.005	1.013	1.040**	1.031
stat	2.112	0.739	0.179	0.871	2.194	1.575
Peru	1.013	1.031	0.999	1.031	1.028	1.028
stat	0.428	1.373	-0.036	0.999	1.175	0.929
Russia	0.948	1.063**	1.065	1.035	1.022	1.029
stat	-0.701	2.172	0.943	1.184	1.487	1.509
S. Africa	1.042	1.026	1.018	1.023	1.046	0.997
stat	1.237	1.391	0.630	1.377	1.033	-0.068

The table reports the RMSFE ratio between the commodity-based model in equation 3.9 (numerator) and a RW model (denominator). We obtain exchange rate and commodity returns orthogonal to the VIX index by running the following regression per country and commodity: $r_t = \alpha_0 + \alpha_1 dVIX_t + \nu_t$, where r_t corresponds to each time-series of exchange rate and commodity log-return, $dVIX_t$ is the change in the VIX index, and α_0 and α_1 are coefficients to be estimated using the OLS method. We interpret the error term of the regression above (ν_t) as the exchange rate and commodity log-returns that are orthogonal to changes in the VIX index. Base currency: USD. Benchmark model: RW without drift. "stat" corresponds to the statistic of the Diebold-Mariano test. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$. Columns correspond to the selected forecast horizons. Lagged observations of explanatory variable. Sample: Monthly data from Jan/2000 to Jun/2018 (222 observations). RMSFE ratios displaying a value lower than one indicate the commodity-based model outperforms the benchmark.

Table 3.5. Panel A: Hill index estimator

	Lower tail	Upper tail
Brazil	0.30 (0.25;0.35)	0.33 (0.27;0.38)
Canada	0.27 (0.23;0.31)	0.26 (0.22;0.30)
Chile	0.25 (0.21;0.29)	0.29 (0.25;0.34)
Colombia	0.32 (0.27;0.37)	0.34 (0.28;0.39)
Mexico	0.32 (0.27;0.37)	0.38 (0.32;0.44)
Peru	0.33 (0.28;0.38)	0.40 (0.34;0.46)
Norway	0.27 (0.23;0.31)	0.25 (0.21;0.29)
Russia	0.38 (0.32;0.44)	0.40 (0.33;0.46)
S. Africa	0.25 (0.21;0.29)	0.30 (0.25;0.34)
Copper	0.32 (0.27;0.37)	0.25 (0.21;0.29)
Gold	0.31 (0.26;0.35)	0.32 (0.27;0.37)
WTI	0.30 (0.25;0.35)	0.31 (0.26;0.35)

The table displays the estimation of the Hill tail index of equation 3.2 for exchange rate returns (top panel) and commodity returns (bottom panel). Confidence intervals at 99% level in parentheses. Daily log-returns, 03/Jan/2000 to 18/Jul/2018 (4270 observations). Threshold corresponds to 2.5% of the data (107 observations). A positive value indicates the distribution of returns follows a fat-tailed distribution.

Table 3.5. Panel B: Tail index estimator of Dekkers et al. (1989)

	Lower tail	Upper tail
Brazil	0.18 (0.02;0.34)	0.23 (0.07;0.39)
Canada	0.22 (0.06;0.38)	0.05 (-0.11;0.20)
Chile	0.11 (-0.05;0.27)	0.24 (0.08;0.41)
Colombia	0.20 (0.04;0.36)	0.11 (-0.05;0.27)
Mexico	0.15 (-0.01;0.30)	0.34 (0.18;0.51)
Peru	0.20 (0.04;0.36)	0.30 (0.13;0.46)
Norway	0.30 (0.14;0.46)	0.07 (-0.09;0.22)
Russia	0.21 (0.05;0.37)	0.29 (0.13;0.45)
S. Africa	0.04 (-0.11;0.19)	0.29 (0.13;0.45)
Copper	0.06 (-0.10;0.22)	0.16 (0.00;0.33)
Gold	0.14 (-0.02;0.29)	0.23 (0.08;0.39)
WTI	0.16 (0.00;0.32)	0.16 (0.00;0.32)

The table displays the estimation of the Dekkers et al. (1989) tail index of equation 3.4 for exchange rate returns (top panel) and commodity returns (bottom panel). Confidence intervals at 99% level in parentheses. Daily log-returns, 03/Jan/2000 to 18/Jul/2018 (4270 observations). Threshold corresponds to 2.5% of the data (107 observations). A positive value indicates the distribution of returns follows a fat-tailed distribution.

Table 3.6. Panel A: ADI using contemporaneous commodity returns

	Case 1	Case 2	Case 3	Case 4
Brazil	0.111 (0.031;0.191)	0.148 (0.063;0.233)	0.046 (-0.006;0.100)	0.028 (-0.014;0.070)
Canada	0.223 (0.126;0.320)	0.161 (0.070;0.251)	0.018 (-0.014;0.051)	0.000 (-0.015;0.016)
Chile	0.104 (0.028;0.179)	0.189 (0.093;0.284)	0.028 (-0.013;0.070)	0.028 (-0.013;0.070)
Colombia	0.196 (0.098;0.294)	0.215 (0.113;0.316)	0.009 (-0.018;0.037)	0.009 (-0.019;0.038)
Mexico	0.152 (0.067;0.236)	0.170 (0.082;0.257)	0.009 (-0.014;0.033)	0.009 (-0.017;0.036)
Norway	0.179 (0.087;0.269)	0.188 (0.091;0.284)	0.027 (-0.012;0.066)	0.009 (-0.013;0.031)
Peru	0.055 (-0.003;0.113)	0.101 (0.026;0.176)	0.055 (0.000;0.109)	0.046 (-0.006;0.098)
Russia	0.214 (0.111;0.317)	0.170 (0.084;0.255)	0.009 (-0.018;0.036)	0.009 (-0.013;0.032)
S. Africa	0.092 (0.023;0.160)	0.133 (0.058;0.208)	0.058 (0.002;0.114)	0.058 (0.001;0.115)

The table reports the estimation results of the asymptotic dependence indicator (ADI) of equation 3.7 using daily log-returns from 03/Jan/2000 to 18/Jul/2018 (4270 observations). Confidence intervals in parentheses at 99% level and obtained by bootstrap using 5000 resampling iterations. Threshold corresponds to 2.5% of the data (107 observations). **Case 1**: NER appreciation (∇S) and increase in comm. price (ΔP_{comm}). **Case 2**: NER depreciation (ΔS) and reduction in comm. price. (∇P_{comm}) **Case 3**: ∇S and ∇P_{comm}. **Case 4**: ΔS and ΔP_{comm}. The ADI value indicates the likelihood of observing extreme values simultaneously in both exchange rate and commodity returns.

Table 3.6. Panel B: ADI using contemporaneous commodity returns and exchange rates with EUR as a base currency

	Case 1	Case 2	Case 3	Case 4
Brazil	0.110 (0.031;0.189)	0.101 (0.025;0.176)	0.055 (-0.003;0.113)	0.037 (-0.009;0.083)
Canada	0.161 (0.073;0.248)	0.098 (0.024;0.172)	0.009 (-0.016;0.034)	0.009 (-0.017;0.035)
Chile	0.053 (0.000;0.106)	0.070 (0.008;0.132)	0.061 (0.004;0.118)	0.061 (0.001;0.121)
Colombia	0.108 (0.032;0.183)	0.153 (0.066;0.240)	0.036 (-0.011;0.084)	0.027 (-0.012;0.066)
Mexico	0.107 (0.033;0.180)	0.116 (0.034;0.197)	0.036 (-0.014;0.086)	0.063 (0.004;0.121)
Norway	0.188 (0.095;0.279)	0.196 (0.104;0.288)	0.027 (-0.013;0.067)	0.009 (-0.014;0.032)
Peru	0.053 (0.001;0.104)	0.018 (-0.015;0.051)	0.071 (0.007;0.134)	0.080 (0.014;0.144)
Russia	0.179 (0.081;0.275)	0.152 (0.060;0.243)	0.018 (-0.014;0.050)	0.036 (-0.008;0.080)
S. Africa	0.050 (-0.001;0.102)	0.100 (0.033;0.167)	0.058 (0.001;0.115)	0.067 (0.006;0.127)

The table reports the estimation results of the asymptotic dependence indicator (ADI) of equation 3.7 using daily log-returns from 03/Jan/2000 to 18/Jul/2018 (4270 observations). Confidence intervals in parentheses at 99% level and obtained by bootstrap using 5000 resampling iterations. Threshold corresponds to 2.5% of the data (107 observations). **Case 1**: NER appreciation (∇S) and increase in comm. price (ΔP_{comm}). **Case 2**: NER depreciation (ΔS) and reduction in comm. price. (∇P_{comm}) **Case 3**: ∇S and ∇P_{comm}. **Case 4**: ΔS and ΔP_{comm}. Euro as a base currency. The ADI value indicates the likelihood of observing extreme values simultaneously in both exchange rate and commodity returns.

Table 3.6. Panel C: ADI using contemporaneous commodity returns and exchange rates with GBP as a base currency

	Case 1	Case 2	Case 3	Case 4
Brazil	0.099 (0.027;0.171)	0.117 (0.044;0.190)	0.054 (-0.001;0.110)	0.027 (-0.015;0.070)
Canada	0.125 (0.047;0.203)	0.125 (0.046;0.203)	0.063 (0.003;0.122)	0.009 (-0.017;0.036)
Chile	0.079 (0.014;0.143)	0.096 (0.023;0.170)	0.079 (0.015;0.142)	0.044 (-0.007;0.095)
Colombia	0.090 (0.019;0.160)	0.153 (0.063;0.242)	0.063 (0.003;0.123)	0.009 (-0.016;0.034)
Mexico	0.063 (-0.001;0.128)	0.117 (0.037;0.196)	0.072 (0.009;0.134)	0.027 (-0.012;0.067)
Norway	0.071 (0.010;0.132)	0.125 (0.046;0.203)	0.054 (-0.003;0.110)	0.027 (-0.013;0.067)
Peru	0.027 (-0.015;0.068)	0.035 (-0.009;0.080)	0.124 (0.047;0.201)	0.062 (0.001;0.122)
Russia	0.162 (0.071;0.252)	0.143 (0.059;0.226)	0.076 (0.007;0.145)	0.029 (-0.016;0.073)
S. Africa	0.050 (-0.006;0.107)	0.075 (0.014;0.136)	0.058 (0.003;0.114)	0.050 (-0.002;0.103)

The table reports the estimation results of the asymptotic dependence indicator (ADI) of equation 3.7 using daily log-returns from 03/Jan/2000 to 18/Jul/2018 (4270 observations). Confidence intervals in parentheses at 99% level and obtained by bootstrap using 5000 resampling iterations. Threshold corresponds to 2.5% of the data (107 observations). **Case 1**: NER appreciation (∇S) and increase in comm. price (ΔP_{comm}). **Case 2**: NER depreciation (ΔS) and reduction in comm. price. (∇P_{comm}) **Case 3**: ∇S and ∇P_{comm}. **Case 4**: ΔS and ΔP_{comm}. Pound Sterling as a base currency. The ADI value indicates the likelihood of observing extreme values simultaneously in both exchange rate and commodity returns.

Table 3.6. Panel D: ADI using contemporaneous commodity returns and exchange rate returns orthogonal to VIX, correcting for heteroskedasticity

	Case 1	Case 2	Case 3	Case 4
Brazil	0.083 (0.015;0.151)	0.065 (0.004;0.125)	0.019 (-0.015;0.053)	0.037 (-0.007;0.082)
Canada	0.188 (0.093;0.281)	0.080 (0.008;0.152)	0.018 (-0.014;0.051)	0.018 (-0.012;0.049)
Chile	0.104 (0.025;0.182)	0.132 (0.046;0.217)	0.019 (-0.013;0.051)	0.009 (-0.014;0.034)
Colombia	0.093 (0.022;0.164)	0.150 (0.065;0.234)	0.000 (-0.018;0.019)	0.028 (-0.013;0.070)
Mexico	0.107 (0.033;0.180)	0.071 (0.010;0.132)	0.027 (-0.010;0.064)	0.027 (-0.011;0.065)
Norway	0.161 (0.072;0.249)	0.098 (0.025;0.171)	0.027 (-0.015;0.069)	0.009 (-0.012;0.030)
Peru	0.073 (0.009;0.138)	0.073 (0.008;0.138)	0.037 (-0.006;0.080)	0.009 (-0.014;0.033)
Russia	0.143 (0.057;0.228)	0.107 (0.033;0.181)	0.009 (-0.014;0.032)	0.018 (-0.014;0.051)
S. Africa	0.092 (0.024;0.158)	0.083 (0.018;0.148)	0.017 (-0.016;0.050)	0.017 (-0.013;0.047)

The table reports the estimation results of the asymptotic dependence indicator (ADI) of equation 3.7 using daily log-returns from 03/Jan/2000 to 18/Jul/2018 (4270 observations). Confidence intervals in parentheses at 99% level and obtained by bootstrap using 5000 resampling iterations. Threshold corresponds to 2.5% of the data (107 observations). **Case 1**: NER appreciation (∇S) and increase in comm. price (ΔP_{comm}). **Case 2**: NER depreciation (ΔS) and reduction in comm. price. (∇P_{comm}) **Case 3**: ∇S and ∇P_{comm}. **Case 4**: ΔS and ΔP_{comm}. We obtain exchange rate and commodity returns orthogonal to the VIX index by running the following regression per country and commodity: $r_t = \alpha_0 + \alpha_1 dVIX_t + \nu_t$, where r_t corresponds to each time-series of exchange rate and commodity log-return, $dVIX_t$ is the change in the VIX index, and α_0 and α_1 are coefficients to be estimated using the OLS method. We interpret the error term of the regression above (ν_t) as the exchange rate and commodity log-returns that are orthogonal to changes in the VIX index. Exchange rates and commodity returns corresponds to the standardised residual obtained from a *ARCH(1)* where the conditional variance is modelled as $\sigma_t^2 = \omega + \alpha u_{t-1}^2$. u_t corresponds to the residuals of the mean equation for returns, and the standardised residuals are computed as $\varepsilon_t = u_t/\sigma_t$. The ADI value indicates the likelihood of observing extreme values simultaneously in both exchange rate and commodity returns.

Table 3.7. Panel A: ADI using lagged commodity returns

	Case 1	Case 2	Case 3	Case 4
Brazil	0.056 (-0.003;0.114)	0.074 (0.008;0.140)	0.139 (0.050;0.227)	0.019 (-0.016;0.053)
Canada	0.063 (0.001;0.123)	0.071 (0.011;0.132)	0.054 (-0.002;0.110)	0.027 (-0.013;0.067)
Chile	0.028 (-0.013;0.071)	0.104 (0.032;0.176)	0.057 (0.000;0.114)	0.057 (-0.002;0.116)
Colombia	0.037 (-0.013;0.089)	0.065 (0.003;0.128)	0.056 (-0.001;0.114)	0.056 (-0.001;0.114)
Mexico	0.071 (0.008;0.135)	0.045 (-0.008;0.098)	0.107 (0.031;0.183)	0.063 (0.002;0.123)
Norway	0.054 (-0.002;0.110)	0.098 (0.026;0.170)	0.045 (-0.009;0.099)	0.027 (-0.011;0.065)
Peru	0.056 (-0.003;0.115)	0.065 (0.004;0.126)	0.037 (-0.014;0.088)	0.046 (-0.005;0.098)
Russia	0.089 (0.022;0.156)	0.054 (-0.002;0.110)	0.045 (-0.006;0.095)	0.036 (-0.009;0.081)
S. Africa	0.050 (-0.002;0.103)	0.025 (-0.014;0.065)	0.067 (0.005;0.128)	0.050 (0.000;0.101)

The table reports the estimation results of the asymptotic dependence indicator (ADI) of equation 3.7 using daily log-returns from 03/Jan/2000 to 18/Jul/2018 (4270 observations). Confidence intervals in parentheses at 99% level and obtained by bootstrap using 5000 resampling iterations. Threshold corresponds to 2.5% of the data (107 observations). **Case 1**: NER appreciation (∇S) and increase in comm. price (ΔP_{comm}). **Case 2**: NER depreciation (ΔS) and reduction in comm. price. (∇P_{comm}) **Case 3**: ∇S and ∇P_{comm}. **Case 4**: ΔS and ΔP_{comm}. The ADI value indicates the likelihood of observing extreme values simultaneously in both exchange rate and commodity returns.

Table 3.7. Panel B: ADI using lagged commodity returns and exchange rates orthogonal to VIX, correcting for heteroscedasticity

	Case 1	Case 2	Case 3	Case 4
Brazil	0.037 (-0.009;0.084)	0.046 (-0.006;0.099)	0.056 (-0.002;0.113)	0.028 (-0.010;0.066)
Canada	0.054 (0.001;0.106)	0.054 (-0.002;0.109)	0.009 (-0.014;0.032)	0.018 (-0.012;0.048)
Chile	0.028 (-0.016;0.073)	0.075 (0.008;0.142)	0.019 (-0.015;0.053)	0.019 (-0.015;0.053)
Colombia	0.065 (0.003;0.128)	0.047 (-0.004;0.098)	0.019 (-0.020;0.058)	0.028 (-0.013;0.069)
Mexico	0.027 (-0.015;0.069)	0.045 (-0.007;0.096)	0.045 (-0.007;0.097)	0.045 (-0.004;0.094)
Norway	0.036 (-0.014;0.086)	0.098 (0.027;0.169)	0.036 (-0.011;0.083)	0.018 (-0.013;0.049)
Peru	0.056 (-0.001;0.113)	0.019 (-0.016;0.054)	0.037 (-0.014;0.089)	0.037 (-0.010;0.085)
Russia	0.054 (-0.001;0.108)	0.063 (0.004;0.121)	0.018 (-0.017;0.054)	0.009 (-0.012;0.030)
S. Africa	0.050 (0.000;0.101)	0.033 (-0.007;0.074)	0.050 (0.000;0.100)	0.050 (0.000;0.101)

The table reports the estimation results of the asymptotic dependence indicator (ADI) of equation 3.7 using daily log-returns from 03/Jan/2000 to 18/Jul/2018 (4270 observations). Confidence intervals in parentheses at 99% level and obtained by bootstrap using 5000 resampling iterations. Threshold corresponds to 2.5% of the data (107 observations). **Case 1**: NER appreciation (∇S) and increase in comm. price (ΔP_{comm}). **Case 2**: NER depreciation (ΔS) and reduction in comm. price. (∇P_{comm}) **Case 3**: ∇S and ∇P_{comm}. **Case 4**: ΔS and ΔP_{comm}. We obtain exchange rate and commodity returns orthogonal to the VIX index by running the following regression per country and commodity: $r_t = \alpha_0 + \alpha_1 dVIX_t + \nu_t$, where r_t corresponds to each time-series of exchange rate and commodity log-return, $dVIX_t$ is the change in the VIX index, and α_0 and α_1 are coefficients to be estimated using the OLS method. We interpret the error term of the regression above (ν_t) as the exchange rate and commodity log-returns that are orthogonal to changes in the VIX index. Exchange rates and commodity returns corresponds to the standardised residual obtained from a *ARCH(1)* where the conditional variance is modelled as $\sigma_t^2 = \omega + \alpha u_{t-1}^2$. u_t corresponds to the residuals of the mean equation for returns, and the standardised residuals are computed as $\varepsilon_t = u_t/\sigma_t$. The ADI value indicates the likelihood of observing extreme values simultaneously in both exchange rate and commodity returns.

Table 3.8. Panel A: ADI using contemporaneous commodity returns at different frequencies

	Daily				Monthly				Quarterly			
	C1	C2	C3	C4	C1	C2	C3	C4	C1	C2	C3	C4
Brazil	0.08	0.06	–	–	–	–	–	–	–	–	–	–
Canada	0.19	0.08	–	–	–	–	–	–	–	–	–	–
Chile	0.10	0.13	–	–	–	0.60	–	–	–	–	–	–
Colombia	0.09	0.15	–	–	–	–	–	–	–	–	–	–
Mexico	0.11	0.07	–	–	–	–	–	–	–	–	–	–
Norway	0.16	0.10	–	–	–	–	–	–	–	–	–	–
Peru	0.07	0.07	–	–	–	–	–	–	–	–	–	–
Russia	0.14	0.11	–	–	–	–	–	–	–	–	–	–
S. Africa	0.09	0.08	–	–	–	–	–	–	–	–	–	–

The table reports the estimation results of the asymptotic dependence indicator (ADI) of equation 3.7 at different frequencies using log-returns from 03/Jan/2000 to 18/July/2018. Number of observations by estimation as follows, daily: 4270, monthly: 222, quarterly: 74 observations. '—' indicates no statistical significance at 1%. Commodity and exchange rate returns both corrected by the VIX index and heteroscedasticity. **Case 1**: NER appreciation (∇S) and increase in comm. price (ΔP_{comm}). **Case 2**: NER depreciation (ΔS) and reduction in comm. price. (∇P_{comm}) **Case 3**: ∇S and ∇P_{comm}. **Case 4**: ΔS and ΔP_{comm}. The ADI value indicates the likelihood of observing extreme values simultaneously in both exchange rate and commodity returns.

Table 3.8. Panel B: ADI using lagged commodity returns at different frequencies

	Daily				Monthly				Quarterly			
	C1	C2	C3	C4	C1	C2	C3	C4	C1	C2	C3	C4
Brazil	–	–	–	–	–	–	–	–	–	–	–	–
Canada	0.05	–	–	–	–	–	–	–	–	–	–	–
Chile	–	0.08	–	–	–	–	–	–	–	–	–	–
Colombia	0.07	–	–	–	–	–	–	–	–	–	–	–
Mexico	–	–	–	–	–	–	–	–	–	–	–	–
Norway	–	0.10	–	–	–	–	–	–	–	–	–	–
Peru	–	–	–	–	–	–	–	–	–	–	–	–
Russia	–	0.06	–	–	–	–	–	–	–	–	–	–
S. Africa	–	–	–	–	–	–	–	–	–	–	–	–

The table reports the estimation results of the asymptotic dependence indicator (ADI) of equation 3.7 at different frequencies using log-returns from 03/Jan/2000 to 18/July/2018. Number of observations by estimation as follows, daily: 4270, monthly: 222, quarterly: 74 observations. '—' indicates no statistical significance at 1%. Commodity and exchange rate returns both corrected by the VIX index and heteroscedasticity. **Case 1**: NER appreciation (∇S) and increase in comm. price (ΔP_{comm}). **Case 2**: NER depreciation (ΔS) and reduction in comm. price. (∇P_{comm}) **Case 3**: ∇S and ∇P_{comm}. **Case 4**: ΔS and ΔP_{comm}. The ADI value indicates the likelihood of observing extreme values simultaneously in both exchange rate and commodity returns.

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Appendix

Table A.3.1. World's top oil producers

	2000	2005	2010	2015	2016
United States	9,058	8,327	9,691	15,139	14,829
Saudi Arabia	9,476	11,496	10,908	12,072	12,387
Russia	6,724	9,511	10,290	11,040	11,250
China	3,389	3,871	4,572	5,146	4,863
Canada	2,753	3,096	3,442	4,511	4,594
Iraq	2,582	1,889	2,398	4,039	4,443
Iran	3,765	4,239	4,243	3,485	4,364
United Arab Emirates	2,572	2,845	2,815	3,673	3,765
Brazil	1,534	2,038	2,723	3,183	3,240
Kuwait	2,201	2,672	2,449	2,880	2,991

The table displays the number of barrels (in thousands) per day. Source: U.S. Energy Information Administration (EIA).

Chapter 4

Commodity market spillovers? Revisiting the impact on financial markets

Abstract

This paper revisits the relationship between commodity and stock exchange markets. Based on a sample of commodity-exporting economies between 2000-2019, we find the correlation between price movements in these two markets increases around episodes of financial distress. Prior research attributes this increase to the effect of contagion initiated by commodity price shocks. However, we find that the documented increase in correlation during crisis episodes does not originate from shocks impacting on commodity markets. Indeed, after controlling for the effect of time varying investor risk aversion, we cannot reject the no contagion hypothesis. Our findings suggest that controlling for the effect of time varying investor risk aversion, and other factors having the potential to cause common variation across price movements in commodity and stock markets, is a key element in accurately capturing the relationship between asset returns in these markets.

4.1 Introduction

The literature documenting financial contagion in asset markets is extensive and covers a variety of financial instruments. Many studies analyse how asset price shocks propagate across markets, offering a variety of proven empirical approaches to assess the degree of financial contagion between asset classes. The relationship between international commodity markets and equity returns, in particular those in emerging economies, is currently one of the least explored areas and we believe is an interesting avenue for further examination. Recent papers (see Creti et al., 2013; Mensi et al., 2013; Roy and Roy, 2017; Xu et al., 2019) analyse the issue from the perspective of commodity-exporting emerging economies and conclude there is some evidence of what they term a contagion effect between these markets. We survey this literature in detail below, but one relevant issue which characterises these studies is that several distinct concepts such as interdependence, comovement, volatility spillovers and synchronisation are often confounded when attempting to provide evidence for the existence of market contagion. Moreover, simply documenting evidence of comovement or synchronisation between markets must be seen as being quite distinct from concluding that contagion is present (Forbes and Rigobon, 2002; Corsetti et al., 2005). It follows that many past studies which evidence a lack of robustness in the procedure they utilise to test for contagion may reach misleading conclusions.

The contributions of this study can be summarised as follows. First, we inform existing findings relating to shock propagation/contagion across commodity and equity markets by including controls for the presence of important systemic global factors that may affect the documented empirical relationships in these markets. In particular, building upon the extensive methodological discussion and evidence surveyed below, we include controls for the presence of time variation in global risk aversion, or equivalently, global investor sentiment. These factors are known to be responsible for generating comovement between asset returns in various markets, and we wish to ascertain if current findings reporting evidence of contagion between commodity and equity returns are robust to their inclusion. This helps to overcome the limitations and biases inherent in previous studies which omit to include such variables in their analysis. Second, many studies which analyse contagion and spillovers across commodity markets and other asset classes either focus upon the role of oil returns in transmitting shocks to advanced domestic stock markets, or analyse markets within a single emerging economy. We extend the scope of previous analysis by including the most relevant commodity export for nine commodity exporting countries, including developed and emerging economies. This facilitates comparisons and allows us to develop and incorporate a more structured narrative in terms of the transmission channel through which shocks to internationally traded commodities affect domestic equity returns. The key results of our study are the following. Unlike other related research that omits to control for global factors known to be important for explaining the dynamics of asset prices, namely time variation in global risk aversion/investor sentiment, we find no evidence of contagion between international commodity markets and the domestic equity returns in our sample of nine commodity-exporting economies. The findings are consistent across developed and emerging economies. They are also consistent with other evidence examining similar issues from the perspective of advanced economies, some of which control for the presence of systemic global factors. We conclude that those studies simply documenting comovement in asset returns between commodity and equity markets without considering the effect of global determinants, fail to provide a convincing rationale for this relationship and may reach misleading conclusions in terms of claiming the existence of so-called contagion effects.

Specifically, when we talk about asset price comovement in financial markets what precisely do we mean? Does the comovement between prices provide sufficient evidence to claim there is some degree of interdependence between markets? Can we interpret this interdependence as financial contagion being transmitted from one market to the other? The answers to those questions have been matters of extensive discussion not least because of the contrasting definitions of financial market contagion that various studies offer. The seminal work of Forbes and Rigobon (2002) illuminates this issues by contributing fundamental theoretical foundations to underpin this discussion. They demonstrate that empirically testing for contagion is not equivalent to simply documenting some form of comovement between asset prices across financial markets, arguing that a high pairwise correlation between markets does not provide evidence of either interdependence or contagion. Karolyi and Stulz (1996) establish the foundations of their analysis by emphasising the role of global and domestic factors in explaining financial contagion. In particular, they find that global shocks exhibit a central role in explaining the comovement between the U.S and Japanese equity returns. More recently, other related studies also highlight the relevance of controlling for underlying global systemic factors or other common external elements that may lead the observed comovement across markets, a literature summarised in the surveys by Karolyi (2003) and Dungey et al. (2005). These studies show that controlling for confounding factors is a necessary condition for distinguishing between mere correlation and financial contagion.

In this study of commodity and equity markets, we address the problems described above and show that related studies that confound key concepts often produce misleading conclusions which lack robustness. We revisit the analysis of contagion between international commodity markets and a set of emerging and developed market equity returns focusing on the role of global factors in analysing the propagation of shocks. In particular, we investigate if there is any additional role for commodity markets in transmitting shocks to equity returns in commodity-exporting economies beyond a global risk aversion/investor sentiment channel. Therefore, we interpret financial contagion as the comovement between markets that goes beyond the acknowledged asset pricing influence of time variation in global risk aversion/investor sentiment. This financial contagion definition closely follows the interpretation of previous literature describing financial contagion as the market comovement which is not explained by fundamentals (Bekaert et al., 2005; Dornbusch et al., 2000; Pindyck & Rotemberg, 1993). Moreover, most of the recent literature on financial contagion adopts a similar definition (Dungey et al., 2005; Karolyi, 2003). Under this interpretation, when two markets exhibit a high degree of comovement, this does not provide any evidence of financial contagion if there exists a third underlying factor which is responsible for the observed correlation between them. Building upon the foundations of related literature, controlling for global common factors is a fundamental part of any analysis detecting evidence of contagion rather than merely empirically documenting market comovement. In our research, the risk aversion/investor sentiment is proxied by the VIX index (VIX) in the first instance, and subsequently by the St. Louis Fed Financial Stress (SLFFE) index and the Economic Policy Uncertainty (EPU) index in robustness exercises.[1] Following Bekaert et al. (2011), we interpret the VIX index as a variable that is able to capture common global investor sentiment across markets in terms of its ability to reflect time variation in investor risk aversion. During episodes of acute financial stress, when the VIX index tends to be high, uncertainty increases and the performance of financial markets deteriorate. As a consequence, risk averse investors become more reluctant to expose their wealth by investing in risky assets. This logic lies behind the idea of colloquially dubbing the VIX index, the “fear index” (see Diebold and Yılmaz, 2014) and may explain its role in the comovement of returns in financial markets. Related studies that focus on controlling for the effect of global factors pay particular attention to the effect of global risk aversion in financial markets. Passari and Rey (2015) argue the VIX index is a key element in characterising the dynamics of risky returns, documenting that a single factor, which correlates highly with the VIX index, explains a high proportion of the volatility exhibited by risky assets. Coudert and Gex (2008) also support this idea by showing that key indicators of changes in investor risk aversion, such as the VIX, are useful in predicting episodes of financial stress in equity markets.

In the present context, several studies also share this approach and document that the link between commodity and international stock markets increases during episodes of financial distress. For instance, Cheng et al. (2014) document the role of financial traders and hedgers in commodity future markets during periods characterised by enhanced global risk aversion.

[1]The VIX index estimates the implied market volatility obtained from option prices of S&P 500 traded equities. The St. Louis Fed Financial Stress (SLFFE) index tracks financial stress levels based on a set of U.S. financial indicator variables. The Economic Policy Uncertainty (EPU) index measures economic uncertainty based on the content of news reports relating to economic uncertainty and the extent of disagreement evident across economic forecasts.

Using the VIX index as a measure of risk aversion, they demonstrate that a fall in commodity futures prices correlates with financial market downturns during times when the VIX is high. Similarly, Silvennoinen and Thorp (2013) use the VIX index, again interpreted as a risk aversion indicator, as a latent transition variable to detect different regimes in a smooth transition DCC-GARCH model. The authors conclude that the correlation between the U.S. equity market and commodity returns is enhanced during periods of financial and investor stress, as indicated by values of the VIX index.

Extending the analysis to incorporate commodity-exporting, emerging economies is potentially interesting, since international commodity markets are often claimed to be a relevant channel through which external shocks affect such economies. This is particularly noteworthy in countries where a significant proportion of economic activity is associated with the performance of the commodity sector. An increasing number of studies focus on the financialisation of commodity markets in an attempt to contribute to understanding the relationship between international commodity markets and other asset classes. Research in this field analyses the role of investment capital, emphasising investors' exposure to commodity price fluctuations and portfolio capital flows. Cheng and Xiong (2014) survey these studies, analysing how investors' commodity market perceptions have evolved during the last two decades. They show how investors interpret commodity markets as another asset class which is available to diversify risk and analyse the implications in terms of price discovery mechanisms. Moreover, Tang and Xiong (2010), argue that the recent phenomena of financialisation has increased the correlation between commodity and equity returns, especially in emerging markets.

An alternative interpretation of the relationship between commodity and stocks markets relates to the ability of commodity prices to capture expectations about the economic outlook of commodity exporting economies. In this context, if the commodity sector is an accurate gauge of domestic economic activity, then investors may interpret current economic conditions as a function of commodity sector activity. Therefore, changes impacting commodity markets may trigger changes in the composition of investors' equity market portfolios within commodity-exporting economies. Several studies build upon this idea and document the transmission mechanism. For instance, Cheng and Xiong (2014) claim that copper and oil, among other commodities, have been useful barometers for tracking economic activity during recent years. Similarly, Hamilton (2009) maintains that oil demand shocks correlate with high levels of business sentiment. Kilian and Park (2009), show that positive oil demand shocks correlate with positive shocks to U.S. stock returns. In this sense, enhanced returns in commodity markets may translate into favourable news for domestic equity returns as investors interpret commodity prices as an instrument to gauge domestic economic prospects. This transmission mechanism is consistent with the asymmetric information channel in King and Wadhwani (1990), where less informed investors extract information concerning the prospects in one market of interest from price changes in related asset markets.

4.2 Literature review

Research relating to financial contagion mainly focuses on stock returns, currencies, government debt indicators and credit markets. A relatively less explored area in this field focuses on examining the relationship between commodity and stock markets. Due to the relevance of mature financial markets for the global financial sector, most research in this regard focuses on advanced economies and consider the international oil price as a reference of commodity markets, while relatively fewer studies investigate the case of emerging stock markets. For instance, Kilian and Park (2009) analyse the case of the U.S. stock market, Zhang (2017), Hatemi-J et al. (2017), and Apergis and Miller (2009) study the case of the major six, seven and eight economies, respectively, with all these studies considering the relationship between stock returns and the oil market. Other authors, such as Hammoudeh and Choi (2006) and Janabi et al. (2010) investigate the relation between the oil market and a group of oil-exporting economies. Silvennoinen and Thorp (2013) is slightly more extensive in its selection of commodities, analysing the relationship between the U.S. stock market and a set of commodity markets covering the oil, metals and food international markets. Among the results of these studies, a common finding is that, even though commodity and stock markets exhibit some degree of relationship, in general other factors, such as other financial and global variables, tend to exhibit a higher comovement with stock markets.

Several previous studies capture the comovement between commodity and stock returns using conditional volatility models. They argue that the evolution of the estimated conditional correlation represents a measure of comovement between markets. Consequently, observing increases in this measure provides evidence of financial spillovers and contagion between markets, especially during episodes of financial stress, such as the global financial crisis in the period 2007-2009 (GFC 2007-09). Some studies focus on advanced economies. For instance, Creti et al. (2013) estimate a DCC-GARCH using data from 2001 to 2011 of 25 commodity markets and S&P500 returns. They document an increase in the comovement between stock returns and commodity markets which tends to intensify during the GFC 2007-09, interpreting this as evidence of a strengthening of the links between markets as a result of the financialisation of commodity markets. Importantly, they do not offer any approach for capturing the effect of global factors. Similarly, Mensi et al. (2013) use a univariate GARCH model to provide evidence of volatility transmission between the S&P500 index and a set of commodity markets (energy, food, gold and beverages) during the period 2000-2011. Their results suggest significant interdependence across markets, with the volatility spillovers being stronger during the GFC 2007-09. Once again, their empirical approach incorporates no variables to control for global factors. Other studies focus on emerging markets. For instance, Sadorsky (2014) employs a DCC-GARCH model for the period between 2000 and 2012 to analyse the relationship between an aggregate, emerging market stock index and a set of commodity markets (oil, copper and wheat). The study documents an intensification of the comovement between stock returns and commodity markets which is notably enhanced after the GFC 2007-09. The author interprets the results as reflecting an intensification of the spillovers across markets. Once again, this study does not control for the presence of global factors. More recently, using a DCC-GARCH model, Roy and Roy (2017) analyse the comovement between an aggregate commodity index and the Indian foreign exchange rate and the Indian equity markets, concluding that the documented increase in correlations during the GFC 2007-09 provides evidence of contagion across markets. Again, there is no attempt to control for the effect of other global variables, while in attempting to capture contagion between markets the authors employ the univariate volatilities of each respective market to explain the increase in the estimated conditional correlation during the GFC 2007-09. This has methodological issues, as the DCC-GARCH model uses estimated univariate volatilities to compute correlations. Therefore, using univariate volatilities once again to explain the estimated time-varying correlation seems inappropriate as it uses the same piece of information twice.

Another family of related studies employs the vector autoregressive (VAR) approach proposed by Diebold and Yilmaz (2009) to analyse the effect of comovement between commodity market prices and equity returns. For instance, Xu et al. (2019) examine volatility spillovers between the oil market and equity returns in China and the U.S. concluding that the relationship tends to intensify during periods of financial distress, such as the GFC 2007-09. Similarly, Awartani and Maghyereh (2013) also document the spillovers between oil markets and equity returns in the Gulf Cooperation Council Countries. They document that oil market spillovers tend to increase during the GFC 2007-09. They interpret those findings as a contagion effect transmitted from oil markets to stock returns. However, their conclusions may be a misleading interpretation in terms of reference to a contagion phenomenon since they do not include any controls to account for global factors that may be leading the increase in comovement between markets.

The previous set of studies finds a significant relation between stock and commodity returns, either in terms of contagion or spillovers across markets. However, we believe they exhibit limitations in the methodological procedure they adopt which cast doubt upon their conclusions in terms of their ability to capture contagion/spillovers effects. In particular, the lack of any controls for underlying variables that may initiate the comovement across markets makes the methodology in these studies open to debate. It is not possible to eliminate concerns that the documented increase in asset market comovement is the result of omitted global factors, leading to concerns that the findings may be less than robust. Following Karolyi (2003) and Dungey et al. (2005), controlling for global factors is a crucial element in characterising how shocks propagate across markets and, therefore, is necessary in order to capture the presence of contagion. Controlling for an underlying common element that may dictate the pricing relationship exhibited among markets allows one to differentiate between a mere asset comovement and a contagion effect. The difference between these two concepts is a key part of the motivation of the present study.

There are a some articles that account for additional financial channels and global factors in examining claims of asset market contagion. These articles tend to find either no effect or a limited role for commodity market spillovers. The majority of these articles either examine oil price shocks, focus upon one country, or exclude the GFC 2007-09, all of which gives our analysis more scope and generality. For instance, Hammoudeh and Choi (2006), using a VEC model and data from 1994 to 2004, find a limited effect of oil prices on the domestic equity returns of five oil exporting-countries from the Gulf Cooperation Council. Importantly, they find that the impact of other international financial shocks, such as U.S. equity returns and U.S. interest rate shocks, account for most of the effect upon the domestic stock returns of the oil exporting-countries. Apergis and Miller (2009) estimate a structural VAR using data from 1981 to 2007 for eight advanced economies to examine the relationship between oil prices and stock returns. Although they find a significant effect from oil prices transmitting shocks to equity markets, the effect tends to be quite small in comparison to idiosyncratic shocks originating in the financial sector. Similarly, Kilian and Park (2009) uses a structural VAR form 1973 to 2006 to disentangle the effect of oil price shocks to the U.S. financial market. They find that oil price shocks explain a fifth of the variability of U.S. stock returns. Importantly, non-oil related shocks emanating from the U.S. financial market explain the majority (80%) of equity return variability. Unlike the present study, none of these authors examine periods of acute global financial market stress when contagion effects are documented to be at their peak, so their findings maybe biased to periods characterised by more tranquil markets.

Janabi et al. (2010) investigate the role of oil and gold prices in relation to market efficiency for six Gulf Cooperation Council economies during the 2006 – 2008 period. They find that gold and oil returns provide no additional information for explaining equity market efficiency in the sample, concluding that those markets exhibit a semi-strong-form of efficiency. Hatemi-J et al. (2017), using stock returns from 1975 to 2013 for the case of the G7 economies, find there is no causal relationship between oil prices and the equity markets for the countries in the sample. However, they document an asymmetric causality between the oil price and U.S., Japanese and German equity returns, but no relationship in the other four economies. Silvennoinen and Thorp (2013) analyse a wider number of commodity markets by estimating the correlation of each commodity returns with the U.S. stock returns. The authors use a smooth transition DCC-GARCH model including the VIX as a threshold variable to capture the effect of global risk aversion. The study finds that U.S. stock and commodity returns are related, however the relationship does not apply to all commodity markets and depends on the period under analysis. The results show there is an increase in correlation during the GFC 2007-09 between U.S. equity market and oil returns. They also document an increase in correlation, albeit to a lesser extent, between commodity metals and U.S. stock returns, however that increase in correlation is not associated with the GFC 2007-09. There is no correlation evident between commodity food and U.S. equity returns. Although Silvennoinen and Thorp (2013) include a wider range of commodity markets in the analysis, the analysis only focuses on the impact on the U.S. stock market. In our study, we aim to extend the evidence to a broader sample of economies. Zhang (2017), using the Diebold-Yilmaz approach, documents that the effect of price dynamics in the oil market on a set of the major stock market returns appears very limited in comparison to the effect of other financial variables on equity returns in their sample.

4.3 Methodology and data

We utilise two different methodologies in order to capture the contagion effect between commodity and stock markets. The first approach adopts a Multivariate DCC-GARCH model originally formulated by Engle (2002). The second follows the Diebold and Yilmaz (2009) methodology employing a forecast error variance decomposition obtained from a vector autoregressive model. Each approach possesses advantages and disadvantages, which we discuss below. However, the inclusion of both methodologies allows us to obtain a clearer perspective on the existing results and we believe it strengthens the conclusions of the analysis.

4.3.1 Multivariate DCC-GARCH

The Multivariate DCC-GARCH methodology estimates the time-varying correlation between commodity and stock returns in order to better understand how the relationship evolves over time. This approach facilitates an analysis of non-linearities in the covariance matrix of log-returns. Importantly, this methodology also controls for the effect of heteroscedasticity in log-returns. Following the discussion in Forbes and Rigobon (2002), the presence of heteroscedasticity in log-returns imparts bias to any correlation estimates during a period of high return volatility. Therefore, by using this multivariate GARCH framework, we can incorporate the effect of time-varying return volatility and avoid problems relating to the presence of bias in the estimates of the correlations between markets. A drawback of this methodology is that it does not enable the identification of the source of the shocks. In this sense, this approach only provides an indicator capturing market comovement and is unable to determine the origin of any shock(s) causing returns in the two markets to move together.

Following Engle (2002), the DCC-GARCH model implemented in this paper is as follows:

$$r_t = Bx_t + u_t$$

$$u_t = H_t^{1/2}\epsilon_t$$

Let r_t be a $n \times 1$ vector of asset returns at time t, while u_t is a $n \times 1$ vector of residuals that follow a conditionally normal distribution, with mean equal to zero and covariance matrix H_t, such that:

$$u_t|\Omega_{t-1} \sim N(0, H_t)$$

where Ω_{t-1} represents the information set at time $t-1$. The covariance matrix is given by:

$$H_t = D_t R_t D_t \tag{4.1}$$

where D_t is a $n \times n$ diagonal matrix of time-varying standard deviations with the ith term in the diagonal corresponding to $\sqrt{h_{it}}$ with estimates undertaken using the univariate GARCH models of equation 4.2.

$$h_t = \alpha_0 + \alpha_1 u_{t-1}^2 + \beta h_{t-1} \tag{4.2}$$

R_t is a $n \times n$ matrix containing the time-varying conditional correlations. The dynamic of the conditional correlation matrix (R_t) is given by:

$$R_t = diag\left(Q_t\right)^{-1} Q_t diag\left(Q_t\right)^{-1}$$

Where $diag\left(Q_t\right)$ is a $n \times n$ diagonal matrix in which the ith diagonal element corresponds to the ith diagonal element of matrix Q_t given in equation 4.3.

$$Q_t = \left(1 - \lambda_1 - \lambda_2\right) \overline{Q} + \lambda_1 \epsilon_{t-1} \epsilon_{t-1}' + \lambda_2 Q_{t-1} \tag{4.3}$$

In equation 4.3, λ_1 and λ_2 are coefficients to be estimated. $\epsilon_{it} = \frac{u_{it}}{\sqrt{h_{iit}}}$ corresponds to the standardised residual obtained from the univariate model in equation 4.2. $\overline{Q} = T^{-1} \sum_{t=1}^{T} \epsilon_t \epsilon_t'$ is the $n \times n$ unconditional correlation matrix of u_t.

We estimate the proposed model using quasi-maximum likelihood procedures.[2] We estimate the multivariate DCC-GARCH model in three different stages to avoid computational issues which may arise due to the number of coefficients to be optimised. The first stage estimates the univariate mean and volatility equations. In the second stage, we estimate the DCC coefficients given the estimates from the first step. The final stage uses the results of the first and second steps as initial values, and estimates the whole model simultaneously. This process guarantees a better algorithm convergence in order to optimise the maximum likelihood function and avoids generating inconsistent estimates of the conditional correlation matrix (Engle & Sheppard, 2005).

4.3.2 Diebold and Yilmaz approach

The second methodological approach we employ is developed by Diebold and Yilmaz (2009) and offers a flexible way to model the relationship between markets. Based on a VAR model formulation, the advantages of this methodology are: first, it provides a measure of the spillovers from one market to another and second, it facilitates identification of the direction of the shock and an understanding of which asset generates any shocks which are transmitted to other markets. Due to the nature of the VAR approach, endogeneity is not an issue. The downside of this approach is that it can only capture linear effects between markets. The

[2]Cappiello et al. (2006) show that even when assumptions concerning the normal distribution of the error term are invalid, the results of the model still have a valid quasi maximum likelihood estimation interpretation.

reduced form of our VAR models is as follows:

$$Y_t = AY_{t-1} + u_t$$

We adopt two specifications in this section. Our initial model formulation does not control for the effects of time variation in global investor risk aversion and is given by:

$$Y = [r_{1,t}, r_{2,t}] \tag{4.4}$$

where $r_{1,t}$ and $r_{2,t}$ stand for the time-series of stock and commodity returns, respectively.

The second model includes a control for the presence of such investor sentiment effects as captured by the VIX index, and is given by:

$$Y = [dVIX_t,\ r_{1,t}, r_{2,t}] \tag{4.5}$$

where $dVIX_t$ represents the change in the VIX index.

Following Diebold and Yilmaz (2009), we compute the spillover indices which we obtain from the forecast error variance decomposition (FEVD) from a VAR model. In the context of a VAR model, the FEVD corresponds to the percentage of the variance of the forecast error which is explained by orthogonal shocks of the variables in the model. A natural interpretation of the FEVD relates to the key question of interest, namely how shocks originating within one asset market impact the other (i.e.: the spillover effect). If we compute the VAR using a sample of rolling windows, then we are able to calculate the spillovers between markets through time. As adopting this methodology enables us to identify the source of shocks, it becomes possible to specify which asset market, if any, impacts the other.

4.3.3 Contagion test

Our earlier literature survey questions certain existing approaches capturing asset market contagion, attributable to the fact that some empirical models may omit the influence of important external underlying factors that may be responsible for the measured correlations between commodity and stock returns. We conclude that testing for market contagion without controlling for the presence of such systemic global factors generates conclusions which may not be robust. This issue arises both when evaluating the dynamic correlations of the DCC-GARCH model and the spillovers of the Diebold and Yilmaz methodology. Furthermore, as Diebold and Yilmaz (2009) emphasise, their spillovers measure is only intended to capture the interdependence between markets and they explicitly avoid considering it as a contagion measure. Our approach to overcoming these drawbacks is to incorporate the effect of global risk aversion (investor sentiment) as an important determinant of the measured relationship between markets. In particular, we employ the VIX index as our control variable for the underlying global risk factors that may explain the documented correlation/spillovers between equity and commodity markets. Our objective is to study if the higher measured comovement between markets around episodes of crisis can be accounted by the presence of this underlying systemic factor.

Employing the first methodology, the DCC-GARCH, we compare the time-varying correlation between two specifications of the mean equation. One specification (equations 4.6 and 4.7) excludes and the other (equations 4.8 and 4.9) includes the VIX index. Subsequently, we compute the average correlation between each specification during sub-samples covering two episodes of financial crisis we specify below. When using the second approach, the Diebold-Yilmaz variance decomposition framework, we compute the spillovers emanating from returns on the country-specific commodity market to the associated domestic equity market. In similar fashion to the previous methodology, we compute two specifications of the model, one excluding and the other including the VIX index in the VAR model, as shown in equations 4.4 and 4.5, respectively, and calculate the average spillovers from each specification during crisis episodes.

We carry out this comparison for two recent significant episodes of financial distress, the global financial crisis in the period 2007-09 (GFC 2007-09) and the European sovereign debt crisis during 2009-12 (ESDC 2009-12). We select these periods for two main reasons. First, this paper focuses upon periods of financial vulnerability, which are precisely those in which previous literature documents that the correlation/spillovers between asset markets tends to increase. It is during such episodes that we wish to discern if there is any additional impact of price changes in commodity markets that exert an influence on equity returns beyond the impact of factors influencing global investor sentiment. In addition, by analysing these episodes we are able to benchmark our results against other studies which focus on the same episodes. Second, related studies mostly evaluate the effect of commodity markets on the relevant domestic equities during the GFC 2007-09 episode. Therefore, by also analysing the contagion effect during the European Sovereign crisis, we aim to extend the existing evidence that relates stock return movements to those in commodity markets. We date the beginning of the GFC 2007-09 to be 17 July 2007 when Bear Stearns disclosures information alerting markets to the sharp plunge in the value of its investment funds (leading to the company's bankruptcy days later). We set the end date of the GFC 2007-09 crisis as 31 August 2009, when according to the U.S. Financial Crisis Inquiry Commission, asset markets stabilised. According to the U.S. National Bureau of Economic Research, this end date is also consistent with the end of the U.S. recession. Following relevant milestones in the crisis timeline, we set the beginning of the ESDC to be 2 October 2009, when the Greek government disclosed that the budget deficit represents 12.5% of the GDP, which is twice its expected level. The end date corresponds to 26 July 2012 when the European Central Bank's president, Mario Draghi, announced the Outright Monetary Transactions program following his "whatever it takes to save the Euro" speech. This programme enabled the ECB to purchase eurozone member countries sovereign bonds in the secondary market. Following this ECB commitment, the spreads associated with sovereign bonds of the most severely affected eurozone economies exhibit a pronounced downward trend, relieving the pressure in European sovereign debt markets and helping to alleviate the negative investor perception of the Eurozone.

In order to establish our contagion test between asset markets we follow Celik (2012) who conducts a test for differences in means to compare model determined time-varying correlations at different points in time obtained from a DCC-GARCH model.

In this study, for each of the two empirical approaches we introduce above (i.e.: the DCC-GARCH model and the Diebold-Yilmaz variance decomposition approach), the contagion test we carry out is a two-sample test for equal means of the outcomes of the two methodologies (i.e.: time-varying correlation or spillovers) under two alternative specifications: including and excluding the effect of the global risk aversion. Therefore, we run the contagion test for each methodology. First, comparing the mean of the time-varying correlations with and without including the effect of global factors. Second, comparing the mean of the spillovers between markets with and without including the effect of global factors. We compute the difference in mean focusing on two sub-sample crisis episodes, the GFC 2007-09 and the ESDC 2009-12. The null and alternative hypothesis are as follows:

$$\mathrm{H}_0 : \mu_{vix} \leq \mu_{no\ vix} \quad (no\ \ contagion)$$

$$\mathrm{H}_a : \mu_{vix} > \mu_{no\ vix} \quad (contagion)$$

where μ_{vix} and $\mu_{no\ vix}$ stand for the means of the sample correlation/spillover computed during the financial crisis episode using the model specification with and without including the effect of the global risk aversion (i.e.: the VIX index), respectively. Under the null hypothesis the average correlation/spillover obtained from the model including the effect of global factors, captured by the VIX index in this case, is less than or equal to the correlation/spillover obtained from the model excluding the VIX, meaning that the effect of the global factors accounts for the increase in correlation/spillovers between markets during crisis episodes, therefore, no contagion evidence between stock and commodity markets. The alternative hypothesis states the average correlation/spillover obtained from the model including the effect of global factors is strictly greater than the one obtained from the model excluding the effect of global factors, suggesting there is an additional link between commodity and stock markets that goes beyond the effect of global risk aversion, therefore, we interpret it as contagion across markets. The test statistic (t) is as follows:

$$t = \frac{\bar{\rho}_{vix} - \bar{\rho}_{no\ vix}}{\sqrt{s^2_{vix}/n_{vix} + s^2_{no\ vix}/n_{no\ vix}}}$$

where $\bar{\rho}_{vix}$ and $\bar{\rho}_{no\ vix}$ are the means of the sample correlation/spillovers during each crisis episode using the specifications with and without the VIX index, respectively. s_{vix} and $s_{no\ vix}$ are the sample variances and n_{vix} and $n_{no\ vix}$ are the sample sizes. We reject H_0 at α% significance level if $t > t_{(1-\alpha,v)}$, where we calculate the degrees of freedom (v) as:

$$v = \frac{(s^2_{vix}/n_{vix} + s^2_{no\ vix}/n_{no\ vix})^2}{(s^2_{vix}/n_{vix})^2/(n_{vix}-1) + (s^2_{no\ vix}/n_{no\ vix})^2/(n_{no\ vix}-1)}$$

Other studies use the same logic to test for asset market contagion but follow a slightly different approach. For instance, Longin and Solnik (1995) test for the statistical difference in the variance-covariance matrix when comparing returns during times of financial distress and more tranquil periods.

4.3.4 Data

The data consists of daily observations of stock and country-specific commodity returns, from 7 January 2000 to 29 March 2019. The stock returns correspond to the following nine commodity-exporting economies, in parentheses the major country-specific export-commodity: Brazil (oil), Canada (oil), Chile (copper), Colombia (oil), Mexico (oil), Peru (copper), Norway (oil), Russia (oil), and South Africa (gold). The selection of countries meets two criteria: first, the country must be one in which commodities represent a significant percentage of its exports (see table 4.1), and second, the country must adopt a "floating" or a "free-floating" exchange rate regime throughout the sample period.[3] The source of data of stocks, commodities and VIX indices is Bloomberg. We obtain the data of the St. Louis Fed financial stress index from the Federal Reserve Economic Data website and the Economic Policy Uncertainty index from Baker et al. (2016)'s website.[4] We compute stock returns from stock indices measured in local currency, while we obtain commodity returns from commodity price indices denominated in U.S. dollars. We follow this strategy to avoid introducing confounding factors associated with fluctuations in the value of the U.S. dollar.

[Table 4.1 in here]

Given the relevance of commodity exports for the countries in the sample, we analyse contagion between the domestic stock market of a particular country and the international price of the main export-commodity of this economy. Our goal is to assess how specific commodity shocks to economically important, country-specific commodity markets affect domestic stock returns. The relevance of analysing the relationship between country-specific commodity markets and domestic stock returns lies in the evidence documenting differentiated price dynamics across international commodity markets. For instance, Daskalaki et al. (2014), document that commodity markets are segmented from one another, meaning that each commodity follows its own price dynamics and its price movements are not generally related to another commodity type. Similarly, Erb and Harvey (2006) and Kat and Oomen (2007) document that the evolution of commodity prices tend to display heterogeneity through time. They conclude that the differentiated dynamic of commodity prices substantially improves portfolio diversification based on commodity assets. In addition, Adams and Glück (2015) show that commodity prices respond differently to the effect of financialisation of commodity markets. They find that, since 2008, oil and copper prices tend to increasingly relate to

[3]The exchange rate regimes we employ correspond to the most flexible categories under "De Facto Classification of Exchange Rate Arrangements" elaborated by the International Monetary Fund. See International Monetary Fund (2019).

[4]Federal Reserve Economic Data website: https://fred.stlouisfed.org. Economic Policy Uncertainty website https://www.policyuncertainty.com.

changes in the U.S. stock market, while others, such as aluminium and wheat prices, show no increase in correlation with U.S. equity returns. Considering the relevance and economic importance of a particular commodity market within a particular country, the documented heterogeneity of price dynamic of international commodity markets lead us to focus on one country-specific commodity price for each economy in our sample.

4.4 Results

4.4.1 DCC-GARCH model

This section presents the estimates of the DCC-GARCH model following the methodology introduced in section 4.3.1. We estimate two different model specifications, with and without controlling for the effect of variations in global investor sentiment (risk aversion) to account for its effects during periods of financial distress.

4.4.1.1 DCC-GARCH without global risk aversion

Under this model specification which excludes changes in the VIX index, we estimate the mean equation of the model for each country as follows:

$$r_{1,t} = \gamma_0 + \gamma_1 r_{1,t-1} + u_{1,t} \tag{4.6}$$

$$r_{2,t} = \gamma_0 + u_{2,t} \tag{4.7}$$

Where $r_{1,t}$ and $r_{2,t}$ stand for equity and country-specific commodity returns, respectively. We include the one period lag of equity returns in its own-return equation to capture the effect of persistence in stock returns. There is no requirement to include lagged values of commodity returns as these asset returns do not exhibit any persistence.[5] We model the univariate volatilities for stocks and commodity returns using a GARCH(1,1) model as shown in equation 4.2. We compute the conditional correlation from the estimated time-varying covariance matrix in equation 4.1. Table 4.2 presents the results of this model specification. Panel A documents the results for the mean equation. The constant term is statistically significant and close to zero for all countries in the sample. The persistence coefficient is statistically significant in the case of Brazil, Chile, Colombia, Mexico, Peru, Russia, and South Africa which is consistent with the idea that emerging stock market returns tend to exhibit some degree of persistence. Panel B shows the results for the univariate volatility models. All coefficients are statistically significant and take values comparable to those reported in related studies. The constant term (α_0) is close to zero, the α_1 and β coefficients, in both equations, sum to less than one, confirming that the univariate volatilities follow a stationary process.

[5] See table A.4.1 in appendix

Panel C presents the estimated coefficients for the time-varying correlation. The estimates are highly significant for all countries and again display similar values when compared to the results documented in related studies. Using this specification, we proceed to estimate the time-varying correlation for each country.[6]

[Table 4.2 in here]

4.4.1.2 DCC-GARCH including global risk aversion

In order to include the global risk aversion component in our model, we now proceed to estimate an alternative specification incorporating variations in investor sentiment, as captured by the VIX index in the mean equation as follows:

$$r_{1,t} = \gamma_0 + \gamma_1 r_{1,t-1} + \gamma_2 dVIX_t + u_{1,t} \tag{4.8}$$

$$r_{2,t} = \gamma_0 + \gamma_2 dVIX_t + u_{2,t} \tag{4.9}$$

where $dVIX$ denotes the change in the VIX index. We estimate the univariate volatilities and the conditional correlation, in the same way as in the previous specification.

[Table 4.3 in here]

Table 4.3 presents the results. The mean equation estimation in panel A confirms the results are similar to the previous specification. In the case of stock returns, the global risk aversion factor, the coefficient associated with the $dVIX$, displays a negative value indicating that increases in investor risk aversion is associated with a decline in stock returns in all countries. In the case of country-specific commodity return equations, the coefficient on $dVIX$ is negative for all countries with the exception of South Africa where it is positive, although only marginally significant. This finding makes certain intuitive sense as the country-specific commodity we associate with South Africa is gold, whose role as a safe-haven asset means its value tends to increase during episodes of financial distress. The remainder of the model estimations for the univariate volatilities in panel B and the conditional correlation in panel C, reveals similar results in comparison to previous specifications of the model.

[Figure 4.1 in here]

Figure 4.1 displays the plots for the conditional correlation under the two different specifications. The grey line represents the conditional correlation, which is estimated using the

[6] In addition, for comparison purposes, we plot our correlation measure with a pairwise correlation measure we estimate using a rolling window. The figure A.4.1 in the appendix presents the results. Both correlation measures, the DCC dynamic correlation in black and the rolling window correlation in grey, are highly similar. On the basis of the above results, we maintain that our DCC-GARCH model is suitable for modelling the time-varying correlation between commodity and stock markets.

model without $dVIX$, while the black line corresponds to the one estimated from the model which includes $dVIX$. In almost all cases, we observe conditional correlations tend to increase during periods of financial distress, namely the GFC 2007-09 and later during the ESDC 2009-12. However, both correlations tend to exhibit discernible differences through time. In particular, the pattern of conditional correlation that includes the global risk aversion factor (black line) tends to be lower in magnitude than the one which excludes it (grey line). The difference between the two conditional correlations is accounted by the influence of the VIX factor. This global component appears to account for a significant element of the co-movement between stock and commodity markets. After controlling for its effect, the correlation between markets tends to be somewhat lower in magnitude (black line), particularly around the crisis episodes we analyse in this paper (GFC 2007-09 and ESDC 2009-12). We interpret this as a reflection of the relevance of recognising the important role played by global factors in the mechanism underlying the propagation of shocks across domestic asset markets.

To this point, we interpret the conditional correlation dynamics as an appropriate measure to capture comovement between stock and commodity markets. The increase in such comovement between commodity and stock markets is something we intuitively expect, given that evidence of increased correlation between asset markets tends to arise during periods of financial turmoil. In this sense, it is not satisfactory simply to compare the evolution of the conditional correlation in order to reach definitive conclusions regarding the degree of contagion across markets. This is because there may be confounding factors, in the present case time varying risk aversion as captured by movements in the VIX index, which causes both markets to co-move simultaneously. This is a key difference to incorporate when comparing our results with those from related studies that omit controlling for the influence of global factors. We discuss this idea more formally in the next section.

4.4.1.3 Contagion test based on the DCC-GARCH methodology

In order to construct an appropriate test for contagion during stress episodes, we compare the conditional correlation between the two model specifications, namely those including and excluding the VIX index risk aversion factor. In particular, following the contagion test we propose, we compute the means of the conditional correlations during the GFC 2007-09 and the ESDC 2009-12. Using the test outlined in section 4.3.3 we determine whether the average correlation when the model excludes the global risk factor is greater or equal to the average correlation obtained when this factor is included. In other words, the specification tests the null hypothesis of no contagion between stock and commodity markets.

[Table 4.4 in here]

Table 4.4 panel A displays the contagion test under two crisis episodes, the GFC 2007-09 and the ESDC 2009-12, using the VIX index to control for the influence of global factors. The results indicate that for most countries we cannot reject the null hypothesis of no contagion for both crisis episodes under analysis. Canada during the GFC 2007-09 is the one exception, where we find some evidence of contagion (at 5% level of significance). Overall, these results point to the conclusion that after controlling for the effect of this channel of global risk aversion, there is no additional influence exerted by commodity markets on equity returns for the majority of the countries in the sample. In terms of the existence of contagion between markets, we interpret these results as revealing no significant excess of comovement between markets after controlling for the impact of time variation in global risk aversion, capturing a global investor sentiment factor. Using this methodology, we soundly reject the contagion hypothesis for the majority of the countries in the sample.

4.4.1.4 Robustness to Alternative Global Risk Factors

Now we undertake robustness exercises by reformulating the specifications to include alternative measures which have been proposed in the literature to capture the effect of systemic global factors which are known to assume particular relevance during times of financial stress. We conduct two alternative specifications. Initially, we replace the VIX by the St. Louis Fed's Financial Stress (SLFFE) index, and subsequently with the Economic Policy Uncertainty (EPU) index proposed and computed by Baker et al. (2016).[7] Both indices are global in their orientation, attempting to capture the effect of systemic shocks which impact upon the global economy. This makes them appropriate indicators of the presence of common factors which may affect the dynamics of returns in both equity and commodity markets. Kliesen et al. (2012) provide a comprehensive comparison of the SLFFE index with other financial stress indicators, including the VIX index. The authors show the appropriateness of the SLFFE index to capture episodes of financial distress. Baker et al. (2019) discuss the difference among the EPU index and other financial related indices. The authors document that the EPU index mostly captures global economic-related uncertainty and appears relatively less sensitive to financial market developments in comparison to financial stress indices.

The SLFEE and the EPU indices differ from each other in terms of what they specifically measure. The SLFFE index is constructed using a principal components methodology and captures variation in a set of indicators relating to the overall financial health of the U.S. economy, reflecting both corporate and governmental concerns. Given the performance of U.S. financial markets is universally acknowledged to be important for the health of the global economy, the SLFEE maybe taken to be a reasonable indicator of performance expectations in asset markets worldwide. The EPU index employs a text search methodology with the objective of capturing the level of global economic uncertainty. This index has been demonstrated to be able to accurately track episodes of economic crisis and financial distress, such as the GFC 2007-09 and the ESDC 2009-12. Several studies incorporate the EPU index to capture how the effect of economic uncertainty affects macroeconomic performance in terms of countries' economic growth (Aisen & Veiga, 2013), stock market returns (Antonakakis

[7] Both indices, the SLFFE index and the EPU index, are available at a monthly frequency, so we conduct the robustness exercises using observation at monthly frequency for the same sample period as before, January 2000 to March 2019. Figure A.4.2 in the appendix displays a comparison of these two indices and the VIX. The correlation matrix below the figure confirms the correlation between the VIX and SLFFE indices is high, while the association turns much weaker when comparing both these indices with the EPU index.

et al., 2013), oil prices (Antonakakis et al., 2014), and output, employment and investment (Baker et al., 2016).

Table 4.4 panels B and C depict the results of the contagion tests using the SLFFE index and the EPU index, respectively. When we use the SLFFE index to capture global factors (panel B), we cannot reject the null hypothesis of no contagion for all countries and both crisis episodes under investigation. These results are consistent with those using the VIX index to measure global factors (panel A). Similarly, when we use the EPU index as the global factor variable, we do not reject the no contagion hypothesis for the majority of countries during both crisis episodes. Two exceptions appear in the cases of Peru and South Africa during both crisis periods, where we are indeed able to reject the null hypothesis of no contagion. In general, both alternative measures support the conclusions we obtain using the VIX index that findings of contagious spillovers between commodity and equity markets are not robust to including controls for time variation in systemic global factors which may exert an influence upon all asset markets.

4.4.2 Diebold-Yilmaz approach

This section presents the estimates of spillovers from one market to another using the Diebold-Yilmaz methodology we introduce in section 4.3.2. Spillovers correspond to the percentage of the forecast error variance of a particular variable explained by orthogonal shocks of the variables in the VAR model. We estimate a VAR model of order 2, setting the forecast error variance decomposition horizon to be 12 periods ahead using the Cholesky decomposition. We use a rolling window length of 1000 observations to undertake the forecasts. The model structure is comparable to that in related studies adopting similar estimation procedures (see Diebold and Yilmaz, 2009). As with the previous methodology, our Diebold-Yilmaz framework involves estimating two VAR specifications, one including (equation 4.5) and the other excluding (equation 4.4) the effect of time varying risk aversion as reflected in the VIX index. We conduct both VAR specifications for each country in the sample.

4.4.2.1 Spillovers between markets

Figure 4.2 depicts the decomposition of the relevant shocks impacting equity markets. Those emanating from the commodity markets when the model excludes the effect of the VIX index are indicated by the grey line. The black line indicates the measured shocks coming from commodity markets when we include the VIX in the empirical specification, while the dashed line captures the shocks attributable to the VIX index itself when it is incorporated into the model.

[Figure 4.2 in here]

Our analysis yields the following results. First, it is evident that for all countries in the sample, the spillovers transmitted from commodity to equity markets vary over time. Before the GFC 2007-09, such spillovers exhibit a similar evolution and tend to be close to zero under both model specifications. After the GFC 2007-09, figure 4.2 indicates that commodity markets spillovers increase and are higher in absolute magnitude in the absence of the global risk aversion factor. Second, the impact of the global risk aversion exceeds that of commodity markets for all countries under analysis and tends to strengthen after the GFC 2007-09 displaying additional increases during the ESDC 2009-12 episode.

These results reflect the fact that before the GFC 2007-09, when the financialisation of commodity markets was less in evidence, time variations in global risk aversion were less relevant for commodities as they were not widely considered to be an important asset class in the portfolio composition of international investors. This translates to the negligible differences we obtain in measuring the impact of commodity market spillovers on equity returns whether or not we control for global risk aversion. This evidence is consistent with the narrative accompanying the observed financialisation of commodity markets documented by previous studies. For instance, Adams and Glück (2015) document a significant change in the relationship between commodity and stock markets from 2008. They argue that commodity markets substantially increase the interest of investors willing to diversify portfolios, consequently the episode of financial distress during the GFC 2007-09 not only negatively impacted stocks but also commodity markets which were considered as an alternative asset class. In this context, Cheng and Xiong (2014) highlight investors' risk appetite as one of the main channels through which commodity and stock markets relate to each other. After the GFC 2007-09, when financialisation of commodity markets intensifies, controlling for risk aversion becomes relevant as it is during financial distress episodes that it manifests itself to be an important transmission channel through which external shocks to major commodity export markets impact equity returns. Of relevance is the fact that the estimated global risk aversion spillovers, represented by the dashed line in figure 4.2, becomes a more relevant source of shocks during both the GFC 2007-09 and the ESDC 2009-12. The difference between commodity market spillovers in both models (grey as compared to black line), also reveals the effect of global risk aversion. After the GFC 2007-09, we observe both a lower impact in the magnitude of the commodity market spillovers to equity markets and noticeable less sharp increases during episodes of financial distress when we include the VIX index in the model (black line) in comparison to the case when we exclude it (grey line). We interpret this feature as evidence of the importance of the global risk aversion component in transmitting shocks. Consequently, the commodity market channel appears less relevant once we include global factors in the analysis. In section 4.4.2.3 we provide a formal tests of these findings in terms of evidence of contagion between markets.

4.4.2.2 Net spillovers between markets

In this section we analyse the net effect between commodity and stock market spillovers. Net spillovers between commodity and domestic stock markets are defined as follows:

$$S_{net} = S_{comm \to stock} - S_{stock \to comm}$$

Where S_{net} represents the net spillovers between commodity and domestic stock markets, $S_{comm \to stock}$ denotes the spillovers of commodity returns on stock returns (i.e.: the percentage of the forecast error variance of stock returns explained by orthogonal shocks of commodity returns), and $S_{stock \to comm}$ corresponds to the spillovers of stock returns on commodity returns (i.e.: the percentage of the forecast error variance of commodity returns explained by orthogonal shocks of stock returns).

[Figure 4.3 in here]

Figure 4.3 depicts the net spillovers between commodity and domestic stock markets. Here, the black (grey) line corresponds to the model which includes (excludes) the VIX index. Our findings are as follows. First, focusing on net spillovers which include the effect of the VIX (black line), the figures reveal the net spillovers tend to be positive for all countries during the entire sample period. Some exceptions appear in the first part of the sample for Chile and Peru, where net spillovers tend to be marginally negative. These results indicate that commodity markets are net *providers* of shocks, while stock markets are net *receivers* (i.e.: spillovers from commodity to stock returns are greater than spillovers from stock to commodity returns). This evidence suggests that, even though stock returns may impact commodity markets, the stronger effect goes from commodity to stock markets. Results without considering the effect of VIX (grey line) show similar patterns, although they come from a less robust approach as the model specification does not include the effect of global risk aversion. Second, for both model specifications, including and excluding the VIX index, we find that net spillovers increase during episodes of financial distress, displaying a sustained increment starting from the GFC 2007-09 and further hikes during the ESDC 2009-12. However, and consistent with the results of the previous section, net spillovers obtained from the model including the VIX index appears lower in magnitude and display less pronounced spikes during crisis episodes in comparison to net spillovers without considering the effect of the VIX. This confirms our previous findings highlighting the relevance of global risk aversion as a source of shocks to domestic financial markets and, consequently, a less relevant impact of commodity returns on stock returns.

4.4.2.3 Contagion test of spillovers effects

In this section we statistically assess the contagion hypothesis between stock and commodity markets considering the spillovers obtained using the Diebold-Yilmaz methodology of section 4.4.2.1. In particular, based on the contagion test we describe in section 4.3.3, we compare the mean spillovers from commodity to stock markets obtained under two model specifications, including and excluding the VIX index. We undertake the comparison of mean spillovers during crisis episodes, therefore, we test contagion twice, first for the GFC 2007-09 and second for the ESDC 2009-12 episode. Our goal is to statistically evaluate whether the average spillovers obtained from the model excluding the VIX index is greater or equal to the average spillover estimated based on the model including this global risk aversion factor. That is, the test evaluates the null hypothesis of no contagion between stock and commodity markets. The results are as follows.

[Table 4.5 in here]

Table 4.5 panel A shows that, for all countries and both crisis episodes, the average commodity market spillovers without the VIX (column 'No global factor') is statistically greater than the average commodity market spillovers including the effect of the VIX (column 'Global factor'). Thus, we do not reject the null hypothesis of no contagion. In addition, we implement two separate robustness exercises by replacing the VIX index by the SLFFE index and then by the EPU index. Table 4.5 panel B presents the results using the SLFFE index to capture the effect of global factors. As we can see, results remain the same for all countries and both crisis episodes, namely we do not reject the no contagion null hypothesis. Table 4.5 panel C displays the results using the EPU index as a proxy of the global factor. The results hold for most of the countries and both crisis episodes with the exception of South Africa during the GFC 2007-09. Overall, the results displayed in table 4.5 show there is a substantial lower impact of commodity markets on stock returns after considering the effect of global factors. Our evidence suggests commodity markets play a secondary role in transmitting shocks to domestic returns. Specifically, commodity market spillovers during episodes of financial distress become less relevant after including the effect of global risk aversion factors. Considering our results, we reject the contagion hypothesis between commodity and stock markets. The conclusions of this section are consistent with the ones we previously obtain in section 4.4.1.3 using the DCC-GARCH methodology.

4.5 Conclusions

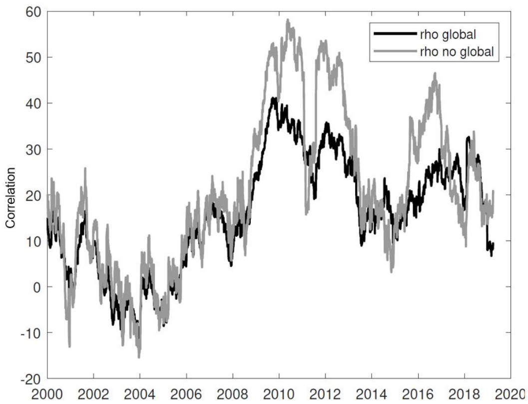
This study revisits current explanations for the measured contagion effects between commodity and equity markets documented in several previous studies. We use a sample of nine major commodity exporting economies, which includes emerging market and advanced economies, between 2000 and 2019. Our main findings are as follows. Using both a DCC-GARCH model and the approach pioneered by Diebold and Yilmaz (2009), we find that the conditional correlation/spillovers between commodity and equity returns varies over time. The correlation/spillovers tends to increase during periods of acute financial distress, namely the GFC 2007-09 and ESDC 2009-12. The evidence applies to all countries in the sample and corroborates previous findings. However, unlike many previous studies, we maintain that this is not evidence of contagion. Specifically, when we account for the effect of time variation in global investor sentiment and risk aversion by including suitable proxies such as the VIX index, the EPU index and the SLFFE index, we observe the impact of commodity markets on equity returns reduces greatly in magnitude in almost all countries in the sample.

These findings lead us to believe that there may be a more limited role for commodity markets in transmitting shocks which impact equity returns than previously maintained. Once we incorporate measures of global risk aversion, we largely reject the idea of contagion arising between these markets. This conclusion is in stark contrast with the findings of related studies which analyse such contagion effects in emerging equity markets and omit the effect of global factors. Failing to control for time variation in such global factors may generate misleading conclusions in terms of attribution of financial contagion. We support this conjecture by employing a statistically robust procedure to further test for evidence of contagion. Our findings lend support to the conclusion of related papers that control for the effect of global risk factors and conclude there is a limited role for commodity markets in propagating shocks to the domestic equity markets of commodity exporting economies.

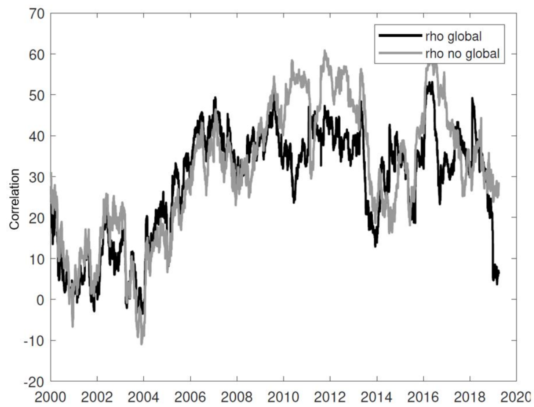
Although we reject the contagion hypothesis, we observe an increase in the correlation/spillovers between commodity and stock markets during episodes of financial distress, even after controlling for the effect of global risk aversion. We interpret these facts as evidence of a greater integration between markets and the growing importance of a global investor sentiment, especially as commodities have become more important as a separate asset class in the portfolios of international investors. As such, our conclusions may be relevant in informing the investment policy of commodity market participants, especially institutional investors and hedge funds. Our results highlight a dynamic, changing pattern between market co-movements over time. However, our interpretation does not strongly support the existence of an additional component in risk transmission from commodity to stock markets that goes beyond the impact of global factors affecting financial markets during episodes of distress. Therefore, unlike those studies which suggest employing additional hedging strategies related to those markets (see Choi and Hammoudeh, 2010), our analysis provides no basis for encouraging revisions to hedging strategies for investors with portfolios exposed to commodity and stock markets of the nine countries we include in our analysis. Overall, given the rejection of the contagion hypothesis, these results suggest there is still an important role for portfolio diversification between the commodity and equity markets of the economies in our sample, a recommendation aligning with other related studies (Arouri & Nguyen, 2010).

Figures

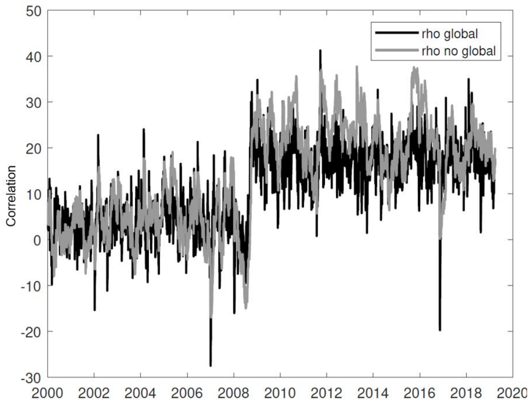
Figure 4.1. Time varying correlation from DCC-GARCH model with and without VIX



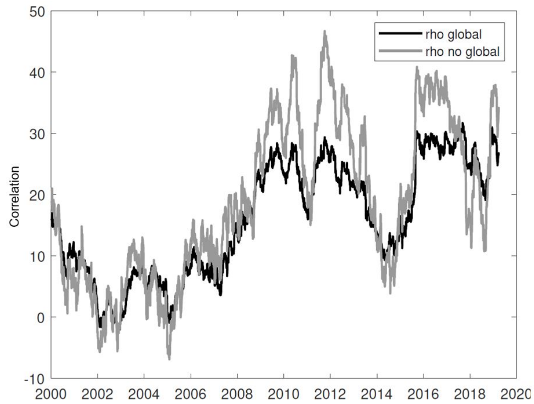
(a) Brazil



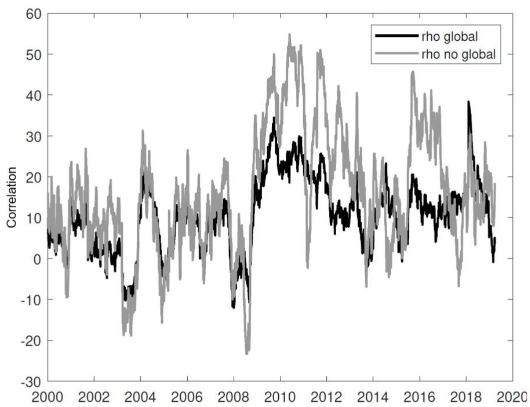
(b) Canada



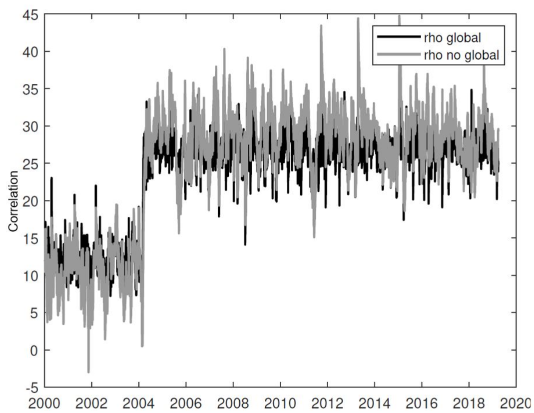
(c) Chile



(d) Colombia

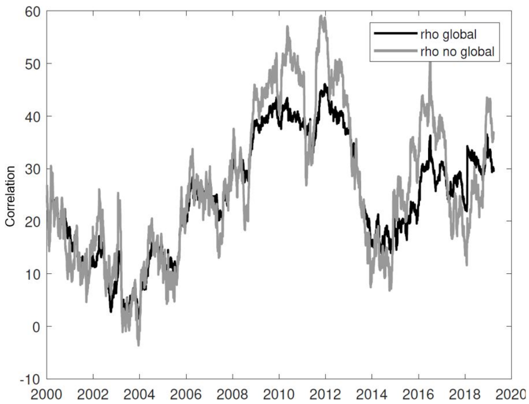


(e) Mexico

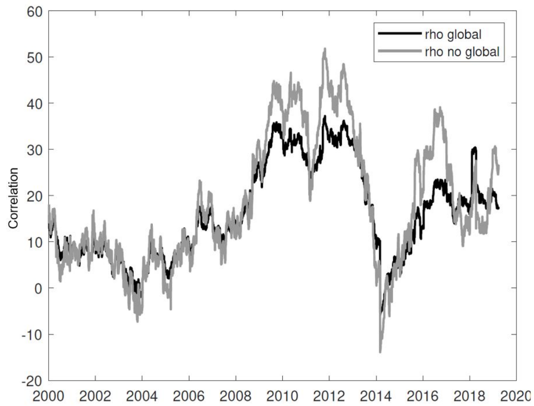


(f) Peru

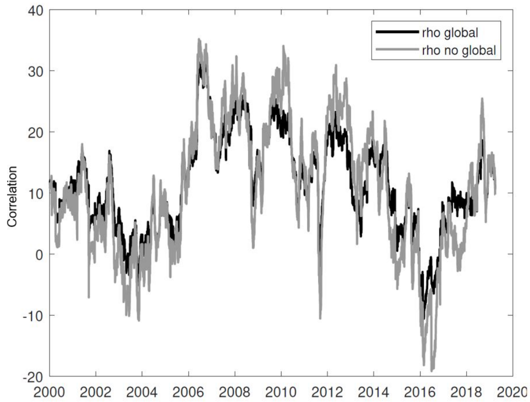
figure 4.1 continued from previous page



(g) Norway



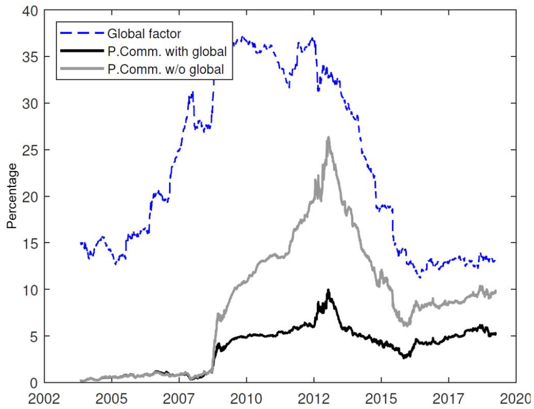
(h) Russia



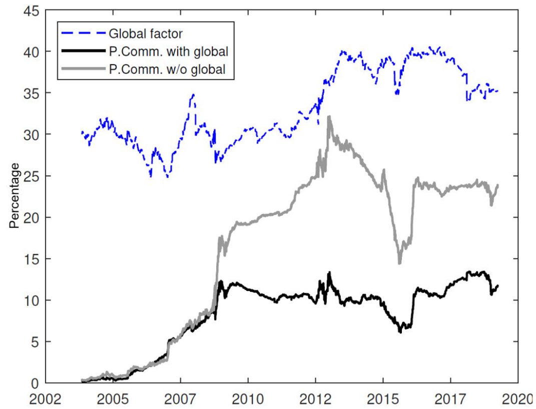
(i) South Africa

The grey (black) line corresponds to the time-varying correlation computed using the DCC-GARCH model introduced in section 4.3.1 excluding (including) the global factor in the model specification. The global factor is proxied by the VIX index. DCC-GARCH model estimated by maximum likelihood using a sample of daily observations from 07 January 2000 to 29 March 2019. Time-varying correlation (ρ_{it}), expressed as a percentage displaying values in the interval (-100,+100), for country i at time t obtained from the covariance matrix in equation 4.1 as follows: $\rho_{it} = \{cov(r_{i,1}, r_{i,2})_t / [\sigma(r_{i,1})_t \cdot \sigma(r_{i,2})_t]\} \cdot 100$, where $r_{i,1}$ denotes domestic stock returns of country i, and $r_{i,2}$ represents the country-specific commodity returns of country i. $cov(.)_t$ and $\sigma(.)_t$ represent the time-varying covariance and time-varying standard deviation, respectively.

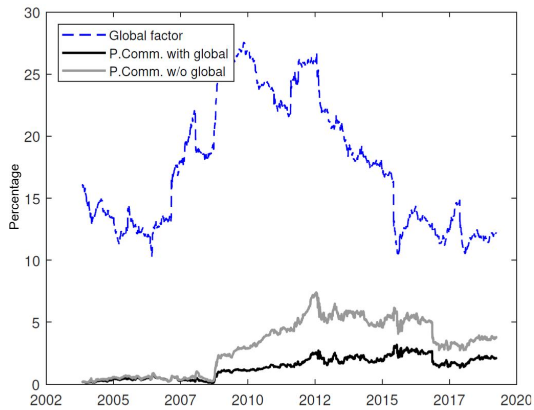
Figure 4.2. Spillovers to domestic stock returns, based on the Diebold-Yilmaz approach, with and without VIX



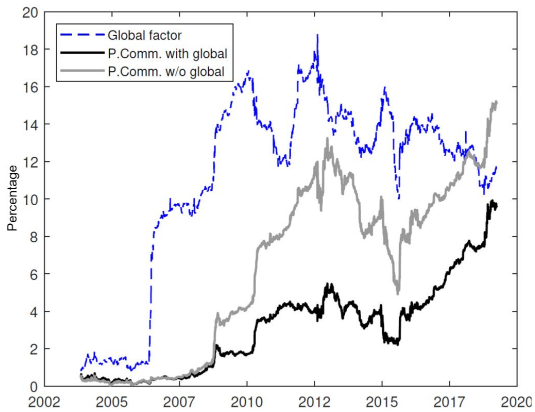
(a) Brazil



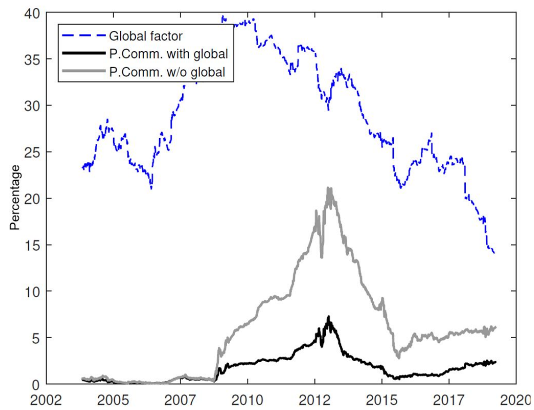
(b) Canada



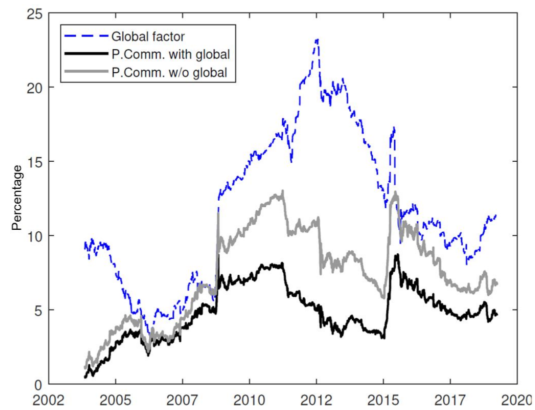
(c) Chile



(d) Colombia



(e) Mexico



(f) Peru

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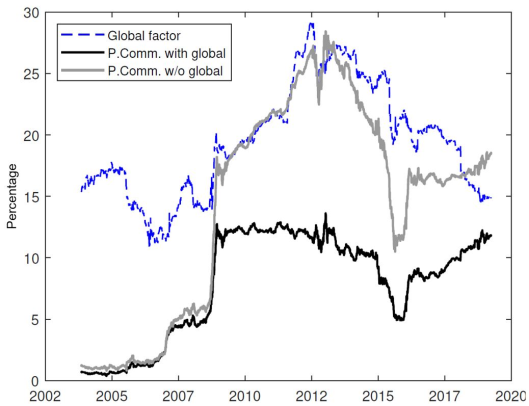
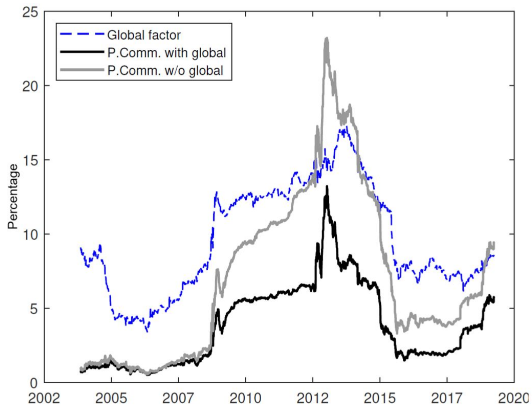
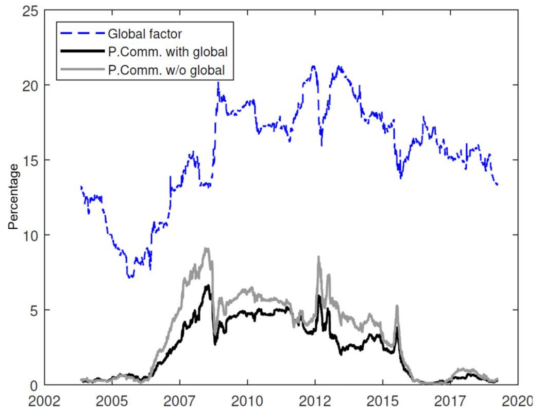


(g) Norway



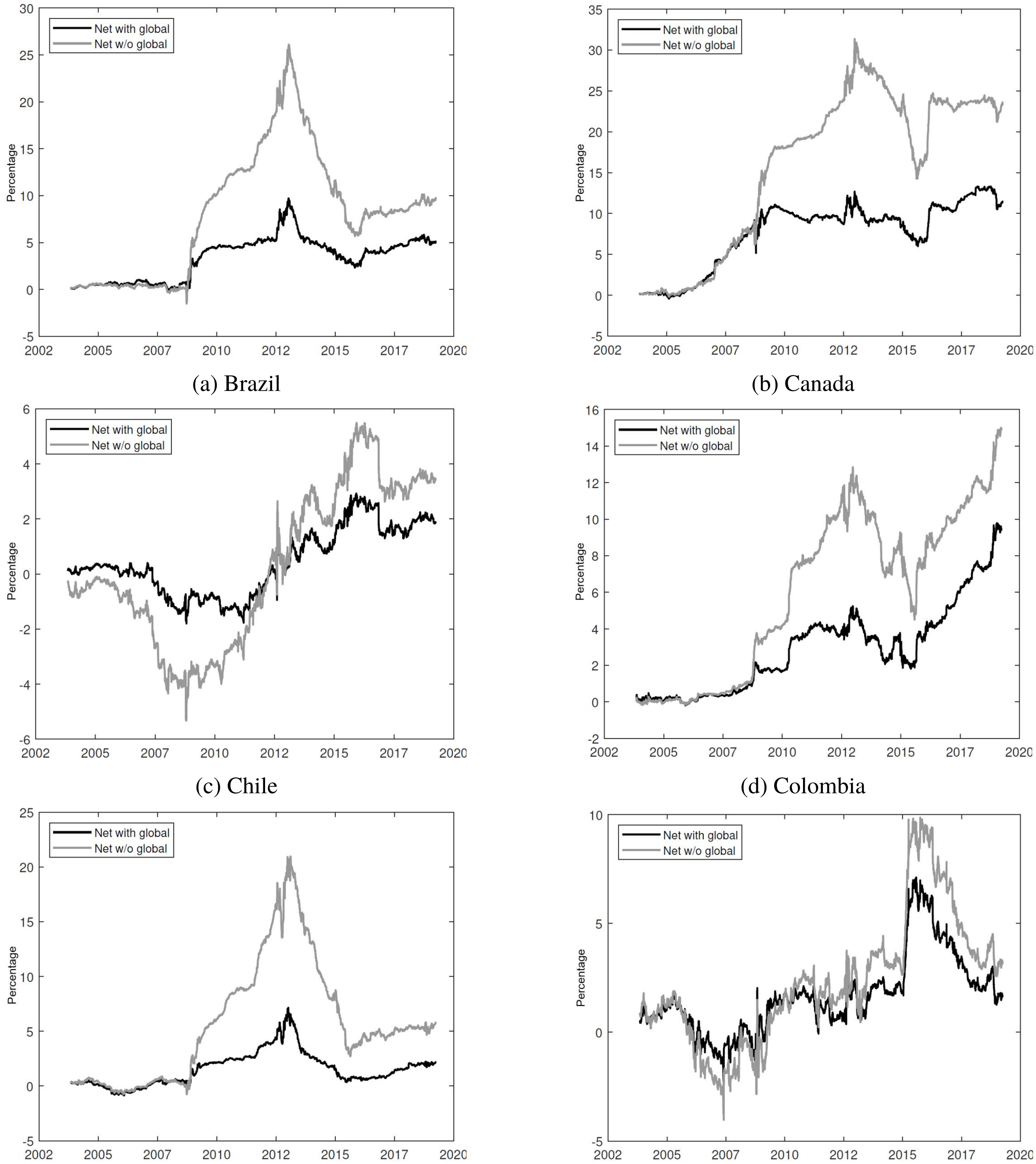
(h) Russia



(i) South Africa

The lines in the figure represent the decomposition of shocks (i.e.: spillovers) received by domestic stock returns. y-axis measures the effect of each spillover as a percentage of the total spillovers received by domestic stock returns. The dashed line depicts the spillovers of the global risk aversion factor, captured by the VIX index, on domestic stock markets. Spillovers of the global factor corresponds to the percentage of the forecast error variance of stock returns explained by orthogonal shocks of changes in the VIX index. Black and grey lines represent the commodity market spillovers to domestic stock markets under two model specifications. The black (grey) line includes (excludes) the effect of global risk aversion factor, captured by the VIX index, in the model specification. Commodity market spillovers correspond to the percentage of the forecast error variance of stock returns explained by orthogonal shocks of commodity returns. Spillovers obtained from the OLS estimation of the vector autoregressive model introduced in section 4.3.2 using a sample of daily observations from 07 January 2000 to 29 March 2019.

Figure 4.3. Net spillovers between commodity and domestic stock markets



(a) Brazil

(b) Canada

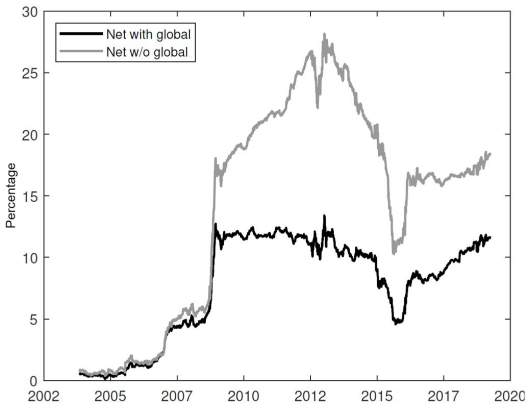
(c) Chile

(d) Colombia

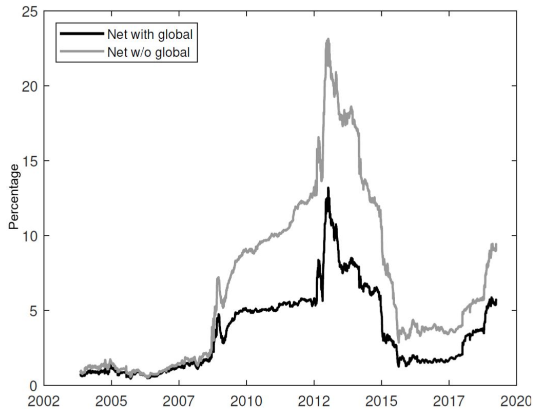
(e) Mexico

(f) Peru

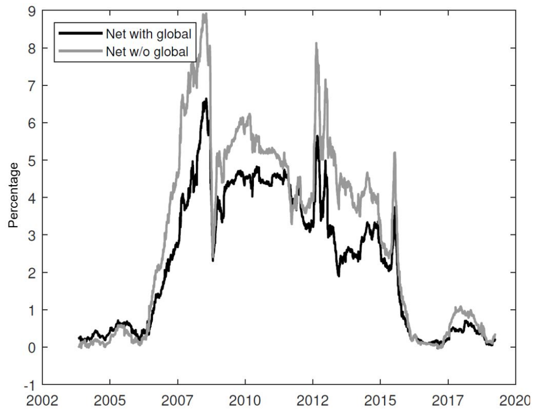
figure 4.3 continued from previous page



(g) Norway



(h) Russia



(i) South Africa

Lines represent the net spillovers between commodity and domestic stock markets under two model specifications: Black (grey) line represents the net spillover including (excluding) the effect of global risk aversion factor, captured by the VIX index, in the model specification. y-axis measures the effect of each net spillover as a percentage of the total shocks received by domestic stock returns. Net spillovers between commodity and domestic stock markets are defined as follows: $S_{net} = S_{comm\rightarrow stock} - S_{stock\rightarrow comm}$, where S_{net} represents the net spillovers between commodity and domestic stock markets, $S_{comm\rightarrow stock}$ denotes the spillovers of commodity returns on stock returns (i.e.: the percentage of the forecast error variance of stock returns explained by orthogonal shocks of commodity returns), and $S_{stock\rightarrow comm}$ corresponds to the spillovers of stock returns on commodity returns (i.e.: the percentage of the forecast error variance of commodity returns explained by orthogonal shocks of stock returns). Spillovers obtained from the OLS estimation of the vector autoregressive model introduced in section 4.3.2 using a sample of daily observations from 07 January 2000 to 29 March 2019.

Tables

Table 4.1. Commodity exporting economies

	Main export	Commodity exports (% of total exports)			Commodity exports (% of GDP)			Main commodity export (% of commodity exports)		
		2000-05	2006-11	2012-17	2000-05	2006-11	2012-17	2000-05	2006-11	2012-17
Brazil	Oil	46	58	64	5	6	6	9	16	13
Canada	Oil	36	49	50	11	12	12	22	33	37
Chile	Copper	84	87	86	24	31	23	55	70	63
Colombia	Oil	65	71	81	9	11	11	43	46	55
Mexico	Oil	18	26	21	4	7	7	59	59	42
Norway	Oil	79	82	80	25	26	22	64	54	44
Peru	Copper	83	88	89	13	23	18	42	49	46
Russia	Oil	76	83	81	22	20	18	57	64	63
S. Africa	Gold	49	55	57	9	12	14	51	62	59

Column “main export” shows the main commodity export by country. The rest of the columns in the table report the average percentage value of exports for each of three periods: 2000-2005, 2006-2011, and 2012-2017 for each country.Source: United Nations Conference on Trade and Development (UNCTAD) website (https://unctadstat.unctad.org/EN/Index.html).

Table 4.2. DCC-GARCH model

		Brazil	Canada	Chile	Colombia	Mexico	Peru	Norway	Russia	S. Africa
Panel A: Univariate mean										
Stock	constant	6.72E-04***	2.28E-04	3.61E-04***	4.69E-04***	3.54E-04**	4.97E-04***	5.66E-04***	7.63E-04***	4.37E-04**
		(2.48E-04)	(1.52E-04)	(1.33E-04)	(1.71E-04)	(1.79E-04)	(1.83E-04)	(2.09E-04)	(2.82E-04)	(1.83E-04)
	Stock (-1)	0.0448***	0.0174	0.1894***	0.2182***	0.1229***	0.1953***	0.0087	0.0559***	0.0612***
		(0.014)	(0.014)	(0.014)	(0.014)	(0.014)	(0.014)	(0.014)	(0.014)	(0.014)
Comm	constant	7.49E-04**	7.49E-04**	4.35E-04*	7.49E-04**	7.49E-04**	4.35E-04*	7.49E-04**	7.49E-04**	3.71E-04**
		(3.42E-04)	(3.42E-04)	(2.26E-04)	(3.42E-04)	(3.42E-04)	(2.26E-04)	(3.42E-04)	(3.42E-04)	(1.52E-04)
Panel B: Univariate volatility										
Stock	α_0	6.38E-06***	7.26E-07***	2.96E-06***	7.53E-06***	1.64E-06***	4.48E-06***	3.01E-06***	4.62E-06***	2.42E-06***
		(9.13E-07)	(1.09E-07)	(3.70E-07)	(6.24E-07)	(2.34E-07)	(3.74E-07)	(4.64E-07)	(3.11E-07)	(4.12E-07)
	α_1	0.07***	0.086***	0.142***	0.213***	0.089***	0.16***	0.106***	0.09***	0.095***
		(0.006)	(0.005)	(0.009)	(0.01)	(0.006)	(0.007)	(0.007)	(0.004)	(0.007)
	β	0.908***	0.908***	0.825***	0.735***	0.902***	0.812***	0.879***	0.898***	0.892***
		(0.007)	(0.005)	(0.011)	(0.011)	(0.005)	(0.007)	(0.008)	(0.004)	(0.008)
Comm	α_0	5.25E-06***	5.25E-06***	2.82E-06***	5.25E-06***	5.25E-06***	2.82E-06***	5.25E-06***	5.25E-06***	7.76E-07***
		(1.03E-06)	(1.00E-06)	(4.43E-07)	(1.03E-06)	(1.02E-06)	(4.36E-07)	(1.03E-06)	(1.04E-06)	(1.26E-07)
	α_1	0.067***	0.067***	0.067***	0.067***	0.067***	0.067***	0.067***	0.067***	0.037***
		(0.004)	(0.004)	(0.005)	(0.004)	(0.004)	(0.005)	(0.004)	(0.004)	(0.002)
	β	0.925***	0.925***	0.921***	0.925***	0.925***	0.921***	0.925***	0.925***	0.956***
		(0.005)	(0.005)	(0.006)	(0.005)	(0.005)	(0.006)	(0.005)	(0.005)	(0.003)
Panel C: DCC										
	λ_1	0.011***	0.01***	0.012***	0.008***	0.017***	0.014***	0.01***	0.008***	0.009***
		(0.002)	(0.002)	(0.003)	(0.002)	(0.003)	(0.005)	(0.002)	(0.001)	(0.002)
	λ_2	0.988***	0.988***	0.975***	0.991***	0.978***	0.928***	0.988***	0.991***	0.986***
		(0.002)	(0.002)	(0.008)	(0.002)	(0.004)	(0.037)	(0.003)	(0.002)	(0.003)

The table reports the results of the maximum likelihood estimation of the DCC-GARCH model we introduce in section 4.3.1. We use daily observation from 07 January 2000 to 29 March 2019. Panel A shows the coefficients of the univariate mean equation for both stock (stock) and commodity (comm) return of equations 4.6 and 4.7, respectively. Panel B exhibits the coefficients of the univariate volatility of equation 4.2 for both stock (stock) and commodity (comm) returns. Panel C presents the estimation results of the coefficients modelling the time-varying correlation of equation 4.3. Standard error in parentheses. (*), (**), (***) indicates statistical significance at 10, 5, and 1% level, respectively. For the case of Chile and Peru, we include a break in the unconditional correlation on 04/09/2008 and 27/02/2004, respectively.

Table 4.3. DCC-GARCH model with VIX

		Brazil	Canada	Chile	Colombia	Mexico	Peru	Norway	Russia	S. Africa
Panel A: Univariate mean										
Stock	constant	8.91E-04***	3.86E-04***	4.63E-04***	5.56E-04***	5.26E-04***	6.04E-04***	7.02E-04***	8.79E-04***	5.38E-04***
		(2.24E-04)	(1.31E-04)	(1.23E-04)	(1.66E-04)	(1.58E-04)	(1.75E-04)	(1.98E-04)	(2.77E-04)	(1.77E-04)
	Stock (-1)	0.0539***	0.0377***	0.2008***	0.218***	0.1298***	0.2056***	0.0194	0.0583***	0.0687***
		(0.013)	(0.012)	(0.013)	(0.013)	(0.012)	(0.013)	(0.013)	(0.014)	(0.014)
	$dVIX$	-0.0011***	-0.0008***	-0.0005***	-0.0004***	-0.0008***	-0.0005***	-0.0007***	-0.0006***	-0.0005***
		(3.13E-05)	(1.83E-05)	(1.72E-05)	(2.32E-05)	(2.21E-05)	(2.44E-05)	(2.77E-05)	(3.86E-05)	(2.47E-05)
Comm	constant	8.77E-04***	8.77E-04***	4.83E-04**	8.77E-04***	8.77E-04***	4.83E-04**	8.77E-04***	8.77E-04***	3.63E-04**
		(3.37E-04)	(3.37E-04)	(2.25E-04)	(3.37E-04)	(3.37E-04)	(2.25E-04)	(3.37E-04)	(3.37E-04)	(1.52E-04)
	$dVIX$	-0.0006***	-0.0006***	-0.0002***	-0.0006***	-0.0006***	-0.0002***	-0.0006***	-0.0006***	0.00004*
		(4.71E-05)	(4.71E-05)	(3.14E-05)	(4.71E-05)	(4.71E-05)	(3.14E-05)	(4.71E-05)	(4.71E-05)	(2.12E-05)
Panel B: Univariate volatility										
Stock	α_0	5.96E-06***	1.54E-06***	3.21E-06***	7.98E-06***	2.45E-06***	4.91E-06***	3.94E-06***	5.03E-06***	2.40E-06***
		(8.58E-07)	(1.04E-07)	(4.03E-07)	(6.60E-07)	(2.40E-07)	(3.90E-07)	(4.12E-07)	(3.95E-07)	(4.33E-07)
	α_1	0.08***	0.101***	0.137***	0.201***	0.096***	0.158***	0.097***	0.085***	0.09***
		(0.006)	(0.006)	(0.01)	(0.01)	(0.005)	(0.007)	(0.007)	(0.004)	(0.007)
	β	0.896***	0.878***	0.821***	0.735***	0.885***	0.808***	0.879***	0.9***	0.895***
		(0.007)	(0.005)	(0.012)	(0.011)	(0.006)	(0.008)	(0.008)	(0.005)	(0.008)
Comm	α_0	5.49E-06***	5.49E-06***	2.60E-06***	5.49E-06***	5.49E-06***	2.60E-06***	5.49E-06***	5.49E-06***	7.40E-07***
		(1.00E-06)	(9.95E-07)	(4.38E-07)	(9.87E-07)	(1.01E-06)	(4.28E-07)	(1.00E-06)	(1.01E-06)	(1.22E-07)
	α_1	0.067***	0.067***	0.065***	0.067***	0.067***	0.065***	0.067***	0.067***	0.037***
		(0.004)	(0.004)	(0.005)	(0.004)	(0.004)	(0.005)	(0.004)	(0.004)	(0.002)
	β	0.924***	0.924***	0.924***	0.924***	0.924***	0.924***	0.924***	0.924***	0.956***
		(0.005)	(0.005)	(0.006)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.002)
Panel C: DCC										
	λ_1	0.007***	0.011***	0.025***	0.004***	0.009***	0.016**	0.005***	0.005***	0.006***
		(0.002)	(0.002)	(0.008)	(0.001)	(0.002)	(0.008)	(0.001)	(0.001)	(0.002)
	λ_2	0.993***	0.987***	0.848***	0.996***	0.986***	0.799***	0.994***	0.994***	0.989***
		(0.002)	(0.003)	(0.069)	(0.001)	(0.004)	(0.168)	(0.002)	(0.001)	(0.004)

The table reports the results of the maximum likelihood estimation of the DCC-GARCH model we introduce in section 4.4.1.2. We use daily observation from 07 January 2000 to 29 March 2019. Panel A shows the coefficients of the univariate mean equation for both stock (stock) and commodity (comm) returns of equations 4.8 and 4.9, respectively. Panel B exhibits the coefficients of the univariate volatility equation 4.2 for both stock (stock) and commodity (comm) returns. Panel C presents the estimation results of the coefficients modelling the time-varying correlation of equation 4.3. Standard error in parentheses. (*), (**), (***) indicates statistical significance at 10, 5, and 1% level, respectively. For the case of Chile and Peru, we include a break in the unconditional correlation on 04/09/2008 and 27/02/2004, respectively.

Table 4.4. Contagion test based on the DCC-GARCH estimation

Panel A: VIX index as a global factor

	Global financial crisis				European sovereign default crisis			
	No global factor	Global factor	Diff	p-val	No global factor	Global factor	Diff	p-val
Brazil	24.8	19.1	-5.7	1.0	45.4	32.2	-13.2	1.0
Canada	36.5	37.4	0.9**	0.011	50.5	38.0	-12.5	1.0
Chile	11.7	10.4	-1.3	1.0	23.2	17.2	-6.0	1.0
Colombia	22.6	18.4	-4.2	1.0	32.6	23.9	-8.7	1.0
Mexico	13.5	8.0	-5.5	1.0	35.8	20.8	-15.0	1.0
Peru	29.5	27.3	-2.2	1.0	29.7	26.9	-2.8	1.0
Norway	34.0	32.5	-1.5	1.0	47.4	40.1	-7.2	1.0
Russia	22.7	20.6	-2.1	1.0	39.1	31.0	-8.1	1.0
S. Africa	20.0	19.8	-0.2	0.7	18.3	15.8	-2.5	1.0

Panel B: SLFFE index as a global factor

	Global financial crisis				European sovereign default crisis			
	No global factor	Global factor	Diff	p-val	No global factor	Global factor	Diff	p-val
Brazil	41.0	26.4	-14.6	1.0	42.1	24.4	-17.7	1.0
Canada	43.5	32.2	-11.3	1.0	56.6	44.8	-11.8	1.0
Chile	23.2	10.1	-13.2	1.0	33.4	22.5	-11.0	1.0
Colombia	23.9	8.3	-15.6	1.0	28.7	3.9	-24.7	1.0
Mexico	24.1	14.1	-10.0	1.0	25.2	14.1	-11.0	1.0
Peru	55.0	46.3	-8.8	1.0	55.0	46.3	-8.8	1.0
Norway	39.9	35.2	-4.7	1.0	54.9	37.8	-17.1	1.0
Russia	36.9	29.0	-7.9	1.0	45.1	32.4	-12.8	1.0
S. Africa	16.4	11.3	-5.1	1.0	16.4	15.1	-1.3	0.8

Panel C: EPU index as a global factor

	Global financial crisis				European sovereign default crisis			
	No global factor	Global factor	Diff	p-val	No global factor	Global factor	Diff	p-val
Brazil	41.0	40.1	-0.9	0.6	46.9	46.1	-0.8	0.9
Canada	43.5	43.4	-0.2	0.5	47.2	46.5	-0.7	0.9
Chile	23.2	20.6	-2.7	1.0	33.4	33.7	0.3	0.3
Colombia	23.9	19.8	-4.1	1.0	28.7	24.8	-3.9	1.0
Mexico	24.1	22.9	-1.2	1.0	25.2	22.9	-2.2	1.0
Peru	55.0	56.3	1.3***	0.0	55.0	56.3	1.3***	0.0
Norway	39.9	39.2	-0.7	0.6	54.9	52.7	-2.2	0.9
Russia	36.9	34.7	-2.2	0.9	45.1	43.7	-1.4	0.9

Table 4.4 continued from previous page

S. Africa	16.4	20.9	4.6***	0.0	16.4	20.9	4.6***	0.0

Column 'No global factor' ('Global factor') corresponds to the average time-varying correlation during the selected crisis episodes computed using the DCC-GARCH model excluding (including) the global factor in the model specification. Time-varying correlation (ρ_{it}), expressed as a percentage displaying values in the interval (-100,+100), for country i at time t obtained from the covariance matrix in equation 4.1 as follows: $\rho_{it} = \{cov(r_{i,1}, r_{i,2})_t/[\sigma(r_{i,1})_t \cdot \sigma(r_{i,2})_t]\} \cdot 100$, where $r_{i,1}$ denotes domestic stock returns of country i, and $r_{i,2}$ represents the country-specific commodity returns of country i. $cov(.)_t$ and $\sigma(.)_t$ represent the time-varying covariance and time-varying standard deviation, respectively. Panel A, B and C depict the estimation results using the VIX index, the St. Louis Fed Financial Stress (SLFFE) index, and the Economic Policy Uncertainty (EPU) index, respectively, as a proxy of the global factor. Column 'Diff' corresponds to the difference between the columns 'Global factor' and 'No Global factor'. Column 'p-val' corresponds to the p-value of the contagion test of section 4.3.3, where H_0: No contagion, and H_a: Contagion. Global financial Crisis dates: 17 July 2007 to 31 August 2009. European Sovereign Default Crisis dates: 2 October 2009 to 26 July 2012. (*), (**), (***) indicates statistical significance at 10, 5, and 1% level, respectively.

Table 4.5. Contagion test based on the Spillovers estimation

Panel A: VIX index as a global factor

	Global financial crisis				European sovereign default crisis			
	No global factor	Global factor	Diff	p-val	No global factor	Global factor	Diff	p-val
Brazil	3.8	2.1	-1.7	1.0	14.4	5.4	-8.9	1.0
Canada	11.6	8.9	-2.7	1.0	21.3	10.6	-10.7	1.0
Chile	1.3	0.6	-0.7	1.0	4.4	1.5	-2.9	1.0
Colombia	2.0	1.1	-0.9	1.0	8.0	3.6	-4.5	1.0
Mexico	2.1	1.0	-1.1	1.0	10.2	3.0	-7.2	1.0
Peru	7.4	5.6	-1.8	1.0	11.2	6.8	-4.5	1.0
Norway	10.5	7.7	-2.8	1.0	22.3	12.2	-10.1	1.0
Russia	4.1	2.8	-1.4	1.0	11.1	5.9	-5.1	1.0
S. Africa	6.5	4.5	-2.0	1.0	5.3	4.5	-0.8	1.0

Panel B: SLFFE index as a global factor

	Global financial crisis				European sovereign default crisis			
	No global factor	Global factor	Diff	p-val	No global factor	Global factor	Diff	p-val
Brazil	14.4	5.6	-8.8	1.0	35.9	14.3	-21.6	1.0
Canada	28.4	15.0	-13.4	1.0	44.3	16.2	-28.2	1.0
Chile	5.7	2.8	-2.9	1.0	12.4	4.6	-7.8	1.0
Colombia	5.8	2.8	-3.0	1.0	12.6	1.3	-11.3	1.0
Mexico	8.3	7.1	-1.2	0.9	11.2	6.9	-4.3	1.0
Peru	28.9	12.6	-16.3	1.0	39.6	12.0	-27.6	1.0
Norway	19.8	12.7	-7.2	1.0	33.7	14.4	-19.3	1.0
Russia	18.7	14.1	-4.6	1.0	27.0	13.4	-13.6	1.0
S. Africa	4.8	3.7	-1.1	0.9	4.0	2.0	-2.0	1.0

Panel C: EPU index as a gloabal factor

	Global financial crisis				European sovereign default crisis			
	No global factor	Global factor	Diff	p-val	No global factor	Global factor	Diff	p-val
Brazil	14.4	9.6	-4.8	1.0	35.9	23.5	-12.4	1.0
Canada	28.4	20.3	-8.0	1.0	44.3	24.2	-20.2	1.0
Chile	5.7	3.2	-2.4	1.0	12.4	5.2	-7.2	1.0
Colombia	5.8	2.8	-2.9	1.0	12.6	3.2	-9.4	1.0
Mexico	8.3	7.2	-1.1	0.9	11.2	8.5	-2.8	1.0
Peru	28.9	22.9	-6.0	1.0	39.6	25.4	-14.2	1.0
Norway	19.8	13.7	-6.1	1.0	33.7	13.0	-20.7	1.0
Russia	18.7	12.8	-6.0	1.0	27.0	12.0	-15.1	1.0

Table 4.5 continued from previous page

S. Africa	4.8	6.9	2.1***	0.0	4.0	4.2	0.2	0.3

Column 'No global factor' ('Global factor') corresponds to the average spillover, expressed as a percentage, from country-specific commodity returns to domestic stocks returns during the selected crisis episodes computed using the Diebold-Yilmaz methodology excluding (including) the global factor in the model specification. Country-specific commodity return spillovers correspond to the percentage of the forecast error variance of stock returns explained by orthogonal shocks of country-specific commodity returns. Panel A, B and C depict the estimation results using the VIX index, the St. Louis Fed Financial Stress (SLFFE) index, and the Economic Policy Uncertainty (EPU) index, respectively, as a proxy of the global factor. Column 'Diff' corresponds to the difference between the columns 'Global factor' and 'No Global factor'. Column 'p-val' corresponds to the p-value of the contagion test of section 4.3.3, where H_0: No contagion, and H_a: Contagion. Global financial Crisis dates: 17 July 2007 to 31 August 2009. European Sovereign Default Crisis dates: 2 October 2009 to 26 July 2012. (*), (**), (***) indicates statistical significance at 10, 5, and 1% level, respectively.

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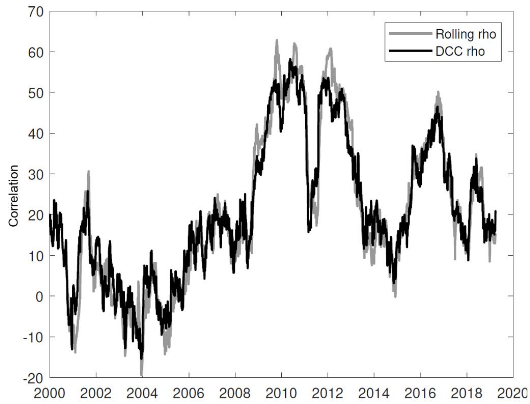
Appendix

Table A.4.1. No persistence in commodity returns

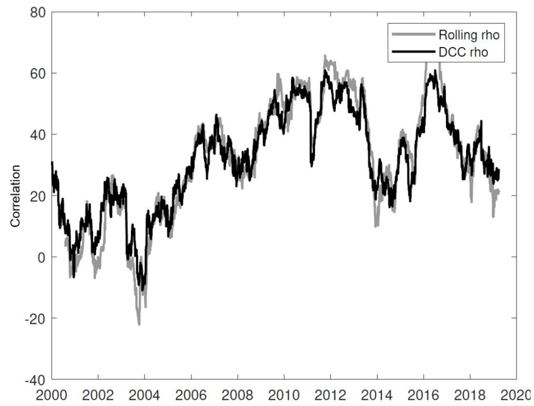
	Copper	**Oil**	**Gold**
Constant	4.89E-04**	8.81E-04***	3.70E-04**
	(2.25E-04)	(3.37E-04)	(1.52E-04)
Persistence	-0.0119	0.0046	-0.0218
	(0.0141)	(0.0139)	(0.0141)
VIX	-0.0002***	-0.0006***	0.00004*
	(3.14E-05)	(4.71E-05)	(2.12E-05)

The table displays the results of time-series regressions estimated by OLS, where the dependent variable corresponds to copper, oil and gold daily returns (in columns). Persistence corresponds to the one-day lagged values of the dependent variable. Standard error in parentheses. Statistical significance: (*) $p < 0.1$, (**) $p < 0.05$, (***) $p < 0.01$.

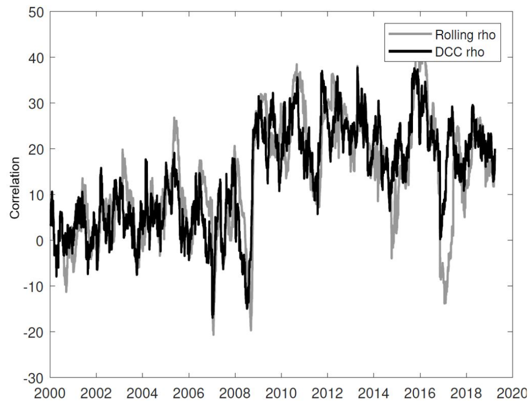
Figure A.4.1. Comparison of rolling pairwise correlation and time-varying correlation (DCC-GARCH model)



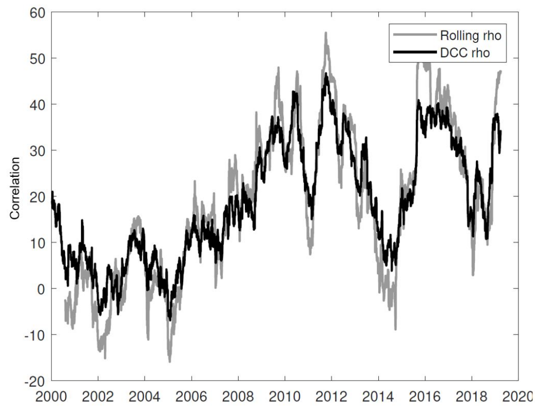
(a) Brazil



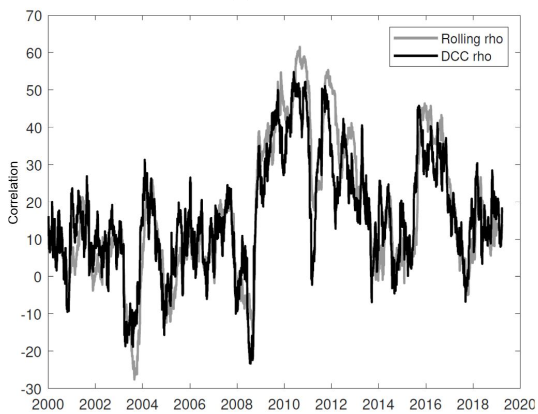
(b) Canada



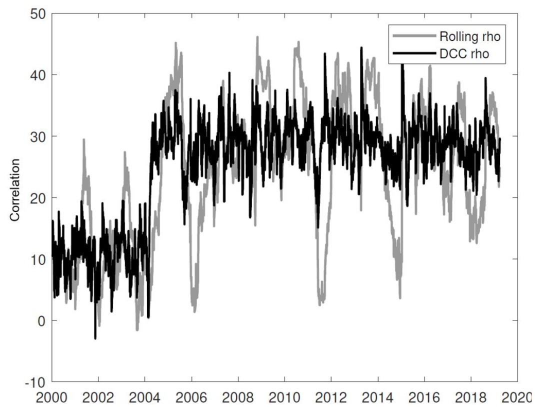
(c) Chile



(d) Colombia

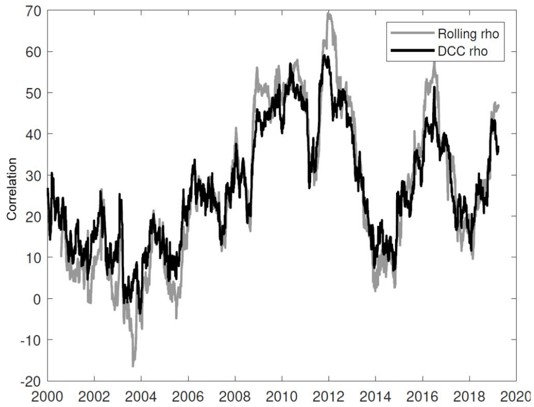


(e) Mexico

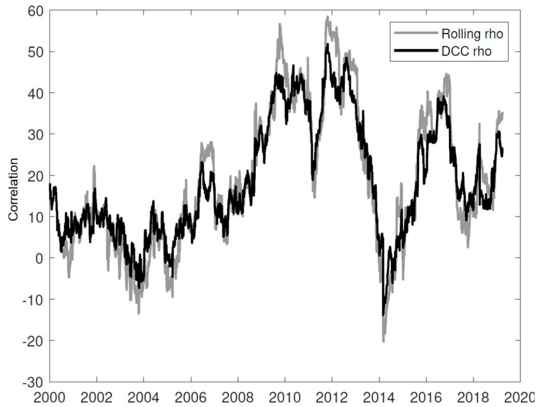


(f) Peru

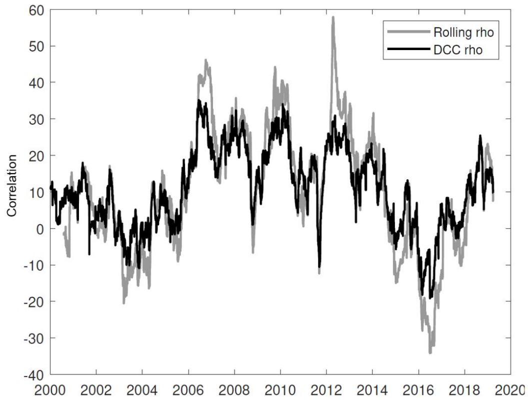
figure A.4.1 continued from previous page



(g) Norway



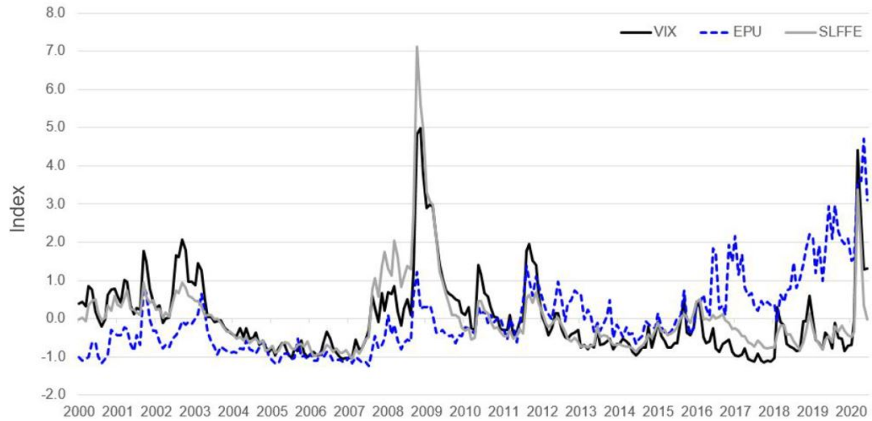
(h) Russia



(i) South Africa

The figure displays a time-series comparison of the time-varying correlation obtained using the DCC-GARCH model (black line) and a rolling window correlation (grey line), both expressed in percentage displaying values in the interval (-100,+100). The DCC-GARCH model corresponds to the one introduced in section 4.3.1 and it is estimated by maximum likelihood using a sample of daily observations from 07 January 2000 to 29 March 2019. We compute the time-varying correlation (ρ_{it}), expressed as a percentage, for country i at time t from the covariance matrix in equation 4.1 as follows: $\rho_{it} = \{cov(r_{i,1}, r_{i,2})_t / [\sigma(r_{i,1})_t \cdot \sigma(r_{i,2})_t]\} \cdot 100$, where $r_{i,1}$ denotes domestic stock returns of country i, and $r_{i,2}$ represents the country-specific commodity returns of country i. $cov(.)_t$ and $\sigma(.)_t$ represent the time-varying covariance and time-varying standard deviation, respectively. We obtain the rolling window correlation using a window length of 150 daily observations.

Figure A.4.2. Comparison of global factor proxies



VIX corresponds to the volatility index computed by the CBOE. EPU corresponds to the Economic Policy Uncertainty index of Baker et al. (2016), available on https://www.policyuncertainty.com/. SLFFE index corresponds to the St. Louis FED Financial Stress Index. For comparison purposes, all series are standardised using their respective mean and standard deviation.

Global factor correlation matrix

	VIX	SLFFE	EPU
VIX	100		
SLFFE	89	100	
EPU	10	6	100

The table reports the correlation matrix. Correlation expressed as percentage. VIX corresponds to the volatility index computed by the CBOE. EPU corresponds to the Economic Policy Uncertainty index of Baker et al. (2016), available on https://www.policyuncertainty.com/. SLFFE index corresponds to the St. Louis FED Financial Stress Index. All indices measured in their original unit of measure.

Chapter 5

Conclusion

The main objective of this thesis is to improve our understanding of the impact of international investment capital flows in a sample of commodity exporting economies, with Chile as the particular centre of attention. We focus on three areas of interest, all of which have implications for policy-makers in such economies. First, we document and provide empirical evidence on how the recent emergence of an important group of financial advisory firms in the Chilean pension fund industry has influenced the investment asset allocation between domestic and overseas markets, and has generated significant concomitant effects on trading activity and price movements in the Chilean FOREX market. Second, we analyse the role of tail dependence in the distributions of exchange rate and commodity price returns as a key factor which underpins the documented ability of commodity prices to predict exchange rates at contemporaneous and short-term forecasting horizons. Finally, we examine the role of global risk aversion in transmitting shocks to financial markets, with a focus on the implications of the existence of such a channel for assertions of so-called contagion existing between commodity and equity markets in commodity exporting economies.

Chapter 2 explores the motivation for the increasing attention which has been received in the past decade or so by financial advisory firms operating in the Chilean pension fund market and analyses its concomitant impact on the Chilean Peso FOREX market. Pension fund companies (PFCs) in the Chilean pension fund industry manage national pension savings totalling $200 thousand million dollars, corresponding to about 80% of the Chilean gross domestic product in 2020. About 40% of the PFCs' balance sheet corresponds to asset held in overseas markets. In this context, one particular company, *Felices y Forrados* (F&F), has positioned itself as the market leader and the most popular advisory firm providing recommendations to pension fund investors in relation to short-term investment strategies for rebalancing the risk exposure of their pension fund portfolios. We demonstrate that investors who follow F&F recommendations rebalance their pension investment fund portfolios, in the process generating large asset reallocations in the pension fund system. These large asset reallocations, following F&F recommendations, subsequently lead to massive, coordinated transactions by PFCs in the Chilean Peso FOREX market. The analysis in this chapter documents the fact that F&F recommendations generate a significant impact on trading activity and price movements in the Chilean Peso FOREX market. We show that after recommendations, on average, the Chilean peso depreciates 0.86% (1.57%) one (five) days after the F&F announcement, while exchange rate volatility sharply increases, then quickly dissipates the day after recommendations. Considering our evidence showing F&F recommendations introduce significant price pressure in the Chilean FOREX market, we also find that other types of market participants appear able to anticipate the subsequent massive, coordinated volume of transactions of PFCs in the Chilean FOREX market. Importantly, our evidence suggests that F&F recommendations generate substantial price pressure in the FOREX market pushing the Chilean Peso's exchange rate beyond the value suggested by economic fundamentals, in the process exacerbating its volatility.

This study constitutes the first piece of evidence evaluating the effect of unregulated agents, such as F&F, that provide investment advice, on the Chilean FOREX market. Our findings contribute to an on-going policy-making discussion by providing important empirical evidence relating to the impact of unregulated financial advisory firms on overall financial stability objectives. We also highlight the fact that additional research is needed to investigate the extent other classes of participants in the Chilean FOREX market may benefit from strategic trading complementarities by front-running the massive transaction of PFCs in this market. Due to data restrictions, our current dataset does not allow us to investigate this idea further, as we only have access to a sample of aggregate FOREX trading volume by agent. However, more detailed and disaggregated data available from the Central Bank of Chile may facilitate further examination of this phenomenon.

Chapter 3 further analyses the non-linear relationship which prior literature has documented to exist between exchange rate and commodity returns using a sample of nine commodity-exporting countries between 2000 and 2018. Using a specific non-parametric estimator, the asymptotic dependence indicator (ADI), we analyse tail dependence between commodity and exchange rate returns to determine if it is indeed the presence of such tail dependence which is critical in understanding any non-linear relationships and the documented ability of commodity prices to forecast exchange rates. Unlike other closely related studies focusing on exchange rate forecasting (Ferraro et al., 2015; Kohlscheen et al., 2017), we provide an explicit rationale which may serve to explain the source of the forecasting ability of commodity returns on exchange rate returns based on an extreme value framework. We believe our contribution lies in this novel approach to synthesising these two literatures. Our results show that both measured asymptotic dependence between exchange rate and commodity returns and any documented ability of commodity returns to predict exchange rates only manifests itself when we use daily, contemporaneous observations. Conversely, we find no asymptotic dependence exists between the returns nor do we observe any forecasting ability for commodity prices for exchange rates using lagged commodity returns or when using observations at a lower data frequency, such as monthly or quarterly. Our evidence suggests that the non-linear relationship between the variables is a key element that underpins the ability of commodity returns to contemporaneously forecast exchange rate returns. Our results also reveal that any documented asymptotic dependence and also the forecasting ability of commodity returns are both transitory and short-lived phenomena, as only contemporaneous, daily observations are able to capture both elements. We believe that the transitory and short-lived relationship between commodity and exchange rate returns reflects the impact of price relevant news and its ability to convey information from one market to another. The fact that only daily contemporaneous information capture both the asymptotic dependence and the forecasting ability of commodity returns reflects the pricing efficiency of FOREX markets.

Our findings relating to the tail behaviour of exchange rate and commodity returns contribute to improving the understanding of the transmission mechanism of shocks between markets. Our evidence may help policy-makers in developing countries when moving towards more flexible exchange rate regimes, by directing them to examine the source of shocks to which flexible exchange rates are exposed. Additionally, understanding the nature of the relationship between exchange rates and commodity markets is relevant for those commodity-exporting economies planning to adopt inflation target regimes, considering that shocks affecting the exchange rate subsequently pass-through to domestic inflation.

Future research on this topic may explore the varying degree of asymptotic dependence exhibited among countries. An interesting point to evaluate relates to the need to analyse the elements explaining the asymptotic dependence across countries and to further evaluate if country-idiosyncratic factors have a role to play in explaining the forecasting ability of commodity prices for exchange rates. Another future research avenue may focus on analysing the effect of COVID-19 pandemic on the transmission of shocks between commodity and exchange rate markets. As some recent studies argue (Ji et al., 2020; OECD, 2020), the transmission of shocks to FOREX markets changes subsequent to the pandemic. This suggests that interesting further analysis may focus on extending the sample date to incorporate the COVID-19 pandemic episode in order to assess its impact on asymptotic dependence and the forecasting ability of commodity returns.

Chapter 4 revisits the relationship between commodity and equity markets. Several recent studies claim the enhanced comovement between these markets, particularly during episodes of financial distress, reflects evidence of financial contagion (Creti et al., 2013; Mensi et al., 2013; Roy & Roy, 2017; Xu et al., 2019). Based on a sample of commodity-exporting economies, we analyse the links between their domestic stock exchanges and international commodity markets. We employ both a DCC-GARCH model and the spillover analysis proposed by Diebold and Yilmaz (2009) to examine the relation between these asset markets. Unlike previous studies documenting financial contagion between commodity and domestic stock returns, we account for the effect of time variation in global investor sentiment and risk aversion by including suitable proxies for these phenomena, such as the VIX index, the economic policy uncertainty index and the St. Louis Fed financial stress index. Among our main findings, we document that the correlation/spillovers between commodity and stock returns varies over time, appearing to intensify during episodes of financial distress, such as the 2007-09 Global Financial Crisis and the 2009-12 European sovereign debt crisis. However, in contrast to many previous studies, we argue this evidence does not constitute financial contagion between markets.

In fact, after controlling for the effect of global factors capturing investors' risk appetite and economic uncertainty, we observe the impact of commodity markets on equity returns reduces substantially in magnitude. Moreover, after undertaking a statistically rigorous approach, we find no evidence of financial contagion between markets in the case of most of the countries in the sample. Our evidence suggests that the role of commodity markets in conveying shocks to the domestic stock market of commodity-exporting economies is more limited than many related papers argue. We find that the variables we employ to proxy for variation in global sentiment and risk aversion capture much of the comovement between markets, and appear to be a critical component in explaining the intensification of the relationship during financial crisis episodes. This result is consistent with the evidence analysing financial contagion in related studies focusing on a sample of advanced economies (Apergis & Miller, 2009; Kilian & Park, 2009; Zhang, 2017). When considering the implications of our findings, we argue that failing to control for time variation in such global factors may generate misleading conclusions in terms of financial contagion. The contribution in this chapter is to improve our understanding of shock transmissions across assets, focusing on commodity and stock markets. In particular, we reassess evidence documenting financial contagion from commodity to stock returns by controlling for additional factors which may account for the comovement between markets. Additionally, we further extend the existing evidence to consider not only the crude oil markets which have been the focus of the majority of prior studies, but also other additional commodity classes, in particular copper and gold. Our findings are relevant for portfolio managers looking for asset diversification. Our results show that stock markets exhibit a milder response to commodity market shocks than documented in many previous studies, therefore, portfolio diversification opportunities are still possible in relation to both asset classes.

Future analysis may aim to assess the transmission of shocks between international commodity and stock exchange markets during the recent COVID-19 episode. Unlike the two crisis we analyse in this chapter, the COVID-19 crisis did not emanate from financial markets, although it exerted a substantial impact upon them. In this sense, evaluating the shock transmission between markets during the current pandemic may shed further light on the relevance of commodity markets for transmitting shocks to stock exchanges in an environment where investor risk appetite plays a less pivotal role.

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